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Social and Demographic Consequences of Health insurance

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Social and Demographic Consequences of Health insurance

Abstract
Scholars are eager to evaluate the effects of health policy on health, but they often neglect that policies are intricately connected to marriage, family structure, and social standing. The three chapters of this dissertation study unintended consequences of health insurance policies in the United States. How people gain private health insurance is connected to divorce (chapter 1) and availability of a public insurance program at birth is associated with lower mortality not only in infancy but also in adulthood (chapter 2). US health policies that tie health insurance coverage to socioeconomic status add dimensions to racial and ethnic inequality. Minorities spend more years without insurance due to their greater probabilities of losing coverage (chapter 3). These chapters provide national landscapes of health insurance coverage and inequality in the years prior to the Patient Protection and Affordable Care Act of 2010 setting the baseline for post-reform comparisons.

In Chapter 1, I apply hazard models to the nationally representative longitudinal Survey of Income and Program Participation (2004 panel) to find lower divorce rates among people who are enrolled in their spouses' health insurance policies. Women who depend on their husbands for health insurance had the lowest rates of divorce. This chapter highlights how family- and employment-based insurance coverage could create inequalities between families and between men and women.

In Chapter 2, I use US Vital Statistics data to compare changes in age-specific mortality rates between cohorts born in states with Medicaid and cohorts born in states without Medicaid. I exploit the variation in the timing of States' Medicaid participation to establish a connection between the availability of public insurance at birth to improvements in later life mortality. This chapter underscores the lasting consequences of having access to medical care during critical periods in the life-course.

Chapter 3 examines the dynamics of gaining and losing health insurance across the life course and how it contributes to racial and ethnic disparities in coverage. African Americans, Hispanics, and Asians have high uninsurance rates mostly due to their greater likelihoods of losing the insurance that they already have. Life-table simulations show that simply increasing the accessibility of health insurance does surprisingly little to reduce disparities in insurance coverage. This paper draws attention to the importance of developing policies that stabilize existing insurance.

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SOCIAL AND DEMOGRAPHIC CONSEQUENCES OF HEALTH INSURANCE

Heeju Sohn

A DISSERTATION

in

Demography and Sociology

Presented to the Faculties of the University of Pennsylvania

in

Partial Fulfillment of the Requirements for the

Degree of Doctor of Philosophy

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SOCIAL AND DEMOGRAPHIC CONSEQUENCES OF HEALTH INSURANCE

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Heeju Sohn
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ABSTRACT

SOCIAL AND DEMOGRAPHIC CONSEQUENCES OF HEALTH INSURANCE

Heeju Sohn
Jason Schnittker

Scholars are eager to evaluate the effects of health policy on health, but they often neglect that policies are intricately connected to marriage, family structure, and social standing. The three chapters of this dissertation study unintended consequences of health insurance policies in the United States. How people gain private health insurance is connected to divorce (chapter 1) and availability of a public insurance program at birth is associated with lower mortality not only in infancy but also in adulthood (chapter 2). US health policies that tie health insurance coverage to socioeconomic status add dimensions to racial and ethnic inequality. Minorities spend more years without insurance due to their greater probabilities of losing coverage (chapter 3). These chapters provide national landscapes of health insurance coverage and inequality in the years prior to the Patient Protection and Affordable Care Act of 2010 setting the baseline for post-reform comparisons.

In Chapter 1, I apply hazard models to the nationally representative longitudinal Survey of Income and Program Participation (2004 panel) to find lower divorce rates among people who are enrolled in their spouses’ health insurance policies. Women who
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Chapter 3 examines the dynamics of gaining and losing health insurance across the life course and how it contributes to racial and ethnic disparities in coverage. African Americans, Hispanics, and Asians have high uninsurance rates mostly due to their greater likelihoods of losing the insurance that they already have. Life-table simulations show that simply increasing the accessibility of health insurance does surprisingly little to reduce disparities in insurance coverage. This paper draws attention to the importance of developing policies that stabilize existing insurance.
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Introduction: Socioeconomic Class and Health in the Context of Public Policy in the United States

A well-developed literature in the social sciences shows how families share social, economic, and human capital and how it translates into an intergenerational transmission of social class. Health and mortality are established as both causes and consequences of socioeconomic class. The chapters of this dissertation aim to contribute to the literature that connects health and socioeconomic status across family ties.

This dissertation examines family processes and social inequality and their interactions with health policy in the United States. People get married, bear children, and make daily medical decisions in the context of the policy environment. Here, I demonstrate how public policy plays a significant role in creating and shaping the boundaries between social classes.

Health insurance coverage allows access to better medical care and mitigates the adverse consequences of poor health for individuals and families. It is considered an
important, if not crucial, form of social insurance (Marmor et al. 2013). In the United States, health insurance coverage is very closely tied to a person’s socioeconomic status. Families with high-paying, stable earnings are more likely to enjoy consistent, higher quality health insurance. Those that lack the means to afford private coverage turn to arguably lower-quality public options such as Medicaid (Quesnel-Vallee 2004) or forgo insurance altogether.

Health insurance coverage is more than a mere reward of employment and wealth. It can in turn, affect the socioeconomic status of individuals and his or her family. Going without insurance limits access to medical care, health, and the financial well-being of the entire household (Himmelstein et al. 2009; Kasper et al. 2000; Nelson et al. 1999; Zuvekas and Weinick 1999).

The current literature on the consequences of unequal health insurance coverage predominantly focuses on short-term health outcomes and its resulting disparities. The chapters of this dissertation contribute to current scholarship by drawing attention to previously understudied demographic processes that may be influenced by health insurance policies (divorce and long-term mortality) and by making improvements to how researchers can measure and compare insurance coverage.

This dissertation bridges three major scholarly themes: transmission of socioeconomic status, health, and public policy. Before summarizing each chapter, I briefly review the major theories that guide the empirical analyses of this dissertation.

Bourdieu’s work on intergenerational transfer of social capital provided the theoretical framework for much scholarship on the transmission of socioeconomic status
within families. Parents would endow children with social capital that they would then use to establish their own class on reaching adulthood. Families with greater socioeconomic status had more capital to bestow on their offspring. And through this process, social class was maintained across generations. The current literature uncovers mechanisms beyond direct economic transfers that reinforce class boundaries. Higher-income parents cultivated their children to develop a social habitus that gave them an advantage for academic and financial success in adulthood (Laureau 2003). These parents are likely to better prepare their children for college with greater material resources for education (Conley 2001; Schoeni et al. 2005) and social connections and information to place them in high status occupations (Lubrano 2005). Children of different social classes have disparate transitions into adulthood (Furstenberg 2010) completing the replication of social class. Diverging demographic trends between social classes (McLanahan 2004) reinforce these intergenerational disparities. Individuals with less education and lower income are more likely to raise children in single-parent families (Edin and Kefalas 2005; Martin 2006) arguably reducing the social capital that these children receive even more (Amato and Gilbreth 1999; Amato and Rivera 1999).

Health and socioeconomic class are intertwined throughout the life-course and across generations. Social and environmental mechanisms for intergenerational transmission of health are also well-studied in social demography. The Barker and fetal origins hypotheses motivated many studied that linked prenatal and early health conditions to later adult outcomes. Being born into a high mortality environment was associated with increased risk of ischemic heart disease in adulthood (Barker and
Osmond 1986). Being born during a recession was also linked to lasting heightened mortality (Van den Berg et al. 2006). Quasi-experimental studies find that cohorts in utero during the 1918 Influenza Pandemic displayed increased incidence of physical disability, lower educational achievement, and lower income in adulthood (Almond 2006). Similarly, children of mothers who were pregnant during the 1944 Dutch Famine showed worse health outcomes throughout adulthood (Roseboom et al., 2001). More direct explorations of the relationship between socioeconomic class and health consistently find that health, education, and income are correlated with each other. Children from low SES households were more likely to have health problems. The correlation between parents’ income and child health becomes stronger as the child grows older even after they leave the household (Case et al. 2005). In adulthood, poor health led to diminished earnings and reductions in wealth (Cutler et al. 2008; Smith 2007). These empirical studies work to establish long-term, intergenerational connections between health and social class.

This dissertation explores the relationship between health and social inequality in the context of health systems and public policy. The three chapters examine unintended consequences of health insurance policy in the United States. Scholars and policymakers are eager to evaluate the effects of health policy on health, but they often neglect that policies are intricately connected to family processes and have lasting effects on health and social class. I employ demographic tools to quantify the unintended consequences of health insurance on family processes and demographic trends. Bane and Ellewood (1986) changed how to think about poverty by examining spells of poverty. Through
demographic methods, I reveal new insight into the interplay of health policy and social processes that static, cross-sectional pictures may obscure.

The first chapter finds evidence that people who are enrolled in their spouses’ health insurance policies have lower risks of divorce. I also find that this association is stronger among women than men. I apply hazard models to a nationally representative longitudinal data (the Survey of Income Program Participation) to compare divorce risks between insurance-dependent, that is, adults who are insured through their spouses, and insurance-independent adults, that is, adults who have health insurance policies available to them through their employment. The findings draw attention to the influence that a marriage and employment-based health insurance system can have on divorce rates. Establishing connections between health policy and family formation/dissolution are especially important as the United States undergoes one of the largest health care reforms of the past century.

The second chapter explores the enduring effects of Medicaid. Medicaid is a means-tested federal program delivering public health insurance program to mostly low-income women and infants. In this study, I show that the start of Medicaid was not only associated with reducing infant mortality rates across states but was also associated with health outcomes even after aging out of eligibility. For the research design, I exploit variation in the timing of States’ Medicaid participation to establish a connection between the initiation of Medicaid availability at birth to later-life mortality. This chapter underscores the lasting consequences of having access to medical care during critical periods in the life-course.
The third chapter examines the dynamics of gaining and losing health insurance across the life course and how it contributes to racial and ethnic disparities in coverage. African Americans, Hispanics, and Asians have high uninsurance rates mostly due to their greater likelihood of losing the insurance that they already have especially during early adulthood and middle age. Furthermore, life-table simulations show that simply increasing the accessibility of health insurance does surprisingly little to reduce uninsurance prevalence among minorities. This paper draws attention to the importance of developing policies that stabilize existing insurance. This chapter motivates further research on whether precarious insurance coverage among minority groups can explain disparities in interactions with health care providers, medical decisions, and health outcomes.

These are interesting topics to study in the wake of the Affordable Care Act (ACA). The chapters set baselines for pre- and post- ACA comparisons once the rollout is complete. State-level variation in ACA adoption provides an interesting setting for a natural experiment to observe changes in family processes and social inequality that coincide with the ACA. Medicaid expansions in select states will allow all low-income persons regardless of marital and parental status to become eligible. Individual mandates will further incentivize people to become insured and to stay insured. State-sponsored insurance exchanges will diminish the strong connection between employment, marriage, and access to health insurance in participating states.
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Chapter 1: Health Insurance and Risk of Divorce: Does Having Your Own Insurance Matter?

ABSTRACT

Most American adults under 65 obtain health insurance through their employers or their spouses’ employers. The absence of a universal healthcare system in the United States puts Americans at considerable risk for losing their coverage when transitioning out of jobs or marriages. Scholars have found evidence of reduced job mobility among individuals who are dependent on their employers for healthcare coverage. This paper finds similar relationships between insurance and divorce. I apply the hazard model to married individuals in the longitudinal Survey of Income Program Participation (N=17,388) and find lower divorce rates among people who are insured through their partners’ plans without alternative sources of their own. Furthermore, I find gender differences in the relationship between healthcare coverage and divorce rates: insurance dependent women have lower rates of divorce than men in similar situations. These
findings draw attention to the importance of considering family processes when debating and evaluating health policies.

Love and commitment are often what couples believe secure and protect marriages from divorce. Sociologists of the family however, find that practical considerations are probably more important (Kalmijn, 1998). Married couples with high incomes are more likely to stay married (Amato, 2010; Gibson-Davis, Edin, & McLanahan, 2005).

Educational attainment—an indicator of earnings potential—is also associated with greater marital stability (Amato, 2010). Education and income are consistently stronger predictors of divorce than sentiment-driven indicators. Recognizing the significance of these factors, researchers are careful to take household income, the couple’s educational attainment, wealth, and other resources into consideration as they study divorce patterns. Researchers have yet to study insurance coverage as a factor that can influence divorce.

This paper examines the relationship between health insurance and divorce by asking three main research questions. (1) Is there an association between being insured by a spouse and divorce? (2) Does this association get stronger when one partner does not have an independent source of health insurance? (3) Do these associations differ by gender?

The United States is among the few and perhaps the only developed country that does not provide universal healthcare to its residents (Jost, 2003). While seniors over 64 years of age are assured coverage under Medicare, the majority of non-elderly adult men and women are left responsible for securing their own health insurance. Employment and marriage are the top two sources of health insurance for American adults; 24% of non-
elderly adult women and 14% of men are covered as a dependent (Kaiser Family Foundation, 2011). In comparison, 55.3% of the adult population gains health insurance though employment (US Census, 2012).

Insurance is consequential. Being uninsured is associated with lower healthcare utilization, increased morbidity, and higher mortality (Institute of Medicine, 2004). Not having insurance creates barriers to adequate access to healthcare (Institute of Medicine, 2002). Uninsured individuals are more likely to be diagnosed with late-stage cancer—a disease that is often detected at early states during routine doctor visits (Halpern, War, Pavluck, Schrag, Bian, & Chen, 2008)—and are less likely to have an ongoing relationship with a health care provider (Holahan & Spillman, 2002). In addition, medical expenses contribute to a large portion of personal bankruptcies in the United States (Himmelstein, Thorne, Warren, & Woolhandler, 2009). Considering these known risks of being uninsured, it is not surprising that Americans treat health insurance as a valuable commodity.

Even brief periods of uninsurance can be costly. As American adults transition into and out of jobs and marriages, they risk losing health insurance coverage (Lavelle & Smock, 2012; Meyer & Pavalko, 1996). No social infrastructure guarantees continued health insurance through these transitions. The Consolidated Omnibus Budget Reconciliation Act (COBRA) allows people to purchase insurance at the lower group rate under certain conditions, but is only available for limited periods and its costs can be prohibitive, especially for recently unemployed or divorced individuals. Many cannot help but experience a gap in insurance coverage as they change jobs or go through
divorce (Lavelle & Smock, 2012; Swartz & McBride, 1990). These gaps are significant. They can increase premiums or limit payouts even when individuals gain coverage, particularly for those who have on-going health care needs (van de Ven, van Vliet, Schut, & van Barneveld, 2000). Gaps can also have significant financial consequences for those who fall ill while uninsured.

Researchers who study employment and wages have established that the health insurance benefits are an important part of an employee’s compensation; a job that provides health insurance yields substantially higher total compensation than a job with the same salary but no health benefits (Woodbury, 1983). Studies also show that insurance can act as a significant motivator for people to seek and retain employment, at times even deterring workers from pursuing otherwise better opportunities (Cooper & Monheit, 1993; Madrian, 1994; Monheit & Cooper, 1994). If health insurance plays such a large role in the labor market, we can expect it to also play a role in the marriage market. After all, approximately 36 million American adults under the age of 65 rely on a family member to provide their health insurance (KFF, 2011). This paper finds that married individuals who are dependent on their spouses for health insurance have lower rates of divorce. Having private health insurance is highly correlated with socio-economic characteristics such as high education, stable employment, and high income. These factors all lower the risk of divorce, but the analyses show health insurance does not follow the same pattern. This paper shows that divorce risk is lowest among people who do not have the option of employment-based coverage and are insured under their spouses’ plans. I use the longitudinal 2004 Panel of Survey of Income Program
Participation (SIPP). The 2004 SIPP tracks a nationally representative sample of over 43,500 households over time with great granularity recording each member’s marital and healthcare status every month from October 2003 through December 2007. I employ the Cox proportional hazard model to estimate the relative risk of divorce between insurance-dependent and non-dependent individuals.

BACKGROUND

In his economic models of marriage, Gary Becker (1974) posits that positive gains to marriage and negative consequences of divorce motivate two individuals to stay married. His theory argues that couples with greater economic resources have lower risks of marital dissolution as these resources will increase the gains that individuals would derive from their unions. Higher levels of income can protect married couples from financial stressors (Sawhill, 1975). Economic assets, such as homeownership, can also have stabilizing effects on marriage (Becker, Landes, & Michael, 1977; Levinger, 1979; South & Spitze, 1986). By the same token, couples with higher levels of educational attainment are less likely to divorce (Martin, 2006).

The decision to divorce is not always unilateral. Becker and his coauthors (1977) explain that couples will remain married if the combined gains of staying married exceed the combined benefits of separating. If person A wants to divorce while their partner B does not, the latter can ‘compensate’ the former to make it worthwhile for both parties. If the combined net benefit of the union is not enough to satisfy both partners, however, the marriage will result in dissolution.
This study examines health insurance as a potential gain to a marriage. Because health insurance pertains to the household, it is perhaps a prominent feature of the combined gains of marriage. A family health insurance plan that covers both spouses can be considered a shared good that benefits from economies of scale. It is less costly than two people being insured independently. A person who is not covered by their spouse must either purchase an individual plan at a higher price or gain access to a group plan, usually through employment. In this way, securing private health insurance for a married couple is cheaper than two individuals obtaining their own and, therefore, it can be considered an economic ‘gain’ of marriage.

While health insurance has not been explicitly studied as an economic gain of marriage, it has been examined in the context of employment and job mobility (Cooper & Monheit, 1993; Madrian, 1994; Monheit & Cooper, 1994). Studies find that jobs that offer health insurance plans have lower turnover rates (Cooper & Monheit, 1993; Madrian, 1994; Monheit & Cooper, 1994). Likewise, being married to someone who can provide a health insurance policy may increase the gains of the marriage, leading to lower rates of divorce. My first hypothesis is:

**Hypothesis 1**: Married individuals who are insured through their spouses’ health plans have lower rates of divorce

The second hypothesis examines the risk of losing health insurance coverage as a potential negative consequence of divorce. Does the association between insurance coverage and divorce get stronger when one partner does not have an independent source of health insurance? Researchers of employer-specific insurance plans and job-mobility
find that people who are expected to be worse off if they lose current coverage—for example, employees with greater medical expenditures and those without spouses who can provide coverage during transitions—have lower rates of job-mobility (Buchmueller & Valletta, 1996; Cooper & Monheit, 1993; Gruber & Madrian, 2002; Madrian, 1994). The economic consequence of divorce can also be influenced by whether those in a marriage have access to group coverage independently. The cost of divorce for people who can switch over to their employers’ plans on divorce will not be as high as those whose only sources of insurance coverage is from their spouses. Not having a comparable source of health insurance outside the marriage increases the negative consequences of divorce that may lead to lower divorce rates. I use employer-sponsored health insurance as a comparable alternative to spouse-provided health insurance plans. There are alternatives, but they are not as relevant. Public means-tested cash transfers, for example, do not appear to lower the cost of divorce. (Hoffman & Duncan, 1995). Likewise, needs-based public health insurance, such as Medicaid, would not be a comparable substitute to a private, family insurance plan. Needs-based public insurance is often available only at very low levels of income and is sometimes considered to be of lower quality than private insurance plans (Quesnel-Vallee, 2004).

Hypothesis 2: Not having an employment-based source of health insurance coverage outside the marriage further lowers the risk of divorce for people enrolled in their spouses’ plans.
Gender, Health Insurance, and Marriage

The story of divorce and health insurance is further complicated by the issue of gender. Historically, the division of labor in a US household has fallen along gendered lines (Cherlin, 1995; Greenstein, 1995; Kalmijn, Loeve, & Manting, 2007; Nock, 1995; Presser, 2000; Sayer, England, Allison, & Kangas, 2011). While spousal roles and duties have become more flexible and negotiable over time, the traditional male breadwinner model of the family persists (Greenstein, 1995). Women still do most housework and childrearing (Cherlin, 1995; Nock, 1995) and marriages are more stable when the husband earns more than his wife (Kalmijn et al., 2007). The responsibility of financial contribution still primarily falls onto the man. A husband’s unemployment is strongly related to divorce whereas a wife’s is not (Sayer et al., 2011). By the same token, Teachman (2010) found that health-related work limitations among men but not among women were associated with higher rates of marital disruption. The gains of a marriage increase when the gendered expectations of their husbands resuming the role of the primary breadwinner are met. For wives, spouses’ incomes have a positive relationship with their self-reported levels of marital commitment (Nock, 1995).

The same relationship applies to health insurance as well. If traditional gender roles put pressure on husbands to contribute financially via economic activity in the labor market, it is also likely that men are expected to provide the family with health insurance. Thus, I expect to see gender differences in the divorce rates associated with insurance dependence in my analysis. My third hypothesis is:
**Hypothesis 3:** Women who are insured through their spouses have lower rates of divorce than men who are insured through their spouses

The fourth hypothesis examines whether the association between divorce and having an alternative source of insurance is stronger for women than it is for men. With men taking on the primary responsibility of the household income, it is not surprising the economic consequences of divorce are less favorable for women than men (Burkhauser, Duncan, Hauser, & Berntsen, 1991). Burkhauser and Duncan (1989) found that divorce or separation was the single most financially detrimental event that could happen to non-elderly women. Divorce’s consequences for health insurance are equally striking (Bernstein, Cohen, Brett, & Bush, 2008; Lavelle & Smock, 2012). Lavelle and Smock (2012) find that, divorce leads to an eight percentage-point decline in women’s private health insurance coverage, net of changes in employment, economic resources, and other factors. This decline was even sharper for women insured as a dependent. A corollary of these findings is that wives’ economic independence through labor force participation may lower barriers against divorce (Duncan & Hoffman, 1985; Lavelle & Smock, 2012). Duncan and Hoffman (1985) found that women’s human capital investments have some modest effects on mitigating the decline in their economic situation following a divorce. Similarly, Lavelle and Smock (2012) also found that just-divorced women who were insured through their own plans were largely protected from the risk of losing private coverage.

**Hypothesis 4:** Not having an alternative source of health insurance outside the marriage lowers divorce risk for women more so than for men.
METHOD

Data

I use the 2004 panel of the Survey of Income and Program Participation (SIPP) to examine the relationship between health insurance status and divorce. The SIPP is a nationally representative series of longitudinal panels whose survey duration ranges from 2.5 to 4 years. The first SIPP panel was sampled in the early 1980s and a new panel was re-sampled from the non-institutionalized population in the U.S. every one to four years. The 2004 SIPP panel is a longitudinal dataset that follows its respondents for about four years from October 2003 until December 2007. The strength of this dataset lies in its large size, its wide range of household insurance and demographic variables, and its longitudinal structure. The SIPP survey is divided into core questionnaires and topical modules. The core questionnaire collects data for the same variables every month throughout the study period. The SIPP administers a topical module containing a different theme (i.e. marital history) once every four months. The SIPP then creates longitudinal panels checking for data inconsistencies and missing values. The Census Bureau accounts for missing values by logically deriving from other available information when possible. They rely on several imputation techniques to fill in the remaining missing data. Thus, if a respondent completed a questionnaire or a module, all variables had valid values. Most of the variables that I use in my analyses are from the core questionnaires. Since these variables are recorded every month, I can get relatively close estimates of when changes occurred. Life events such as divorce and changes to insurance rarely happen multiple times within a single month.
The 2004 SIPP collected data for 45,540 individuals. The hypotheses in this paper pertain to the risk of divorce. Therefore, I only include married individuals in the analytic sample. I dropped 14,239 individuals (33% of total initial 2004 SIPP dataset) who were not married during the study period. Furthermore, I only include married individuals whose marriage duration is known. The sample includes people who were already married when the SIPP started to collect data in October 2003 and people who became married during the study period before it ended in December 2007. For those who were already married in October 2003, I derive their marriage durations from their marital history topical module that the SIPP administered in May 2004. I exclude 1,354 (3% of initial 2004 SIPP dataset) married respondents who did not complete the marital history topical module from the sample. Lastly, I limit my analysis to non-elderly adults. Those over 65 are almost universally insured through Medicare and divorce may affect their insurance status differently than the rest of the adult population. I exclude 10,559 persons (24% of initial data) over the age of 64 or under the age of 18 from the sample.

My analysis sample consists of 17,388 individuals who were married at some point during the study period between October 2003 and December 2007. They collectively experienced over 500 divorces. I weight the observations using sampling weights to compensate for SIPP’s different selection probabilities into the 2004 panel across subpopulations.

**Variables**

*Marital Status.* The 17,388 individuals in my sample, weighted to represent approximately 46 million US residents, were married at some point during the study
period. A change in a respondent’s marital status reflects a marriage or divorce/separation sometime during the previous 30 days. Only the respondents who are married are at risk of divorce or separation and I limit hazard calculations to the currently married at any point in time. The incremental hazards of separation in response to all covariates are similar to the hazards of divorce. Sensitivity tests result in similar hazard patterns for all three events: divorce only, separation only, and divorce or separation. The analyses in this paper examine the hazard of either divorce or separation.

**Health Insurance Status.** I combine two SIPP variables asking health insurance type and coverage source to create a nominal scale to differentiate individuals who are covered by their own plan from those covered under someone else’s, government sponsored need based insurance (Medicare or Medicaid), or none at all. The SIPP specifically asks if the primary subscriber is the respondent’s spouse at only one point in time during the survey period. More than 95% of married people on June 2005 who are insured on someone else’s plan are insured on their spouses’ plan. I assume that this extremely high rate of spousal coverage among married individuals who are not primary subscribers is approximately constant throughout the study period. I lag insurance status by one month to associate the insurance status prior to the divorce, as this marital event is often accompanied by a simultaneous change in health care status. I have conducted sensitivity test using a two, three, and six month lag time between insurance status and marital disruption. There is little difference in the coefficients and the hypotheses tests up to a lag time of three months. At six months, the coefficients in the hazard models
become smaller in size, but even with a lag time that is one-eighth of the total survey period the tests of hypotheses yield the same results.

*Alternative Source of Health Insurance.* In my analysis, I consider individuals who are unemployed, contingent workers, or employed part-time at small companies with less than 100 employees to be without access to employer-sponsored health insurance plans outside the marriage. Full-time employment status and the size of the firm are two indicators that predict whether an employer offers health insurance to an employee (Kaiser Family Foundation, 2013). Throughout the study period, 97 to 99 percent of firms with over 200 workers offered health insurance to their employees, whereas only 59 to 65 percent of smaller firms offered insurance plans (Kaiser Family Foundation, 2013). Very few firms offered insurance plans to temporary workers (three to six percent). Less than a quarter of small firms offered plans to their part-time staff but almost half (47%) of large firms offered insurance to workers who were not full time employees (Kaiser Family Foundation, 2013). Lavelle and Smock (2012)’s study also find that full time workers were largely protected from insurance loss after divorce.

I confirm from the 2004 SIPP that its respondents’ employment statuses and their employers’ firm sizes are indeed good predictors of whether the respondents have access to health benefits. I test how closely these variables capture respondents’ access to employment-based health insurance by examining single, never-married individuals. Almost 80% of single, never-married respondents who I assigned as having “access to employer-sponsored health insurance” are insured as the primary subscriber. About half of the never-married people who I categorized as having access but are in fact uninsured,
voluntarily elected not to enroll in their employers’ plans. The SIPP asked specific questions on whether a respondent’s employer offers insurance plans and his or her reasons for not enrolling only once in June 2005. Therefore, I could not use this variable to create a better measurement of the availability of employment-based insurance throughout the entire period.

Control Variables. I incorporate demographic and socioeconomic variables that prior research identified as determinants or predictors of divorce—education, race and ethnic origin, number of children, age, marital history, and household income (Amato, 2010; Casper & Bianchi, 2001; Cherlin, 1992). My models include gender, race/ethnicity, educational attainment, the number of children, and a polynomial term for age as controlling covariates. Many households in the SIPP report income that fluctuates significantly from month to month. Some households report their entire annual income in one month leaving the monthly income for the rest of the year at zero. To smooth the income flow, I use the average of monthly total family income to measure the family’s level of financial standing. I expect higher educational attainment, higher income, and having more children to be associated with lower rates of divorce (Amato, 2010). Marital history is another predictor of divorce or separation (Becker et al., 1977; Lehrer, 1988; Presser, 2010). I include an indicator for whether or not the current marriage is a first marriage. Higher order marriages may be more likely to end in divorce or separation (Becker et al., 1977; Lehrer, 1988). I also include a polynomial term for the length of marriage. Probability of dissolution decreases with the duration of marriage (Becker et al., 1977; Presser, 2000). Couples who would end up divorced tend to end their marriage
earlier than later (Becker et al., 1977) and people make more marriage-specific investments (e.g. children, sexual compatibility) the longer they are together (Lehrer, 1988). All covariates with the exception of gender and race are time-varying.

**Analytic Strategy**

I use Cox's proportional hazard model (Cox, 1972) to measure the effect of health insurance on marriage. This model estimates the incremental risk of an event happening to one group relative to a reference group. A person’s hazard is a multiplicative replica of the baseline hazard based on his or her set of covariates. In this way, the model can identify characteristics that are associated with greater or lower hazards of events such as divorce (Bumpass, 1990). I use STATA13’s stcox package to estimate all hazard coefficients. I test the model’s assumption that divorce hazards of the comparison groups are proportional over time. I interact the key variables of the analysis (insurance status and access to employer-sponsored health plans) with time (reference month) and test for statistically significant time-varying effects. A joint test of significance revealed no systematic change in the association between insurance status and access to employer-sponsored health plans.

I estimate four hazard models to test each of the four hypotheses in this paper. The first model estimates the divorce hazards associated with insurance status. Model 2 adds access to employment-based insurance to model 1 and interacts it with the respondent’s insurance status. Model 3 interacts gender with insurance status to examine any gender differences in the association between insurance and divorce. The fourth model is a three-way interaction between insurance status, access to an employment-
based option, and gender. I present the hazards associated with the interactions in Models 2-4 with a series of dummy variables. I report these hazards in odds ratios. Each odds ratio indicates the relative hazard of divorce or separation of a respondent with a particular characteristic relative to the reference group. All models include covariates for age, race, education, children, higher-order marriage, marriage duration, and logged average monthly income as controls. I then select the relevant coefficients from the hazard models to explicitly test the four hypotheses of this paper. I calculate p-values from one-sided t-tests adjusted for false discovery rate (Benjamini & Hochberg, 1995). These p-values are more conservative than unadjusted p-values as they take into account that the probability of falsely rejecting a null condition increases with the number of tests performed. I present the four Cox hazard models and their hypothesis tests in the main results section of this paper.

RESULTS

Descriptive Results
I estimate divorce hazards from the 17,388 individuals in the 2004 SIPP who were married at some point between October 2003 and December 2007. Table 1.1 shows the basic descriptive statistics of my analysis sample. The percentages and averages presented in Table 1.1 are weighted to represent the adult US married population under age 65. About half the sample was insured under their own names and a third were insured as dependents. Less than 20% of the analysis sample were insured by need-based government plans or were uninsured. About a third received four-year college degrees.
and about another third did not receive any education beyond high school. Roughly half
the respondents were women.

[Table 1.1. Descriptive Statistics for Study Measures]

Being insured under one’s own name was correlated with other measures of
socioeconomic standing. Over 85% of people with their own health insurance were
employed full-time and contributed over half of their households’ total incomes on
average. They were also more likely to have attended college. I show in Table 1.2 that
more married women than men were insured as a dependent. In concordance with the
findings by the Kaiser Family Foundation (2011) women in my sample were more likely
than men to be insured on another’s plan. About 44% of married women, compared to
16% of married men, were insured as dependents. The educational attainment and income
of these men were also not too different from the married men who were insured under
their own names. Men who were enrolled as a dependent still have higher earnings and
were more likely to have attended college than their female counterparts. While the
insurance-dependent men earned less on average, the proportion with some post-
secondary education was slightly higher.

[Table 1.2. Descriptive Statistics of Risk Population by Insurance Status and Gender]

While individuals who were insured under their own plans contributed proportionally
more to total household income, those who were insured through their spouses also made
economic contributions to the household. The sizes of these contributions however,
differed by gender. Men who were enrolled on another’s insurance plan contributed 41%
of the total household income, on average. Women who were enrolled on another’s
insurance plan, on the other hand, contributed on average, less than 20%. Table 1.3 shows the differences in economic contribution to the household by insurance status and gender. These percentages represent the average monthly earnings of an individual as a proportion of the monthly total household income during the marriage. Total household income includes income from means-tested cash transfers and income from property.

[Table 1.3. *Earnings Contribution to Total Household Income by Insurance and Gender (%)*]

**Main Results**

The following section tests and reports the results of the four main hypotheses of this paper. I test each hypothesis with a separate Cox proportionate hazard model (Models 1 to 4) in the same order that I presented in the background section.

*Hypothesis 1*: Married individuals who are insured through their spouses’ health plans have lower rates of divorce

[Table 1.4. *Hypothesis Test 1: Divorce Hazard on Insurance Status*]

At any point in time, married individuals who were insured on their spouses’ plans were indeed significantly less likely to divorce or separate than those who were covered under their own policies. Being dependent on one’s spouse for health insurance lowered the divorce hazard by almost 70%. The odds of divorce associated with being insured by a spouse (0.321) were statistically significant at the 0.01 level. Table 1.4 shows the odds ratios for covariates related to insurance and family income. Logged family monthly
income had a significant negative association with divorce hazard consistent with prior findings (Bumpass, Martin, & Sweet, 1991). The insurance coefficients remain significant indicating that their association with divorce could not be entirely explained by the family’s income.

**Hypothesis 2**: Not having an employment-based source of health insurance coverage outside the marriage further lowers the risk of divorce for people enrolled in their spouses’ plans.

[Table 1.5. *Hypothesis Test 2: Divorce Hazard on Insurance Status and Access to Employment-based Option*]

Model 2 confirms that not having an employment-based source of insurance was associated with further declines in divorce hazards among individuals who were insured by their spouses. This model interacts insurance status with access to employer-sponsored plans as people may have had the option to enroll in their own employers’ health plans but had to forgo them in favor of their spouses’ family policies. Table 1.5 reports the divorce odds ratio for each insurance status group who had and who did not have an employment-based option. The divorce ratio of someone insured under their spouses’ plans with an option for an employer-sponsored plan was 0.433 relative to persons who were insured under their own names. Individuals insured under their spouses’ plans without access to employer-sponsored plans had divorce hazard ratios of 0.179. Both coefficients were significant at the alpha 0.005 level. I test whether not having access to an employer-sponsored plan significantly lowered the divorce hazards among people who
were insured by their spouses in the lower panel of Table 1.5. Not having an insurance option outside the marriage further lowered divorce odds by 0.412 and this difference is statistically significant at the alpha 0.005 level. Logged family monthly income had a significant negative association with divorce in Model 2. However, similarly to the first model, family income did not completely moderate the relationship between insurance and divorce.

**Hypothesis 3:** Women who are insured through their spouses have lower rates of divorce than men who are insured through their spouses

[Table 1.6. Hypothesis Test 3: Divorce Hazard on Insurance Status and Gender]

Model 3 shows that divorce hazards associated with being covered by a spouse’s health insurance did not differ by gender. Men who were insured on their spouses’ plans had divorce risks that were 0.475 that of men who were primary subscribers. Women who were insured by another had hazards of 0.373 of the same reference group (Table 1.6). The difference in hazard ratio between men and women were not statistically significant. Model 3 reveals another interesting gender difference. A woman who had her own source of health insurance was associated with a significantly higher hazard of divorce than a man who was the primary subscriber. This is consistent with some findings from prior research on the positive relationship between women’s financial independence and likelihood of divorce (South and Spitze, 1986; Greenstein, 1995). While a wife’s earnings can increase gains and stability to a marriage by augmenting the household income (McLanahan, 2004), empirical evidence also shows that wives who earned more than
their husbands were more likely to experience divorce (Kalmijn et al., 2007). Because the divorce risks of women who were primary subscribers were so high, the differences in divorce hazards between primary and dependent insurance subscribers were greater among women (1.580 vs 0.373) than among men (1.000 vs 0.475).

**Hypothesis 4**: Not having an alternative source of health insurance outside the marriage lowers divorce risk for women more so than for men

[Table 1.7. Hypothesis 4: Divorce Hazard on Insurance Status, Access to Employment-based Option, and Gender]

The divorce hazards of spouse-insured women who did not have an employment-based source were significantly lower than their male counterparts. The divorce hazards of men who were insured by their spouses and did not have employment-based options were 0.489. The hazards for similar women were 0.179 (Table 1.7). The difference in divorce odds between these two groups was statistically significant at the alpha 0.05 level. The reference group for Model 4 consists of male primary subscribers who were most likely to have employer-sponsored insurance.

Having an outside option for insurance was not associated with higher divorce hazards among men when they were already insured by their wives (Table 1.7). The story is different for women. The divorce hazards of women who were insured by their spouses but also had access to employment-based plans outside their marriages were significantly higher than spouse-insured women who did not have alternative sources through employment. The odds ratio of 0.412 associated with having access to employer-
sponsored insurance plans among women who were insured through their spouses was statistically significant (Table 1.7). These coefficients’ significance persists even when family income is included as a covariate in Model 4. Taken together, these findings indicate that the insurance that a spouse provides may act as a deterrent to marital disruption in addition to other economic predictors of divorce such as employment and income.

DISCUSSION

This paper tests two main ideas. Does being dependent on a spouse for health insurance lower the hazard of divorce? And, does this relationship between health insurance dependency and divorce differ between men and women? The results affirm that on average, people who were insured through their spouses’ health plans had lower rates of divorce. Incorporating employment status into the baseline model shows that higher levels of dependency on their spouses for health insurance coverage further led to diminished risks of divorce or separation. These results are in concordance with the findings from researchers of employer-provided health benefits and job mobility. Not having an alternative source for health care outside their current arrangement—employment and marriage—made individuals less likely to terminate their jobs and marriages.

The results further demonstrate that the relationship between insurance dependency and marital stability differed by gender. The gendered relationship between health insurance and divorce mirrors the dynamics of income and marital stability. Risk of divorce rises along with wives’ contribution to the family income in excess of their
husbands’ with the most stable marriages being those where the husband is the primary earner (Heckert, Nowak, & Snyder, 1998; Kalmijn, Loeve, & Manting, 2007; Ono, 1998). Likewise, I find divorce rates are the highest among women who had access to health insurance independent of their husbands. While women’s employment may have transitioned from being a marital destabilizer to a stabilizer in recent decades (Oppenheimer, 1994; Sayer et al., 2011), it appears that securing the family health insurance still remains within the male domain.

There are several limitations to this study. I recognize that marital decisions may not always be unilateral and often result from joint decision-making between the two spouses. A wife may be motivated to stay together in consideration for the husband’s lack of health insurance. Whatever mechanisms at play, my results show the different divorce outcomes based on an individual’s insurance situation. While the monthly health insurance and marital status measurements in the SIPP are strengths in determining a relationship between the two, I also note that couples often obtain legal divorce decrees months after they make their decisions. The health insurance situation of the two individuals involved in the failing marriage may have changed since beginning divorce proceedings. Research has shown increases in married women’s labor force participation in the periods prior to divorce (Gray, 1995). Similarly, the insurance-dependent partner may be motivated to secure other sources for health insurance in anticipation of the change in marital status. Applications for divorce specifically address the issue of healthcare coverage and divorcing individuals are fully aware of the termination of benefits through their soon-to-be former spouse. If this is the case, the effects on divorce
rates associated with insurance-dependence may be an upper-bound more relevant to couples who were not able to secure independent coverage the month immediately prior to the finalization of the divorce. My sensitivity analyses with lagged insurance statuses of two, three, and six months indeed show that the longer the lag, the coefficient decreases in magnitude. Nevertheless, hypotheses tests still yield the same conclusion; insurance-dependence lower risk of marital disruption and that this association is more pronounced for women than for men.

The health status of an individual may change the association between insurance and divorce. Having poor health may increase the dependency on a spouse’s health insurance plan as having continuous coverage become more important and poor health can also negatively impact employment prospects. The SIPP, unfortunately, does not have good measures of health to adequately capture the respondents’ need for health insurance. In a sensitivity analysis, I used self-rated health rated on a five-point scale both as a control and as a moderator for the relationship between insurance and divorce. This variable in the 2004 SIPP was problematic as less than 3% of the analytic sample indicated that they were not in good health and the SIPP only recorded the variable twice during the four-year data collection period. Self-rated health in this case, was neither a predictor of divorce nor an influential moderator.

Lastly, the SIPP’s short study period of 48 months can also be another limitation of the data. The study can only prospectively observe the risk of divorce of a couple only for a four-year window of their marriage. It cannot account for the entire history of all couples’ marriages by tracking them from their wedding till its dissolution through
divorce or death. The divorce and separation hazards are estimated from this longitudinal data’s four-year study period and are extrapolated throughout couples’ marital life courses.

Despite these limitations, this paper contributes to the broader literature examining the determinants of marital stability. It distinguishes health insurance’s influence on divorce from other measures of socioeconomic status such as education, income, and employment. If private insurance coverage was simply an artifact of these characteristics, we would expect that those with stable employer-based coverage would have the lowest rates of divorce. On the contrary, the analyses demonstrate that spouse-insured persons who have less economic options are the ones who are the least likely to divorce. While the traditional economic resources contribute to lower divorce rates by making the marriage more attractive, it is the aversion to the risk of losing health insurance that deters people away from divorce. It is a subtle but an important distinction especially when studying health policies that aim to guarantee health coverage to more people. The incentive to stay in marriages for health insurance may be stronger for those with lower prospects of securing and maintaining private health insurance through employment. The negative association between insurance dependence and divorce that we see in this paper may likely to be stronger among people with low socioeconomic status; these people may rely more on their spouses to protect themselves from the risk of losing health coverage.

This paper also underscores the gendered patterns in marriage and economic dependence. The reduction in divorce rates associated with insurance dependence is
stronger among women than for men. It is consistent with research showing that
American marriages are still governed by gendered social norms; men are often expected
to resume the responsibility of financially providing for the family through labor force
participation outside the household.

This paper draws attention to the strong connection between health care and
marriage. Acquiring health insurance in the United States is largely dependent on work
and marriage—two things that are valued in American society (Kaiser Family
Foundation, 2011). The system rewards adults who seek and maintain good employment
or who remain married to partners who can provide spousal coverage. When affordable
and dependable health care is not guaranteed, people are incentivized to conform to the
social behaviors that American policies promote. Whether or not sociologists and health
researchers agree that access to good health care should be so intricately tied to family
values, they cannot neglect the inevitable influences that they have on each other.
Divorce is only one of many family processes that could be affected by health care
policies in the United States. The findings in this study call attention to the importance of
taking into consideration family dynamics when developing and evaluating health care
policies. Marriage, childbirth, divorce, remarriage, and transitions to adulthood are all
significant life events that could be shaped by policies. Understanding the relationship
between health policies and family processes is crucial to working toward a more
effective health care system that improves the overall health and happiness of the
population.
REFERENCES


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Quesnel-Vallee, A. (2004). Is it really worse to have public health insurance than to have no insurance at all? *Journal of Health and Social Behavior, 45*, 376-392.


Table 1.1 Descriptive Statistics for Study Measures (N = 17,388)

<table>
<thead>
<tr>
<th>Population at Risk</th>
<th>M or %</th>
<th>SD</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Insurance Status</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Insured under own name</td>
<td>50.84</td>
<td></td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>31.64</td>
<td></td>
</tr>
<tr>
<td>Gov't Insurance (Medicare, Medicaid)</td>
<td>5.64</td>
<td></td>
</tr>
<tr>
<td>Uninsured</td>
<td>11.88</td>
<td></td>
</tr>
<tr>
<td><strong>Race/ethnicity</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Non-Hispanic White</td>
<td>73.55</td>
<td></td>
</tr>
<tr>
<td>African American</td>
<td>7.66</td>
<td></td>
</tr>
<tr>
<td>Hispanic</td>
<td>12.58</td>
<td></td>
</tr>
<tr>
<td>Asian</td>
<td>3.84</td>
<td></td>
</tr>
<tr>
<td>Other</td>
<td>2.47</td>
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</tr>
<tr>
<td><strong>Educational Attainment</strong></td>
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<td></td>
</tr>
<tr>
<td>Less than High School</td>
<td>9.43</td>
<td></td>
</tr>
<tr>
<td>High School Diploma or Equiv.</td>
<td>21.06</td>
<td></td>
</tr>
<tr>
<td>Associate degree or some college</td>
<td>37.47</td>
<td></td>
</tr>
<tr>
<td>Bachelors' degree</td>
<td>20.95</td>
<td></td>
</tr>
<tr>
<td>Advanced degree</td>
<td>11.09</td>
<td></td>
</tr>
<tr>
<td><strong>Children</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Not living with Children</td>
<td>36.85</td>
<td></td>
</tr>
<tr>
<td>One child</td>
<td>23.09</td>
<td></td>
</tr>
<tr>
<td>Two children</td>
<td>24.90</td>
<td></td>
</tr>
<tr>
<td>Three children</td>
<td>15.16</td>
<td></td>
</tr>
<tr>
<td><strong>Age</strong></td>
<td>41.85</td>
<td>10.61</td>
</tr>
<tr>
<td><strong>Family Monthly Income</strong></td>
<td>6,712.19</td>
<td>5,116.87</td>
</tr>
<tr>
<td><strong>Gender: 0 = male, 1 = female</strong></td>
<td>0.53</td>
<td></td>
</tr>
</tbody>
</table>

*Note:* Population at risk at first reference month (November 2003 if already married or first month of marriage). Values weighted to represent the US population.
Table 1.2 Descriptive Statistics of Risk Population by Insurance Status and Gender (N = 17,388)

<table>
<thead>
<tr>
<th>Insurance Status</th>
<th>Sample Size (%)</th>
<th>Mean Age</th>
<th>Avg. Monthly Earnings (USD)</th>
<th>Proportion with some college education (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Men</td>
<td>Women</td>
<td>Men</td>
<td>Women</td>
</tr>
<tr>
<td>Insured under own name</td>
<td>66.8</td>
<td>36.8</td>
<td>42.9</td>
<td>41.9(^a)</td>
</tr>
<tr>
<td>(n=8,091)</td>
<td>(n=9,297)</td>
<td>(n=8,091)</td>
<td>(n=9,297)</td>
<td>(n=8,091)</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>17.0</td>
<td>44.6</td>
<td>43.9(^b)</td>
<td>41.8(^a)</td>
</tr>
<tr>
<td>(n=8,091)</td>
<td>(n=9,297)</td>
<td>(n=8,091)</td>
<td>(n=9,297)</td>
<td>(n=8,091)</td>
</tr>
<tr>
<td>Gov't Insurance</td>
<td>4.4</td>
<td>6.8</td>
<td>45.0(^b)</td>
<td>38.2(^a,b)</td>
</tr>
<tr>
<td>(Medicare, Medicaid)</td>
<td></td>
<td></td>
<td>737(^b)</td>
<td>390(^a,b)</td>
</tr>
<tr>
<td>Uninsured</td>
<td>11.9</td>
<td>11.9</td>
<td>38.9(^b)</td>
<td>37.9(^b)</td>
</tr>
<tr>
<td></td>
<td>(n=8,091)</td>
<td>(n=9,297)</td>
<td>(n=8,091)</td>
<td>(n=9,297)</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>1,873(^b)</td>
<td>742(^a,b)</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>42.4(^b)</td>
<td>42.9(^b)</td>
</tr>
</tbody>
</table>

Note: Population at risk at first reference month (November 2003 if already married or first month of marriage). Values are weighted to represent the US population.

\(^a\)Denotes statistical difference between men and women at significance level, 0.05. \(^b\)Denotes difference from being self-insured at significance level 0.05.
Table 1.3 Earnings Contribution to Total Household Income by Insurance and Gender in Percentages

<table>
<thead>
<tr>
<th>Insurance Status</th>
<th>Men</th>
<th>Women</th>
<th>Overall</th>
</tr>
</thead>
<tbody>
<tr>
<td>Insured under own name</td>
<td>63.11</td>
<td>41.91</td>
<td>54.96</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>41.96</td>
<td>18.64</td>
<td>24.52</td>
</tr>
<tr>
<td>Gov't Insurance (Medicare, Medicaid)</td>
<td>26.39</td>
<td>13.45</td>
<td>18.14</td>
</tr>
<tr>
<td>Uninsured</td>
<td>58.97</td>
<td>22.47</td>
<td>39.58</td>
</tr>
</tbody>
</table>

*Note:* Population at risk at first reference month (November 2003 if already married or first month of marriage). N = 17,388 (men n = 8,091; women n = 9,297). Values are weighted to represent the US population. Total household income includes income from property and means-tested cash transfers.
Table 1.4 Test of Hypothesis 1

Model 1. Cox Regression of Divorce Hazard on Insurance Status (hazards in odds ratios)

<table>
<thead>
<tr>
<th>Insurance Status</th>
<th>Coefficient</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Insured under own name</td>
<td>(reference)</td>
<td>a</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>0.32***</td>
<td>b</td>
</tr>
<tr>
<td>Gov't Insurance (Medicare, Medicaid)</td>
<td>0.59***</td>
<td></td>
</tr>
<tr>
<td>Uninsured</td>
<td>0.73*</td>
<td></td>
</tr>
</tbody>
</table>

Logged family monthly income 0.44***

Note: Model includes age, age-squared, race, education, children, higher-order marriage, and marriage duration as controls. Coefficients are not shown. N=17,388 (men n=8,091; women n=9,297). Values are weighted to represent the US population. *p < .05. **p < .01. ***p < .005.

Test of Hypothesis 1: Married individuals who are insured through their spouses’ health plans have lower rates of divorce

<table>
<thead>
<tr>
<th>Key Coefficient for Hypothesis Test</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>a Insured under own name (reference group)</td>
<td>1.00</td>
</tr>
<tr>
<td>b Insured under someone else's plan</td>
<td>0.32***</td>
</tr>
</tbody>
</table>

Ratio of b to a 0.32***

Note: P-values of one-sided t-tests are corrected adjusted for False Discovery Rate (Benjamini and Hochberg 1995). *p < .05. **p < .01. ***p < .005.
Table 1.5 Test of Hypothesis 2

Model 2. Cox Regression of Divorce Hazard on Insurance Status and Access to Employment-based Option (hazards in odds ratios)

Two-way interaction between insurance status and access to employment-based option

<table>
<thead>
<tr>
<th>Insurance Status</th>
<th>Access to Employment-based Option</th>
<th>Coefficient</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Insured under own name</td>
<td>Yes</td>
<td>(reference)</td>
<td></td>
</tr>
<tr>
<td>Insured under own name</td>
<td>No</td>
<td>0.59*</td>
<td></td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>Yes</td>
<td>0.43***</td>
<td>a</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>No</td>
<td>0.18***</td>
<td>b</td>
</tr>
<tr>
<td>Gov't Insurance (Medicare, Medicaid)</td>
<td>Yes</td>
<td>1.03</td>
<td></td>
</tr>
<tr>
<td>Gov't Insurance (Medicare, Medicaid)</td>
<td>No</td>
<td>0.37***</td>
<td></td>
</tr>
<tr>
<td>Uninsured</td>
<td>Yes</td>
<td>1.03</td>
<td></td>
</tr>
<tr>
<td>Uninsured</td>
<td>No</td>
<td>0.39***</td>
<td></td>
</tr>
</tbody>
</table>

Logged family monthly income 0.41***

*Note: Model includes age, age-squared, race, education, children, higher-order marriage, and marriage duration as controls. Coefficients are not shown. N=17,388 (men n=8,091; women n=9,297). Values are weighted to represent the US population.

*p < .05. **p < .01. ***p < .005.

Test of Hypothesis 2: Not having an employment-based source of health insurance coverage outside the marriage further lowers the divorce risk of people enrolled in their spouses’ plans.

Key Coefficients for Hypothesis Test

<table>
<thead>
<tr>
<th>Coefficient</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>a Insured under someone else's plan &amp; has employment-based option</td>
<td>0.43***</td>
</tr>
<tr>
<td>b Insured under someone else's plan &amp; has no employment-based option</td>
<td>0.18***</td>
</tr>
</tbody>
</table>

Ratio of b to a 0.41***

*Note: P-values of one-sided t-tests are corrected adjusted for False Discovery Rate (Benjamini and Hochberg 1995).

*p < .05. **p < .01. ***p < .005.
### Table 1.6 Test of Hypothesis 3

**Model 3. Cox Regression of Divorce Hazard on Insurance Status and Gender (hazards in odds ratios)**

Two-way interaction between insurance status and gender

<table>
<thead>
<tr>
<th>Insurance Status</th>
<th>Gender</th>
<th>(reference)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Insured under own name</td>
<td>Male</td>
<td>(reference)</td>
</tr>
<tr>
<td>Insured under own name</td>
<td>Female</td>
<td>1.58***</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>Male</td>
<td>0.48**</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>Female</td>
<td>0.37***</td>
</tr>
<tr>
<td><strong>Gov't Insurance (Medicare, Medicaid)</strong></td>
<td>Male</td>
<td>0.41*</td>
</tr>
<tr>
<td><strong>Gov't Insurance (Medicare, Medicaid)</strong></td>
<td>Female</td>
<td>0.88</td>
</tr>
<tr>
<td>Uninsured</td>
<td>Male</td>
<td>0.91</td>
</tr>
<tr>
<td>Uninsured</td>
<td>Female</td>
<td>0.89</td>
</tr>
<tr>
<td>Logged family monthly income</td>
<td></td>
<td>0.44***</td>
</tr>
</tbody>
</table>

*Note: Model includes age, age-squared, race, education, children, higher-order marriage, and marriage duration as controls; coefficients are not shown. N=17,388 (men n=8,091; women n=9,297). Values are weighted to represent the US population. *p < .05. **p < .01. ***p < .005.*

**Test of Hypothesis 3: Women who are insured on their spouse’s health plans have lower rates of divorce than men who are insured by their spouse**

<table>
<thead>
<tr>
<th>Key Coefficients for Hypothesis Test</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>a Insured under someone else's plan &amp; male</td>
<td>0.48**</td>
</tr>
<tr>
<td>b Insured under someone else's plan &amp; female</td>
<td>0.37***</td>
</tr>
</tbody>
</table>

**Ratio of b to a**

| 0.78 |

*Note: P-values of one-sided t-tests are corrected adjusted for False Discovery Rate (Benjamini and Hochberg 1995). *p < .05. **p < .01. ***p < .005.*
### Table 1.7 Test of Hypothesis 4

**Model 4. Cox Regression of Divorce Hazard on Insurance Status, Access to Employment-based Option, and Gender (hazards in odds ratios)**

Three-way interaction between insurance status, access to employment-based option, and gender

<table>
<thead>
<tr>
<th>Insurance Status</th>
<th>Access to Employment-based Option</th>
<th>Gender</th>
</tr>
</thead>
<tbody>
<tr>
<td>Insured under own name</td>
<td>Yes</td>
<td>Male (reference)</td>
</tr>
<tr>
<td></td>
<td>Yes</td>
<td>Female 1.70***</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Male 0.75</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Female 0.75</td>
</tr>
<tr>
<td>Insured under someone else's plan</td>
<td>Yes</td>
<td>Male 0.44*</td>
</tr>
<tr>
<td></td>
<td>Yes</td>
<td>Female 0.61*</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Male 0.49</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Female 0.19***</td>
</tr>
<tr>
<td>Gov't Insurance (Medicare, Medicaid)</td>
<td>Yes</td>
<td>Male 0.22</td>
</tr>
<tr>
<td></td>
<td>Yes</td>
<td>Female 2.03*</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Male 0.41*</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Female 0.49***</td>
</tr>
<tr>
<td>Uninsured</td>
<td>Yes</td>
<td>Male 1.13</td>
</tr>
<tr>
<td></td>
<td>Yes</td>
<td>Female 1.49</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Male 0.52*</td>
</tr>
<tr>
<td></td>
<td>No</td>
<td>Female 0.47**</td>
</tr>
<tr>
<td>Logged family monthly income</td>
<td></td>
<td>0.40***</td>
</tr>
</tbody>
</table>

*Note: Model includes age, age-squared, race, education, children, higher-order marriage, and marriage duration as controls. Coefficients are not shown. N=17,388 (men n=8,091; women n=9,297). Values are weighted to represent the US population.

*p < .05. **p < .01. ***p < .005.

Test of Hypothesis 4: Not having an alternative source of health insurance outside the marriage lowers divorce risk for women more so than for men

<table>
<thead>
<tr>
<th>Key Coefficients for Hypothesis Test</th>
</tr>
</thead>
<tbody>
<tr>
<td>c Insured under someone else's plan &amp; has no employment-based option &amp; male</td>
</tr>
<tr>
<td>d Insured under someone else's plan &amp; has no employment-based option &amp; female</td>
</tr>
</tbody>
</table>

Ratio of d to c 0.37*

Not having an alternative source of health insurance outside the marriage lowers divorce risk **for men**

<p>| | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>a Insured under someone else's plan &amp; has employment-based option &amp; male</td>
<td>0.44*</td>
</tr>
<tr>
<td>c Insured under someone else's plan &amp; has no employment-based option &amp; male</td>
<td>0.49</td>
</tr>
</tbody>
</table>

Ratio of c to a 1.12

Not having an alternative source of health insurance outside the marriage lowers divorce risk **for women**

<p>| | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>b Insured under someone else's plan &amp; has employment-based option &amp; female</td>
<td>0.61*</td>
</tr>
<tr>
<td>d Insured under someone else's plan &amp; has no employment-based option &amp; female</td>
<td>0.18***</td>
</tr>
</tbody>
</table>

Ratio of d to b 0.30***

*Note: P-values of one-sided t-tests are corrected adjusted for False Discovery Rate (Benjamini and Hochberg 1995).

*p < .05. **p < .01. ***p < .005.
Chapter 2: Life-Course Mortality Consequences of Public Health Insurance in the United States

ABSTRACT

Forty-four percent of the Affordable Care Act’s 1.1 trillion-dollar expenditure is projected to be spent on expanding Medicaid. The impact of Medicaid on its recipients and on the broader population has been a matter of contentious debate since its inception in 1965. My contribution to this debate is two-fold. First, I examine infant mortality improvements between 1959 and 1979 that accompanied states’ Medicaid participation. Here, I exploit the variation in when each of the 50 States adopted Medicaid to estimate its impact on national infant mortality rates. The annual rate of infant mortality decline doubled with Medicaid participation. Mortality predictions without Medicaid result in 78,000 excess infant deaths between 1965 and 1980. Second, I examine Medicaid’s lasting consequences that persist into adulthood. Cohorts born post-Medicaid experienced greater improvements in mortality even after aging out of eligibility. Medicaid is still shaping the US population today. A cohort born in 1970 without Medicaid would have 13,700 fewer people aged 40 in 2010.
Title II of the Affordable Care Act (ACA) aims to expand Medicaid eligibility to an estimated 17 million more Americans (Congressional Budget Office 2011). This is the largest initiative for public health insurance since the creation of Medicaid and Medicare as a part of the 1965 Social Security Amendments. Medicaid is a means-tested government insurance program aiming to provide health insurance to people who cannot afford to pay for private health policies. Researchers and policy-makers debate over whether this single-payer government-run insurance program is an effective way to deliver health care services to low-income and medically needy people (Kaiser Family Foundation 2013). Critics of Medicaid claim that this public insurance not only fails to adequately benefit its target population, but it also has detrimental effects on the privately insured (McDonough 2011). A large body of literature examining the effects of Medicaid on its recipients has emerged since the program’s inception (Baldwin et al. 1998; Braveman et al. 1993; Copeland and Meier 1987; Currie and Grogger 2002; Devany et al. 1992; Guyer 1990; Howell 2001; Lykens and Jargowsky 2002; Moss and Carver 1998; Schor et al. 2007). However, few if any studies examine Medicaid’s contribution to general population health and its long-term consequences. In this paper, I examine changes in the rates of decline in states’ infant, childhood, and adult mortality associated with the availability of Medicaid at birth. I exploit the variation in States’ timing of Medicaid participation between 1966 and 1979 to address two research questions. (1) Did Medicaid reduce infant mortality rates at the population level? (2) Did the availability of Medicaid at birth have lasting consequences on mortality into adulthood? Macro-level views of the consequences of large-scale health care reforms are informative as the policies often influence individuals outside their immediate targeted population. And
these health effects during early life may endure into adulthood. This paper is the first to follow birth cohorts over time to examine lasting mortality consequences associated with Medicaid.

THEORY AND LITERATURE

The first research question adds to the literature by examining the overall changes in infant mortality rate—recipients and non-recipients alike—associated with States’ Medicaid participation. Medicaid was one of the largest U.S. federal initiatives to disseminate medical innovations and health information to pregnant women and infants in the 1960s. By increasing the number of women who seek pre- and postnatal care, Medicaid aimed to improve nutrition and health behaviors among expectant and new mothers (David and Seigel 1983; Lee et al. 1980). This in turn would lead to healthier infants and lower mortality rates among women who would not otherwise have had access to health care. Medicaid targeted low-income women and infants, a group that was more susceptible to infant mortality. Improving the health of this disadvantaged group would effectively lower the average infant mortality for the overall population.

Medicaid may also have had spillover effects on women and infants who did not gain insurance coverage from the public program. Theories in diffusion suggest that as a greater proportion of women gain access to pre- and post-natal care, the health of all women would improve. Health knowledge spreads through interpersonal networks as well as institutional organizations (Backer 1991; Tarde 1962; Green et al. 2009). Interaction with peers, friends, and family as well as health care providers influence how individuals approach their health (Christakis and Fowler 2007, 2009; Glanz et al. 2008). Social network effects on health behavior are not trivial. Groups of socially connected
people quit smoking in concert at different times despite the already ‘generalized knowledge’ of smoking’s adverse effects on health (Christakis and Fowler 2009). Medicaid connects more women with health care providers and encourages transfer of health knowledge from physicians to the public. As more women adopt good nutrition and refrain from risky health behaviors, the more likely they are to influence other women in their social networks, even those who always had access to health care services.

Researchers, for the most part, agree that Medicaid increased insurance coverage and prenatal care utilization, but many disagree on the program’s impact on population health. While Medicaid fails to ensure that all persons who are eligible have health insurance (Braveman et al. 1993; Currie and Grogger 2002; Holahan and Zedlewski 1991; Kenney and Haley 2001; Wilensky and Berk 1982), the introduction of the public health insurance program visibly improved insurance coverage and health care utilization in the United States. Prior to 1965, only half of low-income Americans had medical-coverage (Copeland and Meier 1987) and despite having poorer health than their wealthier counterparts, low-income individuals and families used fewer medical services (Copeland and Meier 1987; Rowland et al. 1988; Wilensky and Berk 1982). In particular, 88% of upper-income women reported seeing a physician during the first trimester of pregnancy compared to 58% of lower-income pregnant women (Andersen and Andersen 1967; Copeland and Meier 1987). By the late 1970s when most states had adopted Medicaid, 24 million Americans and over a third of people in households with incomes below 125% FPL received health insurance through Medicaid (Oberg and Polich 1988; Wilensky and Berk 1982). Medicaid specifically targeted low-income pregnant women
and infants and their enrollee characteristics reflected this trait (Wilensky and Berk 1982). Medicaid covered about three-quarters of low-income pregnant mothers and children under the age of 5, reducing the rate of uninsurance for these target groups to below the national average (Holahan and Zedlewski 1991). By 1980, the low-income group was utilizing medical services (hospital stays and physician visits) at an equivalent rate as the higher-income groups (Copeland and Meier 1987; Rowland et al. 1988; Wilensky and Berk 1982).

Critics claim that greater insurance coverage and health care utilization through Medicaid do not necessarily translate into improved population health. Medicaid recipients may receive inferior medical care than those who are privately insured. The low pay-out to health care providers who treat Medicaid patients incentivize doctors to limit the number of Medicaid patients they are willing to accept (Decker 2012). Medicaid patients find themselves waiting longer for care and having a narrower selection of doctors (Bisgaier and Rhodes 2011; Merrick, et al. 2001). Several studies also show that Medicaid patients receive poorer quality care and face worse outcomes from the same doctor relative to comparable individuals with private insurance (Hwang et al. 2005; Wang et al. 2004). Criticisms against Medicaid’s effectiveness do not stop at the uninsured, low-income population that the program is designed to target. Medicaid induces eligible people who already have private insurance to switch to the arguably inferior public health insurance policy (Blumberg et al. 2000; Dubay and Kenney 1996; Wilensky and Berk 1982). Some further claim that health care providers shift the cost of caring for Medicaid patients onto privately insured patients (McDonough 2011).
Empirically measuring the health outcome of Medicaid is tricky as socioeconomic class—a well-established determinant of health (Cutler et al. 2008, Elo 2009; Link and Phelan 1995; Lynch et al. 2004; Smith 2007)—is explicitly tied to eligibility (Lyken and Jargowsky 2002). Copeland and Meier (1987) use Medicaid expenditures in their time-series model to distinguish infant mortality decline attributable to increasing access to care from technological advances in medicine. They find a significant negative association between federal Medicaid expenditures and the infant mortality rate. However, their simple study design was inadequate to establish a convincing relationship between Medicaid and infant mortality. Grossman and Jacobowitz (1981) utilize disparate Medicaid eligibility rules to estimate the change in county-level neonatal mortality rate associated with the state’s Medicaid coverage. The paper attempts to control for the composition differences between counties by estimating the proportion of births to low-income mothers. They also include proportion of women with high-school degrees and the number of physicians per 1000 to control for socioeconomic differences between counties. Their analyses find little support for Medicaid in contributing to the decline in neonatal mortality rates between 1964 and 1977. The limitation of this study lies in its assumption that counties that share these few characteristics are essentially equal, apart from the availability of Medicaid and other maternal and infant health programs. Currie and Gruber (1994) find the most convincing evidence that Medicaid expansions in the 1980s contributed to reductions in infant mortality rates. They standardized States’ Medicaid expansions by simulating the proportion of women who would be eligible from a nationally representative sample of 3,000 women from the Current Population Survey each year. This method allowed the authors to effectively
isolate the effects of States’ extent of Medicaid expansions from the composition of their residents.

In sum, the introduction of Medicaid in 1965 led to a greater proportion of pregnant women seeking prenatal care but the program’s contribution to overall population health is unclear. The literature is mostly focused on determining Medicaid’s health effects on its recipients. This paper contributes to this literature by examining Medicaid’s impact on the entire population.

Population-level analyses of other high-income OECD countries show significant associations between infant mortality and the country’s health care system. Countries with publicly funded health care had lower infant mortality rates than similar countries whose health care services are generally private (Elola et al. 1995). National health care systems were associated with greater improvements in infant mortality after accounting for secular declines and changes in GDP (Macinko et al 2004). Between 1970 and 1996, public health care systems were associated with over 3 less infant deaths per 1000 live births as IMR fell from 16.63 to 6.20 per 1000. These publicly funded national health systems also attenuated the positive association between income inequality and infant mortality (Macinko et al 2004). While Medicaid is not a national health system, it gave access to public health care to a large subset of the US population. The program specifically targeted groups who made large contributions to the national IMR. I expect Medicaid to have similarly positive associations with infant mortality improvements in the US.

The second part of this paper shows that cohorts who were born after their states adopted Medicaid had greater improvements in probability of survival throughout
childhood and into adulthood compared to cohorts born immediately prior to Medicaid. Women who receive prenatal care and thus eat better and refrain from risky behaviors are more likely to give birth to healthier babies (Almond 2005; Oreopolous et al. 2008). These babies not only have higher chances of survival, but would also lead healthier lives throughout childhood and adulthood. If Medicaid improves infant health by disseminating health knowledge and behaviors to pregnant women, it would then lead to healthier newborns who would grow up to be healthier adults.

While no study has examined the long-term consequences of being born into an era of Medicaid, many studies have made the connection between prenatal and early-life environmental conditions to later-life health outcomes. One of the most well-known of these studies finds a positive link between ischemic health disease mortality and infant mortality rates at the place and time of birth (Barker and Osmond 1986). Their study suggests that early-life living conditions and nutrition (measured by local infant mortality rates) have long-term consequences well into adulthood even when they move to another region. Another study find higher mortality rates throughout the life course among cohorts born during a macroeconomic recession (Van den Berg et al. 2006). Quasi-experimental studies also find poorer health outcomes among birth cohorts born during the 1918 Influenza Pandemic and the 1944 Dutch Famine relative to cohorts born immediately before or after these sudden deteriorations in living conditions (Almond 2006; Roseboom et al. 2001). The literature is strongly suggestive of initial conditions at birth having long-lasting health consequences. Medicaid explicitly aims to improve prenatal health and birth outcomes. Thus, I expect these wide-spread efforts to translate into improved health throughout the life-course.
HISTORICAL CONTEXT

Before delving into the analyses, I briefly describe the historical decline in infant mortality rates leading up to the advent of Medicaid and the policy context of States’ Medicaid participation during the period after 1965.

A rise in living standards, better nutrition, public sanitation, clean water supply, and public health campaigns contributed to dramatic reductions in infant mortality rates in the United States during the first half of the Twentieth Century (Condran and Crimmins-Gardner 1978; Cutler and Miller 1995; Deaton and Paxton 2001; Elo and Preston 1996; Ewbank and Preston 1990; Fogel 2004; McKeown 1976, 1979; Meeker 1972; Preston and Haines 1991; Szreter, 1988). After several decades of experiencing significant improvements, the year-to-year decline in infant mortality stalled to 0.5 percent by 1950 (Corman and Grossman 1985). While mortality from common childhood infectious diseases had reached very low levels by 1950 (Armstrong et al. 1999) considerable national attention was focused on how infant health improvements in the US were lagging behind other developed nations (Committee on Maternal and Child Care 1965; Falkner 1969; Lee et al. 1980; Shapiro et al. 1968). Many efforts were made to improve prenatal and neonatal medical technology as well as to make services more accessible (Corman and Grossman 1985). Medicaid, a means-tested public insurance program targeting low-income pregnant women and infants, was created by Title XIX of the Social Security Act in 1965. The period beginning 1965 experienced twentieth century’s second surge of infant mortality decline (Grossman and Jacobowitz 1981; Lee et al. 1980). Infant mortality declined 4.5 percent per year between 1965 and 1982 (Corman and Grossman 1985). Increased access to care through Medicaid is often
credited as one of the contributors to the rapid decline in infant mortality rates in the decade after 1965 (Corman and Grossman 1985; Mason 1991). But no prior study has attempted to quantify mortality improvements due to Medicaid.

The analyses in this paper estimate the change in mortality decline associated with Medicaid by comparing state’s annual mortality improvements pre- and post-Medicaid participation. This regression discontinuity framework utilizes States’ differences in when they adopted Medicaid. States began to join the federal Medicaid program over sixteen years beginning in 1966, a year after Medicaid was created by the Social Security Amendments of 1965.

The program gave federal grants to states to provide health insurance to eligible persons. The federal government gave considerable flexibility in when or whether each state could participate in the program. The federal government mandated participating states to provide coverage to pregnant mothers and infants that met income and asset requirements. The State and the federal government would be jointly responsible for the costs associated with Medicaid. States began to participate in Medicaid quickly after the Title XIX of the Social Security Act in 1965. Six states adopted Medicaid in January 1966. 20 more followed later that year. Eleven joined in 1967, one joined in 1968, and by 1970, most of the states had adopted Medicaid. Appendix A shows the dates of when each State implemented Medicaid. Arizona is the last of the States to adopt Medicaid in 1982. Appendix B displays the same data on the US map. The first states to adopt Medicaid in January 1966 were located throughout the entire nation. States in the Northeast were relatively early adopters except New Jersey who was among the last. Southern states were generally late in adopting Medicaid with some exceptions such as
Georgia, Oklahoma, and Louisiana. States’ timing of Medicaid adoption does not appear to follow the 1964 presidential election results. Heavily republican states such as Louisiana, Nebraska, Utah, Georgia, and Oklahoma were among the first to implement Medicaid in 1966. Democratic New Jersey and Alaska were among the last. I later show in my analysis that the timing had little bearing on states’ mortality improvements after Medicaid adoption.

States’ timing of Medicaid implementation was also different from other public policies that the Johnson administration introduced to combat poverty in the 1960s and 1970s. The Food Stamp Act of 1964 was the other major initiative targeting women and children to improve prenatal health and reduce infant mortality (Almond et al. 2011). However, the rollout of the Food Stamp Program did not coincide with states’ Medicaid implementation dates. The Food Stamp Program was implemented at the county-level and its implementation stretched out between 1961 and 1976. Thirty percent of the US population already had access to food stamps by the time the first states began to offer Medicaid in January 1966. About a quarter of the population gained access to food stamps after 1970 when all but two states had joined the Medicaid program. Furthermore, states’ Medicaid participation had little bearing on its counties’ food stamp participation start date. Most notably, most counties in New York, Oklahoma, and Massachusetts began to offer food stamps relatively late, despite being early adopters of Medicaid. Many counties in Texas and California did not offer food stamps till after 1970. Thus, the exact timing of Medicaid implementation appears exogenous to other prenatal and infant health policy initiatives that occurred during this period.
DATA

I rely on the US Vital Statistics micro-data from the National Center for Health Statistics (NCHS) for my analyses. I examine the mortality patterns for birth cohorts born between 1959 and 1979—the 20-year period surrounding the introduction of Medicaid in 1965. I start my analysis in 1959 when Hawaii became the last state to join the US. I use the natality micro-data for births between 1959 and 1979 to calculate the number of births by year for each state. The NCHS allocates births to the state of the mother’s residence and does not include births to US citizens outside the United States. To derive the age-specific mortality rates for cohorts born between 1959 and 1979, I use the mortality micro data from 1959 to 2010. The NCHS constructed the mortality microdata from death certificates filed in vital statistics offices of each State. Mortality data for New Jersey is missing for the years, 1962 and 1963. In 1972, NCHS processed only a 50 percent sample of death records. I have multiplied the number of deaths in 1972 by a factor of two in my analyses.

I limit the analyses in this paper to the white US population. The US Vital Statistics changed the way it categorized race in their birth and death certificates in 1968. Prior to 1968, people were categorized as either white or non-white. After 1968, the non-white population was classified into sub-groups. The exact list of sub-groups changed multiple times between 1968 and 2010. The changes in how the data classifies race makes it difficult to compare minority groups across the years. In addition, the number of non-white being born and dying in some states were so small that the age-specific mortality rates quickly became unreliable.
ANALYTIC STRATEGY

I exploit the variation in State’s timing of Medicaid participation to examine the improvements in US mortality rates attributed to the public insurance program. I estimate the change in the annual rate of mortality decline after Medicaid adoption using state fixed effects weighted by the number of births from each state. All regression models have panel-corrected standard errors and they also correct for heteroskedasticity. This method relies on Medicaid adoption timing to be exogenous to other factors that may influence mortality rates. I evaluate several factors that may undermine this framework.

First, states’ baseline infant mortality in 1965 had no relation to when they joined the Medicaid program. Utah who had the lowest infant mortality rate adopted Medicaid at the same time as West Virginia who had the highest infant mortality. A low correlation coefficient across all states further confirms that states’ infant mortality rates in 1965 did not determine its Medicaid participation.

Second, Medicaid adoption year did not determine how much improvement a state experienced after participation. I separately examined each state’s change in the rate of infant mortality decline after Medicaid implementation. Implementation year was not correlated with the magnitude or significance of their IMR improvements. For example, Illinois, Indiana, and Iowa all experienced statistically significant IMR improvements of about -0.53 deaths per 1000 per year after they adopted Medicaid. Their adoption dates were 1966, 1970, and 1968 respectively. Thirty-three out of the forty-eight states that joined Medicaid prior to 1970 had a statistically significant infant mortality improvement that coincided with their Medicaid participation years. And, these thirty-three states’ Medicaid adoption dates were spread across all four years between 1966 and 1970.
Third, systematic migration of would-be mothers of healthy babies into states offering Medicaid is unlikely. Interstate migration flows of a demographic group that would produce the healthiest babies (college-educated persons between the ages 25 to 39) have no relations to States’ timing of Medicaid participation. California, an early adopter, and Florida, a late adopter, were among the top recipients of the young, college educated population in the late 1960s (Goworowska and Gardner 2012). Young, single, college-educated people consistently out-migrated from Minnesota and Alabama despite one being an early Medicaid adopter and the other, a late adopter. Furthermore, the magnitudes of these flows are small. Thirteen percent of persons aged 25 to 39 moved to a different state between 1965 and 1970. Out of the thirteen percent, less than thirty percent were college educated.

Lastly, infant mortality rate declines associated with state Medicaid participation remain significant, independent of period-specific effects. The stalling infant mortality improvements in the 1950s and early 1960s drew the attention of policy makers. The federal enactment of the Medicaid program was accompanied by increased funding to the NIH for prenatal and neonatal research (Shapiro 1981). An influential medical innovation that coincides with several states’ Medicaid adoption can overstate IMR improvements attributed to Medicaid. To test this, I add a term to indicate whether the birth occurred before or after a particular year between 1960 and 1970 in a series of regressions. For example, the first of this series will add infant mortality rate on a dummy variable for post-1960 births to the original model. The inclusion of these terms did not substantially change the magnitude or the significance of infant mortality reductions attributed to State Medicaid participation.
RESULTS

Research Question 1: Is Medicaid associated with greater declines in US infant mortality rates?

I examine the change in the annual rate of infant mortality decline after States’ implementation of Medicaid to address my first research question. I calculate the infant mortality rate for each state and year between 1959 and 1979 as follows.

\[ IMR_{y,s} = \frac{\text{Number of deaths for infants under age one residing in state } s, \text{ in year } y}{1000 \text{ live births to mothers who are residents of state } s, \text{ in year } y} \]

Figure 2.1 shows the declining trend of US infant mortality rates between 1959 and 1979. Confirming prior studies (Corman and Grossman 1985; Grossman and Jacobowitz 1981; Lee et al. 1980) the national infant mortality rate shows an accelerated decline during the period after 1965. The infant mortality rate decreased on average 1.2 percent (at a rate of 0.28 infant deaths /1000 live births) per year between 1959 and 1965. Between 1965 and 1979, the infant mortality rate decreased 4.4 percent (0.72 infant deaths/1000 live births) per year. It is also during this latter period that States began to participate in the federal Medicaid program. Six States were the first to implement Medicaid in January 1966. Twenty more joined by the end of the year and eleven States implemented Medicaid the following year in 1967. By 1970, 48 States were offering Medicaid to their residents.

In Figure 2.2, I show the changes in the annual rates of infant mortality after Medicaid implementation for each state. The sizes of the circles in the graph represent the number of births. Both large and small states saw significant mortality improvements after Medicaid. And these improvements do not depend on which year the states joined.
the program. Thirty-three out of 50 states experienced faster declines in infant mortality after Medicaid. Montana, Idaho, Mississippi, West Virginia and the New England states experienced the largest improvements. The two largest states, New York and California, saw moderate improvements. The 17 states that did not have significant improvements in infant mortality are scattered throughout the West, Midwest, and the South. Texas is the largest state that did not appear to have faster mortality declines after Medicaid adoption.

Figure 2.3 descriptively shows faster improvements in infant mortality after each state’s Medicaid implementation relative to the years leading up to Medicaid. The graph shows average logged ratios (weighted by number of live births) of IMR relative to the IMR the year before Medicaid took effect. Figure 2.3 accounts for variations in states’ timing of Medicaid and their baseline IMR levels. The rate of IMR decline appears to be almost three times as fast during the seven years following Medicaid than during the seven year leading up to the program’s implementation.

Table 2.1 describes the weighted yearly rate of decline in infant mortality during the years before and after States’ Medicaid participation. Infant mortality declined at a rate of 0.38 deaths per 1000 births per year during the years before Medicaid implementation. This rate increased to 0.71 deaths per 1000 births per year after the introduction of Medicaid. The model estimates incremental mortality improvement of 0.33 deaths per 1000 births per year after Medicaid implementation. All coefficients are significant at the 0.001 level. The mortality improvements associated with Medicaid adoption is not trivial. Had infant mortality improvements stayed constant at pre-Medicaid rates, the IMR among US whites would be 15.57 in 1979—about 4.11 deaths per 1000 in excess of the estimated 11.46 with the presence of Medicaid (Figure 2.2).
This translates to approximately 78,000 white babies between 1966 and 1979 that would have not survived infancy.

Research Question 2: Did the availability of Medicaid at birth have lasting consequences on mortality throughout childhood and into adulthood?

To address my second research question, I examine changes in the cumulative mortality between cohorts born before and after Medicaid in ten-year age periods. The eldest age group in this analysis is the 30 to 39 age period. I calculate age-specific mortality probabilities from life tables of cohorts born between 1959 and 1979. I then derive the cumulative mortality probabilities for each ten-year age period from each birth cohort.

Figure 2.4 shows the mortality trends of cohorts born between 1959 and 1979 by ten-year age periods. Unlike infant mortality, mortality for older age groups did not experience steady and constant declines during this period. In fact, mortality rates appear to increase after 1970 (Murphy et al. 2013). Mortality data ends in 2010 when the 1970 birth cohort is 40 years old. Constricted by data limitations and possible confounding effects from later mortality trends, I limit this section of the analysis to the 1959-1970 birth cohorts.

All states except Arizona and Alaska had implemented Medicaid by the end of 1970. I estimate the change in mortality probability associated with the availability of Medicaid in the state at the time of birth. The regressions also include state fixed-effects with panel-corrected standard errors and corrections for heteroskedasticity. All age groups displayed improvements in mortality associated with the availability of Medicaid at birth. Table 2.2 presents the change in the annual rate of mortality improvement associated with Medicaid implementation at the state level. All coefficients in Table 2.2
are significant at the 0.001 level. Cohorts born after Medicaid became available in their states experienced greater rates of mortality improvements at all age groups than cohorts born prior to Medicaid implementation. The largest differences between the pre- and post-Medicaid cohorts appear in the 1-9 and 20-29 age groups. The cumulative mortality improvement rate increased by over 300 percent from 0.005 to 0.022 and 0.01 to 0.034 percent per year respectively. While not as large, the rate of mortality decline at ages 10 to 19 and 30 to 39 still accelerated significantly after Medicaid implementation. The annual rate of improvement increased from 0.011 to 0.017 and 0.023 to 0.044 percent per year respectively. Figure 2.5 presents expected mortality probabilities for cohorts born between 1964 and 1970 with and without the estimated effects of Medicaid. The mortality probabilities are lower in predictions that account for Medicaid in all four age groups and the differences becomes larger in more recent cohorts.

These percentage differences may appear small but their implication on the US population is substantial. The final component of this paper show how the 1970 birth cohort would differ without the mortality effects of Medicaid estimated in the prior models. Table 2.3 presents the 1970 cohort’s predicted under-40 mortality schedule with and without the estimated effects of Medicaid. The number of survivors at the beginning of each age group (lx) is based on the actual number of white births in the United States in 1970. I apply the different age-specific mortality probabilities to the 1970 birth cohort and compare the resulting number of survivors at age 40 in 2010. Without mortality improvements associated with Medicaid, out of the 3.1 million people born in 1970, 13,681 fewer people would be alive in 2010. 13,681 people represent an 8.7 percent
increase in under-40 mortality for the 1970 cohort and 0.4 percent of all live births in 1970.

LIMITATIONS

The analysis is presented in the paper has several limitations. First, the analysis is limited to the US white population. The poor data quality for the non-white population in the late 1950s and 1960s does not produce dependable results. Many states during this period had very few non-white births and even fewer deaths. Age-specific death rates deteriorate into noisy trends for many smaller states. In addition, the NCHS changed the categorization of non-whites several times during this period. As the immigrant population grew in the 1960s and 1970s, the Vital Statistics separated Hispanics and Asians from the African American population. During the early years, all non-white racial and ethnic groups were reported as simply, non-white. Medicaid likely had a greater influence on the non-white population. A greater proportion of non-whites would have been eligible for Medicaid as poverty and single motherhood were more prevalent. Infant mortality among non-whites during this period was also substantially higher and increasing their access to medical care may have had a greater impact on infant mortality. Given these considerations, the changes in US mortality attributed to Medicaid in this paper may be a conservative estimate of its actual effect.

The second limitation is the analyses’ assumption that deaths occurred in the state of the person’s birth. I derive the age-specific death rates using actual number of deaths by year, age, and state and the cohort’s corresponding number of births in the same state. These age-specific death rates inflate from deaths of persons born out-of-state and deflate when the net-migration of people born in those states are below zero. Limiting the
analysis to the white population has substantially reduced inflation from the large growth of immigration into the United States during the 1970s and 1980s. Immigration from Europe and Canada remained relatively constant during this period at rate of approximately 100,000 per year (Fix et al. 1994). The cumulative number of these immigrants would have been less than 1% of the US population in 1990. Furthermore, they would have contributed even less to mortality at younger ages which is the focus of this paper. Moving to another state after birth can also affect States’ mortality schedules. It is unlikely that migration patterns coincide with the availability of Medicaid at a person’s time of birth. Migration between states would more likely dilute the estimated effects of Medicaid on mortality making this paper’s results conservative.

This study is also limited by the length of its data. The 1965 birth cohort is only 45 years old in 2010 and the lasting effects of Medicaid could only be observed at relatively low mortality ages. Prior literature suggests that the effects of early-life health environment become more prominent at older ages when mortality rates increase. Similarly, the association between Medicaid availability at birth and adult mortality may become stronger in later years.

DISCUSSION

Faster declines in infant mortality accompanied the advent of Medicaid in the late 1960s. These improvements associated with the federal public insurance program were substantial. The results from this paper’s analysis show that the annual rate of infant mortality decline doubled after states’ Medicaid participation. It made an impact on the overall population health by targeting women and infants who had the least access to health care and the highest levels of infant mortality. Improving the health of the neediest
group raised the average survival rate of the overall population. Moreover, Medicaid may also have had a positive health impact on women and infants who did not directly benefit from the public program. It served as a conduit for health knowledge to travel from researchers and physicians to the general public. As more women interacted with the medical system and changed their health beliefs and behavior, they reinforced the health knowledge of other women. Medicaid enhanced concurrent public health initiatives to reduce infant mortality. Combined, these factors lead to the large infant mortality improvements observed in the United States in the period after 1965.

The health impact of Medicaid in-utero and in infancy was significant enough to yield later-life consequences. Cohorts born after their States adopted Medicaid experienced greater mortality decline throughout childhood and into adulthood, well past their eligible age. In fact, mortality improvements associated with Medicaid are greater at older age groups. This paper draws attention to the significance of early life access to health care in-utero and during infancy in improving population health throughout the life-course.

These findings suggest what the United States may expect after the rollout of the Affordable Care Act. Proposed Medicaid expansions will provide access to estimated 17 million adults who currently live without health care coverage. Younger adults who comprise a large proportion of the uninsured would be most impacted by these expansions. These adults will have better access to care and have more frequent interactions with health care providers. In a similar manner to pregnant women and prenatal care, people who have greater interaction with health care providers are more likely to have better personal health behaviors and practices (Andersen 1995; Kiefe et al.
Increasing access to health care through the ACA will connect more people to health care providers and services. These people will adopt healthier behaviors and in turn, will reinforce good practices and behaviors in their communities.

Better health behaviors in earlier adulthood also have lasting consequences in later life health and mortality. Adapting healthy behaviors such as refraining from smoking, engaging in physical activity, and having sound dietary habits lead to lower morbidity and mortality (Daneai et al. 2009). Just as an increase in insurance coverage among pregnant women had lasting mortality consequences on their children, an increase in insurance coverage among young, healthy adults may lead to a healthier older population.
REFERENCES


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Wang, E., Choe, M., Meara, J., & Koempel, J. (2004). Inequality of access to surgical specialty health care: why children with government-funded insurance have less
access than those with private insurance in Southern California. Pediatrics, 114, 584–590.

Fig. 2.1  Infant mortality rates in the United States, 1959-1979

Notes: Calculations are limited to the white population. Scatter plot represents fifty states excluding the District of Columbia. National Infant Mortality Rate weighted by the number of births in each state.
Fig. 2.2 Annual improvements in infant mortality associated with Medicaid implementation by State

Notes: Calculations are limited to the white population. Scatter plot represents fifty states excluding the District of Columbia. Data point sizes represent the number of births in 1965.
Fig 2.3 Logged ratio of infant mortality rate surrounding Medicaid start

Notes: Calculations are limited to the white population. Calculations are weighted by the number of births. Values are average logged ratio of states’ infant mortality rates relative to the year immediately prior to Medicaid implementation.
Table 2.1  Annual change in state infant mortality rate regressed on timing of Medicaid participation

<table>
<thead>
<tr>
<th></th>
<th>Coefficient</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Number of years since Medicaid implementation in state at birth(^a)</td>
<td>-0.326</td>
<td>***</td>
</tr>
<tr>
<td>Birth year (^b)</td>
<td>-0.382</td>
<td>***</td>
</tr>
<tr>
<td>Constant</td>
<td>23.22</td>
<td></td>
</tr>
<tr>
<td>R-Squared</td>
<td>0.98</td>
<td></td>
</tr>
</tbody>
</table>

Infant mortality trends before and after Medicaid implementation estimated from coefficients

- Annual rate of IMR decline in years before Medicaid: \(-0.382\)
- Annual rate of IMR decline in years after Medicaid \(^c\): \(-0.708\)

Notes: IMR units are in infant deaths per 1000 live births. Regression model includes state fixed-effects and weighted by the number of births in given year. Data is limited to the US white population. Standard errors are corrected for heteroskedastic panels.

\(^a\) Years prior to Medicaid implementation is set to zero

\(^b\) Represented as years since 1959

\(^c\) Calculated as the sum of coefficients, \(a\) and \(b\)

\(*p < 0.5, **p < 0.01, ***p<0.001\)
Fig. 2.4  Expected US infant mortality rates with and without Medicaid

Notes: Estimates are based on coefficients from Table 2.1. Data is limited to the US white Population. Values are weighted by the number of births in each state.
Fig. 2.5  Cumulative mortality probabilities by ten-year age group and birth year

Notes: Data is limited to the US white population. Mortality probabilities are cumulative probabilities of dying for each birth cohort living through each ten-year age period.
Table 2.2  Annual change in cumulative mortality probabilities regressed on timing of Medicaid participation (%)

<table>
<thead>
<tr>
<th></th>
<th>Age 1-9</th>
<th>Age 10-19</th>
<th>Age 20-29</th>
<th>Age 30-39</th>
</tr>
</thead>
<tbody>
<tr>
<td>Number of years since Medicaid implementation in state at birth (^a)</td>
<td>-0.017</td>
<td>-0.006</td>
<td>-0.034</td>
<td>-0.021</td>
</tr>
<tr>
<td>Birth year (^b)</td>
<td>-0.005</td>
<td>-0.011</td>
<td>-0.010</td>
<td>-0.023</td>
</tr>
<tr>
<td>Constant</td>
<td>0.561</td>
<td>0.718</td>
<td>1.215</td>
<td>1.622</td>
</tr>
<tr>
<td>(R^2)</td>
<td>0.87</td>
<td>0.87</td>
<td>0.85</td>
<td>0.70</td>
</tr>
</tbody>
</table>

Mortality trends for cohorts born before and after Medicaid implementation

<table>
<thead>
<tr>
<th></th>
<th>Age 1-9</th>
<th>Age 10-19</th>
<th>Age 20-29</th>
<th>Age 30-39</th>
</tr>
</thead>
<tbody>
<tr>
<td>Annual rate of mortality probability decline for cohorts born before Medicaid</td>
<td>-0.005</td>
<td>-0.011</td>
<td>-0.010</td>
<td>-0.023</td>
</tr>
<tr>
<td>Annual rate of mortality probability decline for cohorts born after Medicaid (^c)</td>
<td>-0.022</td>
<td>-0.017</td>
<td>-0.043</td>
<td>-0.044</td>
</tr>
</tbody>
</table>

Notes: All coefficients are significant at the alpha 0.001 level. Mortality probabilities are cumulative probabilities of dying within each ten-year age group. Regression model includes state fixed-effects and is weighted by the number of births in given year. Data is limited to the US white population. Standard errors are corrected for heteroskedastic panels.

\(^a\) Years prior to Medicaid implementation is set to zero

\(^b\) Represented as years since 1959

\(^c\) Calculated as the sum of coefficients, \(a\) and \(b\)
Fig. 2.6  Expected cumulative mortality probabilities with and without Medicaid by age group

Notes: Analysis is limited to US white population. Mortality probabilities are cumulative probabilities of dying for each birth cohort living through each ten-year age period. Estimates are based on coefficients presented in Table 2.2.
Table 2.3  Predicted population with and without Medicaid, 1970 birth cohort

<table>
<thead>
<tr>
<th>x</th>
<th>Without Medicaid</th>
<th>With Medicaid</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>nqx (%)</td>
<td>ndx</td>
</tr>
<tr>
<td>0-1</td>
<td>2.06</td>
<td>63,488</td>
</tr>
<tr>
<td>1-9</td>
<td>0.50</td>
<td>15,159</td>
</tr>
<tr>
<td>10-19</td>
<td>0.60</td>
<td>18,022</td>
</tr>
<tr>
<td>20-29</td>
<td>1.11</td>
<td>33,171</td>
</tr>
<tr>
<td>30-39</td>
<td>1.37</td>
<td>40,454</td>
</tr>
<tr>
<td>40</td>
<td>1.37</td>
<td></td>
</tr>
</tbody>
</table>

Excess under 40 mortality in 1970 cohort without Medicaid (2,918,917 - 2,932,598) 13,681

Notes: Data is limited to the US white population. Life tables use actual number of births in 1970 as the radix. nqx is based on mortality probabilities estimated in Tables 2.1 and 2.2.
<table>
<thead>
<tr>
<th>Implementation Month and Year</th>
<th>State</th>
</tr>
</thead>
<tbody>
<tr>
<td>January 1966</td>
<td>Hawaii, Illinois, Minnesota, North Dakota, Oklahoma, Pennsylvania</td>
</tr>
<tr>
<td>March 1966</td>
<td>California</td>
</tr>
<tr>
<td>May 1966</td>
<td>New York</td>
</tr>
<tr>
<td>July 1966</td>
<td>Connecticut, Idaho, Kentucky, Louisiana, Maine, Maryland, Nebraska, Ohio, Rhode Island, Utah, Vermont, Washington, West Virginia, Wisconsin</td>
</tr>
<tr>
<td>September 1966</td>
<td>Massachusetts</td>
</tr>
<tr>
<td>October 1966</td>
<td>Delaware, Michigan</td>
</tr>
<tr>
<td>December 1966</td>
<td>New Mexico</td>
</tr>
<tr>
<td>July 1967</td>
<td>Iowa, Kansas, Montana, Nevada, New Hampshire, Oregon, Wyoming</td>
</tr>
<tr>
<td>September 1967</td>
<td>Texas</td>
</tr>
<tr>
<td>October 1967</td>
<td>Georgia, Missouri, South Dakota</td>
</tr>
<tr>
<td>July 1968</td>
<td>South Carolina</td>
</tr>
<tr>
<td>January 1969</td>
<td>Colorado, Tennessee</td>
</tr>
<tr>
<td>July 1969</td>
<td>Virginia</td>
</tr>
<tr>
<td>January 1970</td>
<td>Alabama, Arkansas, Florida, Indiana, Mississippi, New Jersey, North Carolina</td>
</tr>
<tr>
<td>July 1972</td>
<td>Alaska</td>
</tr>
<tr>
<td>October 1982</td>
<td>Arizona</td>
</tr>
</tbody>
</table>

Source: Gruber (2003)
Notes: States joined the Medicaid program between January 1966 and October 1982. 48 out of 50 states implemented Medicaid prior to 1971. Light shades indicate early adoption and dark shares indicate late adoption. Source: Gruber (2003)
Chapter 3: Racial and Ethnic Disparities in Health Insurance Coverage:
Dynamics of gaining and losing coverage over the life-course

ABSTRACT

Health insurance coverage varies substantially between racial and ethnic groups in the United States. Current health care reforms attempt to reduce this gap by offering government subsidies and expanding Medicaid to eligible persons. By centering the discourse on insuring the uninsured, policymakers and scholars neglect to consider the dynamic nature of gaining and losing insurance. I simulate the expected number of years without health insurance coverage before becoming eligible for Medicare at 65 for non-Hispanic whites, African Americans, Hispanics, and Asians. I derive age-specific rates of insurance gain and loss from the 2008 Panel of the Income and Program Participation (N= 114,345) to construct increment-decrement life tables. The expected number of years to live uninsured for whites is 8 years. In comparison, African Americans are expected to live 13 years, Hispanics 22 years, and Asians 11 years uninsured before reaching Medicare eligibility. I decompose this racial and ethnic disparity in expected coverage and find that the disparity is largely driven by minority groups’ greater propensity to lose the health insurance that they already have. Increasing insurance-security to non-Hispanic white levels among minority groups will reduce the number of years uninsured by about 50% for African Americans and Hispanics, and by 23% for Asians.
Racial and ethnic disparities in health insurance coverage rates account for a sizable share of the difference in access to health care (Lillie-Blanton and Hoffman 2005). African American and Hispanic individuals in the United States are more likely to be uninsured at any given moment than non-Hispanic individuals (Kirby and Kaneda 2010). Without insurance, people face considerable barriers in receiving health services. The uninsured have a limited source of care as many health care providers require insurance coverage from their patients and face prohibitively high costs when they do receive care (Himmelstein et al. 2005; Institute of Medicine 2002; Kasper et al. 2000; Nelson et al. 1999; Zuekas & Weinick 1999). Inconsistent or unstable insurance coverage also have negative consequences on the health care. Patients who frequently change health care providers due to insurance loss or change experience more interruptions in their care and are less likely to establish ongoing relationships with their physicians.

Efforts to decrease health disparities between racial and ethnic groups must identify and reduce factors that cause African Americans, Hispanics, and Asians to have greater uninsurance rates relative to non-Hispanic whites. Prior literature has identified socioeconomic characteristics—income, employment, citizenship, and language—associated with uninsurance that are more prevalent in minority populations. All these factors are presented as barriers to acquiring health insurance. Few studies acknowledge that high uninsurance rates can occur in populations due to the greater frequency in which people lose insurance. Even fewer studies, if any, account for how the changing dynamics of gaining and losing insurance across the life-course contributes to overall racial and ethnic disparities in insurance coverage rates.
This paper simulates insurance coverage for African Americans, Hispanics, and Asians from birth till age 65 and examines how the differences in rates of loss and gain at various ages contribute to their lower insurance coverage relative to non-Hispanic whites. Specifically, I address three research questions.

1. How much of between-race/ethnic disparity is explained by differences in rates of insurance gain? How much of it is explained insurance loss?

2. How do the dynamics of losing and gaining insurance contribute to the disparity across the life-course?

3. What is the expected number of years to live uninsured among minorities if they had the same rates of insurance gain or loss of non-Hispanic whites?

The findings shed light on how much recent health reforms will reduce coverage gaps between minority groups and the white population. I decompose the overall racial and ethnic disparity into insurance gain and insurance loss controlling for mortality rates to determine the impact of policies targeted towards making insurance more accessible to disadvantaged populations.

Overall, about 19 percent of the non-elderly US population is uninsured (Clemens-Cope et al. 2012) but the prevalence of uninsurance differs substantially by race or ethnic group. About twenty-percent of African Americans are uninsured. In comparison non-Hispanic whites have an uninsurance rate of about thirteen percent (KFF 2013). About 18 percent of Asians are not insured. Hispanics have the higher prevalence of uninsurance; about a third of Hispanics living in the United States are without health insurance. Researchers cite low income and propensity to work in jobs with no health insurance benefits as the primary cause for high uninsurance rates among African
Americans (Institute of Medicine 2003). Studies say their low-income jobs pay too much to qualify for public assistance but pay too little to be able to afford private insurance policies leaving individuals and families to live without coverage (Edin and Kefalas 2011). Lack of job-based insurance is also a reason why Hispanics have high uninsurance rates. In addition, language barriers and immigration rules that prevent undocumented and recent immigrants from enrolling in public plans prevent Hispanics from being getting insurance (DeNavas-Walt et al. 2013; Goldman, Smith, and Sood 2005). Low take-up of public insurance has been cited along with employment in jobs without health benefits as the cause of high uninsurance rates among Asians (Institute of Medicine 2003). The Affordable Care Act (ACA) and other health policy measures are taking step to address these issues. The ACA is offering government subsidies to help lower-income working families without employer benefits afford private insurance plans. Medicaid aims to expand eligibility beyond children and the medically needy to reduce uninsurance rates among low-income, healthy adults. Outreach in multiple languages aims to lower linguistic barriers to enrolling in both public and private insurance among Hispanics and Asians.

The focus is clearly on insuring the uninsured. Cross-sectional evaluations of the policy’s impact, while informative, overlook the beneficiaries’ potentially greater risk of losing their new insurance coverage. Studies expect the ACA to increase enrollment among racial and ethnic minorities (Clemans-Cope et al. 2012; Holahan and McGrath 2013) but this paper shows increasing enrollment does not necessarily translate into proportionately higher insurance coverage rates.
THEORY AND LITERATURE

Racial and ethnic disparities in insurance coverage rates result from differences in their tendencies to lose or gain health insurance. In this section, I first briefly discuss the necessity of examining health insurance status as a dynamic process. Second, I review the literature on how the rates of gain or loss contribute to racial and ethnic disparities in coverage rates. Third, I discuss the importance of comparing age-specific patterns of insurance transitions to understand and to reduce coverage disparities.

*Dynamic Nature of Insurance Coverage*

I examine insurance coverage rates as a function of the population’s rates of insurance gain and loss. This methodology of examining the uninsured stems from research on the persistence of spells of poverty and unemployment (Bane and Ellewood 1985; Corcoran et al. 1985). Similarly to poverty and unemployment, a person’s insurance status or a change in status is not permanent. Point-in-time estimates of the uninsured are also over-represented by the proportion that has been uninsured for a long time and masks the heterogeneity of the group (Swartz and McBride 1990; Swartz, Marcotte, and McBride 1993; Monheit and Schur 1988). Examining insurance status as a dynamic process can capture how frequently groups lose insurance and how long it takes to gain insurance coverage again. This approach allows us to differentiate people who are uninsured because they are more likely to lose insurance from people who are uninsured because they are less likely to find insurance.

When researchers began to examine poverty as a dynamic process, scholarly understanding of who experience poverty changed. The perception of the “underclass” popularized by poverty debates in the 1960s (Harrington 1962; Willis 1977) gave way to
new research in the 1970s and 1980s that showed the dynamic and heterogeneous nature of falling into and out of poverty. People from various socioeconomic backgrounds experienced poverty often coinciding with life events such as the birth of a child, starting a new household, job loss, and divorce (Edwards 2014; Corcoran 1995).

Researchers have been applying these methods on longitudinal data to examine the dynamics of health insurance coverage (Swartz and McBride 1990; Fairlie and London 2009). The vast majority of people live without insurance in short spells; only a small fraction of uninsured had been living without insurance for more than two years (Swartz and McBride 1990; Congressional Budget Office 2003). The literature on unemployment is beginning to evolve from cross-sectional examinations of the uninsured to studying the dynamics of insurance.

Young adults, individuals with less education, the unemployed, and the unmarried have higher rates of losing health insurance. Trigger events such as losing employment, changing jobs, losing a spouse are also connected to insurance loss (Lavelle and Smock 2012; Peters, Simon, and Taber 2014). Once an individual loses health insurance the person’s demographic and socioeconomic characteristics also determine how quickly they will regain coverage. Individuals with higher income, full-time employment, and greater educational attainment have higher rates of gaining insurance which result in shorter spells without insurance (Swartz, Marcotte, and McBride 1993).

Health insurance policies in the United States attempt to provide a safety net for individuals who are experiencing life events that may trigger insurance loss. COBRA allows individuals to temporarily maintain coverage from a private insurance plan after
divorce or job loss. The Affordable Care Act aims to make health insurance more accessible through exchanges and expanding Medicare eligibility.

**Racial and Ethnic Disparity in Uninsurance**

The African American and Hispanic population has a greater prevalence of trigger events and socioeconomic characteristics that are associated with greater insurance loss and slower insurance gain. Access to private health insurance coverage is tied to employment and marriage in the United States. Minority groups are disadvantaged in both areas. Rates of unemployment are higher among African American men and women than their non-Hispanic white counterparts and job loss is more prevalent in among minority groups (Bureau of Labor Statistics 2014). African American and Hispanic individuals are less likely to marry than non-Hispanic whites. Among those who did marry, their first marriages are more likely to end in divorce and the proportion remarrying is lower than non-Hispanic white men and women (Aughinbaugh, Robles, and Sun 2013; Bulanda and Brown 2007).

Insurance policies that aim to provide safety nets during events associated with insurance loss (COBRA) and needs-based public insurance options that makes insurance more accessible (Medicaid) do not completely mitigate the insurance consequences of socioeconomic differences.

Furthermore, Fairlie and London (2009) find that insurance coverage disparities cannot be completely explained by compositional differences in educational attainment, income, and employment. Prior studies show that racial or ethnic background did not exert a significant impact on the rate of exiting an uninsurance spell (Swartz, Marcotte,
and McBride 1993). Rather, African Americans and Hispanics’ had greater rates of insurance loss than non-Hispanic whites (Fairlie and London 2008).

**Age-Dependent Nature of Insurance Coverage**

Health insurance coverage varies distinctly by age. Empirically, the insurance coverage rate is below 10 percent for children under 18. The rate of uninsurance increases to around 20 percent between 18 and 24 and reaches its peak in early adulthood between 25 and 35. The uninsurance rate decreases in later adulthood but does not reach under-18 levels until age 65 when the vast majority of US residents become eligible for Medicare (Cohen and Martinez 2014).

Age-dependent demographic, economic, and policy factors lead to this age-pattern of insurance coverage. Children under 18 can be eligible for needs-based insurance coverage through state-sponsored programs such as Medicaid and Children’s Health Insurance Program (CHIP). Due to state-level eligibility rules, younger children are more likely to have access to state-sponsored health insurance coverage than older children. Children age out of public health insurance plans after turning 19 and can no longer be enrolled in their parents’ private health plans after turning 26. Eligible pregnant women and mothers of young children (generally affects people aged between 15 and 45) can enroll in Medicaid. Getting married and having children is also associated with great insurance gain among the general population (Fairlie and London 2008). Full-time employees are more likely to gain and maintain health insurance coverage (Fairlie and London 2008) and the proportion of the population with full-time employment steady increases throughout adulthood till retirement at 65 (Bureau of Labor Statistics 2015).
From these age-dependent factors, I expect the rates of insurance gain to be highest among children under 18 and the rates of insurance loss to peak in early adulthood.

DATA

I use the 2008 Panel of the Survey of Income Program Participation (SIPP) to derive the age-specific rates of losing and gaining health insurance that served as the basis of my life-table calculations. The SIPP is a nationally representative series of longitudinal panels whose survey duration ranges from 2.5 to 4 years. The first SIPP panel was sampled in the early 1980s and a new panel was re-sampled from the non-institutionalized population in the U.S. every one to four years. The SIPP revisits respondent every four months and collects information on their insurance status for the preceding four months. Each four-month period is known as a wave. SIPP’s 2008 panel collected 14 waves for available respondents (non-institutionalized, US residents) covering information across 56 months throughout 2008 to 2012.

Heaping is a known problem in the SIPP and respondents are biased towards reporting changes to their insurance status at the beginning of each wave rather than at the actual month that the change occurred. While monthly insurance status is available in the SIPP, I have chosen to consider only the first reference month of each wave to evaluate respondents’ insurance statuses and record changes. This method makes the assumption that changes in insurance status can only happen up to once in a four-month period. All calculations are weighted by SIPP’s person-level weights that account for sampling and attrition.
ANALYTIC STRATEGY

I utilize insurance data from the 2008 SIPP to calculate the rates of gaining and losing insurance as well as rates of mortality. I derive these transition rates separately for each race or ethnic group to compare their overall rates of insurance loss and gain. Rates of insurance gain or loss depend on age. Thus, I calculate age-specific rates of insurance transitions by race and create two-state increment-decrement life tables for each group to describe the dynamics of living with and without insurance. I then compare the differences in the proportion uninsured between race/ethnic groups by age. Lastly, I decompose this difference across the life course to determine how much of the racial and ethnic disparity can be explained by the differences in the rates of insurance gain or loss. All analyses are limited to persons under 65 years of age. I describe this process in more detail.

Calculating Transition Rates

To address the first research question, I calculate the rates of losing and gaining insurance for each racial or ethnic group. The pattern of gaining and losing insurance resembles a Poisson distribution (Swartz and McBride 1990). Using this property, I derive the probability of losing or gaining insurance within a year of being insured or uninsured. I took all persons whose insurance statuses were recorded in two consecutive waves and calculated the proportion of insured in the former wave that was uninsured in the latter wave. I repeat a similar calculation to derive the proportion of uninsured who gained insurance in the later wave. From these proportions, I converted them into annual rates using the assumption that these transitions occur in a Poisson process with a constant rate.
where \( i \) is the initial insurance status and \( j \) is the insurance status after 4 months for group \( r \). \( d \) denotes the number of people who transitioned from state \( i \) to \( j \) and \( p \) denotes the number of persons in insurance state \( i \) at the beginning of the four-month period.

Using these rates I convert them into annual transition probabilities for each group.

\[
p_{rij} = 1 - \exp\left(-m_{rij}\right),
\]

where \( i \) is the initial insurance status and \( j \) is the insurance status after 4 months group \( r \).

These numbers represent the probability of losing or gaining insurance within one year of being insured or uninsured. We can compare these rates between groups and examine how fast one group loses or gains health insurance relative to another.

*Creating Multi-state Increment-Decrement Models*

These models describe how each racial and ethnic group transition between being insured and uninsured throughout the life course taking into account differential rates of gaining and losing insurance by age. They also include differential mortality by age and insurance status.

First I derive the age-specific insurance transition and mortality probabilities by group from the 2008 SIPP. I took all persons whose insurance or death status were recorded in two consecutive waves and calculated the proportion who transitioned into another state since the prior wave. I calculated these proportions separately by age at the beginning of the prior wave. I derived the proportion that lost/gained insurance among
those who had/did not have insurance at the beginning of the prior wave. In a similar fashion, I calculate the proportion that gained insurance and the proportion that died from each insurance state. Again using the properties of the Poisson distribution, I convert these 4-month transitions probabilities into annual rates.

\[ n m_{x, r, i, j} = \ln \left( 1 - \frac{\binom{n d_{x, r, i, j}}{p_{x, r, i}}}{\binom{4}{12}} \right), \]

where \( i \) is the originating state (insured or uninsured) at age \( x \) and \( j \) is the transition state (insured, uninsured, or dead) after 4 months, for group \( r \). Using these rates, I calculate transition probabilities for each age.

\[ n p_{x, r, i, j} = 1 - \exp\left( -n m_{x, r, i, j} \cdot n \right), \]

where \( i \) is the originating state (insured or uninsured) at age \( x \) and \( j \) is the transition state (insured, uninsured, or dead) at age \( x+1 \), for group \( r \).

These transition probabilities serves as the basis for the multi-state increment-decrement life table that I created employing the methodology described in Schoen (1975). This framework accounts for age-specific different transition forces between states to estimate the number of person insured and uninsured across the life-course.

Comparing Standardized Life-Tables between Groups

By comparing the proportion uninsured at each age, we can determine the disparity in insurance rates between racial and ethnic groups by age. Because these life tables are solely derived from five factors—initial proportion without insurance at birth, age-specific rates of losing insurance, age-specific rates of gaining insurance, age-specific rates of mortality for uninsured, and age-specific mortality for insured—we can compare
the effects of the difference in one of the factors while standardizing the remaining four. Taking the difference between equivalently standardized life-tables from two racial or ethnic groups will display the age-specific disparity attributable to one of the five factors. In this analysis, I specifically examine the life-course disparity in uninsurance prevalence due to difference in age-specific rates of gaining insurance and difference in age-specific rates of losing insurance. I standardize all other factors using the methodology described in Gupta (1993).

**Decomposing the Disparity across the Life Course**

The last component of this analysis calculates the proportion of the overall racial and ethnic disparity that is caused by each of the five factors. I adapt the decomposition methodology described in Gupta (1993) and apply the standardized life-tables created in the previous section onto a common population distribution (US 2009 population). I then derive the difference in the proportion uninsured attributable to each of the five factors. These differences sum to the overall between-group difference. In addition, I examine the disparity in terms of expected number of years lived without insurance between birth and age 64. In a similar manner, I decompose the difference to determine how many more years each factor contributes to a group living without insurance relative to non-Hispanic whites.

**RESULTS**

I present the results in answer to this paper’s three research questions.

*How much of between-race/ethnic disparity is explained by difference rates of insurance gain? How much of it is explained insurance loss?*
Non-Hispanic whites have the smallest proportion (.12) living without health insurance among African Americans, Hispanics, and Asians. In concordance with prior research, the rate of uninsurance is very high among the Hispanic population (.35). A third of non-elderly Hispanic persons live without health insurance coverage. Table 3.1 presents these results. The disparity in the expected number of years to live without insurance under current conditions is also quite large. Whites are expected to live on average a little less than eight years without insurance before reaching 65. In comparison, African Americans are expected to live almost 13 years and Asian Americans, 10 years without health insurance. Hispanics are expected to live over 22 years without health insurance coverage before reaching 65.

The last two lines of Table 3.1 compare the dynamics of losing and gaining insurance between the groups. The disparity in the rates of losing insurance is large. Non-Hispanic whites have a probability of .12 of losing health insurance within one year. African Americans are twice as likely to lose insurance with a probability of .24. Hispanics have a greater probability still at .31. The probability of losing insurance is not as high for Asians at .15. In contrast, the disparity in the rates of gaining insurance between groups is not as high. In fact, African Americans are more likely to get insured within one year of losing insurance (.65) than non-Hispanic whites (.59). Hispanics are about 13 percent less likely than whites to gain insurance after one year of living without coverage.

A decomposition analysis of the race or ethnic difference in insurance coverage confirms that differences in the rates of loss accounts for much of the disparity. Table 3.2 shows how much of the overall disparity is caused by each of the five factors. Negative
numbers indicate that the minority group has an advantage over non-Hispanic whites for the corresponding factor. All groups are compared to the non-Hispanic white population.

African Americans’ advantageous rates of insurance gain relative to non-Hispanic whites are completely offset by their very high rates of loss. Their higher rates of insurance gain alone would yield a lower prevalence of uninsurance relative to whites but their significantly greater rates of loss, slightly higher rates of infants born without insurance, and mortality patterns results in a difference of .08. That is, after accounting for differences in population distribution, eight percent more African Americans live without health insurance than whites. Seventy-eight percent of the Hispanic-white disparity in insurance coverage is explained by their greater rates of coverage loss. Twenty-two percent of the disparity is caused by their lower rates in obtaining health insurance. The rates of insurance gain among Hispanics are the lowest relative to whites. Together, they contribute to a greater proportion of Hispanics being uninsured (22% more) than whites. Almost all of the Asian-white disparity—a difference of 4 percent—is caused by Asians’ greater probability of losing insurance (98%).

*How do the dynamics of losing and gaining insurance contribute to the disparity across the life-course?*

Figure 3.1 shows the proportion uninsured by race/ethnic group from birth to age 64. Children under 18 have lower rates of uninsurance relative to adults within their race/ethnic group. This reflects the availability of state-sponsored insurance options for lower-income children. The age-specific patterns and levels of insurance coverage differ substantially by group. The prevalence of uninsurance is the highest among Hispanics at
all ages. The difference is particularly high after the age of 30. African American and Asian children have similar coverage to non-Hispanic whites during childhood but they diverge in young adulthood. The 20s is a period of high uninsurance for all groups but it is particularly higher for minorities. African Americans’ coverage increases in later adulthood but remains lower than whites until everyone qualifies for Medicare at age 65. Asians’ coverage exceeds that of whites in their early 30s but steadily falls till 65. Much of the coverage disparity between Asians and non-Hispanic whites originate from greater uninsurance among middle-aged Asians.

Figure 3.2’s age-specific rates of insurance loss loosely mirrors the age-patterns of insurance coverages rates. Insurance loss spikes in the early 20s for all groups but Hispanics have the greatest loss rate at almost all ages. Insurance coverage is notably precarious for African American infants and young adults; their rates of loss are similar to that of Hispanics’ during these age groups. Asians have relatively higher rates of loss during young adulthood and after age 40 reflecting their age-patterns of uninsurance prevalence. Surprisingly, age-specific rates of insurance gain (Figure 3.2) are the highest among African Americans. This gain-advantage over whites is particularly prominent during childhood. Except for a brief period in early adulthood, African Americans have advantageous insurance gain rates relative to whites throughout adulthood. The converse is true for Hispanics. Their rates of insurance gain are lower than their white counterparts at all ages. Asians’ rates of insurance gain are equivalent to that of whites after early childhood.

How do these disparate age-specific rates of insurance gain and losses contribute to the age patterns of uninsurance? Figures 3.3 to 3.5 examines the age-specific disparity
in uninsurance rates of the three minority groups relative to the non-Hispanic population. In each of these graphs, I show three lines. The solid line is the actual difference in the proportion uninsured. Values above 0 indicate greater proportion of the minority group without insurance. The dashed line is the simulated difference in proportion due to differences in age-specific rates of insurance loss standardized for all other factors. This can be interpreted as the disparity in uninsurance prevalence had only the rates of loss been different. The dotted line is the simulated difference in proportion due to differences in age-specific rates of insurance gain.

The prevalence of uninsurance among African American is higher than whites throughout all ages. They are also more likely than whites to gain insurance at all ages with the exception of a brief period in their early 20s. The dotted line in Figure 3.3 shows that African Americans’ gain-advantage would yield lower rates of uninsurance than whites. However, the high rates of insurance loss among African Americans more than offset this gain-advantage. The dashed line in Figure 3.3 shows that African Americans would have had an even higher uninsurance rate without their advantageous rates of insurance gain.

The coverage disparity between Hispanics and whites are explained both by Hispanics’ lower rates of insurance gain and higher rates of insurance loss (Figure 3.4). Differences in rates of loss account for most of the coverage disparity in childhood but, Hispanics’ increasing difficulty in gaining insurance becomes a greater contributor to coverage disparity in adulthood.
Asians’ rates of insurance loss account for almost all their lower coverage rates relative to whites. Figure 3.5 shows that the simulated coverage disparity due to difference in insurance loss closely follows the age-patterns of actual coverage disparity.

What is the expected number of years to live uninsured among minorities if they had the same rates of insurance gain or loss of non-Hispanic whites?

Table 3.3 represents the findings in terms of number of years without insurance. These numbers indicate the number of years a person would expect to live without coverage before age 65 if he or she is exposed to the age-specific rates of insurance gain, insurance loss, and mortality observed from cross-sectional data. African Americans are expected to spend almost 5 more years without insurance throughout the life-course than non-Hispanic whites. If they had the same rates of insurance loss as whites, African Americans would spend 6.2 less years uninsured. From our previous graphs we saw that the rates of insurance gain among African Americans were higher than non-Hispanic whites. If the gain rates of African Americans’ were lowered to that of whites, the African Americans would spend 1.3 more years without coverage. Ensuring that insured Hispanics do not lose their coverage more than whites would reduce the expected number of years without insurance by half from 22 years to 11 years. Increasing the rates of insurance gain for uninsured Hispanics to white-levels would only decrease the expected number of years spent without insurance to about 19 years. Asians are expected to live about 2.5 years longer without insurance than whites. A small portion of that difference is also due to lower under-65 mortality rates among Asians relative to whites (Table 3.3).
Asians were to have the same level of insurance security as whites, their expected years without insurance would decrease to 8 years.

DISCUSSION

The results demonstrate that racial and ethnic disparities in overall health insurance coverage are predominantly driven by differences in rates of insurance loss. Across the life course, non-Hispanic whites are least likely to lose health insurance and thus, have the greatest coverage. Early adulthood appears to be a time when minority groups are doubly disadvantaged. They are more likely to lose insurance and take longer to find new coverage than non-Hispanic whites. Differences in employment and marital status may account for this effect. Young white adults are more likely to have an employer who offers health insurance and to be married. During early adulthood, not only are the rates of uninsurance the highest, but racial and ethnic inequalities are also the greatest.

The analyses in this paper do not distinguish the different forms of insurance: employment-based private, marriage-based private, Medicaid, or Medicare. Employment-based private insurance plans may be more stable than Medicaid. Economic, demographic, and social inequalities between racial/ethnic groups determine the types of health insurance that groups enroll in. The type of health insurance is a large determinant in the likelihood of losing coverage. The analyses of this paper calculate the inequalities in insurance gains and losses that result from these factors.

The SIPP does not include persons who are incarcerated. Almost ten percent of African American adult men under 40 were incarcerated at any given day in 2010 (Neal and Rick 2014). In comparison, less than three percent of non-Hispanic white men were incarcerated. The disproportionately high incarceration rate among African American
men may bias the results drawn from the SIPP. If these men were not incarcerated, the insurance coverage disparity would increase as the uninsurance rate would likely be high among those at risk for incarceration. More African American men without stable health insurance coverage would be included in the analysis pool increasing the already large coverage gap in early adulthood. The current results would be understating the racial inequality.

It is very likely that high rates of insurance gain and loss will have negative consequences on a person’s health care. After insurance loss, patients may need to stop ongoing care. And when they re-gain insurance coverage, they may need to seek new health care providers that accept the new plan. The frequent changes in sources of care prevent patients from developing an ongoing, established relationship with health care providers. Physicians have less knowledge of the medical history of new patients than established patients. Levels of trust between physicians and patients may also be low. These factors could contribute patients with unstable health insurance coverage to receive poorer care compared to their continuously insured counterparts even when insured. Unstable health insurance coverage may contribute to the empirically observed lower levels of physician trust among patients of minority racial and ethnic backgrounds (Blendon et al. 1995; Gamble 1993; Peterson 2002; Stepanikova et al. 2006).

Patients who have unstable insurance may make their medical decisions with the expectation of losing insurance coverage. They may have a preference for shorter-term solutions or treatments plans that requires less follow-up. Greater expectation of insurance loss by either the patient or the physician may contribute to biases in referrals to specialists and in receiving surgical procedures (Einbinder and Schulman 2000).
While establishing a direct connection between insurance instability and health care delivery is beyond the scope of this paper, the results draw attention to a potentially large and significant mechanism through which health inequality persists between racial and ethnic groups in the United States. Social and economic factors create unstable and precarious insurance coverage among minority groups compared to non-Hispanic whites. This greater insurance instability may translate into disparities in health care delivery and inequalities in health outcomes.
REFERENCES


Table 3.1. Differences in prevalence and dynamics of uninsurance between groups

<table>
<thead>
<tr>
<th></th>
<th>Non-Hispanic White</th>
<th>African American</th>
<th>Hispanic</th>
<th>Asian</th>
</tr>
</thead>
<tbody>
<tr>
<td>Proportion without Insurance</td>
<td>0.12</td>
<td>0.20</td>
<td>0.35</td>
<td>0.16</td>
</tr>
<tr>
<td>Expected Years without Insurance</td>
<td>7.97</td>
<td>12.81</td>
<td>22.25</td>
<td>10.41</td>
</tr>
<tr>
<td>Probability of losing insurance</td>
<td>0.12</td>
<td>0.24</td>
<td>0.31</td>
<td>0.15</td>
</tr>
<tr>
<td>Probability of gaining insurance</td>
<td>0.59</td>
<td>0.65</td>
<td>0.51</td>
<td>0.58</td>
</tr>
</tbody>
</table>

Notes

Proportions without insurance and expected years without insurance derived from persons under 65.

Values represent uninsurance prevalence of a hypothetical cohort exposed to the age-specific rates of insurance gain, loss, and mortality observed from cross-sectional data.

Data Source: 2008 SIPP
Table 3.2. Decomposition of disparity in proportion without health insurance prior to age 65

<table>
<thead>
<tr>
<th></th>
<th>African American</th>
<th></th>
<th>Hispanic v. non-Hispanic</th>
<th></th>
<th>Asian v. non-Hispanic</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>v. non-Hispanic</td>
<td>Hispanic white (%)</td>
<td></td>
<td></td>
<td>Hispanic white (%)</td>
<td></td>
</tr>
<tr>
<td>Disparity due to insurance coverage at birth</td>
<td>0.00</td>
<td>0.20</td>
<td>0.00</td>
<td></td>
<td>0.16</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>Disparity due to greater rate of insurance loss</td>
<td>0.10</td>
<td>125.81</td>
<td>0.17</td>
<td>77.67</td>
<td>0.04</td>
</tr>
<tr>
<td></td>
<td>Disparity due to lower rates of insurance gain</td>
<td>-0.02</td>
<td>-26.29</td>
<td>0.05</td>
<td>22.05</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>Disparity due to mortality among insured</td>
<td>0.00</td>
<td>0.11</td>
<td>0.00</td>
<td>0.04</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>Disparity due to mortality among uninsured</td>
<td>0.00</td>
<td>0.17</td>
<td>0.00</td>
<td>0.07</td>
<td>0.00</td>
</tr>
<tr>
<td>Overall Difference</td>
<td>0.08</td>
<td>100</td>
<td>0.22</td>
<td>100</td>
<td>0.04</td>
<td>100</td>
</tr>
</tbody>
</table>

Values are standardized using the US population distribution in 2009 (Source: CDC/NCHS, National Vital Statistics System.)
Fig. 3.1. Age-specific patterns of insurance coverage by race and ethnicity
Fig. 3.2. Age-specific patterns of insurance gain, insurance loss, and mortality by race and insurance coverage
Fig. 3.3. Difference in proportion without health insurance between African Americans and non-Hispanic whites

Notes

Values greater than 0 indicate greater proportion of African Americans without insurance relative to the non-Hispanic White population. Difference from loss/gains rates are derived by taking the difference in simulated proportion uninsured controlling for differences due to gain/loss, population distribution, mortality, and initial proportions.

Population distribution standardized to the 2009 US population.

Data Source: 2008 Survey of Income Program Participation
Fig. 3.4. Differences in proportion without health insurance between Hispanics and whites

Notes
Values greater than 0 indicate greater proportion of Hispanics without insurance relative to the non-Hispanic White population. Difference from loss/gains rates are derived by taking the difference in simulated proportion uninsured controlling for differences due to gain/loss, population distribution, mortality, and initial proportions. Population distribution standardized to the 2009 US population.

Data Source: 2008 Survey of Income Program Participation
Fig. 3.5. Differences in proportion without health insurance between Asians and whites

Notes

Values greater than 0 indicate greater proportion of Asians without insurance relative to the non-Hispanic White population. Difference from loss/ gains rates are derived by taking the difference in simulated proportion uninsured controlling for differences due to gain/loss, population distribution, mortality, and initial proportions. Population distribution standardized to the 2009 US population.

Data Source: 2008 Survey of Income Program Participation
### Table 3.3. Actual and simulated expected number of years to live without insurance from birth to age 65

<table>
<thead>
<tr>
<th></th>
<th>African American (%)</th>
<th>Hispanic (%)</th>
<th>Asian (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Actual expected years to live uninsured</strong></td>
<td>12.81</td>
<td>22.25</td>
<td>10.41</td>
</tr>
<tr>
<td>Excess uninsured years relative to non-Hispanic whites (a - 7.97 years)</td>
<td>4.84</td>
<td>14.29</td>
<td>2.45</td>
</tr>
<tr>
<td><strong>Years saved from reducing insurance loss to non-Hispanic white levels</strong></td>
<td>6.19</td>
<td>48.3</td>
<td>11.04</td>
</tr>
<tr>
<td>Simulated expected years uninsured with white insurance loss rates (a - b)</td>
<td>6.62</td>
<td>51.7</td>
<td>11.21</td>
</tr>
<tr>
<td><strong>Years saved from increasing insurance gain to non-Hispanic white levels</strong></td>
<td>-1.29</td>
<td>-10.1</td>
<td>3.12</td>
</tr>
<tr>
<td>Simulated expected years uninsured with white insurance gain rates (a - c)</td>
<td>14.10</td>
<td>110.1</td>
<td>19.13</td>
</tr>
</tbody>
</table>

### Notes

Values represent expected years without insurance for a hypothetical cohort exposed to the age-specific patterns of insurance gain, loss, and mortality observed from cross-sectional data. Expected number of years without insurance for non-Hispanic whites is 7.97 years. Negative values indicate that expected years uninsured will increase. Values are standardized using the US population distribution in 2009 (Source: CDC/NCHS, National Vital Statistics System.)

Data Source: SIPP 2008