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Marital Disruption and Health Insurance

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Abstract Despite the high levels of marital disruption in the United States and the fact that a significant portion of health insurance coverage for those less than age 65 is based on family membership, surprisingly little research is available on the consequences of marital disruption for the health insurance coverage of men, women, and children. We address this shortfall by examining patterns of coverage surrounding marital disruption for men, women, and children, further subset by educational level. Using the 1996, 2001, and 2004 panels of the Survey of Income and Program Participation (SIPP), we find large differences in health insurance coverage across marital status groups in the cross-section. In longitudinal analyses that focus on within-person change, we find small overall coverage changes but large changes in type of coverage following marital disruption. Both men and women show increases in private coverage in their own names, but offsetting decreases in dependent coverage tend to be larger. One surprising result is that dependent coverage for children also declines after marital dissolution, even though children are still likely to be eligible for that coverage. Children and (to a lesser extent) women show increases in public coverage around the time of divorce or separation. We also find that these patterns differ by education. The most vulnerable group appears to be lower-educated women with children because the increases in private,

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own-name, and public insurance are not large enough to offset the large decrease in dependent coverage. As the United States implements federal health reform, it is critical that we understand the ways in which life course events—specifically, marital disruption—shape the dynamic patterns of coverage.

Keywords Health insurance · Divorce · Separation · Children · Family economics

Introduction

In the United States, most individuals younger than age 65 are covered by employer-provided health insurance. Because a substantial proportion of individuals covered by employers are covered as dependents (where eligibility for adults depends on marital status), this form of insurance is heavily family-based and thus has the potential to expose those undergoing marital instability to coverage vulnerabilities. Our article investigates the impact of marital disruption on health insurance for women, men, and children, a topic of surprisingly little prior research. About 1 million divorces occur in the United States each year (National Center for Health Statistics 2013), leaving a large number of individuals at risk for health insurance change associated with marital disruption. Adults who are divorced or separated are less likely to be insured than married individuals (Berk and Taylor 1984; Willis and Weir 2002; Zimmer 2007), and children with divorced or separated parents are less likely to be insured than those with married parents (Monheit and Cunningham 1992; Weinick and Monheit 1999).

New provisions in the Affordable Care Act (ACA) will increase availability of health insurance, especially through Medicaid expansions and subsidized exchange-based private coverage, and may mitigate some of the detrimental impacts of marital disruption. However, because employers are expected to remain the main source of coverage for the nonelderly population (Congressional Budget Office 2012) and because of lingering uncertainty regarding state compliance with expansions (Kaiser Family Foundation 2013), marital disruption is likely to remain a cause of instability in insurance coverage. Thus, it is critical that we understand how life course events such as marital disruption shape the dynamic patterns of coverage.

In this article, we conduct a systematic longitudinal study of marital disruption and health insurance for nonelderly populations. Specifically, we examine (1) how coverage varies across individuals by current marital status; (2) how health insurance coverage changes after marital disruption for men, women, and children; (3) how types of health insurance coverage (own employer, dependent, or public) change after marital disruption; and (4) how these experiences differ by education. We use a method similar to that used in the literature studying the pattern of earnings changes before and after job loss (Couch and Placzek 2010; Jacobson et al. 1993) and the impact of divorce on earnings (Couch et al. 2013). Our findings in this area contribute to the literature on the consequences of family marital transitions and to the literature on access to health insurance.

Prior Literature

The Link Between Health Insurance and Marital Status

In 2010, 56.5 % of the U.S. population was covered by employer health insurance, of which 34.8 % was own-name and 21.7 % was dependent (Janicki 2013). Employer health insurance is the only coverage type that is explicitly family–based; eligibility for Medicaid or other public programs is determined on an individual basis. The family-based nature of employer health insurance has been used to shed light on topics such as joint health insurance and labor market decision-making between married partners (Abraham and Royalty 2005), the effect of spousal health insurance availability on married women's labor supply (Buchmueller and Valletta 1999), and on retirement behavior (Karoly and Rogowski 1994).

Previous research, which is largely descriptive, supports the hypothesis that insurance rates decrease after divorce, as we would expect from a family-based system in which those who were previously covered as dependents on their spouse's insurance are no longer eligible after divorce. Berk and Taylor (1984) found that in 1977, married women were one-half as likely to be uninsured compared with divorced women. Willis and Weir (2002) confirmed this result among the near-elderly, finding that divorced or never-married women were less likely to be insured than married women. Monheit and Cunningham (1992), Weinick and Monheit (1999), and Heck and Parker (2002) found that children in single-parent families were more likely to be uninsured than children with both parents present.

We are aware of only two publications that examined the relationship between changes in marital status and changes in health insurance. Zimmer (2007) found that women are 13 percentage points more likely to lose insurance after a divorce compared with women who remain married. In a more recent study, Lavelle and Smock (2012) found estimates that are somewhat smaller than those of Zimmer (2007), with the loss of health insurance being in the range of 5 to 6 points. More specifically, they examined women's changes in health insurance after divorce for the period 1996–2007 using multivariate fixed-effects models with data from the Survey of Income and Program Participation (SIPP). They found that approximately 115,000 American women lose private health insurance each year immediately after divorce and that slightly more than one-half become uninsured as a result. They also showed that baseline factors—including the source of coverage while married, employment status, education, job tenure, race, poverty status, age, presence of children, and health—moderate the loss of coverage after divorce.

Our article uses the same data source that Lavelle and Smock (2012) used, and one of our contributions to the literature is to examine this question for men and children as well as for women. The gender comparison is important: health insurance in the United States is largely employment-based, and employment experiences differ for men and women. Furthermore, no research to date has examined how type of coverage changes over time relative to marital disruption for men as compared with women, how patterns and differences across men and women depend on their level of education, or how children's coverage changes when their parents experience marital disruption. We also document the patterns of change in each type of coverage for the period 12 months before and after disruption. Finally, our analysis includes those who experience a

separation as well as those who divorce, and we find similar consequences among the two groups.

Conceptual Model of Health Insurance and Marital Disruption

For adults and children, the past literature shows that health insurance status depends on individual, family, and social policy characteristics, such as education, health, employment, marital status, family income, state Medicaid eligibility rules, and child support policies (e.g., Currie et al. 2008; Cutler and Lleras-Muney 2010; DeNavas-Walt et al. 2012; Kaestner and Kaushal 2005; Kapur et al. 2008; Pollack and Kronebusch 2005; Taber 2011).

In the most direct link between marital disruption and health insurance for adults, a former spouse will no longer be covered as a dependent on an employer health insurance policy. This results from a combination of tax laws that provide benefits to employees and their qualified dependents (which do not include former spouses) as well as from employer and insurer rules that limit coverage to defined family members (Sections 152 and 105 (b) of the Internal Revenue Code). Although a few states have passed laws requiring that dependent coverage be available to divorcing spouses, the effectiveness of such coverage is complicated by the federal Employee Retirement Income Security Act (ERISA) (McGorrian 2012). Generally, only temporary and expensive employer coverage through the Consolidated Omnibus Budget Reconciliation Act (COBRA) is available to former spouses of workers (Landers 2012; U.S. Department of Labor 2012).[1] Because 25 % of women are covered as dependents on employer health insurance policies compared with 13 % of men (Salganicoff 2008), the loss of family health insurance as the result of marital disruption is likely to affect women more than men.

The direct effect of marital disruption on health insurance may also depend on whether the disruption is a separation or a divorce. Spouses who are separated are often eligible for coverage as dependents until they are divorced. However, the acrimony that often accompanies marital disruption might lead a spouse with primary coverage to drop a dependent spouse from his or her policy. Alternatively, a spouse with dependent coverage might have access to employer coverage through his or her own job and may switch to that coverage upon separation. Children can still be legally covered as dependents after parents' divorce, regardless of whether they actually live with that parent, but such coverage may not be cost effective if a parent's employer plan is tied to a local provider network but his or her child(ren) live elsewhere.

Other mechanisms that link marital disruption and health insurance are indirect and more likely to apply to both divorce and separation. One obvious mechanism involves changes in employment. Women's employment increases after marital disruption (Couch et al. 2013; Johnson and Skinner 1986). Some women may find employment

[1] The 1986 Consolidated Omnibus Budget Reconciliation Act (COBRA) stipulates that former spouses are allowed to continue to purchase employer health insurance through the employer for 18 months following divorce in firms with 20 or more employees (provided that notification is given within 60 days of the divorce), but the individual is responsible for paying the full cost of the policy, plus a small administrative fee, that in total will not exceed 102 % of the employer cost. As a result, COBRA is unaffordable for many who are recently divorced or newly unemployed.

in marginal jobs that do not provide insurance; but, especially because the evidence shows that hours as well as labor force participation increase (Johnson and Skinner 1986), we expect that jobs held after a marital disruption are more likely to include health insurance as a benefit. This pathway may have the largest impact for women, but because children can be included as dependents on their mother's insurance, increased employment of mothers may help maintain health insurance coverage for children, too.

Another pathway that links changes in marital status and health insurance is the general decline in income that accompanies marital disruption, especially for women. Bianchi et al. (1999) found that the income-to-needs ratio for mothers falls by more than 25 % following marital disruption. Because men's own earnings are generally higher than women's earnings and because children live with their mothers in the vast majority of cases (82 %, according to Grall 2013), the standard of living for men falls much less than it does for women following marital disruption (Stevenson and Wolfers 2007). Thus, based on the link between low income and lack of health insurance, we expect to find a larger effect of marital disruption on health insurance coverage for women.[2] Declines in income may be mediated by public assistance programs such as Temporary Assistance for Needy Families (TANF) and the Earned Income Tax Credit (EITC), which may encourage work and increase the probability of holding employer coverage (Kaestner and Kaushal 2005). Child support payments and entry into the labor force may also reduce the drop in income associated with marital disruption.

The detrimental effect of a decline in family income on children's health insurance and, to a lesser extent, on the health insurance of low-income parents may also be mediated by eligibility for public health insurance through Medicaid and the State Children's Health Insurance Program (SCHIP). These programs, whose exact rules vary by state and over time, expanded substantially during the early years of our study period (Gruber 2008). We control for these expansions in our work using a measure of the generosity of a state's program in a particular year. (See Cutler and Gruber 1996 for more details on measuring the generosity of public health insurance.) However, as public coverage has become more available in recent years, children's coverage might have been more detrimentally affected by family disruption in an earlier period.

Having health insurance might also affect marital status changes, especially when ill health is involved. Similar to the "job-lock" phenomenon—whereby workers do not leave employers that provide them health insurance even though it may be otherwise preferable to do so (e.g., Madrian 1994)—perhaps some marriages survive because health insurance availability is tied to one spouse. In this article, we study only those patterns of health insurance that follow marital disruption; we do not study the extent to which the existence of health insurance may reduce divorce rates.

[2] The breakup of a cohabiting relationship is another type of household disruption that might lead to changes in health insurance coverage through this indirect mechanism, which involves the loss of income sharing and economies of scale. This issue is beyond the scope of this article, in part because of the difficulty in identifying the timing of dissolution among cohabiting couples. In addition, cohabiting couples are not likely to have access to dependent coverage from employers. In 1997, only 7 % of employers offered such coverage, although this figure rose to 19 % in 2002 and to 27 % by 2004 (Employee Benefits Research Institute 2004; Schaefer 2009). Furthermore, cohabitating couples report less income sharing, and both members of cohabitating couples are more likely to work than are both marital partners (Addo and Sassler 2010; Heimdal and Houseknecht 2003); cohabiting couples are thus more likely to have access to their own health insurance than married couples.

In summary, the fact that many individuals obtain health insurance through marriage to a worker with access to employer coverage implies a direct impact of marital disruption on presence and type of health insurance. Marital disruption may also have secondary impacts that have health insurance consequences: for example, labor market entry may lead to increases in private own-name coverage, while declines in income resulting from disruption may result in eligibility for public coverage, particularly for women and children. All these effects are likely to be mediated by educational status. Those with greater labor market opportunities are better able to provide coverage in their own name through a new or existing job, while those with less education are more likely to be eligible for public coverage. Factors that affect health insurance for children differ from those for adults because of differences in legal and institutional frameworks that lead employers to cover children regardless of custodial relationships, and because of the greater availability of public health insurance for children.

Hypotheses

The previous discussion leads us to propose three sets of testable hypotheses. First, we expect to see significant variation in presence and type of health insurance across marital statuses. Further, we expect that these cross-sectional differences will be larger than estimates from longitudinal data comparing health insurance before and after marital disruption because those who are divorced or separated are likely to differ from those who remain married in unobserved ways.

Second, we expect to find systematic differences in how men, women, and children are affected. In particular, we expect women's health insurance to suffer greater declines after marital disruption because we observe in our data (Fig. 1) that women are more likely to have been dependents under husband's policies and because of women's weaker labor market ties. For women, we also expect to see changes in type of coverage (from dependent to own-name employer coverage and from private to public coverage) that are larger than those for men. We expect to find differences according to whether women have children: women with children are less likely to work and receive coverage on their own, and they also have more opportunities to receive public coverage. We expect that children's coverage will not suffer as much, partly because of access to coverage from either parent following divorce and also because of public policy. Medicaid and SCHIP cover children to higher income levels than adults, and state laws require health insurance to be addressed in child support agreements in most states (Taber 2011).

Third, in addition to expecting differences between men, women with and without children, and children, we expect to see differences related to educational attainment. Those with lower levels of education will see greater declines in health insurance following marital disruption because of poorer labor market opportunities. Both higher- and lower-education groups are likely to lose dependent coverage around the time of a marital disruption, but those with more education are more likely to be able to obtain private coverage either through their original job or through a new job. Those with lower levels of education are more likely to see increases in public coverage, although this mainly impacts children and, to a lesser degree, women with children.

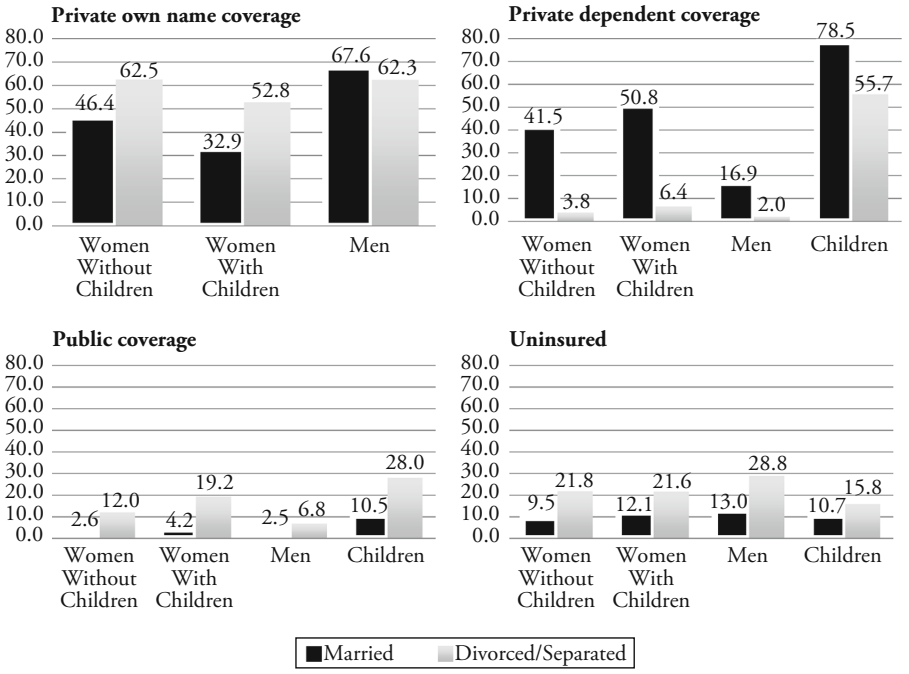

Fig. 1 Current health insurance by current marital status: 1996, 2001, and 2004 panels for adults ages 23–55 and children ages 0–15. All figures are weighted means from the first month and first wave of the 1996, 2001, and 2004 panels of the SIPP. $N = 22{,}200$ women without children; 29,382 women with children; 45,755 men; and 53,317 children. All the differences between the married and divorced/separated groups are statistically significant at the 1 % level

Data

We use data from the 1996, 2001, and 2004 panels of SIPP, which cover the years 1996 to 2007. Each SIPP panel interviews a nationally representative sample of households in the United States over a period of 2.3 to 4 years. For each panel, all individuals in the household age 15 and older are interviewed; if they are not available, a proxy response is obtained. The interviews take place every four months, and information is collected about certain variables for each of the four months in the wave. Original household members age 15 and older who move to a new address are followed as well. One advantage of using SIPP relative to other data sets is that it can follow individuals before and after a change in marital status with month-specific observations. A second advantage is that the SIPP collects information on marital status, labor market status, income, and program participation for all individuals in the household aged 15 and older; and it collects information on education, age, race/ethnicity, and health insurance coverage and type for all individuals, regardless of age.

The sample size for our cross-sectional analysis is 45,755 men and 51,582 women between the ages of 23 and 55 who are married, divorced, or separated (29,382 without children and 22,200 with children) as well as 53,317 children between the ages of 0 and

15 whose parents are married, divorced, or separated.[3] We limit our analysis to age 55 and younger to reduce the possibility of including those who have retired or are near retirement; we also limit our adult population to age 23 and older given that many individuals between the ages of 18 and 23 are in college and likely have different factors impacting health insurance availability and choices.

For our longitudinal analysis, in which we examine how health insurance changes as marital disruption occurs, we use the same age restrictions as noted earlier and further limit our sample to all individuals (or children of these individuals) married during the first month of the panel who subsequently divorced or separated. We can observe data for 1,289 men, 1,709 women, and 1,831 children before and after marital disruption.

Based on these numbers, we observe fewer men than women both before and after divorce because attrition from SIPP for the divorced and separated is higher for men than for women. This difference is perhaps due to men's greater likelihood of moving to a new address after marital disruption; although the SIPP makes every attempt to follow those who move out of originally sampled households, it is not always successful. We investigate whether selective attrition may bias our findings, given that those who attrit may have different health insurance patterns than those who remain in the data. We find that men who are no longer in the data one month after their wife reports a divorce or separation are significantly more likely to have no college education, less likely to have private coverage at baseline, more likely to be uninsured at baseline, more likely to have family incomes below poverty, and less likely to have family incomes of more than 300 % of the federal poverty level (Table S1 in Online Resource 1). We also find some selective attrition for women, albeit to a lesser degree. Women who attrit are significantly less likely to have children.

Two articles that investigated the bias resulting from differential attrition in longitudinal data sets found minimal bias. Zabel (1998:479) found selective attrition in the SIPP but concluded, "(T)he estimation results for a model of attrition and labor market behavior show little indication of bias due to attrition." Lillard and Panis (1998:437) found that attrition in the Panel Study of Income Dynamics (PSID) was higher among those with a marital disruption, but they concluded that "the biases that are introduced [in estimating the effect of marital status on mortality risk] by ignoring selective attrition are very mild." Nevertheless, we are cognizant that selective attrition could lead us to underestimate the effect of marital disruption on health insurance loss if the most vulnerable leave the sample.

Marital Status Variables

In each wave, the SIPP asks about current marital status and about marital status for each of the previous three months, yielding monthly data on marital status. A complication to studying marital status dynamics is the inherent ambiguity of the survey question about current marital status for those whose marital status is in transition. The respondent's answer could refer to the legal marital status (e.g., those not divorced are still legally married) or to a perception about the situation (e.g., those whose divorce is

[3] For the cross-sectional analysis, which includes only the first month of each panel, age is measured in that month. For the remaining analysis, adults must be between the ages of 23 and 55 at the month of marital disruption. Children must be between the ages of 0 and 15 at the month of the parents' marital disruption.

pending may report that they are divorced). As a result, we also consider an alternative way of capturing a marital disruption that is based on two people sharing the same residence. Specifically, we know the initial address of each person in the household at the start of the panel and can construct an indicator for the month at which the husband and wife start to live apart.[4]

There are two important issues to consider when choosing which type of marital transitions to study. The first is whether to focus on divorce, separation, or both. A spouse who is separated but not divorced is legally entitled to be covered as a dependent on employer coverage. However, because of a number of transactional and emotional reasons, a change in coverage might occur during the period of separation prior to the date of legal divorce. Thus, we might expect to see important changes in health insurance resulting from separation as well as from divorce.

A second issue relates to measurement error regarding the date of separation. Because separation is a more fluid and ambiguous construct than divorce, it is more difficult to measure, and the meaning of this status may differ across individuals.[5] For example, some individuals may have a legal separation agreement and may report being separated only after that date, while others may report being separated when they stop living together. As evidence of the ambiguous reporting of separation, 30 % of the divorces observed in our longitudinal data sample transition directly from marriage to divorce, without an intervening period of separation. Although this is legally possible in a few states, most states require some waiting period or period of separation; and in practice, it takes time to legally end a marriage. Thus, many individuals who report transitioning directly from marriage to divorce have, in reality, some period of separation prior to legal divorce that they do not report in the data. For these cases, a measure of when a separated couple stopped living together would enable us to capture the period of separation.

Each measure of marital disruption (divorce, separation, and address change) captures valuable and unique information, so there are no clear correct definitions for our purposes. In the results that we present later, we do not examine separation independently from divorce for three additional data reasons. First, 35 % of individuals with a marital disruption are censored at separation, forming a large and diverse group of all those in the SIPP who experienced marital disruption. Some of these individuals experienced a separation near the end of the panel and are likely to have had short-term separations that will convert to divorce, while others may be long-term separations that never become legal divorces. Because of this within-group heterogeneity, it is not clear whether we would gain much insight from treating the separated group differently from the divorced group. Second, sample sizes become too small for reliable estimates in some subanalyses when we distinguish those who report being separated from those who divorce. Third, our earlier discussion suggests that measurement error for separation dates is likely to be much higher than for divorce dates. Thus, any difference in results about health insurance coverage before and after divorce compared with

[4] We limit our sample to those initially living at the same address who had reported that they were married and who then report being divorced or separated at some point during the survey.

[5] For example, there are differences in the timing of reporting of separation between husband and wife, with 21 % of our sample reporting that they became separated in a different month than the month reported by their spouse.

separation could be due to differences in the measurement of those two types of transitions.[6]

Because of the pros and cons of different ways of measuring marital disruption, we take a hybrid approach, using the event that occurred first: the first date of reported marital status change (divorce or separation) or the date of the address change. Most individuals report living at a different address than their spouse at the same time that they first report divorce or separation. We find that 97 % of adults live at different addresses the month of reported marital disruption when both spouses are still in the data. By contrast, only 12 % live at different addresses the month before a reported disruption.[7]

Health Insurance Variables

In the SIPP, each household member's health insurance status and type (e.g., Medicaid/SCHIP, Medicare, an employer-sponsored plan, or a nongroup plan) during each month is recorded from interviews conducted every four months. For those covered by an employer-sponsored plan (including both private and public employers), we know whether they are initially covered as dependents. We can observe how health insurance status changes in the months or years preceding or following a divorce or separation and for both members of the divorced or separated couple as well as for any children present in the household.

Education Variables

The SIPP provides detailed information on the highest grade completed or degree attained for adults. The lower-education group constitutes those who have at most a high school diploma or the equivalent. The higher-education group consists of those who have at least some college education. Children are categorized according to their mother's level of education.

Policy and State Contextual Variables

Our regressions control for state- and year-specific labor market variables that capture potential access to insurance through changes in employment after marital disruption. These include the unemployment rate from the U.S. Bureau of Labor Statistics and the average full cost of employer health insurance plans from the Agency for Health Care Research and Quality.[8] Cawley and Simon (2005) found that the impact of unemployment rates on health insurance differs among men, women, and children.

[6] In our empirical results discussed later, we show that there are only small differences in patterns of health insurance coverage before and after marital disruption for those who are only observed to be separated and for those who transition directly from marriage to divorce.

[7] In our analysis, a measure of marital disruption based solely on the first reported marital status change yields similar results to the hybrid measure we use in reported results, although the latter is more closely related to health insurance loss in some analyses.

[8] The average cost of employer health insurance plans can be found on the Medical Expenditure Panel Survey website (http://meps.ahrq.gov/mepsweb/data_stats/quick_tables_search.jsp?component=2&subcomponent=2).

We also include three measures of public programs that affect income, the probability of being on welfare, and incentives to work. As a proxy for the generosity of welfare programs and incentives to be on welfare, we include the maximum state- and year-specific TANF guarantee for a family of four. We also include a variable measuring the phase-in rate for the EITC, which varies by year, state, and whether there are one or more children in the family. A higher phase-in rate is likely to encourage employment by low-income individuals, which will affect their health insurance opportunity set. In addition, we include a measure of SCHIP eligibility, which varies at the state and year levels, in our children's regressions. We do so using an eligibility calculator that assesses the fraction of children from a representative national population eligible for coverage in a certain state and year. This produces an index that measures policy generosity toward Medicaid, which has been used in prior Medicaid/SCHIP research (e.g., Gruber and Simon 2008).

Methods and Results

We investigate the relationship between marital status and health insurance using three methods: a cross-sectional analysis, a longitudinal descriptive analysis, and fixed-effects multivariate regressions subset by educational attainment. Our cross-sectional analysis shows differences in health insurance between those with different marital statuses at a point in time, whereas our longitudinal analysis examines within-person changes in health insurance around the time of marital disruption. Our regression analysis controls further for time-varying policy and economic variables and examines how higher- and lower-education groups differ in health insurance trends around the time of marital disruption. All our analyses are conducted separately for four demographic groups: women without children, women with children, men, and children.

Cross-Sectional Analysis

As a baseline, and for comparison with the few previous studies on health insurance and marital status, we first tabulate health insurance coverage by marital status (or parent's marital status) separately for our four demographic groups. We use data from the first month of each SIPP panel (1996, 2001, and 2004), pooling data from all three panels together for the tabulations.

The cross-sectional results in Fig. 1 show large differences in insurance coverage rates and types of insurance by marital status and subgroup. Rates of uninsurance are much higher for those who are separated or divorced than for those who are married (Fig. 1, bottom right panel). The difference is largest for men, who are 16 percentage points more likely to be uninsured if they are divorced or separated compared with those who are married; the difference is smallest for children, who have a gap of only 5 percentage points (both of these differences in differences are statistically significant). Women with and without children are 20 and 16 percentage points more likely, respectively, to have private own-name coverage if they are divorced or separated than if they are married (Fig. 1, top left panel). However, divorced or separated men are 5 percentage points less likely to have coverage in their own name compared with men who are married. Men, women with and without children, and children are all much

less likely to have private coverage as a dependent on a family member's plan if they or their parents are divorced or separated (Fig. 1, top right panel). The differences range from 15 percentage points for men to 44 percentage points for women with children. All four groups are also more likely to have public coverage if they are no longer married (Fig. 1, lower-left panel). These differences range from 4 percentage points for men to 18 percentage points for children. Although these cross-sectional differences in health insurance are interesting, we know that the married differ from those who are divorced or separated in many ways. Thus, the results in Fig. 1 do not tell us how health insurance status changes as a result of marital disruption. For example, those who are married may have more advantageous characteristics in observed and unobserved ways (Goldman 1993). Additionally, people who are currently divorced or separated will have had that status for different durations, producing yet another type of heterogeneity. To answer our questions about how health insurance changes after marital disruption, we turn to estimates that take advantage of the SIPP's rich longitudinal data.

Longitudinal Analysis

Our second method follows individuals and their children longitudinally and documents health insurance changes before and after marital disruption. For the analysis of the transition from marriage to separation or divorce, we begin by identifying the sample of men and women who report being married in the first month of each panel and who subsequently report being divorced or separated before the end of the panel; we also identify their children. We focus on the first marital transition event for a given individual, setting as the "zero date" the month in which they either first reported being divorced or separated or the month in which they first ceased sharing an address with their spouse, depending on which happened first. Note that because the zero date can occur at any time during our SIPP panels (which span approximately two to four years), the number of individuals we observe in our data set is greatest at the zero date and smallest when considering dates furthest from the marital disruption date (either before or after that date). Table S2 in Online Resource 1 reports sample sizes by month relative to the zero date and shows that sample sizes at plus or minus 12 months are 29 % to 44 % smaller than those in the month of disruption. Because sample sizes are largest closest to the zero date, we concentrate our analysis on the 12 months before and after disruption.

Figure 2 shows the changes in coverage rates experienced at various points in time relative to the date of marital disruption (the zero date) for the four groups experiencing marital disruption as well as for those who remain married. We find that the declines in coverage 12 months before and after a marital disruption are substantially smaller than the cross-sectional differences in coverage by marital status, suggesting that many of the differences in the cross-sectional analysis are due to selection. Specifically, over this two-year time span, the data show a 6.4 percentage point drop for women with children, a 4.9 percentage point drop for men, and a 6.6 percentage point drop for children; these differences are statistically significant. Women without children experience a short-term decline in coverage, but eventually they show some recovery. These changes are in contrast to the continually married sample, whose health insurance rates remained fairly constant throughout the time frame.

For most groups, coverage rates begin to decline several months prior to the disruption date, perhaps reflecting the ambiguity in defining the exact time of disruption

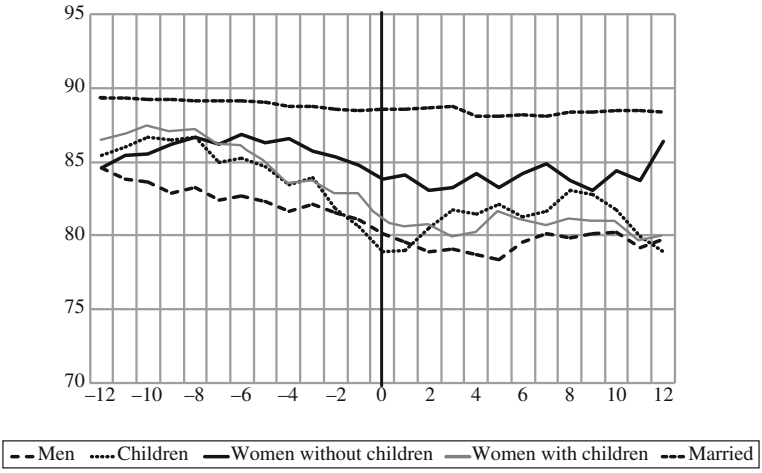

Fig. 2 Any coverage relative to marital disruption date. All figures are weighted means from the 1996, 2001, and 2004 panels of the SIPP. The sample is those married (or whose parents were married) in Wave 1 of the panels. $N = 513$ women without children, 1,196 women with children, 1,289 men, and 1,831 children

discussed earlier; coverage rates generally improve in the months following disruption. The decline could also reflect changes in health status or in employment status that might affect both insurance coverage and the likelihood of marital disruption. One should thus be cautious in interpreting these changes in coverage as being caused by marital disruption. We discuss this issue in more detail at the end of the article.

The relatively small decreases in coverage seen in Fig. 2 mask greater changes in the composition of coverage for all four groups. Comparing coverage in the 12th month before and after disruption, Fig. 3 shows that private own-name coverage increases by 17 percentage points for women without children, 13 percentage points for women with children, and 4 percentage points for men. Meanwhile, the prevalence of dependent coverage during the same period declines for all four groups, ranging from 25 percentage points for women with children to 7 percentage points for men. Although children would generally still be eligible for dependent coverage after disruption, this source of coverage falls by about 15 percentage points. During this same time frame, public coverage increases by 4 to 7 percentage points for children and both groups of women, but does not increase for men.

Multivariate Analysis

We use regression analysis with individual fixed effects, controlling for time-varying state policy and contextual factors to isolate the impact of marital status change on insurance:

$$HI_{it} = \sum_k \left(D_{it}^k \delta_k \right) + \mathbf{X}_{it}\beta + \alpha_i + \gamma_t + \varepsilon_{it}, \qquad (1)$$

where *HI* represents health insurance status for person *i* at month *t*. *K* indexes a set of monthly dummy variables, D_{it}, that indicate the time period around marital disruption. This provides a flexible functional form for capturing the association of marital disruption in each month before and after disruption through the parameters δ_k. A

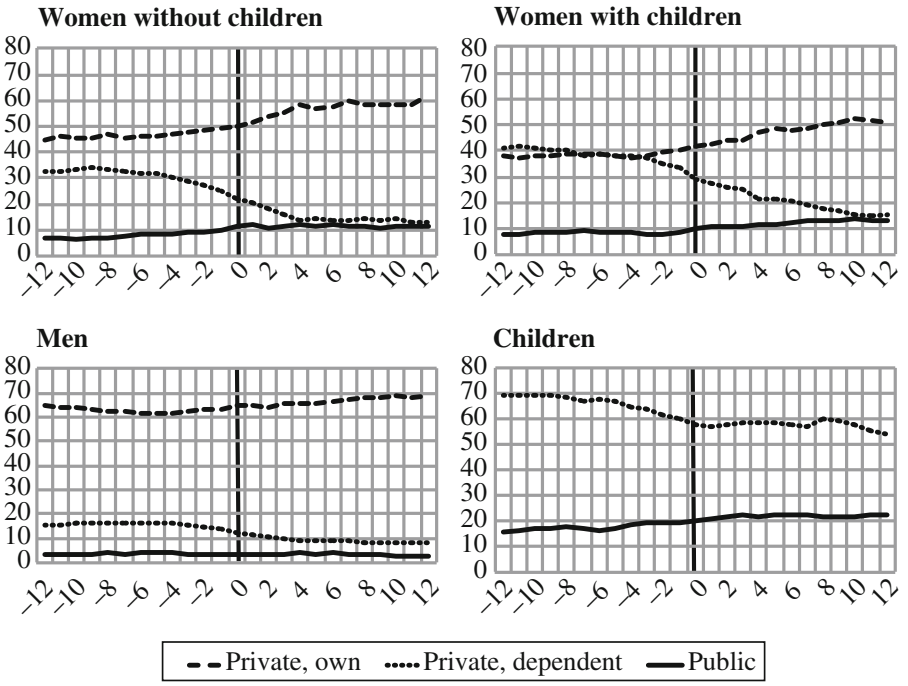

Fig. 3 Type of coverage relative to marital disruption date. All figures are weighted means from the 1996, 2001, and 2004 panels of the SIPP. The sample is those married (or whose parents were married) in Wave 1 of the panels. $N = 513$ women without children; 1,196 women with children; 1,289 men; and 1,831 children

vector of time-varying controls at the state level $\mathbf{X}_{it}$ includes the unemployment rate, the average full cost of employer health insurance plans, proxies for welfare program generosity, a variable capturing the EITC, and a measure of Medicaid generosity for children; the child specifications also control for a state/year Medicaid/SCHIP eligibility generosity index. α_i represents an individual fixed effect, γ_t is a set of calendar year dummy variables, and ε_{it} is a stochastic error term. Because we include person fixed effects, we do not include individual time-invariant characteristics, such as race/ethnicity. Additionally, we do not include endogenous time-varying personal characteristics, such as income. These models are estimated separately for different types of coverage and for different subpopulations of women, men, and children. Each model is also estimated separately by higher and lower educational attainment, defined as having at least some college education and high school graduation or less, respectively. For ease of estimation given the large number of person and time fixed effects, we use linear probability models. Standard errors are clustered at the person level.

We present coefficients and confidence intervals of the regression results in Figs. 4, 5, 6, and 7.[9] The coefficients on each dummy variable, δ_k, indicate how health insurance outcomes differ in each of the 12 months before and after marital disruption relative to the zero date reference point (month of disruption) after accounting for individual fixed effects and time-varying state policy and labor market measures.

[9] We show a full set of regression results in Table S3 in Online Resource 1 for one representative specification: women with children who have completed no more than a high school diploma.

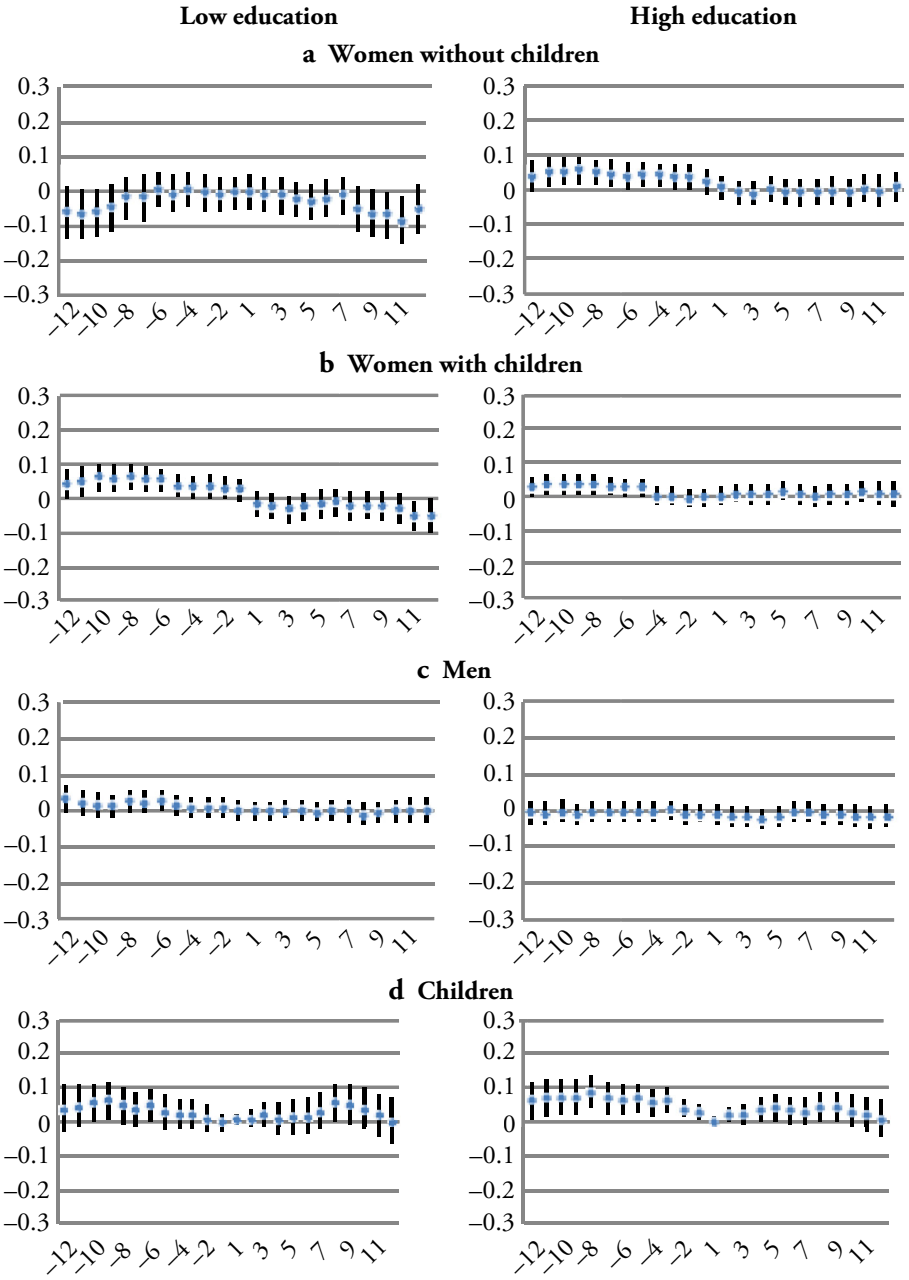

Fig. 4 Any coverage regression coefficients by education. Low education includes those with a high school diploma or less. High education includes those with some college or more. For children, this refers to the educational attainment of the mother. *Source:* 1996, 2001, and 2004 panels of the SIPP

Standard errors are shown in the form of "whiskers," which represent the 95 % confidence interval for that point estimate. Plotting the results this way enables us to parsimoniously display the regression coefficient results from many dummy variables,

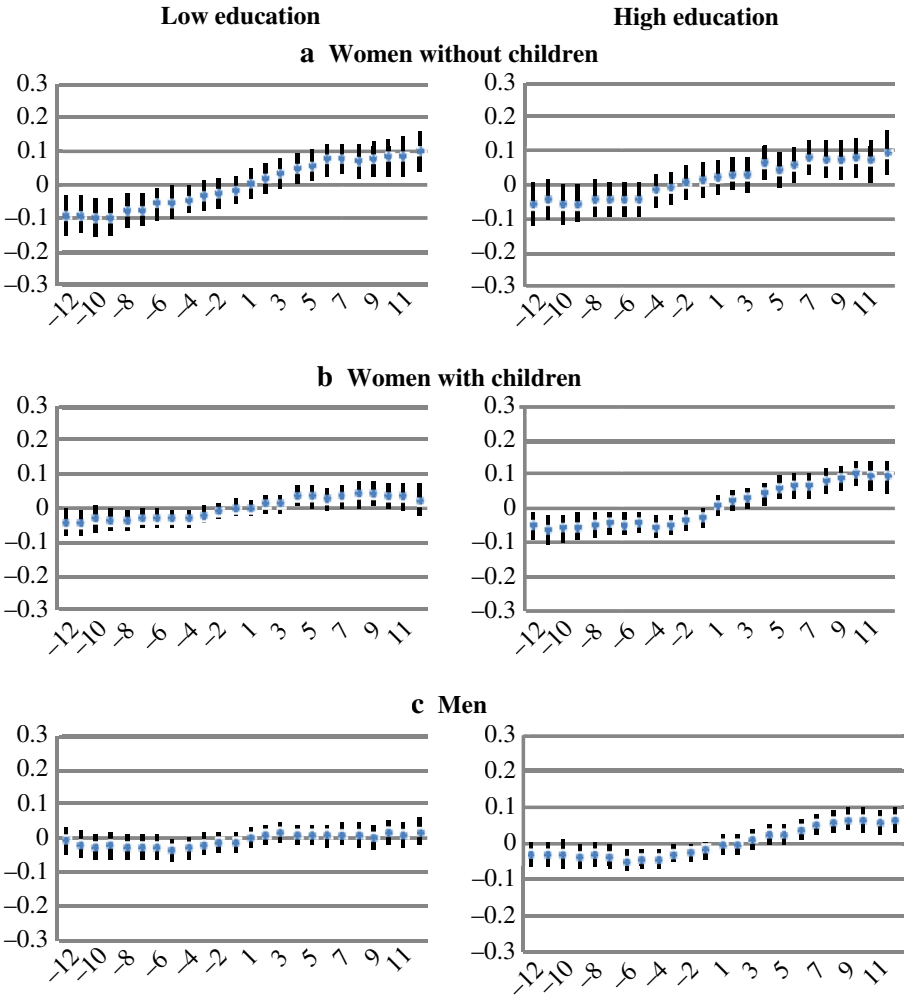

Fig. 5 Private own-name coverage regression coefficients by education. Low education includes those with a high school diploma or less. High education includes those with some college or more. For children, this refers to the educational attainment of the mother. *Source:* 1996, 2001, and 2004 panels of the SIPP

to analyze trends, and to assess their statistical significance. We also compare the coverage coefficient in the 12th month prior to marital disruption with the coverage coefficient in the 12th month following marital disruption and present statistical tests of differences (see Table 1). These are comparisons across coefficients from dummy variables that we measure relative to the zero date.

Figures 4, 5, 6, and 7 illustrate any coverage, private own-name coverage, private dependent coverage, and public coverage, respectively. Each figure contains two columns of graphs: one for those with a high school education or less, and the other

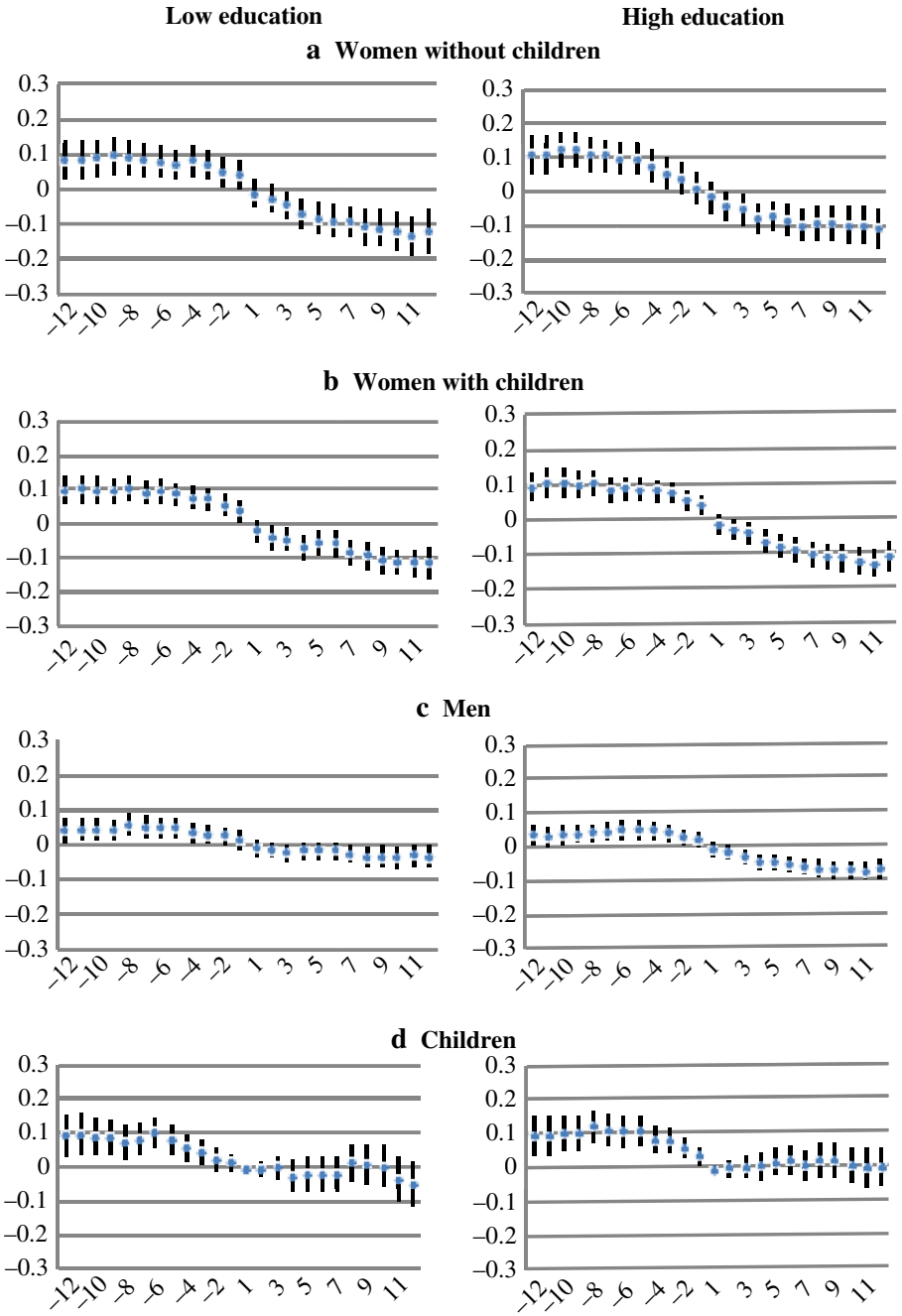

Fig. 6 Private dependent coverage regression coefficients by education. Low education includes those with a high school diploma or less. High education includes those with some college or more. For children, this refers to the educational attainment of the mother. *Source:* 1996, 2001, and 2004 panels of the SIPP

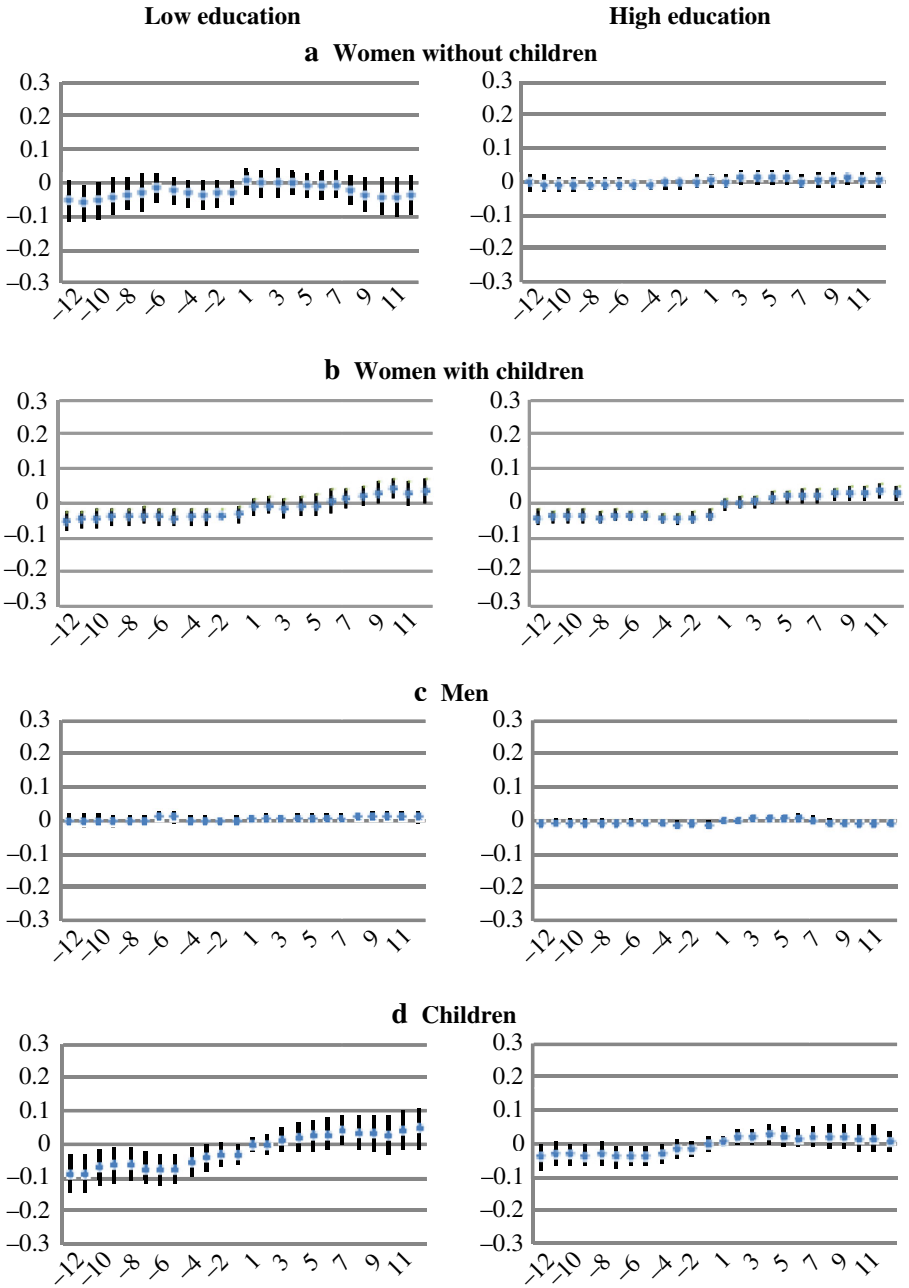

Fig. 7 Public coverage regression coefficients by education. Low education includes those with a high school diploma or less. High education includes those with some college or more. For children, this refers to the educational attainment of the mother. *Source:* 1996, 2001, and 2004 panels of the SIPP

for those with some college or more. Within each column, we provide estimates for subsamples of women without children, women with children, men, and children. We report the total number of individuals in each regression at the bottom of Table S4 in

Table 1 Difference in coefficients for the –12-month and 12-month dummy variables

	Any Coverage	Private, Own Name	Private, Dependent	Public
Women Without Children				
Lower education	0.008	0.192**	–0.206**	0.016
	(0.052)	(0.041)	(0.043)	(0.042)
Higher education	–0.033	0.152**	–0.220**	0.031*
	(0.029)	(0.043)	(0.039)	(0.013)
Women With Children				
Lower education	–0.093**	0.068*	–0.215**	0.054*
	(0.032)	(0.030)	(0.030)	(0.022)
Higher education	–0.023	0.145**	–0.207**	0.043**
	(0.023)	(0.031)	(0.029)	(0.015)
Men				
Lower education	–0.034	0.026	–0.076**	0.016
	(0.025)	(0.029)	(0.023)	(0.012)
Higher education	–0.008	0.095**	–0.108**	0.005
	(0.021)	(0.024)	(0.021)	(0.008)
Children				
Lower education	–0.037	NA	–0.143**	0.085 †
	(0.050)		(0.050)	(0.046)
Higher education	–0.053	NA	–0.099*	0.026
	(0.040)		(0.044)	(0.028)

Notes: The sample is adults age 23–55 and children age 0–15 who were married (or whose parents were married) in month 1 of the survey and subsequently divorced or separated (or whose parents divorced or separated). All means are from the first month of the survey.

†$p < .10$; *$p < .05$; **$p < .01$

Online Resource 1, and sample sizes range from 212 for lower-educated women without children to 938 for children with higher-educated parents.

Similar to our descriptive longitudinal results, the regression-adjusted results shown in Fig. 4 and Table 1 show relatively modest declines in coverage. The only group with a statistically significant decline in coverage across the two-year period is women with children and less education (9 percentage point decline). However, as we show later, this result masks larger (and offsetting) changes in specific types of insurance.

Figure 5 and Table 1 show a statistically significant increase in private own-name coverage (ranging from 7 to 19 percentage points) for all groups except for men with lower levels of education. Generally, the increase in private own-name coverage is larger for the groups with more education, which may be due to the better labor market opportunities that they face (i.e., better jobs are more likely to offer coverage).

Figure 6 and the third column of Table 1 show significant declines in dependent coverage across all groups, and the decline is relatively similar across education groups. The decline is smallest for men, who are less likely to have dependent coverage in general. Children also experience smaller declines than do women, most likely because they are still eligible to be covered after dissolution. However, the decline for children

is still 10 percentage points for the higher-educated group and 14 percentage points for the lower-educated group. Perhaps families with lower levels of education (and thus lower economic resources) find the high premiums charged for dependent coverage too expensive to maintain after a marital disruption, especially when alternative sources of coverage may be available (e.g., public). Another possible reason for this difference is that higher-educated families undergoing marital disruption are more likely to be divorced than separated, and divorce involves a more formal process of addressing children's health insurance in child support agreements (Taber 2011).

Changes in public coverage are presented in Fig. 7 and in the last column of Table 1. As expected, the largest increase in public coverage (8.5 percentage points) is for children whose mothers have less education. Women with children also have modest increases in public coverage (4 to 5 percentage points). It is, however, somewhat surprising that we see significant increases in public coverage for both the low and high education groups.

Heterogeneity of Effects

Our primary analysis focuses on assessing patterns in health insurance coverage and type of coverage before and after dissolution for four different demographic groups, with each group estimated separately for two levels of education. The patterns we find, however, may differ by other characteristics of individuals, such as baseline values for age and insurance or work status, or by whether the dissolution was a divorce or separation. Because our results show within-person changes over time, the only way to show the effect of time-invariant individual characteristics is to interact those characteristics with each of our 24 monthly time dummy variables (or to run our regressions separately by each of the relevant characteristics, as we did for our different education groups). We can then test whether the set of interactions is jointly significantly different from zero. In this section, we briefly describe the results of such an analysis for a small set of relevant characteristics.

As we discuss earlier, the pattern of results might be different for those who remain separated for a long time compared with those who transition quickly from marriage to divorce. To investigate that hypothesis, we compare the group of individuals who transitioned immediately from marriage to divorce with the group whose transition was from married to separated and who remained separated when last observed.[10] We estimate 24 regressions, one for each of three adult demographic groups stratified by high and low education as well as for each of the four types of insurance coverage. We find that the set of dummy variables is significantly different for separated and divorced individuals in only a few regressions for any, private-own, and public coverage. As might be expected, we find the most difference in dependent coverage: the set of coefficients on the dissolution dummy variables is significantly different for four of the six groups. This is more likely to be the case for the higher-education group, and the result is that the divorced group is therefore generally less likely to have dependent

[10] Because of sample size constraints, we eliminate the group who first transitioned from marriage to separation and then from separation to divorce.

coverage after dissolution compared with the separated group. This result is consistent with the hypothesis that dependent coverage is not available after divorce.

To assess the importance of other time-invariant characteristics, we estimate similar interaction models for a small set of characteristics: median age, initial insurance status (covered or not), race/ethnicity, and initial employment status. When we subset our regressions by age, race/ethnicity, or initial employment status, we see a similar pattern across the two subgroups, with a few exceptions. We see more substantial differences when we subset the regressions by initial insurance status, especially when the outcome is "any coverage." However, this is expected given that those who start without coverage can only gain coverage.

Threats to Causal Inference

Absent a natural experiment, the patterns we observe before and after marital disruption may not be causal because we are not able to fully rule out other factors that may cause marital disruption and may also affect health insurance independently. Two types of confounding mechanisms may cause challenges for our analysis: initial conditions and changes. First, initial conditions in employment and health may matter for how likely individuals are to divorce and for their experiences after divorce. For example, our results may be biased if poor health causes a person to be more likely to experience marital change and more likely to maintain health insurance because of the increased demand for it. Although these factors may play an important role in explaining the cross-sectional results, individual fixed effects in the longitudinal and multivariate analyses should adequately control for these factors if they do not vary over time.

The larger concern regards biases introduced by changes in health and employment as well as other variables. For example, those who lose a job for some exogenous reason may be more likely to experience a marital disruption as well as loss of health insurance. Alternatively, a health problem could increase the probability of both job loss and marital disruption. Although we can observe some of these changes, such as job loss, controlling for these factors can eliminate biases only if they are exogenous. However, employment and health status change could be caused by a third factor also associated with divorce or may be jointly determined with marital status change (Mincy et al. 2009).

To shed some light on potential bias resulting from job loss, we find that about 11 % of individuals in the lower-education sample and 6 % of those in the higher-education sample who are employed 12 months prior to disruption transition to no employment by the disruption date. This compares with about 4 % of lower-educated adults (of comparable age) who switch from full-time employment to no employment in a one-year period and 4 % for higher-educated adults. Thus, job loss prior to divorce seems a bit more prevalent in the population with a marital disruption than in the population as a whole, especially for those with less education.

However, just knowing that job loss takes place does not help us determine the direction of the potential bias of the impact of marital disruption on health insurance in our analysis. For example, if higher job loss is due to an individual-specific characteristic that affects both employability and marital stability, our fixed-effects specification should eliminate that bias. Similarly, if job loss occurs because of trauma in a marriage, then any loss in health insurance that might occur would appropriately be attributed to

the marital disruption. In contrast, if job loss (and subsequent loss of health insurance) creates strain in a marriage, we would be overestimating the impact of marital disruption on health insurance coverage. Although it is difficult to account for all these potential biases, our data suggest that job loss prior to disruption is only slightly higher than for the general population and that our specifications are able to account for many possible confounding factors.

Discussion and Conclusions

Despite the ubiquity of marital disruption in modern society and the vast policy attention paid to the consequences of uninsurance, surprisingly little research has explored the consequences of marital disruption for health insurance coverage of men, women, and children. We address this shortfall by examining the patterns of health insurance coverage surrounding marital disruption for these subpopulations, further subset by education level.

Our conceptual model provides a framework for understanding the ways in which marital disruption impacts health insurance through access to spousal dependent coverage, labor market participation, health status, public insurance, and other factors. Although it is difficult to unambiguously identify the causal impact of an exogenous marital disruption on health insurance coverage, we are able to control for many of the confounding factors and gain insight into the key relationships.

Analyzing longitudinal data to isolate changes within individuals suggests that the gaps in health insurance coverage found in cross-sectional data are primarily due to unobserved differences between individuals across marital status rather than the effects of marital disruption per se. When we analyze specific sources of coverage, however, we find that smaller differences in the presence of coverage mask larger shifts in type of coverage. We also find that the pattern of our results differs across education groups.

In general, our results are consistent with increased independence after marital disruption and a partial compensation for lost marital resources. For example, we observe greater reliance on own-name coverage after marital disruption even though declines in dependent coverage are not fully compensated through this avenue. Public insurance makes up for some of the shortfall in the cases of children and women with children; as expected, however, public insurance is not a factor for men. The results are also consistent with a world in which access to better health insurance benefits at work is greater for those with more education (Currie and Yelowitz 2000). For example, we find that the increase in private own-name coverage is larger for more-educated men and women with children. Lower-educated women with children appear to be the most vulnerable in terms of loss in coverage. For that group, coverage falls by 9 percentage points after dissolution because the increases in private, own-name, and public insurance are not large enough to offset the large decrease in dependent coverage. One surprising result is that dependent coverage for children falls after marital dissolution even though they are still likely to be eligible for that coverage. This is especially true for children with lower-educated parents. This decline in coverage may be related to decreased involvement of nonresidential fathers after disruption, or perhaps insurance coverage by a nonresidential father is less cost-effective if children no longer live in the same geographic area as their fathers.

For women, we can compare our results to those of Lavelle and Smock (2012), who used the same SIPP data to examine patterns of coverage following divorce. We find

broadly similar results in the magnitude of the decline in insurance after divorce for women (about 6 percentage points). We also confirm their finding that part of the cross-sectional gap in insurance between married and divorced women is present even when both groups were married, and we find smaller declines in any coverage than in private coverage because of the ability of some women to switch to public coverage. Similarly, dependent private coverage declines more than any private coverage because of women's ability to switch to own-name private coverage.

New provisions in the ACA scheduled to take place in 2014 are expected to make health insurance more widely available, especially outside the workplace, through Medicaid expansions and subsidized exchange-based private coverage. However, because employers will remain the main source of coverage for the non-elderly population (Congressional Budget Office 2012), marital disruption is likely to continue to lead to substantial instability in insurance coverage. Moreover, documenting the gaps in insurance caused by marital disruption in a period entirely prior to the ACA allows us to form a baseline against which to judge the future impact of the ACA in lessening this social problem. For both these reasons, it is critical that we understand the ways in which life course events, specifically marital disruption, shape the dynamic patterns of coverage.

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