Employer-Sponsored Health Insurance and the Gender Wage Gap

Benjamin Cowan[†]
School of Economic Sciences
Washington State University
Hulbert Hall 101
Pullman WA 99164 USA
Email: ben.cowan@wsu.edu

Benjamin Schwab[†]
Department of Agricultural Economics
Kansas State University
310 Waters Hall
Manhattan, KS 66506 USA
Email: benschwab@ksu.edu

September 21, 2015

Abstract

During prime working years, women have higher expected healthcare expenses than men. However, employees' insurance rates are not gender-rated in the employer-sponsored health insurance (ESI) market. Thus, women may experience lower wages in equilibrium from employers who offer health insurance to their employees. We show that female employees suffer a larger wage gap relative to men when they hold ESI: our results suggest this accounts for roughly 10% of the overall gender wage gap. For a full-time worker, this pay gap due to ESI is on the order of the expected difference in healthcare expenses between women and men.

Keywords: Gender, Wages, Employer-sponsored health insurance, Compensating differential
JEL Codes: I1, J3, J7

[†]We are grateful to Andrew Cassey, Donna Ginther, Robert Rosenman, David Slusky, Jon Yoder, and seminar participants at Montana State University for insightful comments and suggestions. The usual disclaimer applies.

1 Introduction

A vast literature attempts to explain why women earn less than men for comparable work. Recent estimates suggest that among full-time U.S. workers, women earn about 20 percent less than men, on average (OECD, 2014). Although the gender wage gap in the U.S. narrowed considerably in the 1980's and early 1990's, a sizable gap persists in spite of the continued growth in human-capital accumulation among women relative to men (Blau and Kahn, 2006; Goldin, 2014). The slowdown in wage convergence that began during the 1990's is largely due to the persistence of factors that are unexplained by typical wage regressions (Blau, 2012). Traditional explanations for the gender wage gap range from differences in pre-market characteristics or labor-market attachment to various forms of discrimination (see Altonji and Blank, 1999). Theories of why the gap has persisted in recent decades include changes in labor-market selectivity by gender (e.g. Blau and Kahn, 2007), non-linear remuneration with respect to hours worked (Goldin, 2014), and a burgeoning literature on innate differences in non-cognitive factors such as preferences for competition and risk-taking (see Fortin, 2008 and Gneezy et al., 2009 for two examples).

This paper proposes and examines an alternative explanation for the persistent female wage gap: gender differences in healthcare costs. Differences in health insurance charges by gender have come into sharp policy focus, culminating in the passage of the 2010 Patient Protection and Affordable Care Act (ACA), which bans gender rating in the U.S. individual health insurance market by 2014 (CMS, 2013). Traditionally, U.S. health insurers have claimed that women are costlier to insure than men, on average, and have charged women more for similar coverage.[1] This is consistent with Cylus et al. (2011), who find that female health spending per capita was \$1,500 higher than that of males in 2004. Employed women who obtain insurance in the group (employer-sponsored) insurance market face a different situation: within a firm, employee contributions to insurance premiums are not adjusted by gender (presumably out of legal considerations–see, for example, Marks, 2011). However, inasmuch as these women predictably consume more healthcare resources than men, their elevated costs are paid for by their employers (Bhattacharya and Bundorf, 2009). All else equal, this results in an incentive for firms to have male-dominated workforces unless female wages or some other form of compensation adjust. This leads us to examine whether gender differences in healthcare bills are contributing to the female wage gap.

The arbitrage opportunity that puts downward pressure on women's relative wages in firms that provide employer-sponsored insurance (ESI) disappears in firms that provide no health benefits to their employees. We exploit this dichotomy to examine whether female wages are lower, on average, as a result of their elevated health liabilities. Specifically, we use a difference-in-differences (DD) strategy that is similar to that of Bhattacharya and Bundorf (2009) (who analyze obese workers) and Cowan and Schwab (2011) (who analyze smokers) by comparing workers who receive insurance from their own employer to workers who do not.

Using data from the National Longitudinal Survey of Youth (NLSY79) and the Medical

[1] 13 states had already banned gender rating in their individual health insurance markets before the ACA took effect. A study by the National Womens Law Center (Garrett, 2012) found that in states that permitted gender rating, the vast majority of "best-selling" plans (as defined by ehealthinsurance.com) charged women more than men (of the same age) for identical coverage. In states where gender rating was permitted, modal premiums for a 40 year-old woman were found to be between 15 and 39 percent higher than those for an identical man. A similar study based on insurance records found that 30 year-old females on the individual market paid 20 percent more in premiums than their male counterparts (Coleman, 2013).

Expenditure Panel Survey (MEPS), our DD estimates indicate that the gap between male and female wages is approximately \$0.50–\$1.50 an hour larger among workers with ESI than it is among workers without ESI. Over the course of a standard full-time work year (roughly 2,000 hours), this translates to an annual pay gap of \$1,000–\$3,000, which is in line with estimates of women's additional healthcare costs. Thus, gender rating of individual healthcare expenses appears to be implicitly just as strong in the U.S. group insurance market as it is in the individual market. A full regression decomposition of men's versus women's wages indicates that women's higher average medical expenditures account for roughly 10% of the gender wage gap.

Our paper is similar in spirit to Daneshvary and Clauretie (2007), who also examine the compensating differential associated with health insurance by gender. Daneshvary and Clauretie (2007) seek to obtain unbiased estimates of the effect of receiving health insurance from one's employer on wages. Because insurance receipt is endogenous, the authors instrument for it using characteristics of spouses' jobs (firm size, whether spouse has family coverage) on a sample of married men and women from the Medical Expenditure Panel Survey (MEPS). They find that even though women experience a larger wage tradeoff in taking ESI than men do (by 3.5 percentage points), health insurance does not contribute to the unexplained part of the gender wage gap.

We take a different approach to that of Daneshvary and Clauretie (2007). In particular, instruments for health insurance of the type used in Daneshvary and Clauretie (2007) are likely to be endogenous themselves (of course, the assumption of exogeneity is fundamentally untestable). Furthermore, depending on which combination of the instruments are used, the results in Daneshvary and Clauretie (2007) differ substantially. As a result, we begin by recognizing that although ESI receipt is very likely correlated with unobserved determinants of wages, one can purge these effects by differencing across gender *as long as* the unobserved effects are the same across gender. If this assumption holds, the difference between the effect of ESI on men's wages and that on women's wages is due to women's higher expected healthcare costs.

Because identification of women's healthcare costs' contribution to the gender wage gap rests on the assumption that unobserved differences in wages by worker ESI status are not correlated with gender, we examine its plausibility with three distinct robustness exercises. First, women may be more likely than men to alter their labor supply decisions–for example, by entering the labor force or searching exclusively for jobs that offer health benefits–to procure health insurance for their families (Buchmueller and Valletta, 1999). Differential selection into ESI and non-ESI jobs by gender could itself contribute to the relatively large wage gap faced by women with ESI. To address this possibility, we examine our results by marital status and presence of children in the home. We find a wage penalty for women in jobs with ESI among all of these samples, lending confidence in the notion that differential healthcare costs by gender are driving our results.

Second, it is possible that the provision of ESI is correlated with other job characteristics that lead to a larger female wage gap in those firms that provide it than in firms that do not. For example, even after inclusion of extensive controls, ESI jobs may tend to offer more training opportunities for employees that women are less likely to take up (perhaps because of less labor market attachment or discrimination on the part of employers). Once again, this could lead to bias in our results. Consequently, we conduct a falsification exercise similar to that of Bhattacharya and Bundorf (2009) and Cowan and Schwab (2011) in which we examine the gender wage gap by receipt of several other kinds of fringe benefits. Benefits such as vocational training, profit sharing, and retirement benefits are also likely to be correlated with unobserved firm characteristics that may lead to a differential female wage penalty that is not due to healthcare costs. However, we find almost no evidence that these benefits are associated with the gender wage gap, and controlling for their presence does not alter the wage offset associated with ESI. Thus, we conclude that our results are driven by higher healthcare spending among women.

As a final way to scrutinize our results, we note that if women's higher healthcare spending contributes to the gender wage gap, the differential wage penalty in jobs with ESI should be smaller in jobs with less generous health insurance. Though coverage parameters for individual insurance plans are not available in our data, we do find that the gender wage gap associated with ESI is indeed smaller in jobs for which ESI is provided via a Health Maintenance Organization (HMO) than in non-HMO settings. As HMO plans cost employers less, on average, than other forms of ESI (Glied, 2000), this is further support that women's healthcare costs play a role in the gender wage gap.

2 Background

Though it has declined in recent years, ESI remains the largest source of private health insurance in the United States (Gruber, 2010). ESI coverage for female and male workers in 2010 was roughly 70 and 67 percent, respectively (Gould, 2012). Firms who offer ESI to their workers are a combination of those who "self-insure," whereby the risk pool is exactly composed of the firm's workforce, and firms who purchase health plans from third-party insurers. In the second case, prices charged to individual firms are generally "experience-rated," meaning they are raised or lowered depending on the projected costs of each firm's own employees (Gruber, 1998).[2] Thus, if women cost more to insure because they tend to generate larger healthcare bills, firms with a greater proportion of female employees will generally have higher expenses regardless of whether they act as their own insurer or purchase insurance on the open market.

There is a growing body of literature that documents that women have higher average healthcare expenses than men during prime adult years. While the precise size of the gender difference in healthcare spending varies by year, dataset, and methodology, several recent studies show average annual healthcare spending by women exceeds that of men by between \$1,000 and \$2,000. Among working-age adults (19-64), the Centers for Medicare and Medicaid Services (CMS) calculates a \$1,342 average spending bump for women during 2004, which falls in the middle of the time period examined in our study.[3] Cylus et al. (2011) similarly find a 37 percent spending difference (roughly \$1,500) and note that the gender cost differential occurs in every major medical spending category. Another study estimated the median gender difference in 2006 for working-age adults to be \$997 (Woolhandler and Himmelstein, 2007).

Using data from patients in a single medical system, Bertakis et al. (2000) and Bertakis and Azari (2010) find that women have mean annual health expenditures that exceed men's (by around \$2,000), owing in large part to women's greater use of primary care visits and

[2] The Affordable Care Act does abolish gender rating in the "small" group market, which is defined as firms with 100 or fewer employees with some exceptions (CMS, 2012).

[3] The 2010 CMS estimate of the gender medical spending difference for the same population is \$1,539 (CMS, 2010).

diagnostic services.[4] Gender differences in expenses remain even after controlling for socioeconomic, demographic, and health-related variables. These differences are not due solely to costs associated with the short-term effects of childbearing: Cylus et al. (2011) and Woolhandler and Himmelstein, 2007 note that large gender disparities in healthcare spending persist among older adults.

Employee contributions to ESI premiums could in theory be gender-rated as rates have been in the individual market. However, this does not occur in practice. Gruber (1998) and Marks (2011) have noted that there are strong legal and tax incentives for offering similar health benefits to all employees within a firm (EEOC, 2000).[5]

Bhattacharya and Bundorf (2009) note that because employee health insurance contributions do not vary by health status (or propensity to consume healthcare) within firms, competitive firms could earn positive profits by hiring healthy (or less costly) individuals exclusively unless other forms of compensation were adjusted to reflect differences in healthcare costs. Though adjusting every employee's compensation to account for his/her expected healthcare expenses is likely administratively burdensome, Gruber and Madrian (2002) note that firms might be able to shift costs across broad, easily identifiable groups. Indeed, Gruber (1994) finds that the incidence of mandated maternity benefits is concentrated among women of childbearing age, and Sheiner (1999) finds that older workers bear the financial

[4] Other papers have found gender differences in medical expenditures related to specific diseases (e.g. hypertension; Basu et al., 2010).

[5] This states that "health insurance benefits must be provided without regard to the race, color, sex, national origin, or religion of the insured. An employer must non-discriminatorily provide to all similarly situated employees the same opportunity to enroll in any health plans it offers. An employer must also ensure that the terms of its health benefits are non-discriminatory." In addition, Mello and Rosenthal (2008) state that under the Health Insurance Portability and Accountability Act (HIPAA), "...no person can be denied group health insurance or charged more for coverage than other "similarly situated" persons because of health status, genetic history, evidence of insurability, disability, or claims experience. The phrase "similarly situated" refers to an employment-based classification, such as full-time or part-time, not a classification based on health factors" (p. 193).

burden of their higher health risks. Similarly, Bhattacharya and Bundorf (2009) find that the incidence of the healthcare costs associated with obesity are fully internalized by obese workers (through lower wages), and Cowan and Schwab (2011) obtain a similar result with respect to smokers. In this paper, we add to the literature by examining whether women's wages differ systematically with respect to ESI status and whether gender differences in healthcare expenses are in line with the magnitude of our estimates.

Gruber and Madrian (2002) also argue that because healthcare costs vary substantially across firms, and because firms have a tax advantage in offering most of their employees the same health insurance options (if they offer any at all), there is likely to be "job lock"– individuals who could obtain a job in which they would be a better match may pass on that opportunity because the offering firm has an inferior health plan (e.g. no offer of insurance). Women, who generally have greater healthcare expenses than men, are likely to place a higher value on insurance and be more inclined to pass up jobs for insurance-related reasons. This will lower the probability of obtaining jobs that pay higher wages directly and decrease a woman's bargaining power with her current employer. Indeed, Buchmueller and Valletta (1996) find that health insurance has a larger (negative) effect on the job mobility of women, which they attribute to women's elevated healthcare expenses. This alternative hypothesis for gender differences in wages can also be tracked back to healthcare costs, but in our empirical model it is indistinguishable from the competitive equilibrium hypothesis discussed earlier.

3 Empirical Framework

Bhattacharya and Bundorf (2009) specify a conceptual model on which our empirical analysis is based. They assume a competitive spot market for labor such that in jobs with no fringe benefits, the wages of worker i, w_i, are equal to the worker's marginal revenue product, MRP_i. Let $E[m_i]$ denote expected healthcare (medical) costs. In jobs that provide health insurance, wages are equal to marginal revenue product minus expected healthcare costs, $MRP_i - E[m_i]$.[6] Since $E[m_i|female] > E[m_i|male]$, the model predicts a larger gender wage gap among workers who receive insurance through their employer.

To examine the possibility that women are paid less on account of their higher healthcare liability, we estimate models of the following form:

$$w_i = \alpha + X_i\beta + \delta * HI_i + \gamma * fem_i + \lambda * HI_i * fem_i + \varepsilon_i,$$

where w_i represents the hourly wage of worker i (or, alternatively, the log of hourly wages), HI_i indicates whether worker i receives health insurance through her employer, fem_i indicates whether worker i is female, X_i represents other observable characteristics that affect wages, and ε_i is the regression error. The difference-in-difference parameter is given by λ, which represents the differential wage offset for women with ESI compared to women without ESI.

Our key identifying assumption is that any unobserved differences in wages between men and women are not correlated with ESI status. If this is the case, λ represents the wage

[6] This assumes that the incidence of a worker's healthcare costs is borne at the individual level and that insurance premiums are actuarially fair, though fixed loading charges would not alter the analysis.

penalty due to women's higher healthcare costs. Unobserved differences between ESI and non-ESI workers alone–or between men and women alone–is not enough to threaten our strategy. Rather, for example, it is only if women are systematically less productive–or receive more intense discrimination–in ESI jobs (relative to non-ESI jobs) that the interpretation of our results is jeopardized. We provide evidence on the plausibility of the DD assumption in Section 5.

4 Data

We use the National Longitudinal Survey of Youth, 1979 cohort (NLSY79) in our main analysis. The NLSY79 is a nationally representative sample of youths who were between 14 and 22 years old in 1979. The survey was conducted annually until 1994, after which it has been conducted biennially. We chose the NLSY79 for this study because of the rich labor-market data it contains; many of its variables serve as important controls in our analysis (we highlight these below).

We use data from 2002, 2004, 2006, and 2008 waves of the NLSY79 in this paper. In 2002, the NLSY79 began classifying workers according to the 2000 Census industry and occupation codes (the 1970 and 1980 definitions had been used prior to that). Because of this change, and because industry and occupation indicators represent important control variables in our analysis, we focus on the most recent years of the survey in our empirical work. As a result, our sample is composed of individuals who are between 37 and 51 years old, ages at which women born in the 1950's and 60's have greater labor-market attachment (Lee, 2014), presumably since most are no longer having children. This largely eliminates maternity expenses as a potential difference between men's and women's healthcare costs in our data, but as is shown in other papers (e.g. Daneshvary and Clauretie, 2007), women still generate significantly higher medical spending at this stage of life.

We follow Bhattacharya and Bundorf (2009) by restricting our analysis to full-time workers (those who usually worked at least 7 hours a day at their primary job). Since the wage-setting process is likely different in government or self-employed/family business jobs, we also restrict attention to those employed in private for-profit firms or non-profit organizations. 16,772 person-year observations fit this criteria. After excluding observations that are missing data on wages, health insurance coverage, or key control variables, we are left with 13,687 person-year observations in our main study sample.

The dependent variable in our analysis is the worker's hourly wage (calculated for all workers regardless of time unit of pay in the NLSY79) for the respondent's current or most recent job.[7] We also use log wages, rather than wages in levels, in some specifications. We adopt level wages as our "primary" dependent variable because the gender wage gap associated with employer-sponsored insurance should be a function of women's higher average healthcare spending, which should be constant, rather than proportional, with respect to wages (Bhattacharya and Bundorf, 2009).

With respect to health insurance, respondents are asked whether they are covered by any kind of private, governmental, or health maintenance organization (HMO) plans. Those with coverage of some type are then asked whether that coverage is through a current employer, former employer, spouse's current employer, spouse's former employer, individually

[7]To correct errors in coding, we top code wages at $290 per hour and bottom code wages at $1 per hour. Bhattacharya and Bundorf (2009) perform the same procedure.

purchased plan, government plan, or other source (respondents may selected more than one type of coverage). We define an indicator for ESI coverage in one's own name that is equal to "1" if the respondent has coverage through his/her current employer and is "0" for those who only receive coverage through a different source or are uninsured.[8]

In our baseline model, we include the following set of covariates: the survey year, race (white, black, and other), an indicator for whether there are any children in the household and its interaction with gender, marital status (never married, married with spouse present, and other), age, age squared, education level measured by highest grade completed (0-8 years, 9-12 years, and 13 or more years), AFQT score (0-24th percentile, 25th-50th percentile, 51st-75th percentile, 76th-100th percentile), job tenure (0-52 weeks, 53-156 weeks, 157-312 weeks, and 313 or more weeks), location of residence (urban or rural), number of employees at workplace (fewer than 10 people, 10-24 people, 25-49 people, 50-999 people, and 1000 or more people), industry category, and occupation category.[9] NLSY79 summary statistics by gender and ESI coverage are presented in Table 1.

In addition to the NLSY79, we use the 2002-2008 waves of the Medical Expenditure

[8]Results in the paper are robust to eliminating individuals who receive insurance through a source other than their current employer and focusing on those with ESI versus those who are uninsured in the analysis.

[9]A similar set of control variables are used in Bhattacharya and Bundorf (2009) and Cowan and Schwab (2011). Industry categories are defined according to the 2000 Census and include Agriculture, Forestry, Fishing, and Hunting; Mining; Utilities; Construction; Manufacturing; Wholesale Trade; Retail Trade; Transportation and Warehousing; Information; Finance and Insurance; Real Estate and Rental and Leasing; Professional, Scientific, and Technical Services; Management, Administrative and Support, and Waste Management Services; Educational Services; Health Care and Social Assistance; Arts, Entertainment, and Recreation; Accomodations and Food Services; Other Services (Except Public Administration); Public Administration and Active Duty Military. 2000 Census occupation categories include Management; Business and Financial Operations; Computer and Mathematical; Architecture and Engineering; Life, Physical, and Social Services; Community and Social Services; Legal; Education, Training, and Library; Arts, Design, Entertainment, Sports, and Media; Healthcare Practitioners and Technical; Healthcare Support; Protective Service; Food Preparation and Serving Related; Building and Grounds Cleaning and Maintenance; Personal Care and Service; Sales and Related; Office and Administrative Support; Farming, Forestry, and Fishing; Construction and Extraction; Installation, Repair, and Maintenance; Production; Transportation and Material Moving; Military.

Panel Survey (MEPS) for certain supplementary analyses in the paper. MEPS collects nationally representative data on healthcare spending and health insurance coverage for the non-institutionalized population in the U.S. The recorded expenditures include both respondents' out-of-pocket costs and expenditures paid on their behalf by third parties (such as insurance companies). In addition, MEPS contains information on wages, insurance status, and many (but not all) of the controls that are available in NLSY79.[10]

There are three main reasons we supplement our NLSY79 analysis with MEPS in this paper. The first is simply that it allows us to verify that our results are robust to a different data source, with the added advantage that we can analyze a much broader range of ages (18-64) using MEPS than we can in the NLSY79 over the same time period (2002-2008). Second, because MEPS contains data on respondents' healthcare expenditures, we can compute the gender difference in expenses in our regression samples ourselves (and then directly compare those to the female wage penalty for ESI). A final advantage to MEPS is that it contains information on not just the presence of ESI, but the type of insurance plan offered by the employer. This allows us to examine whether the ESI-driven gender wage penalty is smaller when plans are less generous (i.e. HMO versus non-HMO plans), as we would predict.

Table 2 contains information on healthcare spending by sex for several different groups in MEPS. Gender cost differences among the privately insured for both the full sample (ages 18-64) as well as a sample matched to the NLSY79 (ages 37-51) are in line with CMS figures over the same time frame at roughly $1,200-$1,300 (CMS, 2010). Furthermore, the gender difference is a little over $200 smaller among those on "HMO" plans, which we define as plans that are explicitly labeled as HMO's or are other gatekeeper plans that do not pay or

[10] AFQT score and job tenure are notable exceptions.

severely restrict out-of-network coverage (roughly 52% of those with ESI in our data have an HMO-like plan by this definition). This translates to a 13-15% difference in total costs that is right in line with the 10-15% figure Glied (2000) finds in her survey of the literature on the expenditure effects of managed care. We return to this point in the next section, which examines the gender pay gap by ESI status in a regression framework.

5 Empirical Results

5.1 Baseline Model

Our main results on the gender wage offset by insurance status using the NLSY79 are contained in Table 3. Because we use repeated observations on NLSY79 respondents, we account for intra-person correlation in error terms by clustering standard errors at the individual level. All columns of Table 3 show results from an ordinary least squares regression of hourly wages on an indicator that is equal to 1 if the respondent receives health insurance through his/her employer and is zero otherwise, an indicator that is equal to 1 if the respondent is female and is zero otherwise, their interaction, and a set of control variables as discussed below.

The first column of Table 3 shows the estimation results of a model that contains all control variables discussed in Section 4 (though only selected coefficients are shown for the sake of presentation). Women who do not hold ESI earn \$1.80 less than their male counterparts. Women who hold ESI in their own name earn \$3.33 (\$1.80 plus \$1.53) less than men in the same position; thus, the regression-adjusted difference-in-difference (DD) estimate of the female wage penalty associated with ESI is \$1.53 (which is significant at the 1% level). This represents roughly an 8% reduction in hourly wages at the mean for women with ESI. Table 3 also shows that workers who hold ESI earn higher wages than workers who do not, a finding that is common in the literature (Bhattacharya and Bundorf, 2009; Levy and Feldman, 2001) but is inconsistent with the theory of compensating differentials. We note that this is likely a result of unobserved worker and firm characteristics that are correlated with both wages and fringe benefits such as ESI. The presence of such unobservables is not, *a priori*, a threat to identifying the effect of women's elevated health insurance expenses on the gender wage gap. Rather, the validity of our results depends on whether unobserved differences by ESI status are correlated with gender (after controlling for observable characteristics).

In the second column of Table 3, we add two health-related variables to the model that have been shown in the literature to have a differential effect on wages in jobs with ESI: obesity and smoking. The "smoker" variable takes a value of 1 if an individual reports being a daily smoker in a given year (and is zero otherwise), and the "obese" variable takes a value of 1 if an individual has a BMI greater than 30 in a given year (and is zero otherwise).[11] Consistent with Bhattacharya and Bundorf (2009), we find a larger wage gap for obese workers who hold ESI (relative to those who do not). The same is true of smoking as in Cowan and Schwab (2011). The model displayed in Column 2 of Table 3 also adds a set of indicators for other fringe benefits (and their interactions with gender) to the control variables associated with Column 1. Because the presence of ESI is correlated with other job benefits that may affect the gender wage gap for reasons other than women's higher average healthcare spending, inclusion of these variables represents an important check on

[11] Smoking information was collected in 2008, but prior to that it was only collected in 1998. We define smoking status in 2002 according to the 1998 values and smoking status in all other years according to 2008 values. Self-reported weight is collected in every year of our sample, and we combine this information with height in 1985 (when NLSY respondents were at least 21 years old) to calculate BMI in each year.

the robustness of our main results. The NLSY collects information on whether one's employer provides training or educational opportunities (e.g. tuition reimbursement), profit sharing, retirement plan, company-provided or subsidized childcare, dental benefits, flexible hours or work schedule, life insurance, and maternity/paternity leave (that allows one to retain his/her job).[12]

As shown in Column 2 of Table 3, inclusion of dummies for obesity and smoking status and fringe benefits (combined with a full set of interactions with "female") has almost no effect on the DD estimate associated with ESI and gender. Furthermore, the joint hypothesis that the interactions between "female" and each of the other job benefits are all zero cannot be rejected at conventional levels (the p-value associated with this Wald test is 0.9). In Section 5.4, we examine the separate effect of each of these fringe benefits on wages and do not find evidence that health insurance is merely proxying for the presence of other job characteristics that actually drive a larger gender wage gap among workers with ESI.

We also examine the possibility that unobserved gender differences in selection into different industries/occupations are correlated with ESI status. Though we include industry and occupation controls in our baseline models, they may be too coarse to pick up all of these differences. To evaluate whether this type of selection is affecting our results, we try dropping the industry and occupation dummies from our model entirely (i.e. evaluating the potential for selection on unobservables by examining selection on observables, in the spirit of Altonji et al., 2005). These results are contained in the second (rightmost) pair of columns in Table 3. Leaving out industry and occupation controls affects our estimates of the DD

[12] In contrast with employer-provided health insurance, information on the offer, rather than the take-up, of these other benefits is all that is collected in the data. Nevertheless, the offer of these other benefits is a good proxy for unobserved job characteristics that we want to control for in our regressions.

parameter only slightly (with coefficients that are a bit smaller than they are in our baseline specification). We view this as evidence that unobserved differences in industry/occupation choice by gender are not responsible for our results.

Table 4 shows the results from models in which log wages takes the place of level wages as the dependent variable. In the leftmost pair of columns (both of which include the same sample as their respective column in Table 3), the DD coefficient correspoding to the differential wage gap experienced by women with ESI is negative, but neither estimate is statistically significant at conventional levels. The DD coefficient ranges between 2.3 and 3.5 percent of wages. One problem associated with including the full range of wages in our analysis is that as nominal wages get very low (either approaching the minimum wage, if it is binding, or zero, if it is not binding), employers' scope for reducing wages to make up for increases in healthcare costs is reduced or eliminated (Marks, 2011). To account for this, we re-run our models including only those individuals whose wages fall strictly above the federal minimum wage in a given year (this reduces our sample by less than 2 percent). These results are contained in the last two columns of Table 4 (under the heading "log wages (only those above the federal minimum wage)").

The results in the two rightmost columns of Table 4 indicate that dropping outliers with extremely low wages strengthens the DD coefficients associated with the ESI-driven gender wage gap. Effects are now 3.8 and 4.8 percent of wages, respectively, and both are signficant at the 5% level.[13]

Overall, the results presented in Tables 3 and 4 lead us to conclude that women endure a

[13] We also ran our level wage specifications after dropping individuals at or below the federal minimum wage, but it made very little difference in the point estimates (thus, we do not present the results here).

larger wage penalty when they receive health coverage through their own employer, which we attribute to their elevated healthcare costs. A caveat is that our baseline DD analysis does not account for potentially different selection into the labor market by gender. A Heckman selection model (Heckman, 1979) is an appropriate tool for dealing with sample selection, though it is highly desirable to use an exclusion restriction for identification (a variable that affects the choice to participate in the labor market but does not affect the second stage, or wage process). Mulligan and Rubinstein (2008) and others have used the presence of a child under age 6 as such an instrument (controlling for total number of children in the home). The problem with that variable in our context is twofold: first, fertility and labor-supply decisions are potentially made jointly; second, the presence of young children may affect not only participation in the labor force, but also the value of health insurance (which could in turn affect wages). These issues may render such an exclusion restriction invalid; unfortunately, we do not have another instrument readily available. As a result, we run a Heckman selection model using this restriction, but the results should be treated with significant caution.

The set of results from these models are contained in Appendix Table 1.[14] Accounting for selection via a standard Heckman two-step procedure reduces the size of the DD coefficient when level wages is the dependent variable, but it remains statistically significant at the 5% level (in Section 5.5, we show that the magnitude of this effect is similar to the one we get when allowing all coefficients in the model to vary by gender). When log wages are employed, the DD coefficients are similar to the ones displayed in Table 4.

[14]We allow for selection by both genders into our regression sample (full-time workers in private and non-profit firms) rather than just into the pool of workers with non-missing wages.

5.2 Results using MEPS

Summary statistics on our MEPS wage regression sample are provided in Appendix Table 2. Restrictions on the MEPS sample follow those applied to the NLSY79 (as described in Section 4). As noted earlier, one reason for using MEPS to supplement the NLSY79 in this paper is that the full range of typical working ages (18-64) is available in MEPS over our sample frame of 2002-2008. With this in mind, we produce an analysis that is similar to the one in Tables 3 and 4 but for MEPS. The results of this exercise are contained in Table 5. Model 1 (Columns 1 and 3 in the table) uses a dummy for any kind of ESI (and its interaction with “female”) as in the NLSY79 analysis. Model 2 (Columns 2 and 4) includes separate dummies for HMO and non-HMO employer plans (and their interactions with “female”) since this information is available in MEPS.

Even though the set of MEPS controls is not exactly the same as in the NLSY79, the gender wage gap associated with ESI in Model 1 is very similar to the one in Table 3 for level wages. This is consistent with our finding in Table 2 that cost differences by sex are similar for the two age groups (37-51 and 18-64). The MEPS DD coefficient is somewhat larger in absolute value than the NLSY79 one when log wages is the dependent variable (Column 3), but overall we view the results in Table 5 as being highly consistent with our NLSY79 findings.

Since MEPS differentiates by type of insurance plan, we can also look at whether the gender wage gap due to ESI is smaller in HMO-style health plans, which tend to limit healthcare expenditures of enrollees and thus the gap between men’s and women’s healthcare costs (as shown in Table 2). A question that arises when we parse our data by type of insurance plan is whether the endogeneity of insurance plan choice for firms and workers (typically through the job choice, although some employers might offer both types of plans) might affect our results. For example, healthier people may select into HMO-style plans and also have higher wages. Once again, however, this selection is only problematic for our strategy if there are systematic differences in behavior across sex. Furthermore, using data from the RAND health insurance experiment, Manning et al. (1984) find no difference in healthcare expenditures for those randomly assigned to an HMO and those who were voluntarily enrolled in an HMO.

Columns 2 and 4 of Table 5 display these results for level wages and log wages, respectively. The estimates indicate that the female penalty associated with ESI is substantially larger when she is on a non-HMO plan through her employer. The difference in wages between women on HMO and non-HMO plans would actually imply a cost difference that is quite a bit larger than the one we found using the MEPS cost data (of a little more than $200 per year). Thus, the magnitude of the difference should be treated cautiously. We do not fully understand why the female wage penalty is disproportionately large for more generous (non-HMO) insurance, but we note that to the extent that insurance plan types offered by firms are correlated with individual and job characteristics, some of those characteristics that are observed in the NLSY79 are not available in MEPS. Nevertheless, we believe that overall our MEPS analysis confirms our conclusions from the last section, which is that women endure a larger wage offset when they receive health coverage through their own employer. Furthermore, this gap appears to be due to their elevated healthcare expenses, since it is smaller when the insurance type curtails the gender cost differential.

Another margin on which we can examine our results using the MEPS is age. The gap in female versus male healthcare costs does shrink at later ages and closes entirely by around age 60 (Yamamoto, 2013). Since the cost gap is smaller for older adults, we might expect the ESI-driven gender wage gap to be smaller for this group. To test this, we split our sample into two age categories: ages 18-40 and 41-64. We display the results of our regressions run separately for these two groups in Table 6. We find little difference in the DD coefficients for younger and older workers. This is puzzling, but one potential limitation to our study design in this respect is that younger and older workers in our sample largely come from different cohorts (in fact, older females experience a much larger wage penalty overall in our data, as seen in Table 6). Future work could examine the gender wage gap by ESI status within cohorts over time.

5.3 Results by Marital and Child Status

As discussed in Section 1, women and men may have different propensities to search for or remain in jobs with health insurance benefits. Outside of differences in expected healthcare usage, married women and/or women with children may suffer a wage penalty in jobs with health benefits because their first priority is to obtain insurance for their families, leading to a greater willingness to trade off health benefits for wages and possibly a greater degree of job lock (Buchmueller and Valletta, 1999). If such is the case, the additional wage burden associated with ESI is likely to be lower among single women and/or childless women. On the other hand, if married men are generally more able than married women to obtain family health coverage through their employer, their spouses may be less tied to jobs offering ESI and thus experience a lower amount of job lock. In either case, the effect of ESI on wages may differ across gender for reasons other than women's higher average healthcare expenses. We examine this possibility by performing our regression models separately by marital status and the presence of dependent children in the household.

Table 7 shows the effects of our baseline model for each group–single, married, without children, with children–separately. We first note that the DD coefficient representing the differential wage offset for women with health insurance is negative and economically large for all groups. The estimates for unmarried and married individuals are very similar and both are statistically significant at the 5% level. Among childless individuals, the effect is smaller (about $1 per hour) and does not achieve statistical significance at conventional levels. As noted, however, it is still economically meaningful. Furthermore, the hypothesis that it is the same as the effect for individuals with children (which is roughly twice as large) cannot be rejected at conventional levels.

Though the evidence should be interpreted cautiously, we broadly conclude that since a negative, sizable interaction between "female" and "ESI" persists regardless of marital status or the presence of children, women bear a unique wage burden in jobs with ESI on account of their elevated healthcare expenses rather than as a result of differences in labor supply decisions generated by family considerations.[15]

[15] We also considered examining the interaction of gender and ESI status by family versus individual coverage, since men who hold family coverage often have female spouses on their plans (and vice versa), which could eliminate costs differences between men and women with family/dependent coverage. In preliminary results, we did not find that the gender pay gap for ESI was smaller among those holding family coverage. However, these results are not necessarily inconsistent with our theory. In particular, since individuals can often switch coverage types even within the same firm over time, firms may not condition on coverage type when making wage offers that are persistent (sticky) over time (even if they do condition on gender, which is time-invariant). Bhattacharya and Bundorf (2009) make a similar argument to suggest that obese workers who have employers who offer ESI will likely endure an ESI-related wage hit even if they do not take up that insurance, since they can usually opt in at any time. An unrelated issue with this analysis is that we found evidence of non-trivial measurement error in family/individual coverage status in MEPS (e.g. 8% of our full-time privately employed sample claim they have "family" coverage yet do not appear to report having any dependents supported by that coverage). As a result of these issues, we leave important questions

5.4 Falsification Exercise: Other Fringe Benefits

This section of the paper examines whether women experience a larger wage penalty in jobs that provide benefits other than health insurance. We show in Table 3 that the gender wage offset associated with ESI is not affected by using other fringe benefits as controls and that the overall effect of these benefits on wages is not statistically different for women and men. However, due to the high degree of correlation between the presence of one kind of benefit and another, in this section we examine how each benefit affects wages by gender separately. Following Bhattacharya and Bundorf (2009), we use this as a falsification test of our main hypothesis that the larger gender wage gap among workers with ESI is a result of higher average medical bills for women. Since the cost of providing many other kinds of benefits should not vary much by sex, if a larger wage offset for women relative to men were observed in jobs with such benefits, it would cast doubt on our hypothesis. Rather, it would suggest that unobserved differences between male and female workers are not uniform across jobs based on their likelihood of providing a variety of fringe benefits (including ESI).

We perform separate regressions in which an indicator for one of the fringe benefits described in Section 5.1 and its interaction with "female" is added to our baseline model. This is then repeated for each of those benefits. The resulting DD coefficients are shown in Table 8. In no case does the effect on wages of the fringe benefit in question depend on gender in a statistically significant way, and the largest negative interaction term (for training/education) is only about 28 cents per hour. The DD coefficient associated with ESI varies between $1.20 and $1.70 an hour (depending on the model) and is always significant

related to dependent coverage and the ESI-related gender wage gap over the life cycle for future work.

at the 5% level.[16] Overall, these results provide evidence that the presence of employer-sponsored health insurance coverage is associated with a wage gap for women due to higher expected healthcare costs rather than differences in unobserved job or worker characteristics.

5.5 Regression Decomposition of the Gender Wage Gap

Thus far, we have shown that the gender wage gap is significantly larger among workers who receive health insurance from their employer. In this section, we examine how much ESI contributes to the overall gender wage gap. To do so, we follow the approach outlined in Fortin (2006) and Jann (2008) as a modification of the method of Oaxaca and Ransom (1994). In particular, we estimate separate models for men and women including all of the relevant covariates in the baseline model discussed in Section 5.1. We then estimate a pooled model that includes a "female" dummy (as recommended in Jann, 2008). The estimates from this model serve as the coefficients from a (counterfactual) non-discriminatory wage structure. Using estimates from the male, female, and pooled regressions, the overall wage differential is decomposed into an "explained" portion (due to differences in attributes) and an "unexplained" portion (due to differences in returns to attributes).[17]

The results of this exercise are contained in Table 9. ESI contributes positively to the explained portion of the gender wage gap because men are more likely than women to hold ESI (which is positively associated with wages in our OLS regressions). The contribution of ESI to the unexplained part of the wage differential is also positive, indicating that the return to ESI is smaller for women than for men. In the case of level wages, the estimated ESI

[16] The models in Table 8 allow the effects of each industry and occupation to also vary by gender to account for the possibility that some industries or fields are more likely to provide certain fringe benefits but may also vary in their level of gender discrimination.

[17] For details, see Jann (2008).

effect in determining the unexplained gap is 0.49 and narrowly misses statistical significance at the 10% level. With respect to log wages, the same effect is a statistically significant (at 5%) 0.03. In both cases, the gender difference in the return to ESI–which we interpret as being due to the higher average costs associated with insuring women–is roughly 10 percent of the overall gender wage gap.[18]

6 Conclusions

In this paper, we examine the role that healthcare costs play in the well-documented gender wage gap. We find a significantly higher wage penalty for female workers with employer-sponsored insurance (ESI) than we do for women without ESI. Our body of results leads us to conclude that this phenomenon is due to predictable differences in the use of healthcare resources by gender. One way to examine the plausibility of our results is to compare the annual loss in wages for women with ESI to estimates of actual healthcare cost differences by gender. The estimates in Table 3 suggest the hourly loss in wages is roughly \$1-\$1.50, while the full decomposition results in Table 9 imply that the difference is smaller, or about \$0.50. Over the course of a year for a full-time worker, this results in a range for the annual (extra) pay gap of \$1,000–\$3,000. This range overlaps several recent estimates of the annual healthcare cost difference between men and women, including those in Bertakis et al. (2000)

[18]We include Appendix Table 3 to show the decomposition described above by marital and child status. In each case, the table shows the contribution of ESI to the unexplained part of the gender wage gap. The results indicate that the level contribution of ESI is somewhat larger for married women than single women, though they are similar as a fraction of the unexplained gap for each group. The level gap is similar for childless women and women with children, though the percentage of the unexplained gap is higher for childless women (since they have a much lower unexplained gap than do mothers). Similar to our baseline analysis by age, there was no difference in the contribution of ESI to the gender pay gap for younger and older workers.

and Woolhandler and Himmelstein (2007).

Much recent policy attention has focused on the use of gender rating in the individual health insurance market. The ACA bans gender rating in the individual market but does not affect the employer-provided market (with the exception of very small firms). Our results imply that even though employee contributions to insurance rates do not vary by gender within a firm, wage offsets account for women's higher insurance risks overall.

Our results also have implications for explaining the persistence of the gender wage gap over time. Because the price of medical care has risen in real terms over time, it is possible that the gap between men's and women's healthcare expenses has also grown. As a result, wage gains that women have made as a result of increased schooling and experience and lower barriers to high-paying occupations may have been partially offset by divergence between their healthcare bills and those of men. The results also imply that countries that rely more heavily on employer contributions to pay for employee healthcare expenses should exhibit large gender wage gaps, all else equal (it is provocative that the United States and Japan, two developed countries in which employers pay a large share of healthcare bills, also have the largest adjusted wage gaps, as identified by Blau and Kahn, 2003). Careful analyses of the relationship between healthcare spending and the gender wage gap across time and country are left to future research.

References

Altonji, J. and R. Blank (1999): "Race and Gender in the Labor Market," *Handbook of Labor Economics*, 3, 3143–3259.

Altonji, J. G., T. E. Elder, and C. R. Taber (2005): "Selection on Observed and Unobserved Variables: Assessing the Effectiveness of Catholic Schools," *Journal of Political Economy*, 113, 151–184.

Basu, R., L. Franzini, P. Krueger, and D. Lairson (2010): "Gender Disparities in Medical Expenditures Attributable to Hypertension in the United States," *Women's Health Issues: Official Publication of the Jacobs Institute of Women's Health*, 20, 114.

Bertakis, K. and R. Azari (2010): "Patient Gender Differences in the Prediction of Medical Expenditures," *Journal of Women's Health*, 19, 1925–1932.

Bertakis, K., R. Azari, L. Helms, E. Callahan, and J. Robbins (2000): "Gender Differences in the Utilization of Health Care Services." *The Journal of Family Practice*, 49, 147.

Bhattacharya, J. and M. K. Bundorf (2009): "The Incidence of the Healthcare Costs of Obesity," *Journal of Health Economics*, 28, 649–658, rX: 850816 (on Apr 14, 2010).

Blau, F. (2012): "Gender, Inequality, and Wages," .

Blau, F. and L. Kahn (2006): "The US Gender Pay Gap in the 1990s: Slowing Convergence," *Industrial and Labor Relations Review*, 45–66.

——— (2007): "Changes in the Labor Supply Behavior of Married Women: 1980–2000," *Journal of Labor Economics*, 25.

Blau, F. D. and L. M. Kahn (2003): "Understanding International Differences in the Gender Pay Gap," *Journal of Labor Economics*, 21, 106–144.

Buchmueller, T. and R. Valletta (1996): "The Effects of Employer-provided Health Insurance on Worker Mobility," *Industrial and Labor Relations Review*, 439–455.

——— (1999): "The Effect of Health Insurance on Married Female Labor Supply," *Journal of Human Resources*, 42–70.

CMS, C. (2010): "Total Personal Health Care Spending by Gender and Age Group," http://www.cms.gov/Research-Statistics-Data-and-Systems/Statistics-Trends-and-Reports/NationalHealthExpendData/Downloads/2010GenderandAgeTables.pdf, accessed: July 15, 2014.

——— (2012): "SHOP and the Small Group Market Policies," https://www.cms.gov/CCIIO/Resources/Presentations/Downloads/hie-shop-and-the-small-group-market-policies.pdf, accessed: July 14, 2014.

——— (2013): "Overview: Final Rule for Health Insurance Market Reforms," https://www.cms.gov/CCIIO/Resources/Files/Downloads/market-rules-technical-summary-2-27-2013.pdf, accessed: July 14, 2014.

Coleman, K. (2013): "Age Gap Bigger than Gender Gap for Health Insurance Premiums," http://www.healthpocket.com/healthcare-research/infostat/age-gap-bigger-than-gender-gap-for-health-insurance-premiums#.U7_xGuKTH0q, accessed: July 11, 2014.

Cowan, B. and B. Schwab (2011): "The Incidence of the Healthcare Costs of Smoking," *Journal of Health Economics*, 30, 1094–1102.

Cylus, J., M. Hartman, B. Washington, K. Andrews, and A. Catlin (2011): "Pronounced Gender and Age Differences are Evident in Personal Health Care Spending per Person," *Health Affairs*, 30, 153–160.

Daneshvary, N. and T. M. Clauretie (2007): "Gender Differences in the Valuation of Employer-Provided Health Insurance," *Economic Inquiry*, 45, 800–816.

EEOC, U. (2000): "Directives Transmittal No. 915.003 EEOC Compliance Manual Chapter 3: Employee Benefits," http://www.eeoc.gov/policy/docs/benefits.html, accessed: July 15, 2014.

Fortin, N. (2008): "The Gender Wage Gap among Young Adults in the United States: The Importance of Money versus People," *Journal of Human Resources*, 43, 884–918.

Fortin, N. M. (2006): "Greed, Altruism, and the Gender Wage Gap," *Unpublished manuscript, Department of Economics, University of British Columbia.*

Garrett, D. (2012): "Turning to Fairness: Insurance Discrimination against Women Today and the Affordable Care Act," National Women's Law Center Publication.

Glied, S. (2000): "Managed Care," *Handbook of Health Economics*, 1, 707–753.

Gneezy, U., K. Leonard, and J. List (2009): "Gender Differences in Competition: Evidence from a Matrilineal and a Patriarchal Society," *Econometrica*, 77, 1637–1664.

Goldin, C. (2014): "A Grand Gender Convergence: Its Last Chapter," *American Economic Review*, 104, 1091–1119.

Gould, E. (2012): "A Decade of Declines in Employer-Sponsored Health Insurance Coverage," Economic Policy Institute Briefing Paper 337.

Gruber, J. (1994): "The Incidence of Mandated Maternity Benefits," *The American Economic Review*, 622–641.

——— (1998): "Health Insurance and the Labor Market," SO: National Bureau of Economic Research, Inc, NBER Working Papers: 6762, 1998 RX: 850816 (on Jun 10, 2010).

——— (2010): "The Tax Exclusion for Employer-Sponsored Health Insurance," SO: National Bureau of Economic Research, Inc, NBER Working Papers: 15766, 2010 RX: 850816 (on Jun 10, 2010).

Gruber, J. and B. Madrian (2002): "Health Insurance, Labor Supply, and Job Mobility: A Critical Review of the Literature," Tech. rep., National Bureau of Economic Research.

Heckman, J. J. (1979): "Sample Selection Bias as a Specification Error," *Econometrica: Journal of the Econometric Society*, 153–161.

Jann, B. (2008): "The Blinder-Oaxaca Decomposition for Linear Regression Models," *The Stata Journal*, 8, 453–479.

Lee, J. Y. (2014): "The Plateau in US Women's Labor Force Participation: A Cohort Analysis," *Industrial Relations: A Journal of Economy and Society*, 53, 46–71.

Levy, H. and R. Feldman (2001): "Does the Incidence of Group Health Insurance Fall on Individual Workers?" *International Journal of Health Care Finance and Economics*, 1, 227–247, rX: 850816 (on Sep 25, 2010).

Manning, W. G., A. Leibowitz, G. A. Goldberg, W. H. Rogers, and J. P. Newhouse (1984): "A Controlled Trial of the Effect of a Prepaid Group Practice on Use of Services," *The New England Journal of Medicine*, 310, 1505–1510.

Marks, M. (2011): "Minimum Wages, Employer-Provided Health Insurance, and the Non-discrimination Law," *Industrial Relations: A Journal of Economy and Society*, 50, 241–262.

Mello, M. M. and M. B. Rosenthal (2008): "Wellness Programs and Lifestyle Discrimination - The Legal Limits," *New England Journal of Medicine*, 359, 192–199, pT: J.

Mulligan, C. B. and Y. Rubinstein (2008): "Selection, Investment, and Women's Relative Wages over Time," *The Quarterly Journal of Economics*, 123, 1061–1110.

Oaxaca, R. L. and M. R. Ransom (1994): "On Discrimination and the Decomposition of Wage Differentials," *Journal of Econometrics*, 61, 5–21.

OECD, O. (2014): "Gender Wage Gap," http://www.oecd.org/gender/data/genderwagegap.htm, accessed: July 14, 2014.

Sheiner, L. (1999): "Health Care Costs, Wages, and Aging," Tech. rep., Board of Governors of the Federal Reserve System (US).

Woolhandler, S. and D. Himmelstein (2007): "Consumer Directed Healthcare: Except for the Healthy and Wealthy Its Unwise," *Journal of General Internal Medicine*, 22, 879–881.

Yamamoto, D. H. (2013): "Health Care Costs–From Birth to Death," *Health Care Cost Institute Report*, 1.

Table 1: Selected descriptive statistics, 2002-2008 NLSY79

	All individuals (N=13,687)		*Men without ESI (N=2,274)*		*Men with ESI (N=5,167)*		*Women without ESI (N=2,466)*		*Women with ESI (N=3,780)*	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Hourly wage ($2002)	20.68	17.29	16.67	13.07	26.11	20.92	13.37	9.01	18.94	13.95
Female	0.43		---		---		---		---	
Employer health coverage in own name (ESI)	0.68		---		---		---		---	
Female*employer coverage (own)	0.26		---		---		---		---	
Black	0.12		0.17		0.09		0.13		0.14	
Hispanic	0.06		0.07		0.05		0.06		0.05	
Non-black, Non-Hispanic	0.82		0.75		0.85		0.81		0.81	
Any children in household	0.62		0.53		0.62		0.71		0.63	
Never married	0.13		0.17		0.12		0.06		0.15	
Formerly married	0.23		0.23		0.17		0.24		0.31	
Currently married	0.64		0.60		0.70		0.69		0.54	
Age	43.86	3.14	43.65	3.19	43.98	3.12	43.75	3.05	43.86	3.19
Education: <9	0.01		0.03		0.01		0.02		0.00	
Education: 9-12	0.49		0.61		0.46		0.53		0.45	
Education: 13 and over	0.50		0.36		0.53		0.46		0.55	
AFQT: 0-25	0.24		0.38		0.21		0.30		0.18	
AFQT: 25-50	0.28		0.26		0.22		0.33		0.33	
AFQT: 50-75	0.24		0.20		0.25		0.22		0.25	
AFQT: 75-100	0.24		0.16		0.31		0.16		0.23	
Job tenure: 0-1 years	0.14		0.25		0.09		0.23		0.09	
Job tenure: 1-3 years	0.20		0.25		0.17		0.25		0.17	
Job tenure: 3-6 years	0.19		0.19		0.18		0.19		0.20	
Job tenure: 6+ years	0.48		0.31		0.57		0.32		0.54	
Urban residence	0.66		0.65		0.66		0.64		0.69	
Employer size: 0-9	0.18		0.37		0.11		0.28		0.10	
Employer size: 10-24	0.14		0.18		0.13		0.18		0.12	
Employer size: 25-49	0.12		0.13		0.12		0.13		0.10	
Employer size: 50-999	0.43		0.27		0.48		0.36		0.49	
Employer size: 1000+	0.13		0.06		0.16		0.05		0.19	

Notes: Estimates are weighted according to the NLSY79 sample weights. Unit of observation is a person-year.

Table 2: Estimates of annual medical expenditures by gender, 2002-2008 MEPS

	Women	Men	Difference
All working ages			
Ages 18-64, all privately insured	$3,435	$2,096	$1,339***
Ages 18-64, insured via HMO plan	$3,290	$1,938	$1,352***
Ages 18-64, insured via non-HMO plan	$4,001	$2,444	$1,557***
Matched to NLSY79 sample			
Ages 37-51, all privately insured	$3,294	$2,065	$1,229***
Ages 37-51, insured via HMO plan	$3,119	$1,871	$1,248***
Ages 37-51, insured via non-HMO plan	$3,769	$2,297	$1,473***

*Significant at 10%; **significant at 5%; ***significant at 1%. Note: Estimates are denoted in $2002.

Table 3: Estimates of the gender wage offset for health insurance, 2002-2008 NLSY79

	Dep. Var.: level wages (includes industry and occupation controls)		Dep. Var.: level wages (excludes industry and occupation controls)	
	1	2	1	2
Female*employer coverage	-1.529***	-1.520**	-1.194**	-1.387**
	(0.483)	(0.632)	(0.524)	(0.691)
Female	-1.803***	-1.383**	-1.825***	-2.179***
	(0.473)	(0.624)	(0.468)	(0.600)
Employer health coverage in own name	3.676***	4.479***	4.187***	4.858***
	(0.365)	(0.622)	(0.399)	(0.682)
Daily smoker	---	-0.434	---	-0.600*
		(0.306)		(0.336)
Daily smoker*employer coverage	---	-2.076***	---	-2.522***
		(0.524)		(0.569)
Obese	---	-0.960***	---	-0.911***
		(0.319)		(0.340)
Obese*employer coverage	---	-1.772***	---	-2.251***
		(0.503)		(0.546)
Observations	13,687	10,457	13,687	10,457
R-squared	0.348	0.360	0.274	0.293
Controls for smoking and obesity	No	Yes	No	Yes
Controls for other fringe benefits and their interactions with female	No	Yes	No	Yes
P-value associated with test that effect of fringe benefits is the same across gender	---	0.90	---	0.86

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. The dependent variable in all models is the worker's hourly wage in $2002. Model 1 includes controls for children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry, and occupation. Model 2 adds controls for smoker (current daily smoker), obese (BMI>=30), and their interactions with employer-sponsored health insurance, plus controls for other fringe benefits (education/training opportunities, profit sharing, retirement plan, company-provided or subsidized childcare, dental benefits, flexible hours/work schedule, life insurance, maternity/paternity leave) and their interactions with female. In the right panel of the table, industry and occupation dummies are excluded from both models.

Table 4: Estimates of the gender wage offset for health insurance, log wage specification, 2002-2008 NLSY79

	Dep. Var.: log wages (full sample)		Dep. Var.: log wages (only those above federal min. wage)	
	1	2	1	2
Female*employer coverage	-0.023	-0.035	-0.038**	-0.048**
	(0.021)	(0.025)	(0.019)	(0.023)
Female	-0.128***	-0.171***	-0.116***	-0.131***
	(0.022)	(0.032)	(0.020)	(0.027)
Employer health coverage in own name	0.208***	0.174***	0.193***	0.165***
	(0.015)	(0.022)	(0.014)	(0.020)
Daily smoker		-0.069***		-0.062***
		(0.019)		(0.016)
Daily smoker*employer coverage		-0.017		-0.028
		(0.024)		(0.021)
Obese		-0.054***		-0.064***
		(0.018)		(0.016)
Obese*employer coverage		-0.023		-0.020
		(0.022)		(0.020)
Observations	13,687	10,457	13,422	10,252
R-squared	0.494	0.520	0.530	0.553
Controls for smoking and obesity	No	Yes	No	Yes
Controls for other fringe benefits and their interactions with female	No	Yes	No	Yes
P-value associated with test that effect of fringe benefits is the same across gender	---	0.87	---	0.92

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. The dependent variable in all models is the worker's log hourly wage in $2002. Model 1 includes controls for children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry, and occupation. Model 2 adds controls for smoker (current daily smoker), obese (BMI>=30), and their interactions with employer-sponsored health insurance, plus controls for other fringe benefits (education/training opportunities, profit sharing, retirement plan, company-provided or subsidized childcare, dental benefits, flexible hours/work schedule, life insurance, maternity/paternity leave) and their interactions with female.

Table 5: Estimates of the gender wage offset for health insurance, 2002-2008 MEPS

	Dep. Var.: level wages		Dep. Var.: log wages	
	1	2	1	2
Female*employer coverage	-1.602***	---	-0.040***	---
	(0.223)	---	(0.012)	---
Female* HMO coverage	---	-1.315***	---	-0.033**
	---	(0.262)	---	(0.014)
Female *non-HMO coverage	---	-2.031***	---	-0.052***
	---	(0.275)	---	(0.014)
Employer coverage	3.707***	---	0.226***	---
	(0.173)	---	(0.009)	---
HMO coverage	---	3.209***	---	0.211***
	---	(0.202)	---	(0.010)
Non-HMO coverage	---	4.513***	---	0.254***
	---	(0.211)	---	(0.010)
Female	-2.829***	-2.815***	-0.154***	-0.154***
	(0.188)	(0.189)	(0.011)	(0.011)
Observations	40,724	38,626	40,724	38,626
R-squared	0.44	0.44	0.47	0.45

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. Model 1 includes a dummy for ESI, while Model 2 includes separate dummies for ESI with an HMO-style plan and ESI with a non-HMO plan (in each case, the omitted category is "no ESI"). Plans assigned as "HMO" include those explicitly reported as such as well as gatekeeper plans that do not pay or severely restrict out-of-network coverage. Model 2 excludes those whose ESI status cannot be determined as HMO or non-HMO because of missing data and/or conflicting information on plan characteristics. The dependent variable in all models is the worker's hourly wage (or log wage) in $2002. All models include controls for the presence of children in the household, race, marital status, age, education, urban residence, region, employer size, interview year, industry, and occupation.

Table 6: Estimates of the gender wage offset for health insurance by age group, 2002-2008 MEPS

	Ages 18-40	Ages 41-64
Female*employer coverage	-1.545***	-1.654***
	(0.277)	(0.358)
Female	-2.173***	-3.255***
	(0.22)	(0.317)
Employer health coverage in own name	3.895***	3.728***
	(0.211)	(0.284)
Observations	19,667	21,057

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. The dependent variable in all models is the worker's hourly wage in $2002. All models include controls for the presence of children in the household, race, marital status, age, education, urban residence, region, employer size, interview year, industry, and occupation.

Table 7: Estimates of the gender wage offset for health insurance by martial status and presence of children, 2002-2008 NLSY79

	Unmarried	Married	No children in household	At least 1 child in household
Female*employer coverage	-1.281**	-1.526**	-0.970	-1.994***
	(0.606)	(0.703)	(0.638)	(0.707)
Female	-1.471***	-1.914**	-1.491***	-5.038***
	(0.513)	(0.825)	(0.429)	(0.635)
Employer health coverage in own name	3.400***	3.983***	3.502***	4.098***
	(0.439)	(0.539)	(0.371)	(0.580)
Observations	5,649	8,038	5,165	8,522
R-squared	0.322	0.341	0.388	0.334

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. The dependent variable in all models is the worker's hourly wage in $2002. All models include controls for children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry, and occupation.

Table 8: Estimates of the gender wage offset for other fringe benefits, 2002-2008 NLSY79

Training/education	-0.282							
	(0.544)							
Profit sharing		-0.112						
		(0.625)						
Retirement			0.109					
			(0.497)					
Childcare				0.352				
				(0.964)				
Dental insurance					0.162			
					(0.610)			
Flexible schedule/hours						0.106		
						(0.497)		
Life insurance							0.559	
							(0.541)	
Maternity/paternity leave								0.320
								(0.606)
Employer health coverage in	-1.201**	-1.235**	-1.338***	-1.139**	-1.352**	-1.343***	-1.627***	-1.685***
own name	(0.506)	(0.482)	(0.497)	(0.479)	(0.529)	(0.474)	(0.490)	(0.541)
Observations	13,232	13,362	13,463	13,034	13,539	13,571	13,432	12,500
R-squared	0.367	0.366	0.366	0.372	0.366	0.366	0.366	0.362

*Significant at 10%; **significant at 5%; ***significant at 1%. Note: Estimates in the table are the coefficient and standard error on the interaction term between female and whether one's current employer offers a particular fringe benefit, with each model representing a different fringe benefit (see row headings). Standard errors (in parentheses) are adjusted for repeated observations of individuals. The dependent variable in all models is the worker's hourly wage in $2002. Estimates include controls for the main effects of female and the fringe benefit in question, employer-sponsored health insurance and its interaction with female, children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry dummies and their interactions with female, and occupation dummies and their interactions with female.

Table 9: Regression decomposition of the gender wage gap, 2002-2008 NLSY79

	Level wages	Log wages
Total gender wage differential	5.228***	0.259***
	(0.391)	(0.016)
Explained differential	0.389	0.058***
	(0.324)	(0.014)
Contribution of:		
Employer-sponsored health insurance	0.271***	0.018***
	(0.042)	(0.003)
Human capital variables	-0.020	-0.005
	(0.152)	(0.006)
Job characteristics	0.138	0.045***
	(0.238)	(0.010)
Unexplained differential	4.839***	0.200***
	(0.412)	(0.014)
Contribution of:		
Employer-sponsored health insurance	0.492	0.029**
	(0.323)	(0.014)
Human capital variables	43.585	0.054
	(36.195)	(1.278)
Job characteristics	0.546	0.024
	(0.677)	(0.029)
Constant	-39.784	0.093
	(36.109)	(1.278)
Observations	13,687	13,687

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. The dependent variable in all models is the worker's hourly wage (or log wage) in $2002. Both models include controls for children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry, and occupation.

Appendix Table 1: Heckman two-step estimates of the gender wage offset for health insurance, 2002-2008 NLSY79

	Dep. Var.: level wages	Dep. Var.: log wages (full sample)	Dep. Var.: log wages (only those above federal min. wage)
Female*employer coverage	-0.535**	-0.024	-0.048***
	(0.248)	(0.021)	(0.018)
Female	-3.114***	-0.148***	-0.151***
	(0.565)	(0.024)	(0.023)
Employer health coverage in own name	2.179***	0.208***	0.196***
	(0.189)	(0.015)	(0.013)
Uncensored observations	13,681	13,681	13,416
Censored observations	15,777	15,777	16,042
P-value associated with test that effect of excluded variable (children under age 6 in the household) is zero in first-stage probit	0.001	<0.001	<0.001

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals and the two-step procedure. The dependent variable in all models is the worker's hourly wage (or log wage) in $2002. The first-stage probit includes dummies for female, presence of children in the household and its interaction with female, race, marital status, age, education, AFQT score, interview year, and presence of a child age 6 or less in the household (this variable is excluded from the second stage). The second-stage model includes the variables shown above as well as controls for children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry, and occupation.

Appendix Table 2: Selected descriptive statistics, 2002-2008 MEPS

	All individuals (N=40,724)		*Men without ESI (N=6,609)*		*Men with ESI (N=15,271)*		*Women without ESI (N=5,791)*		*Women with ESI (N=13,053)*	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Hourly wage ($2002)	21.95	12.07	17.27	12.15	25.48	11.89	16.25	10.99	21.39	10.44
Female	0.46		---		---		---		---	
Employer health coverage in own name (ESI)	0.75		---		---		---		---	
Female*employer coverage (own)	0.34		---		---		---		---	
Black	0.11		0.09		0.09		0.11		0.14	
Other non-white	0.07		0.07		0.06		0.07		0.07	
Any children in household	0.49		0.58		0.49		0.59		0.42	
Never married	0.22		0.26		0.22		0.16		0.25	
Formerly married	0.17		0.09		0.13		0.13		0.26	
Currently married	0.61		0.65		0.65		0.72		0.50	
Age	41.34	11.14	38.39	12.80	41.65	10.03	40.92	12.37	42.24	10.87
Education: 13 and over	0.44		0.46		0.44		0.49		0.43	
Education: 12	0.45		0.26		0.48		0.37		0.52	
Education: <12	0.11		0.28		0.08		0.14		0.05	
Urban residence	0.84		0.84		0.85		0.81		0.86	
Employer size: 0-9	0.14		0.35		0.10		0.22		0.09	
Employer size: 10-24	0.14		0.21		0.12		0.19		0.11	
Employer size: 25-49	0.12		0.13		0.11		0.14		0.11	
Employer size: 50-499	0.38		0.23		0.41		0.32		0.42	
Employer size: 500+	0.22		0.08		0.25		0.13		0.26	

Notes: Estimates are weighted according to the MEPS sample weights. Unit of observation is a person-year.

Appendix Table 3: Contribution of ESI to the unexplained portion of the gender wage gap, 2002-2008 NLSY79

	Level wages	Log wages
Full sample	0.492	0.029**
	(0.323)	(0.014)
	[10%]	[15%]
	N=13,687	
Unmarried	0.325	0.009
	(0.395)	(0.021)
	[11%]	[6%]
	N=5,649	
Married	0.701	0.035*
	(0.469)	(0.019)
	[12%]	[16%]
	N=8,038	
No children in household	0.584	0.029
	(0.416)	(0.022)
	[27%]	[22%]
	N=5,165	
At least 1 child in household	0.597	0.022
	(0.480)	(0.019)
	[9%]	[9%]
	N=8,522	

*Significant at 10%; **significant at 5%; ***significant at 1%. Notes: Standard errors (in parentheses) are adjusted for repeated observations of individuals. Bracketed terms show the contribution of ESI as a percentage of the total unexplained gap. The dependent variable in all models is the worker's hourly wage (or log wage) in $2002. All models include controls for children in the household and its interaction with female, race, marital status, age, education, urban residence, AFQT score, job tenure, employer size, interview year, industry, and occupation.