

## Let Them Have Choice: Gains from Shifting Away from Employer-Sponsored Health Insurance and Toward an Individual Exchange<sup>†</sup>

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*Most nonelderly Americans purchase health insurance through their employers, which sponsor a limited number of plans. Using a panel dataset representing over ten million insured lives, we estimate employees' preferences for different health plans and use the estimates to predict their choices if more plans were made available to them on the same terms, i.e., with equivalent subsidies and at large-group prices. Using conservative assumptions, we estimate a median welfare gain of 13 percent of premiums. A proper accounting of the costs and benefits of a transition from employer-sponsored to individually-purchased insurance should include this nontrivial gain. (JEL G22, I13, J32)*

Nearly 60 percent of nonelderly Americans purchase employer-sponsored health insurance (ESI).<sup>1</sup> Although there are no legal impediments to offering a broad array of plans, in practice employers offer a very limited set of choices: a 2005 survey by the Kaiser Family Foundation/Health Retirement Education Trust finds 80 percent of employers who offer insurance provide only one option. The restriction of employee choice may prevent individuals and families from selecting the healthplan that best suits their needs, and from trading off added benefits against the associated premium increases.

As the United States embarks on the most aggressive healthcare reform since the introduction of Medicare in 1965, the possibility of leveling the playing field between group and individual insurance (through a variety of means) has come to the fore. Many have expressed concerns about the erosion of employment-based coverage, but researchers have not systematically examined, let alone quantified, the benefits associated with individual choice of insurance. In this paper, we use a

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<sup>†</sup> To comment on this article in the online discussion forum, or to view additional materials, visit the article page at <http://dx.doi.org/10.1257/pol.5.1.32>.

<sup>1</sup> Using data from the March 2010 Current Population Survey, Fronstin (2010) reports 59 percent of the non-elderly population had employment-based health benefits in 2009. The analogous figure for 2000 is 68.4 percent (Fronstin 2009).

large panel dataset on employer offerings and employee choices to infer the gains consumers would enjoy were they able to select from a broader spectrum of plans in their local market, holding constant employers' spending on employee subsidies and the full tax-deductibility of premiums. By quantifying the gain to individuals from being able to select any plan available in their local market, we back out the amount by which prices would have to increase to fully offset this gain. In so doing, we provide policymakers with guidance regarding the implementation and design of reforms that bolster individual choice. In a companion paper (Dafny, Ho, and Varela 2010), we examine the distributional consequences of expanded choice and contrast the characteristics of plans selected by employers and those that would be selected by employees if they were available on the same terms.

We use a unique dataset of employer plan offerings and employee plan selections for a national sample of 800+ large US employers during the period 1998–2006, representing over ten million insured lives in every year. Our approach consists of three distinct components. First, we estimate a discrete choice model of employee demand for health plans, conditioning on the set of plans offered by the relevant employer in the relevant geographic market and year. The parameters from this model reflect the values placed by employees on individual plan characteristics. Second, we estimate a hedonic model of premiums that permits us to predict the premiums a given employee would face for each plan offered in her local market. Third, we use the demand estimates, together with the predicted premiums, to predict employee choices of plans and their expected utility when offered additional plans currently existent in that market and year. The counterfactuals are budget-neutral for employers; that is, their total contributions to health insurance are held constant. Conceptually, the counterfactual is akin to granting employees a voucher equal to their employer's present contribution to health insurance, valid for the purchase of insurance plans on the individual market (which could be a regulated "exchange"). We use the results to estimate the amount by which premiums would need to increase (relative to the levels predicted by our hedonic model, which implicitly assumes group-based pricing due to the underlying data) to fully offset the net gain in consumer surplus.

We find choice is worth quite a bit for most individuals: in our most conservative hypothetical scenario the median employee would enjoy a surplus gain of roughly 13 percent of combined employer and employee premium contributions. In year 2000 dollars, this gain is approximately \$310 per individual or \$1,240 for a family of four. Combining these figures with data on employer subsidies, we find the median employee would be willing to forego 16 percent of her employer subsidy simply for the right to use what remains toward a plan of her choosing.<sup>2</sup> (As an analogy, consider the employer who offers her employee a choice of heavily subsidized vehicles: the Ford Focus or the Cadillac Escalade. The employee would trade a non-trivial percent of the employer subsidy in exchange for the freedom to use the subsidy toward her most-preferred vehicle, assuming it is available at the same price as currently paid by employers who buy in bulk.) Of course, we do not anticipate

<sup>2</sup>This estimate is obtained from our preferred specification, described below; estimates from other models are also presented.

the net benefits of choice to be as large as our estimates, since premiums are likely to increase with the devolution of insurance to the individual marketplace, and there are costs associated with the proliferation of choice. However, a complete accounting of an exchange proposal should include some estimate of the value of choice.

We caution that our results provide a conservative estimate of this value (or, equivalently, a low estimate of the premium increases that would offset the benefits of choice). First, our data enable us to build a very rich logit model of choice for a given set of employees, but we do not incorporate differences in preferences across individuals *within* employee groups except through a random error term. There may be substantial gains from better matching of individuals to plans. Second, due to the well-known limitations of the logit choice model, we do not expand the choice set to include *all* plans we observe in a given market, except to provide an upper-bound estimate of the value of choice. In fact, the conservative scenario mentioned above holds constant the number of plans in the choice set and simply switches the observed options with those that are most preferred by employees in the relevant firm and market.

The paper proceeds in seven sections. Section I discusses the recent trends in the degree of choice in employer-sponsored plans and summarizes related research. Section II describes the data. Section III presents the estimation strategy and results for the demand and hedonic models used as inputs into the simulations presented in Section IV. Sections V and VI discuss the implications of the results and some limitations, and Section VII concludes.

## I. Background

### *A. Employer-Sponsored Insurance Plans: How Much Choice Is There and How Is This Changing?*

Most workers who receive insurance through their employers have a choice of plans but this choice can be quite limited. The Kaiser Family Foundation/Health Research and Educational Trust Employer Health Benefits Annual Survey (hereafter Kaiser/HRETsurvey) studies the percentage of workers with job-based coverage, the kinds of plans employees offer and the choices made. Approximately 2,000 randomly selected employers are surveyed, covering a range of industries and both public and private firms. The survey indicates that 60 percent of firms and 98 percent of firms with over 200 workers offered health benefits in 2005. As mentioned earlier, 80 percent of firms sponsoring insurance offered a single plan. However, **Figure 1** shows that larger firms offered more choice than smaller firms: 27 percent of large firms (those with 1,000–4,999 workers) and only 17 percent of firms with over 5,000 workers offered a single plan. Overall, 63 percent of covered workers could choose from multiple health plans.

The most common healthplan offered to workers in 2005 was a PPO plan: 82 percent of covered workers had access to this type of plan. **Figure 2** documents that 28 percent of covered workers had access to a POS plan, 44 percent had access to an HMO and only 12 percent had access to a conventional indemnity plan. Indemnity plans have become less widely available over time while the availability of PPO plans has increased dramatically since 1988. The patterns in the dataset we

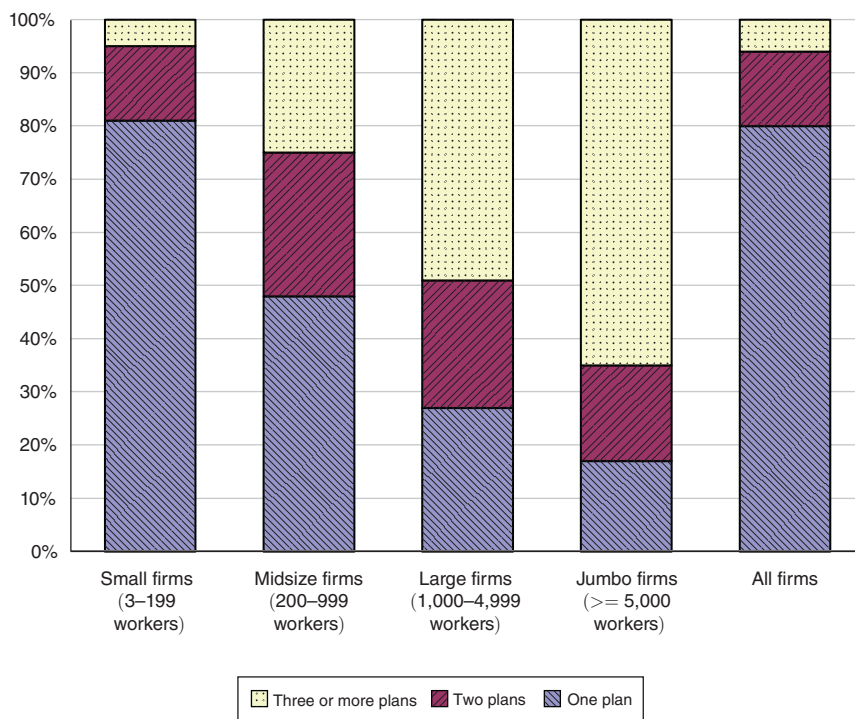


FIGURE 1. DISTRIBUTION OF FIRMS PROVIDING A CHOICE OF HEALTH PLANS, 2005

Source: Kaiser/HRET Survey of Employer-Sponsored Health Benefits

use are similar to those in the survey, although our sample is skewed towards larger firms so that choice is less limited than is the case for the average employee. For example, in 2005, about half of the employee groups in our sample are offered a single option. The choice sets observed in our data are discussed further in Section III.

### B. How Do Employees and Employers Choose among Plans?

Several studies in the health economics and health policy literatures investigate the factors influencing employees' choice of health plans. A much smaller set of papers examine employer decision making, with an emphasis on whether healthplan quality affects employer choices. To our knowledge, no study combines empirical analysis of both decisions, preventing any quantitative assessment of the tradeoffs associated with allocating decision rights to one or the other party. In the review that follows, we focus exclusively on studies that pertain to the working population, as our data includes only active employees.

*Employee Choice of Health Plans.*—Most studies in this category focus on the sensitivity of employees to variations in plan price and quality, as measured by items included in the Health Plans Employer Data and Information Set (HEDIS) and the

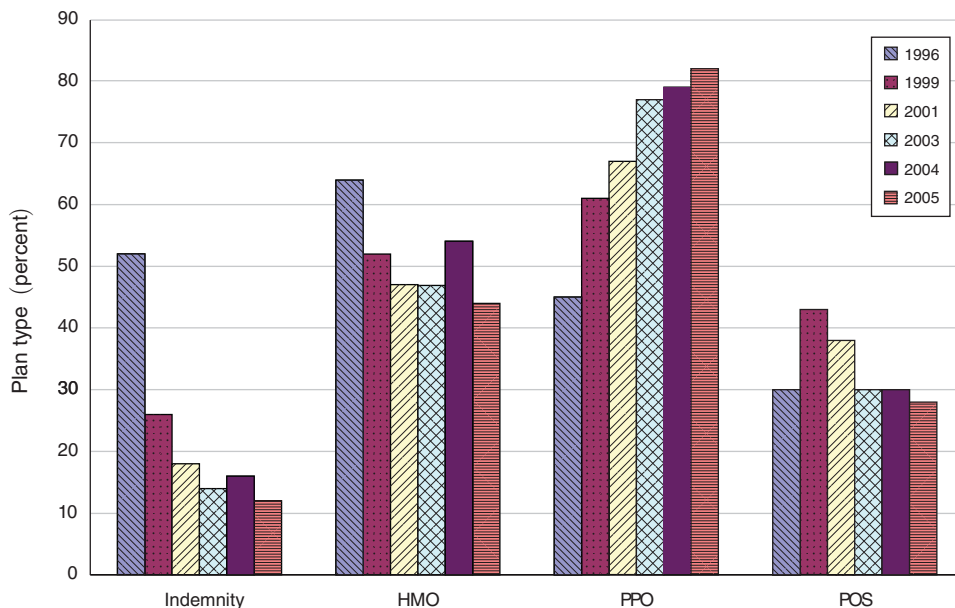


FIGURE 2. PERCENTAGE OF COVERED WORKERS WITH A CHOICE OF INDEMNITY, HMO, PPO OR POS PLANS, 1996–2005

Consumer Assessment of Healthcare Providers and Systems (CAHPS) survey.<sup>3</sup> Elasticities are reported for various definitions of price (or premium): with/without employer contributions; pre- or post-tax. Here we discuss estimates of the elasticity of within-employer-group enrollment with respect to pre-tax employee contributions, which corresponds to our model below.

Two studies identify these elasticities using the plan choices of university employees following implementation of a fixed-contribution arrangement, under which employees bear the incremental cost of their plan choices. Royalty and Solomon (1999) report estimates between  $-0.28$  and  $-0.62$  for Stanford employees; this compares closely with Cutler and Reber's (1998) estimates for Harvard employees ( $-0.30$  to  $-0.60$ ).<sup>4</sup> Two recent studies give more divergent estimates. Using data on plan choices of employees in eleven small and midsize firms in the Western United States during 2004–2005, Levin, Bundorf, and Mahoney (forthcoming) find elasticities in the upper tail of the range cited above (approximately  $-0.57$ , per our calculations from reported semi-elasticities). In contrast, Carlin and Town (2009) report an elasticity of demand around  $-0.06$ , using an autoregressive, multinomial probit choice model estimated on data from a large, self-insured employer between 2002–2005. Levin, Bundorf, and Mahoney (forthcoming) observe this lower estimate may be due to the well-known fact that elasticities are higher for initial plan

<sup>3</sup>This survey primarily addresses consumer satisfaction, and is maintained by the governmental Agency for Healthcare Research and Quality (AHRQ).

<sup>4</sup>Royalty and Solomon also report much higher estimates (exceeding  $-1$ ) for models with individual fixed effects, however in this case identification stems from post-implementation price changes, the sources of which are unclear. The authors refer to  $-0.55$  as their average estimate from "preferred" specifications.

choice (the key decision in Levin, Bundorf, and Mahoney forthcoming), whereas Carlin and Town identify the response to price through plan switchers. Significantly, Carlin and Town report somewhat higher elasticities from multinomial logit specifications, the model most commonly utilized in this literature. Our own estimates (from a multinomial logit model) vary by industry and other employee characteristics; across the entire sample, however, our estimate is  $-0.28$ . As with Carlin and Town, identification relies in part (although not exclusively) on plan switchers.

The relevant studies that consider the sensitivity of employee decisions to health plan quality include Wedig and Tai-Seale (2002) on federal employees; Beaulieu (2002) on Harvard employees; and Scanlon (2002) and Chernew, Gowrisankaran, and Scanlon (2008) on General Motors employees. Generally speaking, these studies find modest reactions to quality information. It is possible these aggregate effects mask larger responses by populations with stronger incentives to respond, however, the evidence to date on this matter is mixed.<sup>5</sup>

*Employer Choice of Health Plans.*—Research on how employers make decisions regarding which plans to offer, and how many, is limited by comparison. We focus here on empirical analyses of plan offerings, as opposed to analyses of surveys that ask employers to report what factors affect their decisions (e.g., Rosenthal et al. 2007). The most relevant papers for our purposes include Bundorf (2002) and Chernew (2004). These papers focus on whether employers' decisions reflect the assumed needs of their employees. For example, Chernew (2004) uses data on the HMO plans offered by 17 large employers in 2000 to see whether CAHPS scores affect the propensity any given plan is offered; they find that employers are more likely to offer plans with strong absolute and relative CAHPS performance measures. In related work, Bundorf (2002) finds employers' offerings correlate with employee characteristics. For example, firms whose employees have greater variation in healthcare expenditures are more likely to offer a choice of plans.

Our project builds on this research by quantifying—in dollars—the loss to consumers associated with restricted choice, and comparing these estimated losses to premium increases likely to occur if employees are free to apply their employer subsidies to other plans offered in their marketplace.

## II. Data

We use a proprietary panel database on health plans offered by a sample of large, multi-site employers from 1998–2006. The dataset, which we call the “Large Employer Health Insurance Dataset” (LEHID), was provided by a major benefits consulting firm which assists employers with designing or purchasing benefits from

<sup>5</sup>Scanlon (2002) finds new hires and plan switchers are more responsive to quality measures as well as price. Using the same study population, Chernew, Gowrisankaran, and Scanlon (2008) report “no significant evidence of heterogeneity” in the valuation of plan attributes based on observable or unobservable employee characteristics. Evidence from a different population—namely Medicare enrollees—is mixed as well. Using enrollment data surrounding the release of Medicare HMO report cards in 2000 and 2001, Dafny and Dranove (2008) find no differences in responses by demographic characteristics at the county level, but stronger evidence of non-report-card-related learning about quality (“market-based learning”) in counties with greater HMO penetration, more private report card data, and more stable populations.



health insurers.<sup>6</sup> The unit of observation is the plan-year. A plan is defined as a unique combination of an employer, geographic market, insurance carrier and plan “type” (HMO, POS, PPO, and indemnity), e.g., Company X’s Chicago-area Aetna HMO. The full dataset contains information from 813 employers and 139 geographic markets in the United States. The markets are defined by the data source and typically delineate metropolitan areas and ex-metropolitan areas within the same state, e.g., Arkansas—Little Rock and Arkansas—except Little Rock.<sup>7</sup> The number of enrollees covered in the data averages 4.7 million per year. Given an average family size above 2, this implies more than 10 million Americans are represented in the sample in a typical year. After excluding observations with missing or problematic data<sup>8</sup>, the sample contains 811 employers, 139 markets and 356 carriers. Most employers are active in a large number of markets (45 for the median employer-year). Descriptive statistics are set out in Table 1. For additional details of the data, see Dafny (2010).

*Premium* is the average annual charge, normalized to year 2000 dollars using the CPI, per person-equivalent covered by a plan.<sup>9</sup> It combines employer and employee contributions. The definition of premium depends on whether a plan is self-insured or fully insured. Many large employers choose to self-insure, outsourcing benefits management and claims administration but paying realized costs of care. Such employers can spread risk across large pools of enrollees, and often purchase stop-loss insurance to limit their exposure. Per ERISA (the Employee Retirement Act of 1974), these plans are also exempt from state regulations and state insurance premium taxes, enabling firms to reduce their insurance costs and/or standardize plan benefits across multiple sites. Reported self-insured plan “premiums” are actually estimates of employers’ projected healthcare expenditures, including any administrative fees and stop-loss premiums.<sup>10</sup>

Demographic factor is a measure that captures the family size, age, and gender of enrollees in a given plan-year. It can be construed as the mean number of “person equivalents” per enrollee. Plan design captures the generosity of benefits for a particular plan-year, including the level of copayments required of enrollees. Both factors are calculated by the source, and the formulae were not disclosed to us.

Our empirical analyses use the employer-market-year as the unit of observation. If an employer appears in the sample in a given year, all health plans it offers in any market are included in the data. However, the panel is unbalanced: on average, 240 employers appear in the sample each year. Of the unique employer-market pairs in the data, 46 percent appear only once, and 17 percent appear twice. We do not

<sup>6</sup>Using a survey of 21,545 private employers, Marquis and Long (2000) find that external consultants were employed by nearly half of the smallest firms (<25 workers) and nearly two-thirds of the largest firms (>500 workers). This suggests that the results of our study will be generalizable beyond our specific sample.

<sup>7</sup>Dafny (2010) includes a map of the geographic markets, which occasionally span state lines.

<sup>8</sup>We drop 347 observations with a missing industry code, 2752 observations associated with employer-market-years in which the employee share of premiums for one or more plans is negative, and 304 observations with missing data. We also consolidate the four plans that appear twice in the data because the employer self-insures some enrollees and fully insures others.

<sup>9</sup>The original data reports the average premium per enrollee. Thus, this average premium is larger for employee groups whose enrollees cover more dependents. We follow the practice of our data source and divide this figure by demographic factor to obtain the premium “per effective enrollee.”

<sup>10</sup>This definition of premiums for self-insured plans is common to all employer surveys, including the KFF/HRET survey described in Section II.

TABLE 1—DESCRIPTIVE STATISTICS

Premium (\$)	2,436 (704)
Employee contribution (percent)	0.212 (0.122)
Enrollment	175 (601)
Demographic factor	2.243 (0.449)
Plan design	1.038 (0.087)
Self-insured (percent)	0.682
Employer-markets in the panel	
Appear in every year	3%
Appear once	46%
Appear twice	17%
Appear thrice	13%
Other	22%
Total number of EM's	39,633
Energy production/transmission	2.9
Number of employers	811
Number of markets	139
Number of carriers	356
Number of EMY	115,440
Observations (EMCPY)	237,253
Plan type (percent)	
HMO	37.8
POS	35.3
PPO	16.5
Indemnity	10.3
Industry (percent)	
Manufacturing	13.4
Retail	12.0
Financial	11.4
Technology	9.4
Consumer products	7.3
Pharmaceuticals	6.4
Insurance	6.3
Transportation	5.6
Telecommunications	5.1
Entertainment and hospitality	5.1
Professional services	3.5
Health care	3.4
Energy production/transmission	2.9
Printing and publishing	2.8
Utilities (gas and electric)	2.4
Chemicals	1.7
Government/education	0.8
Unclassified	0.3

*Note:* Industry breakdown percentages obtained using the employer-market-year as the unit of observation, as this is the unit of interest for the choice models.



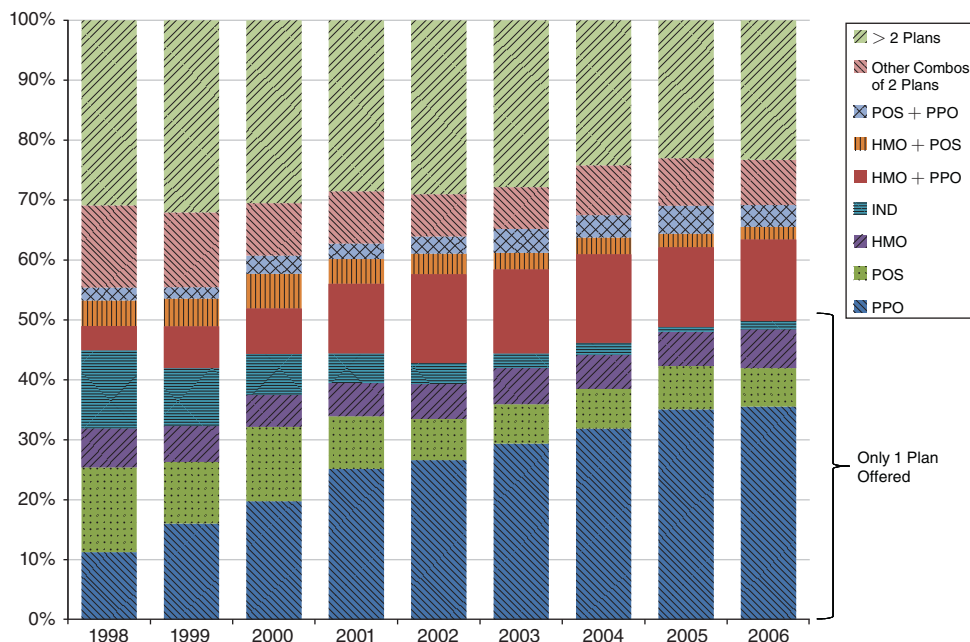


FIGURE 3. OBSERVED CHOICE SETS IN LEHID

observe the total number of employees offered insurance, hence our analyses are limited to employees who do take up coverage. As additional employees may take up coverage if more options are available, this likely understates the total gains from expanding choice.<sup>11</sup>

Before moving to the empirical analysis, we present statistics on the state and evolution of choice within our sample and study period (**Figure 3**). As expected, choice is more common among employers in our sample than among the universe of employers sampled by the Kaiser/HRET survey, however nearly half of the employer-market-years offer access to only one plan. Over 75 percent offer at most two plans. Fifty percent of those offering two plans offer an HMO and a PPO; 14 percent offer a POS plan and a PPO. The figure also shows that the amount of choice offered has fallen over time and (consistent with the survey evidence) that PPO plans have increased in popularity while indemnity plans have become less popular.

<sup>11</sup> To estimate the share of employees who do not take up employer-sponsored insurance, we matched our data to total employment figures reported in the Compustat Financial Database. However, Compustat is limited to large, publicly-traded firms, substantially reducing our sample size. In addition, the employment figures are very noisy, particularly as some firms report employment for North America rather than for the United States. The implied mean enrollment rate across employer-years was 46 percent, much lower than the 67 percent reported by the Kaiser-HRET survey for large firms (200+ workers) offering health benefits in 2005. We concluded that the analyses using this matched data sample are less informative than those utilizing the entire sample, notwithstanding the loss of an “outside option” in our choice model.

### III. Empirical Strategy

We conduct our analysis in three steps. First we use our data on consumer choices of health plans conditional on the options offered by their employers to estimate a utility equation describing employee preferences for plan characteristics. Second, we estimate a hedonic equation that describes the relationship between the premiums we observe in the data and plan, employer, and market characteristics. We use the coefficient estimates from this equation to predict the combined employer and employee premiums that employees in a given firm, market and year would face for every plan offered in their market and year, assuming large-group pricing prevails. Last, we use the results of both analyses to predict employee choices and expected utility under our counterfactual scenarios in which additional plans are made available on the same terms (i.e., a fixed percentage subsidy for a given set of employees, group rates and full tax-deductibility).

Although we are interested in the effect of expanding consumers' choice sets to encompass all possible options, the structure of our utility equation (which includes a logit error term with unbounded support) implies that adding all available plans to the choice set would overestimate the welfare gains of choice. We therefore investigate three counterfactual scenarios. First, we maintain the same number of plans in the choice set for each employer-market-year, but we substitute the most preferred plans for those currently offered (that is, if the employer does not choose optimally for its employees). We call this the "plan swapping" scenario. Second, we assume that employees within each employer-market-year triple gain access to their preferred option within each of three plan types: HMO, POS, and PPO plans.<sup>12</sup> (We exclude indemnity plans because they are rarely offered in our data. Employers already offering indemnity plans receive their most preferred indemnity plan in the counterfactual to ensure a strictly expanded choice set.) We call this the "all plan types" scenario. Third, we make all plans in the market-year available to all employees (the "all plans" scenario). The changes in consumer surplus predicted by the "plan swapping" and "all plans" scenarios provide lower and upper bound estimates, respectively, of the value of greater choice, with the "all plan types" scenario falling in between.

#### A. Demand Model

The first step is to estimate a model of consumer demand for health plans. We use a logit model, including in the consumer's choice set only the plans that are offered by the relevant employer in the relevant market and year. We denote a "plan" as a unique employer-market-carrier-plan-type-year quintuple, the unit of observation for our data. Consumer  $i$ 's utility from plan  $emcjt$  in year  $t$  is modeled as

$$(1) \quad u_{imcjt} = \delta_{emcjt} + \varepsilon_{imcjt},$$

<sup>12</sup>If an employer previously offered more than one option within a given plan type, we retain the same number of options within that plan type in the simulation.

where  $\delta_{emcjt} = x_{emcjt}\beta_{emt} - \alpha_{emt}p_{emcjt} + \xi_{emcjt}$ : a linear combination of observed characteristics of the plan (denoted  $x$ ), premium ( $p$ ), and an unobserved quality variable ( $\xi$ ). The coefficients on plan characteristics and premium are permitted to vary across employee groups (described in detail below). The term  $\varepsilon_{imcjt}$  is consumer  $i$ 's idiosyncratic preference for carrier  $c$  and plan type  $j$  in market  $m$  at time  $t$ .

Before discussing the details of our estimation, we offer remarks on our use of a simple logit model rather than a nested logit or random coefficients model. The most intuitive nested logit model, in which the first nest is the choice of plan type (such as HMO or non-HMO) and the second is the choice of plans within type, requires eliminating most of the data because choice sets typically contain at most one of each plan type. A random coefficients model, though feasible, is unnecessary as our data afford us the opportunity to explicitly model heterogeneity in preferences through a large set of fixed effects and interactions between plan characteristics and observable characteristics of the relevant employee population. These terms, which we discuss in detail below, permit the coefficients on the key explanatory variables to differ across observably different groups of consumers.

Berry (1994) shows that the parameters in equation (1) can be estimated using the following linear equation, which explicitly lists all covariates:

$$\begin{aligned}
 (2) \ln(s_{emcjt}) - \ln(s_{em0t}) &= \alpha + \xi_e + \nu_m + \psi_c + \eta_j + \delta_t + \varsigma_{em} + \omega_{mc} \\
 &+ \varphi_{mt} + \chi_{mj} + \sum_i \lambda_{ijt} I(\text{industry}_e = i) \\
 &+ \alpha_1 p_{emcjt} + \alpha_2 p_{emcjt} \times \text{demographic factor}_{emcjt} \\
 &+ \sum \alpha_{3i} I(\text{industry} = i) \times p_{emcjt} \\
 &+ \sum_i \alpha_{4i} I(\text{industry} = i) \times \text{demographic factor}_{emcjt} \\
 &\times p_{emcjt} + \psi \text{plan design}_{emcjt} + \sum_i \mu_i I(\text{industry} = i) \\
 &\times \text{plan design}_{emcjt} + \pi \text{self-insured}_{emcjt} + \xi_{emcjt}.
 \end{aligned}$$

In equation (2),  $s_{emcjt}$  is the market share of plan  $emcjt$  and  $s_{em0t}$  is the market share of the outside option in the relevant employer-market-year triple. We define the “outside option” to be the most frequently offered plan in the employer-market-year triple, which implies normalizing its unobserved quality to zero. Other plans’ observed characteristics are measured relative to those of this baseline plan.<sup>13</sup> For robustness, we also report results obtained when the outside option is the least generous plan in the relevant employer-market-year triple.<sup>14</sup>

<sup>13</sup>In the case of a tie the most frequently offered plan is designated as the plan with the largest number of enrollees in the market-year.

<sup>14</sup>The least generous plan is defined using plan type and premium. Indemnity plans are the most generous, followed by PPOs, POS plans and HMOs in that order. Within a particular plan type, the cheaper of a pair of plans is defined as the less generous plan.

The covariates include several fixed effects, two continuous measures—plan design and the employee’s contribution to premiums (“price”)—and interactions including these two measures. We discuss each in turn.

The fixed effects include all of the “main effects,” that is, dummies for each employer, market, carrier, plan type, and year. However, the dummies for employer, market, and year are differenced out when we normalize the characteristics of each plan with respect to those of the baseline plan, which implies we obtain coefficient estimates only on carrier and plan type dummies. Carriers and plan types with the largest coefficient estimates generate higher utility, *ceteris paribus*, for enrollees.

It is also possible to include second-, third-, and fourth-order fixed effects. Such terms have the advantage of enabling a better fit of the model, but there are four important disadvantages. First, they absorb variation in continuous regressors of interest such as price, leaving little to identify the coefficients on these measures. Second, many of these terms cannot be included in the counterfactual scenarios. For example, employer-carrier fixed effects would capture the mean utility of different carriers to employees of a particular firm. In a counterfactual that expands the choice set to include carriers not presently offered by that firm, it would not be possible to estimate the utility of the new options. Third, even if a coefficient could technically be estimated for a third or fourth-order interaction term, the number of observations identifying it would be small and therefore unlikely to yield a representative estimate valid for counterfactual simulations. Last, some of these terms raise endogeneity concerns. For example, employer-year interactions would capture the fixed utility associated with the set of plans offered by an employer in a given year, but this depends on choices currently on offer, and would presumably change when the choice set changes.

In recognition of these issues, we include a parsimonious set of second-order interaction terms that control for the most important unobservable correlates of utility while permitting estimation of our counterfactual scenarios.<sup>15</sup> Two of five interactions we include will be “differenced out” in our specifications: employer-market fixed effects and market-year fixed effects. We mention them here to clarify the sources of identification for coefficients in the demand model, and because these terms appear (and are not differenced out) in the hedonic premium model we discuss below. Conceptually, employer-market interactions absorb time-invariant differences across specific sets of employees. For example they capture the fact that employees of a firm in some markets are particularly well-educated, have a particularly high income or are particularly risk averse and therefore place a high value on health insurance. They also absorb any fixed variation in price for a set of enrollees that may be correlated with time-invariant differences in risk profiles and demographics. The market-year fixed effects pick up market-specific shocks to utility such as a reduction in provider quality due to the closure of a hospital. Importantly, their inclusion implies that changes in market-level prices for plans do not identify

<sup>15</sup>Of the ten possible second-order interaction terms, we include five for reasons detailed in the text that follows. We exclude employer-year interactions due to endogeneity concerns, and employer-carrier and employer-plan type interactions because these are not compatible with our counterfactual simulations. Finally, carrier-plan type and carrier-year interactions are unlikely to be important determinants of unobserved quality. Indeed, we find our price elasticities are unaffected by excluding these terms.

the price coefficient. Rather, identification relies on changes in the relative prices of plans offered to employees in a given market. Because we also include plan type-year fixed effects, relative price changes associated with national plan type-specific trends also do not identify the price coefficient.<sup>16</sup> In terms of the utility model, the plan type-year interactions capture changes in consumer preferences (such as the decline in popularity of HMOs in the 1990's) and in plan management (such as HMOs' decision to engage in less utilization review) over time. By interacting the plan type-year terms with industry dummies, we capture heterogeneity in preferences across employer groups.

We include market-carrier fixed effects to capture the "fixed utility" associated with each market-carrier combination. For example, Blue Cross Blue Shield of Illinois may be an especially attractive Blue Cross Blue Shield carrier because it has a very large network of hospitals. This interaction is therefore important to capture unobserved plan quality. Finally, we include market-plan type interactions to capture differences across markets in the utility associated with particular plan types. For example, HMOs are more highly-valued in areas where they have a long history and this may be important for demand.

In addition to these fixed effects, we include two continuous measures: plan design and the employee's contribution to the annual premium (hereafter "price"), denoted  $p_{emcjt}$ . We interact both with dummies for industry categories to incorporate potentially different valuations of these characteristics by employee populations in different industries. We also include interactions between our measure of family size (demographic factor) and price. Finally we interact both price and the price-demographic factor interaction with industry category dummies. This functional form exploits the richness of the dataset, allowing, for example, Firm X's employees in Industry Y to have less price sensitivity than Firm Z's employees in Industry Y due to their larger hypothetical family size.

Our model takes price to be exogenous to unobserved plan quality, conditional on the many covariates included. The rich set of fixed effects and interaction terms we include mitigates concerns about endogeneity, specifically that price will be positively correlated with unobserved quality, yielding a downward-biased coefficient estimate. For example, unobserved quality of a particular carrier is absorbed in the carrier fixed effects and the market-carrier interactions. Unobserved differences in quality across types of plan are absorbed in the plan type variable, the market-plan type interactions and the plan type-year-industry category interactions. We considered several instruments, including for example the average price of plans offered by the same employer-year in different markets and different plan types, but conditional on all of the fixed effects in our model there is insufficient variation in these potential instruments to predict the remaining variation in price. As noted earlier, our estimates of price elasticity fall in the lower end of the range of estimates from other studies of healthplan choice.

<sup>16</sup>The industry categories are: chemicals, consumer products, energy, entertainment and hospitality, financial services, government and education, health care, insurance, manufacturing, pharmaceuticals, printing and publishing, professional services, retail, technology, telecommunications, transportation, utilities, and unclassified.

TABLE 2—DEMAND ESTIMATES  
(Reported for manufacturing industry)

	MF model	LG model
Price (\$000's)	1.351 (0.394)	0.313 (0.387)
Price × demographic factor	-1.121 (0.164)	-0.727 (0.161)
Plan design	1.384 (0.209)	1.947 (0.212)
Self-insured	0.270 (0.010)	0.279 (0.010)
Mean demographic factor for manufacturing	2.52	2.52
Implied price coefficient for manufacturing	-1.47	-1.52
Adjusted $R^2$	0.320	0.330

Notes: "LG model" uses the least-generous plan within each employer-market-year as the outside option, while "MF model" uses the most-frequently offered plan within a market-year as the outside option. Estimates in this table are for the manufacturing sector, as price, price × demographic factor, and plan design are interacted with industry dummies. All implied price coefficients are statistically different from zero at  $p \leq 0.05$ .  $N = 237,253$ .

Last, we include an indicator for whether plan  $j$  is self-insured (SI). Although enrollees are unlikely to know whether a plan is self-insured, to the extent that self-insurance is correlated with unobserved healthplan attributes, it may be an important determinant of utility. A priori the sign of the coefficient estimate is ambiguous. On the one hand, SI plans could be less appealing than observably identical fully insured (FI) plans because SI plans need not cover state-mandated benefits—although these differences should be captured in plan design. On the other hand, according to our source there are two important unobserved benefits of SI plans. First, such plans often have priority or dedicated customer service lines for handling member calls and resolving issues promptly. Second, plan administrators may be more lax with utilization review, as their incentives to minimize outlays are muted or nonexistent. As we discuss in the following section, we find there is positive utility associated with self-insurance—utility which will not be accessible in our simulations of individual choice. Importantly, including the self-insurance indicator does not affect estimates of other parameters in the utility equation, suggesting it truly captures unobserved quality and is not correlated with included variables.

*Demand Results.*—The demand estimates are summarized in Table 2. Columns 1 and 2 display results for models using the least generous plan ("LG model") and the most frequent plan ("MF model") as the outside options, respectively. The coefficients for price, the price-demographic factor interaction, and plan design differ across industries because all three are interacted with industry category dummies. In Table 2, we display the estimates for the manufacturing industry, the largest in the data; price elasticities for other industries are displayed in Table 3.

We begin by observing that estimates of the parameters of interest are similar in both models. The interaction of price with the demographic factor makes the price coefficients difficult to interpret from the simplest table of results. The mean

TABLE 3—AVERAGE IMPLIED PRICE COEFFICIENTS AND ELASTICITIES FOR SELECTED INDUSTRIES

	MF model (1)	LG model (2)
Price coefficients		
Manufacturing	-1.47	-1.52
Financial	-0.74	-0.73
Retail	-1.55	-1.53
Technology	-1.04	-1.13
Consumer products	-1.38	-1.43
Telecommunications	-1.16	-1.13
Pharmaceuticals	-0.94	-0.64
Elasticities		
Manufacturing	-0.14	-0.15
Financial	-0.28	-0.28
Retail	-0.46	-0.45
Technology	-0.21	-0.23
Consumer products	-0.18	-0.19
Telecommunications	-0.08	-0.08
Pharmaceuticals	-0.13	-0.09
Percent of employee-market-years with positive implied price coefficients	1.5	1.6
Number of employer-market-years	115,440	115,440

Notes: "LG model" uses the least-generous plan within each employer-market-year as the outside option, while "MF model" uses the most-frequently offered plan within a market-year as the outside option. All implied price coefficients are statistically different from zero at  $p < 0.05$ . Average elasticities across employer-market-year-plan combinations within each industry are reported. Elasticity is defined as price coefficient  $\times (1 - \text{share}) \times \text{price}$ .

demographic factor for the manufacturing industry, together with the implied price coefficient for each specification, is provided beneath the coefficient estimates.

Table 3 reports the implied price coefficients, together with estimated price elasticities, for the seven largest industries in the data. The price coefficients are negative and significant at  $p = 0.05$  for all industries and specifications. The elasticities in the LG model vary from  $-0.08$  in the telecommunications industry to  $-0.46$  in the retail industry; the estimates from the MF model are nearly identical. The weighted average estimate,  $-0.28$  falls comfortably in the range of prior estimates, though on the lower rather than the higher side of that range. Below, we present simulations using demand estimates from both models; these do not reveal meaningful differences in the distribution of estimated utility gains overall.

As expected, we find that plan design has a significant positive effect on utility. Although the coefficient estimate is larger in the MF model, given the magnitude of the standard deviation in plan design (0.09), this difference is not economically important. The coefficient on the SI indicator is positive and statistically significant: the coefficient of 0.27, divided by the weighted average price coefficient for all employees in the sample (1.13), implies that SI status generates utility equivalent to a decrease of \$239 in annual contributions (as compared to the weighted average mean premium of \$2,436). Though sizeable, this estimate is not implausibly large given the potential benefits of SI plans enumerated earlier. To confirm that this SI finding is not indicative of broad misspecification of the demand model, we re-estimated our models excluding the SI indicator. These estimates yielded very similar coefficients on all included



terms, implying SI status is orthogonal to included covariates. Although encouraging, we revisit this issue in the discussion of simulation results below.

### B. Hedonic Equation

We use a hedonic regression model to predict the price at which each plan will be made available to the population in a particular employer-market-year in our simulations. Simply using the average of observed premiums for each plan is undesirable because premiums vary with the composition of the relevant employee population. It is worth noting that we do not expect our estimates to approximate the price that would prevail on an “exchange” for individually-purchased plans; the reduction in group size implied by individual shopping may lead to a substantial price increase, a subject we address in Section V. Instead we use our predicted prices to estimate the consumer surplus increase from expanding choice, *ceteris paribus* (that is, with continued price-setting at the employer-market-year level). This model implicitly assumes that all buyers are treated similarly. For example if insurance carrier A’s HMO carries a 10 percent premium relative to insurance carrier B’s HMO then all aspiring enrollees will also face a 10 percent premium for this plan (they may also face a price increase or reduction due to the characteristics of their employer group and market).

Our model takes the following form:

$$\begin{aligned}
 (3) \quad \ln(\text{premium})_{emcjt} = & \alpha + \xi_e + \nu_m + \psi_c + \eta_j + \delta_t + \varsigma_{em} + \omega_{mc} \\
 & + \varphi_{mt} + \chi_{mj} + \kappa_{jt} + \eta \text{plan design}_{emcjt} \\
 & + \sum_i \varphi_i I(\text{industry} = i) \times \text{plan design}_{emcjt} \\
 & + \gamma \text{self-insured}_{emcjt} + \varepsilon_{emcjt}.
 \end{aligned}$$

We regress log premium per effective enrollee (combined employer and employee premium contributions) on plan design (interacted with industry dummies), a self-insurance indicator, and the same first and second-order fixed effects included in the utility equation.<sup>17</sup> We considered, but do not ultimately include, indicators for the number of plans offered to each employee group. Insurers reportedly price differently for “slice business,” in which their products compete with plans offered by competitors, both due to adverse (or favorable) selection within an employee group and to reduced economies of scale. In practice, these indicators were superfluous to the model, which already incorporates employer-market fixed effects. In such a model, these indicators will be identified solely off changes in the number of plans offered by a given employer-market over time, controlling for market-specific trends in this number. In addition, we would not anticipate a slice pricing effect

<sup>17</sup>Note that, compared to the utility model, the hedonic model excludes interactions between industry dummies and the plan type  $\times$  year interactions; this omission is intended to reduce “overfitting” of the data, which could result in misleading predictions of premiums.

TABLE 4—HEDONIC RESULTS

	Panel A.	Panel B.			Panel C.			
	Residual	Predicted premiums			Grouped by		Grouped by	
	ratio	Uncensored	Censored	Observed	Actual	All	Actual	All
	(1)	(1)	(2)	data	choice set	plans	choice set	plans
				(3)	(1)	(2)	(3)	(4)
5%	-0.316	1,591	1,612	1,507	0.01	0.10	0.09	0.23
25%	-0.086	1,989	1,997	1,929	0.04	0.15	0.20	0.44
50%	0.000	2,373	2,372	2,339	0.08	0.19	0.28	0.49
75%	0.076	2,813	2,805	2,817	0.14	0.25	0.37	0.55
95%	0.198	3,719	3,651	3,726	0.29	0.35	0.48	0.67
Smallest	-315	209	645	182	0.00	0.00	0.00	0.00
Largest	141	13,580	8,323	13,980	0.88	0.75	0.80	1.10
Mean	-0.03	2,464	2,460	2,436	0.10	0.21	0.28	0.49
SD	1.08	665	632	704	0.09	0.08	0.12	0.12
Observations	237,253	1,676,268	1,676,268	237,253	62,387	62,387	18,768	18,768

Notes: Premiums are per effective enrollee. Panel A: Residual Ratio = Residuals/Actual Premiums. Panel B: Censored predicted premiums are censored at the 5 percent and 95 percent values within each market-year. Panel C: Span Ratio = (Largest Premium in group–Smallest premium in group)/Average Premium in Group. All figures in panel C use censored predicted premiums. E-M-Ys with only 1 plan are excluded from panel C, column 1. M-C-P-Ys offered by fewer than 3 E-M-Ys are only included in column 2 of panel C.

for employee groups where all plans on offer are self-insured, which accounts for 60 percent of all groups.

We anticipate a negative coefficient estimate on the self-insurance dummy. Self-insurance should be cheaper, *ceteris paribus*, because the employer bears some (or all) of the risk of medical expenditures, self-insured plans are exempt from state mandates and premium taxes, and employers fulfill some of the administrative functions that insurers perform for fully-insured plans (such as explaining benefit coverage to enrollees). Finally, we expect the employer-market interactions to be particularly important because they capture unobserved demographic information that is likely to affect health risk and therefore the cost of insurance.

*Hedonic Results.*—The results of the hedonic regression are summarized in Table 4. As a measure of the fit of our model, panel A describes the distribution of the ratios of the regression residuals to the actual premiums. The fit is good: the fifth percentile of this distribution is  $-0.32$  and the ninety-fifth percentile is  $0.20$ . That is, the smallest residuals are roughly  $-32$  percent of premiums and the largest residuals are roughly one-fifth of the corresponding premiums. The adjusted  $R^2$  of the regression is  $0.792$ .

The discussion thus far pertains to goodness of fit of the regression within sample. However, we are interested in predicting premiums out of sample; goodness of fit for this purpose is illustrated by panels B and C. Column 1 of panel B gives the distribution of predicted premiums for all (hypothetical and observed) employer-plan

combinations.<sup>18</sup> This distribution compares very favorably to the distribution of observed premiums, reported in column 3. For example, the mean predicted premium is \$2,464, compared to an observed mean premium of \$2,436. However, to ensure that our simulations are not overly sensitive to outliers, we take the extra precaution of censoring predicted premiums at the 5 percent tails before performing our simulations.<sup>19</sup> The distribution of censored premiums is given in column 2.

Our final summary of the predictions implied by the hedonic model is given in panel C, which presents the distribution of a statistic we term the “span ratio.” The span ratio equals the difference between the largest and smallest predicted premiums, divided by the mean predicted premium, for a given set of observations.<sup>20</sup> Columns 1 and 2 define this set as the employer-market-year, providing a snapshot of the range of premiums from which an employee can choose given their current set of options (column 1, which only includes employer-market-years in which more than one option is available) or could choose if all plans were made available (column 2). The median figure in column 2 is 19 percent, as compared to 8 percent in column 1, implying that in a market with full choice, employees would have a wider range of price points from which to choose. Of course, even the 8 percent figure overstates the current range of price points as only 55 percent of employer-market-years offer any choice at all. We note that “span” is defined for combined employer and employee premiums; the span of employee contributions may certainly differ.

Columns 3 and 4 of panel C also report span ratios calculated using the set of observed plans and “all plans,” respectively, but here the set of underlying observations is grouped by market-carrier-plan type-year. Thus, these columns illustrate the variation in premiums for, say, the Aetna POS plan in Chicago in 2003, due to employer-specific characteristics (apart from family size, age, and gender, which are already accounted for as premiums are reported per effective enrollee). The median span ratios are 28 percent (using actual plans on offer, and associated predicted premiums) and 49 percent (all plans, predicted premiums). The sizeable spans are not surprising: the risk profiles of employee populations are very different, and premiums are experience-rated for large groups. As expected, the span ratio in the “all plans” scenario is larger even within the same market-carrier-plan type and year, as we have expanded the range of employee groups for which each product is available.

In the interest of space we do not report the coefficient estimates from the hedonic model, but we note here that the sign of the coefficient estimate on the self-insured dummy is positive, contrary to expectations. Though statistically significant, the coefficient estimate of 0.005 is economically small: a self-insured plan typically costs 0.5 percent more than a fully-insured plan, *ceteris paribus*. Together with the estimates from the demand model, in which self-insured plans were found to be more attractive

<sup>18</sup>By construction, the number of observations is very large: the average market-year has 15 carrier-plan type combinations offered by at least 3 employers. Given there are 115,440 employer-market-year units the total number of observations exceeds 1.7 million.

<sup>19</sup>Premiums in the low and high tails are replaced by the fifth percentile and ninety-fifth percentile of premiums within the relevant market-year, respectively.

<sup>20</sup>We report all span ratios using predicted, rather than actual, premiums as our simulation results use predicted premiums to estimate both current and predicted utility under the various scenarios.

to consumers all else equal, this implies any cost savings associated with self-insurance may be passed on to employees in the form of higher quality.

### C. Simulations

The next step is to use the estimated coefficients from the demand models to predict employee choices and the resulting consumer surplus if employees are permitted to select among a wider set of health plans than that offered by their employers, and premiums for these plans are estimated using the hedonic model.

As noted earlier, because our utility equation includes a logit error term which has unbounded support, expanding the choice set to include all observed health plans in each market-year will overestimate the value to consumers of increased choice. Thus we also provide a conservative estimate (the “plan swapping” scenario that holds constant the number of choices available to each set of employees, but substitutes the most preferred plans for those currently offered) and an intermediate estimate, the “all plan types” scenario that includes access to the most preferred option within each of the HMO, POS, and PPO plan types. We note that both the “plan swapping” and the “all plan types” scenarios are subject to the statistical problem that we swap or add plans that are estimated to be the most preferred for employees within the relevant employer-market-year, so the estimated utility gains are likely to include positive estimation errors on average.<sup>21</sup> We conduct a robustness test in Section IV which demonstrates that the resulting bias is small.

We define a health insurance plan or “option” as an MCPY combination, for example United Healthcare’s Chicago-based PPO in 2005, and we exclude from the counterfactuals plans that are offered by fewer than three employers in the relevant market-year.<sup>22</sup> We also ensure that our estimates are conservative by excluding plans whose predicted average utility is below the fifth or above the ninety-fifth percentile of the estimated utility distribution.<sup>23</sup> Finally, we drop the small share of employer-market-years for which the estimated price coefficient is positive; the exact percentages are reported in Table 3 and vary depending on the demand specification.

To measure consumer surplus, we use the approach delineated by Nevo (2001) and based on McFadden (1981). Consumer  $i$ ’s expected gain from a change in the set of health plans available to him is:

$$(4) \quad \Delta_i = u_i^t - u_i^{t-1},$$

<sup>21</sup>This error is distinct from that associated with the logit error in the utility equation.

<sup>22</sup>These plans may be offered by carriers who are not truly active in a market, but who rent the networks of active participants in order to provide service in markets where they are not otherwise present. In the counterfactual simulations we allow employer-market-years where these small products are currently offered to keep them; that is, employees never lose MCPY combinations that are currently offered to them.

<sup>23</sup>Specifically, plans added in counterfactual scenarios may not fall in the 5 percent tails of the utility distribution for the relevant market and year. To construct this distribution, we calculate the weighted average utility for each plan across all employer-market-year observations. Any plans falling at either extreme of this distribution within the relevant market and year are not included in the choice set for any counterfactual, unless such plans were offered in the original choice set. This reduces the influence of outliers on our estimated surplus gains.

where  $u_i^t$  and  $u_i^{t-1}$  are defined by

$$(5) \quad u_i^t = E_e \max_j (\delta_{jt} + \varepsilon_{ijt}).$$

Note that this is the expected welfare gain from the perspective of the econometrician given the available data. We measure the change in consumer surplus from the hypothetical change in choice sets using the compensating variation (CV). McFadden (1981) shows that

$$(6) \quad CV_{it} = (u_i^t - u_i^{t-1})/\alpha,$$

where  $\alpha$  is the coefficient on price in the plan utility equation. Integrating analytically over the extreme value distribution of  $e$  implies that the compensating variation of consumers in employer  $e$ , market  $m$ , and year  $t$  is given by

$$(7) \quad CV_{emt} = (1/\alpha_{emt}) \left[ \ln \sum_{c,j \in J_{emt,cft}} \exp(\delta_{emcjt,cft}) - \ln \sum_{c,j \in J_{emt,obs}} \exp(\delta_{emcjt,obs}) \right],$$

where  $J_{emt,obs}$  and  $J_{emt,cft}$  are the choice sets available to employees of firm  $e$  in the observed and counterfactual scenarios respectively and  $\delta_{emcjt,obs}$  and  $\delta_{emcjt,cft}$  are the values predicted by the demand model.

The inputs to  $\delta_{emcjt,obs}$  and  $\delta_{emcjt,cft}$  are price, plan design, the fixed effects from the demand model and the unobserved quality  $\xi_{emcjt}$ . To calculate  $\delta_{emcjt,cft}$  for plans that were originally offered by more than one employer we use the weighted (by number of enrollees) mean plan design and the median unobserved quality ( $\xi_{emcjt}$ ) for the relevant market-carrier-plan type-year (MCPY). Using the median  $\xi_{emcjt}$  reduces the noise caused by particularly high or low estimated unobserved quality for particular employers, although in practice this has very little effect on our estimates.<sup>24</sup>

In all counterfactual scenarios we set the SI indicator to zero; that is, we assume the utility generated by SI status is no longer available to employees. The premiums in both the observed and the counterfactual scenarios are the values predicted by the hedonic regression described above. We incorporate employer subsidies to health insurance by assuming budget neutrality for employers: for every employer-market-year, spending equals actual employer contributions to premiums in the original data. Given most employers in our data appear to pay the same percentage of premium across plans (the median is 79 percent) rather than the same dollar amount, we retain this feature in our simulations. Thus, in each scenario we solve for the fixed percentage subsidy that yields budget neutrality, and apply this subsidy to plan premiums to determine the price faced by the employee.<sup>25</sup>

<sup>24</sup>There is one exception to these rules. If an employer offers a plan to its employees in the data, in the counterfactual those employees are offered exactly the same plan design, with the same unobserved quality, even if it is offered by fewer than three employers and even if there are multiple plans in the MCPY. This design results in strictly superior choice sets in the counterfactual scenarios.

<sup>25</sup>This requires us to solve for the fixed point of an equation, since the share of a given plan depends on how much it is subsidized, which in turn depends on how many consumers are choosing that plan.

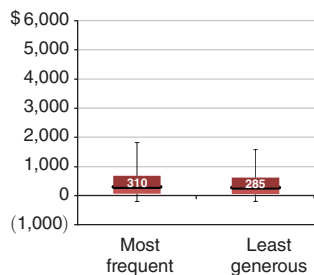


FIGURE 4A. UTILITY GAINS FROM CHOICE SET INCLUDING EMPLOYEES' PREFERRED PLAN(S), HOLDING THE NUMBER OF OPTIONS CONSTANT

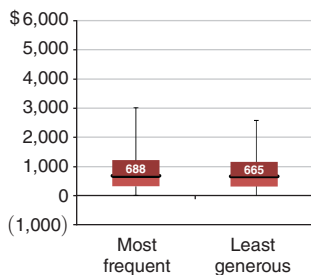


FIGURE 4B. UTILITY GAINS FROM CHOICE SET INCLUDING EMPLOYEES' PREFERRED PLANS, ENSURING AT LEAST 1 PPO, 1 HMO AND 1 POS

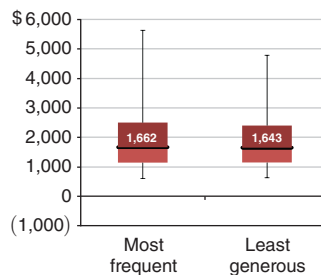


FIGURE 4C. UTILITY GAINS INCLUDING ALL POTENTIAL PLANS IN CHOICE SET

#### IV. Simulation Results

The three panels in Figure 4 summarize the utility gains from the three counterfactuals: plan swapping (4A), all plan types (4B), and all plans (4C). Each panel includes two boxplots that present the distribution of utility changes corresponding to the two distinct demand equations described earlier: MF, denoting the model that uses the most-frequently offered plan in a given market-year as the baseline option, and LG, denoting the model that uses the least-generous plan in a given employer-market-year as the baseline option. The boxes are bounded by the twenty-fifth and seventy-fifth percentiles of the relevant distribution, and the ends of the vertical lines define the fifth and ninety-fifth percentiles.

The results reveal sizeable gains from all scenarios, and the choice of the baseline plan matters little. As expected, the plan swapping scenario (Figure 4A) yields lower estimated gains than the all plan types scenario (Figure 4B): \$310 for the median covered person as compared to \$688, using the MF estimates. The gains from the all plans scenario (Figure 4C) are the highest at \$1,662 for the median covered person.

As a check for misspecification of the demand model, we considered an alternative approach, in which we excluded the SI indicator from the demand estimation and calculated welfare gains using the resulting estimates. We compared these predicted gains to those generated from the baseline demand model (including the SI variable) when, in contrast to the simulations presented above, we set the SI indicator to zero in the “observed” as well as the “counterfactual” scenario. The robustness test therefore compared predictions from simulations where SI status could not generate utility; the estimated gains would differ if including the SI indicator in the utility specification affected other parameter estimates. In fact, the estimates were very similar, implying that the SI indicator does not “soak up” utility generated by observables, biasing the estimates of the coefficients on other variables.

As noted above, the results reported for the “plan swapping” and “all plan types” scenarios may be biased upwards because we swap in the plans that are estimated rather than observed to be most preferred. We approximated the bias by taking 100 draws from the estimated parameter distribution and for each draw calculating the distribution of welfare gains from the plan-swapping scenario using the MF model.

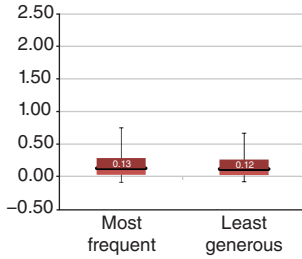


FIGURE 5A. PERCENT INCREASE IN PREMIUMS REQUIRED TO OFFSET GAINS FROM EMPLOYEES' PREFERRED PLAN(S), HOLDING THE NUMBER OF OPTIONS CONSTANT

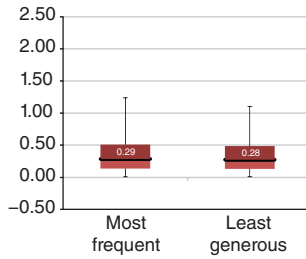


FIGURE 5B. PERCENT INCREASE IN PREMIUMS REQUIRED TO OFFSET GAINS FROM CHOICE SET OF EMPLOYEES' PREFERRED PLANS, ENSURING AT LEAST 1 PPO, 1 HMO AND 1 POS

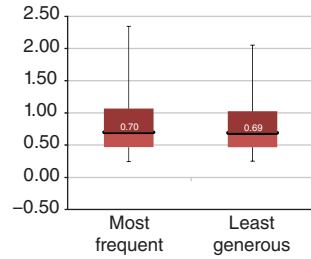


FIGURE 5C. PERCENT INCREASE IN PREMIUMS REQUIRED TO OFFSET GAINS WHEN INCLUDING ALL POTENTIAL PLANS IN CHOICE SET

The cross-draw average of the median welfare gain was \$321, quite close to the \$310 estimated in the baseline analysis. The cross-draw standard deviation of that median value was just \$10, and the cross-draw minimum and maximum were \$297 and \$344 respectively. We conclude that statistical bias on the estimated median values is small.

## V. Discussion

Our findings reveal that, on average, restricting employee choice yields substantial amounts of deadweight loss. This loss is due both to poor matching between employees' preferred plans and employers' offerings, holding the absolute number of choices constant, and to the reduced variety of plans that are offered. In this section we assess whether premium increases that might occur in an expanded-choice scenario are likely to fully offset projected surplus gains.

Our estimates of consumer surplus are calculated under the unrealistic assumption that individuals would enjoy group pricing when choice is expanded. For reasons we detail below, premiums are likely to rise if employer involvement in plan sponsorship is limited to a subsidy. We begin by presenting data on the amount by which premiums would have to increase to fully offset the gains from choice. We express this figure as a percentage of the average predicted premium for each employer-market year, and present boxplots of the resulting distribution in Figure 5. As in Figure 4, panel A corresponds to the "plan swapping" scenario, panel B to the "all plan types" scenario, and panel C to the "all plans" scenario.

As expected, the numbers reflect the gains reported in Figure 4. The median increase in premiums needed to offset surplus gains is roughly 13 percent, 29 percent, and 70 percent in the plan-swapping, all plan types, and all plans scenarios, respectively. Of course, to interpret these results we require projections regarding the likely premium increase when the choice set is expanded. We discuss both current differences in loading fees for individual/small group versus large group plans, and projected loading fees for plans to be offered on a hypothetical "individual exchange," as reported by organizations performing evaluations of recent healthcare reform proposals. As the loading fee represents the difference between dollars paid



in as premiums and dollars paid out to providers of healthcare services, we implicitly abstract away from absolute changes in medical spending that may result when the same set of individuals enrolls in different plans.

According to the National Health Expenditure Accounts, which produces estimates of private premiums and insurer outlays on an annual basis, loading fees increased from 10.5 percent of premiums in 1998 to 12.8 percent of premiums in 2006.<sup>26</sup> These figures include self- and fully-insured plans of all sizes. Other sources report similar aggregate estimates.<sup>27</sup> Loading fees can be divided into administrative and non-administrative components, although the categorization of expenses is somewhat arbitrary.<sup>28</sup> Non-administrative components include corporate taxes, profits, and additions to capital reserves. Notably, self-insured plans are exempt from premium taxes and solvency requirements, so non-administrative costs for individual and small group plans—which are fully insured—are clearly higher.

Administrative costs are also higher for small plans. According to a white paper by the American Society of Actuaries (2009), there are four key components to administrative costs of health plans: marketing (including broker commissions), provider and medical management (e.g., provider network management), account and member administration (includes billing, customer service, and claims processing), and corporate services (including underwriting and associated risk premiums). All but the second of these components will clearly be higher for small plans.

There are few sources that compare the loading fees for large group versus individual/small group plans, and none to our knowledge that separate these estimates by expense category. Before discussing available figures, we note three reasons the difference in current loading fees is likely to overstate premium increases in our hypothetical scenario: (i) the risk premium due to adverse selection in the current individual marketplace exceeds that which would likely prevail if all employer-sponsored enrollees were included in the pool of insured; (ii) as the pool grows, statutorily-required capital reserves should decline as a percent of premiums; (iii) state premium tax rates should also decline as the taxable base increases.

These caveats notwithstanding, the best available estimate of the difference in loading fees between the smallest groups (fewer than 100 employees) and the largest (>1,000 employees) is 10 percent of premiums (Karaca-Mandic, Abraham, Phelps).<sup>29</sup> A 2006 study by the Council for Affordable Health Insurance reports a

<sup>26</sup> Source: Table 12, <http://www.cms.hhs.gov/NationalHealthExpendData/downloads/tables.pdf>.

<sup>27</sup> The Congressional Budget office puts the figure at 11 percent. The Sherlock Company, a health care financing firm, reports the median BCBS plan spends 10.4 percent of premiums on administration, and the Lewin Group, a health policy and management consulting firm, estimates the figure is 13.4 percent.

<sup>28</sup> For example, risk premiums are viewed as an administrative expense, while additions to capital reserves are not. The NHEA views premium taxes and profits as “non-administrative” expenses.

<sup>29</sup> These estimates are based on insurance plans selected by 6,115 individuals with employer-sponsored insurance from 2,842 different employers, who appear in the 1997, 1998, 1999, and 2001 linked Medical Expenditure Panel Survey-Insurance Component (MEPS-IC). The authors estimate two regression models—one to predict insurer payments on behalf of each individual, and a second to predict premiums at the individual level. In the latter model, they regress total (employer and employee) premiums on firm size dummies, projected insurer payments from the first model, interactions of firm size dummies and projected insurer payments, demographic and health status measures, employer characteristics, market covariates, and state and year dummies. The indicators for firm size are statistically significant, but the interactions with projected payment are not. Using the parameter estimates from both models, they calculate loading fees for a “typical” employer in each size category. Because healthcare utilization is underreported in the MEPS-IC, they believe their estimates of loading fees may be overstated. Thus,

load difference between individual and large group policies of 17.5 percent, although the methodology is not provided. Finally, the Congressional Budget Office reports that administrative costs range between 7 percent of premiums for firms with at least 1,000 employees to nearly 30 percent of premiums in the individual insurance market, yielding a maximal loading fee differential of 23 percent of premiums. Again, the source of these figures is not reported.

Evaluations of healthcare reform proposals constitute an alternative source of cost estimates for plans offered through an exchange. The Lewin Group estimated that administrative costs for an exchange with only private plans would be 10.7 percent if all workers are eligible to participate (as opposed to only small groups and individuals, as some reform proposals specify). However, the figures underlying these estimates are also not reported.<sup>30</sup>

Although the range of estimated premium increases associated with a transition to an individual marketplace is large, those most relevant to our setting are comparable in magnitude to the median estimated benefit from our most conservative “plan swap” scenario. We surmise that even a modest increase in choice, coupled with the improved matching of choices to employee preferences that is modeled in this conservative scenario, is likely to generate surplus gains that outweigh the associated premium increases.

## VI. Limitations

Our analysis does not incorporate some important costs that would reduce the estimated gains from increased choice in our less conservative counterfactual scenarios. Consumers may incur disutility from having to choose from a larger set of options or may bear a personal cost of shopping which increases with the number of health plans available to them. Abaluck and Gruber (2011) find that seniors choosing Medicare Part D prescription drug plans often make choices that are inconsistent with optimization under full information, suggesting confusion when faced with large choice sets. This finding could also apply in our setting. As noted in Handel (2010), inertia or switching costs can be substantial: there is very little switching between plans from year to year even when plan characteristics and prices change substantially. These costs may help explain why an existing program that permits workers to maintain their tax exemption while choosing from a larger subset of plans is little utilized. Under Section 125 of the Internal Revenue Code, employers may set up “cafeteria plans” through which employers and employees contribute tax-free dollars for use toward benefits of the employees’ choosing. Of course, problems with adverse selection and underwriting in the individual and small group market, which would be addressed by pooling in our hypothetical scenario, may also be important explanations. In addition, it is also notable that a 2007 proposal by Senators Ron Wyden (D-Oregon) and Robert Bennett (R-Utah) to eliminate direct

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they also report figures that adjust expenditures to levels estimated by NHEA. These are the figures we use to construct loading fees as a percent of premium.

<sup>30</sup>“The Impact of House Health Reform Legislation on Coverage and Provider Incomes,” The Lewin Group Testimony before the Energy and Commerce Committee, US House of Representatives, June 25, 2009. Downloaded 9/23/2009 from <http://www.lewin.com/content/publications/June25TestimonyUpdate.pdf>.

employer subsidies for health insurance, and replace these with tax deductions for individually-purchased insurance on regulated exchanges, did not receive serious consideration in the debate over healthcare reform.

Conversely, there are reasons why our estimates may represent a lower bound on the value of choice. Employers, like consumers, bear a cost of shopping and this would be reduced if employees made their own selections. In addition we observe only a subset of plans available in the market, and the US experience with the introduction of Medicare Part D suggests that even more choice would become available in a subsidized individual exchange setting (Abaluck and Gruber 2011). (Of course, some of the plans currently provided by carriers to selected firms might disappear, particularly if adverse selection arises. However, most exchange proposals are accompanied by insurance reforms prohibiting selection tools such as exclusion of pre-existing conditions.) Our estimates also understate the benefits of choice because, aside from the stochastic error term, we do not model consumer heterogeneity within employer groups. This technical shortcoming precludes estimation of surplus gains associated with better matching of plans to individuals within a given employee group.

Last, we note that The Patient Protection and Affordable Care Act, signed into law in March 2010, will establish state-level insurance exchanges for individuals and select small employee groups. Our estimates of the value of choice are based on employees of large firms, and may not be representative of the gains for this group of individuals.

## VII. Conclusions

In its current incarnation, employer-sponsored insurance in the United States is characterized by a very limited choice, if any, for workers fortunate enough to be eligible to enroll. Our research makes use of a large panel of employer healthplan offerings and employee plan selections to quantify the surplus foregone as a result of restricted choice in the employer-sponsored system. By examining employees' choices among the set of plans they are offered, we obtain estimates of their preferences that enable us to identify their most preferred plans (and corresponding dollar-valued utility) from the entire set available in their marketplace. We estimate the median employee would be willing to forego roughly 16 percent of her subsidy for the right to apply the remainder to any plan she chooses.

In a companion paper (Dafny, Ho, and Varela 2010), we analyze the distributional effects of expanding choice and explore the differences between plans that are offered to workers in our data and those they would select under expanded choice. Importantly, we do not find evidence that employers "overweight" premiums when making healthplan selections (that is, offer cheaper plans than employees would be willing to pay for), as surveys of employers suggest (e.g., Gabel, Hunt, and Hurst 1998, and Maxwell, Temin, and Watts 2001). Our analyses indicate that employees would choose similarly-priced plans, but these plans would differ along other dimensions such as plan type and insurance carrier. In that paper, we discuss the possible reasons for the (apparent) misalignment of employer and employee preferences, an important subject for future research.

Before concluding, we note that a significant body of literature, reviewed in Gruber and Madrian (2004), documents another benefit of severing the link between employment and health benefits: a reduction in labor-market frictions, in particular “job lock” arising from the lack of insurance portability between jobs and in or out of the labor force. Like the value of reduced job lock, the value of choice is difficult to quantify and cannot be included in formal “scoring” or budgetary estimates of legislation performed by the Congressional Budget Office. Nevertheless, we estimate the value of choice is a nontrivial benefit from a widescale transition to an individual insurance system, and may more than offset the higher costs associated with an individual insurance marketplace.

## APPENDIX

APPENDIX TABLE 1A—DISTRIBUTION OF UTILITY GAINS

	All plans scenario		All plan types scenario		Plan swapping scenario	
	Least generous	Most frequent	Least generous	Most frequent	Least generous	Most frequent
5%	627	607	22	19	-185	-195
25%	1,144	1,138	317	326	62	71
50%	1,643	1,662	665	688	285	310
75%	2,396	2,493	1,161	1,222	628	687
95%	4,780	5,624	2,611	3,013	1,585	1,823
Mean	3,108	3,115	1,424	1,512	616	762
SD	176,200	91,300	94,720	59,150	22,710	31,560
Smallest	-928	-2,942	-28,440	-23,690	-535,400	-146,300
Largest	5.75E+07	2.37E+07	3.12E+07	1.52E+07	5.01E+06	7.70E+06
Observation	113,586	113,696	113,586	113,696	113,586	113,696

Note: This table corresponds to Figure 4.

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