How Responsive are Quits to Benefits?

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ABSTRACT

Economists have argued that one function of fringe benefits is to reduce turnover. However, the effect on quits of the marginal dollar of benefits relative to wages is underresearched. We use the benefit incidence data in the 1979 National Longitudinal Survey of Youth and the cost information in the National Compensation Survey to impute benefit costs and estimate quit regressions. The quit rate is much more responsive to benefits than to wages, and total turnover even more so; benefit costs are also correlated with training provision. We cannot disentangle the effects of individual benefits due to their high correlation.

I. Introduction

There is a sizable literature analyzing the relationship between fringe benefits and turnover. One reason that has been advanced as to why employers might use in-kind compensation in addition to money wages is that fringe benefits have a stronger negative effect on turnover. For example, employers might use benefits of more value to mature adults, such as health insurance with family coverage, in order to attract a more stable workforce.

A major limitation of previous work is that authors have only had access to information on whether a particular benefit has been offered to a worker, and not on the employer's expenditure on the benefit (for example, Mitchell 1983, 1982; Barron and Fraedrich 1994; Madrian 1994).[1] It would truly be surprising if holding wages, work-

1. Some previous papers utilize information on fringe benefit expenditures, but at the industry and not the establishment or individual worker level. For example, Parker and Rhine (1991) find that quits by major industry are negatively related to the share of pensions in total compensation.

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ing conditions, and other benefits the same, the presence of a fringe benefit did not lower a worker's quit probability since all that is necessary is that workers place some positive valuation on the fringe benefit. The more interesting question is whether the negative relationship between fringes and quits persists when one controls for total compensation: Does a dollar spent by an employer on benefits reduce quits by more than a dollar spent on wages? This question is the focus of the current paper.

Our analysis is based on a unique data source. The 1979 National Longitudinal Survey of Youth (NLSY79) contains information on the presence of five different fringe benefits. In order to calculate the Employment Cost Index (ECI), the National Compensation Survey obtains information on both the wages that an employer pays and the amounts he spends on fringe benefits. Using job characteristics that are contained in both the NLSY79 and the ECI data, we impute the cost to employers of the benefits received by the NLSY79 recipients. The value of imputed benefits is then entered as an explanatory variable in a mobility equation that is estimated using turnover information in the NLSY79.

Our estimated mobility equations have two appealing features. First, all fringes are included in the equation, so that, for example, the estimated health insurance coefficient does not capture the effect of an omitted leave variable. Second, the explanatory fringe benefits variable is not a binary variable, but the employer's spending on the fringe benefit. Thus, we are able to directly compare the effect of an increase in fringe benefits on quits with the effect of an increase in wages. We find that the quit rate is much more responsive to fringe benefits than to wages, and total turnover even more so.

The benefits receiving by far the most attention in regard to turnover have been pensions and health insurance. It has been well established that these benefits are negatively correlated with turnover, although the precise interpretation of this relationship is open to question. The earliest studies examining the effect of pensions on mobility utilize a binary variable for pensions (for example, Mitchell 1982, 1983). As discussed by Gustman, Mitchell, and Steinmeier (1994) in their survey paper, subsequent studies attempted to estimate the effect of the actual pension capital loss (Allen, Clark, and McDermed 1993) and to distinguish between the effects of defined benefit and defined contribution plans (Gustman and Steinmeier 1993). Much of the literature on health insurance and turnover has been concerned with "job-lock" and has focused on differing effects of health insurance coverage on different types of workers, with no attention to the costs of the coverage to the employer. (For examples, see Madrian 1994; Holtz-Eakin 1994; Buchmueller and Valletta 1996; Berger, Black, and Scott 2004).

The analysis of the effect of a particular benefit on turnover is complicated by the high correlation of various fringe benefits. Employers that offer health insurance are also more likely to offer pensions and paid leave. The estimated coefficients on the fringe benefits that are included in a mobility equation will be biased by the ones that are omitted. Most studies focus on the effect of one fringe benefit, with the effects of the other benefits being picked up by the error term.[2] In contrast, the analysis in this

2. Some studies have information on two fringes. Mitchell (1983) has the most comprehensive information on fringes. Included as explanatory variables in her job change equations are binary variables for pension, medical insurance, life insurance, stock options, and profit sharing. Baughman, DiNardi, and Holtz-Eakin (2003) have information on a number of "family-friendly" fringe benefits such as family leave, flexible sick leave policies, flexible work scheduling arrangements, and childcare for their sample of 120 employers in upstate New York.

paper uses imputed cost information on five benefits – pensions, health insurance, sick leave, vacation leave, and life insurance; in addition, we impute the total cost of all other benefits for which we do not have incidence information.

As is well recognized in the literature, a negative coefficient on a fringe benefit in a mobility equation may reflect either of two channels by which the fringe benefit has an effect on turnover. First, the benefit may directly influence employee behavior; defined benefit pensions, which act as a form of deferred compensation, are the most familiar example of this. In addition, the benefit may also reduce turnover through a selection effect: more stable workers may be attracted to employers offering pensions, health insurance, or leave benefits.[3] From the point of view of the employer it is not clear that this distinction matters very much, as the end result is reduced turnover in either case. We do not focus on this distinction in our empirical work (although our results do suggest that sorting considerations may not be terribly important).

A recent paper by Dale-Olsen (2006) using Norwegian data obtains findings similar to our paper. Dale-Olsen has access to administrative records with information on the value of fringes that were reported to tax authorities. Carrying out a fixed effect analysis that estimates the effect on turnover of wage and fringe benefit expenditures above those paid by other firms to similar workers, Dale-Olsen finds that fringe benefits have a large negative effect on separations. Indeed, when the log of total compensation and the log of fringes are both included in his turnover equation, the coefficient on fringes is large in absolute value and negative, and the coefficient on total compensation, although negative, is not statistically significant. Unlike our data, Dale-Olsen's data do not distinguish between layoffs and quits.

One question that arises from our results (and Dale-Olsen's) is whether firms' behavior is consistent with profit maximization given that firms could reduce turnover costs by shifting compensation from wages to benefits. In the next section of this paper, we develop a theoretical framework for interpreting the strong negative relationship between fringe benefits and quits. We show that this relationship is consistent with competitive equilibrium. We also show that firms with higher turnover costs will tend to be those with higher benefit expenditures. Proxying turnover costs by training, we test this implication in our empirical work. Section III of the paper describes our data and empirical methodology and Section IV presents our estimation results. Concluding comments appear in the final section.

II. A Simple Model of the Relationship Among Benefits, Wages, and Quits

We develop a simple static model to explain how the effect on quits of a dollar of benefit expenditures can be greater than that of a dollar of wages in a competitive equilibrium. Consider a labor market where each firm employs one

3. Analyzing federal government employees, Ippolito (2002) presents evidence that workers who choose to contribute to defined contribution pension plans tend to have lower quit rates. In addition, Ippolito finds that savers contributing to pension plans are likely to be better workers (as evidenced by higher job ratings and promotion rates), which he suggests might help explain why turnover is lower and wages are higher at employers offering pensions.

worker. Employers offer workers a compensation package that consists of wages W and benefits B. A worker's quit probability depends on the compensation package he receives and on his type α, which for simplicity is assumed to be observable to the employer. An employer cares about quits because it is costly to replace a worker who turns over. Turnover cost χ varies across firms depending on the type of output they produce. Output price $P(\chi)$ varies with χ, so that in equilibrium all firms earn zero profit and workers are content with their allocation among employers.

A worker's utility $\tilde{U}$ depends on the wages and benefits he receives and on a random shock that is not revealed until some time after he has started the job:

(1) $\tilde{U}(W,B,\alpha,\varepsilon) = W + f(B,\alpha) + \varepsilon, f_B > 0, f_{BB} < 0.$

The function f indicates the dollar value a worker places on the benefits he receives. If $f_B >$ (<) 1, an extra dollar of benefits is worth more (less) to a worker than an extra dollar of wage compensation. Among other things, f reflects tax considerations. A tax policy that gives preferential tax treatment to fringe benefits raises f and f_B. (See Woodbury 1983 and Woodbury and Hamermesh 1992 for empirical analyses of the effect of taxes on the choice of benefits.) The parameter α is inversely related to a worker's quit propensity Q. To capture the idea in the introduction that more stable workers place a higher value on benefits than less stable workers, we assume that $f_{B\alpha} > 0$. The random shock ε reflects the fact that the worker learns about the nonpecuniary aspect of an employer's job after some period of employment.

As discussed above, benefits deter quits. More formally, let $\phi(B,\alpha)$ denote the cost of changing jobs, and $V(\alpha,\psi)$ the expected utility a type α worker with productivity ψ can obtain elsewhere in the market. A worker quits if his utility at the employer's job falls below that which he could obtain by switching jobs or $U(W,B,\alpha) + \varepsilon - V(\alpha,\psi) + \phi(B,\alpha) < 0$, where $U(W,B,\alpha) = W + f(B,\alpha)$. This implies that the probability of a quit can be expressed as a function of B, U-V, and α, or $Q = \zeta(B,U\text{-}V,\alpha)$. By assumption, $\zeta_\alpha < 0$ and $\zeta_{B\alpha} \leq 0$: Other things the same, higher α workers are less likely to quit and are at least as responsive to benefits as low α workers. If $\phi_B = 0$, then benefits only affect quits through their effect on the worker's utility and thus the effect on quits of a marginal dollar of benefits relative to a dollar of wages equals the marginal rate of substitution between benefits and wages: $\partial Q/\partial B = f_B(\partial Q/\partial W)$. However, in addition to their effect on a worker's utility at a point in time, benefits such as pensions can be thought of as deferred compensation, which can be represented in our model as an increase in mobility costs, implying that $\zeta_B < 0$ and $|\partial Q/\partial B| < f_B\, |\partial Q/\partial W|$.

An employer chooses the wage-benefits package (W^*, B^*) to maximize expected profit $\pi = P(\chi)\psi - W^* - B^* - \chi Q$ subject to the constraint that workers with characteristics (α,ψ) receive the same expected utility available elsewhere in the market or

(2) $W^* + f(B^*,\alpha) = V(\alpha,\psi).$

It is straightforward to show that the choice of W and B must satisfy the condition[4]

(3) $-\chi\zeta_B = 1 - f_B\,.$

Note that the left-hand side of Equation 3 is the marginal benefit to the employer from

4. The derivation of this result and those that follow can be found in our longer working paper Frazis and Loewenstein (2009).

the direct reduction in quits from an additional dollar of benefits, while the righthand side is the cost to the employee from switching the marginal dollar of compensation from wages to benefits.

To gain additional insight into the choice of B, note that firms will offer the compensation package minimizing costs (including turnover costs) for workers with given characteristics (α, ψ). Let $\xi = W + B$ denote total compensation and let (W',B') be the wage-benefits package yielding the reservation utility level $V(\alpha, \psi)$ and satisfying $f_B = 1$ (workers value an extra dollar of benefits the same as an extra dollar of wages). If benefits have no effect on quits other than through their effect on utility, then an employer will choose the wage-benefits package (W',B') since this is the lowest-cost way of generating utility level V. However, if benefits deter worker quits—that is, if $\zeta_B < 0$—then it follows from Equation 3 that the employer's optimal wage-benefits package (W^*,B^*) must be such that $f_B < 1$, which in turn implies that $B^* > B'$ and $W^* < W'$—compensation is shifted from wages to benefits to the point where the marginal dollar of benefits is worth less than a dollar of wages to the worker. Total compensation $\xi^* = W^*+B^*$ exceeds $\xi' = W'+B'$, as benefits must be increased sufficiently to hold utility constant, but the worker is less likely to quit. Equilibrium requires that the increase in total compensation from a further increase in B must just equal the expected reduction in the cost of turnover.[5]

Note that the implication of our model that $f_B \leq 1$ is at odds with Dale-Olsen's (2006) inference "that workers have stronger preferences for the reported values of fringe benefits than for the equivalence in money wages." The findings in Royalty (2000) are consistent with our model. Royalty estimates workers' valuation of health insurance using data on workers' choices among fringe benefits packages. Her results indicate that families value health benefits substantially more than singles, but still far less than one-for-one with wage dollars.

An employer's choice of benefits and wages will obviously depend on turnover cost χ, worker quit propensity α, and worker productivity ψ, which we may represent as

(4) $B^* = B^*(\chi,\alpha,\psi)$.

In particular, it is straightforward to show that employers with higher turnover costs offer more benefits, as do employers hiring more stable workers. Employers' choices of benefit-wage packages in turn induce worker sorting among jobs. Specifically, firms with different values of χ will offer different wage-benefit packages, so workers will choose among values of χ. Let

(5) $\chi^* = \chi^*(\alpha,\psi)$

indicate the turnover cost associated with the job chosen by a worker with quit propensity α and productivity ψ. One can show that in a competitive equilibrium, more stable and more productive workers choose to work in jobs with higher turnover costs, so that $\partial\chi^*/\partial\alpha > 0$ and $\partial\chi^*/\partial\psi > 0$.

In the empirical work that follows, we examine the empirical relationship among

5. Note that we have simplified the analysis by assuming that α is costlessly observable to employers. If α is not observable, there arises the possibility that high turnover cost employers may screen out low α workers by offering a high B, analogous to the situations analyzed by Spence (1973) and Rothschild and Stiglitz (1976). This too results in a situation where $f_B < 1$ - that is, where workers value an extra dollar of benefits less than an extra dollar of wages.

benefits, wages, and quits. This relationship reflects both the causal effect of benefits and wages on quits and the sorting of workers across jobs.

Combining Equations 5 and 4, in equilibrium, the benefits received by a worker with quit propensity α and productivity ψ is given by

(6a) $B^E = B^*(\chi^*(\alpha,\psi),\alpha,\psi) \equiv G^B(\alpha,\psi)$.

The equilibrium wage received by a worker with quit propensity α and productivity ψ can therefore be written as

(6b) $W^E = V(\alpha, \psi) - f(G^B(\alpha,\psi), \alpha) \equiv G^W(\alpha,\psi)$.

Application of the implicit function theorem reveals that $\partial\alpha/\partial B^E > \partial\alpha/\partial W^E$, $\partial\psi/\partial B^E < \partial\psi/\partial W^E$, and $\partial\chi^*/\partial B^E > \partial\chi^*/\partial W^E$. That is, an observed shift in compensation away from wages toward benefits is associated with a higher α, a higher χ, and a lower ψ. Holding compensation constant, a larger share of compensation in the form of benefits implies a high-turnover-cost employer who is hiring more stable, but less productive, workers. While stability α and productivity ψ are difficult to observe, we can proxy turnover costs by measures of employee training, so we can test the implication that $\partial\chi^*/\partial B^E > \partial\chi^*/\partial W^E$.

To compare the observed effect of benefits on quits with that of wages, differentiate $Q = \zeta(B,U\text{-}V,\alpha)$, noting that $U = V$ in equilibrium, to obtain

(7) $$\frac{\partial Q}{\partial B^E} - \frac{\partial Q}{\partial W^E} = \zeta_B + \zeta_\alpha\left(\frac{\partial\alpha}{\partial B^E} - \frac{\partial\alpha}{\partial W^E}\right) < 0.$$

Quits are lower at employers with a higher proportion of compensation in the form of benefits, reflecting the fact that benefits raise the cost of quitting and attract more stable workers. Thus, in equilibrium, an extra dollar of benefits will be observed to be associated with a greater reduction in quits than an extra dollar of wages.

Our analysis has emphasized the fact that an employer can reduce quits by increasing the share of compensation that is in the form of benefits as opposed to wages, partly due to the deferred compensation nature of (some) benefits. However, as discussed extensively in the literature (for example, see Becker 1962, Salop and Salop 1976, and Hashimoto 1981), an employer can also reduce quits by deferring wage compensation from the present to the future. In reality, of course, employers need to choose both the tenure profile of compensation and the division of compensation into wages and benefits. For simplicity, we have focused solely on the second consideration. Extending the analysis to incorporate both considerations requires a multiperiod model rather than the single-period model that we have presented, but is otherwise straightforward. To the extent that benefits more strongly reflect deferred compensation than do wages and are preferred by more stable employees, one would still obtain the result that quits are more responsive to benefits than to wages.

Another extension is to let the cost of providing benefits vary across firms. Oyer (2008) and others have noted that some employers can provide benefits more cheaply than others. The most important determinant of the cost of providing benefits is likely to be employer size, which we control for in our regression analysis. Controlling for employer size, one suspects that variations in turnover cost are much more important than variations in the cost of providing benefits. Consequently, in our empirical work, we implicitly rule out any unobserved variations in the cost of providing benefits that may be correlated with B.

Finally, we have simplified by assuming that an employer can tailor a unique wage-benefits package for each of his workers. However, in practice, it is prohibitively costly to set up and administer a fringe benefits plan for every worker; federal tax rules also limit within-firm inequality in benefits. As predicted by our model, this problem is vitiated by workers sorting among employers on the basis of their preferences for benefits (for example, see Scott, Berger, and Black 1989). Of course, this sorting is undoubtedly imperfect (for example, see Carrington, McCue, and Pierce 2002), which simply means that there is a public good aspect to the choice of fringe benefits so that Equation 3 holds on average among an employer's employees.

III. Empirical Methods and Data

In the previous section we described the market equilibrium implied by our assumptions about the effect of benefits. In our empirical work we investigate whether reduced-form regressions of quits and turnover costs (proxied by training) are compatible with the model.

We do not give the estimated coefficients a causal interpretation. Our basic regression is the following:

$$(8) \quad Q_{t+1} = f\left(W_t, B_t\right)\boldsymbol{\beta}_1 + \boldsymbol{X}_t\boldsymbol{\beta}_2 + e_t$$

where Q_{t+1} denotes whether the respondent observed in a given job in year t quit that job by year t+1; W denotes wages, B denotes the imputed cost of benefits, $\boldsymbol{X}$ denotes other control variables, e is a residual, $\boldsymbol{\beta}_1$ and $\boldsymbol{\beta}_2$ are vectors of coefficients, and f is a specified function. We estimate Equation 8 as a linear probability model. (This was chosen rather than logit or probit to reduce computation time as all standard errors are estimated using bootstrap replications.) In addition to our main results using quits, we also estimate regressions with turnover rather than quits as the dependent variable.

We estimate the quit and turnover equations using NLSY79 data for 1988 through 1994. The NLSY79 is a data set of 12,686 individuals who were aged 14 to 21 in 1979. These youths were interviewed annually from 1979 to 1994, and every two years since then. The NLSY79 contains data on the incidence of many fringe benefits from 1988 through 1994, including five also included in the NCS data: health insurance, pensions, vacation, sick leave, and life insurance.

The NLSY79 data contain information on the incidence of various benefits, but not on their dollar value. We therefore impute the value of benefits conditional on the characteristics of the job held at the time of the interview. The imputations use the microdata collected to produce the Employment Cost Index (ECI). The ECI, which constitutes the index component of the National Compensation Survey (NCS), measures changes in wage and benefit costs over time. Establishments are the primary sampling units. A field economist visiting an establishment randomly chooses one to eight jobs, with jobs being distinguished on the basis of job title and such employment attributes as full-time status, union coverage, and incentive-based pay. Wage and nonwage compensation costs are obtained by averaging over the employees in the job. Nonwage compensation categories include pension and saving plans, health and life insurance, several forms of leave, and legally required expenditures on Social Security.[6]

Ideally, our quit measure would come from a data set representative of the general population, rather than the restricted age range of the NLSY79. However, we are unaware of any data set representative of the general population that has the extensive information on fringe benefit incidence and quits contained in the NLSY79. The NCS benefit imputations will yield biased benefits estimates for the NLSY79 if there remain significant differences in the benefits received by workers of different ages after controlling for observable job characteristics. This potential bias is ameliorated by the fact that NCS wage and benefit costs are for specific jobs rather than by broad occupation. Furthermore, we impute NLSY79 benefit costs partly on the basis of wages, which are a good predictor of work level within the NCS.[7] Our benefits imputation equations also include dummies for occupation, industry, establishment size, union coverage, full-time, calendar year, and the incidences of various benefits. The estimated NCS benefit equation has a high R^2, indicating that the residual effects of unobservable factors including age cannot be too large.

We use as our turnover measure whether the job is held at the time of the next interview; quits are measured from a question about the reason the respondent left the job. There are 360 observations in our main regression sample where the reason the respondent left is missing. We assign these a value of 0.7 in our quit equation, as quits comprise 70 percent of turnover.

We impute the value of benefits conditional on the characteristics of the job held at the time of the interview. Our imputations are based on job characteristics that are contained in both the NLSY79 and the NCS data. We start by totaling benefit costs $B = \sum B_i$ in the NCS, where B_i denotes a particular benefit i. In addition to the five benefits on which we have information in both the NLSY79 and the NCS data, there are benefits in the NCS data for which we have no information in the NLSY79. We divide these benefits into mandatory and nonmandatory benefits[8] and include nonmandatory benefits in our measure of total benefit costs.

Included as control variables in the regression are dummies for the incidences of each of the five benefits in the NCS and interactions of the incidence dummies with the log of real wages and its square; the log of establishment size; and dummies for union, full-time, and one-digit occupation. The latter variables are also all entered separately, along with dummies for calendar year. Dummies for two-digit industry are also included in the regression.

One additional complication in using the NCS data is that there are many observations for which we observe that a particular benefit is offered, and thus has a positive

6. In addition to providing estimates of employment cost trends over time, the National Compensation Survey (NCS) also provides information on occupational wages and employee benefits. For more information on the NCS and the ECI, see the BLS Handbook of Methods (http://www.bls.gov/opub/hom/home.htm).

7. The NCS field economist rates the level of work for a selected job by evaluating its duties and responsibilities. A job's work level is very highly correlated with its wage.

8. Nonmandatory benefits include sickness and accident insurance and holidays. Mandatory benefits include Social Security taxes, state Unemployment Insurance, and Worker's Compensation. Other benefits included in the NCS are treated as follows: "Other paid leave" is combined with vacations. Nonproduction bonuses, severance pay, supplemental unemployment pay, Federal unemployment insurance, other legally mandated benefits, and various railroad benefits are of small magnitude and/or do not fit our concept of fringe benefits and are omitted.

cost, but information on its cost is missing. (The NCS imputes missing values, but the imputed values have a much different—and weaker—relation to the covariates than reported values.) Omitting these observations would bias estimates of average benefit costs, as observations with positive cost would be omitted but not observations with zero costs.

In assigning values to missing benefit cost observations, our goal is to impute $f(B,W)$ in Equation 8 by substituting a consistent estimate of $E(f(B,W) \mid \mathbf{Z})$, where $\mathbf{Z}$ is our vector of control variables. In specifying f, we find that specifications using logs in benefits and wages fit better than linear specifications in explaining quits. In our preferred quadratic specification, the R^2 of a quadratic in logs is 0.0986, while the R^2 for a quadratic in linear benefits and wages is 0.0950. Thus our goal in the imputation is to obtain an estimate of $E(\ln B \mid \mathbf{Z})$ that can be used as a regressor in the quit regression. Note too that $E(\ln B \mid \mathbf{Z}) = E(\ln \sum B_i \mid \mathbf{Z}) \neq \ln \sum E(B_i \mid \mathbf{Z})$ due to the nonlinearity of the log function, so simply substituting predicted values $\hat{B}_i$ for missing values of individual benefits will not yield a consistent estimate of $E(\ln B \mid \mathbf{Z})$; we must also impute residuals e_i. The details of this imputation can be found in our longer working paper Frazis and Loewenstein (2009).

Unlike the NLSY79, the NCS separates the straight-time wage rate from overtime payments and the shift differential. We construct a wage measure W in the NCS by adding the straight-time wage rate, overtime payments, and the shift differential. (We omit observations where information on overtime or the shift differential is missing.) In addition, we add the mandatory benefits to create an augmented wage rate $\tilde{W} = W + MB$, where MB are the mandatory benefits. As we have no information on these benefits in the NLSY79, we impute them in the same manner we impute nonmandatory benefits, by regressing the log of $\tilde{W}$ in the NCS data on the control variables (including ln W and its square) and using the coefficient vector to predict log $\tilde{W}$ in the NLSY79. We denote this predicted value $\widehat{\ln \tilde{W}}$.

After filling in missing values for benefits and constructing the augmented wage rate, ln B is regressed on our control variables in the NCS:

(9) $\ln B = \mathbf{Z}\boldsymbol{\gamma} + v.$

We use the coefficients from this regression to generate the predicted value $\widehat{\ln B}$ from the NLSY79 data, which we use as our regressor in the quit equation.

For some purposes it will be useful to estimate the distributions of the individual B_i and of B in the NLSY79 data. We simulate distributions for the three benefits health insurance, pensions, and "other" consisting of all the other nonmandatory benefits. We adopt a method similar to that for imputing missing values. Once again, the details of the imputation can be found in our working paper. While conceptually similar issues arise for $\tilde{W}$, the R^2 for the regression of ln $\tilde{W}$ on $\mathbf{Z}$ is 0.9962, so we treat the distribution of $\tilde{W}$ in the NLSY79 as being equivalent to the distribution of $\exp(\widehat{\ln \tilde{W}})$.

Other than the incidence dummies and their interactions, all variables used in the construction of $\widehat{\ln B}$ are included in our quit regressions. Thus, identification of the effect of a dollar of benefits comes largely from the incidence dummies. Another way of viewing our procedure is that we are scaling the NLSY79 incidence dummies so that they are comparable with wages, with the scaling dependent on the other independent variables.

The quit regressions also include the log of weeks of tenure (plus one) at the time of the interview. Tenure is not available in the NCS data set and so cannot be used to impute benefits. The presence of a variable in the quit regression that is not used to impute benefits will bias our estimates relative to what their values would be if benefit costs were observed in the NLSY79. We show in the Appendix that the magnitude of the difference between the estimated effects of a marginal dollar of wages and a marginal dollar of benefits will be underestimated in our data (given the values we observe in the data of other parameters) under the assumption that tenure is positively correlated with $\ln B - \widehat{\ln B}$ net of the other covariates. We believe this assumption is plausible. To take one benefit, days of vacation and sick leave commonly increase with tenure. For example, the 1993 Employee Benefit Survey showed that for medium and large private establishments, the average number of vacation days granted to full-time employees increased from 9.4 days at one year of service to 16.6 days at ten years (Bureau of Labor Statistics, 1994).

Most of our regressions do not include controls for demographic variables. The omission is intentional. The object of interest is how a firm's compensation policy affects turnover. As highlighted by our theoretical model, part of the effect of a compensation policy designed to minimize turnover might be to attract workers with low rates of turnover (high α), workers who may predominantly come from specific demographic groups. From the firm's point of view, the demographic composition of its labor force is endogenous. We control for what job characteristics we can by including major occupational group in our regression, as well as firm characteristics. All regressions are weighted using the sample weights supplied by the NLSY79.

IV. Estimation Results

We restrict our sample to private sector workers (jobs in the case of the NCS) whose hourly wages are greater than one dollar and less than 100 dollars in 1982–84 dollars. Descriptive statistics for both the NCS and NLSY79 samples are shown in Table 1. Our NCS sample has 348,392 observations from 7,826 establishments over the period 1988–93. Our NLSY79 sample consists of 23,119 observations from 7,178 different individuals over the same period. Note that NLSY79 sample members are ages 22 through 36 during the sample period. Mean log wages and benefit incidence are surprisingly similar between the two samples, although vacation and sick leave are more frequently reported in the NCS. NLSY79 respondents report more professional, managerial, and skilled blue-collar occupations than is indicated in the NCS. (One typically finds that the incidence of managerial and professional jobs is higher in household than in establishment surveys; see Abraham and Spletzer 2010).

A. Quits

As discussed above, we first regress ln (B+.01) (hereafter referred to as ln B for simplicity) on our control variables using the NCS data, and then use the estimated equation to predict benefits for the individuals in the NLSY79 sample. (Using ln(B+.05) instead yields point estimates very close to those in the text.) Our initial NCS regression, which is weighted using the NCS sample weights, has an R^2 of 0.861, so the fit is quite good.

Table 1
Descriptive Statistics

Variable	N	Mean	Standard Deviation	Minimum	Maximum
NCS					
Benefits	348,088	1.90	2.00	0	40.36
Wages	348,392	8.46	5.93	1.00	99.94
Augmented wages	347,355	9.48	6.38	1.17	112.41
Ln (Benefits +.01)	348,088	–0.22	1.79	–4.61	3.70
Log wage	348,392	1.96	0.59	0.00	4.60
Log augmented wage	347,355	2.08	0.57	0.16	4.72
Sick leave incidence	348,392	0.72	0.45	0	1
Vacation incidence	348,392	0.93	0.25	0	1
Life insurance incidence	348,392	0.70	0.46	0	1
Health insurance incidence	348,392	0.78	0.42	0	1
Pension incidence	348,392	0.61	0.49	0	1
Pension cost	348,392	0.34	0.68	0	23.81
Health insurance cost	348,392	0.70	0.72	0	17.23
Other benefits	348,088	0.87	0.99	0	24.24
Ln (pension + 0.01) (pension offered)	223,665	–1.46	1.59	–4.61	3.17
Ln (health insurance + 0.01) (insurance offered)	285,050	–0.51	1.16	–4.51	2.85
Ln (other benefits + 0.01)	348,088	–0.92	1.59	–4.61	3.19
Union	348,392	0.16	0.37	0	1
Fulltime	348,392	0.80	0.40	0	1
Log establishment size	348,392	4.75	2.07	0	13.22
Year = 1988	348,392	0.17	0.37	0	1
Year = 1989	348,392	0.17	0.38	0	1
Year = 1990	348,392	0.17	0.38	0	1
Year = 1991	348,392	0.16	0.37	0	1
Year = 1992	348,392	0.16	0.37	0	1
Year = 1993	348,392	0.16	0.37	0	1
Professional/technical	348,392	0.11	0.32	0	1
Executive/administrative/ managerial	348,392	0.09	0.28	0	1
Sales	348,392	0.11	0.32	0	1
Administrative/clerical	348,392	0.16	0.37	0	1
Precision production/craft/repair	348,392	0.11	0.31	0	1
Operators/assemblers/inspectors	348,392	0.11	0.32	0	1
Transportation	348,392	0.05	0.21	0	1
Handlers/cleaners/laborers	348,392	0.08	0.28	0	1
Service	348,392	0.17	0.38	0	1

(*continued*)

Table 1 (*continued*)

Variable	N	Mean	Standard Deviation	Minimum	Maximum
NLSY79					
Quits	23,119	0.19	0.39	0	1
Turnover	23,119	0.26	0.44	0	1
Log wage	23,119	1.98	0.49	0.00	4.57
Imputed augmented wages	23,119	9.27	5.06	0.72	98.76
Imputed Log augmented wages	23,119	2.11	0.48	0.16	4.63
Imputed Ln (benefits +.01)	23,119	–0.32	1.64	–5.32	3.53
Sick leave incidence	23,119	0.62	0.49	0	1
Vacation incidence	23,119	0.84	0.37	0	1
Life insurance incidence	23,119	0.69	0.46	0	1
Health insurance incidence	23,119	0.80	0.40	0	1
Pension incidence	23,119	0.59	0.49	0	1
Imputed Ln pension costs (pension offered)	13,263	–1.62	0.73	–4.74	2.65
Imputed Ln health insurance (insurance offered)	18,093	–0.66	0.64	–3.51	1.37
Imputed Ln other benefits	23,119	–1.04	1.43	–6.44	2.70
Imputed pension costs	23,119	0.29	0.40	–0.42	5.56
Imputed health insurance	23,119	0.66	0.49	–0.50	3.37
Imputed other benefits	23,119	0.76	0.72	–0.97	8.71
Union	23,119	0.14	0.35	0	1
Fulltime	23,119	0.92	0.28	0	1
Log establishment size	23,119	4.14	2.25	0	11.51
Year = 1988	23,119	0.16	0.36	0	1
Year = 1989	23,119	0.18	0.38	0	1
Year = 1990	23,119	0.17	0.38	0	1
Year = 1991	23,119	0.16	0.37	0	1
Year = 1992	23,119	0.16	0.37	0	1
Year = 1993	23,119	0.17	0.37	0	1
Professional/technical	23,119	0.14	0.34	0	1
Executive/administrative/ managerial	23,119	0.14	0.35	0	1
Sales	23,119	0.11	0.32	0	1
Administrative/clerical	23,119	0.17	0.38	0	1
Precision production/craft/ repair	23,119	0.15	0.35	0	1
Operators/assemblers/ inspectors	23,119	0.09	0.29	0	1
Transportation	23,119	0.05	0.21	0	1
Handlers/cleaners/laborers	23,119	0.05	0.21	0	1

(*continued*)

Table 1 (*continued*)

Variable	N	Mean	Standard Deviation	Minimum	Maximum
Service	23,119	0.10	0.30	0	1
Tenure (weeks)	23,119	201.67	183.90	0	824
Log (weeks tenure + 1)	23,119	4.76	1.22	0	6.72
Log (hours training +1)	24,337	0.48	1.30	0	9.15
Female	23,119	0.43	0.50	0	1
Black	23,119	0.11	0.32	0	1
Hispanic	23,119	0.06	0.23	0	1
Highest grade completed	23,086	13.12	2.25	0	20
AFQT (residual)	23,119	6.52	18.98	–65.48	45.94
Experience at start of job	23,119	327.01	179.89	0	951
Married	23,018	0.43	0.50	0	1

Our results for quits using the NLSY79 data are shown in Table 2. For comparison purposes, and to verify that the presence of fringe benefits actually reduces quits, we first estimate a specification using ln W (not augmented by mandatory benefits) and dummies for the incidence of the five benefits in the NLSY79. All of the benefit coefficients are negative and of substantial size, and three out of five are significant at the 5 percent level (using a one-tail test).

The fourth column shows results for a specification using ln B and $\ln\tilde{W}$.[9] As the specification is nonlinear the relative magnitude of the effects of marginal dollars of B and $\tilde{W}$ will vary by their specific values. We handle this issue in two ways—by calculating effects at specific points such as the median, and by estimating the proportion of workers in the sample cohorts for which the effect of a marginal dollar of B is greater in magnitude than the effect of a marginal dollar of $\tilde{W}$. For both purposes we use the simulated distribution of B rather than the fitted values.

The coefficients on $\ln\tilde{W}$ and ln B are approximately equal. As $\tilde{W}$ is always greater than B, and $d \ln X / dX = 1 / X$, the magnitude of the effect of benefits on quits is greater than the magnitude of the effect of wages for essentially all points in our sample (or, more precisely, for more than 99.9 percent of the sample on a weighted basis). As shown in the bottom panel of Table 2, at the median values of wages and benefits the effect of the marginal dollar of benefits is approximately five times that of wages. These findings are consistent with our theoretical model in Section II.

For comparison purposes, in Columns 2 and 3 we show results for specifications omitting benefits, using ln W in Column 2 and ln $\tilde{W}$ in Column 3. The coefficients in both columns are approximately double the estimated wage coefficient in Column 4,

9. To account for the randomness of the imputation, all standard errors for regressions containing imputations are from bootstraps. Bootstrap samples are drawn from both the NCS and the NLSY79; we cluster by individual respondents in both data sets. Bootstrapping will take into account the randomness of our missing-data imputation in the NCS; see Little and Rubin (2002, p. 79–81). Standard errors in Table 2 and Table 6 are derived from 200 bootstrap replications.

Table 2
Regression Coefficients, Linear Probability Model, Quits
(Bootstrap standard errors in parentheses)

Regression Coefficients	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Ln wages	–0.044	–0.051					
	(0.008)	(0.007)					
Ln augmented wages			–0.057	–0.027	–0.069	–0.073	–0.070
			(0.008)	(0.009)	(0.060)	(0.060)	(0.039)
Pension offered	–0.016						
	(0.008)						
Health insurance offered	–0.031						
	(0.013)						
Life insurance offered	–0.011						
	(0.010)						
Sick leave offered	–0.029						
	(0.008)						
Vacation offered	–0.015						
	(0.012)						
Ln (benefit costs+0.01)				–0.023	–0.028	–0.027	–0.027
				(0.003)	(0.019)	(0.019)	(0.014)
Ln benefits x Ln wages					0.0045	0.0042	0.0041
					(0.0064)	(0.0063)	(0.0047)

(Ln wages)^2					0.0092	0.0119	0.0094
					(0.0129)	(0.0130)	(0.0097)
(Ln benefits)^2					0.0011	0.0010	0.0011
					(0.0024)	(0.0024)	(0.0021)
Employee contributions for health insurance included in benefits?	No	No	No	No	No	No	Yes
Demographics included?	No	No	No	No	No	Yes	No
R^2	0.1021	0.1012	0.0976	0.1012	0.1016	0.1024	0.1016
N	23,119	26,169	23,119	23,119	23,119	22,985	23,119
Effect of wages (or augmented wages)	–0.006	–0.007	–0.007	–0.003	–0.003	–0.003	–0.004
at median	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)
Effect of benefits at median				–0.017	–0.013	–0.012	–0.010
				(0.002)	(0.007)	(0.007)	(0.005)
Difference in effect at median				–0.013	–0.009	–0.010	–0.006
				(0.003)	(0.008)	(0.008)	(0.006)
Effect of wages (or augmented wages)			–0.010	–0.005	–0.007	–0.006	–0.007
at 25th percentile			(0.001)	(0.002)	(0.002)	(0.003)	(0.002)
Effect of benefits at 25th percentile				–0.045	–0.041	–0.041	–0.041
				(0.006)	(0.013)	(0.013)	(0.010)
Difference in effect at 25th percentile				–0.041	–0.035	–0.035	–0.034
				(0.007)	(0.015)	(0.015)	(0.010)

showing that one needs to take into account fringe benefits when estimating the effect of wages on turnover. Moreover, using the dollar cost of fringes reduces the wage coefficient by substantially more than using the incidence of specific fringes—compare Columns 1 and 4.

Next we allow a more general functional form. Specifically, Column 5 presents estimates for the quadratic specification given by

$$(10) \quad E(Q|W,B,\boldsymbol{X}) = \beta_{1w}\widehat{\ln \tilde{W}} + \beta_{2w}\widehat{\ln(\tilde{W})^2} + \beta_{1b}\widehat{\ln B} + \beta_{2b}\widehat{\ln B^2} + \beta_3\widehat{\ln(\tilde{W})(\ln B)} + \boldsymbol{X}\beta_{\boldsymbol{x}}$$

The three quadratic terms are jointly significant at the 10 percent level (p =0.069). As the specification with ln B implies a large marginal effect of small amounts of benefits, we also experimented with a specification quadratic in log compensation and ln $\tilde{W}$, but it did not fit as well as Equation 10.

The interpretation of the coefficients in the quadratic specification is not transparent, but we may note that the estimated effect of the marginal dollar of benefits is greater than the effect of the marginal dollar of wages for essentially all (99.7 percent) of the NLSY79 sample.[10] However, the p value for the hypothesis that the effect of benefits is greater than the effect of wages for a majority of the sample is only 0.17.[11] At the median values of wages and benefits, the marginal dollar of benefits reduces quits by 1.3 percentage points, while the marginal dollar of wages reduces quits by 0.3 percentage points. The 0.9 percentage point difference is not significant at conventional levels (t = 1.09).

To the extent that reductions in quits from higher benefits and wages are proportionate to the quit rate (and quit rates in turn are higher at lower compensation levels), the difference in terms of percentage point reductions will be greater at lower compensation levels and it may be easier to discern the larger effect of benefits on quits at lower levels. We find that the estimated difference between the effects of benefits and wages is especially large for low values of wages and benefits. At the 25th percentile for both wages and benefits, a marginal dollar of benefits reduces quits at the rate of 4.1 percentage points while the marginal dollar of wages only reduces quits by 0.7 percentage points (the t statistic of the difference is 2.34). As another way of examining differences across the compensation distribution, we divide the sample into halves by simulated compensation $\tilde{W} + \tilde{B}$. In our sample the quit rate is 24.9 percent when imputed compensation $\hat{C}$ is less than or equal to the sample median and 12.9 percent when compensation is greater than the median. A marginal dollar of benefits has a greater effect on quits for over 99 percent of the observations in each half of the sample. However, while the p value for the half with compensation less than or equal to the median is 0.035, the p value for the half with compensation above the median is 0.265. One obtains almost identical results if instead of using the simulated $\tilde{B}$ one takes predicted compensation to be $\hat{C} = \exp(\ln \hat{B}) + \tilde{W}$.

The next to last column in Table 2 shows the effect of adding demographic variables

10. Once again, the stronger effect of benefits on quits is due to the dollar cost of benefits being less than wages. Indeed, the estimated quit function in Equation 10 is symmetric with respect to W and B if $\beta_{1w} = \beta_{1b}$ and $\beta_{2w} = \beta_{2b}$. The data show no evidence against symmetry, as the p value of the relevant chi-square test is 0.81.

11. This p value is computed from the bootstrap distribution of the quadratic coefficients. Letting $\hat{\boldsymbol{\beta}}_j$ denote the estimated vector of coefficients from bootstrap replication j, j= 1, . . . , 200, for 17 percent of the replications the effect of benefits was less for a majority of the sample using $\hat{\boldsymbol{\beta}}_j$ to estimate the effects.

and other personal characteristics. Specifically, this regression includes age, highest grade completed, Armed Forces Qualifying Test (AFQT) score,[12] labor market experience at the start of the job, and dummies for female, black, Hispanic, and married. The coefficients in the table are little changed. Thus, the strong negative relationship between benefits and turnover is not due to sorting on characteristics of workers that are observable to us. Of course, we cannot ascertain the importance of sorting on unobservables, but if sorting considerations were truly very important, one might expect the inclusion of demographic variables and an ability proxy to have a larger effect on the benefits coefficients.

Our NCS data count as benefit expenditures only employer contributions, not contributions by employees from wages even though such contributions might be mandatory. Employees may have wages deducted from their paycheck to pay for health insurance or retirement benefits. Mandatory deductions for defined benefit pensions are relatively rare in the private sector in the period covered by our data — for example, only 3 percent of defined benefit plans in medium and large private establishments required an employee contribution in 1993 (Bureau of Labor Statistics 1994). Furthermore, voluntary deductions for defined contribution plans have close substitutes in the form of Individual Retirement Accounts, so the distinction between these contributions and wages is not clear. However, employees cannot typically easily purchase health insurance at rates comparable to those that firms can purchase, so employee contributions for health insurance arguably should be classified as benefits and not wages.

Accordingly, we estimate an alternative specification in which estimated employee contributions for health insurance are deducted from wages and added to benefits. The NCS does not have information on employee contributions for health insurance for all of the years of our analysis. However, it does contain partial information on such contributions for the years 1993 and 1994. We use these data to impute the percentage of health benefits paid by the firm, so that total health expenditures are $\hat{H} = H / \hat{P}$, where H is the cost to the employer of providing health insurance as recorded in the NCS data and $\hat{P}$ is the imputed proportion of total contributions paid for by the employer. The predicted value of total health insurance contributions is substituted for H in adding up total benefits, and the estimated employee contribution is subtracted from wages. Details of the imputation procedure are given in an appendix in our longer working paper.

Results for the quadratic specification are shown in the last column of Table 2. These results are broadly similar to those in the previous specification. The difference at the 25th percentile remains great, although there is some diminution of the effect at the median. The proportion of the NLSY79 sample for which the estimated effect of the marginal dollar of benefits is greater than the marginal dollar of wages is reduced to 96.1 percent. For sample respondents with compensation below the median, this proportion is 93.3 percent, but it remains significantly different from 50 percent at the 2 percent level.

One caveat to the above results is that firms with higher fringe benefits may also have greater nonpecuniary compensation such as more comfortable working conditions. Such nonpecuniary compensation, which can be thought of as unobserved fringe

12. More precisely, the residual from a regression of AFQT score on dummies for years of age at the time of the test.

Table 3
Effects of Individual Benefits on Quits, Log and Quadratic Specifications

	Log specification			Quadratic specification		
		Effect of marginal dollar on quits at			Effect of marginal dollar on quits at	
	p Value, Inclusion of variable	Median	25th percentile	p Value, Inclusion of variables	Median	25th percentile
Wages	0.051	–0.002	–0.003	0.162	–0.003	–0.003
		(0.001)	(0.002)		(0.004)	(0.004)
Health ins.	0.074	–0.015	–0.085	0.001	0.038	0.033
		(0.008)	(0.047)		(0.038)	(0.039)
Pensions	0.343	–0.070	–0.247	0.375	–0.015	–0.015
		(0.074)	(0.261)		(0.025)	(0.026)
Other ben.	0.001	–0.037	–0.115	0.019	–0.064	–0.077
		(0.011)	(0.036)		(0.031)	(0.035)
R^2	0.1015			0.1015		
N = 23,119						

benefits, would imply that our estimates exaggerate the effect of benefits on turnover. However, note that we estimate the effect of each dollar of wages and benefits to be roughly equal, with the larger marginal effect of benefits occurring because the effects are nonlinear and benefit costs are much lower than wages. In order to explain the larger marginal effect of benefits, unobservable nonpecuniary compensation would have to be of a sufficient magnitude to bring benefits and wages into rough equality, which is implausible.

Finally, we attempt to estimate the effects of individual benefits. To simplify our task somewhat and to focus on the most widely researched benefits, we aggregate vacations, sick leave, life insurance, and the benefits with incidence not collected in the NLSY79 into a single "other" category, thus estimating the effects for the three benefits: pensions, health insurance, and "other." Log pension and health costs are imputed as $I_i \widehat{\ln B_i}$, where I_i is an indicator for benefit i (i = [pension, health insurance]) and $\widehat{\ln B_i}$ is the imputed value of the log of benefit i using the coefficients from a regression estimated with NCS observations where benefit i is present. The R^2 for the regression for log pension costs (log health insurance costs) where pensions (health insurance) are offered is 0.308 (0.381). The R^2 for the log of other benefits is 0.813.

We first discuss results from a specification linear in log benefits. (The p value of the additional terms in a quadratic-in-logs specification is 0.168.) Standard errors are derived from 100 bootstrap replications. In this specification, shown in the left side of Table 3, the coefficients on wages and health insurance are significantly different from zero at the 10 percent level and other benefits are significant at the 1 percent level. Significance levels are similar for the difference between the effect of a marginal dol-

Table 4
Wage and Benefit Correlations, NLSY79

Correlations, Log Wage and Benefit Incidence						
	Ln Wage	Pension offered	Health Insurance offered	Vacation offered	Life Insurance offered	Sick Leave offered
Ln wage	1					
Pension offered	0.31	1				
Health insurance offered	0.31	0.52	1			
Vacation offered	0.18	0.40	0.56	1		
Life Insurance offered	0.30	0.58	0.69	0.48	1	
Sick leave offered	0.25	0.36	0.43	0.49	0.40	1

Correlations, Log Augmented Wages and Imputed Log Benefit Costs				
	Ln Augmented Wage	Ln Pension costs	Ln Health Insurance costs	Ln Other benefit costs
Ln augmented wage	1			
Ln pension costs	0.48	1		
Ln health insurance costs	0.48	0.68	1	
Ln other benefit costs	0.61	0.66	0.83	1

lar of wages and a marginal dollar of these two benefits. The pension coefficient is not significantly different from zero. Not surprisingly, standard errors are for the most part larger than in specifications aggregating benefits, and given the large standard error on the marginal dollar of pensions—0.074 at the median—it would be surprising if the effect was large enough to be significant at conventional levels.

A specification quadratic in benefits is shown on the righthand side of Table 3. The fit is similar for the two specifications. However, the effect of the marginal dollar of health insurance is now substantially wrong-signed, though not significant. The difference between the effect of the marginal dollar of wages and other benefits is significant at the 10 percent level at the median and at the 5 percent level at the 25th percentile. The equivalent differences for pensions and health insurance are not significant.

Our overall conclusion is that the estimates for the effects of individual benefits are imprecise and volatile. Part of the reason for this is that the incidences of the individual benefits are strongly correlated with each other and with wages. This is demonstrated in the first panel of Table 4, which shows the correlation matrix for ln *W* and the individual benefit incidences in the NLSY79. Benefit cost is even more highly correlated across benefits, as shown in the bottom panel of Table 4. The higher correlation of cost relative to incidence is due to the association of the incidence of individual benefits with higher costs for other benefits. An appendix to our longer working paper shows the coefficients on cross-benefit incidence dummies (and log wages) from our NCS

Table 5
Regression Coefficients for Health Insurance and Pension Costs, Quit Regression

Ln health insurance	–0.008	–0.018		–0.017
	(0.004)	(0.003)		(0.004)
Ln pension	–0.002		–0.011	–0.004
	(0.003)		(0.002)	(0.003)
Ln other benefits	–0.021			
	(0.007)			
R^2	0.1015	0.1007	0.0987	0.1008
N	23,119	23,119	23,119	23,119

regressions on individual benefits. These coefficients are generally positive and often of substantial magnitude. The combined effect of the positive association of the incidence of individual benefits with both the incidence and cost of other benefits is to make it difficult to disentangle the effects of different benefits upon quits. Note that these positive associations are consistent with our theory. If all types of benefits had a greater effect on quits than wages, firms especially concerned with reducing quits would be expected to offer several types of benefits and to more generously fund those they did offer.

The large correlations between benefits also imply that examining the effect of benefits individually may greatly exaggerate their effect on quits. Table 5 shows examples of this with specifications using logs of (augmented) wages and individual benefits (without quadratic terms). Entered separately, both health insurance and pensions have large and significant effects on quits. Entered together but without other benefits, the effect of pensions is cut by almost three-quarters and is no longer significant; the effect of health insurance drops slightly. Entered with other benefits, the effect of pensions drops further and the effect of health insurance is cut by more than half. Papers dealing with the effect of individual benefits on turnover should be read with this in mind.[13]

B. Turnover

Table 6 shows the results with turnover rather than quits as the dependent variable. For aggregate benefits, the results are similar to but stronger than the results for quits. Again, there is no evidence against symmetrical wage and benefit effects. Note that the difference between the estimated effects of wages and benefits is larger than for quits at various points in the compensation distribution, with a difference of 8.2 percentage points at the 25th percentile and 3.0 percentage points for the median (both are significant at the 1 percent level). Specifications using health, pensions, and "other" benefits show results that are similarly volatile and imprecise to those for quits.

13. As noted above, papers analyzing turnover typically analyze the effect of only one individual benefit. Mitchell (1983) has the most comprehensive information on fringes. Unlike us, Mitchell finds that including other fringes has only a small effect on the pension coefficient in a quit or turnover equation.

C. Training

Our model predicts that firms that pay a higher proportion of compensation in the form of benefits will predominantly be firms with greater hiring and training costs. One obvious proxy for training costs is the amount of formal training provided to the employee, which the NLSY79 provides data on. We regressed the log of hours (plus one) of formal training with the current employer in the previous year on wages and benefits, using the same quadratic specification as in Equation 10 (with the exception that we use a cubic in tenure rather than log tenure on the basis of fit). The results, shown in Table 7, support our model. Both wages and benefits are significantly associated with training, but in the range of most of the data the effect of benefits is much larger. At the median, a dollar increase in benefits is associated with six times the increase in log training that a dollar increase in wages is. Similarly, we find that for 86.6 percent of the sample the marginal effect of benefits exceeds that of wages (with a standard error of 4.3 percent).

V. Conclusion

It has been argued that one of the functions of fringe benefits is to reduce turnover. We have investigated this question both theoretically and empirically. Our theoretical model shows how it is possible in a competitive equilibrium that the marginal dollar of benefits would reduce quits more than the marginal dollar of wages.

For our empirical work, we turned to an untapped data source, the National Compensation Survey, to analyze the responsiveness of quits to fringe benefits. Specifically, by combining information in the NCS on the cost of benefits with information on worker quits and fringe benefit incidence in the NLSY79, we have been able to estimate the quit probability as a function of a worker's wage and the dollar value of his fringe benefits.

While our estimation procedure is reduced-form and thus sheds light on causality only indirectly, our results are consistent with employers using fringe benefits to reduce quits. Our estimates indicate that an additional dollar of fringe benefits is more strongly associated with lower quits than is an additional dollar of wages. Consistent with our theoretical model, which predicts a positive association between benefits and turnover costs, we find that employers who provide more training and presumably have greater turnover costs offer greater benefits.

If a number of benefits have strong effects on quits, firms especially concerned with reducing quits would be expected to offer them simultaneously and to fund them relatively generously. This is borne out empirically: The incidence of individual benefits is positively correlated with both the incidence and cost of other benefits. Consequently, the effect of an individual benefit on quits is greatly exaggerated when other benefits are not included in the estimated equation. Unfortunately, the high correlations among individual benefits coupled with the fact that we impute rather than observe individual benefits in the NLSY79 means that we are not able to tease out the separate effects of the individual benefits.

Table 6
Regression Coefficients, Linear Probability Model, Turnover
(Bootstrap standard errors in parentheses)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Regression coefficients							
Ln wages	–0.051	–0.063					
	(0.009)	(0.008)					
Ln augmented wages			–0.071	–0.023	–0.015	0.011	–0.057
			(0.009)	(0.010)	(0.067)	(0.067)	(0.043)
Pension offered	–0.032						
	(0.009)						
Health insurance offered	–0.042						
	(0.014)						
Life insurance offered	–0.019						
	(0.011)						
Sick leave offered	–0.031						
	(0.009)						
Vacation offered	–0.043						
	(0.013)						
Ln (benefit costs+.01)				–0.037	–0.066	–0.060	–0.052
				(0.004)	(0.021)	(0.022)	(0.016)
Ln benefits x ln wages					0.0111	0.0098	0.0066
					(0.0073)	(0.0073)	(0.0055)
$(\text{Ln wages})^2$					0.0006	–0.0008	0.0076
					(0.0145)	(0.0147)	(0.0106)

(Ln benefits)^2					–0.0020	–0.0017	–0.0010
					(0.0025)	(0.0025)	(0.0023)
Employee contributions for health insurance included in benefits?	No	No	No	No	No	No	Yes
Demographics included?	No	No	No	No	No	Yes	No
R^2	0.1579	0.1559	0.1492	0.1565	0.1568	0.1606	0.1566
N	23,119	26,169	23,119	23,119	23,119	22,985	23,119
Marginal Effects							
Effect of wages (or augmented wages) at median	–0.007	–0.009	–0.009	–0.003	–0.001	0.001	–0.003
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)
Effect of benefits at median				–0.027	–0.031	–0.029	–0.022
				(0.003)	(0.007)	(0.008)	(0.006)
Difference in effect at median				–0.024	–0.030	–0.030	–0.019
				(0.004)	(0.009)	(0.009)	(0.007)
Effect of wages (or augmented wages) at 25th percentile			–0.012	–0.004	–0.003	0.000	–0.006
			(0.001)	(0.002)	(0.003)	(0.003)	(0.002)
Effect of Benefits at 25th percentile				–0.073	–0.085	–0.078	–0.075
				(0.008)	(0.015)	(0.015)	(0.011)
Difference in effect at 25th percentile				–0.069	–0.082	–0.079	–0.069
				(0.009)	(0.017)	(0.017)	(0.012)

Table 7
Regression Coefficients, Log (Training Hours + 1), Current Employer in Previous Year

Ln augmented wages	0.483
	(0.144)
Ln (benefit costs+.01)	0.086
	(0.055)
Ln benefit x ln wages	0.013
	(0.018)
(Ln wages)^2	–0.088
	(0.033)
(Ln benefits)^2	0.020
	(0.006)
Effect of augmented wages at median	0.014
	(0.005)
Effect of benefits at median	0.091
	(0.021)
Difference in effect at median	0.077
	(0.024)
Effect of augmented wages at 25th percentile	0.028
	(0.006)
Effect of benefits at 25th percentile	0.161
	(0.040)
Difference in effect at 25th percentile	0.134
	(0.044)
R^2	0.0553
N	24,337

Appendix 1

Signing the Imputation Bias from the Omission of Tenure

In this appendix we evaluate the bias resulting from the imputation of benefits given that tenure is included in the quit regressions but not the imputation procedure. We show that if tenure and the benefit incidence dummies all have positive coefficients in a regression of benefits on the covariates observed in the NCS plus tenure, we will underestimate the effect of benefit expenditure on quits relative to wages.

We take as an example the specification in the fourth column of Table 2:

(A1) $$Q = \beta_w \ln \tilde{W} + \beta_b \ln B + \mathbf{X\beta_X} + \beta_T \ln T + e$$

where T denotes tenure and e is a residual orthogonal to the covariates. Substituting imputed benefits for the unobserved actual benefits (and ignoring the difference between imputed and actual $\tilde{W}$, as in the text), Equation A1 becomes:

(A1') $$Q = \beta_w \ln \tilde{W} + \beta_b \widehat{\ln B} + \boldsymbol{X\beta_X} + \beta_T \ln T + [\beta_b(\ln B - \widehat{\ln B}) + e],$$

where the term in brackets is the residual of the equation as estimated.

To evaluate the bias stemming from estimating Equation A1' instead of Equation A1, consider $U \equiv \ln B - \widehat{\ln B}$ as an omitted variable and apply the formula for omitted-variable bias:

(A2) $$b_w = \beta_w + \beta_b \delta_w$$

$$b_b = \beta_b + \beta_b \delta_b$$

where b_w and b_b denote the estimated coefficients on $\ln \tilde{W}$ and $\widehat{\ln B}$ in Equation A1' and where δ_w, δ_b, and δ_T denote the coefficients on $\ln \tilde{W}$, $\widehat{\ln B}$, and $\ln T$ in a regression of U on $[\ln \tilde{W}, \widehat{\ln B}, \boldsymbol{X}, \ln T]$. (Here and in what follows, we ignore the distinction between sample and population for convenience.)

Claim 1: Let $\chi \equiv \beta_b / B - \beta_w / \tilde{W}$ denote the difference between the estimated effect of a marginal dollar of benefits and a marginal dollar of wages on quits at given levels of B and $\tilde{W}$ were one able to estimate Equation A1 using actual benefits and let $D \equiv b_b / B - b_w / \tilde{W}$ denote the difference between the estimated effects of benefits and wages when one estimates Equation A1' using predicted benefits. Then

(A3) $$b_w - \beta_w = -\frac{\delta_T \theta_w b_b}{(1+\delta_b)}$$

$$b_b - \beta_b = -\frac{\delta_T \theta_b b_b}{(1+\delta_b)}$$

$$D - \chi = \frac{-\delta_T b_b}{1+\delta_b}\left(\frac{\theta_b}{B} - \frac{\theta_w}{\tilde{W}}\right),$$

where θ_b and θ_w denote the coefficients on $\widehat{\ln B}$ and $\ln \tilde{W}$ in a regression of $\ln T$ on $[\ln \tilde{W}, \widehat{\ln B}, \boldsymbol{X}]$.

Proof: The system A2 can be simplified as follows. Note that $\widehat{\ln B}$ is constructed as a predicted value from a regression on $\boldsymbol{X}$, $\ln W$, and benefit incidence dummies. Consider the regression

(A4) $$U = d_w \ln \tilde{W} + d_b \widehat{\ln B} + \boldsymbol{X d_x} + \eta .$$

Since OLS prediction errors are uncorrelated with regressors, the coefficients d_w, d_b, and $\boldsymbol{d}_x$ are all zero. The δ_w and δ_b coefficients in Equation A2 are related to the d_w and d_b coefficients through the omitted-variable bias formula:

(A5) $$d_w = \delta_w + \delta_T \theta_w = 0$$

$$d_b = \delta_b + \delta_T \theta_b = 0 .$$

From Equation A5, it follows that $\delta_w = -\delta_T\theta_w$ and $\delta_b = -\delta_T\theta_b$. Substituting these results into Equation A2 and rearranging, we obtain Equation A3. Q.E.D.

To determine the bias in b_w, b_b, and D, we must sign δ_T and $-\delta_T/(1-\delta_b)$. Let $\boldsymbol{V}$ denote the vector of control variables in the predicted benefits equation that are not included in the quit Equation A1 (that is, the vector $\boldsymbol{Z}$ in the paper is given by $\boldsymbol{Z} = [\boldsymbol{X}\ \boldsymbol{V}]$). Specifically, in our empirical work, $\boldsymbol{V}$ consists of the benefit incidence variables. As a first step, write predicted benefits as a least squares projection on $\boldsymbol{V}$, W, and $\boldsymbol{X}$:

$$\text{(A6)}\quad \widehat{\ln B} = \boldsymbol{V}(\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V) + (\alpha_W + \alpha_T\gamma_W)\ln\tilde{W} + \boldsymbol{X}(\boldsymbol{\alpha}_X + \alpha_T\boldsymbol{\gamma}_X),$$

where α_T is the coefficient that we would see on ln T were it also included in Equation A6 and where $\boldsymbol{\gamma}_V$, γ_w, and $\boldsymbol{\gamma}_X$ denote the coefficients on $\boldsymbol{V}$, $\ln\tilde{W}$, and $\boldsymbol{X}$ in a regression of ln T on $[\boldsymbol{V}, \ln\tilde{W}, \boldsymbol{X}]$.

To simplify the ensuing analysis, let $\boldsymbol{V}^*$ and ln T^*, respectively, denote the residuals from regressions of $\boldsymbol{V}$ and ln T on $\ln\tilde{W}$ and $\boldsymbol{X}$. From the discussion in Goldberger (1991, pp. 185–86), we know that that the residual of the regression of $\widehat{\ln B}$ on $\ln\tilde{W}$ and $\boldsymbol{X}$ is given by $\widehat{\ln B}^* = \mathbf{V}^*(\boldsymbol{\alpha}_V + \alpha_T\boldsymbol{\gamma}_V)$. Also, denoting the least squares projection of U on $\widehat{\ln B}^*$ and ln T^* as

$$\text{(A7)}\quad U = \delta_{b*}\widehat{\ln B}^* + \delta_{T*}\ln T^*,$$

we know that $\delta_{b*} = \delta_b$ and $\delta_{T*} = \delta_T$.

We now relate δ_{b*} and δ_{T*} to the parameters in Equation A6.

Claim 2: Let $\varepsilon \equiv \ln T^* - \boldsymbol{V}^*\boldsymbol{\gamma}_V$ denote the residual from a regression of ln T^* on $\boldsymbol{V}^*$ and let λ denote the coefficient on $\widehat{\ln B}^*$ from a simple regression of ln T^* on $\widehat{\ln B}^*$. Then

$$\text{(A8)}\quad \delta_{b*} = D_1 D_2,$$

where

$$D_1 \equiv \frac{-\mathrm{Var}(\varepsilon)}{\mathrm{Var}(\ln T^*) - \lambda\mathrm{Cov}(\widehat{\ln B}^*, \ln T^*)},$$

$$D_2 \equiv \frac{(\boldsymbol{\alpha}_V + \alpha_T\boldsymbol{\gamma}_V)\boldsymbol{V}^{*\prime}\boldsymbol{V}^*(\alpha_T\boldsymbol{\gamma}_V)}{(\boldsymbol{\alpha}_V + \alpha_T\boldsymbol{\gamma}_V)\boldsymbol{V}^{*\prime}\boldsymbol{V}^*(\boldsymbol{\alpha}_V + \alpha_T\boldsymbol{\gamma}_V)}.$$

Proof: The coefficient b_1 of a variable X_1 from a regression of Y on two variables X_1, X_2, and a constant is given by

$$b_1 = \frac{\mathrm{Cov}(X_1, Y)\mathrm{Var}(X_2) - \mathrm{Cov}(Y, X_2)\mathrm{Cov}(X_1, X_2)}{\mathrm{Var}(X_1)\mathrm{Var}(X_2) - \mathrm{Cov}(X_1, X_2)^2}$$

(for instance, see Gujarati 1978, p. 103). Correspondingly, the estimated coefficient on $\widehat{\ln B}^*$ in (A7) is given by

$$\text{(A9)}\quad \delta_{b*} = \frac{\mathrm{Cov}(\widehat{\ln B}^*, U)\mathrm{Var}(\ln T^*) - \mathrm{Cov}(U, \ln T^*)\mathrm{Cov}(\widehat{\ln B}^*, \ln T^*)}{\mathrm{Var}(\widehat{\ln B}^*)\mathrm{Var}(\ln T^*) - \mathrm{Cov}(\widehat{\ln B}^*, \ln T^*)^2}$$

$$= \frac{-\mathrm{Cov}(U, \ln T^*)\mathrm{Cov}(\widehat{\ln B}^*, \ln T^*)}{\mathrm{Var}(\widehat{\ln B}^*)\mathrm{Var}(\ln T^*) - \mathrm{Cov}(\widehat{\ln B}^*, \ln T^*)^2}$$

as $\text{Cov}(\widehat{\ln B}^*, U) = 0$.

Note that $U = \alpha_T \varepsilon + \eta$, where η is orthogonal to ***X***, *W*, and ln *T*, and hence ln *T**. Thus,

(A10) $\text{Cov}(U, \ln T^*) = \alpha_T \text{Var}(\varepsilon)$.

Substituting Equation A10 into Equation A9 and dividing the numerator and denominator by $\text{Var}(\widehat{\ln B}^*)$, one gets

(A11) $$\delta_{b^*} = \frac{-\alpha_T \lambda \, \text{Var}(\varepsilon)}{\text{Var}(\ln T^*) - \lambda \, \text{Cov}(\widehat{\ln B}^*, \ln T^*)},$$

where the denominator is the residual variance from a regression of ln *T** on $\widehat{\ln B}^*$.

Expressing ln *T** and $\widehat{\ln B}^*$ in terms of ***V****, the coefficient λ evaluates to

(A12) $$\lambda = \frac{(\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V) \boldsymbol{V}^{*\prime} \boldsymbol{V}^* \boldsymbol{\gamma}_V}{(\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V) \boldsymbol{V}^{*\prime} \boldsymbol{V}^* (\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V)}.$$

Substituting Equation A12 into Equation A11 yields Equation A8. Q.E.D.

We can now evaluate δ_{b^*}. The coefficients on $\widehat{\ln B}$ and $\ln \tilde{W}$ in a regression of ln *T* on $[\ln \tilde{W}, \widehat{\ln B}, \boldsymbol{X}]$—$\theta_b$ and θ_w—are both positive in our data, reflecting the fact that tenure is positively related to wages and imputed benefits. It is reasonable to believe that $\alpha_T > 0$; that is, the coefficient on ln *T* would be positive were ln *T* included in the benefits regression. It also reasonable to believe that each element of α_V is greater than zero: were ln *T* included in the benefits regression, the coefficients on the benefit incidence variables would still be positive. We observe in the data that λ is positive and that each element of $(\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V)\boldsymbol{V}^{*\prime}\boldsymbol{V}^*$ is positive. This leads to:

Claim 3: Given $\alpha_T > 0$, each element of $\boldsymbol{\alpha}_V > 0$, and other parameters mentioned in the previous paragraph are as in the data, $-1 < \delta_{b^*} < 0$ and $\delta_{T^*} > 0$.

Proof: As mentioned, $\text{Var}(\ln T^*) - \lambda \, \text{Cov}(\widehat{\ln B}^*, \ln T^*)$ is the residual variance from a regression of ln *T** on $\widehat{\ln B}^*$. Consequently, $\text{Var}(\ln T^*) - \lambda \text{Cov}(\widehat{\ln B}^*, \ln T^*) \geq \text{Var}(\varepsilon)$. Thus $-1 \leq D_1 < 0$. Since $\boldsymbol{V}^{*\prime}\boldsymbol{V}^*$ is positive definite, the denominator in the expression for D_2 is positive. Since α_T and λ are both positive, the numerator must also be positive. The denominator will be less than the numerator if $(\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V)' \boldsymbol{V}^{*\prime} \boldsymbol{V}^* \boldsymbol{\alpha}_V > 0$, which is a weighted average of the elements of the $\boldsymbol{\alpha}_V$ vector with weights the corresponding elements of $(\boldsymbol{\alpha}_V + \alpha_T \boldsymbol{\gamma}_V)\boldsymbol{V}^{*\prime}\boldsymbol{V}^*$. Since these weights are positive and the elements of $\boldsymbol{\alpha}_V$ are positive, the numerator in the expression for D_2 is less than the denominator. Thus $0 < D_2 < 1$.

Similarly, we have

(A13) $$\delta_{T^*} = \frac{\text{Cov}(U, \ln T^*) \text{Var}(\widehat{\ln B}^*) - \text{Cov}(U, \widehat{\ln B}^*) \text{Cov}(\widehat{\ln B}^*, \ln T^*)}{\text{Var}(\widehat{\ln B}^*) \text{Var}(\ln T^*) - \text{Cov}(\widehat{\ln B}^*, \ln T^*)^2}$$

$$= \frac{\alpha_T \, \text{Var}(\varepsilon)}{\text{Var}(\ln T^*) - \lambda \, \text{Cov}(\widehat{\ln B}^*, \ln T^*)} > 0. \quad \textit{Q.E.D.}$$

Since b_b is negative; δ_T, θ_b, and θ_w are all positive; and θ_b is between −1 and 0, we can conclude from Equation A3 that our estimates of the effects of benefits and wages on quits are both biased downward in absolute value. The bias expression for the difference in marginal effects between benefits and wages, $D - \chi$ from Equation A3, evaluates to

$$D - \chi = \frac{-\alpha_T \operatorname{Var}(\varepsilon)}{\operatorname{Var}(\ln T^*) - \lambda \operatorname{Cov}(\widehat{\ln B}^*, \ln T^*) - \alpha_T \lambda \operatorname{Var}(\varepsilon)} \left(\frac{\theta_b}{B} - \frac{\theta_w}{\tilde{W}} \right) b_b.$$

As discussed in the text, $(\theta_b / B - \theta_w / \tilde{W})$ is positive in our data. We can conclude that the difference in the estimated marginal effects of benefits and quits on wages is biased downward in absolute value.

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