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Abstract

It has been argued that one of the functions of fringe benefits is to reduce turnover. However, due to a lack of data, the effect on quits of the marginal dollar of benefits relative to the marginal dollar of wages is an under-researched topic. This paper uses the benefit incidence data in the 1979 Cohort of the National Longitudinal Survey of Youth (NLSY79) and the cost information in the National Compensation Survey to impute benefit costs. The value of imputed benefits is then entered as an explanatory variable in a mobility equation that is estimated using turnover information in the NLSY. We find that the quit rate is much more responsive to fringe benefits than to wages; this is even more the case with total turnover. We also find that benefit costs are correlated with training provision. Due to the high correlation of the costs of individual benefits, it is not possible to disentangle the effects of separate benefits. An interesting feature of the model that we develop for interpreting the strong negative relationship between fringe benefits and turnover is that abstracting from heterogeneity, workers must at the margin place a higher valuation on a dollar of wages than a dollar of benefits since otherwise an employer could profit by switching compensation from wages to fringes. Worker heterogeneity modifies this result and reinforces any causal relationship between fringe benefits and turnover provided that more stable workers have a greater preference for compensation in the form of fringes.

I. Introduction

There is a sizable labor economics 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 reduce turnover more than money wages of the same value to the employer. 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. The benefits receiving by far the most attention in this regard 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.

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 [e.g., Mitchell (1983), Mitchell (1982), Barron and Fraedrich (1994), Madrian (1994)].[1] It would truly be surprising if holding wages, working 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.

Pensions are the fringe benefit that has received the most scrutiny. As discussed in the survey paper by Gustman, Mitchell, and Steinmeier (1994), there is substantial

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[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.

evidence of a significant negative correlation between pensions and turnover. For example, using the Survey of Income and Program Participation, Gustman and Steinmeier (1993) find one-year mobility rates of about 20% for male workers with no pension and 6% for those with pensions.[2] Most of the studies examining the effect of pensions on mobility utilize a 0-1 variable for pensions. However, Gustman and Steinmeier are able to estimate the backloading associated with an individual's defined benefit plan; they find that the backloaded pension compensation and the non-backloaded compensation premium paid to a worker have similar effects on the worker's quit probability.[3] In the same vein, Gustman and Steinmeier find that defined benefit and defined contribution plans have similar effects on quit rates. In a similar spirit, Allen, Clark, and McDermed (1993) merge information from the Employee Benefit Survey into the PSID to estimate the pension benefits that job leavers forego and the resultant effect on turnover. Their estimates indicate that the capital loss from foregone pension benefits is associated with a substantial reduction in turnover, but this is mainly due to a reduced layoff probability. Their estimates also provide some evidence of self-selection into pensions by employees on the basis of observable characteristics (but not on the basis of unobservables).

A second fringe benefit that has received a fair amount of attention is health insurance. The 'job-lock' literature analyzing the effect of health insurance on quits attempts to sort out the compensation effect indirectly (as pointed out, for example, by

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[2] Using the Quality of Employment Survey, Mitchell (1983) too obtains a large negative effect of pensions on quits for males. For females, however, the effect is much smaller and not statistically different from zero.

[3] When not controlling for workers' alternative compensation, Gustman and Steinmeier find that backloaded pension compensation has a very large negative effect on mobility. Ippolito (1991) obtains this result using the Pension Benefit Amounts Survey.

Buchmueller and Valletta (1996)). Instead of focusing on whether workers with health insurance are more likely to quit their jobs than individuals without health insurance, papers in this literature compare whether workers whose employers provide health insurance and whose spouses' employers do not provide health insurance have a lower quit rate than workers who receive health insurance and whose spouses' employers also provide health insurance. This approach essentially differences out the compensation effect of employer-provided health insurance.[4]

Complicating the interpretation of studies of particular benefits is the high correlation of different 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.[5] In contrast, we have incidence and imputed cost information on five benefits – pensions, health insurance, sick leave, vacation leave, and life insurance; in addition, we impute the cost of benefits where we do not have incidence information.

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[4] Of course, as noted by Madrian (1994), the effect of other job attributes that tend to be associated with health insurance are also differenced out; we return to this point below. Using the National Medical Expenditure Survey, Madrian finds a significant job-lock effect for married men. In contrast, Holtz-Eakin (1994) finds no effect using PSID data. Buchmueller and Valletta find a strong effect for married women, but only weak effects for married men. Examining other interactions (specifically, whether the quit rates of individuals whose employers offer health insurance and whose family members have health problems are lower than the quit rates of individuals whose employers offer health insurance and whose family members do not have health problems), Berger, Black and Scott (2004) do not find evidence for job lock. See Berger, Black, and Scott for a review of other papers in the literature.

[5] Some studies have information on two fringes. Mitchell (1983) has the most comprehensive information on fringes. Included as explanatory variable in her job change equations are 0-1 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 child care for their sample of 120 employers in upstate New York.

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.[6] From the point of the 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).

The analysis in this paper is based on a unique data source. The NLSY79 contains information on the presence of five different fringe benefits. In order to calculate the Employment Cost Index, the National Compensation Survey obtains information on both the wages that an employer pays and the amounts he spends on fringe benefits. We impute the value of benefits by using job characteristics that are contained in both the NLSY79 and the ECI data.[7] The value of imputed benefits is then entered as an

[6] 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.

[7] Pierce (forthcoming and 2001) has used the ECI micro-data to analyze how the distribution of compensation (that is, wages plus fringe benefits) has changed over time. Gruber and Lettau (2004) have used the ECI data to estimate the effect of taxes on firms' demand for health insurance. Carrington, McCue, and Pierce (2002) use the ECI micro-data to analyze the effect of benefit nondiscrimination rules in the tax code. The Employee Benefit Survey (EBS) used by Allen, Clark, and McDermed (1993) is now collected along with the ECI data as part of the National Compensation Survey. The EBS has information on the features of the benefit plans offered by employers. Combining the information on pension formulas with the information in the PSID, Allen, Clark, and McDermed impute pension compensation and the capital loss from turnover. In contrast, we use the information in the ECI micro-data on the cost to employers of the benefits they provide.

explanatory variable in a mobility equation that is estimated using turnover information in the NLSY.

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 0-1 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.

A recent paper by Dale-Olsen (2006) obtains similar findings for Norway. 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.[8] 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.

In the next section of this paper, we develop a theoretical framework for interpreting the strong negative relationship between fringe benefits and turnover. Our model is at odds with Dale-Olsen's inference "that workers have stronger preferences for

[8] The same result obtains for "excess turnover," defined as hires if the employment at a firm is falling and as separations if it is increasing.

the reported values of fringe benefits than for the equivalence in money wages." If the negative relationship between fringe benefits and turnover is simply due to worker sorting, each employer will offer a level of benefits such that his workers place equal value on a dollar of fringe benefits and a dollar of wages. And if benefits directly reduce quits, then in equilibrium an employer must be offering a level of benefits such that workers place a *lower* value on a dollar of fringe benefits than wages and a dollar of fringe benefits because otherwise he could profit by switching compensation from wages to fringes.[9]

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 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.

[9] The findings in Royalty (2000) are consistent with this prediction. Royalty estimates workers' valuation of health insurance using data on workers' choices of fringe benefits packages offered by the employer. Her results indicate that families value health benefits substantially more than singles, but still far less than one-for-one with wage dollars.

A worker's utility 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:[10]

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

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.[11] 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 $f_{B\alpha} > 0$. For simplicity, we assume that the random shock ε is distributed uniformly. This shock reflects the fact that the worker learns about the non-pecuniary aspect of an employer's job after some period of employment.

As discussed above, benefits deter quits. More formally, let $\varphi(B, \alpha)$ 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 $U(W,B,\alpha) - V(\alpha, \psi) + \varepsilon + \varphi(B, \alpha) < 0$. This implies that the probability of a quit can be expressed as a function of B, U-V, and α:

(2) $Q = \zeta(B,U\text{-}V,\alpha).$

[10] Note that we have simplified the analysis by eliminating income effects, which are unlikely to be a major consideration in the current context.

[11] Empirical analyses of the effect of taxes on the choice of benefits include Woodbury (1983) and Woodbury and Hamermesh (1992), who assume that employers care only about the total compensation paid to workers. This contrasts with our analysis, which allows for the possibility that a dollar spent by an employer on benefits may reduce quits by more than a dollar spent on wages.

Note that $\zeta_{U\text{-}V} < 0$: benefit – wage combinations yielding a higher level of ex ante utility in the current job relative to the market alternative result in a lower quit probability. And by assumption $\zeta_\alpha < 0$: other things the same, higher α workers are less likely to quit. In addition,

(3a) $\partial Q/\partial W = \zeta_{U\text{-}V}$

(3b) $\partial Q/\partial B = \zeta_B + f_B\zeta_{U\text{-}V} = \zeta_B + f_B(\partial Q/\partial W)$

If $\varphi_B = 0$, then benefits only affect quits through their effect on the worker's utility and $\partial Q/\partial B = f_B(\partial Q/\partial W)$. 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 $\varphi_B > 0$, $\zeta_B < 0$ and $\partial Q/\partial B < f_B(\partial Q/\partial W)$. We also assume that $\varphi_{\alpha B} \geq 0$, which implies that $\zeta_{B\alpha} \leq 0$: high α workers are at least as responsive to benefits as low α workers.

Now consider an employer's choice of W and B for workers with given characteristics (α, ψ). An employer's expected profit is given by

(4) $\pi = P(\chi)\psi - W - B - \chi Q.$

The employer chooses W and B to maximize expected profit subject to the constraint that the wages and benefits offered to a worker provide expected utility equal to at least that which the worker can obtain elsewhere in the market: $U(W,B,\alpha) \geq V(\alpha, \psi)$.[12] In a symmetric equilibrium where all employers with the same hiring cost offer the same wages and benefits, the inequality will be binding so that

(5) $W = V(\alpha, \psi) - f(B, \alpha).$

[12] 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 some employers might screen out low α workers by offering a high B, analogous to the situations analyzed by Spence (1973) and Rothschild and Stiglitz (1976).

Substituting (5) into (4) and differentiating, one finds that the choice of *B* satisfies the first-order condition:

(6) $$f_B(1+\chi\frac{\partial Q}{\partial W})-1-\chi\frac{\partial Q}{\partial B}=0\,.$$

Using (3), equation (6) can be rewritten as

(7a) $$-\chi\zeta_B = 1 - f_B\,.$$

The corresponding second order condition is given by

(7b) $$\eta_{BB} \equiv f_{BB} + \chi\zeta_{BB} < 0.$$

To gain insight into the choice of *B*, let $\xi = W + B$ denote total compensation and let (W',B') be the wage-benefits package satisfying (5) and $f_B = 1$. If benefits only deter quits through their effect on utility, then an employer will choose the wage-benefits package (W',B') since this is the lowest-cost way of offering a worker the utility level *V*. However, if benefits deter worker quits, that is, if $\zeta_B < 0$, then it follows from (7a) 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 $W^* < W'$. Total compensation $\xi^* = W^*+B^*$ exceeds $\xi' = W'+B'$, but a worker is less likely to quit than if he were to receive the wage-benefits package (W',B'). 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.

An employer's choice of *B* depends on turnover cost χ, worker quit propensity α, and worker productivity ψ: $B^* = B^*(\chi,\alpha,\psi)$ Differentiating (7a) yields

(8a) $$\frac{\partial B^*}{\partial \chi} = \frac{\zeta_B}{\eta_{BB}} > 0\,,$$

(8b) $$\frac{\partial B^*}{\partial \alpha} = \frac{-\eta_{B\alpha}}{\eta_{BB}} > 0$$

(8c) $$\frac{\partial B^*}{\partial \psi} = 0 \, ,$$

where $\eta_{B\alpha} \equiv f_{B\alpha}$ - $\chi\zeta_{B\alpha}$.. Employers with higher turnover costs offer more benefits, as do employers hiring more stable workers.

We now analyze how workers are sorted among jobs. Firms with different values of χ will offer different wage- benefit packages, so workers will choose among values of χ. Note that in equilibrium, profits equal zero, so that

(9) $$W^* = P(\chi)\psi - B^* - \chi Q \, .$$

Substituting (9) into (1) yields

(10) $$U(W^*,B^*,\alpha) = f(B^*, \alpha) + P(\chi)\psi - B^* - \chi Q.$$

In determining where to work, a worker chooses the job offering the highest expected utility. Differentiating (10) with respect to χ and noting that $\partial U / \partial B^* = f_B - 1 - \chi\zeta_B = 0$, one obtains the first-order condition for a worker's choice of χ:

(11a) $$P_\chi\psi - Q = 0 \, .$$

Using (8a), the corresponding second-order condition is given by

(11b) $$\psi P_{\chi\chi} - \frac{\zeta_B^2}{\eta_{BB}} < 0 \, .$$

Let $\chi^* = \chi^*(\alpha,\psi)$ indicate the turnover cost associated with the job chosen by a worker with quit propensity α and productivity ψ. Differentiating (11a) with respect to α and ψ, one obtains

(12a) $$\frac{\partial \chi^*}{\partial \alpha} = \frac{\zeta_\alpha \eta_{BB} - \zeta_B \eta_{B\alpha}}{\psi P_{\chi\chi}\eta_{BB} - \zeta_B^2} > 0$$

(12b) $$\frac{\partial \chi^*}{\partial \psi} = \frac{-P_\chi \eta_{BB}}{\psi P_{\chi\chi}\eta_{BB} - \zeta_B^2} > 0 \ .$$

Thus, in equilibrium more stable and more productive workers choose to work in jobs with higher turnover costs.

To complete the characterization of market equilibrium, note that $V(\alpha, \psi) = U(W^*(\chi^*, \alpha, \psi), B^*(\chi^*, \alpha, \psi), \alpha)$ and differentiate (10) to obtain

(13a) $V_\alpha(\alpha,\psi) = f_\alpha + \chi^*\varphi_\alpha$

(13b) $V_\psi(\alpha, \psi) = P(\chi)$.

High α workers are rewarded for the lower turnover cost they impose on employers and high Ψ workers are rewarded for their higher contribution to revenue. Further differentiation of (13), using (12) and (8) yields

(14a) $$V_{\alpha\alpha} = f_{\alpha\alpha} + \chi\varphi_{\alpha\alpha} + \frac{\zeta_\alpha(-\zeta_\alpha\eta_{BB} + \zeta_B\eta_{B\alpha}) + \eta_{B\alpha}(\zeta_B\zeta_\alpha - \psi P_{\chi\chi}\eta_{B\alpha})}{\psi P_{\chi\chi}\eta_{BB} - \zeta_B^2} \ ,$$

(14b) $$V_{\psi\psi} = \frac{-P_\chi^2\eta_{BB}}{\psi P_{\chi\chi}\eta_{BB} - \zeta_B^2} > 0 \ ,$$

(14c) $$V_{\psi\alpha} = P_\chi \frac{\zeta_\alpha\eta_{BB} - \zeta_B\eta_{B\alpha}}{\psi P_{\chi\chi}\eta_{BB} - \zeta_B^2} > 0 \ .$$

In the empirical work that follows, we examine the empirical relationship between benefits and quits. Specifically, we compare the responsiveness of the quit rate with respect to benefits and wages. Differentiating (9), (13a), and (13b), using (14) and noting from (9) and (11a) that $P - P_\chi\chi^* = \mathrm{P} - \chi^*Q/\psi = (W^*+B^*)/\psi > 0$, one obtains

(15a) $$\frac{\partial \alpha}{\partial B^*} - \frac{\partial \alpha}{\partial W^*} = \frac{P(P_{\chi\chi}\eta_{BB} - \zeta_B^2) + \chi^*\zeta_B^2 P_\chi}{\zeta_\alpha\zeta_B(P - \chi^* P_\chi) - P\eta_{B\alpha}\psi P_{\chi\chi}} > 0,$$

$$\text{(15b)} \quad \frac{\partial \psi}{\partial B*} - \frac{\partial \psi}{\partial W*} = \frac{\chi * \psi P_{\chi\chi}(\zeta_\alpha \eta_{BB} - \zeta_\alpha \eta_{B\alpha})}{\zeta_\alpha \zeta_B (P - \chi * P_\chi) - P \psi P_{\chi\chi} \eta_{\alpha B}} < 0$$

$$\text{(15c)} \quad \frac{\partial \chi *}{\partial B*} - \frac{\partial \chi *}{\partial W*} = \frac{(P - \chi * P_\chi)(\zeta_\alpha \eta_{BB} - \zeta_B \eta_{\alpha B})}{\zeta_\alpha \zeta_B (P - \chi * P_\chi) - P \psi P_{\chi\chi} \eta_{\alpha B}} > 0 .$$

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 large share of compensation in the form of benefits implies a high-turnover-cost employer who is hiring more stable, but less productive, workers.

Two different effects contribute to the result that holding compensation constant, a higher share of compensation in the form of benefits is associated with a lower ψ. There is a direct effect stemming from the fact (discussed above in regard to (7a)) that $\zeta_B < 0$ implies that in equilibrium workers must prefer wages to benefits on the margin, so that a given increase in wages will attract more productive workers than the same increase in benefits. In addition, there is an indirect effect stemming from the fact that jobs offering a higher share of compensation in the form of benefits, but the same total compensation, are filled by more stable workers who place a higher value on benefits. Worker productivity must be lower to offset the fact that other things the same, more stable workers receive higher compensation as a reward for their lower quit rate.

To compare the observed effect of benefits on quits with that of wages, recall that $U = V$ in equilibrium and differentiate (2) to obtain

$$\text{(16)} \quad \frac{\partial Q}{\partial B*} - \frac{\partial Q}{\partial W*} = \zeta_B + \zeta_\alpha \left(\frac{\partial \alpha}{\partial B*} - \frac{\partial \alpha}{\partial W*}\right) < 0.$$

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 is associated with a greater reduction in quits than an extra dollar of wages. (An employer who shifts compensation from wages to benefits to reduce quits will end up hiring less productive workers.)

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 multi-period 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.[13]

Another extension is to let the cost of providing benefits vary across firms. Oyer (2005) and others have noted that some employers can provide benefits more cheaply than others. One way to capture this is to let $b = \delta B$, where B denotes benefits received by workers and b denotes benefit expenditures by an employers; employers with a cost advantage in providing benefits have a lower δ. Adopting this specification and

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[13] It turns out that in our simplified model, $\frac{\partial Q}{\partial W^*} = \zeta_\alpha \frac{\partial \alpha}{\partial W^*} > 0$. If wages were allowed to vary over time, the relationship between quits and wages would be ambiguous.

differentiating the resulting first-order conditions, one finds that $\frac{\partial B^*}{\partial \delta} < 0$ and $\frac{\partial \alpha^*}{\partial \delta} < 0$. However, although employers who can provide benefits more cheaply offer more benefits, the effect of a higher δ on benefit expenditures is unclear. The most important determinant of δ 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 δ that may be correlated with *B*.

Finally, we have simplified our theoretical analysis by assuming that an employer can tailor a unique wage-benefits package for each of his workers. This is not feasible in practice in part because setting up and administering a fringe benefits plan for every worker would be costly and in part because of federal tax rules that limit within-firm inequality in benefits. One way around this problem is for workers to sort among employers on the basis of their preferences for benefits. Scott, Berger, and Black (1989) do find evidence of such sorting. And Carrington, McCue, and Pierce (2002) find evidence that workers whose desired benefits differ from their coworkers' are more likely to be employed part-time, perhaps as a way for employers to get around benefit nondiscrimination rules.

However, sorting by workers in accordance with their preferences for benefits is undoubtedly imperfect. Consistent with this, Carrington, McCue, and Pierce present evidence consistent with the hypothesis that employers are not entirely free to vary benefits among their workers. Specifically, they show that low-pay workers with high-pay co-workers receive more benefits than they would if their co-workers had low pay while high-pay workers with low-pay co-workers receive fewer benefits than they would if their co-workers were high-pay. This consideration complicates, but does not vitiate, our theoretical analysis. It simply means that there is a pubic good aspect to the choice of fringe benefits so that an employer's choice of benefits must balance workers' average valuation of benefits against the average effect of benefits on quits.

III. Empirical Methods and Data

Our basic regression is the following:

$$(11) \qquad Q_{t+1} = f(W_t, B_t, \delta) + X_t\beta + e_t$$

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, *X* denotes other control variables, e is a residual, and δ and β are vectors of coefficients. We estimate (11) as a linear probability model. In addition to our main results using quits, we also estimate regressions with turnover rather than quits as the dependent variable. (Unless obvious from the context, references to "quit regressions" also include turnover regressions.)

We estimate the quit and turnover equations using NLSY79 data for 1988 through 1994. The NLSY79 is a dataset of 12,686 individuals who were aged 14 to 21 in 1979. These youth 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 five 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.[14]

Ideally, our quit measure would come from a dataset representative of the general population, rather than the restricted age range of the NLSY79. However, we are unaware of any dataset with both the extensive information on fringe benefit incidence and quits contained in the NLSY79 and representative of the general population. The NCS benefit imputations will yield biased benefits estimates for the NLSY79 if after controlling for observable job characteristics, there remain significant differences in the benefits received by workers of different ages. This potential bias is ameliorated by the fact that 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,

[14] 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 (https://www.bls.gov/opub/hom/home.htm).

which are a good predictor of work level within the NCS.[15] Our benefits imputation equations also include dummies for occupation, industry, establishment size, union coverage, full-time, calendar year, and for 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 where 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 non-mandatory benefits[16] and include non-mandatory benefits in our measure of total benefit costs.

[15]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.

[16] Non-mandatory 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. Non-production 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.

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 1-digit occupation. The latter variables are also all entered separately, along with dummies for calendar year. Dummies for 2-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 cost, but information on its cost is missing.[17] Omitting these observations would bias estimates of average benefit costs, as observations with positive cost would be omitted but not observations with zero costs.

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 .0986, while the R^2 for quadratic in linear benefits and wages is .0950. Thus our goal in the imputation is to use an estimate of E(ln B | Z) as a regressor in the quit regression, where Z is our vector of control variables in the imputation equation. Note too that $E(\ln B \mid Z) = E(\ln \sum B_i \mid Z) \neq \ln \sum E(B_i \mid Z)$ due to the non-linearity of the log function, so simply substituting predicted values of B_i for missing values will not yield a consistent estimate of $E(\ln B \mid Z)$. To obtain a consistent estimate, we also impute residuals; our imputations of missing values for B_i are thus $\hat{B}_i + e_i$. To more closely approximate normality, we first estimate a Box-Cox transform $C(B_i) = \frac{B_i^{\lambda_i} - 1}{\lambda_i}$

[17] Missing values are imputed in the NCS, but the imputed values have a much different (and weaker) relation to the covariates than reported values.

by maximum-likelihood, where $C(B_i)$ is assumed normally distributed, over observations where B_i is reported to be offered and the observation is valid.[18] We then regress $B_i^{\hat{\lambda}_i}$ (where $\hat{\lambda}_i$ is the estimated value of λ_i) on our covariates (excluding the interaction of the incidence dummy for benefit *i* with the other variables). Missing values are imputed as $(\hat{P}_i + e_i)^{1/\hat{\lambda}_i}$, where $\hat{P}_i$ is the predicted value of $B_i^{\hat{\lambda}_i}$ using the regression coefficients, and e_i is drawn from a $N(0, \sigma_i^2)$ distribution where σ_i^2 is the regression variance. These observations are then included in our NCS measure of (non-mandatory) benefits $B = \sum B_i$.

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 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 $(\overline{\ln \tilde{W}})$.

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

[18] A small number of these observations have a reported cost of zero even though the establishment indicates that it offers the benefit. We include these observations. To ensure positive values, we add .001 to all observations.

(12) $\ln B = Z\gamma + v$.

We use the coefficients from this regression to generate the predicted value $\overline{\ln B}$ 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 NLS data. We adopt a method similar to that for imputing missing values. We simulate distributions for the three benefits health insurance, pensions, and "other" consisting of all the other non-mandatory benefits. As we did in imputing missing values in the NCS data, we generate $\tilde{B}_i = (\hat{P}_i + e_i)^{1/\hat{\lambda}_i}$, where $\hat{P}_i$ is the predicted value of $B_i^{\hat{\lambda}_i}$ using the regression coefficients from the NCS and covariates from the NLS, and e_i is drawn from a $N(0, \sigma_i^2)$ distribution where σ_i^2 is the regression variance estimated from the NCS. For health insurance we generate $\tilde{B}_H$ for NLS observations with health insurance incidence. For pensions we generate $\tilde{B}_P$ for observations with pension incidence. As there are a large number of zeros in the NCS data even where pensions are offered, we also estimate a probit model in the NCS $\Pr(B_P > 0) = \Phi(Z_1\delta_p)$ and generate $\tilde{B}_P* = I(Z_1\hat{\delta}_P + u_P)\tilde{B}_P$ where $I(x)$ is an indicator function equal 1 if $x > 0$ and 0 otherwise and u_P is a random draw from a standard normal distribution. The vector Z_1 does not contain industry dummies due to small cell sizes for some industries. Other benefits $\tilde{B}_O$ are generated similarly. Total imputed benefits $\tilde{B} = \tilde{B}_H + \tilde{B}_P * + \tilde{B}_O$.

Conceptually similar issues arise for $\tilde{W}$. However, the R^2 for the regression of ln $\tilde{W}$ on Z_2 is .9962, so we treat the distribution of $\tilde{W}$ in the NLSY79 as being equivalent to the distribution of $\exp(\overline{\ln \tilde{W}})$.

It is important to note that other than the incidence dummies and their interactions, all variables used in the construction of $\overline{\ln B}$ are included in our quit regressions. Thus, identification of the effect of a dollar of benefits comes largely from the incidence dummies. The quit regressions also include a cubic in tenure at the time of the interview.

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 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 - 1993. Our NLSY79 sample is 23,119 observations from 7,178 different individuals over the same period. Note that NLSY79 sample member 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.[19]

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. 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.

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 four out of five are significant at the five-percent level (using a one-tail test).

The fourth column shows results for a specification using ln *B* and $\ln\tilde{W}$.[20] 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 all points in our sample. This finding is consistent with our theoretical model in Section II.

For comparison purposes, in the second and third columns we show results for specifications omitting benefits, using ln W in column 2 and ln $\tilde{W}$ in column 3. The

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[19] One typically finds that the incidence of managerial and professional jobs is higher in household than in establishment surveys. See Abraham and Spletzer (forthcoming).

[20] 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 datasets. 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.

coefficients in both columns are more than double the coefficient in column 4, 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 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

$$(18)\qquad \begin{aligned} E(Q \mid W, B, X) = {} & \beta_{1w}\overline{\ln(\tilde{W})} + \beta_{2w}\overline{\ln(\tilde{W})^2} + \beta_{1b}\overline{\ln(B)} \\ & + \beta_{2b}\overline{\ln(B)^2} + \beta_3\overline{\ln(\tilde{W})\ln(B)} + X\beta_x \end{aligned} .$$

For benefits the quadratic term $\overline{(\ln B)^2}$ is imputed as $\overline{(\ln B)^2} = (Z\hat{\gamma})^2 + (Z\hat{\gamma}_2)$, where γ_2 represents the vector of coefficients from a regression of the squared residuals from (12) on Z. (In cases where $Z\hat{\gamma}_2 < 0,$ we set $\overline{(\ln B)^2} = (Z\hat{\gamma})^2$.)[21] As before we ignore the deviation of $\overline{\ln \tilde{W}}$ from $\ln \tilde{W}$, so that the other quadratic terms are calculated as $\overline{\ln \tilde{W}}^2$ and $\overline{\ln \tilde{W}} \times \overline{\ln B}$. The three quadratic terms are jointly significant at the 5 percent level (p = .010).[22]

The interpretation of the coefficients in the quadratic specification is not transparent, but we may note that the 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

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[21] Note that $E(X^2) = E(X)^2 + Var(X)$, so $E((\ln B)^2 \mid Z) \geq E(\ln B \mid Z)^2$. We also experimented with estimating the expected value of the squared residuals as $\exp(Z\gamma_2)$. Point estimates of the quit equation were similar and the fit of the quit equation was identical. Bootstrap standard errors were not estimated due to computation time.

[22] 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 the specification in the text.

NLS sample.[23] 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 .12.[24] 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 4.8 percentage points while the marginal dollar of wages only reduces quits by 0.7 percentage points (the *t* statistic of the difference is 2.75). At 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 difference of 1.0 percentage points is not signficant at conventional levels ($t = 1.15$).

Another way of comparing between low and high values of wages and benefits is to divide the sample into halves by compensation $\tilde{W} + \tilde{B}$ (note that $\tilde{B}$ is simulated). Over 99 percent of each half of the sample has greater estimated effects for a marginal dollar of benefits. However, the *p* value for the half with compensation less than or equal to the median is .02, while the *p* value for the half with compensation above the median is .26. It is not surprising that it is easier to discern the larger effect of benefits on quits at lower compensation levels. The quit rate is 24.9 percent when $\hat{C}$ is less than or equal to the median and 12.8 percent when $\hat{C}$ is greater than the median. Thus to the extent that reductions in quits from higher benefits and wages are proportionate to the quit rate, the

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[23] 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 (18) 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 0.72.

[24] This *p* value is computed from the bootstrap distribution of the quadratic coefficients. Letting $\hat{\beta}_j$ denote the estimated vector of coefficients from bootstrap replication *j*, *j*= 1,…,200, for 12 percent of the replications the effect of benefits was less for a majority of the sample using $\hat{\beta}_j$ to estimate the effects.

difference in terms of percentage point reductions will be greater at lower compensation levels.[25]

The next to last column in table 2 shows the effect of adding demographic variables and other personal characteristics. Specifically, this regression includes age, highest grade completed, Armed Forces Qualifying Test (AFQT) score,[26] job experience at the start of the job, and dummies for female, black, Hispanic, and married. Wages and benefits both have a slightly smaller effect on quits, but there is essentially no change in their relative effects. 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 counts 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

[25] One obtains almost identical results if instead of using the simulated $\tilde{B}$ one takes predicted compensation to be $\hat{C} = \exp(\overline{\ln B}) + \tilde{W}$.

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

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 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.

Results for the quadratic specification are shown in column 6 of Table 2. These results are broadly similar to those in the previous specification. The difference at the 25^{th} percentile remains great, although there is some diminishing of the effect at the median. The proportion of the NLS sample for which the estimated effect of the marginal dollar of benefits is greater than the marginal dollar of wages is reduced to 96.3 percent. However, for sample respondents with compensation below the median the equivalent percentage is 99.1 percent, signficantly 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 non-pecuniary compensation such as more comfortable working conditions. Such non-pecuniary compensation, which can be thought of as unobserved fringe 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 benefit costs are much lower than wages and the effects are non-linear. In order to explain the larger marginal effect of benefits, unobservable non-pecuniary 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 NLSY 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 $I_i \overline{\ln B_i}$, where I_i is an indicator for benefit *i* (*i* = [pension, health insurance]) and $\overline{\ln B_i}$ 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 .308 (.381). The R^2 for the log of other benefits is .813.

We first discuss results from a specification quadratic in log benefits. (The *p* value of the quadratic terms is .051.) Standard errors are derived from 100 bootstrap replications. The quadratic-in-log-benefits specification is:

$$E(Q\,|\,W,B,X) = \beta_{1w}\overline{\ln(\tilde{W})} + \beta_{2w}\overline{\ln(\tilde{W})^2} +$$
$$\sum_i [\beta_{1i}\overline{\ln(B_i+.01)} + \beta_{2i}\overline{\ln(B_i+.01)^2} + \beta_{3i}\overline{\ln(\tilde{W})\ln(B_i+.01)}] + \sum_i\sum_{j>i}\beta_{4ij}\overline{\ln(B_i+.01)\ln(B_j+.01)} + X\beta_x.$$

As above, in imputing the quadratic terms variances around E(B_i|Z) are set to zero where regression predictions are negative. In addition, the cross-product terms for different benefits imply correlations between these benefits; where predicted cross-products imply correlations greater than one in magnitude they are set to the value implying a correlation of 1 or -1.[27]

Table 3 shows that it is difficult to disentangle the effects of individual benefits. Of the terms including the three benefits, only the terms for "other benefits" are jointly significant at the 10 percent level (p =.036), and wages are not significant. The standard errors for the marginal effects of the individual benefits are .030 or greater at the median and .071 or larger at the 25th percentile, quite large in this context. It is consequently not a surprise that at the median, none of the estimated effects of the marginal dollar of benefits are significant for any benefit (and similarly for wages); the effect of pensions is wrong-signed. At the 25th percentile, the effect of "other benefits" is significantly different from that of wages at the 10 percent level using a one-tailed test (t = 1.56). Symmetry of the terms involving the three different benefits is not rejected, nor is symmetry of the benefit terms with the terms for wages.

Consistent with the large standard errors, different specifications yield somewhat different results. A quadratic specification where benefits and wages are entered linearly rather than as logs yields an only slightly worse fit than the log specification (0.0982 vs.

[27] For a small number of observations the implied variance-covariance matrix of the three benefits is still not positive semi-definite even after correcting the variances and correlations between any two benefits; the average violation is small and these were left uncorrected.

0.0994). Relative to the log specification, the estimated signs change for health insurance and pensions at the median and 25th percentiles. Of the terms for the separate benefits only the terms for pensions are not jointly significant at the 5 percent level.[28]

In summary, the estimates for the effects of individual benefits are imprecise and volatile. Part of the reason for this is because 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. Appendix table 1 shows the coefficients on cross-benefit incidence dummies (and log wages) from our NCS 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.

These large correlations between benefits also imply that examining the effect of benefits individually may greatly exaggerate their effect on quits. Table 5 shows

[28] Similarly, a specification where the variance of benefit *i* around its expected value is estimated as $\exp(Z\gamma_{2i})$ in the first stage (as in footnote 24) yields an identical fit to 4 decimal places but a different sign for the effect of health insurance at the median and substantially different estimates for pensions.

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.[29]

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 wages and benefits effects. The differences between the estimated effects of wages and benefits at various points in the distribution are also larger than for quits, with differences of 9.3 percentage points at the 25th percentile and 3.2 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.

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 (18). The results, shown

[29] 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.

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 86.5 percent of the sample the marginal effect of benefits exceeds that of wages (with a 95 percent confidence interval of 78.5 to 88.5 percent).

V. Conclusion

It has been argued that one of the functions of fringe benefits is to reduce turnover. We 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.

Our estimates indicate that quits are much more responsive to an additional dollar of fringe benefits than to an additional dollar of wages. Consistent with our theoretical model, which predicts a positive association between benefits and turnover costs, we find that employers providing more training, who 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 NLSY9 means that we are not able to tease out the separate effects of the individual benefits.

While both the quit and training equations are consistent with the hypothesis that employers use fringe benefits to reduce quits, the estimation is reduced-form due to lack of an exogenous instrument for fringe benefits. Interestingly, however, adding demographic characteristics and an ability proxy to our quit equation has a fairly small effect on the estimates. Our quit regressions also control for major occupation, tenure, and employer size. While we cannot formally rule out the possibility that the reduction in quits associated with benefits may be caused by some omitted variable that is correlated with benefits, it is hard to imagine what that variable might be.

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Table 1

Descriptive Statistics

Variable	N	Mean	Standard Deviation	Min.	Max.
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 + .01) (Pension offered)	223,665	-1.46	1.59	-4.61	3.17
Ln (Health Insurance + .01) (Health insurance offered)	285,050	-0.51	1.16	-4.51	2.85
Ln (Other benefits + .01)	348,088	-0.92	1.59	-4.61	3.19
Union	348,392	0.16	0.37	0	1
Full time	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

Table 1 (continued)

Variable	N	Mean	Standard Deviation	Min.	Max.
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 (Health 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
Full time	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
Service	23,119	0.10	0.30	0	1
Tenure (weeks)	23,119	201.67	183.90	0	824
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

Table 2
Regression coefficients, quits

Ln Wages	-.045	-.055					
	(.008)	(.007)					
Ln Augmented Wages			-.059	-.025	-.074	-.076	-.079
			(.008)	(.010)	(.060)	(.060)	(.040)
Pension offered	-.015						
	(.008)						
Health Ins. offered	-.033						
	(.013)						
Life Ins. offered	-.013						
	(.010)						
Sick Leave offered	-.027						
	(.008)						
Vacation offered	-.026						
	(.012)						
Ln (Benefit Costs+.01)				-.026	-.031	-.030	-.030
				(.003)	(.019)	(.019)	(.014)
Ln Ben. x Ln Wages					.0056	.0052	.0051
					(.0063)	(.0063)	(.0047)
$(\text{Ln Wages})^2$					.0107	.0132	.0118
					(.0131)	(.0121)	(.0098)
$(\text{Ln Benefits})^2$					.0016	.0016	.0017
					(.0023)	(.0024)	(.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
Effect of Wages (or Augmented Wages) at median	-.005	-.010	-.007	-.003	-.003	-.002	-.004
	(.001)	(.001)	(.001)	(.001)	(.002)	(.002)	(.001)
Effect of Benefits at median				-.019	-.013	-.013	-.010
				(.002)	(.007)	(.007)	(0.005)
Difference in effect at median				-.016	-.010	-.010	-.007
				(.003)	(.008)	(.008)	(.006)
Effect of Wages (or Augmented Wages) at 25th percentile			-.010	-.004	-.007	-.006	-.007
			(.001)	(.002)	(.003)	(.003)	(.002)
Effect of Benefits at 25th percentile				-.054	-.048	-.047	-.048
				(.007)	(.013)	(.013)	(.010)
Difference in effect at 25th percentile				-.050	-.041	-.041	-.041
				(.008)	(.015)	(.015)	(.011)
R^2	.0987	.0952	.0933	.0979	.0986	.0996	.0986
n	23,119	26,169	23,119	23,119	23,119	22,985	23,119

Table 3

Effects of Individual Benefits on Quits, Quadratic-in-logs Specification

	p Value, Inclusion of variables	**Effect of marginal dollar on quits at:**	
		Median	25^{th} percentile
Wages	.279	-0.002	-0.005
		(0.003)	(0.004)
Health Insurance	.256	-0.023	-0.099
		(0.066)	(0.096)
Pensions	.170	0.098	0.725
		(0.218)	(2.174)
Other Benefits	.036	-0.025	-0.120
		(0.030)	(0.071)
R^2	.0994		

n = 23,119

Table 4

Correlations, Log wage and benefit incidence, NLSY79

	Ln Wage	Pension offered	Health Ins. offered	Vacation offered	Life Insurance offered	Sick Leave offered
Ln Wage	1					
Pension offered	0.31	1				
Health Ins. 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, NLSY79

	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

Table 5

Regression coefficients for health insurance and pension costs, quit regression.

Ln Health Insurance	-0.007	-0.020		-0.019
	(0.004)	(0.003)		(0.004)
Ln Pension	-0.001		-0.012	-0.003
	(0.003)		(0.002)	(0.003)
Ln Other Benefits	-0.027			
	(0.007)			
R^2	0.0983	0.0970	0.0944	0.0971
n	23,119	23,119	23,119	23,119

Table 6
Regression coefficients, turnover

Ln Wages	-.055 (.009)	-.071 (.008)					
Ln Augmented Wages			-.077 (.009)	-.021 (.011)	-.028 (.068)	.006 (.069)	-.075 (.045)
Pension offered	-.031 (.009)						
Health Ins. offered	-.045 (.014)						
Life Ins. offered	-.023 (.011)						
Sick Leave offered	-.029 (.009)						
Vacation offered	-.061 (.014)						
Ln (Benefit Costs+.01)				-.043 (.004)	-.071 (.022)	-.064 (.022)	-.056 (.016)
Ln Ben. x Ln Wages					.0123 (.0074)	.0115 (.0075)	.0078 (.0056)
$(\text{Ln Wages})^2$					.0037 (.0149)	.0012 (.0151)	.0119 (.0109)
$(\text{Ln Benefits})^2$					-.0011 (.0025)	-.0006 (.0025)	-.0002 (.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
Effect of Wages (or Augmented Wages) at median	-.005 (.001)	-.009 (.001)	-.009 (.001)	-.003 (.001)	-.001 (.002)	.002 (.002)	-.003 (.001)
Effect of Benefits at median				-.031 (.003)	-.033 (.008)	-.029 (.008)	-.023 (.006)
Difference in effect at median				-.028 (.004)	-.032 (.009)	-.031 (.009)	-.020 (.007)
Effect of Wages (or Augmented Wages) at 25th percentile			-.013 (.002)	-.004 (.002)	-.004 (.003)	.000 (.003)	-.007 (.002)
Effect of Benefits at 25th percentile				-.089 (.008)	-.097 (.015)	-.089 (.015)	-.087 (.011)
Difference in effect at 25th percentile				-.085 (.009)	-.093 (.017)	-.090 (.017)	-.080 (.012)
R^2	0.1478	0.1407	0.1368	0.1464	0.1469	.1513	.1467
n	23,119	26,169	23,119	23,119	23,119	22,985	23,119

Table 7

Regression coefficients, log (training hours + 1), current employer in previous year

Ln Augmented Wages	0.487
	(0.154)
Ln (Benefit Costs+.01)	0.083
	(0.056)
Ln Ben. x Ln Wages	0.014
	(0.020)
$(\text{Ln Wages})^2$	-0.089
	(0.034)
$(\text{Ln Benefits})^2$	0.020
	(0.006)
Effect of Augmented Wages at median	0.015
	(0.005)
Effect of Benefits at median	0.091
	(0.020)
Difference in effect at median	0.076
	(0.023)
Effect of Augmented Wages at 25th percentile	0.028
	(0.007)
Effect of Benefits at 25th percentile	0.164
	(0.039)
Difference in effect at 25th percentile	0.137
	(0.043)
R^2	.0553
n	24,337

Appendix

Data on employee contributions to health plans are available from the Employee Benefit Survey (EBS) for 1993 for large and medium establishments (establishments of greater than 100 employees) and 1994 for small establishments. The EBS is conducted jointly with the ECI survey used in the main empirical work. The unit of observation is the health plan. Each plan has a single and family option with associated single and family employee monthly contribution. In the combined 1993-94 data, 57,577 out of 65,485 plans in the EBS data can be matched to observations in the NCS data. These plans are matched to 12,243 benefit observations from 2,182 employers. Observations are matched to the nearest quarter to the EBS reference month.

Plans in the 1993 data come in four categories: medical, dental, vision, and prescription. The prescription category is omitted in the 1994 data. The categories are not mutually exclusive—a plan can be in multiple categories. Over 99 percent of plans with multiple categories are medical along with some other type or types. Each plan has a participation rate associated with it representing the proportion of employees in the job choosing that plan. These rates do not distinguish between choice of the single or family option. Because of the separate categories, participation rates can sum to more than one. Because some employees do not choose a plan (presumably mostly those covered by a spouse), participation rates can also sum to less than one; there is no separate participation rate for "declined all plans". No information on cost is available for plans where the amount per month is not fixed.

There are two basic steps to our imputation procedure. First, we estimate a value of employee contributions for observations with matches to our NCS data. Second, we impute values of employee contributions for all our NCS observations with health insurance offered.

Estimating employee contributions:

We can distinguish four types of plans:

1. No employee contribution.
2. Fixed employee contribution with data available.
3. Employee contribution not fixed or unknown positive contribution.
4. Unknown if there is employee contribution.

If all plans were of types 1 and 2 (with either zero or known positive contributions), average employee contributions for a quote would simply be $\sum r_i c_i$, where *r* denotes participation rate, *c* denotes employee contribution (including zero), and *i* denotes plan. With unknown contributions, we need to impute.

A plan's category or categories (medical, etc.) conveys information on employee contribution. We simplify by assigning each plan to a single category. All plans with the medical indicator checked are considered medical. For plans without the medical indicator, plans with the dental indicator are considered dental. Next in priority is vision, and last is prescription. Aside from medical plans, of which 96 percent have other indicators, close to 99 percent of other plans have only a single indicator checked. (When categorized this way, participation rates for all plans summed within their category sum to one or less, with the exception of the medical category for 76 out of 42,000 plans.)

With this preliminary, the formula for employee contributions for a given job is:

(A1) $$E = \sum_j \frac{\sum_i r_{2ij} c_{2ij} + r_{3j} \left(\sum_i r_{2ij} c_{2ij} \Big/ \sum_i r_{2ij} \right)}{r_{1j} + \sum_i r_{2ij} + r_{3j}}$$

where *i* denotes plan and *j* denotes category (medical, dental, vision, prescription), r_{1j} is the sum of the participation ratios for plans of type 1 above (i.e., no employee contribution) in category *j*, r_{2ij} is the participation rate for plan *i* of type 2 in category *j*, and r_{3j} is the sum of participation ratios for type 3 (employee contribution not fixed) plans of category *j*. The above formula implicitly imputes contributions for type 3 plans as the average for fixed positive employee contributions (type 2 plans) in the category and contributions for type 4 plans as the average inclusive of plans with no contribution (type 1 and type 2 plans).

If there are no plans of type 2 for a given job and category, but plans of type 3 exist, E_j (where $E = \sum_j E_j$) is imputed as the average predicted value for that category across plans from a regression of observed contributions on the log of establishment size, full-time status, union, and the plan's participation rate. If there are only type 4 observations (no information about employee contributions) for a particular combination of job and category, values are predicted from a regression on log employment, union, and full-time status using the estimated E_j from other observations.

The foregoing ignores the distinction between single and family plans. As stated, the proportion choosing the family option is unavailable in the data. We use a figure from Levy (1998), who uses March CPS data and finds that approximately 55 percent of workers covered through their employer chose the family option in this period. Accordingly, our estimate is:

(A2) $E = .55\, E_F + .45\, E_S$,

where E_F denotes (A1) applied to the family option and E_S the same for the single option.

Imputing values for observations in the NCS:

Most of our NCS observations are in years that do not match the plan data. Presumably employee contributions will track employer contributions. Accordingly, we estimate the proportion of total health insurance expenditures that is paid by the employer.

For each observation we match, total hourly expenditures are $H = hE + F$, where h is a factor relating monthly expenditures to hourly costs and F is the employer's hourly contribution. (We set the factor h converting monthly to hourly expenditures at 12/2000 for full-time quotes and 12/1000 for part-time). The proportion paid by the employer is $P = F/H$. We regress P on our usual vector of covariates, excluding year dummies, to generate a predicted value $\hat{P}$, and impute total expenditures as $\hat{H} = F / \hat{P}$. The imputed $\hat{H}$ is used as a summand for total benefits in the first stage, and the employee contribution is subtracted from augmented wages.

Appendix 2

Table

Coefficients on benefit incidence, log benefit costs, NCS data

	Ln Pension*	Ln Health Insurance*	Ln Other Benefits
Pension offered		0.369	0.248
		(0.043)	(0.026)
Health Ins. offered	-0.058		0.492
	(0.151)		(0.053)
Vacation offered	0.000	-0.019	1.588
	(0.172)	(0.205)	(0.066)
Life Insurance offered	0.113	0.309	0.320
	(0.106)	(0.058)	(0.032)
Sick Leave offered	-0.065	0.132	0.287
	(0.074)	(0.038)	(0.030)
Ln Wage	1.013	1.347	1.336
	(0.221)	(0.180)	(0.137)
(Ln Wage) squared	0.059	-0.145	-0.058
	(0.045)	(0.037)	(0.028)
n	223,665	285,050	348,088

* Sample restricted to respondents offering benefit.