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ABSTRACT

The Changing Distribution of Job Satisfaction*

The distribution of job satisfaction widened across cohorts of young men in the U.S. between 1978 and 1988, and between 1978 and 1996, in ways correlated with changing wage inequality. Satisfaction among workers in upper earnings quantiles rose relative to that of workers in lower quantiles. An identical phenomenon is observed among men in West Germany in response to a sharp increase in the relative earnings of high-wage men in the mid-1990s. Several hypotheses about the determinants of satisfaction are presented and examined using both cross-section data on these cohorts and panel data from the NLSY and the German SOEP. The evidence is most consistent with workers' regret about the returns to their investment in skills affecting their satisfaction. Job satisfaction is especially responsive to surprises in the returns to observable skills, less so to surprises in the returns to unobservables; and the effects of earnings shocks on job satisfaction dissipate over time.

JEL Classification: J28

Keywords: Job satisfaction, wage inequalitiy returns to human capital, earnings shocks

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In the end, economics is not about wealth -- it s about the pursuit of happiness (Krugman, 1998).

I. Introduction and Rationale

A rapidly growing literature has identified rising earnings inequality as the premier labor-market problem of the 1980s and 1990s in the United States (e.g., Bound and Johnson, 1992; Juhn et al, 1993; OECD (1996), and its importance in other countries has also been examined.[1] Recently some attention has been paid to the changing distribution of nonwage monetary returns to work (Pierce, 1998); and the changing distribution of a variety of nonmonetary aspects of work has also been analyzed (Hamermesh, 1999a). All of these studies concentrate on some part of what workers derive from their jobs, and the student of inequality can use them to piece together a jigsaw puzzle describing much of the total change in inequality of the returns to work. But any study of nonwage monetary or even nonmonetary returns will necessarily ignore some nonpecuniary aspects of these returns. What is needed is an understanding of how the overall distribution of the returns to work has changed.

Only one measure, the satisfaction that workers derive from their jobs, might be viewed as reflecting how they react to the entire changing panoply of job characteristics. As our theoretical arguments and empirical analyses will demonstrate, even this measure cannot in the long run reflect changing inequality in the overall returns to work. In the short run, however, it can provide a reflection of all the returns to labor-market activity, and as such can allow us to infer whether and how the relative well-being of different kinds of workers has changed in response to shocks to the labor market.

Economists have traditionally been loathe to deal with subjective outcomes describing work, feeling that these cannot be linked to any underlying concept of utility and that, even if they could, their subjective nature renders them too noisy to be of analytical value. We have not, however, been entirely aloof from an area that has chiefly been the domain of social psychologists. Hamermesh (1977) constructed and tested a theory of overall

[1]See OECD (1996) for an examination of time-series patterns in wage inequality in a number of developed economies.

job satisfaction; and Freeman (1978) and Borjas (1979) examined the effects of unionization on overall job satisfaction. More recently Farber (1990) continued the analysis of unions and job satisfaction, and Clark and Oswald (1996), Gerlach and Stephan (1996) and others have considered cyclical variations in satisfaction.

In this study I examine how the distribution of satisfaction from work has changed over time in the United States and elsewhere. The first part of the empirical analysis takes subsamples of male workers ages 26-31 in 1978 from the National Longitudinal Survey cohort of young men (NLSYM) who were asked questions about their satisfaction with their jobs and compares them to men the same age in 1988 who were asked the same question in the National Longitudinal Survey of Youth (NLSY). In the second part of the study I control for the heterogeneity in workers' satisfaction that is inherent in the first part by following the job satisfaction of members of the NLSY over a long period (1984-96) as they matured and as the distribution of their pay changed, presumably in part unexpectedly. In both of these empirical analyses we expect to observe a widening of the distribution of satisfaction, one that is correlated with the growing inequality of pay and of nonwage benefits and nonpecuniary amenities. As a check on the validity of this approach, a third set of empirical analyses examines longitudinal data on job satisfaction and pay inequality in Germany from 1984 to 1996.

II. Motivation

Job satisfaction is a slippery concept, one of those subjective outcomes that many economists view as outside the purview of economic analysis. If job satisfaction were a monotonic transformation of full income, and thus a true measure of utility, it would be extremely interesting for use in welfare comparisons. Regrettably, as I argue, that is not likely to be the case. Despite that disappointment, however, while it is not likely to indicate their well-being in any sense that is useful for an outsider evaluating welfare, how workers perceive their work does affect economic outcomes. A more satisfied worker, even one whose economic situation appears to an outsider to be no better than that of otherwise identical workers, is less likely to leave his/her job voluntarily. Depending upon current job satisfaction, a worker will be more or less likely to invest in firm-specific human

capital that will increase his/her commitment to the employer (with the same argument applying to occupation- or industry-specific investment). Indeed, one might even reasonably imagine that the fluctuations in the animal spirits that are a major Keynesian motivation for business cycles arise in part from variations in workers perceptions of their well being. Presumably more satisfied workers, who are secure in their jobs, have a reduced motive to undertake precautionary saving. In short, even though it may be a substantial stretch to link expressed job satisfaction (or satisfaction generally) to utility, studying job satisfaction is still important for understanding labor-market behavior and perhaps economic activity more generally.

All of the available sets of data describe job satisfaction (JS) as a categorical response that presumably maps the worker s underlying feelings about his/her job to a few discrete choices. Assume that there are J such choices, and let S be a continuous index of the worker s satisfaction. Then we will observe worker i s responses on JS as:

(1) $JS_i = J$, if $S_{it} > S_J$;
$JS_i = J-1$, if $S_{it} > S_{J-1}$;
.
.
.
$JS_i = 1$, if $S_{it} \leq S_2$.

Recognizing that this is a tremendous simplification, to make the notation here easier I assume that S_{it} is a linear function of its argument(s).[2] In each case I assume that these arguments are measures of objective characteristics of the job.

The determination of job satisfaction depends on workers expectations about their current earnings and working conditions. Consider a series of cases describing workers perceptions.

[2]All that is needed is that S be a monotonic transformation in its arguments.

I. Complete forgetfulness and complete surprise. Workers care only about their current full earnings E_{it} and make no comparisons based on their characteristics (or compare themselves only to the average worker) or on their past histories or those of other workers. In this case:

(2a) $S_{it} = E_{it}$.

The transformation from current earnings to job satisfaction is monotonic, and comparisons of job satisfaction are indicative of utility comparisons. In reality even in this case the transformation from wages to S will be nonlinear: Although we expect wages and amenities to be related to a worker s full earnings, the demands for each are not unit income-elastic. Indeed, given evidence that the demand for job-market amenities is quite income-elastic (Hamermesh, 1999a), one might expect the distribution of job satisfaction to be wider than that of wages. Under this assumption greater (lesser) wage inequality will generate correlated increases (decreases) in the inequality of job satisfaction.

II. Knowledge of average current rates of return. This assumption stems from expectancy theory (stated by Lewin, 1938, and summarized nicely by Vroom, 1964). Hamermesh (1977) showed that this approach to the determination of job satisfaction dominated empirically the assumption implicit in (2a) in a cross-section of American workers in 1969.[3] If workers compare themselves to others who had made the same investments at the same time, differences in satisfaction will arise out of heterogeneity in the returns to those investments. Only supernormal returns and quasi-rents will generate higher job satisfaction. In a cross section this implies that job satisfaction will be related to the residual from equations relating earnings to measures of investment in general human capital, and to measures of quasi-rents (e.g., returns to firm tenure). The satisfaction index is determined as:

(2b) $S_{it} = E_{it} - E^*_{it} \mid X_i$,

[3]A journalistic description of the problems of managers and professionals earning only $100,000 to $200,000 (Wall Street Journal, August 3, 1998, p. 1) illustrates the role of expectancy theory in affecting satisfaction.

where $E^*_{i\,t} \mid X_i$ are the full earnings of the average worker with characteristics X identical to those of worker i.

Increasing (decreasing) inequality in the distribution of earnings over time will increase (decrease) the dispersion in the distribution of job satisfaction only if the change in inequality results from greater dispersion in the returns to unobservable skills. This requires separating out changes in this distribution from those in wages generally (see Di Nardo et al, 1996).

III. Disappointing returns. As is implicit in Lévy-Garboua and Montmarquette (1998), a rational individual will base his/her job satisfaction on comparisons of outcomes to expectations that were formed at the time that the investment decisions that generated those returns were undertaken. Thus in this view job satisfaction is determined by the worker comparing his/her full earnings to what would have been expected upon entering the labor market at time t_0, having made the investments the worker made and with the returns to the worker s other characteristics. No extra satisfaction is generated by unusually high or low returns received by heterogeneous workers if those were expected at t_0; but temporal changes in the means and variances of the distributions of returns will alter the distribution of job satisfaction, so that:

$$(2c) \qquad S_{i\,t} = E_{i\,t} - {}_{t_{i0}}E^*_{i\,t} \mid X_i ,$$

where ${}_{t_{i\,0}}E^*_{i\,t} \mid X_i$ is the full earnings that a worker with i s characteristics would have expected to receive at time t had labor-market conditions remained as they were when i entered the labor market at time $t_{i\,0}$.

Admittedly this is backward-looking, in the sense that the worker retains at time t full memory of the expectation at $t_{i\,0}$; but it is consistent with substantial evidence in the literature in psychological economics of the role of regret in affecting behavior (for example, Thaler, 1992, Chapter 6). Rising (declining) earnings inequality will raise (lower) the variance of the distribution of job satisfaction to the extent that today s workers

formed their expectations when the distribution of earnings was different. This will persist until all the workers who entered the labor market at a time when earnings inequality differed from its current state have retired.

IV. Rational expectations with learning and heterogeneity. It is difficult to believe that regret lasts an entire working life: If the demand for a worker s skills dropped unexpectedly and permanently in the first year of the stream of returns, one might expect that worker to be more regretful in that year than thirty years later. After some period of time the worker s expectations of future returns may have adjusted to the likely reality. This consideration suggests that job satisfaction will be determined by the deviation of the returns to the worker s skills over a continually adjusting forecast of those returns. The forecast will be based on what the worker expected at t_{i0}, but it will be modified by events. The satisfaction index will be determined as:

$$(2d) \qquad S_{it} = E_{it} - \sum_{k=t_0}^{t} {}_k E^*_{ik} \,\Big|\, X_i ,$$

with the weights ${}_k$ presumably greater as $k \to t$. In this case increasing (decreasing) inequality of earnings can increase (decrease) the dispersion of job satisfaction, but only if the changes in inequality are unexpected. If earnings inequality is greater and has been so for some years, under this view the distribution of satisfaction will approach what it was before the initial shock occurred.

We can go beyond the four Cases to examine how job satisfaction responds to unexpected returns to investment in human capital as compared to unexpected returns to a worker s unobservable (to the econometrician) characteristics. There is no reason to expect workers to react the same way to changes in the returns to their investments in education and on-the-job training as they do to changes in the returns to their native capacities, since their ability to perceive these two sources of the wage outcomes may differ. Quite possibly (2c) and (2d) should also be decomposed into two components, one showing the deviation of returns to those measurable skills in which the worker had invested, the other showing the deviation in returns to unobservables.

Also in (2d) the paths of adjustment of expectations (the $_k$) of the responses to surprises need not be the same for both components of the unexpected returns.[4]

III. Cohorts of American Men

A. Data

The ages of the respondents and the timing of information on job satisfaction in the National Longitudinal Survey of Young Men and the National Longitudinal Survey of Youth fortuitously enable us to compare workers within a narrow, fairly young age range just before and immediately after the most rapid increase in earnings inequality in the past 75 years occurred in the United States. I take nonhispanic white male workers 26-31 in each survey, with those in the NLSYM observed in 1978 (one of only two years in which questions about job satisfaction were asked in that study) and those in the NLSY observed in 1988. In both surveys workers were asked the question, How do you feel about your job? Do you like it: very much (4); like it fairly well (3); dislike it somewhat (2); dislike it very much (1)? [5] Because the last response is very rare (1.73 percent in 1978, 1.51 percent in 1988), and the third is not common, I combine these two categories in the subsequent analyses here and in Section IV.

The sample used in most of this section consists of all men in the required range of ages whose annual hours of work averaged at least 35 per week in the year preceding the survey (so that I restrict the analysis to full-time workers). This restriction eliminates 10 percent of working men of this age in 1978, and 14 percent of the sample in 1988. I use annual earnings (for 1977 in the NLSYM, 1987 in the NLSY) as the measure of wages in the comparisons in the remainder of this section and in Section IV to reduce problems of measurement error (see Bound et al, 1999).

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[4]The theoretical issue throughout this discussion is essentially one of examining the extent of cognitive dissonance between where the worker currently is and what he/she perceives the correct place to be (see Festinger, 1957). The empirical issue is the determination of what that perception is and how it evolves.

[5]The responses are coded in the data in reverse order. I have transformed them here to make the ordering consistent with general usage.

B. The Relationship Between Changing Earnings Inequality and Job Satisfaction

Table 1 presents the distributions of the crucial variables in this analysis, reported job satisfaction and actual annual earnings. The data on earnings corroborate for these very narrow age groups the well-known results on changing earnings inequality: The distribution widened substantially during the 1980s, as shown by the changing ratios of W_p / W_{50}, annual earnings at percentile p relative to median annual earnings at the time. Moreover, as comparisons of the second part of the table show, the greatest change that occurred was the result of rapid growth of earnings at and within the top decile of the distribution.

The first thing to note from the distribution of job satisfaction is that there is no apparent evidence that average satisfaction increased during this decade. Indeed, if anything there is some indication that the responses became more negative during this period. Other than Case I the explanations of job satisfaction outlined in Section II do not imply any necessary trend in the responses, and even Case I might be interpreted as requiring only comparisons to other workers' actual wages at each point in time. Thus the absence of an upward trend in satisfaction during a time of rising real wages is not surprising. The Table also shows no indication of a widening dispersion in the distribution of job satisfaction. This too is not surprising, as the arguments in Section II imply that we should observe changes in the distribution of job satisfaction of workers arrayed by their earnings, not necessarily in the raw distribution of satisfaction. (For example, if average earnings rise steadily, and some shock lowers the return to skill, the earnings distribution will narrow even though under Case III and, at least until expectations adjust, also under Case IV we would observe the distribution of satisfaction to have widened.)

We cannot test the predictions of Section II even indirectly without examining how job satisfaction changed at different points of the widening distribution of earnings. For that purpose I array the respondents by annual earnings (by quartiles, to circumvent the random variation induced by the small sample sizes within deciles) and examine how average job satisfaction changed over time within each quartile. Figure 1 presents:

$$(3) \qquad {}^{2}_{qj} = \left\{[f_q = j] - [f_1 = j]\right\}_{88} - \left\{[f_q = j] - [f_1 = j]\right\}_{78}, \; j = 4, 3,$$

where f indicates a fraction and q denotes a particular quartile. In all cases over 40 percent of the sample members like their jobs very much, and over 40 percent like them fairly well, so that the most useful comparison is of how these two fractions differ over time by earnings quartile.

Figure 1 presents ( ) the double-differences in the logarithm of average earnings by quartile compared to the lowest quartile of workers sorted by earnings. As was implied by the data in Table 1 at various percentiles, these double-differences are positive, with the largest value being that for the double-difference between the top and bottom quartiles. During this time of rising earnings inequality the fraction of workers in each of the highest three earnings quartiles who said that they liked their jobs very much ( ) increased relative to the change in the lowest quartile, while the fractions stating they only liked their jobs somewhat ( ) decreased relatively. Moreover, these changes were greatest in precisely that part of the earnings distribution, the top quartile, that saw the biggest relative increase in earnings. These results are consistent with Cases I, III and IV in Section II, and with Case II to the extent that the rising inequality of returns to unobservables accounts for part of the overall increase in earnings inequality.

C. Distinguishing the Cases

To determine which of the alternative explanations of job satisfaction that I outlined in Section II describes these data better, we need to move from those abstract models to an estimable model. Ideally I would like to approximate E_{it}, $E^*_{it} \mid X_i$ and ${}_{t_{i0}}E^*_{i\,t} \mid X_i$ by measures that account for a large variety of returns to work, not just by wages. Regrettably, good measures of the nonwage monetary returns to work are not available in this data set; and obtaining even incomplete measures of just some of the nonpecuniary aspects of these workers jobs is very difficult. Accordingly, I proxy E by the worker s wage, W, throughout the empirical work. While this neglects substantial aspects of the returns to work, the wage, the nonwage pecuniary returns, and nonpecuniary amenities are all increasing functions of full earnings (none is inferior). Thus one can view W as a monotonic function of E whose elasticity probably is not unity, and need not even be constant.

Table 2 shows the parameter estimates describing various characterizations of the cases in Section II using the data on job satisfaction for young men ages 26-31 in 1988. I present the results of estimating these ordered probits as the impact of a ten-percent increase (actually 0.1 log points) in the independent variable on the probabilities of being in each of the three job-satisfaction categories (since by themselves the coefficients from ordered probits are uninformative (Hamermesh, 1999b)). Thus the first coefficient in each triad is $.1\partial \Pr\{JS_t = 4\}/\partial Z\}$, and the second is $.1\partial \Pr\{JS_t = 3\}/\partial Z\}$, where Z is a wage measure. The sign of the t-statistic in parentheses below the derivative indicates the sign of the coefficient in the ordered probit.

The coefficients in the first row are from an ordered probit in which the only independent variable is the worker s annual earnings (in 1987), as in (2a). Those in the second row are on the residual u_{88} from a regression of annual earnings reported in 1988 (for 1987) on education, actual job experience, job tenure, annual hours, and indicator variables for marital status, location in the South and in an SMSA, union status, and bad health. This ordered probit is the econometric analog of (2b). Unlike in Hamermesh (1977) actual earnings describe the distribution of job satisfaction better than does the residual from a cross-section earnings equation.

Neither of these models describes job satisfaction in 1988 as well as does a simple version of Case III, in which I try to capture the idea of regret/disappointed expectations. The estimated effects from the ordered probits that are presented in the third row of Table 2 are on annual earnings in 1988 and on $_{78}\hat{W}_{i,88}$, what the average earnings in 1988 of a worker with individual i s observable characteristics would have been had the distribution of returns to someone with those characteristics remained unchanged between 1978 and 1988. This equation is thus essentially identical to (2c), except actual earnings replace full earnings. By using $_{78}\hat{W}_{i,88}$ to proxy $_{ti\,0}E^*_{i\,t} \mid X_i$ I am implicitly assuming that when they entered the labor market in (roughly) 1978 those

workers who were 26-31 in 1988 based their expectations about their labor-market success at those ages on the experiences of workers who were 26-31 in 1978.[6]

This model performs better than the simpler models describing Cases I and II: Not only are the pseudo-R^2 higher and each of the coefficients above its standard error; the coefficient on ${}_{78}\hat{W}_{i,88}$ has the expected negative sign. The results suggest that, if a worker s current wages are above (below) what his observable skills would have generated based on his expectations when he entered the labor market, he is unusually satisfied (dissatisfied) with his job. The importance of this variable suggests that workers do exhibit some regret in their feelings about their work. Comparing the results to those of Case II implies that workers are more concerned with how the labor market treats them relative to their initial expectations than how it treats them relative to its current treatment of other workers with the same (observable) skills.

While the data from these two cross-sections of young men do not allow distinguishing between Cases III and IV, they can be used to examine the relative importance of regrets about the returns to observable and unobservable characteristics in the determination of job satisfaction. The remaining effects presented in Table 2 are from ordered probits that include as independent variables combinations of measures that decompose earnings shocks into those to the returns to the observable and unobservable components of workers skills. The first term is $[u_{88} - \hat{u}_{p,78}]$, the residual from an equation describing earnings in 1988 minus what that residual would have been at the p th percentile of the earnings distribution if the distribution of observed earnings had remained unchanged since 1978. The second term is $[\hat{W}_{88} - {}_{78}\hat{W}_{88}]$, the earnings predicted in 1988 for a person with worker i s characteristics minus what would have been predicted for those characteristics in 1978. The first term thus

[6]In all of these equations I experimented with including other variables, among them each of the measures included in the equations used to generate u_{88}. Among these none achieved a t-statistic above 1 in absolute value. Including a vector of indicator variables for one-digit industry generated only one new result, namely that workers in professional services were more satisfied (given their wages, or the residuals of their wages) with their jobs than were other workers. Even here, this effect was only marginally significant statistically. The inclusion of these other variables never lowered the statistical significance of the wage measures.

measures how well the worker was rewarded in 1988 relative to otherwise observationally identical workers then, compared to the returns that those unmeasurable skills would have generated in 1978. The second term measures the gain (or shortfall) of the returns to his/her measurable skills (demographics, education, years of tenure, union status, etc.) relative to what could have been expected in 1978.

The fourth and fifth rows of Table 2 present the results of including these two measures separately, while the sixth set of ordered-probit coefficients includes them both. Clearly, this decomposition of the measures that are essentially summed in Case III matters, since the pseudo-R^2 rises. More important, the estimates suggest that the unanticipated returns to a worker s observable characteristics -- education, job tenure, etc. -- are the more important determinants of job satisfaction. Workers are either less aware of the unexpected returns to their idiosyncratic characteristics (those parts of their drive, innate ability, etc., that are not determinants of the education and job tenure that they accumulate); or the returns to these unobservables simply do not affect how they perceive their jobs.[7]

The final equation in Table 2 includes an interaction of $[u_{88} - \hat{u}_{p,78}]$ and $[\hat{W}_{88} - {}_{78}\hat{W}_{88}]$ to test whether unusually high (low) unexpected returns to both observables and unobservables interact to generate unusually high (low) job satisfaction. With the inclusion of this interaction term all three coefficients are significantly nonzero; and the results suggest that unusual returns along both observable and unobservable dimensions affected workers job satisfaction in 1988. Between 1978 and 1988 the returns to both types of skills rose; and those workers who were fortunate enough to have accumulated a lot of education and who were highly able and driven reaped the benefits of those apparently unexpected higher returns and were especially well satisfied with their jobs.

D. Cohorts of Men 31-33, 1978 and 1996

[7]While I present only the results for nonhispanic whites in Table 3, estimates for the entire sample of males 26-31 years old were generally comparable and demonstrated the superiority of Case III over Cases I and II as an explanation of job satisfaction in 1988.

Only longitudinal data can both circumvent the problems of heterogeneity and enable us to distinguish empirically between Cases III and IV; but using a pair of cross sections still further apart than the decade that separates the workers in the samples on which the estimation in Parts B and C was based can at least provide some hint of whether workers' expectations about their treatment in the labor market do adjust to the realities that confront them as their careers progress. I take 31-33 year-old nonhispanic white males observed in 1978 in the NLSYM and compare them to men in the same age group in the NLSY observed in 1996. The sample for 1978 contains the same men age 31 who were also included in the data used previously in this section. I reuse these observations in order to prevent an already small sample (400 observations in 1978, 627 observations in 1996) from becoming smaller still.

Figure 2 shows the double differences in the logarithms of annual earnings by quartile. The results indicate unsurprisingly that over this nearly twenty-year period the distribution of earnings in this narrow age range became less equal, with the biggest change being the sharply widening gap between workers in the top quartile of earnings and others. The other two lines in Figure 2 present the double differences in the fractions of workers responding that they are very satisfied or somewhat satisfied with their jobs. These results are directly comparable to those in Figure 1 for slightly younger nonhispanic white men observed over the decade 1978-88. As in that Figure the double differences on JS = 4 are positive, suggesting that even over this longer period workers' expectations about their labor-market success did not fully adjust to the change in outcomes.

A comparison of the results in the two Figures reveals one crucial difference, however: The largest positive double difference in Figure 1 was in the top quartile, mirroring the very sharp rise in relative earnings in that quartile compared to others. In Figure 2, even though the top-bottom quartile double difference in log-earnings over the eighteen-year period far exceeded that in Figure 1 and is the largest in this Figure, the double difference in JS in the top quartile is smaller than those in the third and second quartiles. While the changing distribution of job satisfaction matched the changing distribution of earnings (by quartile) perfectly between 1978 and 1988, the match was less complete, although still in the expected direction, between 1978 and 1996. This

distinction provides evidence that there was at least some adjustment in workers' expectations about the returns to their skills after they entered the labor market, and it thus hints that Case IV may describe the path of job satisfaction better than Case III.

IV. Longitudinal Changes in Job Satisfaction When Earnings Inequality Widens -- U.S. 1984-96

The attempt to distinguish among the theories of job satisfaction is plagued by problems of heterogeneity: Those individuals who are inherently more satisfied may also be those who invest more in themselves, whose wages are higher (or higher than an econometrician would expect) and whose wages grow more rapidly once they cease formal education. The ordered probits describing Cases I-III may thus have been misspecified because they ignored a common component in both the job satisfaction measure and the wage measure(s). For example, (2a) might more appropriately be specified as:

(2a') $S_{i\,t} + v_i = E_{i\,t} + \quad _i$,

where v_i and $\quad _i$ are individual effects, with $E(v_i)=E(\quad _i)=0$, and $E(v_i \quad _i)>0$. The $v_{\,i}$ and $\quad _i$ might be viewed as similar to positively correlated measurement errors that are unchanging over time. Clearly, estimation based on (2a') or its analogs will impart a positive bias to the impact of earnings measures on satisfaction because of heterogeneity rather than any causation from E to S. This difficulty has no effect on our inferences about the changing distribution of job satisfaction, but it does prevent us from drawing structural inferences.

To circumvent this potential difficulty we need to examine longitudinal data that allow us to difference out the v_i and $\quad _i$. The NLSY is the only American data set that contains repeated observations on large numbers of individuals' job satisfaction, and I use it again in this section. Since W is represented by annual earnings, I again restrict the sample to full-time workers (in this case, those who worked at least 1500 hours in a year). In order to avoid observing the sample members during a time when many were acquiring additional education I start the panel only in 1984.[8] Moreover, to prevent sample sizes from becoming minuscule because of the restriction

[8]By 1984 the number of sample members whose education was still increasing grew much more slowly than before. Indeed, while the fraction of the sample with 16+ years of schooling rose from 12.2 percent in 1982 to 18.2 percent in 1984, it grew to only 22.9 percent in 1990 and to 22.7 (lower because of differential attrition) in

to full-time employees, I examine the panel members' job satisfaction over the twelve-year period 1984-96, but only in the three years 1984, 1990 and 1996. (Three years of data should suffice to discriminate between Cases III and IV.) This restriction and the requirement that information be available on satisfaction in each year and on the observables used to predict W reduce the sample to 1280 individuals.

I estimate standard log-earnings equations for each of t = 1984, 1990 and 1996 to obtain u_t for use in estimating empirical analogs to (2b). In each case the explanatory variables include the usual human capital measures (education, actual experience and job tenure), indicators of union status, a variety of demographic measures, and indicators of location, with separate earnings equations estimated each year for men and women. For each sex the earnings equations for the three years are estimated as a system of seemingly-unrelated equations. The more difficult choice is the estimation of $_{t_0}\hat{W}_t$. An early cross section of the NLSY cannot be used for this purpose, since the age range in any cross section is too narrow to allow within-sample prediction of the growth of earnings in the NLSY over the twelve-year period that we use.

The best alternative is to predict earnings growth in the NLSY sample by examining a cross section of workers from the CPS Outgoing Rotation Groups on whom we have some of the same information as in the NLSY. This approach has the virtue of providing very large samples covering the entire age range of full-time workers, but the disadvantage that it lacks the richness of information in smaller samples such as the NLSY. Nonetheless, it captures the majority of the determinants of earnings differences (especially among full-time workers) that we observe in these smaller samples. I use the 1979 CPS-ORG, because that year roughly corresponds to t_0, the time when the NLSY panel members might have been forming their expectations about their profiles of future earnings. Using the CPS-ORG I estimate the log-earnings equations for 1979, the results of which are presented in Appendix Table A1. The estimated coefficients are completely unsurprising and require

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1996.

no comment here. I then use these coefficients and the data on the NLSY panel members (including their computed potential experience) to obtain $_{79}\hat{W}_t$ in each of the three years t -- 1984, 1990 and 1996.

The final issue in specifying the longitudinal models is the treatment of the measures of job satisfaction. In the panel data these can be viewed as allowing us to form a pair of 3x3 transition matrices, one showing transitions in satisfaction from 1984 to 1990, the other from 1990 to 1996. While I present these matrices below, by themselves they are not readily usable in models describing the determinants of changing job satisfaction. I thus estimate ordered probits for JS_{90} and JS_{96}, in each case using as additional independent variables a vector of two indicator variables that spans JS_{t-6}. Obviously there are many other ways to describe the changes implicit in these 3x3 transition matrices, but this one is the simplest.[9] Coupled with six-year changes in the various wage measures, it allows us to remove the v_i and $_i$ from the equations; at the same time, for JS_{96} we can distinguish between Cases III and IV by examining the relative effects of recent and past changes in the wage measures on the satisfaction outcomes.

Table 3 presents the characteristics of both the full subsample of 1280 workers and, for purposes of closer comparison to the results in Section III, for a still smaller subsample consisting of nonhispanic white males whose educational attainment was unchanged from 1984 to 1996. Panel A of the table demonstrates the expected rise in earnings over the twelve years. More interesting for our purposes is the pattern of wage growth in the smaller subsample shown in the first row of Panel B. The returns to schooling in this subsample rose from 1984 to 1990, but stayed essentially constant between 1990 and 1996.[10] That earnings differentials by education first widened in this sample and then stopped widening provides substantial room for examining the time path responses of job satisfaction to wage shocks.

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[9]In this case, however, ordered probit does lead to estimates of the parameters that are biased in unknown directions.

[10]The rates of return shown in the table are coefficients of years of schooling from the log-earnings equations that are used to calculate u_t.

The remainder of Table 3 presents the distribution of earnings within this subsample broken down by education (at most 12 years, more than 12 years) as well as the distributions of job satisfaction by education for each of the three years sampled. Perhaps the most interesting thing to note is that workers with at least some college education were more satisfied with their jobs than were high-school graduates. This difference suggests the important role of current earnings; but it is also quite remarkable that, between 1990 and 1996, exactly that time when the rate of return to schooling ceased rising in this sample (and perhaps in the United States more generally), the distribution of job satisfaction among the more educated workers began to approximate that of less educated workers more closely. This conjunction of a growing similarity in the responses of workers with different educational attainment and a halt in the growing returns to schooling suggests that changes in job satisfaction are more subtle than might be implied by any theory based on current outcomes alone.

Table 4 lists the transition matrices for the 1280 continuous full-time employees in the NLSY sample. The ^{2}-statistics for both the six-year transition matrices and the twelve-year matrix (1984 to 1996) show that current satisfaction is not independent of previous job satisfaction ($^2_{.01}(4) = 13.28$). This may be evidence for the assumption that $v_i \neq 0$, or it may show that some individuals earn high wages, or unexpectedly high wages, and are more satisfied with their jobs because of that. Nonetheless, that the off-diagonal elements in each matrix contain nearly half of the sample members suggests that there is ample room for describing changes in job satisfaction over the six-year intervals. The corresponding elements of the two six-year transition matrices are remarkably similar, with the only difference being the somewhat smaller off-diagonal terms in the later matrix, perhaps indicating greater stability of workers responses as they settle into longer-term jobs. The lesser significance of the ^{2}-statistic describing the twelve-year transition matrix may suggest that the impacts of shocks that affect job satisfaction do dissipate over time.

Table 5 contains the principal results of this section, the ordered probits describing job satisfaction in 1990 and 1996. In each equation I also include indicator variables for JS = 4 and JS = 3 at time t-6. The results are presented as in Table 2: The first two numbers are the impacts of ten-percent increases in the independent

variables on the probabilities of responding JS_t =4 or JS_t = 3, while the t-statistic on the ordered-probit coefficient is in parentheses below these derivatives. The first thing to note is that the hypotheses implicit in Cases I and II both make at least some sense in these estimates that account for individual heterogeneity. For both 1990 and 1996 the coefficient on dW_t is positive and significant, as is that on du_t. Moreover, when we add $d(_{79}\hat{W}_t)$ to the equation containing dW_t to embody Case III, the impact of the actual wage remains unchanged and still highly significant, while the wage predicted at t_0 is either negative, or positive and insignificant. Case II seems to describe the data best in 1990, while in the ordered probits without the lagged wage terms Case I or III does better in 1996.[11]

Because we have many years of longitudinal information on these workers, we can at least distinguish crudely between Cases III and IV by including lagged changes (changes between 1984 and 1990) in the various wage measures in the ordered probits for job satisfaction in 1996. The second row in each pair of ordered probits describing JS_{96} includes lagged terms (measuring percentage changes in the wage variable between 1984 and 1990). Considering the first equation (containing only dW_t and dW_{t-6}), the insignificant term in the lagged wage change suggests very clearly that more distant lags in a stimulus to job satisfaction have little effect (implicitly that the shocks are dissipated over time). A similar result is apparent in the equations embodying Case II (including du_t and du_{t-6}). The results become more interesting in the final specification in Table 5, the only one presented thus far that enables us to differentiate Cases III and IV. While the lagged wage change term is insignificant, the lagged term in the change in expected wages is highly significant statistically (and has the expected negative sign). None of the results showing the unimportance of the lagged changes in the wage measures changes qualitatively if we delete the lagged satisfaction measures from the ordered probits. Except

[11]Both here and in the next Section I examined whether job satisfaction responds asymmetrically to wage growth (or wage residuals) depending on whether this is positive or negative. Except for some slight indication in one case in the next Section, there is no evidence of any asymmetry in the responses of changes in job satisfaction to changes in wages (or in wage residuals).

for the apparent importance of initial changes in expected wage growth, none of the lagged changes in wages is important. This finding implies that past shocks to earnings have no effect on current satisfaction.

As a check on these estimates I present in Appendix Table A2 the results of the same specifications estimated over the reduced subsample of full-time continuous nonhispanic white male workers. The patterns of the coefficients are somewhat different from those in Table 5: As before, in the equations describing JS_{90} Case II seems to fit best; but in the description of JS_{96} Case I seems to fit best. There is no evidence that wages predicted at labor-force entry affect current job satisfaction, and none that past wage shocks persist in the determination of current job satisfaction.

The appropriate general conclusion from estimates based on longitudinal data from the United States is the same as from the comparison across cohorts of young adult men: Job satisfaction is not simply based on either actual wages or on apparent unexpected returns (excesses of wages over those received by observationally identical individuals). Instead, it may be based on comparisons of actual wages to what the individual might have expected based on the state of the labor market when he/she was making decisions about investing in human capital. The evidence on whether early labor-market experience has a long-term effect on job satisfaction is mixed, although the general impression is that early expectations and wage shocks become less important as workers age. The results in this section also indicate (since they implicitly relate changes in job satisfaction to changes in wages) that growing wage inequality generates at least a temporary increase in the inequality of the distribution of job satisfaction.

V. Longitudinal Changes in Job Satisfaction -- Germany 1984-96

Germany is one of the few countries with a long household panel containing information on a large number of workers job satisfaction and other characteristics. This is the German Socioeconomic Panel (GSOEP) begun in West Germany in 1984 and expanded to a panel in the former East Germany in 1991. Using data on Germany (actually, on the former West Germany) has the additional virtue that, unlike the United States, the country experienced no change in earnings inequality, at least through 1994 (Steiner and Wagner, 1997,

1998). Thus we may be able to examine the hypotheses discussed in Section II in a labor market characterized by a substantially different experience from the American one.

Because the GSOEP covers Germans of all ages, we cannot focus on two entirely different cohorts as in Section III.B. Instead, for 1984, 1994 and 1996 I construct data describing all men who were ages 30-59 in the particular year, who worked 12 months in the previous year and at least 30 hours in the survey week, and on whom earnings information is available. The sample includes West Germans only. The annual earnings measure is a broad one that adds to 12 times gross monthly earnings the 13^{th}-month pay, 14^{th}-month pay, Christmas bonuses, vacation pay, professional pay and other pay. The data on job satisfaction in the GSOEP are scaled more finely than those in the NLSY, with the worker asked to rate his/her job on a scale ranging from very satisfied -- 10 -- to very dissatisfied -- 0. For purposes of analysis and presentation here I combine the responses on this very fine scale into four categories.

The upper parts of Table 6 show the information on earnings and its distribution in this sample of full-time German male workers ages 30-59. Between 1984 and 1994 there was essentially little change in earnings inequality, with the slight lengthening of the right tail of the distribution being due partly to increases in nonwage compensation of higher-wage workers, partly to a slight increase in monthly earnings inequality.[12] Between 1994 and 1996 inequality of earnings in Germany expanded rapidly. As the data in Table 6 indicate, there was little change in the distribution of the left tail of earnings; indeed, even workers at the 75^{th} percentile did not see their earnings increase greatly relative to the median. Rather, the right tail of the distribution stretched out rapidly during this period.[13] Given this growth in earnings inequality, rather than providing a different pattern against

[12]For example, the 90-50 ratios of monthly earnings were 1.667 in 1984 and 1.706 in 1994.

[13]This is not an artifact of some error in the data for 1996. Already in 1995 the 90-50 earnings ratio had increased to 1.762, essentially half of the increase between 1994 and 1996. Nor is it a result solely of my using a broad definition of earnings, although that is part of the explanation. The 90-50 ratio of monthly earnings rose to 1.750 in 1996 after fluctuating between 1.65 and 1.70 between 1984 and 1994.

which to reflect the American results, changing German earnings inequality between 1984 and 1996 allows us to attempt to replicate the American results over a period in which inequality widened suddenly near its end.

The distributions of job satisfaction for all workers in the sample in each of the years are presented in the bottom part of Table 6. By themselves they are not informative, other than to underscore the perils of comparing trends in average job satisfaction: It is very difficult to believe that the percentage of men viewing their jobs as yielding satisfaction below 8 on a ten-point scale rose from 35 percent to 52 percent at a time when real earnings were rising, just as it was difficult to believe the apparent decline in average satisfaction implicit in Table 1. Comparisons of the effects of changing wage inequality on inequality in job satisfaction must abstract from aggregate changes in satisfaction that may arise from slight changes in survey design, question structure or even general attitudes about life outside the labor market.

The appropriate examination of how job satisfaction by earnings quantile changed over this period uses the double-difference calculation presented in (3). Figure 3 presents calculations for 1996 compared to 1984 exactly analogous to those in Figure 1. As is implicit in Table 6, the main change in earnings inequality over the twelve years was the rise in average earnings in the right tail (shown by the large positive double difference in Ln(W) at the top quartile). As the double differences in job satisfaction show, this rise in their relative earnings was accompanied by a sharp increase in the relative fraction of workers in the top earnings quartile responding that they were highly satisfied (9 or 10), and a sharp drop in the fraction stating they were not very satisfied (less than 6 on the ten-point scale). Beyond this result the relative changes in satisfaction at the second and third earnings quartiles were quite small, unsurprisingly given the small relative changes in earnings in these quartiles over the twelve years.

Figure 4 presents results like those in Figure 3, but calculated only over the two-year interval 1994 to 1996. It shows that most of the change in the inequality of satisfaction occurred only in the mid-1990s. On a per-annum basis the rise in the relative wages of workers in the top earnings quartile was four times faster in this biennium than over the entire twelve years. The double differences comparing the percentage stating they are

highly satisfied in the top and bottom quartiles of earnings in 1996 relative to 1994 are about as large as the change observed in Figure 3: Most of what we observed in that Figure is the result of the very rapid rise in earnings in the top quartile of German workers in the mid-1990s. Given the recentness of the rise in relative earnings in the top quartile, it is not surprising that this finding mirrors the strong results for the U.S. between 1978 and 1988 rather than the attenuation that occurred in the U.S. over the longer period 1978-1996.

The GSOEP also lends itself for use in an examination of the determinants of job satisfaction like that conducted on the NLSY in Section IV. Because it contains the same kind of information needed to construct wage residuals, it allows for distinguishing between Cases I and II. Unfortunately, because of the broad age range of the respondents, there is no information that would allow constructing empirical analogs to $_{t_i 0}E^*_{i t} \mid X_i$ (and no extraneous information that would enable me to form these expectations), and thus no way to test the validity of Cases III and IV. The panel nature of the data does, however, allow testing the importance of lagged terms in Cases I and II and thus whether wage shocks have long-term consequences on job satisfaction.

Here I construct a sample of respondents who were full-time (defined above) workers in each of 1984, 1990 and 1996, the same years used in Section IV for the American NLSY. That dynamic models of job satisfaction have something to explain is indicated by the transition matrices in Table 7. There is clearly some persistence of the responses: All the ^{2}-statistics strongly reject the hypothesis that job satisfaction in one year is independent of the response six years earlier, although as in the U.S. the stability becomes less apparent as the interval between observations increases (notice the lower ^{2}-statistics for the matrix linking 1984 and 1996). Nonetheless, the mass of the distribution in the off-diagonal elements of the matrices is substantial. Indeed, between 1984 and 1990 29 percent of the respondents changed their expressed satisfaction by more than two (of the four) categories. Between 1990 and 1996 the percentage was 24 percent, still substantial, and, as in the NLSY, lower than in the previous six-year period (during which members of the panel were younger and more mobile).

Table 8 replicates for the GSOEP the estimates of the dynamic job-satisfaction equations that were presented for the NLSY in Table 5. (Similar estimates, but for the reduced sample of native-born workers who acquired no additional education between 1984 to 1996, are presented in Appendix Table A3.) The earnings residuals are from equations that contain the usual education, experience and tenure terms. For each ordered probit I present the effect of a 10-percent increase in earnings on the average worker s probability of responding that he or she is in the top (9 or 10), the second (8), or the third (6 or 7) category on the scale of job satisfaction. The results for both JS_{90} and JS_{96} suggest that Case I, in which changes in job satisfaction depend on earnings, describes the data somewhat better than does Case II, in which changes in the unexplained component of earnings determine changes in satisfaction. This partly contradicts the results from the NLSY panel in Section IV, but we found this same result for levels of job satisfaction in the 1988 NLSY cross section.

The most interesting results in this section are those for JS_{96} that include the lagged terms in changes in wages (wage residuals). As in Table 5, these terms, changes in the returns to work between 1984 and 1990, have little effect on the change in job satisfaction between 1990 and 1996. Nor is their unimportance an artifact of having included the indicators describing JS_{90} in the equations: When these are removed, the t-statistics on the two wage change terms become 3.87 and 0.69, while those on the changes in the wage residuals are 3.55 and 0.83. While we cannot test for the importance of early-career expectations (test Case III), this evidence reinforces the conclusion from the NLSY estimates that past earnings shocks have little effect on workers current job satisfaction. They thus imply that Case IV would be a better description of behavior than Case III.

VI. Synthesis and Inferences

With so many different samples and several ways of examining them it is unlikely that we will find complete unanimity in their implications for the determinants of job satisfaction. Nonetheless, some conclusions seem quite strongly supported by the analyses: 1) Inequality in job satisfaction tracks shocks to earnings inequality; but 2) There is little relation between job satisfaction and persistent inequality of earnings. In terms of underlying models of job satisfaction, all the evidence suggests that recent shocks to earnings matter more to

current and recent changes in job satisfaction than do distant earnings shocks; and the evidence implies that the influence of early-career expectations on job satisfaction diminishes with age.

That the distribution of job satisfaction widens (temporarily) in response to shocks that widen the distribution of the returns to work has an interesting implication for the dynamics of those returns. We know that voluntary turnover and job satisfaction are negatively related (e.g., Clark et al, 1999). That being the case, our results imply that an exogenous shock that widens the distribution of earnings also raises turnover among low-wage workers. This change in turn leads them and their employers to invest less in firm-specific human capital, with the opposite occurring among high-wage workers and their employers. Since these investments generate future returns, the linkage through changes in the distribution of job satisfaction, even though these are temporary, can generate permanent hysteresis effects on the distribution of earnings.

From the studies presented here it is clear that changes in earnings affect job satisfaction. What I have not answered, and what is not answered in the literature, is the exact mechanism through which these effects operate. While it seems likely that workers pay more attention to comparisons of the returns to observable than to unobservable skills, the results are mixed about whether workers reference others with similar observable characteristics, reference all workers, or what. Only with longer panels of data, preferably from several countries, are we likely to be able to disentangle any further the various explanations of patterns in job satisfaction.

Limitations on the data have prevented going beyond examining the impacts of transformations of wages on job satisfaction. Yet the evidence that the demand for nonpecuniary and nonwage pecuniary returns to work is income-elastic suggests that it would be very worthwhile to examine a broader set of economic determinants of satisfaction. That analysis awaits a set of longitudinal data that contains objective measures of these nonwage aspects of work.

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Table 1. Global Job Satisfaction and Selected Characteristics, Nonhispanic White Male Full-time 26-31 Year-Olds, 1978 and 1988

	1978	1988
Annual Earnings		
Mean	15154	24260
SD	(6419)	(12407)
Distribution of W_p / W_{50}		
Percentile:		
5th	0.488	0.425
10th	0.589	0.533
25th	0.767	0.733
75th	1.254	1.333
90th	1.549	1.733
95th	1.742	1.978
Job Satisfaction (percent distributions)		
Like very much	50.67	41.42
Like fairly well	42.12	48.49
Dislike somewhat or very much	7.21	10.09
N	1040	1060

Table 2. The Determinants of 1988 Job Satisfaction, NLSY, Nonhispanic White Male Full-time Workers (N = 1056)*

	Independent Variable:							
Case:	W_{88}	u_{88}	${}_{78}\hat{W}_{88}$	$[u_{88} - \hat{u}_{p,78}]$	$[\hat{W}_{88} - {}_{78}\hat{W}_{88}]$	Interaction	Pseudo-R^2	
I	.0006 .0055 (2.85)							.0041
II		.0082 -.0046 (2.38)						.0028
III	.0083 -.0046 (3.06)		-.0069 .0004 (-1.23)					.0048
III, decomposed								
				.0133 -.0075 (1.72)				.0015
					.0360 -.0034 (3.22)			.0052
				.0105 -.0004 (1.72)	.0362 -.0035 (3.24)			.0067
				-.0656 -.0142 (-2.36)	.0372 -.0028 (3.37)	.2287 -.0839 (2.34)		.0109

*The underlying earnings equations include education, actual job experience, job tenure, annual hours, and indicator variables for marital status, location in the South and in an SMSA, union status, and bad health. Here and in Tables 5 and A2 the first number is the effect of a 10-percent increase on Pr{JS=Like Very Much}; the second is the effect of this increase on Pr{JS=Like Fairly Well}, and the number in parentheses is the t-statistic on the underlying coefficient.

Table 3. Sample Characteristics and Time-series Comparisons of the Distribution of Job Satisfaction, NLSY

A. Means and Standard Deviations of Ln(Annual Earnings), Full-time Workers All Three Years (N=1280)

1984:	9.347	1990:	9.976	1996:	10.281
	(0.575)		(0.490)		(0.563)

B. Job Satisfaction and Means and Standard Deviations of Annual Earnings, Nonhispanic White Males with Unchanging Schooling, Full-time Worker All Three Years (N=423)

Return to Schooling (estimate and standard error), seemingly-unrelated estimates

1984:	.069	1990:	.102	1996:	.108
	(.013)		(.011)		(.013)

By Education:

≤12 Years of Schooling (N=302):

Annual Earnings

1984:	9.401	1990:	9.969	1996:	10.290
	(0.506)		(0.478)		(0.509)

Percent Distribution of Job Satisfaction

	1984	1990	1996
Like Very Much	39.74	41.72	41.06
Like Fairly Well	50.33	52.98	50.66
Dislike	9.93	5.30	8.28

>12 Years of Schooling (N=121):

Annual Earnings

1984:	9.550	1990:	10.380	1996:	10.744
	(0.651)		(0.410)		(0.528)

Percent Distribution of Job Satisfaction

	1984	1990	1996
Like Very Much	54.55	56.20	45.45
Like Fairly Well	39.67	38.84	47.11
Dislike	5.79	4.96	7.44

Table 4. Transitions of Job Satisfaction, NLSY, Continuous Full-time Workers (N = 1280)

	Like Very Much	Like Fairly Well	Dislike
		1990	
1984			
Like Very Much	25.23	16.64	2.11
Like Fairly Well	16.02	25.39	4.69
Dislike	2.66	5.55	1.72

$^2(4) = 85.70$

		1996	
1990			
Like Very Much	26.41	15.23	2.27
Like Fairly Well	4.53	27.66	4.53
Dislike	1.80	5.08	1.64

$^2(4) = 125.61$

		1996	
1984			
Like Very Much	24.22	16.56	3.20
Like Fairly Well	16.33	26.25	3.52
Dislike	3.05	5.16	1.72

$^2(4) = 66.92$

Table 5. Ordered Probit Estimates of the Determinants of Longitudinal Changes in Job Satisfaction, NLSY Continuous Full-time Workers (N = 1280)*

	Independent Variable:						
Dependent Variable:	dW_t	dW_{t-6}	du_t	du_{t-6}	$d(_{79}\hat{W}_t)$	$d(_{79}\hat{W}_{t-6})$	Pseudo-R^2
JS_{90}:	.0046 -.0004 (2.50)						0.0377
			.0065 -.0014 (2.97)				0.0388
	.0042 -.0003 (2.26)				.0042 -.0003 (0.73)		0.0379
JS_{96}:	.0068 .0003 (3.11)						0.0523
	.0067 .0001 (2.96)	.0011 .00003 (-.67)					0.0525
			.0055 -.0003 (2.29)				0.0505
			.0058 -.0003 (2.35)	.0058 -.0001 (.56)			0.0506
	.0061 .0006 (2.83)				.0072 .0006 (1.62)		0.0535
	.0072 -.0009 (2.93)	.0020 -.0002 (.12)			.0040 -.0005 (.79)	-.0197 .0011 (-3.32)	0.0584

*The ordered probits here and in Tables 8, A2 and A3 also include pairs of indicator variables describing JS_{t-6}.

Table 6. Global Job Satisfaction and Selected Characteristics, Full-time 30-59 Year-Old Men, 1984, 1994 and 1996, German Socioeconomic Panel

	1984	1994	1996
Annual Earnings			
Mean	43831	66723	72346
SD	(19684)	(37100)	(46841)
Distribution of W_p / W_{50}			
Percentile:			
5th	0.649	0.646	0.631
10th	0.715	0.720	0.700
25th	0.838	0.830	0.830
75th	1.259	1.293	1.326
90th	1.662	1.702	1.814
95th	2.000	2.075	2.165
Job Satisfaction (percent distributions)			
10 or 9	41.40	19.15	20.13
8	23.82	29.19	28.26
7 or 6	18.14	29.00	30.60
<6	16.64	22.66	21.01
N	2338	1514	1366

Table 7. Transitions of Job Satisfaction, GSOEP, Continuous Full-time Male Workers (N = 970)

1990

1984	9 or 10	8	7 or 6	≤6
9 or 10	14.02	13.51	8.87	3.30
8	5.67	9.28	6.91	2.99
7 or 6	2.68	4.74	6.70	3.71
<6	2.99	3.51	4.95	6.19

$^2(9) = 107.40$

1996

1990				
9 or 10	10.93	7.11	4.74	2.58
8	4.74	12.16	9.79	4.33
7 or 6	2.68	7.53	10.10	7.11
<6	0.72	3.30	4.43	7.73

$^2(9) = 210.52$

1996

1984				
9 or 10	10.31	12.47	10.93	5.98
8	4.43	8.76	7.53	4.12
7 or 6	1.75	5.36	5.67	5.05
<6	2.58	3.51	4.95	6.60

$^2(9) = 62.18$

Table 8. Ordered Probit Estimates of the Determinants of Longitudinal Changes in Job Satisfaction, GSOEP Continuous Full-time Male Workers (N = 970)*

	Independent Variable:				
	dW_t	dW_{t-6}	du_t	du_{t-6}	Pseudo-R^2
Dependent Variable:					
JS_{90}:	.0063				0.0345
	.0229				
	-.0251				
	(1.94)				
			.0055		0.0343
			.0250		
			-.0274		
			(1.76)		
JS_{96}:	.0060				0.0698
	.0047				
	-.0188				
	(3.44)				
	.0057	.0012			0.0706
	.0032	-.0034			
	-.0188	-.0025			
	(3.62)	(1.42)			
			.0064		0.0690
			.0086		
			-.0210		
			(3.10)		
			.0064	.0015	0.0695
			.0089	-.0033	
			-.0210	-.0016	
			(3.29)	(1.18)	

*The underlying earnings equations include education, actual job experience, job tenure, annual hours, and an indicator variable for marital status. Here and in Table A3 the first number is the effect of a 10-percent increase in the variable on Pr{JS=9 or 10}; the second is the effect of this increase on Pr{JS=8}; the third is the effect of this increase on Pr{JS=6 or 7}, and the number in parentheses is the t-statistic on the underlying coefficient.

Figure 1. Double Differences in Jobsat
and Wages, 1988- 1978

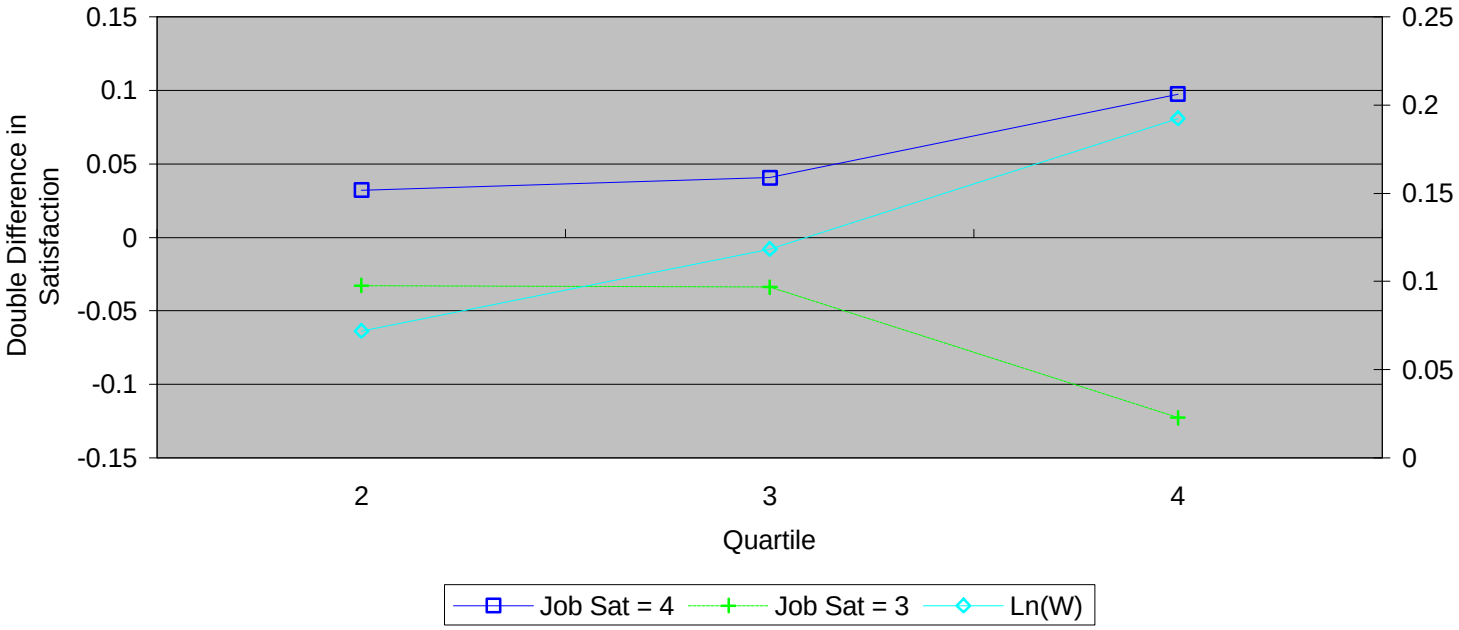

Figure 2. Double Differences in Jobsat and Wages, 1996 - 1978

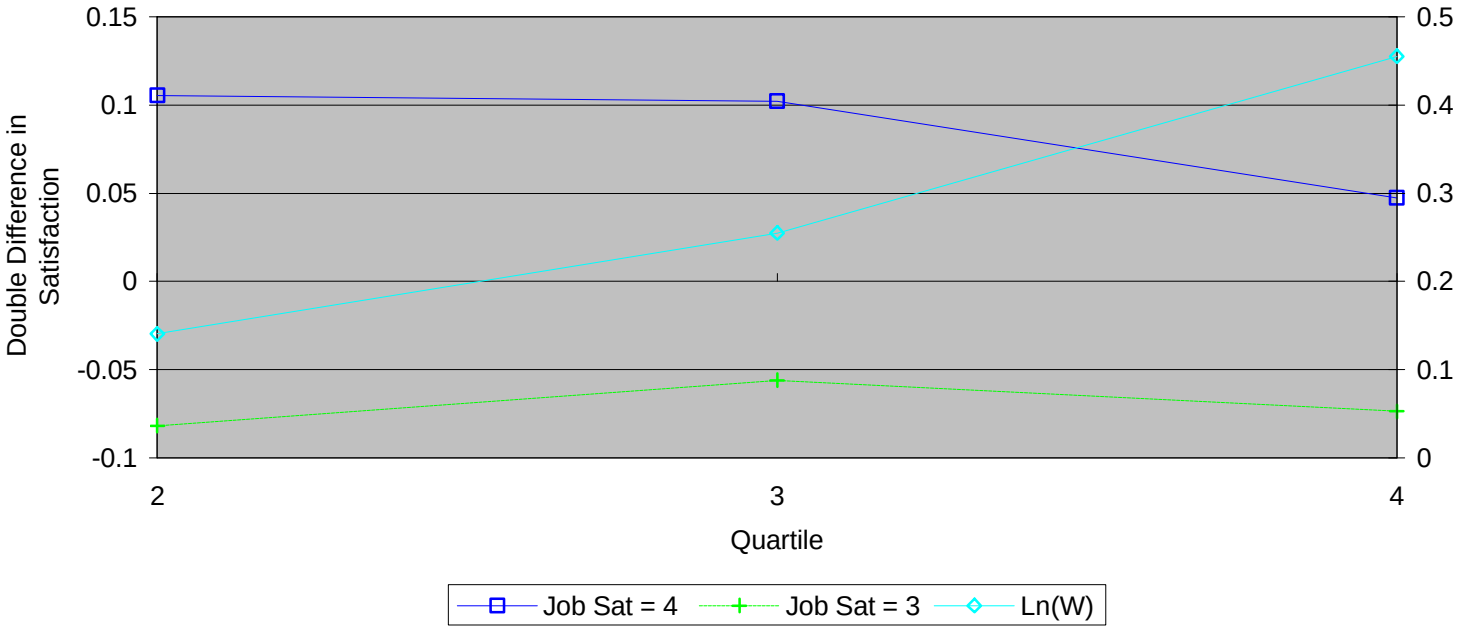

Figure 3. Double Differences in Jobsat and Earnings, 1996 - 1984, GSOEP

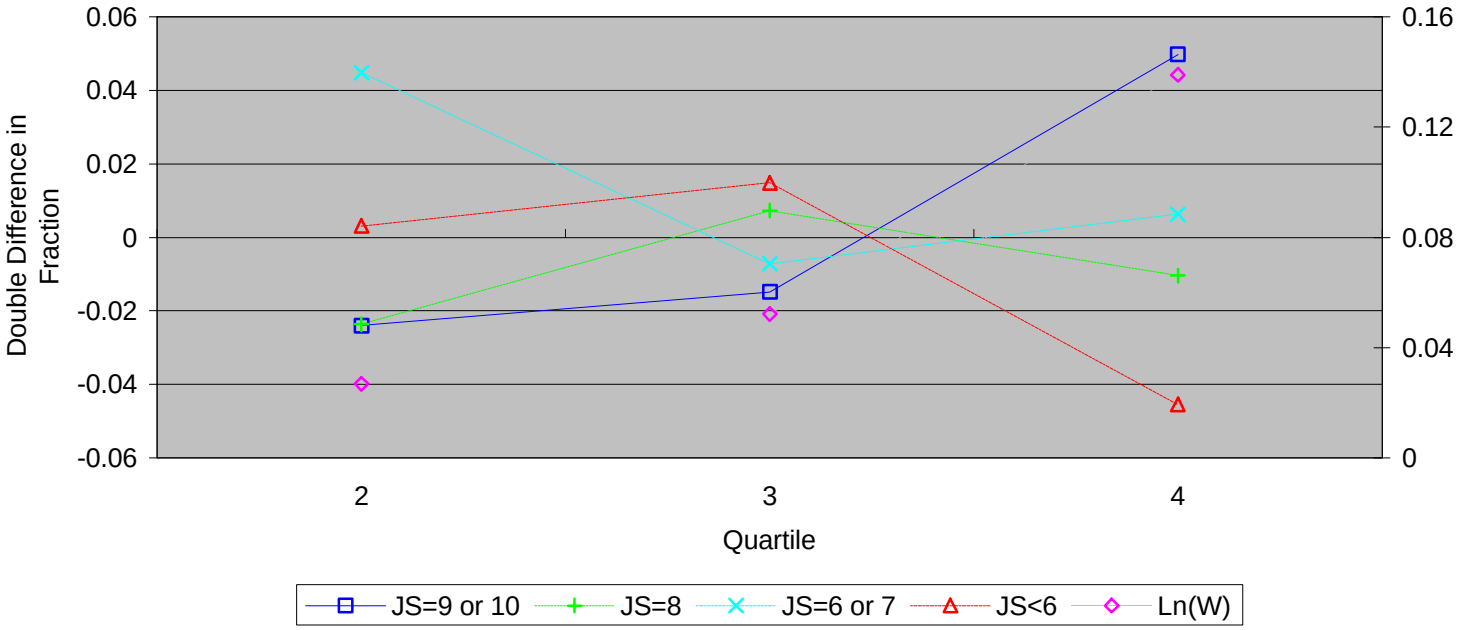

Figure 4. Double Differences in Jobsat
and Earnings, 1996 - 1994, GSOEP

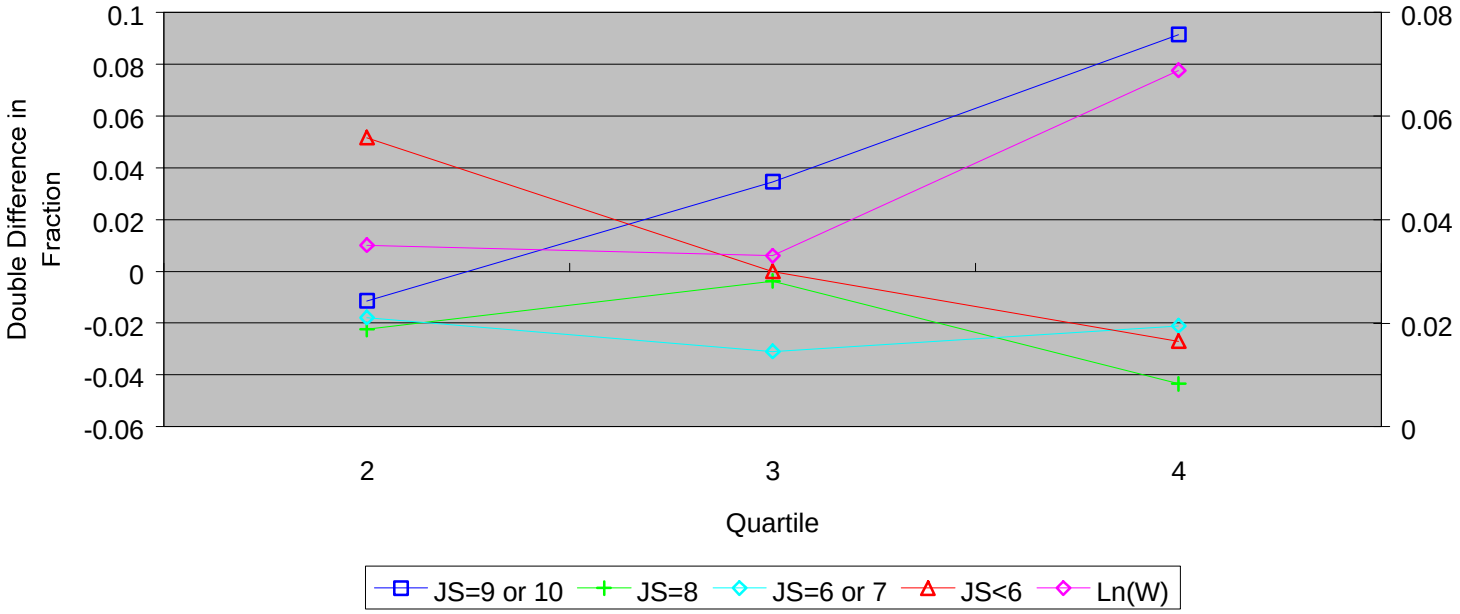

Table A1. OLS Coefficients of Human Capital Variables, CPS 1979*

	Men	Women
Schooling	.060	.082
	(.001)	(.002)
Experience	.026	.024
	(.002)	(.002)
$Experience^2/100$	-.022	-.020
	(.003)	(.004)
Schooling*Experience/10	.011	-.00002
	(.001)	(.002)
$Schooling*Experience^2/100$	-.004	-.002
	(.0003)	(.0003)
$\bar{R}^2$	.338	.264
N	86,934	58,368

*The dependent variable is the logarithm of usual weekly earnings. Also included in the equations are a continuous measure of usual weekly hours and indicator variables describing race (black or not), ethnicity (Hispanic or not), location in the South, location in an MSA, and marital status (married or not).

Table A2. Ordered Probit Estimates of the Determinants of Longitudinal Changes in Job Satisfaction, NLSY Continuous Full-time Nonhispanic White Male Workers with Constant Education (N = 423)

	Independent Variables:						
Dependent Variable:	dW_t	dW_{t-6}	du_t	du_{t-6}	$d(_{79}\hat{W}_t)$	$d(_{79}\hat{W}_{t-6})$	Pseudo-R^2
JS_{90}:	.0073 -.0048 (1.82)						0.0329
			.0105 -.0076 (2.28)				0.0355
	.0071 -.0045 (1.73)				.0028 -.0018 (.19)		0.0329
JS_{96}:	.0122 .0067 (3.91)						0.0577
	.0124 .0059 (3.87)	-.0010 -.0005 (-.44)					0.0579
			.0109 .0020 (2.98)				0.0492
			.0111 .0019 (2.98)	.0007 .0002 (.25)			0.0493
	.0118 .0069 (3.82)				.0033 .0021 (.44)		0.0579
	.0125 .0053 (3.77)	-.0009 -.0005 (-.39)			.0033 .0016 (.39)	-.0028 -.0014 (-.31)	0.0584

Table A3. Ordered Probit Estimates of the Determinants of Longitudinal Changes in Job Satisfaction, GSOEP Continuous Full-time West-German Born Male Workers with Constant Education (N = 558)

	Independent Variables:				
Dependent Variable:	dW_t	dW_{t-6}	du_t	du_{t-6}	Pseudo-R^2
JS_{90}:	. 0066 .0193 -.0258 (2.17)				0.0399
			.0056 .0223 -.0317 (2.42)		0.0406
JS_{96}:	.0057 .0024 -.0178 (2.67)				0.0850
	.0057 .0028 -.0178 (2.60)	-.0004 .0011 .0008 (-0.33)			0.0851
			.0062 .0063 -.0199 (2.41)		0.0841
			.0062 .0062 -.0188 (2.24)	-.0008 .0020 .0011 (-0.52)	0.0843