METHODS OF PAY AND EARNINGS: A LONGITUDINAL ANALYSIS

DANIEL PARENT*

Using data from the National Longitudinal Survey of Youth (1988–90), the author investigates the relationship between methods of pay, including piece rates and bonuses, and the level and variance of wages. Among men, piece rate workers earned a premium compared to other workers, but the evidence on bonuses is mixed. The author finds evidence that female piece rate workers earned more than other female workers once a control variable for the presence of dependents is interacted with the piece rate variable. With controls for the wage effects of schooling and experience, unobserved worker productivity is found to have accounted for most of the wage variance among both male and female piece rate workers; wage variance among workers not having explicit pay for performance schemes, in contrast, was predominantly due to other factors.

Over the past twenty years or so, a number of theoretical models have been developed to analyze the employment relationship, with a particular focus on how contracts can be structured to provide appropriate incentives. Despite the rapid pace at which information economics has progressed, empirical investigations of the implications of these models (let alone direct testing of the models themselves) have been rather rare. The obvious reason for this neglect is the lack of data.

The objective in this paper is to examine how piece rates and bonuses affect wages by looking at three consecutive waves from the National Longitudinal Survey of Youth (NLSY) for the years 1988–90. Over that three-year period, workers were asked if part of their labor earnings was based on job performance. They were also asked what form of incentive schemes their employer used. Thus, at the first level, we can examine the extent to which incentive pay was used in a relatively representative sample of young workers.

Most previous studies have used cross-sectional data to analyze the same ques-

*The author is Assistant Professor at McGill University, Researcher at the Centre de recherche et developpement en economique (C.R.D.E.), and researcher at the Centre interuniversitaire de recherche en analyse des organisations (CIRANO). He thanks Bentley MacLeod for many discussions on this topic and seminar participants at the Princeton Labor Lunch, Université Laval, and École des Hautes Études Commerciales, Montréal, for useful comments. He also thanks Princeton University for its hospitality during a visit at which he did some of the work for this study.

The data and computer programs used in this paper can be obtained from the author at McGill University, 855 Sherbrooke St. W., Montréal, Québec H3A 2T7, Canada.

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tions. The longitudinal structure of the NLSY data, however, opens up the possibility of disentangling incentive effects from pure selection effects by controlling for unobserved worker and job-match heterogeneity. Also, having three years of data allows the estimation of a simple model of the covariance structure of wages by type of contract.

Modeling the covariance structure of earnings has two benefits. First, it allows one to determine what fraction of the unexplained variance in earnings can be attributed to unobserved worker heterogeneity, since pay is closely tied to individual performance when explicit incentives are used. In addition, it allows an exploration of the reasons for the frequent finding that workers whose earnings are explicitly tied to performance earn more, on average, than salaried or hourly paid workers. To investigate whether that pattern simply reflects the payment of compensating differentials for the increased risks faced by piece rate workers (Seiler 1984), for example, one needs to "purge" the total cross-sectional wage variance of the part that comes from unobserved worker heterogeneity—an exercise that requires the use of longitudinal data.

Previous Research

The studies most comparable to the present one are Ewing (1996), Brown (1992), and Seiler (1984). In each of those papers the authors used a large data set—a wave of the NLSY (Ewing) or data from the U.S. Bureau of Labor Statistics' Industry Wage Surveys (Brown and Seiler)—to study the use and wage impact of pay-for-performance schemes.[1] The main finding from these papers is that "incentive workers," piece rate workers in particular, earned more, on average, than time-rated workers.[2]

A recent paper by Lazear (1996) used data from one company to study the effects on productivity and wages of switching hourly rated workers to piece rate contracts (see also Pencavel 1977). Consistent with previous results, Lazear found that moving from an hourly wage to a piece rate caused a 41% increase in productivity, as measured by units of glass installed on automobiles. The switch to piece rates also had the effect of decreasing absenteeism, which is another measure of worker productivity. In addition, output variance increased with the move to piece rates. Exploiting the fact that he could follow the same workers before and after the change, Lazear found that sorting and incentives both played an important role in this productivity gain, with more than half the gain being due to incentive effects.[3] Finally, wages also increased by an average of 9% following the firm's decision to move to a piece rate scheme.

Theoretical Considerations

Relationship between Piece Rates and Earnings

Lazear's (1986) two-period model analyzes the conditions under which piece rate contracts are preferred to fixed salaries. Lazear defines salaried workers as individuals for whom

$$w_t = g(E_t), \tag{1}$$

where E_t is some measure of effort (such as hours worked) in period t. For piece rate workers,

$$w_t = f(q_t), \tag{2}$$

[1]The data sets employed in these three studies were more representative of the labor force than were the samples used in some of the recent work studying the interaction between incentives, firm performance, and top executive salaries. See, for example, Gibbons and Murphy (1992) and Jensen and Murphy (1990).

[2]For a more structural approach to the study of worker incentives, see also the work of Ferrall and Shearer (1996) and Shearer (1996).

[3]Lazear distinguishes incentive effects from pure selection effects by using fixed-effects regressions to purge the data from unobserved fixed worker-firm heterogeneity.

where q_t is output in period t. A crucial assumption in Lazear's model is the existence of a monitoring cost. Without such an assumption, no fixed salary jobs would exist. Let C be the monitoring cost, R the piece rate, and q the (lifetime) output of the worker. For piece rate jobs, we have

$$w = Rq - C, \tag{3}$$

and for fixed salary jobs,

$$w = E(q). \tag{4}$$

It is assumed that people can work in other jobs at w^*. If $E(q) > w^*$ and there are no piece rate firms, then all workers work at this firm. But this result is inefficient: those for whom $q < w^*$ should take the alternative job. Separating them from the others causes expected output to increase. If both parties are equally uninformed as to the productivity of the worker, then separation can be achieved by putting everybody in the piece rate job for a period of time in order for both parties to learn q. After they learn q, those with $Rq > w^*$ stay in the piece rate job. With asymmetric information, less productive workers select themselves into the salary jobs. The firms know that they have attracted low-ability workers and they pay accordingly. Thus, mean earnings should be higher for piece rate workers than for salaried workers. Pencavel (1977), Seiler (1984), Brown (1992), and Lazear (1996) each have found evidence that piece rate workers are paid more than salaried workers. Also, since pay and productivity are more closely linked in the case of piece rate contracts, the variance of wages for piece rate workers should be greater than for other workers. Seiler (1984) found evidence from the Industry Wage Surveys that wages are indeed more dispersed for piece rate workers than for other workers. He interpreted this finding as evidence that piece rate workers are paid compensating wage differentials for the increased risks they face.

Intertemporal Strategic Behavior

Piece rates can be negatively related to past output, because employers may use past performance as an indicator of the difficulty of a particular task. If workers perform very well early on, employers will come to believe that the task is relatively easy and will, therefore, lower the pay rates (or increase the performance standard). This is the standard "ratchet effect." Realizing this, workers may depress one period's output in order to realize a better rate in the next period. Lazear argued that this strategic behavior can be undone by inflating the piece rate in the first period. Workers will respond with a higher effort level even though they anticipate that the employer will cut their rates in the second period. Thus, if workers and employers behave as Lazear suggests, we should see a stronger effect of piece rates on earnings early in the employment relationship than later.

That prediction hinges completely on the assumption that workers can fully commit not to quit at the end of period one. Employers must be in a position to exploit some ex post monopsony power. Otherwise, a competitive firm can attract workers whose rates have been decreased.

This idea was formally developed by Kanemoto and MacLeod (1991), who showed that the existence of outside opportunities can eliminate the ratchet effect, even when workers incur mobility costs. In the context of a two-period model, firms compete for a fixed number of workers. Competition among firms thus reduces ex ante profits to zero. A cost C is borne by the workers each time they take on a job (mobility cost or cost of retraining). This cost captures the fact that once a worker is in a job, the firm has some monopsony power over that person. The firm tries to exploit this power by lowering rates (or increasing sales targets) in the second period. Kanemoto and MacLeod showed that for sufficiently low mobility costs, competition for what they call second-hand workers can eliminate the ratchet effect. Unlike Lazear's model, in their model efficient piece rate contracts are possible without perfect commitment.

Three predictions stem from these models. First, more productive workers will

choose piece rate jobs (Lazear 1986). In other words, any measured positive relationship between piece rates and wages is attributed to self-selection on the part of workers. Thus, using fixed effect methods should result in the elimination of all positive effects. Second, since piece rate schemes gear the worker's compensation directly to output, the variance of wages should be larger for piece rate jobs than for salaried jobs. Finally, Lazear's model predicts that the effect of piece rates should be stronger early in the employment relationship, while Kanemoto and MacLeod predict no such time decreasing pattern.

Relationship between Bonuses and Earnings

Piece rate or commission contracts are explicit in nature: one gets paid a certain contractually specified amount per unit produced. Bonuses can be explicit as well, such as when workers get rewarded for achieving or surpassing a sales target. But employers can also award bonuses on a more discretionary basis. An implicit contract between the worker and the firm might specify that performances deemed satisfactory by the employer will be rewarded by the payment of a discretionary end-of-period bonus. Such an implicit agreement will be carried out only if both parties gain by doing so. If they do not, nothing can prevent them from reneging: workers will not offer superior performance and firms will not pay any bonus even if the worker provides a high level of effort. For both parties to keep their promises, the contract has to be self-enforcing.[4] (See Carmichael 1989 for a review of self-enforcing contracts.)

One of the basic features of such an employment relationship is the inherent conflict between short-term and long-term interests. In the short term, the firm would rather not pay a bonus and the worker would rather not provide a superior effort level. However, if either party reneges on the implicit agreement, it will be punished: shirking workers will not receive a bonus and will get fired, while firms that do not pay the bonus as promised will see their workers quit.

These ideas were spelled out by MacLeod and Malcomson (1989) in the context of a repeated game model. They first showed that for a self-enforcing contract to exist, there has to be a surplus from continuing the relationship: both parties are at least as well off in the current contract as they would be in market alternatives, and one of them must be strictly better off. Second, incentives can be provided either by paying so-called "efficiency wages" (Shapiro and Stiglitz 1984) or by paying end-of-period bonuses. Since in that model bonuses occur when workers receive no rent from the relationship, it follows that total earnings (base salary plus bonus) should be set equal to the market alternative. Thus, there is no reason to believe that in a cross-section of workers, those paid bonuses should earn more, on average, than other workers.[5] Worker risk aversion could also imply the payment of compensating differentials. Since bonuses are tied to performance, which in turns depends on ability, if workers are uncertain about their ability, they may ask for higher pay as insurance against the possibility that they are low-productivity individuals.

[4]Given the subjective nature of the performance evaluation, such contracts are not legally binding, that is, although observable by both parties, a "satisfactory" performance level is not generally verifiable in courts (or, at least, it is costly to verify).

[5]Perhaps it would be more appropriate to say that even if we find bonus workers are paid more on average than time-rated workers, that result would still be consistent with MacLeod and Malcomson (1989). MacLeod and Malcomson's implicit contract model predicts that when workers have no surplus from the relationship, incentives are provided by paying bonuses. Although total earnings (salary plus bonus) do not have to be higher than for straight salary workers, they may be higher for other reasons, such as compensating differentials for higher effort levels required from bonus workers.

Data

The National Longitudinal Survey of Youth data set covered 12,686 young males and females who were between the ages of 14 and 21 in 1979. In 1988, 1989, and 1990, respondents were asked whether all or part of their earnings were based on job performance. They were also asked a few questions about their work environment. For instance, we know if the respondents were supervising other employees and whether they had received a promotion since the last interview. Unfortunately, we do not know the precise dollar amounts of incentive pay received by workers, nor do we know the proportion of their earnings that is due to pay-for-performance. These represent significant data limitations that constrain the interpretation of the results.

The question pertaining to pay-for-performance is the following:

> The earnings on some jobs are based all or in part on how a person performs the job (*hand card d*). On this card are some examples of earnings that are based on job performance. Please tell me if any of the earnings on your job (are/were) based on any of these types of compensation. Please do not include profit sharing or employee stock purchase plans.
>
> 1. Piece rates.
> 2. Commissions.
> 3. Bonuses (based on job performance).
> 4. Stock options.
> 5. Tips.
> 6. Other.

Note that it is not possible to tell *a priori* whether the bonuses refer to amounts paid at the discretion of the employer when the latter subjectively considers that the performance of the employee is worthy of a cash reward ("merit pay"), or whether they merely represent another form of piece rate, or even whether they are team bonuses.

Some summary statistics are presented in Table 1.[6] I restrict the sample to individuals who were in the labor market on a full-time basis, meaning (i) those whose primary activity was either working full-time, on a temporary lay-off, or looking actively for a job, and (ii) those who had worked at least half the year since the last interview and who were working at least 20 hours per week. Individuals excluded from the sample are those who had been in the military at any time, the self-employed, and all public sector employees. Also, I exclude observations for which real (1979-dollar) hourly earnings are less than $1.00 or greater than $100.00. These restrictions leave an unbalanced sample of 8,137 observations (3,844 workers). Of some importance is the fact that the workers in this data set are aged 23 to 33. Thus, the proportions of managerial positions and true merit pay jobs are smaller than in a representative sample of the entire population.[7]

For comparison purposes, note that in the Earnings Supplement to the January 1977 Current Population Survey (CPS), in which workers are matched to their employers through a Validation Survey, 4% of the workers report being paid a piece rate (out of 4,945 worker-observations), and their employers also report that 4% of the workers are paid a piece rate (out of 4,468 employer-observations), compared to the 9.4% reported in the NLSY. When I restrict the CPS sample to workers who are less then 34 years old, the percentage of piece rate workers is even lower, at about 3.5%. Part of the discrepancy in the percentage of workers reporting they are paid a piece rate stems, no doubt, from the different wording of the questions themselves in the NLSY and the CPS. In the NLSY, workers are asked if any of their earnings are based on

[6]Note that the statistics are not weighted to reflect the fact that the NLSY sample I am using includes the minority and poverty subsamples. Doing so increases the average hourly earnings and the average schooling, but leaves the percentage of the sample paid either a bonus or a piece rate essentially the same.

[7]Like Lazear (1986), I define piece rates as the method of pay used by firms when they measure each worker's output and tie wages to measured output. Thus, in that sense, commissions are also piece rates and are therefore labeled as such in the sample.

Table l. Mean Sample Statistics.

Statistic	*Mean*
Hourly Earnings (1979 dollars)	5.5
Age	28.0
Tenure	3.2
Schooling	12.4
Percentage Unionized	15.2
Percentage Female	43.4
Percentage Paid Piece Rates	9.4
Percentage Paid Bonuses	14.2
Sample Size	8,137
Number of Workers	3,844

piece rate pay. In the CPS, workers are asked whether they are paid "on a piecework basis" (as opposed to an hourly rate or a salary), and the different pay methods are mutually exclusive categories, which is not the case for the NLSY. In other words, someone reporting that she is paid on a piecework basis in the CPS cannot report being paid also on an hourly basis. Consequently, it is likely that CPS workers report the pay method from which they earn most of their labor income.

Concerning bonuses in the CPS sample, a little over 8% of the workers reported receiving a bonus, and about 9.5% of the corresponding employers declared having paid one, compared to a figure of 14.2% for my NLSY sample. Thus, while both bonuses and piece rates apparently occurred at a higher rate among workers in the CPS than among those in the NLSY, the difference is especially noticeable in the case of piece rates. Note, however, that the CPS, unlike the NLSY, contains no direct question on bonuses based on job performance. To arrive at the above figures, I used a question on the extra amounts that salaried workers may have earned in tips, commissions, or bonuses, deriving the percentage of workers who earned bonuses by excluding workers on commissions or who received tips.

The Panel Study of Income Dynamics (PSID) also contains a question on the amounts earned in bonuses, commissions, or overtime. Thus, one can also construct a measure of bonus incidence using the PSID by deleting commission and overtime workers. As it turns out, the percentage of workers in the PSID reporting a bonus in 1976 is very close to the percentage reported in the CPS (8.6% and 8.0%, respectively) and also, for the years 1988–90, very close to the percentage reported in the NLSY (12.4% in the PSID, compared to 14.2% in my NLSY sample).

Turning to the incidence of piece rates and bonuses by occupation or industry, over 38% of sales workers in my NLSY sample report being paid a piece rate, followed by operatives at less than 14%.[8] Industries where piece rates are frequently observed are retail and wholesale trade (11.8% of the observations in that industry), finance, insurance, and real estate (11.0%), and manufacturing (10.8%). The frequency of bonuses is more evenly distributed across occupations and industries than is the frequency of piece rates. While sales workers and managers stand out with reported rates of 25.7% and 28.5%, respectively, other 1-digit occupations report an incidence ranging from 8% for service workers to 14.9% for professional and technical workers. The distribution across industry shows that construction is the only 1-digit industry with a bonus incidence of less than 10%, all other industries having a bonus incidence ranging from 10.6% (agriculture, forestry, and fisheries) to 18.2% (mining).

Although it seems reasonable to think that many of the bonuses paid to managers and especially to sales workers are in effect another form of piece rate, in which pay is based on the number of units sold or produced, this conjecture is not so obvious for professionals and clerical workers. The fact that bonuses are more evenly distributed than piece rates across occupations and industries suggests that the bonuses reported in these data are, to a degree, a different form of incentive scheme.

[8]Remember that the label "piece rate" includes commissions as well. Distinguishing between piece rates and commissions as is done in the questionnaire would reveal that commission contracts are most common for sales workers while piece rate contracts are most common for operatives.

Table 2. Average Change in Log-Earnings by Gender and Type of Transition.

	Pay Method in Year T+l					
	Men			*Women*		
Pay Method in Year T	*Piece Rate*	*Bonus*	*Salary or Hourly Only*	*Piece Rate*	*Bonus*	*Salary or Hourly Only*
			Panel A: Job Stayers Only			
Piece Rate	0.061	0.143	–0.129	0.018	0.133	–0.101
(Obs.: Men, 177; Women, 117)	[0.571]	[0.339]	[0.282]	[0.658]	[0.162]	[0.282]
Bonus	0.129	0.057	–0.009	0.179	0.144	–0.007
(Obs.: Men, 303; Women, 186)	[0.155]	[0.505]	[0.413]	[0.167]	[0.423]	[0.475]
Salary or Hourly Only	–0.063	0.113	0.020	–0.054	0.127	0.028
(Obs.: Men, 1,396; Women, 1,081)	[0.031]	[0.095]	[0.881]	[0.028]	[0.079]	[0.897]
			Panel B: Job Changers Only			
Piece Rate	0.227	0.228	–0.25	–0.021	0.116	0.035
(Obs.: Men, 79; Women, 32)	[0.383]	[0.213]	[0.510]	[0.313]	[0.219]	[0.594]
Bonus	0.221	0.083	–0.134	–0.245	0.021	–0.041
(Obs.: Men, 118; Women, 45)	[0.125]	[0.347]	[0.583]	[0.182]	[0.228]	[0.606]
Salary or Hourly Only	0.121	0.174	0.002	–0.102	–0.019	0.038
(Obs.: Men, 693; Women, 278)	[0.102]	[0.107]	[0.820]	[0.071]	[0.091]	[0.842]

Notes: Each cell entry represents the weighted average change in earnings for workers in year $T+1$ ($T=1988$, 1989) who were paid either one of the pay methods in year T. The number of observations refers to Year T. The numbers in brackets represent the transition rates between pay methods from Year T to Year $T+1$.

To further explore these data, Table 2 shows the average change in log earnings between pay methods from one year to the next as well as the associated rate of transition, for men and women. As noted in the theory section, self-selection of high-ability people into piece rate contracts might help explain the finding that workers paid piece rates have higher earnings than salaried workers. We can see that even within jobs, there is movement from one pay method to the next, although more for job changers than for stayers.[9]

Among men, stayers and movers differ considerably in the average change in log earnings. Workers who change employers while still being paid piece rates have a much larger increase, on average, than stayers do. The same is true for salaried and hourly paid workers who move to a piece rate contract with another firm. In fact, stayers who experience the same type of transition lose, on average. For both movers and stayers, changing from a piece rate scheme to a salary or hourly rate is associated, on average, with a decrease in earnings. These patterns are suggestive that some selection is going on, with more able workers who change firms and pay method gaining when they move to a piece rate contract.

Interestingly, the average changes in log earnings for movers and stayers apparently do not fit the same patterns for women as for men, especially in the case of job changers. In fact, as is shown below, some of these gender differences are striking.

[9]The reason each row's percentages do not sum to one is that some workers are paid both a bonus and a piece rate. Also, the subsamples used to compute these figures consist of workers who are observed in two consecutive years. Workers who are present in 1988, absent in 1989, and back in the sample in 1990 are not included.

Estimation and Results

Methods of Pay and the Effect on Earnings

Consider the following log-wage equation:

$$(5) \qquad \ln w_{it} = X_{it}\beta + p_{it}\delta_1 + b_{it}\delta_2 + \eta_{it}$$

$$\eta_{it} = \alpha_i + \varepsilon_{it},$$

where w represents labor earnings per hour of worker i at time t, X is a vector of controls,[10] p_{it} is equal to one if worker i is paid a piece rate at time t and b_{it} is a similar indicator for the presence of a bonus, α_i represents unobserved worker productivity, and ε_{it} is an error term that is independently and identically distributed.

Previous studies, such as Brown (1992), Seiler (1984), and Ewing (1996), have all found that piece rate workers earn more than either salaried workers or hourly paid workers. For comparison purposes, I first estimate equation (5) using OLS to determine whether the same empirical pattern is present in these data. It is not clear *a priori* what sign should be expected for δ_2. If most of the bonuses paid have more in common with piece rates than with merit pay, we should expect an effect similar to that for piece rates. If, instead, these bonuses are more like merit pay cash rewards, then according to models such as the one in MacLeod and Malcomson (1989), there is no reason to expect workers paid bonuses to earn more, on average, than other workers.

According to Lazear (1986), the empirical regularities found in previous studies can be attributed to workers selecting the type of contract they prefer according to their productivity. More able workers will self-select into piece rate jobs, while other workers choose time-rated jobs (or straight salaries). Thus, any measured positive effect of piece rates on earnings cannot be attributed to an incentive effect. In other words, switching all time-rated workers to piece rate jobs would not cause an increase in productivity (and wages). In terms of equation (5), this means that p_{it} is positively correlated with α_i. Consequently, if the positive effect of piece rates on wages is due solely to a selection bias, estimating (5) with fixed-effects should result in the estimated δ_1 being close to zero. On the other hand, if piece rates (and bonuses) provide workers with better incentives than a salary or an hourly rate, the measured effects should not disappear once fixed-effects are used.

However, even if piece rate workers are found not to enjoy any wage premium once fixed-effects are used, it would be premature to conclude that incentives do not play a role. Firms choose the pay method that maximizes profits, given all the constraints, be they technological or informational. If incentives can be provided by paying salaries or hourly rates, then one should not be so surprised that in a cross-section of workers, no pay methods seem to bring about higher wages compared to the others. It could just mean that all pay methods are equally successful in getting workers to put forth a satisfactory level of effort. Essentially, finding a wage effect even after controlling for selection indicates that incentive effects are there, but finding no such wage effect does not imply that incentives are not provided.

A second prediction from Lazear's model is that firms adjust rates upward in the first period of the employment relationship to provide sufficient incentives for workers to give the first-best level of effort, even if the workers know that employers cannot be prevented from behaving opportunistically in the second period by decreasing the rates once they know the productivity of their workers. If the bonus measure included in the NLSY data is like a form of piece rate for which workers get rewarded if they surpass a quantitative target, then this strategic interaction should also be at work.[11] If, instead, bonuses reported in this

[10]This vector includes the number of years of schooling, actual labor market experience and its square, employer tenure and its square, and dummies for gender, race, industry, occupation, region, and health status. Also included is a dummy indicator for increase in responsibilities.

[11]See Oyer (1998) for an analysis of the interaction between incentives and sales seasonality. He showed how salespersons may manipulate an incentive scheme by varying their performance over their firms' fiscal years.

Table 3. Earnings Function Estimates.
(Dependent Variable: Log of Real Hourly Labor Income)

	Men			*Women*		
Independent Variable	*1* *OLS*	*2* *Fixed-Eff. (Within-Worker)*	*3* *Fixed-Eff. (Within-Job)*	*1* *OLS*	*2* *Fixed-Eff. (Within-Worker)*	*3* *Fixed-Eff. (Within-Job)*
Piece Rate	0.131	0.071	0.061	−0.015	−0.017	0.004
	(0.027)	(0.027)	(0.033)	(0.031)	(0.035)	(0.050)
Piece Rate × Tenure	−0.014	−0.019	−0.017	−0.000	0.013	0.007
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.009)
Bonus	0.105	−0.016	−0.005	0.038	0.040	0.020
	(0.023)	(0.022)	(0.024)	(0.027)	(0.026)	(0.031)
Bonus × Tenure	−0.011	0.008	0.006	0.012	−0.006	−0.004
	(0.005)	(0.005)	(0.005)	(0.006)	(0.006)	(0.007)
R-Squared	0.438	0.892	0.943	0.430	0.894	0.931
Sample Size	4,582	4,582	4,582	3,555	3,555	3,555

Notes: Other covariates include the number of years of schooling, experience and its square, and dummies for race, industry, occupation, year, region, health status, and increase in responsibilities.

data set are more akin to merit pay, then no such strategic interaction is at play. Although I do not have the actual rates paid to these workers and hence cannot directly see whether these rates are high early in the employment relationship only to be cut later, one can get an idea of such strategic behavior by interacting the incentive pay dummies with employer tenure. According to Lazear, if rates are set higher early in the employment relationship than subsequently, the measured effect of being paid a piece rate (and, possibly, bonuses) should be a negative function of tenure. We should thus expect a negative sign for the interaction term.

The results are reported in Table 3. Looking at the results for men, we can see in column 1 that OLS estimates indicate, consistent with prior evidence, that piece rate workers and bonus workers earn higher average wages than workers paid on a hourly basis or paid salaries. Interestingly, once fixed-effects are used, the positive correlation between the use of a piece rate and wages does not disappear, and the effect of being paid a bonus vanishes completely. The same conclusion basically holds when all variables are measured in deviation from job-match means, thereby sweeping out any fixed unobserved matching effect (column 3). The only difference between columns 2 and 3 is that the combined effect of a slightly reduced coefficient with a larger standard error makes the result borderline in statistical significance (p-value of 0.0661).

These results suggest that piece rates, but not bonuses, provide extra incentives to increase productivity and wages. Note, however, that while the total effect of piece rates using OLS, taking into account the negative interaction term with tenure, is 0.086 with an associated standard error of 0.026 when evaluated at the average tenure level for men (3.2 years), the within-job effect is virtually zero when evaluated at 3.2 years of tenure.

In the case of bonuses among male workers, the positive impact measured with OLS merely seems to reflect omitted ability variable bias. Thus, controlling for unobserved worker productivity, there is no evidence that workers paid bonuses earn a premium over those on an hourly or salary pay scheme. This is consistent with MacLeod and Malcomson's (1989) model of how labor market conditions affect the form of the contract (that is, base pay plus a bonus

versus straight salary) while the total amount paid to workers remains the same.

However, this evidence should not be construed as showing that bonuses have no effect. Given the wording of the question on bonuses, an affirmative answer to that question does not necessarily imply that a bonus was actually received by the worker, only that the respondent's earnings are based partly on bonuses. Hence, an attenuation bias caused by measurement error cannot be ruled out. Furthermore, this problem is compounded when I use fixed-effects.[12]

As for women, the results are quite striking and strongly suggest that whether one uses ordinary least-squares or fixed-effects methods, "incentive contract" workers do not earn more than other workers. It has long been known that women's patterns of labor force attachment differ from men's due to (for example) fertility decisions. Whether these forces also affect the role incentives play for female workers is an unexplored question to which I return below.

I now turn my attention to the possibility of strategic behavior between workers paid piece rates and their employers (that is, the ratchet effect). Looking first at the OLS results for men, we can see that the effect on wages seems to decline with tenure with the employer. However, it would be rash to interpret the negative interaction term as reflecting the ratchet effect. As Table 2 shows, the self-selection of workers who move from one pay method to another could bias the results in various ways. The descriptive statistics in the left portion of Table 2 (the estimates for men) show that workers who change employers and move to a piece rate contract gain on average, while those who stay with their employer and move from a salary or hourly rate to a piece rate lose on average. In fact, the general pattern among men is for job changers to gain more than stayers from switching to piece rates. The fact that those gainers who move to a new employer are then observed at lower tenure levels whereas stayers who lose from switching to a piece rate contract are observed at higher levels of tenure would, by itself, tend to produce a negative interaction term.

On the other hand, there are other transitions in Table 2 that would tend to bias the interaction coefficient in the other direction for men. In Lazear's symmetric information case (that is, neither the firm nor the worker knows the worker's productivity ex ante), only higher ability people stay in piece rate jobs. The estimates for men in the first row of Table 2 show that for both job changers and job stayers, moving from a piece rate to a salaried job or hourly rated job is associated with a substantial decrease in earnings. Thus, looking at stayers only, if the pool of piece rate workers tends to be of higher quality as tenure increases, that would tend to produce a *positive* interaction term in an OLS regression. Overall, it is not clear which form of bias would dominate.

The results in columns 2 and 3 of Table 3 indicate that the coefficient is actually slightly larger in absolute value with the use of fixed-effects.[13] Thus, differencing out all unobserved individual and job match heterogeneity reinforces the OLS result and tends to point toward positive selection of workers into piece rate jobs as tenure increases.[14] I must emphasize again that these data do not contain information on the actual rates paid to workers (or the production standards), which would make it easier to analyze the ratchet effect and perhaps make the results more convincing. Still, the fact that in a cross-section of workers, those paid piece rates tend to see their wage premium (be it an ability premium or an incentive premium) decrease with tenure compared to time-rated workers seems

[12]See Card (1996) for an analysis of the effect of union misclassification on the bias of OLS estimates.

[13]The difference, however, is not statistically significant.

[14]By positive selection, I mean the process by which higher-ability people select themselves into the piece rate jobs.

to be consistent with some ratcheting of the rates/standards.

Another reason for the existence of such a strong effect of piece rates on earnings at low tenure levels could be career concerns (Gibbons and Murphy 1992). Workers may respond to incentives not only in order to maximize current earnings but also to affect other people's perceptions. Workers who are successful in doing so may improve their chances of getting promotions or future wage increases. Thus, a prediction of career concerns models is that since relatively little is known about workers new to their positions, effort will be higher at lower tenure levels. Note that career concerns considerations apply to all workers, not only to those who are paid through explicit incentive schemes. Consequently, if it were true that career concerns affect everyone irrespective of pay method, the interaction term in Table 3 should be zero. The fact that it is not may mean that it is relatively easier for a young worker to affect the perception of others when that worker is paid a piece rate.

In the case of workers paid bonuses, there is no evidence of a negative relationship between tenure and the occurrence of bonuses once fixed-effects are used.

Again, women have totally different patterns. In terms of the selection explanation given above, it is interesting to note that the right-hand portion of Table 2 shows no substantial difference in the average log earnings change between female job changers and job stayers: both lose, on average, from moving to a piece rate contract, at least when their previous pay method was a salary or an hourly rate.

Why Are the Results Different for Women?

Two potential explanations come to mind. The first one relates to the fact that women may work in occupations in which ability does not matter quite as much as it does for men, either because women choose to work in those occupations or because of occupational segregation. Controlling for one-digit occupation is but a crude way of trying to condition on such considerations. To see whether refining the occupation categories would make any change to the results, I re-estimated the model with controls for, respectively, two- and three-digit occupation. In both cases, the results were unchanged. In other words, even for workers in the same closely defined occupation, those who are paid a piece rate do not earn more, on average, than others.

A second avenue worth exploring is the role of fertility. Is having one or many children at home influencing female workers' responsiveness to incentive schemes differently from men's? I examine that possibility by interacting the piece rate dummy with a dummy for the presence of dependents at home. I construct two measures for the dummy indicating the presence of at least one dependent. From 1979 to 1988, respondents were directly asked the following question:

> Not counting (yourself/yourselves), but including your children, how many persons are dependent upon you [or your (husband/wife)] for at least one-half their support?

Unfortunately, the following interviews only asked for the number of children in the household, which may not be as sharply focused a question as the previous one, given that there may be children in the household for which the respondent is not responsible. In 1989 respondents were also asked whether they had had a child since the previous interview. Thus, the first measure uses the 1988 question on dependents for those that are in the sample in 1988 and then, for 1989 and 1990, I combine the question on the number of children at home with the 1989 question on having had a child since the last interview to construct a dummy variable equal to one if the respondent has what I call at least one dependent at home, and zero otherwise.

There are three potential problems with using that measure: the presence of children in the household does not necessarily imply that those children are dependents; many women who do give birth to a child (and, possibly, men who declare having had a child) are simply excluded from the

Table 4. Further Results: The Impact of Dependents.
(Dependent Variable: Log of Real Hourly Labor Income)

	Men				*Women*			
Independent Variable	*1* OLS	*2* *Fixed-Eff. (Within-Job)*	*3* OLS	*4* *Fixed-Eff. (Within-Job)*	*1* OLS	*2* *Fixed-Eff. (Within-Job)*	*3* OLS	*4* *Fixed-Eff. (Within-Job)*
Piece Rate	0.142	0.096	0.110	0.098	0.078	0.034	0.141	0.031
	(0.030)	(0.039)	(0.034)	(0.034)	(0.038)	(0.065)	(0.044)	(0.063)
Piece Rate × Dependent	–0.047	–0.07	–0.048	–0.041	–0.163	–0.047	–0.213	–0.049
	(0.035)	(0.039)	(0.039)	(0.042)	(0.040)	(0.061)	(0.045)	(0.061)
R-Squared	0.430	0.944	0.437	0.936	0.429	0.931	0.425	0.926
Sample Size	4,582	4,582	3,667	3,667	3,555	3,555	2,730	2,730

Notes: Other covariates are the same as in Table 3. See text for details on the specifications.

sample because they are not in the labor force at the time of the interview or they have not worked a sufficient number of weeks during the year; and the decision to have children is endogenous, and simply interacting the piece rate dummy and the presence of dependents dummy establishes no more than a statistical relationship.

For comparison purposes, I also construct the following measure: only those workers who are present (for the first time in the sample) in 1988 are included in the estimation, and the answer they give to the 1988 question on the number of dependents is attributed to their record for interview years 1989 and 1990. One reason for using only that question is that it may largely reflect past fertility decisions, which can be taken to be exogenous. The other reason is that the question itself is more focused. Also, given that these workers are very young, it is unlikely that one who has dependents at home in 1988 does not have any in the following years.[15]

The results are presented in Table 4, where specifications 1 and 2 refer to the first measure while specifications 3 and 4 refer to the second measure discussed in the preceding paragraph. The results using OLS provide strong evidence that women who have dependents may not be able to respond quite as much to an explicit incentive scheme as can (a) women who have no dependents or (b) men. When the interaction term is included, the positive relationship between piece rates and earnings emerges quite clearly. This is especially apparent in specification 3, where the dummy for the presence of dependents is constructed from the sharper 1988 question.[16]

Note that adding the interaction term for men does not markedly change the results shown in Table 3, although it tends to increase the effect of piece rates when the within-job estimator is used. To be sure, the impact of the interaction term is nowhere near what it is for women. Thus, given the results in columns 2 and 4 of Table 4, it appears that selection according to ability matters also for women and that this selection effect is largely hidden when we ignore the presence of children. Concerning the fixed-effects results (columns 2 and 4), part of the puzzle still remains, although it must be noted that the coefficients are not very precisely estimated: while it is true that one cannot reject the null of a zero effect, neither can one reject the null of a 10% effect. Also, controlling for family responsibilities does not change the result

[15]Of course, had the same question been asked in 1989 and 1990, one could check that conjecture.

[16]In column 3, the piece rate effect for women is 0.029 when the interaction term is excluded and is not statistically significant.

Table 5. Minimum Distance Estimation of Covariance Structure, by Gender. (Equally Weighted Minimum Distance Estimation; Standard Errors in Parentheses)

	Men			*Women*		
Parameter	*Salaried and Hourly Paid Workers*	*Piece Rate Workers*	*Bonus Workers*	*Salaried and Hourly Paid Workers*	*Piece Rate Workers*	*Bonus Workers*
Variance of Unobserved Worker Ability	0.115 (0.004)	0.210 (0.012)	0.099 (0.010)	0.097 (0.002)	0.142 (0.026)	0.089 (0.016)
Variance of i.i.d Component	0.050 (0.005)	0.028 (0.017)	0.052 (0.014)	0.044 (0.002)	0.061 (0.037)	0.060 (0.022)
Chi-Square Statistic	8.7	2.5	5.1	0.5	6.1	3.2
Degrees of Freedom	4	4	4	4	4	4
Number of Workers	1,832	317	433	1,554	214	284

of an essentially zero effect for the tenure-piece rate interaction term, even with OLS.

Methods of Pay and Wage Dispersion

The closer link between productivity and pay that is a characteristic of piece rate contracts should be reflected in the fact that most of the unexplained variance resulting from the estimation of equation (5) is attributable to the variance of unobserved worker productivity in the case of piece rate workers. The first step to verifying this prediction is to break the sample into three subsamples according to method of pay (piece rate, bonus, others).[17] Then, exploiting again the longitudinal aspect of the data set, I use the log wage observations adjusted for schooling and labor market experience to estimate a simple model of the covariance structure of wages for each type of pay method, using Chamberlain's (1982, 1984) minimum distance estimator adapted to unbalanced data.[18] In other words, the approach used is first to regress (by year) the log wage on labor market experience and schooling and then to use the estimated residuals to compute estimates of the unrestricted covariance matrix of residuals. I then impose the restrictions implied by the simple error component structure given in equation (5) to estimate the parameters of interest and to test the fit of the model (see the appendix for details).[19]

The results in Table 5 show that the simple two-component model fits the data quite well, as all goodness-of-fit statistics are unsurprising values coming from a χ^2 (4) distribution. Looking first at men, note that the relative contribution of unobserved worker productivity is much larger for piece rate workers than for other workers. This is consistent with the view that pay is sensitive to differences in productivity for piece rate workers. Thus, although the simple explanation of worker self-selection into piece

[17]More precisely, if a worker is paid according to a piece rate contract in 1988 (say), then that worker is part of the "piece rate subsample." If, in 1989, the same worker is paid a bonus, then he or she belongs to the "bonus subsample." If the worker is paid both a piece rate and a bonus, whether during the same interview year or in different years, I include that person in both subsamples. If a worker reports a piece rate one year and neither a piece rate nor a bonus in another year, then that worker belongs to the piece rate subsample for the year in which he or she reports a piece rate, and belongs to the "others" subsample for the year in which she or he reports only a salary or an hourly rate.

[18]See Abowd and Card (1989) and Farber and Gibbons (1996) for applications of these techniques.

[19]Note that the results are not sensitive to different specifications of the log wage equation estimated in the first step. Various specifications including controls for tenure, occupation, and industry were used without altering the basic conclusions.

rate jobs does not seem to be the whole story behind the wage premium that these workers earn, the prediction of Lazear's (1986) model concerning the dispersion of wages for incentive workers is strongly supported. Note also that since controlling for unobserved ability accounts for a sizable portion of the variance, Seiler's (1984) conjecture that piece rate workers are rewarded for facing higher income risks is not consistent with the results reported here. The results, at least for men, show that much of the cross-sectional earnings variance for piece rate workers is due to unobserved worker heterogeneity, not to random shocks.[20]

In Lazear's model, the reason earnings are more dispersed for piece rate workers is that output is more tightly linked to variation in individual productivity. A related explanation for the higher variance could also be that for firms paying piece rates, the underlying production technologies allow heterogeneity in individual productivity. However, if the work environment is characterized by teamwork (Holmstrom 1982) or multitasking (Holmstrom and Milgrom 1991, 1994), then it might not be possible for firms to accommodate large differences in productivity. The presence of teamwork is, in some sense, equivalent to having higher costs for monitoring individual output, which, in Lazear's framework, makes piece rates less desirable. However, the provision of incentives in a multitasking environment is a different problem. It might be very cheap to monitor each worker's output in such an environment, but piece rates are likely not to be used if firms care about other aspects of the job, such as equipment maintenance or output quality.[21]

[20]Seiler's explanation could still be true if workers did not know their own productivity levels. Given the possibility that they may turn out to be low-productivity individuals, risk-averse workers would want to be insured against that risk by asking for a premium.

[21]For an analysis of the impact of these and other factors on the form of the contract, see MacLeod and Parent (1998).

Again, the results are not so clear-cut for women, although it is still true that the relative contribution of worker heterogeneity to the total variance is larger for piece rate workers than for other workers.

Conclusion

I have found that the wage premium enjoyed by male piece rate workers over salaried or hourly paid workers does not completely vanish once proper controls are entered for unobserved worker and job-match heterogeneity. For women, there is no such evidence of a premium when fixed-effects are used, although controlling for family responsibilities does make the OLS results similar to the results for men.

The second most important result of the study is that cross-sectional earnings dispersion for piece rate workers is largely accounted for by unobserved worker characteristics, especially for men. As for the part played by tenure, the finding that for men the wage premium tends to decrease with tenure seems to result at least in part from different selection patterns for job stayers compared to movers.

Some aspects of the data on methods of pay contained in the NLSY restrict the freedom with which the results presented in this paper can be interpreted. Notably, because the dollar amounts paid as "incentive" pay are not known, it may be that for many workers, much of their labor income is in the form of hourly rates or salaries, which are *supplemented* by a certain amount of income earned through pay-for-performance schemes.

Still, the results provide evidence that, consistent with Lazear's (1996) findings with company data, the wage effect of piece rates is not all a product of selection: there seems to be an influence of incentives as well, at least for men. In fact, an incentive effect accounts for close to half of the total wage effect identified using ordinary least squares analysis.

APPENDIX

Let e_{it} be the residual from a first-stage regression of the log-wage on schooling and labor market experience. Then we have

$$e_{it} = \ln w_{it} - X'_{it}\hat{\gamma}, \tag{6}$$

where X'_{it} is the vector of controls (experience and schooling) and $\hat{\gamma}$ is the vector of the estimated parameters. Then a natural estimator of the covariance between the residuals at time t and at time s is given by

$$\hat{\sigma}_{t,s} = \frac{1}{N}\sum_{i=1}^{N} e_{it}e_{is}, \tag{7}$$

where N represents the number of workers.[22] A consistent estimator for the standard error of $\hat{\sigma}$ is then given by

$$Std(\hat{\sigma}_{t,s}) = \sqrt{\frac{1}{N}\sum_{i=1}^{N}(e_{it}e_{is} - \hat{\sigma}_{t,s})^2}. \tag{8}$$

Turning now to the restrictions imposed by the error component structure suggested by equation (5), I assume that the observed residual is given by

$$e_{it} = \alpha_i + \varepsilon_{it}. \tag{9}$$

Therefore, the variance of the observed residuals at period t is

$$var(e_{it}) = E(\alpha_i)^2 + E(\varepsilon_{it})^2 = \sigma_\alpha^2 + \sigma_\varepsilon^2, \tag{10}$$

and the covariance between residuals at periods t and s is

$$cov(e_{it}e_{is}) = E(\alpha_i)^2 + E(\varepsilon_{it}\varepsilon_{is})^2 = \sigma_\alpha^2,$$

since I am assuming that ε is i.i.d. Thus, the model to be estimated has the following linear structure:

$$\sigma_{ts} = \theta_1 + \theta_2 D_{ts}, \tag{12}$$

where $D_{ts} = 1$ if $t = s$ and $D_{ts} = 0$ otherwise, while $\theta_1 = \sigma_\alpha^2$ and $\theta_2 = \sigma_\varepsilon^2$. With 3 years of data, that means I have 6 distinct elements (per method of pay) to try to fit with the two-component of variance representation given by equations (10) and (11). The covariance model given by these two equations militates that all diagonal elements should be equal and that all off-diagonal elements should also be equal. This gives a total of four restrictions to be imposed on the data.

To test these restrictions, I make use of the minimum distance estimator proposed by Chamberlain (1982, 1984). Succinctly, let m_i be the column vector of the unique elements of the 3×3 cross-product matrix of residuals for workers i. Then I calculate, for each method of pay, the sample means of the elements of m_i, giving us the column vector m. The model I want to estimate is a model for m. A consistent estimator of the covariance matrix of the covariance elements in m is given by

$$\Omega = \frac{1}{N}\sum_{i=1}^{N}(m_i - m)(m_i - m)'. \tag{13}$$

The equally weighted minimum distance estimator (EWMD) minimizes the following quadratic form with respect to Θ:

$$\Pi = [m - D\Theta]'[m - D\Theta], \tag{14}$$

where D is the design matrix made of ones and zeros while $\Theta = (\sigma_\alpha^2, \sigma_\varepsilon^2)$ is the vector of parameters to be estimated. On the other hand, the optimally weighted minimum distance estimator (OMD) minimizes with respect to Θ the quadratic form

$$\Pi = [m - D\Theta]'\Omega^{-1}[m - D\Theta], \tag{15}$$

where Ω is the covariance matrix of the vector of covariance elements. Under the null hypothesis of a correct specification in equation (12), the value of the objective function (15) is distributed as a χ^2 with degrees of freedom given by the number of restrictions (the difference between the number of sample moments to fit [6] and the number of parameters [2]).[23] Altonji and Segal (1994) discussed how small sample size can lead to a substantial bias when the OMD estimator is used in comparison to using the identity matrix as the weighting matrix. Results are thus presented for the more robust equally weighted minimum distance estimator, which uses an identity weighting matrix.[24]

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[22]Note that because the data are unbalanced, the number of workers used to compute the estimate is not the same for each covariance element. To simplify the exposition of the methodology, I assume that the data are balanced.

[23]The computation of the χ^2 statistic for the EWMD estimator differs somewhat different from that for the omd. See the appendix in Abowd and Card (1989) for the details.

[24]Results with the optimally weighted minimum distance estimator are very similar.

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