Performance Pay and Earnings Dynamics*

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Abstract

Using data from the Panel Study of Income Dynamics, we study how the autocovariance structure of wages and earnings differs under different contractual arrangements. We divide jobs on the basis of whether they pay for performance, and whether they are covered by collective bargaining agreements. While cross-sectional wage inequality is larger in performance-pay than non-performance-pay jobs, precisely the opposite happens in the case of annual earnings. This suggests that hours of work respond more to demand shocks when wages are inflexible (in non-performance-pay jobs) than when they are flexible because of performance-pay schemes. This result only holds, however, for purely cross-sectional measures of inequality. Since the variation in hours is mostly transitory, from a long-run perspective inequality remains larger in performance-pay than non-performance-pay jobs.

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1 Introduction

Numerous studies suggest that pay-setting institutions and contractual arrangements in the labor market play an important role in labour market outcomes. For instance, Blau and Kahn (1996) have shown that differences in union coverage are related to cross-countries differences in inequality across countries. Card (1996), Freeman and Needels (1993), and DiNardo, Fortin, and Lemieux (1996)) find that a substantial part of the growth in U.S. wage inequality in the 1980s and 1990s is attributable to the decline in the rate of unionization. While part of the relationship between unions and wage dispersion may be due to other unmodelled factors such as technological change, recent work by Fransen (2012) does find evidence that the effect is "causal" by comparing that wage distribution within firms where unions have barely won or lost a union-organizing election.

That said, data on aspects of labor contracts besides union coverage are not generally available for large and representative cross-sections of the workforce. Using a proxy for the presence of performance-pay contracts, Lemieux, MacLeod, and Parent (2009) find that performance pay increases wage inequality, and that it has contributed to the growth in wage inequality in the United States, especially at the top-end of the distribution.[1] Intuitively, performance-pay workers should have wages that better reflect productivity than workers who are paid a fixed, or union-set wage. Consistent with this view, Lemieux, MacLeod, and Parent (2009) find that the return to education and broader measure of wage inequality are indeed larger in performance-pay than fixed-wage jobs.

But while unions and fixed-wage contracts may yield a lower level of cross-sectional wage inequality, it is not clear they necessarily reduce broader measure of welfare inequality. In particular, when wages are rigid workers may be more likely to experience unemployment or reduced hours in response to negative demand shocks than in cases where wages are more closed tied to productivity through the use of performance-pay, profit-sharing, or other flexible wage arrangements. Once the hours dimension is taken into account, the dispersion of total earnings or income could actually be larger under fixed-wage than performance-pay contracts.

A second important issue is that, from a life-cycle point of view, inequality comparisons under different contractual arrangements depend on the dynamic response of wages and hours of work to demand or productivity shocks. If these shocks are very transitory in nature, they

[1] See also Lazear and Shaw (2009) and Guadalupe and Cunat (2009).

only yield short-term variability in earnings that has little impact on welfare over the whole life cycle. In such a setting, one has to look at the whole autocovariance structure of earnings and hours to assess the full impact of contractual arrangements on welfare.

In this paper, we use data from the Panel Study of Income Dynamics (PSID) to look at how dimensions of contractual arrangements –performance-pay and, to a smaller extent, unionization– affect the whole autocovariance structure of earnings and hours. A first important finding is that the cross-sectional variance of annual earnings is smaller under performance-pay contracts, which is the opposite of what Lemieux, MacLeod, and Parent (2009) found in the case of hourly wages. We interpret this as evidence that fixed-wage contracts result in involuntary unemployment, thereby increasing the dispersion in annual earnings among workers.

We then turn to a full analysis of the autocovariance structure of hourly wages and annual earnings. Descriptive evidence indicates that while the variance of annual earnings is large in fixed-wage than performance-pay jobs, this is all due to a higher transitory dispersion in earnings that lasts for no more than a few years. Estimates based on an error-components model confirm this descriptive evidence. We find that the variance of the permanent component of either hourly wages of annual earnings is larger in performance-pay than fixed-wage jobs because skills are more rewarded in the former type of jobs. By contrast, the variance of the transitory component of annual earnings is larger in fixed-wage jobs. This is consistent with transitory demand or productivity shocks generating a large variability in hours of work in fixed-wage jobs. These results are robust to a variety of specifications allowing for time-varying factor loadings and worker-firm match components. We also find a similar pattern of results when comparing union and non-union jobs.

The rest of the paper proceeds as follows. In Section 2, we propose a simple model where contracts are used to protect specific investments. Because of asymmetric information, employers can either offer a fixed-wage contract, which potentially generates involuntary unemployment, or pay a fixed monitoring cost and offer a performance-pay contract instead. We discuss how wages and hours respond to various shocks under these two contractual arrangements, and derive the implications of the model on the covariance structure of hours wage wages.

In Section 3 we present the error-component model we use to parametrize the autocovariance structure of hours and wages, and test the predictions of the model under different contractual arrangements. In Section 4 we discuss how we measure contract form using the PSID data, and present some descriptive evidence on wages, earnings, and hours of work for performance-pay and fixed-wage jobs. The main estimation results are presented in Section 5, and we conclude in Section 6.

2 Theory

Before discussing how we expect the distribution of earnings and wages and their autocovariance structure to differ under various contractual arrangements, it is important to explain why we expect these contractual arrangements to arise in the first place. In the first part of this section, we present a simple explanation based on the need for contractual arrangements in the presence of relationship specific investment. This simple model also has straightforward predictions on how wages and hours are expected to respond to demand or productivity shocks. We then discuss a more general case with heterogenous workers to get a full set of implications for the whole distribution of earnings and wages, as well as their autocovariance structure.

2.1 Specific investments, asymmetric information, and monitoring costs

This part of the paper relies heavily on Lemieux, MacLeod, and Parent (2012). We build a very simple labor contracting model using three ingredients. The first is the existence of relationship specific investment that both Becker (1962) and Mincer (1962) emphasized as central to the wage setting process. The second ingredient is contract law for which the protection of specific investments, or what legal scholars call the "reliance interest" (see Fuller and Perdue (1936)). In particular, if a worker makes an investment, such as moving to take up a job or acquiring some job-specific skills, an employer cannot after the fact unilaterally reduce wages. Legal doctrines, such as good faith and fair dealing, explicitly limit the scope of firms to reduce wages.[2] When employment contracts cannot be enforced, and firms can unilaterally lower wages, this leads to what Goldberg (1977) calls "holdup", which can result in an inefficiently low level of reliance by both the worker and the firm.

[2] In Rigby v Ferodo [1988] ICR 29, [1987] IRLR 516 the U.K. courts ruled that employers could not unilaterally lower wages.

The final ingredient is asymmetric information. Hart and Moore (1988) have shown that when there is symmetric information, and contracts are incomplete, then rational parties would always renegotiate inefficient agreements, leading to efficient employment *ex post*, which in turn leads to *inefficient* investment *ex ante*. MacLeod and Malcomson (1993) apply these ideas to the employment contract and show that the optimal contract entails fixed wages that are renegotiated in the face of better market alternatives for the worker or the firm. Contracts can be designed to achieve efficiency in a wide variety of cases, including the case of *cooperative investment* - one party's investment affecting both parties' payoff. Such a model, like implicit contract theory, cannot by itself explain inefficient unemployment. (Hart (1988) explicitly makes the point that asymmetric information is central to a theory of unemployment). Following Hall and Lazear (1984), we suppose that there is asymmetric information, which makes it impossible for firms to efficiently modify wages *ex post.* Here, we combine asymmetric information regarding worker productivity with relationship-specific investments to produce a simple model where wages are set in advance, and employment is lower than the first best.[3]

When the measurement of worker productivity is possible, inefficient unemployment linked to fixed wages can be eliminated by introducing bonus pay. Our key identifying assumption is that monitoring costs vary across firms.[4] This allows us to compare empirically performance-pay and fixed-wage jobs, holding worker quality fixed. In this case, the model predicts that bonus-pay workers are more likely to be employed than workers on fixed-wage jobs. Furthermore, demand shocks should have a larger impact on wages and a smaller impact on employment for bonus-pay workers.

More formally, consider a two-period model that illustrates the consequences of specific investment, enforceable wage contracts and asymmetric information, the details of which are found in the model appendix. The timing of choices is as follows:

1. In period 1 the worker can accept a wage offer w^1 from firm A or B and then make a specific investment k^1 - in order to avoid hold-up the firms agrees not to lower the wage in period 2. Firm A pays a monitoring cost c to produce a publicly observable

[3]This is consistent with Card (1986) who finds that wages are rigid in the short run.

[4]Brown (1992) found that the assumption that there is a common fixed cost of monitoring is inconsistent with the evidence, while MacLeod and Parent (1999) show that bonus pay is wide spread in the labor market, with its incidence varying with job characteristics.

signal $s \in \{0, 1\}$ that is positively correlated with firm productivity. Firm A agrees to pay a bonus b if $s = 1$.

2. If the worker rejects both offers, then she waits until the firms realize their productivities and make new wage offers based upon these realized productivities. She can take up either offer at cost $k^2 > k^1$.

3. If she accepts the offer from say A, then in period 2 firm B can offer a new wage w^2 based on its productivity that she can choose to accept at a cost k^2.

4. Before production begins, firms can choose to layoff workers. Since firm productivity is private information, workers will refuse to renegotiate wages down in the event the firm has low productivity. Such behavior is a necessary ingredient of any wage contract consummated in period 1.

Suppose that *ex ante* the worker is better matched at A. This model illustrates in the most simple form the trade-off between early investment into firm A, against the option value of delaying investment if it turns out that either firm B or exit from the labor market is *ex post* optimal. Notice that asymmetric information is a key ingredient here. If worker productivity is information that is private to the firm, then workers would never agree to wage cuts because the firm may simply be misrepresenting its costs. Wages can only respond to credible signals, such as a wage offer from another firm.[5]

The model generates a number of empirical implications -discussed in the appendix- on how contract form matters for wages and employment. Most importantly, we expect employment to be higher and less responsive to demand shocks in performance-pay jobs since fixed-wage workers get laid off whenever productivity falls below w^1. The presence of a bonus makes the wage rate (inclusive of bonus) more sensitive to demand shocks than under fixed wages, and hence the worker faces a lower probability of layoff under performance pay.

2.2 Heterogenous workers

Now consider what happens when workers are heterogenous, and both firms and workers only have imperfect information about workers' productive ability. Lemieux, MacLeod, and

[5] See Carmichael and MacLeod (2003) for a discussion of the credibility of such behavior when contracts are not legally enforceable.

Parent (2009) have looked at this issue in a simple model where information is imperfect but symmetric, and where performance pay also provides an incentive for workers to provide effort. While generalizing this model to the dynamic case with specific investments and asymmetric information about demand shocks is beyond the scope of this paper, we believe that most of the predictions derived by Lemieux, MacLeod, and Parent (2009) would still hold here.

In the presence of imperfect information, fixed-wage contracts will be based on the expected productivity of the worker given the information available at the time the worker is hired. In principle, the worker and the employer could learn about the productivity of the worker with the passage of time, as in a standard Jovanovic learning model. In the absence of an explicit measurement technology, however, learning about workers' productivity is expected to happen at a much slower rate than in performance-pay jobs. For the sake of simplicity, we ignore these dynamic considerations here and simply assume that fixed-wage contract pay a wage equal to the expected value of output at the time the worker is hired. Likewise, the base wage in a performance-pay contract would also depend on expected productivity, but the bonus part of the contract would depend on actual productivity that is (imperfectly) captured by the signal s.

The important implication of this assumption is that wages will be more dispersed in performance-pay than fixed-wage jobs. For a given information set available at the time the worker is hired, more productive workers will get a positive productivity signal more often than less productive workers. As a result, they will be more likely to receive a bonus, which will increase wage (inclusive of bonus) dispersion among workers with similar characteristics. Intuitively, wages are more dispersed among performance-pay workers because they are more closely tied to productivity than in fixed-wage jobs.

2.3 Wage setting equations

Consider the (log) wage of worker i at time t under a performance-pay contract:

$$w_{it}^{p} = a^{p} + x_{it}b^{p} + c^{p}u_{it} + d^{p}\theta_{i} + \varepsilon_{it}^{p}, \tag{1}$$

where x_{it} represents standard observable characteristics such as potential experience, education, occupation, θ_i is the unobservable ability component; u_{it} is a productivity (or other demand-side) shock; ε_{it}^{p} is an idiosyncratic error term. Similarly, the wage under a fixed-wage contracts (i.e. on a "non-performance-pay jobs") is given by

$$w_{it}^{n} = a^{n} + x_{it}b^{n} + c^{n}u_{it} + d^{n}\theta_{i} + \varepsilon_{it}^{n}. \tag{2}$$

When wages are completely fixed in advance, we expect the shock u_{it} to have no effect on wages at all, and c^{n} to be equal to zero. By contrast, under performance pay at least part of the shock should be passed on to wages, so that $c^{n} > 0$. As we discuss in the data section, however, we only have imperfect measures of whether workers are on performance-pay or fixed-wage contracts. Therefore, we don't generally expect to find that $c^{n} = 0$ in the data. That said, if the way we divide performance-pay and fixed-wage jobs is informative we should still expect to find $c^{p} > c^{n}$. In light of our discussion of the case of heterogenous workers, we also expect that the factor loading on both observable characteristics and unobserved ability to be larger in performance-pay and fixed-wage jobs, i.e. $b^{p} > b^{n}$ and $d^{p} > d^{n}$.

Now at consider the case of annual earnings for performance-pay

$$y_{it}^{p} = a_{y}^{p} + x_{it}b_{y}^{p} + c_{y}^{p}u_{it} + d_{y}^{p}\theta_{i} + \xi_{it}^{p}, \tag{3}$$

and fixed-wage workers:

$$y_{it}^{n} = a_{y}^{n} + x_{it}b_{y}^{n} + c_{y}^{n}u_{it} + d_{y}^{n}\theta_{i} + \xi_{it}^{n}. \tag{4}$$

As we discussed above, a shock u_{it} should have a larger impact on annual hours of work under fixed-wage contracts since none of the adjustment to, say, a negative shock is absorbed in terms of lower wages. Depending on how large the hours adjustment is relative to the wage adjustment, c_{y}^{n} may either be larger or smaller than c_{y}^{p}. By contrast, we still expect the factor loadings on x_{it} and θ_{i} to be larger in performance-pay than fixed-wage jobs ($b_{y}^{p} > b_{y}^{n}$ and $d_{y}^{p} > d_{y}^{n}$). The intuition is that, if anything, paying a higher wage to higher ability workers should make them work more hours if the labor supply elasticity is positive.

Note that since we are working in logs, annual hours in performance-pay and fixed wage jobs are given by $h_{it}^{p} = y_{it}^{p} - w_{it}^{p}$ and $h_{it}^{n} = y_{it}^{n} - w_{it}^{n}$. As a matter of convenience, we prefer working with wages and earnings, but the exact same results would be obtained working with wages and hours instead.

2.4 Unions

At this stage, we simply view unionization as an additional indicator of the type of contract involved. While the above model draws a sharp contrast between performance-pay and fixed-wage contracts, we will see below that there are some difficulties in measuring these two concepts empirically. For simplicity, we expect that unionization gets us even closer to a fixed-wage setting, since collective bargaining agreements indeed tend to pre-specify wages over the duration of the contract (2-3 years, often more in recent years). Interestingly, however, some union contracts allow for a limited amount of pay-for-performance, which generates more flexibility in response to labor market shocks.

In practice, we divide contracts into up to four categories based on union and performance-pay status. We expect union contracts without performance pay to exhibit the least wage flexibility and the largest hours response, and non-union contracts with performance pay to be the most flexible.

3 Empirical Model for the Autocovariance Structure of Wages and Earnings

The c and d parameters (factor loadings) in equations (1), (2), (3), and (4) determine how demand shocks u_{it} and unobserved ability θ_i map into wages and earnings. Since we expect demand shocks to be transitory while unobserved ability is a permanent component of the error term, the factor loadings have important implications for the nature of wage and earnings inequality. A larger d factor loading increases the permanent component of inequality, while a larger c factor loading increases transitory inequality.

Starting with Moffitt and Gottschalk (1994), there is a large literature that seeks to examine whether the large increase in U.S. wage inequality is driven by the transitory or permanent component of inequality. Knowing which of the two components accounts for the bulk of the growth inequality has important implications in terms of welfare.

Our focus is slightly different here since we are not primarily interested in understanding why inequality has increased over the last few decades. Our main focus is instead in estimating the impact of contract form on the transitory and permanent components of wage and earnings inequality. That said, our analysis has some indirect implications for this debate since contract form has been changing over time with the decline of unionization and the growth of performance-pay contracts (Lemieux, MacLeod, and Parent (2009), Bloom and Van Reenen (2011), and Dube and Freeman (2011)).

We explore these issues formally by estimating parametric models for the autocovariance structure of wages and earnings. As is common in this literature (e.g. Abowd and Card (1989)), we first "residualize" wages and earnings by running regressions on a rich set of covariates such as experience, education, etc.[6] Note that we are focusing our analysis on the "residual" or "within-group" component of inequality, as opposed to the more systematic "between-group" component. Lemieux, MacLeod, and Parent (2009) show that performance-pay also tends to increase between-group inequality (higher return to education) and, thus, the permanent component of inequality. We don't look at this inequality component here since we are more interested in the impact of idiosyncratic demand/productivity shocks that affect the within- as opposed to the between-group component of wages and earnings.

Following Lemieux, MacLeod, and Parent (2009), we also consider a possible job-match component ν_{ij} in the wage and earnings equation. In their model, the variance of the job-match component is expected to be lower in performance-pay than other jobs. The idea is that when workers are paid for performance, the particular job they have should not matter as much as their own productive ability. Our main focus here is simply to control for that variance component to make sure it does not confound the estimates of the variance of unobserved heterogeneity ($Var(\theta_i)$) and demand shocks ($Var(u_{it})$).

Consider the residual wage for performance-pay jobs, $\widetilde{w}^p_{ijt}$:

$$\widetilde{w}^p_{ijt} = c^p u_{it} + d^p \theta_i + \nu^p_{ij} + \varepsilon^p_{it}, \tag{5}$$

and for non-performance-pay jobs, $\widetilde{w}^n_{ijt}$, is:

$$\widetilde{w}^n_{ijt} = c^n u_{it} + d^n \theta_i + \nu^n_{ij} + \varepsilon^n_{it}, \tag{6}$$

where the subscript j refers to the job (i.e. the employer-employee or job-match). The corresponding equations in the case of earnings are:

$$\widetilde{y}^p_{ijt} = c^p_y u_{it} + d^p_y \theta_i + \mu^p_{ij} + \xi^p_{it}, \tag{7}$$

[6] As we explain in the data section, because of data limitations we only estimate our models for men. The covariates used to residualize the wage and earnings data are polynomials (cubic) in potential experience and tenure, years of completed schooling, and dummies for occupation, industry, race, marital status, collective bargaining, and calendar year.

and

$$\widetilde{y}_{ijt}^{n} = c_y^n u_{it} + d_y^n \theta_i + \mu_{ij}^n + \xi_{it}^n, \tag{8}$$

We estimate the model under the simplifying assumption that the idiosyncratic error terms ε_{it}^p and ε_{it}^n are uncorrelated over time. These error terms could either reflect measurement error, or purely transitory shocks. A simple approach for distinguishing the demand/productivity shocks u_{it} from the idiosyncratic errors is to assume some (limited) persistence in u_{it}. For instance, we consider the AR(1) model:

$$u_{it} = \rho u_{it-1} + \omega_{it}, \tag{9}$$

where ω_{it} is uncorrelated over time.

In some specifications, following the existing literature (e.g. Baker and Solon (2003)) we also allow the factor loadings d and the variance of the idiosyncratic errors to vary over time. This helps capture changes in returns to (unobserved) skills and related factors that have been linked to the secular growth in wage and earnings inequality.

Following Parent (2002), we estimate the variance components by fitting regression models to all the cross-products of residuals for the same individual.[7] This procedure is similar to the equally-weighted minimum distance approach of Abowd and Card (1989), but provides an easy way of dealing with an unbalanced sample like ours.

To see how the variance components models are identified, consider the expected value of the different cross-products of the wage residuals in the case where the factor loadings (the d's) do not change over time. For individuals on performance-pay jobs, the expected value of the squared residuals is

$$E(\widetilde{w}_{ijt}^p \cdot \widetilde{w}_{ijt}^p) = (c^p)^2\, var(u_{it}) + (d^p)^2 \cdot var(\theta_i) + var(\nu_{ij}^p) + var(\varepsilon_{ijt}^p).$$

The expected value of cross-products for two observations (at time t and time s) on the same job j is

$$E(\widetilde{w}_{ijt}^p \cdot \widetilde{w}_{ijs}^p) = \rho^{t-s}\,(c^p)^2\, var(u_{it}) + (d^p)^2 \cdot var(\theta_i) + var(\nu_{ij}^p),$$

and the expected value of cross-products for two observations on different jobs j and k is

$$E(\widetilde{w}_{ijt}^p \cdot \widetilde{w}_{ijs}^p) = \rho^{t-s}\,(c^p)^2\, var(u_{it}) + (d^p)^2 \cdot var(\theta_i)$$

[7] See Parent (1999) for a related analysis with the NLSY comparing piece-rate/commission workers and those receiving bonuses to salaried and hourly paid workers.

In this simple example, we can estimate the four variance components $(c^p)^2 \, var(u_{it})$, $(d^p)^2 var(\theta_i)$, $var(\nu^p_{ij})$ and $var(\varepsilon^p_{ijt})$ as well as the AR(1) parameter ρ by fitting the model using non-linear least squares. The same procedure can then be used to estimate the corresponding parameters for fixed-wage workers, and for the annual earnings models.

If the variance of u_{it} and θ_i were the same for performance-pay and fixed-wage workers, it would be straightforward to estimate all the c and d loading factors subject to a normalization such as $c^p = d^p = 1$. With $c^p = 1$ we can identify $var(u_{it})$ from the wage model for performance-pay workers. The other loading factors (c^n, c^p_y, c^p_y) can then be estimated as the ratio of the squared root of the variances.

The variance of u_{it} and θ_i being the same for performance-pay and fixed-wage workers is a strong assumption, however. Unless workers are randomly selected into performance-pay jobs, we expect the variance of unobserved ability, $var(\theta_i)$, to be different for the two types of jobs. We deal with the problem empirically by restricting the estimation to "switchers" who are observed on both types of jobs. This ensures that $var(\theta_i)$ is the same on the two types of jobs, and that difference in these variances is due to the loading factors as opposed to unobserved heterogeneity.

It is harder to deal with potential differences in the variance of shocks, $var(u_{it})$, in the two types of jobs since the choice of contract form may also depend on that parameter. In light of this, differences in the variance of shocks in the two types of jobs has to be interpreted with caution. That said, under the weaker assumption that wages and earnings for a given worker are subject to the same shock u_{it}, we can still identify the ratio of factor loadings c^p_y/c^p and c^n_y/c^n as the square root of the estimated variances. For instance, if most of the shocks in performance-pay jobs are absorbed in terms of lower (or higher) wages and hours are relatively unaffected, we will have $c^p_y/c^p \approx 1$. By contrast, if hours are the main margin of adjustment in fixed-wage jobs, we should find that $c^n_y/c^n >> 1$ since shocks will generate large variance in annual earnings (factor loading c^n_y), and little variance in hourly wages (factor loading c^n).

4 Data and motivating evidence

Our analysis is conducted using data from the PSID. The main advantage of the PSID is that it provides a representative sample of the workforce for a relatively long time period. One disadvantage of the PSID is that our constructed measures of performance pay are relatively crude for reasons discussed below.

4.1 The Panel Study of Income Dynamics (1976-2010)

The PSID sample we use consists of male heads of households aged 18 to 65 with average hourly earnings between \$4 and \$300 (in 1979 dollars) for the years 1976-2010, where the hourly wage rate is obtained by dividing total labor earnings from all jobs by total hours of work, both reported retrospectively for the previous calendar year.[8,9] Given our focus on performance pay, this wage measure based on total yearly earnings, inclusive of performance pay, is preferable to "point-in-time" wage measures that would likely miss infrequent payments (e.g. bonuses) of performance pay.

Individuals who are self-employed are excluded from the analysis since our measure of performance pay based on receiving bonuses, commissions, or piece-rates is defined for employed workers only.[10] We also exclude workers from the public sector. This leaves us with a total sample of 39,718 observations for 4,633 workers. Note that we only keep observations where individuals have worked a positive number of hours during the year, and record positive earnings. While this sample exhibit a large amount of variation in annual hours due to spells of unemployment, etc., we exclude the cases where people don't work at all during the year since wages are not observed in that case. In some specifications we also focus on a more "stable" subsample of workers who are employed at the time of the interview. The sample

[8] In the PSID, data on hours worked during year t, as well as on total labor earnings, bonuses/commissions/overtime income, and overtime hours, are asked in interview year t+1. Thus we actually use data covering interview years 1976-2011. Annual earnings were top coded at \$99,999 until 1982 (and not top coded since then), but only a handful of individuals were at the top code. We trim very high values of wages (above \$300.00 in 2011 dollars) but do not otherwise adjust for top coding.

[9] Our focus on male heads of households stems from the fact that only heads are asked about their income derived from bonuses, commissions, or overtime. In the PSID, males are designated as the head in all husband-wife pairs. The same is true if the female has a boyfriend with whom she has been living for at least a year, even if the female is the person with the most financial responsibility in the family unit. Consequently, the sample of female heads is relatively small. Perhaps more importantly, issues of representativeness would arise as those female heads are disproportionately nonwhite and are much less likely to be married.

[10] Self-employed workers can be viewed as being, by definition, paid for performance regardless of the mode of payment (earnings, dividends, etc.) they use to remunerate themselves.

size drops to 37,157 observations, meaning that in 2561 cases we have workers unemployed at the interview who, nonetheless, still end up working a positive number of hours during the year. All of the estimates reported in the paper are weighted using the PSID sample weights.

Identifying Performance Pay We construct a performance-pay indicator variable by looking at whether part of a worker's total compensation includes a variable pay component (bonus, commission, or piece-rate). For interview years 1976-1992, we are able to determine whether a worker received a bonus or a commission over the previous calendar year through the use of multiple questions. First, workers are asked the amount of money they received from working overtime, from commissions, or from bonuses paid by the employer.[11] Second, we sometimes know only whether or not workers worked overtime, and if they are working overtime in a given year, not the amount of pay they received for overtime. Thus, we classify workers as not having had a variable pay component if they worked overtime. Third, workers not paid exclusively by the hour, or not exclusively by a salary, are asked how they are paid: they can report being paid commissions, piece-rates, etc., as well as a combination of salaried/hourly pay along with piece-rates or commissions.[12] Through this combination of questions, we are thus able to identify *all* non-overtime workers who received performance pay in bonus, commission, or piece-rate form.

Starting with interview year 1993, there are separate questions about the amounts earned in bonuses, commissions, tips, and overtime for the previous calendar year. Thus, there is no need to back out an estimate of bonuses from an aggregate amount since the question is asked directly. For the sake of comparability with the pre-1993 years, we nevertheless classify as receiving no performance pay all workers who report any overtime work. In this way we

[11] Note that the question refers specifically to any amounts earned from bonuses, overtime, or commissions in addition to wages and salaries earned.

[12] In many survey years workers are not asked if their compensation package involves a mixture of salary/hourly pay and a variable component. All they are asked is how they are paid if not by the hour or with a salary. Although there is no way to directly verify it, this likely results in understating the incidence of any form of variable pay because workers are not allowed to answer that they are paid, say, a salary, and then report a commission: they have to choose. Our assertion that this response likely understates the extent of variable pay is motivated in part by the fact that workers in the NLSY are not restricted in describing the way they are paid. We find that workers in the NLSY are more likely to report having part of their compensation package contain a performance-pay component.

are able to determine whether a worker's total compensation included a performance-pay component for each year of the survey. One obvious drawback is that it is likely that the performance-pay component we construct will be noisy for hourly workers, though not for salaried workers who are not eligible for overtime payments. However, due to our treatment of overtime workers, we conservatively lean on the side of misclassifying workers as receiving no performance pay even when they do.[13]

Defining Performance-pay Jobs We define performance-pay jobs as employment relationships in which part of the worker's total compensation includes a variable pay component (bonus, a commission, piece-rate) at least once during the course of the relationship.[14] Since we use actual payments of bonuses, commissions or piece rates to identify performance-pay jobs, we are likely to misclassify performance-pay jobs as non-performance-pay jobs if some employment relationships are either terminated before performance pay is received, or partly unobserved for being out of our sample range. This source of measurement error is problematic because of an "end-point" problem in the PSID data. Given our definition of performance-pay jobs, we may mechanically understate the fraction of workers in such jobs at the beginning of our sample period because most employment relationships observed in 1976 started before 1976, and we do not observe whether or not performance pay was received prior to 1976. Similarly, jobs that started toward the end of the sample period may be performance-pay jobs but are classified otherwise because they have not lasted long enough for performance pay to be observed.

The problem is that, conditional on job duration, we tend to observe a given job match fewer times at the two ends of our sample period than in the middle of the sample. Consider, for example, the case of a job that lasts for five years. For jobs that last from 1985 to 1989, all five observations on this job match are captured in our PSID sample. For jobs that last

[13] In an earlier version of the paper, we re-did the analysis for 1992 to 1998 using the finer measure of performance pay that allows us to identify the performance-pay status of overtime workers. Doing so had little impact on the results. It only increased the fraction of workers on performance-pay jobs (for 1992-98) by one percentage point, and regression coefficients were essentially unchanged.

[14] We use "jobs", "employment relationship", and "job match" interchangeably. Although the PSID does have information on tenure in the position in most of the survey years spanning the sample period, we do not use it. As is well known, simply determining employer tenure in the PSID can be problematic (see Brown and Light (1992)). As a result, what we call a "job match" could be called an "employer match" instead. We generally use the word "job" for the sake of simplicity.

from 1973 to 1977, however, only two of the five years of the job match are observed, which mechanically reduces the probability of classifying the job as one with performance pay.

Because of this end-point problem, we get an unbalanced distribution of the number of times job matches are observed at different points of the sample period. One simple solution to the problem is to "rebalance" the sample using regression or other methods. In practice, we adjust measures of the incidence of performance pay over time by estimating a linear probability model in which dummies for calendar years and for the number of times the job-match is observed are included as regressors (estimating a logit gave almost identical results). We then compute an adjusted measure of the incidence of performance pay by holding the distribution of the number of times the job-match is observed to its average value for the years 1982 to 1990, which are relatively unaffected by the end-point problem.

The end-point problem could also affect the estimates of the effect of performance pay on both wage, hours, and earnings because the sample of non-performance-pay jobs is being contaminated by observations from performance-pay jobs for which performance-based payments are never observed. Lemieux, MacLeod, and Parent (2009) have investigated this issue in detail and concluded that, if anything, this measurement problem biases downward the estimated effect of performance pay. For the sake of clarity and simplicity, the wage samples we work with in the next sections are unadjusted for these measurement issues.

4.2 Descriptive Statistics from the PSID

Table 1 compares the mean characteristics of workers on performance-pay and non-performance-pay jobs, respectively. First, notice that 34 percent of the 39,718 observations are in performance-pay jobs. Workers on performance-pay jobs tend to earn more and have higher levels of education than workers on non-performance-pay jobs. Note that the hourly wage rate includes both regular wage and salary earnings and performance pay in the case of workers on performance-pay jobs. Annual hours worked and employer tenure also tend to be higher for workers on performance-pay than non-performance-pay jobs.

The unionization rate (percent covered by a collective bargaining agreement) is much lower among performance-pay workers. This suggests that, as expected, the pay structure in union firms corresponds more closely to the fixed-wage contracts discussed in Section 2. Another important difference is that there is a much higher fraction of workers paid by the hour in non-performance-pay than performance-pay jobs. Conversely, workers on performance-pay jobs are more likely to be salaried workers than those on non-performance-pay jobs.

An important point illustrated at the bottom of the table is that, of the 4633 workers, 1850 are observed on a performance-pay job, and 4113 are observed on a non-performance-pay job. So 1330 workers (1850+4113-4633) are "switchers" observed on both types of jobs. As discussed above, we present estimates of the variance components models for this subsample to control for potential differences in the unobserved ability of workers on the two types of jobs.

Figure 1 shows the incidence of performance pay over the 1976-2010 sample period. Note that we correct for the end-point problem using the procedure described above. Figure 1 shows that the overall incidence of performance-pay jobs has increased from about less then 40 percent in the late 1970s to around 45 percent in the 1990s, and has remained relateively constand since then. The figure also shows the simpler measure based on the fraction of workers actually reporting performance pay in a given year. This alternative measure clearly understates the incidence of performance-pay jobs since workers on performance-pay jobs will not necessarily receive a performance payment (like a bonus) in each year on the job. One advantage of this simple measure, however, is that it is not affected by the end-point problem and provides additional evidence of the robustness of the underlying trends in performance pay. Indeed, even this crude measure of performance pay clearly increases over time, especially in the 1980s.

Figure 1 also shows the fraction of workers covered by a collective bargaining agreement. Interestingly, the decline in unionization and the growth in performance pay are both concentrated in the same period (the 1980s). Figure 2 presents kernel density estimates of the distribution of annual hours for performance-pay and non-performance-pay jobs. The figure shows that annual hours have a higher mean and median, and are less evenly distributed among performance-pay than non-performance-pay jobs.

Note that performance pay represents a relatively modest share of total earnings (Figure 3). However, this does not mean that performance pay has a limited impact on total compensation since we expect the straight wage component to be more sensitive to workers' characteristics on performance-pay than non-performance-pay jobs. In order to pay for performance, the employer must evaluate the worker, which then affects the straight wage through promotions and job assignment. Hence, even though performance pay is a relatively small fraction of compensation for most workers, the fact that it exists is a signal of more careful monitoring.

4.3 Local unemployment rate as a measure of shocks

Before turning to the main empirical analysis based on the estimation of variance components models, we provide some motivating evidence using an observable measure of shocks. Following Lemieux, MacLeod, and Parent (2012), we use the local unemployment rate as a measure of shocks. A simple implication of the model in Section 2 is that the local unemployment rate should have a larger impact on wages in performance-pay than fixed-wage jobs, while the opposite should happen in the case of hours. It is not clear, however, how the impact on the two types of jobs should compare in the case of earnings. As discussed in Section 2, this depends on the elasticity of labor supply.

The results are reported in Appendix Tables 1 and 2.[15] All models include a large set of covariates that are not reported in the table: polynomials (cubic) in potential experience and tenure, years of completed schooling, and dummies for occupation, industry, race, marital status, collective bargaining, and calendar year. Standard errors are clustered at the county-year level.

In all regression models we show three sets of models for two different sample of workers. We start with simple OLS estimates, then move to models with worker-specific fixed effects that capture unobserved ability θ_i, and then report models with a full set of job match effects. Estimates with job-match effects are particularly credible as they solely rely on differential variation in the local unemployment rate, after controlling for year effects and job-match effects to identify the differential responsiveness to unemployment shocks in the different types of jobs.

The two sample of workers used are based on whether or not the worker is unemployed at the time of the interview. In our PSID sample, we only keep workers with at least some positive earnings and hours of work in the previous calendar year. These workers may or may not be working at the time of the interview. We first report results for the more "stable" sample of workers employed at the time of the interview in Appendix Table 1 (columns 1-3),

[15] For the sake of comparability with Lemieux, MacLeod, and Parent (2012), we use the same sample as theirs which stops in 1998. Note also that the local (county) unemployment rate is no longer available in recent years.

and then for the broader sample that also includes workers unemployed at the time of the interview (columns 4-6).

Panel A of Appendix Table 1 shows that, as expected, the unemployment rate has a negative and significant effect of wages in performance-pay jobs, but no significant effect on non-performance-pay jobs. The estimated coefficient for performance-pay jobs varies across specification but is generally close to -0.01, suggesting that a one percentage point increase in the unemployment rate is associated with a one percent decline in the hourly wage.

Panel B of Appendix Table 1 shows that precisely the opposite happens in the case of hours of work. The unemployment rate has a negative and significant impact on hours of work for workers not paid for performance, but an insignificant impact for performance-pay workers. The latter effect is consistent with a fairly inelastic labor supply elasticity for performance-pay workers. By contrast, since wages fail to adjust for non-performance-pay jobs, employers have little choice but to cut back on hours and employment in the presence of adverse productivity shocks.

The results for hours are in levels. They are not directly comparable to those for wages (in logs). Since average yearly hours is about 2000, the -10 estimate reported in Panel B corresponds to a 0.5 percent decline in hours. This is fairly similar in terms of magnitude to the estimated effect on the wages of performance-pay workers (Panel A of Appendix Table 1). It suggests that the total effect of the unemployment rate on earnings (wages time hours) should be roughly comparable for performance-pay and non-performance-pay jobs. The only difference is that the adjustment happens along the wage margin for performance-pay workers, but along the hours margins for non-performance workers.

This conjecture is confirmed in Panel C of Appendix Table 1, which shows the estimated effect of the local unemployment rate on the log of annual earnings. In our preferred specification with job-match fixed effects, the effect of the unemployment rate on annual earnings is equal to about -.008 for both performance-pay and non-performance pay jobs. This means that a one percentage point increase in the local unemployment rate reduces earnings by close to one percent in both sectors. The difference is that hourly wages account for essentially all the earnings adjustment in performance-pay jobs, while hours account for the bulk of the adjustment in non-performance-pay jobs.

Appendix Table 2 presents similar estimates except that we now divide jobs both in terms of performance-pay and union status. As discussed in Section 2, we expect wages to be most responsive to shocks in performance-pay jobs that are not unionized, and least responsive to shocks in unionized non-performance-pay jobs. We also expect the exact opposite to happen for hours of work. The two other types of contractual arrangements (union/performance-pay and non-union/non-performance pay) should fall somewhere in between these two extreme cases.

Looking once again at our preferred specification (column 3), we see that the results are consistent with these expectations. Panel A of Appendix Table 2 shows that the unemployment rate has the largest impact on non-union performance-pay jobs (-0.0076) and the smallest (and not statistically significant) impact on union non-performance-pay workers (-0.0001). By contrast, exactly the opposite happens in the case of hours of work (Panel B). As a result, the overall impact on annual earnings is more or less similar for all four types of contractual arrangements (Panel C).

One potential concern with these results is that some of the differential responsiveness to shocks under differential contractual arrangements is due to composition effects. For example, performance-pay workers tend to be more concentrated in occupations such as managers and professionals (see Lemieux, MacLeod, and Parent (2009)) that may be less sensitive to the business cycle than blue collar occupations. One simple way of checking for this is to rebalance the performance-pay and non-performance pay samples so that they have the same distribution of observed characteristics. We did so using a reweighting procedure and this did not substantially changed the results.

5 Main Results

Before turning to the estimation of the variance components model, we show the raw autocovariance structure for wages and earnings in performance-pay and non-performance-pay jobs in Appendix Table 3 (employed workers at the time of the interview) and 4 (all workers). We also show in Figure 4 that, consistent with the well documented increase in inequality, the variances of wages and earning increase substantially over time. The autocovariances are averaged over all sample years to simply the exposition.

Comparing columns 1 and 2 of Appendix Table 3, it is clear that going from wages to annual earnings has a relateively modest impact on the autocovariances in performance-pay jobs. By contrast, there is a large difference between the autocovariances of wages and earnings in non-performance-pay jobs (columns 3 and 4). This is consistent with our finding in Appendix Tables 1-2 that there is much more variation in hours in response to shocks in non-performance-pay jobs. The difference also diminishes quickly as we increase the order of the autocovariances, suggesting that most of the variation in hours is transitory. As a result, the autocovariance in earnings in performance-pay jobs is larger than in non-performance-pay jobs for all autocovariances except the zero order autocovariance (the variance).

The estimates of the variance components models are reported in Tables 2-7 for various samples and specifications. At this stage we focus on estimates from simple models where shocks u_{it} are assumed to be purely transitory ($\rho = 0$). In that context, we cannot separately estimate the variance of the u_{it} and ε_{it} terms (or ξ_{it} for earnings), and just report the sum of the two variances in the tables. This corresponds to a conventional variance components model with a pure transitory and a pure permanent variance component, though we also include a job-match component in some specifications. Table 6 and 7 go one step further by reporting results where u_{it} is modeled as a MA(3) process. Preliminary estimates indicate that estimates of ρ for the various specifications are tightly clustered around 0.8-0.85. While allowing for $\rho \neq 0$ greatly improves the fit of the model, it does not change the qualitative conclusions presented below on the effect of shocks on wages and earnings under the different types of contracts.

In Tables 2-5, we first report estimates from a simple model where the factor loadings d_t^p and d_t^n are assumed to be fixed over time, and the job-match component is set to zero. We then add the job-match component in a second specification, and free up the factor loadings (return to unobserved ability) and the variance of ε_{ijt}^p and ε_{ijt}^n to reflect the well-known growth in inequality over time.

In all tables, we first report (Panel A) results estimated over the whole sample. As we discussed earlier, one potential pitfall of using the whole sample is that some individuals are only observed on performance-pay jobs, while others are only observed on non-performance-pay jobs. To control for these composition effects, we report in Panel B the results for the subsample of "switchers" who are observed on both performance-pay and non-performance-pay jobs.

Table 2 shows the results of the wage decomposition for the subsample of more "stable" workers employed at the time of the survey (this table is an updated version of Table 5 in Lemieux, MacLeod, and Parent (2009)). Consistent with the pattern observed in the empirical autocovariances (Appendix Table 3), the results confirm that the permanent component of wages (variance of θ_i) is substantially larger in performance-pay than non-performance-pay jobs.

For performance-pay jobs, the variance decomposition of earnings reported in Table 3 is fairly similar to the decomposition for wages in Table 2. Both the permanent and transitory variances (variance of u_{it} and ε^p_{ijt}) are slightly larger than in the wage models. By contrast, the transitory variance is almost twice as large for earnings than wages in non-performance-pay jobs, while the permanent variance is only slightly larger. This is consistent with the pattern of results documented in the raw data reported in Appendix Table 3. The transitory variance becomes even larger when workers unemployed at the time of the interview are also included in the sample in Table 5. Even in that case, however, the variance linked to the permanent wage component θ_i remains larger in performance-pay than non-performance-pay jobs.

Tables 6 and 7 present a variety of robustness checks. All the specifications now include a MA(3) in u_{it}. We also allow the transitory variance to freely vary over time in columns 2-3 and 4-5, and introduce a factor loading in the permanent component in columns 3 and 6.

As in the other tables, the variance of the permanent component is systematically higher in performance-pay than non-performance-pay jobs. The results are very robust across the disfferent specifications. Qualitatively speaking, the results for the MA components are broably similar to those based on the models with only a pure transitory component. To simplify the exposition, we report in Tables 6-7 the fitted covariance of u_{it} obtained using a MA(3) model. For instance, in column 1 of Table 6 (Panel A) we see that the fitted autocovariance goes from 0.030 for the first order autocovariance to 0.017 for the third order autocovariance. Fitting those with an AR(1) process instead (not reported in the tables) yields an estimate of ρ of about 0.8.

The main result in Tables 6 and 7 is that while the variance of u_{it} is systematically higher in non-performance-pay jobs (especially for earnings), this source of wage and earnings variability dies off over time because of limited persistence in u_{it}.

5.1 Implications for welfare inequality

Looking at earnings instead of just hourly wages first suggests that inequality is actually smaller in performance-pay than non-performance-pay jobs, contrary to what Lemieux, MacLeod, and Parent (2009) found for hourly wages only. This result only holds, however, for purely cross-sectional measures of inequality. Since the variation in hours is mostly transitory, from a welfare point of view inequality likely remains larger in performance-pay than non-performance-pay jobs. The fact that hours of work are quite volatile and respond more to shocks in non-performance-pay jobs has a substantial impact on earnings in the short run, but little impact on long-run measures of both the level and the inequality of earnings.

6 Conclusion

In this paper, we use data from the Panel Study of Income Dynamics to study how the aucovariance structure of wages and earnings differ under different contractual arrangements. We divide jobs on the basis of whether they pay for performance, and whether they are covered by collective bargaining agreements. Using the county unemployment rate as a proxy for local labor market shocks, we find that wages and hours of work respond very differently to shocks depending on contractual arrangements. Wages are most flexible under non-union performance-pay contracts, and least flexible under non-performance-pay union contracts. Precisely the opposite happens in the case of hours of work that are the least sensitive to shocks under non-union performance-pay contracts, and the most sensitive under union non-performance-pay contracts.

We then perform a full-fledge variance components analysis. Looking at earnings instead of just hourly wages suggests that inequality is actually smaller in performance-pay than non-performance-pay jobs. This result only holds, however, for purely cross-sectional measures of inequality. Since the variation in hours is mostly transitory, from a welfare point of view inequality likely remains larger in performance-pay than non-performance-pay jobs. The fact that hours of work are quite volatile and respond more to shocks in non-performance-pay jobs has a substantial impact on earnings in the short run, but little impact on long-run measures of both the level and the inequality of earnings.

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Figure 1. Performance Pay Job Incidence
PSID 1976–2011

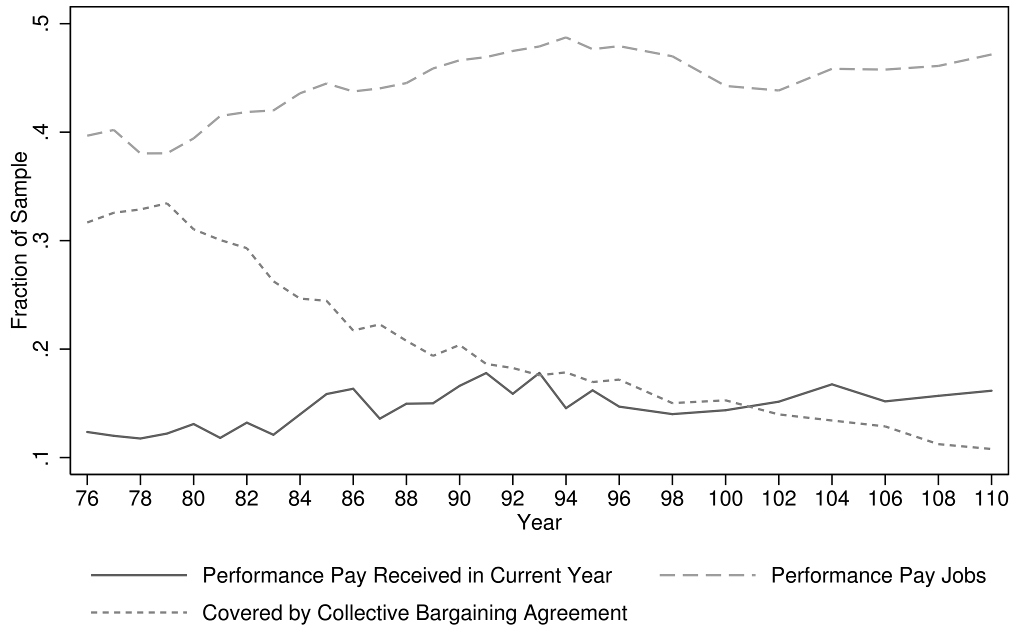


Figure 2. Distribution of Hours Worked
PSID 1976–2011

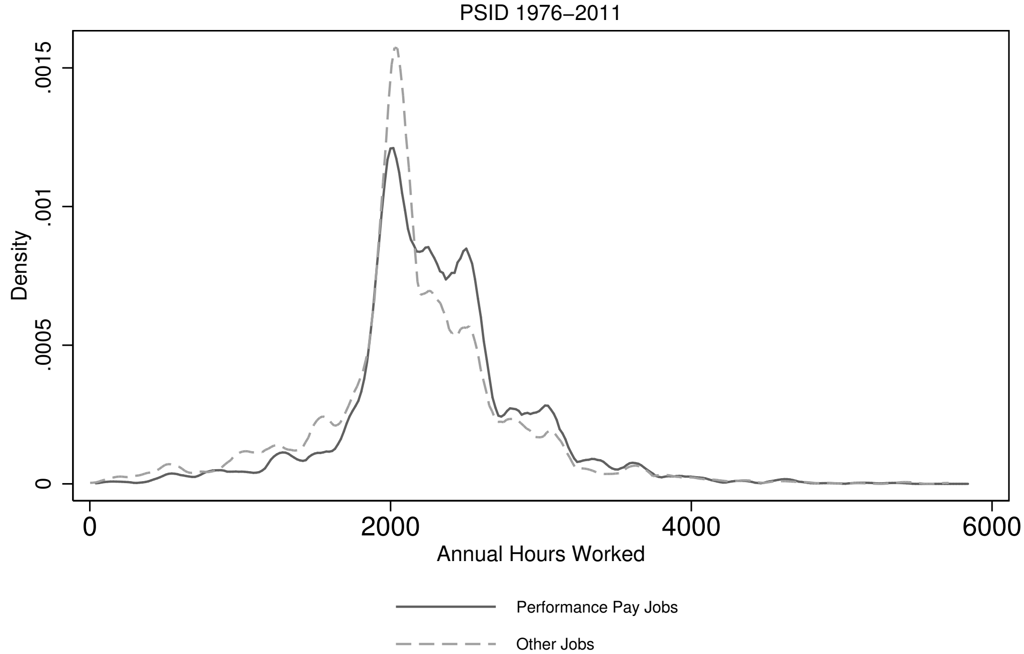


Figure 3

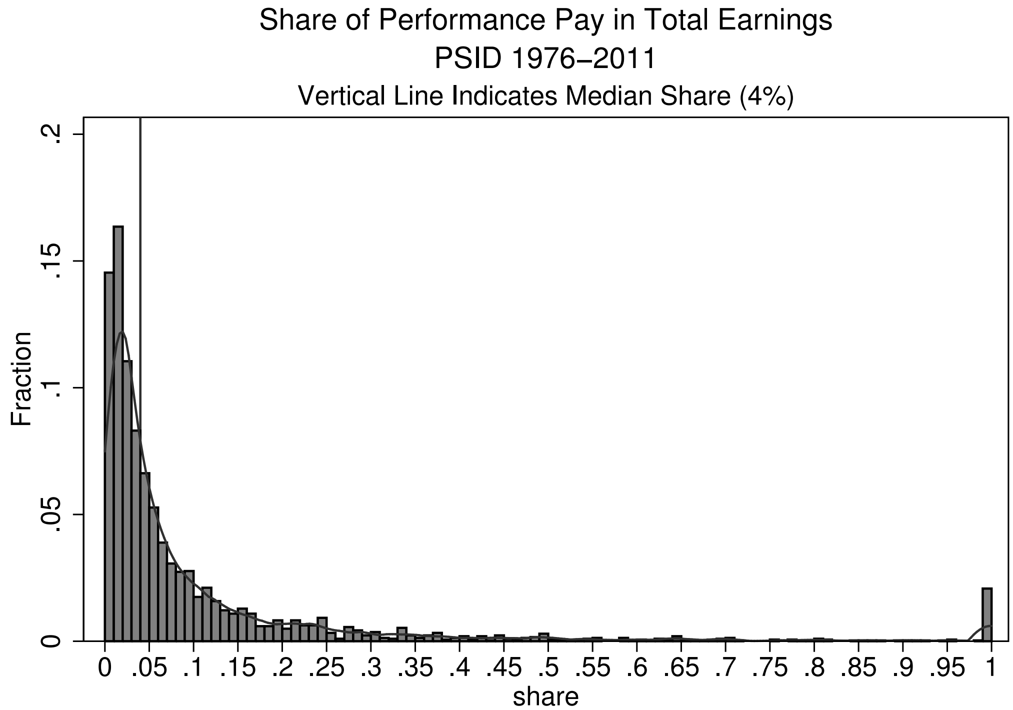
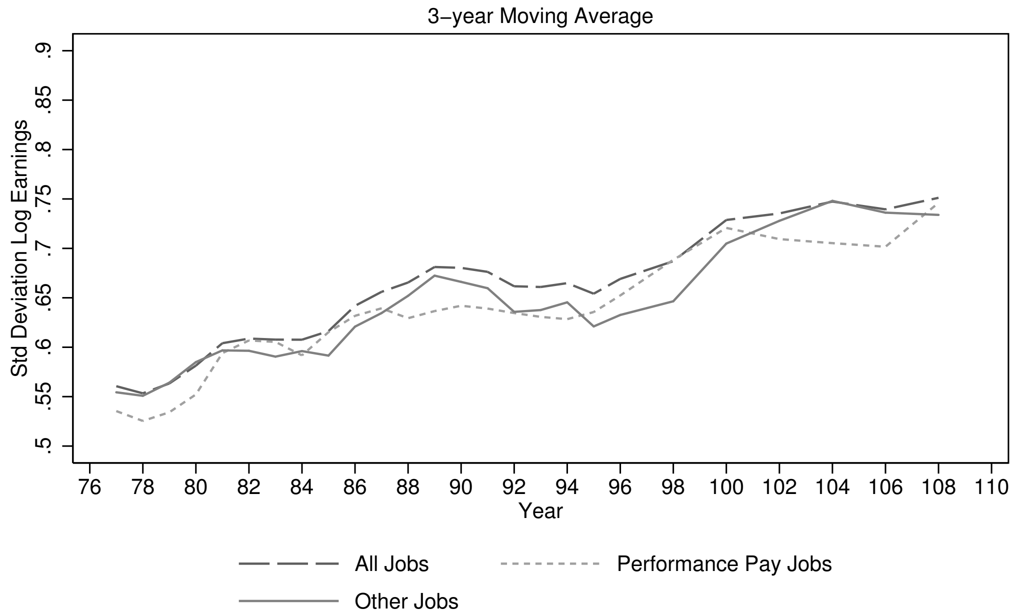

Share of Performance Pay in Total Earnings
PSID 1976−2011
Vertical Line Indicates Median Share (4%)
Fraction
0
.05
.1
.15
.2
0 .05 .1 .15 .2 .25 .3 .35 .4 .45 .5 .55 .6 .65 .7 .75 .8 .85 .9 .95 1
share

Figure 4. Total Log Earnings Inequality

Panel A: Sample Includes Only Employed Workers



Panel B: Sample Includes Employed and Unemployed Workers

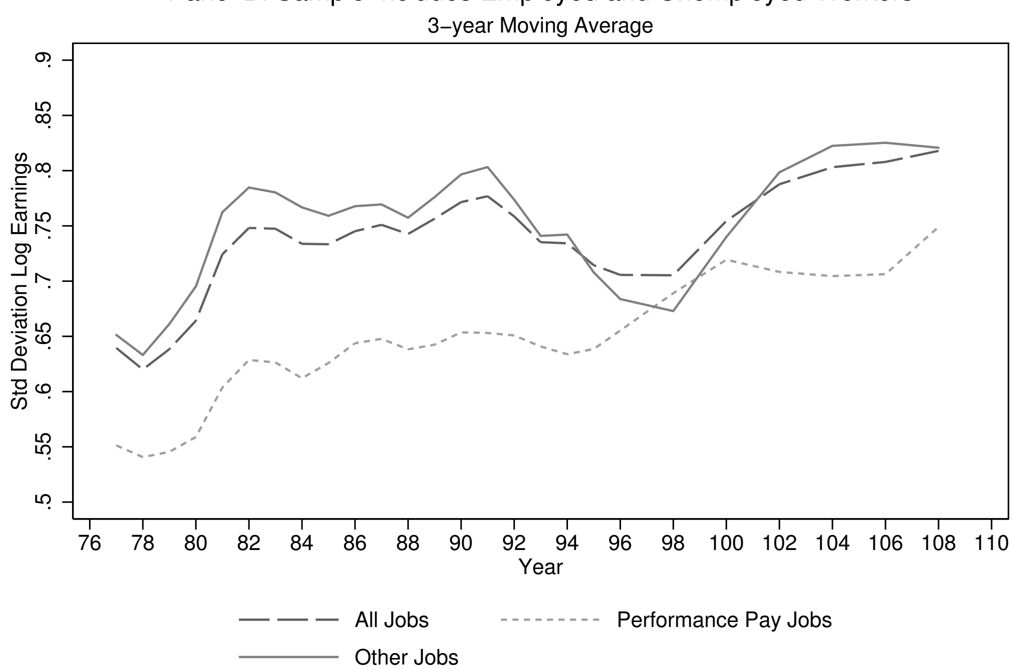


Table 1 Summary Statistics: Panel Study of Income Dynamics 1976-2011

	Non-Performance Pay Jobs	Performance Pay Jobs
Average Hourly Earnings ($2011)	22.91	30.41
Education	12.55	13.25
Potential Experience	24.05	23.83
Tenure	6.94	9.80
Married	0.59	0.64
Non White	0.26	0.23
Paid by the Hour	55.3	37.0
Paid a Salary	20.1	36.3
Fraction Unemployed at Interview	0.082	0.016
Annual Hours Worked	2101.6	2286.9
# Workers (Tot: 4633)	4113	1850
# Job Matches (Tot: 13670)	10937	2733
# Observations (Tot: 39718)	26360	13358

Notes: The sample consists of male household heads aged 18-64 working in private sector, wage and salary jobs. All figures in the table represent sample means. Education and potential experience are measured in years. Potential experience is defined as age minus education minus 6. Temporarily laid off workers are included among the employed. The number of workers in either type of jobs is larger than the total number of workers because a subset of individuals are employed in both types of jobs.

Table 2
Error Component Models of Hourly Wages by Type of Job: Employed Workers at Time of Interview

Panel A: full sample

	Performance-pay jobs			Non-performance-pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.112	0.106	0.074	0.080	0.068	0.057
worker component	(0.001)	(0.002)	(0.004)	(0.001)	(0.001)	(0.002)
Factor loading: 1990-94	-	-	1.254	-	-	1.116
relative to 1976-79			(0.034)			(0.026)
Factor loading: 2006-2010	-	-	1.138	-	-	1.005
relative to 1976-79			(0.045)			(0.036)
Variance of job-	-	0.015	0.011	-	0.021	0.019
match component		(0.003)	(0.003)		(0.001)	(0.001)
Variance of	0.097	0.094	0.102	0.112	0.103	0.087
idiosyncratic error	(0.003)	(0.003)	(0.008)	(0.002)	(0.002)	(0.004)
Change in variance,	-	-	-0.009	-	-	0.038
1976-79 to 1990-93			(0.010)			(0.006)
Change in variance,	-	-	0.019	-	-	0.056
1976-79 to 2006-2010			(0.013)			(0.007)
Number of workers	1 778	1 778	1 778	3 890	3 890	3 890
Number of cross-products	87 337	87 337	87 337	141 067	141 067	141 067

Panel B: workers who worked in both types of jobs

	Performance pay jobs			Non performance pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.111	0.101	0.052	0.079	0.065	0.051
worker component	(0.001)	(0.003)	(0.005)	(0.002)	(0.002)	(0.005)
Factor loading: 1990-93	-	-	1.453	-	-	1.147
relative to 1976-79			(0.071)			(0.072)
Factor loading: 2006-2010	-	-	1.554	-	-	1.139
relative to 1976-79			(0.092)			(0.090)
Variance of job-	-	0.014	0.011	-	0.034	0.033
match component		(0.003)	(0.003)		(0.003)	(0.003)
Variance of	0.104	0.100	0.130	0.134	0.114	0.096
idiosyncratic error	(0.004)	(0.004)	(0.011)	(0.004)	(0.004)	(0.010)
Change in variance,	-	-	-0.036	-	-	0.036
1976-79 to 1990-93			(0.014)			(0.013)
Change in variance,	-	-	-0.005	-	-	0.049
1976-79 to 2006-2010			(0.019)			(0.016)
Number of workers	1 196	1 196	1 196	1 196	1 196	1 196
Number of cross-products	50 882	50 882	50 882	31 825	31 825	31 825

Note: Standard errors in parentheses. Models in columns 3 and 6 allow the variance of the idiosyncratic errors and the factor loadings on the worker component to vary across the 1976-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1998, 2000-2004, and 2006-2010 periods, while models in columns 1, 2, 4, and 5 do not. These equally weighted covariance structure models are fit to the cross-products of the residuals of an OLS regression of log wages on the same set of covariates described in Table 3. Note that the factor loadings in columns 3 and 6 are normalized to 1 in the base period (1976-1979), so that the changes in factor loadings can be interpreted as the percentage changes in the return to the worker component.

Table 3
Error Component Models of Annual Earnings by Type of Job: Employed Workers at Time of Interview

Panel A: full sample

	Performance-pay jobs			Non-performance-pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.121	0.109	0.074	0.105	0.096	0.088
worker component	(0.001)	(0.003)	(0.004)	(0.001)	(0.002)	(0.004)
Factor loading: 1990-94	-	-	1.287	-	-	1.066
relative to 1976-79			(0.039)			(0.032)
Factor loading: 2006-2010	-	-	1.399	-	-	0.960
relative to 1976-79			(0.054)			(0.044)
Variance of job-	-	0.015	0.013	-	0.015	0.012
match component		(0.003)	(0.003)		(0.003)	(0.003)
Variance of	0.113	0.110	0.107	0.187	0.181	0.145
idiosyncratic error	(0.003)	(0.003)	(0.009)	(0.003)	(0.003)	(0.008)
Change in variance,	-	-	-0.010	-	-	0.064
1976-79 to 1990-93			(0.012)			(0.011)
Change in variance,	-	-	0.025	-	-	0.130
1976-79 to 2006-2010			(0.016)			(0.014)
Number of workers	1 778	1 778	1 778	3 890	3 890	3 890
Number of cross-products	87 337	87 337	87 337	141 067	141 067	141 067

Panel B: workers who worked in both types of jobs

	Performance pay jobs			Non performance pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.116	0.100	0.055	0.095	0.084	0.072
worker component	(0.002)	(0.003)	(0.006)	(0.003)	(0.003)	(0.008)
Factor loading: 1990-93	-	-	1.356	-	-	1.113
relative to 1976-79			(0.077)			(0.080)
Factor loading: 2006-2010	-	-	1.919	-	-	1.017
relative to 1976-79			(0.120)			(0.101)
Variance of job-	-	0.021	0.022	-	0.027	0.024
match component		(0.004)	(0.004)		(0.005)	(0.005)
Variance of	0.123	0.118	0.128	0.198	0.182	0.123
idiosyncratic error	(0.004)	(0.004)	(0.013)	(0.006)	(0.007)	(0.016)
Change in variance,	-	-	-0.026	-	-	0.048
1976-79 to 1990-93			(0.016)			(0.021)
Change in variance,	-	-	-0.013	-	-	0.143
1976-79 to 2006-2010			(0.023)			(0.025)
Number of workers	1 196	1 196	1 196	1 196	1 196	1 196
Number of cross-products	50 882	50 882	50 882	31 825	31 825	31 825

Note: Standard errors in parentheses. Models in columns 3 and 6 allow the variance of the idiosyncratic errors and the factor loadings on the worker component to vary across the 1976-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1998, 2000-2004, and 2006-2010 periods, while models in columns 1, 2, 4, and 5 do not. These equally weighted covariance structure models are fit to the cross-products of the residuals of an OLS regression of log wages on the same set of covariates described in Table 3. Note that the factor loadings in columns 3 and 6 are normalized to 1 in the base period (1976-1979), so that the changes in factor loadings can be interpreted as the percentage changes in the return to the worker component.

Table 4
Error Component Models of Hourly Wages by Type of Job: Employed and Unemployed Workers at Time of Interview

Panel A: full sample

	Performance-pay jobs			Non-performance-pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.111	0.100	0.074	0.082	0.072	0.059
worker component	(0.001)	(0.002)	(0.004)	(0.001)	(0.001)	(0.002)
Factor loading: 1990-94	-	-	1.251	-	-	1.131
relative to 1976-79			(0.033)			(0.026)
Factor loading: 2006-2010	-	-	1.134	-	-	0.980
relative to 1976-79			(0.044)			(0.033)
Variance of job-	-	0.014	0.010	-	0.018	0.017
match component		(0.003)	(0.003)		(0.001)	(0.001)
Variance of	0.099	0.097	0.104	0.125	0.116	0.098
idiosyncratic error	(0.003)	(0.003)	(0.007)	(0.002)	(0.002)	(0.004)
Change in variance,	-	-	-0.009	-	-	0.045
1976-79 to 1990-93			(0.010)			(0.006)
Change in variance,	-	-	0.020	-	-	0.058
1976-79 to 2006-2010			(0.013)			(0.007)
Number of workers	1 837	1 837	1 837	4 070	4 070	4 070
Number of cross-products	89 672	89 672	89 672	161 306	161 306	161 306

Panel B: workers who worked in both types of jobs

	Performance pay jobs			Non performance pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.113	0.101	0.059	0.073	0.061	0.047
worker component	(0.001)	(0.003)	(0.005)	(0.001)	(0.002)	(0.005)
Factor loading: 1990-93	-	-	1.367	-	-	1.091
relative to 1976-79			(0.059)			(0.068)
Factor loading: 2006-2010	-	-	1.369	-	-	1.072
relative to 1976-79			(0.075)			(0.086)
Variance of job-	-	0.016	0.012	-	0.034	0.033
match component		(0.003)	(0.003)		(0.003)	(0.003)
Variance of	0.104	0.100	0.118	0.157	0.134	0.110
idiosyncratic error	(0.003)	(0.004)	(0.011)	(0.003)	(0.004)	(0.009)
Change in variance,	-	-	-0.025	-	-	0.058
1976-79 to 1990-93			(0.013)			(0.013)
Change in variance,	-	-	0.015	-	-	0.059
1976-79 to 2006-2010			(0.018)			(0.015)
Number of workers	1 314	1 314	1 314	1 314	1 314	1 314
Number of cross-products	55 758	55 758	55 758	41 663	41 663	41 663

Note: Standard errors in parentheses. Models in columns 3 and 6 allow the variance of the idiosyncratic errors and the factor loadings on the worker component to vary across the 1976-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1998, 2000-2004, and 2006-2010 periods, while models in columns 1, 2, 4, and 5 do not. These equally weighted covariance structure models are fit to the cross-products of the residuals of an OLS regression of log wages on the same set of covariates described in Table 3. Note that the factor loadings in columns 3 and 6 are normalized to 1 in the base period (1976-1979), so that the changes in factor loadings can be interpreted as the percentage changes in the return to the worker component.

Table 5
Error Component Models of Annual Earnings by Type of Job: Employed and Unemployed Workers at Time of Interview

Panel A: full sample

	Performance-pay jobs			Non-performance-pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.120	0.109	0.074	0.119	0.122	0.102
worker component	(0.001)	(0.003)	(0.004)	(0.002)	(0.002)	(0.005)
Factor loading: 1990-94	-	-	1.292	-	-	1.121
relative to 1976-79			(0.040)			(0.035)
Factor loading: 2006-2010	-	-	1.390	-	-	0.962
relative to 1976-79			(0.055)			(0.048)
Variance of job-	-	0.014	0.012	-	-0.006	-0.010
match component		(0.003)	(0.003)		(0.003)	(0.003)
Variance of	0.122	0.119	0.114	0.265	0.268	0.210
idiosyncratic error	(0.003)	(0.003)	(0.009)	(0.004)	(0.004)	(0.010)
Change in variance,	-	-	-0.011	-	-	0.108
1976-79 to 1990-93			(0.012)			(0.015)
Change in variance,	-	-	0.022	-	-	0.154
1976-79 to 2006-2010			(0.016)			(0.018)
Number of workers	1 837	1 837	1 837	4 070	4 070	4 070
Number of cross-products	89 672	89 672	89 672	161 306	161 306	161 306

Panel B: workers who worked in both types of jobs

	Performance pay jobs			Non performance pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.120	0.104	0.060	0.108	0.104	0.067
worker component	(0.002)	(0.003)	(0.006)	(0.004)	(0.005)	(0.010)
Factor loading: 1990-93	-	-	1.349	-	-	1.294
relative to 1976-79			(0.070)			(0.119)
Factor loading: 2006-2010	-	-	1.724	-	-	0.874
relative to 1976-79			(0.101)			(0.135)
Variance of job-	-	0.021	0.020	-	0.011	0.006
match component		(0.004)	(0.004)		(0.008)	(0.008)
Variance of	0.129	0.124	0.128	0.308	0.301	0.233
idiosyncratic error	(0.004)	(0.004)	(0.013)	(0.009)	(0.010)	(0.023)
Change in variance,	-	-	-0.025	-	-	0.090
1976-79 to 1990-93			(0.016)			(0.031)
Change in variance,	-	-	-0.002	-	-	0.188
1976-79 to 2006-2010			(0.023)			(0.036)
Number of workers	1 314	1 314	1 314	1 314	1 314	1 314
Number of cross-products	55 758	55 758	55 758	41 663	41 663	41 663

Note: Standard errors in parentheses. Models in columns 3 and 6 allow the variance of the idiosyncratic errors and the factor loadings on the worker component to vary across the 1976-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1998, 2000-2004, and 2006-2010 periods, while models in columns 1, 2, 4, and 5 do not. These equally weighted covariance structure models are fit to the cross-products of the residuals of an OLS regression of log wages on the same set of covariates described in Table 3. Note that the factor loadings in columns 3 and 6 are normalized to 1 in the base period (1976-1979), so that the changes in factor loadings can be interpreted as the percentage changes in the return to the worker component.

Table 6
Error Component Models of Hourly Wages by Type of Job: Additional Results

Panel A: Employed (at time of interview) workers who worked in both types of jobs

	Performance-pay jobs			Non-performance-pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.098	0.098	0.086	0.056	0.058	0.053
worker component	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)
Linear trend of	-	-	0.005	-	-	0.003
worker component (X 100)			(0.001)			(0.001)
Variance of job-	0.006	0.006	0.008	0.016	0.016	0.020
match component	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)
MA(1) loading on	0.030	0.029	0.024	0.046	0.046	0.040
transitory component	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)
MA(2) loading on	0.025	0.025	0.021	0.036	0.036	0.031
transitory component	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)
MA(3) loading on	0.017	0.017	0.014	0.025	0.025	0.021
transitory component	(0.005)	(0.005)	(0.005)	(0.006)	(0.006)	(0.003)
Single transitory comp.	Yes	No	No	Yes	No	No
for all years						
Unrestricted yearly	No	Yes	Yes	No	Yes	Yes
transitory component						
Number of workers	1 196	1 196	1 196	1 196	1 196	1 196
Number of cross-products	50 882	50 882	50 882	31 825	31 825	31 825

Panel B: Employed and unemployed workers who worked in both types of jobs

	Performance pay jobs			Non performance pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.098	0.098	0.090	0.053	0.053	0.050
worker component	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)
Linear trend of	-	-	0.004	-	-	0.001
worker component (X 100)			(0.001)			(0.001)
Variance of job-	0.008	0.008	0.010	0.016	0.016	0.021
match component	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
MA(1) loading on	0.028	0.027	0.023	0.048	0.048	0.043
transitory component	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)
MA(2) loading on	0.024	0.024	0.020	0.039	0.039	0.035
transitory component	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)
MA(3) loading on	0.017	0.017	0.014	0.028	0.028	0.025
transitory component	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Single transitory comp.	Yes	No	No	Yes	No	No
for all years						
Unrestricted yearly	No	Yes	Yes	No	Yes	Yes
Number of workers	1 314	1 314	1 314	1 314	1 314	1 314
Number of cross-products	55 758	55 758	55 758	41 663	41 663	41 663

Note: Standard errors in parentheses. Models in columns 3 and 6 allow the variance of the idiosyncratic errors and the factor loadings on the worker component to vary across the 1976-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1998, 2000-2004, and 2006-2010 periods, while models in columns 1, 2, 4, and 5 do not. These equally weighted covariance structure models are fit to the cross-products of the residuals of an OLS regression of log wages on the same set of covariates described in Table 3. Note that the factor loadings in columns 3 and 6 are normalized to 1 in the base period (1976-1979), so that the changes in factor loadings can be interpreted as the percentage changes in the return to the worker component.

Table 7
Error Component Models of Annual Earnings by Type of Job: Additional Results

Panel A: Employed (at time of interview) workers who worked in both types of jobs

	Performance-pay jobs			Non-performance-pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.096	0.096	0.078	0.074	0.074	0.073
worker component	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.004)
Linear trend of	-	-	0.009	-	-	0.001
worker component (X 100)			(0.001)			(0.001)
Variance of job-	0.011	0.011	0.013	0.001	0.001	0.013
match component	(0.004)	(0.004)	(0.004)	(0.006)	(0.006)	(0.004)
MA(1) loading on	0.043	0.043	0.036	0.070	0.070	0.061
transitory component	(0.005)	(0.005)	(0.005)	(0.008)	(0.008)	(0.008)
MA(2) loading on	0.031	0.031	0.025	0.053	0.053	0.046
transitory component	(0.005)	(0.005)	(0.005)	(0.009)	(0.008)	(0.008)
MA(3) loading on	0.020	0.020	0.015	0.036	0.036	0.030
transitory component	(0.005)	(0.005)	(0.004)	(0.009)	(0.009)	(0.009)
Single transitory comp.	Yes	No	No	Yes	No	No
for all years						
Unrestricted yearly	No	Yes	Yes	No	Yes	Yes
transitory component						
Number of workers	1 196	1 196	1 196	1 196	1 196	1 196
Number of cross-products	50 882	50 882	50 882	31 825	31 825	31 825

Panel B: Employed and unemployed workers who worked in both types of jobs

	Performance pay jobs			Non performance pay jobs		
Parameter	[1]	[2]	[3]	[4]	[5]	[6]
Variance of	0.101	0.101	0.087	0.086	0.086	0.093
worker component	(0.003)	(0.003)	(0.004)	(0.005)	(0.005)	(0.006)
Linear trend of	-	-	0.006	-	-	-0.003
worker component (X 100)			(0.001)			(0.002)
Variance of job-	0.011	0.011	0.013	-0.027	-0.027	-0.014
match component	(0.004)	(0.004)	(0.004)	(0.009)	(0.009)	(0.008)
MA(1) loading on	0.040	0.040	0.035	0.108	0.108	0.099
transitory component	(0.005)	(0.005)	(0.005)	(0.011)	(0.011)	(0.011)
MA(2) loading on	0.030	0.029	0.024	0.081	0.080	0.073
transitory component	(0.005)	(0.005)	(0.005)	(0.012)	(0.012)	(0.012)
MA(3) loading on	0.020	0.020	0.015	0.053	0.053	0.047
transitory component	(0.006)	(0.006)	(0.006)	(0.013)	(0.013)	(0.013)
Single transitory comp.	Yes	No	No	Yes	No	No
for all years						
Unrestricted yearly	No	Yes	Yes	No	Yes	Yes
Number of workers	1 314	1 314	1 314	1 314	1 314	1 314
Number of cross-products	55 758	55 758	55 758	41 663	41 663	41 663

Note: Standard errors in parentheses. Models in columns 3 and 6 allow the variance of the idiosyncratic errors and the factor loadings on the worker component to vary across the 1976-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1998, 2000-2004, and 2006-2010 periods, while models in columns 1, 2, 4, and 5 do not. These equally weighted covariance structure models are fit to the cross-products of the residuals of an OLS regression of log wages on the same set of covariates described in Table 3. Note that the factor loadings in columns 3 and 6 are normalized to 1 in the base period (1976-1979), so that the changes in factor loadings can be interpreted as the percentage changes in the return to the worker component.

Appendix Table 1. The Effect of Local Labor Market Conditions: PSID, 1976-1998

Panel A: Log Hourly Earnings

Sample:	Employed Individuals			All Employed and Unemployed Individuals		
	[1]	[2]	[3]	[4]	[5]	[6]
Variable	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer
Unemployment Rate in County X Non Performance Pay Job	-0.0030	-0.0066	-0.0021	-0.0027	-0.0075	-0.0023
	(0.0037)	(0.0020)	(0.0020)	(0.0035)	(0.0021)	(0.0019)
Unemp. Rate in County X Performance Pay Job	-0.0113	-0.0148	-0.0072	-0.0097	-0.0153	-0.0082
	(0.0050)	(0.0026)	(0.0029)	(0.0049)	(0.0026)	(0.0028)
P-Value of Test of Equality	0.0793	0.0101	0.1225	0.1165	0.0196	0.0714
Unemployed at Interview	-	-	-	-0.1110	-0.0400	-0.0447
				(0.0213)	(0.0180)	(0.0310)
Number of Observations	26146	26146	26146	27899	27899	27899

Panel B: Annual Hours Worked

Sample:	Employed Individuals			All Employed and Unemployed Individuals		
Variable	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer
Unemployment Rate in County X Non Performance Pay Job	-6.04	-10.76	-10.73	-6.79	-11.59	-9.80
	(3.18)	(2.94)	(3.19)	(3.84)	(3.45)	(3.89)
Unemp. Rate in County X Performance Pay Job	-0.11	1.68	1.22	0.47	1.43	1.41
	(3.92)	(3.52)	(3.92)	(3.14)	(2.84)	(2.96)
P-Value of Test of Equality	0.2148	0.0052	0.0236	0.1141	0.0028	0.0258
Unemployed at Interview	-	-	-	-773.72	-654.51	-617.25
				(27.62)	(28.34)	(47.15)
Number of Observations	26146	26146	26146	27899	27899	27899

Panel C: Log Annual Earnings

Sample:	Employed Individuals			All Employed and Unemployed Individuals		
Variable	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer
Unemployment Rate in County X Non Performance Pay Job	-0.0070	-0.0165	-0.0076	-0.0072	-0.0110	-0.0064
	(0.0038)	(0.0023)	(0.0021)	(0.0039)	(0.0027)	(0.0020)
Unemp. Rate in County X Performance Pay Job	-0.0116	-0.0125	-0.0078	-0.0096	-0.0163	-0.0085
	(0.0050)	(0.0027)	(0.0029)	(0.0051)	(0.0027)	(0.0028)
P-Value of Test of Equality	0.3582	0.2321	0.9527	0.6200	0.3694	0.5233
Unemployed at Interview	-	-	-	-0.7627	-0.5904	-0.5514
				(0.0336)	(0.0310)	(0.0489)
Number of Observations	26146	26146	26146	27899	27899	27899

Notes. Estimates come from unrestricted regressions in which all covariates are interacted with the performance pay job dummy. Other covariates include polynomials (cubic) in potential experience and tenure, years of completed schooling, and dummies for occupation, industry, race, marital status, collective bargaining, calendar year, and having part of current year's earnings based on performance pay. Standard errors are clustered at the county X year level.

Appendix Table 2. The Effect of Local Labor Market Conditions: PSID, 1976-1998; All Sectors

Panel A: Log Hourly Earnings

Sample:	Employed Individuals			All Employed and Unemployed Individuals		
	[1]	[2]	[3]	[4]	[5]	[6]
Variable	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer
U. Rate X Unionized Workers	0.0162	0.0052	-0.0001	0.0175	0.0047	-0.0001
X Non Performance Pay Job	(0.0034)	(0.0021)	(0.0021)	(0.0034)	(0.0022)	(0.0021)
U. Rate X Non-Unionized Workers	-0.0124	-0.0122	-0.0036	-0.0117	-0.0128	-0.0036
X Non Performance Pay Job	(0.0038)	(0.0021)	(0.0021)	(0.0037)	(0.0022)	(0.0021)
U. Rate X Unionized Workers	0.0095	-0.0101	-0.0040	0.0100	-0.0115	-0.0060
X Performance Pay Job	(0.0071)	(0.0040)	(0.0041)	(0.0064)	(0.0048)	(0.0039)
U. Rate X Non-Unionized Workers	-0.0145	-0.0154	-0.0076	-0.0138	-0.0160	-0.0084
X Performance Pay Job	(0.0053)	(0.0027)	(0.0033)	(0.0053)	(0.0027)	(0.0032)
Number of Observations	26146	26146	26146	27899	27899	27899

Panel B: Annual Hours Worked

Sample:	Employed Individuals			All Employed and Unemployed Individuals		
Variable	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer
U. Rate X Unionized Workers	-13.32	-14.63	-13.90	-14.71	-15.48	-13.31
X Non Performance Pay Job	(3.36)	(3.59)	(3.68)	(3.32)	(3.13)	(3.33)
U. Rate X Non-Unionized Workers	-2.43	-8.62	-7.84	-3.36	-10.09	-6.91
X Non Performance Pay Job	(3.54)	(3.13)	(3.53)	(3.51)	(3.06)	(3.42)
U. Rate X Unionized Workers	3.19	-2.64	2.77	5.65	0.19	1.64
X Performance Pay Job	(10.99)	(7.63)	(6.32)	(9.91)	(6.94)	(5.85)
U. Rate X Non-Unionized Workers	-0.14	1.96	1.27	0.05	1.97	0.78
X Performance Pay Job	(3.84)	(3.83)	(4.28)	(3.84)	(3.75)	(4.27)
Number of Observations	26146	26146	26146	27899	27899	27899

Panel C: Log Annual Earnings

Sample:	Employed Individuals			All Employed and Unemployed Individuals		
Variable	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer	OLS	Fixed-Effects Within Worker	Fixed-Effects Within Employer
U. Rate X Unionized Workers X Non Performance Pay Job	0.0094 (0.0035)	-0.0021 (0.0022)	-0.0070 (0.0023)	0.0104 (0.0035)	-0.0016 (0.0026)	-0.0055 (0.0022)
U. Rate X Non-Unionized Workers X Non Performance Pay Job	-0.0150 (0.0040)	-0.0172 (0.0026)	-0.0079 (0.0023)	-0.0151 (0.0041)	-0.0180 (0.0029)	-0.0067 (0.0022)
U. Rate X Unionized Workers X Performance Pay Job	0.0106 (0.0086)	-0.0124 (0.0049)	-0.0065 (0.0038)	0.0125 (0.0077)	-0.0131 (0.0050)	-0.0060 (0.0033)
U. Rate X Non-Unionized Workers X Performance Pay Job	-0.0149 (0.0053)	-0.0171 (0.0030)	-0.0077 (0.0032)	-0.0138 (0.0054)	-0.0170 (0.0030)	-0.0090 (0.0032)
Number of Observations	26146	26146	26146	27899	27899	27899

Notes. Estimates come from unrestricted regressions in which all covariates are interacted with the performance pay job dummy. Other covariates include polynomials (cubic) in potential experience and tenure, years of completed schooling, and dummies for occupation, industry, race, marital status, collective bargaining, calendar year, and having part of current year's earnings based on performance pay. Standard errors are clustered at the county X year level.

Appendix Table 3: Autocovariances in Wages and Earnings
(workers employed at the time of the interview)

	Performance-pay		Non-Performance-Pay	
Order of autocov:	[1] Wage	[2] Earnings	[3] Wage	[4] Earnings
0	0.218	0.244	0.215	0.295
1	0.131	0.150	0.118	0.145
2	0.126	0.137	0.105	0.127
3	0.118	0.127	0.092	0.111
4	0.115	0.125	0.088	0.104
5	0.113	0.117	0.076	0.093
6	0.109	0.113	0.075	0.081
7	0.108	0.111	0.070	0.082
8	0.105	0.107	0.060	0.069
9	0.104	0.106	0.068	0.082
10	0.096	0.100	0.050	0.062
11	0.094	0.097	0.055	0.071
12	0.091	0.095	0.039	0.048
13	0.084	0.084	0.036	0.058
14	0.082	0.086	0.026	0.038
15	0.076	0.080	0.037	0.044
16	0.078	0.086	0.043	0.033
17	0.078	0.078	0.042	0.036
18	0.076	0.087	0.048	0.047
19	0.090	0.107	0.025	0.032
20	0.094	0.127	0.040	0.027
22	0.085	0.118	0.033	0.043
23	0.097	0.132	0.039	0.031
24	0.074	0.123	0.056	0.056
25	0.119	0.172	0.073	0.067
26	0.100	0.062	0.066	0.043
27	0.152	0.077	0.045	0.112
28	0.191	0.046	0.081	0.124

Appendix Table 4: Autocovariances in Wages and Earnings
(workers employed and unemployed at the time of the interview)

	Performance-pay		Non-Performance-Pay	
Order of autocov:	[1] Wage	[2] Earnings	[3] Wage	[4] Earnings
0	0.219	0.253	0.232	0.416
1	0.132	0.151	0.113	0.173
2	0.127	0.139	0.101	0.149
3	0.119	0.130	0.089	0.125
4	0.116	0.127	0.084	0.111
5	0.114	0.120	0.068	0.102
6	0.111	0.118	0.068	0.089
7	0.110	0.117	0.065	0.091
8	0.108	0.113	0.053	0.074
9	0.106	0.108	0.061	0.090
10	0.101	0.106	0.044	0.072
11	0.097	0.104	0.049	0.078
12	0.093	0.100	0.034	0.052
13	0.085	0.088	0.030	0.062
14	0.087	0.093	0.022	0.025
15	0.082	0.087	0.036	0.031
16	0.083	0.094	0.038	0.029
17	0.084	0.090	0.032	0.030
18	0.080	0.093	0.038	0.061
19	0.088	0.103	0.014	0.035
20	0.087	0.116	0.027	0.020
22	0.082	0.109	0.026	0.039
23	0.094	0.118	0.036	0.030
24	0.061	0.102	0.048	0.047
25	0.095	0.139	0.060	0.051
26	0.063	0.036	0.051	0.041
27	0.097	0.052	0.056	0.120
28	0.146	0.044	0.087	0.164