Shareholder Wealth and Wages: Evidence for White-Collar Workers

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We present empirical evidence on the relationship between individual wages, conditional on worker characteristics, and equity returns using a unique survey from the Bureau of Labor Statistics. Equity returns affect the wages only of workers with three or more years of tenure. A 4 percent increase in a firm's market value raises pay by 0.3 percent within three years. Our estimates suggest that each \$10 increase in shareholder wealth raises the present value of a firm's wage bill by \$1. The elasticity of white-collar wages with respect to equity returns is one-third smaller than the CEO salary elasticities in our sample.

I. Introduction

How much higher are wages in more profitable companies, ceteris paribus? Although the standard competitive model predicts that wages depend on a worker's skills and not on the employer's financial performance, implicit contract, incentive contract, principal-agent, and bargaining models suggest that profit sharing is an important part of the employment relationship. Labor market frictions and positively sloped labor supply schedules may also cause a positive short-run correlation between wages and profits. Few studies have analyzed the relationship between individual pay and a firm's profits because of data limitations. In this paper we present some of the first empirical evidence on the relationship between a worker's wage and firm profitability, con-

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ditional on the worker's demographic characteristics, for white-collar workers in the United States.

Several empirical studies focus on the relationship between interindustry wage differentials and industry profitability. Dickens and Katz (1987), Krueger and Summers (1988), and Katz and Summers (1989) find that interindustry wage differentials are positively related to differences in average profits across industries.[1] Allen (1995) finds that the effect of industry profits on industry average wages is sensitive to the choice of profit measure used. Blanchflower, Oswald, and Sanfey (1996), using individual wage data from the Current Population Survey and industry average profits in U.S. manufacturing, find that wages increase about three years after an increase in profits. These studies are limited by their inability to control for job tenure, which is omitted from these regressions and may be correlated with profitability. In addition, industry average profits ignore the substantial within-industry variation in firm performance.

A few papers examine the relationship between union wage bargains and quasi rents or profits in the United States and Canada. Abowd and Lemieux (1993) find a positive relationship between wages and quasi rents per worker in Canadian labor contracts. Christofides and Oswald (1992) find that union wages are positively related to lagged industry profits in Canada. Currie and McConnell (1992) find that collectively bargained wages are negatively related to profits per employee in the United States. Data limitations precluded these authors from controlling for skill differences across wage bargains, but this omitted variable problem is crucial in interpreting their results. Higher union wages in more profitable firms may reflect more skilled workers rather than higher pay for workers of a given ability. In addition, these results may not generalize to nonunion workers, who constitute the vast majority of the private-sector workforce in the United States.

Several studies estimate the relationship between employer average wages and employer-specific measures of profits or productivity in Great Britain. Nickell and Wadhwani (1990), Nickell, Vainiomaki, and Wadhwani (1994), and Hildreth and Oswald (1997) use panel data on British manufacturing firms and find a positive relationship between a firm's average wage and price, profit, and productivity variables. These studies are limited by their inability to control for interemployer differences in workers' skills. More successful firms may pay higher wages because they employ more skilled workers, not because their wages are higher for a given quality of worker. In Nickell et al. (1994), the positive coefficients on output price and Solow residuals in wage regressions are reduced

[1] Holmlund and Zetterberg (1991) and Nickell and Kong (1992) find that "insider variables" such as revenue per worker are positively related to wage differentials.

by two-thirds when a firm's average wages are disaggregated by skill group. Thus much of the positive coefficient on price and productivity variables is probably due to interfirm skill differences. Blanchflower, Oswald, and Garrett (1990) use a cross-section British data set with average wages at an employer disaggregated by skill group and a self-reported categorical measure of firm performance. Firms that report that their financial performance is "a lot better than average" pay significantly higher wages.

Abowd, Kramarz, and Margolis (1999) estimate employer wage and wage growth differentials, conditional on workers' experience and job tenure, using a large panel data set of French workers and firms. This comprehensive data set allows the authors to measure the wage-profitability relationship holding constant observed and unobserved worker and firm characteristics. They find that enterprises that hire high-wage workers are more productive, but not more profitable, and enterprises that pay higher wages, when they control for person effects, are more productive, profitable, and capital intensive. Unfortunately, this important work may not generalize to the United States, where labor markets are considerably less regulated and wages are much less likely to be collectively bargained.

The studies cited above use a variety of profit measures including accounting profits per worker, the rate of return on capital (accounting profits/capital), quasi rents (profits adjusted by alternative opportunity costs for capital and labor) per worker, Tobin's q, and Solow residuals. Each of these profitability variables is likely to be measured with error and correlated with variables omitted from the wage regression. Accounting profits understate the opportunity cost of capital and are therefore strongly positively correlated with the capital intensity of the firm or industry. A positive correlation between wages and accounting profits can result from the complementarity of capital and skilled labor and omitted skill variables. This omitted variable bias is exacerbated in many studies in which the capital intensity of the firm or industry has been omitted from the wage regression. Solow residuals are also likely to be positively correlated with unobserved skills. An employer that generates more revenues for a given number of workers and book value of capital, and has a larger Solow residual, is likely to use capital with a higher market value and employ more skilled workers, who appear more expensive ceteris paribus.

There are numerous studies of executive compensation (e.g., Jensen and Murphy 1990; Rosen 1992; Garen 1994) that use equity returns as a measure of firm profitability. The most convincing studies use panel data to difference out time-invariant individual and firm characteristics and estimate the effect of changes in equity values on changes in compensation to chief executive officers (CEOs). Although these studies

are important for understanding the executive labor market, they may provide little information about the relationship between wages and profits for the typical worker, who has substantially less wealth and less influence on firm performance than the CEO.

Our estimates of the relationship between wages and profitability have a number of advantages and provide some important contrasts to the empirical studies cited above. In contrast to nearly all the previous work using firm-specific profits, we use individual worker wages rather than a firm's average wage and estimate employer wage differentials conditional on a worker's education, job tenure, age, race, and sex. We also estimate the effect of profitability on wages after differencing out unobserved job match–specific effects by utilizing starting pay information that is available for a fraction of our sample. The wage regressions we estimate include equity returns, rather than accounting profits, as the key profitability variable. Equity returns represent unanticipated innovations in profits and are weakly correlated with other firm characteristics. Firm-specific equity returns also vary substantially over time and within industries and cause fewer aggregation problems than the industry average profit variables used in many studies.[2] The typical member of our sample is a nonunion white-collar worker. Finally, we compare the equity return elasticity for white-collar wages to the equity return elasticity for CEO salaries, using the same empirical methodology, over the same time period, in the same sample of firms.

II. Data

A. *Worker Characteristics*

The data set used in this study is derived from the Bureau of Labor Statistics' White Collar Pay (WCP) survey, which measures private-sector wages in white-collar occupations that match those in the federal government.[3] The WCP collects the straight-time salary and detailed occupation of full-time workers in a nationwide sample of private-sector employers. Our data set is based on a supplement to the WCP conducted in 1989 and 1990. In this test survey, 354 establishments were asked to report current and starting pay, age, race, sex, education, and tenure for a random sample of their employees in "matched" white-collar occupations. Three hundred establishments provided complete informa-

[2] An aggregation problem remains because not all of a firm's establishments are equally profitable and establishment-specific equity shares are not traded.

[3] Occupations in the WCP include accountants, personnel specialists and supervisors, attorneys, buyers, computer programmers and systems analysts, chemists, engineers, nurses, medical and engineering technicians, drafters, computer and key entry operators, photographers, clerks, messengers, secretaries, and typists.

TABLE 1
DISTRIBUTIONS OF WORKERS BY INDUSTRY AND OCCUPATION

	WCP Workers in Traded Firms (1)	WCP Workers in Nontraded Firms (2)	White-Collar Workers in CPS Outgoing Rotation Groups (3)
		Industry	
Nondurable manufacturing	.252	.168	.073
Durable manufacturing	.558	.440	.145
Mining and construction	.102	.074	.028
Services	.037	.235	.428
Other	.052	.083	.326
		Occupation	
Professional	.354	.286	.274
Administrative	.310	.254	.112
Technical	.172	.168	.215
Clerical	.164	.293	.399
Number of workers	807	949	16,424

tion on current pay, tenure, and demographic characteristics for 1,756 workers between the ages of 18 and 64.[4]

We matched 109 establishments and 807 workers in the WCP to 92 different publicly traded parent corporations in the Compustat database using the names and addresses of establishments in the WCP. There are 34 firms with fewer than five workers, 27 firms with five to nine workers, 23 firms with 10–19 workers, and eight firms with 20 or more workers. The median worker in a Compustat firm has 14 coworkers in the sample. There are 949 workers in the 191 establishments that we were unable to match to Compustat corporations.

We first assess whether our WCP sample is representative of the population of white-collar workers in the United States. Table 1 compares the distribution of workers by occupation and industry in our sample of publicly traded firms in the WCP to two comparison groups. The first group includes workers in the WCP pilot survey who are not employed in publicly traded firms, and the second includes private-sector workers in WCP occupations in the 1989 outgoing rotation group samples of the Current Population Survey (CPS). The WCP occupations account for 39 percent of all white-collar workers in the CPS. The publicly traded WCP sample is disproportionately employed in professional

[4] See Bronars and Famulari (1997) for a more complete description of the full data set.

and administrative occupations and manufacturing industries. The WCP sample is concentrated in manufacturing industries because the pilot survey was primarily conducted in 1990, when goods-producing industries were surveyed in the WCP.

Column 1 of table 2 reports sample means and standard deviations of worker characteristics for the publicly traded WCP sample. To facilitate comparison across samples, we use sample weights that restrict the weighted distribution of workers in the comparison groups to match the distribution of workers in publicly traded firms across industry/occupation cells. Columns 2 and 3 present weighted sample statistics for WCP workers at non–publicly traded employers and for CPS workers. Columns 4 and 5 present mean differences across the publicly traded and comparison samples. The WCP workers in publicly traded firms earn 8.5 percent more, have longer job tenure, and are more likely to be in large establishments than WCP workers in nontraded firms. Workers in traded firms earn 16.2 percent more, are more educated and experienced, and are more likely to be male than CPS workers in similar industries and occupations. Standard decompositions from log wage regressions reveal that differences in worker characteristics account for half of the 16.2 percent pay gap between workers in traded firms and CPS workers and one-third of the 8.5 percent pay differential between WCP workers in traded and nontraded firms.[5] Finally, although the WCP does not report union status, few workers in this study are likely to be unionized because only 5.1 percent of workers in the CPS comparison sample are covered by a collectively bargained contract.

B. *Firm Characteristics*

The Compustat database annually reports a firm's capital expenditures, net operating income, sales, employment, and the book value of its capital stock (plant and equipment) net of depreciation. The log of annual sales and log of the ratio of the book value of capital to employment, averaged over the two years preceding the WCP survey, measure a firm's size and capital intensity. The book value of a firm's capital stock, net of depreciation, reported by Compustat reflects historical, nominal costs and accounting measures of depreciation. We also imputed values for a firm's capital stock for the 80 companies in our sample

[5] The regression pooling WCP workers in traded and nontraded firms includes polynomials in education, experience, and tenure; female interactions with these polynomials; race dummy variables; and establishment fixed effects. The regression pooling CPS and WCP data excludes tenure variables and establishment fixed effects because they are unavailable for the CPS sample.

TABLE 2
SAMPLE MEANS AND STANDARD DEVIATIONS

	WCP				
	Traded (1)	Nontraded[a] (2)	CPS[a] (3)	DIFFERENCE (1)−(2) (4)	DIFFERENCE (1)−(3) (5)
Monthly real wage	2,255.6	2,086.6	1,974.7	169.0**	280.9**
	(944.5)	(930.0)	(982.3)	(44.9)	(35.4)
Log(real wage)	7.632	7.546	7.471	.085**	.162**
	(.432)	(.444)	(.502)	(.021)	(.018)
Tenure	10.104	8.474		1.631**	
	(8.794)	(7.913)		(.399)	
Education	14.657	14.573	14.376	.084	.287**
	(2.212)	(2.329)	(2.130)	(.109)	(.077)
Experience	18.634	18.330	16.325	.304	2.310**
	(10.213)	(10.571)	(10.948)	(.498)	(.394)
Female	.400	.429	.447	−.028	−.047**
				(.024)	(.018)
Black	.056	.046	.057	.010	−.001
				(.010)	(.008)
Other	.063	.095	.053	−.032**	.011
				(.013)	(.008)
Covered by union contract			.051		
Northeast	.233	.246	.239	−.013	−.007
				(.020)	(.015)
Midwest	.349	.255	.226	.094**	.124**
				(.022)	(.015)
South	.342	.300	.291	.042*	.051**
				(.022)	(.016)
West	.076	.199	.243	−.123**	−.168**
				(.016)	(.015)
In a metropolitan area	.779	.813	.846	−.034*	−.066**
				(.019)	(.013)
Establishment size:					
Under 500 employees	.249	.552		−.303**	
				(.022)	
500–999 employees	.145	.219		−.074**	
				(.019)	
1,000 or more employees	.606	.229		.377**	
				(.022)	
Workers	807	949	16,424		
Establishments	109	191			

NOTE.—All workers typically work 35–40 hours per week, and nominal wages are converted to 1983 dollars using the employer cost index for private-sector wages and salaries. In cols. 4 and 5, the numbers in parentheses are standard errors.
[a] Observations are weighted so that the distribution of workers across major industry/occupation cells is identical to the distribution of WCP workers in publicly traded firms.
* Significant at the .10 level.
** Significant at the .05 level.

that reported capital investment from 1972 to 1989.[6] Although the mean value of our imputed capital stock is about 2.5 times as large as the book value of capital, the correlation between these variables is .98 in 1989. We use the book value of capital because it has fewer missing values, and our empirical results are insensitive to the capital measure used.

We control for interfirm differences in growth by calculating employment growth rates over several time periods prior to the WCP survey. Wages may be correlated with lagged employment growth because demand shifts cause positive comovements between wages and employment due to positively sloped labor supply schedules. Alternatively, the composition of a firm's workforce may vary with expansions and contractions of employment, so that lagged employment growth rates are correlated with unobserved skills.

Operating income per employee measures a firm's accounting profitability and probably reflects quasi rents on sunk capital investments as well as economic profits. We use corporate income per worker rather than a rate of return on capital because it is difficult to measure the market value of capital and wages may be a function of quasi rents per worker. The within-firm serial correlation in operating income per employee is so high that we are unable to distinguish between alternative lag structures for the relationship between wages and accounting profitability.[7]

Stock returns from the Center for Research in Security Prices are available for 86 firms in the sample, and data for the entire six-year period preceding the WCP survey are available for 78 firms. Annual rates of return on equity and dollar changes in shareholder wealth per worker are calculated for the periods one to two years, three to four years, and five to six years prior to the WCP survey.

Table 3 presents sample statistics for firm characteristics. Columns 1 and 4 report the number of workers and firms for which each characteristic is observed. Columns 2 and 3 present weighted means and standard deviations in which each firm's weight is its number of WCP workers. Columns 5 and 6 report unweighted firm means and standard deviations. The mean firm has 36,250 workers, $4.85 billion in sales, $51,900 of capital equipment per worker, and operating income of

[6] We impute a current value of capital installed prior to 1972 by deflating the nominal book value of net plant and equipment in 1972 by the consumer price index (CPI) and depreciating by 10 percent per year. We impute the value of net additions to capital by summing nominal investment in capital equipment from 1972 to 1989, deflating by the CPI, and depreciating by 10 percent per year. Our imputed value of capital stock is the sum of these two measures.

[7] For example, the correlation between operating income per worker at lags of one to two and three to four years is .979, and the correlation between equity returns at lags of one to two and three to four years is .053.

TABLE 3
FIRM CHARACTERISTICS

	Weighted Sample Statistics			Unweighted Sample Statistics		
	Number of Workers (1)	Mean (2)	Standard Deviation (3)	Number of Firms (4)	Mean (5)	Standard Deviation (6)
Employment (thousands)[a]	784	59.94	67.14	88	36.25	47.43
Capital (billions of dollars)[a]	789	4.533	10.136	89	1.944	5.487
Sales (billions of dollars)[a]	807	9.477	15.548	92	4.847	9.072
Capital/labor (thousands of dollars per worker)[a]	777	72.22	128.0	87	51.93	94.34
Log(sales)[a]	807	8.077	1.670	92	7.404	1.623
Log(capital/labor)[a]	777	3.469	1.077	87	3.233	1.075
Real annual equity return:						
Lagged 1–2 years	733	.080	.244	82	.064	.246
Lagged 3–4 years	721	.088	.227	80	.083	.209
Lagged 5–6 years	711	.112	.189	78	.116	.235
Lagged 3–6 years	711	.100	.155	78	.100	.170
ΔShareholder wealth/worker (hundreds of thousands of dollars/worker) lagged 3–6 years	675	.554	1.016	72	.459	.870
Annual operating income/worker (hundreds of thousands of dollars/worker) lagged 3–6 years	751	.217	.279	83	.173	.214
Annual employment growth:						
Lagged 1–2 years	782	.028	.110	87	.018	.151
Lagged 3–4 years	767	.028	.169	85	.035	.177
Lagged 5–6 years	733	.031	.116	80	.062	.160
Lagged 1–6 years	733	.026	.096	80	.034	.123

NOTE.—All dollar values are deflated to constant 1983 dollars using the CPI.
[a] Mean value over the two years prior to the WCP survey.

$17,300 per employee. Below we find the strongest relationship between wages and equity returns lagged three to six years. The mean firm had a real average annual equity return of 10 percent over the period three to six years prior to the WCP survey, which generated increases in shareholder wealth of $45,900 per worker over the same four-year period.

III. Wages and Equity Returns

A. *Empirical Specification*

Our primary goal is to measure the wage differential across more and less profitable firms, holding constant as many other worker and firm characteristics as possible. Consider the individual log wage regression[8]

$$\ln W_{ij} = \mathbf{X}_{ij}\beta + \mathbf{Z}_{ij}\gamma + \mathbf{F}_j\alpha + \epsilon_{ij}, \qquad (1)$$

where W_{ij} is the current straight-time wage of worker i at employer j, and $\mathbf{X}_{ij}$ is a vector of worker characteristics that includes polynomials in education, experience, and tenure; interactions between these polynomials and a female dummy variable; and race dummy variables. The term $\mathbf{Z}_{ij}$ is a vector of dummy variables for major industry and dummy variables including female interactions, and $\mathbf{F}_j$ is a vector of employer attributes such as equity returns (or other profit variables), firm size, growth rates, and capital intensity.[9] The error term, ϵ_{ij}, is expected to be correlated across workers at the same employer. Ordinary least squares (OLS) regressions that ignore the within-group correlation in ϵ_{ij} report standard errors that are likely to be biased downward. Employer fixed effects control for this correlation but eliminate the possibility of estimating the coefficients on firm-specific factors ($\mathbf{F}_j$).

We use random effects to account for the within-group correlation in ϵ_{ij}. Our approach allows for the following error structure:

$$\epsilon_{ij} = \delta_j^F G_{ij} + \delta_j^M(1 - G_{ij}) + u_{ij},$$

where $G_{ij} = 1$ if worker i at employer j is female, δ_j^F and δ_j^M are the random effects for females and males at employer j, and u_{ij} is an independently and identically distributed error. This specification assumes that δ_j^F and δ_j^M are uncorrelated with the independent variables in the

[8] The regression in (1) is equivalent to running separate male and female wage regressions and pooling the coefficients on region dummy variables and firm attributes **F**. In results not reported here, we allowed for separate coefficients on these variables by sex and failed to reject the hypothesis of equal coefficients by sex.

[9] Although some of the firm characteristics are not reported for a few of the corporations in our sample, we replace missing values of each characteristic with zero and generate a dummy variable that equals one if the characteristic is missing. We use the entire sample of workers and establishments in our regressions by including as explanatory variables both firm characteristics and the corresponding missing dummy variables.

regression, and below we show that specification tests generally fail to reject this hypothesis.[10] In results not reported here, we also estimate OLS wage regressions with standard errors that are corrected for the within-group correlation in ϵ_{ij}. In general, the coefficients on profitability are quite similar in the OLS and random-effects specifications.

B. *Using Equity Returns as a Measure of Profitability*

Column 1 of table 4 reports selected coefficients from random-effects regressions of equation (1) that include equity returns lagged one to two, three to four, and five to six years. The coefficient on returns lagged three to four years is positive and significant at the .05 level. Equity returns lagged five to six years have a similar impact on wages, but this effect is imprecisely measured; returns lagged one to two years have a small and insignificant impact on wages. The result that wages adjust to profits with a three-year lag confirms the findings of Blanchflower, Oswald, and Sanfey (1996), who used CPS wages and industry average profits.

More rapidly growing firms pay lower wages. The coefficient on employment growth lagged five to six years is significantly negative, and employment growth lagged one to two years has a similar, but statistically insignificant, effect on wages. The pattern of coefficients on lagged equity returns and employment growth rates suggests that a more parsimonious specification of the model that imposes restrictions on these coefficients will yield more precise parameter estimates.

In column 2 we include average equity returns over the period three to six years prior to the WCP survey.[11] This specification restricts the coefficients on equity returns to be equal across lags of three to four and five to six years. A one-percentage-point increase in a firm's annual return over this four-year period, which causes the firm's market value of equity to rise by about 4 percent, is associated with a 0.25 percent increase in white-collar wages. Our estimate is similar in magnitude to the wage-profit elasticity of .08 reported in Blanchflower, Oswald, and Sanfey (1996), which predicts that a 4 percent increase in profits per worker causes a 0.32 percent increase in wages.

The specification in column 2 includes a firm's average employment growth rate over the six-year period prior to the survey, which restricts

[10] The specification test we use is described in detail in Hausman (1978). The test compares the estimated β coefficients across random-effects and fixed-effects models. Under the null hypothesis of no correlation between the δ's and the **X**'s, both estimators are consistent but the random-effects estimator is more efficient.

[11] We also included the average equity return over the entire six-year period prior to the survey and estimated slightly larger coefficients on equity returns than those reported here.

TABLE 4
EFFECT OF EQUITY RETURNS ON CURRENT PAY (Sample Size 807)

	(1)	(2)	(3)	(4)
Annual equity returns:				
Lagged 2 years	.0490			
	(.0675)			
Lagged 3–4 years	.1450**			
	(.0640)			
Lagged 5–6 years	.1342			
	(.0846)			
Lagged 3–6 years		.2507**		
		(.0982)		
Dummy variable: top quartile of 3–6 year lagged equity return distribution			.0658	
			(.0411)	
Dummy variable: bottom quartile of 3–6 year lagged equity return distribution			−.0339	
			(.0413)	
P-value: *t*-test of equality of dummy variable coefficients across quartiles			.0442	
ΔShareholder wealth/worker lagged 3–6 years (hundreds of thousands of dollars/worker)				.0387**
				(.0195)
Annual employment growth rate:				
Lagged 1–2 years	−.2128			
	(.1485)			
Lagged 3–4 years	−.0127			
	(.0925)			
Lagged 5–6 years	−.2902**			
	(.1431)			
Lagged 1–6 years		−.4013**	−.3441**	−.3446**
		(.1567)	(.1576)	(.1564)
Log(*K*/*L*)	.0736**	.0761**	.0727**	.0662**
	(.0200)	(.0196)	(.0202)	(.0212)
Log(sales)	.0120	.0176*	.0210**	.0208**
	(.0097)	(.0094)	(.0103)	(.0095)
Hausman specification test: *p*-value	.8448	.8606	.9010	.9294

NOTE.—The random-effects regressions include polynomials in education, experience, and tenure and interactions between these polynomials and a female dummy variable; race dummy variables; region dummy variables; a dummy variable for working in a metropolitan statistical area (MSA); four broad industry dummy variables; and interactions between female and the industry dummies. Standard errors are in parentheses.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

the growth rate coefficients to be equal across lag lengths of one to two, three to four, and five to six years. A 1 percent increase in a firm's annual growth rate over the past six years, with its size held constant, is associated with a 0.4 percent decrease in wages. This somewhat surprising result is consistent with the notions that larger high-wage firms tend to grow more slowly and shrinking firms lay off workers with lower wages and unobserved skills. The elasticity of wages with respect to the capital/labor ratio is .076, and the firm size elasticity is .018. The four firm attributes of size, growth, capital intensity, and equity returns jointly explain about 30 percent of interemployer wage differentials. In results not reported here, we find that the estimated elasticities of wages with respect to size, growth, and capital intensity are fairly sensitive to the model specification because of the high degree of collinearity between these firm attributes. In contrast, the estimated coefficient on equity returns is robust to the exclusion of other firm attributes from the regression. This follows from the fact that equity returns are weakly correlated with all the other explanatory variables in the wage regression.

The estimates in column 2 imply that a one-standard-deviation increase in annual equity returns lagged three to six years (a 17 percent increase) causes a 4.27 percent increase in straight-time pay. The large dispersion in equity returns causes some concern that this significant coefficient estimate is due to wage differences at a few firms that are outliers in the return distribution. Column 3 presents results from a regression that includes dummy variables for whether a firm's lagged equity return is in the top or bottom quartile of the return distribution. There is substantial variation in profitability across the top quartile, where the mean annual equity return is 32.3 percent, and the bottom quartile, where the mean annual return is −11.7 percent over the period three to six years prior to the wage survey. The regression in column 3 exploits this across-quartile variation but ignores any within-quartile variation in returns and therefore mitigates the impact of outliers. We find that the highest-return employers pay 9.97 percent more than the lowest-return firms, ceteris paribus.

Previous studies hypothesize that the relationship between wages and profitability is best captured by including profits per worker as an explanatory variable in wage regressions. We implement this approach by measuring the dollar change in shareholder wealth per worker over the period three to six years prior to the WCP survey. The results in column 4 show that a $100,000 increase in shareholder wealth per worker (approximately one standard deviation) raises wages by 3.87 percent, or about $1,050 per year, for the mean worker with an annual salary of $27,000. The present value of this wage increase, with a 10 percent real interest rate (the mean firm's rate of return), is roughly one-tenth the size of the gain in shareholder wealth. These wage effects are separate from any explicit bonuses or employee stock ownership plans and hence are conservative estimates of the elasticity of wages with respect to innovations in equity values. Finally, note that whether we measure profitability in terms of percentage rates of return on equity or dollar increases in shareholder wealth per worker, a one-standard-deviation increase in equity values leads to approximately a 4 percent increase in straight-time pay.

TABLE 5
EFFECT OF ACCOUNTING PROFITS ON CURRENT PAY (Sample Size 807)

	(1)	(2)
Average corporate income/worker lagged 3–6 years (hundreds of thousands of dollars/worker)	−.0802	.2143**
	(.1316)	(.0804)
Annual employment growth lagged 1–6 years	−.3220**	−.4168**
	(.1567)	(.1563)
Log(K/L)	.0874**	
	(.0321)	
Log(sales)	.0211**	.0285**
	(.0094)	(.0091)
Hausman specification test: p-value	.8533	.9068

NOTE.—The random-effects regressions include polynomials in education, experience, and tenure and interactions between these polynomials and a female dummy variable; race dummy variables; region dummy variables; a dummy variable for working in an MSA; four broad industry dummy variables; and interactions between female and the industry dummies. Standard errors are in parentheses.
** Significant at the 5 percent level.

C. *Using Accounting Profits per Worker as a Measure of Profitability*

A number of studies find a significant positive correlation between wages and lagged accounting profits or quasi rents per worker. A firm's net operating income proxies for both current economic profits and quasi rents on sunk capital investment: the correlation between operating income per worker and the log capital/labor ratio is .90, and the correlation between accounting profits and equity returns is .30. A positive correlation between wages and corporate income per worker is inconclusive evidence of rent sharing because it may merely reflect the fact that capital and skilled labor are complements. Suppose that capital-intensive firms, with greater quasi rents, tend to employ workers with more unobserved skills. It will then appear as though capital-intensive firms pay higher wages for a given quality of labor.

Table 5 reports estimates of log wage regressions that use accounting profits (operating income) per worker, averaged over the period three to six years prior to the WCP survey, as a measure of firm profitability. The coefficient on accounting profits per worker in column 1 is insignificantly different from zero.[12] Multicollinearity between a firm's income per employee and capital intensity probably explains the difference between this result and those in previous studies. In column 2, accounting profits have a significant positive impact on wages if capital intensity is excluded from the model. This suggests that earlier findings

[12] In regressions not reported here, we also failed to find a significant coefficient on accounting profits per worker lagged one to two, three to four, and five to six years.

of a positive correlation between wages and accounting profits per worker may merely indicate that more capital-intensive firms tend to have higher accounting profits per worker and employ more skilled labor.

IV. Extensions

A. *Market-Correlated Risk and Idiosyncratic Returns*

Risk-sharing contract models suggest that wages should be related to the market-correlated component of equity returns, whereas incentive contracts predict that pay should vary more with the firm-specific or idiosyncratic component of returns. We attempt to distinguish between the risk-sharing and incentive contract explanations of the wage–equity return relationship by decomposing each firm's return into market-correlated and idiosyncratic components and including each component into the log wage regression.

Let R_{jt} and R_{mt} denote firm j's equity return and the return on the market portfolio in year t, respectively. The systematic risk of firm j's equity share is $\beta_{jt} = \mathrm{Cov}(R_{jt}, R_{mt})/\mathrm{Var}(R_{mt})$.[13] The market-correlated component of firm j's return is $\beta_{jt}E(R_{mt})$, and the idiosyncratic component is $R_{jt} - \beta_{jt}E(R_{mt})$. Risk-sharing contract models predict that wage variation is likely to be greater in high-β firms, the wage risk should be proportional to $\beta_j^2\mathrm{Var}(R_{mt})$, and workers should receive a compensating differential for this risk.

The mean firm in our sample has a β of 1.13, and the standard deviation of β across firms is .31. Idiosyncratic returns have a mean of −.033 and a standard deviation of .165 over the period three to six years before the WCP survey in our sample. Most of the sample variation in equity returns is due to idiosyncratic risk (e.g., the correlation between overall and idiosyncratic returns is .95) because all observations occur in 1989 or 1990 and there is little variation in $E(R_{mt})$ in our sample. Therefore, it is difficult to determine whether the idiosyncratic and market-correlated components of returns have different effects on wages. An ideal sample design, in which there is substantial variation in both components of equity returns, could provide information on the relative importance of risk sharing and incentives in the wage–equity return relationship.

In regressions not reported here, we include both the market-correlated and idiosyncratic components of equity returns in the regression

[13] We use the value-weighted portfolio of New York Stock Exchange (NYSE) stocks as the market portfolio, realized returns on the NYSE index to proxy for $E(R_{mt})$, and estimate β_{jt} using monthly data over a period from five years prior to the WCP survey until five years after the survey.

model in (1). Both components of returns have significant positive effects on wages, but the difference in effects is not statistically significant because of the imprecise measurement of the coefficient on the market-correlated component of returns. We also estimated versions of the wage equation that included either $\beta_j^2\mathrm{Var}(R_{mt})$ or β_j and found no evidence that high-β employers pay higher wages, all else equal.

B. *Is the Wage–Equity Return Relationship Due to Inelastic Labor Supply?*

Wages and equity returns may be positively related because demand and productivity shocks cause wage changes that lag changes in equity prices. If increases in equity values signal future increases in labor demand, future wages will rise whenever labor supply schedules are upward sloping. The less elastic the short-run supply curve, the greater the change in future wages due to an innovation in equity returns. If general and specific skills are imperfect substitutes, jobs that require specific human capital will have the least elastic labor supply. This suggests that workers with high job tenure who have made the most specific investments should have the least elastic labor supply schedules and the largest elasticities of wages with respect to equity returns.[14]

Column 1 of table 6 presents results from a regression that allows the coefficient on lagged equity returns to differ across workers with high (above-median) and low job tenure. Although the impact of stock returns on wages is slightly larger for high-tenure workers, the difference in coefficients across occupations is not statistically significant. Column 2 presents results that include separate coefficients, by high- and low-tenure workers, on the change in shareholder wealth per worker lagged three to six years. The shareholder wealth effect on wages is nearly twice as large for high-tenure workers as it is for low-tenure workers, but the difference in coefficients across groups is insignificantly different from zero. In regressions not reported here, we also allowed the equity return and shareholder wealth coefficients to differ by the broad occupation of the worker. Equity return and shareholder wealth coefficients are larger for professional and administrative workers, compared to technical and clerical workers, but these differences are insignificantly different from zero.

[14] High-tenure workers will also have higher wage–equity return elasticities if the amount of rent sharing increases with job tenure.

TABLE 6
DOES THE WAGE–EQUITY RETURN RELATIONSHIP DIFFER BY JOB TENURE?
(Sample Size 807)

	(1)	(2)
Equity return (lagged 3–6 years) × low-tenure dummy	.2706**	
	(.1082)	
Equity return (lagged 3–6 years) × high-tenure dummy	.2263*	
	(.1253)	
ΔShareholder wealth/worker (lagged 3–6 years) × low-tenure dummy		.0441**
		(.0200)
ΔShareholder wealth/worker (lagged 3–6 years) × high-tenure dummy		.0226
		(.0248)
p-value: test of equal coefficients across groups	.7209	.2758
Annual employment growth previous 1–6 years	−.4030**	−.3470**
	(.1567)	(.1565)
Log(K/L)	.0759**	.0664**
	(.0195)	(.0212)
Log(sales)	.0175*	.0207**
	(.0093)	(.0095)
Hausman specification test: p-value	.8660	.9507

NOTE.—The random-effects regressions include polynomials in education, experience, and tenure and interactions between these polynomials and a female dummy variable; race dummy variables; region dummy variables; a dummy variable for working in an MSA; four broad industry dummy variables; and interactions between female and the industry dummies. Standard errors are in parentheses.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

C. *Do Equity Returns and Other Firm Characteristics Proxy for Unobserved Skills?*

Innovations in shareholder wealth are unlikely to be correlated with observed firm or worker characteristics (such as average skills), ex ante, because of efficient capital markets. It is quite possible, however, that realized equity returns are correlated with establishment characteristics such as the average level of skills at the employer. This correlation could arise if demand or productivity shocks increased the profitability of firms that tended to employ relatively more skilled workers in the 1980s. If this is true, the significant coefficients on equity returns may be due to omitted variable bias. Wages may be higher in more profitable firms because unobserved skills are higher in these firms. Although we cannot speculate how much of the wage-profits relationship is due to a correlation between profitability and unobserved skills, in this subsection we examine whether average *observed* skills at an employer are positively correlated with various measures of firm profitability.[15] In the results below we use a skill index, the mean of $\mathbf{X}_{ij}\beta_{OLS}$ at an employer, to measure average observed skills at the employer.[16] The standard deviation of the skill index is .25 across establishments, so there is substantial variation in skills across employers.

[15] Our failure to reject the random-effects specification tests in Sec. III implies that there is no significant correlation between the employer-specific component of the error term in (1) and worker characteristics ($\mathbf{X}_{ij}$). Firm-specific characteristics ($\mathbf{F}_j$), such as equity returns, may be significantly correlated with the employer-specific component of the error term and not result in a rejection of the random-effects specification test.

TABLE 7
ARE FIRM CHARACTERISTICS CORRELATED WITH OBSERVED WORKER SKILLS?
(Sample Size 109)
Dependent Variable: Mean Skill Index by Establishment

Independent Variable	(1)	(2)	(3)
Equity returns lagged 3–6 years	−.0972		
	(.1502)		
ΔShareholder wealth/worker lagged 3–6 years		.0289	
		(.0268)	
Operating income/worker lagged 3–6 years			.2333
			(.2552)
Employment growth lagged 1–6 years	−.7378**	−.7693**	−.7669**
	(.2173)	(.2154)	(.2043)
Log(sales)	.0435**	.0416**	.0448**
	(.0123)	(.0121)	(.0125)
Log(K/L)	−.0063	−.0159	−.0380
	(.0397)	(.0422)	(.0655)
R^2	.3869	.3873	.4033

NOTE.—Standard errors are in parentheses and are heteroskedasticity-consistent.
** Significant at the 5 percent level.

Table 7 presents results from regressions with one observation per establishment, where the dependent variable is the average skill index at an establishment, and explanatory variables include firm size, growth, capital intensity, profitability, and industry, region, and metropolitan area dummy variables.[17] Observed skills at an employer are insignificantly correlated with each of the profitability variables. Therefore, it is less plausible that the wage–equity return relationship in table 4 merely reflects high stock market gains in the 1980s at firms that employ relatively skilled workers. Average skills are significantly positively correlated with firm size and significantly negatively correlated with the firm's growth rate. The fact that average observed skills differ substantially across large and small and growing and shrinking firms suggests that caution should be used in interpreting estimated elasticities of wages with respect to firm size or growth rates or both.

[16] This equals the wage that the mean worker at employer j would expect to earn at the mean employer in our sample. In regressions not reported here, we obtain similar results using average education, tenure, or experience at an employer as proxies for skills.

[17] We present heteroskedasticity-consistent standard errors because the dependent variable, average skills at an employer, is estimated with varying precision across establishments.

V. Wage Changes and Equity Returns

It is still possible that innovations in shareholder wealth proxy for factors that influence wages but are omitted from our model. Wage differences over time would purge the error term of time-invariant job match–specific factors that are omitted from the model. Our ability to difference the data is limited because the WCP is a cross-section data set, and wages at the beginning of the job match are reported for a relatively small fraction of our sample.[18] An analysis of wage changes may also be problematic because the length of time between current and starting wage observations ranges from less than one year to over 30 years for workers in our sample. Despite these data limitations, we use two approaches to mitigate the influence of time-invariant unobserved factors in our regression model. First, we include a worker's starting wage as a conditioning variable in our current wage regressions. By estimating the effect of equity returns on current wages, conditional on starting pay at an employer, we hold constant time-invariant job match–specific variables. Our second approach is to explicitly difference wage observations over time and estimate the effect of changes in equity values on wage changes over a worker's tenure for job matches that have lasted no more than 10 years.

A. *The Relationship between Wages and Equity Returns Conditional on Starting Pay*

Columns 1 and 2 of table 8 report the means and standard deviations of worker characteristics for the 331 workers in our sample with valid starting wage data. These workers are employed in 57 establishments owned by 52 different Compustat firms. The mean current monthly wage in this sample is $2,121, and the mean real starting wage is $1,695. The typical worker has 7.25 years of tenure and real wage growth of 24.1 percent with his or her current employer.

Table 9 reports results of log current wage regressions that include all the explanatory variables in equation (1) and two additional variables: the worker's log wage at the start of the job match and an interaction between tenure and the log starting wage. Columns 1 and 2 report results from random-effects regressions that are similar to those reported in columns 2 and 4 of table 4. It is not surprising that specification tests reject the hypothesis that the employer-specific component of the error term is uncorrelated with worker characteristics, which now include the wage at the start of the job match. We therefore present OLS estimates in columns 3 and 4, where the standard errors have been corrected for the within-employer correlation in the error term. A one-percentage-point increase in the annual equity return over the past three to six years leads to a 0.268 percent increase in wages, conditional on the worker's wage at the start of the job match. A $100,000 increase in shareholder wealth per worker over the same period causes a 3.89 percent increase in wages, conditional on starting pay. These estimates are remarkably similar to the random-effects results in table 4. Even after one controls for a worker's wage at the start of the job match, firms pay a substantial wage premium if their equity shares performed well three to six years ago.

[18] In Bronars and Famulari (1998), we found that when starting pay is missing for one worker at an employer, it tends to be missing for all workers at an employer. Missing starting pay has a regional pattern, but no other worker or firm characteristics are significantly related to the probability of nonresponse for this question.

TABLE 8
STARTING PAY SAMPLES

	All Workers with Reported Starting Pay (N=331)		Workers Hired in Past 10 Years with Reported Starting Pay (N=245)	
Variable	Mean	Standard Deviation	Mean	Standard Deviation
Real wage	2,120.8	931.8	1,992.2	909.5
Log(real wage)	7.562	.450	7.495	.458
Real starting wage	1,695.0	822.9	1,730.2	849.0
Log(start wage)	7.321	.485	7.341	.481
Real wage growth	.241	.319	.154	.220
Tenure	7.245	6.449	3.992	2.788
Education	14.616	2.236	14.702	2.265
Experience	17.338	9.549	14.951	8.945
Starting experience	10.094	8.385	10.959	8.435
Female	.438		.457	
Black	.051		.049	
Other	.060		.069	
Surveyed in 1990	.870		.853	
Census region:				
Northeast	.338		.331	
Midwest	.314		.286	
South	.257		.290	
West	.091		.094	
Lives in an MSA	.761		.767	
Establishment size:				
Under 500 employees	.299		.314	
500–999 employees	.190		.192	
1,000 or more employees	.511		.494	

NOTE.—All wage values have been converted to 1983 dollars using the employment cost index for wages and salaries.

TABLE 9
EFFECT OF EQUITY RETURNS ON WAGES CONDITIONAL ON STARTING PAY
(Sample Size 331)

	RANDOM EFFECTS		OLS WITH CORRECTED STANDARD ERRORS	
	(1)	(2)	(3)	(4)
Annual equity return lagged 3-6 years	.2840**		.2681**	
	(.1115)		(.1005)	
ΔShareholder wealth/worker lagged 3-6 years		.0384**		.0389**
		(.0176)		(.0156)
Annual employment growth lagged 1-6 years	−.3898**	−.2911*	−.3666**	−.2746**
	(.1702)	(.1628)	(.1432)	(.1334)
Log(*K*/*L*)	.0320	.0069	.0338	.0085
	(.0274)	(.0285)	(.0217)	(.0219)
Log(sales)	.0020	.0084	.0003	.0064
	(.0093)	(.0091)	(.0084)	(.0090)
Log(real starting wage)	.9168**	.9293**	.9154**	.9323**
	(.0590)	(.0593)	(.0758)	(.0739)
Log(real starting wage) × tenure	−.0449**	−.0448**	−.0437**	−.0443**
	(.0052)	(.0052)	(.0056)	(.0056)
Hausman specification test: *p*-value	.0297	.0029		

NOTE.—Regressions include broad industry and region dummy variables and a dummy variable for working in an MSA. Standard errors are in parentheses.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

B. *Wage Change Regressions*

In order to difference all the variables in our log wage regression, we must observe worker and firm characteristics in the current year and the worker's starting year. This sample selection criterion further limits our sample because (1) the Compustat data pertain to 1972 to the present, and (2) changes in the ownership of establishments make it difficult to calculate changes in firm-specific attributes over the duration of long job spells. We therefore limit our analysis of wage changes to workers with reported starting wages and job matches that have lasted no more than 10 years. This sample selection restriction limits our sample to 245 of the 331 workers with valid starting wage data but includes all 57 establishments and 52 firms from the starting wage subsample. Columns 3 and 4 of table 8 present the means and standard deviations of key variables for the 245 workers with valid starting wage data and 10 or fewer years of job tenure. The mean worker in this sample has about four years of tenure on the job and real wage growth of 15.4 percent over her job spell.

We use this sample of 245 workers with 10 or fewer years of tenure to estimate the following wage growth regression:

$$\begin{aligned}\ln W_{ijt} - \ln W_{ijs} = {} & \pi_0 + \pi_1(t-s) + \pi_2(t-s)^2 + \pi_3 R_{jst-1} \\ & + \pi_4 \mathrm{EXP}_{ijs} + \pi_6 \mathrm{FEMALE}_i + \pi_7 I_j \\ & + \pi_8[\log(\mathrm{SALES}_{jt-1}) - \log(\mathrm{SALES}_{js-1})] \\ & + \pi_9\left[\log\left(\frac{K}{L}\right)_{jt-1} - \log\left(\frac{K}{L}\right)_{js-1}\right] + u_{ij}. \end{aligned} \quad (2)$$

In this regression, t is the current year and s is the year in which the job match began; $\ln W_{ijt} - \ln W_{ijs}$ is worker i's cumulative wage growth at employer j; EXP_{ijs} is i's potential experience at the start of the job; $t - s$ is the length of the job match; and R_{jst-1} is either firm j's annual rate of return from year s through year $t-1$ or the annual gain in shareholder wealth per worker (measured in hundreds of thousands of dollars per worker per year) over the job match. The regression also includes a dummy variable for the sex of the worker (FEMALE_i), four broad industry dummy variables (I_j), and changes in firm size and capital intensity between years $s-1$ and $t-1$. Finally, we assume that the error term u_{ij} has an establishment-specific component and estimate (2) using random effects.

Table 10 presents results from the wage growth regression in (2). The coefficients reported in column 1 show a positive but insignificant relationship between a worker's wage growth and the annual equity return since the worker was hired by the firm. The results in column 3 show that for each additional $100,000 per year of shareholder wealth accumulated over the duration of the job match, wages increase by 8.3 percent, on average. Although this estimated effect is large, it is insignificantly different from zero.

Our cross-section results suggest that innovations in wealth affect straight-time wages only after a lag of three years. The regressions in columns 2 and 4 of table 10 allow for separate coefficients on equity returns and changes in shareholder wealth for workers with three or more years of tenure and for workers with two or fewer years of tenure. In both specifications we find a significant positive relationship between wage growth and shareholder returns for workers with three or more years of job tenure. The estimated elasticity of wage changes with respect to equity returns in column 3 is quite similar to the cross-section results in tables 4 and 9. A 1 percent increase in annual equity returns over the duration of the job increases the wages of workers with three or more years of tenure by 0.304 percent. The coefficients in column 4 imply that a $100,000 increase in shareholder wealth per worker, accumulated over a four-year job match, leads to a wage premium of 4.2

TABLE 10
RANDOM-EFFECTS WAGE CHANGE REGRESSIONS: WORKERS HIRED IN THE PAST 10 YEARS
(Sample Size 245)

	(1)	(2)	(3)	(4)
Annual equity return over duration of job match	.0587			
	(.0793)			
Annual return × dummy variable for 3 or more years of tenure		.3040**		
		(.1359)		
Annual return × dummy variable for 1–2 years of tenure		−.0196		
		(.0838)		
ΔShareholder wealth over duration of job match			.0831	
			(.0545)	
ΔAnnual shareholder wealth × dummy variable for 3 or more years of tenure				.1723**
				(.0733)
ΔAnnual shareholder wealth × dummy variable for 1–2 years of tenure				.0060
				(.0683)
ΔLog sales since hired by the firm	−.0555	−.0708*	−.0621*	−.0666*
	(.0382)	(.0376)	(.0360)	(.0352)
ΔLog(K/L) since hired by the firm	.0286	.0168	.0232	.0166
	(.0463)	(.0443)	(.0445)	
Years of tenure	.0705**	.0567**	.0739**	.0629**
	(.0190)	(.0200)	(.0185)	(.0196)
Years of tenure squared	−.0027	−.0017	−.0030*	−.0022
	(.0018)	(.0018)	(.0017)	(.0018)
Female	.0258	.0281	.0193	.0175
	(.0249)	(.0248)	(.0249)	(.0247)
Years of experience at the start of the job match	−.0037	−.0037	−.0039	−.0039
	(.0047)	(.0047)	(.0047)	(.0047)
Years of experience squared/100 at the start of the job match	.0012	.0018	.0026	.0026
	(.0149)	(.0149)	(.0150)	(.0149)
Hausman specification test: p-value	.6230	.5228	.3679	.1501

NOTE.—Standard errors are in parentheses.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

percent, or approximately $1,000 per year.[19] We find no evidence that equity returns affect straight-time wages during the first two years of the job match.

VI. CEO Salary Changes and Equity Returns

We now compare pay–equity return elasticities for white-collar workers and CEOs. Although many previous studies have estimated the elasticity of CEO salaries with respect to equity returns, these results are typically based on first differences in log CEO salaries (or total compensation). Our white-collar wage elasticities are based on one-year to 10-year differences in log wages, with a mean of about four years between pay observations. In this section we use a measure of CEO salary change that is similar to the one used for white-collar workers. We then estimate regressions on CEO log salary changes that are similar to the wage change regression in (2) and compare the magnitude of the equity return coefficients across samples.

[19] This describes the impact of a one-standard-deviation increase in shareholder wealth per worker evaluated at sample means. The mean job match has lasted four years, and the mean change in shareholder wealth per worker per year is $12,730 with a standard deviation of $24,400 per year.

The CEO salary data are available for 60 of the firms in our WCP sample from the annual executive compensation information published in *Forbes* magazine. For the 47 executives who became CEOs within the past 10 years, we construct a salary change measure that is identical to wage changes in the WCP sample (i.e., the CEO's log salary in the WCP survey year minus the log salary in his first year as CEO). For the remaining executives, with more than 10 years on the job, we used pay in 1981 as a substitute for starting CEO salary.[20] The mean tenure for the initial executive salary observation in our sample is 3.43 years. Mean CEO tenure during the WCP survey year is 7.23 years, so our rule for generating CEO salary change data yields observations that are 3.8 years apart, on average.

The average CEO has an annual salary of $828,000 (in 1983 dollars) and is 55 years old in the WCP survey year. The dependent variable in our regressions is the change in the CEO's log salary between our initial observation and the WCP survey year. The mean of the dependent variable is .320, with a standard deviation of .374. Thus the typical CEO experienced a 32 percent increase in his real salary over the 3.8 years between initial and current pay observations.

The explanatory variables in the CEO salary change regression include two alternative measures of changes in shareholder wealth: the annual equity return between CEO salary observations and the value of shareholder wealth accumulated per year over the same time period, in billions of 1983 dollars. Other explanatory variables include changes in log sales and capital intensity between the initial and current salary observations, age at the beginning of the CEO's term, and the number of years between pay observations. The estimated CEO salary elasticity in column 1 of table 11 is .429. A 1 percent increase in a firm's annual return during the typical CEO's tenure of four years increases the firm's value by about 4 percent, raises the CEO's salary by 0.43 percent, and increases white-collar wages (for workers with three or more years of tenure) by 0.30 percent.[21] The estimates in column 2 imply that a $4

[20] *Forbes* first reported four of the CEO salaries after 1981 for CEOs who began their term before 1981. For these CEOs we used the initial observation in the *Forbes* salary survey as a substitute for their starting pay.

[21] The impact of changes in equity values on total CEO wealth, due to bonuses, stock options, and the executive's stock portfolio, is substantially larger than the salary elasticities presented here. See Garen (1994) for a comparison of these effects for a large sample of CEOs.

TABLE 11
CEO Salary Change Regressions (Sample Size 60)

	(1)	(2)
Annual equity return between CEO salary observations	.4294**	
	(.0975)	
Annual Δshareholder wealth between CEO salary observations		.0842**
		(.0399)
Change in log sales	−.0296	−.0330
	(.1292)	(.1464)
Change in log(*K*/*L*)	−.0554	−.1786
	(.1684)	(.1888)
CEO's age for the first salary observation	.0074	.0043
	(.0079)	(.0090)
Number of years between salary observations	.0579**	.4284**
	(.0168)	(.0190)
R^2	.3740	.2082

Note.—Standard errors are in parentheses.
** Significant at the 5 percent level.

billion increase in shareholder wealth over a four-year period would increase the CEO's salary by 8.4 percent. The same $4 billion increase in equity values would raise white-collar wages by 4.8 percent over four years according to the estimates in table 10.[22]

VII. Conclusion

The straight-time wages of white-collar workers are significantly higher in firms with higher equity returns. A 1 percent increase in a firm's annual return over the previous three to six years raises wages by 0.25–0.30 percent. This wage elasticity is remarkably similar whether it is based on cross-section or wage change regressions. Equity returns have an impact on wage changes only for workers with three or more years of job tenure, but the percentage effect on pay is only one-third smaller than the corresponding elasticity for CEO salaries. When profitability is measured as changes in dollars of shareholder wealth per worker, each $100,000 increase in wealth per worker raises white-collar pay by about 4 percent, or $1,000 per year. This suggests that a $10 increase in shareholder wealth causes the present value of a firm's wage bill to increase by $1. The magnitude of our results is surprising given that few of the workers in our sample are likely to be covered by collectively bargained contracts, and our wage effects are separate from explicit bonuses or employee stock ownership plans.

Our results refine the findings of earlier studies in several important ways. We find that wages are insignificantly related to lagged accounting profits per worker, after we control for a firm's capital intensity. We conjecture that the positive correlation obtained in earlier studies may be explained, in part, by mismeasurement of the opportunity cost of capital, the complementarity of skilled labor and capital, and omitted variable bias.

It is unlikely that wages are positively related to equity returns because more successful firms employ more skilled labor that is also more expensive. We estimate the impact of equity returns on wages after controlling for the worker's age, race, sex, education, and job tenure. Moreover, wage change regressions that difference out time-invariant firm and worker characteristics, including unobserved skills, yield nearly identical coefficients on equity returns. Finally, equity returns are insignificantly related to observed worker characteristics, which suggests that the bias due to omitted skill variables in our sample is small.

Our empirical results also demonstrate that the positive relationship between wages and shareholder wealth is not explained by demand shocks and inelastic labor supply schedules. More rapidly growing firms employ less skilled workers and pay lower wages, all else equal. In addition, estimated equity return elasticities are nearly identical whether wage regressions include or exclude employment growth variables. Finally, wage–equity return elasticities are larger at longer lags, and these elasticities are not significantly higher for workers with more specific human capital, which contradicts the predictions of a simple model with demand shocks and inelastic labor supply.

[22] The mean firm has 36,000 employees, so a $4 billion increase in wealth over four years translates into a $27,778 increase in shareholder wealth per worker per year.

References

Abowd, John M.; Kramarz, Francis; and Margolis, David N. "High Wage Workers and High Wage Firms." *Econometrica* 67 (March 1999): 251–333.

Abowd, John M., and Lemieux, Thomas. "The Effects of Product Market Competition on Collective Bargaining Agreements: The Case of Foreign Competition in Canada." *Q.J.E.* 108 (November 1993): 983–1014.

Allen, Steven G. "Updated Notes on the Interindustry Wage Structure, 1890–1990." *Indus. and Labor Relations Rev.* 48 (January 1995): 305–21.

Blanchflower, David G.; Oswald, Andrew J.; and Garrett, Mario D. "Insider Power in Wage Determination." *Economica* 57 (May 1990): 143–70.

Blanchflower, David G.; Oswald, Andrew J.; and Sanfey, Peter. "Wages, Profits and Rent-Sharing." *Q.J.E.* 111 (February 1996): 227–51.

Bronars, Stephen G., and Famulari, Melissa. "Wage, Tenure, and Wage Growth Variation within and across Establishments." *J. Labor Econ.* 15 (April 1997): 285–317.

———. "Employer-Provided Training, Wages, and Capital Investment." In *Labor Statistics Measurement Issues,* edited by John Haltiwanger, Marilyn E. Manser, and Robert Topel. Chicago: Univ. Chicago Press (for NBER), 1998.

Christofides, Louis N., and Oswald, Andrew J. "Real Wage Determination and Rent-Sharing in Collective Bargaining Agreements." *Q.J.E.* 107 (August 1992): 985–1002.

Currie, Janet, and McConnell, Sheena. "Firm-Specific Determinants of the Real Wage." *Rev. Econ. and Statis.* 74 (May 1992): 297–304.

Dickens, William T., and Katz, Lawrence F. "Interindustry Wage Differences and Theories of Wage Determination." Working Paper no. 2271. Cambridge, Mass.: NBER, June 1987.

Garen, John E. "Executive Compensation and Principal-Agent Theory." *J.P.E.* 102 (December 1994): 1175–99.

Hausman, Jerry A. "Specification Tests in Econometrics." *Econometrica* 46 (November 1978): 1251–71.

Hildreth, Andrew K. G., and Oswald, Andrew J. "Rent-Sharing and Wages: Evidence from Company and Establishment Panels." *J. Labor Econ.* 15 (April 1997): 318–37.

Holmlund, Bertil, and Zetterberg, Johnny. "Insider Effects in Wage Determination: Evidence from Five Countries." *European Econ. Rev.* 35 (July 1991): 1009–34.

Jensen, Michael C., and Murphy, Kevin J. "Performance Pay and Top-Management Incentives." *J.P.E.* 98 (April 1990): 225–64.

Katz, Lawrence F., and Summers, Lawrence H. "Industry Rents: Evidence and Implications." *Brookings Papers Econ. Activity, Microeconomics* (1989), pp. 209–75.

Krueger, Alan B., and Summers, Lawrence H. "Efficiency Wages and the Inter-industry Wage Structure." *Econometrica* 56 (March 1988): 259–93.

Nickell, Stephen, and Kong, Paul. "An Investigation into the Power of Insiders in Wage Determination." *European Econ. Rev.* 36 (December 1992): 1573–99.

Nickell, Stephen; Vainiomaki, J.; and Wadhwani, Sushil. "Wages and Product Market Power." *Economica* 61 (November 1994): 457–73.

Nickell, Stephen, and Wadhwani, Sushil. "Insider Forces and Wage Determination." *Econ. J.* 100 (June 1990): 496–509.

Rosen, Sherwin. "Contracts and the Market for Executives." In *Contract Economics,* edited by Lars Werin and Hans Wijkander. Oxford: Blackwell, 1992.