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The Employer Size–Wage Effect

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We consider six explanations for the positive relationship between employer size and wages: large employers (1) hire higher-quality workers, (2) offer inferior working conditions, (3) make more use of high wages to forestall unionization, (4) have more ability to pay high wages, (5) face smaller pools of applicants relative to vacancies, and (6) are less able to monitor their workers. We find some support for the first of these, but there remains a significant wage premium for those working for large employers.

There is much evidence that "large" employers pay more than "small" employers even when their union status is the same (Lester 1967; Masters 1969; Antos 1981; Mellow 1982; Atrostic 1983; Oi 1983). There is, however, much less information that can help us answer a number of key questions concerning this wage differential: Is it company size or establishment size that matters for wages, or does each have an independent effect? If employer size is treated in a continuous fashion, precisely how big is the size-wage effect? How much of

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this employer size–wage differential can be explained by the fact that employers who are larger hire better workers? How much can be explained by a size differential in working conditions? How much by the fact that, in the nonunion sector, larger employers do more to avoid unions than smaller employers? How much by a differential in product market power? Why else might larger employers pay more?

This study addresses each of these questions. Section I discusses the four factors that dominate previous research on determinants of the employer size–wage differential: labor quality, working conditions, union avoidance, and product market power.

The size-wage differential is one of the key differentials observed in labor markets. It is particularly interesting because, unlike the union wage differential, it exists in the absence of an obvious agent, one of whose goals is its existence. Hence, if employers of different sizes pay very differently for the same quality of labor working in a similar environment, there is no readily available deus ex machina to save the day: our knowledge of the labor exchange must ultimately be relied on.

Section II provides the stylized facts of the matter. The evidence reveals that the size-wage effect is quite large: the wage gain associated with moving from an employer whose size is one standard deviation below average to an employer one standard deviation above average is roughly the same as the gain associated with moving from a nonunionized to a unionized employer. Company size and establishment size have independent effects on pay. Finally, the findings presented strongly suggest that, while a size differential in labor quality can explain about one-half of the total size-wage differential, the other three factors under consideration can explain little of the remainder. Since the residual size-wage effect is large, the question of why size matters for wages is much more perplexing than it may have appeared at first blush.

Section III presents a number of additional hypotheses about the origins of employer size–pay differentials, and the likely explanatory power of each of these theories is assessed. Section IV presents conclusions.

I. Some Traditional Explanations

Both neoclassical and institutional labor economists have offered explanations of why larger employers pay more than smaller employers. As would be expected, the neoclassicists have focused on size differentials in labor quality or working conditions. While the institutionalist approaches are more varied, they often turn to factors such as

union avoidance and product market power. While each of the explanations alluded to is plausible, neither their individual nor their collective power has yet been tested.

The Neoclassical Explanations

The theory of "compensating differentials" or "equalizing differences" is at the heart of neoclassical labor economics. It is for this reason that many of the existing discussions of size-wage differentials focus on size differentials in the quality of labor or the conditions of work.

The labor quality explanation of the size-wage relationship can be simply put: larger firms or establishments employ higher-quality workers. There are several reasons why larger employers might make greater use of higher-quality labor, all else the same. Greater capital intensity of larger establishments and capital-skill complementarity provide one explanation (Hamermesh 1980, p. 386). An alternative is presented by Oi (1983), who suggests that large firms employ higher-quality workers in order to reduce the costs of monitoring a given quantity of labor services. In deriving this result, Oi makes the very strong assumption that greater entrepreneurial ability (which is what generates larger firms in the model) increases the quantity of decision making that can be achieved in an hour of the entrepreneur's time, but it does not affect the number of workers whose output can be monitored per hour by the entrepreneur.[1]

The labor quality explanation of the size-wage effect lends itself to a number of statistical tests. The first involves the estimation of wage equations with cross-sectional data on individuals. Very simply put, the analysis asks whether or not the estimated size-wage effect can be explained in terms of measured dimensions of labor quality. While unmeasured dimensions of quality clearly exist, one might hope that they will be correlated with measured variables such as schooling, age, and the like.

The labor quality explanation can be addressed in a different fashion with longitudinal data, by comparing the wages of the same individual when he or she is working for different-sized employers. To the extent that unmeasured dimensions of worker quality are fixed over time, looking at wage rate change as a function of change in employer size (and other measured factors) will give an estimate of the size-wage effect that is not biased by constant dimensions of labor

[1] For a model in which firm size is related to the ability of the entrepreneur but not necessarily to the ability of other workers, see Rosen (1982).

quality. It should be noted, however, that the downward bias resulting from classical measurement error in the size variable will be exacerbated by differencing if the ratio of error variance to true variance is greater for the change in size than for its level.

Other pieces of evidence may help to distinguish *among* competing labor quality hypotheses. For instance, the explanation that links skill to scale on the basis of capital-skill complementarity tends to take *establishment* size as the relevant measure of scale and hence predict a relationship between wages and establishment size. Oi's monitoring model focuses on the effects of *firm* size. While the correlation between firm and establishment size is strong enough to produce similar results when only one measure is used, including both firm and establishment size can clearly refine the "stylized facts" that must be explained.

If the higher wages of larger employers are due to differences in worker quality, then those working for larger employers would earn no more than they would earn elsewhere. This would imply that the quit rates of larger employers would be no different from the quit rates of smaller employers when wages and nonwage benefits (including a larger menu of potential jobs) are held fixed. Hence, information on size differentials in quit rates, company tenure, and "job" tenure can shed light on the size-wage effect puzzle.

While the details differ, each of the labor quality hypotheses presented above has larger employers and smaller employers paying the same wage for workers of given quality. Alternatively, undesirable working conditions generally associated with larger workplaces—such as greater reliance on rules and less freedom of action and scheduling (Masters 1969; Stafford 1980), more impersonal work atmosphere (Lester 1967), or longer commuting (Scherer 1976, p. 111)—may force larger employers to pay higher wages to get a given quality of labor.

The first step in testing this compensating differentials explanation involves isolating the unattractive aspects of larger workplaces. Including variables for these "job characteristics" in a wage equation should reduce or eliminate the wage premium associated with employer size. Unfortunately, some job characteristics are hard to measure directly. However, since it seems likely that a substantial fraction of the total variation in such working conditions occurs across industries and occupations, an analysis that fits wage equations without and with detailed industry and occupation controls would provide valuable information about the validity of the working conditions explanation. Finally, if the compensating differentials view is correct, size should be positively related to quits if the wage is held constant but working conditions are not.

The Institutional Explanations

One institutionally oriented explanation for the differences in labor market behavior between large and small employers is that large nonunion employers act in many ways as if they were unionized in order to avoid unionization. There is considerable evidence that employers that follow a strategy of "positive labor relations" to avoid unionism will pay higher wages, offer more benefits, and provide better working conditions than otherwise similar nonunion employers (see Curtin 1970, p. 60; Foulkes 1980; Freeman and Medoff 1984, p. 153). Since it is primarily the large employers that adopt such personnel policies, one result is that union wage and benefit differentials vary inversely with size. Freeman and Medoff describe just such a variation, with a union wage differential of 5 percent for workers in firms with more than 1,000 workers compared with 22 percent for workers in firms with fewer than 100 workers. They find a similar pattern for fringe benefits.

The importance of union avoidance efforts can be assessed by more detailed investigation of the size-wage relationship. We can determine whether size-wage differentials exist even within the union sector and in occupations or industries for which there is a near-zero threat of unionism. For both already-organized and unorganizable workers, union avoidance by large employers is unlikely.

Each of the theories discussed so far assumes (or is consistent with) cost-minimizing behavior by firms. Alternatively, large firms or establishments are sometimes said to engage in different labor market behavior than smaller ones because they possess product market power. One common argument is that firms with "monopoly power" may share with their workers some of the "excess" profits or rents that such power yields (Weiss 1966; Mellow 1982). However, even if large employers did use their excess profits to overpay their workers, one must still explain why they pay more than market wages and why competition for these choice jobs does not lead to a work force that is overqualified but not overpaid.

It is not clear whether the product market power explanation refers to large firms or simply to *industries* in which the typical firm is large. The latter view can be tested by controlling for detailed industry in wage regressions. In any case, the premise of the argument—the product demand curves of large employers are less elastic—can be checked directly.

II. Stylized Facts

The tables presented in this section shed light on the likely validity of the explanations outlined above. In addition to documenting the inability of this set of explanations to resolve the employer size–wage effect puzzle, the section provides a new set of facts with which the ultimate explanation must reckon.

Evidence on the Labor Quality Explanation

Tables 1 and 2 relate to the "labor quality" explanation of the size-wage effect. Table 1 offers estimates of size-wage differentials based on five data files: The Current Population Survey (CPS) and Quality of Employment Survey (QES) give data for individuals, while the Survey of Employer Expenditures for Employee Compensation (EEEC), the Wage Distribution Survey (WDS), and the Minimum Wage Employer Survey (MWES) contain data for establishments.[2]

Three of these files have information on company size in addition to establishment (or location) size. These make it possible to determine the wage differential associated with establishment size, with company size held constant, and vice versa. The WDS company size variable is, however, only an estimate, calculated as employment in the company's surveyed establishment(s) times the ratio of enterprise (company) to surveyed establishment sales.[3] Establishment size is reported by size category (e.g., 500–1,000 workers) in the CPS and QES (and for company size for multiestablishment companies in the EEEC); each set of categories was converted to a continuous variable using the estimated mean employment by size category and broad industry (based on *County Business Patterns* data [U.S. Bureau of the Census 1977] for establishments and unpublished Small Business Administration data for companies).[4]

For each data set, size effects are given for the total sample and for various subsamples, defined in terms of broad occupational category and unionization. As the column headed "Other Independent Variables" indicates, we control as much as the data allow for variation in labor quality across different-sized employers.[5]

Despite the variety of data sets used, table 1 provides consistent support for these conclusions: (1) For the private-sector wage and salary work force as a whole, there is a substantial wage differential associated with establishment size (with company size not controlled for) even in the presence of controls that would be expected to capture much of the cross-employer differences in labor quality: an employee working at a location with ln(employment) one standard deviation (which equals about two) above average can be expected to earn 6–15 percent more than a similar employee at a location with ln(employment) one standard deviation below average. (2) For the same work force, there appears to be a company size–wage effect when establishment size is controlled for and vice versa.[6] The company size effect is weaker statistically as well as practically. This *may* be due to less accurate measurement of company size, although the very indirect evidence on this possibility is not clear.[7] (3) There is clear evidence of a size-wage effect in each of the three subgroups of workers. Subtler questions—the relative ranking of the three groups' size-wage effects or the relative importance of establishment and company size for each group—receive different answers with different surveys. (4)

[2] Our CPS sample consists of respondents to the May 1979 supplement to the Current Population Survey. To maintain comparability with other data files, we limited our sample to private wage and salary workers. The QES, conducted by the Survey Research Center at the University of Michigan, interviewed those employed 20 or more hours per week (thus excluding many part-time workers) in 1972–73. A subset was reinterviewed in 1977, and we use the file consisting of those interviewed in both waves. The QES wage is annual earnings divided by 52 times hours worked per week (weeks worked in the previous year were not available). Both the EEEC and the WDS are probability samples of private, nonagricultural establishments, conducted by the Bureau of Labor Statistics. The probability of selection is approximately proportional to employment in the establishment. The WDS excludes supervisory workers. The MWES is a survey of establishments conducted by the Survey Research Center in 1980. In addition to oversampling large establishments, it also oversampled those with minimum-wage workers. Consequently, we use the weights calculated by the center in weighting the MWES wage equations. The MWES gives wage distributions (fraction of workers in each of seven intervals) from which we calculated an average wage. The following sources present detailed information on these data sets: CPS, Mellow (1982); QES, Scherer (1976) and Kwoka (1980); EEEC, U.S. Bureau of Labor Statistics (1982); WDS, Gilroy (1981); and MWES, Converse et al. (1981).

[3] If company sales/(sum of reported establishment sales) was less than one or greater than 100, we deleted the observation as an outlier.

[4] When the category boundaries differed (e.g., QES has categories of 1,000–1,999 and 2,000+, while *County Business Patterns* reports only 1,000+), we assumed a Pareto upper tail in estimating mean employment.

[5] We use ordinary least squares (OLS) estimation in table 1 and later tables. We also tried correcting for possible (unspecified) heteroskedasticity using White's (1980) procedure. For the CPS equations in table 1, the standard errors of the size variables were less than 2 percent higher than those computed when possible heteroskedasticity was ignored.

[6] This was previously noted in the CPS data (with categorical size variables) by Mellow (1982) and Oi and Raisian (1985), and by Antos (1981) and Atrostic (1983) for white-collar workers. Dunn (1980, 1984) reported generally similar findings (positive firm size effects but inconsistent establishment size effects) with continuous size variables in smaller, less representative samples.

[7] One piece of evidence supporting the measurement error conjecture is the fact that, in the WDS data, measuring company size by the logarithm of company sales (which is better measured but probably less appropriate than estimated company employment) leads to appreciably larger company size effects. On the other hand, measurement errors are probably least pronounced in the EEEC, where both company employment category and establishment employment are employer reported; yet the size effects are smaller in the EEEC than in the CPS.

TABLE 1

Estimates of the Size-Wage Effect Using Cross-Section Data

Data Set and Year (Sample Size)	Dependent Variable	Other Independent Variables*	Size Variable†	Coefficient of Size Variable‡ TOT	WC	NUBC	UBC
		Individual Analyses					
1a. May CPS, 1979 (13,829; 6,901; 4,591; 2,337)	Ln(usual hourly earnings)	Union coverage, sex, race, schooling, experience and its square, tenure and its square, SMSA (2), region (3), industry (41), occupation (8)	*E*	.027 (.002)	.028 (.003)	.026 (.003)	.013 (.004)
1b. Same	Same	Same	*E*	.015 (.002)	.018 (.003)	.011 (.005)	.001 (.005)
			C	.013 (.002)	.012 (.002)	.013 (.003)	.016 (.004)
2. QES, 1973 cross-section of 1973–77 panel (921; 451; 251; 219)	Ln(hourly earnings)	Same as CPS except union member instead of union coverage, SMSA (1), industry (40)	*E*	.038 (.007)	.059 (.010)	.028 (.014)	.019 (.013)
		Establishment Analyses					
3a. WDS, 1979 data for nonsupervisory workers only (1,440;—; 1,243; 197)	Ln(hourly earnings)	Union coverage, sex (2),§ age (4), SMSA, region (3), industry (40), pay type (2), average production workweek	*E*	.032 (.005)	...	.032 (.005)	.030 (.011)
3b. Same	Same	Same	*E*	.031 (.006)	...	.034 (.007)	.017 (.014)
			C	.001 (.004)	...	−.002 (.004)	.015 (.011)
4a. EEEC, 1974 (4,396; 4,033; 2,240; 1,175)	Ln(total hourly wage)‖	Union coverage, SMSA, region (3), industry (40)	*E*	.020 (.003)	.019 (.003)	.030 (.004)	.032 (.005)
4b. Same	Same	Same	*E*	.014 (.004)	.030 (.005)	.027 (.006)	.032 (.006)
			C	.005 (.003)	−.011 (.004)	.003 (.004)	−.000 (.004)
4c. Same	Ln(total hourly compensation)‖	Same	*E*	.032 (.003)	.033 (.003)	.046 (.004)	.045 (.005)
4d. Same	Same	Same	*E*	.021 (.004)	.036 (.005)	.038 (.007)	.041 (.007)
			C	.011 (.003)	−.004 (.004)	.007 (.004)	.004 (.004)
5a. MWES, 1980 (1,041)	Ln(average hourly wage)	Union coverage, region (3), industry (60)	*E*	.016 (.007)	...	...	...
5b. Same	Same	Same	*E*	.008 (.007)	...	...	...
			M	.079 (.021)	...	...	...

* Numbers in parentheses are the number of dummy variables representing categories of this variable; e.g., region (3) means that there are dummy variables for three main regions plus one omitted (reference) region.

† *E* = ln(establishment employment), *C* = ln(company employment), and *M* = dummy variable for multiple-establishment companies.

‡ The letters TOT refer to all workers, WC to white-collar or nonproduction workers, UBC to unionized blue-collar plus service or production workers, and NUBC to nonunionized blue-collar plus service or production workers. Standard errors are in parentheses.

§ Sex variables include percentage male and percentage of unreported sex; pay type variables include percentage tipped and percentage paid by hourly rate.

‖ Total hourly wage equals total wage/total hours worked plus paid hours leave; total hourly compensation equals total compensation/total hours worked.

When wages are measured by "wages plus fringes per hour worked," size-wage effects are stronger than when the more common "wages per hour paid" are used. Thus the effects of establishment and firm size on fringe benefits and paid vacations and holidays are stronger than their effects on hourly wages.[8]

Table 1 implies that measured dimensions of labor quality cannot fully explain the size-wage effect. Compared with regressions that control only for broad occupation, the additional labor quality variables in table 1 reduce the CPS and QES establishment size and CPS company size effects by roughly one-half.

We also estimated equations similar to those in table 1 for six broad industries[9] using CPS and EEEC data. The pattern of size differentials (particularly the sum of the two size coefficients) was quite similar across broad industries. Moreover, there was no tendency for an industry with larger- or smaller-than-average size effects in one data set to show a similar result in the other one.

Table 1 controls for labor quality by holding constant the worker characteristics that are most obviously related to earnings. An alternative complementary strategy is to look within very narrowly defined occupations. We have explored this approach using data from the Area Wage Surveys (AWS) and the Professional, Administrative, Technical, and Clerical Worker Survey (PATC), both conducted by the Bureau of Labor Statistics. The AWS covers 32 cities over the period 1968–82. The PATC data are based on nationwide surveys for 1965–82. In both data sets, average wages in each occupation and average employment per establishment can be calculated for two size classes, "large" and "medium."[10]

The AWS and PATC data are complementary in that the AWS includes blue-collar occupations while the PATC provides more white-collar detail.[11] Three conclusions emerge from these by-

[8] Positive effects of employer size on fringe benefits have been documented previously by Antos (1981), Freeman (1981), and Atrostic (1983).

[9] The industry groups were mining; manufacturing; transportation and public utilities; trade; finance, insurance, and real estate; and services.

[10] We have avoided using "small" to characterize the smaller size classes because really small establishments are not surveyed. In the AWS, large establishments employ 500 or more workers, while medium establishments employ 50–499. In the PATC survey, large establishments employ 2,500 or more workers, while medium establishments employ 100–2,499 workers. The lower bound is, in some cases, 100 rather than 50 in the AWS and 50 or 250 rather than 100 in the PATC.

[11] The AWS data also include the fraction of office and nonoffice workers covered by collective bargaining in each city-year, but these are not tabulated separately by size class. In analyzing the AWS data, we assumed that the logarithm of the wage depended on a city-year-specific fixed effect, the logarithm of establishment size, and interactions of size with unionization and time. The linear effects of unionization and time are

occupation analyses (available from the authors): (1) There is clear evidence of higher wages in larger establishments; the size effects are centered roughly on .05, which is not very different from the estimates in table 1 when firm size is not held constant.[12] (2) The AWS and especially the PATC provide information on different grade levels (corresponding to different levels of responsibility) for white-collar occupations. A striking regularity among the professional, technical, and managerial workers is the tendency for the wage differential to decline with increasing skill level. Whether one interprets this as a true difference in size-wage effects or as a difference in levels of unmeasured skill within grade levels, it seems to be regular enough to warrant attention.[13] (3) Both data files show a general pattern of *increasing* size differentials between the late 1960s and early 1980s.

We also analyzed salary and fringe benefit data for professional and managerial employees of different-sized firms made available by Hay Associates. Here the occupational stratification was based on the Hay rating of individual jobs for compensation purposes (100 points, 200 points, etc.); at each occupational level we regressed the logarithm of compensation on the logarithm of firm employment and 27 industry dummies. We again find smaller differentials at higher occupational levels, at least up to the lower managerial ranks, and this pattern persists when fringe benefits and incentive pay are added to salary. At higher managerial levels, the size differential becomes larger.

Table 2 addresses the question of whether the size-wage differential can be explained in terms of unmeasured dimensions of labor quality whose effect on wages is fixed over time. Such omitted quality dimensions should not be a source of bias when the earnings function is fit using changes in worker characteristics to explain changes in wages (fixed-effect estimates). The size-wage differential observed with cross-sectional data is reduced by 5–45 percent by estimating the

captured in the fixed effect. Taking differences for the two available size classes in any city-year gives

$$\Delta \ln(W_{ct}) = B_1 \Delta \ln(\text{size}_{ct}) + B_2 \text{union}_{ct} \times \Delta \ln(\text{size}_{ct}) + B_3 t \times \Delta \ln(\text{size}_{ct}).$$

To make each "size effect" as comparable to the others as possible, we have evaluated the size effect for each occupation at the mean value of unionization and for $t = 1982$. With the PATC data a similar procedure was followed, except that there is only one (national total) observation for each size class each year, and there are no unionization data.

[12] The blue-collar occupations seem to show slightly larger size effects, but given that the blue-collar occupations in the table are predominantly skilled maintenance work, one should not make a great deal of this difference.

[13] The trend × size interactions tend to be slightly larger in higher-skilled occupations, suggesting that the pattern of size effects declining with grade level was even more pronounced before 1982.

TABLE 2

Estimates of the Size-Wage Effect Using Longitudinal Data: QES, 1973–77 Differences

Independent Variable	Coefficient (Standard Error) (1)	(2)
Δ ln(establishment size)	.032	.021
	(.010)	(.011)
Δ union status	.108	.132
	(.045)	(.048)
Δ schooling, Δ experience and Δ experience squared, Δ tenure and Δ tenure squared	Yes	Yes
Δ SMSA, Δ region (3), Δ industry (41), and Δ occupation (8)	No	Yes
Ln(establishment size)*	.034	.038
	(.006)	(.007)
Union status*	.105	.102
	(.027)	(.030)

* These estimates are from "level" models that are analogues to the "change" models except that they include dummy variables for sex, race, and year; the estimates were derived with the 1973 and 1977 data for the pooled 1973–77 QES sample used in fitting the change models ($N = 982$).

earnings function with fixed individual effects.[14] This finding strongly[15] suggests that the size-wage differential cannot be explained solely by appealing to the size–labor quality differential. The fixed-effect estimates of the size-wage effect are large in practical terms: if a typical worker went from an establishment with employment one standard deviation below average to an establishment with employment one standard deviation above average, the employee would enjoy a wage increase of 8–12 percent, about as large as the union-nonunion differential in these data.[16]

[14] We also divided the sample into those who changed employers and those who did not. Changes in size in the latter group reflect establishments growing or shrinking and measurement error. The effects of changes in size were larger for the employer-changers than in table 2 and essentially zero for those who did not change employers.

[15] As Griliches and Hausman (1984) note, fixed-effect estimators are likely to intensify the downward bias because of measurement error, and in practice the resulting coefficient estimates are often implausibly small or statistically insignificant. As they suggest, if increasing the period spanned by the two years of data increases the signal/noise ratio by increasing the amount of "real" change in the independent variables, the 5-year span of the QES is an attractive feature of this file. Another concern is based on the observation that workers choose different-sized employers only if it is advantageous to do so. Consequently, self-selection of size changers could bias the coefficient of the change in employer size reported in table 2. If all changes are voluntary and workers do not care about employer size per se, Freeman's (1981) argument suggests that self-selection will lead us to underestimate the true size effect. Solon (1986) considered a model formally equivalent to the case in which size per se matters. If d is the wage gap between large and small employers and a is the compensating differential required by workers, the change regression underestimates the true wage premium as long as $d > a$. Since these results assume voluntary job changing, it is worth considering further evidence. If there is a Mills ratio term that belongs in our wage change equation, it should be larger for voluntary job changers. Adding dummy variables for all job changers and for voluntary job changers left the estimated size coefficient undisturbed. (It increased in the third decimal place.) Similar results were obtained replacing the voluntary change dummy with a dummy for those who had lined up a new job before leaving their old job.

[16] Evans and Leighton (1987) estimated wage change equations using National Longitudinal Survey (Young Men) data. Their results are broadly similar to ours, if a bit less clear. They find that workers whose *firm* size increases experience a (statistically significant) 5.5 percent wage gain, while those whose firm size decreases suffer a (statistically insignificant) 0.6 percent loss. Adding a dummy variable for job changers makes the wage gain only marginally significant ($t = 1.88$).

Evidence on the Working Conditions Explanation

In a purely competitive labor market, a wage differential not explained by labor quality differences must be due to behind-the-scenes differences in working conditions. Can the size-wage effect be explained in terms of differences in the conditions of work?

While the question is simple, providing a convincing response is not. The reason is that "working conditions" are a very complex phenomenon—hard to define and even harder to measure. Since no survey provides an index of the quality of working conditions that would be widely accepted and could be entered into an earnings function that included a size variable, we must conduct a number of less direct but reasonable investigations that, taken together, should permit us to judge the working conditions explanation of the size-wage effect.

Table 3 presents the results of two of these attempts. The first examines the extent to which the size-wage differential is affected by more detailed controls for industry and occupation. More detailed controls should be capturing a greater amount of the variation in working conditions since presumably much of this variation is across industries and occupations.

Our experimentation with industry and occupation controls shows that essentially all the size-wage differential occurs within detailed industries and occupations and thus cannot be explained in terms of a cross-industry or a cross-occupation correlation between establishment size and conditions of work. If differences in working conditions explain the size-wage differential, there must be sizable partial correlations between establishment size and working conditions within detailed industries and occupations.

To deal with this possibility, our second investigation focused on

TABLE 3

ESTIMATES OF THE SIZE-WAGE EFFECT WITH VARIOUS SETS OF CONTROLS FOR NONWAGE WORKING CONDITIONS

Data Set and Year (Sample Size)	Independent Variable (Same as Table 1 Except)	Size Variable	Coefficient of Size Variable*
1a. May CPS, 1979 (13,829)	No industry or occupation dummies	E	.019 (.002)
		C	.015 (.002)
1b. Same	Two-digit census industry dummies (41); "major" census occupation dummies (8)	E	.015 (.002)
		C	.013 (.002)
1c. Same	Three-digit census industry dummies (195); detailed census occupation dummies (37)	E	.016 (.002)
		C	.015 (.002)
2a. QES, 1973 and 1977 (878)	No industry or occupation dummies; year dummy	E	.037 (.006)
2b. Same	Two-digit census industry dummies (41); "major" census occupation dummies (8); year dummy	E	.043 (.007)
2c. Same	Two-digit census industry dummies (41); "major" census occupation dummies (8); year dummy; working conditions variables (10)†	E	.044 (.007)
3a. QES, 1973–77 longitudinal file (439)	Change analogue to model 2a	E	.037 (.010)
3b. Same	Change analogue to model 2b	E	.033 (.011)
3c. Same	Change analogue to model 2c	E	.028 (.011)

NOTE.—See tables 1 and 2 for information about the variables used in the analyses summarized in this table. All results are for the TOT sample.

* Standard errors are in parentheses.

† The working conditions variables are described in the text.

the 1973–77 QES, a longitudinal file that contains information on location size, job conditions, wages, and other factors in both 1973 and 1977. We focused on job conditions that seemed most closely related to issues mentioned in the literature as sensitive to employer size: weekly hours; dummy variables for working on the second or third shift; two variables indicating extent of choice concerning overtime work; variables indicating dangerous or unhealthy conditions on the job and whether the danger/threat problem is serious; catchall variables indicating whether more comfortable, pleasant working

conditions are desired; variables indicating whether any of the entire set of job conditions creates a sizable problem; and variables giving commuting time. In order to make clear the impact of adding these variables, *all* the QES regressions in table 3 are limited to observations for which these variables are available.

Rows 2c and 3c of table 3 present the results of our analysis of the impact of (stated) job conditions on the wage differential associated with location size. The findings indicate that the direct information on the conditions of people's jobs collected in the QES can explain very little of the size-wage effect. We also experimented with additional variables intended to measure more elusive working conditions such as pace of work (Oi and Raisian 1985), relationships with co-workers and supervisors, perceived job security, and so forth. Their collective impact on the size coefficient in row 3c was to *increase* it trivially.

Given the potential problems from measurement errors in the working conditions, which would bias their coefficients downward and reduce their impact on the size coefficient, it is worth asking whether there is any persuasive evidence that working conditions are, in fact, worse in larger employment settings. The lack of such a relationship in the QES data is striking. Of the 42 job characteristic variables included in the regression equation described above, only 21 showed a negative relationship between good characteristics and establishment size, with the other (nonjob characteristic) variables controlled for. Of the four significant negative relationships, three related to promotion issues: perceived unfairness in promotions or lack of opportunity to advance. While it is perhaps too much to claim that a difference in working conditions cannot explain the size-wage relationship, our results suggest that it is an unlikely explanation.[17]

Turnover, Tenure, and Wage-Tenure Profiles

The first two rows in table 4 indicate that the quit rate declines with employer size even when the wage rate is held constant. Rows 3 and 4 show that years of tenure with employer (which reflect absence of quits and discharges) grow significantly with employer size, independent of the size-wage effect; in row 4, a two-standard-deviation difference in size implies a 1.6-year or 20 percent differential in employer

[17] Dunn (1980, 1984) tried to assess the disutility of work by looking at the number of dollars workers would pay for (hypothetical) fringe benefits compared to the number of unpaid hours they would work to obtain the same fringes. She found that this disutility rose with firm size in one sample but not in the other; even in the first sample, the wage premium more than offset the increased disutility. Her results are therefore consistent with our result that, taken together, variations in working conditions are at best a partial explanation for the size-wage relationship.

TABLE 4

ESTIMATES OF THE SIZE-QUIT AND SIZE-TENURE EFFECTS

Data Set and Year (Sample Size)*	Dependent Variable	Mean of Quit Rate or Years of Tenure†	Other Independent Variables	Size Variable‡	Coefficient of Size Variable§
1. Three-digit SIC manufacturing industries, 1958–71 (89)	ln[quit rate/(1 − quit rate)]‖	.019 (.008)	Percentage covered, percentage production, percentage male, four-firm shipments concentration ratio, ln(mean hourly wage)	*C*	−.181 (.069)
2. State × two-digit manufacturing industries, 1972 (151)	ln[quit rate/(1 − quit rate)]‖	.027 (.012)	Percentage union members, index of labor quality, industry (19), region (3), ln(mean hourly wage)	*E*	−.438 (.172)
3. May CPS, 1979 (13,829)	Tenure with employer	6.34 (8.2)	Union coverage, sex, race, schooling, experience and its square, SMSA (2), region (3), industry (41), occupation (8), ln(hourly wage), year dummy	*E*	.242 (.042)
				C	.189 (.029)
4. QES, 1973 and 1977 (1,522)	Tenure with employer	8.51 (8.50)	Same as CPS except union member instead of union coverage, SMSA (1)	*E*	.422 (.096)
5. QES, 1973 and 1977 (1,522)	Time on current job	6.05 (7.07)	Same	*E*	.105 (.086)
6*a*. QES, 1973 and 1977; same three-digit census occupation in 1973 and 1977 (238)	Change employer	.184 (.388)	1973 values; same as in row 4 except no year dummy	*E*	−.032 (.016)
6*b*. QES, 1973 and 1977; different three-digit census occupation in 1973 and 1977 (291)	Change employer	.446 (.498)	Same	*E*	−.041 (.015)

* These analyses are for the TOT samples used in table 1; sample sizes may differ from those in table 1 because of missing data.
† Standard deviations are in parentheses.
‡ Rows 1 and 2 are mean values of total hours per company or establishment; rows 3–6*b* use company and establishment size as in table 1.
§ Standard errors are in parentheses.
‖ Rows 1 and 2 use weighted least squares on aggregated data; rows 3–6*b* use ordinary least squares.

tenure. These results again suggest that the size-wage differential is not simply due to some nonwage "bad" whose prevalence grows with employer size.

It should be noted that one reason why employees might be less likely to leave a large employer is that there is greater opportunity with a large employer to move from one assignment to another without quitting. (Note also, however, that if the worker did not like largeness per se, he or she *would* have to quit.) If this were all that lay behind the size-tenure relationship, that relationship could not be taken to mean that those who work for large employers remain on their jobs longer because the package of wages and working conditions they receive is more attractive than that typically available elsewhere.

This issue is addressed in the last three rows of table 4. The length of time that QES respondents report working "on a job like this one" is only weakly related to employer size (row 5); this is consistent with the claim that there is more internal job movement among the employees of larger employers. However, even for those who did not change (three-digit) occupation between the 1973 and 1977 surveys, employer size is significantly negatively related to the probability of changing employers (row 6*a*), and the relationship is nearly as large as the one for those who did change occupation (row 6*b*). Thus, even among those who remain in the same "job" (as measured by census occupation), those working for large employers are more likely to continue working for them.

Given that table 4's results control for the wage rate, which is positively related to size, we interpret the greater tenure at larger establishments as indicating that large employers offer, if anything, superior working conditions to workers of given quality. An alternative is that they offer packages similar in overall value to those offered by smaller employers but that larger employers' wage profiles are steeper, and so quits are less common.

To test this alternative, we interacted the logarithm of company size (and sometimes the logarithm of establishment size) with tenure and tenure squared (and sometimes experience and experience squared). We used the May 1979 CPS data, with the same control variables and groups that were used in table 1. The estimated coefficients of the interaction terms were sometimes nontrivial, although as often as not they were statistically insignificant.

The results with company size interacted with tenure, tenure squared, experience, and experience squared are representative of the specifications with the largest size-tenure interaction. A two-standard-deviation difference in company size increased log-wage growth per year of tenure by .003, .004, and .012 for the total, white-collar, and nonunion blue-collar samples, and it had no effect for union blue-collar workers. These compare with "average" log-wage growth per year of tenure of .013, .015, .015, and .004, respectively. However, new workers (those with zero tenure) still receive higher wages if they work for larger firms or establishments, and these differentials are very similar in magnitude to those reported in table 1.

The previously discussed finding that in AWS, PATC, and the Hay Associates data the size premium is larger in the lower grades of (white-collar) occupations than in the higher grades is hard to square with the idea that larger employers offer steeper profiles. Given the mixed results in previous studies (Oi and Raisian 1985; Pearce 1985), it would be fair to conclude that if large employers do offer steeper wage profiles, the difference is probably not very large.[18]

Evidence on the Union Avoidance Explanation

Since employer size is related to higher wages for *union* workers (table 1), union avoidance efforts cannot be the *only* reason for size-wage differentials. In table 5, we attempt to determine how important they are for understanding the size-wage relationship for nonunion workers. The first row shows by-now-familiar establishment size– and company size–wage effects for nonunion workers. The next four rows report analogous coefficients for four groups of workers for whom the threat of unionization is minimal. If union avoidance efforts are an important part of the size-wage relationship, that relationship should be much weaker for workers who seem very unlikely to seek unions. We find, however, that the size-wage relationship for these workers is about as strong as that for all nonunion workers. However important union avoidance efforts may be, they are not an important part of the size-wage story.[19]

[18] A more complicated explanation for lower quit rates among those working for larger employers is that those employers' training is more firm-specific. If small employers offer more general training while large ones offer more specific training, it is possible for the two types of employers' wage-tenure profiles to have similar slopes, but the gap between wages and alternative wages to be growing faster in large firms. We cannot test this hypothesis with the data used in this study. It is worth noting, however, that by itself the hypothesis does not explain why those working for large employers earn more *initially*.

[19] That managers of larger firms have higher earnings is not surprising given the literature that asks whether managers' salaries depend on sales or profits, especially since an important challenge of this literature is to deal with the high correlation between these variables (see Ciscel and Carroll 1980). Notice, however, that we use a less restrictive definition of manager than these studies (which focus on executives) tend to use.

TABLE 5

ESTIMATES OF THE SIZE-WAGE EFFECT ACROSS EMPLOYEE GROUPS WITH VERY LOW RATES OF UNIONIZATION USING MAY 1979 CPS DATA

Group	Sample Size	Coefficient of Size Variables* E	Coefficient of Size Variables* C
All nonunion private nonfarm wage and salary workers	10,753	.019 (.003)	.010 (.002)
Managers and administrators	1,317	.033 (.008)	.011 (.005)
Professional, technical, and kindred workers	1,576	.022 (.008)	.010 (.006)
Nonunion workers in occupations with union membership percentage ≤ 5%†	2,845	.020 (.006)	.009 (.004)
Nonunion workers in industries with union membership percentage ≤ 5%†	2,722	.026 (.006)	.010 (.004)

NOTE.—The list of independent variables used in each model is the same as the one used for row 1*b* of table 1.
* Standard errors are in parentheses.
† The union membership percentages used are from Freeman and Medoff (1979).

Evidence on the Product Market Power Explanation

Previous studies have shown that the size-wage relationship survives even when more direct measures of market power (concentration ratios [Weiss 1966; Mellow 1982] or industry profits [Pugel 1980]) are held constant. One might still wonder whether these are ideal measures of market power. As long as one accepts the premise of these earlier studies—that market power depends on *industry* characteristics—industry dummies are a reasonable way of avoiding controversies about the correct characteristic(s) to hold constant. Table 3, however, shows that industry dummies down to the three-digit level of detail have no effect on the size-wage relationship.

It is, of course, possible that the products of larger employers are sufficiently differentiated from those of smaller firms *in the same industry* that the larger producers have less elastic product demand curves and, hence, greater potential profits to share with their workers. To test this conjecture, we analyzed data from the MWES, which asked employers to estimate how their sales would respond to a 10 percent increase in the price of their product, with their competitors' prices held constant. On average, their estimates implied a demand elasticity of -2.3. With two-digit industry controlled for, the demand was less elastic for multiestablishment firms (by .66, with a standard error of .18) than for single-establishment firms, but larger establishment size was associated with *more* elastic demand: $\partial\eta/\partial$ ln(establishment size) = $-.31$; standard error = .06. Measuring size by establishment size alone also suggested more elastic demand at larger establishments (coefficient of $-.24$, with a standard error of .06). Thus there is little support for the hypothesis that (within two-digit industries) larger employers face less elastic product demands. In any case, adding the estimated demand elasticity had little effect on the size coefficients in table 1 because its effect on wages was small.

A very different way of looking at the ability to pay explanation is to study wage rates of local government employees because for local governments, credible exogenous measures of ability to pay (income or wealth per capita) are available. In a companion paper (Brown and Medoff 1988), we find that controlling for these measures of ability to pay has little effect on the positive relationship between size of local-government employers and the wages they pay.

III. Additional Explanations

A good deal of attention has recently been devoted to formal modeling of employers' strategies for recruiting workers and for monitoring and motivating those who are hired. It is not surprising that these models have been used to explain the relationship between employer size and wage rates.

Labor Pools, Worker Selection, and Employer Size

Weiss and Landau (1984) focus on recruitment and selection strategies that minimize the per unit cost of labor and how these differ for employers that differ in the number of units they employ. Each employer chooses *a* wage rate to offer to *all* the workers (in an occupation) it wishes to hire and a minimum level of worker quality. The wage offer determines the quality of the best worker it can hope to attract. The wage offer and the hiring standard must be chosen jointly to minimize labor costs while obtaining the desired number of units of labor.

The key assumption of Weiss and Landau's model is that, as the number of units of labor to be employed increases, the size of the available labor pool does not increase in the same proportion, so the number of applicants per vacancy falls. Consequently, at any given minimum qualification level, the larger employer will be forced to pay higher wages in order to satisfy the greater labor input requirement.

If the distribution of worker quality in the firm's area satisfies certain conditions, this mechanism produces a positive relationship between employer size and wage rates. When positive hiring costs are introduced, the model becomes *very* complicated, although Weiss and

Landau demonstrate a tendency for wages to fall initially and then rise with employer size.

As Weiss and Landau note, their model explains the general tendency for wages to rise with employer size.[20] If hiring costs per worker are more important at higher skill levels, the relationship between employer size and wages is likely to be weaker at high skill levels, which is consistent with the AWS, PATC, and Hay data presented in Section II. While it is most natural to think of *establishment* size as the relevant variable in their model, they argue that it can explain at least some positive company size effects as well.

The model is too complicated to have derivable predictions about the relationship between employer size and quality of worker hired. Thus it is consistent with the positive size-quality relationship we reported earlier, but it would be as consistent with the opposite result.

Our reading of the Weiss-Landau model is that it predicts that wages will rise with employer size, eventually, but for reasonably small employers this relationship will be ambiguous. Since "size" here is measured relative to the relevant hiring pool, we expect a weaker relationship between size and wages when the employer is very small relative to that hiring pool. Thus we expect the size-wage relationship to be weaker in metropolitan areas or in occupations with national hiring markets. We find little evidence of such patterns.[21]

We also tried to investigate the key premise of the model directly, by analyzing data from the MWES. Employers with minimum-wage workers were asked, "If you were to have an opening of a minimum-wage job now, how many qualified applicants would you get?" Whether or not we controlled for a limited set of demographic characteristics of the workers in such jobs or for characteristics of the job such as length of workweek and turnover rates, the elasticity of applicants per vacancy with respect to establishment size is positive, "small" but statistically significant (typically about .11 with a *t*-ratio of 3). Establishments of given size that are part of larger firms also get about 10 percent more applicants per vacancy.[22]

Holzer, Katz, and Krueger (1988) report an elasticity of (actual) applicants per opening for the last position filled with respect to establishment size of .16 and mixed results for the elasticity with respect to firm size, holding the wage constant in a broader sample of workers from the Employment Opportunity Pilot Project (EOPP) data. Hollenbeck and Mahle (1985, table 2.3) report an elasticity of applicants with respect to employer size of about .25, but with weaker controls for offered wages.

These relatively consistent findings[23] can be interpreted in several ways. First, one can accept them at face value and conclude that larger employers have more applicants per vacancy than smaller ones. Second, they might reflect the influence of unmeasured fringes or working conditions (better at larger employers), though this seems less plausible in the MWES data, which apply to minimum-wage jobs and control explicitly for turnover rates. Third, one can question how successfully the elasticities that are estimated correspond to the relationship that is at issue. If larger employers have several vacancies at the same time, the number of applicants for a hypothetical vacancy (MWES) may overstate the number of applicants *per vacancy*. Similarly, if (as Holzer et al. suggest) large employers are more likely to reconsider applicants for previous openings or convert casual inquiries into formal applications, the EOPP applicant per vacancy measure may be inflated upward for these employers.

Monitoring

An alternative approach to explaining the size-wage relationship used by some authors is based on the premise that larger employers have

[20] Weiss and Landau find some evidence in previous work of a flat or even downward-sloping size-wage relationship among relatively small establishments, and (as noted above) their model with positive hiring costs can generate this result. We investigated this possibility with the EEEC and WDS data, which provide continuous rather than categorical measures of establishment size. We allowed the coefficient of ln(employment size) to take on a different value at high than at low values of ln(establishment size), with the two segments joined at either 25 or 100 workers. There was no evidence that the additional term mattered in any consistent way.

[21] When we added an interaction between ln(establishment size) and metropolitan area to the CPS, EEEC, and WDS equations in table 1, there was very little difference in establishment size effects in metropolitan and nonmetropolitan areas. The interaction term was typically right-signed (i.e., negative) but statistically insignificant and a third or less (typically much less) of the ln(establishment) coefficient. As regards workers in national markets, table 5 provides no evidence of smaller size premia for professional, technical, and kindred workers. Even among these workers, however, there are some (especially technicians) whose labor markets may be more local than national. We therefore examined the effect of excluding precollege teachers, technicians, and similar occupations. This reduced the establishment size coefficient to .007 (.009) but increased the company size coefficient to .017 (.008); so their sum, .024, was very close to the table 5 value. With establishment size as the only size measure, its coefficient was .021 (.007).

[22] The fact that the question referred to *qualified* applicants may introduce an element of ambiguity to the results if larger employers respond to the problem of attracting workers by lowering qualifications. We regard this as a minor problem for two reasons. First, the average number of "qualified" applicants per vacancy is five; so respondents obviously used "qualified" loosely, rather than using it to refer to the quality of the one applicant they typically hire. Second, when we added a very limited set of worker characteristics to control for "qualifications," the results did not change.

[23] These results are consistent with Ochs's (1984) experimental study of buyers (≈ job searchers) when the amounts of merchandise available for sale (≈ job vacancies) are known to vary across locations. He finds that buyers choose locations so that, if anything, the buyer/merchandise (≈ applicant/vacancy) ratio is greatest at the location with the most stock.

more difficulty monitoring workers. Unlike Oi's (1983) paper, in which larger employers choose high-quality workers to economize on fixed per worker monitoring, these papers argue that large firms' disadvantage in monitoring leads them to monitor less closely. As a result, they are less able to detect the subtler aspects of worker quality (such as effort) and they pay more for workers of given quality.

It is perhaps ironic that a discussion of "recent" work on the subject should begin with Stigler's (1962) classic paper:[24]

> Wage rates and skilled search are substitutes for the employer: the more efficiently he detects workers of superior quality the less he need pay for such quality.
>
> The small company has distinct advantages in the hiring process, so far as judging the quality of workers is concerned. The employer can directly observe the performance of the new worker and need not resort to expensive and uncertain rating practices to estimate the workers' performance. It is well known that wage rates are less in small plants than in large, and the difference reflects at least in part (and perhaps in whole) the lower costs to the small-scale employer of judging quality. A similar result [negative correlation with firm size] obtains with respect to dispersion of wages . . . . Men should in general enter smaller companies the greater their ability. [Pp. 102–3]

Garen (1985) presents a more formal version of Stigler's model, in which large firms' disadvantage in monitoring leads to different offered-wage schedules, and workers' choice of employers takes this difference into account.

In order to evaluate Stigler's model, we investigated whether it is fully consistent with the results reported in tables 1 and 2 and whether its wage structure predictions are accurate. Because diseconomies in monitoring cannot be measured directly, the empirical tests are necessarily less direct than those used for other hypotheses.

We showed in table 1 that both establishment and company size have independent effects on wages. In the quotation above, Stigler shifts from "company" to "plant" (i.e., establishment) and back to "company," without clearly distinguishing between them. If we restrict attention to single-establishment firms, larger size is associated with greater monitoring difficulties. But when company and establishment size are not the same, the implications of the monitoring model are less clear, particularly for the effect of establishment size on wages.

[24] Although Stigler refers to small companies' advantage in the "hiring" process, his analysis really deals with the greater ability to monitor those who already have been hired so that the best workers can be rewarded and, hence, retained. Indeed, large firms have obvious scale advantages in hiring (Hamermesh 1980, p. 387), a larger sample of observations for detecting the relationship between worker characteristics and productivity, and economies of scale in studying such relationships.

In order to explain the observed *partial* effect of establishment size on wages, one would have to argue that, if we hold firm size constant but increase establishment size (say, by consolidating the work force into fewer establishments), monitoring of workers has become more difficult. However, it might well be *easier* to monitor 1,000 workers in one location than 1,000 workers spread across 10 100-worker locations. Stafford (1980, p. 340) has suggested an alternative possibility: if larger establishments have larger work groups, determining the productivity of individual workers may be more difficult in larger establishments. Alternatively, it may be that it is really "profit center" size that matters, which might mean that both firm and establishment size are associated with monitoring difficulties.

Stigler also noted that small employers' greater ability to judge worker quality should lead higher-ability workers to select such employers. This argument does not necessarily apply to measures of ability such as years of schooling or years of experience, which are easily observed by employers of all sizes.[25] Rather, it refers to subtler abilities that require careful observation (monitoring) to detect. Garen finds that proxies for intelligence are more highly rewarded by smaller employers. However, in table 2 we showed that the coefficient of employer size fell when we moved from OLS to the fixed-effect estimator. Thus those working for small employers appear to have *fewer* of the subtler virtues that are not captured by the readily observed variables in the OLS equation but implicitly held constant in the fixed-effect model.[26]

The prediction that monitoring difficulties will lead to a relationship between employer size and wage *structure* appears to have received little subsequent attention because of the difficulty of obtaining the necessary data. One neglected source of such data is the Industry Wage Survey (IWS), conducted by the Bureau of Labor Statistics. The

[25] Garen (1985) argues that the relationship between employer size and the return to schooling is theoretically ambiguous, and he finds that large employers reward extra years of schooling less than small employers do, although this difference is not statistically significant. Stolzenberg (1978) found significantly higher returns to years of schooling in large firms.

[26] One might argue that Stigler's hypothesis holds only after one takes account of differences in "skill requirements" of jobs in large and small firms. But this explanation requires differences in such requirements *within* two-digit occupations, given the occupation dummies in our fixed-effect models.

IWS, which surveys individual establishments in selected industries,[27] collects the usual data on establishment characteristics (industry, location, unionization, and employment) and considerable detail on wage structure for production workers. For example, establishments classify the method(s) of pay they use. Important for our purposes, the IWS distinguishes time-rated from incentive pay systems and, for time-rated systems, permits us to distinguish those in which pay is merit related from standard rate systems in which wages depend only on one's job (and perhaps seniority).[28] The main disadvantage of the IWS is that firm size is not recorded.

Stigler's argument is based on the premise that larger employers have greater difficulty monitoring the performance of their workers through *judgmental* rating schemes. One would expect that those judgmental rating schemes would receive less weight—in the limit, negligible weight by avoiding them altogether—in salary setting by large employers. Such employers should also be more willing to undertake the (more costly but more accurate) nonjudgmental evaluation implicit in a piece-rate system.[29] The evidence in the first two rows of table 6 is quite consistent with these predictions: larger establishments are significantly more likely to use both standard rates and incentive pay (and less likely to use merit pay systems).[30]

More generally, if a larger employer's estimate of the productivity of a given worker is less reliable than that of a smaller employer, the larger employer should place less weight on that estimate, and its wage distribution should be relatively compressed (Garen [1985]; see Aigner and Cain [1977] for a similar result in the statistical discrimination literature). Our examination of this prediction, using the two establishment level data sets in table 1, which both have information on the establishment's wage distribution and identify firm size (or, in the case of MWES, multiestablishment firms), is presented in rows 3

[27] Our IWS sample is a set of 10 manufacturing industries previously analyzed by Freeman and Medoff (1984). Their selection criterion was that industries have sizable union and nonunion sectors, which should not impart any particular bias for our purposes.

[28] The IWS has nine relevant method of pay categories (plus commissions, which is not relevant to our sample). Of these, four are types of incentive pay, which we grouped together. The five time-rated categories are single rates (paying everyone in a job the same wage); "range of rates" systems, with progression through the range based on seniority, merit, or a combination of seniority and merit; and individual determination. We combined "single rates" and "range of rates: seniority" to form our "standard rate" category.

[29] For a similar argument, see Goldin (1986) or Lazear (1986).

[30] Larger establishments probably also have lower per worker costs of setting up (and updating) piece-rate systems (International Labor Office 1984), so the latter result by itself could simply reflect that advantage rather than a disadvantage in using judgmental schemes.

TABLE 6

EMPLOYER SIZE AND WAGE STRUCTURE

Data Set (Sample Size)	Dependent Variable	Other Independent Variables	Size Variable	Coefficient of Size Variable*
1. IWS (3,216)	Proportion of production workers paid standard rates†	Union coverage, sex, SMSA, region (3), industry (21), wage-weighted occupation index§	E	.022 (.007)
2. IWS (3,216)	Proportion of production workers receiving incentive pay†	Same	E	.025 (.004)
3*a*. WDS (1,355)‡	Standard deviation of ln(hourly earnings)	Union coverage, sex (2), age (4), SMSA, region (3), industry (40), pay type (2), average production workweek	E	.021 (.003)
3*b*. Same	Same	Same	E	.032 (.004)
			C	−.011 (.002)
4*a*. MWES (978)‡	Same	Union coverage, region (3), industry (60)	E	−.003 (.004)
4*b*. Same	Same	Same	E	.005 (.004)
			M	−.073 (.012)
5*a*. IWS (3,185)	Same	Union coverage, sex, SMSA, region (3), industry (21)	E	.002 (.001)
5*b*. Same	Same	Same as 5*a* plus s(occ)‖	E	−.005 (.001)

* Standard errors are in parentheses.

† "Standard rates" includes single-rate systems and range-of-rates systems in which progression through the range is based on seniority. "Incentive pay" includes individual incentive pay, individual bonus pay (incentive pay beyond some target level of output), group incentive pay, and group bonus pay.

‡ These sample sizes are smaller than those in table 1 because analysis of wage dispersion requires deletion of establishments with only one worker.

§ This is equal to ln(wage) for the establishment if it paid each worker the industry average wage for that worker's occupation.

‖ s(occ) = standard deviation of ln wage for the establishment if it paid each worker the industry average wage for that worker's occupation.

and 4 of table 6. If employer size is measured by establishment size alone, its effect on the standard deviation of ln(wage) is wrong-signed (positive), though not statistically significant in row 4*a*. If we try to distinguish establishment from firm size effects (or presence of multiple establishment effects), the results remain anomalous. Greater establishment size is associated with *greater* wage dispersion; larger firm size reduces wage dispersion. While the former result is not entirely inconsistent with a monitoring story (we noted the ambiguity of the partial effect of establishment size on monitoring difficulties earlier), it is hard to reconcile opposite-signed effects on wage dispersion with similar-signed effects on wage levels. Moreover, the point estimates in row 3*b* imply that if we increase firm size by making each establishment larger (keeping the number of establishments fixed), the effect (measured by the sum of the E and C coefficients) is to increase wage dispersion, which *is* inconsistent with Stigler's analysis.

One possible explanation for these essentially negative results is that larger establishments have a wider range of jobs in them. If this is true and the larger between-job wage dispersion dominates the (hypothesized) smaller within-job dispersion, the problem may lie in our inability to separate the two types of variation.

While we cannot pursue this with either the WDS or the MWES (which have no occupational detail), we can do so with the IWS (rows 5*a* and 5*b*). In row 5*a*, we show the coefficient of ln(establishment size) in a regression, similar to those in rows 3*a* and 4*a*, to explain the standard deviation of ln(wage).[31] As was true in the other data sets, larger establishments have greater wage dispersion. However, in row 5*b* we add as a control variable the standard deviation of ln(wage) that the establishment would have if it had its actual occupational distribution but paid each worker the mean wage for his or her occupation. The added variable controls for the impact of varying occupational distributions on wage dispersion, and with this refinement the coefficient of E is now negative and statistically significant. Whether controlling for occupational distribution would improve the results when both E and C are used to measure employer size cannot be determined, however, since the IWS does not include a company size measure.

An alternative class of monitoring models posits a different relationship between employer size and monitoring and wages. Faced with an imperfect ability to monitor the effort of workers, firms may offer a wage that exceeds that needed to attract the desired quality of workers. In doing so, they provide a greater penalty to any worker discharged for shirking: severity of punishment substitutes for certainty of detection. While the models surveyed by Yellen (1984, pp. 201–3) do not appear to distinguish types of firms, Bulow and Summers (1986) have recently developed a model with two groups of firms that differ in their monitoring costs. The most obvious predictions—higher wages where monitoring is most difficult and, because wages exceed reservation wages, lower quit rates—are consistent with the evidence on different-sized firms. However, the usual reasons for using above-market wages instead of steep wage-tenure profiles to deter shirking—workers' inability to borrow during the initial period of underpayment and lack of "reputation" to deter *firm* shirking—are perhaps least persuasive for large firms.[32] Clague's (1977) argument that large firms pay above-market wages to create loyalty and esprit de corps as a substitute for monitoring is less vulnerable to this objection.

Common to all these explanations, in which large employers pay higher wages because they have difficulty monitoring workers, is the idea that where monitoring is not difficult (or where large employers are at no disadvantage) there should be no size-wage premium. It seems plausible to us that, *among piece-rate workers*, large employers face no such disadvantage. We pursued this idea using the IWS data. We regressed ln(wage) on the control variables in row 5*a* of table 6, the logarithm of establishment size (E), the proportion of the establishment's blue-collar work force paid standard rates (S) and incentive pay (I), and interactions of E with S and I. The following coefficients resulted: E, .034 (.004); S, .109 (.026); I, .031 (.055); $E \times S$, −.010 (.005); and $E \times I$, .015 (.011). Thus, even among piece-rate workers, those working for larger employers receive higher wages;[33] indeed, the size-wage premium is, if anything, larger for piece-rate workers than for standard-rate workers or those paid by the reference group "merit pay."

We have, in this section, investigated whether the size-wage link is due to monitoring problems experienced by larger employers. The evidence is not all in one direction, but on balance we are skeptical that monitoring is the correct explanation. In particular, the size-wage link among piece-rate workers leads us to question the role of monitoring difficulties in explaining the more general size-wage relationship.

[31] The IWS obtains wage rates for all workers in selected occupations rather than all workers. However, the "selected occupations" are chosen by industry and generally account for a majority of blue-collar employment in the industry.

[32] Recall that when we explored size-tenure interactions, we found that even for newly hired workers there is a sizable size-wage premium.

[33] The result in the text defines "incentive pay" as including both individual and group incentives. Restricting this category to the former does not change the qualitative results. However, the interaction between size and incentive pay is smaller, .006 (.012).

IV. Conclusion

We began by asking why large employers pay their workers more than small employers do. We do not have "the" answer to this question. We do have a much more detailed set of stylized facts than was previously available. Among these finer results are the following: (1) The effect of employer size on wages is *both* an establishment and a firm size effect. (2) Even for subsets of workers grouped by "collar color," union status, or industry, those who work for larger employers receive higher wages. (3) Within detailed professional, technical, and managerial occupations, employer size premia are smallest (in percentage terms) in the highest pay grades. (4) The employer size effects are not greatly reduced by looking at changes in wages for particular workers as they move to different-sized employers. (5) Differences in working conditions that we can measure seem not to explain much of the size-wage differential. Moreover, even when wages are held constant, worker attachment to large employers is greater than that to small employers. (6) The size premium occurs even in contexts in which the threat of unionization is implausible and in the union sector. Thus the *threat* of unionization to large employers does not explain the size premium. (7) The size-wage relationship is almost unaffected by controlling for two- or three-digit industry. Within industries, product market power (demand inelasticity) does not explain the size-wage premium. (8) Large employers are more likely to have single-rate wage policies within job categories for their blue-collar workers, and they are less likely to opt for a separately determined wage for each worker. (9) When establishment size is controlled for, wage dispersion is smaller for larger firms; wage dispersion does not seem to decline with establishment size, apparently because between-occupation dispersion is greater. (10) Even among piece-rate workers, larger employers pay higher wages.

Our bottom line is that the size-wage differential appears to be both sizable and omnipresent; our analysis leaves us uncomfortably unable to explain it, or at least the part of it that is not explained by observable indicators of labor quality.

In lieu of a more positive conclusion, we offer two observations that might contribute to an explanation. First, large employers pay more for their labor but less for their other inputs because of lower interest rates on borrowed funds and quantity discounts on purchased inputs. This may well explain how large employers survive despite paying higher wages, but it does not explain why they offer higher wages in the first place. Second, large firms are also older firms. Is it possible that the size-wage premium is really a relationship between employer age and wages? Do firms that treat their employees well live longer, or is it the other way around? Thus the employer size–wage effect remains a fact in need of an empirically based theory. It is hoped that the facts presented in this paper will guide that search.

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