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Testing for employer monopsony in turn-of-the-century coal mining

William M. Boal*

Isolated company towns are often cited as likely examples of labor monopsony. This article tests for monopsony power by estimating inverse labor supply elasticities using a county-level panel dataset on nonunion West Virginia coal mining from 1897 to 1932. The model specification incorporates dynamics in such a way that an estimate of the gap between marginal revenue product and the wage can easily be computed as a weighted average of short- and long-run inverse elasticities. Modest estimated short-run inverse elasticities and very small long-run inverse elasticities imply that coal operators enjoyed little, if any, monopsony power over their workers.

1. Introduction

■ The classic textbook example of labor monopsony is the isolated mining town in the nineteenth and early twentieth centuries (Ehrenberg and Smith, 1994; Fleisher and Kniesner, 1984; Hamermesh and Rees, 1992; Kaufman, 1993; McConnell and Brue, 1989). West Virginia coal mining at the turn of the century probably fits this archetype as well as any industry. Most mines were located in remote, sparsely populated areas, so miners wishing to change employers had to relocate. Roughly 80% of miners lived in company housing in company towns (see Table 1). Company stores and the issuance of scrip-money by employers seem to have been much more common in West Virginia than in the rest of the coal industry (Fishback, 1986, 1992b; Timberlake, 1987; U.S. Coal Commission, 1925). Transportation was costly due to rough terrain, few roads, and essentially no navigable rivers except the Ohio and the lower Kanawha. (Even today, the cost of building roads in West Virginia is several times the U.S. national average (DeParle, 1991).)

Despite the remote location of West Virginia mines, Table 1 shows that turnover rates there were actually higher than in the rest of the coal industry, whose rates were comparable to other industries with establishments of similar size (U.S. Coal Commission, 1925). Turnover rates in West Virginia coal were several times higher than recent turnover rates in health care and old-age homes, but similar to recent turnover rates in restaurants, which various authors have characterized as monopsonized.[1] Whether high turnover is evidence against monopsony may depend on its nature.

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[1] Turnover rates in health care, old-age homes in England, and restaurants are given respectively by Miner (1977), Machin, Manning, and Woodland (1993), and Cohen and Schwartz (1980).

TABLE 1 Company Housing and Labor Turnover in Nonunion Bituminous Coal Fields

State(s)	Company Housing Rate (%)	Field	Labor Turnover Rate (%)
West Virginia[a]	78.8	Logan	230.0
		Kenova-Thacker	234.0
		Pocahontas-Tug River	196.0
		New River-Winding Gulf[b]	148.0
Pennsylvania[c]	50.7	Somerset	121.5
		Westmoreland	81.5
		Connellsville	152.0
Maryland, Virginia, Kentucky, and Tennessee[d]	64.4	Cumberland-Piedmont[e]	77.0
		Virginia	198.0
		Northeast Kentucky	212.0
		Southeast Kentucky	198.0
		Tennessee[f]	122.0
Alabama	65.9	Alabama-Georgia	80.0
Western states[g]	54.6	Utah-New Mexico	138.0
		Colorado	125.0

Source: Company housing rate is percent of mine workers housed in company-owned family dwellings, 1922–1923, from U.S. Coal Commission (1925) Volume III, p. 1467, Table 14. Labor turnover rate is separations as percent of average number on roll, 1921, from same source, p. 1264, Table 1.

[a] Includes unionized Kanawha, Panhandle, and Fairmont fields.

[b] New River portion is union and Winding Gulf portion is nonunion.

[c] Includes unionized Central Pennsylvania and Pittsburgh fields.

[d] Includes unionized Western Kentucky field and mixed Tennessee field.

[e] Field is partly in West Virginia.

[f] Field is mixed union and nonunion.

[g] Includes union fields in Montana, North Dakota, Oklahoma, Arkansas, Texas, Wyoming, and Washington.

Turnover in restaurants seems to be mostly movement in and out of employment by secondary wage earners (Alpert, 1986). In contrast, turnover in coal seems to have been mostly movement between employers, in response to wage and nonpecuniary differentials, by primary wage earners (Corbin, 1981). Miners and their families relocated frequently, apparently moving by train along the same rail lines built to carry the coal itself from each mine to distant markets. In any case, the turnover data suggest that the case for monopsony is not closed and that an empirical test is in order.

The possible existence of employer monopsony has important welfare implications for policy. Wage-raising institutions that might have a distortionary effect in a competitive labor market, such as labor unions and minimum wage laws, can increase welfare and employment in a monopsonized market by imposing a fixed wage and thereby lowering marginal factor cost. The source of monopsony power is important for both policy and empirical analysis. Following Robinson (1969), it is useful to distinguish between monopsony arising because employers are "acting in concert," i.e., from collusion, and monopsony arising because "the supply of labor to each firm is less than perfectly elastic," i.e., from differentiation. A large literature addresses collusion in such fields as nursing and professional sports. However, the sheer number of coal mining companies in West Virginia during this period makes collusion unlikely. For example, in 1925 there were 793 coal companies operating 1,208 mines, and the largest four companies together produced only 10% of the total state coal output (West Virginia Department of Mines, 1925). Although operators' associations existed, they were apparently formed for the purpose of fighting or negotiating with the union

(Corbin, 1981) and in any case had difficulty controlling their members (Fishback, 1992b). Moreover, entry was fairly easy (Tams, 1963).

Instead, this article tests for differentiation across firms. In doing so, it exploits the pattern of coal employment growth in this period. The West Virginia coal industry developed rapidly during the late nineteenth and early twentieth centuries, apparently constrained only by the pace of geological discovery and railroad construction (see Figure 1).[2] Growth was not uniform across the state, however. Early development along navigable rivers (in the Panhandle and Kanawha fields) was followed by development in remote areas accessible only after the construction of railroads (in the Fairmont, New River, Winding Gulf, Logan, Mingo, Pocahontas, and Greenbrier fields).[3] The industry thus experienced large shifts in both aggregate levels and geographical distribution of employment. Figure 2 illustrates this uneven growth in employment for the five largest counties (by employment) from 1897 to 1932. This article may be interpreted as investigating the size of the wage changes required to generate these employment shifts. If large wage changes were required to generate these shifts, employers may have had significant monopsony power from differentiation.

The literature on labor monopsony from differentiation is smaller than the literature on collusion. Early econometric studies of differentiated employer monopsony power tend to rely on cross-sectional data. Bunting (1962) develops measures of employer concentration at the county level using data for the entire United States in 1948. Bunting finds a positive (though insignificant) correlation between wages and concentration in the southern textile industry, thus appearing to refute the possibility of monopsony in this case. Bunting makes no attempt to estimate structural parameters of labor supply or firm behavior. Nelson (1973) develops a spatial model in which workers' transportation costs endow employers with market power. Fitting this model to Census data produces very high elasticities of labor supply for most of the United States. Machin, Manning, and Woodland (1993) use data on flows of hires and separations, as well as the employment size-wage relationship across firms, to infer an upward-sloping supply of labor to individual homes for the aged in England.

Cross-sectional studies suffer from some inherent limitations. They cannot identify how workers actually respond to wage differentials except by assuming special functional forms for labor supply (Nelson, 1973; Machin, Manning, and Woodland, 1993). Furthermore, wage levels alone cannot measure labor market imperfection unless one controls for other, possibly unobserved influences on labor supply. For example, an employer operating in a remote location might well behave like a monopsonist but still pay above-average wages due to the disamenities of the location of employment.

A panel study can address these problems more easily. Multiple observations on the same employer allow estimation of dynamic labor supply while controlling for employer-specific fixed effects. D. Sullivan (1989) uses panel data to analyze the supply of nurses to individual hospitals. He estimates inverse elasticities for various time intervals, under several alternative assumptions about the interaction between nearby employers. His estimates suggest that hospitals have considerable short-run market power in nursing.

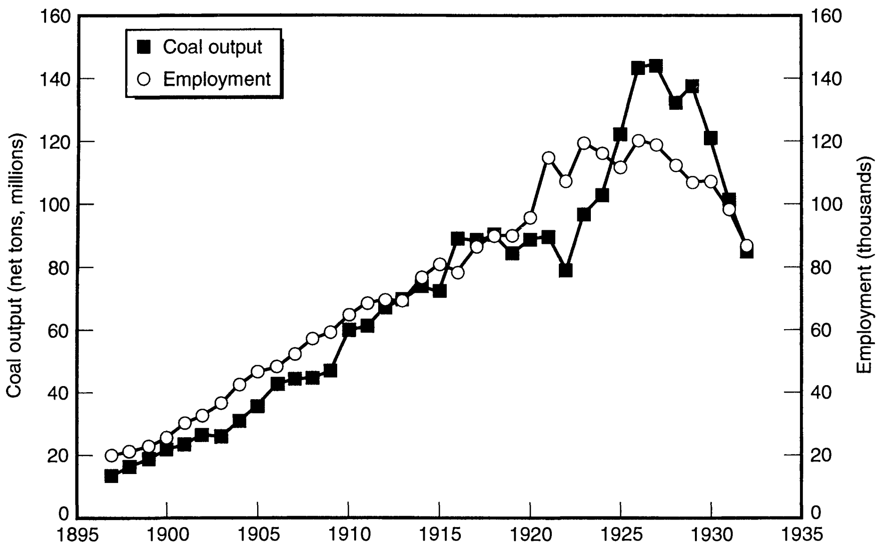
In this article I use panel data to test for the existence of employer monopsony power in West Virginia coal mining during the early twentieth century. I do not test whether that market power was actually exercised. In particular, I do not estimate the actual gap between marginal revenue product and wages but only the potential gap,

[2] Figure 1 shows little if any increase in output per worker over time. However, days of operation per year declined throughout this period (Boal, Fishback, and Kantor, 1994). Output per worker-day therefore shows an upward trend, even for pick mining (Dix, 1977).

[3] See Massay (1970) on early development of the Fairmont field. See C.K. Sullivan (1989) and Tams (1963) on the New River, Winding Gulf, and Pocahontas fields.

FIGURE 1

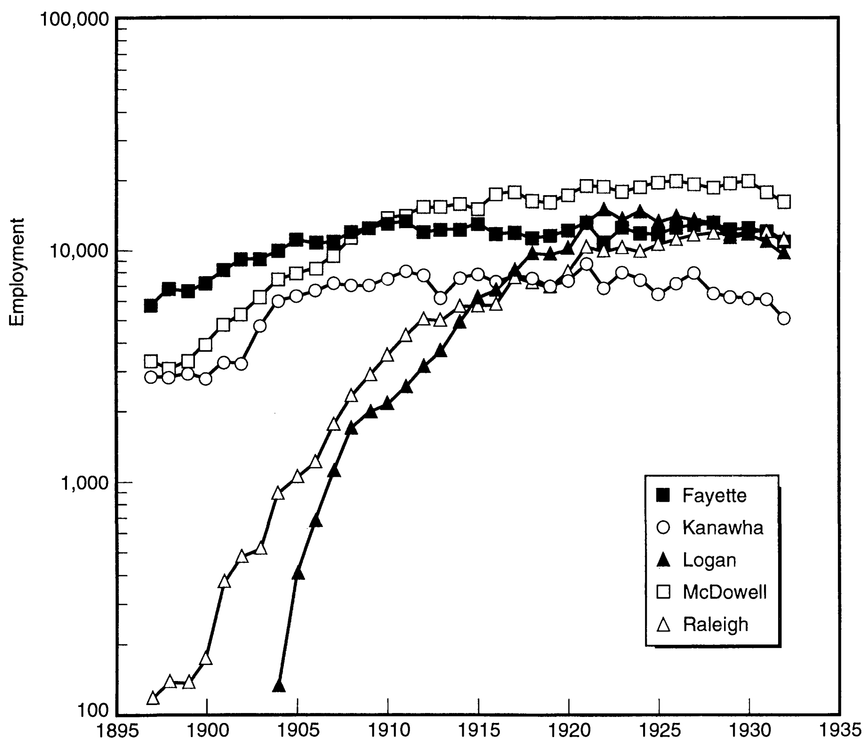
COAL OUTPUT AND EMPLOYMENT IN WEST VIRGINIA, 1897–1932



SOURCE: West Virginia Department of Mines.

FIGURE 2

EMPLOYMENT IN COAL MINING IN FIVE WEST VIRGINIA COUNTIES, 1897–1932



SOURCE: West Virginia Department of Mines.

conditioned on the equilibrium concept for employer interaction and measured by the perceived inverse elasticity of labor supply to the firm. The present analysis thus uses a methodology similar to Baker and Bresnahan's (1985, 1988) analysis of product market power in the brewing industry, Scheffman and Spiller's (1987) analysis of product market definition in gasoline refining, and D. Sullivan's (1989) analysis of employer market power in nursing. The present article contributes an improved dynamic specification to this methodology. In particular, it shows that a partial adjustment model can yield estimates of the Lerner index of market power, given the solution concept and the firm's discount rate. The substantive contribution of the article is to show that for West Virginia coal, the gap between wages and marginal product would have been very small under either Bertrand or Cournot open-loop equilibria, assuming firms discounted the future at least somewhat. This is so because while short-run inverse elasticities of supply were positive, long-run inverse elasticities were essentially zero.

The article is organized as follows. In Section 2 I present a tractable dynamic model of labor supply from which a simple expression for the gap between wages and marginal revenue product is derived under alternative assumptions. In Section 3 I present the data and the estimates. Section 4 concludes.

2. Model specification

■ Any investigation of employer monopsony power should address two key factors on which such power must depend, namely "on the ease with which men can move, and on the extent to which they and their employers consider the future, or look only to the moment" (Hicks, 1935). Put differently, it should address cross-elasticities and dynamics of labor supply to the individual firm, as well as the dynamics of firm optimization and strategic interaction.

□ **Labor supply.** Consider the supply of labor to two firms. To address "the ease with which men can move," assume that the two firms are differentiated by locational amenities or working conditions so that own- and cross-elasticities of supply may be finite (Robinson, 1969). Assume that workers do not respond instantaneously to changes in relative wages due to, say, mobility costs, and that the dynamics of response can be usefully approximated by a simple, tractable partial-adjustment model. Let L_t^i and W_t^i represent the logarithms of labor supply and wages, respectively, at firm i in period t. Let workers' desired (log) labor supply to each firm i at time t be given by the following:

$$L_t^{i*} = \alpha_0^i + \alpha_1^i W_t^i + \alpha_2^i W_t^j, \qquad \alpha_1^i > 0, \qquad \alpha_2^i < 0, \tag{1}$$

where $i = 1, 2$ and j indexes the other firm. Assume that workers adjust their labor supply only partially from the level of the previous period to the current desired level:

$$L_t^i - L_{t-1}^i = \delta^i(L_t^{i*} - L_{t-1}^i), \qquad 0 < \delta^i < 1. \tag{2}$$

Substituting (1) into (2) gives

$$L_t^i = \delta^i(\alpha_0^i + \alpha_1^i W_t^i + \alpha_2^i W_t^j) + (1 - \delta^i)L_{t-1}^i. \tag{3}$$

Moving the firm's own wage to the left-hand side gives

$$\begin{gathered} \mathrm{W}_t^i = \beta_1^i + \beta_2^i L_t^i + \beta_3^i L_{t-1}^i + \beta_4^i W_t^j, \\ \beta_2^i > 0, \qquad \beta_3^i < 0, \qquad \beta_4^i > 0, \end{gathered} \tag{4}$$

where the β parameters are functions of the α and δ parameters of (1) and (2).[4]

The parameters β_2^i and β_3^i measure monopsony power. The parameter β_2^i is the inverse short-run elasticity of labor supply with respect to firm i's own wage, given the wage chosen by the other firm. If β_2^i is zero, labor supply is perfectly elastic in the short run and the firm has no short-run monopsony power. The inverse long-run elasticity of labor supply also takes a simple form. In this partial-adjustment framework, current labor supply L_t^i depends on an exponentially weighted moving average of current and past wages as shown by solving (4) for L_t^i and repeated substitution:

$$\begin{aligned} L_t^i &= \frac{1}{\beta_2^i} W_t^i - \frac{\beta_3^i}{\beta_2^i} L_{t-1}^i + \ldots \\ &= \frac{1}{\beta_2^i} W_t^i - \frac{\beta_3^i}{\beta_2^i}\left(\frac{1}{\beta_2^i} W_{t-1}^i - \frac{\beta_3^i}{\beta_2^i} L_{t-2}^i\right) + \ldots \\ &= \frac{1}{\beta_2^i} \sum_{s=0}^{\infty} \left(\frac{-\beta_3^i}{\beta_2^i}\right)^s W_{t-s}^i + \ldots, \end{aligned} \tag{5}$$

where "..." represents constants and terms in the current and lagged wages of the other firm. The n-period elasticity is thus

$$\frac{1}{\beta_2^i} \sum_{s=0}^{n} \left(\frac{-\beta_3^i}{\beta_2^i}\right)^s \tag{6}$$

and the n-period inverse elasticity is

$$\frac{\beta_2^i}{\sum_{s=0}^{n} \left(\frac{-\beta_3^i}{\beta_2^i}\right)^s}. \tag{7}$$

The limit of the series in the denominator is $\beta_2^i/(\beta_2^i + \beta_3^i)$, so the long-run inverse elasticity is just $(\beta_2^i + \beta_3^i)$, which equals $1/\alpha_1^i$, the inverse elasticity of desired labor supply with respect to firm i's own wage, again given the wage path of firm j. If the sum $(\beta_2^i + \beta_3^i)$ is zero, labor supply is perfectly elastic in the long run and the firm has no long-run monopsony power.

Alternatively, the wage of the other firm may be eliminated from each equation, replaced by the other firm's current and lagged employment. Simply substitute (4) for firm j into the same equation for firm i to get

$$\begin{aligned} &W_t^i = \beta_1^{i\prime} + \beta_2^{i\prime} L_t^i + \beta_3^{i\prime} L_{t-1}^i + \beta_4^{i\prime} L_t^j + \beta_5^{i\prime} L_{t-1}^j, \\ &\beta_2^{i\prime} > 0, \qquad \beta_3^{i\prime} < 0, \qquad \beta_4^{i\prime} > 0, \qquad \beta_5^{i\prime} < 0, \end{aligned} \tag{8}$$

where the β' parameters are functions of the α parameters and the δ parameters of (1) and (2). Because the β' parameters are more numerous than the α and δ parameters on which they depend, nonlinear cross-equation restrictions must relate the β' parameters to each other, but these restrictions cannot be imposed if only the equation for one firm is to be estimated.[5] As before, the parameter $\beta_2^{i\prime}$ is the inverse short-run elasticity of

[4] In particular, $\beta_2^i = 1/(\delta^i \alpha_1^i)$ and $\beta_3^i = (\delta^i - 1)\beta_2^i$.

[5] Additional restrictions relate the parameters within each equation if $\delta^i = \delta^j$, i.e., equal speed of adjustment. This assumption will not be imposed here, however.

labor supply with respect to firm i's own wage, and the sum $(\beta_2^{i\prime} + \beta_3^{i\prime})$ is the inverse long-run elasticity of labor supply with respect to firm i's own wage, this time given the employment path chosen by the other firm. In general, these inverse elasticities will be larger than those in (4). In particular, $\beta_2^{i\prime} = \gamma\beta_2^i$ and $\beta_3^{i\prime} = \gamma\beta_3^i$, where

$$\gamma = \alpha_1^i\alpha_2^j/(\alpha_1^i\alpha_2^j - \alpha_2^i\alpha_1^j).$$

Note that γ is greater than one if own-elasticities exceed cross-elasticities in absolute magnitude, so in general the Cournot coefficients should exceed the Bertrand coefficients in absolute magnitude. However, the Bertrand and Cournot coefficients should match for two polar cases. On the one hand, if employers are completely undifferentiated and numerous, the inverse elasticities in (4) and (8) will both be zero. Wages will be determined completely by variables outside the individual firm. On the other hand, if employers are completely differentiated, the inverse elasticities in (4) and (8) will be identical and greater than zero. Coefficients corresponding to the other firm's wages or employment will be zero, because the first firm's wages will then be unaffected by the behavior of the other firm.

□ **Equilibrium.** It is worth emphasizing that (4) and (8) are equivalent descriptions of the same labor supply behavior. Under the basic assumptions in (1) and (2), either wage equation (4) or (8) should fit the data equally well, no matter how employing firms interact with each other, provided equilibrium occurs on the labor supply curve. On the other hand, whether the parameters of (4) or the parameters of (8) are more relevant in measuring monopsony power depends on the assumed equilibrium concept for employer interaction. If the firms play a Bertrand game, taking each other's wages as given, then each will maximize profits while perceiving its labor supply curve to be (4). If the firms play a Cournot game, taking each other's employment as given, then each will maximize profits while perceiving its labor supply curve to be (8).[6] This article will not attempt to choose between Bertrand and Cournot behavior, but it will estimate both (4) and (8). However, interpretation of the coefficients in these equations requires specifying how firms "consider the future." Several specifications suggest themselves.

The simplest specification is the myopic or one-shot equilibrium, where firms "look only to the moment." Under this specification, profit maximization implies that the percentage gap between marginal revenue product (MRP) and wages is just the perceived short-run inverse elasticity of labor supply: β_2^i for the Bertrand case and $\beta_2^{i\prime}$ for the Cournot. This specification is obviously unrealistic if labor supply adjusts slowly and firms have finite discount rates.

A more realistic but still tractable specification is the open-loop equilibrium, where each firm anticipates the effect of its current action on its future labor costs (but not the effect on future actions of the other firm). Suppressing firm superscripts for the moment, let l_t and w_t represent employment and wages in levels (not logs as before), let $R_t(l_t)$ be the firm's revenue function, and let r be its discount rate. Then the profit-maximizing firm maximizes

$$\sum_{t=0}^{\infty} \frac{R_t(l_t) - w_t l_t}{(1 + r)^t} \tag{9}$$

[6] A slight elaboration of the model, permitting the adjustment to depend on the difference between actual and desired labor supply at both firms, was also investigated. The results were indistinguishable from those reported below. (See Boal, 1993).

subject to the relevant perceived labor supply equation and taking the other firm's entire sequence of wages (Bertrand) or employment levels (Cournot) as given. The first-order conditions for a maximum imply

$$MRP_t = \frac{dR_t}{dl_t} = w_t + \frac{dw_t}{dl_t} l_t + \frac{dw_{t+1}}{dl_t} \frac{l_{t+1}}{1 + r}. \tag{10}$$

Thus, for the Bertrand case, suppressing the firm subscripts on the parameters,

$$MRP_t = w_t \left(1 + \beta_2 + \frac{\phi_t \beta_3}{1 + r} \right),$$

where

$$\phi_t = \frac{l_t w_{t+1}}{l_{t+1} w_t}. \tag{11}$$

If a stationary-state solution applies, then ϕ_t equals one[7] and the percentage gap between MRP and the wage—the Lerner index—is given by

$$\frac{MRP_t - w_t}{w_t} = \beta_2 + \frac{\beta_3}{1 + r}. \tag{12}$$

This expression can be viewed as a weighted average of the short-run and long-run inverse elasticities of labor supply, with weights $r/(1 + r)$ and $1/(1 + r)$ respectively. A similar expression holds for the Cournot case. Note that this expression for the Lerner index depends only on the first derivatives (or local elasticities) of the perceived labor supply functions. Even if these log-linear supply functions are viewed as local approximations only, the values of their parameters can be used to compute the Lerner index.

A more complicated and perhaps even more realistic specification is the closed-loop equilibrium, where each firm anticipates both the effect of its current action on future labor costs (the so-called cost effect) and the effect on future actions of the other firm (the so-called strategic effect). For the constant-elasticity perceived labor supply functions derived above, it can be shown that in a two-period game, the strategic effect in the first period is negative, implying that firms will "underinvest" in wages and employment, setting them lower than in an open-loop equilibrium. For the more general case, however, the sign of the strategic effect is ambiguous because it depends on the second derivatives of the perceived labor supply functions (see Boal, 1993). If (4) and (8) are viewed as local approximations only, the values of their parameters cannot be used to compute the Lerner index. Moreover, the perceived labor supply parameters are not sufficient—even for the simple constant-elasticity case, the size of the strategic effect depends on the slopes of both firms' MRP curves. For these reasons, I will compute the Lerner index below only for the open-loop equilibrium solution.[8]

[7] In the dataset described below, the mean of ϕ_t ranges from .969 to .995, while the standard deviation ranges from .358 to .413, depending on the definition of wages.

[8] However, if both firms have constant-elasticity supply curves and flat (constant) MRP curves, the strategic effect is zero under both Bertrand and Cournot behavior (Boal, 1993), so the estimates computed below may be interpreted alternatively as reflecting the closed-loop equilibrium solution with flat MRP curves. Flat MRP curves are implied by constant returns to scale, perfectly elastic demand for output, and perfectly elastic supply of nonlabor inputs.

□ **Estimation strategy.** In the above analysis I show that the coefficients of employment and lagged employment in wage equations (4) or (8) can be used to estimate the Lerner index. However, this estimate is conditional on the employer's discount rate and the assumed equilibrium concept. An alternative approach might compute the Lerner index directly from wage data and estimates of marginal revenue products, as in Scully (1974). Unfortunately, direct estimates are inconsistent in the presence of systematic measurement error on wages, such as is probably present here. In particular, the wage data in this study are best interpreted as wage indices only, and in any case they omit such important components of labor cost as company house rentals, company-store purchases priced above or below cost, and company-town amenities.[9] In contrast, the log-linear labor supply approach used here is robust to such measurement error, provided that true labor costs are proportional to observed wages. With the addition of individual dummies, that constant of proportionality can even vary across firms.[10]

3. Results

■ **The West Virginia coal data.** The Bertrand and Cournot perceived labor supply curves are estimated on a panel dataset of West Virginia coal mining for the years 1897 through 1932. Most of the data are taken from the *Annual Reports* of the West Virginia Department of Mines. Employment is measured as the number of workers. Two measures of wages are used in this article, as follows.

Tonnage wage of pick miners. This series is given directly by the *Annual Report.* Although the *Report* gives wages for many categories of workers, these categories change over time and missing values abound. The wage paid to pick miners per ton of coal represents the most plentiful single wage series available. Pick miners represented more than half of the total mine workforce in West Virginia at the turn of the century. But by 1932 they represented fewer than 8% and in some counties were completely replaced by mechanized methods. Hence the pick miners' wage is plagued by small-sample noise and missing values toward the end of the sample period.

Index of tonnage wages. To reduce the noise in the pick miners' wage and increase the number of observations toward the end of the sample period, I constructed an index to exploit data on other tonnage workers where available. Components of the index include the tonnage wages of machine miners and machine runners as well as pick miners, who together represented roughly one-half to two-thirds of the mine labor force. The index was anchored by setting the value for every county to 100 in 1925, the year of the most plentiful data. Next, I computed indices for 1924 and 1926 by growing this value at the average growth rate of all available tonnage wages in each county. (Thus the pick miners' wage might be missing for 1924, but if, say, the machine miners' wage were available for both 1924 and 1925, the index could still be computed.) The index was extended year by year in this fashion back to 1897 and forward to 1932.

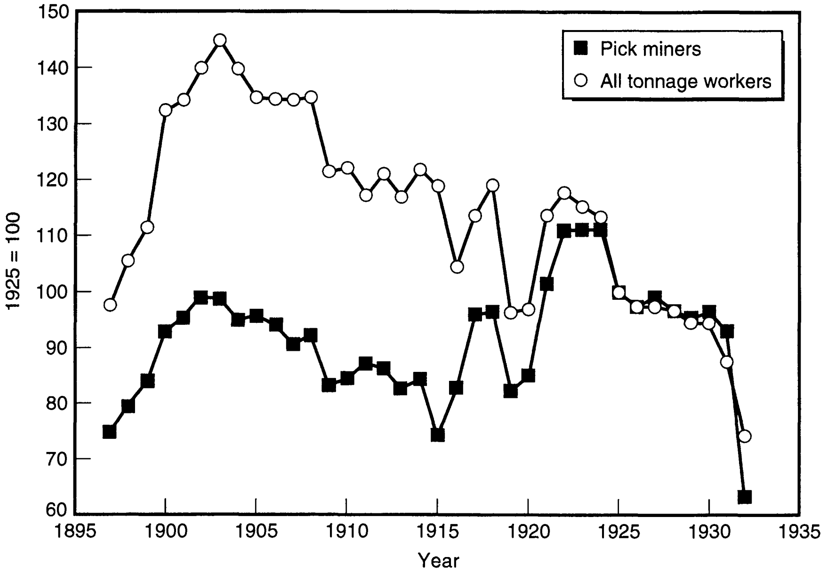
Piece rates are a good index of earnings to the extent that productivity is constant. In fact, productivity differed across mines due to differences in seam thickness and other natural conditions (Fishback, 1992b). These conditions were unlikely to change quickly over time, however, so individual fixed effects and individual trends should account for most of these productivity differences in regression analysis. Statewide averages for these wage measures are plotted in Figure 3. The downward trend visible in the tonnage wage index, relative to the wages of pick miners, is most likely due to

[9] Fishback and Lauszus (1989) find a negative effect of company-town sanitation amenities and a positive effect of company-housing rents on wages.

[10] However, see Boyd (1994) for an attempt at direct measurement of the Lerner index.

FIGURE 3

MEASURES OF REAL WAGES IN WEST VIRGINIA COAL, 1897–1932



SOURCE: See text. Pick miners' wage scaled to equal 100 in 1925.

technical progress in machine mining, which raised productivity and reduced piece rates.

An important constraint on this article is that the Department of Mines reports wages at the level of the county, not the mine or the firm. In consequence, the estimates below actually measure labor supply to entire counties. Most coal-producing counties contained at least a dozen firms, often operating in close proximity. Cross-elasticities between firms within the same county were surely negative. Therefore, the county-level inverse elasticities reported below are overestimates of the corresponding firm-level inverse elasticities. Consequently, the estimates below measure the true Lerner index only if all the firms in each county colluded with respect to wages. If collusion within counties was imperfect, however, then the estimates are best interpreted as upper bounds to the true Lerner index. I examine this aggregation bias in detail after presentation of the estimates.

The dataset is unbalanced in the sense that not all counties are observed in every year, for three reasons. First, some data (usually wages) are missing in the source. Second, some counties began or ceased production of coal during the sample period. Finally, many West Virginia counties were partly or entirely unionized during the teens and early twenties. Most models of unionism predict that unions raise wages by pushing labor markets off the supply curve. Observations experiencing significant unionism according to Boal (1994) were therefore dropped from the sample.[11] Descriptive statistics for the sample are given in Table 2.

□ **Estimation method.** The simple two-equation scheme described above is infeasible for the thirty-odd coal-producing counties in West Virginia. With fewer than thirty observations for most counties, cross-elasticities with respect to all other counties are obviously not estimable. Instead, I estimate the equations (4) and (8) for firm i over

[11] "Significant unionism" here means more than 5% of county tonnage produced under union contract.

TABLE 2 County Data on West Virginia Coal Mining, 1897–1932: Summary Statistics

Variable	Mean	Standard Deviation	Minimum	Maximum	Number of Observations
Tonnage wage of pick miners	.431	.072	.220	.790	685
Index of all tonnage wages (1925 = 100)	109.476	23.944	59.358	232.369	721
Average annual earnings in manufacturing[a]	641.302	106.430	491.566	801.393	731
Price of coal (county)	.962	.307	.463	2.532	719
Price of coal (national)[b]	1.204	.193	.963	1.873	731
Coal employment	2,476.731	3,853.177	4	19,925	731
Population[c]	29,201.532	21,648.293	7,420	165,286	731
Number of mines	25.375	32.359	1	163	731

Source: All data from West Virginia Department of Mines, *Annual Report,* except as noted below.
Notes: Wages, earnings, and prices are in real terms (1914 dollars). All observations are nonunion.
[a] Series D-781 in U.S. Department of Commerce (1976) through 1926. Thereafter, Series D-740 in same source.
[b] Series M-97 in same source.
[c] From decennial Census. A piecewise log cubic Bessel interpolation (Conte and Boor, 1972) is used to impute population between census years.

all counties, defining firm j for each county as simply the rest of the state. Thus, the county's own wage is regressed against its own current and lagged employment and also either the average wage in the rest of the state (Bertrand) or the total current and lagged employment in the rest of the state (Cournot). The constant term in either equation is expanded to include other exogenous variables tending to shift the labor supply curve. These include wages paid elsewhere in the economy (measured by national average annual earnings in manufacturing) and the population of the county.[12] County dummies and county-specific time trends are included to reflect the effect of other unobserved differences across counties, such as the nonwage labor costs mentioned above and productivity effects on piece rates.[13]

Whichever equilibrium concept applies, any random shift in a county's labor supply curve would affect both equilibrium wages and equilibrium employment in that county. Moreover, one might reasonably suspect that these two variables are measured with error. Hence I treat these two variables as endogenous and estimate the Bertrand and Cournot equations by two-stage least squares. Lagged employment in the county also must be treated as endogenous if the error term is serially correlated, or if employers anticipate values of the error term. However, if one assumes that any particular county exerts negligible influence on the rest of the state, then wage and employment variables for the rest of the state may be treated as exogenous. Appropriate instruments should be variables that influence the endogenous regressors without being themselves influenced by the error term. One set of instruments satisfying this requirement comes from the first-order conditions of firm optimization. Whichever equilibrium concept

[12] Average annual earnings in manufacturing are represented by Series D-781 ("Average annual earnings of wage earners in manufacturing") in U.S. Department of Commerce (1976) through 1926. To this series is spliced Series D-740 ("Average annual earnings per full-time employee, manufacturing") for later years. County population is from the decennial Census, with piecewise log cubic Bessel interpolation (Conte and de Boor, 1972) used to impute population between census years.

[13] The estimates reported below show, as expected, that pick miners' wages were negatively correlated with productivity. The correlation of the wage differentials implied by the Bertrand dummy and trend coefficients with log output per worker-day was −.47 in 1925. For the Cournot specification, the correlation was −.27. Fishback (1992b) obtains similar results.

applies, these first-order conditions will surely include the price of coal. The estimation results here use the county average price of coal as reported by the West Virginia Department of Mines. Because this figure is most likely measured with error, I included a national price and GNP as well.[14] Price, wage, and employment variables for the rest of the state would not be separately correlated with the endogenous regressors under myopic or open-loop equilibrium concepts, but they are nevertheless uncorrelated with the error term and are included here. (Their omission did not change the results much.) Finally, because lagged employment in the county is endogenous, I also included lagged values of all exogenous variables and above-named instruments. All prices and wages are converted to real terms using the consumer price index.[15] All variables except county dummies and trends are measured in logarithms. Non-identically and independently distributed errors were accommodated using the standard error formula of Newey and West (1987), which is robust to heteroskedasticity and two-period serial dependence of arbitrary form.[16] (Usual standard errors were slightly smaller; see Boal (1993).)

Estimates of equations (4) and (8) are given in Table 3. It should be emphasized that these results cannot be used to choose between Bertrand and Cournot, because either equation should describe the data under either equilibrium concept. The parameters of firm interaction have not been estimated. If one knew the correct equilibrium concept, one could pin down the Lerner index more precisely. Nevertheless, the estimates in Table 3 can together help bound that gap.

□ **Bertrand estimates.** The estimates for Bertrand perceived labor supply are shown in the first two columns of Table 3. The positive coefficients for current employment (.15 and .22, depending on the measure of wages) indicate small positive short-run inverse elasticities and small short-run monopsony power. However, the fact that the coefficients for lagged employment are of roughly the same magnitude indicates near-zero long-run inverse elasticities and near-zero long-run monopsony power. Similarly, for a benchmark employer with a discount rate of, say, 20%, the estimated Lerner indices defined by equation (12) are very close to zero.

The Bertrand results seem fairly robust. The following alternative estimation approaches produced almost identical results: limited-information maximum likelihood (LIML) estimation instead of two-stage least squares; including the number of mines in the county as an instrument; partitioning the wages in the rest of the state into separate measures for nearby and more distant counties; estimation on first-differenced data; excluding employment and coal prices in the rest of the state from the list of instruments; and excluding county population as a regressor (see Boal, 1993).

For comparison, I also estimated the Bertrand specifications using two alternative wage measures constructed for this study, with roughly similar results. An index of time wages produced an estimated Lerner index of about 7%, and a combined index of all wages produced an estimated Lerner index of less than 1%. However, these estimates are less reliable than those presented in Table 3 because the data are inferior. In principle, time wages should provide a better index of labor costs than piece wages, but this time-wage index suffers from missing data and narrow coverage. In particular,

[14] The national coal price was taken from Series M-97 ("Bituminous coal—Average value per net ton, f.o.b. mine") in U.S. Department of Commerce (1976). GNP is taken from Series F-3 ("Gross national product, 1958 prices") in the same source.

[15] The CPI is reported as Series D-727 and E-186 in U.S. Department of Commerce (1976). The first is a Bureau of Labor Statistics series beginning in 1914. The second is an extension by Albert Rees back to 1890.

[16] In panel data, the kernels for robust estimators sometimes span different units, but this does not affect consistency (White and Domowitz, 1984).

TABLE 3 **County Labor Supply in West Virginia Coal Mining, 1897–1932**

Dependent Variable	Bertrand: Tonnage Wage of Pick Miners	Bertrand: Index of All Tonnage Wages	Cournot: Tonnage Wage of Pick Miners	Cournot: Index of All Tonnage Wages
Coal employment in county	.147482	.221575	.527399	.522038
	(.079328)	(.082476)	(.119964)	(.127446)
Lagged coal employment in county	−.140957	−.290560	−.517763	−.538496
	(.074786)	(.078040)	(.147170)	(.131878)
Wages in rest of state	.727164	.775129		
	(.085807)	(.108295)		
Coal employment in rest of state			.205288	.175810
			(.156590)	(.169445)
Lagged coal employment in rest of state			.029168	.034619
			(.173429)	(.174675)
Average annual earnings in manufacturing	.323115	.501622	1.64061	.827651
	(.171359)	(.138636)	(.208881)	(.198567)
Population of county	.060242	.330487	.024709	.281963
	(.133461)	(.135415)	(.174392)	(.164296)
R^2 (adjusted)	.580987	.648969	.208488	.437008
Functions of coefficients				
Long-run inverse elasticity	.00652496	−.068985	.00963584	−.016458
	(.040152)	(.037963)	(.086837)	(.088656)
Lerner index (at 20% discount rate)	.030018	−.020558	.095930	.073291
	(.040193)	(.038586)	(.075328)	(.082714)

Notes: Coefficient estimates are computed by two-stage least squares, with county employment and its lag treated as endogenous. Standard errors (in parentheses) are computed using the Newey-West estimator allowing for heteroskedasticity and serial dependence for two periods. Instruments include county coal price, coal price in rest of state, national coal price, and real GNP. All variables are in logarithms. Wages and prices are in real terms. All equations include county-specific intercepts and county-specific time trends. All observations are nonunion. The estimates using the tonnage wage of pick miners reflect 637 observations on 33 counties. The estimates using the index of all tonnage wages reflect 673 observations on 34 counties.

it is based heavily on data for machine runners and helpers, the "aristocrats of the industry," a highly skilled and highly paid group that never represented more than 6% of the West Virginia mine labor force and that was usually paid on piece (Goodrich, 1925; Fishback, 1992b). The combined index shares some of these defects too, because it is defined as an employment-weighted average of the tonnage-wage and time-wage indices. Nevertheless, it is reassuring that the estimated Lerner indices are not greatly affected. (See Boal (1993) for details.)

□ **Cournot estimates.** The estimates for Cournot perceived labor supply are shown in the last two columns of Table 3. The coefficients for current employment (.53 and .52, depending on the measure of wages) are larger than the corresponding Bertrand estimates, as expected, and show more substantial short-run monopsony power. However, the coefficients for lagged employment are almost identical in magnitude, again indicating essentially zero long-run monopsony power. For an employer with a discount rate of 20%, the estimated Lerner indices are 10% and 7%—larger than the Bertrand estimates, but still fairly small and not statistically different from zero.

The Cournot estimates are somewhat less robust to alternative estimation approaches than are the Bertrand estimates. The estimates of the short-run inverse elasticity were especially unstable, varying from .35 to .99, but this is unsurprising given the close correlation between county employment, employment in the rest of the state, and their lags. The estimates of the long-run inverse elasticity and the benchmark Lerner index were somewhat more stable. For these quantities, the following alternative estimation approaches produced almost identical results: LIML estimation instead of two-stage least squares; including the number of mines in the county as an instrument; partitioning employment levels in the rest of the state into separate measures for nearby and more distant counties; and excluding wages and coal prices in the rest of the state from the list of instruments (see Boal, 1993). Estimates excluding county population as a regressor raised the coefficient of county employment and pushed the (negative) coefficient of lagged employment closer to zero (as might be expected, since population is positively correlated with both variables). This resulted in estimated long-run inverse elasticities of .10 and .12 and estimated benchmark Lerner indices of 18% and 19%, depending on the measure of wages. Estimates using first-differenced data had a qualitatively similar but stronger effect, partly because first differencing reduces population, which is strongly trended, virtually to a constant. However, first-differenced estimation is highly vulnerable to time-shifting measurement error, a potential problem because the precise date(s) at which employment and wages are measured are not given in the *Annual Reports,* so the first-differenced results must remain suspect.

Cournot estimates using the alternative wage measures produced similar short-run elasticities, but larger long-run elasticities and estimated Lerner indices. The index of time wages produced a Lerner index of 18%, while the combined-wage index produced a Lerner index of 20%. The difference between these estimates and those in Table 3 may partly reflect defects in the time-wage data described above.

□ **Bias from variation in local cost of living.** One possible source of bias in these estimates arises from the use of a national consumer price index to convert nominal wages to real wages. Many miners relied on company stores, whose prices probably differed systematically from national averages (U.S. Coal Commission, 1925). It is important to note that the inverse elasticity estimates are unaffected if company store prices were above or below national prices, provided they remained proportional. The constant of proportionality is simply absorbed into the county-specific fixed effect. Bias arises only if employers strategically varied store prices directly (or inversely) with nominal wages over time, to diminish (amplify) the effects of nominal wage changes. In that case, inverse elasticity estimates that ignore such pricing behavior are biased up (down) because they overestimate (underestimate) the real wage changes associated with the observed employment shifts.

This source of bias is difficult to evaluate because few data are available on company-store prices in coal mining, and most studies of other industries, such as Johnson (1952), focus on static price levels, not dynamic pricing behavior. However, some reassuring evidence is given by Fishback (1986), who analyzes prices at the Stonega Coal Company stores in Virginia from 1918 to 1932. Fishback reports that store prices were strongly correlated with a national consumer price index and uncorrelated with the firm's nominal wages. If the Stonega Coal Company's behavior was typical of firms in West Virginia, strategic company-store pricing is not a major source of bias in these estimates.

□ **Aggregation bias.** A second source of bias results from the aggregate nature of the data. County-level data yield consistent estimates of market power only if cross-elasticities of labor supply between firms within any county are zero—implausible—or if all firms within a county collude in setting wages—implausible in most cases.[17]

[17] The *Annual Report* for 1925 shows only two counties—Randolph and Tucker—where the majority of mines were owned by a single company. Both counties were small, producing less than 1% each of total West Virginia tonnage.

Otherwise the estimated inverse elasticities are biased upward by an amount that varies inversely with the absolute value of the cross-elasticities within counties.

Cross-elasticities within counties cannot be measured with these data, but their effect on the estimates can be computed under alternative assumptions. At one extreme, if the cross-elasticities within counties are assumed infinite, then the biases are large. A firm playing Bertrand obviously has no market power, and the true Bertrand firm-level inverse elasticities are zero. A firm playing Cournot faces an inverse elasticity equal to the county inverse elasticity multiplied by the firm's employment share in the county. Assuming a rough average of 17 firms in a county,[18] the Cournot estimates in Table 3 imply a short-run inverse elasticity of at most .031 and a benchmark Lerner index of less than 1%.

At the other extreme, if the cross-elasticities within counties are assumed to be merely identical to the cross-elasticities between firms in different counties, then the biases are quite small. This is because cross-effects are then proportional to numbers, and firms within any county are not as numerically important as firms in the rest of the state. In particular, one can show that the Bertrand estimates are biased upward by a multiplicative factor of $(1 + \beta_4^i N)$, where β_4^i is the coefficient of wages in the rest of the state and N is the ratio of the number of other firms in the county to the number of other firms in the rest of the state (roughly .03, given 34 counties). Clearly, this bias is small. Assuming symmetry, one can show that the Cournot estimate of $\beta_2^{i'}$, the short-run inverse elasticity, is biased upward by an additive term equal to $(\beta_4^{i'} N)$, where $\beta_4^{i'}$ is the coefficient of current employment in the rest of the state. Similarly, the Cournot estimate of β_3^i, the coefficient of lagged employment in the county, is biased away from zero by an additive term equal to $(\beta_5^{i'} N)$, where $\beta_5^{i'}$ is the coefficient of lagged employment in the rest of the state. Clearly, the biases in the Cournot coefficients are also very small (because N is small) under the assumption of identical cross-elasticities. (See Boal (1993) for derivations.)

Limited evidence suggests that cross-elasticities are not identical, that firms actually compete for labor more intensively with nearby firms than with more distant firms. When for the Bertrand equation wages in the rest of the state were partitioned into separate measures for nearby counties and more distant counties, the coefficients for nearby counties were larger than their relative numbers would predict. In this procedure, the state's counties were divided into six regions, corresponding roughly to major coal fields. In place of the single variable "wages in the rest of the state," two regressors were entered: one for average wages outside the county but in the same region, and the other for average wages in other regions. If cross-elasticities were identical, the coefficients for these variables would be roughly proportional to their relative numbers of constituent counties—that is, in a proportion of 1:5. In fact, the coefficients were in proportions of 1:1.25 or 1:2, depending on the measure of wages, suggesting that nearby counties exert greater influence on wages, relative to their numbers, than did more distant counties. If competition between nearby counties is more intense than competition between more distant counties, then competition between firms in the same county is probably even more intense. Therefore, some upward bias in the estimated elasticities and benchmark Lerner index due to the aggregate nature of the data seems likely.

4. Conclusion

■ In this article I estimated the parameters of labor supply to differentiated firms in a setting popularly associated with employer monopsony. Although some short-run

[18] In 1925, a total of 793 companies operated 1,208 mines in West Virginia, according to West Virginia Department of Mines (1925), giving an average of 1.52 mines per company. Dividing this figure into 25.4, the mean number of mines per county in the sample (see Table 2) gives 17 firms per county.

monopsony power was detected, this power was sharply attenuated if employers "considered the future," that is, if they foresaw the effect of current wages and employment levels on their own future labor supply. If one assumes that firms played a Bertrand game, taking other firms' wages as given, then the data indicate there was probably no substantial gap between marginal revenue product and the wage. If one assumes that firms played a Cournot game, taking other firms' employment levels as given, then small amounts of monopsony are consistent with the data, but the Lerner index was probably less than 10%. This article does not attempt to infer the correct equilibrium concept, but the results are clearly not very different for Bertrand and Cournot.

The results in this article demonstrate that employer monopsony played at best only a small part in determining wages in turn-of-the-century coal mining. In contrast, real wage fluctuations over time, driven primarily by product demand shocks, show frequent movements of 20% or more over intervals of just a few years (see Figure 3). Lewis (1963) estimated union relative wage effects ranging from 33% to over 100% in the coal industry during this period, although union effects within West Virginia were much smaller (Boal and Pencavel, 1994). Even intraindustry differences in state average accident rates have been estimated to cause as much as a 7% wage differential in favor of West Virginia, relative to coal mining in the rest of the United States.[19]

Moreover, the upper-bound Cournot estimates reported in Table 3 are much smaller than comparable estimates for other labor markets that various authors have characterized as monopsonized. D. Sullivan (1989) estimates inverse elasticities in the short run of .79 and, after a two-year lag, of about .26 for nurses. (In comparison, (7) above implies that the two-year lag estimates are about .05 for Bertrand and about .18 for Cournot before correcting for any aggregation bias.) Scully (1974) estimates Lerner indices for baseball pitchers and hitters in 1968 and 1969 under the reserve clause, a form of collusion, of roughly 100% to 400%. Machin, Manning, and Woodland (1993) estimate a Lerner index of about 13% at U.K. old-age homes. Katz and Krueger (1992) estimate an inverse elasticity of roughly .38 to .59 for fast-food restaurants.

In summary, the econometric evidence indicates that West Virginia coal is not a good example of monopsony and that miners moved relatively quickly in response to wage differences across employers. Of course, a few individual employers in unusually remote locations may have enjoyed substantial monopsony power, but the econometric evidence shows that, on average, labor supply to individual employers must have been very elastic. In light of this evidence, the traditional view that company towns were a source of monopsony power must be reconsidered. In fact, the opposite may have been true. The simple correlation of company housing and turnover across coal fields in Table 1 is .57 (or .22 excluding West Virginia). Perhaps company housing, by reducing a worker's transactions cost of relocating to a new job (Fishback, 1992a), actually increased labor mobility. Moreover, any extractive industry, no matter how remotely located, must have had sufficient transportation facilities to move the product to market. Perhaps remoteness should not automatically be viewed as a barrier to labor mobility either.

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[19] Data reported in Fishback (1986) show that small-scale accidents (the most frequent type) were 33.5% more common in West Virginia than for the United States as a whole. His econometric estimates show that a 1% increase in wages was associated with a 4.8% increase in the accident rate. Dividing 33.5 by 4.8 gives 7.0%.

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