This PDF is a selection from an out-of-print volume from the National Bureau of Economic Research

Volume Title: The Measurement of Labor Cost

Volume Author/Editor: Jack E. Triplett, ed.

Volume Publisher: University of Chicago Press

Volume ISBN: 0-226-81256-1

Volume URL: http://www.nber.org/books/trip83-1

Publication Date: 1983

Chapter Title: Intermetropolitan Wage Differentials in the United States

Chapter Author: George Johnson

Chapter URL: http://www.nber.org/chapters/c7381

Chapter pages in book: (p. 309 - 332)

8 Intermetropolitan Wage Differentials in the United States

George E. Johnson

8.1 Introduction

The purpose of this paper is to investigate the nature and causes of wage differentials between large metropolitan areas in the United States. Most of the recent literature on this subject has concerned the question of whether or not wages in the southern United States are lower than elsewhere.[1] The current consensus is that, although there is a wide disparity in *nominal* wage levels between the South and elsewhere, there is virtually no difference in *real* wage levels. The results in the present paper do not contradict this conclusion; indeed, real wage levels in Atlanta, Dallas, and Houston are estimated to be slightly higher in 1973–76 than wages in comparable cities in the North.

There is, however, considerable variation in wage rates in large metropolitan areas throughout the United States. For example, the estimated nominal wage of a private sector, nonunion, white, full-time male paid by the hour is 10 percent less in Boston than in Detroit (the real wage is 23 percent less). My purpose is to sort out why these differences exist. Are they best explained as a disequilibrium phenomenon, as the result of regional differences in the extent and nature of unionism, or as compensating differentials to reflect differences in the nonpecuniary attributes of areas?

The first task of the study is to estimate the "area effects" on nominal wages for four different types of workers (full-time male and female, hourly and salaried) from the May Current Population Survey data for 1973–76. These estimated area effects, which control for the standard human capital variables, race, unionism, and public sector employment,

George E. Johnson is with the Department of Economics, University of Michigan, Ann Arbor, and the National Bureau of Economic Research.

are then converted into real area effects by subtracting the logarithm of the price level for each area from the nominal area effect. For women the dispersion of area real wage levels across regions is smaller than the dispersion of nominal wages, but for men it is greater. There is, in addition, a positive correlation between both the nominal and real area effects for the four groups of workers. In other words, whatever set of factors makes wages high or low in a particular city for one group of workers also makes them high or low for other groups of workers.

There is some rather weak evidence that the size of the real nonunion area effect is negatively related to the extent of unionization in the area. This is consistent with the standard hypothesis that individuals will accept nonunion jobs in an area with a high degree of unionization at a lower wage than they would accept nonunion jobs in an area with a low degree of unionization. The reason for this is that their long-run income prospects may be better if they reside in an area with a greater probability of eventual attainment of a high-rent job. This runs counter to two hypotheses, which are reviewed in section 8.2, that predict that a large union sector will cause high union wages to "spill over" to the nonunion sector. The evidence on the distribution of real *nonunion* area effects is not encouraging to either of these hypotheses.

8.2 Conceptual Issues

Suppose we observe that nominal wage levels, after adjusting for human capital, in different areas across the United States are subject to substantial variation. To what could this phenomenon be attributed? Since, as will be shown subsequently, interarea wage levels *are* subject to variation that is both large and persistent, it is useful at the outset to state the various hypotheses that might explain it: (a) Wage differences represent a *compensating variation* to offset differences in price levels, nonpecuniary attributes, or both, and (b) wage differences are caused by *institutional rigidities*, primarily by differences in the nature and extent of unionism across areas.

8.2.1 Compensating Variation

If there are no differences in the assessments by individuals of the nonpecuniary attributes associated with residence in different areas, as well as no institutional impediments to the adjustment of relative wages, we would expect the "real" wage in all areas to be equal. In its most straightforward terms, this implies that

$$\frac{w_i}{w_j} = \frac{p_i}{p_j}, \tag{1}$$

where the w's are nominal wage levels and the p's are price levels for areas i and j. If, for example, we observed that the nominal wage in area i

were greater than that level that satisfies equation (1), we would expect that there would be migration into the area sufficient to drive w_i/p_i down to the nationwide value of the real wage rate $(w/p_i)*$.[2] Thus, by this simple specification, the long-run supply curve to an area is perfectly elastic at the $(w/p)*$. It is, therefore, a labor supply condition rather than a conventional particular supply curve.

In fact, the effect of price variation across areas on the equilibrium regional wage structure is slightly more complicated than this. The preceding discussion implies that, other things equal, a 1 percent increase in the cost of living in an area will increase the equilibrium wage in that area by 1 percent. This may not be true for two reasons: First, the nature of the tax and transfer system, and second, people have the option of retiring in areas with low price levels.

To take a simple case, assume that the utility of each person is a function of consumption during his working life (c_1) and during retirement (c_2) and that the utility function is Cobb-Douglas, i.e.,

$$U = \alpha \log c_1 + (1 - \alpha) \log c_2. \tag{2}$$

Second, assume that federal income taxes may be approximated by a linear function, $t_f(w_i - X)$, where t_f is the marginal federal tax rate, w_i the gross earnings the individual earns in area i during his working life, and X the constant tax deduction (cumulated over his working life). In addition, the individual pays a proportional local income tax, $t_i w_i$, which is not deductible from his federal tax base. Finally, upon his retirement the individual receives a social security payment of S from the government, which is assumed to be independent of w_i. If the individual chooses to remain in area i during the years of his retirement, his lifetime budget constraint is

$$\begin{aligned} O &= [w_i(1 - t_f - t_i) + t_f X - p_i c_i] + \left(\frac{1}{1 + z}\right)(S - p_i c_2) \\ &= V_i + V_T - p_i c_1 - \frac{p_i}{1 + z} c_2, \end{aligned} \tag{3}$$

where z is the interest rate (net of taxation, which is assumed to be solely federal), and p_i is the price level in the area (including local sales and property taxes). $V_i = w_i(1 - t_f - t_i)$ is the present value of net labor earnings (evaluated at the marginal rather than the average federal tax rate), and V_T is the present value of the income tax deduction and the social security payment.

Maximization of equation (2) with respect to c_1 and c_2 subject to equation (3) yields a utility-maximizing path of consumption over the life cycle. Plugging these values back into the utility function yields the utility associated with location in area i, and this depends on w_i and p_i as well as the various tax and transfer parameters. Labor supply equilibrium requires that the lifetime utility associated with all areas be equal. Of particular interest for present purposes is the effect of changes in the two parameters subject to interarea variation, p_i and t_i, on the eqilibrium area wage level. This can be shown to be

$$(4) \qquad d(\log w_i) = \frac{V_i + V_T}{V_i} d(\log p_i) + \frac{dt_i}{1 - t_f - t_i}.$$

The coefficient on $d(\log p_i)$ is *greater* than one if the tax system is progressive ($X > 0$) or if there is a social security system (and the tax system not too regressive). The reason for this is that higher values of X and S diminish the relative importance of labor earnings in lifetime net income, thus requiring greater variation in w_i to compensate for a given variation in p_i. Variation in local (proportional) income taxes, however, have the same effect on w_i as one would expect on the basis of the equalization of after-tax wage rates.

The second complication arises from the fact that an individual may move from area i to area j (where $p_i > p_j$). Moving costs, which are incurred in period 2, are M, so for a person who does move the budget constraint becomes

$$(5) \qquad O = V_i + V_T - \frac{M}{1+z} - p_i c_i - \frac{p_j}{1+z} c_2.$$

Repeating the procedure followed above, the effect of variations in p_i on the equilibrium value of w_i is equal to ($\partial[\log w_i])/\partial t_i$ is the same as in equation [4])

$$(6) \qquad \frac{(\log w_i)}{(\log p_i)} = \alpha \left(\frac{V_i + V_T - \frac{M}{1+z}}{V_i} \right).$$

This may be greater or equal to one. If V_T and M were zero, it would simply equal the share of first-period consumption in the utility function, which is less than one. The more important transfers are and the more progressive the tax system is, the larger the effect of variations in area price levels is on the variation in area nominal wage levels.

A second source of variation in area real wage levels arises from the possibility that individuals may receive utility from specific nonpecuniary attributes of different areas. If, say, the San Francisco area is considered (over the relevant range) to be y more desirable in terms of climate, physical beauty, public services (net of taxes), and the like than the averge area in the country, then the equilibrium real wage in San Francisco would be $1 - y$ of the average real wage for all areas. The resultant supply curve would be horizontal at a lower real wage than that for the typical area (certainly lower than that of Buffalo). Compensation variations due to nonpecuniary attributes are subject to the modification arising from the possibility of locational change upon retirement that was discussed above with respect to price level differences. This merely implies that nonpecuniary differences are less important than if people never moved.

Another variant of the nonpecuniary model stresses that individuals have different tastes for different areas. For example, some individuals abhor climatic variation (and hence would sacrifice real income to live in San Diego) while others enjoy the "change of seasons" (and would, other things equal, prefer to live in Buffalo). This specification yields an upward-sloping equilibrium supply curve for each area that is the more elastic the less individuals' assessments of the nonpecuniary attributes of different areas vary. The general model makes no prediction about the eqilibrium wage structure—if real wages are high in area *i*, it is not viewed as an attractive area. The general specification, however, predicts that, given stability over time in the distribution of preferences, increases in the relative employment of an area, which arise due to relative shifts in the demand functions, will be associated with increases in relative wages. Both of the other specifications—nonpecuniary attributes are (a) not important or (b) evaluated identically by everyone—predict that there will be no relation between changes in relative wages and employment.

8.2.2 Institutional Rigidities

A second set of reasons why nominal and real wage levels may vary across areas is the failure of markets to adjust because of institutional rigidities. The primary candidate for such a rigidity is, of course, trade unionism—although the federal government and some local governments have wage structures that are a similar source of rigidity.[3]

Now if some areas were 100 percent unionized and unions were able to obtain a wage premium for their members, we would expect that the real wages of all the nonunion areas would be equal and the real wages of each union area to vary with the bargaining strength and success of the particular unions in that area. Labor mobility in this case would equilibrate the expected incomes (or utility) of each unionized area with that of the nonunion areas, and the equilibrating variable would be the unemployment rate in each unionized area.[4]

In fact, no areas are 100 percent unionized; there is not even a major metropolitan areas of the United States in which a majority of the labor force is unionized. Some areas (especially in the New York to Chicago industrial belt) are relatively heavily (25–45 percent) unionized, while other areas (especially in the South and Southwest) have very little unionization (10–15 percent). In this situation of *partial* unionization, we can observe equilibration through wage adjustment.

First, if individuals must live in an area to obtain a union job in that area, the equilibrium nonunion real wage in area *i* should depend negatively on the extent of unionism in that market. The reason for this is that the reservation wage of a person in a market with a high probability of future high wages will be lower than that in a market with a low probability of future rents. If, however, it is not possible to move from a nonunion to a union job (because unemployed persons are able to corner the search market), the nonunion wage rate in a heavily unionized labor market will not be influenced by the union wage. Instead, the nonunion wage will equal the nonunion wage in all other labor markets as well as the expected value of the income (or utility) associated with attaching oneself to the unionized sector.

Second, if individuals do not have to live in an area to obtain a union job in a highly unionized area, they can search for high-rent jobs (by, say, telephone or a two-day trip to Cleveland) without sacrificing their nonunion jobs in their initial area. In this circumstance, as in the case of the inability to move from a nonunion to a union job, there would be no tendency for nonunion jobs in highly unionized markets to offer real wages that are lower than those in relatively unionized markets.

The existence of a relatively high degree of unionization in an area may, therefore, lower the nonunion wage in an area or, under certain circumstances, have no effect on it. By the above arguments, however, it will not raise the nonunion wage. There are, however, two ways in which the existence of a large union sector could increase the nonunion wage above that in relatively unorganized labor markets.

First, assume that a nonunion employer could hire all the labor he wants at a wage of w_0. However, the probability that his firm will be organized (that an NLRB certification election will be held and won by the union) is $U = \phi(w/w_U, U)$, where w is the wage he offers and U is the extent of organization in the area as a whole.[5] Presumably, $\phi_1 < 0$ (the firm is more likely to be organized the lower its wage offer is relative to the prevailing union wage), $\phi_2 > 0$ (organization is more likely the greater the extent of unionization in the area), and $\phi_{12} < 0$ (the reduction in the threat of organization as w/w_U is increased the greater the extent of organization in the area). The expected wage rate that the firm will pay then depends on the wage it offers relative to the union wage, $R = w/w_U$, and the extent of unionization in the area, that is

$$\bar{w} = w_U[\phi(R, U) + R(1 - \phi(R, U))], \; R \geq R_0. \tag{7}$$

If there is an interior solution (i.e., $d\bar{w}/dR > 0$ at $R = R_0$), the value of R satisfies

$$1 - \phi(R, U) + (1 - R)\phi_1(R, U) = 0, \tag{8}$$

and, given that $\phi(1, U) = 0$, the cost-minimizing relative wage offer is between R_0 and 1.

An example of a functional form that satisfies the assumptions of the model is the quadratic:

$$\phi(R, U) = a_0 U(1 - R)^2. \tag{9}$$

In this case the cost-minimizing relative wage offer is

$$R_* = 1 - \left(\frac{1}{3a_0 U}\right)^{1/2}, \tag{10}$$

which is binding so long as $R_* > R_0$ obviously increases as U increases; a decrease in a_0 (which would result from, say, passage of a right-to-work law in the area), would diminish the probability that the threat effect is operative.

To the extent that the threat effect is operative in *any* areas of the country, it should be most important in those areas with relatively high degrees of unionization. In those areas (if, again, there are any), the only way that the market can equate the attractiveness of the area with that of other markets is for excess normal unemployment to occur. Jobs will be rationed in both the union and nonunion sectors—although it is possible that the nonunion wage in a highly unionized area is still lower than that in a weakly unionized market. If the threat effect is operative, the nonunion wage is simply higher than that value that clears the market.

The threat effect model may have had great general relevance for wage determination in the United States during the 1930s and 1940s—and there are still large firms that reputedly pay union wages in order to stave off organization. It may, however, seem a trifle unrealistic to attempt to apply the model to the larger part of the nonunionized sector. The union movement in the United States, it could be argued, it not very interested in attempting to organize most currently unorganized firms that are small, pay low wages, and are characterized by rapid labor turnover. The marginal cost of servicing such bargaining units would, in most instances, be less than the marginal revenue.

An argument could be made, however, that wages in the union sector could, even in the absence of the threat of organization, have a direct influence on the wage levels of many nonunion firms. Suppose that a nonunion firm can hire as many workers as it wants at a wage w_0, but it realizes that the effort expended by the typical employee (a) will depend positively on the wage the firm offers relative to the nonunion wage,[6] say $a = \psi(R, U)$, where $\psi_1 > 0$, $\psi_2 < 0$, and $\psi_{12} > 0$. The cost of a unit of effort is then

$$c = \frac{R w_U}{\psi(R, U)}, \quad R \geq R_0, \tag{11}$$

and it is *possible* that the cost-minimizing relative wage offer is between R_0 and 1. This will be the case if the elasticity of a with respect to R evaluated at R_0 exceeds unity.

This *contamination effect* model has roughly the same implications as the threat effect model. If there is anything at all to the hypothesis, the work effort of nonunion members will depend more significantly on their wage relative to the union wage in highly unionized areas than in those areas that have little unionization because high union wages are much more visible in the former than the latter. For example, the typical nonunion employee working in a highly unionized area is more likely to have held a high-rent union job (a higher wage for the same work) than is an equivalent person residing in an area with low unionization and thus would be more sensitive to the size of the union/nonunion differential.

8.2.3 Demand-Determined Versus Supply-Determined Area Wage Levels

To this point the discussion of area wage levels has been cast solely in terms of the supply side. A wage level in area i can deviate from that of other areas because a compensating variation is required or because of institutional considerations. For example, an area that is unattractive on nonpecuniary grounds will require high wage rates to attract workers. Given a finite long-run demand elasticity, the employment level in that area will be lower than if it were an attractive area, but there will still be a positive equilibrium employment level.

If, however, output were a function solely of labor and capital, the assumption of a finite demand elasticity would be wrong.[7] If the underlying production function were linear homogeneous, the condition of equal returns to capital would imply that all wage rates must be equal in the long run. If an area had a wage higher than any other area, its capital would flee and its employment would disappear. In other words, the demand curve, like the supply curve, would be horizontal.

On the other hand, suppose that the underlying production function is linear homogeneous in three factors: labor (E_i), capital (K_i), and resources (R_i^1). The last of these would include the industrial and commercial use of land, water, locally produced energy, and the like. Households, whose number is proportional to E_i, also demand resources—land for houses, water for swimming pools, etc., and the aggregate use of resources by households is R_i^2. I will assume, for the sake of simplicity, that each area has a fixed stock of resources (R_i) for both uses, i.e., $R_i = R_i^1 + R_i^2$.

The utility function for each household is given by $U_i = A_i\phi(c_i, r_i)$, where A_i is an area shift parameter reflecting the nonpecuniary attributes of the area, r_i is lifetime consumption of resources, and c_i lifetime consumption of all other goods. Ignoring taxes and area variations in the price of c_i (as well as the possibility of movement to low price level areas upon retirement), U_i is maximized subject to the budget constraint $w_i = c_i + b_i r_i$, where b_i is the price of a unit of resources. Thus, each household will demand $r(b_i, w_i)$ units of resources, and the total demand for resources by households in the area is

$$R_i^2 = r(b_i, w_i)E_i, \tag{12}$$

where $\partial(\log r)/\partial(\log b_i) = -\eta_b$ and $\partial(\log r)/\partial(\log w_i) = \eta_w$ are the price and income elasticities of demand.

Equalization of net advantages of all areas implies that the total derivative of the utility function with respect to A_i, w_i, and b_i be equal to zero, or

$$dU_i = \phi dA_i + A_i\phi_1(dw_i - r_i db_i) = 0. \tag{13}$$

Upon manipulation of this, we have

$$d(\log w_i) = -\frac{1-\beta}{\theta} d(\log A_i) + \beta d(\log b_i), \tag{14}$$

where $\theta = c_i\phi_1/\phi$ is the elasticity of utility with respect to c_i and $\beta = b_i r_i/w_i$ is the share of household income going to the consumption of resources.

On the factor demand side, the aggregate production function for the area is $Q_i = F(E_i, K_i, R_i^1)$. It is assumed that F is linear homogeneous and, for simplicity, that the elasticities of substitution between each of the three factors are identical (σ). Thus, the logarithmic derivatives of the three factor prices are given by

$$d(\log w_i) = -\frac{1-\alpha_1}{\sigma} d(\log E_i) + \frac{\alpha_2}{\sigma} d(\log K_i) + \frac{\alpha_3}{\sigma} d(\log R_i^1), \tag{15}$$

$$O = \frac{\alpha_1}{\sigma} d(\log E_i) - \frac{1-\alpha_2}{\sigma} d(\log K_i) + \frac{\alpha_3}{\sigma} d(\log R_i^1), \tag{16}$$

and

$$d(\log b_i) = \frac{\alpha_1}{\sigma} d(\log E_i) + \frac{\alpha_2}{\sigma} d(\log K_i) - \frac{1-\alpha_3}{\sigma} d(\log R_i^1), \tag{17}$$

where α_1, α_2, and α_3 are the three factor shares (which sum to one). The left-hand side of equation (16) is set equal to zero, reflecting the fact that the return to capital in all areas must be equal.

Since the fixed stock of resources in the area is divided between use by firms and households, i.e., $R_i = R_i^1 + R_i^2$, it follows that

$$d(\log R_i) = kd(\log R_i^1) + (1-k)d(\log R_i^2), \tag{18}$$

where $k = R_i^1/R_i$ is the fraction used by firms. Differentiating (12) logarithmically,

$$\alpha(\log R_i^2) = -\eta_b d(\log b_i) + \eta_w d(\log w_i) + d(\log E_i). \tag{19}$$

Substituting (19) into (18) and solving for (18), we obtain

$$(20)\quad d(\log R_i^1) = \frac{1}{k} d(\log R_i) + \frac{1-k}{k} \eta_b d(\log b_i) - \frac{1-k}{k} \eta_w d(\log w_i) - \frac{1-k}{k} d(\log E_i).$$

R_i is allowed to vary to see how changes in the supply of resources in an area influence the other variables.

Equations (20), (14), (16), (17), and (18) can be manipulated to see how variations in the two exogenous variables, A_i and R_i, influence each of the five endogenous variables, w_i, b_i, E_i, K_i, and R_i^1. The determinant resultant system is

$$(21)\qquad \Delta = \frac{1}{\sigma^2}(\alpha_3 + \beta\alpha_1),$$

which is positive if $\alpha_3 > 0$ (firms use resources) or $\beta > 0$ (households use resources). This implies that there will be a unique solution of the endogenous variables of the model unless land, water, and the like are in infinite supply in each area.

An increase in the aggregate supply of resources in an area has no effect on the equilibrium levels of w_i and b_i. The two variable prices are only determined by the shift parameter A_i, that is

$$(22)\qquad \frac{\partial(\log w_i)}{\partial(\log A_i)} = -\frac{\alpha_3}{\alpha_3 + \beta_1}\frac{1-\beta}{\theta},$$

and

$$(23)\qquad \frac{\partial(\log b_i)}{\partial(\log A_i)} = \frac{\alpha_1}{\alpha_3 + \beta\alpha_1}\frac{1-\beta}{\theta}.$$

The effect of changes in the two exogenous variables on the level of employment is seen to be

$$(24)\qquad d(\log E_i) = -\frac{(1-k)(\alpha_1\eta_b + \alpha_3\eta_w) + k\sigma(\alpha_1 + \alpha_3)}{\alpha_3 + \beta\alpha_1} \frac{1-\beta}{\theta} d(\log A_i) + d(\log R_i).$$

Although w_i is an endogenous variable in the model, a quasi-elasticity of labor demand, $\partial(\log E_i)/\partial(\log w_i)$ holding R_i constant, can be obtained by dividing the coefficient on $d(\log A_i)$ in (24) by the negative of (22). This yields

$$(25)\qquad -\frac{\partial(\log E_i)}{\partial(\log w_i)} = -\frac{(1-k)(\alpha_1\eta_b + \alpha_3\eta_w) + k\sigma(\alpha_1 + \alpha_3)}{\alpha_3},$$

which is finite if $\alpha_3 > 0$, i.e., if firms use scarce natural resources as well as capital and labor.

The preceding model is merely an attempt to justify why I assume that there *could* be a regional labor market equilibrium with different area wage levels. It could also be extended in several directions—addition of variations in the prices of other goods due to transportation costs, taxes, the possibility of movement upon retirement, and the like.

8.3 Data and Initial Results

In order to estimate wage differentials between regions, it is necessary to adjust for the other factors that influence wages. To do this I shall employ the standard technique for analyzing the determinants of wages from cross-sectional data: the earnings function. The hourly wage of each worker is assumed to depend on four sets of variables: (a) *skill*, (b) *compensating*, (c) *discrimination*, and (d) *rent* variables. In terms of the CPS data I shall use in the analysis, the specification of the model is, for each sex,

$$\log w = \alpha_0 + \alpha_1 S + \alpha_2 X + \alpha_3 X^2 + \alpha_4 \mathrm{BL} + \alpha_5 \mathrm{OTH} + \alpha_6 U + \alpha_7 \mathrm{PUB} + \alpha_8 U \times PUB + \sum_{i=1}^{I-1} \gamma_i \mathrm{AR}_i + \sum_{j=1}^{J-1} \mu_j \mathrm{IND}_j + \sum_{k=1}^{K-1} \gamma_k \mathrm{OCC}_k + \epsilon, \tag{26}$$

where:

W = hourly nominal wage rate of a person
S = years of schooling attended
X = years of potential labor market experience (age $- S - 5$)
BL = one/zero dummy variable for blacks
OTH = dummy variable for race other than black or white
U = dummy variable for union membership
PUB = dummy variable for public employment
AR_i = set of dummy variables for geographic location
IND_j = set of dummy variables for industry
OCC_k = set of dummy variables for occupation.

The skill variables are proxies for the individual's stock of human capital and typically include S, X and its square, as well as some measure of innate ability. The Current Population Survey data set, however, includes no estimate of ability. $\alpha_1 = \partial(\log W)/\partial S \approx (W_S - W_{S-1})/W_{s-1}$ is (approximately) the rate of return to schooling without allowance for its resource cost. Past earnings function estimates have always found that

$\alpha_2 > 0$ and $\alpha_3 > 0$, presumably reflecting a diminishing rate of investment in human capital over the life cycle.

The compensating variables include several factors. First, W is measured in nominal terms, so, other things equal, wages should vary more or less in proportion to the price level of the region of residence of the worker. Second, some jobs are more onerous or dangerous than others, and persons in the "bad" jobs should receive a compensating differential. Similarly, employment in certain industries is subject to severe seasonal (e.g., construction) or cyclical (e.g., durable goods manufacturing) fluctuations, so persons in these industries should receive a higher hourly wage than persons in industries with secure employment. Third, areas that are attractive in terms of climate, physical characteristics, the net quantity and quality of public services, and the like should offer lower wages, ceteris paribus, than unattractive places.

In terms of the CPS data set, the second set of factors may be proxied (albeit somewhat imperfectly) by the industry and occupational dummy variables (IND and OCC). To a certain extent, however, these variables, especially the latter, are proxies for skill and luck, and I will present estimates of the major coefficients based on a *basic model* (without IND and OCC) and a *full model* (including them).

The first and third sets of factors are related to the interpretation of the set of coefficients that is central to this paper, those on the AR variables. The coefficient γ_i is the logarithmic difference, after accounting for the influence of the other variables, between the wage level in area i and the (arbitrarily excluded) area l. Thus, the ratio of what a person would earn in area i relative to area i' is $W_i/W_i' = \exp(\gamma_i - \gamma_i')$. The *nominal* area effect is γ_i, and the *real* area effect is γ_i minus some function of the area price level, p_i. The discussion in section 8.2 suggested that, in the absence of either government transfers or the possibility of postretirement migration, the appropriate function is $\log p_i$. If one used $\log W - \log p_i$ instead of $\log W$ in the earnings function, the area coefficients would be interpreted in real rather than in nominal terms—exactly what one would get by subtracting $\log p_i$ from the estimated nominal area effects. Since it is useful to hold open the question of how variations in area prices influence equilibrium area wage levels, I will estimate the earnings function in nominal terms.

The discrimination variables are represented in the United States by sex and race. Because (1) the preferences for different types of jobs may differ between men and women and (2) the potential experience variable is a much worse proxy for actual experience for women than for men, the model is estimated separately for the two sexes. Some of the difference between the predicted earnings of men and women for a given set of values of the independent variables may represent direct labor market discrimination against women, but it is impossible to tell how much.

Similarly, differences in the area coefficients between the sexes are consistent with both differential degrees of labor market discrimination against women and differences in the tastes of men and women for particular areas.

The coefficients on BL and OTH in equation (26) represent the logarithmic difference between the wages of each group relative to whites, other factors held constant. Thus, blacks earn $\exp(\alpha_4)$ of the wages of whites with the same observed qualifications, or, by one interpretation, employers behave as if they taxed black workers by $1 - \exp(\alpha_4)$ of their wage bill. This specification assumes that the proportional black/white differential is identical in all regions.[8]

The principal rent variable is unionism, and much attention in labor economics has focused on estimation of the union/nonunion relative wage advantage, $\exp(\alpha_6) - 1$, for private sector employees. In addition, it is possible that public employees earn more or less than their private sector counterparts, and the union/nonunion wage differential may be different in the public and private sectors. These last two possibilities can be tested by seeing if α_7 and α_8 in (26) are significantly different from zero. As with the case of the race variables, equation (26) assumes that unionism has the same proportionate impact on the wage in all areas. It is possible, however, that unions create a national wage scale, implying that the coefficients will vary less for union workers than for nonunion workers. To test this hypothesis, the model can be run separately for union and nonunion workers.

The data on which equation (26) is estimated are from the Current Population Survey for May of 1973 through 1976. The sample consists of all persons during each sample week who were (a) employed (but not self-employed or farmers), (b) between the ages of 17 and 72, inclusive, (c) had a positive wage, (d) were employed on a full-time basis, and (e) resided in one of the thirty-four large Standard Metropolitan Statistical Areas (SMSA) that are identifiable in the data set. In testing for the consistency of the results for the two wage measures, it became clear that for both men and women the estimated parameters of the basic model differed greatly with respect to the method by which the individual was paid (hourly versus salaried). Thus, the total sample of 43,940 persons during the four years was divided into four subsamples: (A) male hourly, (B) female hourly, (C) male salaried, and (D) female salaried.

The estimated coefficients of the basic model (that does not include industry and occupational dummy variables) are presented in table 8.1. These regressions do include thirty-three dummy variables for SMSA (Detroit is the excluded area), and these coefficients are discussed below. The results on the skill variables suggest that schooling and potential experience have a greater effect on the earnings of salaried workers than on those of hourly workers. As expected on the basis of several past

Table 8.1 **Estimated Coefficients for Basic Model (estimated standard errors in parentheses)**

Variables	Hourly (A) Men	Hourly (B) Women	Salaried (C) Men	Salaried (D) Women
S	.039	.052	.074	.069
	(.001)	(.002)	(.001)	(.002)
X	.030	.014	.042	.022
	(.001)	(.001)	(.001)	(.001)
X^2	−.00049	−.00022	−.00067	−.00038
	(.00002)	(.00002)	(.00002)	(.00002)
BL	−.146	−.028	−.158	−.088
	(.009)	(.011)	(.012)	(.012)
OTH	−.167	−.057	−.174	−.045
	(.025)	(.028)	(.024)	(.026)
U	.252	.184	.019	.078
	(.006)	(.010)	(.010)	(.015)
PUB	.151	.145	−.013	.136
	(.014)	(.014)	(.010)	(.010)
PUB × U	−.178	−.012	.007	−.050
	(.018)	(.026)	(.017)	(.021)
$D74$	.079	.095	.062	.072
	(.008)	(.011)	(.009)	(.010)
$D75$	.150	.165	.134	.137
	(.008)	(.011)	(.009)	(.010)
$D76$	.204	.237	.201	.208
	(.008)	(.011)	(.009)	(.010)
Constant	.627	.276	.325	.245
	(.02)	(.03)	(.03)	(.03)
R^2	.353	.277	.344	.330
SEE	.31	.31	.39	.35
N	12,191	6,760	15,355	9,634

studies, experience is a much more important determinant of earnings for men than for women. The estimated differential between the wages of white and nonwhite workers is larger for men than women.

The estimated effect of union membership on the earnings of full-time workers is quite large (a 28.7 percent advantage for men and a 20.2 percent advantage for women in the private sector), but it is much smaller for full-time salaried workers (1.9 percent for men and 8.1 percent for women). Being employed in the public sector increases the wages of women workers and male hourly workers by approximately 15 percent, but it has no effect on the earnings of salaried males. The estimated impact of public sector unions on wages, the sum of the coefficients on U and PUB × U, is, for full-time workers, greater for women than men and for hourly than salaried workers.

The predicted hourly wage rates of a typical worker—white, private sector, twelve years of schooling, age 35—in Detroit in 1973 in each of the four subsamples are as follows:

	Predicted Hourly Wage		
	Nonunion	Union	Percent Private Sector Unionized
A. Full-time hourly male	4.31	5.54	56
B. Full-time hourly female	2.90	3.49	22
C. Full-time salaried male	5.67	5.78	17
D. Full-time salaried female	3.78	4.09	9

Notice that the only group which private sector union membership is extensive is A, hourly males. In fact, whereas group A workers compose only 28 percent of total private sector employment, they have 58 percent of total private sector union membership. The impact of unionism on the wage rates of this group is such that a unionized male hourly worker has almost as high a wage as a salaried male worker.

8.4 Differences in Wage Levels between Areas

The four regressions in table 8.1 also include thirty-three dummy variables for SMSA of residence of the individual. The null hypothesis that the presence of these does not add sufficiently to the explanation of log w to justify the sacrifice of 33 degrees of freedom—i.e., that nominal wage levels for each of the groups do not differ among the thirty-four areas—is decisively refuted for the four groups (F values between 9 and 15 compared to $F_{.05}(33, \infty) = 1.44$).

Table 8.2 reports the point estimates of the coefficients on the area variables for the four groups of full-time employees. The "basic" model refers to the standard earnings functions, whose other coefficients were given in table 8.1, and the "full" model includes dummy variables for both industry and occupation at the one-digit aggregation. Each of the coefficients represents the estimated logarithmic deviation of the area effect for that area less that for Detroit. Thus, from the basic model, male hourly workers in New York, given their education, experience, race, union membership, and public/private status, earn $\exp(-.109) = 89.7$ percent of what comparable workers in Detroit earn. They earn $\exp(-.109 + .249) = 115.0$ percent of what comparable workers in Tampa earn. The estimated standard errors of the differences in the area coefficients range from a low of about .015 for areas with large samples of a subgroup to almost .050 for areas with very small samples. For example, the New York/Detroit relative for group A workers has a standard error

Table 8.2 **Estimated Logarithmic Difference between Wage Levels in Thirty-four SMSA's and Detroit for Full-Time Hourly and Salaried Workers, by Sex, 1973–76**

	Hourly				Salaried			
	Men		Women		Men		Women	
SMSA	Basic	Full	Basic	Full	Basic	Full	Basic	Full
East								
1. New York	−.109	−.116	.031	−.049	−.144	−.104	−.005	−.015
4. Philadelphia	−.082	−.116	−.071	−.180	−.128	−.117	−.160	−.172
8. Boston	−.103	−.146	−.016	−.020	−.158	−.129	−.077	−.090
9. Nassau-Suffolk	.020	−.033	.008	−.088	−.033	.002	−.046	−.047
10. Pittsburgh	−.183	−.187	−.172	−.129	−.188	−.164	−.172	−.184
15. Newark	−.110	−.124	−.087	−.078	−.094	−.086	−.139	−.140
23. Paterson	−.048	−.085	.010	−.118	.016	.032	−.033	−.025
25. Buffalo	−.161	−.143	−.098	−.149	−.230	−.200	−.264	−.248
Midwest								
3. Chicago	−.009	−.036	.018	−.026	−.056	−.034	−.049	−.066
5. Detroit	.000	.000	.000	.000	.000	.000	.000	.000
11. St. Louis	−.058	−.088	−.096	−.121	−.117	−.120	−.178	−.172
13. Cleveland	−.075	−.097	−.097	−.149	−.081	−.072	−.164	−.160
16. Minneapolis	.000	−.038	−.025	−.087	−.059	−.044	−.137	−.143

20. Milwaukee	−.038	−.051	−.021	−.022	−.089	−.081	−.117	−.135
22. Cincinnati	−.119	−.145	−.096	−.202	−.227	−.209	−.261	−.234
27. Kansas City	−.083	−.107	−.132	−.166	−.192	−.168	−.226	−.211
30. Indianapolis	−.080	−.116	−.068	−.054	−.127	−.124	−.214	−.201
South								
7. D.C.	.071	−.012	.003	−.015	.018	.007	−.032	−.135
12. Baltimore	−.070	−.116	−.038	−.050	−.151	−.138	−.150	−.146
14. Houston	−.031	−.120	−.167	−.116	−.060	−.057	−.215	−.216
17. Dallas	−.168	−.202	−.098	−.189	−.115	−.103	−.239	−.218
21. Atlanta	−.073	−.111	−.040	−.142	−.112	−.096	−.126	−.130
26. Miami	−.149	−.179	−.141	−.132	−.285	−.232	−.190	−.179
32. New Orleans	−.115	−.190	−.227	−.144	−.142	−.140	−.290	−.274
33. Tampa	−.249	−.283	−.216	−.264	−.300	−.278	−.303	−.290
West								
2. Los Angeles	−.103	−.109	−.015	−.076	−.115	−.087	−.087	−.050
6. San Francisco	.024	.020	.034	.061	−.089	−.058	−.044	−.049
18. Seattle	−.066	−.077	−.075	−.094	−.093	−.082	−.169	−.135
19. Anaheim	−.053	−.073	−.060	.004	−.069	−.063	−.120	−.128
24. San Diego	−.089	−.127	−.105	−.079	−.187	−.168	−.158	−.149
28. Denver	−.028	−.083	−.081	−.082	−.144	−.141	−.143	−.133
29. San Bernardino	−.071	−.099	−.146	−.175	−.212	−.059	−.227	−.186
31. San Jose	.011	−.003	.017	.053	−.071	−.076	−.061	−.045
34. Portland	−.067	−.075	−.074	.000	−.182	−.146	−.163	−.131

NOTE: Numbers before SMSAs indicate population ranking in 1976.

of 1.7 percent, the New York/Tampa relative a standard error of 2.9 percent.

To adjust these estimated nominal wage effects for prices, I used the BLS index of comparative living costs for an "intermediate" living standard for 1974.[9] Most of the variation in this price index is due to variations in housing costs and taxes.[10] The elasticity of the equilibrium wage level in a region with respect to the local price level may be, as shown in section 8.2, greater or less than one, but I initially constrained its impact to be one. Thus, the estimated real area effect is the estimated nominal effect less the logarithm of the price level, and these are shown for hourly workers in table 8.3. (For salaried workers the real area effect can be calculated by taking the nominal area effect in table 8.2 and subtracting the natural logarithm of the price level—normalized at one rather than 100.) To obtain the estimated average area effect, one averages the union and nonunion wage levels, which is done by adding the estimated coefficient on union membership (.25 for men and .18 for women) times the proportion of hourly private sector workers who are unionized. This is then exponentiated and normalized at Detroit equal to 100.

One index of the dispersion of area wage rates is the (weighted) standard deviation of the logarithm of the area effects. This is shown in table 8.4 for the nonunion area effect and the average (i.e., including unionism) for both nominal and real area effects. (Since unionism has little direct impact on the wages of salaried workers, the dispersion of average wages of salaried workers is not reported.) One might suspect that the dispersion of real area effects would be lower than the dispersion of nominal area effects if workers move between areas to equalize net returns. In fact, this is true for both groups of women workers, but it is *greater* for both groups of men workers.

What this implies is difficult to tell. First, the dispersion of real area effects will only be lower than the dispersion of nominal area effects if the slope coefficient of a regression of the nominal area effect on the log of area price level is greater than .5.[11] As shown in section 8.2, $\partial(\log w_i)/\partial(\log p_i)$ can be less than one even in a world of real income-maximizing suppliers of labor. Further, to the extent that area price levels are measured with error, the implicit coefficient of log w_i on log p_i will be biased down, thus increasing the estimated variance of $\log(w_i/p_i)$.

If there are no threat or contamination effects, the nonunion wage for a group for which there is a significant union effect will depend negatively on the extent of unionism in the area. Since the estimated union/nonunion effect is positively only for hourly workers, we would therefore expect that the extent of unionism would have a negative effect on the area effect for nonunion wages. On the other hand, if either the threat or contamination effect were operative, we would expect that the relation between the nonunion area effect and unionism would be much weaker—posibly even positive. For male hourly workers the coefficient of the logarithm of the real area effect for nonunion workers on the extent of unionism is −.15 (.09). This implies that the real nonunion wage level is about 9 percent lower in a heavily unionized area (about 75 percent) than in a weakly

Table 8.3 **Estimated Real Wage Levels (w/p) for Private Sector Hourly Workers in Thirty-three SMSA's (Detroit = 100)**

		Men		Women	
SMSA	Price Level	Nonunion Wage	Average Wage	Nonunion Wage	Average Wage
East					
1. New York	116	77	85	76	90
4. Philadelphia	103	89	88	85	90
8. Boston	117	77	71	84	81
9. Nassau-Suffolk	116	88	84	87	86
10. Pittsburgh	97	86	87	86	86
15. Newark	116	77	75	79	81
23. Paterson	116	82	79	87	88
25. Buffalo	107	80	85	79	83
Midwest					
3. Chicago	103	96	99	94	98
5. Detroit	100	100	100	100	100
11. St. Louis	97	97	98	94	93
13. Cleveland	102	91	89	90	88
16. Minneapolis	104	96	94	92	93
20. Milwaukee	105	92	93	90	92
22. Cincinnati	96	92	88	95	91
27. Kansas City	97	95	89	92	89
30. Indianapolis	99	93	94	90	92
South					
7. D.C.	105	102	98	93	93
12. Baltimore	100	93	89	96	93
14. Houston	90	108	97	87	91
17. Dallas	90	94	95	83	97
21. Atlanta	91	102	91	106	102
26. Miami	89	97	87	98	94
32. New Orleans	90	99	91	89	85
33. Tampa	89	88	76	91	87
West					
2. Los Angeles	98	92	86	101	99
6. San Francisco	106	97	98	95	99
18. Seattle	101	93	92	92	92
19. Anaheim	98	97	89	96	93
24. San Diego	98	93	89	92	89
28. Denver	95	102	95	97	94
29. San Bernardino	98	95	92	88	84
31. San Jose	106	95	90	96	94

NOTE: Numbers before SMSAs indicate population ranking in 1976.

Table 8.4 Standard Deviation of Estimated Area Wage Effects for Four Subgroups

	Average Wage		Average Wage	
	Nominal	Real	Nominal	Real
Male hourly	.066	.084	.085	.082
Female hourly	.070	.061	.080	.057
Male salaried	.077	.085	—	—
Female salaried	.084	.069	—	—

unionized area (about 15 percent), but this estimate is subject to a large standard error. For female hourly workers this coefficient is $-.20$ (.10), which, given the range of the extent of unionism, is actually a smaller effect.

One implication of both the threat and contamination effect models of the spillover of union wages to the nonunion sector is that the nonunion wage will depend positively on the extent of organization only when the extent of unionism is fairly large. The threat or contamination effect should be fairly small (or nonexistent) until unionism reaches a certain proportion—then they will be rather extensive. To test this I added quadratic terms so that the logarithm of the nonunion real area effect was a function of U and U^2 for both men and women hourly workers. The above argument would suggest that their coefficients would be negative and positive, respectively. The results, for both men and women, were the opposite—although the negative coefficient on the U^2 term dominates the positive coefficient on U. Adding dummy variables for certain "troublesome" areas (specifically, Detroit, which seems to be a special case, and Washington, D.C., which is dominated by a high-rent public sector), did nothing to upset the conclusion that there is little to the spillover hypotheses.

Notes

1. See, e.g., Scully (1969), Coelho and Ghali (1971), Bellante (1979), and Goldfarb and Yeager (1981).

2. It is, of course, possible that this migration takes a long time to occur so that there is at any time a large *disequilibrium* component in any observed distribution of regional wages. In the United States, however, there is a considerable amount of interarea labor mobility. For example, between March 1969 and March 1979, 2.9 percent of the male population from 35 to 44 years of age moved between states; for males ages 45 to 64 the mobility rate was 1.5 percent. Now even if some of this mobility is not related to economic migration (e.g., a move from Cos Cob, Connecticut to Short Hills, New Jersey), there appears to be enough movement to eliminate disequilibrium rather quickly.

3. Specifically, the federal government offers wage rates that are considerably in excess of the reservation prices of potential employers. For an empirical study of this phenomenon, see Smith (1976).

4. The original application of the Harris-Todaro framework to the explanation of regional differentials in unemployment was Hall (1970).

5. The original formulation of the "threat effect" model was done by Rosen (1969).

6. For a discussion of some further implications of the assumption of "interdependence" of different workers' utility functions, see Hamermesh (1975).

7. This point is stressed in Richard Muth's comment on this paper. The following section is a reply to that part of his comments.

8. It is, of course, likely that this assumption is not correct, as indeed has been demonstrated with respect to broad regional groups by Kiefer and Smith (1977). I did run my basic regressions separately for blacks and whites, but, despite the fact that estimated differentials between blacks and whites were somewhat larger in southern SMSA's than elsewhere, it made very little difference in the estimation of area effects.

9. The price levels for New York, Nassau-Suffolk, Newark, and Paterson are the New York/New Jersey rate; Miami and Tampa are assigned the level for Orlando; Anaheim and San Bernadino, the Los Angeles level (which equals the San Diego level); New Orleans, the Baton Rouge level; and San Jose, the San Francisco/Oakland level. No price level was available for Portland, Oregon.

10. The standard deviation of the logarithm of the housing component of the index for 1974 was .132; for taxes it was .204. All other categories—except medical, which had a relatively low weight—varied by .06 or less.

11. Let A_i and N_i be the real and nominal area effects and p_i the logarithm of the area price level. Then $A_i = N_i - p_i$, and $\text{var}(A_i) = \text{var}(N_i) + \text{var}(p_i) - 2\,\text{cov}(N_i, p_i) = \text{var}(N_i) + \text{var}(p_i)(1 - 2b_{N \cdot p})$. Thus, $\text{var}(A_i) < \text{var}(N_i)$ only if $b_{N \cdot p} > .5$.

References

Bellante, Don. 1979. The North-South differential and the migration of heterogeneous labor. *American Economic Review* 69: 166–75.

Coelho, Philip R. P., and Moheb A. Ghali. 1971. The end of the North-South wage differential. *American Economic Review* 61: 932–37.

Freeman, Richard B., and James L. Medoff. 1979. New estimates of private sector unionism in the U.S. *Industrial and Labor Relations Review* 32: 143–74.

Goldfarb, Robert S., and Anthony Yeager. 1981. Has the North-South wage differential really disappeared?—Some new evidence. George Washington University. Mimeo.

Hall, Robert E. 1970. Why is the unemployment rate so high at full employment? *Brookings Papers on Economic Activity* 3:369–402.

Hamermesh, Daniel S. 1975. Interdependence in the labor market. *Economica* 42: 420–29.

Kiefer, Nicholas M., and Sharon P. Smith. 1977. Union impact and wage discrimination by region. *Journal of Human Resources* 12:521–34.

Rosen, Sherwin. 1969. Trade union power, threat effects, and the extent of organization. *Review of Economic Studies* 36: 185–96.

Scully, Gerald W. 1969. Interstate wage differentials: A cross-section analysis. *American Economic Review* 59: 757–73.
Smith, Sharon P. 1976. Pay differentials between federal government and private sector workers. *Industrial and Labor Relations Review* 29: 179–97.

Comment Richard F. Muth

George Johnson's paper is the most thorough, carefully done study of intercity wage differentials that I have seen. Because of the great detail in the data set he uses, Johnson is able to eliminate the effects of a variety of other influences on wages and salaries which other studies have only partially controlled for. Among these other factors are schooling, potential labor force experience, race, union membership, public employment, industry, and occupation. Moreover, separate regressions are run for men and women and for hourly and salaried workers. The area effects which Johnson estimates are therefore remarkably free from the effects of possible correlation with omitted variables.

Yet, because of the attention given earlier in the paper to the effects of intercity differences in the prices of consumer goods on nominal wage levels in equilibrium, I was somewhat surprised that Johnson didn't include the BLS intercity living cost index as an explanatory variable in his nominal wage regressions. Earlier he argues correctly that progressive taxation of nominal earnings and social security payments upon retirement may make the partial derivative of the equilibrium wage rate with respect to consumer goods prices greater than unity. The option of retirement to a low-price area coupled with moving costs can make this partial derivative either greater or less than one. Surely then, it is not correct to divide the estimated area effect on nominal wages by the area's price index to obtain the real area effect, as Johnson does. Indeed, his finding of a greater area dispersion of real area effects for men than for nominal effects may merely reflect the fact that the former were incorrectly estimated.

In the earlier part of his paper, Johnson devotes considerable space to the effect of nonpecuniary advantages and disadvantages on equilibrium wage levels. I would have found it interesting if measures of such effects had been included among the explanatory variables and their effects on area wage levels calculated. A variety of such variables have been included in various spatial studies, especially studies of intermetropolitan migration. Variables related to weather, such as heating degree days (essentially the absolute difference between average temperature and

Richard F. Muth is professor of economics at Stanford University.

some level such as 50 degrees Fahrenheit if the former is smaller summed over the year), cooling degree days (similarly defined), humidity, and annual rainfall, are obvious candidates for inclusion. Other such variables might include proximity to oceans and to mountains and the presence of a symphony orchestra, or, if my tastes are indicative, a professional football team. Not only would the effects of such factors on (supposedly equilibrium) wage levels be of considerable interest but any remaining areal effects would approximate disequilibrium wage differences.

My greatest single criticism of Johnson's paper, however, is its neglect of demand-side variables. Implicit in the paper, it seems to me, is the hypothesis that area demand curves for labor are downward sloping to the right. This would be the case if products produced in a particular place were either unique or sold to a limited market area surrounding the city in which they are produced. If such were the case and there were no important differences among workers in the value placed upon nonpecuniary advantages, long-run horizontal labor supply curves would fix equilibrium wage levels. Forces influencing the area's demand curve for labor would then affect its total population and employment but would have no influence on long-run equilibrium wage levels. The above, it seems to me, is the predominant view among urban and regional economists, and Johnson's paper is certainly consistent with it.

There is an alternative view, however, attributable to Borts and Stein (1964), which suggests demand factors are all important in determining areal wage levels. Suppose that in any urban area there exist firms in significant numbers which, in the aggregate, have a negligible effect on the prices for their products which prevail on national or world markets. Furthermore, let the rental values of capital be fixed by external conditions to firms in the urban area, a condition most urban and regional economists would accept. Then, if production functions are homogeneous of degree one in labor and capital,[1] there exists a fixed nominal wage level in the urban area consistent with equilibrium for producers facing fixed product prices and selling in outside markets. Competition for labor would require that firms producing for domestic consumption or others facing downward sloping product demand schedules pay the same wage. The latter would be affected only by changes in capital rental values or changes in f.o.b. export prices of firms facing perfectly elastic product demand schedules, and not by the total level of employment. Factors affecting labor supply schedules, which would have to be upward sloping for equilibrium to be determinate, would, under these conditions, influence only an area's total population and employment; they would have no impact on nominal wages. Admittedly, this alternative view is a minority one among urban and regional economists. My own work (Muth 1968), though, suggests it more closely approximates the U.S. economy during the 1950s than the more conventional view sketched above.

Johnson certainly can't be faulted for not having studied intermetropolitan wage differences over time, for the basic data to be used probably don't exist. Yet, changes in wage differentials over time are of even greater interest than their level at a moment in time. One of the most striking features of the U.S. economy is the convergence of per capita incomes over time, especially the increase in the South relative to the rest of the United States. My earlier analysis of Easterlin's data (Perloff et al. 1960, chap. 28) together with more recent examinations of Census earnings data for white urban males for 1950–70 suggest that regional earnings differentials have been remarkably constant for half a century. Rising relative per capita incomes in the South would appear to have resulted primarily from a declining relative importance of agriculture in southern states. It would be nice, though, if we had a study as well done as Johnson's for some earlier period, such as 1950, from which we could better appraise the influence of changes in wage differentials on the regional covergence of per capita incomes.

Note

1. In the revised version of his paper, Johnson correctly argues that resources as an input into production make the demand for labor less than perfectly elastic. If resources as an input into the production of exportable commodities are relatively unimportant, however, the less than perfectly elastic demand may be of little practical significance.

References

Borts, George H., and Jerome L. Stein. 1964. *Economic growth in a free market*. New York: Columbia University Press.

Muth, Richard F. 1968. Differential growth among large U.S. cities. In *Papers in quantitative economics*, ed. James P. Quirk and Arvid M. Zarley. Lawrence, Kansas: The University Press of Kansas.

Perloff, Harvey S. et al. 1960. *Regions, resources, and economic growth*. Baltimore: The Johns Hopkins Press.