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ABSTRACT

Factor Price Equality and the Economies of the United States*

We develop a methodology for identifying departures from relative factor price equality across regions that is valid under general assumptions about production, markets and factors. Application of this methodology to the United States reveals substantial and increasing deviations in relative skilled wages across labour markets in both 1972 and 1992. These deviations vary systematically with labour markets' industry structure both in cross section and over time.

JEL Classification: C14, F16, J30 and R23
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Submitted 06 June 2005

1. Introduction

Variation in relative factor prices across labor markets is influential in determining workers' susceptibility to international trade shocks, regional income convergence and the spatial location of industries. Persistence in relative factor reward variation over time also sheds light on the degree to which factor mobility is sufficient to arbitrage away wage gaps. This paper develops a methodology for identifying relative factor price differences across regions and uses it to test whether relative skilled wages are equal in the United States' 181 labor markets.

Identifying relative factor price equality is a difficult problem for two reasons. First, any methodology must account for the possibility that factors vary in terms of unobservable quality or composition across labor markets. Regions with superior educational systems or worker training programs, for example, might possess higher-productivity skilled workers than regions without these attributes, thereby inducing higher observed relative skilled wages even if quality-adjusted skill premia are equal. Second, a useful methodology must correctly identify failures of relative factor price inequality in the face of variation in market structure across industries and regions. Such variation in market structure is difficult for econometricians to discern.

Tests of relative factor price inequality across countries are common in the international trade literature.[1] However, the scarcity of internationally comparable wage data has motivated the creation of tests that verify the implications of relative wage variation (e.g. production specialization) rather than differences in relative wages directly.[2] The outcomes of these tests suggest significant relative factor price differences across developed and developing economies. This paper focuses on relative factor price equality within a single country, which is generally thought to be more likely.

In contrast with the international trade literature, we develop a methodology for identifying departures from relative factor price equality that is based upon general optimality conditions for producer equilibrium. This methodology possesses a number of important advantages over traditional methods. First, it is valid under

[1] Empirical tests of factor price equality across countries include Trefler (1993), Repetto and Ventura (1998), Davis and Weinstein (2001), Cunat (2000), Debaere and Demiroglu (2003) and Schott (2003). Tests for factor price equality within countries include Davis et al. (1997) and Debaere (2004) who study prefectures in Japan, Debaere (2004) who examines administrative regions in the United Kingdom, and Hanson and Slaughter (2002) who analyze U.S. states.

[2] Theoretical conditions necessary for factor price equality have been explored by Samuelson (1949), McKenzie (1955), Dixit and Norman (1980), Wu (1987), Courant and Deardorff (1994) and Deardorff (1994).

a wide range of assumptions regarding production, markets and factors, including imperfect competition and increasing returns to scale. Second, because it makes no assumptions about the preferences and costs of living faced by different groups of workers, it is robust to unobserved variation in consumer price indices specific to locations. Third, it controls for a variety of measurement issues that can cause observed factor prices to vary even if true, unobserved factor prices are identical, in particular region-factor-industry variation in the quality or composition of factors. Finally, it is easy to implement and can be used in a variety of contexts. The only data required are total payments to factors by industry and region.

Application of our methodology to the United States reveals several surprising facts about the geographic variation of quality-adjusted relative skilled wages in manufacturing over time. We find that the United States' 181 labor markets exhibit statistically significant and economically meaningful differences in non-production worker wages relative to production worker wages in both 1972 and 1992. In 1972, for example, the quality-adjusted skill premium was 30 percent higher in Nashville than in New York City. By 1992, this differential had risen to 36 percent. Overall, labor markets exhibit increasing relative-wage polarization: dividing U.S. labor markets into three groups according to the significance of their relative skilled wage differences in both years, we find that the number of labor markets in the "middle" declines with time as the two groups at either end expand.

The economic importance of these relative wage differences is signalled by their strong relationship to labor markets' industry structure. In cross-section, we find that the larger the difference in two labor markets' relative skilled wages, the smaller the number of industries they produce in common. Within labor markets across time, we find that greater changes in relative skilled wages are associated with a larger number of added and dropped industries: as labor markets' skill premia evolve, their industry mix adapts.

Neoclassical trade theory provides a useful intuition for these trends. In that framework, sufficient heterogeneity of regional factor endowments combined with factor immobility across regions can give rise to an equilibrium in which regions offer different relative factor prices and attract different sets of industries: skill-scarce Nashville offers a high skill premium and attracts skill-scarce industries, skill-abundant New York City offers a low skill premium and attracts skill-intensive industries, and too few workers move between New York and Nashville to bid these relative wage differences away.[3] Though the assumption of labor immobility upon

[3] In the neoclassical model, this outcome is referred to as a multiple cone equilibrium. See, for

which this equilibrium depends is strong, it is not without independent empirical support.[4]

Variation in labor markets' industry participation is noteworthy because it implies potential asymmetric exposure of otherwise identical U.S. workers to domestic and international shocks. In particular, it may insulate unskilled workers in skill-intensive regions from the well-known distributional consequences of trade liberalization implied by the factor proportions framework. Because skill-scarce labor markets are more likely to produce goods in common with labor-abundant trading partners like Mexico and China, the real wages of unskilled workers in these regions may respond more readily – and negatively – to the price declines associated with falling trade costs. Our findings suggest that standard implications of international trade theory may apply within as well as across nations.

Relative factor price inequality is also informative about the possibility of regional income convergence within countries. Research in the macroeconomic literature, for example, has found sluggish equilibration of relative per worker income levels across U.S. regions over time.[5] Those findings suggest that either relative factor endowments or relative factor prices are at best converging slowly. Our demonstration of persistent and increasing relative wage disparities provides evidence of the importance of factor prices, while our use of local labor market areas gives a much higher level of spatial resolution than is typical in the literature.

An alternate interpretation of our results is that they are driven by unobserved variation in the relative cost of living across factors *and* regions. This explanation, in contrast to the neoclassical model, is consistent with perfect labor mobility: if the consumer price index for skilled workers relative to unskilled workers were lower in skill-abundant regions, real relative consumption wages could be equal across labor markets even if nominal relative wages were not. In that case workers would have no incentive to relocate, but production specialization – and asymmetric susceptibility of U.S. workers to macroeconomic shocks – would still occur as industry location depends upon nominal rather than real relative wage differences.[6] Though this explanation does not rely on labor immobility, it does depend upon enduring, and in some cases increasing, differences in factor-region specific costs of living across the

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example, Leamer (1987).

[4] Bound and Holzer (2000), for example, find that imperfect mobility of unskilled workers in the United States contributed toward increased income inequality in the 1980s.

[5] See, for example, Barro and Sala-i-Martin (1991) and Carlino and Mills (1993).

[6] In another contrast with the neoclassical interpretation, this explanation suggests asymmetric shocks will be more readily transmitted across regions via factor mobility.

United States. Because such differences are also in principle subject to arbitrage, the reasons for this persistence are not obvious.

Our methodology and results contribute to the large literature on U.S. income inequality. A number of papers have demonstrated that U.S. skill premia have risen precipitously over the past few decades.[7] These studies generally document trends either for the U.S. as a whole or for relatively aggregate regions or states within the United States.[8] Our examination here of differences in relative wages across the full set of U.S. labor markets is, to our knowledge, unique. It suggests that previously observed increases in aggregate U.S. income inequality may obscure important regional heterogeneity that could be exploited to determine its ultimate cause.

Our paper is organized as follows. Section 2 details the relevant propositions on relative factor price equality and develops their testable implications. In Section 3, we outline our empirical methodology. Section 4 provides an overview of U.S. regional variation and presents results for our test of relative factor price equality in 1972 and 1992. Section 5 offers evidence on the relation between industry structure and factor prices. Section 6 discusses possible explanations for our findings and Section 7 concludes.

2. Relative Factor Price Equality

Factor price equality can be either absolute or relative. If absolute factor price equality holds (AFPE), regions must have identical nominal factor rewards for identical quality-adjusted factors at a point in time. If relative factor price equality holds (RFPE), regions must have identical relative factor rewards for identical quality-adjusted factors even though absolute factor prices may differ.

We devote our theoretical and empirical attention in this paper to a test of *relative* factor price equality for three reasons. First, there is a natural and rich link between variation in regions' relative factor prices and their industry structure, e.g., skill-intensive industries have an incentive to locate in skill-abundant regions. Second, as we demonstrate below, testing for relative factor price equality, unlike absolute factor price equality, is robust to potential variation in production technologies across regions and industries. Finally, a test of relative factor price equality

[7] See, for example, Katz and Murphy (1992) and Juhn et al. (1993).

[8] Topel (1994), for example, documents a rise in U.S. income inequality across nine U.S. Census regions. An exception is Bound and Holzer (2000), which examines relative wage trends within U.S. metropolitan statistical areas (MSAs).

is more stringent in the sense that relative factor prices can be equal even if absolute factor price equality fails. Nevertheless, for interested readers we provide in an Appendix a complementary test for absolute factor price equality.

Our method for identifying departures from factor price equality emphasizes the importance of potential unobserved variation in region-industry-factor quality that can bias traditional wage comparisons. We demonstrate how total payments to each factor, i.e., wagebills, can be exploited to control for this unobserved variation.

2.1. Basic Setting

Let

$$Y_{rj} = A_{rj} F_j \left(S_{rj}, U_{rj}, K_{rj} \right), \tag{1}$$

be a value-added production function for industry j and region r, where A_{rj} is a Hicks-neutral productivity shifter that allows technology to vary across regions and industries and S_{rj}, U_{rj}, and K_{rj} are quality-adjusted inputs of skilled workers, unskilled workers, and capital, respectively. Individual factors enter production through the function F_j which varies across industries but is the same across regions within an industry. Firms in region r and industry j choose factor usage to minimize costs,

$$\begin{array}{ll} \min\limits_{S_{rj}, U_{rj}, K_{rj}} & w_r^S S_{rj} + w_r^U U_{rj} + w_r^K K_{rj} \\ \text{such that} & A_{rj} F_j \left(S_{rj}, U_{rj}, K_{rj} \right) = Y_{rj} \end{array} \tag{2}$$

which defines the total cost function,

$$B_{rj} = A_{rj}^{-1} \Gamma_j (w_r^S, w_r^U, w_r^K) Y_{rj}. \tag{3}$$

In this specification, firms may act either as price-takers in product markets (perfect competition; this section) or choose prices subject to a downward sloping demand curve (imperfect competition; next section). Here we begin by assuming constant returns to scale; later we extend the analysis to allow for internal and external increasing returns to scale. In factor markets, we assume that firms choose employment taking factor prices as given.[9] Though we write down the model with

[9] The analysis is consistent with both competitive factor markets and 'right to manage' models of union behavior, where firms and unions bargain over wages within an industry but firms choose employment (see, for example, Farber 1986 and Layard et al. 1991). For clarity of exposition, we focus on the competitive case in the text, where wages are equalized across industries. With industry-specific bargaining, wages will generally vary across industries. Inter-industry wage differentials are consistent with our approach, which only exploits variation across regions within industries. We return to discuss this point further in the empirical section below.

three factors of production, the analysis can be extended to an arbitrary number of industries or factors.

Let a tilde (~) signify observed quantities that have not been adjusted for quality, and let θ_{rj}^{z} denote a quality adjustor for industry j, region r and factor z. Note that θ_{rj}^{z} allows for unobserved variation in quality that is specific to factors, regions and industries. The quality-adjusted employment level and wage of factor $z \in (S, U, K)$ in region r equals the observed variable scaled by the quality adjuster, i.e.

$$z_{rj} = \theta_{rj}^{z} \widetilde{z}_{rj} \quad \text{and} \quad w_{rj}^{z} = \widetilde{w}_{rj}^{z} / \theta_{rj}^{z}. \tag{4}$$

Without loss of generality, assume there are two regions, r and b, where region b is the reference region whose factors are taken to be the baseline quality benchmarks, i.e. $\theta_{bj}^{z} = 1$.

The demand for quality-adjusted factor z may be obtained using Shephard's Lemma,

$$z_{rj} = A_{rj}^{-1} Y_{rj} \frac{\partial \Gamma_j(\cdot)}{\partial w_r^z}. \tag{5}$$

Dividing one first-order condition by another provides an expression for the relative demand for any two quality-adjusted factors of production. The relative demand for skilled workers in terms of unskilled workers is

$$\frac{S_{rj}}{U_{rj}} = \frac{\partial \Gamma_j(\cdot)/\partial w_r^S}{\partial \Gamma_j(\cdot)/\partial w_r^U}. \tag{6}$$

Notice that terms in region-industry productivity, A_{rj}, do not appear in equation (6) as the direct effect of variation in technology on the marginal revenue product is identical for each factor. Similarly, region-industry variation in relative goods prices has symmetric direct effects on the marginal revenue product of every factor and thus does not affect the relative factor demands.[10] Using the relationship between quality-adjusted and observed values in (4), this implies the following relative demand for observed factors of production,

$$\frac{\widetilde{S}_{rj}}{\widetilde{U}_{rj}} = \frac{\theta_{rj}^{U}}{\theta_{rj}^{S}} \frac{\partial \Gamma_j(\cdot)/\partial w_r^S}{\partial \Gamma_j(\cdot)/\partial w_r^U}. \tag{7}$$

[10] In general equilibrium, variation in goods prices or technologies across regions and industries can cause variation in relative factor prices; see Section 6 for further detail. Our test correctly rejects relative factor price equality in such cases.

Under the null of relative factor price equality, quality-adjusted relative wages and factor usage across regions r and b must be equal[11],

$$(\mathrm{H}_0\text{: RFPE}) \qquad \frac{w_r^S}{w_r^U} = \frac{w_b^S}{w_b^U} \quad \text{and} \quad \frac{S_r}{U_r} = \frac{S_b}{U_b}, \tag{8}$$

where the second equation follows directly from equation (6).[12]

Observed relative wages and observed factor usage in the two regions, under the null of RFPE, are given by

$$\frac{\tilde{w}_r^S}{\tilde{w}_r^U} = \frac{\theta_{rj}^S}{\theta_{rj}^U}\frac{\tilde{w}_b^S}{\tilde{w}_b^U} \quad \text{and} \quad \frac{\widetilde{S}_{rj}}{\widetilde{U}_{rj}} = \frac{\widetilde{S}_{bj}/\theta_{rj}^S}{\widetilde{U}_{bj}/\theta_{rj}^U}. \tag{9}$$

These relationships demonstrate the difficulty of using either observed relative wages or observed factor usages to test for factor price equality: even under the null hypothesis of RFPE, observed relative wages and observed usages can vary across regions within industries if there are differences in unobserved factor quality (i.e. $\theta_{rj}^S \neq 1$ or $\theta_{rj}^U \neq 1$).[13] We solve this problem by combining observed wages and employment into wagebills, where the wagebill for factor z is equal to $w_{rj}^z z_{rj} \left(= \tilde{w}_{rj}^z \tilde{z}_{rj}\right)$. As is evident from equation (9), multiplying wages and employment causes region-industry-factor quality adjustors to drop out. As a result, observed relative wagebills, which are generally available to empirical researchers, are equal under the null hypothesis of relative factor price equality,

$$(\mathrm{H}_0\text{: RFPE}) \qquad \frac{\widetilde{\text{wagebill}}_{rj}^S}{\widetilde{\text{wagebill}}_{rj}^U} = \frac{\widetilde{\text{wagebill}}_{bj}^S}{\widetilde{\text{wagebill}}_{bj}^U}. \tag{10}$$

If RFPE does not hold, the quality-adjusted relative w^S/w^U wage differs across regions r and b by a multiplicative factor, γ_{rb}^{SU},

$$(\mathrm{H}_1\text{: No RFPE}) \qquad \frac{w_r^S}{w_r^U} = \gamma_{rb}^{SU}\frac{w_b^S}{w_b^U} \tag{11}$$

[11] RFPE holds if the quality-adjusted relative wages are equal for any $M-1$ of the M factors of production.

[12] Homogeneity of degree one of the cost function implies that the derivatives $\partial\Gamma_j/\partial w_r^\varphi$ are homogenous of degree zero in factor prices. It follows immediately from equation (6) that, with identical quality-adjusted relative factor prices, regions will employ quality-adjusted factors of production in the same proportions.

[13] As the factor quality of the base region has been normalized to equal one, $\theta_{bj}^\varphi = 1$, $\theta_{rj}^\varphi \neq 1$ indicates that factor quality differs in industry j between the base region and region r.

Here, again, we let region b be the benchmark region, so that $\gamma_{rb}^{SU} = \gamma_r^{SU}/\gamma_b^{SU}$, where $\gamma_b^{SU} = 1$. Across regions, observed relative wages now vary because of differences in factor quality and because of variation in true wages,

$$\frac{\tilde{w}_r^S}{\tilde{w}_r^U} = \gamma_{rb}^{SU} \frac{\theta_{rj}^S}{\theta_{rj}^U} \frac{\tilde{w}_b^S}{\tilde{w}_b^U}. \tag{12}$$

With quality-adjusted relative wage differences across regions, observed factor usage also varies because of differences both in factor quality and in factor demand driven by relative wage differences,

$$\frac{\widetilde{S}_{rj}}{\widetilde{U}_{rj}} = \frac{\theta_{rj}^S}{\theta_{rj}^U} \left(\frac{\partial \Gamma_j(\cdot)/\partial w_r^S}{\partial \Gamma_j(\cdot)/\partial w_r^U} \right) \left(\frac{\partial \Gamma_j(\cdot)/\partial w_b^S}{\partial \Gamma_j(\cdot)/\partial w_b^U} \right) \frac{\widetilde{S}_{bj}}{\widetilde{U}_{bj}}. \tag{13}$$

Multiplying the expressions for observed relative factor prices and observed relative employments (equations 12 and 13), the terms in unobserved factor quality again cancel. Under the alternate hypothesis of no RFPE, relative wagebills vary across regions because of differences in factor prices and variation in factor usage,

$$(\text{H}_1\text{: No RFPE}) \qquad \frac{\widetilde{\text{wagebill}}_{rj}^S}{\widetilde{\text{wagebill}}_{rj}^U} = \eta_{rbj}^{SU} \frac{\widetilde{\text{wagebill}}_{bj}^S}{\widetilde{\text{wagebill}}_{bj}^U}, \tag{14}$$

where

$$\eta_{rbj}^{SU} = \gamma_{rb}^{SU} \left[\left(\frac{\partial \Gamma_j(\cdot)/\partial w_r^S}{\partial \Gamma_j(\cdot)/\partial w_r^U} \right) \left(\frac{\partial \Gamma_j(\cdot)/\partial w_b^U}{\partial \Gamma_j(\cdot)/\partial w_b^S} \right) \right]. \tag{15}$$

If RFPE fails, there are two effects on the relative wagebill for an industry across regions. The first is given in equation (15) directly by the difference in relative wages, γ_{rb}^{SU}. The second effect, inside the brackets, is due to differences in relative factor usage caused by the variation in relative wages, and thus is also a function of γ_{rb}^{SU}. For example, if the region-industry production function is CES with $\sigma_j = 1/(1-\rho_j)$ as the elasticity of substitution between factors, then,

$$\eta_{rbj}^{SU} = \gamma_{rb}^{SU} \left[\left(\gamma_{rb}^{SU} \right)^{1/(\rho_j - 1)} \right] = \left(\gamma_{rb}^{SU} \right)^{\rho_j/(\rho_j - 1)}. \tag{16}$$

Such an assumption about the form of the production function enables the researcher to recover the underlying relative wage difference, γ_{rb}^{SU}, from the estimates of η_{rbj}^{SU}.

Together equations (10) and (14) provide the basis for a test of the null hypothesis of RFPE that is robust to unobserved region-industry variation in factor quality. The intuition underlying this methodology is that, although the empirical researcher cannot observe factor quality or quality-adjusted factor prices, observed factor prices contain information about the quality of observed factors when firms minimize costs. Multiplying observed factor prices by observed factor quantities enables us to control for unobserved variation in factor quality.

In our empirical work below, we test whether $\eta_{rbj}^{SU} = 1$ across regions. RFPE asserts all relative factor prices are equal. A rejection of relative factor price equality for any pair of factors, e.g. skilled and unskilled labor, is sufficient to reject the null hypothesis of RFPE.[14]

We caution that $\eta_{rbj}^{SU} \neq 1$ is sufficient to reject RFPE, but not necessary. Under CES production, for example, even if $\gamma_{rb}^{SU} \neq 1$ so that quality-adjusted relative wages are not equalized, the parameter $\eta_{rbj}^{SU} = \left(\gamma_{rb}^{SU}\right)^{\rho_j/(\rho_j-1)}$ equals unity for the special case of a Cobb-Douglas cost function ($\rho_j = 0$). In the implementation of our methodology below, we test the null hypothesis $\eta_{rbj}^{SU} = 1$ and, in so far as this hypothesis is rejected, this result is *sufficient* for us to reject RFPE. Under CES, the fact that $\left(\gamma_{rb}^{SU}\right)^{\rho_j/(\rho_j-1)}$ is close to 1 for ρ_j close to 0 actually makes it *harder* for us to reject the null hypothesis and strengthens any finding of a rejection of RFPE.

In the remainder of this section, we demonstrate the robustness of the relative wagebill test to the existence of imperfect competition, to production exhibiting increasing returns to scale, and to differences in factor composition.

2.2. Imperfect Competition

If firms maximize profits subject to a downward sloping inverse demand curve, $v_{rj}(Y_{rj})$, under conditions of imperfect competition, the first-order condition for profit-maximization is

$$\frac{dv_{rj}(Y_{rj})}{dY_{rj}} Y_{rj} + v_{rj}(Y_{rj}) - \frac{\Gamma_j(\cdot)}{A_{rj}} = 0. \tag{17}$$

Defining the elasticity of demand as $\varepsilon_{rj}(Y_{rj}) \equiv -(dY_{rj}/dv_{rj})v_{rj}/Y_{rj}$ where v_{rj} denotes price, we obtain the standard result that equilibrium price is a mark-up over marginal cost,

$$v_{rj}(Y_{rj}) = \left(\frac{\varepsilon_{rj}(Y_{rj})}{\varepsilon_{rj}(Y_{rj}) - 1}\right) \frac{\Gamma_j(\cdot)}{A_{rj}}. \tag{18}$$

By Shephard's Lemma, equilibrium demand for each quality-adjusted factor of production continues to be given by the derivative of the total cost function with respect

[14] With perfect capital mobility, the rate of return to capital may be equalized across regions. However, as long as there is a degree of immobility for at least one other factor of production, quality-adjusted relative factor prices will generally vary.

to the factor price as specified in equation (5). The derivation of the test for relative factor price equalization is thus identical to that provided above.

2.3. External Economies of Scale

It is straightforward to introduce external economies of scale into the framework above in either perfectly or imperfectly competitive market structures. External economies of scale correspond to the assumption that technical efficiency in a region-industry is a function of scale. In the most general case, we have,

$$A_{rj} = A_{rj}(Y_{rj}, Y_{r,-j}, Y_{-r,j}, Y_{-r,-j}) \tag{19}$$

where $Y_{r,-j}$ is the vector of outputs in all other industries in a region, $Y_{-r,j}$ is the vector of all other regions' outputs in the industry, and $Y_{-r,-j}$ is the vector of all other regions' outputs in all other industries. Because the cost-minimization behavior of the firm is the same (see equation 2), the derivation of the test for relative factor price equality remains unchanged.

2.4. Internal Economies of Scale

Internal economies of scale must clearly be combined with imperfect competition and imply that the cost function (3) is no longer linearly homogenous of degree one in output. Equilibrium price continues to be a mark-up over marginal cost,

$$v(Y) = \left(\frac{\varepsilon(Y)}{\varepsilon(Y)-1}\right)\frac{1}{A_{rj}}\frac{\partial\Gamma_j(w_r^U, w_r^S, w_r^K, Y)}{\partial Y}. \tag{20}$$

Equilibrium demand for quality-adjusted factors of production may again be obtained using Shephard's Lemma. Using the relationship between quality-adjusted and non quality-adjusted values, relative demand for observed skilled and unskilled workers will be given by,

$$\frac{\widetilde{S}}{\widetilde{U}} = \frac{\theta_{rj}^U}{\theta_{rj}^S}\frac{\partial\Gamma_j(w_r^S, w_r^U, w_r^K, Y)/\partial w_r^S}{\partial\Gamma_j(w_r^S, w_r^U, w_r^K, Y)/\partial w_r^U}. \tag{21}$$

Multiplying the expressions for observed relative factor prices and observed relative employments, the terms in unobserved factor quality again cancel. The expression for relative wagebills becomes,

$$\frac{\widetilde{\text{wagebill}}_{rj}^S}{\widetilde{\text{wagebill}}_{rj}^U} = \gamma_{rb}^{SU}\left(\frac{\partial\Gamma_j(\cdot)/\partial w_r^S}{\partial\Gamma_j(\cdot)/\partial w_r^U}\right)\left(\frac{\partial\Gamma_j(\cdot)/\partial w_b^U}{\partial\Gamma_j(\cdot)/\partial w_b^S}\right)\frac{\widetilde{\text{wagebill}}_{bj}^S}{\widetilde{\text{wagebill}}_{bj}^U} \tag{22}$$

where the terms in brackets capturing relative unit factor input requirements are now a function of output, Y.

In the standard case of trade under internal economies of scale in the theoretical literature, firms within an industry face the same constant elasticity of substitution ε_j, cost functions are homothetic and identical within industries, and there is free entry so that price equals average cost. Combining free entry with the pricing relationship in (20), the equilibrium ratio of average to marginal cost will equal a constant $\varepsilon_j/(\varepsilon_j - 1)$, which with homothetic cost functions defines a unique equilibrium value of output for all firms in the industry.

Under the null hypothesis of relative factor price equalization, $\gamma_{rb}^{SU} = 1$, and with all firms in the industry facing the same factor prices and producing the same output, the terms in parentheses in (22) cancel, so that we again obtain the prediction that relative wagebills are equalized under the null.[15]

More generally, in the presence of internal economies of scale, variation in firm size across regions and industries may influence factor demand and lead in general equilibrium to a violation of relative factor price equality.[16]

2.5. Factor Quality and Factor Composition

Our methodology for uncovering a failure of relative factor price equality is robust to region-industry variation in the mix of factors used within factor groups, e.g. variation in the relative use of skilled managers versus skilled engineers within the skilled worker factor group.[17] We assume that the production technology is weakly separable in skilled and unskilled workers, so that firms first choose optimal quantities of skilled and unskilled workers before choosing optimal amounts of worker types within these categories. We demonstrate the point formally for skilled workers, but, without loss of generality, the argument applies for any factor of production. Though, for simplicity, we consider two types of skilled workers, the

[15] Helpman and Krugman (1985) provide an analysis of theoretical models of monopolistic competition and increasing returns to scale with factor price equalization.

[16] We revisit the implications of internal increasing returns to scale in discussing the empirical results in Section 6.4.

[17] Our methodology also accounts for the misclassification of workers across factor categories if the misclassification is either random or occurs systematically along industry lines. However, if assignment of employees to factor categories differs systematically across regions – e.g. if all industries in a region report non-production workers as production workers – our methodology, and other techniques, will register a "spurious" rejection of RFPE. While misclassification between non-production and production factor categories may occur, it is likely to be less severe than variation in unobserved factor quality or composition within categories. Furthermore, it is also unlikely that such misclassification would be along regional rather than industry lines.

analysis goes through for any number of skill types. For notational convenience, we suppress region and industry subscripts throughout this section.

Assume the quality-adjusted flow of skilled labor services is a constant returns to scale function of the quality-adjusted flow of labor services supplied by managers and engineers:

$$\begin{aligned} S &= \phi(S_1, S_2) \qquad (23) \\ &= \phi\left(\frac{S_1}{(\widetilde{S}_1 + \widetilde{S}_2)}, \frac{S_2}{(\widetilde{S}_1 + \widetilde{S}_2)}\right)(\widetilde{S}_1 + \widetilde{S}_2) \\ &= \phi\left(\theta^{S_1}\widetilde{n}_1, \theta^{S_2}\widetilde{n}_2\right)\widetilde{S}, \end{aligned}$$

where S is quality-adjusted skilled labor services, S_1 is quality-adjusted manager labor services, S_2 is quality-adjusted engineer labor services, $\phi(\cdot)$ is assumed to be linearly homogenous of degree one, $\widetilde{S} = \widetilde{S}_1 + \widetilde{S}_2$ is the observed number of skilled workers, θ^{S_1} is the quality of managers, θ^{S_2} is the quality of engineers, and $\widetilde{n}_1$ and $\widetilde{n}_2$ are observed shares of engineers and managers in skilled employment. Equation (23) may be re-written more compactly as:

$$S = \widetilde{S}\theta^S, \qquad \theta^S \equiv \phi\left(\theta^{S_1}\widetilde{n}_1, \theta^{S_2}\widetilde{n}_2\right) \qquad (24)$$

where the term for the unobserved quality of skilled workers (θ^S) now captures the quality of managers, the quality of engineers, and the composition of skilled workers between managers and engineers.

The quality-adjusted wage of skilled workers is now a price index, defined as the dual to equation (23):

$$w^S = \psi(\omega_1, \omega_2) \qquad (25)$$

where ω_1 is the quality-adjusted wage of managers and ω_2 is the quality-adjusted wage of engineers.

Expenditure on quality-adjusted skilled worker services is equal to observed expenditure on skilled workers:

$$w^S S = \widetilde{w}^S \widetilde{S} \qquad (26)$$

where w^S is the price index defined above and $\widetilde{w}^S$ is the observed wage per skilled worker. It follows that the quality-adjusted skilled worker price index and the observed skilled worker wage are related according to:

$$w^S = \widetilde{w}^S / \theta^S, \qquad \theta^S \equiv \phi\left(\theta^{S_1}\widetilde{n}_1, \theta^{S_2}\widetilde{n}_2\right). \qquad (27)$$

It is clear from equations (24) and (27) that the factor composition term enters in exactly the same way as factor quality and the derivation of the relative factor price test remains unchanged. Under the null hypothesis that quality-adjusted relative wages are equalized across regions, the ratio of the observed wagebills of skilled and unskilled workers must be equalized across regions. Thus, as well as allowing for variation in factor quality, the wagebill test also controls for differences in factor composition that are specific to regions and industries.

3. Econometric Specification

In Section 2 we showed that under the null of RFPE the ratio of the skilled workers' wagebill to the unskilled workers' wagebill is the same across regions within an industry. This implies that, for an industry j, each region's relative wagebill equals the value for any base region b and, in particular, for the United States as a whole,

$$\frac{\widetilde{\text{wagebill}}^{S}_{rj}}{\widetilde{\text{wagebill}}^{U}_{rj}} = \frac{\widetilde{\text{wagebill}}^{S}_{bj}}{\widetilde{\text{wagebill}}^{U}_{bj}} = \frac{\widetilde{\text{wagebill}}^{S}_{USj}}{\widetilde{\text{wagebill}}^{U}_{USj}}. \tag{28}$$

The simplest test of the null hypothesis is therefore to regress the log of the ratio of wagebills for region r relative to the ratio for the U.S. on a set of region dummies,

$$\ln\left(\frac{RWB^{SU}_{rj}}{RWB^{SU}_{USj}}\right) = \sum_{r} \alpha^{SU}_{r} d_r + \varepsilon^{SU}_{rj} \tag{29}$$

where RWB^{SU}_{rj} denotes the relative wagebill in industry j and region r for skilled workers and unskilled workers ($RWB^{SU}_{rj} = \text{wagebill}^{S}_{rj}$ / wagebill^{U}_{rj}); RWB^{SU}_{USj} is the corresponding relative wagebill for the U.S. as a whole; and the α^{SU}_{r} correspond to the coefficients on the regional dummies d_r. Note that we exclude the own region r when defining the relative wagebill for the U.S. as a whole. Under the null hypothesis of RFPE, $\alpha^{SU}_{r} = 0$ for all regions and factor pairs, and a test of whether the α^{SU}_{r} are jointly equal to zero therefore provides a test of RFPE.

The regression in equation (29) corresponds to a differences in means test. We choose the aggregate U.S. as a base region and test RFPE by comparing the relative wagebill for an industry j across all regions r to the value for the aggregate U.S. in the same industry.

We also test RFPE by allowing individual regions to be the base region. That is, we begin by choosing a region b to be the base (where $\gamma^{SU}_{b} = 1$) and run a regression analogous to equation (29),

$$\ln\left(\frac{RWB_{rj}^{SU}}{RWB_{bj}^{SU}}\right) = \sum_{r} \alpha_{rb}^{SU} d_r + \varepsilon_{rbj}^{SU}. \tag{30}$$

A test of whether the α_{rb}^{SU} are jointly equal to zero provides a test of the null hypothesis of RFPE. Rejecting $\alpha_{rb}^{SU} = 0$ is sufficient to reject the null hypothesis of RFPE, and any pair of regions r and r' face the same relative factor prices if $\alpha_{rb}^{SU} = \alpha_{r'b}^{SU}$. To avoid problems with the choice of the base region, we estimate equation (30) for all possible choices of base region b.

Although regions have the same relative wagebills under the null hypothesis of RFPE (hence $\alpha_{rb}^{SU} = 0$), the theoretical analysis of Section 2 suggests that, under the alternative hypothesis, the coefficients on the regional dummies (η_{rb}^{SU} in equation 14 and α_{rb}^{SU} in equations 29 and 30) may vary across industries. With a constant elasticity of substitution (CES) production technology, this cross-industry variation is associated with different elasticities of substitution between skilled and unskilled workers (equation 16).

We have no strong priors on the industry variation in the elasticity of substitution between different types of labor or in other features of the operator Γ_j in the cost function (equation 3), and therefore we pool observations across industries. Since under the null hypothesis, $\alpha_{rbj}^{SU} = 0$, holds for all industries j, a finding of statistically significant coefficients on the regional dummies when pooling observations is sufficient to reject RFPE.

Our test for relative factor price differences holds for any constant returns to scale production technology. However, if we are willing to assume a CES production technology and choose a value for the elasticity of substitution σ, the estimated coefficients on the regional dummies may be used to derive implied quality-adjusted relative wages and unobserved factor quality across regions via equation (16). We make this assumption in interpreting our empirical results below.

Note that equations (29) and (30) compare the relative wagebill for skilled and unskilled workers in region r to the value in a base region within each industry j. This is a 'difference in differences' specification with a number of attractive statistical properties. Any industry-specific determinant of relative wagebills that is common across regions is 'differenced-out' when we normalize relative to the base region on the left-hand side of the equations (for example, features of the production technology, compensating differentials across industries, other inter-industry wage differentials, and industry-specific labor market institutions such as the degree of unionization). The analysis thus explicitly controls for observed and unobserved heterogeneity in the determinants of relative wagebills across industries.

Similarly, in both region r and the base region we analyze the wagebill of skilled workers *relative* to unskilled workers. Therefore, any region-specific determinant of wagebills that is common to both skilled and production workers is 'differenced-out' when we construct a region's relative wagebill (RWB_{rj}^{SU} = wagebill$_{rj}^{S}$ / wagebill$_{rj}^{U}$). Here potential examples include neutral regional technology differences and compensating differentials across regions that are common to skilled and unskilled workers, e.g. region-specific differences in the cost of living.

4. Empirical Implementation

In this section we apply our methodology to test for relative factor price equality across 181 U.S. labor markets in 1972 and 1992.

4.1. Data

We examine wagebills across the 181 Labor Market Areas (LMAs) that make up the continental United States (Alaska and Hawaii are excluded). LMAs, constructed by the Bureau of Economic Analysis, are aggregations of counties that are based on commuting patterns and therefore correspond closely to the concept of regional labor markets where wages are determined (see Johnson and Spatz 1993 for more detail). LMAs are permitted to cross state lines, and more than one labor market may appear in each state. As a result, LMAs provide greater resolution of relative factor price variation than more aggregate geographic units such as states or Census regions.[18]

Data on total payments to production (unskilled) and non-production (skilled) workers for 1972 and 1992 by industry and labor market area are obtained from the Censuses of Manufactures in the Longitudinal Research Database (LRD) collected by the U.S. Bureau of the Census.[19] We exclude four-digit Standard Industrial

[18] A number of studies (e.g. Topel 1986; Lee 1999, Bound and Holzer 2000, Hanson and Slaughter 2002, and Bernard and Jensen 2000) document variation in income inequality or wages across either the nine U.S. Census regions or across U.S. states. Related work using wage regressions by Heckman et al. (1996) finds that worker characteristics are priced differently across U.S. Census regions.

[19] Our sample covers all manufacturing establishments in the continental United States for which information on production and non-production workers is available. This sample excludes very small plants that do not report information on their inputs. Other data sources, such as the Decennial Census, collect more detailed information on worker wages and observed characteristics than does the LRD. However, these surveys generally record the industry of the worker at a very aggregate level of activity. Furthermore, sampling in these datasets does not ensure proportional

Classification (SIC4) industries that explicitly include miscellaneous products (i.e., SIC4 codes ending in '9') in order to base our examination on industries that are more likely to include comparable products.[20] This pruning leaves us with 401 of the original 458 SIC4 industries covering 88 percent of manufacturing output and 86 percent of manufacturing employment.

Though use of the non-production – production distinction to classify workers as skilled and unskilled is imperfect, this imperfection is mitigated here by our methodology's robustness to unobserved differences in region-industry factor quality and composition.

4.2. Testing RFPE

Table 1 reports the results of testing for relative factor price equality across LMAs using the U.S. average as the base region (equation 29). The data easily reject the null hypothesis of RFPE across regions within the United States for both 1972 and for 1992.[21] In 1972, 37 (55) regions have relative wagebills significantly different from the U.S. average at the 5 (10) percent level of significance. In 1992, 64 regions reject at the 5 percent level and 74 at the 10 percent level.

The relative *wagebill* results in Table 1 can be used to estimate relative skilled *wage* differences in individual labor markets by assuming CES production, as noted in equations (15) and (16) above. Nashville and New York City, for example, have significantly different relative wagebills for non-production and production workers, and thus significantly different relative wages. In 1972 the average relative wagebill across all industries in Nashville is 10 percent below the U.S. average while that for New York is 15 percent above. Twenty years later, the gap between the two labor markets had widened to 34 percent. Assuming CES production technologies and an elasticity of substitution of 2 between production and non-production workers (i.e. $\rho = 0.5$) in both years, these wagebill differences imply that quality-adjusted relative wages were 1.30 and 1.36 times higher in Nashville than in New York in 1972

representation by region-industry limiting their usefulness for testing relative factor price equality.

[20] "Spurious" rejection of RFPE is possible if industries are comprised of heterogeneous products and regions systematically specialize in certain products within industries. Suppose that relative factor prices are equal across all regions and that products vary in terms of skill intensity within all industries. If a region systematically produces skill-intensive products in every industry, its relative wagebill would be larger than the average for the country even if relative factor prices are identical across regions. We have attempted to mitigate the role of intra-industry heterogeneity by exploiting the highly disaggregate, four-digit SIC industry information available in the LRD. These industries are the most detailed industry data available for our sample period.

[21] The hypothesis that all the LMA coefficients are equal to zero is rejected at the 1 percent level in both 1972 and 1992.

and 1992, respectively.[22] In both periods, skilled workers in skill-scarce Nashville received higher relative wages than skilled workers in skill-abundant New York.[23]

We assign LMAs to factor-price cohorts based on the sign and significance (at the 10 percent level) of coefficients reported in Table 1. Figures 1 and 2 display the distribution of regions across cohorts for 1972 and 1992, respectively. Regions with relative skilled wagebills that are significantly higher than those for the aggregate United States are grouped in cohort A (black shading), while those with relative skilled wagebills that significantly lower are assigned to cohort C (cross-hatching). The remaining labor markets, with relative skilled wagebills that are not significantly different from the U.S. as a whole, are placed in cohort B. Regions in cohort A have higher relative wagebills and thus lower relative wages for skilled workers, while regions in cohort C have higher relative wages for skilled workers. New York, with a relatively low skilled wage is in cohort A and shaded black, while Nashville, with a relatively high skilled wage, is in cohort C and cross-hatched. Using the same assumption about the elasticity of substitution between factors as above, we estimate the average quality-adjusted relative skilled wage to be 11 percent higher and 21 percent lower than the national average in cohorts C and A, respectively, in 1972. The comparable percentages for 1992 are, respectively, 10 percent higher for C and 16 percent lower for A.

As indicated in Table 1 and highlighted in Figures 1 and 2, there is substantial movement of labor markets across cohorts between 1972 and 1992. In 1972 there are 9, 126, and 46 labor markets in the A, B and C cohorts, respectively. The corresponding figures are 16, 107 and 58 for 1992. Twenty-seven labor markets jump to a higher relative wagebill cohort over the sample period, while 34 regions drop to a lower relative wagebill cohort. These movements suggest an evolution of U.S. labor markets into extreme relative wage cohorts over time, a somewhat surprising result given the usual assumption that U.S. labor markets are becoming more tightly integrated and thus more likely to exhibit the same relative wages.

Our second specification for testing for relative factor price equality is the complete set of bivariate regressions captured by equation (30). Because there are far too many coefficients to report (32,580 per year when every region is used as a base), we report a summary of rejections in Table 2.[24] In 1972, 19 percent of the region-

[22] An elasticity of substitution between skilled and unskilled workers greater than unity is consistent with empirical estimates in the labor literature (Katz and Autor 1999).

[23] A lower relative wage for skilled workers in New York does not mean that the absolute level of wages is lower.

[24] Disclosure of individual coefficients from Table 2 is also not possible under Title XIII of the

pairs reject relative factor price equality at the 10 percent level, while 13 percent reject at the 5 percent level. Every region rejects with at least 3 other regions. In 1992, 24 percent of the region pairs reject relative factor price equality at the 10 percent level, 17 percent reject at the 5 percent level. Every region rejects with at least 3 other regions.

Both specifications provide strong evidence against the hypothesis that all regions in the United States face the same relative factor prices in either 1972 or 1992. The results show both that labor markets in the U.S. vary significantly in terms of relative wages, and that relative wages are lower in areas with large quantities of skilled workers. In the next section we explore the link between relative wage variation and industry structure.

5. RFPE and Industry Specialization

In this section we examine the link between labor markets' industry structure and the relative wage variation found above. Establishment of such a link highlights the importance of our finding of a breakdown of factor price equality because variation in region industry structure may be a key reason for divergence of regional outcomes over time. This link also provides insight into possible theoretical explanations of the failure of factor price equality, which we explore in Section 6..

Table 3 summarizes industry overlap among labor markets. The first two rows report the minimum, median and maximum percent of regions per industry, i.e., the breadth of industry production across regions. The median industry is produced in 34 percent of regions in 1992, up from 28 percent of regions in 1972. As indicated in the final column, some industries, like cement, are produced in every region.

The middle two rows of the table report the minimum, median and maximum percent of industries per region, i.e., the variety of industrial production within regions. No region produces all industries in either year; the most 'diverse' region manufactures 84 percent of all industries in 1972 and 86 percent in 1992. The median region increases its scope from 13 to 20 percent of all industries between 1972 and 1992.

The final two rows of Table 3 characterize the extent of bilateral industry overlap among regions. The percent of industries that two regions have in common is defined as the number of industries produced in both regions divided by the number of industries produced in the region with the larger number of industries. As

Bureau of Census.

indicated in the table, no two regions produce the same set of industries, though the extent of overlap increases with time.

We now test whether larger differences in relative skilled wages across labor markets are associated with smaller overlaps in the industries they produce, both in cross-section and over time.

5.1. *Industry Mix Across Regions*

We gauge whether the overlap in industry mix between two regions falls with differences in their relative factor prices by running an OLS regression of the number of industries two regions have in common on the distance between regions' relative wagebills,

$$COMMON_{rb} = \delta_0 + \delta_1|\alpha_{rb}^{SU}| + \beta_r I_r + \beta_b I_b + \epsilon_{rb}, \tag{31}$$

where $|\alpha_{rb}^{SU}|$ is the absolute value of the regression coefficient from equation (30), $COMMON_{rb}$ is the number of industries that regions r and b produce in common, and I_r and I_b are the number of industries produced by region r and b, respectively.[25] Separate estimation results for 1972 and 1992 are reported in Table 4. These results indicate that regions with more dissimilar wagebill ratios have fewer industries in common. The point estimates suggest that a pair of regions with the maximum estimated differences in relative wages would have 17 and 28 fewer industries in common in 1972 and 1992, respectively. Two regions with the median number of industries would have few, if any, regions in common if they exhibited the maximum differences in relative wages in each year.[26]

5.2. *Industry Mix Over Time*

If relative wage differences are important for firms seeking to minimize costs, regions experiencing larger changes in their relative wages over time should display greater churning of their industry mix in terms of adding and dropping industries. To check this relationship, we run an OLS regression of the form,

$$CHURN_r = \alpha + \beta_d \left|\alpha_{r,92}^{SU} - \alpha_{r,72}^{SU}\right| + \epsilon_r, \tag{32}$$

where the dependent variable, $CHURN_r$, is the percent of industries either added or dropped by region r between 1972 and 1992 relative to its number of industries in

[25] All our estimates to this point have been based on industries that exist in both regions.

[26] Regional product mixes may not be mutually exclusive because some goods with very high transport costs, such as cement, are essentially untradeable.

1972, and $\left|\alpha_{r,92}^{SU} - \alpha_{r,72}^{SU}\right|$ is the absolute value of the change in region r's wagebill ratio relative to the U.S. between 1972 and 1992 (Table 1). These changes range from 0.005 to 0.6 with a median of 0.07. Results are reported in Table 5. They indicate that industry churning and changes in estimated wagebill ratios are positively and significantly correlated. The implied value of $CHURN_r$ for the median change in relative wagebill ratios is 7.5 percentage points.

Together, the results of Tables 4 and 5 indicate that changes in relative wagebills are associated with varying industry mix both across regions and over time. These relationships suggest that regions have differential exposure to industry-specific shocks and that these asymmetric shocks will have uneven effects on regional labor markets.

6. Theoretical Explanations for the Failure of RFPE

In this section we discuss several potential explanations for our results. Neoclassical trade theory and spatial variation in relative costs of living appear to be the most likely. Alternate explanations based on region-industry productivity differences, increasing returns to scale and variation in industry prices across regions have additional implications that appear implausible.

6.1. Neoclassical Trade Theory

Our findings – rejection of factor price equality across U.S. labor markets, lower estimated relative wages for skilled workers in skill-abundant regions, and a negative relationship between relative wage differences and industry overlap – are consistent with a regional interpretation of the Heckscher-Ohlin model of neoclassical trade theory. Under this framework, inter-regional factor immobility prevents endowment-driven variation in labor markets' relative factor rewards from being arbitraged away. Regions specialize in industries according to comparative advantage, with skill-abundant countries having relatively low skilled wages producing skill-intensive goods while skill-scarce countries exhibiting relatively high skilled wages manufacture more labor-intensive products. Rejection of factor price equality across U.S. labor markets indicates U.S. factor mobility may be insufficient to even out disparities in regional relative factor endowments.

6.2. Spatial Variation in Nominal Wages

Relative wage differences in the neoclassical model are supported by factor immobility. However, even in the presence of perfect labor mobility, nominal relative wages may not be identical – and RFPE may be rejected – if the relative cost of living for skilled and unskilled workers diverges across regions.[27]

If this variation in relative consumer price indexes is sufficient, real consumption wages will be equal across regions even if nominal wages are not, providing workers no incentive to relocate. Note that, even in this case, industry participation will vary across regions because the location of production responds to producer prices and (nominal) factor prices faced by the firm rather than to real consumption wages. This variation in industrial structure means that regions again have differential exposure to industry-specific shocks, though factor mobility may now play a role in transmitting these shocks across regions.

For spatial variation in relative consumer price indexes to be consistent with our empirical findings, the cost of living for skilled workers relative to unskilled workers must vary substantially across regions and must be lower in skill-abundant regions. In addition, the regional disparity in relative costs of living must also increase between 1972 and 1992. The data required to test this additional implication would have to track the relative cost of living for production and non-production workers across LMAs over time.

6.3. Region-Industry Productivity Differences

Region-industry variation in total factor productivity may induce a rejection of relative factor price equality. In general equilibrium, if technology is not common across regions and varies differentially across industries, relative factor prices will vary so long as there is geographical immobility in at least one factor. However, to explain our empirical finding of a lower quality-adjusted skill premium in relatively skill-abundant regions, technical efficiency would have to be systematically relatively high in *low-skill* intensive industries within *high-skill* abundant regions. The intuition for why this relationship is necessary is that an increase in the technical efficiency of low-skill industries acts like an increase in their relative price. Such a price increase reduces the quality-adjusted skill premium and motivates a switch toward more skill-intensive techniques in both sectors.

[27] Divergence in the relative cost of living for skilled and unskilled workers across regions may be due to price differences associated with non-tradeables or region-specific amenities that are valued differentially by skilled and unskilled workers.

Though this explanation is theoretically possible, relatively low technical efficiency in skill-intensive industries within skill-abundant regions appears unlikely. Indeed, consideration of knowledge spillovers and external economies of scale suggest that technical efficiency would be relatively higher in skill-intensive industries located in skill-abundant regions.

6.4. Increasing Returns to Scale

Increasing returns to scale, either internal or external, can also motivate a rejection of relative factor price equality. To match the skill premia we observe, increasing returns to scale must reduce the relative (average) costs of production of low-skill industries in high-skill regions. This region variation in scale economies would raise the relative demand for low-skill workers (in the skill-abundant regions) and reduce the skill premium. Since output is typically larger in high-skill industries in high-skill regions, this explanation again appears implausible.

6.5. Variation in Industry Prices Across Regions

The null hypothesis of relative factor price equality may also be rejected if relative industry prices vary across regions. To match our results, the relative price of high-skill industries would have to be lower in skill-abundant regions. This sort of goods price variation is in fact an implication of Heckscher-Ohlin models that assume costly trade between regions in addition to factor immobility. Factor rewards diverge in these models for the same reason given above, namely differences in underlying factor endowments, though this divergence may be more extreme if trade costs lead to variation in relative goods prices.

7. Conclusions

This paper develops a methodology for testing whether factor prices are equal across geographic regions. It is based on cost minimization by firms and invokes only general assumptions about production, markets and factors. In particular, the method can identify departures from relative factor price equality in the presence of unobserved industry and factor heterogeneity that is likely to be present in any cross-country or cross-region sample, including unobserved variations in production techniques across industries, unobserved Hicks-neutral region-industry productivity differences, unobserved region-industry factor quality differences, unobserved region-industry variation in factor composition, and unobserved differences in regional consumer price indices. The test is relatively easy to implement in that it requires data only on the total payments to factors (e.g. wagebills) by industry and region.

We use our methodology to test for relative factor price equality across 181 U.S. labor markets areas in 1972 and 1992. The data reject the null hypothesis that all regions offer the same relative factor prices in both years. Results indicate substantial relative wage variation across skill-scarce and skill-abundant labor markets. We also find that relative wage differences have real economic impact: the greater the difference in relative wages across a region pair, the greater the difference in the pair's industry structure. This relationship is also evident within regions across time: regions experiencing larger changes in relative wages between 1972 and 1992 undergo larger changes in the set of industries they produce.

The association we find between regions' relative wages and their industry structure suggests U.S. labor markets may be asymmetrically exposed to domestic and external shocks. Further examination of this link may shed light on several literatures in economics, including the ability of skill-scarce regions to catch up with skill-abundant regions, the impact of trade liberalization on U.S. relative wages, and the effects of asymmetric shocks in optimum currency areas. Our finding of substantial relative wage variation across labor markets also suggests that use of industry production functions to estimate country-level productivity may need to be modified to account for regional heterogeneity.

Finally, we note that our approach to characterizing factor price inequality might usefully be applied to other settings where unobserved variation in quality is an important problem for identification. A similar test based on consumer expenditure minimization, for example, could be developed to test the law of one price across geographic areas.

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A Appendix A: Absolute Factor Price Equalization (AFPE)

This appendix develops a test for absolute factor price equality that controls for unobserved factor quality. Like our test for relative factor price equality, it makes use of the result that factor quality terms cancel when observed wages and observed employment levels are multiplied.

To test absolute factor price equalization (AFPE) we analyze variation across regions in the share of total payments to a factor of production in output. Though our demonstration here is for skilled workers, the analysis for other factors of production is analogous. Observed employment of skilled workers may be obtained from equations (4) and (2). Multiplying observed employment by observed wages and dividing by output, we obtain,

$$\frac{\tilde{w}_{rj}^S \widetilde{S}_{rj}}{Y_{rj}} = \frac{w_r^S S_{rj}}{Y_{rj}} = w_r^S A_{rj}^{-1} \frac{\partial \Gamma_j(\cdot)}{\partial w_r^S}. \tag{33}$$

Under the null hypothesis of AFPE, quality-adjusted wages are equal across regions ($w_r^S = w_b^S$) and observed wages vary in direct proportion to unobserved factor quality ($\tilde{w}_{rj}^S = \theta_{rj}^S w_b^S$), where we again choose region b as a reference region so that $\theta_{bj} = 1 \; \forall j$. The equality of the absolute level of factor prices requires identical production technologies across regions and industries ($A_{rj} = A_{bj}$). Using this relationship in equation (33), it follows that, under the null hypothesis of AFPE, factor shares are equalized across regions,

$$(H_0 : \text{AFPE}), \qquad \frac{w_r^S S_{rj}}{Y_{rj}} = \frac{w_b^S S_{bj}}{Y_{bj}}. \tag{34}$$

Under the alternative hypothesis of non-AFPE, technical efficiency may vary across region-industry pairs and regions may be characterized by different equilibrium factor prices. In this case, from equation (34), factor shares in the two regions are related as follows:

$$(H_1 : \text{non-AFPE}), \qquad \frac{w_r^S S_{rj}}{Y_{rj}} = \gamma_{rb}^S \left(\frac{A_{bj}}{A_{rj}}\right) \left(\frac{\partial \Gamma_j(\cdot)/\partial w_r^S}{\partial \Gamma_j(\cdot)/\partial w_b^S}\right) \left(\frac{w_b^S S_{bj}}{Y_{bj}}\right). \tag{35}$$

Together, equations (34) and (35) provide the basis for a test of the null hypothesis of AFPE, with AFPE implying a testable parameter restriction in equation (35).

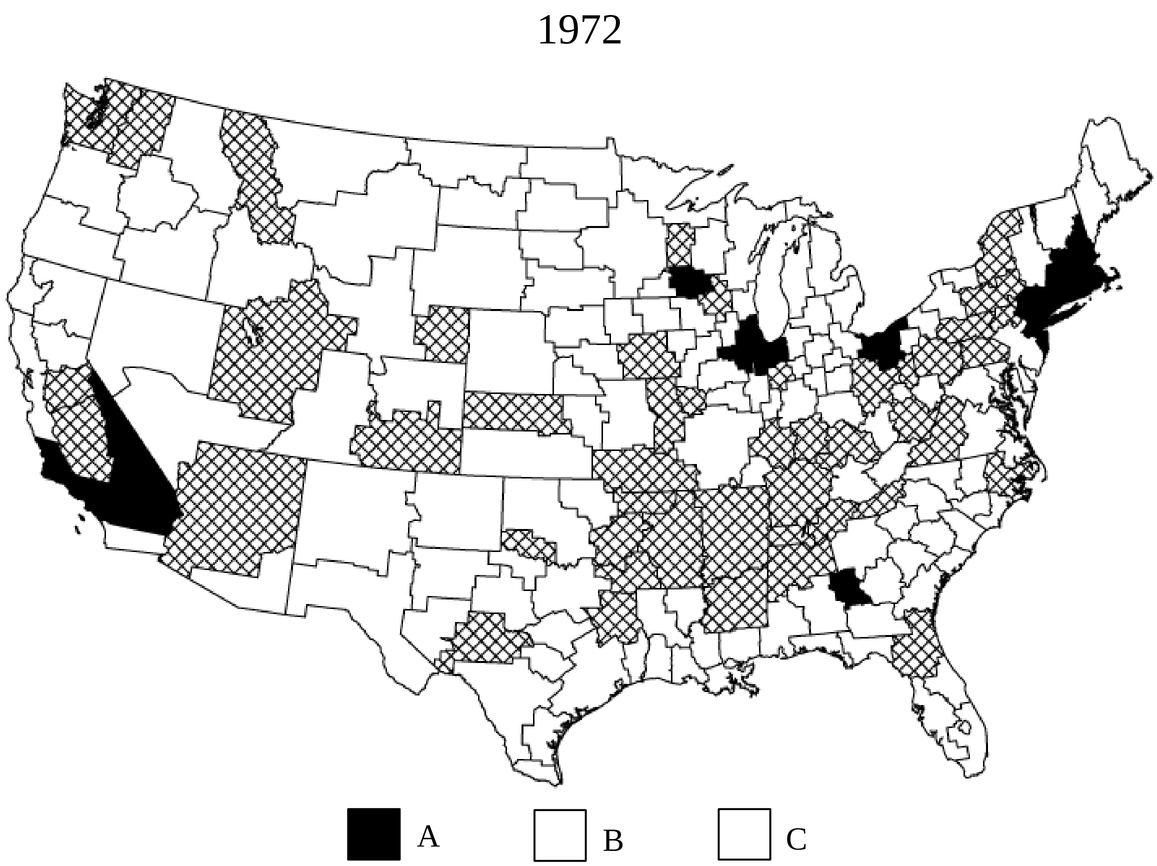

Figure 1: Labor Market Areas and Relative Wagebill Groups - 1972

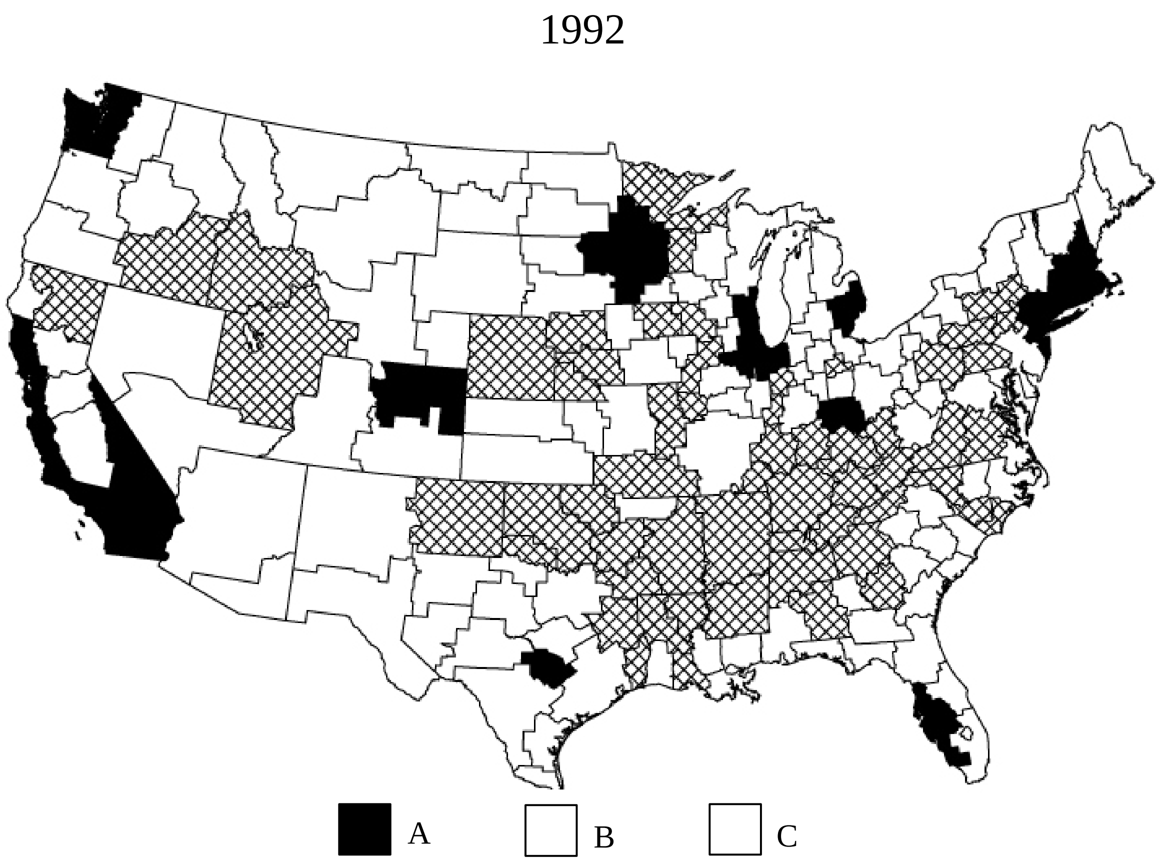

Figure 2: Labor Market Areas and Relative Wagebill Groups - 1992

LMA	Region	1972	1992	LMA	Region	1972	1992	LMA	Region	1972	1992
1	Bangor, ME	-0.10	0.07	62	Parkersburg, WV	0.17	-0.11	123	Austin, TX	0.12	0.11 *
2	Portland, ME	0.02	0.07	63	Wheeling, WV	-0.18 *	-0.13	124	Waco, TX	0.09	-0.10
3	Burlington, VT	-0.08	-0.04	64	Youngstown, OH	-0.07	-0.05	125	Dallas, TX	0.03	0.03
4	Boston, MA	0.10 ***	0.16 ***	65	Cleveland, OH	0.07 *	-0.03	126	Wichita falls, TX	-0.08	0.09
5	Providence, RI	0.12 **	0.11 **	66	Columbus, OH	-0.10 **	-0.07	127	Abilene, TX	0.05	0.01
6	Hartford, CT	0.13 ***	0.10 ***	67	Cincinnati, OH	0.02	0.13 ***	128	San angelo, TX	-0.42 *	0.23
7	Albany, NY	0.01	-0.02	68	Dayton, OH	-0.04	-0.01	129	San antonio, TX	0.02	-0.03
8	Syracuse, NY	-0.12 **	-0.03	69	Lima, OH	-0.11	-0.26 ***	130	Corpus christi, TX	-0.11	0.13
9	Rochester, NY	0.06	0.05	70	Toledo, OH	-0.06	-0.08	131	Brownsville, TX	-0.11	-0.10
10	Buffalo, NY	-0.06	-0.05	71	Detroit, MI	0.03	0.08 **	132	Odessa, TX	0.10	-0.01
11	Binghamton, NY	-0.17 ***	-0.12 *	72	Saginaw, MI	-0.06	0.05	133	El paso, TX	-0.12	-0.03
12	New york, NY	0.15 ***	0.17 ***	73	Grand rapids, MI	0.02	0.07	134	Lubbock, TX	-0.02	0.07
13	Scranton, PA	-0.26 ***	-0.18 ***	74	Lansing, MI	-0.04	-0.07	135	Amarillo, TX	-0.11	-0.24 **
14	Williamsport, PA	-0.19 ***	-0.13 **	75	South bend, IN	-0.01	-0.08	136	Lawton, OK	-0.30 *	-0.36 **
15	Erie, PA	0.02	-0.02	76	Fort wayne, IN	-0.05	-0.06	137	Oklahoma city, OK	-0.03	-0.13 **
16	Pittsburgh, PA	-0.12 ***	-0.11 ***	77	Kokomo, IN	0.04	0.00	138	Tulsa, OK	-0.07	-0.11 **
17	Harrisburg, PA	-0.09 **	-0.23 ***	78	Anderson, IN	-0.11	-0.06	139	Wichita, KS	-0.02	-0.07
18	Philadelphia, PA	-0.03	0.05	79	Indianapolis, IN	0.01	-0.03	140	Salina, KS	-0.28 **	-0.03
19	Baltimore, MD	0.04	0.02	80	Evansville, IN	-0.11 *	-0.21 ***	141	Topeka, KS	-0.09	-0.07
20	Washington, DC	0.01	-0.05	81	Terre haute, IN	0.07	-0.22 *	142	Lincoln, NE	-0.14	-0.20 **
21	Roanoke, VA	-0.14 **	-0.18 ***	82	Lafayette, IN	-0.32 ***	-0.26 **	143	Omaha, NE	-0.01	-0.14 **
22	Richmond, VA	-0.06	-0.14 **	83	Chicago, IL	0.08 **	0.09 ***	144	Grand island, NE	-0.13	-0.31 ***
23	Norfolk, VA	-0.12	0.11	84	Champaign, IL	-0.12	-0.03	145	Scottsbluff, NE	-0.36 *	-0.05
24	Rocky mount, NC	-0.13 *	-0.06	85	Springfield, IL	0.06	-0.08	146	Rapid city, SD	0.25	-0.02
25	Wilmington, NC	-0.03	-0.25 ***	86	Quincy, IL	-0.25 *	-0.31 **	147	Sioux falls, SD	-0.01	-0.11
26	Fayetteville, NC	0.00	-0.20 **	87	Peoria, IL	-0.11	0.05	148	Aberdeen, SD	-0.15	-0.18
27	Raleigh, NC	0.03	-0.04	88	Rockford, IL	-0.05	0.00	149	Fargo, ND	-0.05	0.02
28	Greensboro, NC	0.01	-0.10 **	89	Milwaukee, WI	0.03	0.08 **	150	Grand forks, ND	0.03	-0.17
29	Charlotte, NC	-0.01	-0.01	90	Madison, WI	-0.24 ***	-0.09	151	Bismarck, ND	0.01	0.19
30	Asheville, NC	-0.23 ***	-0.15 **	91	La crosse, WI	0.18 *	0.11	152	Minot, ND	0.19	0.04
31	Greenville, SC	0.00	0.00	92	Eau claire, WI	-0.27 **	-0.28 ***	153	Great falls, MT	-0.17	0.10
32	Columbia, SC	0.02	0.03	93	Wausau, WI	-0.05	-0.11	154	Missoula, MT	-0.39 ***	-0.12
33	Florence, SC	-0.11	-0.11	94	Appleton, WI	0.01	0.00	155	Billings, MT	0.03	-0.03
34	Charleston, SC	0.20	-0.04	95	Duluth, MN	-0.07	-0.17 *	156	Cheyenne, WY	-0.21	-0.03
35	Augusta, GA	0.03	-0.01	96	Minneapolis, MN	0.04	0.11 ***	157	Denver, CO	0.03	0.20 ***
36	Atlanta, GA	-0.02	-0.09 **	97	Rochester, MN	0.01	-0.16	158	Colorado springs, CO	-0.17 *	-0.12
37	Columbus, GA	0.18 **	-0.02	98	Dubuque, IA	-0.05	-0.39 ***	159	Grand junction, CO	-0.06	0.04
38	Macon, GA	-0.09	-0.24 ***	99	Davenport, IL	-0.06	-0.15 **	160	Albuquerque, NM	0.00	-0.05
39	Savannah, GA	0.00	-0.02	100	Cedar rapids, IA	-0.02	0.02	161	Tucson, AZ	-0.16	0.10
40	Albany, GA	-0.10	-0.10	101	Waterloo, IA	-0.07	-0.24 ***	162	Phoenix, AZ	-0.11 *	-0.02
41	Jacksonville, FL	-0.12 **	0.01	102	Fort dodge, IA	0.01	0.07	163	Las vegas, NV	-0.19	-0.12
42	Orlando, FL	0.08	0.06	103	Sioux city, IA	0.07	-0.26 ***	164	Reno, NV	-0.16	0.04
43	Miami, FL	0.05	0.06	104	Des moines, IA	-0.16 **	-0.06	165	Salt lake city, UT	-0.11 *	-0.09 *
44	Tampa, FL	-0.04	0.14 ***	105	Kansas city, MO	-0.02	0.02	166	Pocatello, ID	0.00	-0.16 *
45	Tallahassee, FL	0.09	-0.11	106	Columbia, MO	-0.20 *	-0.25 ***	167	Boise city, ID	-0.14	-0.22 **
46	Pensacola, FL	-0.03	0.00	107	St. louis, MO	-0.04	-0.03	168	Spokane, WA	-0.02	-0.02
47	Mobile, AL	-0.08	-0.05	108	Springfield, MO	-0.15 ***	-0.17 ***	169	Richland, WA	-0.14	-0.16
48	Montgomery, AL	-0.08	-0.11 *	109	Fayetteville, AR	-0.33 ***	-0.13	170	Yakima, WA	-0.26 **	-0.12
49	Birmingham, AL	-0.10 **	-0.11 **	110	Fort smith, AR	-0.17 *	-0.17 **	171	Seattle, WA	-0.08 *	0.07 *
50	Huntsville, AL	-0.13 *	-0.19 ***	111	Little rock, AR	-0.26 ***	-0.15 ***	172	Portland, OR	0.00	-0.02
51	Chattanooga, TN	-0.11 *	-0.14 ***	112	Jackson, MS	-0.14 **	-0.13 **	173	Eugene, OR	-0.14	-0.01
52	Johnson city, TN	-0.11	-0.20 ***	113	New orleans, LA	0.03	-0.04	174	Redding, CA	-0.23	-0.23 *
53	Knoxville, TN	0.01	-0.14 ***	114	Baton rouge, LA	0.03	-0.05	175	Eureka, CA	-0.11	-0.07
54	Nashville, TN	-0.10 **	-0.17 ***	115	Lafayette, LA	-0.07	-0.25 ***	176	San francisco, CA	0.00	0.19 ***
55	Memphis, TN	-0.08 *	-0.20 ***	116	Lake charles, LA	-0.04	-0.19	177	Sacramento, CA	-0.07	0.04
56	Paducah, KY	-0.16	-0.29 **	117	Shreveport, LA	-0.01	-0.17 **	178	Stockton, CA	-0.37 ***	-0.09
57	Louisville, KY	-0.12 **	-0.10 *	118	Monroe, LA	-0.05	-0.20 **	179	Fresno, CA	-0.18 ***	-0.06
58	Lexington, KY	-0.18 **	-0.14 **	119	Texarkana, TX	-0.23 **	-0.25 ***	180	Los angeles, CA	0.09 ***	0.15 ***
59	Huntington, WV	-0.08	-0.34 ***	120	Tyler, TX	-0.17 **	-0.23 ***	181	San diego, CA	0.04	0.18 ***
60	Charleston, WV	-0.18 *	-0.09	121	Beaumont, TX	-0.08	-0.27 **				
61	Morgantown, WV	-0.12	-0.14	122	Houston, TX	0.06	0.04				

*Significant at the 10% level; **Significant at the 5% level; ***Significant at the 1% level.

Table 1: Coefficients of Regression of Region Relative Wagebill on US Average Relative Wagebill

	1972	1992
Percent of Region Pairs Rejecting at 5% Level	13	17
Percent of Region Pairs Rejecting at 10% Level	19	24
Minimum Rejections	3	3
Mean Rejections	35	42
Maximum Rejections	116	128

Notes: Table summarizes rejections of relative factor price equality from estimation of equation 30.

Table 2: Summary of Bilateral Region-Pair RFPEQ Rejections from Estimation of Equation 18

	Year	Minimum	Median	Maximum
Regions Per Industry as a Percent of All Regions	1972	1	28	100
	1992	3	34	100
Industries per Region as a Percent of All Industries	1972	2	13	84
	1992	2	20	86
Bilateral Overlap as a Percent of the Larger Region's Industries	1972	5	32	94
	1992	6	34	93

Table 3: Overlap of Four-Digit SIC Industries Across US Labor Market Areas

	Number of Industries Common to Regions r and s	
	1972	1992
Relative Wagebill Disparity	-12.9	-23.1
	-0.8	-1.3
Industries in Region r	0.2	0.3
	-0.003	-0.003
Industries in Region s	0.3	0.3
	-0.004	-0.004
Constant	-4.9	-13.3
	-0.4	-0.5
Observations	16,290	16,290
R^2	0.68	0.75

Notes: OLS regression results. Dependent variable is number of industries produced in common by regions r and s. Robust standard errors noted below each coefficient.

Table 4: Regional Industry Overlap As a Function of Relative Wagebill Disparity

	$Churn_r$
|1972 to 1992 Change in Wagebill Ratio |	107.6
	-33.3
Constant	79.5
	-4.4
Observations	181
R^2	0.06

Notes: OLS regression results of changes in region industry structure on changes in relative wagebill ratio over time. Dependent variable is the percent of industries added or dropped by region r between 1972 and 1992 relative to its number of industries in 1972. The first idependent variable is the absolute value of the change in region r's wagebill ratio relative to the U.S. between 1972 and 1992 (i.e., the coefficients listed in Table 3). Robust standard errors noted below each coefficient.

Table 5: Industry Churning versus Relative Wagebill Changes, 1992 versus 1972