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Relative Wage Variation and Industry Location in the United Kingdom*

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Abstract

This paper shows that both relative wages and industry structure vary considerably across regions of the United Kingdom. In accordance with the neoclassical model of trade, regions abundant in a factor (i) exhibit lower relative prices of that factor than regions scarce in the factor, and (ii) tend to specialize in a mix of industries intensive in the use of the factor. We show that this specialization leads UK regions to be asymmetrically exposed to external macroeconomic or international trade shocks.

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I. Introduction

Geographic variation in relative factor rewards influences where firms choose to locate, where workers opt to live and how governments manage economic development. Although economic theory identifies two powerful forces – goods trade and factor mobility – that promote factor price convergence, factor price *inequality* may persist if spatial differences in relative factor endowments endure through time. The possible persistence of factor price inequality is relevant for a host of intra- and international public policy issues. In particular, because relative wages help determine industry structure, geographic variation in relative factor rewards may render regions within a country, or countries within a common market or currency union, asymmetrically susceptible to external shocks.

This paper focuses on identifying geographic variation in relative wages and industry mix within the United Kingdom over the period 1978–93. We find that relative wages vary considerably across both relatively aggregate Administrative Regions and more disaggregate counties, and that the pattern of variation is consistent with a key implication of neoclassical trade theory, with regions abundant in a factor exhibiting lower relative prices of the factor than regions scarce in that factor. Adjusted for quality, the 1993 ratio of white-collar to blue-collar wages in the UK's South-East Administrative Region, which is dominated by London, was just two-thirds the value in Wales and the North-East.[1] Moreover, we show that spatial variation in relative wages has widened over time. In 1978, for example, the ratio of white-collar to blue-collar wages in the South-East was three-fourths that in Wales.

The second part of our analysis demonstrates that spatial variation in UK relative wages coincides with geographic differences in industrial production. We show that regions with low relative white-collar wages specialize in a different set of industries from regions with high relative white-collar wages. These differences in patterns of industrial production also accord with the predictions of the neoclassical model: regions with low relative wages of white-collar workers have a higher concentration of industries that use white-collar workers intensively than regions with higher relative wages of white-collar workers. These findings echo recent studies of relative development within the UK that show that GDP per head in the South-East is roughly 140% of the level in the North-East.[2]

One explanation of our findings is that worker mobility across labour markets within the United Kingdom is imperfect and therefore unable to arbitrage away differences in regional relative wages. This hypothesis receives empirical support from research documenting limited labour mobility across regions within European countries.[3] An alternate explanation of our results – also consistent with the low observed

[1]Throughout the paper, we follow standard usage in distinguishing white-collar (non-production) workers from blue-collar (production) workers. The distinction between white-collar and blue-collar occupations is typically found to be positively correlated with measures of educational attainment (see, e.g. Machin and Van Reenen, 1998).

[2]See, for example, Cabinet Office (1999).

[3]See, for example, Hughes and McCormick (1994) and Jackman and Savouri (1992).

movement of labour – is that workers are, in fact, perfectly mobile but choose not to relocate because the nominal relative wage variation we document is offset by substantial differences in factor-region-specific amenities or living costs. The cost of living in the South-East is likely to be higher for both white-collar and blue-collar workers than in other UK regions. However, if the *relative* cost of living for white-collar workers compared with that for blue-collar workers is lower in the South-East (because, e.g. housing is a smaller share of white-collar workers' expenditure), white-collar workers may have no incentive to relocate from the South-East to other regions where the white-collar wage is high *relative* to the blue-collar wage. However, even in this case, production specialization – and the uneven exposure of workers across the UK to macroeconomic shocks – would still occur as industry location depends upon nominal rather than real relative wage differences.

Our findings relate directly to ongoing debates over the effectiveness of regional development assistance in offsetting the distributional consequences of trade liberalization. Because regions with high relative white-collar wages tend to attract industries that use blue-collar workers intensively, they may face greater competition from low-wage countries like China than regions that specialize in more white-collar-intensive industries. To be effective, development policies designed to raise the prospects of under-performing regions must take into account patterns of regional comparative advantage. For example, incentives to induce industries that use white-collar workers intensively to relocate to lagging regions may fail if they are inconsistent with the targeted regions' relative factor prices. Policies engineered to increase levels of education attainment and promote occupational mobility, by contrast, may meet with greater, albeit longer-term, success.

Our analysis complements existing studies of factor price variation both within and across countries in the international trade literature.[4] One of the key advantages of the test that we employ following Bernard, Redding and Schott (2005) is that it controls for region-factor-industry differences in the quality or composition of factors. We are therefore able to test whether relative factor prices differ across regions after controlling for unobserved differences in factor quality. The test holds under a range of assumptions about production, factors and markets, and is based upon the identifying assumptions of cost minimization, constant returns to scale, Hicks-neutral technology differences and weak separability of the production technology in white-collar and blue-collar workers. Our findings also relate to research in labour economics that examines returns to schooling.[5] Relative to that literature, this paper places greater emphasis on how relative factor prices are determined in general equilibrium as well as the manner in which spatial variation in relative factor prices is linked to industrial structure.

[4]Cunat (2001), Debaere and Demiroglu (2003), Debaere (2004), Repetto and Ventura (1998) and Schott (2003) examine factor price equalization across countries. Davis *et al.* (1997), Hanson and Slaughter (2002), Bernard, Robertson and Schott (2004) and Bernard *et al.* (2005) test for its existence within countries.

[5]See, for example, Mincer (1974), Griliches (1977), Jackman and Savouri (1992), Machin (1996), Machin and Van Reenen (1998) and Duranton and Monastiriotis (2002). Other research into earnings variation in the UK includes Cameron and Muellbauer (2001) and HM Treasury and DTI (2001).

The paper is structured as follows. Section II briefly discusses the methodology we use to identify departures from factor price equality. Section III presents evidence of relative wage variation across UK Administrative Regions and counties. Section IV examines the link between industrial structure and relative wages. Section V discusses some policy implications of our findings. Section VI concludes.

II. Identifying departures from RFPE

We identify departures from relative factor price equality (RFPE) in the United Kingdom using a methodology developed by Bernard *et al.* (2005) that is robust to unobserved factor-region-industry heterogeneity in the quality or composition of factors. The methodology exploits cost minimization to control for unobserved factor quality under a reasonably general set of assumptions about production technology and market structure.[6]

Consider a production technology defined over inputs of an arbitrary number of factors of production, including white-collar or non-production workers (N) and blue-collar or production workers (P). Under the assumptions of constant returns to scale and Hicks-neutral technology differences across regions, the production technology can be represented by the following unit cost function:

$$C_{rj} = \frac{1}{A_{rj}} \Gamma_j(\omega_r) \tag{1}$$

where r denotes regions; j corresponds to industries; ω_r is the vector of quality-adjusted factor prices; A_{rj} is the Hicks-neutral technology parameter; and the homogeneous of degree one function $\Gamma_j(\cdot)$ varies across industries.

We allow for unobserved differences in the quality of factors of production across regions, factors and industries. Therefore, the observed prices and employment of factors of production have the following relationship to their true quality-adjusted values: $\tilde{w}^z_{rj} = \theta^z_{rj} w^z_{rj}$ and $\tilde{E}^z_{rj} = E^z_{rj}/\theta^z_{rj}$, where z indexes factors of production, a tilde denotes an observed value, w^z_{rj} denotes the factor price, E^z_{rj} denotes employment of the factor and θ^z_{rj} denotes unobserved factor quality. We assume that the production technology is weakly separable in quality-adjusted white-collar workers, quality-adjusted blue-collar workers and other factors of production.[7]

Using Shephard's lemma and the relationship between observed and quality-adjusted employment, the observed relative demand for any two factors of production, such as white-collar and blue-collar workers, can be written as follows:

[6]Since the methodology exploits cost minimization, it is consistent with both perfect competition and imperfect competition. While the exposition here assumes constant returns to scale, the test can be extended to incorporate external and internal economies of scale.

[7]Under the assumption of weak separability, the quality-adjusted quantities (E^z_{rj}) and factor prices (w^z_{rj}) can be interpreted as indices defined over various types of each factor of production (e.g. types of white-collar workers). In this case, the unobserved factor quality parameters θ^z_{rj} capture differences in the quality and composition of factors of production across regions. See Bernard *et al.* (2005) for further discussion.

$$\frac{\tilde{E}_{rj}^{N}}{\tilde{E}_{rj}^{P}}=\frac{\theta_{rj}^{P}}{\theta_{rj}^{N}}\frac{\partial C_{rj}/\partial w_{r}^{N}}{\partial C_{rj}/\partial w_{r}^{P}}. \tag{2}$$

One of the key insights of our empirical test for RFPE is that the null hypothesis of RFPE implies that the relative prices of *all* factors of production are equalized across regions.[8] Therefore, to reject RFPE, it is sufficient to establish that the relative price of *any* two factors of production is different across regions. For this reason we focus in our empirical analysis on the relative wages of white-collar and blue-collar workers. While in principle one could test the hypothesis that the relative wages of these two factors of production are equalized across regions using data on observed white-collar and blue-collar wages, a major concern in doing so would be the unobserved variation in factor quality. Such unobserved variation in factor quality could induce a rejection of RFPE even though true quality-adjusted relative wages are equalized across regions. To address this concern, our methodology exploits information on observed values of both relative wages and employment to control for unobserved variation in factor quality.

Under the *null hypothesis* of RFPE, observed relative wages only vary across regions because of unobserved differences in factor quality:

$$\frac{\tilde{w}_{r}^{N}}{\tilde{w}_{r}^{P}}=\frac{\theta_{rj}^{N}}{\theta_{rj}^{P}}\frac{\tilde{w}_{s}^{N}}{\tilde{w}_{s}^{P}} \tag{3}$$

where, without loss of generality, we have chosen region s as the benchmark for measuring factor quality ($\theta_{sj}^{z}=1$).

Multiplying observed relative wages and observed relative employments in equations (2) and (3), the terms in unobserved factor quality cancel. Therefore, the null hypothesis of RFPE implies that relative wage bills are equalized across regions:

$$\frac{\widetilde{\text{wagebill}}_{rj}^{N}}{\widetilde{\text{wagebill}}_{rj}^{P}}=\frac{\widetilde{\text{wagebill}}_{sj}^{N}}{\widetilde{\text{wagebill}}_{sj}^{P}}. \tag{4}$$

The logic behind this result is that cost minimization and RFPE together imply that relative unit factor input requirements for quality-adjusted factors of production are the same across regions. Therefore, under the null hypothesis of RFPE, the only reason why relative wages and relative employment vary across regions is differences in unobserved factor quality. By combining the information in the variation in both relative wages and employment, we are thus able to control for differences in unobserved factor quality.

[8]We examine relative rather than absolute factor price equality for several reasons. First, there is a natural and rich link between variation in regions' relative factor prices and their industry structure, e.g. industries intensive in white-collar workers have an incentive to locate in regions abundant in white-collar workers. Secondly, a test of RFPE is more stringent in the sense that relative factor prices can be equal even if absolute factor price equality fails.

In contrast, under the *alternative hypothesis* of non-RFPE, observed relative wages vary across regions because of both unobserved differences in factor quality and differences in quality-adjusted factor prices. These differences in quality-adjusted factor prices themselves lead to differences in relative unit factor input requirements across regions. Combining differences in quality-adjusted factor prices and relative unit factor input requirements, relative wage bills in general differ across regions under the alternative hypothesis of non-RFPE:

$$\frac{\widetilde{\text{wagebill}}_{rj}^{N}}{\widetilde{\text{wagebill}}_{rj}^{P}} = \eta_{rsj} \frac{\widetilde{\text{wagebill}}_{sj}^{N}}{\widetilde{\text{wagebill}}_{sj}^{P}} \tag{5}$$
$$\eta_{rsj} = \gamma_{rs}^{NP}\left[\left(\frac{\partial\Gamma_j(\cdot)/\partial w_r^N}{\partial\Gamma_j(\cdot)/\partial w_r^P}\right)\left(\frac{\partial\Gamma_j(\cdot)/\partial w_s^P}{\partial\Gamma_j(\cdot)/\partial w_s^N}\right)\right]$$

where $\gamma_{rs}^{NP} \neq 1$ denotes the difference in quality-adjusted relative wages, such that $w_r^N/w_r^P = \gamma_{rs}^{NP}(w_s^N/w_s^P)$, and the difference in unit factor input requirements (captured by the term inside the square parentheses) depends in general on the difference in quality-adjusted relative wages and differences in the relative prices of other factors of production.

Comparing equations (4) and (5), the null hypothesis of RFPE can be tested in the presence of unobserved differences in factor quality by testing whether relative wage bills are equalized across regions. Finding that $\eta_{rsj} \neq 1$ is sufficient to reject the null hypothesis of RFPE.[9] Before proceeding with the empirical implementation of this test, we note that our approach does make a number of identifying assumptions: cost minimization, constant returns to scale, Hicks-neutral technology differences and weak separability of the production technology in white and blue-collar workers. Furthermore, the null hypothesis that we are testing is that *all* relative factor prices are equalized, and so one reason why relative wage bills may differ across regions is differences in the prices of other factors of production that have different degrees of complementarity or substitutability with white and blue-collar workers. However, while our test for RFPE is a joint test of our identifying assumptions and the hypothesis that all factor prices are equalized, its ability to control for unobserved differences in factor quality is an important advantage relative to other possible approaches.[10] Additionally, our identifying assumptions are relatively general and consistent with a range of production technologies and market structures. The data

[9]Although $\eta_{rsj} \neq 1$ is sufficient to reject RFPE, it is not necessary. Assuming a CES production technology with a common elasticity of substitution σ across all factors, $\eta_{rsj} = (\gamma_{rs}^{NP})^{\rho/(\rho-1)}$ where $\rho = (\sigma - 1)/\sigma$. In this case, even if $\gamma_{rs}^{NP} \neq 1$, so that quality-adjusted relative wages are not equalized, the parameter η_{rsj} equals unity for the special case of a Cobb–Douglas technology where $\rho = 0$. We test the null hypothesis $\eta_{rsj} = 1$ and, in so far as this hypothesis is rejected, this result is sufficient for us to reject RFPE.

[10]Estimating a production function and using the estimated coefficients on factor inputs to derive predicted wage premia across regions would provide an alternative approach to testing for RFPE. One of the challenges faced by this alternative approach is controlling for unobserved factor quality. The factor inputs are observed rather than quality-adjusted values. Therefore, the predicted wage premia derived from the estimated coefficients will reflect both true differences in wage premia and unobserved variation in factor quality.

requirements of our framework are also quite modest; an ideal data set would include total payments to several factors across geographic regions for each industry.

III. RFPE in the UK, 1976–93

In this section, we test for RFPE across regions within the United Kingdom using the methodology outlined above. We report results for two levels of regional aggregation.

Data description

Data on wages paid to white-collar and blue-collar workers by region and industry for 1976–93 are drawn from the Annual Respondents Database (ARD).[11] This data set includes basic information on the population of plants and establishments in the manufacturing sector, including employment, location, ownership, and presence in one of 209 four-digit UK Standard Industrial Classification (SIC 1980) manufacturing industries.[12] More detailed information on levels of output, employment and total wage bill payments to white-collar (non-production) and blue-collar (production) workers is available for a subset of roughly 13,000 establishments per year. This subset includes all large establishments and a sample of smaller establishments.[13] Because our methodology requires information on total wage payments, we work mainly with the establishment-level data. After presenting our main findings, we demonstrate that they are robust to potential sources of bias related to sampling, and the fact that some establishments report on plants in more than one location. In sections IV and V, when constructing measures of regions' industrial structure, we use the information on employment for the whole population of manufacturing plants so as to avoid any potential problems caused by sampling or establishment-level reporting.

We test for relative-wage differences across two different levels of regional aggregation, alternately breaking the United Kingdom into 10 relatively aggregate Administrative Regions and 63 relatively disaggregate counties and Scottish Regions (counties). A breakdown of counties by Administrative Region is presented in Table 1.[14]

[11]The ARD has two key advantages over labour-force surveys such as the Labour Force Survey (LFS) and the New Earnings Survey (NES) for testing for factor price equality within the UK. First, labour-force surveys use a sampling frame that is not representative at the region-industry level. Secondly, labour-force survey cell sizes become extremely small once one simultaneously conditions on region, industry and skill, especially for the kinds of disaggregated regions and industries one would hope to use.

[12]ARD establishments correspond roughly to a 'line of business'. An establishment can include more than one plant under common ownership. SIC 1980 codes are available in the data from 1979 to 1993. For the years 1976–78 we map SIC 1968 codes into SIC 1980 codes. The change to SIC 1992 from 1994 onwards represents a major change in industrial classification, hence we do not extend the analysis beyond 1993.

[13]For the years we focus on (1976–93), the sampling threshold is typically 100 employees, and is lower in three years. The sample represents roughly 75% of manufacturing employment. For this period the ARD contains information on production activity only and therefore excludes, for example, headquarter services. For more detailed discussions of the ARD, see Devereux, Griffith and Simpson (2004), Disney, Haskel and Heden (2003), Duranton and Overman (2005) and Griffith (1999).

[14]Throughout the paper, Northern Ireland is omitted due to a lack of data. References to the United Kingdom are therefore exclusive of Northern Ireland.

TABLE 1

Administrative Regions, counties and Scottish regions

Region	*County*	*Region*	*County*	*Region*	*County*
South-East	Bedfordshire	West Midlands	Hereford & Worcestershire	Northern	Cleveland
	Berkshire		Shropshire		Cumbria
	Buckinghamshire		Staffordshire		Durham
	East Sussex		Warwickshire		Northumberland
	Essex		West Midlands		Tyne & Wear
	Greater London				
	Hampshire	East Midlands	Derbyshire	Wales	Clywd
	Hertfordshire		Leicestershire		Dyfed
	Isle of Wight		Lincolnshire		Gwent
	Kent		Northamptonshire		Gwynedd
	Oxfordshire		Nottinghamshire		Mid Glamorgan
	Surrey				Powys
	West Sussex	Yorkshire & Humberside	Humberside		South Glamorgan
			North Yorkshire		West Glamorgan
East Anglia	Cambridgeshire		South Yorkshire		
	Norfolk		West Yorkshire	Scotland	Highland
	Suffolk				Grampian
		North-West	Cheshire		Tayside
South-West	Avon		Greater Manchester		Central
	Cornwall		Lancashire		Fife
	Devon		Merseyside		Strathclyde
	Dorset				Lothian
	Gloucestershire				Dumfries & Galloway
	Somerset				Borders
	Wiltshire				

Notes: Table lists UK counties by Administrative Region. Northern Ireland is excluded from the analysis.

We construct industry-region measures of total payments to each type of worker as well as total employment. When presenting results for Administrative Regions and counties across the full sample period, we average annual observations across 3-year intervals to both conserve space and smooth out annual fluctuations. Finally, in all results we exclude industries classified as ‘other manufacturing’ as these categories may include very different sub-industries in different regions. This trimming reduces the set of industries in the sample to 185. These industries represent roughly 80% of manufacturing employment in 1993, and our sample of establishments represents around 80% of employment in these 185 industries.

Econometric specification

Under the null of RFPE, the ratio of the non-production workers’ wage bill to the production workers’ wage bill is the same across regions within an industry. This implies that, for an industry j, each region's relative wage bill equals the value for any base region s and, in particular, for the aggregate United Kingdom,

$$\frac{\widetilde{\text{wagebill}}_{rj}^{N}}{\widetilde{\text{wagebill}}_{rj}^{P}} = \frac{\widetilde{\text{wagebill}}_{sj}^{N}}{\widetilde{\text{wagebill}}_{sj}^{P}} = \frac{\widetilde{\text{wagebill}}_{UKj}^{N}}{\widetilde{\text{wagebill}}_{UKj}^{P}}. \tag{6}$$

The simplest test of the null hypothesis is therefore to regress the ratio of wage bills for region r relative to the ratio for the aggregate UK on set of region dummies,

$$\ln\left(\frac{RW\,B_{rj}^{NP}}{RW\,B_{UKj}^{NP}}\right) = \sum_{r} \alpha_r^{NP} d_r + \varepsilon_{rj}^{NP} \tag{7}$$

where $RW\,B_{rj}^{NP}$ denotes the relative wage bill in industry j and region r for non-production workers and production workers (i.e. $RW\,B_{rj}^{NP} = \text{wagebill}_{rj}^{N}/\text{wagebill}_{rj}^{P}$); $RW\,B_{UKj}^{NP}$ is the corresponding relative wage bill for the UK as a whole; and α_r^{NP} correspond to the coefficients on the regional dummies d_r. We note that, when defining the relative wage bill for the aggregate UK, we exclude the own region r from the aggregate. Under the null hypothesis of RFPE, $\alpha_r^{NP} = 0$ for all regions and factor pairs, and a test of whether α_r^{NP} are jointly equal to zero therefore provides a test of RFPE.

The regression in equation (7) corresponds to a differences in means test. We choose the aggregate UK as a base region and test RFPE by comparing the relative wage bill for an industry j across all regions r to the value for the UK as a whole 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 s to be the base and run a regression analogous to equation (7),

$$\ln\left(\frac{RW\,B_{rj}^{NP}}{RW\,B_{sj}^{NP}}\right) = \sum_{r} \alpha_{rs}^{NP} d_r + \varepsilon_{rsj}^{NP}. \tag{8}$$

The F-test of whether the α_{rs}^{NP} are jointly equal to zero provides a test of the null hypothesis of RFPE. Rejecting $\alpha_{rs}^{NP} = 0$ is sufficient to reject the null hypothesis of RFPE. Any pair of regions r and r' face the same relative factor prices only if $\alpha_{rs}^{NP} = \alpha_{r's}^{NP}$. To avoid having the results driven by the choice of the base region, we estimate equation (8) for all possible choices of base region s.

Note that equations (7) and (8) compare the relative wage bill for non-production and production 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 wage bills that is common across regions is 'differenced-out' when we normalize relative to the base region on the left-hand side of the equations (e.g. features of the production technology, compensating differentials across industries, other inter-industry wage differentials and industry-specific labour market institutions such as the degree of unionization). The analysis thus explicitly controls for observed and unobserved heterogeneity in the determinants of relative wage bills across industries. Similarly, in both region r and the base region we analyze the wage bill of non-production workers *relative* to that of production workers. Therefore, any region-specific determinant of wage bills that is common to both non-production and production workers is 'differenced-out' when we construct a region's relative wage bill ($RWB_{rj}^{NP} = \text{wage bill}_{rj}^{N}/\text{wage bill}_{rj}^{P}$). Here, potential examples include neutral regional technology differences and compensating differentials across regions.

Although regions have the same relative wage bills under the null hypothesis of RFPE (hence $\alpha_{rs}^{NP} = 0$), the theoretical analysis of section II suggests that, under the alternative hypothesis, the coefficient on the regional dummies (η_{rs}^{NP} in equation 5 and α_{rs}^{NP} in equations 7 and 8) varies across industries. Nonetheless, as under the null hypothesis, $\alpha_{rsj}^{NP} = 0$ holds for all industries j, a finding of statistically significant coefficients on the regional dummies when pooling observations is sufficient to reject RFPE.

The key identifying assumptions imposed in deriving our test for RFPE are cost minimization, constant returns to scale, Hicks-neutral technology differences, and weak separability of the production technology in quality-adjusted white-collar and blue-collar workers. While our test of RFPE requires only these assumptions, the implied difference in quality-adjusted relative wages of white-collar and blue-collar workers cannot be derived from the estimated coefficients on the regional dummies α_{rs}^{NP} without making further assumptions. The reason is that the estimated coefficients α_{rs}^{NP} depend in general on the functional form of the production technology and the relative prices of other factors of production that can exhibit different degrees of complementarity and substitutability with white-collar and blue-collar workers. However, under the assumption of a constant elasticity of substitution (CES) production technology with a common elasticity of substitution σ across all factors, the implied differences in quality-adjusted relative wages can be derived. Under the CES assumption, the unit cost function takes the following form:

$$C_{rj} = \frac{1}{A_{rj}} \left[\sum_{z=1}^{Z} w_r^{1-\sigma} \right]^{\frac{1}{1-\sigma}} \tag{9}$$

where Z denotes the number of factors of production.

Applying Shephard's lemma, the relative wage bills for any two factors of production such as white-collar and blue-collar workers can be expressed follows:

$$\frac{\widetilde{\text{wagebill}}_{rj}^{N}}{\widetilde{\text{wagebill}}_{rj}^{P}} = \left(\gamma_{rs}^{NP}\right)^{\rho/(\rho-1)} \frac{\widetilde{\text{wagebill}}_{sj}^{N}}{\widetilde{\text{wagebill}}_{sj}^{P}} \tag{10}$$

where $\gamma_{rs}^{NP} \neq 1$ under the alternative hypothesis of non-RFPE and $\rho = (\sigma - 1)/\sigma$.

From equation (10), in this special case of a CES production technology, the implied differences in quality-adjusted relative wages γ_{rs}^{NP} can be derived from the estimated coefficients on the regional dummies under an assumption of a value for the elasticity of substitution between factors, as $\eta_{rs}^{NP} = \exp(\alpha_{rs}^{NP}) = (\gamma_{rs}^{NP})^{\rho/(\rho-1)}$.

Empirical results

In our first set of results, we test for RFPE across UK regions using the relative wage bill specification in equation (7). This specification compares each region's relative wage bill to a base region that is defined as the aggregate UK less the own region.

Table 2 presents estimation results at the level of Administrative Regions during six 3-year time periods from 1976 to 1993.[15] Results are stable over time. In each year, there is typically one statistically significant rejection above zero (the South-East) and four statistically significant rejections below zero (Yorkshire and Humberside, Northern, Wales and Scotland). An exception is 1988–90, a period of recession, when the two Midlands regions and the North-West also reject below zero. Over time, the positive South-East coefficient increases while the negative Northern region declines. These trends suggest an increase in spatial relative-wage polarization over time.

Estimation results for counties are also persistent over time. Tables 3 and 4 report counties exhibiting statistically significant positive and negative coefficients in each

TABLE 2

Tests of common relative wage across Administrative Regions, UK base

Administrative Region	*1976–78*	*1979–81*	*1982–84*	*1985–87*	*1988–90*	*1991–93*
South-East	0.154***	0.166***	0.233***	0.234***	0.253***	0.267***
East Midlands	−0.052	−0.012	−0.021	−0.058	−0.075*	0.003
South-West	0.001	0.025	−0.016	−0.003	0.012	0.002
East Anglia	−0.029	−0.052	−0.086**	−0.054	−0.059	−0.029
North-West	−0.025	−0.014	−0.045	−0.007	−0.071*	−0.030
West Midlands	−0.068*	−0.050	−0.022	−0.051	−0.109***	−0.057
Scotland	−0.112***	−0.109***	−0.148***	−0.161***	−0.190***	−0.136***
Wales	−0.167***	−0.150***	−0.134***	−0.110***	−0.166***	−0.155***
Yorks & Humberside	−0.171***	−0.091**	−0.100***	−0.089**	−0.096**	−0.159***
Northern	−0.169***	−0.151***	−0.200***	−0.278***	−0.200***	−0.200***
F-stat (p-value)	<0.01	<0.01	<0.01	<0.01	<0.01	<0.01
Observations	1,584	1,544	1,564	1,541	1,581	1,580

Notes: Table displays point estimates from estimation of equation (7) on Administrative Regions. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels, respectively. Statistical significance is based on standard errors that are robust to heteroscedasticity. Dashed lines group Administrative Regions into relative wage cohorts according to relative wage point estimate sign and statistical significance in 1991–93.

Source: Authors' calculations using ARD (source: ONS).

[15]As noted above, we average the data in 3-year increments to both conserve space and smooth out annual fluctuations. None of our results is sensitive to this averaging.

TABLE 3

Counties with positive and significant coefficients, UK base

Administrative Region	*County*	*1976–78*	*1979–81*	*1982–84*	*1985–87*	*1988–90*	*1991–93*
South-East	Berkshire	0.144**	0.279***	0.221***	0.154**	0.251***	
	Buckinghamshire	0.143**			0.189***		
	Essex						0.118*
	Greater London	0.111**	0.120**	0.117**	0.147***	0.176***	0.171***
	Hampshire			0.142**		0.207***	0.161**
	Oxfordshire	0.151*					
	Surrey	0.185***	0.179***	0.231***	0.268***	0.193***	0.189**
	West Sussex		0.136**			0.184**	0.287***
South-West	Wiltshire				0.147*	0.184**	0.134*
East Midlands	Leicestershire						0.117*
North-West	Cheshire		0.104*		0.142**		
F-stat (p-value)		<0.01	<0.01	<0.01	<0.01	<0.01	<0.01
Observations		5,345	5,021	5,097	4,867	5,322	5,093

Notes: Table displays positive and significant point estimates from estimation of equation (7) on UK counties. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels, respectively. Statistical significance is based on standard errors that are robust to heteroscedasticity.

Source: Authors' calculations using ARD (source: ONS).

3-year interval. Positive coefficients on the regional dummies are concentrated in the South-East. But there is some evidence that the areas exhibiting significantly positive coefficients are spreading out geographically over time to areas outside the South-East, for example Wiltshire in the South-West joins the majority of South-East counties from 1985–87 onwards. This movement is consistent with the expansion of white-collar-intensive industries along motorway arteries radiating outwards from South-East, such as the M4 and M3. Negative and statistically significant coefficients on the regional dummies are concentrated disproportionately outside the South-East in the North and West of the United Kingdom (in the Northern, Scotland and Wales Administrative Regions). Figures 1 and 2 provide geographic representations of our results at the county level for the periods 1976–78 and 1991–93.

Under the assumption of a CES production technology with a common elasticity of substitution across factors, the coefficient estimates presented in this section can be used to derive implied quality-adjusted relative wage differentials across regions. Assuming an elasticity of substitution of $\sigma = 2$, positive coefficients on the regional dummies α_{rs} correspond to lower quality-adjusted wages of white-collar workers relative to blue-collar workers, because $\exp(\alpha_{rs}) = (\gamma_{rs}^{NP})^{\rho/(\rho-1)}$ and $0 < \rho < 1$.[16] Under this assumption, Table 5 shows that in 1978 the ratio of white-collar to blue-collar wages in the South-East (i.e. the area around London) was more than 25% lower than that in Wales, Scotland and the Northern and Yorkshire and Humberside regions

[16]The assumption $\sigma = 2$ is consistent with empirical estimates in the labour literature (see, in particular, Katz and Murphy, 1992; Katz and Autor, 1999).

TABLE 4

Counties with negative and significant coefficients, UK base

Administrative Region	*County*	*1976–78*	*1979–81*	*1982–84*	*1985–87*	*1988–90*	*1991–93*
South-East	Bedfordshire			−0.175**			
	East Sussex		−0.273***			−0.176**	
	Isle of Wight				−0.289*		
East Anglia	Norfolk	−0.288***	−0.198***	−0.128*		−0.176**	
	Suffolk		−0.166**	−0.157**	−0.127*	−0.126*	
South-West	Avon			−0.125*			
	Cornwall	−0.217***	−0.163*	−0.280***	−0.332***	−0.218**	−0.279***
	Devon		−0.194***			−0.126*	−0.193**
	Dorset			−0.230***			
	Somerset				−0.140*		
West Midlands	Shropshire	−0.147*			−0.188**	−0.248***	−0.278***
	Staffordshire	−0.156**				−0.130*	−0.115*
	West Midlands	−0.122**	−0.123**	−0.092*			
East Midlands	Derbyshire	−0.163***		−0.126**	−0.153**	−0.125**	
	Lincolnshire	−0.175**		−0.166**		−0.145*	−0.183**
	Nottinghamshire	−0.162***		−0.144**	−0.175***	−0.127*	−0.121*
Yorkshire and Humberside	Humberside	−0.313***	−0.230***	−0.119*	−0.202***	−0.119*	−0.183**
	North Yorkshire	−0.251***	−0.230***				
	South Yorkshire	−0.208***		−0.149**	−0.220***	−0.193***	−0.246***
	West Yorkshire	−0.162***	−0.140***	−0.156***	−0.103*	−0.111**	
North-West	Greater Manchester	−0.090*				−0.114**	
	Lancashire	−0.197***	−0.147**	−0.142**	−0.137**	−0.171***	
	Merseyside					−0.141**	
Northern	Cleveland	−0.290***	−0.272***	−0.294***	−0.362***	−0.407***	−0.433***
	Cumbria		−0.161**	−0.306***	−0.250***		
	Durham		−0.163**	−0.217***	−0.320***	−0.181**	−0.358***
	Northumberland	−0.270***	−0.305***	−0.201**	−0.207*	−0.304***	−0.337***
	Tyne & Wear	−0.241***	−0.120*	−0.151**	−0.237***		

continued overleaf

TABLE 4

(continued)

Administrative Region	*County*	*1976–78*	*1979–81*	*1982–84*	*1985–87*	*1988–90*	*1991–93*
Wales	Clwyd	−0.173**	−0.163**	−0.229***	−0.145*	−0.311***	−0.377***
	Dyfed					−0.204*	−0.216*
	Gwent	−0.205***	−0.178**		−0.142*		
	Gwynedd	−0.466***	−0.297**	−0.319**		−0.411***	−0.298**
	Mid Glamorgan	−0.244***	−0.245***	−0.174**	−0.199***	−0.220***	−0.224***
	Powys	−0.277**	−0.242**	−0.239**		−0.203*	−0.370***
	South Glamorgan				−0.172*	−0.175*	
	West Glamorgan				−0.203**	−0.238**	−0.218*
Scotland	Borders	−0.219*	−0.339**	−0.271**		−0.325**	
	Central	−0.134*	−0.191**	−0.320***	−0.295***	−0.214**	
	Dumfries & Galloway	−0.450***	−0.379***	−0.226*	−0.475***	−0.731***	−0.508***
	Fife		−0.144*	−0.235***	−0.160**	−0.142*	−0.223**
	Grampian	−0.319***	−0.278***	−0.367***	−0.366***	−0.267***	−0.339***
	Highland	−0.486***	−0.374***	−0.301**	−0.334**	−0.358**	−0.301*
	Lothian	−0.215***		−0.123*			
	Strathclyde	−0.097*	−0.092*	−0.112**	−0.131**	−0.152***	
	Tayside	−0.272***	−0.222***	−0.279***		−0.214**	−0.284***
F-stat (p-value)		<0.01	<0.01	<0.01	<0.01	<0.01	<0.01
Observations		5,345	5,021	5,097	4,867	5,322	5,093

Notes: Table displays negative and significant point estimates from estimation of equation (7) on UK counties. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels, respectively. Statistical significance is based on standard errors that are robust to heteroscedasticity.

Source: Authors' calculations using ARD (source: ONS).

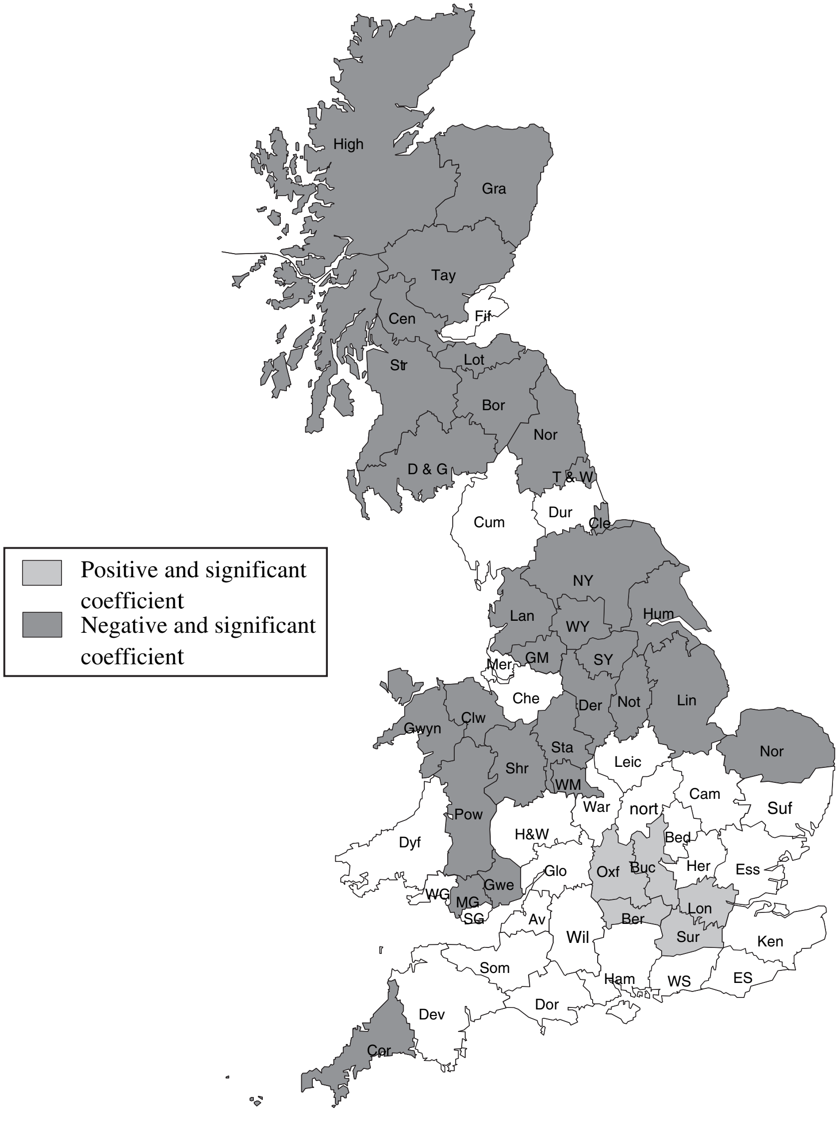

Figure 1. Counties with positive and negative relative wage bill coefficients (1976–78)

of England. The evolution of coefficients over time indicates that the quality-adjusted relative wage of white-collar workers fell quite significantly over time in the South-East compared with the UK as a whole, and rose in Scotland and the Northern region of England compared with the UK as a whole. By 1993, the relative white-collar wage in the South-East was 35% lower than that in the Northern region of the country.

Departures from RFPE can also be identified via equation (8), which uses individual regions rather than the aggregate UK as a basis of comparison. Results for this

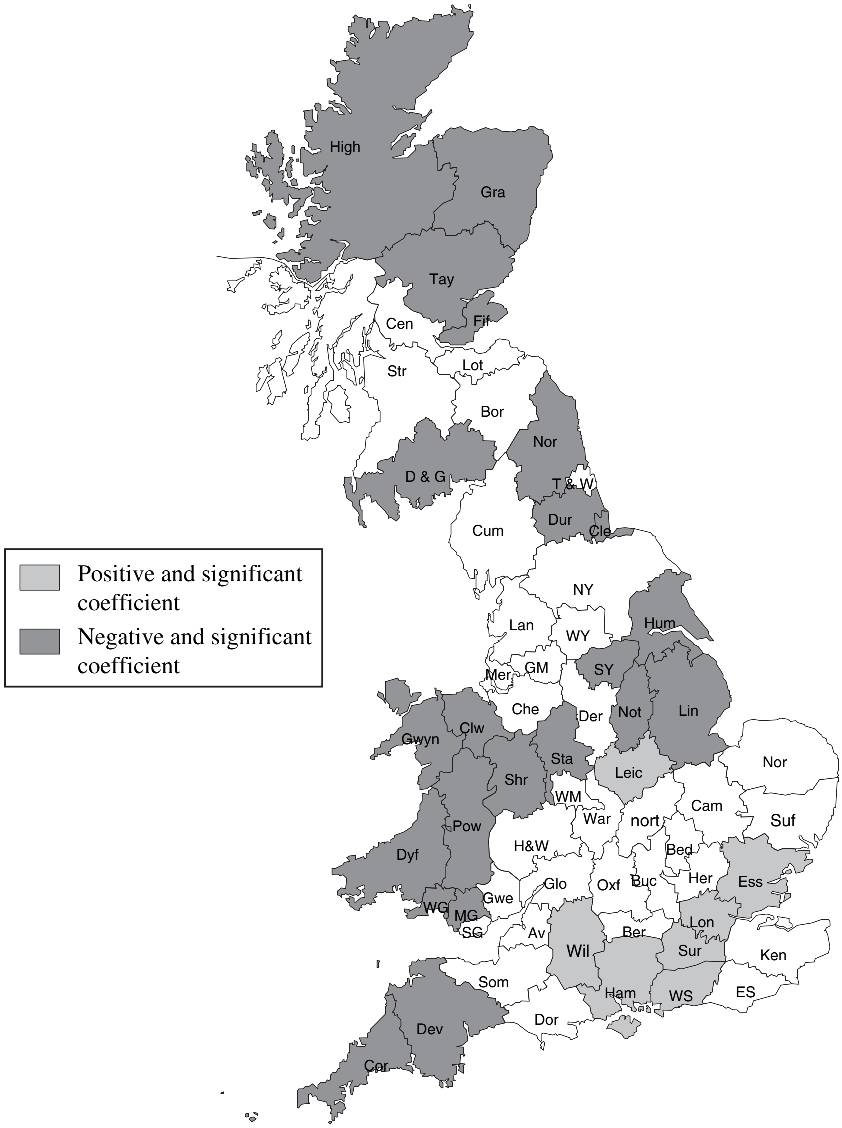

Figure 2. Counties with positive and negative relative wage bill coefficients (1991–93)

estimation are summarized in Table 6 for both Administrative Regions and counties. Approximately 50% of bilateral Administrative Region pairs and 25% of bilateral county pairs reject RFPE at the 10% level of significance. These rejections are not due exclusively to the influence of the South-East. Across all time periods, each Administrative Region rejects with an average of six other regions and a minimum of two. Table 7 examines the number of bilateral rejections by base region, where we again observe a systematic pattern across time periods, with the South-East and Northern regions producing the highest number of rejections.

TABLE 5

Implied quality-adjusted relative wage differences (CES technology)

Administrative Region	*Regression coefficient 1976–78*	*Implied quality-adjusted wage of white-collar workers relative to blue-collar workers*	*Regression coefficient 1991–93*	*Implied quality-adjusted wage of white-collar workers relative to blue-collar workers*
South-East	0.154***	0.860	0.267***	0.770
East Midlands	−0.052	1.050	0.003	1.000
South-West	0.001	1.000	0.002	1.000
East Anglia	−0.029	1.030	−0.029	1.030
North-West	−0.025	1.030	−0.030	1.030
West Midlands	−0.068*	1.070	−0.057	1.060
Scotland	−0.112***	1.120	−0.136***	1.150
Wales	−0.167***	1.180	−0.155***	1.170
Yorkshire and Humberside	−0.171***	1.190	−0.158***	1.170
Northern	−0.169***	1.180	−0.200***	1.220

Notes: Table reports the implied, quality-adjusted wage of white-collar workers relative to blue-collar workers given the point estimates from equation (7), by Administrative Region. See text for details. Results are sorted by the 1991–93 regression coefficients. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels, respectively. Statistical significance is based on standard errors that are robust to heteroscedasticity. Dashed lines group administrative regions into relative wage cohorts according to point estimate sign and statistical significance in 1991–93.

Source: Authors' calculations using ARD (source: ONS).

Robustness

In this section we demonstrate the robustness of the relative wage bill results to potential sources of bias related to the way in which our data are sampled. To conserve space, we report robustness results using Administrative Regions for the end period 1991–93. Results are similar across years.

Results are reported in Table 8, the first column of which reproduces the baseline results for 1991–93 reported in Table 2. These baseline results include all establishments, some of which may report on plants in more than one Administrative Region or county. We therefore undertake the following two robustness tests. In column (2) we report coefficient estimates for the sub-sample of single-plant establishments where overlap does not occur. We find a very similar pattern of results. Second, in column (3), we allocate establishment-level data to plants on the basis of their shares of establishment employment. All plants associated with an establishment are given the same relative wage bills. This robustness test introduces a bias against rejecting RFPE. However, even with this bias, we continue to reject RFPE with a positive and significant coefficient in the South-East and negative and significant coefficients in the Yorkshire and Humberside and Northern regions of England, Wales and Scotland.

The baseline sample includes the population of establishments with more than 100 employees and a sample of establishments with fewer than 100 employees. In

TABLE 6

Bilateral region-pair rejections

Region definition	*Percentage of rejections (%)* *5% significance level*	*10% significance level*	*Distribution of rejections across base regions (10% significance level)* *Minimum*	*Mean*	*Maximum*
Administrative Regions					
1976–78	58	64	3	6	9
1979–81	48	57	2	6	9
1982–84	54	59	3	6	9
1985–87	51	57	2	6	9
1988–90	44	56	2	6	9
1991–93	36	44	2	5	9
Counties					
1976–78	22	30	2	25	53
1979–81	15	23	1	18	41
1982–84	17	24	3	20	40
1985–87	16	24	3	20	41
1988–90	20	28	3	20	41
1991–93	20	26	4	20	41

Notes: Bilateral rejections of relative factor price equality are based on estimation of equation (8) for all possible base regions. Statistical significance is based on standard errors that are robust to heteroscedasticity. Region-industry cells with less than three establishments are excluded from regressions. The number of region-industry cells is higher in 1976–78 (as are the number of rejections) because of finer sampling of establishments.

Source: Authors' calculations using ARD (source: ONS).

TABLE 7

Bilateral rejections by base region

Administrative Region	*1976–78*	*1979–81*	*1982–84*	*1985–87*	*1988–90*	*1991–93*	*Mean*
South-East	9	9	9	9	9	9	9
East Midlands	6	5	5	2	3	4	4
South-West	8	6	5	5	8	5	6
East Anglia	3	4	3	4	3	2	3
North-West	5	4	4	6	4	3	4
West Midlands	5	3	4	5	4	3	4
Scotland	5	5	6	4	6	3	5
Wales	5	2	4	4	2	3	3
Yorks & Humberside	6	6	6	5	4	2	5
Northern	6	7	7	7	7	6	7
Mean	6	6	6	6	6	5	5

Notes: Bilateral rejections of relative factor price equality at the Administrative Region level based on estimation of equation (8) for all possible base regions. There are a total of 10 Administrative Regions. Statistical significance is based on standard errors that are robust to heteroscedasticity. Region-industry cells with less than three establishments are excluded from regressions. Dashed lines group administrative regions into relative wage cohorts according to relative wage point estimate sign and statistical significance (Table 5).

Source: Authors' calculations using ARD (source: ONS).

TABLE 8

Robustness checks for 1991–93 Administrative Region coefficients

Administrative Region (1991–93)	*Base results*	*Single-plant establishments*	*Plant-level estimation results*	*Establishments with >100 employees*	*At least five establishments per region-industry pair*	*Base sample, weighted estimation*
South-East	0.267**	0.219**	0.321**	0.342**	0.259**	0.220**
East Midlands	0.003	0.035	0.003	−0.021	−0.024	−0.038
South-West	0.002	0.074	0.034	0.009	0.003	0.019
East Anglia	−0.029	−0.025	−0.054	0.017	−0.035	0.060
North-West	−0.030	−0.042	−0.032	−0.036	−0.039	−0.037
West Midlands	−0.057	0.006	−0.048	−0.110**	−0.077	−0.050
Scotland	−0.136**	−0.157**	−0.130**	−0.124**	−0.109**	−0.143**
Wales	−0.155**	−0.146**	−0.180**	−0.233**	−0.201**	−0.165**
Yorks & Humberside	−0.158**	−0.153**	−0.158**	−0.210	−0.121**	−0.109**
Northern	−0.200**	−0.114**	−0.185**	−0.196**	−0.176**	−0.183**
F-stat (p-value)	0.000	0.000	0.000	0.000	0.000	0.000
Observations	1,580	1,512	1,599	1,319	1,208	1,580

Notes: Table displays point estimates from estimation of equation (7) on Administrative Regions. Column (1) reproduces results from Table 2. Column (2) reports results based on a sample of single-plant establishments. Column (3) reports results where establishment-level data are allocated to plants on the basis of their shares of establishment employment. Column (4) reports results for the population of establishments with more than 100 employees. Column (5) reports results where region-industry pairs are excluded if they have fewer than five establishments. Column (6) reports weighted estimation results using industry-region employment as weights. ** denotes statistical significance at the 5% level. Statistical significance is based on standard errors that are robust to heteroscedasticity. Dashed lines group Administrative Regions into relative wage cohorts according to relative wage point estimate sign and statistical significance (Table 5).

Source: Authors' calculations using ARD (source: ONS).

order to ensure that our results are not driven by the presence of a non-random sample of smaller establishments, column (4) reports results separately for the population of establishments with more than 100 employees. Again, we find a similar pattern of results.

Our coefficient estimates are mean values across all four-digit industries in each region. Although each region has a large number of establishments across all industries, some individual industries within a region may contain few establishments. As measurement error at the establishment level is a potential concern, in column (5) we drop all four-digit region-industries that contain fewer than five establishments. Once again RFPE is rejected and a similar pattern of estimated coefficients is observed. Finally, to ensure that the results are not driven by small region-industries with low employment levels, column (6) reports similar estimation results from a weighted regression using region-industry employment as weights.

The results from this section show economically and statistically significant departures from RFPE across regions in the UK. These departures from RFPE have persisted and increased over time and are robust across a variety of different specifications.

IV. Relative wage differences and industrial structure

In this section we investigate whether departures from RFPE are associated with differences in the set of industries that regions produce. We run the following ordinary least square (OLS) regression:

$$Z_{rs} = \beta_0 + \beta_1 |\alpha_{rs}^{NP}| + \beta_2 I_r + \beta_3 I_s + u_{rs} \tag{11}$$

where Z_{rs} is a measure of the similarity of the industrial structure for regions r and s, α_{rs}^{NP} are the estimated bilateral relative wage bill differences from equation (8), I_r and I_s are the number of industries produced by region r and s, respectively, and u_{rs} is a stochastic error.[17]

Equation (11) is estimated using two alternative measures of regions' similarity in terms of industrial structure. The first is a measure of industry overlap that is based on the number of industries that are present in both regions. This measure, $Z_{rs} = I_{rs}$, is a count of the number of industries common to both region r and region s. Similarity of industrial structure across regions is *increasing* in this measure.

The second measure of industrial similarity exploits information not just on whether an industry exists in each region but also on its level of economic activity. This measure is widely used in the empirical geography literature (see, in particular, Krugman, 1991), and is the sum of the absolute differences in industry shares of manufacturing employment in the two regions, so

$$Z_{rs} = \sum_j \text{abs}\left(\frac{L_{rj}}{L_r} - \frac{L_{sj}}{L_s}\right). \tag{12}$$

The industry share measure, or specialization index, takes the value zero if regions r and s have exactly the same industrial structure and attains a maximum value of 2 when the regions' employment structures are such that they have no industries at all in common. Similarity of industrial structure across regions is *decreasing* in this measure. In calculating both measures of industrial similarity we make use of the plant-level population data from the ARD, again taking averages over 3-year periods (2 years for the initial period). These data on the population of plants are available from 1980 onwards and so we now focus on the 1980–93 portion of our sample period.

While the Krugman measure has the advantage of taking into account employment shares in industries rather than simply whether or not industries are active in each region, the count measure has the advantage of being closer to the predictions of neoclassical trade theory. In the general equilibrium of neoclassical trade theory, regions in the same cone of diversification (active in the same set of industries) have the same relative factor prices. For these reasons, and to establish that our results are

[17]Up until this point, all estimates have been based solely on industries that exist in both regions r and s.

TABLE 9

Industrial specialization across Administrative Regions and counties

Administrative Regions	*Year*	*Minimum*	*Median*	*Maximum*
Regions per industry as per cent of all regions	1980–81	10%	100%	100%
	1991–93	20%	100%	100%
Industries per region as per cent of all industries	1980–81	90%	94%	99%
	1991–93	91%	98%	100%
Bilateral overlap as a per cent of the larger region's industries	1980–81	89%	94%	98%
	1991–93	92%	96%	99%
Krugman specialization index	1980–81	0.63	0.94	1.20
	1991–93	0.71	0.96	1.10
Counties				
Regions per industry as per cent of all regions	1980–81	2%	73%	100%
	1991–93	5%	86%	100%
Industries per region as per cent of all industries	1980–81	26%	69%	97%
	1991–93	34%	77%	98%
Bilateral overlap as a per cent of the larger region's industries	1980–81	28%	69%	97%
	1991–93	38%	79%	97%
Krugman specialization index	1980–81	0.91	1.43	1.84
	1991–93	0.87	1.40	1.81

Notes: Bilateral overlap is the number of industries that a pair of regions have in common, expressed as a percentage of the number of industries in the region with the largest number of industries. Specialization index is the Krugman (1991) index, which equals the sum across industries of the absolute values of differences in industry employment shares between a pair of regions.

Source: Authors' calculations using ARD (source: ONS).

not driven by the consideration of a particular measure of industrial similarity, we report results using both measures.

Table 9 summarizes the extent of industrial similarity across Administrative Regions and counties in the start and end periods of our analysis. The first two rows in each section of the table show the minimum, median and maximum per cent of regions per industry, i.e. the extent to which industries cover a broad range of regions. For example, looking at the figures for counties, in 1980–81 the median industry was produced in 73% of counties, increasing to 86% by 1991–93. While some industries are produced in all regions, some were only produced in two or three counties.

The next two rows in each section of Table 9 report the minimum, median and maximum per cent of industries per region, i.e. the extent to which regions produce a wide variety of industries. On this measure we see an increase over time in the diversity of production within regions. The next two rows in each section of the table show the extent of bilateral industry overlap across regions. The percentage of industries that two regions have in common is defined as the number of industries produced in both regions as a percentage of the number of industries in the region which produces the larger number of industries. We find evidence of an increase in bilateral industry overlap over time using the measure based on the number of industries present in a region. We find similar results using the Krugman specialization index, although the

TABLE 10

Industrial specialization, counties, pooled 3-year cross-sections (1979–81 to 1991–93)

Dependent variable	*Common industries*		*Krugman specialization index*	
Absolute relative wage bill gap	−1.462	−0.387	0.042	0.019
	(0.354)	(0.301)	(0.006)	(0.005)
Number of industries in *r*	0.591	0.545	−0.002	−0.002
	(0.005)	(0.011)	(0.000)	(0.000)
Number of industries in *s*	0.707	0.625	−0.003	−0.001
	(0.003)	(0.009)	(0.000)	(0.000)
Observations	9,765	9,765	9,765	9,765
Year dummies	Yes	Yes	Yes	Yes
Bilateral county pair dummies	No	Yes	No	Yes

Notes: Specialization index is the Krugman (1991) index, which equals the sum across industries of the absolute values of differences in industry employment shares between a pair of regions. A pair of regions is more similar in industrial structure if they have a larger number of industries in common and a lower value of the Krugman specialization index. Robust standard errors in parentheses.

Source: Authors' calculations using ARD (source: ONS).

increase in the degree of similarity in industrial structure over time using this measure is less marked.

The neoclassical model of trade predicts that departures from RFPE are associated with differences in the set of industries that regions produce. As this correlation is an equilibrium relationship between two endogenous variables, a test for the statistical significance of the coefficient β_1 should be interpreted purely as a test for the statistical significance of this equilibrium relationship, and we caution that the coefficient β_1 cannot be given a causal interpretation. To estimate equation (11) we pool the cross-sections of industrial similarity measures and relative wage bill differences over time, including a full set of time dummies to control for macroeconomic shocks and to abstract from secular trends in the left- and right-hand-side variables. In view of the small number of Administrative Regions, we focus on estimating equation (11) at the county level.

The first and third columns of Table 10 present our baseline estimates of equation (11) for the two measures of industrial similarity. Using industry overlap as the dependent variable, we find a negative and statistically significant coefficient on the estimated relative wage bill differences. This result is consistent with the hypothesis that regions with bigger differences in relative factor prices have fewer industries in common. Using the Krugman (1991) industry-share measure, we obtain a positive and statistically significant coefficient, again as predicted by theory. Region pairs with greater differences in relative factor prices are indeed less similar in terms of industrial structure.

In the second and fourth columns of Table 10 we include region-pair dummies so that the relationship between industrial similarity and differences in relative

TABLE 11

Industry churning and relative wagebill changes

Regressor	*Percentage of industries added or dropped*
Absolute change in relative wage bills, 1979–81 to 1991–93	70.9 (29.60)
Constant	9.7 (3.20)
Obs.	63
R^2	0.13

Notes: Robust standard errors in parentheses.
Source: Authors' calculations using ARD (source: ONS).

factor prices is identified solely from the time-series variation in the data. Again as predicted by theory, we find that regions that experienced diverging relative factor prices became less similar in terms of industrial structure.

Finally, we also look at the relationship between changes in industrial structure and changes in relative factor prices over time within regions. If relative factor prices are important for firms seeking to minimize costs, then regions that experience larger changes in their relative factor prices over time should display greater churning in their industry mix in terms of the entry of new industries and exit of existing industries. To investigate this, we run an OLS regression of the form,

$$\mathrm{CHURN}_r = \alpha + \beta_d |\alpha^{NP}_{r,91-93} - \alpha^{NP}_{r,79-81}| + \epsilon_r \qquad (13)$$

where the dependent variable CHURN_r is the percentage of industries either added or dropped by region r between the periods 1979–81 and 1991–93 relative to its number of industries in the first period, and $|\alpha^{NP}_{r,91-93} - \alpha^{NP}_{r,79-81}|$ is the absolute value of the change in region r's wage bill ratio relative to the UK average between the two periods. We note again that the relationship between industry structure and relative factor prices is an equilibrium relationship between two endogenous variables and therefore caution that the coefficient β_d should not be given a causal interpretation.

As shown in Table 11, there is a positive and significant correlation between the extent of industry churning and the changes in estimated wage bill ratios. The absolute value of the change in a region's wage bill ratio on the right-hand side of the regression varies from 0.002 to 0.3 across counties, with a mean value of 0.1. Therefore, for the mean change in relative wage bill ratios, the results in Table 11 imply a value of industry churning of 7.1 percentage points. As a robustness test we also re-estimated the specification in equation (13) using the Krugman measure, which yielded a similar pattern of results, consistent with the results including region-pair dummies in the second and fourth columns of Table 10.

V. Implications

Spatial variation in industrial composition can leave regions asymmetrically exposed to common external shocks. Workers in regions specializing in industries intensive in the use of blue-collar workers, for example, may be more adversely affected by declines in the world market price of goods that use blue-collar workers intensively than workers in regions where these goods are not produced (who simply enjoy a terms-of-trade gain). In this section we demonstrate that UK regions with low wages of white-collar workers relative to those of blue-collar workers tend to specialize in industries intensive in the use of white-collar workers. We then highlight the link between this specialization and regions' differential susceptibility to competition from low-wage countries.

The first two columns of Table 12 report the share of manufacturing employment in industries with high and low intensity of use of white-collar workers across the 10 Administrative Regions.[18] Regions are sorted by their share of employment in white-collar-intensive industries, from high to low. As indicated in the table, the South-East and South-West in 1991–93 had relatively large shares of their manufacturing employment in high white-collar-intensive industries, while the corresponding percentages were relatively low for the West Midlands, Wales and Yorkshire and Humberside. Shares of employment in low white-collar-intensive industries, on the other hand, display the opposite pattern, being relatively low for the South-East and South-West and relatively high for the East Midlands, Wales and Yorkshire and Humberside.

The second two columns of Table 12 report the share of the population with high (Degree or Higher) and low (No Qualifications) educational attainment by Administrative Region. These data are available for Administrative Regions as a whole from ONS (1993). These shares reveal that regional levels of educational attainment are correlated with specialization in white- and blue-collar-intensive industries. The South-East and South-West have relatively high educational attainment while the corresponding attainment for the West Midlands, Wales and Yorkshire and Humberside is relatively low. Under the assumption that occupation- and education-based measures of skills are positively correlated, these findings together with our results in Table 5 suggest that skill-abundant regions are an attractive location for skill-intensive industries.

We illustrate geographic variation in the United Kingdom's susceptibility to international trade shocks by examining how import competition from relatively labour-abundant economies varies across regions. We consider exposure to two groups of countries: low-wage countries, i.e. those with less than 5% of US per capita GDP; and former communist countries.[19] To assess regional exposure to these imports, we

[18]Industries with high (low) intensity of use of white-collar workers are defined as those with a ratio of white-collar to blue-collar workers in the top (bottom) third of four-digit industries, defined using the establishment-level sample. Shares of employment in these two groups of industries within regions are then calculated using the plant-level population. The lowest white-collar-intensive industries by this measure are Textiles, Footwear and Leather Goods, while the highest white-collar-intensive industries are Instrument Engineering, Computers and Mechanical Engineering.

[19]Low-wage countries include China and India. The former communist countries are comprised of the former USSR, Romania, Hungary, Poland and Bulgaria, former Yugoslavia, former Czechoslovakia and former East Germany.

TABLE 12

Industrial structure and educational attainment

Administrative Region	*% Manufacturing employment in region in industries with high- and low-white-collar intensities, 1991–93*		*% of economically active population in region by qualification, 1992*	
	High-white-collar industries	*Low-white-collar industries*	*Degree (or equivalent) and higher*	*No qualifications*
South-East	52	26	15	21
South-West	46	30	10	22
East Anglia	39	38	10	25
North-West	38	36	10	25
Scotland	34	36	10	23
Northern	34	34	8	25
East Midlands	30	47	9	28
Wales	29	47	9	26
West Midlands	28	40	8	30
Yorkshire and Humberside	27	48	9	26

Notes: In columns (1) and (2), industries with high (low) white-collar intensity are defined as the top (bottom) third of four-digit manufacturing industries in terms of the ratio of white-collar (non-production) to blue-collar (production) workers; employment in these industries by region is calculated using the data on employment in the population of all manufacturing plants. Regions are sorted by their share of employment in high white-collar-intensive industries.

Source: Columns (1) and (2) are based on authors' calculations using ARD (source: ONS). The data on educational attainment in columns (3) and (4) is from (ONS, 1993).

first calculate the share of total UK imports in each four-digit industry in each year coming from each group of countries.[20] We then compute regions' weighted-average exposure to these imports in each year using their four-digit industry employment shares at the beginning of the time period as weights. Time-series variation in regions' exposure, therefore, stems both from a different initial pattern of specialization across industries and from differential rates of import growth across industries.

Figure 3 shows the evolution of import competition for three groups of Administrative Regions. These groups, defined according to the sign and statistical significance of the relative wage bill results in Table 2, are first, South-East; second, East Midlands, South-West, East Anglia, North-West and West Midlands; and third, Scotland, Wales, Yorkshire and Humberside and Northern. The time series for each group is the average of the trade exposure of all of its members. For all three groups, the average share of imports from low-wage economies starts to rise, beginning in the 1980s. While the trends over time are relatively similar for the three groups of regions, the higher percentage of white-collar-intensive manufacturing industries in the South-East leads to lower exposure to this type of import competition in all years. Regions with high relative wages of white-collar workers to blue-collar workers that specialize

[20]This computation relies on data from Feenstra (2000) and a concordance from SITC codes to UK SIC codes provided by Robert Elliott.

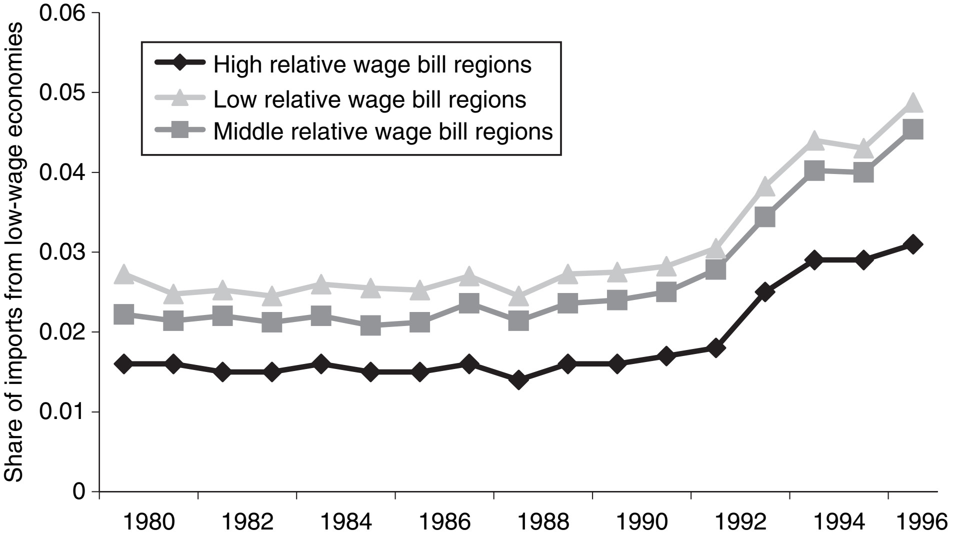

Figure 3. Imports from low-wage countries: each line represents the average share of imports from low-wage countries across the regions in the group. High relative wage bill regions include South-East; low relative wage bill regions include Yorkshire and Humberside, Northern, Wales and Scotland; middle relative wage bill regions include East Anglia, South-West, West Midlands, East Midlands and North-West. Source: authors' calculations using ARD (source: ONS) and Feenstra (2000)

in blue-collar-intensive industries have the greatest exposure. The trends in Figure 4 are similar, with northern regions being the most exposed to import competition from the former Soviet Union and Eastern European economies.[21]

Regional variation in GDP per head across UK regions is well known. The results in this section suggest that public policy designed to boost the performance of lagging regions should take patterns of regional comparative advantage into account. For example, development assistance oriented towards inducing white-collar-intensive industries to move to lagging regions may fail if it contradicts the targeted regions' relative endowments and factor prices. On the other hand, policies promoting human capital accumulation and occupational mobility, which make lagging regions fundamentally more attractive to white-collar-intensive industries, may meet with greater long-term success.

VI. Conclusions

We examine the extent of relative wage variation across geographic areas of the United Kingdom. Despite the UK being a small, densely populated country with highly integrated goods markets and potential for factor mobility, we find strong

[21] Similar trends are found with respect to county-level international trade exposure. We do not report them to conserve space, but they are available from the authors upon request.

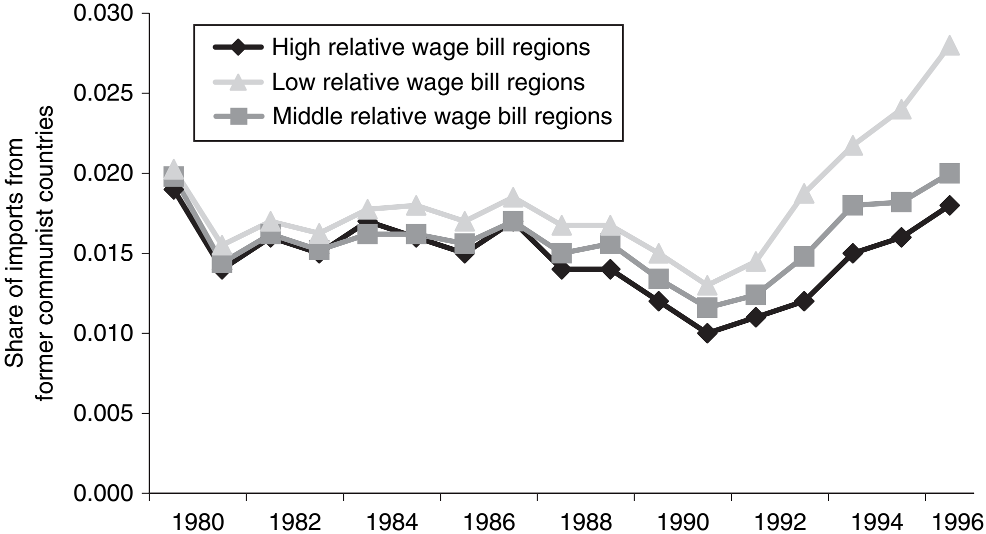

Figure 4. Imports from former Soviet Union and Eastern and European economies: each line represents the average share of imports from former Soviet bloc countries across the regions in the group. High relative wage bill regions include South-East; low relative wage bill regions include Yorkshire and Humberside, Northern, Wales and Scotland; middle relative wage bill regions include East Anglia, South-West, West Midlands, East Midlands and North-West. Source: authors' calculations using ARD (source: ONS) and Feenstra (2000)

evidence against RFPE. Using finely detailed industry and region data, we report statistically significant and economically large relative factor price differences across Administrative Regions and counties within the UK, which persist over time. We find that areas abundant in white-collar workers, predominantly in the South-East of England, exhibit lower wages of white-collar workers relative to blue-collar workers than areas in the Northern and Yorkshire and Humberside regions of England, Wales and Scotland.

We also present evidence of a strong relationship between regional production structure and relative factor prices. We demonstrate that pairs of regions with larger estimated differences in relative factor prices are less similar in industrial structure. We find that regions with low relative wages of white-collar workers produce a more white-collar-intensive mix of industries. Our findings emphasize the way in which industrial structure and relative factor prices are jointly determined in general equilibrium.

The results contribute to our understanding of regional variation in economic outcomes within the UK. Our analysis suggests that variation in relative factor prices plays an important role in shaping firms' location decisions. We also illustrate that variation in industrial structure across regions can imply differential exposure to common external shocks. This relates to a broader debate about the impact of the European

Monetary Union and the extent to which individual regions within a country are more similar than regions across countries.

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