Skill imbalances in the UK labour market: 1979-99

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Abstract

This paper investigates the evolution of skill imbalances in the UK labour market over the past two decades. Movements in the relative ease with which firms can recruit skilled workers can affect unemployment, inflation, and productivity. Any assessment of changes in the skill balance is complicated by the fact that different indicators often send conflicting messages. Such conflicts could reflect the underlying definitions of skilled and unskilled workers, as well as differences in the sensitivities of each measure to alternative market shocks. Our analysis casts doubt on the reliability of standard measures of unemployment dispersion across educational groups, and the Confederation of British Industry ratio of skilled and unskilled labour shortages, as measures of skill imbalance. The gap between the demand for, and the supply of, educated labour has in fact increased steadily over the past two decades, particularly for those workers with graduate-level qualifications. So the apparent decline in the NAIRU over the recent cyclical upswing cannot be attributed to an improvement in the relative ease with which firms can hire educated workers.

Summary

This paper examines the evolution of skill imbalances in the UK labour market over the past two decades. A rise in skill imbalances is defined as an increase in the difficulty of recruiting skilled workers compared with unskilled workers. Our investigation is primarily motivated by the observation that different balance indicators often send conflicting messages. These differences could reflect several factors, including the underlying definitions of skilled and unskilled workers, and the sensitivity of each measure to different types of labour market shock.

We consider three approaches in the literature to measuring skill imbalances:

- Comparing the growth rate of the ratio of the skilled and unskilled wage bill shares with the growth rate of the ratio of the skilled and unskilled labour force shares. This measure, which uses educational attainment to define skill groups, suggests that imbalances have been rising steadily over the past two decades, especially in the market for graduates. This implies that the widely perceived decline in the NAIRU over the recent cyclical upswing does not reflect an improvement in the relative ease with which firms can hire educated workers. Movements of this index seem to reflect genuine reallocations of labour demand/supply across educational groups.

- The dispersion of wages and unemployment rates across education groups. These measures also point to rising imbalances since 1979. However, our analysis suggests that the unemployment measures are primarily driven by aggregate labour demand/supply shocks that have no implications for the skill balance. This undermines their reliability as measures of it.

- The Confederation of British Industry (CBI) ratio of skilled labour shortages and unskilled labour shortages in manufacturing. The CBI ratio indicates that imbalances have declined over 1979-99. This does not necessarily contradict the other measures, because the CBI data are not based upon an explicit definition of skilled and unskilled workers and only cover the manufacturing sector. However, the robustness of the CBI ratio to skill-neutral shocks is also questionable.

1 Introduction

The potential macroeconomic consequences of the demand for skilled labour outpacing its supply include upward pressure on unemployment, pay, and inflation, and downward pressure on productivity.(1) So an accurate picture of rises in skill imbalances, defined as an increase in the difficulty of recruiting skilled labour compared with unskilled labour, is of obvious importance for monetary policy. It is also relevant for understanding more structural matters. For example, the widely perceived decline in the NAIRU over the current upswing (see Astley and Yates (1999), Wadhwani (2000a, b) and Nickell (2001)) could reflect improvements in the supply of skilled labour relative to demand.(2)

One problem with attempting to understand changes in the skill balance is that different measures often send conflicting signals. These differences could reflect several factors, such as the underlying definitions of skilled and unskilled workers and the sectoral coverage of each measure. Another potentially important source of friction is the extent to which the movements of a given indicator are driven by relative or absolute labour demand/supply shocks. Relative shocks can be defined as changes in the distribution of a given level of labour demand/supply across skill groups. Such disturbances will affect the skill balance. Absolute shocks, on the other hand, are changes in the levels of labour demand and/or supply. These shocks have no implications for the skill balance. A reliable indicator should fully reflect skill-specific impulses, while being robust to those that are skill-neutral.

This study covers the key imbalance measures in the literature. It begins with an index proposed by Manacorda and Petrongolo (1999). Given a production function with skilled and unskilled labour inputs, they show that a reallocation of demand towards skilled workers will raise the ratio of the skilled share of the aggregate wage bill to the unskilled share of the wage bill. This is because the relative wages and/or employment rates of skilled workers increase. On the other two balance between the demand for, and supply of, educated workers. The index will be impervious to aggregate disturbances provided that such shocks do not alter the distribution of the wage bill or labour force shares across educational groups. Although we cannot quantify the effects of such hand, a reallocation of labour supply towards the skilled will raise the ratio of the skilled share of decades, particularly in the market for graduates. This implies that the apparent easing of the labour force to the unskilled share. Their measure compares the growth rates of these relative demand and relative supply indicators using educational attainment as a measure of skills. Our estimates based on this methodology show that imbalances have increased steadily over the past NAIRU over the current cyclical upswing cannot be attributed to favourable movements in the skill-neutral impulses, they appear to be relatively unimportant compared with shifts in the distribution of labour demand and supply across the educational groups.

We then turn to the absolute and relative dispersion of wages and unemployment according to educational attainment. Like the Manacorda and Petrongolo index, these measures derive from

(1) For example, an excess demand for skilled workers can reduce productivity growth by forcing firms to substitute unskilled workers on tasks that optimally require skilled labour (Haskel and Martin (1993a, b)). See also Krugman (1994) and Bannock Consulting (2000).
(2) Astley and Yates (1999), Wadhwani (2000a, b), and Nickell (2001) discuss several explanations for the apparent decline in the NAIRU.

the idea that relative shocks will change relative wage and unemployment rates. The wage dispersion indicators, which appear to be robust to aggregate factors, point to worsening imbalances since 1979. However, the unemployment measures are ambiguous. One reason for this indeterminacy is that both unemployment indices are strongly correlated with cyclical movements in aggregate unemployment, albeit in opposite directions. A key innovation of this paper is to explore these patterns by formally investigating the relationship between the unemployment dispersion and Manacorda and Petrongolo imbalance measures. This investigation suggests that the observed movements in the dispersion indices can be attributed to skill-neutral shocks.

Finally, we examine the ratio of the Confederation of British Industry (CBI) indices of skilled and unskilled labour shortages. These indices are based on a sample of manufacturing firms who are asked whether shortages of skilled and 'other' labour are expected to limit output over the next four months. A rise in the ratio can be interpreted as signifying an increase in the relative excess demand for skilled workers. It could also mean that firms are finding it more difficult to fill a skilled job slot compared with an unskilled vacancy. In contrast to the other indicators, which use education to proxy for skills, the CBI ratio implies that imbalances have declined since 1979. This may reflect the possibility that the ratio captures skills that lie outside those associated with formal educational status. Trends in the manufacturing sector may have differed from the rest of the economy. However, there is also some evidence that this decline is driven by aggregate factors.

2 Evidence from shifts in net demand and supply across education groups

2.1 The imbalance measure

The Manacorda and Petrongolo (1999) index is based upon the idea that a relative increase in demand for skilled workers will increase their share of the aggregate wage bill, while a relative increase in the supply of skilled workers will increase their share of the workforce. The index in effect compares these demand and supply changes.

To fix ideas, consider the following constant elasticity of substitution (CES) production function for aggregate output Y with skilled (N_s) and unskilled (N_u) labour inputs:

$$Y = A(a_s N_s^{-r} + a_u N_u^{-r})^{-1/r} \qquad \textbf{(1)}$$

where A represents technological progress, $a_s + a_u = 1$, and $s = 1/(1+r)$ is the elasticity of substitution between skill groups. a_s and a_u are relative productivity indices for each skill group.(3)

If W_s and W_u are the real wage rates of each skill group, then profit maximisation implies:

$$\frac{a_s}{a_u} = \frac{W_s (N_s)^{1/s}}{W_u (N_u)^{1/s}} \qquad \textbf{(2)}$$

Let L_s and L_u signify the labour force of each group. Thus $E_i = N_i/L_i$ is the group i employment rate. Now define $l_i = L_i/L$ as the labour force share of group i, where L is the total workforce. Dividing both sides by $(L_s/L_u)^{1/s}$ and using $l_i = L_i/L$ gives:

$$\frac{a_s}{a_u}\Bigg/\left(\frac{l_s}{l_u}\right)^{1/s} = \frac{W_s (E_s)^{1/s}}{W_u (E_u)^{1/s}} \qquad \textbf{(3)}$$

Jackman *et al* (1999) interpret the left-hand side of equation **(3)** as a measure of skill mismatch, *SM*. Taking logarithms of both sides and totally differentiating throughout leads to:

$$d\ln\left(\frac{a_s}{a_u}\right) - (1/s)\,d\ln\left(\frac{l_s}{l_u}\right) = d\ln\left(\frac{W_s}{W_u}\right) + (1/s)\,d\ln\left(\frac{E_s}{E_u}\right) \qquad \textbf{(4)}$$

The left-hand side of **(4)** can be interpreted as the difference in the growth rates of the demand and supply of skilled labour relative to the unskilled. Thus it identifies the growth of skill imbalances, dln*SM*. A change in the relative demand for skilled workers is captured by $d\ln(a_s/a_u)$, which is the growth of the ratio of skilled and unskilled wage bill shares adjusted for

(3) This production function can be generalised to include capital and other factors of production without affecting the results.

their substitutability in the production process.[4] l_s/l_u is the ratio of the skilled and unskilled labour force shares. So $(1/s)\mathrm{dln}(l_s/l_u)$ traces movements in the supply of skilled labour relative to unskilled labour, again adjusted for input substitutability. The right-hand side of **(4)** expresses the index in terms of relative wages (W_s/W_u) and employment rates (E_s/E_u). An increase in market imbalances translates into either a rise in the relative wages of skilled workers, a rise in their relative employment rates (which is interpreted as a rise in the unskilled unemployment rate relative to the skilled unemployment rate), or a combination of both changes.[5] The index can be easily generalised to the case where the production function has $j>2$ labour inputs. In this case j-1 indices can be constructed where one skill group, typically those with the lowest level of skills, is used as the reference group.

The calculation of the index also requires a value for the elasticity of substitution between skill groups, s. Mismatch growth will decline as s increases. The intuition is that as the elasticity of substitution rises, it becomes easier to satisfy a net demand shift towards skilled labour by reallocating existing labour inputs. Choosing s is discussed below.

2.2 Identifying skilled and unskilled workers

Deriving the index requires the definitions of skilled and unskilled workers. Ideally, these definitions should capture those worker attributes that matter to employers. This is inherently difficult because such attributes vary substantially across employers, and include formal educational qualifications, leadership, reliability, creativity, and punctuality (see Green, Machin and Wilkinson (1998)). There are two main approaches to identifying skilled and unskilled workers in the economics literature—neither of which is fully satisfactory. The first is to allocate workers into skilled and unskilled groups according to the nature of their occupation (see Haskel and Heden (1999)). One problem with this strategy is that many non-manual jobs are unskilled. The other procedure, adopted in this paper, is to use educational attainment.[6] The downside here is that many highly educated workers are in low-skill jobs, which suggests that educational attainment does not capture all the skills that are relevant to employers. Green, McIntosh and Vignoles (1999) estimate that the incidence of 'overeducation' in the UK workforce stood at around 30% between 1986-97. They present evidence that overeducated workers earn less than their similarly educated peers and interpret this as signifying that the overeducated are less productive.[7]

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[4] The adjusted wage bill share of each group, i, can be calculated from the following formula:
$a_i = (W_i\,(N_i)^{1/s})/\Sigma(W_i\,(N_i)^{1/s})$.

[5] Equation **(4)** assumes that the relative demand and relative supply of workforce skills are independent. This helps to simplify the analysis, but is unlikely to hold in practice. Consider a shift in demand towards skilled workers. Any resulting increase in the returns to skills should serve to increase the share of workers who will invest in acquiring the relevant attributes. This will eventually boost the relative supply of skills. Skill demand may also respond to an increase in supply. Machin and Manning (1997) and Acemoglu (1998) develop models where a jump in the relative supply of skilled workers induces firms to create skilled jobs. US evidence suggests that an increase in the relative supply of skilled workers lags movements in relative demand by eight to ten years (see Mincer and Danninger (2000)).

[6] In practice, the results from both approaches are similar, partly because of a positive correlation between being in a non-manual job and educational attainment (see Machin (1996)).

[7] Green, McIntosh and Vignoles (1999) estimate the incidence of overeducation by comparing the qualifications that workers say they need for their current job with those that they actually possess. It might have been better to ask employers about the required education levels.

2.3 *Previous estimates of the growth of skill imbalances*

Before presenting our own estimates of dln*SM* we briefly compare our approach with the existing literature. These studies are summarised in Table A. Notice that:

- All the studies cover only the mid/late 1970s to early 1990s and show that mismatch has grown over the period at an annual average rate of 0.7%-2.2%.[8] We assess the 1979-99 period. This enables us to examine the possibility that any fall in the NAIRU over the current upswing can be attributed to an easing of educational imbalances.

- All of the studies use a dichotomous skilled/unskilled split at either A-level or O-level (or equivalent) educational attainment. In an attempt to isolate better the exact source of any imbalances, we allocate individuals into four educational groups on the basis of highest academic qualification or its notional vocational equivalent: degree, A-level, O-level, and below O-level. The unskilled comprise those whose highest qualification lies in the below O-level category, which includes those with no formal qualifications.[9] The assumption that a given academic qualification can be combined with its vocational counterpart is common in the literature. However, the validity of this procedure is disputed by Robinson (1997), who argues that academic and vocational qualifications should be treated separately because the earnings return to an academic qualification typically exceeds its vocational counterpart. In principle, *SM* could be biased if the academic wage premium has been changing over the past two decades. However, it has been fairly stable over this period (see Robinson (1999)).

- Both Manacorda and Petrongolo (1999) and Jackman *et al* (1999) specify $s=1$. This is on the basis of cross-country panel data regressions of the production function where a common value of s is imposed across countries and time periods. Separate analyses using only British data from the General Household Survey (GHS) also indicate that $s=1$. Nickell and Layard (1999) calculate the index for $s=0.5$, $s=1$, and $s=2$.[10] We also use these values of s as a robustness check.

Table A: Previous studies of imbalance growth

Study	Data	Skilled	Annual mismatch growth
Manacorda and Petrongolo (1999)	GHS 1974-92 (Britain)	A-level +	0.74 ($s=1$)
Jackman *et al* (1999)	GHS 1975-92 (Britain)	A-level +	0.68 ($s=1$)
Nickell and Layard (1999)	NES and LFS 1979-91(UK)	O-level +	0.65 ($s=0.5$), 1.29 ($s=1$), 2.24 ($s=2$)

Notes: GHS=General Household Survey, NES=New Earnings Survey, LFS=Labour Force Survey. All the studies use annual data.

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(8) One reason why the literature focuses on *average* annual growth rates rather than annual growth rates is that the latter are very erratic.

(9) The detailed classifications are shown in the appendix.

(10) The literature offers a wide range of estimates of s across different categories of labour input. For example, Hammermesh (1993) reports estimates for blue and white-collar workers ranging from -0.48 to 6. Estimates of s across education groups (defined by years of schooling) range from 0.61 to 1.34. Autor, Katz and Krueger (1998) point out that s will not only reflect the substitutability of skilled and unskilled workers at the firm level but also the possibilities of outsourcing and substitution between goods and services in consumption. Note that if $s\leq 1$ then no output can be produced without unskilled labour.

To compute the indices we use employment and labour force data from the LFS for 1979-99. Following the literature we measure L_i/L as the share of group i in the working-age population. We use GHS data on weekly wages for 1979-92. Unfortunately the GHS was discontinued in 1997. So to extend the analysis through to 1999, we use Labour Force Survey (LFS) spring quarter wage data for 1993-99.[11]

2.4 Estimating the elasticity of substitution between skill groups

Given the CES production technology, the relative demand for skilled labour inputs can be written as:

$$\ln(N_j / N_u) = -s \ \ln(W_j / W_u) + s \ \ln(a_j / a_u) \qquad \textbf{(5)}$$

where j indexes the degree, A-level and O-level education groups and u indexes those without O-levels.

Under the assumption that relative wages are predetermined, equation **(5)** can be estimated as a three-equation seemingly unrelated regression model where, following Katz and Murphy (1992), the logarithm of relative productivities, $\ln(a_j/a_u)$, is proxied by a linear trend. In line with the theory we constrain the elasticity of substitution to be identical across all three equations.[12]

Table B reports the results. The elasticity of substitution, s, is estimated to be 0.37 and is significantly lower than unity. To check whether this reflected our finer classifications we also used the basic A-level/O-level split. In this case, the estimated value of s increased to 0.47 (with a standard error of 0.24). We also instrumented the relative wages in equation **(5)** with their lags. In this case s=0.25 with a standard error of 0.2 indicating that s is not significantly different from zero, which is implausible. Finally, the model was estimated using only GHS data from 1974-96, but again the results were robust. As an approximation then, we set s=0.5 to derive our benchmark results. It turns out that our core findings are not very sensitive to s.

Table B: Relative labour demand equations: 1979-99

	Degree	**A-level**	**O-level**
Constant	-8.27 (0.29)	-8.13 (0.32)	-6.78 (0.54)
Time trend	0.09 (0.003)	0.08 (0.003)	0.07 (0.009)
Ln(W_s/W_u)	-0.37 (0.18)	-0.37 (0.18)	-0.37 (0.18)
R^2	0.97	0.96	0.89

2.5 Main features of the data

Charts 1 and 2 show that the wage and employment rates of skilled workers have typically grown faster than their unskilled counterparts over the past two decades. Taken at face value, these

[11] The LFS is biannual over 1979-83 and annual from 1984. The relevant values for 1980 and 1982 are interpolated.
[12] This restriction could not be rejected at conventional levels of significance.

results line up with the idea of a sustained reallocation of labour demand towards skilled workers that has not been matched by a commensurate realignment of labour supply. However, as we shall discuss later, they may also reflect aggregate shifts in worker demand and supply.

The most striking feature is that this improvement in the relative fortunes of the skilled has been concentrated among graduates. For example, the wage premium to possessing a degree rose by around 14% between 1979-99. By comparison, the extra return to A-level attainment shows no overall trend, while the O-level premium increased by 5%. The rises in relative employment rates ranged from 16% for graduates, to 6% for those with A-levels.

Chart 1: Relative weekly wages

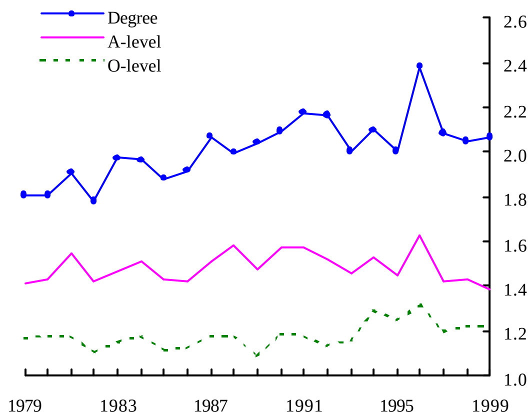


Chart 2: Relative employment rates

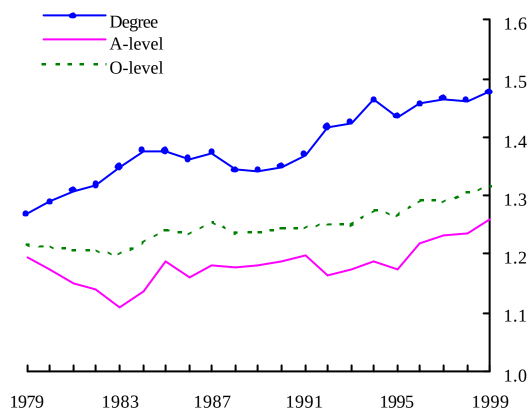


Table C documents the wage bill shares of each group in 1979 and 1999. The shift in the wage bill towards educated workers is clearly dominated by graduates, whose share has increased from 20% to 40%. Over the same period the unskilled (below O-level) share has declined from 48% to 18%. Table D shows that these wage bill changes have been accompanied by a rise in the relative supply of educated workers. Between 1979-99 the percentage of the labour force qualified to degree level rose from 10% to 22%, while the unskilled percentage fell from 68% to 32%.

Table C: Wage bill shares (%)

	Degree	A-level	O-level	Below O-level
1979	19.2	7.2	25.9	47.7
1999	39.3	12.6	30.1	18.0

Table D: Labour force shares (%)

	Degree	A-level	O-level	Below O-level
1979	10	5.3	16.4	68.3
1999	22	12.2	33.3	32.5

2.6 Empirical results

Table E presents the average annual growth rates of the net demand for skilled workers d$\ln(a_s/a_u)$, the net supply of skilled workers $(1/s)\mathrm{d}\ln(l_s/l_u)$, and the imbalance index, dlnSM, for 1979-99 and 1993-99. These estimates assume that s=0.5. Two points emerge:

- The demand for skilled workers has grown faster than the supply of skilled workers at all educational levels. Graduate imbalances have increased by around 2% per year since 1979. This is more than double the rates of those with O and A-levels, which stand at around 1% and 0.4% respectively.

- The results suggest that any decline in the NAIRU over the 1993-99 upswing cannot be attributed to an improvement in the educational balance. Imbalances continued to grow over this period.

Table E: Average annual percentage change in skill demand, supply, and imbalance

	1979-99			1993-99		
	Degree	**A-level**	**O-level**	**Degree**	**A-level**	**O-level**
Demand	16.67	15.31	14.76	13.63	16.29	9.50
Supply	14.59	14.90	13.80	12.15	15.02	7.31
Imbalance	2.08	0.41	0.96	1.47	1.26	2.19

2.7 Sensitivity of the results to s

The benchmark estimates assume that s=0.5. Given the uncertainty surrounding this parameter we also calculated the mismatch index for s=1 and s=2. Table F shows that our conclusions are broadly robust. The most notable qualitative change occurs when s=2. In this case, mismatch among those educated to A-level is estimated to have fallen at an annual rate of around 0.2% since 1993. As expected, the growth rate of educational imbalances declines as s increases.

Table F: Average annual percentage growth in skill imbalance

	1979-99			1993-99		
	Degree	**A-level**	**O-level**	**Degree**	**A-level**	**O-level**
s=**0.5**	2.08	0.41	0.96	1.47	1.26	2.19
s=**1**	1.36	0.16	0.58	0.96	0.28	1.45
s=**2**	1.00	0.04	0.40	0.71	-0.22	1.08

2.8 Interpreting the imbalance index

From equation **(4)** it is clear that the ability of the measure to capture relative shocks will depend upon the extent to which such disturbances feed through to changes in relative wages (W_s/W_u) and/or employment (E_s/E_u) rates. W_s/W_u and E_s/E_u will fully absorb skill-specific shocks if the skilled and unskilled labour markets mimic the textbook Walrasian model. In practice, frictions such as firms' monopsony power in the low-wage labour market (see Card and Krueger (1995)), the costs of changing wages and employment (see Layard *et al* (1991)), the social welfare system, and binding minimum wages, mean that pay and employment may only adjust partially.(13) Such forces could lead to biases in the index, although the direction is ambiguous. For example, suppose the unskilled are paid the minimum wage. Then a demand shift towards skilled workers will raise E_s/E_u by more than it would under full wage flexibility. However W_s/W_u will increase by less than it would have in the absence of a wage floor.

The index will be robust to aggregate shocks if such impulses have no effect on relative wage and relative employment rates. LFS data indicate that workers with below O-level attainment are disproportionately employed in the manufacturing and construction sectors, which are particularly sensitive to aggregate fluctuations (see Ganley and Salmon (1997)).(14) This means that W_s/W_u and E_s/E_u could rise during a general downturn. A similar outcome will emerge if firms hoard skilled labour during recessions but shed unskilled workers more readily (see Bean and Pissarides (1991)).(15)

It is possible to show formally that aggregate shocks can affect the imbalance index. We begin by outlining the relationships between the index, unemployment, and an aggregate shock in the form of a rise in real wages above productivity growth. These are set out in Manacorda and Petrongolo (1999) and Jackman *et al* (1999) who assume that r=0 in equation **(1)**. Thus s=1 and the production function takes the Cobb Douglas form:

$$Y = AN_s^{a_s} N_u^{a_u} \qquad \textbf{(6)}$$

where $a_s + a_u = 1$. The labour demand equation for skilled workers can be written as:

$$\ln W_s = \ln A + \ln a_s + (a_s - 1)\ln(N_s / N_u) \qquad \textbf{(7)}$$

Taking the total derivative of this expression, adding and subtracting $(1-a_s)\mathrm{d}\ln(L_u/L_s)$, and using the fact that $\mathrm{d}\ln a_s = a_u \mathrm{d}\ln(a_s/a_u)$, leads to

(13) Between 1909 and 1993, Wages Councils set minimum pay rates in a wide variety of low-paying sectors. The Councils were abolished in 1993. The National Minimum Wage was introduced in 1999.
(14) Our assessment that workers with below O-level qualifications are disproportionately located in the manufacturing and construction sectors is based upon the following ratio: the employment share of educational group *j* in industry *k*/ the overall employment share of educational group *j*. A ratio above 1 signifies that the educational group is disproportionately employed in industry k. The calculations are based on LFS data over 1979-99.
(15) Such hoarding could be the result of the lower hiring and firing costs of unskilled employees.

$$d\ln W_s = \alpha_u\left(d\ln\left(\frac{\alpha_s}{\alpha_u}\right) - d\ln\left(\frac{\lambda_s}{\lambda_u}\right)\right) + \alpha_u[d\ln(1-U_s) - d\ln(1-U_u)] \tag{8}$$

$$+\left(d\ln A + \ln\left(\frac{N_s}{N_u}\right)d\alpha_s\right)$$

Notice that the term in the first brackets is the change in the logarithm of the mismatch index, dlnSM, when σ=1. A similar expression holds for unskilled labour. Now assume that the wage function of group i can be written as:

$$\ln W_i = z_i - \gamma \ln U_i \tag{9}$$

where z_i includes standard factors that move the wage-setting curve, including the level of benefits, worker bargaining power, and the long-term/short-term composition of the unemployed pool.[16] Taking the total derivative of this wage function gives

$$d\ln W_i = dz_i - \gamma d\ln U_i \tag{10}$$

Let $z=\alpha_s z_s+\alpha_u z_u$. Thus z is the wage bill share weighted average of group-specific wage pressure factors. With this definition, Jackman *et al* (1999) define the change in aggregate wage pressure (AWP) as

$$dAWP = dz - d\ln A - \ln\frac{N_s}{N_u}d\alpha_s \tag{11}$$

where $dz=\alpha_s dz_s+\alpha_u dz_u$. This can be interpreted as the difference between the growth in economy-wide wage pressure, dz, and the feasible growth in real wages, $d\ln A+\ln(N_s/N_u)d\alpha_s$. This feasible growth in real wages captures the growth in total factor productivity (dln$A$) and the growth in output from a productivity shock favouring skilled labour at given employment rates, $\ln(N_s/N_u)d\alpha_s$. Jackman *et al* (1999) define a rise in AWP as an aggregate labour market shock. Manacorda and Petrongolo (1999) argue that $\ln(N_s/N_u)d\alpha_s$ cannot be interpreted as a mismatch factor because it captures the impact of the productivity shock on output which affects both groups. Nonetheless it is worth noting that a rise in SM that is driven by an increase in α_s could also affect aggregate wage pressure. The sign and magnitude of the effect will depend on $\ln(N_s/N_u)$. If $N_s=N_u$, then changes in α_s have no effect on AWP.

Combining **(8)** and **(10)** for skilled and unskilled labour, and using **(11)**, leads to the following equations for dU_s and dU_u:

[16] This specification for the wage-setting function is standard in the labour market literature. See Blanchard and Katz (1997).

$$\mathrm{d}U_s = T\left(-\frac{g a_u}{U_u} (\mathrm{d}\ln SM - (\mathrm{d}z_s - \mathrm{d}z_u)) + \frac{U_u + g(1-U_u)}{U_u(1-U_u)} \mathrm{d}AWP \right) \quad \textbf{(12)}$$

$$\mathrm{d}U_u = T\left(\frac{g a_s}{U_s} (\mathrm{d}\ln SM - (\mathrm{d}z_s - \mathrm{d}z_u)) + \frac{U_s + g(1-U_s)}{U_s(1-U_s)} \mathrm{d}AWP \right) \quad \textbf{(13)}$$

where

$$T = \frac{U_s U_u (1-U_s)(1-U_u)}{g[g(1-U_s)(1-U_u) + a_s U_u + a_u U_s - U_s U_u]} > 0 \quad \textbf{(14)}$$

Using the fact that the aggregate unemployment rate, U, satisfies $U = U_s l_s + U_u l_u$, the change in the aggregate unemployment rate, dU, can be written as:

$$\begin{aligned} \mathrm{d}U &= Tg\left(l_u \frac{a_s}{U_s} - l_s \frac{a_u}{U_u} \right)(\mathrm{d}\ln SM - (\mathrm{d}z_s - \mathrm{d}z_u)) \\ &\quad + T\left(l_s \frac{U_u + g(1-U_u)}{U_u(1-U_u)} + l_u \frac{U_s + g(1-U_s)}{U_s(1-U_s)} \right)\mathrm{d}AWP + (U_s - U_u)\mathrm{d}l_s \end{aligned} \quad \textbf{(15)}$$

We can now examine the comparative statics of the model. Equations **(12)**, **(13)** and **(14)** show that a rise in *SM*, leads to a decline in the skilled unemployment rate and rise in the unskilled unemployment rate. From equation **(15)** it is clear that aggregate unemployment will also rise if skilled workers have higher wages and/or lower unemployment rates. Both conditions hold in the United Kingdom. A rise in aggregate wage pressure (*AWP*) raises all unemployment rates.(17) The final term of equation **(15)** shows that a rise in the skilled share of the labour force reduces aggregate unemployment if skilled workers have lower unemployment rates.

The key point is that a rise in *AWP* is unlikely to have a neutral impact on the imbalance measure because it will affect both W_s/W_u and E_s/E_u. To see this suppose that *AWP* rises because of an upward shift in the wage functions of both groups, so that dz_s=dz_u=Δ>0. Equations **(12)** and **(13)** show that a rise in *AWP* will typically have differential impacts upon U_s and U_u. Thus U_s/U_u and E_s/E_u will change. From equation **(9)** we can see that relative wages satisfy:

$$W_s / W_u = \exp(z_s - z_u - g \ln(U_s / U_u)) \quad \textbf{(16)}$$

Therefore any change in relative unemployment rates will affect relative wages.

The relative influence of skill-specific and skill-neutral disturbances on the Manacorda and Petrongolo index is ultimately an empirical matter. Charts 3 and 4 show that skill imbalances

(17) Notice that the response of these unemployment rates to changes in *AWP* depends on the group-specific unemployment rates and the size of *g*. For example, if the group-specific unemployment rates are all equal so that $U=U_u=U_s$, then the coefficients on *AWP* in **(12)** and **(13)** becomes equal to U/g, which is rising in U and declining in *g*.

have been rising steadily over the past 20 years with no cyclical pattern.[18] Simple regressions confirm that movements in the indices are not significantly related to GDP growth. Although far from conclusive, these findings are consistent with the notion that any impacts of aggregate shocks have been small compared with the net increase in the relative demand for educated workers.

Chart 3: Imbalance indices for s =0.5

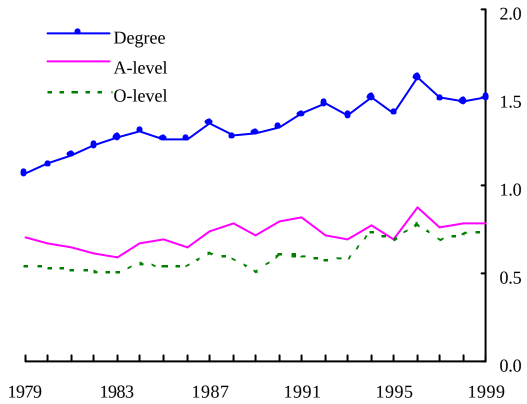
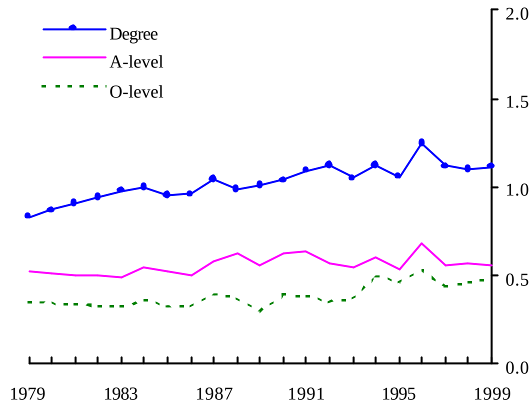


Chart 4: Imbalance indices for s =1



Both charts show ln(*SM*) for the relevant educational group, where SM is defined in equation **(3)**.

3 Evidence from wage and unemployment dispersion across education groups

3.1 The dispersion measures

Measures of wage and unemployment dispersion across skill groups may also be useful for identifying the evolution of the skill balance. We consider two standard dispersion indices: absolute (*AD*) and relative (*RD*). Both indices are popular in the economic literature (see Abraham (1987), Jackman, Layard and Savouri, (1991) and Barwell (2000)).[19]

Each unemployment dispersion index can be calculated from the following formulae:

$$AD_u = \sum_{i=1}^{r} \frac{L_i}{L} \left(U_i - U\right)^2 \tag{17}$$

$$RD_u = \sum_{i=1}^{r} \frac{L_i}{L_t} \left((U_i / U) - 1\right)^2 \tag{18}$$

where L_i is the labour force of skill group *i*, *L* is the aggregate labour force, U_i is the unemployment rate of skill group *i*, *U* is the aggregate unemployment rate, and *r* is the number of groups. AD_u is the labour force share-weighted variance of group-specific unemployment rates. RD_u is the share-weighted variance of the ratio of group unemployment rates to the national

(18) A similar chart for $s=2$ exhibits the same behaviour.

(19) Jackman, Layard and Savouri (1991) develop a model of the relationship between the NAIRU and sectoral unemployment dispersion, which implies that only movements in RD_u are relevant.

average. The corresponding indices for wage dispersion are similarly defined. We define skill groups according to the educational classifications used to compute the Manacorda and Petrongolo imbalance indices. The potential problems with using education to capture skills, discussed in Section 2.2, are also relevant here.

3.2 Empirical results

Charts 5 and 6 display the absolute and relative indices respectively. Pay variation increased steadily between 1979-92, on both measures. This upward trend has continued since 1993 although the data have become more erratic. The unemployment indices offer contradictory messages. AD_u rises to a peak in the mid-1980s, and then generally declines, while RD_u shows a rising trend over the period.

Chart 5: Absolute wage and Unemployment dispersion indices

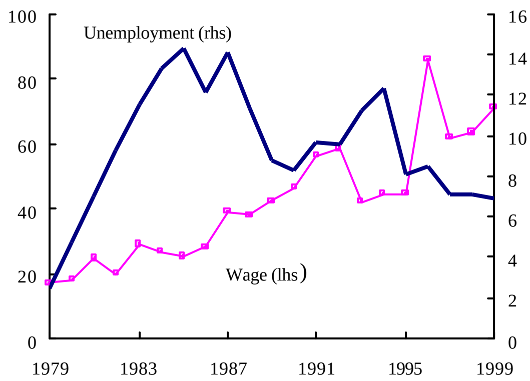
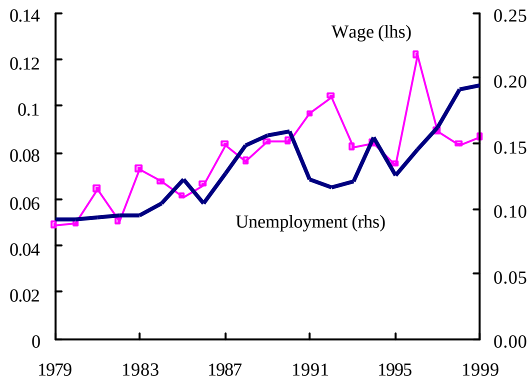


Chart 6: Relative wage and unemployment dispersion indices



Combining equations **(17)** and **(18)** shows that $RD_u \equiv AD_u/(U)^2$ (a similar relationship links RD_w and AD_w). This identity indicates that the relative and absolute measures will move together when the aggregate unemployment (or wage) rate is constant. However their paths can diverge if U (or W) fluctuates. Chart 7 shows that AD_u basically tracks the aggregate unemployment rate, while Chart 8 shows that RD_u and U are inversely related.

Chart 7: Absolute unemployment dispersion and unemployment rate

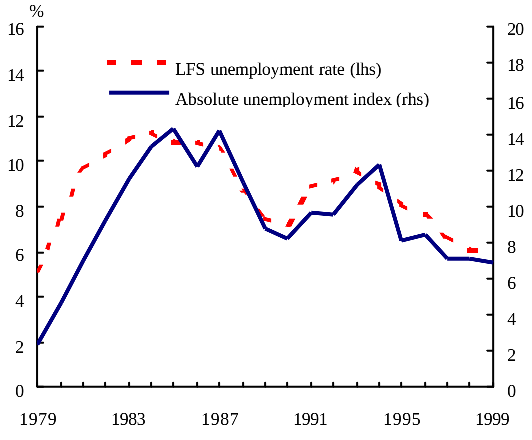
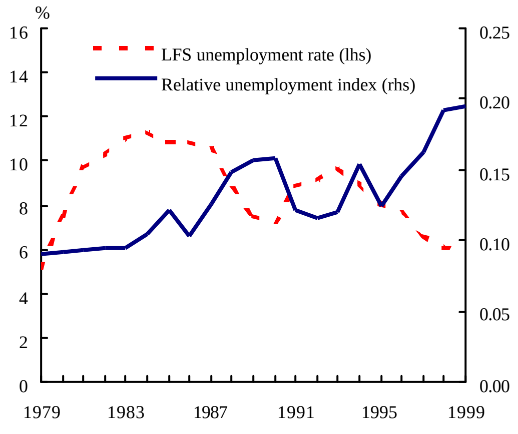


Chart 8: Relative unemployment dispersion and unemployment rate



Charts 9 and 10 plot AD_w and RD_w respectively. The sample average weekly wage, W, is also shown on each chart. Notice that W is relatively stable over the economic cycle compared with aggregate unemployment. The coefficient of variation of aggregate unemployment over 1979-99 is 0.21, while it is 0.13 for wages over the same period. This comparative stability of aggregate wages increases the likelihood that movements in RD_w and AD_w genuinely reflect inter-group movements in pay. Of course, such wage movements need not be the result of changes in the skill balance. Variations in the minimum wage, benefit rates, or a shift in pay bargaining power between firms and different skill groups will also affect wage dispersion.

Chart 9: Absolute wage dispersion and average weekly wage

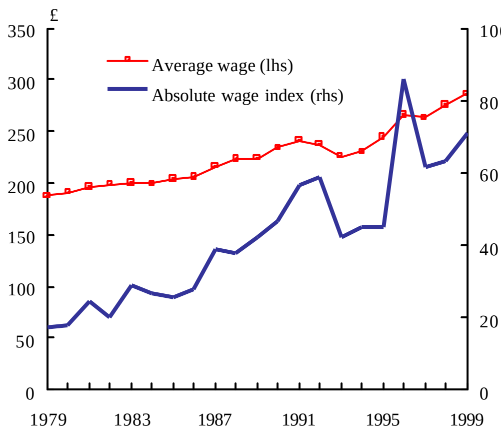


Chart 10: Relative wage dispersion and average weekly wage

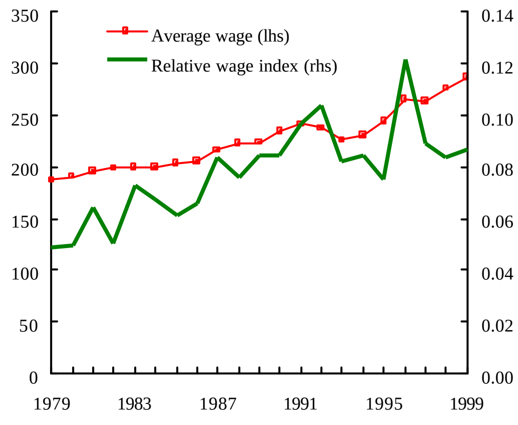


Table G shows the estimated annual growth rate of skill imbalances implied by each dispersion measure. All the indices point to an increase in imbalances between 1979 and 1999. AD_w indicates that wage dispersion accelerated slightly over 1993-99, while RD_w points to a considerable easing of wage variation across educational groups. Both wage indices are very erratic over this period (see Charts 5 and 6), so these conclusions should be treated cautiously. As expected, the unemployment-based dispersion measures move in opposite directions over the cycle. According to AD_u skill mismatch has declined since 1993, while RD_u indicates a marked acceleration in imbalances. This cyclical divergence of the unemployment indices is the focus of the next subsection.

Table G: Annual average percentage change in wage and unemployment dispersion

	1979-99	1993-99
Absolute wage dispersion (AD_w)	6.8	7.5
Relative wage dispersion (RD_w)	2.8	0.7
Absolute unemployment dispersion (AD_u)	5.0	-6.9
Relative unemployment dispersion (RD_u)	3.6	6.9

3.3 Interpreting the behaviour of the unemployment dispersion indices

One reason for the contrasting cyclical behaviour of the unemployment-based indices is that each index is sensitive to particular co-movements of U_i and U. Given some variation in skill-specific unemployment rates, 'dispersion' according to AD_u will rise if U and all the U_i increase by the same proportionate amount, while RD_u will remain unchanged. On the other hand, 'dispersion' as captured by RD_u will decline if the co-movements are equal in percentage point terms, while AD_u will remain unchanged. The intuition for these results is that relative differentials (U_i/U) move in response to a given percentage point or *linear* change in group-specific and aggregate unemployment rates, while absolute differentials (U_i-U) are invariant. In contrast, absolute differentials change in response to *proportional* co-movements between U_i and U, while relative differentials remain fixed.

Does either dispersion measure actually fit the LFS unemployment data? This can be easily examined. If the co-movements of U_i and U are uniformly linear, then the relative measure will vary negatively with U, while absolute dispersion will be completely unrelated to the aggregate unemployment rate. Under proportionality, AD_u will vary positively with U, while RD_u will be independent of U. Finally, if the co-variation of U_i and U lies somewhere between linearity and proportionality then AD_u and RD_u will both move over the unemployment cycle. Table H reveals that neither measure is invariant to U. Consequently, both linearity and proportionality can be rejected. The fact that both indices vary in opposite directions with U signals that co-movements of U_i and U lie somewhere between linearity and proportionality.

Table H: Regression results for relative and absolute unemployment dispersion indices

	RD_u	AD_u
Constant	**0.21***	**-2.48***
U	**-0.01***	**1.37***
R^2	**0.30**	**0.67**

1979-99 annual data. *Significant at 5% level.

We can also use the Manacorda and Petrongolo (1999) framework to assess further the behaviour of the unemployment dispersion indicators. To focus upon essentials we assume that movements in SM reflect relative skill demand and supply shocks.[20] With this assumption we can show that the cyclical patterns of AD_u and RD_u are inconsistent with skill-specific shocks, but can be explained by fluctuations in aggregate wage pressure.

Consider first the absolute index AD_u. Suppose that there are two skill groups with constant labour force shares.[21] In this case, we can write,

$$\mathrm{d}AD_u = 2[l_s(U_s - U)(\mathrm{d}U_s - \mathrm{d}U) + l_u(U_u - U)(\mathrm{d}U_u - \mathrm{d}U)] \qquad \textbf{(19)}$$

Using $\mathrm{d}U=\mathrm{d}U_s l_s+\mathrm{d}U_u l_u$, this simplifies to:

$$\mathrm{d}AD_u = 2l_s l_u(\mathrm{d}U_s - \mathrm{d}U_u)(U_s - U_u) \qquad \textbf{(20)}$$

Equation **(20)** shows that AD_u is *increasing* in the Manacorda and Petrongolo imbalance index as both bracketed terms are negative. If SM rises then $\mathrm{d}U_s$-$\mathrm{d}U_u$ is negative because the skilled unemployment rate falls while the unskilled rate rises (see equations **(12)** and **(13)**). U_s-U_u is negative because skilled workers have lower unemployment rates.[22]

The impact of a rise in aggregate wage pressure (AWP) on AD_u is indeterminate because both skilled and unskilled unemployment rates rise. Thus the sign of $\mathrm{d}U_s$-$\mathrm{d}U_u$ is ambiguous. From equations **(12)** and **(13)** we can write:

$$\mathrm{d}U_s - \mathrm{d}U_u = T\left(\frac{U_u + g(1-U_u)}{U_u(1-U_u)} - \frac{U_s + g(1-U_s)}{U_s(1-U_s)}\right)\mathrm{d}AWP \qquad \textbf{(21)}$$

Since T>0, the sign of this expression depends on the sign of the term in brackets. We calculate this term using 1975-92 GHS data for U_s and U_u presented by Jackman *et al* (1999). U_s is the unemployment rate of those with at least A-level attainment and above, while U_u is the unemployment rate of those with below A-level qualifications.[23] Manacorda and Petrongolo (1999) argue that the true value of the wage-flexibility parameter, g, lies between 0.035 and 0.1. Thus we use both values in the calculations. The results show that a rise in aggregate wage pressure has a negative effect on $\mathrm{d}U_s$-$\mathrm{d}U_u$ in every year over 1975-92. From equation **(20)**, it follows that the absolute measure of dispersion is *increasing* in aggregate wage shocks. Recall that such shocks also lead to a rise in aggregate unemployment (equation **(15)**). Consequently aggregate wage shocks lead to positive co-movements of AD_u and U which is what we observe in the data (see Chart 7 and Table H).[24]

---

(20) Recall that Charts 3 and 4 suggest that the impact of aggregate shocks on SM is negligible.

(21) The assumption of constant labour force shares means that movements in the imbalance index, SM, reflect relative demand shocks. This does not affect the results.

(22) Skilled workers have lower unemployment rates than unskilled workers in all the major industrialised countries except Italy.

(23) This data is set out in Appendix 2 of Jackman *et al* (1999).

(24) Jackman *et al* (1999) also provide similar data for Australia, Canada, Italy, France, West Germany, the Netherlands, Norway, the United States, and Spain. Aggregate wage pressure shocks have a consistently negative impact on $\mathrm{d}U_s$-$\mathrm{d}U_u$ for each country, except Spain.

Now we can consider the relative index. In this case,

$$\mathrm{d}RD_u = 2l_s\left(\frac{U_s}{U}-1\right)\mathrm{d}\left(\frac{U_s}{U}\right)+2l_u\left(\frac{U_u}{U}-1\right)\mathrm{d}\left(\frac{U_u}{U}\right) \tag{22}$$

Using the fact that d$(U_u/U)=-(l_s/l_u)$d(U_s/U), this simplifies to:

$$\mathrm{d}RD_u = 2l_s\left(\frac{U_s-U_u}{U}\right)\mathrm{d}\left(\frac{U_s}{U}\right) \tag{23}$$

This can be simplified further by noting that d$(U_s/U)=(U\mathrm{d}U_s-U_s\mathrm{d}U)/U^2$. Thus,

$$\mathrm{d}RD_u = 2l_s(U\mathrm{d}U_s-U_s\mathrm{d}U)\left(\frac{U_s-U_u}{U^3}\right) \tag{24}$$

Equation **(24)** shows that RD_u is *increasing* in the skills mismatch index. This is because when *SM* rises the skilled unemployment rate declines (dU_s<0) while the aggregate rate rises (dU>0). Thus the first bracketed term is negative. The second bracketed term is negative since $U_s<U_u$. Again a rise in aggregate wage pressure, *AWP*, has an ambiguous effect. From equations **(12)** and **(15)**, the effect of an increase in *AWP* on the relative index depends on the following term:

$$U\mathrm{d}U_s-U_s\mathrm{d}U = T\left[U\left(\frac{U_u+g(1-U_u)}{U_u(1-U_u)}\right)-U_s\left(l_s\frac{U_u+g(1-U_u)}{U_u(1-U_u)}+l_u\frac{U_s+g(1-U_s)}{U_s(1-U_s)}\right)\right]\mathrm{d}AWP \tag{25}$$

Since T>0, the sign of this expression depends on the term in square brackets. Once again it can be computed using data from Jackman *et al* (1999). This time in addition to group-specific unemployment rates we also need the aggregate unemployment rate, U, and the group-specific labour force shares, l_i. Jackman *et al* (1999) provide U. They also give GHS estimates of the group-specific labour force ratio, L_s/L_u, which we use to derive l_s and l_u.[25] The calculations reveal that $U\mathrm{d}U_s-U_s\mathrm{d}U>0$. Therefore RD_u is *decreasing* in aggregate wage shocks. Since such shocks increase aggregate unemployment, it follows that RD_u will co-vary negatively with U, which accords with Chart 8 and Table H.[26]

Table I summarises our findings on the relationships between the Manacorda and Petrongolo (1999) measure of skills imbalance, aggregate wage pressure, and unemployment dispersion. A rise in educational imbalances will lead to an increase in both dispersion indices. However this is not what we observe in the data where, according to the Manacorda and Petrongolo measure, imbalances have increased steadily over the past two decades (see Charts 3 and 4), while unemployment dispersion has been cyclical.[27] The fact that both AD_u and RD_u co-vary with

[25] Notice that $L_s/L_u=l_s/l_u$. Since $l_s+l_u=1$, it follows that $l_u=1/(l_s/l_u+1)$.
[26] $U\mathrm{d}U_s-U_s\mathrm{d}U>0$ for the other OECD countries.
[27] The absolute and relative measures of employment dispersion across skill groups are also cyclical.

aggregate unemployment in a manner that is consistent with the calculated impacts of aggregate wage pressure shocks supports the idea that both dispersion indicators are primarily driven by aggregate impulses. It follows that both AD_u and RD_u are likely to be unreliable. The cyclical patterns of both measures are incompatible with reallocation shocks, but can be explained by a skill-neutral rise in real wages above productivity growth.

Table I: Impacts of skill imbalance and wage pressure on unemployment dispersion

	Absolute dispersion	Relative dispersion
Skill imbalance	*Positive*	*Positive*
Wage pressure	*Positive*	*Negative*

This analysis can also shed some light on the co-movements of skill-group unemployment rates, U_i, and aggregate unemployment, U, following a change in the imbalance index, SM, or aggregate wage pressure, AWP. Equations **(12)**, **(13)** and **(15)** show that a change in SM and AWP leads to co-movements of U_i, and U, are neither linear nor proportional. A rise in SM leads to a decline in the skilled unemployment rate and a rise in the unskilled rate. Although a rise in AWP leads to rises in all unemployment rates, notice that the coefficients on dAWP in equations **(12)**, **(13)** and **(15)** are non-linear functions of the skilled and unskilled unemployment rates. It follows that both the absolute and relative unemployment dispersion measures impose false restrictions upon the co-movements of U_i and U.

4 Evidence from the CBI ratio of labour shortages

4.1 The CBI index

Some studies have used the ratio of the CBI measures of skilled and unskilled labour shortages to capture movements in the skill balance (see Nickell and Bell (1995) and Wadhwani (2000b)). This ratio is based on the quarterly CBI *Industrial Trends Survey* of manufacturing companies. One of the survey questions asks 'What factors are likely to limit your output over the next four months?'. Respondents can choose from a list of potential culprits, including a shortage of skilled labour, and a shortage of 'other' labour (usually interpreted as unskilled labour).[28] The CBI indices of skilled and unskilled worker shortages are the percentage of respondents answering yes to either category. The main features of the survey are shown in the upper panel of Table J.

[28] The factors are: orders or sales, skilled labour, other labour, plant capacity, credit or finance, materials and components, and other. The choices are not mutually exclusive and there are no restrictions on the number of factors that may be considered.

Table J: Main surveys of labour availability

Survey	Frequency	Start	Sample	Issue	Index
CBI	Quarterly	1972	3,000 manufacturing employers. 30% average response rate.	Skilled and 'other' labour likely to limit output over the next four months?	% yes
BCC	Quarterly	1989	8,000 employers in manufacturing and services covering around 800,000 workers. 30% average response rate.	Experienced recruitment difficulties in past quarter? (Broken down by sector).	% yes
REC	Monthly	1997	11,000 recruitment and employment agencies. 5% average response rate.	Availability of staff better, same or worse than one month ago? (Broken down by occupation).	% balance
DfEE	Annual	1990	4,000 establishments in manufacturing and services with over 25 employees. 75% response rate. The survey covers 1990-98.	Have currently, or had over the past twelve months any vacancies that are proving or have proved to be hard to fill?	% yes

Note: The CBI survey actually started in 1960 but quarterly information is available only since 1972.

Reported shortages of labour can be interpreted as measuring the share of firms who demand more labour than is supplied at the wage they are willing to pay. This view underlies the argument of Nickell and Bell (1995) that the CBI skilled/unskilled ratio captures the 'relative excess demand for skilled labour'. Consequently a rise in this value implies that the percentage of firms facing an excess demand for skilled labour has risen relative to its level in the unskilled labour market. An alternative perspective assumes that wages adjust to clear the skilled and unskilled labour markets. However, matching frictions such as the spatial immobility of workers mean that vacancies take time to be filled. Firms report a skill shortage when the duration of a vacancy is higher than normal, or when they have to search more intensively for the appropriate worker (see Haskel and Martin (1993a)). In this case a rise in the ratio is synonymous with an increase in the difficulty firms face in filling a skilled job slot compared with an unskilled vacancy.

Chart 11 shows that the CBI ratio has trended downwards over the past two decades. It is also very cyclical, co-varying positively with the aggregate LFS unemployment rate. According to the index, skill imbalances eased by an average of 2.5% per year over 1979-99 (see Table K). This improvement has accelerated since 1993 in line with the sharp fall in unemployment. These patterns differ markedly from the Manacorda and Petrongolo (1999) and wage dispersion measures, which indicate that imbalances have increased over both periods.

Chart 11: CBI skilled/unskilled ratio

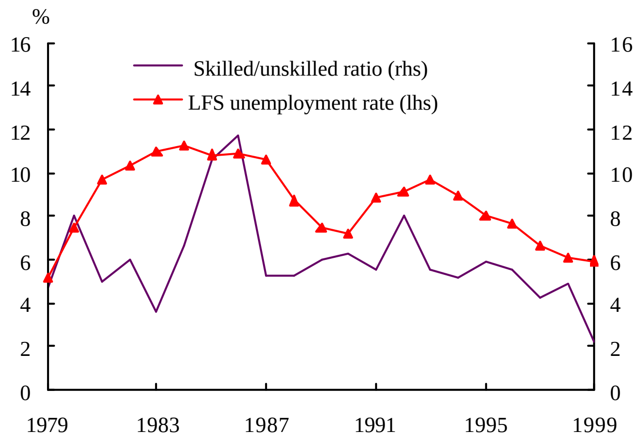


Table K: Average annual percentage change in the CBI ratio

1979-99	**1993-99**
-2.5	-8.5

4.2 Interpreting the behaviour of the CBI index

One possible explanation for this contrasting behaviour of the CBI ratio is that it is fundamentally different from the other measures we have looked at so far which use educational attainment to capture skills. The CBI survey leaves the definitions of 'skilled' and 'unskilled' workers to the discretion of the respondent.(29) So its evolution may reflect changes in the balance of a different set of workforce characteristics.(30) The survey is also restricted to the manufacturing sector, which accounts for only 20% of UK output and 15% of employment. It is possible for an improvement in the skill balance in manufacturing to co-exist with a deteriorating picture elsewhere.(31) Since the CBI survey is not based upon a panel of manufacturing companies, changes in the composition of the sample over time may also be important.

Although we cannot quantify the influence of relative and aggregate shocks on the CBI measure, several pieces of evidence suggest that its movements reflect skill-neutral, rather than skill-specific, developments.

(29) This lack of clear definitions in the CBI questionnaire threatens the consistency of the responses across firms and over time. Moreover, skill deficiencies that are not expected to limit output, but may affect other aspects of company performance, such as product quality, might not be captured.
(30) CBI respondents could be interpreting the question primarily in terms of the relative supply of educated workers, which has improved over the past two decades. However, one problem with this explanation is that this improvement has been continuous over the period, rather than following the cyclical pattern of the CBI ratio.
(31) Such trends might be aggravated by the fact that manufacturing share of output has fallen by 7 percentage points, while its share of employment has fallen by 11 points over the past two decades. Robinson (1996) speculates that the CBI figures are a reasonable proxy for trends in skill shortages across the whole economy. This is based on the observation that other employer surveys that cover recruitment frictions in both the manufacturing and service sectors, such as that conducted by the British Chambers of Commerce, indicate that frictions in both sectors are positively correlated. Our own preliminary analysis suggests that the rise in educational imbalances also took place in manufacturing. CBI survey data may be unrepresentative of the manufacturing sector because it is skewed towards larger companies.

Chart 12 shows that both skilled and unskilled shortages co-vary positively. This pattern challenges the notion that movements in the CBI data capture relative demand and supply shifts, because such impulses should induce a negative relationship between both measures.(32) For example, a redistribution of demand towards skilled workers should lead to a rise in shortages of skilled labour, while unskilled shortages decline. However, these co-movements are the plausible outcome of absolute labour demand/supply fluctuations, which will affect shortages of all types of worker in a similar fashion.

Chart 12: CBI labour shortages

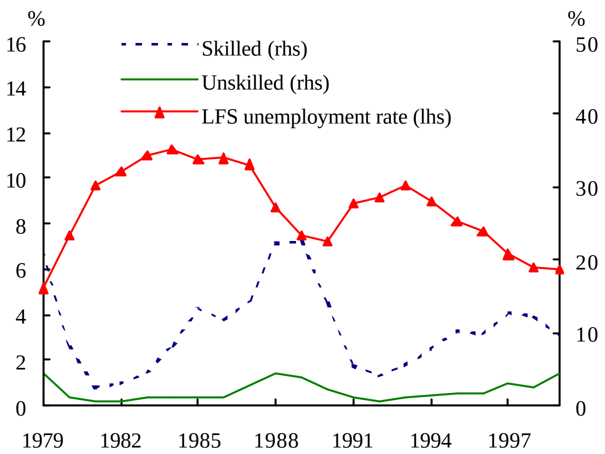


Chart 12 also shows that both indices are inversely related to the aggregate unemployment rate.(33) This is what we might expect if they are driven by changes in the general availability of all types of worker. To explore this further, we compared the CBI data with the results of other employer surveys that attempt to measure the incidence of general recruitment difficulties. These surveys are:

- The British Chambers of Commerce (BCC) quarterly survey of recruitment difficulties;
- The Recruitment and Employment Confederation (REC) monthly survey of staff availability; and
- The Department for Education and Employment Skill Needs (DfEE) annual survey of hard-to-fill vacancies.

Rows 2-4 of Table J outline the main features of these other surveys. Both the BCC and REC indices are backward-looking measures of recruitment frictions. The BCC index begins with all respondents who attempted to recruit staff over the previous quarter and reports the percentage that experienced difficulties. The REC measure covers recruitment and employment agency reports on the availability of staff compared with the previous month. The DfEE survey reports the percentage of respondents who currently have, or have had in the past twelve months, any vacancies that are proving, or have proved, hard to fill.(34)

(32) The correlation coefficient between the skilled and unskilled indices is 0.88.
(33) Simple regressions reveal that these negative relationships are significant.
(34) This survey ran between 1990-98. In 1999 it was replaced by the Employer Skills Survey (ESS), which covered 27,000 establishments in England with at least five employees. Hudson (2000) outlines the main findings of the ESS.

Charts 13 and 14 plot the BCC and DfEE measures respectively over the past decade (the REC only began in 1997). Each chart also shows the corresponding ILO unemployment rate. Simple regressions reveal that, like the CBI measures, these also co-vary negatively and significantly with unemployment. Of course, the fact that the CBI figures also exhibit this inverse pattern does not constitute proof that the CBI shortages capture general rather than skill-specific changes in the availability of labour. Nonetheless, it does suggest that interpreting the CBI data solely in terms of relative shocks could be erroneous.

Chart 13: BCC recruitment difficulties

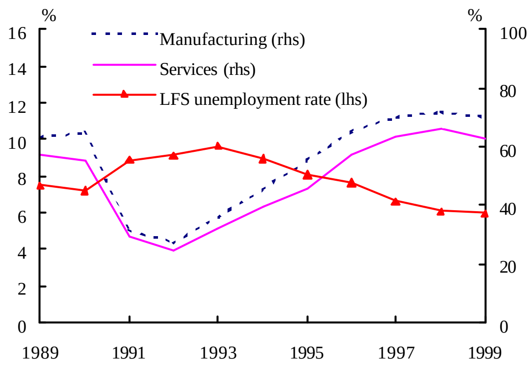
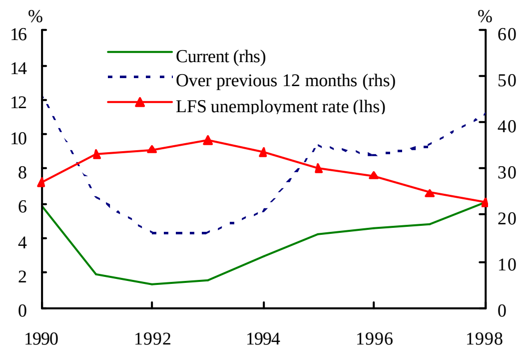


Chart 14: DfEE hard-to-fill vacancies



Even if the numerator and denominator of the CBI ratio reflect aggregate shocks, the ratio will still be a robust measure of imbalances if such disturbances shift skilled and unskilled shortages by the same factor. However, the positive correlation between the CBI ratio and aggregate unemployment (Chart 11) arises because unskilled shortages are proportionately more variable over the unemployment cycle than skilled shortages. The same factors that could plausibly underlie the transmission of aggregate shocks to the Manacorda and Petrongolo index could be at work here. For example, if firms hoard skilled employees during downturns, then the unemployment pool will contain disproportionately more unskilled labour. Consequently, shortages of unskilled workers can be expected to ease compared with those for skilled workers, leading to a rise in the skilled/unskilled shortage ratio. Of course, this begs the question as to why the Manacorda and Petrongolo measure appears to be comparatively immune to such factors. One possibility is that reports of labour shortages are far more sensitive to frictions in worker availability than relative wage or employment rates, especially in the short run.

The upshot is that the easing of skill imbalances over the past 20 years, suggested by the CBI ratio, may well be legitimate. The CBI measure is likely to cover a wide set of attributes such as reliability and interpersonal skills, which will be missed by education-based indicators. It is also possible that the easing has been confined to manufacturing. Nonetheless, there is also some evidence that movements of the CBI ratio are influenced by aggregate shocks. In particular, the decline may simply reflect cyclical differences in the hiring and firing rates of skilled and unskilled employees.

5 Conclusions

This paper has examined several measures of the movements in the balance between the demand for, and supply of, workforce skills in the United Kingdom over the past two decades. Our analysis casts doubt on the usefulness of the absolute and relative dispersion of unemployment across educational groups, and the CBI ratio of skilled and unskilled labour shortages. These indicators appear to be particularly susceptible to skill-neutral shocks, which have no implications for the skill balance. We also find that educational imbalances have increased steadily over 1979-99, particularly in the market for graduates. The idea that the apparent decline in the NAIRU over the 1993-99 upswing can be attributed to favourable movements in the educational balance finds no support.

Appendix: The education classifications

From the LFS data we allocated individuals into one of four skills based upon Labour Force Survey information on their highest formal qualification. These groups are:

Degree or equivalent: Undergraduate or higher degree, nursing or other medical qualification, high vocational qualifications (NVQ levels 4-5, HNC, HND, BTEC higher, Royal Society of Arts higher diploma, and other higher education).

A-Level or equivalent: A-level, Scottish 6th year Certificate, AS Level, SCE highers, mid-vocational qualifications (NVQ level 3, GNVQ advanced, RSA advanced diploma, ONC, OND, BTEC, and SCOTVEC national).

O-Level or equivalent: O-level, GCSE grade A-C and low vocational (NVQ level 2, GNVQ intermediate, RSA diploma, City & Guilds advanced & craft, BTEC/SCOTVEC general diploma, and completed apprenticeship).

Below O-Level: CSE below grade 1, GCSE below grade C, NVQ level 1, GNVQ/GSVQ foundation level, BTEC/SCOTVEC general certificate, SCOTVEC modules, RSA other qualification (including stage I-III), City & Guilds other, Youth Training certificate, other vocational qualifications, and no qualifications.

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