Labour market institutions and skill premiums: an empirical analysis on the UK 1972-2002

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Abstract. This paper analyzes the links between labour market institutions and skill premiums in the UK, controlling for other explanatory variables such as market conditions, international trade and skill-biased technology. We find that the trade union decline in unskilled workers can explain more than half of degree premium' increase over the period 1979-1998 in the private sector, while the overall effect of trade union on degree premiums is only one third during the same period. Decline of trade union has less significant effect on skill premiums in the public sector.

Keywords: skill premiums, trade union, panel data

JEL codes: J31, J51, K31

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Wage premiums of highly educated workers have substantially increased since the end of 1970s in the United Kingdom. Rising wage differentials between education groups have been identified as key feature of rising wage inequality in the UK and other OECD countries (Acemoglu and Autor, 2010, Gosling et al., 2000, Lindley and Machin, 2011, Atkinson, 2007). The existent literature has tried to isolate the causal factors underpinning these market changes. The most popular candidates may be the skill-biased technical change (SBTC) and increased international trade.

First of all, there is strong evidence of the empirical association between proxies for SBTC (computers or other Information and Communication Technology (ICT) facilities) and the widened wage gap of the UK and US in the 1980s (Katz and Autor., 1999, Krueger, 1993, Machin, 2001, Machin and Van Reenen, 1998, O'Mahony et al., 2008, Perugini and Pompei, 2009). Moreover, the trade explanation focuses on changes in product demand largely associated with large trade deficits in the 1980s. Wood (1994, 1995, 1998) argues that the growth of manufacturing imports from newly industrializing economies have led to a sharp decline in unskilled manufacturing employment and a shift in employment toward other skill-intensive sectors. However, the trade explanation is not convincing for many authors (Krugman and Lawrence, 1993, Machin and Van Reenen, 1998, Sachs and Shatz, 1994, Schmitt, 1995) who point that the effect of international trade on relative demand for skill is small. Ghose (2000) confirms that the growth of trade in manufacturing with some developing countries has certainly had adverse effects on employment and wage of low-skilled workers in the industrialized countries, but such effects have been quite small. Hence, on the whole the evidence seems to lean towards the SBTC explanation (Machin, 1996).

At the same time, the widening wage gap in the UK has been accompanied by institutional reform in the labour market since Thatcher-era. Labour policy directed by US-style flexibility may be part of the causation of the widening wage structure. This paper aims to analyze the effects of changes in labour market institutions (such as trade unions, taxation, unemployment benefits and the national minimum wages) on the skill premiums, controlling for changes in technology and trade patterns.

With the same access to technology and international competition, and having had a similar education expansion, the increasing skill premiums in the UK, in contrast to the stable wage structure in continental European countries can only be explained by a different institutional environment. Hence, Acemoglu (2003) argues that changes in the supply and demand for skills are unlikely to fully account for the marked differences in skill premiums across countries. The "Krugman hypothesis" states that the rise in wage inequality in the Anglo-Saxon countries as well as the rise in unemployment in continental Europe are "two sides of the same coin", namely a fall in the relative demand for unskilled workers under different wage setting institutions (Krugman, 1994, Nickell and Bell, 1996, Puhani, 2008).

A substantial amount of research on wage inequality has regarded and examined labour market institutions as important factors that may affect the wage response of markets to shifts in the relative demand for skills (Blau and Kahn, 1996, Card et al., 2003, Gottschalk and Joyce, 1998, Katz et al., 1995, Koeniger et al., 2007, Machin, 1996, Machin, 1997). One strand of this research has studied how specific labour market institutions affect wage differentials in the UK. First of all, the possibility of there being a connection between the wage differentials and trade unions has been studied in a large literature. Casual inspection shows a striking association between movements in union density over time and changes in the earnings dispersion. Schmitt (1995) has calculated that the decline in union density could account for 21 percent of the rise in the pay premium for a university degree and for 13 percent of the increase in the non-manual differential during 1978-1988. Machin (1997) obtains more dramatic results that the male variance would have been 40 percent less if the 1980s levels of union coverage had prevailed in 1991. Bell and Pitt (1998) also conclude the deunionization between the early 1980s and 1990s widened the male earnings distribution by about 20 percent. Card *et al.* (2004) for a comparison of the United States, the United Kingdom, and Canada, and Kahn (2000) for OECD countries have also found that higher union density is associated with lower wage inequality. This paper pushes the discussion further and focuses on the union effect on the skill premiums over the last three decades in the UK.

Moreover, Dickens *et al.* (1999) and papers in the special session on the British minimum wages in the Economic Journal 2004 (Dickens and Manning, 2004, Machin and Wilson, 2004, Stewart, 2004, Metcalf, 2004) have found that national minimum wages reduce wage inequality by increasing the bottom deciles of the pay distribution without a negative impact on employment. DiNardo *et al.* (1996) and Lee (1999) also find the same effect of minimum wages for the United States. For other labour market institutions, tax wedges and unemployment benefits may affect skilled and unskilled workers at different degrees and change the skill premiums. Brewer *et al.* (2008) study about five million income tax returns covering the period 1996-2005 and find that even though the current government has increased taxes on people with high incomes, this has not prevented them from racing further away from the average level of living standards across the country. They think that the outlook for inequality in Britain may depend more on the outlook for the stock market than on Government tax and benefit policies. Thus, this paper also investigates the different effects of the minimum wages, tax wedge and benefits on skill premiums.

Most institution-specific research studies cross-section/longitudinal data at country level (Blau and Kahn, 1996, Koeniger et al., 2007, Wallerstein, 1999). However, the cross-country comparison cannot test for differences within a country, which is our contribution. No previous empirical study has tried to quantitatively assess respective importance of labour market institutions for workers with different education attainments. We construct a balanced panel data of six skill (education) groups over the period of 1972-2002 from several micro datasets, through which we investigate the effect of institutional factors on distinct skill groups and then skill premiums in the UK. This article is divided into four parts. The first reviews the theoretical models that motivate the estimated log-linear equation and provides our empirical specifications. The second part describes the main data sources and measures those variables, and the third represents empirical results. The fourth part concludes with a summary of the main findings and suggestions for policy.

Empirical specifications

In this part, we briefly review a union bargaining model provided by Koeniger *et al.* (2004, 2007), in which labour market institutions alter the outside options of skilled and unskilled workers differently and thus affect relative labour demand as well as the wage differentials. [1] Changes of institutions as well as market conditions, technologies and international competition are reflected in the following skill premium equation:

$$\ln\left(\overline{w_H}/\overline{w_L}\right) \cong \underset{+\ \ -}{f(tud_H, tud_L)} + \underset{+\ \ -}{g(tax_H, tax_L)} + \underset{+\ \ -}{h(repr_H, repr_L)} + \underset{-}{j(MW)} + \underset{-\ \ +}{k(u_H, u_L)}$$
$$+ \underset{+\ \ -\ \ +\ \ +}{l(comp_H, comp_L, ind_H, ind_L)} \qquad (1)$$

where H denotes high skilled workers while L is low skilled workers. The skill premiums, i.e. log form gross wage differentials for skilled workers, $\ln\left(\overline{w_H}/\overline{w_L}\right)$, mainly depends on trade union density (*tud*), the tax wedge (*tax*), benefit replacement ratios (*repr*), unemployment rates (*u*), technology (*comp*) and international trade (*ind*) by skills, with the addition of the minimum wage variable (*MW*). Skill premiums depend on human capital and forgone earnings, and should be remarkably constant over the long run. However, short- and medium- run factors, including variables of institutions, market conditions, technology and international competition in equation (1) also affect skill premiums. Now, we go through the variables in the order they appear in the equation and present our arguments underlying equation (1) as follows.

Koeniger *et al.* (2004) make union bargaining central to their derivation of equation (1), but many of their arguments hold in a competitive market as well, as we will explain. First, the skill premiums will be smaller if unions favour unskilled workers (tud_L) more than skilled workers (tud_H). And, the trade union bargaining model in Koeniger *et al.* (2004) regards earnings tax as a mark-up part of the gross wages for both skilled and unskilled workers. This result also holds in a competitive market model with individual bargaining. A similar analysis can be applied for unemployment benefit (*repr*) and unemployment rates (*u*) in equation (1): higher replacement ratios for skilled workers ($repr_H$) increase the skill premiums, while higher replacement ratios for unskilled workers ($repr_L$) decrease it. And, higher unemployment rates for skilled workers (u_H) are likely to decrease the skill premiums, while higher unemployment rates for unskilled workers (u_L) are likely to increase it. The overall effect of unemployment benefits (or unemployment rates) on the skill premiums depends on a comparison between its respective wage effect on skilled and unskilled workers.

Second, as DiNardo *et al.* (1996) reveals, a minimum wage can directly compress the skill premiums by binding wages of unskilled workers, whereas wages of skilled workers are not directly affected. Hence, the minimum wages will cut off all unskilled wages below it and make the skill premiums smaller.

Last but not least, skill premiums are affected by medium and short run shocks from technology (such as computer usage, *comp*) and international competition (such as industrial shifts, *ind*) in the market. New technologies adopted by skilled workers ($comp_H$) increase their marginal products and push up the skill premiums temporarily, while new technologies adopted by unskilled workers ($comp_L$) also increase their marginal products but decrease the skill premiums. However, if new technologies are complementary to skills (Acemoglu, 1998), total factor productivity of skill-intensive sectors (for example, computer software industry) grows faster than labour-intensive sectors (for example, textile industry). Technology shifts may have higher wage

[1] See Koeniger *et al.* (2004) for derivation details of this model. We omit employment protection legislation (EPL) in their model, since the EPL index has been stable in the UK for the last thirty years (Daniel and Siebert (2005), Figure 4 and 5).

effects on skilled workers than on unskilled workers and would push up the skill premiums.

International competition from newly industrialised countries may decrease the price of labour-intensive goods, as well as the demand for unskilled workers. At the same time, excess demand abroad may increase the domestic price of skill-intensive goods and increase the relative demand for skilled workers. Increasing international competition is good for skilled workers (ind_H) but bad for unskilled workers (ind_L). Thus, international trade effects on both skilled and unskilled workers are likely to increase the skill premiums.

Our empirical work uses a two-step estimation procedure, which is designed to get round the Moulton (1986) problem of explaining earnings based on individual data with variables based on aggregate data.[2] In step 1, we use all individual observations to estimate education wage differentials as proxies of skill premiums over time. This equation is given as:

$$\begin{aligned}\ln w_{it} &= \alpha_0 + \sum_{t=1}^{T} n_t Y_t + \sum_{t=1}^{T} b_t B_{it} Y_t + \sum_{t=1}^{T} o_t O_{it} Y_t + \sum_{t\ t=1}^{T} a_t A_{it} Y_t + \sum_{t=1}^{T} h_t H_{it} Y_t + \sum_{t=1}^{T} d_t D_{it} Y_t + \beta X_{it} + \varepsilon_{it} \\ &= \alpha_0 + \sum_{t=1}^{T} n_t Y_t + \sum_{j=b}^{d} \sum_{t=1}^{T} s_{jt} S_j Y_t + \beta X_{it} + \varepsilon_{it}\end{aligned} \tag{2}$$

where w_{it} is the real gross hourly wage rate. Following the tradition of Blau and Kahn (1996), Dickens (2000), Gosling *et al.* (2000), Koeniger *et al.* (2004, 2007) and Schmitt (1995), we only concentrate on male full time workers (weekly working hours>=35). Y_t denotes a year dummy representing the baseline group of workers without education qualification (*NOQUAL*); B_{it} denotes a dummy variable for workers with below O-level qualifications (*BOLEV*); O_{it} denotes a dummy variable for the O-level group (*OLEV*); A_{it} denotes a dummy variable for workers with A-levels (*ALEV*); H_{it} denotes a dummy variable for workers with higher educational qualifications but not degrees (*HIGHER*); and D_{it} denotes a dummy variable for worker with degree equivalent or above qualifications (*DEGREE*). X_{it} is a vector of controlling variables that may influence wages including potential labour market experience (Katz and Murphy, 1992), marital status, ethnicity, tenure and region; and ε_{it} is a random error term.

Correspondingly, n_t are the estimated coefficients of the NOQUAL group, which are the wages of this group in year t relative to their wages in the first sample year, 1972. Following the same method, b_t, o_t, a_t, h_t and d_t are the estimated incremental wage effects of the different education groups: *BOLEV, OLEV, ALEV, HIGHER* and *DEGREE* over the baseline group *NOQUAL* in the same year. We stack b_t, o_t, a_t, h_t and d_t in equation (2) to form a skill premiums variable s_{jt}, which is the estimated skill premiums of education group j (j=b,…,d) relative to the baseline

[2] Moulton (1986) shows that individuals in the same year/area share some common component of variance that is not entirely attributable either to their measured characteristics (e.g., gender and age) or to any aggregate variable in the year/area. In this case, the error component in an OLS regression will be positively correlated across people in the same year/area, causing the estimated standard error of the aggregated variable to be downward biased. A similar two-stage procedure is used in the wage cyclicality (beginning with Solon et al. (1997)) and wage curve literature (Nijkamp and Poot (2005), p 434).

NOQUAL group. Hence, a panel dataset is built to find the links between the skill premiums and labour market institutions in step 2:

$$s_{jt} = \theta_1 tud_{jt} + \theta_1^n tud_{nt} + \theta_2 tax_{jt} + \theta_2^n tax_{nt} + \theta_3 repr_{jt} + \theta_3^n repr_{nt} + \theta_4 MW + \theta_5 u_{jt} + \theta_5^n u_{nt}$$
$$+ \theta_6 ind_{jt} + \theta_6^n ind_{nt} + \theta_7 comp_{jt} + \theta_7^n comp_{nt} + \upsilon_j + \upsilon_t + \upsilon_{jt}$$

(j= b, o, a, h and d) (3)

where s_{jt} are the estimated skill premiums for education group j in the year t, and labour market institutions indicators and those control variables of market conditions, technology, and international competition are defined in equation (1). All variables of the baseline group (tud_{nt}, tax_{nt} $repr_{nt}$, u_{nt}, ind_{nt} and $comp_{nt}$) are also put into equation (3) to control for changes in the baseline group. υ_j is a vector of education group dummies, υ_t are year dummies, and υ_{jt} is the stochastic error term.

Equation (3) assumes the existence of a long run equilibrium relation between skill premiums and institutions. Also, the adjustment should be contemporaneous. However, much literature shows an increasing trend in the skill premiums (for example, Gosling *et al.* (2000)) as well as a decline of trade unions since the 1970s (for example, Bell and Pitt (1998) and Disney *et al.* (1998)). Since our panel data cover the period of 1972-2002, the skill premiums of each group are probably non-stationary (see ADF tests below). Hence a co-integration problem may exist in the links between skill premiums and institutional variables. If there is some inertia in the adjustment process a re-parameterisation of equation (3) - as in equation (4) below - might be preferable. Thus, we put an Error Correction Mechanism (ECM) into equation (3) to clear the long-term relationship between the *level* of skill premiums and *level* of institutions. We follow the approach used in Ammermueller et al. (2010), but only put the ECM in trade union density variables (tud_{jt} and tud_{nt}) since trade unions are regarded as the most important institutional factor in most of the literature and only union density variable shows non-stationarity over the last thirty years. The error-correction specification is:

$$\Delta s_{jt} = \theta_0 tud_{jt-1} + \theta_1 \Delta tud_{jt} + \theta_0^n tud_{nt-1} + \theta_1^n \Delta tud_{nt} - as_{jt-1}$$
$$+ \theta_2 tax_{jt} + \theta_2^n tax_{nt} + \theta_3 repr_{jt} + \theta_3^n repr_{nt} + \theta_4 MW$$
$$+ \theta_5 u_{jt} + \theta_5^n u_{nt} + \theta_6 ind_{jt} + \theta_6^n ind_{nt} + \theta_7 comp_{jt} + \theta_7^n comp_{nt} + \upsilon_j + \upsilon_t + \upsilon_{jt}$$

(j= b, o, a, h and d) (4)

Thus, in this specification, the long run equilibrium between the *level* of the skill premiums and *level* of trade union density is embodied in an ECM.

Data description

Skill premiums

The principle data used in this paper comes from the series of the annual General Household Survey (GHS) from 1972 to 2002. The GHS is a continuous multipurpose survey of large random samples of households across Great Britain, conducted on an annual basis by the Office for National Statistics (ONS). The survey has been carried out continuously except for two breaks in 1997 when the survey was reviewed and

1999 when the survey was redeveloped. Hence, there are 29 years of data in this paper (T=29) over the period 1972-2002.

We use the highest educational qualification earned by the respondent to measure the skill levels which are either vocational or academic (see more details in Schmitt (1995)). The wage variable used here is the real gross hourly earnings, deflated by the annual Retail Price Index (RPI) with 1995 as 100 which is also provided by the ONS. The wage variable is from a wage sample including all full-time employees aged sixteen to sixty-six. "Full time employee" here is defined as workers (excluding employer and self-employed) with weekly working hours exceeding 35 hours. Self-employed workers, part time workers and those working without pay are excluded from the sample. Extreme cases of earnings are excluded. Therefore, in the first step regression based on equation (2), we have 138,103 observations over the entire period, in which 114,491 workers are in the private sector while 23,323 workers are in the public sector. In the second step, we have 5×29=145 observations in both sectors.

Table 1 presents descriptive statistics of the variables used in the second step regressions by skill level and private/public sector. For the private sector, skill premiums are higher in the groups with more education attainment as expected. The variation of degree premium (7.08 percent) is much higher than that of other groups, reflecting the dramatic increase of about 27 percent from about 56 percent in 1979 to about 83 percent in 1998 (also see panel A of Table 4). Skill premiums of the HIGHER group increase by about 13 percent over the same period of 1979-1998, which also have higher variation (4.59 percent) than medium skilled groups (ALEV, OLEV or BOLEV, around 2.9 percent) but to a less degree. These results are consistent with the findings in Walker and Zhu (2003, 2008) using the Labour Force Surveys (LFS). However, medium skilled groups in the public sector have much higher and more volatile skill premiums than those in the private sector, while degree premium is a little lower in the public sector. Actually, there is no increasing trend for skill premiums in the public sector. Thus, the worsening of wage inequality since the 1970s is perhaps caused by the increasing skill premiums in the private sector, especially from the rapidly increasing degree premiums rather than changes of skill premiums in the public sector.

Next, we use the GHS 1972-2002 to measure those institutional and controlling variables on the right hand side (RHS) of the Step 2 regression, combined with another three datasets: the Family and Working Lives Survey (FWLS 1994/1995), the British Household Panel Survey (BHPS 1991-2002) and the UK Family Expenditure Survey (FES 1972-2002). [3] . Because our dependent variable (skill premiums) have mainly controlled potential labour market experience in the Step 1, we follow the Katz and Murphy (1992) to control the within experience composition biases of these RHS variables by using the fixed weighted average of eight experience sub-groups with skill and sector. Weights are the average

[3] The FWLS is a life and work history data, which provide retrospective and representative information about people living in Britain. The BHPS was designed as an annual survey of each adult (16+) member of a nationally representative sample of more than 5,000 households in the UK, making a total of approximately 10,000 individual interviews yearly. The FES is a continuous survey of household expenditure and income, which has been in existence since 1957. The FES was replaced by a new survey in 2001, the Expenditure and Food Survey (EFS). Thus, the last two years' data are from the EFS 2001 and 2002, in which they have the same definition. We will not differentiate the two surveys in later discussion.

employment share of each experience sub-group over the entire period of 1972-2002, also following the tradition of Katz and Murphy (1992).[4]

Institutions

The main purpose of using these additional datasets is to compile a time series on union density by skill level. Information on union membership (*tud*) since the 1970s, along with worker's characteristics is not available in any single British dataset. The GHS does not provide information about the trade union membership except in one year (1983). Our union density variable is from the FWLS for the period 1972-1994 (Disney *et al.*, 1998), and for the period 1995-2002, it is from the BHPS. Figure 1 compares the trade union density changes in different datasets and illustrates a famous decline of trade union after the 1970s. The change of union density between 1991 and 1995 is very similar in the FWLS and BHPS. This similarity shows that the union density has a consistent pattern between the two datasets. For the BHPS, the union questions were only asked for those who moved job in 1992-1994 (but for everyone in other years), so we did not include the period 1992-1994 in this figure and do not use these data in the analysis. We also compare our data of trade union density with the data from the Certification Office (The "Bain and Price" series, see Disney *et al.* (1998) and Bell and Pitt (1998)) and find that the consistency between datasets is satisfactory.

More descriptive statistics of trade union densities are presented in Table 1 by skill and sector. The combination of the FWLS and BHPS reveals the trade union density in the medium skilled groups (BOLEV, OLEV and ALEV) is higher than the unskilled (NOQUAL) and high skilled groups (HIGHER and DEGREE) in the private sector, which is reasonable. For workers in all skill groups in the private sector, trade union density tends to decline after the end of 1970s, during which the skill premiums of the DEGREE and HIGHER move in the opposite direction. However, the situation in the public sector flips, in which union density of unskilled (NOQUAL) and high skilled groups (HIGHER and DEGREE) is higher than that in medium skilled groups (BOLEV, OLEV and ALEV). And, we do not find a clear decline of unions in the public sector since the 1970s. It is perhaps the reason for the different evolution of skill premiums in these two sectors.

Moving on to the tax and benefit system, the private and public sectors are very similar to each other for these two variables, since the tax wedge and replacement ratios are actually irrelevant to sector. The little difference between two sectors is actually from within experience composition. Concerning the different tax wedge (*tax*) for skilled and unskilled workers, the GHS does not provide information about tax deductions from gross earnings. We therefore use the FES, which is a better dataset for tax expenditure. The FES 1972-2002 in fact provides tax wedges by skill level. The tax rate is defined here as the proportion of income tax deduction (Pay As You Earn amount) relative to normal gross wages. The theoretical model in Koeniger *et al.* (2004) implies that tax wedge is only a mark up factor on the gross wages. The relative tax wedge between skilled and unskilled workers should be positively correlated with the skill premiums. Table 1 shows that the tax wedge increases with skill levels reflecting the increasing tax rates according to "ability to pay" principle.

As for benefit indices, they measure the proportion of unemployment benefits relative to average earnings before tax. The GHS provides information for

[4] Technique details are available if requested.

unemployment benefits over the entire period 1972-2000. After 1996, the British unemployment benefit changed its name to job seeker allowance. We will keep using the unemployment benefit term in the discussion. For practical purpose, we also put income support and incapacity benefit into our benefit indices since both of them will increase the outside option of workers. Since the data about housing benefit (particular for council tax) are not consistent over time in the GHS, we do not include it. Moreover, a problem arises that unemployed workers can only provide the actual amount of benefit received not their earnings. Hence, the replacement ratios of benefits (*repr*) are estimated as the proportion of unemployment benefits they received relative to their estimated earnings in a standard earnings equation.

On the other hand, the Koeniger *et al.* (2004) model implies that the replacement ratios should be negatively correlated with the skill premiums, if unemployment benefit is more generous for unskilled workers. Table 1 shows that the replacement ratios decrease with skill levels as expected. Brewer *et al.* (2008) argue that government has imposed large rises in taxation to fund higher benefit payments and tax credits after 1996. Hence, the combination of tax and benefit system may have larger effects and alleviate the wage inequality in recent years.

As far as the minimum wages (*mw*) are concerned, the breakdown by skill and private/public sector is not necessary. The UK National Minimum Wage (NMW) Act came into force on 1 April 1999. We build a variable being zero before 1998, and taking the log form of national minimum wages after 1998 as a proxy for this policy change (Metcalf (2004), Table 1).

International trade, technology and market conditions

The unemployment rates of skill groups play an important role for the skill premiums because they represent the market conditions and outside options of workers with different education attainment. We calculate the unemployment rates only by skill level over the entire period using the GHS 1972-2002, similarly being due to its irrelevance to sector. The little difference between two sectors is also from within experience composition. Actually, we compare the unemployment rates in the GHS with other data sources such as the LFS and the BHPS, and find no much difference in these three data sources. Hence, we use the GHS here for consistency. The theoretical model of Koeniger *et al.* (2004) implies that there is a negative (positive) relationship between the unemployment rates of skilled (unskilled) workers and the skill premiums, *ceteris paribus*. Table 1 shows the different labour market conditions of skill groups. Obviously, the lower educated workers are more vulnerable when the labour market is loose. We also find that the unemployment gap between unskilled/medium-skilled and high skilled workers became wider in the 1980s and early years of 1990s. Higher unemployment rates of unskilled workers worsen their outside option and also decrease the collective bargaining power of their trade unions. Thus, the skill premiums should increase if the unemployment rate of unskilled workers increases faster than that of skilled workers.

In the model of Koeniger et al. (2004), international trade and technology determine the skill premiums through relative prices and relative total factor productivity. In a two input × two industry world, as frequently used in the literature, the effect of international trade can be represented by the employment shifts from the manufacturing to the services as in Schmitt (1995). In this paper, the proxy of the trade effect is the employment proportion of manufacturing workers within each skill group (*ind*). For its next, international competition may compress the product price and profit of firms in traded sector, and decrease wages there. As more workers within one skill group shift from the manufacturing sector to the services, the pressure from international competition must be bigger.

Table 1 shows the employment shifts mainly happen in the low skilled groups such as NOQUAL, BOLEV, OLEV and ALEV, which have continuous declines in manufacturing employment proportions. For workers in the high skilled groups of HIGHER and DEGREE, there is not much change in the manufacturing ratio. More detailed study shows that the manufacturing employment proportions even increase in high skilled groups in the early years of the 1980s, which may contribute to the increasing skill premiums in the 1980s.

As for SBTC, we use computer usage density (*comp*) as a proxy. Computer usage is a widely applied measure of skill biased technology (Krueger, 1993). The disadvantage of this proxy is that the computer usage variable is not available before 1984 in the GHS. As an alternative, we spliced into the series data from telephone usage using the telephone/computer ratio in 1984. This series gives the approximate computer usage in years before 1984. We have no information to divide the public and private sector, so this variable is the same for two sectors. As we expect, Table 1 shows higher computer usage is associated with higher skill levels.

Empirical results

Basic results

Table 2 presents the fixed effect results from equation (3) by the private and public sector. We focus on the private sector which is the majority of the workforce. First, there are significant associations between the skill premiums and trade union density of skilled groups (tud_{jt}) in the private sector. A point increase of trade union density in the skilled group will increase the skill premiums by 0.15 percent. As the theoretical model predicts, a point increase of trade union density in the baseline unskilled group (tud_{nt}) will decrease the skill premiums by 0.21 percent, but this effect is not significant. Thus, our results suggest that trade unions have different effects on wages of workers at different skill levels.

Second, the tax wedge shows a significant mark up effect for unskilled workers as the theoretical model predicts. A one point increase of the unskilled workers' tax wedge (tax_{nt}) decreases the skill premiums by about 1.7 percent. As the theoretical model predicts, the same change in the skilled workers' tax wedge (tax_{jt}) increases the skill premiums by about 1.03 percent, but insignificant. However, the benefit index has no significant effect on the skill premiums. Neither does the minimum wage variable show significant effects on the skill premiums in Table 2.

Third, unemployment rates should reflect the business cycle and the outside options of workers. Workers can bargain more strongly if the labour market is tight. Yet, from Table 2, there is no significant effect of market conditions on the skill premiums. Solon *et al.* (1994) point out that more unskilled workers would join the employment as labour market is tight and push the overall wages down. Since employment composition within each education group also changes over the business cycle, it is not surprising to see insignificant effect of market conditions on skill premiums. Hence, the insignificant overall wage cyclicality here may just show the composition biases.

As for other variables, we find significant effects of industrial shifts in skilled workers (-0.23) on the skill premiums. Moreover, the computer usage variables show significant positive associations with workers' wages, implying new technologies can improve productivity of all workers. A one point increase in computer usage of skilled workers (*$comp_{jt}$*) is associated with a 0.29 percent increase of the skill premiums, while that of unskilled workers (*$comp_{nt}$*) decreases about 0.57 percent of the skill premiums. This result suggests that adaptation of new technology for unskilled workers can help decrease the wage inequality.

For all estimations in the public sector, there is no significant result except unemployment rate (0.69) and industrial shifts (1.45) of unskilled workers, implying more static skill premiums in the public sector. Wages of unskilled workers in the public sector can respond to the market condition and affect skill premiums, which is consistent with findings in Devereux and Hart (2006) and Peng and Siebert (2007). As the unemployment rates of unskilled workers increase, their wages decrease and push up the skill premiums. Similarly, more unskilled workers are employed in the public manufacturing increase the skill premiums. It seems that the wage setting in the public sector does not follow the model of Koeniger *et al.* (2004). Instead, bureaucratic and administered price models may be needed to explain wage management in the public sector (see a summary in Kaufman (2007) using transaction costs theory).

Results of ECM specification

The fixed effect results in Table 2 would be biased by co-integration problems if the skill premiums and trade union density were non-stationary. Augmented Dickey-Fuller (ADF) Unit root test shows that the degree premiums in the private sector are non-stationary over the entire period, even more non-stationary during the period 1979-1998. And, trade union densities of all education groups in the private sector are non-stationary over the entire period.[5] Hence, results in Table 2 may be biased by co-integration problem. Table 3 tries the fixed effect ECM model using the better specification in equation (4). This improvement in methodology clears up the long term relationship between institutions and the skill premiums. The main improvement is that institutional effects on the skill premiums are more important and significant in the private sector. A one point increase of trade union density in the skilled group (*tud_{jt}*) still increases the skill premiums by 0.18 percent. However, the effect of trade unions on skill premiums becomes bigger and significant for unskilled workers. A one point increase of trade union density in the unskilled group (*tud_{nt}*) will decrease the skill premiums by 0.59 percent. The tax wedge shows the right mark up effect as the model expects, but insignificant; the benefit variable of high skilled workers (*$repr_{jt}$*) is also insignificant as in Table 2. However, the benefit variable of the unskilled group (*$repr_{nt}$*) becomes significant. One point increase of benefit variable of unskilled workers can decrease the skill premiums by about 0.39 percent.

Moreover, as the theoretic model predicts, unemployment rates of skilled workers (*u_{jt}*) now show a negative association with the skill premiums (-0.47). Hence, the higher unemployment rates of skilled workers bring down their wages and decrease the skill premiums. This is consistent with the model of Koeniger *et al.* (2004) and the wage cyclicality literature. More skilled workers are employed in

[5] ADF test shows that the degree premium is non-stationary over the entire period (t value:-2.68, MacKinnon p value: 7.8%), especially during the period 1979-1998 (t value:-1.60, MacKinnon p value: 48.4%). Skill premiums of other groups are all stationary over the entire period. Trade union densities of all groups are non-stationary over the entire period (t value: -0.55 for NOQUAL, -0.199 for BOLEV, 0.709 for OLEV, -0.834 for ALEV, -0.324 for HIGHER and -0.624 for DEGREE). Thus, the co-integration problem is serious for the regression, especially for specific skill groups and sub-periods.

manufacturing will decrease the skill premiums (-0.24) in the private sector. The technology change also shows the right direction for both skilled (0.26) and unskilled groups (-0.44). More computer usage in the unskilled workers appears to increase their wages and decrease skill premiums.

For the public sector, we find that trade union density of unskilled workers is negatively associated with skill premiums in long run (-0.4), but positive in short run (0.64). The long term and short term of effects of trade union offset each other for unskilled workers and keep the skill premiums sticky in the public sector. Hence, skilled workers in the public sector seems not benefit much for wages from their trade unions. Acemoglu *et al.* (2001) argue that trade unions compress wages of skilled members to compensate unskilled members (in the public sector, skilled and unskilled are to a large extent in the same unions, e.g. Unison). Another interesting point worthy of mention is the ECM variable, which is the lagged skill premiums variable, s_{jt-1}. Its coefficient is 0.71 in the private sector but around 1 in the public sector, and both significant. This result confirms our argument that the short run wage adjustments in the private sector are more rapid than in the public sector. In fact, there may be no ECM in the public sector since the skill premiums there appear to be static. A compensation model as in Acemoglu *et al.* (2001), or a bureaucratic and administered price model as in Kaufman (2007) may be better for explaining wage management in the public sector.

Contributions to the degree premium

Table 4 estimates the contribution of each explanatory variable to the changes of the degree premium over three typical periods: 1972-1979 and 1979-1998 and 1998-2002. All figures in Table 4 are calculated by using coefficients in Table 3. For simplicity, we only concentrate on the institutional effects on the degree premium (as a proxy to earnings inequality) in the private sector and ignore insignificant estimates in Table 3

The top panel shows changes in degree premium and those explanatory variables for both groups: NOQUAL and DEGREE. The middle panel shows effects of each explanatory variable on the degree premium. The bottom panel is the overall contribution of explanatory variable in different period. In analysis below, we concentrate on the long period 1979-1998, during which the degree premium has increased to the highest value in 1998.

From Table 4, we can find the decline of trade union is the most important factor for the increasing degree premium during the period 1979-1998. The union decline in the DEGREE group (-30.75 percent) decreases the degree premium by about 5.53 percent (=0.18×30.75). At the same time, however, trade union density decline in the unskilled group (NOQUAL, -25.29 percent) increases the degree premium by about 14.92 percent (=0.59×25.29), that is, about half of total rise in degree premium (27.4 percent). Hence, the union decline in these two groups has a combined positive effect of 9.39 percent (=14.92-5.53) on the degree premium, which can account for about 34.24 percent (=9.39/27.4) of the rise in degree premium. This result is consistent with literature on trade union effect on earnings inequality such as Schmitt (1995) (about 21 percent of the rise in degree premium), Machin (1997) (about 40 percent of the rise in male variance) and Bell and Pitt (1998) (about 20 percent of the male earnings distribution). And, as Tzannatos and Aidt (2006) argues, the weak need trade union as their representation to improve their welfare more than those not weak.

Following the same way, we calculate the overall effect of the tax and benefit system. The increasing benefit replacement ratio in unskilled workers (6.54 percent) can reduce the degree premium by about 2.55 percent, which is about 9.31 percent (=2.55/27.4) of the rise in degree premium. The market condition variable (as a proxy of business cycle) and the industrial shifts variable can only account for a small part, 3-4 percent of the rise in degree premium.

Moreover, the increasing computer usage in the DEGREE group (61.21 percent) increase the degree premium by about 15.91 percent (=0.26×61.21). At the same time, however, the increasing computer usage in the unskilled group (24.31 percent) decreases the degree premium by about 10.69 percent (=0.44×24.31). Hence, the increasing computer usage in these two groups has a combined effect of 5.22 percent (=15.91-10.69) increase on the degree premium, which accounts for about 19.04 percent (=5.22/27.4) of the rise in degree premium during this period. Therefore, our results are consistent with the cross-country results of Koeniger *et al.* (2007), which claim "changes in these institutions can explain a substantial part of observed changes in male wage inequality — at least as much as is explained by our trade and technology measures."

Sensitivity Tests

Results of sensitivity tests are summarised in Table 5. For simplicity, we only concentrate on the institutional effects on skill premiums in the private sector. Column (a) use weekly earnings as the dependent variable; Column (b) still uses hourly wage as the dependent variable and the six-skill-level framework, but only run the regression for two medium skilled groups (BOLEV and OLEV); Column (c) only takes the results from years after 1979. Since both the degree premium and trade union density are non-stationary after 1979, we only test the fixed effect ECM model in equation (4) to avoid the co-integration problem.

In column (a), equation (4) also works on weekly earnings. Trade union density of unskilled workers (-0.57) has bigger effect on the skill premiums than that of skilled workers (0.19). Hence, the similar decline in trade union for skilled and unskilled workers should increase the skill premiums. Moreover, unemployment benefits of unskilled workers can decrease skill premiums (-0.45), while that of skilled workers is insignificant. Thus, our conclusions from hourly earnings estimation still remain for weekly earnings. And this test suggests that weekly working hours are not sensitive to these explanatory variables.

In column (b), a one point increase of trade union density in the medium skilled group (BOLEV and OLEV, tud_{jt}) can increase the skill premiums by 0.23 percent in long run and by 0.27 percent in short run, while a one point increase of trade union density in the unskilled group (tud_{nt}) decrease the skill premiums by 0.31 percent. Hence, the overall effect of union decline also pushes up skill premiums of medium skilled workers.

As far as special periods are concerned, column (c) shows that the effect of trade union is much more prominent in the years after 1979, and only changes in trade union density of unskilled workers are important. A one point increase of trade union density of the unskilled group (tud_{nt}) decreases the skill premiums by 1.58 percent. Unemployment benefits of unskilled workers can decrease skill premiums (-0.4), while that of skilled workers is insignificant. Thus, our conclusion from the entire period still holds for the special period of 1980-2002.

Conclusions

This paper analyzes the links between labour market institutions and the skill premiums in the UK, controlling for labour market conditions, industrial structure shifts and skill-biased technology change. We find the institutional factors such as trade union, tax wedge and benefit system are very important for skill premiums.

For the skill premiums in the private sector, institutions are more important for unskilled workers than those skilled. The trade union decline after 1979 is associated with different effect on wages of skilled and unskilled workers and pushes the skill premiums up. By using the fixed effect ECM model, we find that the trade union decline in unskilled workers can explain more than half of the rise in degree premium over the period 1979-1998. The overall effect of trade union in all workers can explain about one-third of degree premium increase in the same period. Trade union effect is also significant for skill premiums of low skilled workers and higher in years after 1979 than in the 1970s. Although the mark-up effects of tax wedge are not significant in this fixed effect ECM model, unemployment benefits of unskilled workers in the private sector reduce skill premiums by about 9 percent over the period 1979-1998.

For the public sector, we also find the significant effect of the trade union of unskilled workers on skill premiums. However, skill premiums in the public sector appear to be less responsive to institutional factors than in the private sector. A compensation model as in Acemoglu *et al.* (2001) or a bureaucratic and administered price model as in Kaufman (2007) might be better for explaining wage management in the public sector.

Evidence from this study suggests that the decline of trade union has a profound impact on the wages of unskilled workers. The decentralization of wage setting institutions in the private sector withheld the wage growth of unskilled (often low-wage) workers who are left behind by their skilled counterparts. Wage setting institutions may be more inclusive for unskilled worker in the public sector, for example, cleaners and assistant nurses in public sector hospitals in Grimshaw (2009). The combined effects of more coordinated and renewed centralised system of wage-setting institutions, more unemployment benefits of unskilled workers in the private sector and extension of public-sector terms and conditions to unskilled workers can decrease the earnings inequality in the UK.

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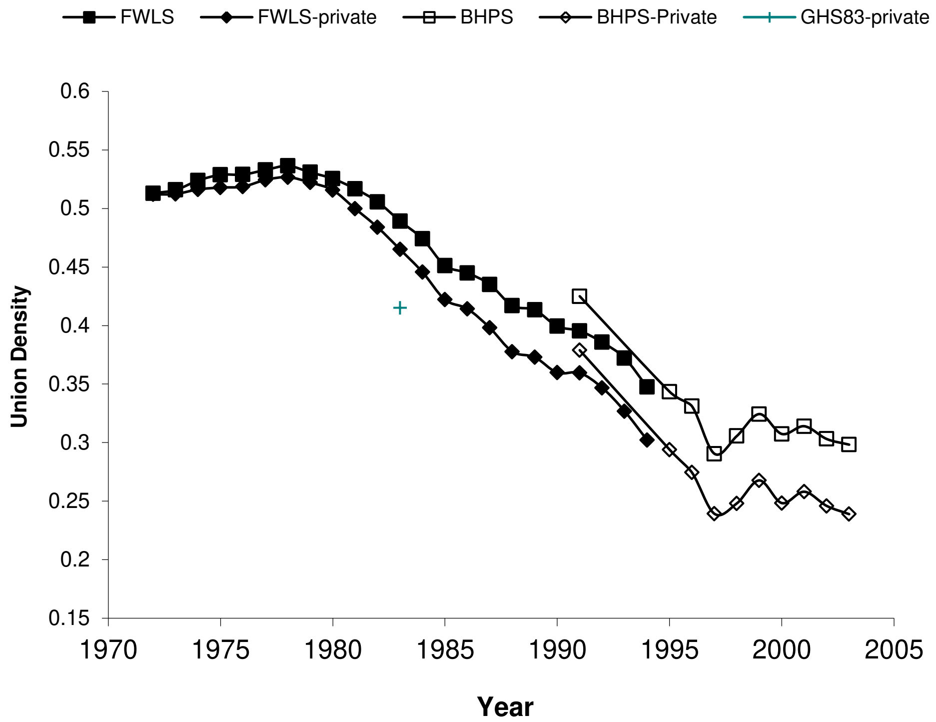
Figure 1 Trade union density in the UK, males 1972-2002



Note: All figures are calculated from the GHS 1983, the FWLS 1994/1995 and the BHPS 1991-2002.

Table 1 Data description by skill and sector, 1972-2002

1A. Descriptive statistics of variables in the private sector, by skill levels (%)

Variables/ skill Levels	NOQUAL Mean	NOQUAL Std. Dev.	BOLEV Mean	BOLEV Std. Dev.	OLEV Mean	OLEV Std. Dev.	ALEV Mean	ALEV Std. Dev.	HIGHER Mean	HIGHER Std. Dev.	DEGREE Mean	DEGREE Std. Dev.
Skill premiums			9.00	2.90	16.81	2.81	27.90	2.87	41.49	4.59	67.21	7.08
Trade union density	40.43	8.87	51.03	12.28	45.46	11.12	47.64	13.96	42.02	9.69	34.21	10.53
Tax wedge	16.60	2.59	16.96	2.76	17.13	2.40	17.57	2.35	18.91	2.21	19.62	2.41
Replacement Ratio	32.77	3.66	30.65	3.68	27.56	3.76	23.75	3.73	19.71	4.20	15.41	3.25
Unemployment rate	12.21	5.44	7.98	4.16	6.70	3.62	5.25	2.73	2.76	1.69	3.14	1.61
Manufacturing proportion	35.91	5.51	35.91	6.65	29.82	3.42	31.16	4.94	32.85	3.41	21.86	2.84
Computer usage	16.85	14.63	21.12	17.78	25.18	20.44	28.90	23.25	31.09	24.23	34.97	27.49

1B. Descriptive statistics of variables in the public sector, by skill levels (%)

Variables/ skill Levels	NOQUAL Mean	NOQUAL Std. Dev.	BOLEV Mean	BOLEV Std. Dev.	OLEV Mean	OLEV Std. Dev.	ALEV Mean	ALEV Std. Dev.	HIGHER Mean	HIGHER Std. Dev.	DEGREE Mean	DEGREE Std. Dev.
Skill premiums			14.61	6.91	28.46	7.45	35.97	7.34	41.85	6.63	61.33	8.44
Trade union density	59.10	7.52	43.28	9.60	43.46	9.51	41.82	13.79	66.42	8.96	65.47	6.57
Tax wedge	16.74	2.67	17.05	2.76	17.37	2.46	17.74	2.32	19.16	2.23	19.92	2.42
Replacement Ratio	32.59	3.97	30.49	3.97	26.92	3.87	23.44	3.75	19.48	4.19	15.41	3.25
Unemployment rate	11.85	5.48	7.60	4.03	6.18	3.37	5.06	2.66	2.79	1.70	2.73	1.45
Manufacturing proportion	35.93	5.77	35.82	6.99	30.21	3.26	31.27	4.63	31.86	3.24	21.36	2.84
Computer usage	16.85	14.63	21.12	17.78	25.18	20.44	28.90	23.25	31.09	24.23	34.97	27.49

Data source: GHS 1972-2002; FWLS 1994/1995; BHPS 1991-2002; and FES 1972-2002.

Table 2 Institutions and skill premiums, male 1972-2002, estimation from equation (3)

Dependent variable: skill premiums (s_{jt}, j = b, o, a, h and d)	**Private**	**Public**
Trade union density of skilled groups (tud_{jt})	0.15*	-0.03
	(0.09)	***(0.06)***
Trade union density of unskilled group (tud_{nt})	-0.21	-0.09
	(0.17)	***(0.15)***
Tax wedge of skilled groups (tax_{jt})	1.03	-0.15
	(0.82)	***(1.13)***
Tax wedge of unskilled group (tax_{nt})	-1.70**	-0.80
	(0.84)	***(1.05)***
Benefit index of skilled groups ($repr_{jt}$)	-0.13	-0.16
	(0.15)	***(0.19)***
Benefit index of unskilled group ($repr_{nt}$)	-0.19	0.39
	(0.24)	***(0.25)***
Unemployment rate of skilled groups (u_{jt})	-0.25	0.50
	(0.28)	***(0.37)***
Unemployment rate of unskilled group (u_{nt})	-0.01	0.69**
	(0.21)	***(0.31)***
Manufacturing proportion of skilled groups (ind_{jt})	-0.23*	-0.08
	(0.14)	***(0.17)***
Manufacturing proportion of unskilled group (ind_{nt})	-0.23	1.45**
	(0.49)	***(0.32)***
Computer usage of skilled groups ($comp_{jt}$)	0.29***	-0.09
	(0.11)	***(0.16)***
Computer usage of unskilled group ($comp_{nt}$)	-0.57***	0.50
	(0.19)	***(0.30)***
Minimum wages (MW)	1.08	-
	(3.24)	-
Observations	140	140
Groups	5	5
R^2 (within)	0.50	0.66
Group dummies	Yes	Yes
Year dummies	Yes	Yes
Year trend	Yes	Yes

Notes: Estimated standard errors are under the coefficients. ***, ** and * denote significance at 1%, 5% and 10% levels for two-tail tests. In the first step, we have 114,491 observations in the private sector and 23,323 observations in the public sector. In the second step, we have 5×29=145 observations in both sectors. Following the wage cyclicality literature (Solon et al. 1997; Peng and Siebert 2008), unemployment rates and unemployment benefit index have actually one year lag. Hence, we have only 140 observations in step 2 regression.

Table 3 Institutions and skill premiums, male 1972-2002, estimation from equation (4)

Dependent variable: growth rate of skill premiums (ds_{jt}, j = b, o, a, h and d)	Private	Public
Lag of trade union density of skilled groups (tud_{jt-1})	0.18**	0.03
	(0.10)	*(0.07)*
Growth of trade union density of skilled groups ($dtud_{jt}$)	0.12	-0.12
	(0.10)	*(0.07)*
Lag of trade union density of unskilled group (tud_{nt-1})	-0.59***	-0.40**
	(0.26)	*(0.20)*
Growth of trade union density of unskilled group ($dtud_{nt}$)	-0.24	0.64***
	(0.23)	*(0.13)*
Lag of skill premiums (s_{jt-1})	-0.71***	-1.12***
	(0.09)	***(0.10)***
Tax wedge of skilled groups (tax_{jt})	0.46	-0.01
	(0.80)	***(1.11)***
Tax wedge of unskilled group (tax_{nt})	-0.72	-0.78
	(0.95)	***(1.04)***
Benefit index of skilled groups ($repr_{jt}$)	-0.05	-0.20
	(0.15)	***(0.19)***
Benefit index of unskilled group ($repr_{nt}$)	-0.39*	0.63**
	(0.26)	***(0.29)***
Unemployment rate of skilled groups (u_{jt})	-0.47*	0.59
	(0.28)	***(0.37)***
Unemployment rate of unskilled group (u_{nt})	0.07	0.47*
	(0.31)	***(0.30)***
Manufacturing proportion of skilled groups (ind_{jt})	-0.24*	-0.15
	(0.14)	***(0.17)***
Manufacturing proportion of unskilled group (ind_{nt})	0.15	0.38
	(0.41)	***(0.35)***
Computer usage of skilled groups ($comp_{jt}$)	0.26***	-0.06
	(0.11)	***(0.16)***
Computer usage of unskilled group ($comp_{nt}$)	-0.44***	0.47
	(0.17)	***(0.30)***
Minimum wages (*MW*)	-	-
	-	-
Observations	140	140
Groups	5	5
R^2 (within)	0.61	0.82
Group dummies	Yes	Yes
Year dummies	Yes	Yes
Year trend	Yes	Yes

Notes: Estimated standard errors are under the coefficients. ***, ** and * denote significance at 1%, 5% and 10% levels for two-tail tests.

Table 4 Contribution of explanatory factors to the degree premium, the private sector

	1972-1979	1979-1998	1998-2002
4A. Changes of each variable (%)			
Degree premium (s_{dt})	-17.67	27.41	-17.56
Trade union density (tud_{dt})	4.55	-30.75	5.78
Trade union density (tud_{nt})	8.41	-25.29	-1.29
Tax wedge (tax_{dt})	3.83	-0.36	-0.55
Tax wedge (tax_{nt})	3.40	-1.40	-1.32
Benefit index ($repr_{dt}$)	-2.60	5.66	-3.97
Benefit index ($repr_{nt}$)	-3.40	6.54	-2.13
Unemployment rate (u_{dt})	-0.72	2.16	-0.42
Unemployment rate (unt)	1.05	5.59	-4.29
Manufacturing proportion (ind_{dt})	7.72	-3.74	-4.18
Manufacturing proportion (ind_{nt})	0.38	-6.70	-4.42
Computer usage ($comp_{dt}$)	11.83	61.21	16.57
Computer usage ($comp_{nt}$)	3.96	24.31	22.49
4B. Effects of changes in each explanatory variable (%)			
Trade union density (tud_{dt})	0.82**	-5.53**	1.04**
Trade union density (tud_{nt})	-4.96***	14.92***	0.76***
Tax wedge (tax_{dt})	-	-	-
Tax wedge (tax_{nt})	-	-	-
Benefit index ($repr_{dt}$)	-	-	-
Benefit index ($repr_{nt}$)	1.32*	-2.55*	0.83*
Unemployment rate (u_{dt})	0.34*	-1.02*	0.20*
Unemployment rate (unt)	-	-	-
Manufacturing proportion (ind_{dt})	-1.85*	0.90*	1.00*
Manufacturing proportion (ind_{nt})	-	-	-
Computer usage ($comp_{dt}$)	3.08***	15.91***	4.31***
Computer usage ($comp_{nt}$)	-1.74***	-10.69***	-9.90***
4C. Overall contribution of each factor (%)			
Trade union density	23.45	34.24	-10.26
Tax wedge	-	-	-
Benefit index	-7.50	-9.31	-4.72
Unemployment rate	-1.92	-3.71	-1.13
Manufacturing proportion	10.49	3.28	-5.71
Computer usage	-7.55	19.04	31.82

Notes: All figures in Table 4 are calculated using estimates in Table 3. ***, ** and * denote significance at 1%, 5% and 10% levels for two-tail tests. Significance of each variable in the middle panel is from Table 3.

Table 5 Institutions and the skill premium (Sensitivity Tests), male 1972-2002, estimation from equation (4)

Dependent variable: growth rate of the skill premiums (ds_{jt}, j = b, o, a, h and d)	**(a) Weekly earnings**	**(b) Semi-skilled**	**(c) 1980-2002**
Lag of trade union density of skilled groups (tud_{jt-1})	0.19**	0.23*	0.14
	(0.10)	***(0.13)***	***(0.15)***
Growth of trade union density of skilled groups ($dtud_{jt}$)	0.11	0.27**	0.05
	(0.10)	***(0.12)***	***(0.14)***
Lag of trade union density of unskilled group (tud_{nt-1})	-0.57**	-0.31**	-1.58***
	(0.27)	***(0.16)***	***(0.69)***
Growth of trade union density of unskilled group ($dtud_{nt}$)	-0.22	0.01	-0.62
	(0.24)	***(0.13)***	***(0.43)***
Lag of skill premiums (s_{jt-1})	-0.64***	-0.68***	-0.82***
	(0.09)	***(0.16)***	***(0.11)***
Tax wedge of skilled groups (tax_{jt})	0.54	-0.45	-0.05
	(0.83)	***(0.42)***	***(1.04)***
Tax wedge of unskilled group (tax_{nt})	-0.83	0.58	-0.51
	(0.98)	***(0.39)***	***(0.57)***
Benefit index of skilled groups ($repr_{jt}$)	0.00	0.03	-0.03
	(0.15)	***(0.16)***	***(0.17)***
Benefit index of unskilled group ($repr_{nt}$)	-0.45*	0.05	-0.40**
	(0.27)	***(0.18)***	***(0.21)***
Minimum wages (MW)	-	-	1.66
	-	-	***(2.32)***
Observations	140	56	105
Groups	5	2	5
R^2 (within)	0.58	0.60	0.58
Group dummies	Yes	Yes	Yes
Year dummies	Yes	No	Yes
Year trend	Yes	Yes	Yes

Notes: as for Table 2.