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INSTITUTIONAL EFFECTS ON SKILL PREMIUM

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Abstract

Skill premium in the United Kingdom has increased substantially since the 1970s. This paper analyzes the link between institutions and the skill premium in the UK controlling for other explanatory variables such as market conditions, international trade and skill-biased technology.

For the private sector, institutions are more important for the unskilled baseline group than the skilled groups. The trade union decline after 1979 brings different effect on wages of skilled and unskilled workers and pushes the skill premium up. We find that the trade union decline in unskilled workers can explain about one third degree premium increase over the period 1979-1998. The overall effect of trade union in all workers can explain about 13.34 percent of degree premium increase during the same period. Our results are insensitive to skill group categorization.

Moreover, we find that minimum wages can only decrease skill premiums of high skilled workers rather than low skilled workers. The increase of unemployment benefit over the period 1979-1998 reduces the increasing the skill premium by about 13.6 percent. But, the mark-up effect of taxation increase skill premium by about 8.26 percent. We find no significant associations between above institutions and skill premiums in the public sector.

Keywords: skill premium, trade union, fixed effect model, GHS

JEL codes: J31, J51, K31

Preliminary: Not for quotation

1. Introduction

Wage inequality in the United Kingdom has substantially increased in the last three decades, which is much higher than that in continental European countries. Analysis of the wage distribution in the UK by Gosling, Machin and Meghir (2000) emphasizes the importance of skill (education) attainment of recent cohorts and changes over time in the skill (education) wage premium. In this paper, we will analyze the links between labour market institutions and the skill (education) premium.

Changes of skill premium are driven by both supply and demand factors. The existing literature has tried to isolate the causal factors underpinning these market changes. The most popular candidates are skill-biased technical change (SBTC) and increased international trade. In fact, there is strong evidence of the empirical association between proxies for SBTC (computers or other ICT facilities) and the widened wage gap of the UK and US in the 1980s (see Krueger 1993, Machin and Reenen 1998, Autor, Katz and Krueger 1998, Katz and Autor 2000, Machin 2001 and O'Mahony et al 2008).

The trade explanation focuses on changes in product demand largely associated with large trade deficits in the 1980s (see trade deficits in Nickell 2006, Table C). Wood (1994, 1995 and 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 labour economists (see Schmitt 1995, and Machin and Reenen 1998). And even trade economists such as Krugman and Lawrence (1993) and Sachs and Shatz (1994) also point that the effect of international trade on relative demand for skill is surprising 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 premium, 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 premium 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 premium 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 and Puhani 2003).

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.[1] 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.[2]

[1] The long reference list includes Katz, Loveman and Blanchflower (1995), Blau and Kahn (1996), Machin (1996, 1997), Gottschalk and Joyce (1998), Card et al (2003) and Koeniger et al (2004 and 2007).
[2] Card (2001) for the United States, 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.

The latest finding of Addison et al (2007) may be the only one to analyze the effect of trade unions on the British wage gap by gender and private/public sector, allowing for worker education. They point out that deunionization is shown to account for surprising little of the increase in earnings dispersion in the private sector for either males or females. The lower union decline in the public sector, however, has actually stronger effect. Additionally, in the public sector, trade unions no longer reduced earnings dispersion as much as they once did by virtue of their growing tendency to organize more skilled groups. But, it is still not very clear why deunionization has such a different effect on the wage differentials in the private and public sector. This paper will push the discussion further and explore the union effect on the skill premium 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 have found that national minimum wages reduce wage inequality by increasing the bottom deciles of the pay distribution without a negative impact on employment (see Dickens and Manning 2004, Machin and Wilson 2004, Stewart 2004 and the summary of Metcalf 2004 in this session). DiNardo et al. (1996) and Lee (1999) also find the same effect of minimum wages for the United States. Again, this paper will push the minimum wages discussion further.

For other labour market institutions, the different effects of the tax wedge and unemployment benefits on skilled and unskilled workers may affect the skill premium in a similar way to employment protection. Brewer et al (2008) study about five million income tax returns covering the period 1996-2005 from two different data sources (the HBAI and the SPI). They argue 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. This paper will push these arguments further and investigate the different effects of the tax wedge and benefits on skill premiums.

Most previous institution-specific empirical studies only use cross-section data at country level (see Blau and Khan 1996). The only two cases of longitudinal data are Wallerstein (1999, for 16 developed countries in 1980, 1986 and 1992), and Koeniger et al (2004 and 2007, for 11 developed countries over 1973-1998). Our analysis builds on Koeniger et al (2004 and 2007), but we construct a balanced panel data of six skill (education) groups in the United Kingdom over 1972-2002 from several micro datasets. Using these data, we can investigate the effect of both aggregated and disaggregated supply-demand-institution factors for distinct skill groups on skill premiums in the UK, and quantitatively assess respective importance. Moreover, cross-national analysis at country level cannot tell us whether the story is the same for the private and public sector. The public-private distinction will also be our contribution.

The remainder of the paper is organised as follows. Section 2 reviews the theoretical model based on Koeniger et al (2004) and motivates the estimated log-linear equation. Section 3 provides our empirical specification and introduces the basic framework of our panel data. Section 4 describes the main data sources and measures those variables used in our empirical specification. Section 5 estimates empirical results. The last section concludes.

2. A model of trade union bargaining

Our empirical work is based on the union bargaining model provided by Koeniger et al (2004). In this section, we review this model, in which labour market institutions alter the outside option of skilled and unskilled workers differently, and thus affect relative labour demand as well as the wage differentials. Changes of institutions as well as market conditions, technologies and international competition are reflected in the following equation[3]:

[3] See Koeniger et al (2004) for derivation details in 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).
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$$\ln\left(\frac{\overline{w_H}}{\overline{w_L}}\right) \cong f(\underset{+}{tud_H}, \underset{-}{tud_L}) + g(\underset{+}{tax_H}, \underset{-}{tax_L}) + h(\underset{+}{repr_H}, \underset{-}{repr_L}) + j(\underset{-}{MW}) + k(\underset{-}{u_H}, \underset{+}{u_L})$$
$$+ l(\underset{+}{comp_H}, \underset{-}{comp_L}, \underset{+}{ind_H}, \underset{+}{ind_L}) \qquad (1)$$

In equation (1), H denotes high skilled workers while L is low skilled workers. The skill premium, i.e. log form gross wage differentials for skilled workers, $\ln\left(\frac{\overline{w_H}}{\overline{w_L}}\right)$, 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 skill, 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.

Let us start with trade unions. 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. Firstly, the skill premium will be smaller if unions favour unskilled workers (tud_L) more than skilled workers (tud_H). Secondly, the trade union bargaining model in Koeniger et al (2004) marks up earnings tax as a 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 premium, 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 premium, while higher unemployment rates for unskilled workers (u_L) are likely to increase it. The overall effect of unemployment benefit (or unemployment rates) on the skill premium depends a comparison between its respective wage effect on skilled and unskilled workers.

Thirdly, as DiNardo et al (1996) reveals, a minimum wage can directly compress the skill premium 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 premium smaller.

Finally, we also need to discuss those technology and international trade variables in equation (1). 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 premium temporarily, while new technologies adopted by unskilled workers ($comp_L$) also increase their marginal products but decrease the skill premium. However, if new technologies are complementary to skills (see 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 effect on skilled workers than on unskilled workers. Hence, the relative demand for skill will increase and push up the skill premium.

International competition from newly industrialized 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 premium.

3. Empirical specification

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.[4] In step 1, we use all individual observations to estimate education wage differentials as proxies of skill premiums over time. This equation is given by:

$$\ln w_{it} = \alpha_0 + \sum_{t=1}^{T} n_t Y_t + \sum_{t=1}^{T} b_t B_{it} Y_t + \sum_{t=1}^{T} o_t O_{it} Y_t + \sum_{t=1}^{T} a_t A_{it} Y_t + \sum_{t=1}^{T} h_t H_{it} Y_t + \sum_{t=1}^{T} d_t D_{it} Y_t + \beta X_{it} + \varepsilon_{it} \quad (2)$$

where w_{it} is the real gross weekly wage rate, Y_t denotes a year dummy representing the base line group of NOQUAL; B_{it} denotes a dummy variable for workers with below O-level qualification; O_{it} denotes a dummy variable for the O-level group; A_{it} denotes a dummy variable for workers with A-levels; H_{it} denotes a dummy variable for workers with higher educational qualification but not degrees; and D_{it} denotes a dummy variable for worker with degree equivalent or above qualification.

X_{it} is a vector of the main additional factors that may influence wages including potential labour market experience, marital status, ethnicity and region, and ε_{it} is a random error term. Correspondingly, n_t (t=1…T, T=29 in this research) 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 year t. These coefficients are shown in Figure 1 below.

In the second step, we estimate the institutional effect on the skill premium, i.e. the incremental wage effect of each educated group in the first step. We stack b_t, o_t, a_t, h_t and d_t from equation (2) to form a skill premium variable s_{jt}, which is the skill premium of each education group relative to the baseline NOQUAL group in the

[4] Moulton (1986) shows that individuals in the same year/area will 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 1994) and wage curve literature (Nijkamp and Poot, 2005, p 434).

same year. Hence, a panel dataset is built to find the links between the skill premium and labour market institutions:

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

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

where s_{jt} is the skill premium for education group j in the year t, and labour market institutions indicators and those control variables of market conditions, technology, are 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.

We concentrate on a study of the skill premium. 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 premium (for example, Gosling, Machin and Meghir 2000, Figure 3.2, p642) as well as a decline of trade unions since the 1970s (for example, Bell and Pitt 1998, Figure 1, p516 and Disney et al 1998, Figure 1-3, p3-4). Since our panel data have a 29-year period, the skill premium of each group is probably non-stationary (see Figure 1 and ADF tests below), hence a co-integration problem may exist in the link between skill premiums and institutional variables. De-trending and simply differencing the data cannot resolve all problems.[5] 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.

[5] Simply de-trending and differencing to remove the non-stationary trend can avoid the spurious regression problem, but it also removes the any long-run information hence is not wiseful (see Harris 1995, p1)

Ammermueller et al (2007) use the same ECM approach in their wage curve research for Italy and Germany with panel data. We follow their approach, but only put the ECM in trade union density variables (tud_{jt} and tud_{nt}) to save degrees of freedom, 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:

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

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

Thus, in the above specification the long run equilibrium, between the *level* of the skill differentials and *level* of trade union density is embodied in an ECM. We use Stata's fixed effect programme (*xtreg, fe*, see Stata 2003) to estimate equation (3) and (4).

4. Data description

4.1 Wage level and skill premiums

The wage variable used here is from the GHS 1972-2002 and defined as the real gross weekly wage in 1995 pounds. Following the tradition of research (see Schmitt 1995, Dickens 2000 and Koeniger et al 2004, 2007), we concentrate on male full-time workers. Then, we have 140,625 observations (117,302 workers in the private sector and 23,323 workers in the public sector) in the first step regression based on equation (2).

The coefficients of year dummies derived from the first step (n_t), is presented in the Panel A of Figure 1 by sector. This graph illustrates the cumulative real wage growth of the NOQUAL group, which displays the log ratio of this group's earnings in each year relative to its level of real earnings in 1972. Hence, as a major part of workers in the NOQUAL group, mean wages have increased by about 40 percent (n_{2002}–n_{1972}≈0.4). However, real wages of unskilled workers in the public sector have only increased about 20 percent since 1972 (n_{2002}–n_{1972}≈0.2).[6] This big difference between two sectors supports the sample division in our further analysis.

Moreover, the coefficients of other year dummies in the first step (b_t, o_t, a_t, h_t and d_t) are stacked up to build the dependent variable (s_{jt}, j=b, o, a, h and d) used in the second step regression. It is a panel data with 5 groups over 29 years. Hence, we have 145 observations (5×29) for both private and public sector in the second step regression based on equation (3), which is presented in the Panel B and C of Figure 1.

Panel B and C shows skill premiums of education groups (b_t, o_t, a_t, h_t and d_t), which are wages of each education groups relative to the wage level of the NOQUAL group in the same year. For example, in Panel B, the line of DEGREE group shows that wages of workers with degrees was about 51.43 percent higher than workers in the NOQUAL group in 1979. But, in 1998, wages of workers with degrees was about 83.48 percent higher than workers in the NOQUAL group. Hence, degree premium had increased about 32 percent from 1979 to 1998 (=83.48-51.43).

Furthermore, we can see that skill premiums in both sectors share the same pattern of "higher skill level equals higher skill premium" in Panel B and C in Figure 1, which is reasonable. And, in the private sector, the skill premiums of the lower skilled groups (BOLEV, OLEV and ALEV) increase more slowly than those of the higher skilled groups (HIGHER and DEGREE) since the 1970s. However, in the public sector, we cannot find an increasing trend of the skill premium for any education group. Thus, the worsening of wage inequality since the 1970s is perhaps caused by the increasing skill premiums in the private sector rather than in the public sector.

4.2 Institutional variables

In this part, we describe the institutional variables such as trade unions, taxation, unemployment benefit and the NMW used in this paper. Besides the GHS 1992-2002, another three datasets have been used to measure those institutions: the UK Family

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[6] Our results show that the real wage of the NOQUAL group was about 84.3 pounds per week in the private sector in 1972, as well as about 73.94 pounds per week in the public sector. It seems unskilled workers in the public sector have a lower wage, perhaps because of different employment insurance for unskilled workers within the two sectors.

Expenditure Survey (FES 1982-2002)[7], the Family and Working Lives Survey (FWLS 1994/1995)[8] and the British Household Panel Survey (BHPS 1991-2002).[9]

The main purpose of using these additional datasets is to compile a time series on union density by education 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). The FES can provide indirect information, via a question on membership of a trade union or professional body. In the income section of the survey, individuals are asked if there are any deductions from pay for subscriptions to friendly societies, trade unions or professional bodies. This measure of trade union density has been used by several studies (e.g. Disney and Cameron 1990, Lanot and Walker 1998, and in particular, Bell and Pitt 1998, Figure 1, p516), which have used response to this question as evidence of union membership.

Presumably, it is possible to falsely classify some union members as non-union workers. Individual who do not pay their union subscription directly at source will not be included in this definition of union membership. Bell and Pitt (1998) argues that this trade union measure is reliable by comparing it with the Workplace Industrial Relations Survey (WIRS) in 1980 and 1990. However, as we will see later, trade union density derived from the FES is very unstable. And, the variable for union membership deduction in the FES is only available after 1981 while our investigation covers the period 1972-2002. A further well-known problem with the FES is that it cannot provide continuous and accurate information about workers' skill level and employment status. Hence, we only use the FWLS and BHPS to derive union density by skill.

[7] 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 EFS 2002, in which they have the same definition. We will not differentiate the two surveys in later discussion.

[8]The FWLS is a life and work history data, which provide representative information about people living in Britain. The FWLS is stored in TDA (Transition Data Analysis) software.

[9] 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.

Figure 2 compares the trade union density changes in different datasets, in which we can find trade union density derived from the FES is almost the same as that from the FWLS in 1982. In the next five years, however, trade union density in the FES dramatically dropped about 15 percent. This big drop cannot be found in the FWLS and the data from the Certification Office (The “Bain and Price” series, see Disney et al 1998 Figure 1, and Bell and Pitt 1998 Figure 1). Hence, the overall union density of all male workers in the FES seems much lower than in the FWLS, the BHPS or the “Bain and Price” series. This result confirms our doubts on the reliability of union density derived from the FES’ union due question. Whether there are deductions from pay for subscriptions seems an inferior indicator for union membership. Thus, we give up trade union density from the FES and use the FWLS and BHPS to build union density by skill level. [10]

Our union density variable is from the FWLS for the period 1972-1994, and for the period 1995-2002, it is from the BHPS. In Figure 2, the change of union density between 1991 and 1995 is very similar in the FWLS and BHPS, for both all workers and for the private sector. This similarity shows that the average union density has a consistent pattern for 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.

Figure 3 presents trade union density by skill level and sector over the last thirty years. The combination of the FWLS and BHPS reveals the trade union density in the semi-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 1979, during which earnings inequality moves 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 semi-skilled groups (BOLEV, OLEV and ALEV). And, we do not

[10] Moreover, in the FES, there is no variable about private and public sector of workers after 1986. Even during the period 1982-1986 with private/public sector information, the union density of the FES-private is also much lower than that of the FWLS-private, just as the overall union density.

find a clear decline of unions in the public sector since the 1970s. It is perhaps the reason for the different wage structure in these two sectors.

Moving on to the tax and benefit system, the division into private and public sectors is not necessary since the tax wedge and replacement ratios do not depend on sector. 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.

As for benefit indices, they measure the proportion of unemployment benefits relative to average earnings before tax. The GHS provides information for unemployment benefit over the entire period 1972-2000.[11] 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.[12] However, 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.

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 wage differentials. From the panel A of Figure 4, we see that the tax wedge gap between high skilled and low skilled group has been wider since the 1970s. This trend is consistent with the findings in Brewer et al (2008) that government has imposed large rises in taxation to fund higher benefit payments and tax credits in recent years.

[11] After 1996, the British unemployment benefit changed its name to job seeker allowance. We will keep using the unemployment benefit term in the discussion.

[12] Since the data about housing benefit (particular for council tax) are not consistent over time in the GHS, we do not include it.

On the other hand, the Koeniger et al (2004) model implies that the replacement ratios should be negatively correlated with the skill differential, if unemployment benefit is more generous for unskilled workers. Panel B of Figure 4 describes a higher benefit index for the low skilled groups. The interesting point is that the increase of benefit index of the DEGREE group during 1980-1985, in contrast to the decline in other groups. This result can help explain the degree premium increase in the 1980s.

As far as the minimum wages (*mw*) are concerned, the UK National Minimum Wage 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 (see NMW values after 1998 in Metcalf 2004, Table 1).

4.3 Control variables for international trade, technology and market conditions

The unemployment rate of each skill group plays an important role for the skill premium because it represents the market conditions and outside options of skilled and unskilled workers. We calculate the unemployment rate by skill level over the entire period using the GHS 1972-2002.[13] The theoretical model of Koeniger et al (2004) implies that there is a negative (positive) relationship between the unemployment rate of skilled (unskilled) workers and the skill premium, *ceteris paribus*.

Panel A of Figure 5 shows the different pattern of movement of each group over the business cycle. Obviously, the lower educated workers are more vulnerable when the labour market is loose. We also see that the unemployment gap between lower skilled and higher skilled worker became wider in the 1980s and early years of 1990s. Higher unemployment rates of unskilled worker worsen their outside option and also decrease the collective bargaining power of their trade unions. Thus, the skill premium should increase if the unemployment rate of unskilled workers increases faster than that of skilled workers.

[13] We compare the unemployment rates in the GHS with other data sources such as the Labour Force Surveys (LFS) and the BHPS. We find no much difference in these three data sources. Hence, we use the GHS here for consistency.

In the model of Koeniger et al (2004), international trade and technology determine the skill premium through relative prices and relative total factor productivity. In a 2×2 world, as frequently used in the literature, the effect of international trade can be represented by the employment shifts from the manufacturing to the service sector 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 service sector, the pressure from international competition must be bigger.

Panel B of Figure 5 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. In the early years of the 1980s, the manufacturing employment proportions even increase in high skilled groups. This result may also contribute to the increasing skill premium 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 (Kruger 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. As we expect, Panel C of Figure 5 shows sparse computer usage and a slow climb during years before 1980, which is consistent with the decline of the skill premium over the 1970s. Then, the acceleration of computer usage in the upper skill groups (ALEV, HIGHER and DEGREE), supports the increased skill premiums in the 1980s. Especially for the DEGREE group, computer usage increased from about 25 percent in 1980 to about 65 percent in 1995, much faster than low skilled groups (for example, computer usage of the NOQUAL group only increased from about 10 percent to about 25 percent during the same period). This pattern is also consistent with the dramatically increasing degree premium in the 1980s and early 1990s.

On the other hand, it is widely realised the diffusion of computers has become so widespread after the 1990s that a simple headcount may no longer measure the SBTC-induced demand shifts (Machin 2001, p772). Indeed, we find that lower skilled groups have a fast convergence process to those high skilled groups for computer usage after 1995. This convergence implies that computer usage may be an inferior indicator of skill biased technology for recent years.

5. Empirical results

In this section, we explore the associations between institutions and skill premiums using equation (3) and (4). Contribution analysis for every explanatory variable to the changes in the skill premium is presented using equation (4). We also test our results for alternative broader education categorization, and for different sup-groups and sub-periods.

5.1 Basic results

Table 1 presents the fixed effect results from equation (3) by the private and public sector. For all estimations in the public sector, there is no significant result, implying a static skill premium in the public sector. It seems that the wage setting in the public sector is against the model of Koeniger et al (2004). Bureaucratic and administered price models are needed to explain wage management in the public sector (see a summary in Kaufman 2007 using transaction costs theory). Thus, we will concentrate on the private sector, since this is the majority of the workforce.

Firstly, there are significant associations between the skill premium and trade union density of high skilled groups (tud_{jt}) and unskilled group (tud_{nt}) in the private sector. A point increase of trade union density in the skilled group will increase the skill premium by 0.19 percent, while a point increase of trade union density in the baseline unskilled group will decrease the skill premium by 0.29 percent. This result is strong evidence of a trade union effect on skill premiums and shows that trade unions have different effects for workers at different skill levels.

Using estimates from equation (3), we can simulate the effect of trade union decline on the degree premium after 1979. Since the DEGREE group is on the upper tail of earning distribution as the NOQUAL group is on the lower tail, the degree premium from equation (2) can be regarded as a proxy of earnings inequality. The degree premium in the private sector has increased by about 20 percent during the period 1979-2002 (see Panel A of Figure 1), while trade union density in both DEGREE and NOQUAL groups has decreased by about 25 percent during the same period (see Panel A of Figure 3). Thus, the union decline in the unskilled group (NOQUAL) increases the degree premium by about 7.5 percent (=0.29×25), which is about 38 percent (=7.5/20) of the overall increase of the degree premium after 1979. Hence, if the trade union density of the unskilled group in 2002 kept the same level as in 1979, degree premium would be about 38 percent lower than the current level.

In the same vein, the union decline in the DEGREE group decreases the degree premium by about 5 percent (=0.19×25). Trade unions are more important for the unskilled workers than skilled workers, which is reasonable. The similar magnitude of union decline in the two groups brings a combined effect on the degree premium, which pushes up the degree premium by about 2.5 percent (=7.5-5). This results implies that about 13 percent (=2.5/20) of increase of the degree premium is from the changes of trade union. Thus, the similar decline of unions for skilled and unskilled workers worsens wage inequality by increasing the skill premium, since wages of unskilled worker will be affected more. This result is consistent with other researches such as Schmitt (1995 Table 5.11, p202) who found union membership losses account for about 21 percent of the rise in the university differential during the 1980s.[14]

Secondly, for the tax and benefit system, the institutional effects are also significant. The tax wedge shows a significant mark up effect for skilled and unskilled workers as the theoretical model predicts. A one point increase of the skilled workers' tax wedge (tax_{jt}) increases the skill premium by about 2.03 percent, and the same change in the unskilled workers' tax wedge (tax_{nt}) decreases the skill premium by about 2.96 percent.

[14] A formal simulation using estimates from equation (4) is presented in Table 3.

Moreover, the benefit index of high skilled workers ($repr_{jt}$) has no significant effect on the skill premium. Wages of skilled workers seem not affected by their benefits. However, the benefit index of the unskilled group is significantly negative as we expect. A one point increase of the benefit index of unskilled workers ($repr_{nt}$) decreases the skill premium by about 0.5 percent. A higher benefit index of unskilled workers means a better outside option for them. Hence, the trade union of unskilled workers may strongly bargain for unskilled workers and decrease the skill premium. In addition, the minimum wage variable does not show significant effects on the skill premium in Table 1. This surprising insignificant effect of the MW variable may be because that the NMWA only covers a few years in our sample, and affect the baseline unskilled group as well as some semi-skilled groups (OBLEV, OLEV and ALEV). We will test this argument later.

Thirdly, unemployment rates should also reflect the outside options of workers. Workers can bargain more strongly if the labour market is tight. Yet, from Table 1, there is no significant effect of market conditions on the skill premium. Many researchers point out that more unskilled workers may join the employment as labour market is tight and push the overall wages down (see Solon et al 1994 for the US and Devereux and Hart for the UK). 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.[15]

As for other variables, we cannot find significant effect of international trade on the skill premium in Table 1, which is consistent with many researches such as Schmitt (1995) and Machin and Reenen (1998). However, the computer usage variables show significant positive associations with workers' wages, proving new technologies can improve productivity of all workers. A one point increase in computer usage of skilled workers ($comp_{jt}$) brings a 0.35 percent increase of the skill premium, while that of unskilled workers ($comp_{nt}$) decreases about 0.51 percent of the

[15] See more details in Peng and Siebert (2007, 2008), in which we discuss the wage cyclicality and control the composition effect using panel data.

skill premium. It seems that adaptation of new technology for unskilled workers is even more important to decrease the skill premium.

As we see, the computer usage of the DEGREE group increased about 65 percent during the period 1979-2002 (see Panel C of Figure 5), while the computer usage of the NOQUAL group only increased about 35 percent during the same period. Hence, computer usage of unskilled workers decreases the degree premium by only about 18 percent (=35×0.51), while that of the DEGREE group increases the degree premium by about 23 percent (=65×0.35). The overall effect of computer technologies on the degree premium is about 5 percent (=23-18), Hence, technology shifts account for about one fourth of the degree premium increase (=5/20) after the 1970s.[16]

5.2 Results of ECM specification

As we know, the fixed effect results in Table 1 may be biased by co-integration problems since the skill premium and trade union density are non-stationary. Augmented Dickey-Fuller Unit root test (ADF) shows that the degree premium is non-stationary over the entire period, while skill premiums of other education groups are stationary over the entire period, but non-stationary during the period 1979-1995. And, trade union densities of all education groups are non-stationary over the entire period. Hence, Table 2 tries the fixed effect ECM model using the better specification in equation (4). This improvement in methodology may clear up the relationship between institutions and the skill premium. In fact, results in Table 2 are not much different from Table 1, implying that the co-integration problem may be not important in the research on the long entire period.[17]

The main improvement is that institutional effects on the skill premium are more important and significant. A one point increase of trade union density in the

[16] However, after 1995, the computer usage in the unskilled groups had a fast convergence to other skilled groups, and may not represent skilled biased technology in recent years (see Machin 2001 and O'Mahony et al 2008).

[17] ADF test shows that only the degree premium is non-stationary over the entire period (t value:-2.22), especially during the period 1980-1995 (t value:-1.619). Skill premiums of other groups are all stationary over the entire period, but also non-stationary during the period 1980-1995. 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, even though the co-integration problem is not serious for the regression on all groups over the entire period, it may be serious for specific sub-groups and sub-periods.

skilled group (tud_{jt}) still increases the skill premium by 0.19 percent, which is the same as in Table 1. However, the effect of trade unions on skill premiums becomes more important and significant for unskilled workers. A one point increase of trade union density in unskilled group (tud_{nt}) will decrease the skill premium by 0.40 percent.

Next, the tax wedge show the right mark up effect as the model expects; the benefit variable of high skilled workers ($repr_{jt}$) is also insignificant as in Table 1. However, the benefit variable of the unskilled group ($repr_{nt}$) becomes more significant and important than that in Table 1. One point increase of benefit variable of unskilled workers can decrease the skill premium by about 0.77 percent. And, the minimum wage variable is still insignificant as before.

Furthermore, unemployment rates of unskilled workers (u_{nt}) now show a positive association with the skill premium (0.44), which is reasonable. Since unskilled workers are the main source of composition biases, the overall wages of unskilled workers may show some procyclicality. Hence, the higher unemployment rate of unskilled workers will bring down their wages and increase the skill premium. This is consistent with the model of Koeniger et al (2004) and the wage cyclicality literature.[18] As in Table 1, there is no significant effect from international trade in Table 2. The technology change also shows the right direction for both skilled and unskilled groups. The improvement of computer usage in the unskilled workers appears to decrease the skill premium continuously.

Another interesting point worthy of mention is the ECM variable, which is the lagged skill premium variable, s_{jt-1}. Its coefficient is 0.68 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, though the public sector model's variables are all insignificant. In fact, there may be no ECM in the public sector since the skill premium there appears to be static as administered price model shows.

[18] Also see Solon et al 1994 for the USA, and Devereux and Hart 2006 for the UK. However, the unemployment rates of skilled workers (u_{jt}) in the public sector shows a significant positive relationship with the skill premium. This result implies that the wage of skilled workers in the public sector cannot adjust with market conditions.

Table 3 estimates the contribution of each explanatory variable in equation (2) to the changes in the degree premium over three typical periods: 1972-1979 and 1979-1998 and 1998-2002.[19] The top panel shows changes in degree premium, and trade union density, the tax wedge, benefit replacement ratios, unemployment rates, computer usage and manufacturing ratios for both the NOQUAL and DEGREE groups. The middle panel shows effects of changes in 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 (see Panel B of Figure 1) and earnings inequality (see Panel A of Figure 2.3) have increased to the highest level in our sample years.

From Table 3, the overall effect of trade union on the degree premium is as important as the technology shifts during the period 1979-1998. For example, the union decline in the unskilled group (NOQUAL, -25.29 percent) increases the degree premium by about 10.12 percent (=0.4×25.29), which is about one third (=10.12/32.05) of the overall increase of the degree premium over this period. At the same time, however, the union decline in the DEGREE group (-30.75 percent) decreases the degree premium by about 5.84 percent (=0.19×30.75). The similar magnitude of union decline in the two groups brings a combined effect of about 4.28 percent (=10.12-5.84) increase on the degree premium, which is about 13.34 percent (=4.28/32.05) of increase of degree premium. Following the same way, the technology shifts account for about 14.44 percent of the degree premium increase during this period.

We also calculate the overall effect of the tax and benefit system, which can reduce the degree premium by about 5.34 percent (=8.26-13.6). The market condition variable, as a proxy of business cycle increases the degree premium by about 6.26 percent. But, contribution of international trade is not significant. Therefore, this

[19]The degree premium (see Panel B of Figure 5) and earning inequality (see Panel A of Figure 2.3) change in a similar pattern in these three periods. We regard the degree premium as a proxy of earnings inequality (90^{th}-10^{th} percentile differential) in this part. Thus, our analysis on the degree premium can be applies on earnings inequality. All figures in Table 3 are calculated using estimates in Table 2. Insignificant variables in Table 2 are also calculated, but actually not strong evidence.

result is consistent with the findings of Koeniger et al (2007, p352) who 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."

5.3 Sensitivity Tests

Results of sensitivity tests are summarised in Table 4. For simplicity, we only concentrate on the institutional effects on skill premiums in the private sector. Column (a) categorizes all workers into only three groups: the unskilled group (below A-level, including NOQUAL, BOLEV and OLEV), the semi-skilled group (ALEV) and the skilled group (HIGHER and DEGREE). The unskilled group is used as the baseline to calculate the skill premium. Column (b) still uses the six-skill-level framework but only run the regression for the sub-group sample of high skilled workers in the HIGHER and DEGREE groups. In addition, column (c) also uses the six-skill-level framework but only gives the results for the sub-group sample of semi-skilled workers in the groups of ALEV, OLEV and BOLEV. Column (d) and column (e) take the results from two different periods: the years before 1980 and years in and after 1980. Since both the skill premium and trade union density are non-stationary after 1980, we only apply the fixed effect ECM model equation (4) to avoid the co-integration problem.

In column (a), the three-skill-level framework shows a similar trade union effect on the skill premium to the six-skill-level framework. Trade union density of unskilled workers (-0.4) still has bigger effect on the skill premium than that of semi-skilled/skilled workers (0.25). But, other institutional factors appear insignificant now.

In column (b), the trade union effect on the skill premium becomes more prominent. A one point increase of trade union density in the skilled group (HIGHERE and DEGREE, tud_{jt}) can increase the skill premium by 1.02 percent, while a one point increase of trade union density in the unskilled group (tud_{nt}) will decrease the skill premium by 1.54 percent. Hence, contribution of overall effect of trade union decline to the degree premium is even higher after the 1970s.

A simple simulation shows that as trade union density declines by 30 percent during the period 1979-2002, the degree premium would increase by about 15.6 percent, that, is, (1.54-1.02)×30. This is about 50 percent (=15.6/30) of the overall increase of degree premium over this period. Column (b) excludes the trade union effects on semi-skilled worker's premiums, and hence is more dramatic than in Table 2 with all education groups. Since the skill premium of high skilled workers is a proxy of earnings inequality, this result is consistent with Machin (1997) which also obtains results that the male variance would have been 40 percent less if the 1980s levels of union coverage had prevailed in 1991.

Column (c) just proves the results in column (b) by showing no trade union effect on skill premiums of semi-skilled workers. Moreover, the tax wedge only affects the skill premium of semi-skilled workers, but not for that of high skilled workers, which is also reasonable and consistent with Brewer et al (2008).

As far as special periods are concerned, column (d) and (e) show that the effect of trade union is much more prominent in the years after 1980 (-1.10) than in the 1970s (-0.23), and only changes in trade union density of unskilled workers are important. Nevertheless, the growth rate of trade union density in unskilled workers also account for the decline of skill premium during the 1970s (-0.26). This result implies that when union density is high as in the 1970s, wages are less sensitive to changes of union density level but more sensitive to the speed of union density changes. However, when union density is low as in years after 1979, wages become more sensitive to changes of union density level.

We only find that minimum wages are significantly negative for skill premiums of high skilled workers in column (b). This result implies that the NMWA only increases wages of unskilled/semi-skilled workers. Hence, skill premiums of high skilled workers as well as the overall earnings inequality are reduced by the NMW, while skill premiums of semi-skilled workers are not affected by the NMW. This result is reasonable since only unskilled/semi-skilled workers at the lower part of wage distribution are likely to benefit form the NWM.

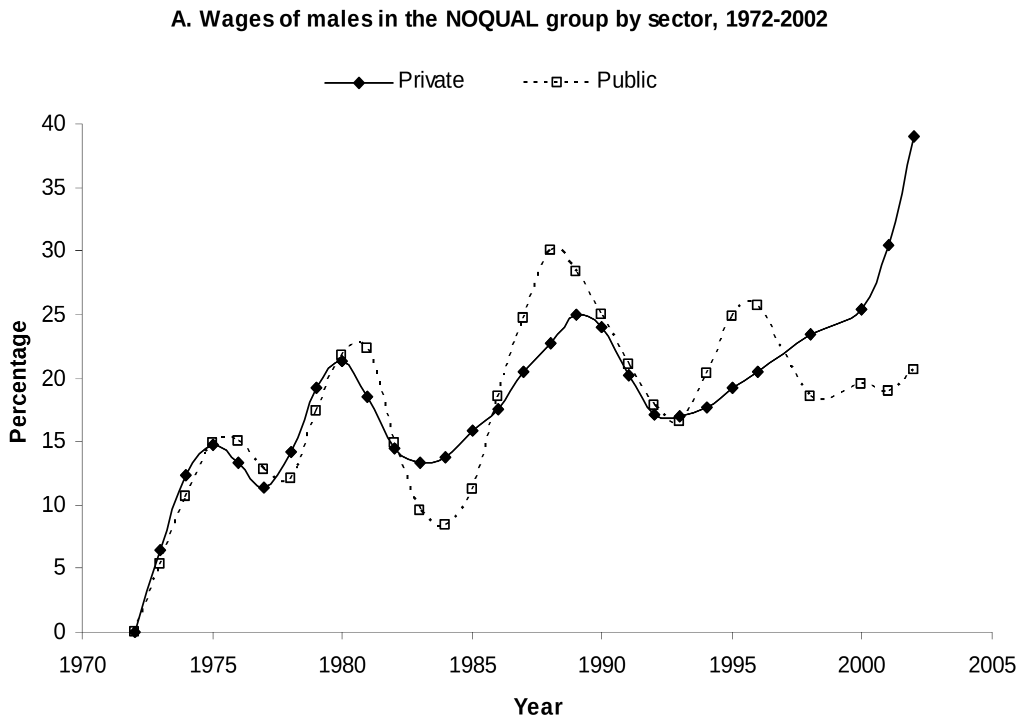
6. Conclusions

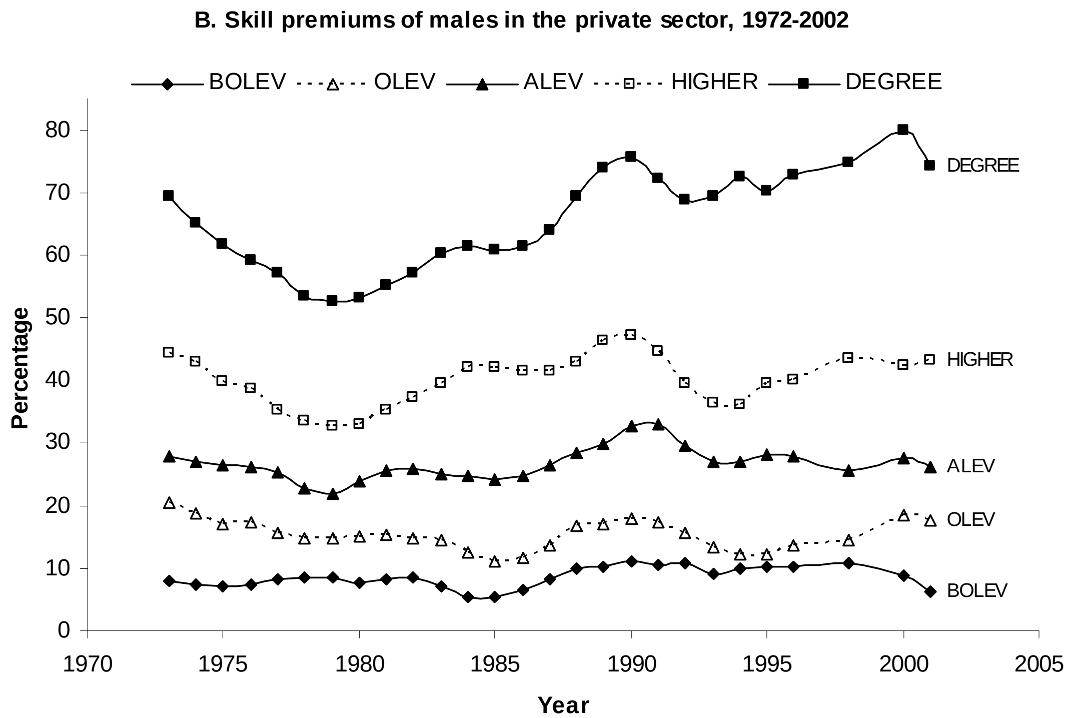
This paper analyzes the link between institutions and the skill premium in the UK controlling for other explanatory variables such as market conditions, international trade and skill-biased technology. We find the institutional factors such as trade union, the tax and benefit system and the national minimum wages are very important for skill premiums and earnings inequality.

For the skill premium in the private sector, institutions are more important for the unskilled baseline group than the skilled groups. The trade union decline after 1979 brings different effect on wages of skilled and unskilled workers and pushes the skill premium up. By using the fixed effect ECM model, we find that the trade union decline in unskilled workers can explain about one third degree premium increase over the period 1979-1998. The overall effect of trade union in all workers can explain about 13.34 percent of degree premium increase in the same period. Trade union effect is higher for skill premiums of high skilled workers than that of semi-skilled workers and higher in years after 1979 than in the 1970s. Our results are insensitive to skill group categorization.

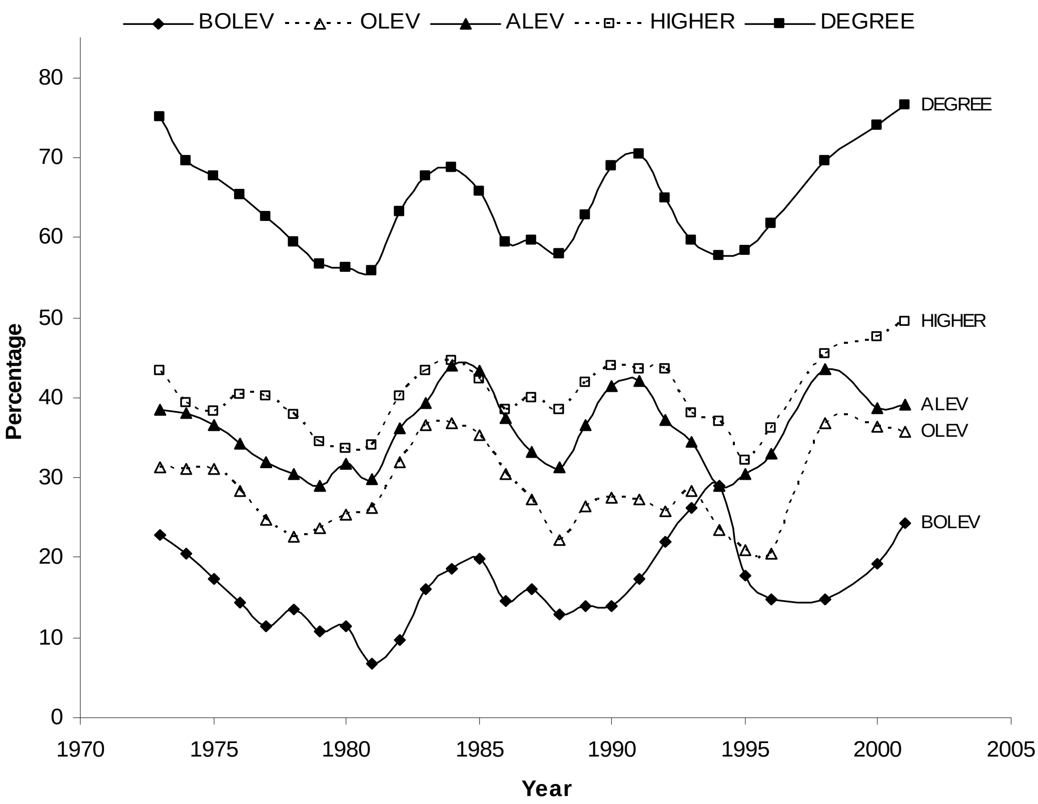
Moreover, we find that minimum wages can only decrease skill premiums of high skilled workers rather than low skilled workers, which is reasonable. The increase of unemployment benefit over the period 1979-1998 reduces the increasing the skill premium by about 13.6 percent. But, the mark-up effect of taxation increase skill premium by about 8.26 percent. We find no significant associations between above institutions and skill premiums in the public sector.

Figure 1: Relative wages of the baseline group and skill premiums by sector, estimates from equation (2)



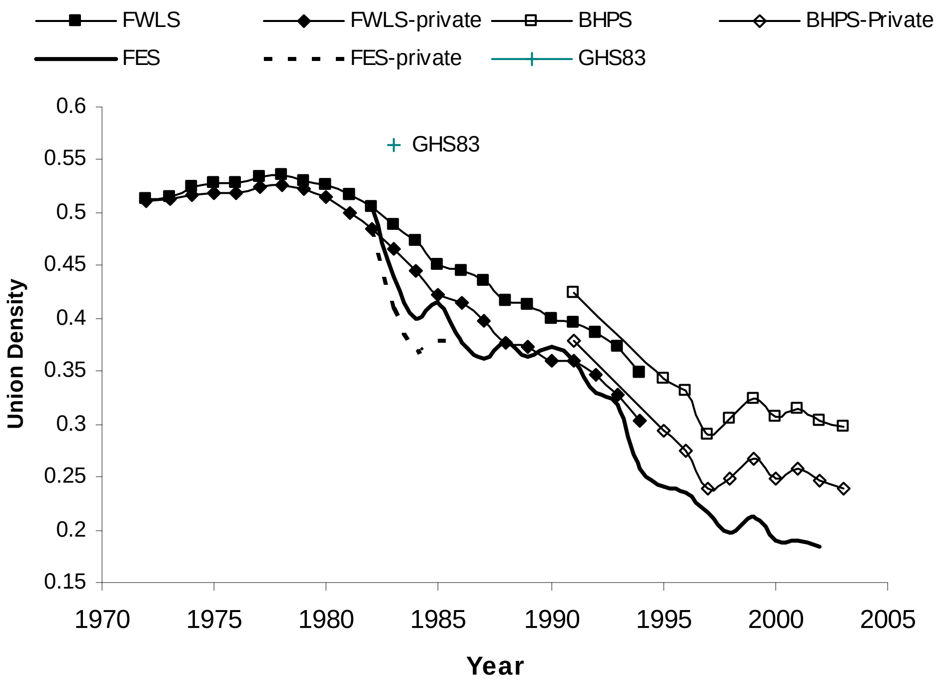


C. Skill premiums of males in the public sector, 1972-2002



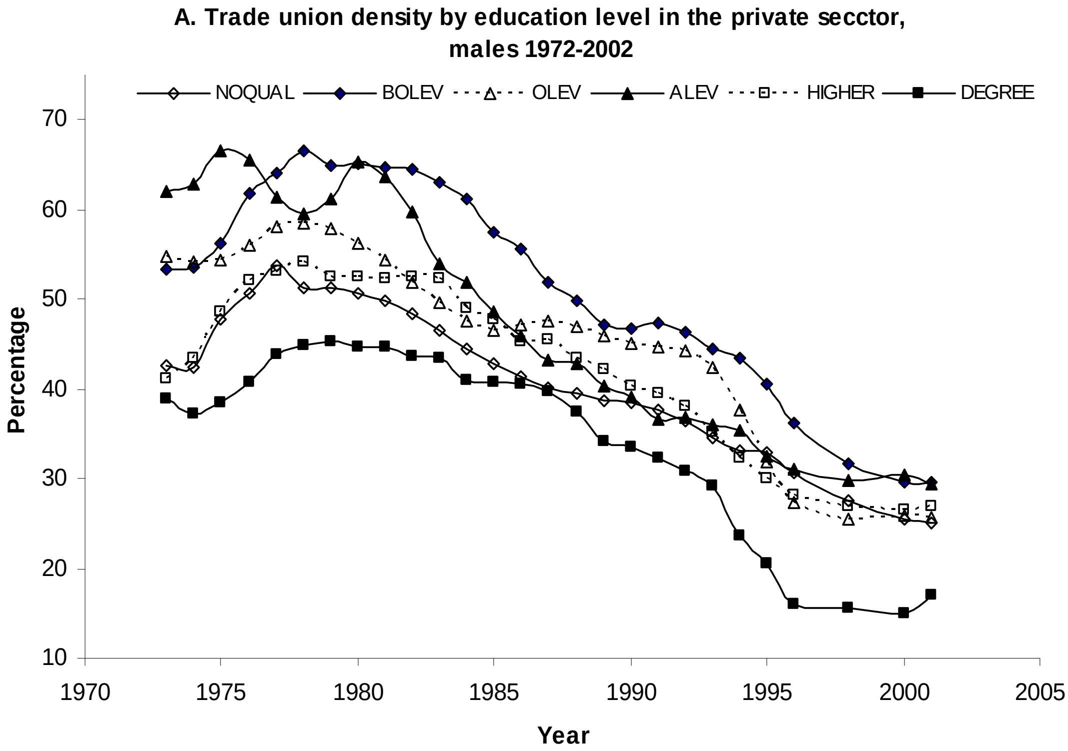
Note: All results are calculated from equation (2) by private/public sector using the GHS 1972-2002. There are 117,302 workers in the private sector and 23,323 workers in the public sector. Wages are deflated based on 1995 pounds. Wage samples include only male full-time (weekly working hours >35) workers aged 16-66 years who were not self-employed. In order to smooth out the trend, the 3-year moving averages are presented.

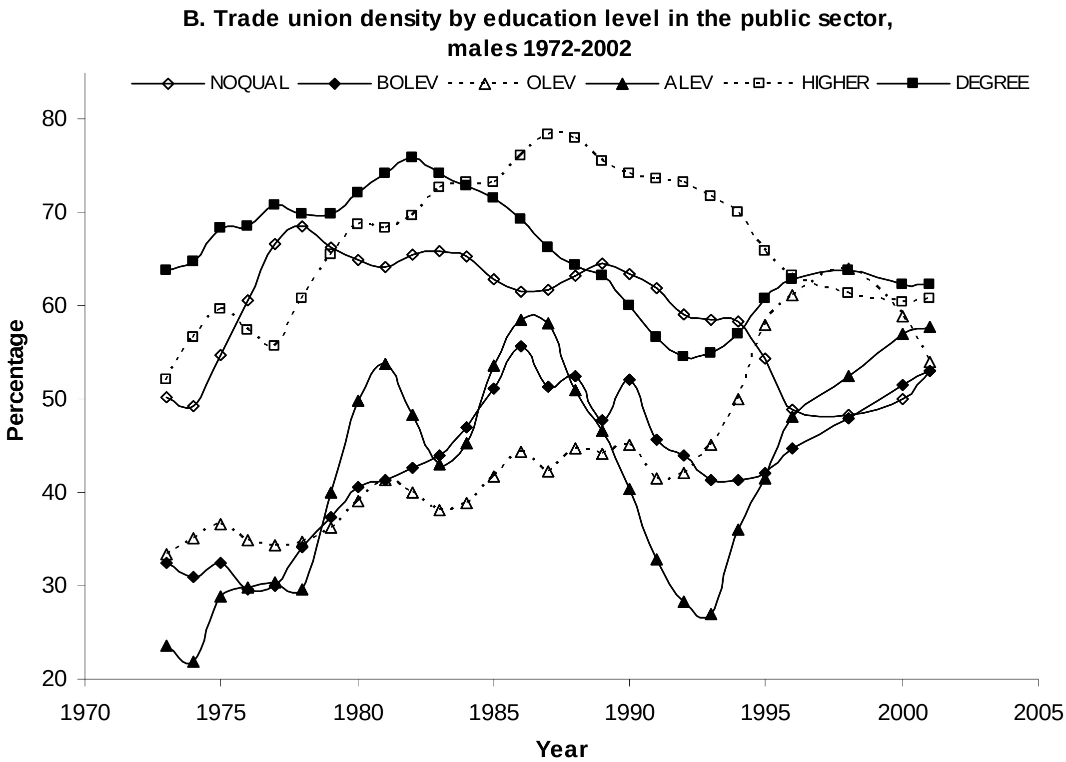
Figure 2: Trade union density in the UK, males 1972-2002



Note: All figures are calculated from the GHS 1983, the FES 1982-2002, the FWLS 1994/1995 and the BHPS 1991-2002.

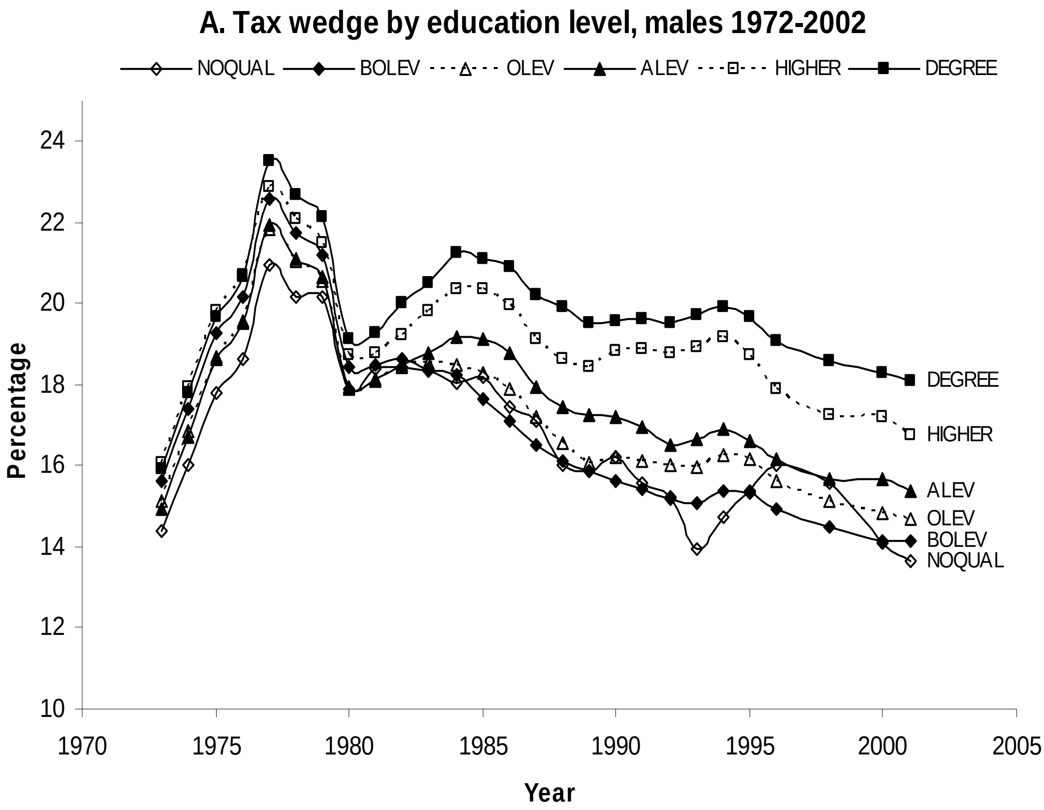
Figure 3: Trade union density in the UK by education level and sector, males 1972-2002

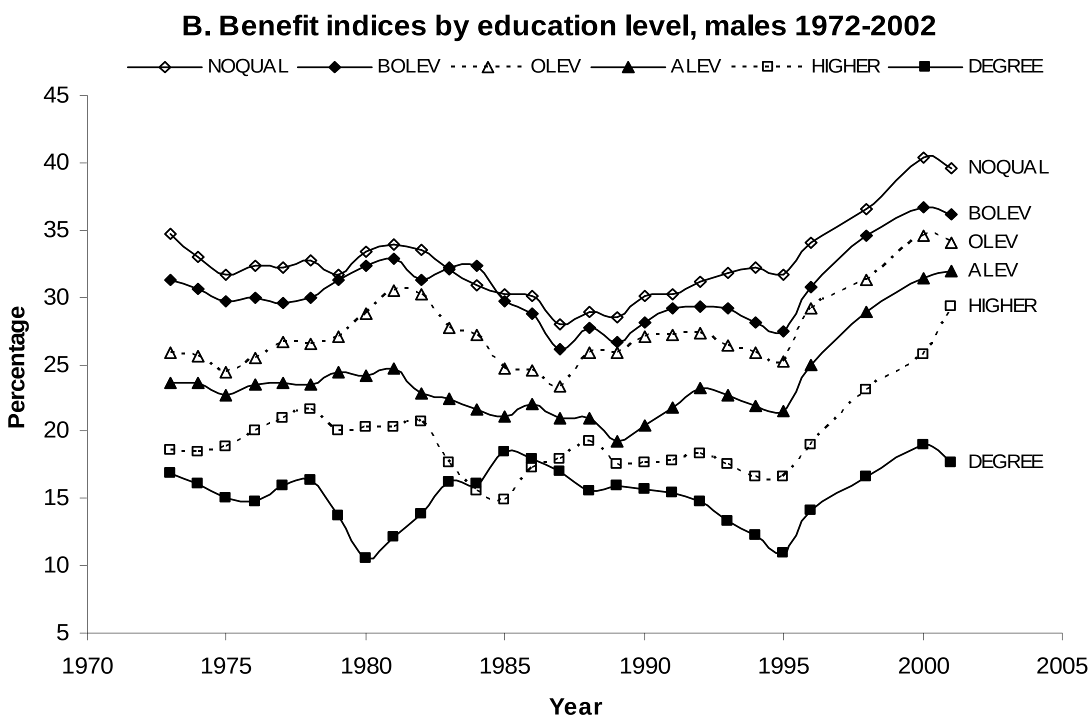




Note: All figures are calculated from the combined dataset of the FWLS 1994/1995 and the BHPS 1991-2002.

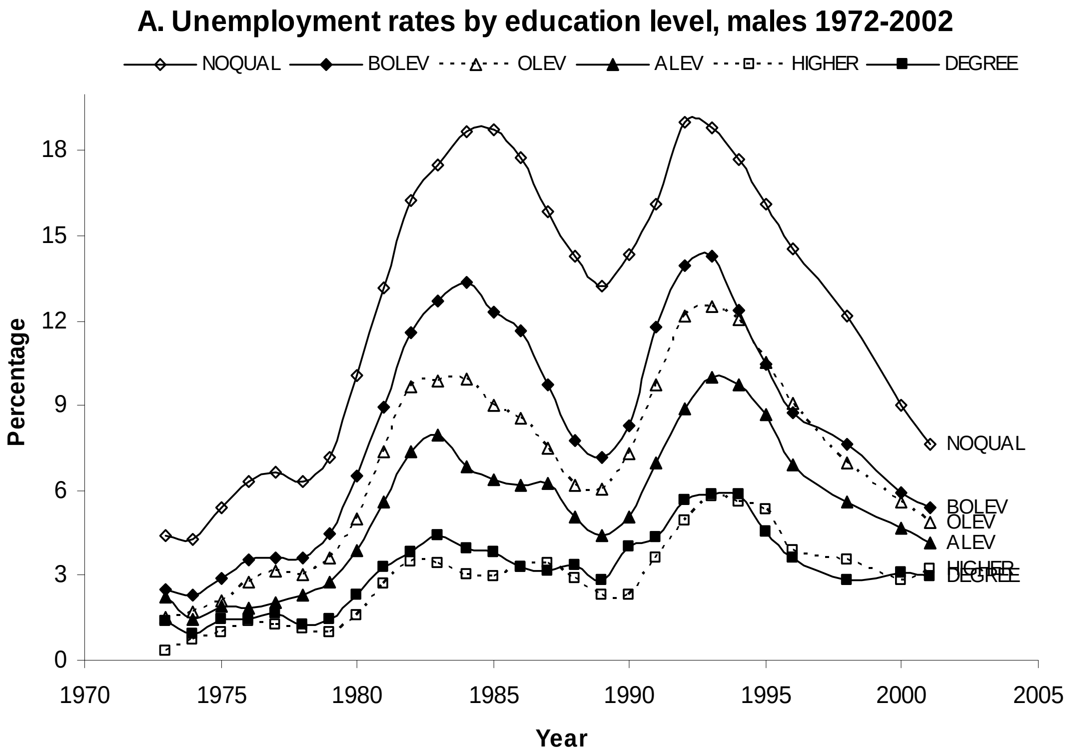
Figure 4: The tax and benefit system in the UK by education level, males 1972-2002

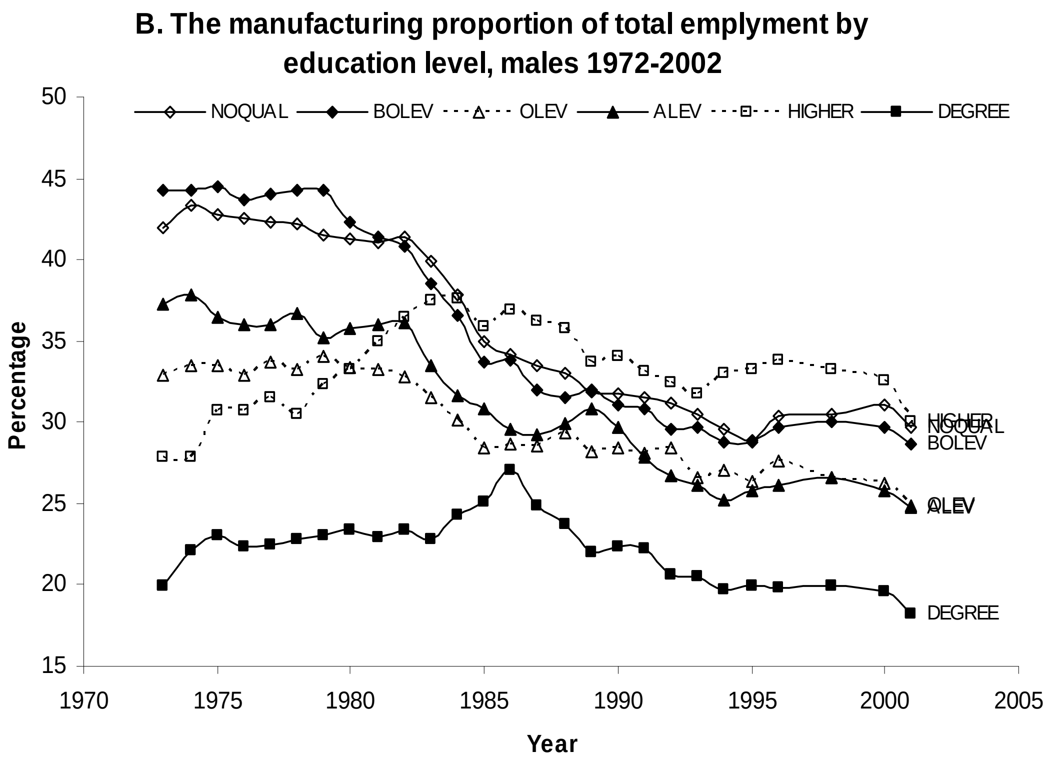


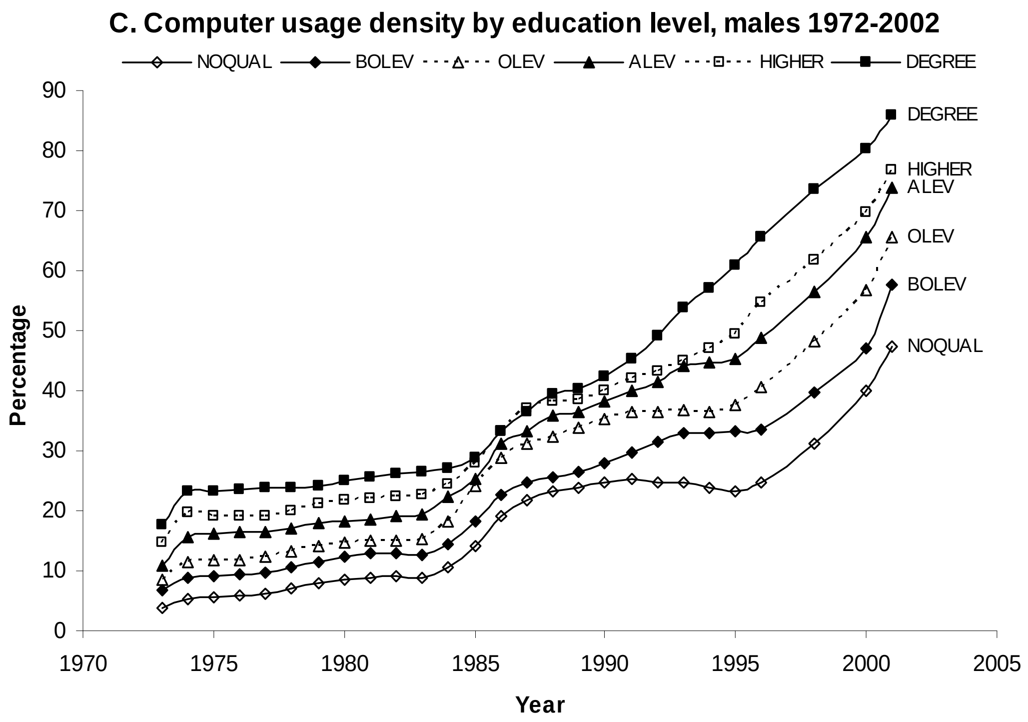


Note: All figures are calculated from the GHS 1972-2002 and the FES 1972-2002.

Figure 5: Unemployment rates, manufacturing ratio and computer usage in the UK by education level, males 1972-2002







Note: All figures are calculated from the GHS 1972-2002.

Table 1: Institutions and the skill premium, male 1972-2002, estimation from equation (3)

Dependent variable: the skill premium (s_{jt}, j = b, o, a, h and d)	Private	Public
Trade union density (tud_{jt})	0.19**	-0.01
	(0.10)	***(0.07)***
Trade union density (tud_{nt})	-0.29*	0.32
	(0.19)	***(0.92)***
Tax wedge (tax_{jt})	2.03***	0.36
	(0.83)	***(1.27)***
Tax wedge (tax_{nt})	-2.96***	-1.72
	(0.87)	***(1.43)***
Benefit index ($repr_{jt}$)	-0.12	-0.25
	(0.17)	***(0.23)***
Benefit index ($repr_{nt}$)	-0.50*	-0.11
	(0.29)	***(0.66)***
Unemployment rate (u_{jt})	-0.28	0.67
	(0.32)	***(0.47)***
Unemployment rate (u_{nt})	-0.06	-1.39
	(0.26)	***(1.40)***
Manufacturing proportion (ind_{jt})	-0.17	0.02
	(0.16)	***(0.21)***
Manufacturing proportion (ind_{nt})	0.24	-0.64
	(0.57)	***(4.99)***
Computer usage ($comp_{jt}$)	0.35***	-0.03
	(0.13)	***(0.21)***
Computer usage ($comp_{nt}$)	-0.51**	-0.59
	(0.24)	***(2.64)***
Minimum wages (MW)	-0.60	18.93
	(3.84)	***(34.44)***
Observations	140	140
Groups	5	5
R^2 (within)	0.59	0.65
Group dummies	Yes	Yes
Year dummies	Yes	Yes

Notes: Estimated standard errors are under the coefficients. ***, ** and * denote significance at 1%, 5% and 10% levels for two-tail tests. Following the wage cyclicality literature (see Peng and Siebert 2007 and 2008 and reference therein), unemployment rates and unemployment benefit index have actually one year lag. Hence, we have only 140 observations. We use Stata's fixed effect programme (*xtreg*, *fe*, see STATA 2003b) to estimate equation (3).

Table 2: Institutions and the skill premium, male 1972-2002, estimation from equation (4)

Dependent variable: growth rate of the skill premium (ds_{jt}, j = b, o, a, h and d)	Private	Public
Trade union density (tud_{jt-1})	0.19*	-0.01
	(0.11)	***(0.08)***
Trade union density ($dtud_{jt}$)	0.10	-0.04
	(0.12)	***(0.09)***
Trade union density (tud_{nt-1})	-0.40**	0.53
	(0.18)	***(1.45)***
Trade union density ($dtud_{nt}$)	0.01	1.21
	(0.16)	***(0.79)***
s_{jt-1}	-0.68***	-1.02***
	(0.09)	***(0.11)***
Tax wedge (tax_{jt})	1.61**	1.04
	(0.83)	***(1.32)***
Tax wedge (tax_{nt})	-2.31***	-2.23
	(0.90)	***(2.10)***
Benefit index ($repr_{jt}$)	0.12	-0.17
	(0.16)	***(0.24)***
Benefit index ($repr_{nt}$)	-0.77***	0.08
	(0.28)	***(0.75)***
Unemployment rate (u_{jt})	-0.21	1.23***
	(0.33)	***(0.49)***
Unemployment rate (u_{nt})	0.44*	0.28
	(0.25)	***(1.53)***
Manufacturing proportion (ind_{jt})	-0.12	0.00
	(0.16)	***(0.23)***
Manufacturing proportion (ind_{nt})	0.50	-0.70
	(0.57)	***(4.39)***
Computer usage ($comp_{jt}$)	0.29**	-0.02
	(0.12)	***(0.22)***
Computer usage ($comp_{nt}$)	-0.46**	0.28
	(0.22)	***(1.53)***
Minimum wages (MW)	2.13	14.45
	(3.21)	***(37.24)***
Observations	135	135
Groups	5	5
R^2 (within)	0.61	0.82
Group dummies	Yes	Yes
Year dummies	Yes	Yes

Notes: Estimated standard errors are under the coefficients. ***, ** and * denote significance at 1%, 5% and 10% levels for two-tail tests. As noted in Table 1, unemployment rates and unemployment benefit index have actually two-year lags here. Hence, we have only 135 observations. We use Stata's fixed effect programme (*xtreg*, *fe*, see STATA 2003b) to estimate equation (4).

Table 3: Contribution of explanatory factors to the degree premium, the private sector

	1972-1979	1979-1998	1998-2002
a. Changes of each variable, percentage			
Degree premium (s_{dt})	-20.19	32.05	-17.55
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 ($comp_{dt}$)	19.95	48.23	16.57
Computer usage ($comp_{nt}$)	5.86	20.35	22.49
b. Effects of changes in each explanatory variable, percentage			
Trade union density (tud_{dt})	0.87*	-5.84*	1.10*
Trade union density (tud_{nt})	-3.36**	10.12**	0.52**
Tax wedge (tax_{dt})	6.17**	-0.58**	-0.88**
Tax wedge (tax_{nt})	-7.85***	3.23***	3.04***
Benefit index ($repr_{dt}$)	-0.31	0.68	-0.48
Benefit index ($repr_{nt}$)	2.62***	-5.04***	1.64***
Unemployment rate (u_{dt})	0.15	-0.45	0.09
Unemployment rate (unt)	0.46*	2.46*	-1.89*
Manufacturing proportion (ind_{dt})	-0.93	0.45	0.50
Manufacturing proportion (ind_{nt})	0.19	-3.35	-2.21
Computer usage ($comp_{dt}$)	5.78**	13.99**	4.80**
Computer usage ($comp_{nt}$)	-2.70**	-9.36**	-10.35**
c. Overall contribution of each factor, percentage			
Trade union density	12.38	13.34	-9.20
Tax wedge	8.32	8.26	-12.33
Benefit index	-11.41	-13.60	-6.61
Unemployment rate	-3.04	6.26	10.25
Manufacturing proportion	3.66	-9.06	9.74
Computer usage	-15.29	14.44	31.57

Notes: All figures in Table 3 are calculated using estimates in Table 2. ***, ** and * denote significance at 1%, 5% and 10% levels for two-tail tests. Significance of each variable in the middle panel is from Table 2.

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

Dependent variable: growth rate of the skill premium (ds_{jt}, j = b, o, a, h and d)	**(a) Three groups**	**(b) High skilled**	**(c) Semi-skilled**	**(d) 1972-1979**	**(e) 1980-2002**
Trade union density (tud_{jt-1})	0.25*	1.02*	0.03	-0.16	0.13
	(0.18)	***(0.63)***	***(0.12)***	***(0.17)***	***(0.16)***
Trade union density ($dtud_{jt}$)	0.01	0.46	0.03	-0.21	0.06
	(0.14)	***(0.45)***	***(0.13)***	***(0.20)***	***(0.15)***
Trade union density (tud_{nt-1})	-0.40**	-1.54**	-0.18	-0.23*	-1.10*
	(0.23)	***(0.74)***	***(0.18)***	***(0.15)***	***(0.67)***
Trade union density ($dtud_{nt}$)	-0.19	-0.77	0.04	-0.26*	0.06
	(0.25)	***(0.55)***	***(0.16)***	***(0.15)***	***(0.32)***
s_{jt-1}	-0.62***	-1.14***	-0.80***	-0.58***	-0.73***
	(0.13)	***(0.23)***	***(0.15)***	***(0.18)***	***(0.11)***
Tax wedge (tax_{jt})	0.84	1.19	2.37**	0.46	1.23
	(0.76)	***(2.10)***	***(1.28)***	***(1.85)***	***(1.13)***
Tax wedge (tax_{nt})	-0.95	-1.80	-2.61**	0.19	-0.24
	(0.91)	***(1.98)***	***(1.34)***	***(1.97)***	***(1.07)***
Benefit index ($repr_{jt}$)	-0.47	0.41	-0.25	-0.38	0.14
	(0.33)	***(0.33)***	***(0.27)***	***(0.59)***	***(0.18)***
Benefit index ($repr_{nt}$)	0.54	-1.05	-0.11	-	-0.77
	(0.40)	***(0.98)***	***(0.29)***	-	***(0.58)***
Minimum wages (MW)	0.00	-19.10*	1.61	-	-
	(0.07)	***(10.70)***	***(3.94)***	-	-
Observations	58	54	81	30	105
Groups	2	2	3	5	5
R^2 (within)	0.94	0.83	0.73	0.83	0.60
Group dummies	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes

Notes: as for Table 2.

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