1. Introduction

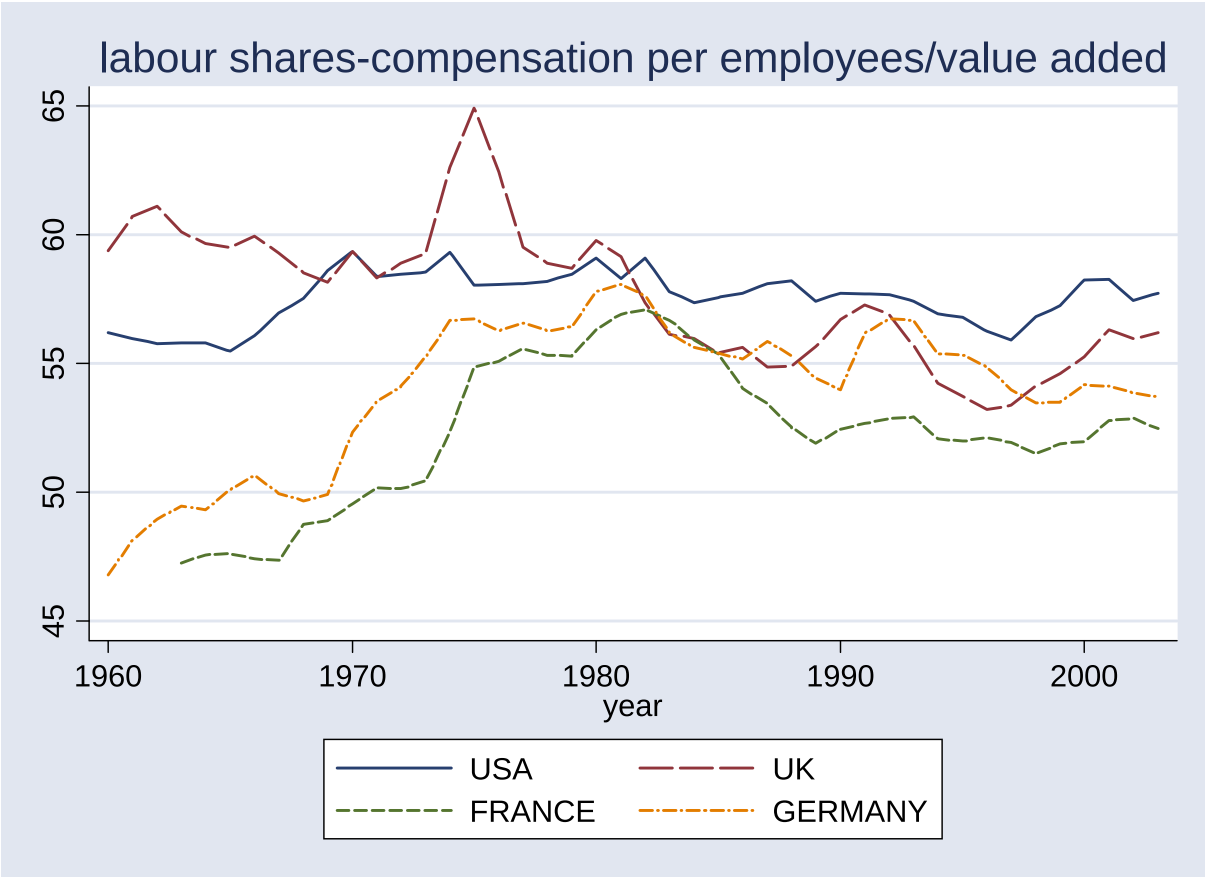
Over the past two decades a large literature has sought to understand the evolution of wage inequality in OECD countries; see, for example, Gottschalk and Smeeding (1997). Different patterns across countries have been documented, and labour market institutions have been shown to play an important role in determining the distribution of wages. In this paper we try to understand which are the determinants of differences in *income* inequality across countries and over time in OECD countries. In doing so, we focus on two aspects largely neglected by the literature. The first one is the role of factor shares as determinants of the personal distribution of income. We argue that wage inequality is only one of the components of personal income inequality, and that both the labour share and unemployment play an important role. The second aspect consists of understanding the impact of labour market institutions on overall income inequality, as opposed to only on relative wages.

Contrary to the textbook approach in macroeconomics where factor shares are taken to be constant, variations in the labour share across countries and over time are large. Figure 1 illustrates the recent experiences of the US, the UK, Germany and France over the period 1960 to 2002. The US has the most stable labour share, which fluctuates between 55 and 59 per cent. France and Germany have, for most of the period, a lower labour share than the two Anglo-Saxon countries, and exhibit a hump-shaped pattern with the labour share increasing up to around 1981 and declining thereafter; while the UK has experienced a decline over the period.[1] Despite a substantial reduction in the differences between those four countries, in 2002 their labour shares ranged from 53% in France to 58% in the US.

The first question we address in this paper is whether these differences in the factor distribution of income can help us explain variations in the distribution of personal incomes in OECD countries over the past thirty years. Indeed, recent work by Piketty (2001, 2003) and Piketty and Saez (2003) has emphasised the importance of capital income for the highest income groups even in recent times, and Atkinson (2003) has suggested that the increase in inequality that took place in a number of OECD countries during the 1980s was in part due to the rise in the return to capital. Our second concern is the impact of labour market institutions. The effect of institutions on relative wages has been well documented. For example, stronger unions tend to compress the wage distribution, which in turn would tend to reduce income inequality. However, these institutions also affect the unemployment rate and, potentially, the labour share, and hence will affect the distribution of income through channels other than wages.

[1] The notable exception is the sharp increase and subsequent fall of the labour share during the labour government of 1974-1976, after a period of major conflict between unions and the conservative government of Heath.

Figure 1 – Compensation of employees over GDP – 1960-2003



We start by presenting the theoretical framework. First, we consider how labour market outcomes –that is, the relative wage, the labour share and the unemployment rate- are determined in an economy with non-competitive labour markets. Crucially for our purposes, we suppose the aggregate production function is CES, implying that the labour share is not constant but rather depends on factor inputs. Since labour markets are not competitive, labour market institutions, by affecting employment levels, become an essential determinant of the labour share. We capture this idea with a model with two types of workers. Skilled workers are subject to efficiency wage considerations, which imply no market clearing. For the unskilled the wage and employment determination process is the outcome of wage bargaining between a union that represents unskilled workers and a firm in a right-to-manage framework. We find that the equilibrium employment levels are a function of union bargaining power, the unemployment benefit, and the capital-labour ratio. The bargained levels of employment and wages will in turn determine the overall labour share, wage ratio, and unemployment rate, making them a function of labour market institutions.

The second step is to decompose the Gini coefficient of the distribution of personal incomes in a model economy. Our highly stylised set up considers four types of agents. The first are the jobless who receive the unemployment benefit. The second are unskilled workers who receive the unskilled wage. Lastly, there are skilled workers, which may own capital or not. Those who do not will simply receive the skilled wage, while those who do (the worker-capitalists) receive both the skilled wage and profits. There are then three sources of inequality: employment versus unemployment, skilled versus unskilled wages, and the distribution of capital. In fact, the Gini index for personal incomes can be expressed as a function of the labour share, the relative wage, the unemployment benefit, and the proportion of the population in each category. A smaller labour share, a higher relative wage, and a lower unemployment benefit, all increase income inequality. This decomposition implies that the effect of labour market institutions on inequality is ambiguous. For example, both higher union power and unemployment benefits increase the unemployment rate, which tends to raise the Gini coefficient, but reduce the relative wage and increase the labour share, both of which tend to lower inequality.

We test these propositions in a panel of OECD countries for the period 1970-96. We find that the labour share remains a fundamental aspect of overall inequality patterns, with an effect roughly as important as that of relative wages. Our results also show that that stronger unions and a more generous unemployment benefit tend to reduce income inequality. The effect of labour market institutions tends to be large, and explains a substantial fraction of the variation across countries. The other variable that emerges from our analysis as having a large impact is the capital-labour ratio. High capital-labour ratios tend to increase the labour share, and hence reduce income inequality. In fact, this appears to have been a major force dampening the increase of income inequality in the US over the last few decades.

The paper adds to the recent revival of interest in the factors shaping the distributions income across countries (Bourguignon and Morrisson, 1990, 1998; Li, Squire, and Zou, 1998; Barro, 2000; Alderson and Nielsen, 2002; Breen and García Peñalosa, 2004). For decades empirical work on cross-country differences in the distribution of income consisted of tests of the Kuznets hypothesis taking the form of regressions of inequality on the level of GDP and its square. Only recently have variables other than the level of income been considered, such as the level of human capital, the degree of democratisation, or the extent of financial development. Although this approach is helpful in understanding the underlying causes of inequality, it leaves little room for policy recommendations as in most cases the particular mechanism through which these variables impact inequality is not understood. By focussing on the basic determinants of the distribution of income we want to understand whether labour market institutions play a role because they affect the unemployment rate, the distribution of wages, or the way in which capital and labour are rewarded.

The paper is also related to the literature on the evolution of inequality in industrial economies over the past three decades. Two features have dominated this literature. One has been the increase in income inequality in a number of countries; the other the sharp rise in the relative wages in the UK and the US (Atkinson, 1997, 2003; Gottschalk and Smeeding, 1997; Bound and Johnson, 1992; Juhn, Murphy, and Brooks, 1993). Our paper emphasises two aspects. First, that although wage inequality is a crucial aspect of the income distribution, the distribution of wealth still plays a substantial role as captured by the negative impact of the labour share in our regressions for the Gini coefficient. Second, our analysis highlights the differences between an increase in the relative wage and in wage inequality. Understanding the evolution of inequality requires knowing the proportions of agents receiving each salary and not only the relative salaries, and looking at the labour share is a (crude) way of capturing both.

A number of recent papers have been concerned with the labour share. The focus of these works has been to understand the determinants of either the evolution of the labour share over time in OECD, or cross-country differences (Blanchard, Nordhaus, and Phelps, 1997; Rodrik, 1999; de Serres, Scarpetta and de la Maisonneuve, 2002; Bentolila and Saint-Paul, 2003). We present a different perspective, trying to understand not the determinants but the effects of differences in the rewards to capital and labour across countries and over time.

The paper is organised as follows. Section 2 presents our theoretical model. Section 3 presents the data and our results. We then perform a number of simulation exercises. Section 4 concludes.

2. Theoretical considerations

2.1. The determinants of the relative wage and the labour share

2.1.1. Technological determinants

We consider an economy with three inputs, capital, denoted by K, skilled workers, H, and unskilled workers, L. Output is produced according to a constant return to scale production function $Y = F(K, L, H)$. As is well known, a Cobb-Douglas production function implies constant labour and capital shares. In order to explain observed variations in labour shares, a more general production function is needed. We assume that output is produced with a CES production function using capital, *K*, and a "labour aggregate". Production is a CES function of K and the labour aggregate, which is in turn a Cobb-Douglas function of skilled and unskilled labour. That is, output is produced according to[2]

$$Y = \left[\alpha K^{-\sigma} + (1-\alpha)\left(H^{\beta} L^{1-\beta}\right)^{-\sigma}\right]^{-1/\sigma} \qquad \text{with } -1 \leq \sigma < \infty,\ 0 < \alpha < 1, 0 < \beta < 1 \qquad (1)$$

This production function allows for different degrees of substitutability across factors. The elasticity of substitution between skilled and unskilled labour is 1, while that between capital and the labour aggregate is $1/(1+\sigma)$. For $\sigma = 0$ the production function would be Cobb-Douglas in the three inputs. In line with existing evidence,[3] we assume that the elasticity of substitution between capital and the labour aggregate is less than one, which requires $\sigma > 0$.

Differentiating the production function we obtain factor demand functions,

$$r = \alpha\left(\alpha + (1-\alpha)x^{-\sigma}\right)^{-(1+\sigma)/\sigma} \qquad (2a)$$

$$w_u = (1-\beta)(1-\alpha)\left(\alpha + (1-\alpha)x^{\sigma}\right)^{-(1+\sigma)/\sigma} x^{\sigma} \frac{K}{L} \qquad (2b)$$

$$w_s = \beta(1-\alpha)\left(\alpha + (1-\alpha)x^{\sigma}\right)^{-(1+\sigma)/\sigma} x^{\sigma} \frac{K}{H} \qquad (2c)$$

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[2] This formulation has been proposed, for example, by Katz and Murphy (1992).

[3] This is consistent with the evidence reported in Hamermesh (1993), Rowthorn (1999), Krusell, Ohanian, Rios-Rull, and Violante (2000), and Antras (2004).

where r is the interest rate, w_s and w_u are respectively the (gross) skilled and unskilled wages, and $x \equiv K / H^{\beta} L^{1-\beta}$.

The labour share, denoted θ, is defined as the ratio of total employee compensation to value added. With two types of workers this is simply

$$\theta \equiv \frac{w_s H + w_u L}{Y} \tag{3}$$

Defining the relative wage as $\omega \equiv w_s / w_u$, and using equations (2) we obtain the inverse relative demand for labour and the labour share as

$$\omega = \frac{\beta}{1-\beta}\frac{L}{H} = \frac{1}{1-\beta}\cdot\frac{1}{h} \tag{4}$$

$$\theta = \frac{(1-\alpha)}{1-\alpha+\alpha x^{-\sigma}} = \left[1+\frac{\alpha}{1-\alpha}\left(k\frac{1+h}{h^{\beta}}\right)^{-\sigma}\right]^{-1} \tag{5}$$

where $k = K/(H+L)$ is the capital-labour ratio and $h = H/L$ the relative skilled employment. The labour share and the relative demand for labour hence depend on the capital-labour ratio and the relative employment ratio, that is, $\omega = \omega(h)$ and $\theta = \theta(k,h)$. The comparative statics are straight forward, with

$$\frac{\partial\omega}{\partial h} < 0,$$

$$sign\left[\frac{\partial\theta}{\partial x}\right] = sign[\sigma], \quad sign\left[\frac{\partial\theta}{\partial h}\right] = -sign[\sigma(\omega-1)], \qquad sign\left[\frac{\partial\theta}{\partial k}\right] = sign[\sigma].$$

A higher relative employment ratio reduces the relative wage, while the impact of the capital-labour ratio and relative employment on the labour share depends on the elasticity of substitution. For $\sigma = 0$, the labour share is simply $\theta = 1-\alpha$, and neither k nor h will affect it. Our assumption of $\sigma > 0$ and supposing, reasonably, that $\omega > 1$, we have $\partial\theta/\partial k > 0$ and $\partial\theta/\partial h < 0$. That is, a higher capital-labour ratio will increase the labour share, while greater relative skilled employment will reduce both the labour share and the relative wage.[4]

2.1.2. Institutional determinants

If labour markets were competitive, (4) and (5) would imply that a country's capital-labour ratio and its relative supply of skills would be the sole determinants of the labour share and the relative wage. However, labour markets are not competitive. Employment levels hence differ from factor supplies, and anything that affects employment would in turn affect θ and ω. In order to understand which are the

[4] The case where $\sigma < 0$ is discussed in Appendix I.

potential determinants of these variables we examine wage and employment determination with two types of labour.

We assume that wages for the two types of workers are determined in different ways. For skilled workers, we suppose that imperfect information on the part of the firm about whether or not employees are shirking forces the former to pay wages above the market clearing level, which in turn lead to unemployment, as in the efficiency wage model of Shapiro and Stiglitz (1985). For unskilled workers, we model the wage and employment determination process as the outcome of wage bargaining between a single union and a single firm in a right-to-manage framework. The union bargains over unskilled wages with the firm, and then the latter sets employment. [5]

Efficiency wages for skilled workers

Consider a simple, one-period efficiency wage model. Suppose skilled agents receive a net wage $\tilde{w}_s = (1-\tau)w_s$, where τ is the tax wedge paid to the government as employer and employee contributions. Workers are assumed to be risk-averse with utility $U(w_i) = w_i^{\rho}$, with $0 < \rho \leq 1$. Then, the utility of shirking is simply $U^S = (1-p)\left((1-\tau)w_s\right)^{\rho} + pB^{\rho}$ and that of not-shirking $U^N = ((1-\tau)w_s - e)^{\rho}$, where p is the probability of being caught if shirking, B is the unemployment benefit (or the monetary equivalent of leisure if the latter is unavailable), and e is the monetary cost of effort.[6] The resulting efficiency wage, $\overline{w}_s$, is given by the solution to

$$((1-\tau)\overline{w}_s - e)^{\rho} = (1-p)\left((1-\tau)\overline{w}_s\right)^{\rho} + pB^{\rho} \tag{6}$$

Simple differentiation shows that $\overline{w}_s$ is increasing in B and e, and decreasing in p. Given $\overline{w}_s$ and the level of unskilled employment, the inverse demand for skilled labour, equation (2c), determines skilled employment, H.

Union bargaining and the unskilled wage

Consider now the determination of the unskilled wage and employment level. We assume that the union represents only the unskilled, and that it has a utilitarian utility function of the form

$$V = \frac{1}{\overline{L}}\left[LU(\tilde{w}_u) + \left(\overline{L} - L\right)U(B)\right] \tag{7}$$

where $\overline{L}$ is the unskilled labour force, $U(.)$ is the workers' utility function, $U(\tilde{w}_i) = \tilde{w}_i^{\rho}$, and the net wage is given by $\tilde{w}_u = (1-\tau)w_u$. The bargaining process is then governed by

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[5] All our results can be obtained if instead of having efficiency wages for the skilled we had a single union that represented both types of workers and bargained simultaneously over both wages.

[6] It would be straightforward to allow for the dynamic flows into and out of employment. For simplicity, we assume here that labour markets are separated by skills, such that an unemployed skilled worker cannot work as unskilled. Note that B goes untaxed, as in most institutional set-ups.

$$\max_{w_u}\left(L\left[((1-\tau)w_u)^{\rho}-B^{\rho}\right]\right)^{\gamma}\left(Y-w_uL-w_sH\right)^{1-\gamma} \tag{8}$$

The bargaining solution is obtained by maximising this expression with respect to w_u, taking into account the fact that, for a given skilled wage, changing the unskilled wage affects both skilled and unskilled employment. The resulting first-order conditions can be expressed as (see Appendix I),

$$\rho(1-\tau)^{\rho}=\left(\frac{1-\gamma}{\gamma}(1-\beta)\frac{\theta}{1-\theta}+\varepsilon_L\right)\left((1-\tau)^{\rho}-\left(\frac{B}{w_u}\right)^{\rho}\right) \tag{9}$$

where ε_L is the elasticity of the demand for unskilled labour. Since ε_L, w_u, and θ are functions of H and L, equation (9) determines w_u as a function of skilled and unskilled employment.

Equilibrium and comparative statics

The equilibrium of the model is then given by equations (2b), (2c), (6), and (9), that is, by

$$\rho(1-\tau)^{\rho}=\left(\frac{1-\gamma}{\gamma}(1-\beta)\frac{\theta}{1-\theta}+\varepsilon_L\right)\left((1-\tau)^{\rho}-\left(\frac{B}{w_u}\right)^{\rho}\right) \tag{9}$$

$$w_u=(1-\beta)(1-\alpha)\left(\alpha+(1-\alpha)x^{\sigma}\right)^{-(1+\sigma)/\sigma}x^{\sigma}\frac{K}{L} \tag{10}$$

$$\overline{w}_s=\beta(1-\alpha)\left(\alpha+(1-\alpha)x^{\sigma}\right)^{-(1+\sigma)/\sigma}x^{\sigma}\frac{K}{H} \tag{11}$$

$$\overline{w}_s=\varphi(B,e,p,\tau) \tag{12}$$

where $\varphi(B,e,p)$ is implicitly defined by (6). Together these four equations determine the equilibrium levels of skilled and unskilled employment, H and L, and the two wages as a function of model parameters: the unemployment benefit, B, the bargaining power of the union, γ, the capital stock, K, as well as the preference parameters, ρ and e, and the technological parameters, α,β,σ, and p.

Let $u\equiv 1-(L+H)/(\overline{L}+\overline{H})$ be the unemployment rate, where $\overline{H}$ is the skilled labour force. Once H and L are determined, we can obtain our three main variables of interest, the labour share, the relative wage, and the unemployment rate, which we can express as functions of the stock of capital and labour market institutions (as well as of the preference and technology parameters):

$$\theta=\theta(K,B,\gamma), \tag{13}$$

$$\omega=\omega(K,B,\gamma), \tag{14}$$

$$u=u(K,B,\gamma). \tag{15}$$

All comparative statics are derived in Appendix I. Consider first the effect of union power. It is possible to show that

$$\frac{dL}{d\gamma}<0,\quad \frac{dH}{d\gamma}<0,\quad \frac{du}{d\gamma}>0,\quad \frac{d\theta}{d\gamma}>0,\quad \frac{dh}{d\gamma}>0,\quad \frac{d\omega}{d\gamma}<0.$$

As in the standard wage bargaining model, the direct effect of greater union bargaining power is to reduce unskilled employment. This reduces the marginal product of skilled labour, and skilled employment falls in order to maintain the skilled wage at $\overline{w}_s$. Since both types of employment are reduced, the unemployment rate increases. Furthermore, under the assumption that $\sigma > 0$, the labour share also increases, the reason being that lower levels of employment result in a higher capital-labour ratio. The effect of an increase in γ on unskilled employment can be shown to be stronger than that on H, implying an increase in relative skilled employment, and hence a reduction in the relative wage.

Concerning an increase in the stock of capital, we have

$$\frac{dL}{dK}>0,\quad \frac{dH}{dK}>0,\quad \frac{du}{dK}<0.$$

A higher capital stock raises the marginal product of labour (both unskilled and skilled), leading to greater employment of both types of workers for a given wage. In the case of unskilled workers, unions react by demanding higher wages, which results in an increase in w_u. In case of skilled workers, given a constant efficiency wage, the increase in the capital stock leads to an expansion of skilled employment so as to maintain the marginal product of labour constant. Moreover, the indirect effects on L through the change in H and vice versa reinforce these direct impacts. Under reasonable conditions (see appendix), we can also show that

$$\frac{d\theta}{dK}>0,\quad \frac{dh}{dK}>0,\quad \frac{d\omega}{dK}<0.$$

A greater capital stock has a direct positive effect on θ, as a higher K increases the marginal product of labour, and indirect negative impacts through the increase in both types of employment. The positive effect dominates, implying that a greater stock of capital increases the labour share. The effect on skilled employment can also be shown to be greater than that on unskilled employment, resulting in a higher h and hence a lower relative wage.

A higher unemployment benefit has two effects. On the one hand, it increases the outside option for unskilled workers, hence unions will bargain for a higher wage and accept a lower level of employment. On the other, it increases the efficiency wages that the firm must pay to skilled workers, which requires the firm to employ fewer skilled workers in order to increase their marginal product. The reduction in H tends to reduce the marginal product of the unskilled and hence partially offsets the reduction in L. If the direct effect dominates, so that $dL/dB < 0$, it is then possible to show that

$$\frac{dH}{dB}<0,\quad \frac{du}{dB}>0,\quad \frac{d\theta}{dB}>0.$$

That is, a higher unemployment benefit reduces both skilled and unskilled employment, increasing the rate of unemployment and raising the labour share. The effect on the relative wage is ambiguous, as both the skilled and the unskilled wage increase.

2.2. The Gini coefficient in a model economy

Having established that labour market institutions affect labour shares, the relative wage, and the unemployment rate, we turn to their impact on the distribution of personal incomes. Our empirical measure of income inequality will be the Gini coefficient. We hence decompose this measure of inequality into its various components for a model economy with four types of agents.

The labour force (or population) is normalised to one, that is, $\overline{L} + \overline{H} = 1$. Following our set-up in the previous section, workers can be either employed and receive the skilled or unskilled wage, $\tilde{w}_s = (1-\tau)w_s$ and $\tilde{w}_u = (1-\tau)w_u$, or unemployed, in which case they receive the unemployment benefit B.[7] Some individuals also own capital and receive profits. We assume that the owners of capital are always skilled workers, and that they are never unemployed. Furthermore, we assume that the revenue raised from employer/employee contributions, τ, is used to finance the unemployment benefit, so that $B = \tau\theta y / u$.[8] This implies that the payment of net wages, capital income, and unemployment benefit exhaust output, and average income is equal to output per capita, y.

We then have four types of agents characterised as follows:

(i) A fraction u of the labour force are unemployed, and receive the unemployment benefit B;

(ii) A fraction l of the labour force are unskilled workers earning a net wage $\tilde{w}_u$;

(iii) A fraction s of the labour force are skilled workers. Of those $s-\kappa$ own no capital and have an income equal to the net skilled wage $\tilde{w}_s$;

(iv) There are κ worker-capitalists, each of whom earns profits π as well as the wage $\tilde{w}_s$.

Our assumptions imply that $s + l + u = 1$. We further suppose that $\tilde{w}_s > \tilde{w}_u > B$, while the profits of each worker-capitalist depend on the capital share, $\pi = (1-\theta)y/\kappa$.

The degree of income inequality is measured by the Gini concentration index computed across subgroups of population. With N subgroups, the definition of the Gini concentration index is:

$$Gini = \frac{1}{2y}\sum_{i=1}^{N}\sum_{j=1}^{N}\left|y_i - y_j\right| \cdot n_i \cdot n_j \tag{16}$$

where y_i is the income in subgroup i, which has relative weight n_i, and y is the average income. Given our assumptions about the population and their incomes, the Gini coefficient can be expressed as

[7] B can also be interpreted as a subsistence wage earned in the informal sector, if an unemployment benefit does not exists.

[8] We are implicitly assuming that profits go untaxed. The implication of this assumption is that the tax rate does not affect the Gini index directly. The alternative assumption (taxing capital income) will make τ appear in the Gini index reported in equation (17). In our dataset, the tax wedge is highly collinear with the unemployment benefit and the unemployment rate, which proved a problem when we introduced it in estimations of the Gini index.

$$Gini = (1-\kappa)(1-\theta) + ls\frac{\tilde{w}_s - \tilde{w}_u}{y} + u(1-u)\frac{\tilde{w} - B}{y} \tag{17}$$

where $\tilde{w}$ is the average net wage. The Gini coefficient is thus a function of population proportions (u, l, s), the number of capital owners κ, the capital share, the wage differential, and the unemployment subsidy. A higher capital share will increase inequality by raising profits and thus the income of the richest individuals. A greater wage differential between the skilled and the unskilled will raise the Gini coefficient as it increases inequality between groups of employed individuals, while a larger unemployment benefit will reduce the Gini coefficient. The effect of the unemployment rate is ambiguous. This is a standard effect when there is inequality within and between groups. The unemployed have a low income but are all equal, while the employed have a higher income but there is inequality within this group. More unemployment, by increasing the number of individuals in the less unequal category, may increase or reduce overall inequality. The equalising effect will, however, only dominate for unemployment rates over 50%, hence we expect higher unemployment to increase income inequality.

Our framework of analysis makes a number of simplifications, which are worth mentioning. First, both the distributions of wealth and of wages have been compressed, since we only have two types of workers (skilled/unskilled) and one type of wealth-owner. Second, two sources of income are missing. One are the rents on assets such as land or intellectual property rights and patents, which we ignore as they are a very minor fraction of the total. The other is pensions. Note, however, that pensions can come from three sources: they can be provided by pension funds, in which case they are capital income; they can be private pensions paid by a company to its former employees, in which case they are (most often) counted as labour payments in the company's balance sheet; and they can be public pensions. It is only the third component that we have not included. This could in principle be an important source of income differences;[9] however the data are rarely available. Third, we do not distinguish between personal income distribution and household income distribution.[10] Lastly, note that we have focussed on gross income inequality, with the only tax we have considered being the unemployment insurance contribution. We also model the tax rate in a naïve way, considering immediate readjustments after a change in unemployment, thanks to the balanced budget constraint; available alternatives not considered here are the lowering of the replacement rate and or a reduction in coverage (see Atkinson and Brandolini 2003).

3. Empirical Analysis

3.1. Empirical specification

We saw in equation (17) that the Gini coefficient of personal incomes could be expressed as a function of the labour share, the wage premium to skill, the replacement rate, and population shares. The model

[9] Indeed Bourguignon et al. (2002) show that a major source of differences in distribution between the US and Mexico is the level of public pensions in those two countries.

[10] Kenworthy (2003) and Esping-Andersen (2004) claim that most of the rising trend in household income inequality is attributable to changing patterns of income distribution within the family, associated with increased labour market participation of women and young people, a question beyond the scope of this paper.

identifies the determinants of employment and wages for both skilled and unskilled workers, and through them of the wage differential, the labour share, and the unemployment rate. We start by estimating these relationships. Denoting by θ_{it} the labour share, by ω_{it} the relative wage, and by u_{it} the unemployment rate for country i in year t, our strategy will consist of estimating the following relationships

$$\theta_{it} = a_0 + \underset{+}{a_1} \cdot \chi_{it} + \underset{+}{a_2} \cdot b_{it} + \underset{+}{a_3} \cdot \gamma_{it} + a_4 \cdot \mu_{it} + \delta_i + \lambda_t + \varepsilon_{it} \tag{18}$$

$$\omega_{it} = c_0 + \underset{-}{c_1} \cdot \chi_{it} + \underset{\pm}{c_2} \cdot b_{it} + \underset{-}{c_3} \cdot \gamma_{it} + c_4 \cdot \mu_{it} + \delta_i + \lambda_t + \varepsilon_{it} \tag{19}$$

$$u_{it} = d_0 + \underset{-}{d_1} \cdot \chi_{it} + \underset{+}{d_2} \cdot b_{it} + \underset{+}{d_3} \cdot \gamma_{it} + d_4 \cdot \mu_{it} + \delta_i + \lambda_t + \varepsilon_{it} \tag{20}$$

where $\chi_{it} = \log\left(\dfrac{K_{it}}{H_{it} + L_{it}}\right)$ denotes the log of capital per worker; $b_{it} = \dfrac{B_{it}}{\tilde{w}_{it}}$ is the unemployment benefit replacement rate; γ_{it} captures wage-push factors, and will be proxied by union membership rates in the labour force and by the so-called Kaitz index (the ratio between the minimum wage and the median wage); and μ_{it} captures additional country specific factors, such as the oil price, educational attainment, and the tax wedge, that have been included in previous analyses of either of these three variables. The signs reported below the coefficients to be estimated indicate our theoretical expectations.

When we move to our variable of interest, the personal distribution of income, we cannot proceed by direct estimation of equation (17). Our expression for the Gini coefficient, although an identity, captures the main components of the distribution of income. Given the distribution of agents in the economy, inequality depends on three factors, namely, the way in which total output is divided between profits and wages, the distribution of wages within the labour force, and welfare provision as captured by the unemployment benefit. If we had information on all the right-hand-side variables we could simply decompose the Gini coefficient into its various components, and examine how much wage inequality or the distribution of wealth contribute to overall income inequality. However, some of the data required, such as the distribution of wealth or the number of employed individuals at each level of education, are not available. Therefore we consider the estimation of the following relationship

$$Gini_{it} = g_0 + \underset{-}{g_1} \cdot \theta_{it} + \underset{+}{g_2} \cdot \omega_{it} + \underset{+}{g_3} \cdot u_{it} + \underset{-}{g_4} \cdot b_{it} + \delta_i + \lambda_t + def_{it} + \varepsilon_{it} \tag{21}$$

where the signs underlying the coefficient are in accordance with equation (17). We also control for different definitions used to compute the Gini index (concerning the nature of the recipient unit and the type of income taken into account) with the variable def_{it}.

The coefficient g_1 captures the relative contribution of the factor distribution of income to personal income inequality, while g_2 measures the contribution of the wage differential to overall inequality. In the highly simplified framework with workers and capitalists, g_1 could be interpreted as a measure of the between-group inequality, where groups are to be defined in accordance to their position in the production process, while the coefficients g_2, g_3, and g_4 can be interpreted as the contribution of inequality within the group of workers.

Equation (21) cannot be directly estimated, since some variables are potentially endogenous and could be correlated with unobservable and/or unmeasured variables (such as the degree of risk-aversion or the level of skilled and unskilled employment) that may also affect personal income inequality through other channels. In order two obtain an unbiased estimate of the impact of the functional distribution of income onto the personal distribution (coefficient g_1), as well as assessing the contribution of earnings inequality (coefficients g_2 and g_3), we will follow two strategies. One consists in estimating the simultaneous equation system given by equations (18), (19), (20) and (21), through three-stage least squares methods. The alternative is to instrument the potentially endogenous variables. Using the estimates obtained from equations (18), (19) and (20), we can then estimate equation (21) as

$$Gini_{it} = g_0 + \underset{-}{g_1} \cdot \hat{\theta}(\chi_{it}, b_{it}, \gamma_{it}) + \underset{+}{g_2} \cdot \hat{\omega}(\chi_{it}, b_{it}, \gamma_{it}) + \underset{+}{g_3} \cdot \hat{u}(\chi_{it}, b_{it}, \gamma_{it}) + \underset{-}{g_4} \cdot b_{it} + \delta_i + \lambda_t + def_{it} + \varepsilon_{it} \quad (22)$$

Since we are also interested in assessing the overall impact of labour market institutions on income inequality, we will estimate as well the reduced form equation obtained when we replace (18)-(20) into equation (21), which yields

$$Gini_{it} = h_0 + \underset{\pm}{h_1} \cdot \chi_{it} + \underset{\pm}{h_2} \cdot b_{it} + \underset{\pm}{h_3} \cdot \gamma_{it} + \delta_i + \lambda_t + def_{it} + \varepsilon_{it} \quad (23)$$

The reduced form equation shows that the overall effect of labour market institutions is ambiguous. Stronger unions tend to increase the labour share and compress the wage distribution, both of which reduce inequality. However, they also increase the unemployment rate, which raises the Gini coefficient. An increase in the unemployment benefit raises both the labour share and unemployment, and these changes have opposite effects on inequality. Lastly, a higher capital-labour ratio unambiguously reduces inequality as it simultaneously increases the labour share and lowers both the unemployment rate and the relative wage.

3.2. The data

Our data cover 16 OECD countries over the period 1960-96. Details on the data and their sources are provided in Appendix II. As is well known, the data on income inequality are problematic and international comparisons difficult (see Atkinson and Brandolini, 2001). For this reason we use two different sources for our income inequality measure: one measure is obtained from Brandolini (2003), who collected comparable measures of income inequality for several OECD countries; the other measure is obtained from Deininger and Squire (1996), which has become the standard dataset for empirical studies of income inequality. Brandolini (2003) provides detailed information on the way in which data were collected, allowing us to build a series that is more comparable over time, and most of our analysis will be based on them. However, as a robustness check, we replicate our regression equations using the Deininger and Squire data. Unfortunately these two datasets on income inequality overlap only partially, and therefore the results are not directly comparable.

Table 1a – Descriptive statistics for main variables – sample means by countries

country	gini1	gini2	p9010	ls1	ur	ben
Australia	32.83	38.08	2.83	0.49	5.32	0.22
Belgium	27.75	26.81	2.34	0.52	6.64	0.41
Canada	36.03	31.32	4.24	0.53	7.42	0.26
Denmark	32.86	32.08	2.17	0.55	5.01	0.44
Finland	21.76	29.77	2.45	0.51	5.68	0.25
France	38.33	42.13	3.44	0.52	6.52	0.30
Germany	36.22	31.23	2.84	0.54	3.85	0.29
Italy	34.71	34.67	2.33	0.46	6.07	0.05
Japan	na	34.86	3.06	0.51	2.20	0.11
Korea	na	34.18	3.97	0.41	na	na
Netherlands	28.55	28.54	2.61	0.55	5.11	0.45
New Zealand	27.23	34.06	3.03	0.48	2.91	0.31
Norway	22.64	34.75	2.08	0.48	2.73	0.23
Sweden	47.12	31.69	2.10	0.58	3.17	0.19
United kingdom	27.52	25.98	3.27	0.58	6.31	0.22
United states	37.58	35.49	4.16	0.58	5.86	0.12
Total	33.98	32.56	3.03	0.52	5.00	0.26

Table 1b – Descriptive statistics for variables in the dataset – sample means

Variable	Obs	Mean	Std.Dev.	Min	Max
gini1	236	33.981	7.295	19.900	54.300
p9010	315	3.028	0.690	1.953	4.640
ls1	651	52.060	5.419	32.268	64.909
ur	593	4.999	3.324	0.000	16.800
ben	600	0.257	0.142	0.003	0.670
udnet	585	0.427	0.181	0.099	0.911
minim	704	0.667	0.275	0.241	1.000
edu	656	9.603	1.799	3.457	12.876
kpw	528	10.187	0.512	7.646	11.173
tw	566	0.492	0.125	0.237	0.831
oil	689	4.436	2.498	-0.333	10.991

Legend:
gini1 = Gini index on personal income distribution, from Brandolini 2003
p9010 = ratio between 90th and 10th percentile in earnings distribution, from OECD
ls1 = labour share on value added at market price, from OECD-Stan database
ur = unemployment rate, from Nickell-Nunziata 2001
ben = unemployment benefit, from OECD 2001
udnet = union density, from Nickell-Nunziata 2001
minim = ratio of minimum wage to median wage, from OECD
edu = average years of schooling of population 25 and over, whether studying or not, from Cohen and Soto 2001
kpw = (log of) capital per worker, from Summer and Heston 1991, updated with mark 5.6 of the Penn tables
tw = tax wedge, from Nickell-Nunziata 2001
oil = (log of) oil price in national currency, from IMF Financial Statistics

Data on labour shares are from the OECD Stan Database. We use the standard definition of total compensation per employee over value added, without any correction for the incomes of the self-employed. This measure fits well our theoretical definition of the labour share, which comprises only the income of employed individuals. The wage differential is proxied by the ratio between 1st and the 9th decile of the earnings distributions (from OECD specific database).[11] Standard datasets were used to obtain information about on labour market institutions, educational attainments, and capital (see Appendix II for details).

Table 1a reports some descriptive statistics for the main variables in our regressions. While the potential sample size is 592 observations (16 countries × 37 years), many observations are missing, thus reducing the available sample to 233 observations, among which the US, the UK, Germany, Sweden, Italy, and Canada have the most observations. Table 1b reports the descriptive statistics of our entire dataset.

3.3. Determinants of labour market outcomes

Table 2 examines the determinants of the labour share and presents three alternative specifications. The strongest impact on the labour share is exerted by the capital/labour ratio (as implied by our model), independently from the specification adopted. In column 1 we find that the labour share is increasing in union density rates. This effect persists when country fixed effects are taken into account (column 2) but disappears when cyclical factors are properly accounted for using year fixed effects (column 3). The results also capture the fact that when minimum wage legislation applies, employment of both skilled and unskilled workers declines, leading to an increase in the wage share.[12] Similarly, the unemployment benefit also has a positive (but weakly significant) impact on the labour share. We include the price of oil in national currency in order to capture exogenous shocks to raw materials prices (this variables also captures the effect of competitive devaluations, and the J-effect on internal inflation).[13] Lastly, we have considered the potential role of the supply of skills. Time series of labour force composition by skills are not available over a long enough time span, therefore we use proxies derived from measures of educational attainment. The one reported in the text is the average years of education in the adult population, but enrolment rates had a similar effect. Once country fixed effects are included, the education variable displays a negative coefficient, suggesting that as the number of skilled individuals increases, the unemployment rate of the skilled rises, reducing the incentives to shirk and hence allowing firms to pay a lower skilled wage.

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[11] We experimented with both the relative difference and the more conventional measure based on percentile ratio, using the latter alternative for better econometric performance.

[12] Using the level of the minimum wage as an explanatory variable is problematic, it is missing for several countries (Denmark, Finland, Germany, Italy, Norway, Sweden and UK for most of the sample period). In order not to loose degrees of freedom, we have replaced the missing observation with a unitary value, which is cleared away with the country fixed effect.

[13] Unfortunately this variable alternates sign depending on whether or not time fixed effects are included. For this reason, we will discard it as potential instrument.

Table 2 – Determinants of labour share – OLS regressions
robust standard errors - t-statistics in parentheses - * significant at 10%; ** significant at 5%; *** significant at 1%

	1	2	3
log capital per worker	-0.036	10.98	8.56
	[0.04]	[16.19]***	[12.49]***
unemployment benefit	1.236	1.583	3.094
	[1.03]	[0.72]	[1.73]*
union density rates	5.529	11.714	-1.054
	[4.59]***	[5.94]***	[0.65]
ratio minimum/median wage	-2.37	9.298	3.768
	[2.99]***	[3.28]***	[1.42]
log oil price in national currency	-0.425	0.957	-0.91
	[3.94]***	[7.30]***	[3.63]***
average years of education	0.537	-4.143	-1.221
	[3.03]***	[13.35]***	[1.78]*
Constant	yes	yes	yes
Country fixed effects		yes	yes
Year fixed effects			yes
Observations	455	455	455
R^2	0.19	0.80	0.88

Our results are in line with those obtained in earlier work. Bentolila and Saint Paul (2003), who consider sectoral data for 12 countries over a shorter time span, find a significant correlation between the labour share (corrected for self-employment) and the capital-output ratio, strike activity, employment adjustment costs (proxied by previous changes in employment) and total factor productivity, whereas the oil price is found to be statistically insignificant. While the sign of the coefficient on the capital-output ratio varies across sectors (depending on the degree of substitutability/complementarity between factors), they find a weakly significant negative coefficient on strike activity, which they interpret as lagged responses to wage push. Blanchard (1997) finds that labour share movements are mainly affected by supply shocks, with significant reaction lags. Our results hence support the traditional view that factor shares respond to relative factor endowments (here proxied by capital per worker) but that there is evidence that wage push factors (union density, minimum wage and unemployment benefit) also have some impact.

In table 3 we report the determinants of the wage differential. The most significant correlations are found with factor endowments: an increase in the capital-labour ratio reduces the wage ratio because unskilled workers exploit relatively better the improved employment situation; while an increase in skill availability in the labour force (proxied by our human capital variable) tends to depress the relative wage. Once again we find a significant impact of labour market institutions. Not surprisingly, the minimum wage reduces wage differentials. The unemployment benefit, which according to the model has an ambiguous effect as it increases both skilled and unskilled wages, also seems to reduce wage dispersion. Some negative impact can also be found for union density, although this may capture some cyclical component as the effect disappear once year fixed effects are included. The time trend exhibits a positive and significant coefficient, capturing the upwards trend in earnings inequality, potentially associated to skill-biased technical change.

Table 3 – Determinants of p90/p10 decile ratio – OLS regressions
robust standard errors - t-statistics in parentheses - * significant at 10%; ** significant at 5%; *** significant at 1%

	1	2	3
log capital per worker	-0.137	-0.58	-0.369
	[0.94]	[4.73]***	[4.11]***
unemployment benefit	-0.683	-0.726	-0.765
	[3.19]***	[2.50]**	[2.36]**
union density rates	-1.385	-0.419	-0.147
	[9.92]***	[1.49]	[0.48]
ratio minimum/median wage	-0.853	-2.405	-1.815
	[5.18]***	[4.65]***	[4.02]***
average years of education	0.11	-0.493	-0.508
	[6.39]***	[5.73]***	[4.08]***
time trend	-0.01	0.066	0.053
	[2.02]**	[6.37]***	[4.81]***
Constant	yes	yes	yes
Country fixed effects		yes	yes
Year fixed effects			yes
Observations	260	260	260
R^2	0.65	0.97	0.98

Koeninger et alt. (2005) study wage inequality in a framework similar to ours, estimating the determinants of the p90/p10 ratio for 11 countries over a similar time interval. Our results are consistent with their analysis, since both papers find that stronger labour market institutions compress wage differentials.[14] In contrast to us, Koeninger at al. (2005) include as regressors import penetration and R&D intensity to account for skill-biased technological change without finding robust effects.[15] We limit ourselves to a linear time trend, which is identical across countries and bears a positive coefficient.

Lastly, table 4 replicates well-known results on the institutional determinants of unemployment, which is positively correlated with union density and the minimum wage. Unemployment declines with capital accumulation, as it increases workers' productivity and hence expands labour demand. Contrary to our theoretical expectation, the coefficient on the unemployment benefit is not significant in this equation,

14 This is also consistent with micro-data analysis; see DiNardo et al. (1996) and more recently Card et al. (2003).

15 A further difference with their analysis is that they consider employment protection. While in their theoretical model they assume that skilled and unskilled workers should face different firing cost, due to the lack of data in the empirical analysis they resort to the unique series available, produced by OECD. However this series exhibit little variation across years, as witnessed by its statistical insignificance when first differences are considered. For this reason we have decided not to include EPL into our regressions.

while the tax wedge has a negative impact.[16] Note that both coefficients are coherent with theoretical expectations and are significant when we do not include country dummies, as found in previous work (see for example Nickell, 1997). More recently, Nickell et al. (2005) have studied the determinants of the unemployment rate for 20 countries over the period 1961-92, including a list of shocks (labour demand, total factor productivity, real import, money supply, and real interest rates) and the lagged dependent variable. They find a positive and significant coefficient on the unemployment benefit (especially the replacement rate) and the rate of change of union density, a weaker effect for the tax wedge, and an absence of statistical significance for the employment protection measure.

Table 4 – Determinants of unemployment rate – OLS regressions
robust standard errors - t-statistics in parentheses - * significant at 10%; ** significant at 5%; *** significant at 1%

	1	2	3
log capital per worker	-0.976	-2.333	-2.097
	[2.28]**	[4.80]***	[3.32]***
unemployment benefit	4.438	-1.212	-0.078
	[5.22]***	[0.73]	[0.05]
union density rates	-1.383	7.912	4.581
	[1.58]	[3.96]***	[1.98]**
ratio minimum/median wage	-1.718	13.484	14.281
	[2.58]**	[4.49]***	[6.05]***
tax wedge	4.634	-7.701	-10.215
	[3.11]***	[2.95]***	[4.01]***
time trend	0.21	0.337	0.247
	[11.30]***	[13.35]***	[4.38]***
Constant	yes	yes	yes
Country fixed effects		yes	yes
Year fixed effects			yes
Observations	448	448	448
R^2	0.45	0.71	0.78

To sum up, our results are consistent with previous work and indicate that labour market institutions are essential determinants of labour market outcomes. Union bargaining power (proxied by union density and the minimum wage) has an impact on the labour share, wage differentials and unemployment rates, while the unemployment benefit mainly affects wage dispersion. Capital accumulation, in terms of both equipment (fixed capital) and educational attainment (human capital) also have a major impact on our dependent variables.

[16] While the standard expectation is of a positive sign (because a higher tax wedge under wage bargaining leads to net wage resistance, and therefore increases labour costs and decreases employment), general equilibrium consideration may lead to the opposite expectation (see Corneo 1995).

3.4. The determinants of personal income inequality

We move next to the determinants of personal income inequality. Table 5 reports our estimates of equations (21) and (22) for the largest available sample.[17] The 1st column of table 5 abstracts from country and year fixed effects, which are subsequently included in the 2nd and 3rd columns; a linear time trend and dummies controlling for changes in definitions are also included.[18] We find that all variables have significant coefficients with the expected signs, with the exception of the unemployment rate. The labour share has a negative coefficient, while the wage differential has a positive one.[19] The unemployment benefit appears negatively related to income inequality, whereas the unemployment rate has an insignificant coefficient. However, the unemployment rate has a positive and significant sign when we move to instrumental variable estimation. The linear time trend bears a negative and significant coefficient, indicating an unexplained decline in inequality over the sample period.

The comparison between the OLS results obtained in the 3rd and 4th columns and the IV estimates reported in 5th and 6th columns indicates that OLS-estimation provides downward-biased estimates of the actual effect of the labour share and wage inequality on income inequality, and an upward bias for the effect of the unemployment benefit.[20] This bias could be merely due to measurement errors, but it could also indicate that some unobservable variable, which correlates with both income inequality and labour market institutions – such as the political orientation of the government or the attitude of the population towards redistribution – has been omitted. It is interesting to note that, while the impact of passive labour market policies remains significant and negative, the unemployment rate and the time trend gain statistical significance under IV estimation. As a robustness check, columns 7 and 8 report the same model estimated in first differences, with and without country fixed effects: the coefficient on the labour share retains its sign and significance, even if the effect is attenuated, whereas the wage differential is close to non-significance, while the variables related to unemployment are both insignificant.

[17] The sample size hinges crucially on the availability of data on wage differentials. If we concentrate on personal income inequality only, the available sample is made of 233 observations. When we consider the overlapping with information on wage differentials, the sample is further reduced to 142 observations. In order not to loose too many observations, we have replaced the missing observation for the P9010 variable with its country-specific sample mean. The sample reduction due to the availability of data on wage differentials (2nd and 5th columns) does not affect sign and significance of the other regressor (details available from the authors). This fictitious enlargement of the sample allows us to retain relevant information that otherwise would be excluded due to missing observations on earnings differentials. Fort his reason, in the sequel we will consider this extended sample.

[18] The controls for definition include whether the income is gross or net, and whether the recipient is household equivalent or person equivalent. We also experimented with errors clustered by countries, without significant changes (available from the authors).

[19] Kenworthy (2003) uses household income inequality and personal earnings inequality (proxied by p90/p10 ratio) computed from LIS (Luxemburg Income Study), with one observation for 14 countries. By regressing the former onto the latter, he finds a coefficient comprised between 0.61 and 0.68, depending on various specifications, which is much lower than our figures. But sample size and countries are not comparable.

[20] The instrument have been selected from the regressors used in tables 2, 3 and 4 so as to satisfy the Sargan test for overidentifying restrictions.

Table 5 – Determinants of personal income inequality – full sample – OLS and IV estimates

robust standard errors – t-statistics in parentheses - * significant at 10%; ** significant at 5%; *** significant at 1%

	1	2	3	4	5	6	7	8
	OLS	OLS	OLS	IV	IV	IV	1st differen	1st differen
labour share at market price	0.229	-0.297	-0.365	-0.261	-0.365	-1.345	-0.141	-0.134
	[1.90]*	[4.19]***	[3.37]***	[0.33]	[2.03]**	[3.58]***	[2.94]***	[2.81]***
extended p90/p10 decile ratio	-1.219	4.212	5.145	0.973	17.942	26.848	0.745	0.697
	[2.51]**	[3.99]***	[3.71]***	[0.34]	[4.52]***	[4.14]***	[1.69]*	[1.58]
unemployment rate	-0.697	0.021	0.033	-2.679	0.784	1.138	0.029	0.025
	[6.11]***	[0.28]	[0.31]	[2.61]***	[2.42]**	[2.45]**	[0.56]	[0.48]
unemployment benefit	-8.52	-21.6	-23.398	2.284	-17.328	-11.079	0.124	1.953
	[3.45]***	[5.47]***	[6.45]***	[0.26]	[3.98]***	[1.73]*	[0.03]	[0.42]
time trend		0.038	0.091	0.636	-0.222	-0.423		
		[1.02]	[1.39]	[3.67]***	[2.46]**	[2.62]***		
Constant	yes	yes	yes	yes	yes	yes	yes	yes
Country fixed effects		yes	yes	yes	yes	yes		yes
Year fixed effects			yes			yes		
Observations	210	210	210	210	210	210	202	202
R^2	0.51	0.93	0.94	0.12	0.87	0.79	0.05	0.15
Sargan test (p-value)				0.05	0.07	0.80		

Endogenous variables: labour share, unemployment rate, p90/p10.
Instruments: (log)capital×worker, tax wedge, years of education, minimum wage, bargaining coordination.

Our preferred specification, in terms of sign, significance and size of the effects, is the 5th column of table 5. Evaluated at sample means, the estimated elasticity of personal income inequality with respect to the labour share is equal to –0.61, which implies that reducing the labour share by one standard deviation would raise the Gini coefficient by 2.47 points. The estimated elasticity with respect to the wage differential is greater, reaching the value 1.63: increasing the decile ratio by one standard deviation would raise the Gini coefficient by 1.27 Gini points. The elasticities with respect to the unemployment benefit and the unemployment rate are much lower, respectively 0.12 and 0.14.

The results are basically unchanged when we computed quinquennial averages of the data, when we restrict the data to the subsample of countries for which we have longer time series, and when we use the labour share measure adjusted for self-employment (details available from the authors). On the contrary, when we use the alternative series for income inequality, the labour share is less significant, with a coefficient that is half in size. The wage differential and the unemployment subsidy are still highly significant, with coefficients of similar sizes, while the unemployment rate is totally insignificant.

Table 6. – Determinants of personal income inequality – 3SLS regressions
Absolute value of z statistics in brackets - * significant at 10%; ** significant at 5%; *** significant at 1%

dependent variable:	income inequality (Gini)	labour share	p90/p10 decile ratio	unemployment rate
labour share at market price	-0.723			
	[4.25]***			
p90/p10 decile ratio	7.389			
	[3.50]***			
unemployment rate	0.294			
	[2.05]**			
unemployment benefit	-5.593	7.864	-0.426	5.771
	[1.35]	[2.20]**	[1.47]	[1.34]
union density rates		1.862	-2.005	23.406
		[0.45]	[6.05]***	[4.59]***
ratio minimum/median wage		3.437	-3.505	41.527
		[0.58]	[7.38]***	[6.50]***
log capital per worker		22.136	0.289	13.821
		[5.71]***	[0.90]	[3.12]***
average years of education		-7.783	-0.104	
		[5.06]***	[0.83]	
log oil price in national currency		-2.567		
		[2.88]***		
tax wedge				-36.702
				[5.94]***
time trend	0.136		0.058	-2.687
	[0.10]		[0.53]	[1.67]*
Constant	yes	yes	yes	yes
Country fixed effects	yes	yes	yes	yes
Year fixed effects	yes	yes	yes	yes
Observations	135	135	135	135
Root mean squared error	1.3	1.06	0.08	1.31
R^2	0.96	0.94	0.99	0.81

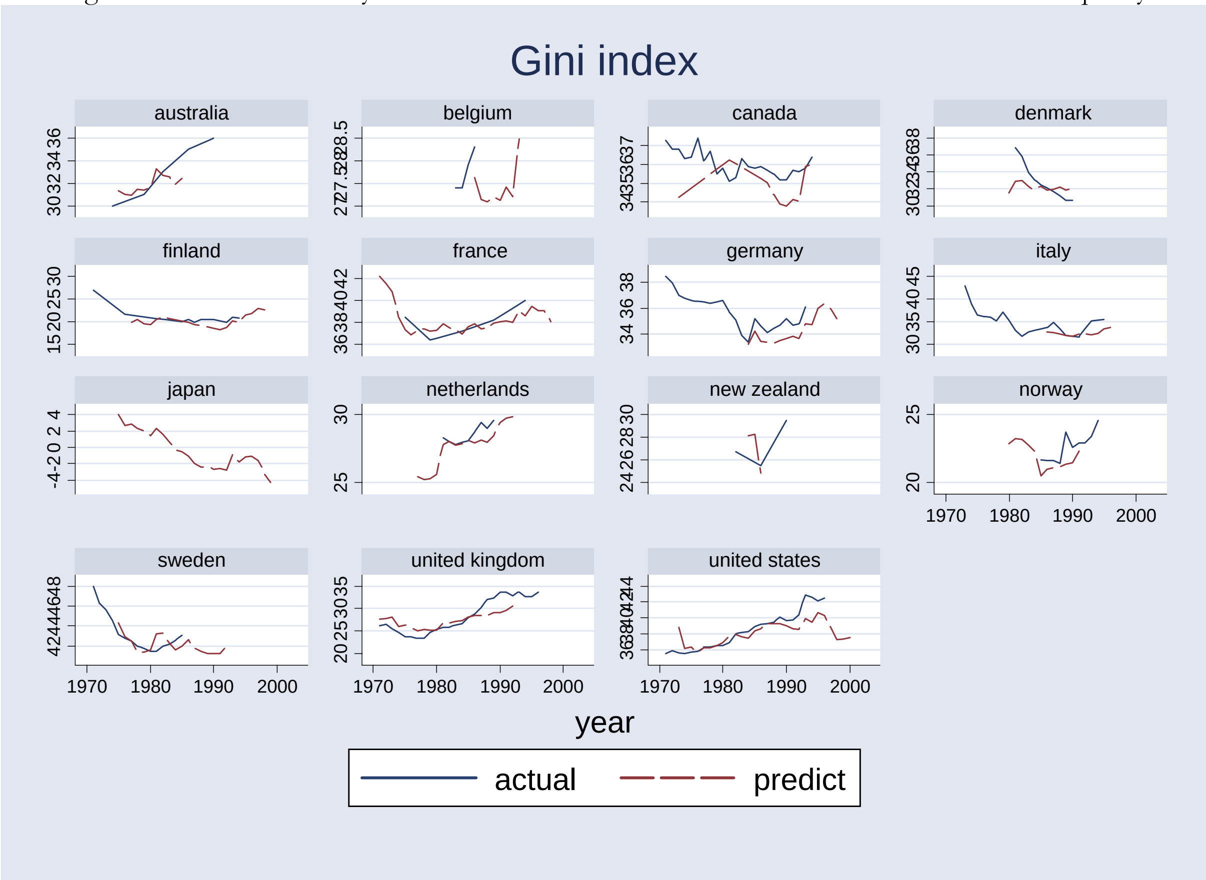
Controls for changes in definition included.

We now consider the alternative strategy of estimating the simultaneous equation system defined by equations (18), (19), (20) and (21) through three-stage least squares. The estimated coefficients are reported in table 6.[21] Previous results are confirmed: our three endogenous variables (the labour share, wage differential and unemployment rate) are correlated with income inequality, whereas the unemployment benefit exerts its impact only indirectly, through the labour share. The equation for the determination of the labour share is consistent with what we have already found in least square estimation (table 2), including country and year fixed effects. Similarly, for the wage differential the pressure for wage compression deriving from union presence and/or from minimum wage legislation is consistent with the OLS estimates reported in table 3, even if capital accumulation and unemployment benefit loose

[21] Note that sample size declines from 210 to 135 observations, because we cannot use the extended series for the wage differential.

significance. What is less satisfying is the unemployment equation, where the tax wedge continues to have a negative impact, while the capital-labour ratio has positive coefficient.[22]

Figure 2 – Predictive ability of the model estimated in table 6 – Gini index on income inequality



The predictive ability of the model is good, as can be seen from figure 2 that compares the actual and the predicted dynamics of the Gini index. This is rather impressive if one considers the block recursive nature of the model, noting that in addition to its own prediction error, the prediction for the Gini index accumulates the prediction errors from the other three endogenous variables. In particular our model captures the trend reversal in income inequality observed in most European countries at the end of the 1970s (notably the Netherlands, France and United Kingdom), which seems to be largely explained by the contemporaneous decline in the labour share.

3.5. Counterfactual exercises

We have performed a number of counterfactual exercises to assess the relative importance of the various labour market variables. To illustrate the process, consider figure 3. Using the estimated coefficients in table 6, we have obtained the predicted values for labour share, wage differential and unemployment rate for all the country/year observations available in the sample (and even outside the estimation sample, as

[22] The former sign reverts to positive when the sample is restricted to 6 countries for which we have more consistent observations (89 observations), while the latter becomes insignificant if year fixed effects are neglected.

witnessed by the predicted Japanese inequality index reported in figure 2). We then use these values to further predict the Gini coefficient.

In figure 3, in addition to the standard prediction (continuous line) we also report the Gini index replacing the predicted US labour share by the French labour share (long-dashed line) and the US wage differential by the French wage differential (short-dashed line). The figure illustrates that inequality in the US would have been 10 Gini points lower if US earnings differentials had been comparable to French ones. However, the higher labour share experienced by US economy helped reduce income inequality, which would have been even higher had the labour share declined as in continental Europe.

The reverse situation occurs when we make use of US labour share and wage differentials in European countries, as we do in figure 4 for the UK and Norway. Not surprisingly, both countries would have experienced greater inequality had they had the US wage differential. The effect of the labour share, however, differs. The UK is characterised by a relatively high and stable labour share and therefore the replacement of US labour share does not affect income inequality, while for Norway its relatively low labour share was an unequalising force over the period.

In figures 6, 7 and 8 we performed counterfactual simulations by replacing the dynamics of the main exogenous variables, labour market institutions (union density and unemployment benefit) and capital accumulation (capital per worker and educational attainment) with those for the US. In figure 5 we use the US union density and unemployment benefit in the prediction of income inequality for Sweden and France, while in figure 6 we do the same for Canada and UK. In general, income inequality would have been higher if European countries had experienced US-type labour market institutions, but there are country specific variations. In the case of Sweden, a country characterised by high density rates (due to the so called "Ghent system", where unions run unemployment benefit schemes on behalf of the state), a decline of density rates to the US level would have induced a 6 Gini points increase in inequality, while a similar move would have left unaffected France, where union coverage is high but membership is even lower than US.[23] The impact of unemployment benefit changes is more limited, because the US-Europe gap is lower.[24] All European countries considered in this exercise would have experienced an increase in income inequality, and the impact seems stronger in Canada, which is typically considered a US-type economy but stronger worker protection.

[23] For a comparison of different models of unionisation see Checchi and Lucifora 2002.

[24] Looking at table 1, the sample-average replacement rate in US is 0.12, to be compared with 0.19 of Sweden, 0.22 of UK, 0.26 of Canada and 0.30 of France.

Figure 3 – Counterfactual 1: US inequality with French labour market outcomes

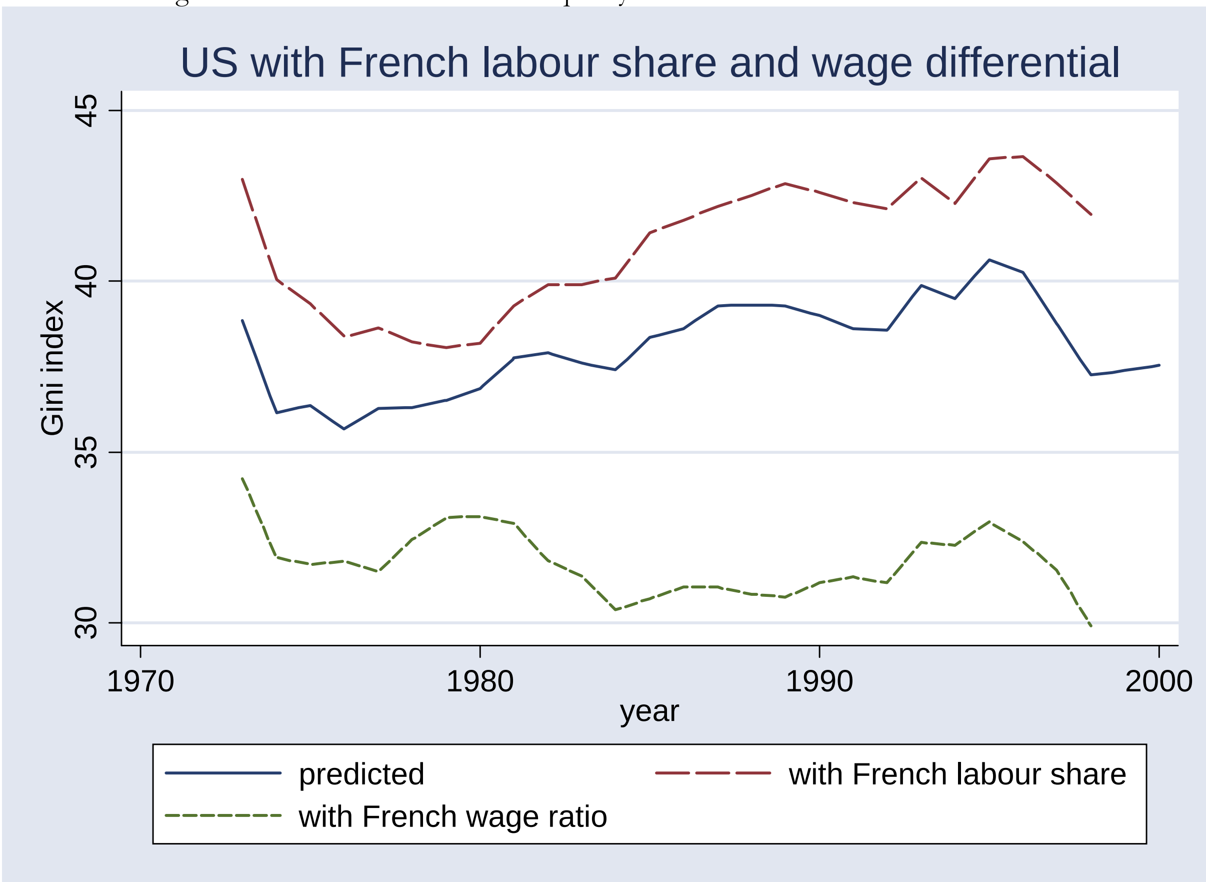


Figure 4 – Counterfactual 2: UK and Norway inequalities with US labour market outcomes

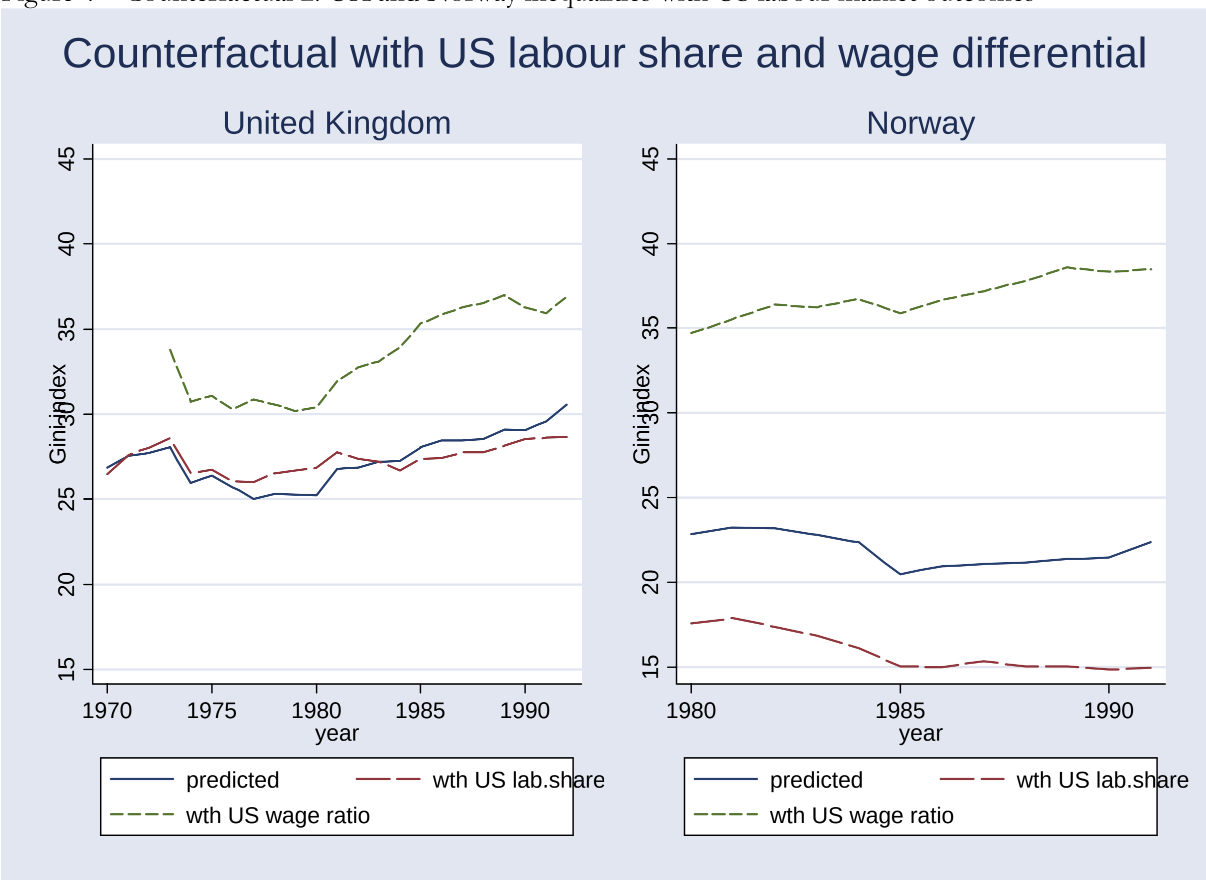


Figure 5 – Counterfactual 3: Sweden and France inequalities with US labour market institutions

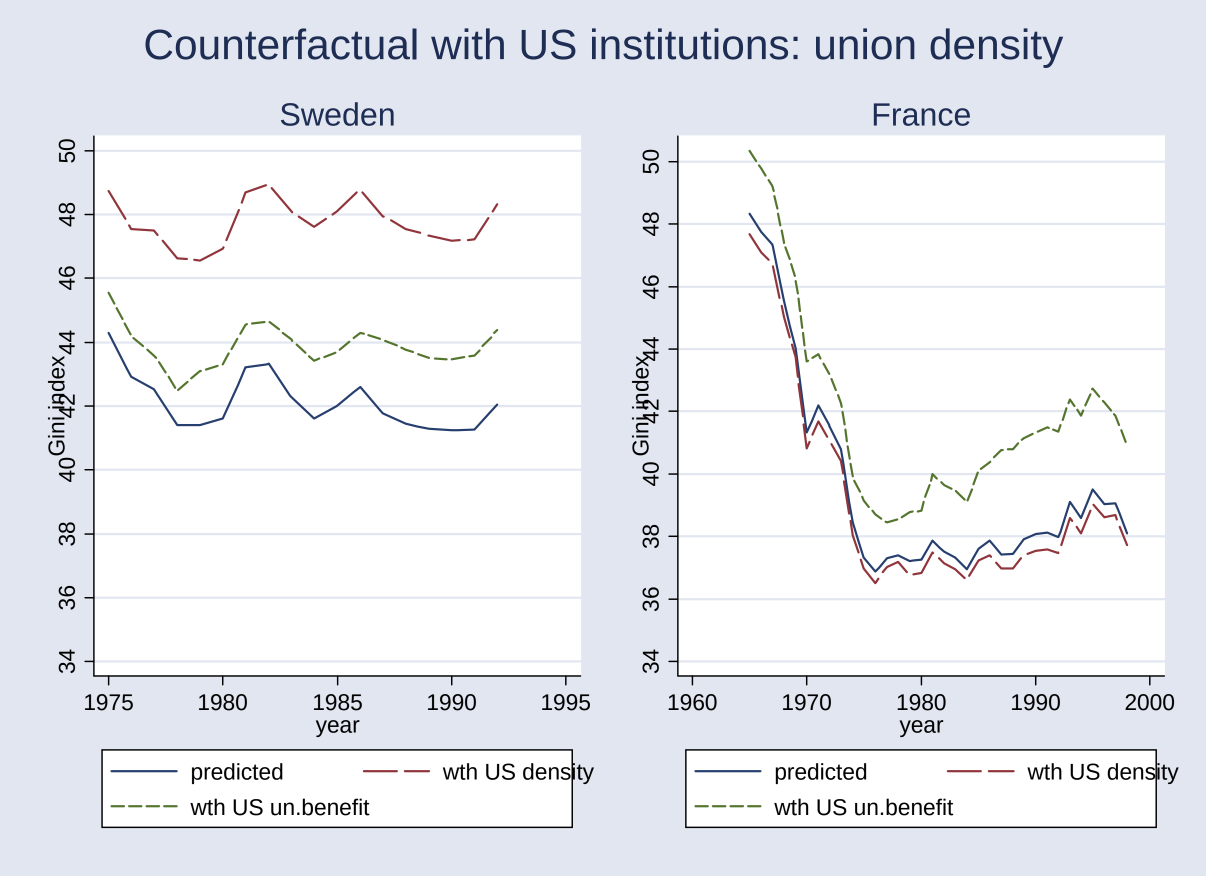
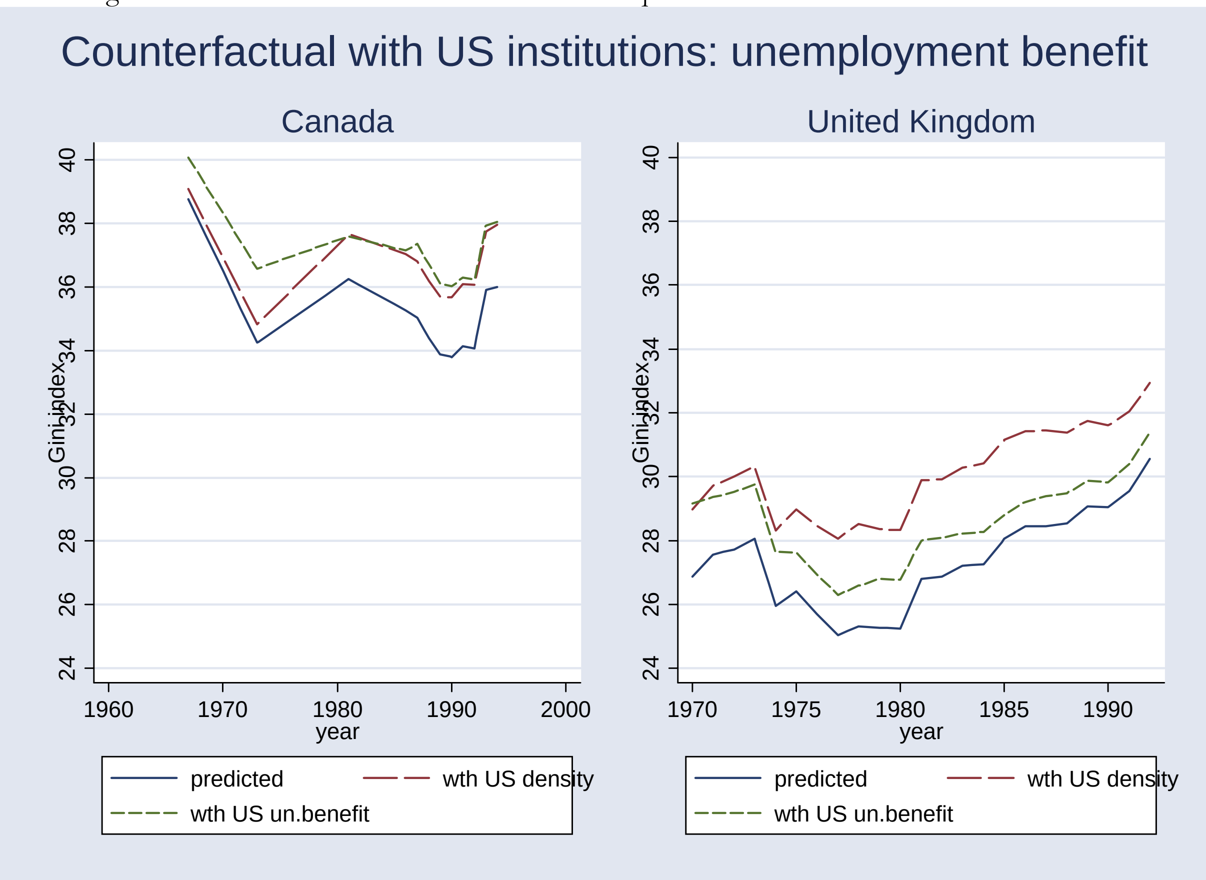
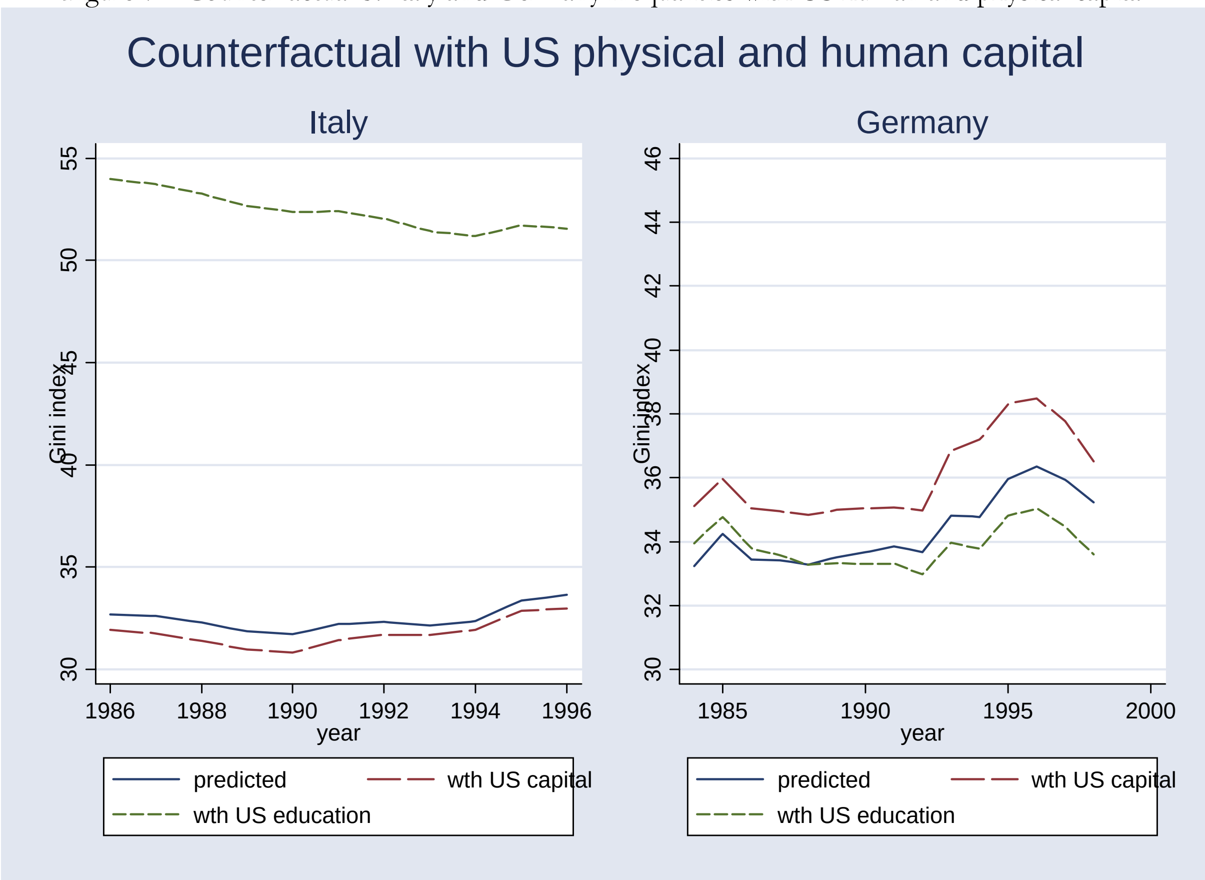


Figure 6 – Counterfactual 4: UK and Canada inequalities with US labour market institutions



We also consider the role of physical and human capital. As can be seen in figure 7, these variables have very significant impact on income inequality. Looking at the left panel, where Italian inequality has been recalculated using US values for capital per worker and average years of education in the population, we notice that inequality would have been much higher if we were to consider the highly educated US labour force instead of the low educated Italian population. According to the estimated model in table 6, educational attainment has two countervailing effects: on the one hand, the increased supply of human capital reduces the wage differential; on the other, by making skilled labour cheaper, it also induces substitution between skilled and unskilled, thus depressing the labour share. The overall effect is therefore ambiguous, and can only by judged case by case. In the case of Germany, a country with similar educational attainment to the US, it is physical capital that generates distributional differences. The high German capital-labour ratio had a strong equalising effect, through its impact on both wage inequality and the labour share.

Figure 7 – Counterfactual 5: Italy and Germany inequalities with US human and physical capital



In order to provide some information about the order of magnitude of these impacts, table 7 reports estimates of the reduced form equation for income inequality corresponding to equation (23). The table reports the OLS standardised beta coefficients, which are to be read as the change, in terms of a fraction of a standard deviation in the dependent variable, induced by a standard deviation change in the exogenous variable. The first two columns consider the entire sample, controlling for country and year effects, while the 3rd and 4th column restrict the sample to 6 countries for which we have longer series of income inequality (US, UK, Germany, Sweden Italy and Canada). The two types of capital have the strongest correlations with income inequality, with the overall impact being negative for capital equipment and positive for human capital.[25] Two labour market institutions, union density and unemployment benefit, negatively affect income inequality and have effects of comparable magnitude, while the minimum wage has a marginally significant, but still negative, effect. Finally, the tax wedge, which probably works as a proxy of the welfare state size, also exhibits a strong and negative correlation with income inequality.

Table 7 – Determinants of personal income inequality – reduced form – OLS
Robust normalized beta coefficients - * significant at 10%; ** significant at 5%; *** significant at 1%

	1	2	3	4
	full sample	full sample	reduced sample	reduced sample
union density rates	-0.328***	-0.423***	-0.053	0.187
ratio minimum/median wage	-0.055	-0.315*	0.405	0.313
unemployment benefit	-0.216**	-0.342***	-0.191*	-0.302***
tax wedge	-0.341***	-0.371***	-0.422***	-0.527***
log capital per worker	-0.422**	-0.965***	-0.831***	-1.918***
average years of education	-0.076	1.393***	-0.061	2.200***
Constant	yes	yes	yes	yes
Time trend	yes	yes	yes	yes
Country fixed effects	yes	yes	yes	yes
Year fixed effects		yes		yes
Observations	211	211	154	154
R^2	0.95	0.96	0.93	0.95

Controls for changes in definitions and oil price included.

4. Conclusions

The recent literature on the determinants of personal income inequality has highlighted the role of a number of factors, including output levels, globalisation, educational attainment, and political stability. Many of these variables help us understand distributional differences in a large cross-section of countries, but have little explanatory power when trying to understand differences across the rather similar OECD economies or variations over time. In this paper we have argued that labour market institutions have played an essential role in explaining differences in inequality within the OECD.

Our analysis highlights the different channels through which labour market institutions affect distribution. In particular, these institutions affect simultaneously relative wages, the labour share, and the unemployment rate, all of which then have an impact on the distribution of personal incomes. Because

[25] Barro (2000) finds a negative correlation between inequality and secondary school enrolment and a positive correlation with tertiary enrolment; these findings are difficult to compare with ours, since we have a stock measure, combining three levels of educational attainment.

they operate through these three mechanisms, their impact is a prioir ambiguous. Our results indicate that the factor distribution of income is still an essential component of personal income inequality, and that the impact of labour market institutions operates in part through the way in which income is shared between capital and labour. The overall effect of stronger institutions is to reduce income inequality, with part of this effect occurring through wage compression and part of it through a reduction in the rewards to capital.

Our analysis has important policy implications. The first one concerns the role of redistribution. The view that a widening wage dispersion has been the major cause of the recent increase in income inequality leaves little role for policy. The increase in wage dispersion is usually seen as the result of trade and innovation. Since both increased openness and technological change are seen as desirable, greater inequality has been perceived as an unavoidable by-product of the growth process. Income redistribution can then be used to reduce net-income inequalities, but would not affect the distribution of market incomes. In contrast, the negative impact of the labour share on the Gini coefficient indicates that the distribution of wealth across agents is still a major source of inequality, and hence leaves room for policy to affect inequality in the long-run. Income redistribution will have the effect of reducing differences in the accumulation of wealth across agents and hence affect gross-income inequalities in the future.

The second aspect concerns the role of labour market institutions as a source of equalisation. We have found that labour market institutions significantly affect income inequality through several channels: stronger unions obtain a larger wage share and compress wage differentials, while higher minimum wages reduce wage differentials. Despite the associated increase in the unemployment, the overall impact is to reduce inequality. A caveat is, however, in order. Our analysis is static and takes the stock of physical capital as given. This implies that we are ignoring the potential impact of labour market institutions on investment. Given the strong equalising effect that a higher capital-labour ratio has, this is an important question that remains to be addressed in future work.

References

Acemoglu, Daron (2003). "Labour- and Capital-Augmenting Technical Change," *Journal of the European Economic Association,* 1(1), pp. 1–37.

Aghion, P., Caroli, E. and García-Peñalosa, C. 1999. "Inequality and Economic Growth: The Perspective from the New Growth Theories", *Journal of Economic Literature*, 37(4): 1615-60.

Alderson, A. and F.Nielsen. 2002. Globalisation and the great U-turn: income inequality trends in 16 OECD countries. *American Journal of Sociology* 107: 1244-1299.

Anand, S and Kanbur, S.M.R. 1993. "Inequality and Development: A Critique", *Journal of Development Economics* 41(1): 19-43.

Antras, Pol (2004). "Is the U.S. Aggregate Production Function Cobb–Douglas? New Estimates of the Elasticity of Substitution," *Contributions to Macroeconomics,* 4 (1).

Atkinson, A. and A.Brandolini. 2003. The Panel-of-Countries approach to explaining income inequality: an interdisciplinary agenda. mimeo

Atkinson, A.B. 1997. "Bringing the Income Distribution in from the Cold", *Economic Journal,* 107(441): 297-321.

Atkinson, A.B. and A. Brandolini. 2001. "Promise and Pitfalls in the Use of "Secondary" Data-Sets: Income Inequality in OECD Countries as a Case Study", *Journal of Economic Literature*, 39(3): 771-99.

Atkinson, Anthony. 1970. "On the Measurement of Inequality," *Journal of Economic Theory* 2, pp. 244-263.

Atkinson, Anthony. 1975. *The Economics of Inequality*. Second edition (1983). Oxford: Clarendon Press.

Barro, R.J. 2000. "Inequality and Growth in a Panel of Countries", *Journal of Economic Growth,* 5: 5-32.

Barro, R.J. and Lee, J.-W. 1993. "International Comparisons of Educational Attainment", *Journal of Monetary Economics,* 32 (3): 363-94.

Bentolila S. and G. Saint-Paul, 2003. "Explaining Movements in the Labor Share", *Contributions to Macroeconomics* 3(1), article 9. http://www.bepress.com/bejm/contributions/vol3/iss1/art9

Blanchard, O. 1997. The Medium Run. *Brookings Papers on Economic Activity,* 2: 89–158.

Bourguignon, F. and Morrisson, C. 1990. "Income distribution, development and foreign trade: a cross-section analysis", *European Economic Review*, 34 (): 1113-32.

Bourguignon, F. and Morrisson, C. 1998. "Inequality and development and: the role of dualism", *Journal of Development Economics*, 57 (): 233-57.

Bourguignon, F., F.H.G. Ferreira and P.G. Leite, 2002. " Beyond Oaxaca-Blinder : Accounting for Differences in Household Income Distributions Across Countries", document de travail DELTA working paper 2002-04.

Brandolini, A. 2003. A bird-eye view of long-run changes in income inequality. Bank of Italy, mimeo.

Breen, R. and García-Peñalosa, C. "Income Inequality and Macroeconomic Volatility: An Empirical Investigation", forthcoming *Review of Development Economics.*

Card, D., T.Lemieux and W.Craig Riddel. 2003. Unionization and wage inequality: a comparative study of the U.S., the U.K., and Canada. NBER working paper n.9473

Champernowne, D.G. 1973. *The Distribution of Income between Persons.* Cambridge: Cambridge University Press.

Checchi, D. 2004. "Does educational achievement help to explain income inequality ?" chapter 4 in A. Cornia (ed), *Inequality, Growth and Poverty in an Era of Liberalization and Globalization*. Oxford University Press 2004

Checchi, D. and C.Lucifora. 2002. Unions and labour market institutions in Europe. *Economic Policy* 17(2): 362-401

Cohen, D. and M. Soto. 2001. Growth and human capital: good data, good results. OECD Development Centre Technical paper n.179

Corneo, G. 1995. National wage bargaining in an internationally integrated product market, *European Journal of Political Economy* 11 (1995), 107-116

Daudey, E. 2004. "The Sharing of Value-Added: Data Sources", mimeo, GREQAM.

De Serres, A., S.Scarpetta and C.De la Maisonneuve. 2002. Sectoral shifts in Europe and the United States: how they affect aggregate labour shares and the properties of wage equations. OECD Economic Department wp.326.

Deininger, K. and Squire, L. 1996. "Measuring income inequality: a new data base", *The World Bank Economic Review* 10 (3): 565-591.

DiNardo, John, Nicole M. Fortin and Thomas Lemieux. 1996. Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semi-Parametric Approach. *Econometrica* 64 (5): 1001-1044.
Esping-Andersen, G. 2004. Income distribution and life chance opportunities. to appear in A.Giddens (ed). *The new egalitarianism: opportunity and prosperity in modern society.*
Förster, M. 2000. Trends and driving factors in income distribution and poverty in the OECD area. Labour Market And Social Policy- Occasional Papers No. 42
Gastil, R. *Freedom in the World*, Westport: Greenwood.
Gollin D., 2002. "Getting Income Shares Right". *Journal of Political Economy*, vol. 110, no. 2
Gottschalk, P. and Smeeding, T.M. 1997. "Cross-National Comparisons of Earnings and Income Inequality", *Journal of Economic Literature*, 35: 633-687.Hamermesh, David S. (1993). *Labor Demand.* Princeton, NJ: Princeton University Press.
Katz L.F. and Murphy, K.M., 1992. "Changes in Relative Wages, 1963-1987: Supply and Demand Factors", *Quarterly Journal of Economics*, vol. 107(1), pp. 35-78
Kenworthy, L. 2003. Explaining comparative trends in income inequality in the 1980s and 1990s. mimeo (appeared as chapter 3 in *Egalitarian Capitalism.* Russel Sage Foundation 2004)
Koeninger, W., M.Leonardi and L.Nunziata. 2005. Labour market institutions and wage inequality. mimeo
Krusell, Per, Ohanian, Lee, Rios-Rull, Victor, and Violante, Giovanni (2000). "Capital Skill Complementary and Inequality." LXIIX, *Econometrica,* pp. 1029–1053.
Li, H., Squire, L. and Zou, H.-F. 1998. "Explaining International and Intertemporal Variations in Income Inequality", *Economic Journal* 108: 26-43.
Nickell, S. 1997, "Employment and labor market rigidities: Europe versus North America", *Journal of Economic Perspectives,* 11/3: 55-74.
Nickell, S. and L.Nunziata. 2001. Labour market institutions database
Nickell, S., L.Nunziata and W.Ochel. 2005. Unemployment in the OECD since the 1960s. What do we know? *Economic Journal* 115: 1-27.
OECD 2002. *Benefits and wages – OECD indicators*. Paris
Ricardo, David. 1821. *Principles of Political Economy and Taxatio*n. 3 rd ed, in *The Works and Correspondence of David Ricard*o, ed. Piero Sraffa, Vol 1. Cambridge: Cambridge University Press.
Rowthorn, R. 1999. Unemployment, wage-bargaining and capital-labour substitution. *Cambridge Journal of Economics* 23, 413-425.
Sen, Amartya. 1973. *On Economic Inequality.* Oxford: Oxford University Press, edition 1997.
Stiglitz, J.E. 1969. "The Distribution of Income and Wealth among Individuals", *Econometrica* 37: 382-397.
Summers,R. and Heston, A. 1991. "The Penn World Table(Mark 5):An expanded set of international comparisons,1950–1988," *Quarterly Journal of Economics* 106(2): 327–368.
Tinbergen, J. 1975. *Income Distribution.* Amsterdam: North Holland.