LABOUR LOSING TO CAPITAL – SUPPORTING MATERIAL FOR CHAPTER 3 OF THE 2012 *OECD EMPLOYMENT OUTLOOK*

The following pages provide supplementary material for Chapter 3 of *OECD Employment Outlook 2012*. This material is organised into 3 annexes (Annexes 3.A2, 3.A3 and 3.A4).

ANNEX 3.A2 THE LABOUR (AND CAPITAL) INCOME OF TOP INCOME EARNERS

Trends in the income share of top earners.

Table 3.A2.1 presents recent data on the evolution of the income share of top 1% earners in total income for the OECD countries covered in the “World Top Incomes Database” (Atkinson, Piketty, and Saez, 2011).

Table 3.A2.1. **The share of top 1% earners’ income in total income, mid 70s-mid2000s**[a]

Percentages

	Mid 70s	1990	Mid 2000s
Australia	5.0	6.4	9.7
Canada	8.2	9.3	12.8
Denmark	4.0	4.1	4.3
Finland	5.7	4.6	8.1
France	8.2	8.0	8.7
Ireland	5.8	7.3	9.8
Italy	7.0	7.8	9.2
Japan	6.9	7.6	9.0
Netherlands	6.1	5.5	5.4
New Zealand	6.7	8.2	9.5
Norway	5.4	4.8	8.2
Portugal	7.1	7.4	9.5
Spain	7.6	8.5	8.8
Sweden	5.0	4.8	6.9
United States	7.9	12.9	18.0
Unweighted average	6.5	7.2	9.2

a) Data refer to three-year rolling averages starting with the earliest year for the mid 70s (generally 1975, except for Denmark, for which they refer to 1980, Portugal, 1976 and Spain, 1981) and 1990 (for all countries); they refer to rolling averages ending with the latest year for mid 2000 (1999 for Netherlands, 2000 for Canada and Ireland, 2004 for Finland and Italy, 2005 for Denmark, France, Japan, New Zealand and Portugal, 2007 for Australia, 2008 for Norway, Spain and the United States, 2009 for Sweden).

Source: World Top Incomes Database.

The data reinforce available evidence (*e.g.* OECD, 2008 and 2011) that top earners have experienced a sizable increase in the share of income in many OECD countries. The rise has been somewhat larger in English-speaking countries, with the biggest shift occurring in the Unites States, where the share has increased by 10 percentage points from the mid-70s and by 5 since 1990. The evolution is less pronounced for European countries and Japan, which have experienced an increase of respectively 2 and 1.4 percentage points since 1990.

In the "World Top Incomes Database" there are also time-series on the share of capital and wage income in total income of top-earners. This allows shedding some light on the evolution of the sources of total top earners' income. In the table below the evolution of the share of wage income on total income for top 1% earners is shown for the OECD countries for which data are available. The data suggest an upward trend between 1990 and the mid 2000s in half of the countries. This is more pronounced in Canada – 7.1 percentage points – and the Netherlands – 13.4 percentage points – less dramatic in France and Japan, with an increase of 0.5 and 2.8 percentage points, respectively. For the Unites States, although decreasing since 1990, no clear long-term trend can be detected as the share fluctuates around 55% between the 70s and the end of the 2000s. The same pattern is observed for Australia. Italy and Spain, by contrast, appear on a steadier declining trend.

Table 3.A2.2. **The share of wage income in total top 1% earners' income, mid 70s-mid2000s**[a]

Percentages

	mid 70s	1990	mid 2000s
Australia	51.1	59.3	39.4
Canada	52.9	58.6	65.6
France	46.4	47.8	48.3
Italy	42.2	38.9	38.8
Japan	73.9	78.6	81.4
Netherlands	43.4	54.2	67.6
Spain	58.9	53.5	48.4
United States	54.5	58.9	54.5

a) Data refer to three-year rolling averages starting with the earliest year for the mid 70s (1975 for Canada, France, Netherlands and the United States, 1976 for Australia, Italy and Japan, 1981 for Spain) and 1990 (for all countries except Japan, 1991); they refer to rolling averages ending with the latest year for mid 2000 (1999 for Netherlands, 2000 for Canada, 2004 for Italy, 2005 for France and Japan, 2007 for Australia, 2008 for Spain and the United States).

Source: World Top Incomes Database.

Decomposing top earners' income shares: the role of wage income, capital income and the labour share

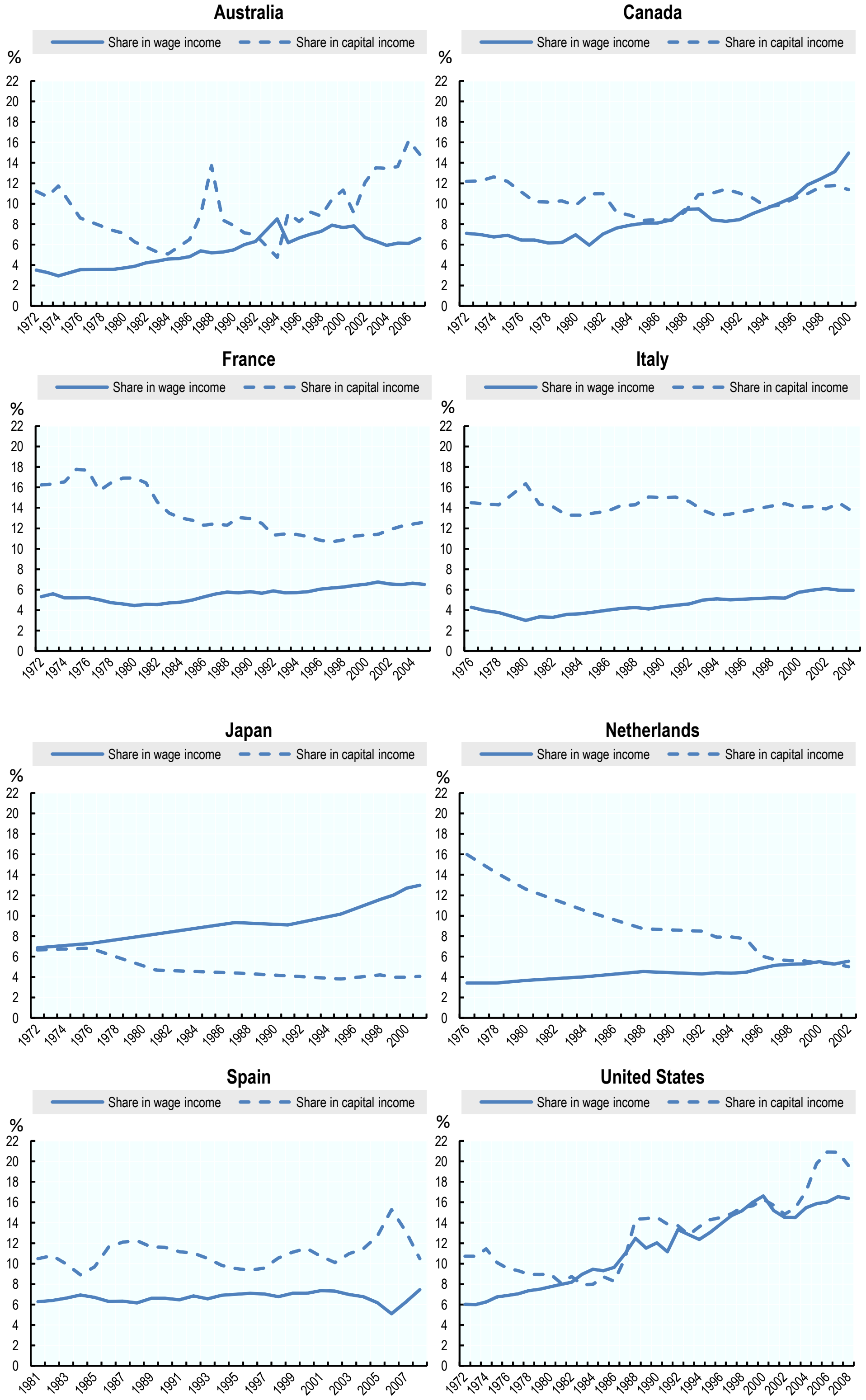
Recent evolution of the income shares of top 1% earners in total wage and capital income

As mentioned above, detailed information on the share of top earners in total capital and wage income is not available in the "World Top Incomes Database". However, one can obtain a reliable estimate of these shares by considering the following equation:

$$I_{ti}^{1} = \frac{W_{ti}^{1}}{INC_{t}^{1}} * \frac{INC_{t}^{1}}{INC_{t}} * \frac{INC_{t}}{W_{ti}} \text{ [A2.1]}$$

where i is wage or capital, t stands for time, I_i is the share of wage/capital top 1% earners income in total wage/capital income, W_i^1 represents the wage/capital income of top earners and INC^1 their total income, while INC_i and W_i are aggregate total and wage/capital income, respectively, of the total population. The first two terms on the right-hand side are available in the "World Top Incomes Database", while the third is the inverse of the wage (capital) share which can be derived from EU-KLEMS and OECD-STAN (as in Figure 3.1) by making the assumption that wage (capital) shares in total value added and domestic income are similar. Figure 3.A2.1 illustrates the results from this exercise for the full period covered by both databases.

Figure 3.A2.1. **Shares of top 1% earners in total wage and capital income, mid 70s-mid2000s**



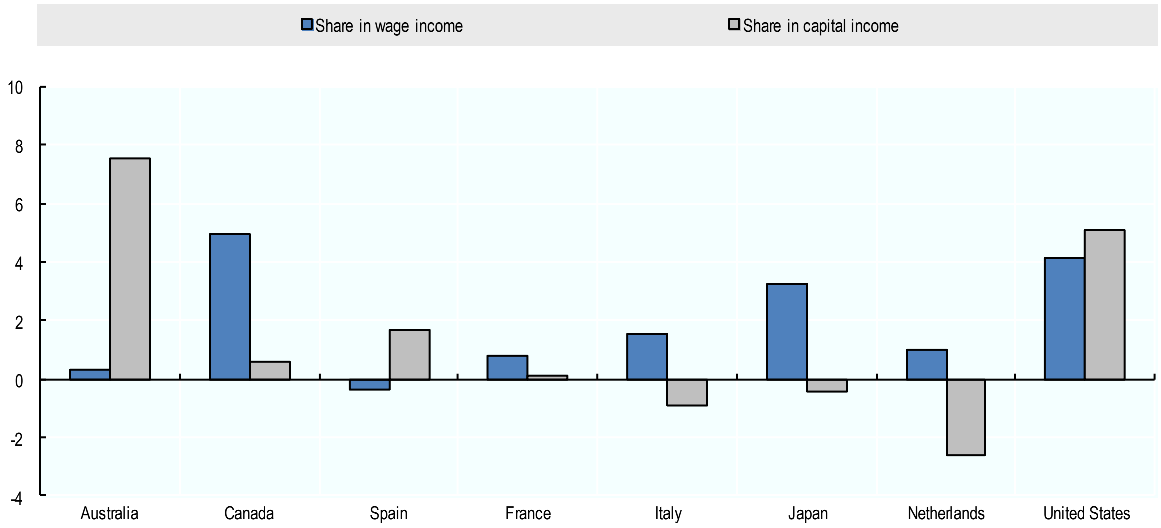
Source: OECD calculations based on the World Top Incomes Database, OECD STAN and EU-KLEMS.

The share of top earners in total capital income exhibits divergent trends across the OECD countries for which data are available. Remarkable upward trends are observed for the United States and Australia, countries in which the share of top 1% earners in capital income has reached respectively 19.5% and 15% of aggregate capital income in the late 2000s. In other OECD countries, trends were not as dramatic: this share was slightly decreasing, in France and the Netherlands, while remaining essentially stable in Canada, Italy, Japan and Spain.

Concerning the share of top earners in wage income, a clear common upward shift emerges for almost all OECD countries, especially focusing on the subsample 1990-mid 2000s (Australia being the only exception). As Figure 3.A2.2 shows, the share of top earners in aggregate wage income has significantly increased on average by 2 percentage points, with Canada, Japan and the United States experiencing the largest growth – 4.9, 3.2 and 4.1 percentage points, respectively. These findings are consistent with what suggested in Section 3.2 of the main text. Capital deepening and skilled-biased technical change have presumably caused a contraction in the share of low and medium-educated workers in labour compensation at the advantage of capital and high-skilled labour.

Figure 3.A2.2. **Changes in the shares of top 1% earners in total wage and capital income, 1990-mid 2000s*[a]***

Percentage points



a) See Table 3.A2.2 for exact dates.

Source: OECD calculations based on the World Top Incomes Database, OECD STAN and EU-KLEMS.

The decomposition of income shares of top 1% earners in within and between components

The income share of top earners in total income can be decomposed into within and between components, by means of a shift-share decomposition similar to the one used for examining the evolution of the overall labour share (see Annex 3.A3). Defining ΔT as the time difference in the share of top earners in total income, one can write it as:

$$\Delta T = T_t - T_{t-1} = \sum_i \bar{F}_i * (I_{ti} - I_{it-1}) + \sum_i \bar{I}_i * (F_{ti} - F_{it-1}) \qquad \text{[A2.2]}$$

where i stands for capital or wage, t for time, F_i represents the share of capital (wage) income in total income (i.e. the aggregate capital and labour shares), and I_i the share of capital (wage) income of top earners' in total capital (wage) income, as derived from [A2.1]. The first term on the right-hand side of

[A2.2] is the within-component, which captures the evolution of top earners' income within capital and labour compensation, and consists of the series discussed in the previous section. A relative gain (loss) in the relative shares of top earners' income results in an increase (decrease) of the top earners' share in total income. Conversely, the second term is the between-component which illustrates the implications of changes in the aggregate labour share on the share of national income appropriated by top earners, holding constant their relative shares within each source of income. A decrease in the aggregate labour share would produce a mechanical redistribution of income in favour of the top earners if capital is the main source of top earners' income in comparison with that of the whole population – that is, if the share of top earners in total capital income is larger than their share in total labour income.

As it can be evinced from Figure 3.A2.3, Panel A, the within component is the main driver of the evolution of the overall share of top earners in total income in all OECD countries considered. Wage income growth accounts, on average, for 60% of the within variation, and for 50% of the total variation. These figures would be higher if Australia and the United States had been excluded from the sample, since in these two countries the contribution of the variation in top earners shares in capital income is large (2.8 and 2.5 percentage points, respectively). By contrast, the contribution of trends in the aggregate labour share – the between component – seems minor, at least for the countries considered, apart from Italy. In sum, the drivers of rising income inequality are mostly trends in the distribution of capital and wage income between top and bottom earners and not trends in factor compensation *per se*. The way total income is distributed between labour and capital does not directly impact on the distribution of income between top earners and the rest of the population. Put another way, these figures suggest that the opposite income trends – and particularly wage income trends – for top and bottom earners are the key determinants of increasing inequality in total income. These considerations do not change if top 10% earners are looked at (Panel B).

Trends in the labour share of top earners

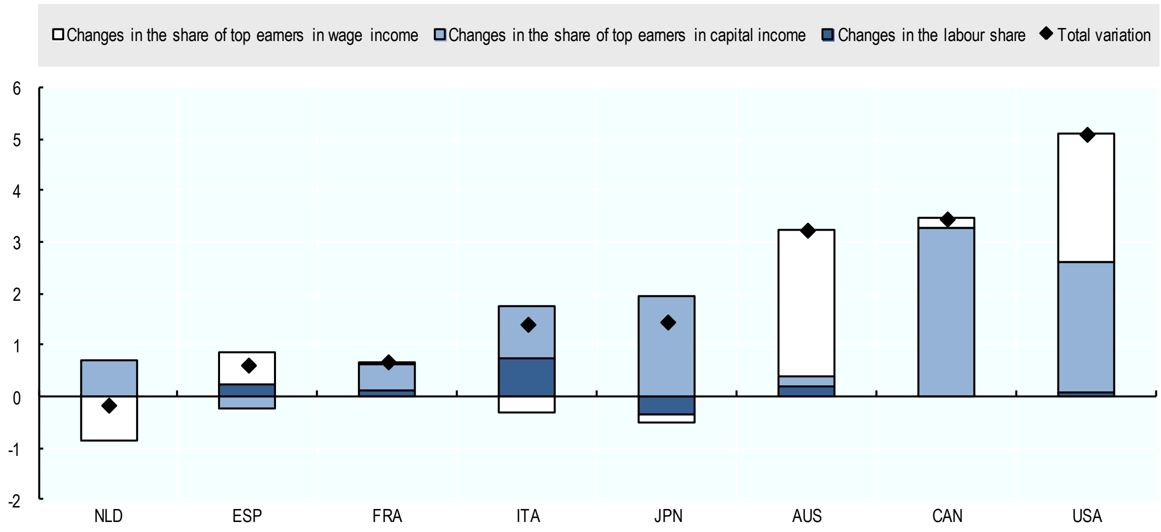
The product of the first two terms on the right-hand side of [A2.1] provides an estimate for the share of top 1% earners in GDP, except for discrepancies between national income and domestic value added. The results from this exercise are summarised in Figure 3.A2.4.

Obviously, subtracting the share of top 1% earners in total output from the aggregate labour share yields an estimate for the share of bottom 99% earners in total output, which is shown in Box 3.1 in the main text. As discussed in there, once top earners are removed from the wage bill, the drop of the labour share appears somewhat greater, especially in Canada and the United States. In many countries, this is due to diverging trends in aggregate and top earners labour shares. Top 1% earners saw their labour share increasing since the mid-70s in most of the countries considered, as Figure 3.A2.4 shows. The rise was substantial for Canada, Japan and the United States, where the shares increased respectively by 4.3, 2.4 and 6 percentage points since the 1970s and by 3.7, 1.5 and 2.3 points since 1990. Conversely the evolution was less pronounced in European countries and Australia.

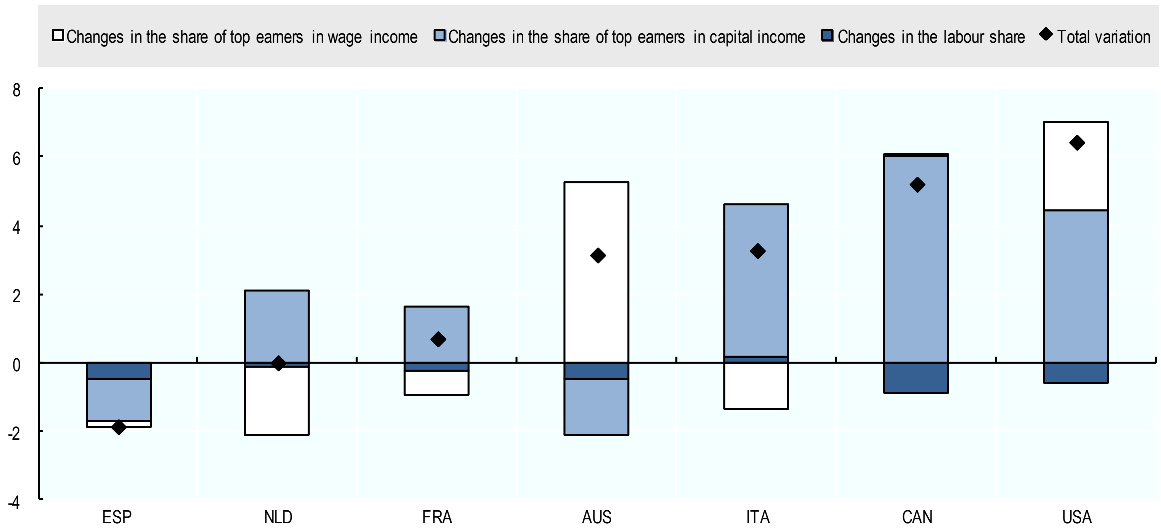
Figure 3.A2.3. **Within and between components of the evolution in the share of top 1% earners in total income**

Percentage-point contributions to changes to the share top 1% earners in total national income, 1990-mid 2000s[a]

Panel A: Top 1% earners



Panel B: Top 10% earners

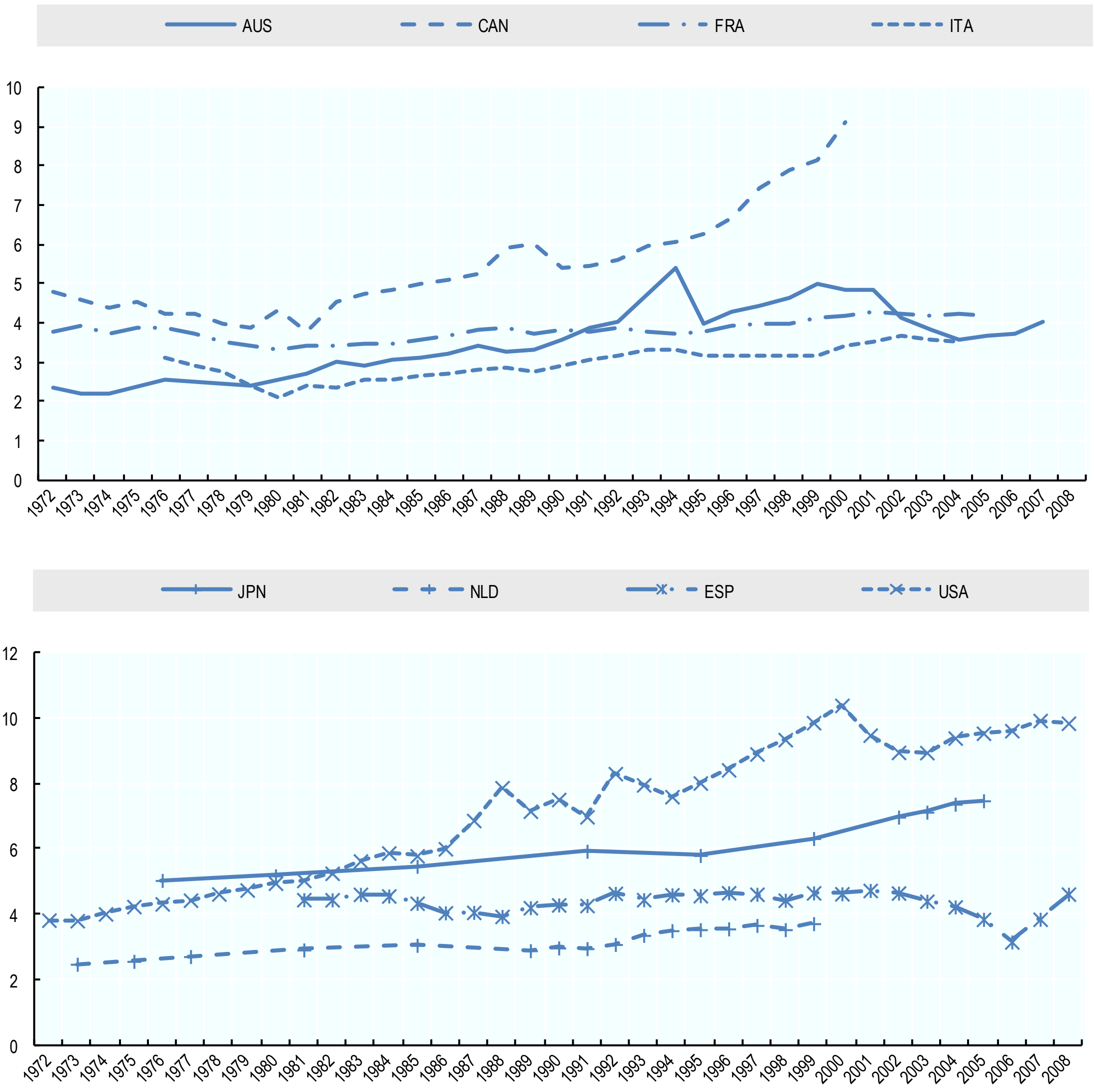


a) See Table 3.A2.2 for exact dates.

Source: OECD calculations based on the World Top Incomes Database, OECD STAN and EU-KLEMS.

Figure 3.A2.4. **Top 1% earners labour share evolution in selected OECD countries, mid-70s-mid 2000s**[a]

Percentages



a) Data have been interpolated for Australia in 1975, 1977, 1978 and 1993; for Italy in 1996-1997; for Japan in 1977-1979, 1981-1984, 1986-1990, 1992-1994, 1996-1998 and 2000-2001; for the Netherlands in 1974,1976, 1978-1980, 1982-1984 and 1986-1988.

Source: OECD calculations based on the World Top Incomes Database, OECD STAN and EU-KLEMS.

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ANNEX 3A.3 SHIFT-SHARE DECOMPOSITIONS

The shift-share methodology allows decomposing aggregate changes of an economic variable in the contribution due to changes of that variable within industries and structural changes in industry composition. Formally, in the case of the labour share, a shift-share decomposition can be written as:

$$F_t - F_{t-1} = \sum_i \bar{s}_i \left(f_{it} - f_{it-1} \right) + \sum_i \left(s_{it} - s_{it-1} \right) \bar{f}_i \qquad \text{[A3.1]}$$

where F and f represent the aggregate and industry labour shares, s is the share of industry i in nominal value added and a bar represents averages between start and end period. The first term on the right-hand side is a weighted average of *within-industry* changes in the labour share while the last term represents the contribution of sectoral reallocation towards or against high-labour-share industries (the so-called *between-industry* component).

The evolution of the labour share in each industry can also be linked to the different evolution of real wages, labour productivity and relative prices (see *e.g.* De Serres *et al.*, 2002, Torrini, 2005). In particular, using logarithmic approximations:

$$\log \frac{F_t}{F_{t-1}} = \log \frac{W_t}{W_{t-1}} - \log \frac{Y_t}{Y_{t-1}} + \log \frac{(P_t / D_t)}{(P_{t-1} / D_{t-1})} \qquad \text{[A3.2]}$$

that is, the percentage change in the aggregate labour share F can be decomposed in the percentage growth of aggregate real gross hourly wage W (deflated with the consumption deflator P) minus the percentage growth in hourly productivity Y (in volumes, that is value added per hour, deflated with the aggregate value added deflator D) and the percentage change in the relative price of consumption with respect to domestic output. As suggested by Böckerman and Maliranta (2012) one can make use of the above formula to extend the standard shift-share decomposition in order to shed light on the relative contributions of wages, productivity and prices to within-industry and between-industry variations of the labour share. More precisely, aggregate real wage and productivity growth can be decomposed as:

$$\log \frac{W_t}{W_{t-1}} \cong \sum_i \bar{h}_i \log \frac{w_{it}}{w_{it-1}} + \sum_i \bar{h}_i \log \frac{w_{it}}{w_{it-1}} \left(\frac{\bar{w}_i - \bar{W}}{\bar{W}} \right) + \sum_i \left(h_{it} - h_{it-1} \right) \frac{\bar{w}_i}{\bar{W}} \qquad \text{[A3.3]}$$

$$\log \frac{Y_t}{Y_{t-1}} \cong \sum_i \bar{h}_i \log \frac{v_{it}}{v_{it-1}} + \sum_i \bar{h}_i \log \frac{v_{it}}{v_{it-1}} \left(\frac{\bar{v}_i - \bar{V}}{\bar{V}} \right) + \sum_i \left(h_{it} - h_{it-1} \right) \frac{\bar{v}_i}{\bar{V}} \qquad \text{[A3.4]}$$

where lowercase letters represent industry-level variables, h stands for the share of industry i in total hours worked and V and v represent nominal value added per hour worked at the aggregate and industry level, *both* deflated by the aggregate value added deflator (the same deflator is used to preserve additivity).[1] The

1 . The formulas are exact for standard growth rates and only approximated for log differences.

first term on the right-hand side of [A3.3] represents a weighted average of *within-industry* wage growth rates. The second term on the right-hand side captures the contribution to aggregate wage growth of the covariance of wage levels and growth: the larger the wage growth in high-wage industries and the larger the aggregate wage growth. This term is often called *convergence/divergence* component. The third term on the right-hand side captures the structural reallocation towards or against high-wage industries, and has the same interpretation as the *between* component in the standard shift-share analysis. The same interpretation holds as regards productivity for the terms on the right-hand side of [A3.4].

From [A3.4] after some manipulations we obtain:

$$\log\frac{Y_t}{Y_{t-1}} \cong \sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}} + \sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right) + \sum_i \left(h_{it} - h_{it-1}\right)\frac{\bar{v}_i}{\bar{V}} +$$
$$\sum_i \bar{h}_i \log\frac{(d_{it}/D_t)}{(d_{it-1}/D_{t-1})} + \sum_i \bar{h}_i \log\frac{(d_{it}/D_t)}{(d_{it-1}/D_{t-1})}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right)$$

where d is the value added deflator of industry i. Taking into account that $\sum_i \bar{h}_i = 1$ and $\sum_i \bar{h}_i \bar{v}_i \cong \bar{V}$, we have:

$$\log\frac{Y_t}{Y_{t-1}} - \log\frac{(P_t/D_t)}{(P_{t-1}/D_{t-1})} \cong$$
$$\sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}} + \sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right) + \sum_i \left(h_{it} - h_{it-1}\right)\frac{\bar{v}_i}{\bar{V}} + \qquad \text{[A3.5]}$$
$$\sum_i \bar{h}_i \log\frac{(d_{it}/P_t)}{(d_{it-1}/P_{t-1})} + \sum_i \bar{h}_i \log\frac{(d_{it}/P_t)}{(d_{it-1}/P_{t-1})}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right)$$

Combining [A3.2], [A3.3] and [A3.5], this implies that the percentage change in the wage share can be decomposed as follows:

$$\log\frac{F_t}{F_{t-1}} \cong$$
$$\left[\sum_i \bar{h}_i \log\frac{w_{it}}{w_{it-1}} + \sum_i \bar{h}_i \log\frac{w_{it}}{w_{it-1}}\left(\frac{\bar{w}_i - \bar{W}}{\bar{W}}\right) + \sum_i \left(h_{it} - h_{it-1}\right)\frac{\bar{w}_i}{\bar{W}}\right] -$$
$$\left[\sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}} + \sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right) + \sum_i \left(h_{it} - h_{it-1}\right)\frac{\bar{v}_i}{\bar{V}}\right] - \qquad \text{[A3.6]}$$
$$\left[\sum_i \bar{h}_i \log\frac{(d_{it}/P_t)}{(d_{it-1}/P_{t-1})} + \sum_i \bar{h}_i \log\frac{(d_{it}/P_t)}{(d_{it-1}/P_{t-1})}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right)\right] +$$
$$\left[\log\frac{(D_t/P_t)}{(D_{t-1}/P_{t-1})}\sum_i \bar{h}_i\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right)\right]$$

The first term in brackets on the right-hand side of [A3.6] represents the real wage effect that can be decomposed into within, convergence/divergence and between components; the second term in brackets represents a real productivity effect that can be again decomposed into within, convergence/divergence and between effects; the third term represents a relative price effect that can be decomposed into within-industry changes of value added prices with respect to the consumption deflator and a convergence/divergence term which has a negative contribution to the wage share dynamics if industries with high value added per hour worked have the largest growth in relative prices; finally, the fourth term represents a residual, which is in practice small in the data.[2]

Re-arranging the terms in [A3.6] one obtains:

$$\begin{aligned}\log\frac{F_t}{F_{t-1}} \cong &\left[\sum_i \bar{h}_i\left(\log\frac{w_{it}}{w_{it-1}} - \log\frac{y_{it}}{y_{it-1}}\right)\right] + \\ &\left[\sum_i \bar{h}_i \log\frac{w_{it}}{w_{it-1}}\left(\frac{\bar{w}_i - \bar{W}}{\bar{\bar{W}}}\right) - \sum_i \bar{h}_i \log\frac{y_{it}}{y_{it-1}}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right)\right] + \\ &\left[\sum_i \bar{h}_i \log\frac{(P_t/d_{it})}{(P_{t-1}/d_{it-1})} + \sum_i \bar{h}_i \log\frac{(P_t/d_{it})}{(P_{t-1}/d_{it-1})}\left(\frac{\bar{v}_i - \bar{V}}{\bar{V}}\right) + o\right] + \\ &\left[\sum_i \left(h_{it} - h_{it-1}\right)\left(\frac{\bar{w}_i}{\bar{\bar{W}}} - \frac{\bar{v}_i}{\bar{V}}\right)\right]\end{aligned} \qquad [A3.7]$$

which is the equation used in Box 3.3 and the rest of the chapter (with *o* indicating the small residual term of [A3.6]). The term in the first bracket represents the contribution to the evolution of the aggregate labour share of the average relative within-industry growth of real wages with respect to productivity. The terms in the second bracket capture the contribution of convergence/divergence patterns in real wages and productivity: real wages provide a greater contribution to the percentage change of the labour share when they grow faster in high-wage industries; conversely when productivity grow faster in high-productivity industries, this compresses the labour share. The terms in the third bracket represents the relative price effect, which is positive if, on average, the consumption deflator grows faster than the output deflators. Finally, the term in the fourth bracket captures the reallocation of labour from/to industries that are relatively more high-wage or high-productivity: it is positive if the difference between expanding industries and the average industry as regards wages is larger than their difference as regards productivity – or, in other words, expanding industries tend to be more high-wage than they are high-productivity industries. The interesting feature of this decomposition is that it allows to single out simultaneously three factors that, to a different extent in different countries, appears to be key in determining the within-industry evolution of the labour share in the business sector: the fact that, on average, within-industry real wage growth did not keep pace with real productivity growth, the role of relative price effects, and the correlation between growth and levels of wages and productivity. The latter factor represents another, more dynamic, type of structural shift within an economy: if the growth rate of real wages is relatively homogeneous across industries while productivity grows faster in high productivity industries, this inevitably depresses the labour share.

2 . Although in practice this term is small in this case, it could, in theory, significantly diverge from zero if both these conditions held true: i) aggregate deflators of value added and consumption diverged and ii) employment growth during the period had markedly opposite correlations with value added shares at the beginning and end of the period.

A similar shift-share analysis can also be used to investigate the relative importance of within-industry changes and sectoral shifts in the evolution of the share of workers with different levels of educational attainment in labour compensation. As for [A3.3] and [A3.4],[3] the percentage change of the labour share for the low-educated (medium-educated) can in fact be decomposed into the variation of the share within industries, a convergence/divergence term, capturing the covariation between levels and changes of the share as above, and the shift in the industrial structure against (or towards) the low-educated (medium-educated):

$$\frac{L_t - L_{t-1}}{\bar{L}} = \sum_i \bar{x}_i \frac{l_{it} - l_{it-1}}{\bar{l}_i} + \sum_i \bar{x}_i \frac{l_{it} - l_{it-1}}{\bar{l}_i}\left(\frac{\bar{l}_i - \bar{L}}{\bar{L}}\right) + \sum_i \left(x_{it} - x_{it-1}\right)\frac{\bar{l}_i}{\bar{L}}$$

where L and l are the aggregate and industry level shares of low-educated (medium-educated) labour in labour compensation and x is the share of industry i in total compensation. However, if workers with different skills were perfectly substitutable and relative productivity were constant (with simply high-educated workers being more productive than low-educated workers by a constant factor), the evolution of the shares by educational attainment will simply match the trends in the relative size of each subpopulation. [4] This implies that if the subpopulation of low-educated (medium-educated) workers contracts, a reduction in their shares in labour compensation might not be a symptom of the worsening of their position in the labour market. Adjusting labour shares by level of education for the relative supply of workers of different types represents therefore an interesting benchmark: if the share of one group falls by more than its reduction in the population, this suggests that the position of that group worsened, no matter what assumption is made on the substitutability across groups. Taking into account that $\sum_i \bar{x}_i = 1$, this yields a decomposition into four terms:

$$\frac{L_t - L_{t-1}}{\bar{L}} = \sum_i \bar{x}_i \left(\frac{l_{it} - l_{it-1}}{\bar{l}_i} - \frac{\mathcal{P}_t - \mathcal{P}_{t-1}}{\bar{\mathcal{P}}}\right) + \sum_i \bar{x}_i \frac{l_{it} - l_{it-1}}{\bar{l}_i}\left(\frac{\bar{l}_i - \bar{L}}{\bar{L}}\right) + \sum_i \left(x_{it} - x_{it-1}\right)\frac{\bar{l}_i}{\bar{L}} + \frac{\mathcal{P}_t - \mathcal{P}_{t-1}}{\bar{\mathcal{P}}}$$

where $\mathcal{P}$ represents the relative share of the low-educated (medium-educated) in the population. The first term in brackets on the right-hand side represents the contribution of within industries changes in compensation shares with respect to changes in the relative size of populations; the second term captures the covariation of levels and changes in compensation shares, the third terms represents the effect of changes in the industry structure, while the last one is the change of the relative size of the subpopulation of the group.

3 . See Böckerman and Maliranta (2012) for the derivation.

4 . The greater the degree of substitutability of different types of workers, the greater the effect of their relative supply in determining their shares in labour compensation. For example, if the elasticity of substitution across workers with different educational attainment is close to 1, the evolution of education shares is independent from the supply of different types of labour.

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ANNEX 3A.4 THE IMPACT OF EMPLOYMENT PROTECTION AND THE MINIMUM WAGE ON PRODUCTIVITY GROWTH AND THE LABOUR SHARE: DETAILED ESTIMATION RESULTS

Empirical evidence presented in the chapter suggests that in the two decades preceding the recent crisis, measured TFP growth reflecting capital-augmenting technical change was one of the most important factors explaining the decline of the labour share. In particular, the spread of the ICT-based technological paradigm is likely to be the key explanation for this rapid growth. Indeed, it appears to have created opportunities for unprecedented advances in innovation and invention of new capital goods and production processes, thereby boosting productivity together with high substitution between capital and labour.

Labour market policies, however, affect the relative prices of capital and labour, and are therefore likely to have played a role in this process. For example, in the past ten years, the ratio of statutory minimum to median wages has increased of about 2 percentage points, in countries where statutory minima exists. In turn, this might have induced firms to overinvest in labour-saving innovation (see *e.g.* Boone, 2000), thereby lowering the labour share through the channels discussed in the previous sections.

Minimum wage

The simultaneous impact of statutory minimum wages on TFP growth and the labour share is analysed here by estimating a variant of the model presented in Box 3.4 (equation [**]) augmented by the ratio of statutory minimum to median wages.[5] However, one complication in estimating the impact of the minimum wage is that its level is the same in all industries; therefore its average impact cannot be identified if country-by-time dummies are included in the specifications. In order to address this issue, one key identifying assumption is made. Minimum wages are more likely to prevent downward adjustment of wages for workers that are paid the minimum wage or only slightly more. As a consequence, it is assumed here that the industries that are more likely to be severely affected by any change in the minimum wage are those that, because of their technological characteristics, have a greater natural propensity to rely on low-wage labour in the absence of a minimum wage. This justifies including in the specification an interaction between this propensity and the statutory minimum wage, while simultaneously controlling for aggregate effects through country-by-time dummies. The advantage of this procedure is that these dummies control for other aggregate effects, including other institutions whose effects do not depend on how much each industry relies on low-wage labour. As a consequence, endogeneity issues and omitted variable problems, typically important in regressions with aggregate institutional variables, are not likely to bias severely estimated results (see Bassanini *et al.*, 2009, for a more extensive discussion). Following Bassanini *et al.* (2010) and Bassanini (2011), in order to reduce biases due to the possible relationship between minimum wages and the distribution of low-wage employment, the propensity of industries to employ low-wage labour is proxied with the incidence of workers with less than secondary education by industry in the

5 . Annual data on bargained minimum wage, as set in collective agreements, are not available. For this reason the analysis of the impact of minimum wages developed in this section relies only on countries with statutory minima. Due to limitations in TFP data availability, the sample is reduced to 8 countries only.

United Kingdom prior to the introduction of statutory minimum wages in 1999 – when there was virtually no floor on wages, except for constraints imposed by collective bargaining.[6]

The same methodology is used to study the impact of the minimum wage on TFP growth. However, as standard in the literature (see *e.g.* Aghion and Howitt, 2006; Griffith *et al.*, 2004), the equation is specified in first differences and the ratio of the TFP level in a given country and industry to the TFP level of the leader of that industry – relative TFP hereafter – is also included in the specifications. This implies that the following equation is estimated:

$$\begin{aligned}\Delta \log TFP_{ijt} = -\phi \log RTFP_{ijt-1} + \beta\Phi_j MW_{it-1} + \gamma\Phi_j \Delta MW_{it} + \\ + X_{ijt}\delta + \mu_{jt} + \mu_{it} + \varepsilon_{ijt}\end{aligned} \qquad \text{[A4.1]}$$

where Δ represents annual changes, *TFP* stands for a measure of level TFP whose changes can noisily proxy for capital-augmenting technical change, *RTFP* and *MW* stands for relative TFP and the ratio of minimum to median wages, respectively, Φ stands for the incidence of low-wage labour (based on UK indicators), X is a vector of other labour-share determinants and controls that vary by country i, industry j and time t, η are industry-by-time and country-by-time effects, ε is an error term and other Greek letters are parameters to be estimated. β and γ capture the effect of the minimum wage on cross-industry differences in TFP growth rates and levels, respectively. Note that, in equation [A4.1], country-by-time dummies control for all aggregate effects, including the average effect of *MW* and ΔMW, which are not therefore included in the specification. In practice, in the spirit of Rajan and Zingales (1998), estimates of β and γ in [A4.1] can be interpreted as difference-in-difference estimates, where industries with low incidence of low-educated workers act as control group for industries where the share of the low-educated is larger. This implies that estimates of β and γ can be used to infer the direction of the average overall impact of *MW* on TFP growth and levels, going beyond simple cross-industry differences in the impact of *MW*, provided that, as assumed, in industries with a lower natural propensity to rely on low-wage labour in the absence of a minimum wage the effect of the minimum wage is less strong but of the same sign (or zero) than in industries with greater propensity.

As one of the motivations of the analysis of this section is that the minimum wage is expected to affect the labour share by raising incentives for labour-saving innovations, it seems preferable to estimate the effect of the former on the labour share without including productivity or TFP as a further control. Therefore, variants of [A4.1], obtained by simply replacing the dependent variable (namely, the labour share, labour productivity, wages and price deflators), are used in the whole analysis of this section.

Last but not least, although most endogeneity concerns are taken care of by the presence of aggregate dummies, one can still worry that the ratio of minimum to median wages, even interacted with industry indicators, could still be endogenous. In particular, TFP growth might raise wages more at the top of the wage distribution. This effect, if any, is likely to be more important in industries that are intensive in high-skilled labour. But these industries typically employ fewer workers with little education. As a consequence an increase in TFP growth in these industries might particularly raise median wages in the whole economy and, in that case, β and γ in [A4.1] might simply capture this reverse causation. Thus, an instrumental variable approach is adopted whenever endogeneity tests reject the null of exogeneity.[7] In that case, the

6 . These statistics are based on 1994-1998 data from UK LFS. Results presented here are robust to two alternative indicators of the incidence of low-wage labour, namely the share of workers with less than upper secondary education and the share of workers in low-pay (less than 2/3 of the median of national wages).

7. Two-stage least squares are used throughout.

logarithm of the deviation of the real minimum wage in 2 000 US dollars PPP from the OECD average is used as an instrument for the ratio of the minimum wage to median earnings.[8]

As shown in Columns 1 and 2 of Table 3.A4.1, in the short-run the impact of minimum wages on TFP and productivity levels appears negative. Columns 3 and 4 provide some possible explanation for this effect: in the short-run, an increase in the minimum wage raises the average wage. However, this effect is entirely reflected in higher prices, which result in lower real sales per hour worked. In nominal terms, however, equilibrium sales also grow (due lower quantities but much greater prices), so that the short-run impact on the labour share is insignificant (Column 5). In the long-run, however, employers react by increasing efficiency levels and productivity.[9] However, no significant long-run effect on prices or wages is estimated, a result consistent with either greater minimum wages inducing faster labour-saving technical change or more firm-sponsored training or both – to the extent that training benefits are mostly reaped by firms.[10] In turn, long-run productivity growth effects are estimated to more than compensate the initial wage effect, which results in a negative long-run impact on the labour share.

Table 3.A4.1**. Minimum wage, productivity and the labour share**

Difference-in-difference estimates

Dependent variables	(1) TFP	(2) Labour productivity	(3) Wage	(4) Price	(5) Labour share	(6) TFP	(7) Labour share
MW x 94-98 UK low-educ. share	0.521***	0.579***	0.151	-0.114	-0.240***	0.563***	-0.244***
	(3.610)	(3.790)	(1.283)	(-0.894)	(-2.668)	(3.861)	(-2.683)
∆MW x 94-98 UK low-educ. share	-3.417**	-3.197**	1.904*	5.016***	0.308		
	(-2.274)	(-1.996)	(1.876)	(3.758)	(0.340)		
Observations	3,294	3,294	3,294	3,294	3,294	3,294	3,294
R-squared	0.349	0.358	0.378	0.467	0.364	0.349	0.364
Country-by-time dummies	yes	yes	yes	yes	yes	yes	yes
Industry-by-time dummies	yes	yes	yes	yes	yes	yes	yes

Notes: All dependent variables are in log differences. All equations control for log-relative TFP. MW: Ratio of minimum to median earnings. 94-98 UK low-educ. share: industry share of those with less than secondary education in the United Kingdom, averaged over 1994-1998 (average value: 0.192). 2SLS estimates. The logarithm of the deviation of the real minimum wage in 2 000 US dollars PPP from the OECD average is used as an instrument for the ratio of the minimum wage to median earnings. Robust t-statistics in parentheses. ***, **, *: significant at the 1%, 5% and 10% level, respectively.

8 . F-test statistics on the significance of the instrument in first-stage regressions are always above 30, which is well above standard minimum thresholds for instrument validity.

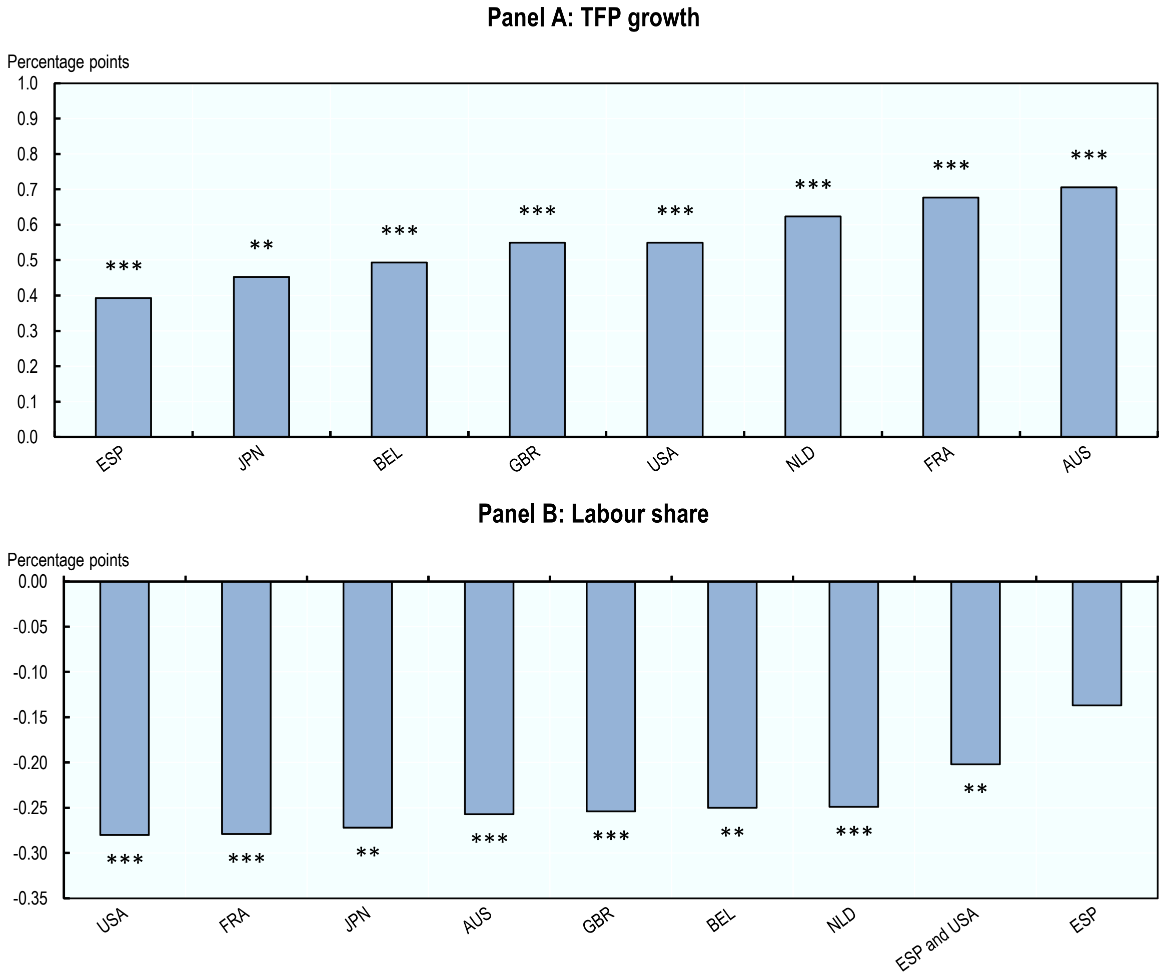
9 . This is consistent with findings of Bassanini and Venn (2007) who, nonetheless, were unable to discriminate whether their finding resulted from from changes in workforce composition induced by the minimum wage or from greater minimum wage inducing greater efficiency. As skill composition is, in principle, taken into account in the TFP measure used here – the TFP measure in EUKLEMS relies on twelve different types of labour, it can be argued that what is estimated here is an efficiency effect.

10 . Theoretical and empirical research suggests that, by compressing the lower tail of the wage distribution without necessarily affecting pre-training individual productivity, minimum wages could increase employers' incentive to pay for training as they can reap the difference between productivity and wage growth after training (see *e.g.* Acemoglu and Pischke, 1999, 2003; Arulampalam *et al.*, 2004). Once the infrastructure for training is in place (that is, the fixed cost has been paid), it is likely to be used also for workers paid above the minimum, in particular if the quality of training can only be imperfectly signalled and valued on the external labour market, so that the firm will reap most of the productivity gains from training.

The key results of Table 3.A4.1 appear also robust to the exclusion of short-run terms in the specification (Columns 6 and 7). Moreover, they are reasonably robust to elimination of countries one-by-one from the sample, taking into account that the number of countries is small (Figure 3.A4.1). In particular, estimates are insignificant if Spain is excluded from the sample but become significant again if other countries such as France or the United States are excluded.

Figure 3.A4.1. **Sensitivity of the effects of the minimum wage to countries included in the sample**

Point-estimates of the coefficient of the interaction between the ratio of minimum to median wage and the 1994-1998 UK share of low-educated workers once the indicated countries are excluded from the sample



Note: Coefficients obtained by re-estimating the specification of Columns 6 and 7 of Table 3.A4.1, excluding the indicated country. **, *** significant at 5% and 1% respectively.

It was argued that one of the key advantages of the difference-in-differences approach adopted here is that it allows controlling for other aggregate confounding factors, including other institutions and policies, some of which are not easy to quantify. This claim is correct provided that there is no reason to believe that the impact of aggregate institutions on different dependent variables varies, on average, across industries with different intensity of low-educated workers. In order to provide evidence supporting this claim, the specifications of Table 3.A4.1 are augmented with interactions between our baseline quantitative indicator of propensity to employ low-wage labour and levels of several aggregate indicators of a number of labour market institutions and product market regulations that are typically used in aggregate unemployment equations (Table 3.A4.2)[11] – the indicator of stringency of employment protection, the average labour tax wedge, the average unemployment benefit gross replacement rates (averaged across different durations and family situations), the level of corporatism in collective bargaining, the share of workers covered by collective agreements (including administrative extension) and a time-varying aggregate indicator of the degree of stringency of anti-competitive product market regulation. Indeed, if the identification assumption made here is valid, all the interactions involving other institutions should turn out insignificant.

Table 3.A4.2**. Minimum wage, TFP and the labour share: including additional institutional controls**

Difference-in-difference estimates

Dependent variables	(1) Labour share	(2) Labour share	(3) TFP	(4) TFP	(5) TFP	(6) TFP	(7) TFP
MW x 94-98 UK low-educ. share	-0.156	-0.323***	0.598**	0.525***	0.622***		
	(-0.841)	(-2.726)	(2.314)	(2.586)	(2.945)		
TW x 94-98 UK low-educ. share	-0.003	0.001	0.006**	0.006***	0.006***	0.001	
	(-1.473)	(0.873)	(2.547)	(3.056)	(3.025)	(1.022)	
EP x 94-98 UK low-educ. share	0.015		0.019				
	(0.789)		(0.720)				
UB x 94-98 UK low-educ. share	-0.000		0.000				
	(-0.258)		(0.140)				
Coord. x 94-98 UK low-educ. share	-0.004		0.006				
	(-0.388)		(0.437)				
CB x 94-98 UK low-educ. share	0.001		-0.003*	-0.002***	-0.003***		-0.000
	(0.561)		(-1.850)	(-2.594)	(-4.078)		(-0.086)
PMR x 94-98 UK low-educ. share	0.010		-0.046*	-0.031			
	(0.551)		(-1.902)	(-1.565)			
Observations	2,736	3,168	2,700	3,024	3,024	3,078	3,348
R-squared	0.351	0.348	0.346	0.355	0.354	0.350	0.349
Country-by-time dummies	yes	yes	yes	yes	yes	yes	yes
Industry-by-time dummies	yes	yes	yes	yes	yes	yes	yes

Notes: All dependent variables are in log differences. All equations control for log-relative TFP. MW: ratio of minimum to median wage. TW: average tax wedge. EP: OECD index of employment protection. UB: average unemployment-benefit gross replacement rate. Coord.: index of coordination/corporatism of the wage bargaining. CB: collective bargaining coverage. PMR: time-varying OECD index of anti-competitive product market regulation. 94-98 UK low-educ. share: industry share of the those with less than secondary education in the United Kingdom, averaged over 1994-1998 (average value: 0.192). 2SLS estimates, except Column 1, where OLS are used. The logarithm of the deviation of the real minimum wage in 2 000 US dollars PPP from the OECD average is used as an instrument for the ratio of the minimum wage to median earnings. Robust t-statistics in parentheses. ***, **: significant at the 1% and 5% level, respectively.

In the case of the labour share, the inclusion of all covariates – interacted with the UK intensity of low-educated labour – results in insignificant coefficients for all variables including the minimum wage indicator (Column 1). Given the limited number of countries in the sample this is not entirely surprising and is likely to reflect multicollinearity.[12] Indeed the cross-country variation of institutional variables plays a key role in determining estimated effects. However, once the least significant variables are progressively excluded, the minimum wage remains the only significant variable (Column 2).[13]

11 . This exercise is only performed for levels of institutional variables in order to lessen multicollinearity problems.

12 . Institutional variables are, in fact, strongly correlated across countries;

13 . In addition, none of these additional covariates turns out significant if included without other institutions in the specification.

The analysis of the case of TFP growth is less straightforward. When other institutions are included, the ratio of minimum to median wages attracts a significant coefficient, which is also close to that of Table 3.A4.1 (Columns 3 to 5). However, when all institutions are included, three of them, beside the minimum wage, appear to have significant coefficients (tax wedge, collective bargaining coverage and product market regulation). Yet, this is likely to result again from multicollinearity. Indeed, once insignificant institutions are excluded, the coefficient of product market regulation appears insignificant (Column 4). Moreover, once the minimum wage variable is excluded, both tax wedge and coverage become insignificant (Columns 6 to 7). Overall these results appear to support the identification assumption adopted in this section.

How large is the potential impact of the minimum wage on the business-sector labour share? For the average industry, it is possible to derive a quantitative estimate of the overall effect of minimum wages by taking the conservative assumption that there are no direct effects in the industry with the lowest incidence. Under this assumption, the effect in the average industry can be obtained by simply multiplying β – or γ, as defined in [A4.1], by the difference between the mean and lowest incidence of labour with less than secondary education in the United Kingdom before the introduction of the national minimum wage. As the average incidence is 19.2% and the lowest incidence is 6.7% (in the energy industry), a change of 10 percentage points of the ratio of minimum to median wages (roughly corresponding to one standard deviation of the distribution), would result in a 0.3% annual contraction of the labour share. However, such a variation of wage floors would be enormous in historical perspective: on average, among OECD countries, the ratio of minimum to median wages contracted by about 1 percentage points in the 1990s (excluding countries that introduced a statutory minimum for the first time – Ireland and the United Kingdom)[14] and then increased by 2 percentage points only between 2000 and 2009. Therefore, it is unlikely that the minimum wage played a key role in the historical contraction of the labour share. Even in the 2000s, taking estimates at face value, the growth of the minimum wage would account for a cumulated effect of 0.2 percentage points over the whole period, which is small in comparison with the contraction of the labour share.

Employment protection

There is clear evidence in the literature that employment protection for regular workers (EPR, hereafter) negatively affects productivity growth (see *e.g.* Autor *et al.*, 2007, Bassanini *et al.*, 2009). By contrast, evidence concerning the impact on wages is more mixed. At the micro level, Leonardi and Pica (2010) analyse the effect of monetary compensation for unfair dismissal on male wages by exploiting an Italian reform that introduced this type of compensation for establishments with less than fifteen employees. They find that the reform had no impact on entry wages, although returns to tenure decreased, as suggested by Lazear (1990). Bassanini *et al*. (2010), using a cross-country/cross-industry difference-in-difference approach similar to that adopted in this section, find that the wage premium to voluntary job changes is smaller when EPR rules are more stringent. However, they also find evidence that involuntary job loss is less frequent in that case, so that the overall impact on wage premia to job changes is ambiguous. By contrast, van der Wiel (2010) identify intra-firm effects of employment protection by exploiting a 1999 Dutch reform, which eliminated age-based terms-of-notice rules but implied the coexistence within the same firm of workers under different rules for a transitory period. She finds that those covered by more stringent rules received higher wages. The possible difference between the effects of EPR on productivity and wages suggests that reforms of dismissal regulations might have an impact on the labour share. Indeed, stringent dismissal regulations might worsen the employer's bargaining position, thereby improving bargaining outcomes for workers. However, there is surprisingly little research studying

14. If a value of 0 is considered for the period preceding the introduction of statutory minimum in the United Kingdom and Ireland, the average growth of the ratio of the minimum to median wage in the 1990s would be 4.9 percentage points.

the effect of employment protection on the wage share. The main exception is perhaps Checchi and Garcia-Peñalosa (2008), who estimate a standard aggregate cross-country/time-series model for OECD countries, and find no impact of employment protection controlling for other institutions.

In order to shed further light on this issue the impact of dismissal regulations on productivity, wages, prices and the labour share is estimated here using a difference-in-difference approach similar to one used above for minimum wages. As standard in the literature, the identifying assumption is in this case that the industries where employment protection regulations – at least those concerning permanent contracts – are more likely to be binding are those where firms typically need to lay off workers to restructure their operations in response to changes in technologies or product demand and where, therefore, high firing costs are likely to slow the pace of reallocation of resources. In these industries – EP-binding industries hereafter, one can expect that dismissal legislation has the greatest impact on productivity and wages. By contrast, in industries where firms can restructure through internal adjustments or by relying on natural attrition of staff, changes in employment protection for open-ended contracts can be expected to have little impact.

In practice, the identification assumption made here implies estimating an equation similar to [A4.1] where *MW* is replaced with the OECD EPR indicator and Φ stands for the natural propensity to adjust through dismissals in the absence of regulations. Following Bassanini *et al.* (2009), industry-level US dismissal rates – defined as the percentage ratio of annual dismissals to total employment –[15] are used here as a proxy for underlying layoff propensity in the absence of EP. The United States appears a natural benchmark in this regard because dismissal regulations are very light in comparison with other OECD countries (the EPR index is close to zero in the United States, see Venn, 2009).

Column 1 of Panel A, Table 3.A4.3 shows that EPR stringency exerts a significant negative impact on TFP growth, consistent with Bassanini *et al.* (2009). [16] The relationship between EPR and TFP growth is also reflected in labour productivity growth (Column 2). Interestingly, however, reforms reducing EPR stringency appear to negatively affect output prices (Column 3), while no clear impact is observed on either real wages or the labour share (Columns 4 and 5). By contrast, the short-run impact of EPR appears insignificant on all dependent variables. Results are also robust to exclusion of insignificant short-time effects (Panel B).

15 . Based on 1996-2006 data, even years, computed from various waves of the CPS Displaced Workers Supplement.

16. This analysis is, however, based on longer time-series and using double-deflated TFP data, which better reflect physical (as opposed to value) productivity.

Table 3.A4.3**. Dismissal regulations, productivity and the labour share**

Difference-in-difference estimates

Panel A: Including short-run effects

	(1)	(2)	(3)	(4)	(5)
Dependent variables	TFP	Labour productivity	Price	Wage	Labour share
EPR x US dismissal rate	-0.144***	-0.125***	0.085**	-0.021	-0.002
	(-3.232)	(-2.613)	(2.234)	(-0.708)	(-0.065)
ΔEPR x US dismissal rate	-0.019	-0.292	0.171	0.092	0.019
	(-0.041)	(-0.578)	(0.582)	(0.326)	(0.067)
Observations	4,806	4,806	4,806	4,806	4,806
R-squared	0.308	0.310	0.371	0.321	0.286
Country-by-time dummies	yes	yes	yes	yes	yes
Industry-by-time dummies	yes	yes	yes	yes	yes

Panel B: Excluding short-run effects

	(1)	(2)	(3)	(4)	(5)
Dependent variables	TFP	Labour productivity	Price	Wage	Labour share
EPR x US dismissal rate	-0.144***	-0.120**	0.082**	-0.023	-0.002
	(-3.271)	(-2.540)	(2.193)	(-0.768)	(-0.077)
Observations	4,806	4,806	4,806	4,806	4,806
R-squared	0.308	0.310	0.371	0.321	0.286
Country-by-time dummies	yes	yes	yes	yes	yes
Industry-by-time dummies	yes	yes	yes	yes	yes

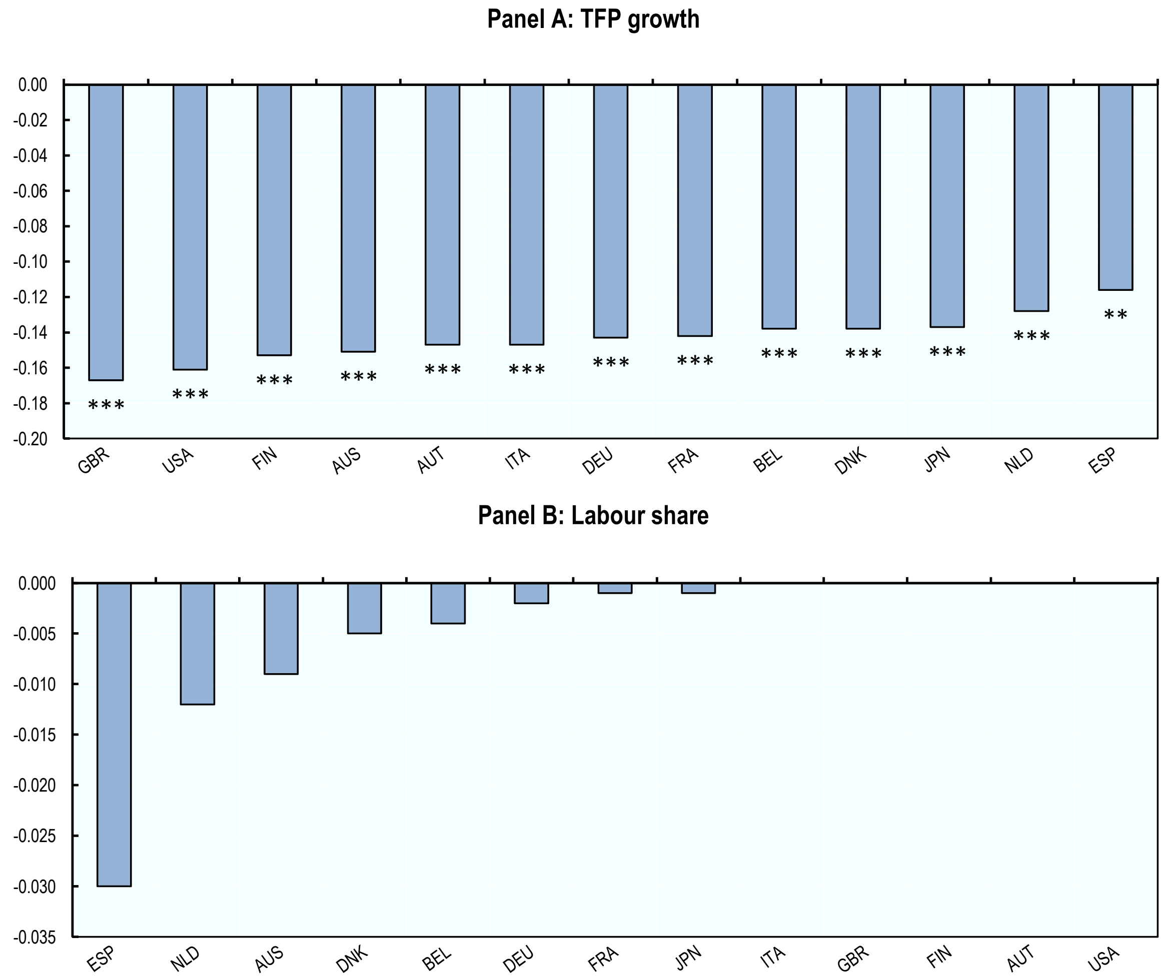
Notes: All dependent variables are in log differences. All equations control for log-relative TFP. EPR: OECD index of employment protection for regular workers. US dismissals: industry dismissal rate in the United States, averaged over 1996-2006 (average value: 0.0518). Robust t-statistics in parentheses. ***, **: significant at the 1% and 5% level, respectively.

These results are also robust to elimination of countries one-by-one from the sample (Figure 3.A4.2). Moreover, as expected, the interactions between other institutions and US dismissals are insignificant, if included, and do not affect the impact of the EPR indicator on TFP (Table 3.A4.4).[17]

17 . Interactions involving collective bargaining coverage are significant at the 10% level. However, they become insignificant if EPR is excluded.

Figure 3.A4.2. **Sensitivity of the effects of EPR to countries included in the sample**

Point-estimates of the coefficient of the interaction between EPR and US dismissal rates once the indicated countries are excluded from the sample



Note: Coefficients obtained by re-estimating the specification of Columns 1 and 5 of Panel B of Table 3.A4.3, excluding the indicated country. **, *** significant at 5% and 1% respectively.

Table 3.A4.4. **Dismissal regulations and TFP: including additional institutional controls**

Difference-in-difference estimates

Dependent variable	(1) TFP	(2) TFP	(3) TFP
EPR x US dismissal rate	-0.186***	-0.208***	
	(-2.818)	(-3.681)	
TW x US dismissal rate	0.001		
	(0.202)		
UB x US dimissal rate	-0.001		
	(-0.366)		
Coord. x US dismissal rate	0.083*	0.078*	-0.023
	(1.914)	(1.893)	(-0.795)
CB x US dismissal rate	-0.003		
	(-1.023)		
PMR x US dismissal rate	0.091		
	(1.251)		
Observations	4,806	4,806	6,660
R-squared	0.309	0.309	0.327
Country-by-time dummies	yes	yes	yes
Industry-by-time dummies	yes	yes	yes

Notes: All dependent variables are in log differences. All equations control for log-relative TFP. EPR: OECD index of employment protection for regular workers. TW: average tax wedge. UB: average unemployment-benefit gross replacement rate. Coord.: index of coordination/corporatism of the wage bargaining. CB: collective bargaining coverage. PMR: time-varying OECD index of anti-competitive product market regulation. US dismissals: industry dismissal rate in the United States, averaged over 1996-2006 (average value: 0.0518). Robust t-statistics in parentheses. ***, **: significant at the 1% and 5% level, respectively.

The overall impact in the average industry can be estimated in the same way as for the minimum wage by assuming that the impact of employment protection reforms on TFP growth is proportional to the difference between the natural level of dismissals of that industry and in the industry with the lowest propensity (as proxied by US dismissals). As the average and lowest US dismissal rates in the sample are 5.18% and 2.22%, respectively, the impact of a 1-point reform of dismissal regulations – approximately leading the average OECD country half-way between the current average stringency of regulations and the level of the United States–[18] would increase TFP growth in the average industry by 0.42 percentage points. This effect is close to that estimated by Bassanini *et al.* (2009) – 0.49 percentage points. By contrast, the impact on the labour share would be negligible (-0.006% per year). Even more, this effects should be compounded with the fact that a 1-point reform of employment protection is large in historical perspective. For example, the 2003 Austrian reform of severance payments, usually considered a radical reform of dismissal regulations, translates into a fall in the indicator by 0.55 points only.

The fact that the impact of dismissal regulations on TFP growth is reflected in prices but not in wages or the labour share can be explained by the fact that those industries in which EP regulations are more likely to be binding are downsizing manufacturing industries where product market competition is typically fierce, dissipating rents arising from efficiency increases. This evidence, however, suggests that policy reforms reducing employment protection benefit all workers, but essentially as consumers, while workers of EP-binding industries bear the brunt of the reforms as their risk of job loss increases.

18. This corresponds to 1 standard deviation in the distribution of the EPR indicator.

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