Universitat Autònoma de Barcelona

Three Essays on the Labour Market Macroeconomics

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by

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Declaration of Authorship

I, Pedro Trivin, declare that this thesis titled "Three Essays on the Labour Market Macroeconomics", and the work presented in it are my own. I confirm that:

- This work was done wholly or mainly while in candidature for a research degree at this University.
- Where any part of this thesis has previously been submitted for a degree or any other qualification at this University or any other institution, this has been clearly stated.
- Where I have consulted the published work of others, this is always clearly attributed.
- Where I have quoted from the work of others, the source is always given. With the exception of such quotations, this thesis is entirely my own work.
- I have acknowledged all main sources of help.
- Where the thesis is based on work done by myself jointly with others, I have made clear exactly what was done by others and what I have contributed myself.

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Last, but the most, this thesis is dedicated to my parents, Alberto and Begoña, and my brother Alberto, which unconditional support and endless trust on me have made this possible. This dissertation belongs to them as much as to me.

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Introduction

> *"The difficulty lies, not in the new ideas, but in escaping from the old ones, which ramify, for those brought up as most of us have been, into every corner of our minds."*
>
> (Keynes, 1936)

This Ph.D. thesis consists of three essays on the macroeconomics of the labour markets, having the goal of contributing to the scientific discussion, and shed new light on the configuration of labour markets.

The remaining of the chapter presents a general motivation for our study together with the main findings and policy implications, which are fully developed along the thesis.

Motivation

Last decades have witnessed striking socioeconomic changes. Beyond the economic distress spread around the world since 2008, recent times are characterised, among other factors, by a secular decrease of the labour income share, an increasing global economic integration, and a larger relevance of the financial sector in the economy. These changes have blurred the effects of traditional nation-based economic policies. The goal of this dissertation is, therefore, to provide an insightful discussion on how these factors have altered the labour markets from a macroeconomic perspective.

Global economic integration, most commonly referred to as "globalisation", has been usually pointed by scholars as one of the crucial factors explaining macroeconomic changes. Growing international integration was made possible by changes in the disposition of countries (trade agreements, financial liberalisation...) along with the last waves of technological change that decreased transportation and monitoring costs.

From a theoretical perspective, standard international trade models (Heckscher-Ohlin) predict a benefical effect of globalisation for all the parties involved. The reasoning is based on a comparative advantage argument. Countries' specialisation, together with larger mobility in production factors and final products, will foster competition, increasing economic efficiency. Empirical evidence supporting this hypothesis comes from Dreher (2006), where a positive relationship between globalisation and economic growth is found.

Nonetheless, there is not little controversy and opponents to the globalisation process. One of the possible explanations for this opposition is the heterogenous impact of globalisation on different economic agents. As Rodrik (1997) has pointed out:

> "Globalization is exposing a deep fault line between groups who have the skills and mobility to flourish in global markets and those who either don't have these advantages or perceive the expansion of unregulated market as inimical to social stability and deeply held norms." (Rodrik, 1997, p.2)

In this sense, Feenstra and Hanson (1996) relate the globalisation process to an increase in the relative demand for skilled labour in the United States. Beyond the conflict among workers, the larger mobility of capital with respect to labour is a potential source of tensions between workers and the owners of capital. Globalisation would contribute not just to increase the level of substitution between production factors, but would also decrease the bargaining power of workers.

Parallel to the globalisation phenomenom, there has been an increase in the weight of the financial sector on economic activity (Stockhammer, 2004). This process has been named "financialisation", and although there is not a unique definition, an insightful one is given by Epstein (2005):

> "Financialization means the increasing role of financial motives, financial markets, financial actors and financial institutions in the operation of the domestic and international economies." (Epstein, 2005, p.365)

Factors such as a more benevolent taxation of the capital income (McGrattan and Prescott, 2005), a reduction of the stock-market costs (McGrattan and Prescott, 2003), and a switch towards a shareholder-oriented corporate model (Kaplan, 1997) are the main candidates to explain this evolution.

According to this literature, corporations have tended since the early 80s to pursue short-term payout policies that increase the equity price at the expense of long-term investment (Lazonick and O'Sullivan, 2000). Therefore, the financial sector would be subtracting productive resources from the economy, having sizeable consequences for the functional distribution of income.

This dissertation is motivated from the scientific urge to learn more about the impact of these factors on the dynamics behind labour market outcomes. The final target of the three papers presented below is to contribute to the scientific knowledge about how labour markets work, but also to question the validity of some of the assumptions widely used in economics. Next sections develop in further detail the main results and policy implications related with this research.

Main results

Chapter 1: The effects of globalisation and technology on the elasticity of substitution

The first chapter analyses the impact of globalisation and technological change on the elasticity of substitution between labour and capital (σ) in both OECD and non-OECD countries. The elasticity of substitution is one of the most relevant parameters in economic models. Previsions about economic growth, the functional distribution of income or the impact of fiscal and monetary policies heavily depend on its value. Given its importance, many authors have explored its magnitude (e.g., Antràs, 2004; and Chirinko, 2008). However, although empirical evidence suggests that the value for σ differs across both periods and regions, little attention has been paid to the factors explaining these heterogeneities.

This chapter borrows the idea of a Variable Elasticity of Substitution (VES) production function (Antony, 2009a,b, 2010) in order to analyse the role played by globalisation and technological change in determining σ. Empirically, the strategy we follow is given by Bentolila and Saint-Paul's (2003) framework, who make use of the relationship between the labour income share and capital intensity (or the capital-output ratio) to study the magnitude of σ.

We show that both globalisation and technological change play a role as σ driving forces. Interestingly, we find heterogeneous impacts for the two groups under analysis. While for the OECD globalisation increases the substitutability between capital and labour, there is no effect for the non-OECD group. Technological change, on the other hand, results in a larger complementary between production factors in OECD countries, but it has the opposite effect in the non-OECD group.

The results on the effects of globalisation support Rodrik's (1997) hypothesis, according to which economic integration increases the substitutability between labour and capital through the additional alternatives available in the economy. The fact that we find an effect only for OECD countries speaks in favour of the argument that offshoring represents a greater threat to workers in developed countries compared to those in developing ones. Concerning technological change, our findings provide empirical support to the capital-skill complementarity hypothesis, implying that capital is more complementary to high-skill than to low-skill workers.

Our reading of these results is related to the functional distribution of income and economic growth. From the economic growth literature, we have learnt that the more flexible a production function is (i.e., the higher the value of σ), the larger the potential growth. However, the value of σ does also affect the functional income distribution. Given that this effect depends on the pattern followed by the capital-output ratio, a higher value of σ has an ambiguous impact on functional inequality. More specifically, an increase in the capital-output ratio will increase the labour share if $\sigma < 1$, but it will have the opposite effect if $\sigma > 1$. Consequently, the intertwined linkage among globalisation, technology and the elasticity of substitution should be taken into account in any policy makers' objective function.

Chapter 2: Finance and the global decline of the labour share

The global decline of the labour income share has attracted growing interest in the literature (Bentolila and Saint-Paul, 2003; Elsby et al., 2013; Karabarbounis and Neiman, 2014; Piketty and Zucman, 2014). Despite the documented importance of globalisation and market regulations, most of the current literature has focused on structural/technological explanations. This branch of the literature relies on a one-to-one relation between the labour share and the capital-output ratio presented in

Bentolila and Saint-Paul (2003), where the direction of the effect depends on the elasticity of substitution between capital and labour.

The main contribution of the existing literature relies on looking for structural drivers of the labour share by making endogenous the dynamics of the capital-output ratio. In this sense, Piketty and Zucman (2014) argue that a persistent gap between the rate of return to capital and the growth rate of output results in a growing accumulation of capital because capitalists save most of their income. On the other hand, Karabarbounis and Neiman (2014), using a version of the model presented in Greenwood et al. (1997), argue that the persistent global decrease in the relative price of investment goods has induced firms to use more capital at the expense of labour, depressing the labour income share.

Although Piketty and Zucman (2014) and Karabarbounis and Neiman (2014) emphasise different channels, both share the view that the increase in the capital-output ratio has been the main driver of the decline in the labour share. However, their results require a degree of substitutability between capital and labour far beyond what has been previously estimated by the literature (Antràs, 2004; Chirinko, 2008; Chirinko and Mallick, 2014).

This chapter contributes to the debate by exploring an alternative structural channel compatible with standard values of the elasticity of substitution ($\sigma < 1$). More specifically, we exploit the relation between financial markets and corporate investment, which we connect with the evolution of the equity Tobin's Q.

Our contribution is both theoretical and empirical. We develop a simple model where a raise in equity Tobin's Q increases equity returns and, importantly, depresses the capital-output ratio. A decrease in the capital-output ratio reduces, therefore, the labour share if $\sigma < 1$. Empirically, based on a common factor model, we estimate several Mean Group-style estimators for a sample of 41 countries finding that the increase in Tobin's Q explains between 41% and 57% of the total decline in the labour income share around the world.

This paper not just implies that financial markets have direct and significant effects on inequality via their impact on the functional distribution of income, but it also reconciles the structural literature with more realistic values of the elasticity of substitution ($\sigma < 1$).

Chapter 3: Labour market dynamics in Spanish regions: Evaluating asymmetries in troublesome times

Spain is a particularly good example of how the global financial crisis severely affected, among other things, countries' unemployment levels. After more than a decade of a downward trend and convergence to the European average, the unemployment rate reached 26% in 2012, surpassing the historical maximum of 24% in 1994. The great variations in unemployment rate go in parallel with an extreme degree of regional persistence. In other words, the Spanish regions with a higher unemployment rate in the expansion phase are the ones with a higher unemployment rate during the recession.

This chapter investigates the different labour market responses to a regional-specific shock in order to identify different patterns along the different phases of the business cycle. We use the framework of analysis developed by Blanchard and Katz (1992) which allows to asses the impact of shocks to employment not only in terms of unemployment rate, but also on the changes in participation rates and regional mobility. Such analysis will enhance the understanding, from a regional perspective, of the labour market adjustment mechanisms in the different scenarios studied. Beyond the study across business cycle phases, we extend the analysis to include price adjustments and heterogeneities among regions.

Our findings are diverse. First, we identify asymmetric labour market responses across business cycle phases. We find that changes in participation rates are the main adjustment mechanism during expansion periods, while unemployment becomes the central one during recessions. Moreover, the long-run impact on employment is larger when the shock hits in a recessive period than when it hits in expansion. This result is indicative of net migration –spatial mobility– being more relevant in troublesome than in good times. We also provide evidence of real wage rigidities in both periods, where neither prices or wages adjust to a regional employment shock. Finally, we find evidence of larger spatial adjustments in high than in low unemployment regions in response to the employment shock. On this account, it seems safe to conclude that people in a region are more willing to migrate (relative to the national average) when the regional shocks take place in a relatively bad scenario.

Conclusions and policy implications

All in all, the goal of this thesis is to contribute to the academic debate about the labour market effects of globalisation, financial markets, or the business cycle from a macroeconomic perspective. The empirical evidence provided carries out important policy implications.

We show how globalisation and technological change have an impact on the elasticity of substitution between labour and capital (σ). The value of this parameter is crucial concerning the implications of several economic policies. However, to the best of our knowledge, so far the literature failed to provide a complete picture of the possible factors driving σ's value. If globalisation and technological change are proved to be determinants of σ, it urges the profession to endogenise its value in both theoretical and empirical studies, in order to correctly address policy making in relevant fields such as economic growth or the functional distribution of income.

With respect to the functional distribution of income, we also outline the close connection between financial markets and the decline in the labour income share observed during the last decades. In particular, contrary to what has been argued, it is not an increase in capital accumulation what explains this decline, but exactly the opposite. Financial markets have subtracted productive resources from the economy, biasing manager decisions towards financial investments to the detriment of physical capital. Policies which encourage physical capital investment, such as tighter financial market regulations or an increase in the income tax obtained from financial products should be promoted in order to change the negative trend followed by the labour income share.

Additionally, we show that Spanish regional unemployment persistence is not different today than 30 years ago, even though it seems that people is more willing to migrate when they face a regional employment shock in a relative negative scenario. Our study shows strong real wage rigidities arising from both nominal wages and consumer prices. This seems to indicate that the product and labour markets are still operating with a substantial degree of imperfect competition, in which case policy measures to foster competition and a larger responsiveness of market prices to the changing (regional) economic environment should be further implemented.

Altogether, this thesis should be viewed as a departing point, which we hope will foster the economic debate on the role played by the aforementioned factors.

Chapter 1

The effects of globalisation and technology on the elasticity of substitution

Abstract

The elasticity of substitution between capital and labour (σ) is usually considered a "deep parameter". This paper shows, in contrast, that σ is affected by both globalisation and technology, and that different intensities in these drivers have different consequences for the OECD and non-OECD economies. In the OECD, we find that the elasticity of substitution between capital and labour is below unity; that it increases along with the degree of globalisation; but it decreases with the level of technology. Although results for the non-OECD area are more heterogeneous, we find that technology enhances the substitutability between capital and labour. We also find evidence of a non-significant impact of the capital-output ratio on the labour share irrespective of the degree of globalisation (which would be consistent with an average aggregate Cobb-Douglas technology). Given the relevance of σ for economic growth and the functional distribution of income, the intertwined linkage among globalisation, technology and the elasticity of substitution should be taken into account in any policy makers' objective function.

JEL Codes: E25, F62, E22, O33.

Keywords: Labour Share, Capital-Output Ratio, Elasticity of Substitution, Globalisation, TFP.

This paper is coauthored with my supervisor Hector Sala.

1.1 Introduction

The elasticity of substitution between capital and labour (σ) is a key macroeconomic parameter. It determines the path of economic growth, affects the functional distribution of income, and conditions the impacts of fiscal and monetary policies (Klump and de La Grandville, 2000; Chirinko, 2008).

The value of σ has received utmost attention in recent years. However, although Constant Elasticity of Substitution (CES) production functions have been widely used to determine its value (Antràs, 2004), such approach delivers the estimation of a constant, which is equivalent to treat σ as a deep parameter.

In contrast, this paper looks at potential determinants of σ and examines their effects. To undertake this analysis, we borrow the idea of a Variable Elasticity of Substitution (VES) production function developed by Antony (2009a,b, 2010),[1] whose main feature is that variations in the relative factor intensity (or capital-labour ratio) may cause changes in σ.

To identify the value of σ, the strategy we follow is given by Bentolila and Saint-Paul's (2003) framework, where the relationship between the labour income share and capital intensity (or the capital-output ratio) is used to study the magnitude of σ.

However, given the intimate relationship between the relative factor intensity and the capital-output ratio (both embedded in any production function), it would be to a large extent tautological to bring the former in the analysis. This is the reason why we directly focus on the determinants of the relative factor intensity, in particular globalisation and technology, which have been highlighted as critical in the literature.

Regarding globalisation, Heckscher-Ohlin-type models predict that economies specialise in the production of goods that use the relatively abundant factor. This would explain the specialisation of OECD countries in capital-intensive goods, and hence their relative increase in capital deepening. At the same time, there is evidence that the vertical fragmentation of the production process between developed and developing countries could cause further capital deepening in the developed ones (Jones and Kierzkowski, 1998; Feenstra and Hanson, 1999, 2001).

[1] This idea was initially developed by Revankar (1971).

With respect to technological change, it may induce capital deepening by increasing labour productivity which will, in turn, boost the marginal productivity of capital. Along these lines, Madsen (2010), for example, extends the model by Abel and Blanchard (1983) to show the causal relationship between total factor productivity (TFP) and capital-deepening.

Finally, there seems to be a close relationship between globalisation and technology as the former can speed the diffusion of technologies (Coe and Helpman, 1995; Falvey et al., 2004), and also spur the innovation, by increasing the extent of the market (Sachs, 2000).

Our main goal is, thus, to shed new light on the effects of globalisation and technology on the elasticity of substitution between capital and labour. More specifically, this paper empirically checks whether globalisation increases the substitutability between capital and labour, and whether new technologies tend to substitute for unskilled labour and complement skilled workers in their qualified tasks –which would confirm the capital-skill complementarity hypothesis. This is done under the premise that OECD economies enjoy, on average, a relatively larger share of skilled workers than non-OECD economies (Krusell et al., 2000).

From a methodological point of view, the fact that σ depends on globalisation and technology leads us to use multiplicative interaction models (see Brambor et al., 2006), which are estimated under two different econometric approaches. First, as standard homogeneous dynamic panel data models. Second, as heterogeneous dynamic panel data using a common factor model approach (see Eberhardt and Teal, 2011), which allows for heterogeneities across countries and to control for the impact of unobservable factors.

Our database contains information for 51 OECD and non-OECD economies with a sample period running from 1970 to 2009.[2] The main findings of our analysis are (1) that both globalisation and technology affect the elasticity of substitution between capital and labour; and (2) that this is expressed in different patterns in the OECD and non-OECD economies.

For the OECD, the results are robust and conclusive. We find that the elasticity of substitution between capital and labour is below unity; that it increases along with the degree of globalisation; but it decreases with the level of technology. The first result confirms previous findings in the literature (Antràs, 2004; Chirinko,

[2]These economies and the corresponding sample periods are listed in Table 1.A1 in the Appendix.

2008). The second one implies that globalisation enhances the substitutability between production factors, also provides empirical support to literature in this area (see among others, Rodrik, 1997; Slaughter, 2001; Saam, 2008; Hijzen and Swaim, 2010). The third result implies that technological change boosts factors' complementarity, and endorses empirically the capital-skill complementarity hypothesis.

Although our results are not so robust for the non-OECD economies, they still yield some useful outcomes. The main one is that the elasticity of substitution between capital and labour becomes larger along with the level of technology. We hypothesize that this enhanced substitutability takes place at early stages of the development process, when standard technologies help mechanizing outdated labour-intensive tasks. With respect to globalisation, however, results are more complex and difficult to interpret, even though we find evidence of a non-significant impact of the capital-output ratio on the labour share irrespective of the degree of globalisation (which implies a unit σ). These inconclusive results could be due to the inherent heterogeneity within this area, as well as to their lower quality data.

Our reading of these findings is related to the functional distribution of income and economic growth. From the economic growth literature we have learnt that the more flexible a production function is (i.e., the higher the value of σ), the larger the potential growth that can be achieved. However, the value of σ does also affect the functional income distribution. Given that this effect depends on the pattern followed by the capital-output ratio, a higher value of σ has an ambiguous impact on functional inequality. Consequently, the intertwined linkage among globalisation, technology and the elasticity of substitution should be taken into account in any policy makers' objective function.

The remaining of the paper is structured as follows. Section 1.2 deals with some basic theoretical issues and the empirical strategy. Sections 1.3 and 1.4 present the econometric methodology and the data. Sections 1.5 and 1.6 show, respectively, the outcomes of the estimated homogeneous and heterogeneous models. Section 1.7 concludes.

1.2 Theoretical issues and empirical strategy

1.2.1 The *share-capital* schedule and σ

Bentolila and Saint-Paul (2003) derive a unique function g relating the labour income share (s_L) and the capital-output ratio (k):

$$s_{L_i} = g(k_i) \tag{1.1}$$

This stable relationship is called the *share-capital* (SK) schedule (or curve), and is unaltered by changes in relative factor prices or quantities, and by labour-augmenting technical progress. Shifts of the SK schedule can be explained by capital-augmenting technology or by the increase of intermediate inputs prices. In turn, factors that generate a gap between the marginal product of labour and the real wage (for example, union bargaining power, and labour adjustment costs), cause shifts off the SK curve.

The analysis of the SK schedule is stressed to deserve special attention because it is intimately related to σ. This is shown analytically by considering a generic production function –with standard assumptions such that it has to be differentiable and with constant returns to scale–, which ends up delivering the following equation:

$$\frac{ds_L}{dk} = -\frac{1+\sigma}{k\eta} \tag{1.2}$$

In equation (1.2), the response of the labour share to changes in the capital-output ratio is related to σ, the labour demand elasticity with respect to wages holding capital constant (η), and the value of the capital-output ratio (k).

This expression allows the estimation of the labour share elasticity with respect to the capital-output ratio (ε_{S_L-k}) to be interpreted in terms of the value of σ:

$$\begin{aligned} \hat{\varepsilon}_{S_L-k} > 0 \implies \sigma < 1 \implies \text{Capital and labour factors are complements.} \\ \hat{\varepsilon}_{S_L-k} < 0 \implies \sigma > 1 \implies \text{Capital and labour factors are substitutes.} \end{aligned} \tag{1.3}$$

In general, the first situation ($\sigma < 1$) has been associated to developed economies, because of their largest proportion of skilled workers making them more complementary to capital (relative to non-skilled workers). In turn, the second scenario

($\sigma > 1$) is more connected to the situation in developing economies, where the larger share of low-skilled workers makes capital and labour more substitutes.[3]

We exploit the relationships represented by expression (1.3) to identify the impact on σ of different driving forces.

1.2.2 A varying σ

Although most of the literature has tended to work with CES type production functions, growing attention has been devoted to relax the assumption of a constant elasticity of substitution. A first important attempt was the VES production function introduced by Revankar (1971).[4] This type of function, however, implies a parametric and strictly monotonic path of σ which has to be above or below 1, but cannot cross.

In this context, a second generation of VES production functions overcoming this restriction has recently been derived by Antony (2009a,b, 2010). In the simplest case, Antony (2009a) refers to a *dual elasticity of substitution* production function, which takes the following general form:

$$y = f(x) = A(\alpha x_b^v + (1-\alpha))^{\frac{1}{v}-\frac{1}{\rho}} \left(\alpha x_b^v \left(\frac{x}{x_b} \right)^{\rho} + (1-\alpha) \right)^{\frac{1}{\rho}}, \tag{1.4}$$

where x represents the relative factor intensity $(\frac{K}{L})$, and x_b is a baseline value of that relative factor intensity.[5] The intuition behind these production functions is that σ can take different values depending on the relative factor intensity.[6]

As noted before, the SK schedule takes into account changes in relative factor prices and quantities, which only affect the position along the curve. The possibility that relative quantities alter σ, however, is not considered in Bentolila and Saint-Paul (2003). This opens up a channel by which the slope of the SK schedule may be affected by relative quantities. In other words, there is a transmission

[3]This is known as the "capital-skill complementarity hypothesis". Formal and empirical evidence can be found in Griliches (1969), Goldin and Katz (1996), and Caselli and Coleman (2001).

[4]Sato and Hoffman (1968) and Jones and Manuelli (1990) have also worked in this production function.

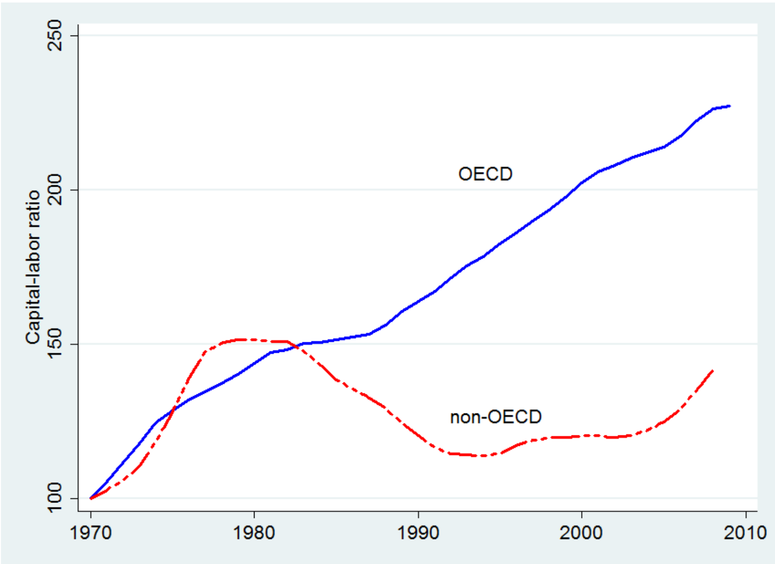
[5]These production functions are based on the idea from de La Grandville (1989) of normalising CES production functions.

[6]More specifically, if x $\neq x_b$; $\sigma = \frac{1}{1-\rho}$, whereas if x $= x_b$; $\tilde{\sigma} = \frac{1}{1-v}$. A generalisation of this type of production function is presented in Antony (2010), where it is extended to allow for multiple values of σ conditioned on a diversity of intervals of the relative factor intensities.

mechanism by which the determinants of the capital-labour ratio end up affecting the impact of the capital-output ratio on the labour income share (see equation (1.2)).

Following this idea, it is crucial to check whether relative factor intensities have changed over last decades. If this is the case, not only it is harder to retain the idea of a constant σ, but it seems also advisable to examine to what extent the elasticity of substitution between capital and labour has responded to these changing patterns.[7]

FIGURE 1.1: Relative factor intensities



Notes: Own calculation (based on raw data from the Extended Penn World Table, EPWT 4.0) obtained as year fixed effects from a GDP weighted regression of the capital-labour ratio including country fixed effects to control for the entry and exit of countries throughout the sample. Initial year is normalised to equal 100 in 1970.

Figure 1.1 shows the evolution of the capital-labour ratio in the OECD and non-OECD economies between 1970 and 2009. A different pattern can be clearly appreciated. Whereas the OECD countries show a continuous (and constant) process of capital deepening (having nowadays a capital-labour ratio more than twice the one in 1970), the non-OECD group has a more heterogeneous pattern. During the 1970s, the increase in the ratio was mostly in parallel with respect to the OECD, but it was followed by an important decrease (40 pp) since then until mid 1990s, when capital gingerly started to rise again with respect to labour. The capital-labour ratio for the non-OECD countries is nowadays almost 1.5 times the value in 1970.

[7]Along these lines, Saam (2008) showed that globalisation tends to rise σ in economies with a larger capital deepening.

Given this information, the next step is to bring the idea of a varying σ into a suitable empirical tool in which the main determinants of relative factor intensity –globalisation and technology– may condition this varying value by affecting the elasticity of the labour share with respect to the capital-output ratio (ε_{S_L-k}). Next, we argue that the estimation of multiplicative interaction terms is this suitable empirical tool.

1.2.3 Empirical strategy

The estimation of standard linear additive models –where all explanatory variables are considered independent– may be inferior to the alternative of estimating multiplicative interaction models –where some explanatory variables are considered dependent (Aiken et al., 1991; Brambor et al., 2006). As explained in Brambor et al. (2006), this is the appropriate tool whenever the hypothesis being tested is conditional in nature. This is our main reason to change the strategy and estimate new empirical versions of the models in Bentolila and Saint-Paul (2003) using multiplicative interaction models.

A simple linear additive model can be expressed as:

$$\hat{Y} = \alpha_0 + \alpha_1 X + \alpha_2 Z + \alpha_3 W, \tag{1.5}$$

where $\hat{Y}$ is the estimated value of the dependent variable, X, Z, and W are the explanatory variables, and the α's are estimated parameters. In turn, the corresponding multiplicative interaction model takes the form:

$$\hat{Y} = \beta_0 + \beta_1 X + \beta_2 Z + \beta_3 W + \beta_4 XZ + \beta_5 XW + \beta_6 ZW + \beta_7 XZW, \tag{1.6}$$

where the β's are estimated parameters, and XZ, XW, ZW, and XZW are interaction terms.

The presence of the interaction terms alters the interpretation of the estimated parameters in a fundamental way. The reason is that in model (1.5) X, Z and W are considered independent of one another, whereas in model (1.6) they are not. In other words, in the additive model the effect, for example, of X on Y is considered to be constant while, in the multiplicative interaction model, this effect depends on the values taken by W and Z. Therefore:

- α_1 is the unconditional marginal effect of X on Y, while β_1 is the conditional marginal effect of X on Y when $W = Z = 0$.

  (The equivalent interpretation holds for α_2 and β_2 regarding Z, or α_3 and β_3 regarding W).

- β_4, β_5, and β_7 capture the impact of X on Y for different values of the modifying variables W and Z, and allow this impact to vary. That is, the overall conditional marginal effect of X on Y is:

$$\frac{\partial Y}{\partial X} = \beta_1 + \beta_4 Z + \beta_5 W + \beta_7 ZW \tag{1.7}$$

Although multiplicative interaction models have been widely used in Political Science, Brambor et al. (2006) acknowledge that just 10% of the papers analysed in their survey estimate and interpret them in the correct way. Two of the most common problems are the lack of one or more constitutive terms, and the lack of computation of "marginal effects and standard errors across a substantively meaningful range of the modifying variable" [Brambor et al., 2006, p.78].

As our empirical model assesses the cross-dependencies among the capital-output ratio, globalisation and technology, the estimation of an equation such as (1.6) is required to handle the first concern. We take care of the second concern by computing the marginal effects and standard errors of X (in our case, the capital-output ratio) across a substantively meaningful range of the modifying variable Z (changes either in globalisation or technology) for given values of W (technology or globalisation).

In this way, we move from a benchmark empirical model in Bentolila and Saint-Paul (2003) taking the form:

$$\ln(s_{L_{it}}) = \beta_0 \ln(TFP_{it}) + \beta_1 \ln(k_{it}) + \beta_2 \ln(q_{it}/p_{it}) + \beta_3 \Delta \ln(n_{it}) + \beta_4 \widetilde{lcr}_{it} + v_{it}, \tag{1.8}$$

where TFP is a measure of TFP; q is the national price of imported oil, which is deflated by p (the national price); $\Delta \ln(n)$ is the growth rate of employment; and $\widetilde{lcr}$ is the number of labour conflicts nationwide representing the factors causing a gap between the marginal product of labour and the real wage, to an augmented model with interactions:

$$\begin{aligned} \ln(s_{L_{it}}) = \boldsymbol{X'_{it}}\gamma_0 + \gamma_1 [\ln(k_{it}) * KOF_{it}] + \gamma_2 [\ln(k_{it}) * \ln(TFP_{it})] \\ + \gamma_3 [\ln(TFP_{it}) * KOF_{it}] + \gamma_4 [\ln(k_{it}) * \ln(TFP_{it}) * KOF_{it}] + \epsilon_{it}, \end{aligned} \tag{1.9}$$

where, KOF is a proxy of globalisation, and vector $\boldsymbol{X_{it}}$ contains all control variables, beyond the interactions.

1.3 Econometric methodology

Our empirical analysis will first rely on the use of standard dynamic panel data techniques. Then, we will use recently developed panel data methods to estimate common factor models, so as to check the robustness of our results on the effects of globalisation on σ, and consider potential asymmetries in these effects at high and low globalisation scenarios. These techniques are described, respectively, in Subsections 1.3.1 and 1.3.2. The first one deals with homogeneous models (those in which homogeneity in the estimated parameters is imposed across countries), while the second one deals with heterogeneous models (those in which cross-country heterogeneity is allowed in the parameters).

1.3.1 Dynamic homogeneous models

Two relevant considerations relate, first, to the use of averages versus yearly data, and, second, to the estimation of static versus dynamic models. Some authors have used five (or three)-years averages arguing that such procedure allows to abstract from the incidence of business cycles. This is accompanied by a static estimation, which is generally interpreted as approaching steady-state relationships. In our view, this aggregation procedure is not only likely to hide precious information on the relationship between variables but, further, it implies strong assumptions when using a wide sample of countries. For example, taking five-year averages imposes that all economies have the same business cycle in terms of length, and also in terms of starting and end years.

Here, we use all available information to estimate dynamic models with annual data. This has the advantage of increasing the number of observations so that the resulting extra degrees of freedom grant the possibility of splitting the analysis into OECD and non-OECD economies.

The estimation of dynamic models, however, rises the issue of inconsistency if the estimation is conducted using traditional panel data techniques. This is so, even in the absence of serial correlation in the error terms (Nickell, 1981; Arellano and Bover, 1995; Blundell and Bond, 1998), and this is the reason why our reference results will be the ones obtained through the estimation by System Generalized Method of Moments (BB).

The System GMM method consists in the estimation of two equations, one in differences and one in levels, such as:

$$\begin{aligned} \Delta y_{i,t} &= \gamma \Delta y_{i,t-1} + \delta \Delta \mathbf{X}_{i,t} + \Delta \epsilon_{i,t} \\ y_{i,t} &= \gamma y_{i,t-1} + \delta \mathbf{X}_{i,t} + \eta_i + \epsilon_{i,t} \end{aligned} \tag{1.10}$$

The key contribution of this method is that it uses further moment conditions than the Difference GMM estimator.[8] Not only the levels of the variables lagged twice and more are used, but also the lags of the variables in differences, which are now added as instruments in the level equation of the system.

GMM estimators were originally developed for panel data with a large number of cross-sections relative to the time dimension of the panel. In contrast, the cross-section and time dimensions of our database are at most similar in the best case. This could cause estimation problems such as the risk of "instruments proliferation", which could bias the Hansen test to generate p-values artificially close to 1 and over fit endogenous variables. As Roodman (2009) explains, this is because the instrument count quartic in the time dimension of the panel.

To avoid this problem, we have followed Roodmans' (2009) recommendations of reducing as much as possible, and make explicit, the number of instruments used in our estimations.[9]

1.3.2 Dynamic heterogeneous models

Some of the caveats arising from the estimation of homogeneous models may be addressed through the estimation of heterogeneous models which, in passing, will also provide a robustness check on our first set of results.

The two main caveats refer to (1) the relative cross-section/time dimension features of the underlying information; and (2) the assumptions of parameters homogeneity

[8]For details on System GMM, and its advantages over Difference GMM, see Bond et al. (2001) and Roodman (2009).

[9]Accordingly, both samples are estimated allowing for just four lags of endogenous variables and using the "collapse" instruments option available in the *xtabond2* Stata command developed by David Roodman. Unfortunately, even in this case, our results on the Hansen test still seem "too good to be real".

(across countries), common impact of unobservable factors, and cross-section independence. If these assumptions are not correct, the results obtained by estimating homogeneous models would be subject to misspecification problems. Beyond that, it is also true that the use of TFP as a measure of technology is to some extent controversial (see Carlaw and Lipsey, 2003).

In order to overcome these potential sources of misspecification, we now undertake a common factor model approach (for details, see Eberhardt and Teal, 2011). Formally, if we assume for simplicity a one-input model, the common factor model is as follows:

$$y_{it} = \beta_i x_{it} + u_i \tag{1.11}$$
$$x_{it} = \delta_i f_t + \gamma_i g_t + \phi_i + e_{it} \tag{1.12}$$
$$u_{it} = \varphi_i f_t + \psi_i + \varepsilon_{it}, \tag{1.13}$$

where y_{it} is the dependent variable, x_{it} is the independent one, and u_{it} contains the unobservables and the error term (ε_{it}). One of the salient features of this model is that it allows for a country-specific impact of the regressor on the dependent variable (β_i). Unobservables in the error term are represented by a country fixed-effect (ψ_i) and a common factor (f_t) with different factor loadings (φ_i), controlling respectively, for time-invariant and time-variant heterogeneity. At the same time, the regressor can be affected by these, or other common factors (f_t and g_t). Unobservables control both for common global shocks and local spillovers (Chudik et al., 2011; Eberhardt et al., 2013), along with cross-section dependence.

Empirically, our estimated equation is:

$$\ln(s_{L_{it}}) = \beta_{i0} + \beta_{i1}\ln(k_{it}) + \beta_{i2}KOF_{it} + \beta_{i3}\ln(k_{it}) * KOF_{it} + \epsilon_{it} \tag{1.14}$$

This equation will be estimated by imposing different restrictions on its coefficients, which vary depending on whether we deal with homogeneous or heterogeneous models. Homogeneous models assume that the impact of the regressors on the dependent variable is common across countries (i.e., $\beta_i = \beta$), while heterogeneous models leave this unconstrained (i.e., β_i is estimated for each country). In both groups, the impact of unobservable factors may be subject to different assumptions leading to different estimation methods.

Within the homogeneous models, we consider the common Pooled Ordinary Least Square (POLS), the Two-way Fixed Effects (2FE), and the Pooled Common Correlated Effects (CCEP) estimators. Regarding the assumptions about the structure of the unobservables, the POLS and 2FE estimators assume a common impact across countries, accounted by time dummies.[10] In contrast, the CCEP estimator allows a different impact of the unobservables across countries. More specifically, it eliminates the possible cross-sectional dependence by augmenting equation (1.14) with the cross-sectional averages of the variables, which have a different impact on each country.

With respect to the second group, we estimate a set of Mean Group-style estimators. We first present the results for the Pesaran and Smith (1995) Mean Group estimator (MG), which accounts for β heterogeneity, but assumes that the unobservables have a constant growth rate (empirically accounted by adding country-specific linear trends). Then, we present those obtained using the Pesaran (2006) Common Correlated Effects Mean Group estimator (CMG), and the Chudik and Pesaran (2015) Dynamic CMG estimator (CMG2), which account for both parameter and unobservable heterogeneities (empirically accounted augmenting equation (1.14) with the cross-sectional averages of the variables).

Once the heterogeneous models have been estimated, we exploit the cross-country heterogeneity from the Mean Group-style estimators' first stage to identify the effect of globalisation on the impact of the capital-output ratio on the labour income share. The result of this alternative approach serves as a robustness check on our previous findings.

Overall, consideration of heterogeneous models allows us to study the impact of globalisation on the *share-capital* (SK) schedule sidelining the debate of whether the TFP is, or not, an accurate proxy of technology.[11] Likewise, given the way they control for unobservables, common factor models are suitable for accounting for structural breaks and business cycle distortions, thus making the use of yearly data perfectly valid.

[10]The inclusion of fixed effects in the 2FE estimators also control for unobservable time-invariant heterogeneities.

[11]The trade-off is, of course, that the impact of technology on the *share-capital* (SK) schedule cannot be studied explicitly, as we do when homogeneous models are estimated.

1.4 Data and stylised facts

1.4.1 Data

Table 1.1 lists the variables used and offers a synoptic definition.[12] Labour shares, capital-output ratios, GDP per capita, and employment growth rates are obtained from the Extended Penn World Table (EPWT 4.0), developed by Adalmir Marquetti and Duncan Foley. From the World Development Indicators (WDI) we get the manufacturing share over GDP, a variable that tries to control for the sectoral economic composition. The proxy for globalisation comes from the KOF index database, which accounts for social, economic, and political globalisation (Dreher, 2006), trade union density from the OECD, and national oil prices from the International Monetary Fund (IMF). The Polity II and TFP indices are obtained, respectively, from the Policy IV and PWT 8.0 databases.

Table 1.1: Data description

Variable	Description	Source
$\ln(s_L)$	(log of) the labour share	(1)
$\ln(k)$	(log of) the capital-to-output ratio	(1)
$\ln(RGDP)$	(log of) real GDP per capita in 2005 US	(1)
Δn	Employment growth rate	(1)
$\ln(MAN\_SHARE)$	(log of) manufacturing, value added (% of GDP)	(2)
KOF	KOF globalisation index	(3)
$\ln(UNION)$	(log of) union members as % of total paid employment	(4)
$\ln(OIL)$	(log of) national oil price	(5)
DEM	Polity II index of democracy	(6)
$\ln(TFP)$	(log of) the Total Factor Productivity index	(7)

Notes: (1) Extended Penn World Table (EPWT 4.0); (2) World Bank Development Indicators; (3) KOF Index; (4) OECD; (5) International Monetary Fund; (6) Policy IV democracy score; and (7) Penn World Table 8.0.

Karabarbounis and Neiman (2014) and Rognlie (2015) have recently made emphasis on the importance for the analysis of the computation of the labour income share and the stock of capital (including/excluding taxes, net vs gross, imputation of mixed income, etc...). The EPWT 4.0 draws information from different United Nations sources and measures the labour income share as the share of total employee compensation in the Gross Domestic Product with no adjustment for mixed rents. In turn, the numerator in the capital-output ratio is the real

[12]Tables 1.A2 and 1.A3 in the Appendix show the main descriptive statistics by group of countries.

net capital stock calculated by the Perpetual Inventory Method (PIM) using the investment series from the PWT 7.0, whereas the denominator is the real Gross Domestic Product. In spite of some caveats surrounding investment data from PWT regarding their quality (Srinivasan, 1995), or the impossibility of abstract from residential structures as there is no disaggregation by type of investment, this database is used on account of the nature of our analysis covering a large number of both developed and developing countries with yearly basis observations.[13]

1.4.2 Stylised facts

Figure 1.2 shows the evolution of our main variables of interest for both the OECD and non-OECD countries. These trajectories are obtained as time dummy coefficients from a GDP weighted regression of the variable of interest against a set of time and country dummies. The initial value is normalised to 100 to facilitate the comparisons.[14]

Figure 1.2a uncovers a parallel falling trend in the labour income shares of the OECD and non-OECD economies, the latter starting before and being steeper. It is, thus, a worldwide phenomenon with different intensities.

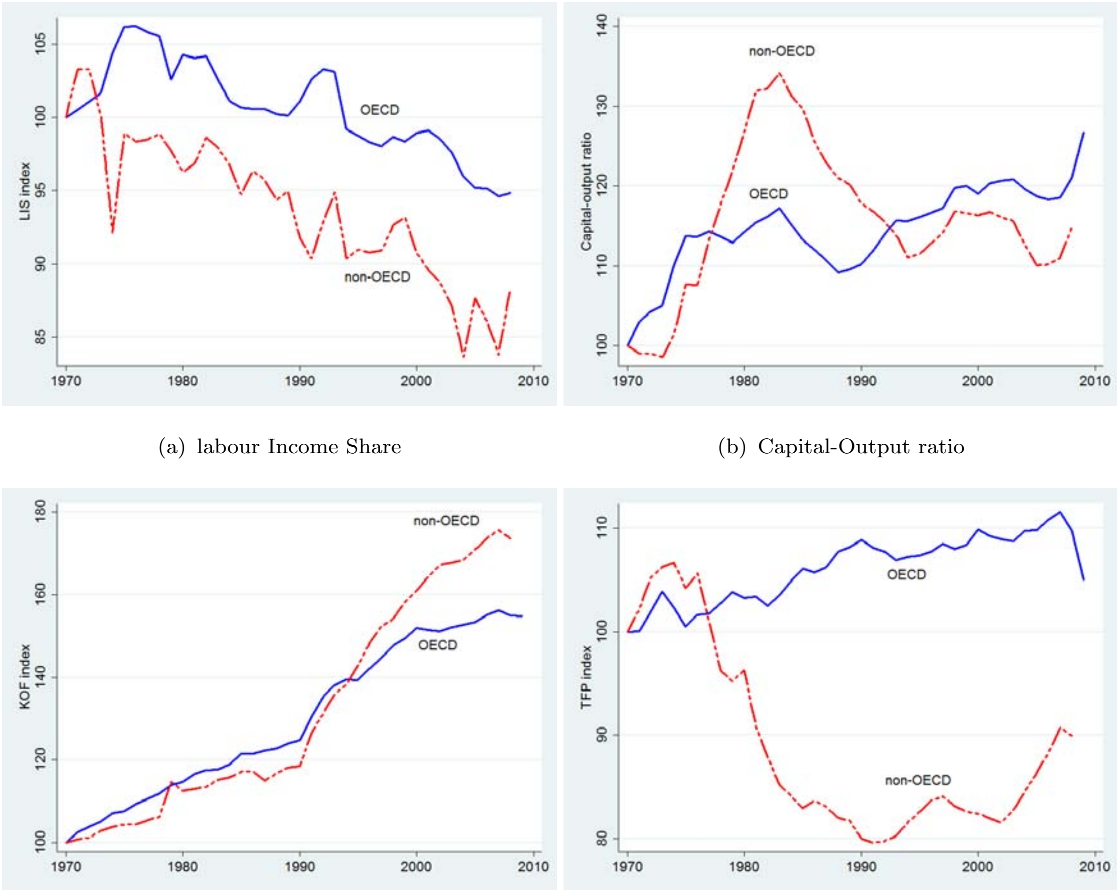
Figure 1.2b shows the evolution of the capital-output ratio, which is more volatile in the non-OECD economies. It grows faster than in the OECD between 1970 and early 1980s, but it deteriorates quicker afterwards (by more than 20 pp) until the mid 1990s, to stay relatively flat thereafter. Overall, the ratio has increased by less than 15 percent, while the continuous positive trend in the OECD area since the late 1980s has led this ratio to grow by around 25 percent.

Regarding globalisation, Figure 1.2c displays a common and positive upward trend between 1970 and 1990. Then the non-OECD economies experienced a faster exposure to globalisation as shown by the much steeper path followed by the KOF index in this area until the Great Recession years. Figure 1.A1c shows that even with the change in the growth rate of globalisation in 1990, non-OECD countries have a degree of globalisation around 20 pp lower than OECD countries.

[13]Both variables have been recently used by Young and Lawson (2014) in their study of the labour income share.

[14]To complement this information, Figure 1.A1 in the Appendix shows the evolution of the labour income share, the capital-output ratio and the KOF index taking as initial value the weighted average at 1970. TFP is not included in this figure, as its variation is within country (the index is equal to 1 for all the countries in 2005), making the different value between groups uninformative.

Figure 1.2: Labour income share, capital-output ratio, KOF and TFP, 1970-2009



(a) labour Income Share

(b) Capital-Output ratio

(c) KOF index of globalisation

(d) TFP index

Notes: Own calculations obtained as year fixed effects from a GDP weighted regression including country fixed effects to control for the entry and exit of countries throughout the sample. Initial year is normalised to equal 100 in 1970.

With respect to technological progress, the pattern is quite different. There is a sort of constant (albeit not large) rate of progress in the OECD countries only crushed by the Great Recession at the end of the sample period. In contrast, non-OECD countries suffered a severe collapse in the aftermath of the oil price shocks lasting until 1990 (with a 25 percent fall) to regain, afterwards, a positive trend, and end up 10 percent below the starting level in 1970.[15]

[15] To complement this general information, Figures 1.A2 - 1.A4 in the Appendix present country-specific correlation coefficients of the capital-output ratio, the KOF index of globalisation, and TFP with respect to the labour income share (to provide the most global picture, these Figures contain information for a wider sample than the one that we could actually use in the analysis due to data limitations). Clear pictures emerge in the first two cases, with worldwide positive and negative correlations in general across all economies. On the contrary, there is a much disperse result regarding TFP, with a negative correlation in most OECD countries, and a not so clear negative relationship in the non-OECD countries.

1.5 Estimated homogeneous dynamic models

1.5.1 Regression analysis

Tables 1.2 and 1.3 present our empirical estimations for the OECD and non-OECD countries. The base-run model is an augmented version of the empirical model in Bentolila and Saint-Paul (2003). Its results are always presented in the left three columns of the tables.

To the specification in Bentolila and Saint-Paul (2003) we have added some additional controls. The central one is the degree of globalisation, a relevant variable in the works by the IMF (2007) and Jayadev (2007), which in our case is proxied by the KOF Index (KOF). The second one is the share of manufacturing production (MAN_SHARE) in order to control for differences in the productive structure of the countries (Young and Lawson, 2014). A third one, is a variable accounting for the degree of democracy (DEM), since differences in this dimension may be specially relevant in the non-OECD countries. Finally, as suggested by Gollin (2002), we also consider real per capita GDP ($RGDP$) to control for the fact that our labour share measure does not adjust for self-employment incomes. This variable is also used by Jayadev (2007) as a proxy of economic development. A final difference with respect to Bentolila and Saint-Paul (2003) is that we use trade union density as a control instead of labour conflicts (due to data limitations, however, it cannot be used in the non-OECD model).

The columns in the right block of Tables 1.2 and 1.3 show the results when the interactions between the capital-output ratio (k), globalisation (KOF), and technology (TFP) are included in the model.

All models are estimated by Pooled OLS (POLS), Two-way Fixed Effects (2FE) and System GMM (BB), all including time dummies. In the System GMM estimation, we control for the potential endogeneity of s_L, k, TFP, Δn, $RGDP$, DEM and the interactions.[16] Our reference estimates are always the ones obtained by System GMM estimation, which are presented in the last column of Tables 1.2 and 1.3.[17]

[16] Ultimately, all variables in the model could be treated as endogenous. However, given the "instruments proliferation" problem, we are constrained to endogenise the most risky group. In this context, $RGDP$ and DEM are treated as predetermined.

[17] As a goodness check, note that the persistence coefficients obtained by the BB estimator lie between the ones estimated by POLS and 2FE (see Bond, 2002). They are the largest ones under the POLS estimation (0.90 in the OECD and 0.96 in the non-OECD areas, respectively), the lowest ones under the 2FE estimation (0.67 and 0.79), and in a middle position when estimated

Table 1.2: Estimated models for the OECD countries

	BASE-RUN			MODEL WITH INTERACTIONS		
	OLS	2FE	BB	OLS	2FE	BB
$\ln(s_{L_{t-1}})$	1.102	0.966	1.081	1.092	0.955	1.04
	(0.070)***	(0.104)***	(0.037)***	(0.071)***	(0.103)***	(0.039)***
$\ln(s_{L_{t-2}})$	-0.24	-0.273	-0.257	-0.235	-0.286	-0.317
	(0.061)***	(0.065)***	(0.057)***	(0.060)***	(0.064)***	(0.058)***
$\ln(k_t)$	0.015	0.11	0.106	-0.011	0.49	0.647
	(0.009)*	(0.044)**	(0.120)	(0.109)	(0.322)	(0.247)***
KOF_t	0.065	-0.064	0.151	0.049	0.136	0.562
	(0.046)	(0.085)	(0.123)	(0.078)	(0.274)	(0.191)***
$\ln(TFP_t)$	-0.031	-0.038	-0.026	0.19	-0.253	-0.287
	(0.028)	(0.035)	(0.093)	(0.177)	(0.481)	(0.272)
$\ln(OIL_t)$	0.001	-0.015	-0.001	0.001	-0.026	-0.005
	(0.003)	(0.011)	(0.003)	(0.004)	(0.014)*	(0.003)
$\ln(UNION_t)$	-0.007	0.005	-0.002	-0.006	0.001	0.0001
	(0.004)*	(0.010)	(0.011)	(0.004)	(0.012)	(0.010)
Δn_t	-0.173	-0.067	0.867	-0.164	-0.073	-0.192
	(0.339)	(0.313)	(0.467)*	(0.345)	(0.300)	(0.519)
$\ln(MAN\_SHARE_t)$	0.024	0.046	0.018	0.022	0.028	0.006
	(0.013)*	(0.024)*	(0.029)	(0.014)	(0.028)	(0.024)
$\ln(RGDP_t)$	0.038	0.041	0.02	0.039	0.009	0.025
	(0.025)	(0.022)*	(0.027)	(0.028)	(0.030)	(0.020)
$\ln(k_t) * KOF_t$				0.051	-0.41	-0.743
				(0.131)	(0.380)	(0.313)**
$\ln(k_t) * \ln(TFP_t)$				-0.704	0.311	0.323
				(0.359)*	(0.591)	(0.489)
$\ln(TFP_t) * KOF_t$				-0.425	-0.039	0.029
				(0.224)*	(0.527)	(0.417)
$\ln(k_t) * \ln(TFP_t) * KOF_t$				1.27	0.425	0.364
				(0.476)***	(0.545)	(0.741)
Observations	621	621	621	621	621	621
Number of id	24	24	24	24	24	24
R-squared	0.95	0.96		0.95	0.96	
CD test	1.26	0.72	1.00	1.20	0.36	0.11
Int	I(0)	I(0)	I(0)	I(0)	I(0)	I(0)
$AR(1)$			0.13			0.15
$AR(2)$			0.46			0.40
Hansen			1.00			1.00
N. instruments			63			83

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%. Endogenous variables: $\ln(s_{L_{t-1}})$, $\ln(k_t)$, $\ln(TFP_t)$, Δn_t, $\ln(k_t) * KOF_t$, $\ln(TFP_t) * KOF_t$, $\ln(k_t) * \ln(TFP_t)$, $\ln(k_t) * \ln(TFP_t) * KOF_t$; Predetermined variables: $\ln(RGDP_t)$; Exogenous variables: KOF_t, $\ln(OIL_t)$, $\ln(TU_t)$, $\ln(MAN\_SHARE_t)$. CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. AR(1), AR(2) show p-values of the Arellano and Bond (1991) test for no serial residual correlation. Hansen reports the p-value of the over-identifying restrictions test.

Table 1.3: Estimated models for the non-OECD countries

	BASE-RUN			MODEL WITH INTERACTIONS		
	OLS	2FE	BB	OLS	2FE	BB
$\ln(s_{L_{t-1}})$	0.983	0.883	0.945	0.969	0.866	0.937
	(0.075)***	(0.074)***	(0.082)***	(0.074)***	(0.073)***	(0.076)***
$\ln(s_{L_{t-2}})$	-0.016	-0.071	-0.122	-0.009	-0.074	-0.12
	(0.074)	(0.070)	(0.064)*	(0.074)	(0.069)	(0.057)**
$\ln(k_t)$	0.001	0.01	-0.024	-0.132	0.032	-0.024
	(0.008)	(0.021)	(0.037)	(0.053)**	(0.066)	(0.139)
KOF_t	0.087	0.08	0.246	-0.045	0.036	0.227
	(0.041)**	(0.083)	(0.150)	(0.062)	(0.108)	(0.204)
$\ln(TFP_t)$	0.049	-0.002	0.031	-0.13	-0.483	-0.655
	(0.016)***	(0.038)	(0.046)	(0.124)	(0.194)**	(0.343)*
$\ln(OIL_t)$	0.001	0.001	0.00	0.001	0.001	0.001
	(0.001)	(0.003)	(0.002)	(0.001)	(0.003)	(0.002)
$DEM2$	0.001	0.001	0.000	0.001	0.001	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Δn_t	0.023	-0.077	-0.89	0.012	0.006	-0.314
	(0.156)	(0.166)	(0.513)*	(0.159)	(0.156)	(0.440)
$\ln(MAN\_SHARE_t)$	0.002	-0.014	-0.012	0.001	-0.032	-0.011
	(0.008)	(0.021)	(0.019)	(0.008)	(0.024)	(0.019)
$\ln(RGDP_t)$	-0.002	0.049	0.02	0.007	0.047	0.02
	(0.005)	(0.029)*	(0.017)	(0.006)	(0.030)	(0.019)
$\ln(k_t) * KOF_t$				0.273	-0.069	0.008
				(0.107)**	(0.144)	(0.266)
$\ln(k_t) * \ln(TFP_t)$				0.162	1.127	1.334
				(0.238)	(0.366)***	(0.647)**
$\ln(TFP_t) * KOF_t$				0.457	1.256	1.775
				(0.314)	(0.460)***	(0.829)**
$\ln(k_t) * \ln(TFP_t) * KOF_t$				-0.435	-2.783	-3.362
				(0.563)	(0.838)***	(1.538)**
Observations	650	650	650	650	650	650
Number of id	27	27	27	27	27	27
R-squared	0.96	0.96		0.96	0.96	
CD test	-3.28	-3.35	-3.50	-3.23	-3.34	-3.59
Int	I(0)	I(0)	I(0)	I(0)	I(0)	I(0)
$AR(1)$			0.00			0.00
$AR(2)$			0.58			0.72
$Hansen$			1.00			1.00
N. instruments			62			82

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%. Endogenous variables: $\ln(s_{L_{t-1}})$, $\ln(k_t)$, $\ln(TFP_t)$, Δn_t, $\ln(k_t) * KOF_t$, $\ln(TFP_t) * KOF_t$, $\ln(k_t) * \ln(TFP_t)$, $\ln(k_t) * \ln(TFP_t) * KOF_t$; Predetermined variables: $\ln(RGDP_t)$, $DEM2$; Exogenous variables: KOF_t, $\ln(OIL_t)$, $\ln(MAN\_SHARE_t)$. CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. AR(1), AR(2) show p-values of the Arellano and Bond (1991) test for no serial residual correlation. Hansen reports the p-value of the over-identifying restrictions test.

To endorse the validity of our results, we conduct a series of specification tests. The AR and Hansen tests check for serial residual correlation and the validity of the instruments; the CD-test corresponds to Pesaran's (2004) test, which examines the cross-section independence of the residuals; finally, the cross-sectional augmented panel unit root (CIPS) Pesaran's (2007) test is used to analyze the residuals' order of integration (Int).

We verify that all the equations are clean of residual autocorrelation, well specified, and deliver stationary residuals. In turn, although the equations for the OECD area show cross-section independence, this is not the case in the non-OECD group. This implies that we have to be careful when interpreting the results for this area.

The base-run models indicate, at first glance, that the main drivers of the OECD countries' labour share are the capital-output ratio, the share of manufacturing production, and the real per capita GDP, all of them with the expected positive influence, together with the employment growth, which surprisingly comes out with a positive impact. Although the results for the non-OECD countries differ in some aspects, once we control for potential endogeneities, just the employment growth appears as significant, with the expected negative sign.[18]

This may reflect either that these variables are not relevant as labour share determinants, or that linear models are not the best tool when the explanatory variables are dependent on one another.

When the interactions are incorporated, the influence of the variables involved in the interactions cannot be assessed unless their marginal effects are properly computed. This is done in Section 1.5.2.

1.5.2 Marginal effects

The computation of the marginal effects allows the impact of the capital-output ratio on the labour share to be assessed at different values of the KOF index of globalisation and the log of TFP.[19]

by System GMM (0.72 and 0.82). We credit the latter and conclude that the labour share in the non-OECD area is more persistent than in the OECD countries.

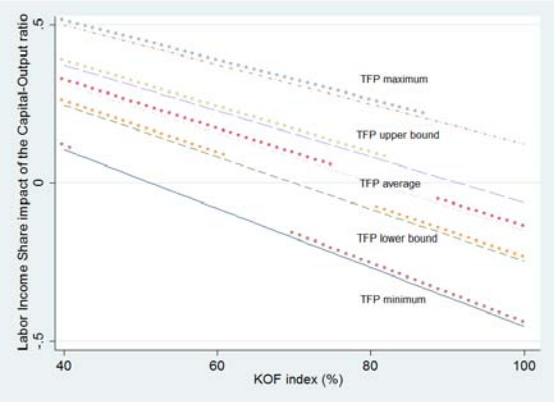
[18] Let us note that in the SK schedule framework, the irrelevance of the capital-output ratio as a regressor is consistent with a Cobb-Douglas production function (i.e., $\sigma = 1$).

[19] To choose these values, we conduct a Kernel density analysis, and select the range of values that fall within the 95% confidence interval given by a two standard deviation from the sample mean. Figure 1.A5 shows the Kernel densities of the KOF and TFP indices in the OECD and non-OECD areas, with the shaded areas indicating the selected values. For the OECD countries,

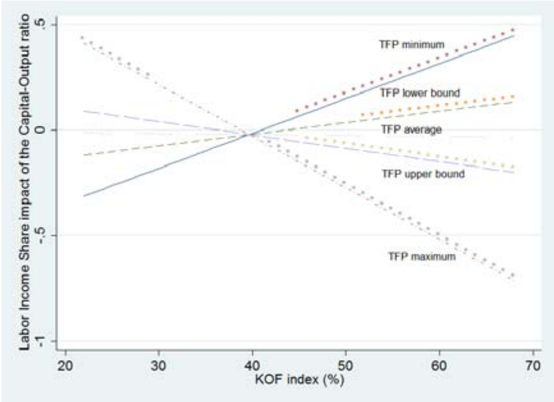
Figure 1.3 presents the estimated impact of the capital-output ratio on the labour share in the OECD and non-OECD areas respectively. Asterisks in these figures denote significance at the 90% confidence interval.

In Figures 1.3a and 1.3b, the continuum of values of the KOF index is presented in the horizontal axis. Then, the impact of the capital-output ratio on the labour share along these values is evaluated in five levels of TFP comprising the minimum, maximum, and average values of technology, and the upper and lower bounds computed as 1 standard deviation from the average.

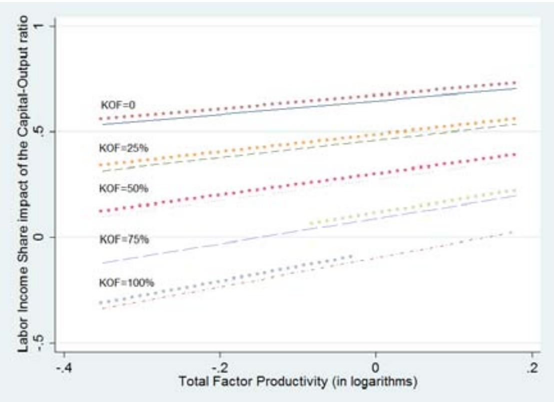
Figure 1.3: Marginal effects across varying levels of globalisation and technology



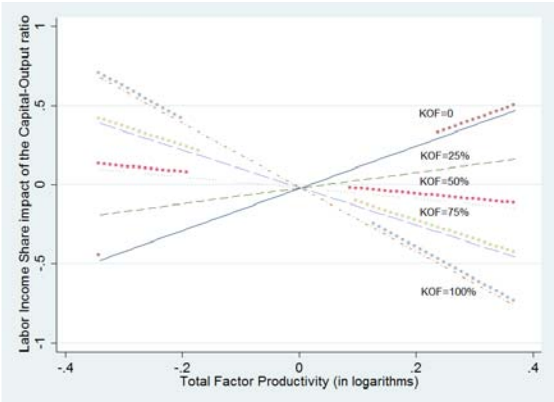
(a) OECD



(b) non-OECD



(c) OECD



(d) non-OECD

In Figures 1.3c and 1.3d, the effect of technology on the elasticity of the labour share with respect to the capital-output ratio is evaluated at selected levels of

they range from 40% to 100% for the KOF index, and from -0.35 to 0.18 for the TFP; for the non-OECD economies, they go from 22% to 68% for the KOF index, and range between -0.34 and 0.38 for the TFP. Note that the wider interval in the non-OECD area implies a larger volatility of the TFP, and does not reflect at all a better technological level.

globalisation. We have the continuum of values of the TFP in the horizontal axis, and 5 different trajectories of the KOF index ranging from one extreme case –a value 0% reflecting autarchy– to the other extreme case –a country 100% globalised. In the intermediate scenarios, we consider KOF index values of 25%, 50% and 75%.

For the OECD area (Figures 1.3a and 1.3c), we find that the larger the degree of globalisation is, the lower the impact of a change in the capital-output ratio on the labour share independently of the countries' technological level. To be more precise, if we take as reference the average value of TFP, an increase in the level of globalisation alters the impact of the capital-output ratio on the labour share from a positive value (0.3 when $KOF = 40\%$) to a negative one (-0.15 when $KOF = 100\%$).

Even though equation (1.2) shows that there are other factors affecting the SK schedule, this change in the sign can only be explained by a change in σ (recall expression 1.3), implying that globalisation has enhanced the substitutability between capital and labour, moving σ from below to above unity. This result supports previous evidence (see among others, Rodrik, 1997; Slaughter, 2001; Saam, 2008; Hijzen and Swaim, 2010), according to which globalisation processes (such as offshoring practices, or a larger market for intermediate inputs) allow companies to substitute easier away from labour in the case of an increase in its price.

With respect to technology, we find that the higher the technological level, the larger the impact of the capital-output is on the labour share, for whatever value taken by the KOF index. This implies a decrease in the degree of substitution between production factors (σ), and gives support to the "capital-skill complementarity" hypothesis (Griliches, 1969; Goldin and Katz, 1996; Caselli and Coleman, 2001), where "new technologies tend to substitute for unskilled labour in the performance of routine tasks while assisting skilled workers in executing qualified work" [Arpaia et al., 2009, footnote 10]. This assertion relies on the presumption that there is a larger share of high skill workers in the OECD economies (Krusell et al., 2000).

In the non-OECD countries, the effects of globalisation are not unanimous (Figures 1.3b and 1.3d). A first issue deserving attention is the presence of an inflection point around a value of 40% in the KOF index of globalisation. Below this point (i.e., for relatively closed economies), the impact of technology is mainly irrelevant. In contrast, for relatively high levels of globalisation (above 40%), the impact of the capital-output ratio on the labour share takes negative values (reflecting $\sigma > 1$)

when the technological level is relatively high. In this context, the more globalised a country is, the smaller the impact of the capital-output ratio on the labour share. This relationship, however, is the opposite at the lowest level of technology, in which case $\sigma < 1$ and the impact of the capital-output ratio on the labour share increases with globalisation.

Regarding technological change, we find evidence that TFP has the opposite role in the non-OECD than in the OECD countries, and increases the substitutability between production factors. This may be due to the growing mechanisation of industries exposed to trade, which are the ones that have received the bulk of foreign direct investment and are more subject to outsourcing and offshoring practices. Progressive substitution of traditional labour-intensive tasks by relatively more capital-intensive ones would explain the enhanced substitutatibility brought by technological progress in this area. Moreover, this effect is larger, the higher the degree of globalisation is (note that for higher levels of globalisation the curve in Figure 1.3d becomes steeper, so that the larger is the decrease in the impact of the capital-output ratio on the labour share due to technological progress).

Summing up, our results provide clear evidence that globalisation in the OECD area has helped to increase the substitutability between production factors, whereas technological progress has, on the contrary, pushed their complementarity. In contrast, the results for the non-OECD economies are more heterogeneous, although we find evidence that technological change increases the substitutability between production factors. In terms of globalisation, the complex relationships uncovered for the non-OECD countries requires further research which, at present, is hindered by data limitations preventing further disaggregation of our analysis.

Next section presents the common factor model results to further examine the role of globalisation on the labour share impact of the capital-output ratio.[20]

[20]In the event of slow capital stock changes and a counter-cyclical behavior of the labour share, yearly data analysis could reflect a spurious positive correlation between the capital-output ratio and the labour income share. To exclude this possibility, Figure 1.A6, in the Appendix, shows the marginal effects for a 3 years average static model estimated by System GMM. It can be observed that our results are robust both for the OECD and non-OECD countries and, thus, we can safely rule out the possibility of a spurious positive correlation.

1.6 Estimated heterogeneous dynamic models

1.6.1 Regression analysis

Table 1.4 presents the results for the whole sample using both homogeneous and heterogeneous models. The first three columns present the estimation using the standard POLS and 2FE models, along with the pooled CCEP. In turn, the last four columns present the Mean Group-style estimators, which do not impose parameter-homogeneity across countries.[21] From now on, the latter constitute our reference regressions.

TABLE 1.4: Homogeneous vs heterogeneous models

	Homogeneous slope			Heterogeneous slopes			
	[1] POLS	[2] 2FE	[3] CCEP	[4] MG	[5] CMG	[6] CMG1	[7] CMG2
$\ln(LIS_{t-1})$	0.967	0.831	0.64	0.547	0.523	0.468	0.358
	(0.010)***	(0.033)***	(0.050)***	(0.042)***	(0.037)***	(0.043)***	(0.057)***
$\ln(k_t)$	-0.029	0.003	0.006	0.057	0.313	0.598	0.564
	(0.019)	(0.035)	(0.084)	(0.240)	(0.285)	(0.392)	(0.462)
$\Delta\ln(k_t)$	0.183	0.17	0.128	0.423	0.315	0.253	0.272
	(0.054)***	(0.050)***	(0.068)*	(0.049)***	(0.075)***	(0.073)***	(0.079)***
KOF_t	0.029	0.022	0.01	0.285	0.486	0.735	0.762
	(0.018)	(0.059)	(0.098)	(0.222)	(0.262)*	(0.290)**	(0.324)**
ΔKOF_t	-0.166	-0.214	-0.272	-0.023	-0.174	-0.167	-0.042
	(0.125)	(0.117)*	(0.137)**	(0.123)	(0.130)	(0.093)*	(0.144)
$\ln(k_t) * KOF_t$	0.062	0.069	0.121	-0.221	-0.722	-1.079	-1.137
	(0.030)**	(0.061)	(0.136)	(0.394)	(0.466)	(0.600)*	(0.705)
Constant	-0.017	-0.158	-0.047	-0.477	-0.449	-0.695	-0.373
	(0.017)	(0.044)***	(0.125)	(0.134)***	(0.287)	(0.314)**	(0.304)
Observations	1586	1586	1586	1586	1533	1397	1341
Number of id	51	51	51	51	47	41	40
CD-test	-0.18	-0.59	-1.77	3.78	-1.63	-1.71	-1.5
Int	I(0)	I(0)	I(0)	I(0)	I(0)	I(0)	I(0)
RMSE	0.0530	0.0516	0.0468	0.0407	0.0326	0.0291	0.0236

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%.
POLS = Pooled OLS (with year dummies), 2FE = 2-way Fixed Effects, CCEP = Pooled Pesaran (2006) Common Correlated Effects (CCE), MG = Pesaran and Smith (1995) Mean Group (with country trends), CMG = Pesaran (2006) CCE Mean Group, CMG1 and CMG2 = CMG with, respectively, one and two extra cross-sectional averages lags, as indicated by Chudik and Pesaran (2015).
CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. RMSE presents the root mean squared error.

[21] There is only one estimated model rejecting the cross-sectional independence of the residuals, the Mean Group estimator (MG). Moreover, none of them show problems related to the integration order of the residuals.

This new estimation delivers a positive impact of the capital-output ratio and globalisation, but a negative one of the interaction, which is robust across the different Mean Group estimators. Given the presence of the interaction term, the corresponding marginal effects (i.e., the influence of globalisation on the elasticity of the labour share with respect to the capital-output ratio) must be assessed at the relevant range of values. Next we explain our strategy regarding the computation of the marginal effects.

1.6.2 Marginal effects

In contrast to Section 1.5, the marginal effects are computed by exploiting the country-specific coefficients, along with the country-specific average globalisation level. The vertical axis in Figure 1.4 presents the country-specific coefficients obtained from the Pesaran (2006) Common Correlated Effects Mean Group estimator (CMG).[22] The horizontal axis shows the country-specific average level of globalisation in our sample period. To explore the potential existence of broad patterns, we fit the sequence of country observations with a fractional polynomial regression line. One standard deviation coefficient intervals are also added to help in the interpretation of the results.[23]

Figures 1.4a, 1.4c and 1.4d show the interaction coefficients in the vertical axis for, respectively, the total, OECD and non-OECD samples.[24] In turn, Figures 1.4b, 1.4e and 1.4f show the impact of the capital-output ratio on the labour share evaluated at the average level of globalisation.[25]

The most general picture, depicted by Figures 1.4a and 1.4b for the total sample, shows substantial heterogeneity. In Figure 1.4a, however, we observe a broadly negative influence of globalisation on the labour share impact of the capital-output

[22]We are aware that some authors have warned against the study of country-specific coefficients in an isolated way (Pedroni, 2007; Eberhardt and Teal, 2013b). For this reason, we do not analyze the specific coefficient of a given country, but the existence of a potential pattern for different average levels of globalisation.

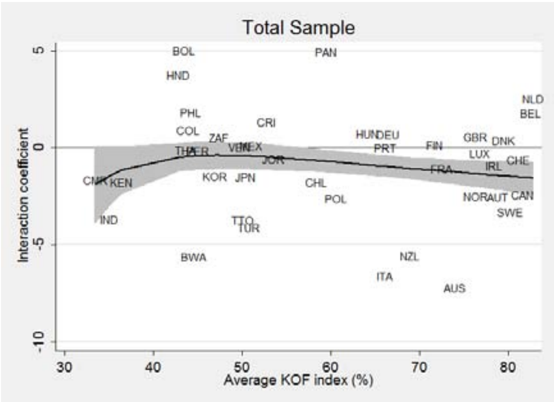
[23]Country-specific coefficients are obtained from the regressions' first stage. These results are robust to the estimation method. No significant differences appear when we use the Chudik and Pesaran (2015) Dynamic CMG estimator (CMG2). Our decision is based on the larger number of countries included in the CMG estimation.

[24]It is important to note that the interaction coefficient is an evaluation of the slope of the SK schedule.

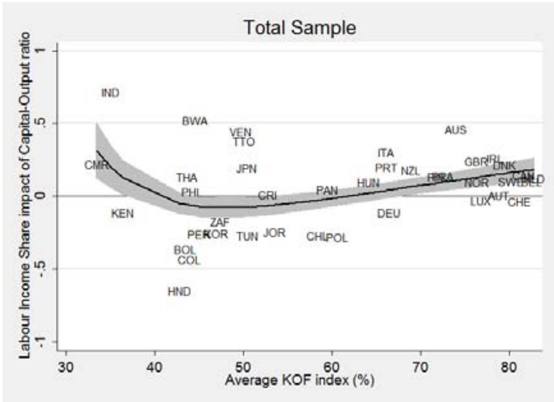
[25]The graphical analysis in this section is based on Eberhardt and Presbitero (2015), using their do-file publicly available at Markus Eberhardt's website (https://sites.google.com/site/medevecon/home).

ratio, which is significant for levels of globalisation larger than 55% (i.e., globalisation decreases the impact of the capital-output ratio on the labour share). In turn, in Figure 1.4b we observe a positive impact of the capital-output ratio on the labour share for globalisation levels above 65%.

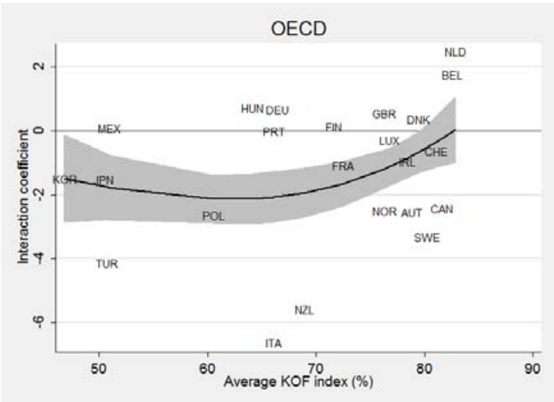
FIGURE 1.4: Heterogeneous coefficients



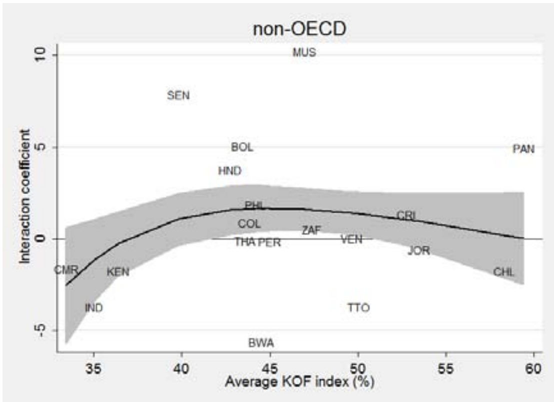
(a) Interaction coefficients, total sample



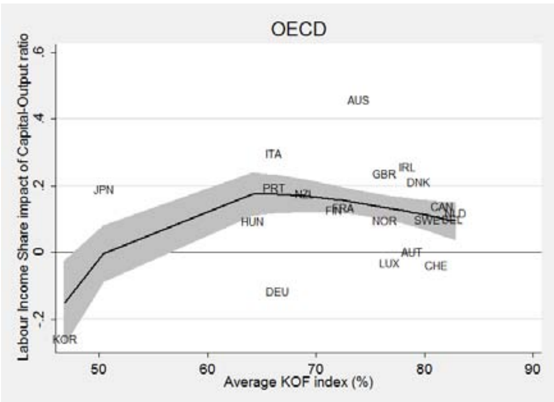
(b) k impact coefficients, total sample



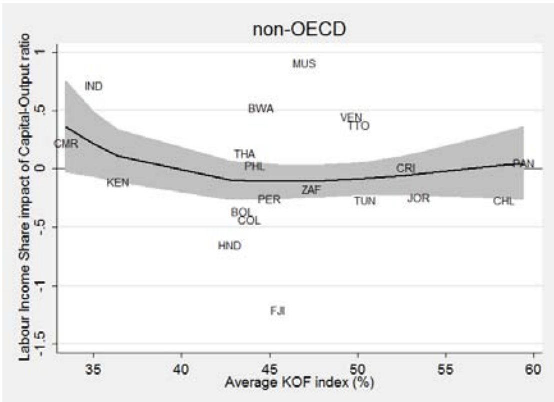
(c) Interaction coefficients, OECD sample



(d) Interaction coefficients, non-OECD sample



(e) k impact coefficients, OECD sample



(f) k impact coefficients, non-OECD sample

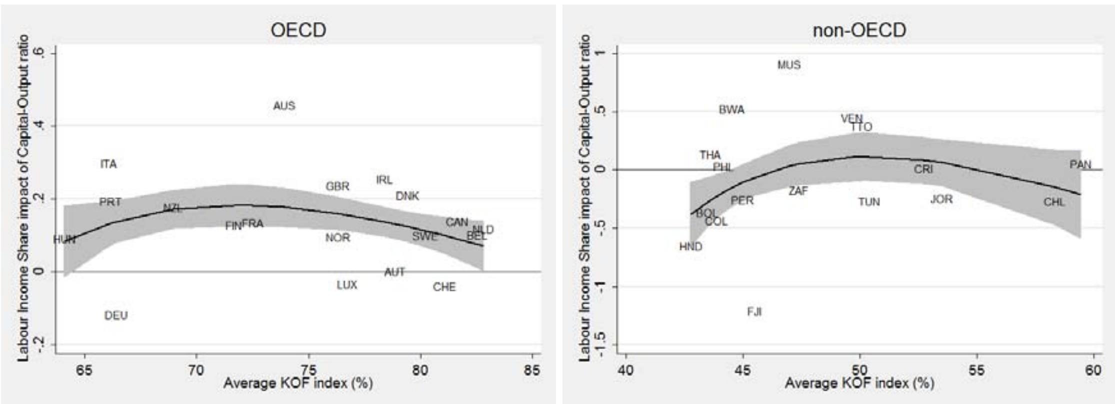
Notes: Country-specific coefficients for the interaction term and the capital-output impact on the labour income share for an average level of globalisation against the average level of globalisation. Coefficients are taken from the first CMG stage using Stata's *rreg* command to account for outliers. A fitted fractional polynomial regression line is added along with $\pm$ one standard deviation (shaded area).

To check the robustness of our results in Section 1.5, we split the sample between OECD and non-OECD countries (Figures 1.4c and 1.4d). We find that the interaction coefficients in the OECD reproduce the finding for the total sample, although the negative influence is somewhat smaller, and eventually non-significant, at very high levels of globalisation (Figure 1.4c). In sharp contrast, in the non-OECD area we find a significant positive influence of globalisation at values between 40% and above 50% (Figure 1.4d). As a final outcome, note that the OECD results drive those for the total sample within the range at which the effects of globalisation on $\frac{ds_L}{dk}$ are significant (non-OECD economies are in general less globalised).

Figures 1.4e and 1.4f, which isolate by areas the information presented in 1.4b, are to some extent comparable to Figures 1.3a and 1.3b (the difference is that this time the variation is obtained from the country-specific average degree of globalisation without accounting for different technological scenarios). These figures disclose a positive impact of the capital-output ratio on the labour share in OECD countries (decreasing when the KOF index is higher than 65%), in contrast to a slightly and non-significant U-shape impact in the non-OECD.

Exclusion of countries with an extremely low-average degree of globalisation (Figure 1.5)[26] confirm our previous results for the OECD: a positive impact of the capital-output ratio on the labour share, which is smaller the larger the degree of globalisation is. For the non-OECD, however, the picture remains non-significant, with a mild inverse U-shape relationship.

FIGURE 1.5: Heterogeneous coefficients, constrained sample



(a) Impact coefficients, OECD sample (b) Impact coefficients, non-OECD sample

Notes: Country-specific coefficients for the labour share impact of the capital-output ratio evaluated at the average level of globalisation, constrained to be larger than 55% in the OECD and 40% in the non-OECD countries. Coefficients are taken from the first CMG stage using Stata's *rreg* command to account for outliers. A fitted fractional polynomial regression line is added along with ± one standard deviation (shaded area).

[26] Japan and Korea from the OECD; Cameroon, India and Kenya from the non-OECD group.

The evidence of a non-significant impact of the capital-output ratio on the labour share, no matter the degree of globalisation, is consistent with a Cobb-Douglas technology since, by construction, it implies a unit value of σ. Of course, this does not exclude divergences by sectors or countries from this specific production function, which would just be reflecting the situation in terms of the average technology across sectors and countries.

These results reassert the conclusions reached in Section 1.5 for the OECD area. In contrast, for the non-OECD countries we find a non-significant impact of the capital-output ratio on the labour share irrespective of the degree of globalisation. Next, we further dig into these results by checking whether the departing point (i.e., from a relatively closed or open economy) is relevant for the effects that changes in globalisation exert on $\frac{ds_L}{dk}$.

1.6.3 Asymmetries

So far, we have studied the influence of globalisation and technology on $\frac{ds_L}{dk}$ by assuming that this influence is constant across degrees of openness and technological levels. Next, we check for the possibility of asymmetries by conducting a new evaluation of this influence at high and low globalisation regimes.

Beyond the potential identification of different patterns when the analysis is conducted for small and large degrees of globalisation, this further exercise will also allow us to check for the possible existence of an inflection point in the non-OECD countries as found in Section 1.5 (Figures 1.3b, 1.3d).

This analysis is based on Shin et al. (2013) and Eberhardt and Presbitero (2015), and requires the estimation of the following model:

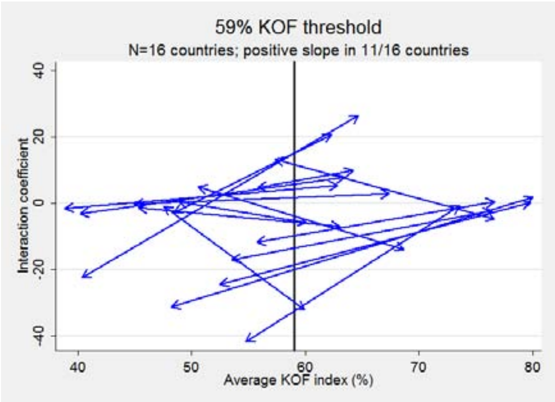
$$\ln(s_{L_{it}}) = \beta_0 + \beta_1 \ln(k_{it}) + \beta_2 KOF_{it}^{+} + \beta_3 KOF_{it}^{-} + \beta_4 \ln(k_{it}) * KOF_{it}^{+} + \beta_5 \ln(k_{it}) * KOF_{it}^{-} + \epsilon_{it}, \tag{1.15}$$

where the globalisation index is decomposed into partial sums above or below a specific threshold. For example, as explained in Eberhardt and Presbitero (2015), if we were to chose a threshold of 0 (i.e., we separate increases from decreases in globalisation) we would have:

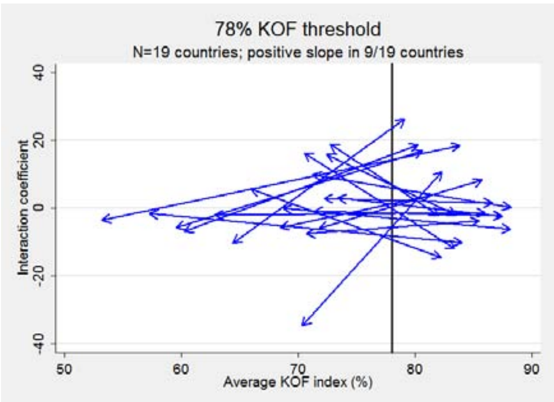
$$\begin{aligned} KOF_{it}^{+} &= \sum_{j=1}^{t} \Delta KOF_{it}^{+} = \sum_{j=1}^{t} max(\Delta KOF_{it}, 0) \\ KOF_{it}^{-} &= \sum_{j=1}^{t} \Delta KOF_{it}^{-} = \sum_{j=1}^{t} min(\Delta KOF_{it}, 0) \end{aligned} \tag{1.16}$$

Given our will to preserve enough degrees of freedom, we use as (ad hoc) threshold the median of globalisation. In order to avoid imprecise coefficient estimations, we only consider countries where at least 10% of all observations are in one regime. We run three different regressions for the total sample (16 countries, threshold 59%), 19 OECD countries (threshold 78%) and 15 non-OECD countries (threshold 45%), and examine whether systematic differences in the interaction coefficients arise when globalisation increases from relatively low or high levels.

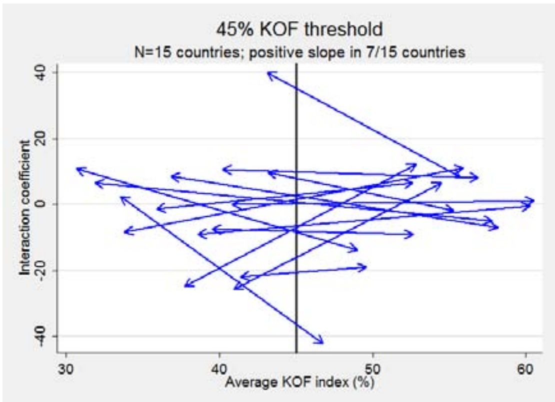
FIGURE 1.6: Marginal effect slopes



(a) Interaction coefficients, Total sample



(b) Interaction coefficients, OECD sample



(c) Interaction coefficients, non-OECD sample

Notes: Interaction coefficients in the low and high globalisation regimes. Coefficients are obtained from a CMG estimation of equation (1.15). x-axis represent the average level of globalisation for the lower and higher regimes.

Figure 1.6 presents the information obtained from the new estimation of the interaction coefficients β_4 and β_5 in equation (1.15).[27] It is important to remark that the interaction coefficients (whose units are in the vertical axis) represent the slope of the relationship between the labour income share and the capital-output ratio

[27]This estimation is a standard CMG model where only the dependent variable and the capital-output ratio introduce dynamics.

studied in previous sections. This information is now presented taking the form of arrows, with left arrow tips reflecting the value of the interaction coefficient in a relatively low globalisation scenario (β_5), and right arrow tips showing the interaction coefficient in a relatively high globalisation scenario (β_4). The horizontal axis shows each country's average level of globalisation for these two globalisation regimes.

To preview a simple case, let us assume that globalisation has no impact on σ when changing in the low regime, but it decreases σ when changing in the high regime. In that case, Figure 1.6 would deliver systematic negatively sloped arrows.

Looking at Figure 1.6, however, we observe an eloquent absence of systematic behaviors. This holds irrespective of the sample under analysis (total, OECD, non-OECD), and does not confirm the existence of an inflection point, as observed through the estimation of homogeneous models.

TABLE 1.5: Descriptive statistics of the interaction coefficients

Threshold	Interaction coefficients					Impact coefficients				
	Obs	Mean	Std	Min	Max	Obs	Mean	Std	Min	Max
Total sample										
lower	16	-8.53	14.94	-41.79	13.11	16	-0.07	0.65	-1.61	1.10
upper	16	0.53	13.30	-32.16	26.36	16	0.91	1.60	-1.05	4.76
OECD										
lower	19	-0.71	11.92	-34.85	18.83	19	0.37	0.56	-0.23	2.26
upper	19	2.48	11.26	-14.64	26.31	19	0.70	1.55	-2.56	4.08
non-OECD										
lower	15	-0.60	16.94	-25.66	40.11	15	-0.25	0.75	-2.17	1.24
upper	15	-2.89	14.28	-42.13	12.22	15	-0.15	1.12	-1.72	1.89

To complement this result, the left block of Table 1.5 provides the descriptive statistics corresponding to the interaction coefficients presented in Figure 1.6. Interestingly, the standard deviation of the interaction coefficients is virtually the same does not matter the regime. However, when we observe the rest of indicators (mean, minimum and maximum), we observe a heterogeneous pattern between OECD and non-OECD countries. While these indicators systematically increase in the OECD area, what implies a decrease in the negative slope or, in other words, a less negative impact of globalisation on the SK schedule, exactly the opposite holds for the non-OECD sample. In addition, information on the corresponding impact coefficients can be found in Figure 1.A7 and in the right block of Table

1.5. The main feature of these results is a larger dispersion of the capital-output impact on the labour share for relatively high degrees of globalisation.

Overall, therefore, we cannot provide evidence of systematic patterns in the influence of globalisation on the SK schedule even when distinguishing low and high globalisation levels.

1.7 Conclusions

We use Bentolila and Saint-Paul's (2003) framework to analyze the interplay between globalisation, technology, and the elasticity of substitution between capital and labour (σ). In this context, we adopt, from Antony (2009a,b, 2010), the possibility that σ varies along with different relative factor intensities. It is through the role of globalisation and technology as key determinants of relative factor intensities that we bring the study of their influence on a varying σ.

We do so by estimating multiplicative interaction models to reappraise the impact of the capital-output ratio on the labour share when globalisation and technology are allowed to influence this impact. We work with the largest possible amount of observations to be able to conduct separate analyses for the OECD and the non-OECD areas. These analyses are first performed through the estimation of homogeneous models. Then, we move to the estimation of a common factor model through Mean Group-style estimators. This allows us to exploit the cross-country heterogeneity, so as to analyze the robustness of our first set of results, and the possibility of asymmetries arising from different scenarios of small and large globalisation levels.

Our findings provide a robust picture for the OECD countries, where we find a positive impact of the capital-output ratio on the labour share; a larger substitutability between production factors along with the globalisation process; and, in contrast, a larger complementarity driven by technological progress. These results are in line, respectively, with international trade literature (Slaughter, 2001; Saam, 2008), and the capital-skill complementarity hypothesis (Arpaia et al., 2009).

The results for the non-OECD area are much less conclusive. Nevertheless, we find evidence of an increase in the substitutability between capital and labour as a consequence of technological improvement, and a non-significant impact of the capital-output ratio on the labour share irrespective of the degree of globalisation (which would be consistent with an average aggregate Cobb-Douglas technology).

One extra result of interest is the absence of evidence of an asymmetric relationship between globalisation and the *share-capital* (SK) schedule. In other words, the fact that globalisation may vary departing from a relatively low or high regime does not systematically alter the labour share response to changes in capital intensity.

The magnitude of σ is critical both for economic growth and factor income distribution. While it has been documented that a larger σ could boost potential growth, it could also put pressure on labour conditions by decreasing workers' bargaining power and rising functional inequality. It follows that the relevance of globalisation and technological change as drivers of σ deserve further attention so as to avoid unexpected and undesirable effects from their progress.

Further research should aim at clarifying the role of globalisation in developing countries, where economic heterogeneities and difficulties in the access to long time series of high quality data hinder the analysis. With respect to the OECD, the natural step forward is to examine whether globalisation and technology have the same influence across sectors and types of workers –skilled/non-skilled– to evaluate their impact from a more disaggregated perspective.

APPENDIX: Supplementary tables and figures

TABLE 1.A1: Selected economies and sample period

OECD		NON-OECD	
Country	Sample	Country	Sample
Australia	1990-2008	Argentina	1993-2007
Austria	1976-2008	Bolivia	1970-2008
Belgium	1995-2008	Botswana	1983-2001
Canada	1981-2008	Brazil	1992-2008
Denmark	1970-2009	Cameroon	1979-2004
Finland	1975-2009	Chile	1970-2008
France	1970-2009	Colombia	1970-2007
Germany	1990-2008	Costa Rica	1983-2008
Hungary	1995-2008	Ecuador	1970-1993
Ireland	1990-2008	Fiji	1983-2001
Italy	1970-2008	Honduras	1980-2005
Japan	1980-2007	India	1980-2008
Korea, Rep	1970-2003	Iran	1994-2007
Luxembourg	1995-2008	Jordan	1970-2003
Mexico	1992-2008	Kenya	1970-2008
Netherlands	1970-2008	Mauritius	1980-2000
New Zealand	1971-2008	Namibia	1990-2003
Norway	1970-2007	Niger	1995-2008
Poland	1995-2008	Panama	1980-2008
Portugal	1988-2009	Peru	1986-2003
Sweden	1980-2009	Philippines	1980-2008
Switzerland	1990-2007	Senegal	1991-2008
Turkey	1986-2003	South Africa	1970-2008
UK	1990-2008	Thailand	1970-2003
		Trinidad & Tobago	1984-2003
		Tunisia	1992-2007
		Venezuela	1970-2006

Table 1.A2: Descriptive statistics, OECD

Panel A: Raw variables OECD						
Variable	Obs	Mean	Median	sd	Min	Max
s_L	621	0.496	0.514	0.077	0.188	0.642
k	621	1.625	1.553	0.328	0.896	2.974
KOF	621	0.738	0.783	0.138	0.279	0.925
TFP	621	0.942	0.973	0.091	0.634	1.139
OIL	621	1821.059	38.6	6600.553	0.023	47561.52
$UNION$	621	0.383	0.335	0.212	0.075	0.839
Δn	621	0.012	0.010	0.017	-0.046	0.200
MAN_SHARE	621	0.195	0.193	0.047	0.079	0.307
$RGDP$	621	25750.31	25752.97	10726.2	3297.037	89814.25

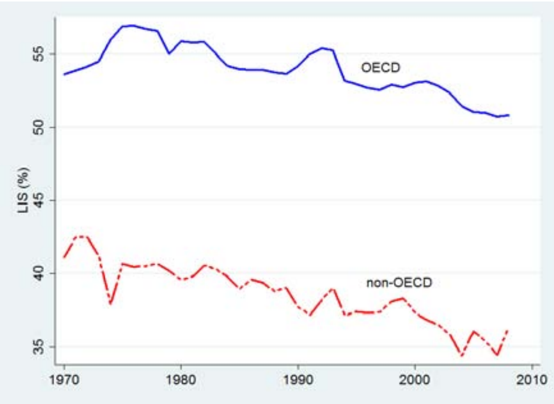
Panel B: regression variables (in logs) OECD						
Variable	Obs	Mean	Median	sd	Min	Max
$\ln(s_L)$	621	-0.715	-0.665	0.182	-1.669	-0.443
$\ln(k)$	621	0.466	0.440	0.194	-0.109	1.090
$\ln(TFP)$	621	-0.064	-0.028	0.102	-0.456	0.130
$\ln(OIL)$	621	4.176	3.653	2.228	-3.791	10.770
$\ln(UNION)$	621	-1.125	-1.094	0.595	-2.584	-0.176
$\ln(MAN\_SHARE)$	621	-1.667	-1.645	0.258	-2.544	-1.182
$\ln(RGDP)$	621	10.059	10.156	0.477	8.101	11.406

TABLE 1.A3: Descriptive statistics, non-OECD

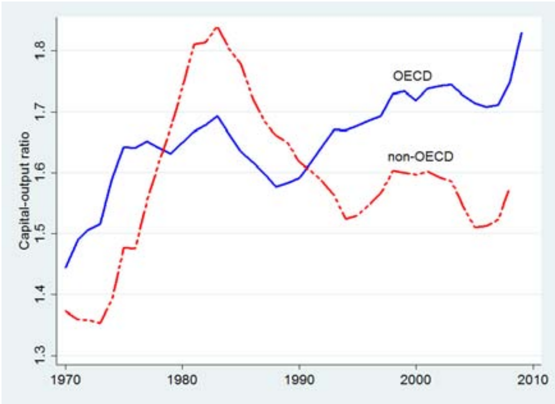
Panel A: Raw variables non-OECD						
Variable	Obs	Mean	Median	sd	Min	Max
s_L	650	0.353	0.355	0.091	0.127	0.628
k	650	1.693	1.501	0.615	0.792	4.368
KOF	650	0.461	0.453	0.108	0.261	0.743
TFP	650	1.009	0.989	0.174	0.615	1.656
OIL	650	22682.99	314.226	143572.8	0.001	1910899
DEM	650	4.197	7.000	6.104	-10.000	10.000
Δn	650	0.029	0.028	0.014	-0.092	0.160
MAN_SHARE	650	0.169	0.165	0.056	0.048	0.348
$RGDP$	650	4881.324	4448.124	2846.24	475.774	18771.87

Panel B: regression variables (in logs) non-OECD						
Variable	Obs	Mean	Median	sd	Min	Max
$\ln(s_L)$	650	-1.079	-1.036	0.280	-2.066	-0.465
$\ln(k)$	650	0.471	0.406	0.326	-0.233	1.474
$\ln(TFP)$	650	-0.005	-0.011	0.166	-0.487	0.505
$\ln(OIL)$	650	5.667	5.750	3.655	-9.320	14.463
$\ln(MAN\_SHARE)$	650	-1.840	-1.800	0.374	-3.042	-1.054
$\ln(RGDP)$	650	8.291	8.400	0.691	6.165	9.840

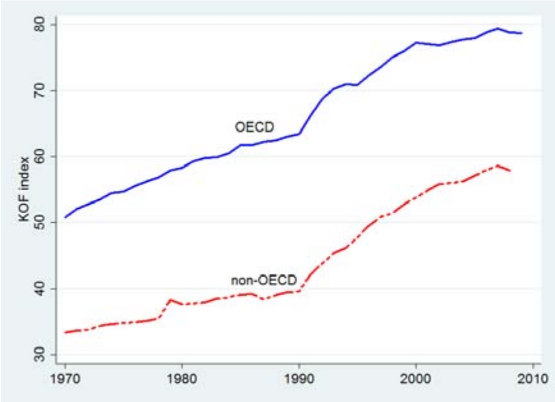
FIGURE 1.A1: Labour income share, capital-output ratio and KOF, 1970-2009



(a) labour Income Share



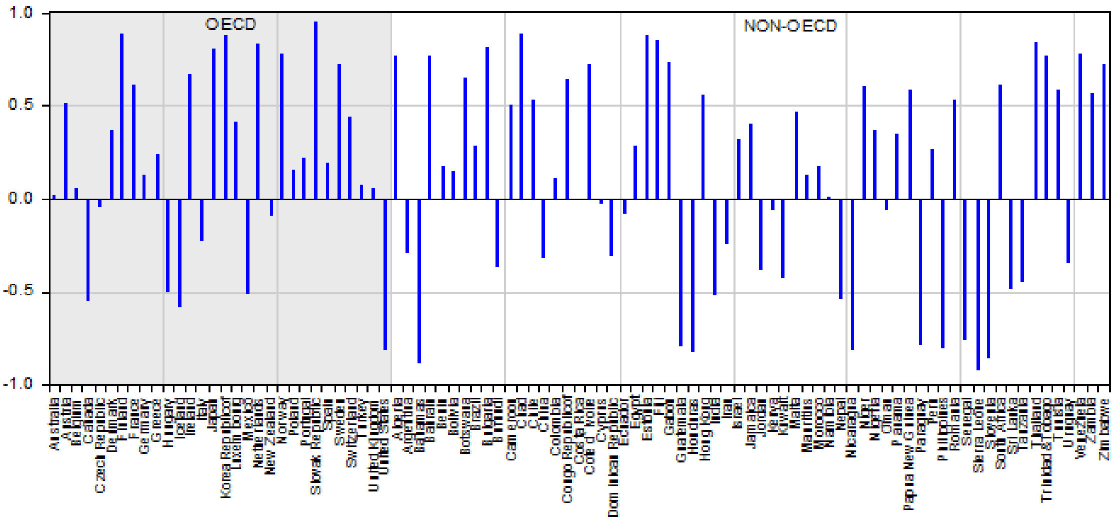
(b) Capital-Output ratio



(c) KOF index of globalisation

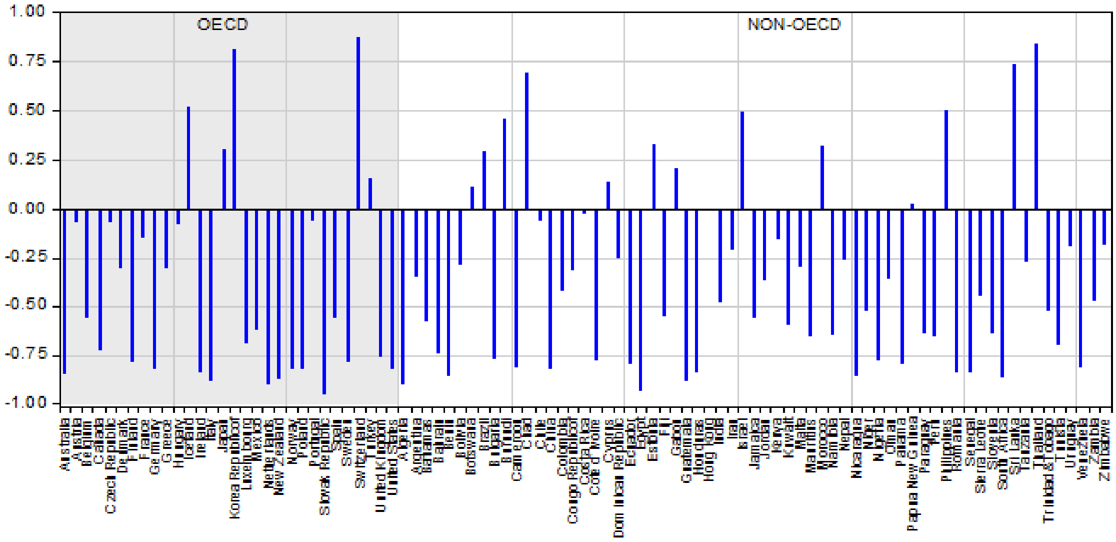
Notes: Own calculations obtained as year fixed effects from a GDP weighted regression including country fixed effects to control for the entry and exit of countries throughout the sample. The initial value equals the weighted average in our dataset in 1970.

FIGURE 1.A2: Correlation coefficients between labour shares and the capital-output ratio



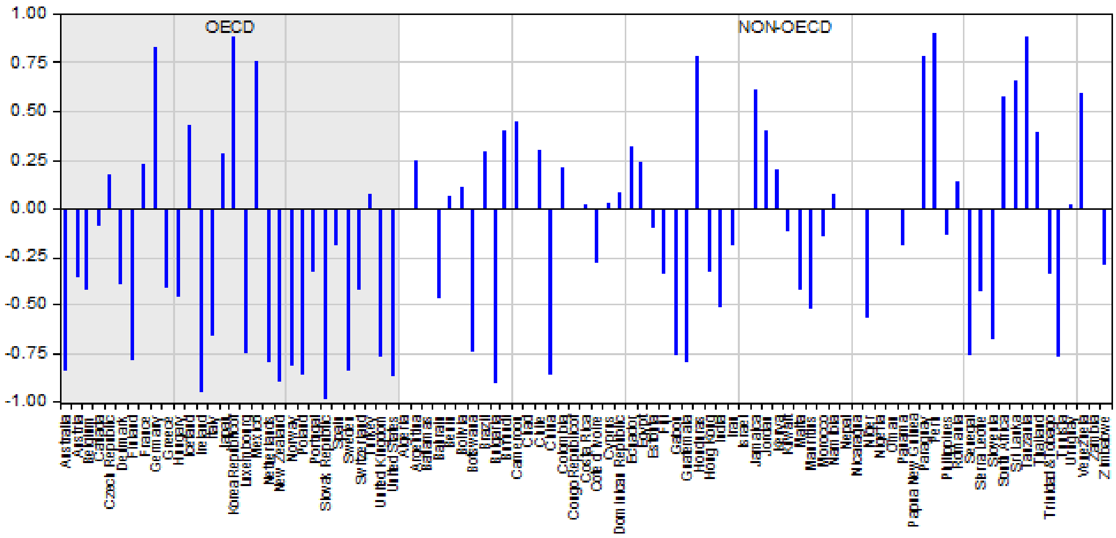
Source: Extended Penn World Table (EPWT 4.0). 1970-2009.

FIGURE 1.A3: Correlation coefficients between the labour shares and the KOF index



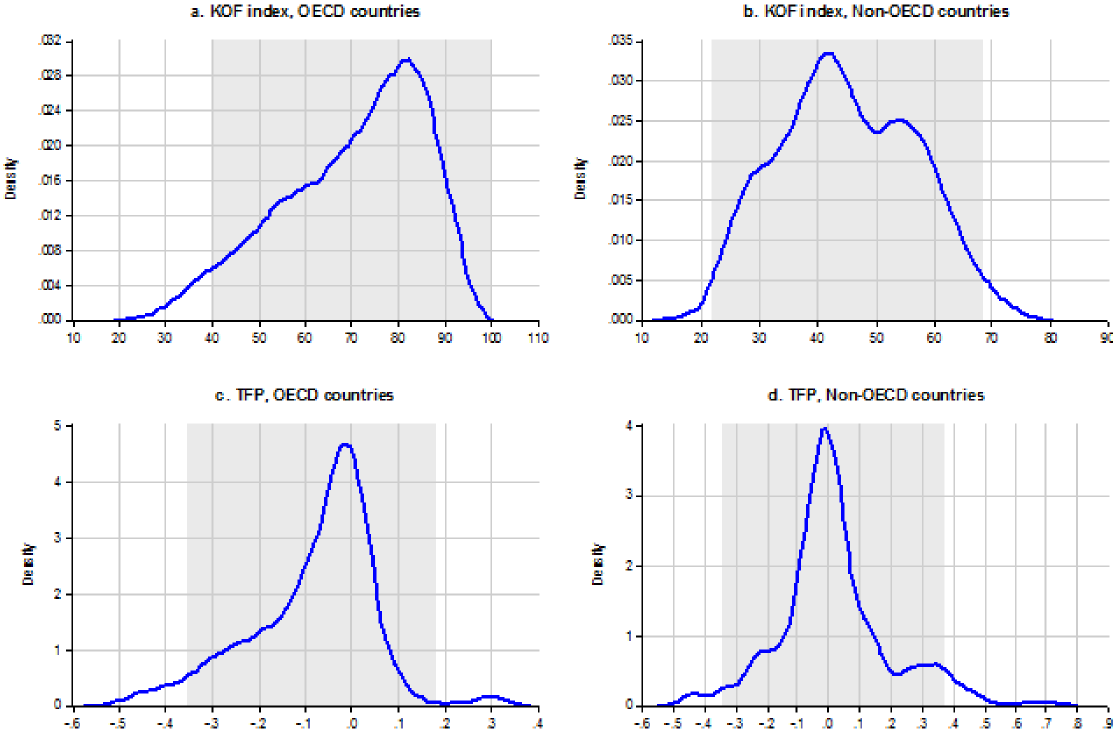
Source: Extended Penn World Table (EPWT 4.0) and KOF Index. 1970-2009.

FIGURE 1.A4: Correlation coefficients between the labour shares and the TFP



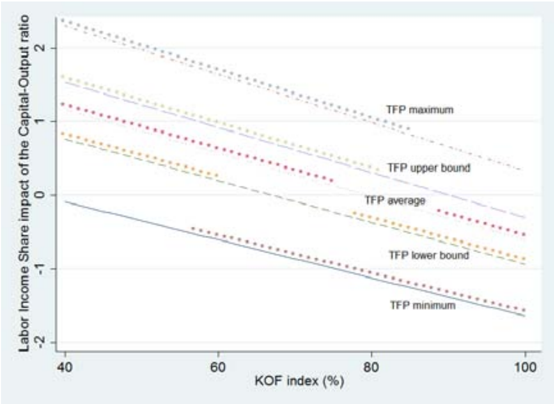
Source: Extended Penn World Table (EPWT 4.0) and PWT 8.0. 1970-2009.

FIGURE 1.A5: Kernel density estimates of the KOF index and TFP

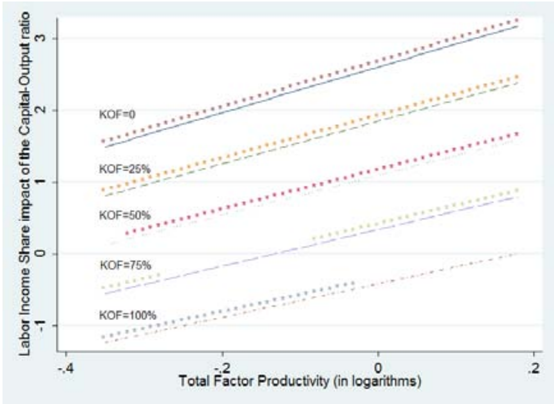


Notes: TFP is measured in logs.

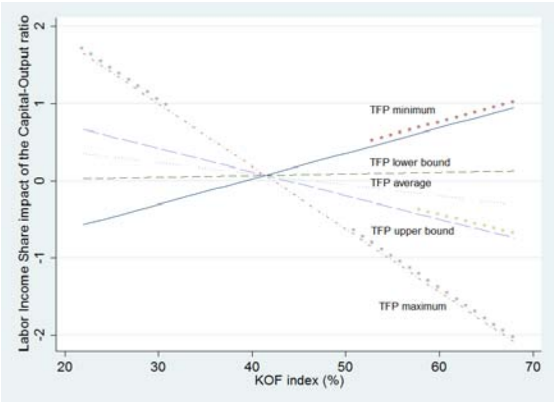
FIGURE 1.A6: Marginal effects: 3 years average



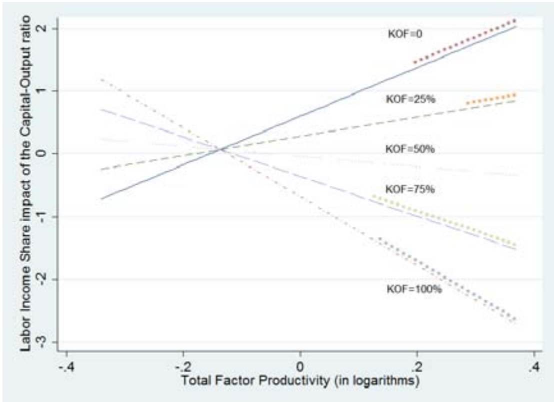
(a) OECD: Across KOF



(b) OECD: Across TFP

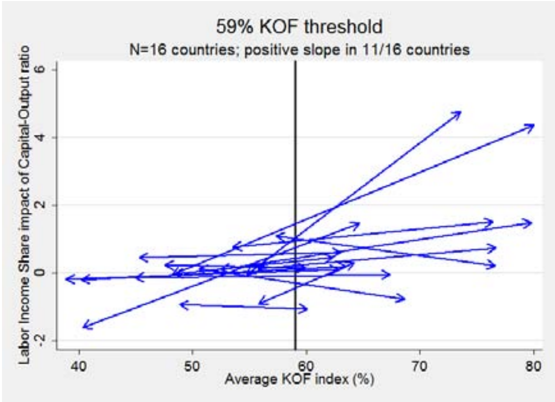


(c) non-OECD: Across KOF

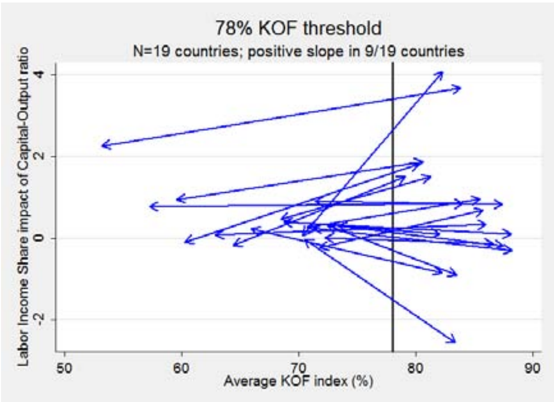


(d) non-OECD: Across TFP

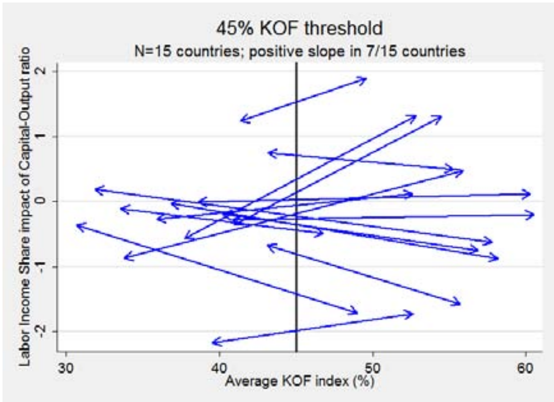
FIGURE 1.A7: k impact coefficients



(a) k impact coefficients, Total sample



(b) k impact coefficients, OECD sample



(c) k impact coefficients, non-OECD sample

Notes: Capital-output ratio impact coefficients for an average level of globalisation in the low and high globalisation regimes. Coefficients are obtained from a CMG estimation of equation (1.15). x-axis represent the average level of globalisation for the lower and higher regimes.

Chapter 2

Finance and the global decline of the labour share

Abstract

The labour income share has been decreasing across countries since the early 1980s, sparking a growing literature about the causes of this trend (Elsby et al., 2013; Karabarbounis and Neiman, 2014; Piketty and Zucman, 2014). At the same time, again since the early 1980s, there has been a global steady increase in equity Tobin's Q which shows -we argue- the increasing role of finance in modern economies. This paper uses a simple model to connect these two phenomena and evaluates its empirical validation. In our model a raise in equity Tobin's Q increases equity returns and, importantly, depresses the capital-output ratio. The impact on the capital-output ratio reduces the labour share for standard values of the elasticity of substitution. Based on a common factor model, we find that the increase in Tobin's Q explains up to 57% of the total decline in the labour income share. This implies that financial markets have direct and significant consequences in inequality through their impact on the functional distribution of income. We highlight the implications of this result and, in the context of the model, we suggest different policies that can revert this declining trend.

JEL Codes: E25, E44, E22.

Keywords: Labour Share, Tobin Q, Finance, Capital-Output Ratio.

This paper is coauthored with Ignacio Gonzalez (Ph.D. Candidate at European University Institute).

2.1 Introduction

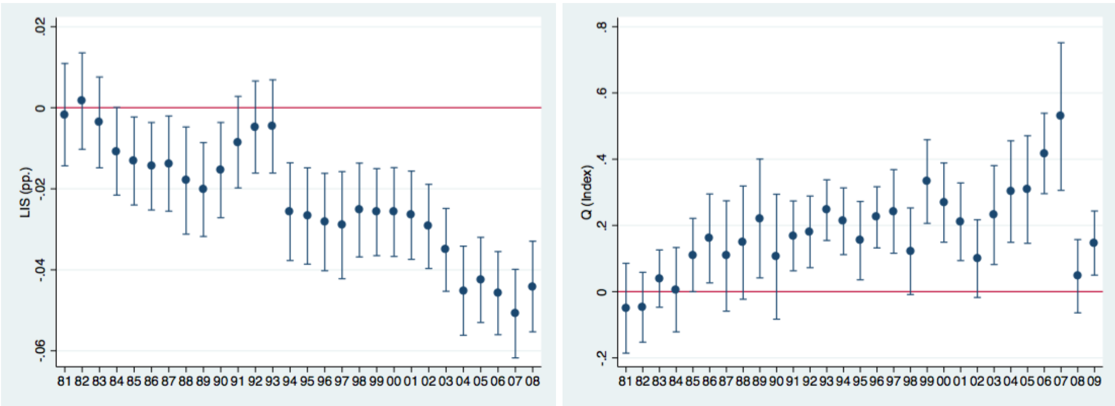
The decline of the labour income share is becoming an increasing popular research topic. The constancy of factor shares, once featured among Kaldor's stylised facts of economic growth, has been challenged by recent literature. For example, Karabarbounis and Neiman (2014) document that the global labour share has declined significantly since the early 1980s, with the decline occurring within the large majority of countries and industries. Elsby et al. (2013) use alternative measures of the labour share and provide convincing evidence of this declining trend for the U.S. economy. More noticeably, Piketty and Zucman (2014) show, for a set of advanced economies, a decreasing (increasing) trend of the labour (capital) income share since the late 1970s.

Figure 2.1.a shows the figure of concern. It displays the evolution of the global labour share according to our data by plotting the year fixed effects from a GDP weighted regression along with its 90% confidence intervals. Country fixed effects are included to eliminate the influence of countries entering and exiting the data set. Taking 1980 as the reference year, we observe that the global labour share has exhibited a clear downward trend only disrupted by the sudden -but short-rise in the early nineties. If we normalised 1980 to equal its weighted average value (57%), labour share reaches a level of roughly 52% at the end of the sample, implying an actual decline of 8.9 per cent during the period considered.

Attempts to explain the non-constant behaviour of the labour share have often departed from reconsidering at least two previously standard assumptions namely - that aggregate technology is Cobb-Douglas and that markets are perfectly competitive. Explanations based on departures from the Cobb-Douglas production function usually assume that technology is characterised by a constant elasticity of substitution (CES) production function. As long as firms produce with a CES technology and the labour market is perfectly competitive, the labour share can be expressed as a function of the capital-output ratio, $LIS = g(K/Y)$. Given this relation, this literature emphasises the role of capital deepening as the main determinant of the labour share. This is the case in Bentolila and Saint-Paul (2003), who refer to this relationship as the share-capital schedule (or curve). This relationship is not altered by changes in factor prices or quantities, or in labour-augmenting technical progress, which are all encompassed in the schedule. Note that within this curve, when everything else is constant, labour share dynamics can only be explained if the economy is not on a balanced growth path, meaning that capital and output are not growing at the same rate, like in Piketty and Zucman (2014) and Karabarbounis and Neiman (2014).

Labour and product market imperfections are also frequently brought up as explanatory factors of the labour share decline. Even when technology is Cobb-Douglas, movements of factor shares can be triggered by changes in the bargaining power of workers and/or in the monopoly power of firms, that is to say, factors that break the equality between marginal costs and marginal products/revenues.

FIGURE 2.1: Labour income share and Tobin's Q, 1980-2009



(a) Labour Income Share (b) Tobin's Q

Notes: Own calculations obtained as year fixed effects from a GDP weighted regression including country fixed effects to control for the entry and exit of countries throughout the sample.

In light of the previous potential explanatory departures, which are the actual drivers of the downward trend of the labour share? The literature has pointed out three potential candidates: (i) globalisation, (ii) the institutional framework, and (iii) structural/technological causes. This paper contributes to the debate by exploring the role played by a new factor: the relation between financial markets and corporate investment, which we connect with the evolution of equity Tobin's Q.

The presence of globalisation as a driving force candidate is not surprising. The unprecedented global integration process that economies have experienced during the last decades has substantially altered many economic relationships. With regard to the distribution of income, from a theoretical perspective, globalisation has an ambiguous effect. On one side, the relative larger capital mobility makes easier for a company to change the location of its production. The change of location can decrease the labour share by the simple elimination of jobs. In addition, given the increasing international competition, firms can also use this as a threat to decrease the bargaining power of workers (Rodrik, 1997) and, thus, the labour income share. On the other side, globalisation has a counterbalance effect by increasing product competition. This increase in competition decreases firms' mark-ups, and this can have a positive impact on the labour income share. Therefore, impact of globalisation is something that has to be empirically evaluated. Guscina (2006) and Jaumotte and Tytell (2007) find a negative relationship between globalisation and the labour income share in developed countries. In their analysis, they include different globalisation proxies such as: international trade, trade with developing countries, offshoring, and the export/import relative prices, finding a robust negative relationship.[1] More recently, Elsby et al. (2013) study the role played by the offshoring process in the U.S. labour share decline. They find that the increased exposure to imported goods accounted for 85% of the total decline in the past quarter century. Therefore, empirical studies suggest a negative impact of globalisation on the labour share of income.

The role of the institutional framework has also received a strong attention in the study of factor share dynamics. The literature has focused on the impact of both labour and product market regulations. Kristal (2010), for example, finds that dynamics of the labour share are largely explained by indicators for workers' bargaining power. Blanchard and Giavazzi (2003) emphasise that labour market regulations have a positive effect in the short-run, but negative in the long-run, because in the long-run employers can substitute capital for relatively more expensive labour. With respect to product market regulations, Raurich et al. (2012) find a negative relationship between imperfect competition and the labour share, showing that estimates of the elasticity of substitution are biased when price mark-ups are ignored. Finally, Azmat et al. (2012), investigating a different channel, find that a fifth of the total labour share decline observed is a consequence of the privatisation of public companies through job shedding.

Despite the documented importance of globalisation and market regulations, most of the current literature has focused on structural/technological explanations. This branch of the literature relies on the aforementioned one-to-one relation between the labour share and the capital-output ratio $LIS = g(K/Y)$, where the direction of the effect depends on the elasticity of substitution between capital and labour. The contribution of this literature relies on looking at structural drivers of the labour share by making endogenous the dynamics of the capital-output ratio. Piketty and Zucman (2014), for example, argue that a persistent gap between

[1] It is worthy to note that both papers find that technology has played a more important role than globalisation.

the rate of return to capital and the growth rate of output results in a growing accumulation of capital because capitalists save most of their income. This would explain the observed movements of the factor shares in advanced economies. Also based on the share-capital schedule, Karabarbounis and Neiman (2014) argue that the persistent global decrease in the relative price of investment goods has induced firms to use more capital at the expense of labour, increasing the accumulation of physical capital and depressing the labour income share. They model capital-biased technological change using a version of the model presented in Greenwood et al. (1997).

Note that although Piketty and Zucman (2014) and Karabarbounis and Neiman (2014) emphasise different channels, both use the share-capital schedule and have the common view that the increase in the capital-output ratio has been the main cause of the recent trend of factor shares. In response to higher capital accumulation, and due to low diminishing returns, the return to capital has not adjusted sufficiently downwards and this has led to an increase in the capital share. In mathematical terms, this is equivalent to say that the elasticity of substitution is larger than one. Only when capital and labour are, in the technological sense, substitutes enough, can capital be accumulated without decreasing much its rate of return.

This degree of substitutability, however, has seldom been found in the empirical literature. Economists have often estimated values of the elasticity of substitution (σ) far below one (Antràs, 2004; Chirinko, 2008). Notably, Chirinko and Mallick (2014) using a sectoral dataset and combining a low-pass filter with panel data techniques, find an aggregate elasticity of substitution of 0.4. Furthermore, when they allow the elasticity to differ across sectors, they find that all the sectoral values are below 1. In the context of the current debate, they convincingly argue that the secular decline in the labour share of income cannot be explained by decreases in the relative price of investment, or by any other mechanism that increases the capital-output ratio.

In this paper we contribute to this recent literature by proposing a new mechanism and by evaluating its empirical validation with recent panel data techniques. Our mechanism emphasises the role of finance and the relation between financial markets and corporate behaviour. In particular, we argue that the widespread increase in equity Tobin's Q has occurred at the expense of corporate investment and the labour income share. We provide a simple model which shows that when equity Tobin's Q raises, the capital-output ratio falls. Importantly, this fall has a standard impact on the labour income share because it requires a value of the elasticity of substitution in line with the estimates traditionally found in the empirical literature.

Our theoretical argument is the following. When equity Tobin's Q raises, financial wealth raises accordingly and, to hold this additional wealth, investors demand a higher return on equity. In any standard model, like ours, equity returns are linked to the marginal productivity of capital. This implies that if firms want to increase the return on equity, they are forced to reduce their capital stock. This depresses the capital-output ratio and has a direct impact on the labour share.

Note that the mechanism of our model makes use of the share-capital schedule: we impact the labour share through changes in the capital-output ratio. In that sense, our paper is close to Piketty and Zucman (2014) and Karabarbounis and Neiman (2014). However, here the share-capital schedule works very differently. In response to an increase in equity Tobin's Q, investment and the capital-output ratio fall, not raise, and equity returns raise, not fall. This way, the mechanism suggests that it is not too much investment what causes the decline of the labour income share, like in Piketty and Zucman (2014) and Karabarbounis and Neiman (2014), but too little, and that is why our model is compatible with standard values of the elasticity of substitution (i.e., $\sigma < 1$).

Our theory gives rise to several questions: Does the raise of Tobin's Q capture the impact of finance on factor shares? What is behind the global evolution of Tobin's Q? And more importantly, is it a relevant mechanism? We certainly do not argue that Tobin's Q is a perfect indicator of financial activity and we neither try to show that Tobin's Q is a variable that captures the whole impact of finance on the capital-output ratio and the factor shares. We simply check the empirical validation of a model that shows that when Tobin's Q raises, the equilibrium capital-output ratio and the equilibrium labour share fall.

Our mechanism is relevant but, to the best of our knowledge, empirically unexplored. It is first relevant because it resembles the widely discussed arguments used by the literature on financialisation (Epstein, 2005; Davis, 2009). This literature studies the increasing role of financial markets and financial motives within the non-financial corporate sector. In particular, it emphasises inequality mechanisms that raise equity wealth and corporate payouts but depress corporate investment, just like in our model.

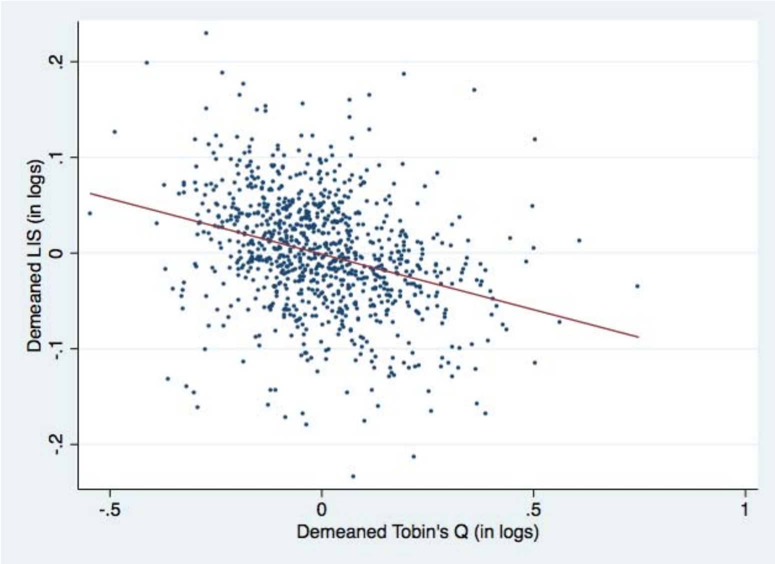
There are different mechanisms whose impact can be encompassed through an increase of equity Tobin's Q. Among them, financialisation literature has emphasised the role played by "shareholder-value ideology". According to this literature, corporations, after the early 80s, tend to pursue short-term payout policies that increase the equity price but that happen to be at the expense of long-term investment (Lazonick and O'Sullivan, 2000). However, there are other mechanisms which can also increase the price of equity and reduce the investment on physical capital. These are, for example, the decrease of dividend income tax rates (Anagnostopoulos et al., 2012; McGrattan and Prescott, 2005), the decline of stock market transaction costs and the raise of monopoly rents (Gonzalez, 2016). Any of these mechanisms can be embodied in a version of the model we present below. The important thing is that they all impact the equilibrium capital-output ratio and the labour share through an increase in equity Tobin's Q. For this reason we prefer to abstract from any specific factor and focus on the impact of Tobin's Q alone.

Figure 2.1.b shows the evolution of global Tobin's Q according to our data by plotting the year fixed effecs from a GDP weighted regression where 1980 is taken as the reference year (1980 = 0). If we consider the weighted average value in 1980, Figure 2.1.b shows a steady Tobin's Q increase from a value below 1.2 to values around 1.7 in 2007.

Figure 2.2 is more illustrative. It shows a negative correlation between the labour income share and the Tobin's Q when we control for country fixed effects. It is therefore the figure that better anticipates the answer to our research question. For our empirical analysis, we rely on recently developed panel time-series techniques that account for macroeconomics data characteristics. In particular, we present different Mean Group-style estimators which rely on a common factor model approach. The main advantage of our empirical approach is that it allows us to control for unobservable heterogeneity in a very tractable way while we can control for the panel time-series characteristics of macro data (i.e., cross-section dependence and nonstationarity) and allow for a country-specific impact of our variables of interest. We opt to further control for the relative price of investment goods to contrast our mechanism with that of Karabarbounis and Neiman (2014).[2]

[2]Changes in the relative price of investment goods impacts the capital-output ratio but they do not change the Tobin's Q.

FIGURE 2.2: Labour income share against Tobin's Q



Notes: Own calculation based on a sample of 41 countries and 911 observations. Variables are demeaned to control for fixed-effects.

Our results show a robust and significant negative impact of the Tobin's Q on the labour income share that can explain up to 57% of its decline since 1980. However, we do not find any significant impact of the relative price of investment goods. Like Chirinko and Mallick (2014), our results suggest that the decline of the labour income share cannot be explained by the secular decline of the relative price of investment.

Since Tobin's Q impacts the labour share through an endogenous decline of the capital-output ratio, our results reconcile the secular decline in the labour income share with standard values of the elasticity of substitution. This is starkly in contrast with the explanations given by Piketty and Zucman (2014) and Karabarbounis and Neiman (2014). We consequently conclude that deep causes for the secular decline of the labour share have to be found not in the mere accumulation of physical capital or in capital biased-technological changes, but in the way financial markets and corporations relate. In particular, the deep causes for functional inequality should be found in those policies or institutional changes that increase financial wealth at the expense of investment.

The remaining of the paper is structured as follows. Section 2.2 develops the theoretical framework relating the Tobin's Q with the labour income share. Section 2.3 introduces and explains the data used in our empirical analysis. Section 2.4 and 2.5 present, respectively, the econometric methodology and the results. Section 2.6 summarises and concludes.

2.2 Theoretical framework

This section presents a model that connects the labour share with financial wealth, physical capital stock and the Tobin's Q. We consider a representative agent economy where households accumulate financial wealth and receive direct utility from the ownership of wealth. The firm accumulates physical capital and distribute dividends to households.

2.2.1 Households

There is a representative household whose maximisation problem is the following (in recursive form):

$$\begin{aligned} V(s) &= \max_{c,s'} \frac{c^{1-\mu}}{1-\mu} + \frac{\gamma(p_{-1}s)^{1-\theta}}{1-\theta} + \beta V(s') \\ s.t. & \quad c + ps' = w + (Qd + p)s, \end{aligned}$$

where, c represents consumption, w is the average wage, d is the amount of dividends distributed in the current period, p is the price of stocks and s represents the amount of stocks owned today. Note that s is a state variable which was decided yesterday. Q represents the fraction of dividend income received by the households. At every given period there is one equity share outstanding. Hence, market clearing in the market for shares requires $s = 1$ for any period.

In a frictionless economy, Q equals one and households receive the total amount of dividends distributed by the firm. In our case, we assume that there is a constant fraction of the dividend income $1 - Q$ which does not go to households. An obvious example of this type of friction is a capital income tax which detracts from households a constant fraction of dividends. However, other frictions, like financial transaction costs, can be thought similarly.

We show later that Q is exactly the equity Tobin's Q. This is the simplest way to have a Tobin's Q different to one. In this particular case, it is also constant along the whole domain of equity returns.[3] In this context, the return on equity is given by $1 + r = \frac{Qd+p}{p_{-1}}$.

The first term of the utility function is the standard Constant Relative Risk Aversion (CRRA) formulation of consumption utility. The second term, proposed by Carroll (1998) and used by Reiter (2004) and Piketty (2011), says that investors derive utility from the ownership of wealth ($p_{-1}s$) and not merely from consumption. A similar specification of wealth in the utility function has been recently used in Kumhof et al. (2015).

We can simplify the problem of the household by using a change of variable. Let a' denote the value of assets acquired by the representative household at the current period. The problem becomes:

$$V(a) = \max_{a'} \frac{[(1+r)(a) - w - a']^{1-\mu}}{1-\mu} + \frac{\gamma a^{1-\theta}}{1-\theta} + \beta V(a'),$$

where $a' = ps'$ and $1 + r = \frac{Qd+p}{p_{-1}}$.

The intertemporal first order condition with respect to a' gives the Euler equation:

$$c^{-\mu} = \beta[(1 + r')c'^{(-\mu)} + \gamma a'^{(-\theta)}],$$

and its corresponding steady state formulation:

$$1 = \beta[(1 + r) + \frac{\gamma a^{-\theta}}{c^{-\mu}}]$$

Note that at the steady state, consumption equals the flow of total interests plus total wage, $c = ra + w$. Given this Euler equation, we can express the steady state demand of financial wealth like:[4]

$$a = \left[\frac{r^{-\mu} - \beta r^{-\mu} - \beta r^{1-\mu}}{\beta\gamma}\right]^{\frac{1}{-\theta+\mu}}$$

[3]There are other modelling strategies to achieve a Tobin's Q different than one, which rationalise other frictions. Gonzalez (2016) shows that monopoly rents within the problem of the firm or different stochastic discount factors between managers and shareholders can also give rise to an equity Tobin's Q different than one. The impact on equilibrium capital-output ratio, however, is similar in the sense that when Tobin's Q raises the capital-output ratio always declines.

[4]For simplicity we do not include w in the steady state equation. This simplification would be equivalent to a model where, in addition to the problem of the shareholder, there is a representative worker with perfect inelastic supply where wages are simply determined by the marginal productivity of labour.

Proposition 2.1. *The steady state demand of financial wealth is an increasing function of the return r for $0 < r < (\frac{1}{\beta} - 1)$ if $\mu \geq 0$ and $\theta > \mu$.*

Proof. The derivative of a with respect to r is:

$$\frac{\partial a}{\partial r} = \underbrace{\left[\frac{r^{-\mu} - \beta r^{-\mu} - \beta r^{1-\mu}}{\beta\gamma}\right]^{\frac{1}{-\theta+\mu}-1}}_{\text{Term A}} \underbrace{\left[\frac{-\mu r^{-\mu-1} + \beta\mu r^{-\mu-1} - \beta(1-\mu)r^{-\mu}}{\beta\gamma}\right]}_{\text{Term B}} \underbrace{\frac{1}{-\theta+\mu}}_{\text{Term C}}$$

For $\beta\gamma > 0$, term A is positive if $0 < r < (\frac{1}{\beta} - 1)$. Term B is negative for any value of r between 0 and $\frac{1}{\beta} - 1$ if $\mu \geq 0$. Accordingly, to have $\frac{\partial a}{\partial r} > 0$ along this range of returns, term C has to be negative and, therefore, $\theta > \mu$ must be satisfied. □

Summing up, when r is below $\frac{1}{\beta} - 1$, the steady state demand of financial assets is monotonically increasing with respect to capital return if $\mu \geq 0$ and $\theta > \mu$. Interestingly, an increasing steady state demand of financial assets shows that using wealth in the utility function within the representative agent framework can be interpreted, for a range of realistic parameter values, as a reduced form for precautionary savings. Indeed, in the standard incomplete markets model, which is often used to model precautionary behaviour, the aggregate demand of assets is also increasing and $r < \frac{1}{\beta} - 1$ is satisfied in equilibrium.[5] Although accumulating wealth for precautionary behaviour is a plausible interpretation for the shape of $a(r)$, other interpretations are also possible. For example, people might also derive direct utility from wealth accumulation due to the social status conferred by wealth (Piketty, 2011). Or people might accumulate wealth for dynastic altruism, that is, to leave a bequest to their descendants. Whatever the interpretation, by using wealth in the utility function, we get an increasing demand for financial assets, which, as shown below, is crucial for the results of the paper. This is in contrast with the standard model where only consumption is included in the utility function and that, according to Carroll (1998), is unable to explain households' saving behaviour. In that case, wealth disappears from the Euler equation and the demand of assets becomes perfect elastic at $\frac{1}{\beta} - 1$. Finally, note that according to our specification, marginal utility is decreasing both in consumption and in wealth, but the restriction $\theta > \mu$ means that it diminishes less rapidly in consumption.

[5]Although concavity is not required for the desired result, it turns out that it is also satisfied, as Figure 2.3 shows.

From now onwards, our demand of financial assets will be expressed as:

$$a = g(r); \quad \text{where} \quad g(r) = \Big[\frac{r^{-\mu}(1-\beta) - \beta r^{1-\mu}}{\beta\gamma}\Big]^{\frac{1}{-\theta+\mu}},$$

where:

$$\frac{\partial g(r)}{\partial r} > 0; \quad \forall r < \frac{1}{\beta} - 1 \quad \text{if} \quad \theta > \mu \quad \text{and} \quad \mu \geq 0.$$

2.2.2 The Firm

The representative firm accumulates physical capital K, pays the wage bill w and uses a CES technology to produce output Y:

$$Y = \Big[\phi K^{(\frac{\sigma-1}{\sigma})} + (1-\phi)L^{(\frac{\sigma-1}{\sigma})}\Big]^{\frac{\sigma}{\sigma-1}},$$

where ϕ is a distributional parameter and σ is the elasticity of substitution between labour and capital. The labour share of income LIS can be expressed in terms of the current period capital-output ratio ($\frac{K}{Y}$):

$$LIS = 1 - \phi\Big(\frac{K}{Y}\Big)^{\frac{\sigma-1}{\sigma}}, \tag{2.1}$$

where the sign of $\frac{\partial LIS}{\partial \frac{K}{Y}}$ depends on whether σ is higher or lower than one. In recursive formulation, the problem of the firm is:

$$\begin{aligned} V(K) &= \max \quad d + m'V(K') \\ s.t. &\qquad d = F(K, L) - (K' - (1-\delta)K) - w, \end{aligned}$$

where δ accounts for the depreciation rate of capital. The firm's discount factor is:

$$m' = \frac{\beta u_{c'}}{u_c - \beta v_{a'}} = \frac{1}{1+r'},$$

which makes the problem of the firm consistent with the problem of the households.

Given that, the firm solves:

$$V(K) = \max F(K, L) - (K' - (1-\delta)K)) - w + \frac{1}{1+r'}V(K'),$$

.

The FOC with respect to K' is

$$F_{K'}(K', L') = \delta + r', \tag{2.2}$$

from where we obtain a standard demand for capital $K'(r)$, which is decreasing in the level of capital returns. Using the transversality condition $\underset{t\to\infty}{K} = 0$, the constant-returns-to-scale assumption (which is satisfied under a CES technology) and abandoning the recursive formulation, we can express total capital stock as a function of future dividends.

$$K_{t+1} = \sum_{j=0}^{\infty} \frac{d_{t+1+j}}{\pi_{h=1}^{j-1}(1 + r_{t+h})},$$

.

From the definition of equity returns, $\frac{Qd'+p'}{p} = 1 + r'$, the stock price ($p$) can be expressed, using forward substitution, as a function of the future stream of received dividends:

$$p_t = \sum_{j=0}^{\infty} \frac{Qd_{t+1+j}}{\pi_{h=1}^{j-1}(1 + r_{t+h})}$$

Tobin's Q is the ratio of the stock market value to the replacement cost of capital. Using the expressions above, Tobin's Q results in:

$$Q_t = \frac{p_t}{K_{t+1}}, \quad \forall t$$

Note that under this specification, Tobin's Q is constant along the whole domain of equity returns. Since the demand of capital has been obtained from the problem of the firm, the value of financial assets is:

$$p(r) = QK'(r)$$

2.2.3 Equilibrium

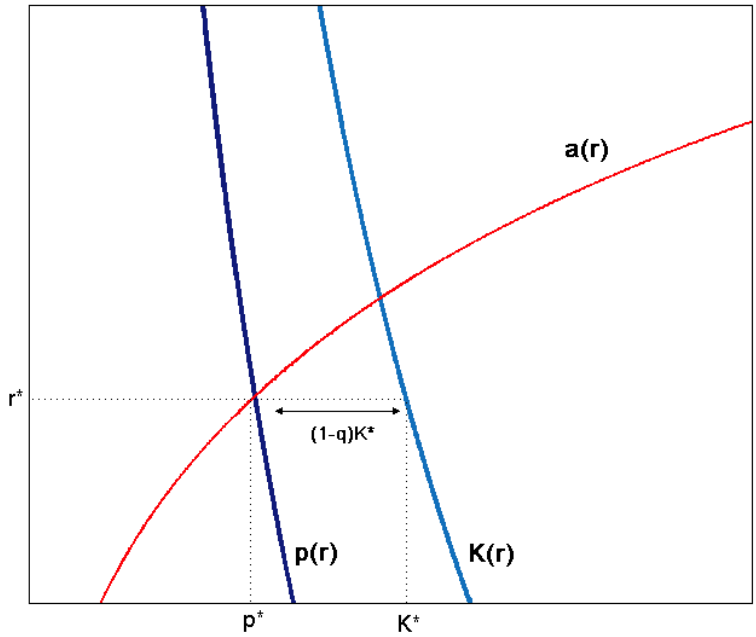
Equilibrium in the capital market would occur at the intersection between $p(r)$ and $g(r)$. Since $g(r)$ is monotonically increasing and $p(r)$ is monotonically decreasing, there is a unique equilibrium characterised by the return to equity (r^*):

$$p(r^*) = g(r^*)$$

Note that since the capital demand is monotonically decreasing with respect to the return, there is a single level of capital corresponding to r^*. We denote this level $K^*(r^*)$.

Importantly, the equilibrium $p(r^*) = g(r^*)$ depends on Q. If Q is larger, then the equilibrium equity return would be higher because investors would demand a higher return to hold the additional financial wealth. Therefore, the equilibrium level of r depends positively on Q.

FIGURE 2.3: Market for capital



When Q changes, there is a change in equilibrium r^*, but also a change in the amount of physical capital demanded by the firm. Figure 2.3 is illustrative at this point. When Q grows, the gap between $p(r)$ and $K(r)$ becomes smaller and r^* raises because the equilibrium is moving upwards along $a(r)$. In response to it, the firm will tend to decrease the amount of physical capital to raise the return on equity, which from equation (2.2) is directly linked to capital productivity. Therefore, we can express the equilibrium level of capital in terms of r^*, and subsequently, in terms of Q:

$$K^*(r^*(Q)) = K^*(Q)$$

Proposition 2.2. *The relation between Tobin's Q and equilibrium capital K^* is negative.*

Proof. Since $p(r^*) = QK(r^*)$, the equilibrium condition $p(r^*) = g(r^*)$ can be expressed as a function $G(K^*, Q)$ where $G(K^*, Q) = \frac{g(r^*(K^*))}{K(r^*)} - Q$. Applying the implicit function theorem, we have that $\frac{dK^*}{dQ} = -\frac{\frac{\partial G}{\partial Q}}{\frac{\partial G}{\partial K^*}} = \frac{-(-1)}{\frac{\partial g}{\partial r}\frac{\partial r}{\partial K}K^{-1} - g(r)K^{-2}}$. Since $\frac{\partial g}{\partial r}$ is positive and $\frac{\partial r}{\partial K}$ is negative, $\frac{dK^*}{dQ}$ has to be negative. □

The equilibrium of the model makes explicit the relation between the capital level of equilibrium and the Tobin's Q. Since the labour share depends on the capital-output ratio (see equation (2.1)), we can make explicit the relation between the labour share and Tobin's Q:

$$LIS^* = 1 - \phi\left[\frac{K^*(Q)}{Y^*(K^*)}\right]^{\frac{\sigma-1}{\sigma}}, \quad \text{where} \quad \frac{\partial LIS^*}{\partial Q} = \frac{\partial LIS^*}{\partial \frac{K^*}{Y^*}} \frac{\partial \frac{K^*}{Y^*}}{\partial K^*} \frac{dK^*}{dQ},$$

therefore:

$$\begin{aligned} \frac{\partial LIS^*}{\partial \frac{K^*}{Y^*}} &> 0 \quad \text{if} \quad \sigma < 1; \\ \frac{\partial \frac{K^*}{Y^*}}{\partial K^*} &> 0 \quad \text{due to CRS;} \\ \text{and} \quad \frac{dK^*}{dQ} &< 0 \quad \text{given by proposition 2.2.} \end{aligned}$$

Importantly, the mechanism proposed here works through the capital-output ratio, that is, Tobin's Q impacts on the labour share through its effect on investment and capital. In that sense, our model is in the spirit of Piketty and Zucman (2014) and Karabarbounis and Neiman (2014). In particular, Karabarbounis and Neiman (2014) build a model to explain the decline of the labour income share with recent movements in the relative price of investment goods. Their mechanism can be easily embedded into our model by adding the relative prices of capital goods (ξ_t) in the budget constraint of the firm.

$$F(K, L) = d + \xi[K' - (1 - \delta)K] + w,$$

where the demand of capital is dependent on ξ, and more specifically, it raises when the relative price of capital goods falls, that is, $\frac{\partial K'(r)}{\partial \xi} < 0$. We empirically evaluate the potential impact of this mechanism compared to ours.

2.3 Data

In order to empirically study the relationship between the Tobin's Q and the labour income share, this paper combines three different databases.

2.3.1 Tobin's Q

Tobin's Q is the market value of capital over its replacement cost. We use data from Worldscope Database to calculate the Tobin's Q as the market value of the sum of equity and non-equity liabilities over the sum of their book value, which is generally acknowledged as the most accurate procedure.

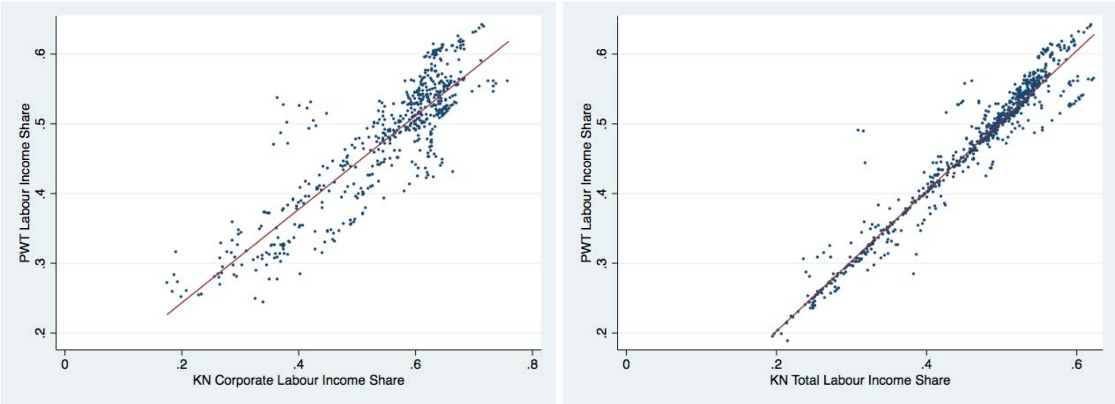
We aggregate firm level data from publicly traded companies following Doidge et al. (2013) methodology. That is, in a first stage firms are clustered in 17 different sectors using the Fama-French 17 industries classification, where a median Q is computed for each industry. Countries' Q are calculated as the market value weighted average of the median industries' Q. The use of industry medians let us overcome the problem of potential outliers in the sample.

2.3.2 Labour income share

Regarding the LIS, Karabarbounis and Neiman (2014) have developed a database of the corporate labour income share for a considerable number of countries obtaining the data from several sources. However, the use of their database would force us to exclude a non-negligible number of countries in our analysis. As an alternative, we lean to use the LIS variable from the Extended Penn World Table 4.0 (EPWT 4.0).

The EPWT 4.0 draws information from different United Nations sources and measures the labour income share as the share of total employee compensation in the Gross Domestic Product with no adjustment for mixed rents, and without distinguishing the corporate sector. Although we are aware of the potential drawbacks from using this LIS definition, the correlation with the corporate labour share and the total labour share used by Karabarbounis and Neiman (2014) is 0.87 and 0.96 respectively (Figure 2.4). We consider this a safe level in order to use our variable.

FIGURE 2.4: EPWT LIS vs KN LIS



(a) EPWT vs Corporate Labour Share

(b) EPWT vs Total Labour Share

2.3.3 Relative prices

The relative price of investment goods with respect to consumption goods is obtained by extending Karabarbounis and Neiman (2014) database.

In order to obtain the relative price in domestic terms, we divide the country-specific relative price obtained from the Penn World Table 7.1 ($\frac{Pi_i}{Pc_i}$), which is calculated using ppp exchange rates, over the relative price of investment in the United States ($\frac{Pi_{US}}{Pc_{US}}$). We then multiply this ratio by the ratio of the investment price deflator to the personal consumption expenditure deflator for the United States ($\frac{ID_{US}}{PCD_{US}}$) obtained from BEA.

$$RP = \frac{\frac{Pi_i}{Pc_i}}{\frac{Pi_{US}}{Pc_{US}}} * \frac{ID_{US}}{PCD_{US}}$$

2.4 Empirical methodology

Beyond the theoretical relationships, we face the challenge of carrying out a robust estimation of the relationship between Tobin's Q and the labour share. This section explains in detail (i) how we go from the theoretical model to an empirical equation, and (ii) the empirical tools which allow us to infer a causal relationship.

2.4.1 Empirical implementation

For empirical purposes, we do not impose a specific production function and, therefore, we do not restrict the functional form of the labour share to be the one derived from a CES technology. We simply assume a general multiplicative form where changes in the capital-output ratio have an impact on the labour share:

$$LIS = g\Big(\frac{K}{Y}\Big) = a\Big(\frac{K}{Y}\Big)^{\alpha}$$

This way, our empirical specification is close to Bentolila and Saint-Paul (2003). Note that we remain agnostic about α and then we do not know ex-ante whether the impact of $\frac{K}{Y}$ on the labour share is positive or negative.

Nevertheless, contrary to Bentolila and Saint-Paul (2003), we further endogenise the capital-output ratio. Our model shows that the equilibrium capital-output ratio depends, among other things, on the Tobin's Q, and that the sign of this relation is negative. However, and again for empirical purposes, we do not impose a particular relation derived from the specifics of the model. Rather, we also assume a general multiplicative form where the capital-output ratio is expressed as a function of Tobin's Q. Following Karabarbounis and Neiman (2014), we also include the relative price of investment goods (RP) as an argument of $\frac{K}{Y}$.

$$\frac{K}{Y} = f(Q, RP) = Q^{\psi_1} RP^{\psi_2}$$

We use these two forms to obtain an estimable equation of the labour share in terms of Q and RP:

$$LIS = g\Big(\frac{K}{Y}\Big) = g(f(Q, RP)) = a(Q^{\psi_1} RP^{\psi_2})^{\alpha}$$

Taking natural logarithms:

$$\log(LIS) = \log(a) + \alpha\psi_1 \log(Q) + \alpha\psi_2 \log(RP) + \Omega_{it},$$

or simplifying:

$$lis_{it} = \beta_0 + \beta_1 q_{it} + \beta_2 rp_{it} + \Omega_{it} \tag{2.3}$$

Where lis, q, and rp are the natural logarithm values of our variables of interest, and Ω is a standard error term.

2.4.2 Econometric methodology

While characterised by a small number of cross-sectional units (N) compared to the time dimension (T), macroeconomics panel data have been traditionally estimated following microeconomics panel data techniques under the assumptions of parameter homogeneity (across countries), common impact of unobservable factors, cross-section independence, and data stationarity.[6] However, if these assumptions are violated, results would be subject to misspecification problems.

In order to overcome these potential sources of misspecification, we rely on relative recently developed panel data techniques (panel time-series), which are especially developed for macroeconomics data characteristics (Pesaran, 2015).[7]

Our empirical framework is based on a common factor model (for details, see Eberhardt and Teal, 2011, 2013a,b).[8] Formally, assuming for simplicity a one-input model, a common factor model is as follows:

$$y_{it} = \beta_i x_{it} + u_{it}, \qquad u_{it} = \varphi_i f_t + \psi_i + \varepsilon_{it}, \tag{2.4}$$

$$x_{it} = \delta_i f_t + \gamma_i g_t + \pi_i + e_{it}, \tag{2.5}$$

$$f_t = \tau + \phi f_{t-1} + \omega_t, \qquad g_t = \mu + \kappa g_{t-1} + \nu_t, \tag{2.6}$$

where y_{it} and x_{it} represent, respectively, the dependent and independent variables, and u_{it}, apart from the error term (ε_{it}), contains the unobservable factors. In particular, unobservable time-invariant heterogeneity is captured through a country fixed effect (ψ_i), while time-variant heterogeneity is accounted through a common factor (f_t) with country-specific factor loadings (φ_i). At the same time, the model allows for the regressor to be affected by these, or other common factors (f_t and g_t). These factors represent both unobservable global shocks that affect all the countries, although with different intensities (i.e., oil prices, financial crisis...), and local spillovers (Chudik et al., 2011; Eberhardt et al., 2013). The presence of the same unobservable process (f_t) as a determinant of both input and output raise endogeneity problems which make difficult the estimation of β_i (Kapetanios et al.,

[6] See Roodman (2009) for a detailed explanation on the potential risks of the popular Difference and System GMM estimators.

[7] Although empirical applications of these methods are still not widespread in the literature, it is worthy to acknowledge the valuable contribution made to the field by Markus Eberhardt and coauthors in the last years. The empirical methodology of this Chapter relies on several of their papers.

[8] This Subsection is partially borrowed from Subsection 1.3.2 in Chapter 1.

2011).[9] β subscript i indicates a country-specific impact of the regressor on the dependent variable.

We can see the previous common factor model as a general empirical framework which encompasses several simpler structures. In particular, a first classification can be done between "Homogeneous models", where the impact of the regressors on the dependent variable is common across countries (i.e., $\beta_i = \beta$), and "Heterogeneous models", which leave this unconstrained (i.e., β_i is estimated for each country). Within each group, the assumptions about the structure of the unobservable factors will lead to different estimation methods.

As in Chapter 1, within the homogeneous estimators, we consider the common Pooled Ordinary Least Square (POLS), the Two-way Fixed Effects (2FE), and the Pooled Common Correlated Effects (CCEP) estimators. While the first two are standard in the literature and account for unobservable heterogeneity through time and country dummies, the CCEP estimator allows for a different impact of the unobservables across countries and time. Empirically, it aims to eliminate the cross-sectional dependence by augmenting equation (2.3) with the cross-section averages of the variables.[10]

Regarding heterogeneous models, we consider different Mean Group estimators. In particular, we present the results for the Pesaran and Smith (1995) Mean Group estimator (MG), the Pesaran (2006) Common Correlated Effects Mean Group estimator (CMG), and the Chudik and Pesaran (2015) Dynamic CMG estimator (CMG2). The common characteristic of these estimators is that they are defined as the simple average of the country-specific estimators (i.e., $\beta^* = N^{-1} \sum_{i=1}^{N} \beta_i$).

Pesaran and Smith (1995) Mean Group estimator (MG) allows for a country-specific impact of both the regressor and the unobservable heterogeneity. The impact of the last one is assumed to be constant, and is empirically accounted by adding country-specific linear trends (t). Therefore, the estimable equation takes the form:

$$lis_{it} = \beta_0^{MG} + \beta_1^{MG} q_{it} + \beta_2^{MG} rp_{it} + \beta_3^{MG} t + \Omega_{it}$$

As explained before, the MG estimator is computed as the simple average of the different country-specific coefficients, which are calculated by regressing the

[9] Equation (2.6) models these factors as a simple AR(1), where no constrains are imposed to get stationary processes. Note that nonstationarity could provoke a spurious relationship between our variables of interest. If our variables are nonstationary, we have to analyse the cointegration relationship among them to infer any causal relationship.

[10] Eberhardt et al. (2013) provide the intuition behind this mechanism.

previous equation for each country. However, although it overcomes the potential misspecification from assuming parameter homogeneity, the introduction of country-specific linear trends could be too simple to rule out all the possible cross-section dependence from the unobserved heterogeneity.

In this sense, Pesaran (2006) proposes the Common Correlated Effects Mean Group estimator (CMG), which is a combination of the MG and the CCEP estimators. In particular, it approximates the unobserved factors by adding the cross-section averages of the dependent and explanatory variables, and then running standard panel regressions augmented with these cross-section averages. Empirically, we estimate:

$$\begin{aligned} lis_{it} &= \beta_0^{CMG} + \beta_1^{CMG} q_{it} + \beta_2^{CMG} rp_{it} \\ &+ \beta_3^{CMG} \overline{lis_t} + \beta_4^{CMG} \overline{q_t} + \beta_5^{CMG} \overline{rp_t} + \Omega_{it} \end{aligned}$$

It is easy to see that the first line is the Pesaran and Smith (1995) MG estimator (without linear trend), and the second line is the way the Pesaran (2006) CMG estimator approximates the unobservable processes.

So far, we have discussed how to deal with different sources of misspecification like the assumption of parameter homogeneity or the existence of cross-section dependence. This paper also analyses potential misspecification arising from a possible dynamic structure of the relation under study by estimating both static and dynamic specifications. However, although Pesaran (2006) CMG estimator yields consistent estimates under a variety of situations (see Kapetanios et al., 2011; Chudik et al., 2011), it does not cover the case of dynamic panels or weakly exogenous regressors. Chudik and Pesaran (2015) propose an extension of the CMG approach (CMG2) to account for the potential problems arising from dynamic panels. In particular, they prove that the inclusion of extra lags of the cross-section averages in the CMG approach gives a consistent estimator of both β_i and β^{CMG}.

Empirically, we present an Error Correction Model (ECM) which has the following form:

$$\begin{aligned} \Delta lis_{it} &= \beta_0^{CMG2} + \beta_1^{CMG2} lis_{i,t-1} + \beta_2^{CMG2} q_{i,t-1} + \beta_3^{CMG2} rp_{i,t-1} + \beta_4^{CMG2} \Delta q_{it} + \beta_5^{CMG2} \Delta rp_{it} \\ &+ \beta_6^{CMG2} \overline{\Delta lis_t} + \beta_7^{CMG2} \overline{lis_{t-1}} + \beta_8^{CMG2} \overline{q_{t-1}} + \beta_9^{CMG2} \overline{rp_{t-1}} + \beta_{10}^{CMG2} \overline{\Delta q_t} + \beta_{11}^{CMG2} \overline{\Delta rp_t} \\ &+ \sum_{l=1}^{p} \beta_{12}^{CMG2} \overline{\Delta lis_{t-p}} + \sum_{l=1}^{p} \beta_{13}^{CMG2} \overline{\Delta q_{t-p}} + \sum_{l=1}^{p} \beta_{14}^{CMG2} \overline{\Delta rp_{t-p}} + \Omega_{it}, \end{aligned}$$

where the first line represents the Pesaran and Smith (1995) MG estimator, the inclusion of the second gives the Pesaran (2006) CMG estimator, and the three lines together are the Chudik and Pesaran (2015) Dynamic CMG estimator (CMG2).[11]

Likewise, given the way they control for unobservables, CMG style estimators are suitable for accounting for structural breaks and business cycle distortions, thus making the use of yearly data perfectly valid in order to infer long run relationships.

2.5 Results

In order to give a systematic view of our results, this section is divided in four subsections. Subsection 2.5.1 presents an exhaustive analysis of the time-series properties of our variables of interest. Subsection 2.5.2 shows the results for a baseline model, where just the Tobin's Q is considered as a regressor. Subsection 2.5.3 further includes the relative price of investment in the analysis, and Subsection 2.5.4 provides evidence supporting the interpretation of our results as a causal relationship.

2.5.1 Time-series properties

The analysis of time-series properties is a central issue in panel time-series econometrics. The order of integration of the variables and the cross-section dependence of the data plays a central role in empirical macroeconomics (Pesaran, 2015). In order to understand the potential problems we could face, Tables 2.1 and 2.2 analyse, respectively, the order of integration and the cross-section dependence of the variables used in our analysis.

Regarding the order of integration, Table (2.1) presents the results for different specifications of the Maddala and Wu (1999) test and the cross-sectional augmented panel unit root (CIPS) Pesaran (2007) test.[12] While Maddala and Wu

[11] Chudik and Pesaran (2015) recommend to set the number of lags equal to $T^{1/3}$. We consider up to 2 extra lags of the cross-section averages.

[12] Panels 2.1.a) and 2.1.c) show the results when a constant is included in the ADF regressions, while 2.1.b) and 2.1.d) include also a deterministic trend.

(1999) test is an exact nonparametric test based on Fisher (1928), which allows for heterogeneity in the autoregressive coefficient of the Augmented Dickey-Fuller (ADF) regression, it ignores the possibility of cross-section dependence (1^{st} generation-type test).[13]

TABLE 2.1: Unit root tests

a) Maddala and Wu (1999): Constant

Lags	*lis*	(*p*)	*q*	(*p*)	*rp*	(*p*)
0	74.090	0.541	147.660	0.000	93.702	0.082
1	105.420	0.014	179.149	0.000	91.131	0.114
2	165.169	0.000	114.449	0.003	82.168	0.294
3	144.513	0.000	114.109	0.003	77.614	0.427
4	95.221	0.067	105.450	0.014	76.572	0.460

b) Maddala and Wu (1999): Constant and deterministic trend

Lags	*lis*	(*p*)	*q*	(*p*)	*rp*	(*p*)
0	55.055	0.966	117.014	0.002	51.116	0.987
1	80.704	0.334	157.053	0.000	84.221	0.243
2	73.616	0.556	107.204	0.011	94.663	0.072
3	91.226	0.112	105.641	0.014	87.151	0.180
4	113.875	0.003	148.840	0.000	73.766	0.551

c) Pesaran (2007) CIPS test: Constant

Lags	*lis*	(*p*)	*q*	(*p*)	*rp*	(*p*)
0	0.431	0.667	-2.744	0.003	-0.118	0.453
1	-0.207	0.418	-2.405	0.008	-0.141	0.444
2	-1.199	0.115	0.103	0.541	0.655	0.744
3	1.802	0.964	2.942	0.998	2.254	0.988
4	5.477	1.000	6.091	1.000	7.211	1.000

d) Pesaran (2007) CIPS test: Constant and deterministic trend

Lags	*lis*	(*p*)	*q*	(*p*)	*rp*	(*p*)
0	1.044	0.852	-2.068	0.019	2.483	0.993
1	0.390	0.652	-1.628	0.052	2.052	0.980
2	-0.033	0.487	1.304	0.904	0.998	0.841
3	5.280	1.000	6.785	1.000	6.006	1.000
4	8.090	1.000	8.949	1.000	9.127	1.000

Notes: Maddala and Wu (1999) test and Pesaran (2007) CIPS test values are obtained respectively from the Fisher statistic and the standardised Z-tbar statistic. H_0 = nonstationarity for both tests. Lags indicates the number of lags included in the ADF regression.

[13] In more technical terms, Maddala and Wu (1999) computes: $\zeta = -2 \sum_{i=1}^{N} p_i \sim \chi(2N)$, where p_i is the probability of the value of the ADF unit root test for the ith unit (country).

Pesaran (2007) CIPS test, besides the heterogeneity in the autorregresive coefficient of the ADF regression, it also permits for potential cross-section dependence of the variables. Similar to Im et al. (2003), Pesaran (2007) CIPS test proposes a standardised average of individual ADF coefficient, where the ADF processes have been augmented by the cross-sectional averages to control for the unobservable component.

Given the results, we can safely assert that the variables under analysis are nonstationarity.

TABLE 2.2: Cross-section dependence tests

a) Levels:				b) Diff:			
Variable	lis	q	rp	Variable	Δlis	Δq	Δrp
CD-test	16.73	29.76	42.37	CD-test	12.99	34.45	6.66
p-value	0.00	0.00	0.00	p-value	0.00	0.00	0.00
corr	0.132	0.250	0.345	corr	0.11	0.296	0.049
abs(corr)	0.472	0.394	0.558	abs(corr)	0.235	0.349	0.223
c) Het. AR(2)				**d) Het. AR(2) CCE**			
Variable	lis	q	rp	Variable	lis	q	rp
CD-test	9.93	33.58	3.40	CD-test	-0.24	-0.66	-2.38
p-value	0.00	0.00	0.00	p-value	0.81	0.51	0.02
corr	0.088	0.301	0.027	corr	-0.006	-0.011	-0.023
abs(corr)	0.243	0.344	0.213	abs(corr)	0.220	0.237	0.213

Notes: CD-test shows the Pesaran (2004) cross-section dependence statistic, which follows a $N(0,1)$ distribution. H_0 = cross-section independence. corr, and abs(corr) report, respectively, the average and average absolute correlation coefficients across the $N(N-1)$ set of correlations.

Table 2.2 shows the Pesaran (2004) CD test for cross-section dependence in panel time-series data. This test uses correlation coefficients between the time-series for each panel member and has proved to be robust to nonstationarity, parameter heterogeneity and structural breaks even in small samples.[14] Table 2.2 is divided in four different panels representing different transformations of the variables. Panels a) and b) show the CD test for the levels and growth rates of our variables. The null hypothesis of cross-section independence is rejected in all the cases. Panels

[14]The test is computed as:

$$CD = \sqrt{\frac{2}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \sqrt{T_{ij}} \rho_{ij},$$

where ρ_{ij} represents the correlation coefficient between country i and j, while T_{ij} is the number of observations used to computed that correlation.

c) and d) present the results when the test is applied to the residuals of an autoregressive regression of order 2 (AR(2)) of each variable. The difference is that, while regressions in panel c) are estimated by the Pesaran and Smith (1995) Mean Group estimator, Panel d) shows the results when the AR process is augmented with cross-section averages in the spirit of Pesaran's (2006) CMG estimator. The difference between both shows the power of the cross-section averages to control for unobservable cross-section dependence.

The presence of nonstationary variables and cross-section dependence in our data make the use of traditional panel data techniques invalid. To be sure that our regression results are not subject to biases due to cross-section dependence or to spurious relationships due to the order of integration, our empirical analysis pays specially attention to regression residuals' characteristics. More specifically, our preferred model will prove that the residuals are stationary (as an informal test for cointegration among the variables) and that they do not have problems of cross-section dependence (as a proof that unobservable heterogeneities are successfully captured in the model).

2.5.2 Baseline results

Tables 2.3 and 2.4, present the results for our baseline model, where just the impact of Tobin's Q on the labour income share is studied. Columns [1] - [4] display the homogeneous-type estimators, where β is constrained to be the same across countries. More specifically, we present the results for the standard OLS with time-dummies (POLS), the Two-way Fixed Effects (2FE), and the Common Correlated Effects pooled estimator (CCEP), with and without including a country-specific linear trend. Each estimator fits in the model described by equations (2.4)-(2.6) by imposing different assumptions on the structure of the unobservable process.[15] The rest of the columns present the heterogeneous-type estimators. In particular, we show the estimates for the Pesaran and Smith (1995) Mean Group estimator (MG), the Cross-Section Demeaned Mean Group estimator (CDMG), and the Pesaran (2006) Common Correlated Effects Mean Group estimator (CMG) without and

[15]POLS and 2FE account for the unobservable heterogeneity in a standard way, that is, the inclusion of time-dummies and cross-section fixed effects control respectively, for the time-varying and invariant heterogeneity. The assumption regarding this process is that time-varying heterogeneity has the same impact across countries for a given year. The CCEP estimator allows for a more flexible structure, where the cross-sectional averages of the variables are included in the regression with a different impact across countries.

with country-specific trends.[16] The common characteristic of these estimators is that all of them allow for a country-specific impact (β_i) of our variable of interest. A country-specific regression is estimated, and the result presented in the table is the average of the country-specific coefficients.[17] Similar to the homogeneous-type estimators, differences within this group depend on the assumptions underlying the unobservable process. While the MG estimator includes a country-specific linear trend to control for the unobservable heterogeneity, the CMG estimator, like the CCEP, includes the cross-sectional averages of the variables, allowing for a much more flexible functional form of the unobservable process.

Table 2.3 presents the estimations corresponding to a static model, where we consider 41 countries and 915 observations.[18] Regarding the homogeneous-type estimators, we find a negative and significant impact of the Tobin's Q on the labour income share in all but the POLS estimator (where the impact is positive and significant). However, before analysing the magnitude of the impact, the results for Pesaran (2007) CIPS and Pesaran (2004) CD tests indicate that the residuals present nonstationarity and cross-section depedence. This warning alerts that our regressions are suffering from some type of misspecification, which from our discussion before could be: (i) the imposition of parameter homogeneity, (ii) an unsuitable structure of the unobservable heterogeneity, or (iii) that the nature of the relationship is not static. The importance of the first two potential sources of misspecification can be tested analysing the Mean Group-style estimators (columns [5] - [8]). A negative and significant impact of the Tobin's Q on the labour income share is still present, ranging from -0.053 to -0.06, where just the result for the CDMG estimator is not significant. However, although the residuals present an improvement in terms of absolute correlation, we still observe cross-section dependence. Stationarity in the residuals is now present in 3 out of the 4 regressions. These results indicate that, although the introduction of parameter heterogeneity improves the specification, it is not enough to solve all the potential misspecification problems.

Table 2.4 analyses the third potential source of misspecification through the estimation of a Partial Adjustment Model (PAM), where the first lag of the dependent variable is included as a regressor. Due to data restrictions, we consider 40

[16]CDMG correspond to Pesaran and Smith (1995) MG estimator when the variables have been transformed as a deviation from the country-average.

[17]Pesaran and Smith (1995) show that the Mean Group-style estimators will produce consistent estimates of the average of the parameters.

[18]Table 2.A1 in the appendix shows the specific countries and period under analysis.

countries with the number of observations ranging from 850 to 885. The first important result is that a clear negative and significant long-run relationship is again observed between the Tobin's Q and the labour share irrespective of the estimator under analysis. The second remarkable fact is that most of the residuals show cross-sectional independence and stationarity, indicating the absence of the previous source of misspecification. Given its flexibility to control for the unobserved factors, our preferred model is the one showed in the last column (CMGt2)) which represents the Chudik and Pesaran (2015) Dynamic CMG estimator, where 2 extra lags of the cross-section averages are included in the regression to control for the potential dynamic bias. We observe that a 1% increase in Tobin's Q decreases the labour income share in the long-run by 0.08%.

2.5.3 Consideration of relative prices

As explained before, Karabarbounis and Neiman (2014) have argued that the global decline in the labour share can be explained, at least in part, by the decrease in the relative price of investment goods. In this section we compare the relevance of their mechanism in our empirical analysis. Tables 2.5 and 2.6 extend our regressions by including the relative price of investment in the estimation.

Results from a static model are presented in Table 2.5, where the inclusion of the relative prices does not alter the negative relationship found between the Tobin's Q and the labour share. Regarding the relative prices, they just present a negative impact under the homogeneous-type estimators. However, once we allow for parameter heterogeneity, they no longer show any kind of influence on the labour income share. Nevertheless, similar to the static model analysed in Table 2.3, residuals present cross-section dependence and nonstationarity.

In order to assess problems arising from the dynamic structure of our equation, this time we estimate an Error Correction Model (Table 2.6), where due to data restrictions we are not able to include more than 30 countries. Although we present the results for different estimators, given its larger flexibility, we focus on the ones obtained from the CMG-style estimators (columns [4]-[7]). The first remarkable fact is the presence of stationarity and cross-section independence in the residuals, indicating the absence of the previous misspecification problems. Regarding the impact of our variables of interest, we observe a negative impact of the Tobin's Q in both the short and long run. If we focus on the long run relationship, our estimations show that an increase of 1% in Tobin's Q would decrease the labour

Table 2.3: Static baseline model

	[1] POLS	[2] 2FE	[3] CCEP	[4] CCEPt	[5] MG	[6] CDMG	[7] CMG	[8] CMGt
q	0.14	-0.083	-0.05	-0.052	-0.057	-0.035	-0.053	-0.06
	(0.052)***	(0.025)***	(0.017)***	(0.016)***	(0.015)***	(0.046)	(0.020)***	(0.016)***
t					-0.003			-0.003
					(0.001)**			(0.001)**
Constant	-0.647	-0.665			-0.656	0.012	-0.483	-0.714
	(0.036)***	(0.017)***			(0.032)***	(0.036)	(0.068)***	(0.105)***
Number Id	41	41	41	41	41	41	41	41
Observations	915	915	915	915	915	915	915	915
R-squared	0.11	0.93	0.99	0.99				
RMSE	0.2244	0.0629	0.0500	0.0474	0.0443	0.0724	0.0435	0.0336
Trends					0.73			0.59
CD test	28.3495	-2.6979	8.688	-2.9706	3.8019	28.6299	9.5781	5.4416
Abs Corr	0.4730	0.4211	0.3710	0.3660	0.3052	0.4158	0.3243	0.2658
Int	I(1)	I(1)	I(0)/I(1)	I(1)	I(1)	I(0)	I(1)/I(0)	I(0)

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%.
POLS = Pooled OLS (with year dummies), 2FE = 2-way Fixed Effects, CCEP = Pooled Pesaran (2006) Common Correlated Effects (CCE), CCEPt = CCEP with linear trend, MG = Pesaran and Smith (1995) Mean Group (with country-specific linear trends), CDMG = Cross-Section Demeaned Mean Group, CMG = Pesaran (2006) CCE Mean Group, CMGt = CMG with country-specific linear trends.
CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. RMSE presents the root mean squared error. Trend show the share of countries where the linear trend is significant at 5%.

TABLE 2.4: Dynamic baseline model

	[1] POLS	[2] 2FE	[3] CCEP	[4] CCEPt	[5] MG	[6] CMG	[7] CMGt	[8] CMGt1	[9] CMGt2
q	-0.009	-0.03	-0.021	-0.023	-0.034	-0.017	-0.024	-0.026	-0.04
	(0.015)	(0.019)	(0.012)*	(0.012)*	(0.010)***	(0.011)	(0.012)**	(0.013)**	(0.010)***
lis_{t-1}	0.977	0.771	0.749	0.718	0.64	0.767	0.608	0.537	0.502
	(0.008)***	(0.078)***	(0.037)***	(0.041)***	(0.034)***	(0.026)***	(0.032)***	(0.037)***	(0.048)***
t					-0.001		-0.001	-0.001	-0.001
					(0.000)		(0.001)*	(0.001)	(0.001)*
Constant	-0.005	-0.141			-0.232	-0.135	-0.358	-0.418	-0.318
	(0.009)	(0.051)***			(0.025)***	(0.026)***	(0.076)***	(0.084)***	(0.078)***
Number of id	40	40	40	40	40	40	40	40	40
Observations	885	885	885	885	885	885	885	868	850
R-squared	0.97	0.98	0.99	0.99					
RMSE	0.0403	0.0384	0.0307	0.0309	0.0338	0.0246	0.0225	0.0202	0.0178
Trend					0.25		0.24	0.33	0.23
lr-q	-0.4112	-0.1288	-0.0854	-0.0815	-0.0944	-0.0725	-0.061	-0.0564	-0.08
se-q	0.7248	0.0613	0.0466	0.0428	0.0296	0.0462	0.0306	0.0281	0.0208
CD test	-0.2311	-1.2989	-0.9679	-2.5497	8.5344	-1.3683	-1.1636	-0.4342	-0.5145
Abs Corr	0.2133	0.2253	0.2309	0.2339	0.2240	0.2171	0.2188	0.2331	0.2426
Int	Int(0)	Int(0)	Int(0)	Int(0)	Int(0)	Int(0)	Int(0)	Int(0)	Int(0)

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%.
POLS = Pooled OLS (with year dummies), 2FE = 2-way Fixed Effects, CCEP = Pooled Pesaran (2006) Common Correlated Effects (CCE), CCEPt = CCEP with linear trend, MG = Pesaran and Smith (1995) Mean Group (with country trends), CMG = Pesaran (2006) CCE Mean Group, CMGt = CMG with country-specific linear trends, CMGt1 and CMGt2 = CMGt with, respectively, one and two extra cross-sectional averages lags, as indicated by Chudik and Pesaran (2015).
CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. RMSE presents the root mean squared error. Trend show the share of countries where the linear trend is significant at 5%. lr-q and se-q represent respectively q's long-run impact and its standard error.

TABLE 2.5: Static model with relative prices

	[1] POLS	[2] 2FE	[3] CCEP	[4] CCEPt	[5] MG	[6] CDMG	[7] CMG	[8] CMGt
q	0.157	-0.08	-0.052	-0.052	-0.067	-0.049	-0.061	-0.052
	(0.051)***	(0.025)***	(0.015)***	(0.015)***	(0.013)***	(0.031)	(0.019)***	(0.015)***
rp	-0.344	-0.113	-0.1	-0.101	0.005	-0.147	-0.001	0.017
	(0.100)***	(0.043)***	(0.048)**	(0.047)**	(0.085)	(0.147)	(0.111)	(0.078)
t					-0.002			-0.002
					(0.001)*			(0.002)
Constant	-0.589	-0.642			-0.678	0.009	-0.681	-0.664
	(0.039)***	(0.019)***			(0.034)***	(0.039)	(0.101)***	(0.078)***
Number of id	41	41	41	41	41	41	41	41
Observations	915	915	915	915	915	915	915	915
R-squared	0.12	0.93	0.99	0.99				
RMSE	0.2229	0.0625	0.0411	0.0399	0.0405	0.0629	0.0311	0.0273
Trend					0.56			0.32
CD test	25.6361	-2.4791	5.4335	-2.3717	3.7041	20.7408	2.4645	4.6826
Abs Corr	0.4506	0.4142	0.3057	0.3102	0.2821	0.3541	0.2522	0.2517
Int	I(1)	I(1)	I(0)/I(1)	I(0)/I(1)	I(0)/I(1)	I(0)/I(1)	I(0)	I(0)

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%.
POLS = Pooled OLS (with year dummies), 2FE = 2-way Fixed Effects, CCEP = Pooled Pesaran (2006) Common Correlated Effects (CCE), CCEPt = CCEP with linear trend, MG = Pesaran and Smith (1995) Mean Group (with country-specific linear trends), CDMG = Cross-Section Demeaned Mean Group, CMG = Pesaran (2006) CCE Mean Group, CMGt = CMG with country-specific linear trends.
CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. RMSE presents the root mean squared error. Trend show the share of countries where the linear trend is significant at 5%.

Table 2.6: ECM with relative prices

	[1] 2FE	[2] CCEP	[3] MG	[4] CMG	[5] CMGt	[6] CMGt1	[7] CMGt2
lis_{t-1}	-0.176	-0.395	-0.449	-0.5	-0.694	-0.72	-0.812
	(0.026)***	(0.049)***	(0.034)***	(0.053)***	(0.061)***	(0.085)***	(0.125)***
q_{t-1}	0.011	-0.012	-0.035	-0.039	-0.067	-0.076	-0.058
	(0.013)	(0.015)	(0.014)**	(0.018)**	(0.026)**	(0.028)***	(0.033)*
rp_{t-1}	-0.032	-0.016	0.064	0.15	0.092	0.129	-0.005
	(0.024)	(0.040)	(0.070)	(0.091)*	(0.115)	(0.166)	(0.186)
Δq	-0.031	-0.033	-0.038	-0.038	-0.051	-0.053	-0.058
	(0.014)**	(0.015)**	(0.009)***	(0.012)***	(0.017)***	(0.019)***	(0.018)***
Δrp	-0.141	-0.214	-0.021	0.049	0.093	0.05	-0.11
	(0.050)***	(0.056)***	(0.065)	(0.108)	(0.099)	(0.107)	(0.095)
t			0.001		0.001	0.001	0.001
			(0.001)		(0.002)	(0.003)	(0.004)
Constant	-0.106		-0.301	-0.273	-0.277	-0.431	-0.356
	(0.018)***		(0.033)***	(0.050)***	(0.084)***	(0.089)***	(0.124)***
Number of id	30	30	30	30	30	29	26
Observations	732	732	732	732	732	700	631
R-squared	0.26	0.59					
RMSE	0.0264	0.0224	0.0191	0.0142	0.0127	0.0101	0.0067
Trends			0.23		0.20	0.21	0.23
lr-q	0.0621	-0.0307	-0.0779	-0.0785	-0.0965	-0.1061	-0.0718
se-q	0.0739	0.0357	0.0327	0.0374	0.0388	0.0405	0.0422
lr-rp	-0.1826	-0.0405	0.1417	0.2999	0.1325	0.1796	-0.0063
se-rp	0.1306	0.1016	0.1573	0.185	0.1661	0.2312	0.2285
CD test	-2.4749	-1.5637	4.9547	-0.0134	-0.2654	1.0079	1.3218
Abs Corr	0.1884	0.217	0.2038	0.2189	0.2216	0.2393	0.2466
Int	I(0)	I(0)	I(0)	I(0)	I(0)	I(0)	I(0)

Notes: Robust standard errors in parenthesis. * significant at 10%; ** significant at 5%; *** significant at 1%. POLS = Pooled OLS (with year dummies), 2FE = 2-way Fixed Effects, CCEP = Pooled Pesaran (2006) Common Correlated Effects (CCE), CCEPt = CCEP with linear trend, MG = Pesaran and Smith (1995) Mean Group (with country trends), CMG = Pesaran (2006) CCE Mean Group, CMGt = CMG with country-specific linear trends, CMGt1 and CMGt2 = CMGt with, respectively, one and two extra cross-sectional averages lags, as indicated by Chudik and Pesaran (2015).
CD-test reports the Pesaran (2004) test statistics, under the null of cross-section independence of the residuals. Int indicates the order of integration of the residuals (I(0) - stationary, I(1) - nonstationary) obtained from Pesaran (2007) CIPS test. RMSE presents the root mean squared error. Trend show the share of countries where the linear trend is significant at 5%. lr-q and se-q represent respectively q's long-run impact and its standard error. lr-rp and se-rp represent respectively rp's long-run impact and its standard error.

income share between 0.072% and 0.11%. However, in contrast to Karabarbounis and Neiman (2014), we do not find any empirical support for the role played by the relative prices. This findings support our theoretical model and, more important, it reconciles the labour income share decline with standard values of the elasticity of substitution.

In order to evaluate the relevance of the Tobin's Q in the secular decline of the labour income share, we undertake a simple simulation exercise. Given the fact that the GDP weighted average Tobin's Q in our sample has increased from a value of 1.15 in 1980 to a value of 1.68 in 2007 (52%), and that the labour income share has evolved from a value of 57% to 52% (-8.9%), our results imply that the increase in Tobin's Q can explain between 41% and 57% of the labour income share decline.

2.5.4 Weak exogeneity test

Our analysis has dealt with the presence of endogeneity from common factors driving both inputs and output. However, it is not uncommon in macroeconomics to suffer from endogeneity due to a reverse causality problem.[19]

Traditionally, the literature has used instrumental variable methods to solve this problem. However, given the nature of our data, it is difficult to find a valid set of instruments (i.e., variables which are correlated with the regressor but not with the error term).[20] Therefore, provided that our series are nonstationary and cointegrated, we follow Canning and Pedroni (2008); and Eberhardt and Presbitero (2015) to estimate an informal causality test based on the Granger Representation Theorem (GRT). The GRT (Engle and Granger, 1987) states that cointegrated series can be represented in the form of an ECM, which in our case is:

$$\Delta lis_{it} = \alpha_{1i} + \lambda_{11}\hat{u}_{i,t-j} + \sum_{j=1}^{K}\phi_{11ij}lis_{i,t-j} + \sum_{j=1}^{K}\phi_{12ij}q_{i,t-j}\sum_{j=1}^{K}\phi_{13ij}rp_{i,t-j} + \epsilon_{1it}, \tag{2.7}$$

$$\Delta q_{it} = \alpha_{2i} + \lambda_{21}\hat{u}_{i,t-j} + \sum_{j=1}^{K}\phi_{21ij}lis_{i,t-j} + \sum_{j=1}^{K}\phi_{22ij}q_{i,t-j}\sum_{j=1}^{K}\phi_{23ij}rp_{i,t-j} + \epsilon_{2it}, \tag{2.8}$$

$$\Delta rp_{it} = \alpha_{3i} + \lambda_{31}\hat{u}_{i,t-j} + \sum_{j=1}^{K}\phi_{31ij}lis_{i,t-j} + \sum_{j=1}^{K}\phi_{32ij}q_{i,t-j}\sum_{j=1}^{K}\phi_{33ij}rp_{i,t-j} + \epsilon_{3it}, \tag{2.9}$$

where $\hat{u}_{it} = lis_{it} - \hat{\beta}_{1i}q_{it} + \hat{\beta}_{2i}rp_{it}$ is the disequilibrium term. In order to identify a long-run equilibrium relationship, the GRT requires at least one of the λ's to be nonzero. If $\lambda_{11} \neq 0$, q and rp have a causal impact on the lis, if λ_{11}, λ_{21}, and λ_{31} are nonzero, then all variables are determined simultaneously, and no causal relationship can be identified.

[19] In our case, reverse causality implies that besides the relative prices and Tobin's Q affecting the labour income share, the labour income share has in turn, a significant impact on their values.

[20] Under the presence of unobservable common factors and parameter heterogeneity, the use of internal instruments (lags of the variables) is not valid anymore.

Table 2.7 presents the results for our weak exogeneity test. Column "Model" refers to the method used to estimate the disequilibrium term ($\hat{u}$). The two big blocks "CA" and "no CA" indicate whether equations (2.7)-(2.9) include, or not, cross-sectional averages of the variables. Within each block, the dependent variable of the system is indicated at the top of the column. The information provided shows the results for the average λ and its respective p-value. Regarding our previous discussion, for a causal effect of the Tobin's Q and the relative prices on the labour share, λ_{11} should be different from 0, while $\lambda_{21} = \lambda_{31} = 0$. We find that just 5 out of 42 cases (highlighted with asterisks) are against a causal relationship in our study. Therefore, we safely conclude that our analysis represents the causal impact of Tobin's Q and the relative price of investment on the labour income share.

TABLE 2.7: Weak exogeneity test

		no CA			CA		
Model		*lis*	*q*	*rp*	*lis*	*q*	*rp*
MG	Avg. λ	-0.52	-0.45	0.02	-0.50	-0.41	-0.04
	ρ	0.00	0.03*	0.48	0.00	0.21	0.60
CMG	Avg. λ	-0.57	-0.40	-0.01	-0.51	-0.54	0.00
	ρ	0.00	0.15	0.83	0.00	0.18	0.94
CMGt	Avg. λ	-0.75	-0.65	0.00	-0.69	-0.74	-0.04
	ρ	0.00	0.01*	0.98	0.00	0.12	0.72
CMG1	Avg. λ	-0.59	-0.23	0.04	-0.51	-0.58	0.03
	ρ	0.00	0.52	0.24	0.00	0.13	0.61
CMGt1	Avg. λ	-0.77	-0.12	0.06	-0.75	-0.60	0.05
	ρ	0.00	0.75	0.19	0.00	0.19	0.38
CMG2	Avg. λ	-0.73	-0.42	-0.07	-0.64	-1.04	-0.05
	ρ	0.00	0.32	0.09*	0.00	0.04*	0.56
CMGt2	Avg. λ	-0.93	-0.46	0.06	-0.82	-1.20	0.05
	ρ	0.00	0.29	0.25	0.00	0.01*	0.44

Notes: Avg. λ shows the robust mean coefficient for the disequilibrium term on the ECM. Asterisks highlight cases which do not support a causality relationship for our analysis.

2.6 Conclusions

The secular labour share decline has received vivid attention in the last years. Beyond the analysis of factors such as globalisation or the market institutional setup, recent contributions (Karabarbounis and Neiman, 2014; Piketty and Zucman, 2014) have focused on the role played by capital accumulation.

In particular, both Karabarbounis and Neiman (2014) and Piketty and Zucman (2014) emphasise that the increase in the capital-output ratio has been the main driver of the decline in the labour income share. However, these results rely on the assumption of both an increase in the capital-output ratio and very low diminishing returns to capital (i.e., $\sigma > 1$). Two facts that have seldom been found in the literature (Chirinko and Mallick, 2014).

This paper reconciles the labour share - capital framework for standard values of the elasticity of substitution ($\sigma < 1$) by relating the labour share decline to the increasing importance of financial markets. More specifically, we develop a model that links the labour share to financial wealth, physical capital stock, and equity Tobin's Q, which can be interpreted as a measure of the impact of financial markets on corporate investment. The intuition is the following. An increase in the equity Tobin's Q boosts financial wealth pushing investors to demand higher equity returns to hold this additional wealth. Given the connection between equity returns and the marginal productivity of capital, firms respond by decreasing the capital-output ratio, which fosters equity returns but drives the labour share down for standard values of σ (i.e., $\sigma < 1$).

Our paper is closely related to the literature on financialisation. This literature emphasises that, starting at the 80s, the goals of non-financial corporations moved from long-term investments towards short-term equity returns aiming at maximising shareholder value (Lazonick and O'Sullivan, 2000). Our mechanism resembles this literature modelling the Tobin's Q in an environment where the demand of assets is increasing with respect to their return. This implies that changes in Tobin's Q would result in an increase in stock market capitalisation and equity returns, but a decrease in corporate investment.

Empirically, we test the validity of our model estimating different Mean Group-style estimators based on a common factor model. Results suggest that the increase of the Tobin's Q since 1980 accounts for between 41% and 57% of the decline in the labour income share. In order to contrast the relevance of our model with the mechanisms recently proposed by the literature (Karabarbounis and Neiman, 2014), we include the relative price of investment in our estimations, finding no significant effects on the labour income share.

Our results show that the relationship between financial markets and corporations, embodied in the equity Tobin's Q ratio, is crucial to understand the dynamics of the capital-output ratio and factor shares. In contrast with what recent literature has argued, the decline in the labour income share is not the irreversible consequence of technological or structural factors, but the result of a change in the paradigm of financial markets, giving now priority to equity wealth and corporate payouts, at the expense of long-term corporate investment. According to our model, policies aiming at reversing the trend in the labour share should have the target of making corporate investment relatively more attractive, even if this is at the expense of equity valuation and equity returns. This could be achieved, for example, by higher capital income taxes.

APPENDIX: Supplementary tables and figures

TABLE 2.A1: Selected economies and sample period

id	Country	Sample period	id	Country	Sample period
1	Australia	1980-2008	22	Luxembourg	1991-2008
2	Austria	1980-2008	23	Mexico	1988-2008
3	Belgium	1980-2008	24	Morocco	1998-2007
4	Brazil	1992-2008	25	Netherlands	1980-2008
5	Canada	1980-2008	26	New Zealand	1986-2008
6	Chile	1990-2008	27	Norway	1980-2007
7	China	1995-2007	28	Peru	1992-2003
8	Colombia	1993-2007	29	Philippines	1988-2008
9	Denmark	1980-2009	30	Poland	1995-2008
10	Finland	1987-2009	31	Portugal	1988-2009
11	France	1980-2009	32	South Africa	1980-2008
12	Germany	1983-2008	33	Spain	1986-2008
13	Greece	1988-2009	34	Sri Lanka	1994-2008
14	Hong Kong	1980-2003	35	Sweden	1982-2009
15	Hungary	1995-2008	36	Switzerland	1980-2007
16	India	1991-2008	37	Thailand	1988-2003
17	Ireland	1981-2008	38	Turkey	1990-2003
18	Israel	1993, 1995-2008	39	UK	1980-2008
19	Italy	1980-2008	40	US	1980-2008
20	Japan	1980-2007	41	Venezuela	1992-2006
21	Korea	1980-2003			

TABLE 2.A2: Descriptive statistics

Panel A: Raw variables					
Variable	Obs	Mean	Std. Dev.	Min	Max
LIS	915	0.468	0.096	0.214	0.636
Q	915	1.241	0.268	0.519	3.229
RP	915	1.041	0.097	0.767	1.413
Panel B: Regression variables (in logs)					
Variable	Obs	Mean	Std. Dev.	Min	Max
lis	915	-0.785	0.234	-1.543	-0.452
q	915	0.195	0.200	-0.655	1.172
rp	915	0.036	0.092	-0.265	0.346

Chapter 3

Labour market dynamics in Spanish regions: Evaluating asymmetries in troublesome times

Abstract

The Spanish labour market disproportionately booms in expansions and bursts in recessions; meanwhile, its regions' relative position persists: those with the highest unemployment rates in 1996 were also in the worse position in 2012. To examine this twofold feature, we apply Blanchard and Katz's (1992) methodology and evaluate how the Spanish labour market reacts to regional employment shocks in a variety of cases. Shock responses are channelled via changes in unemployment, labour market participation, and spatial mobility. Our results provide evidence of asymmetric responses across business cycle phases (1996-2007 and 2008-2012). While changes in participation rates are the main adjustment mechanism in expansion, unemployment and spatial mobility become the central ones in recession. We also provide evidence of real wage rigidities in both periods, due to rigidities in both nominal wages and consumer prices. We conclude with a cluster analysis showing that high and low unemployment regions have similar responses in the short-run while, in the long-run, the former are more reactive in terms of spatial mobility. Overall, we provide evidence that people in a region are more willing to migrate (relative to the national average) when a regional shock occurs in relatively worse economic contexts.

JEL Codes: J20, E24, J61, R11.

Keywords: Employment, Unemployment, Regional Labour Markets, Spain.

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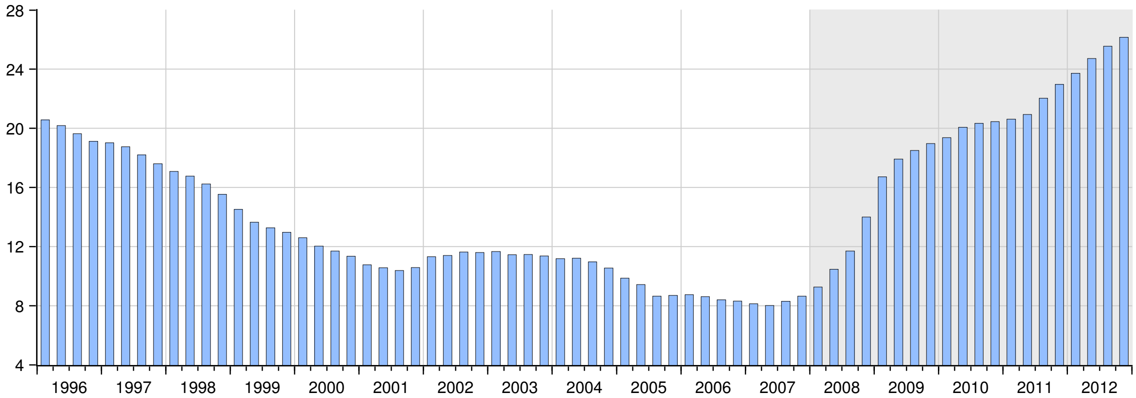
3.1 Introduction

The Great Recession has severely hit Spain in many dimensions. No more than other economies in most of them (economic (de)growth, sovereign debt crisis, banking system collapse), but disproportionately hard on unemployment. After more than a decade trending downwards and converging to the European average, the rate of unemployment reached 8.0% in 2007 –falling from a peak of 24.5% in 1994 and values above 20% still in 1996. In 2012, however, after five years of steep rise, the historical maximum was surpassed reaching a massive 26.0%.

FIGURE 3.1: Quarterly unemployment rate in Spain. 1996-2012



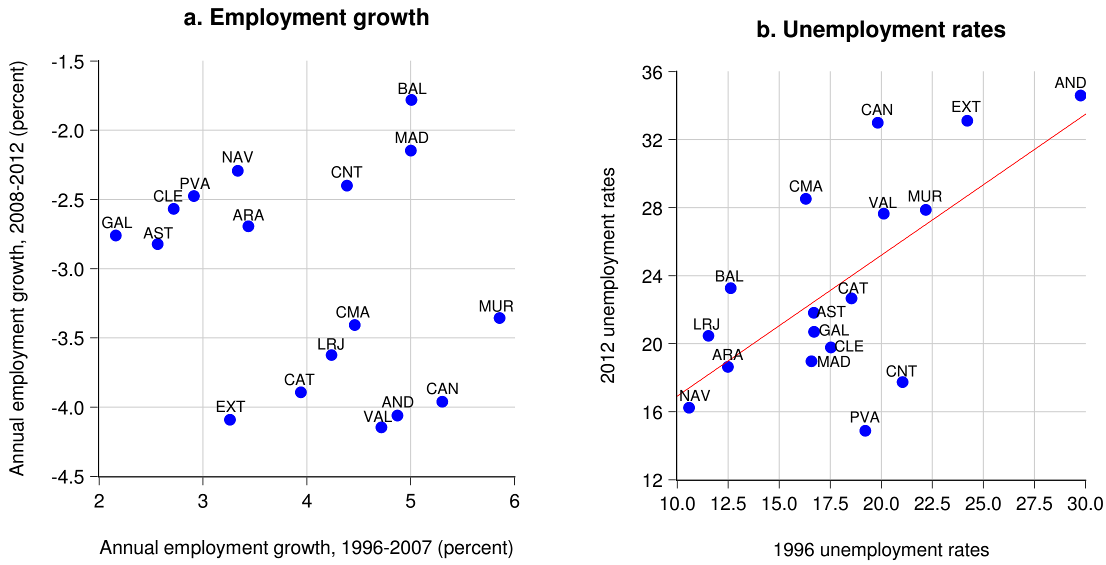
Source: Spanish Labour Force Survey (EPA).

The intense progress first, and deterioration afterwards, of the Spanish labour market goes in parallel with an extreme degree of regional persistence in labour outcomes. This is illustrated in Figure 3.2. Figure 3.2a shows two groups of regions. The first one with employment growth rates around -2.5% in 2008-2012 (with the Balearic Islands and Madrid close to -2.0%), and a second one between -3.3% and -4.1%.[1] The difference between the two groups points to the existence of less responsive regions in the North and North-West of Spain (Galicia, Asturias, Castile and Leon, Basque Country, Navarre, Aragon), and more volatile ones in the South and East part of Spain. Madrid (also the Balearic Islands and to some extent Cantabria) would be a salient exception with top employment performance simultaneously in good and bad times. Figure 3.2b, in contrast, gives a much more homogeneous picture in terms of unemployment rates, with a regression slope of 0.83 and a R^2 of 0.43. When combined, the information supplied by Figures 3.2a and 3.2b discloses two main stylised facts: (i) changes in employment provide just

[1]Given that a simple regression line takes a misleading downward slope, it is not drawn. We should rather think on two upward sloping lines, one per group, indicating that well-performing regions coincide in booms and busts.

a partial explanation of the evolution of unemployment, and (ii) there is a great persistence in regional unemployment over the years.

FIGURE 3.2: Labour market performance of Spanish regions. 1996-2012



Source: Spanish Labour Force Survey (EPA). AND=Andalusia; ARA=Aragon; AST=Asturias; BAL=Balearic Islands; CAN=Canary Islands; CNT=Cantabria; CLE=Castile and Leon; CMA=Castile-La Mancha; CAT=Catalonia; VAL=Valencian Community; EXT=Extremadura; GAL=Galicia; MAD=Community of Madrid; MUR=Region of Murcia; NAV=Navarre; PVA=Basque Country; LRJ=La Rioja.

These facts and the regional specificity of the Spanish labour market may be studied from a variety of perspectives, taking into account, along the lines of Marston (1985), that changes in regional (un)employment may be the outcome of both national and regional driving forces. Elhorst (2003) distinguishes four types of approaches including single-equation models, implicit models (where he places the Blanchard and Katz model), accounting identity models, and simultaneous-equation models dealing with interactions. The strength of the implicit models are their solid theoretical basis, while simultaneous equation models should be chosen from an empirical viewpoint (Elhorst, 2003, p. 741).

Multi-equation structural models have been used in Bande and Karanassou (2009, 2013, 2014) to assess to what extent the evolution of differences in Spanish regional unemployment can be attributed to disparities in the respective regional equilibrium unemployment rates or to the evolution of other key variables such as, for example, capital accumulation.[2] Our aim, however, is to analyse the regional labour market from a regional specific point of view. It would be too demanding,

[2]Other significant articles concerned with Spanish labour market regional disparities are López-Bazo and Motellón (2012, 2013) and, with a specific focus on wage setting, Bande et al. (2008, 2010).

in our context, to conduct a detailed analysis using their Chain Reaction Theory methodology. The reason is that we consider small sample periods of study, as deserved by the unprecedented specificities of the recent economic developments, at the same time that we need information highly disaggregated by regions. On one side, this causes severe restrictions in terms of degrees of freedom. On the other side, it constrains the analysis to a relatively small number of variables quarterly available for all Spanish regions, and with up-to-date coverage.[3]

We are interested in answering questions related to the most recent evolution of the Spanish labour market. What has happened regarding the specific regional responses to labour market shocks? Have they changed relative to previous responses, studied for the period up to the mid 1990s? Are these responses similar in good and bad times? What role do prices play? The framework of analysis developed in Blanchard and Katz (1992) allows us to provide answers to these questions. It yields the possibility of evaluating the impact of employment shocks through the responses they cause, not only in terms of the unemployment rate, but also through changes in participation rates and regional mobility.[4] Such analysis will enhance our understanding, from a regional perspective, of the labour market adjustment mechanisms in the different scenarios studied.

The model of Blanchard and Katz has been used to investigate the dynamics of regional labour markets in the US (Blanchard and Katz, 1992), Europe (Decressin and Fatás, 1995), Sweden (Fredriksson, 1999), the Netherlands (Broersma and van Dijk, 2002), Finland (Mäki-Arvela, 2003) and, more recently, for the German East-West disparities (Alecke et al., 2010). It has also been used to analyse the Spanish labour market by uncovering its regional persistence in 1976-1994 (Jimeno and Bentolila, 1998), and to provide specific analyses on the Southern regions (Murillo et al., 2006) and by level of education (Mauro and Spilimbergo, 1999).

Notwithstanding its wide use, it is important to discuss two of its prominent features since it is a model that relies upon (1) the assumption of regional mobility of workers and firms; and (2) the measurement of regional variables as deviations from the national average, which implies that shocks are regionally idiosyncratic.

[3]Note that National Accounts data at the regional level are issued with severe delays, and other regional databases, such as the BD-Mores, only cover up to 2007. Moreover, none of them provide data at quarterly frequencies.

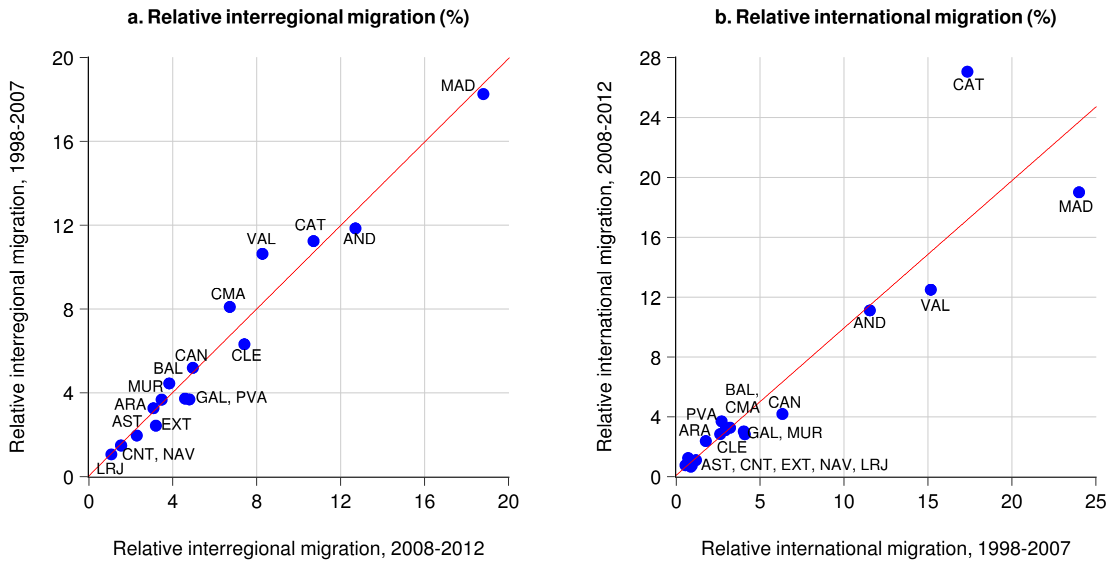
[4]"The *feedback* effects of the regional unemployment rate on labour supply, labour demand and regional wage-setting in simultaneous equations models dealing with interactions are comparable to those in Blanchard and Katz" (Elhorst, 2003, p. 723).

Regarding the first feature, it is important to show that Blanchard and Katz's model can be safely applied to study the behavior of the Spanish regions in a context of relative low interregional mobility (see Figure 1 in Arellano and Bover, 2002), and relative large international flows, regionally heterogeneous, since the end of the 1990s. In particular, it is important to ensure that our analysis of migration responses to specific shocks are not mixed with demographic changes also affecting the patterns of regional population.

Figure 3.3 plots the regional relative evolution of interregional and international migration in Spanish regions since 1998.[5]

Figure 3a shows the average relative internal residential migration rate for each region computed as the ratio of each region's residential migration (across Spanish Autonomous Communities) over total interregional migration. Note that, although there is in general low mobility, regions with a larger relative interregional migration in expansion are also the ones with larger mobility in recession.

FIGURE 3.3: Relative regional migration in Spanish regions (%). 1998-2012



Note: the complete name of each region is provided in the note below Figure 3.2.
Source: National Statistics Institute (Variaciones Residenciales Interiores y Exteriores).

Figure 3.3b shows the relative international migration rate computed as the total flows of international migration flowing into and out of each Spanish region over

[5]For Figure 3.3a, the data used is internal residential migrations by region of origin. For Figure 3.3b, both international immigration and emigration are considered. We have also checked alternative indicators (such as the internal residential migration by region of destiny, and the flows of international immigration and emigration separately), all yielding very similar results. The sample period is 1998-2012 according to available data.

the total flows of international migration in Spain. Note that it is not the relative net inflows what are computed, but the addition of inflows and outflows of international migrants in each region over total international migration flows in Spain. In this way, the computed ratios on relative interregional and international migration are directly comparable. Once again, it can be observed that each region's ratio has remained stable across both periods of analysis.

This analysis, in relative terms, is of course compatible with fluctuations in the absolute values characterised by large net inflows of international migrants during the boom years, especially in the regions with the largest employment growth rates, and the subsequent brake in these flows during the crisis.

It is important to note that these ratios have remained roughly constant between 1998-2007 and 2008-2012 which correspond, broadly, to the two periods of analysis in this work. As there are no important changes in the ratio for each region from one period to another, the regional migration response (both interregional and international) that we estimate and compare across periods is not subject to biases stemming from variations in the regional migration behaviour.[6]

Regarding the fact that we are only evaluating region-specific shocks, we acknowledge that nation-wide shocks may also be relevant, as argued by Bande and Karanassou (2009, 2014) for other periods. As it will be shown below, however, regional characteristics still play an important role in the determination of the labor market variables. Given that the effects of nation-wide shocks cannot be examined within our framework, this study should be interpreted as complementary to the existing ones conducted through the estimation of multi-equation models.

In any case, the novelty of our analysis neither lies in the use of Blanchard and Katz's methodology nor is a mere time extension of the work by Jimeno and Bentolila (1998). The paper contributes to the literature in three main dimensions.

One contribution is the specific evaluation of the effects of average regional employment shocks when hitting in expansion and when hitting in recession. For this, we use quarterly data (as Jimeno and Bentolila, 1998) and consider two subsample periods: 1996-2007, covering the expansion; and 2008-2012, covering the crisis. This disaggregation allows an evaluation of the asymmetries in shock responses across business cycle phases (upward and downward).

[6]The period averages shown in Figure 3.3 are not hiding relevant information. Yearly examination of this evolution yields the same conclusion.

Another key contribution consists in extending the labour market model to include prices. This extension was already present in Blanchard and Katz (1992), but it has generally been disregarded in subsequent literature. Consideration of prices in Spain is a relevant issue both in 1996-2007 and in 2008-2012. In expansion, it allows us to assess the response of wages to the improved economic conditions of the workers. In recession, it allows us to examine to what extent price adjustments have followed the intense quantity adjustments characterising the Spanish economy in recent years. Summing up, we offer new information on how prices respond regionally to labour demand shocks, and the potential asymmetries of these responses in good and bad times.

A third key contribution, finally, is the additional disaggregation by groups of regions based on a cluster analysis. The two resulting groups (one including Catalonia, Madrid, Navarre, and the Basque Country, and the other one grouping the rest of the regions) are used to re-estimate the models and conduct the analysis for the two groups.

Our findings are diverse. First, we identify asymmetric labour market responses across business cycle phases. We find that changes in participation rates are the main adjustment mechanism in expansion, while unemployment becomes the central one in recession. Moreover, the long-run employment impact is larger when the shock hits in a recessive period than when it hits in expansion. This result is an indication that net migration –spatial mobility– is more relevant in troublesome than in good times.

We also provide evidence of real wage rigidities in both periods along the lines of Jimeno and Bentolila (1998).[7]

And there is, finally, evidence of similar labour market dynamics across high and low unemployment regions generated by the one-off employment shocks. This is consistent with the large degree of aggregate (un)employment persistence characterising Spain, and is to some extent reassuring in the sense that consideration of an average Spanish region is not flawing the results. Nevertheless, we still find differences in the relative long-run regional employment impact and unemployment persistence, resulting on larger spatial adjustments in high than in low unemployment regions, which appear as more resilient to the shock. On this account, it

[7] Jimeno and Bentolila (1998) uncovered a high degree of persistence in the Spanish regions, as compared to the US and the EU. This was mainly due to real wage rigidities and low interregional migration.

seems safe to conclude that people in a region are more willing to migrate (relative to the national average) not just when regional shocks take place in a recessive period, but also when they impact in places with larger relative unemployment rates.

The remaining of the paper is structured as follows. In Section 3.2, we outline the analytical framework and its empirical implementation. In Section 3.3 we present the data used and the econometric methodology. In the following four sections we show our findings related, respectively, to the aggregate analysis, the disaggregation by business cycle phases, the inclusion of price responses, and the consideration of two groups of regions. Section 3.8 concludes.

3.2 Analytical framework

To conduct our analysis, we use the framework developed in Blanchard and Katz (1992). This framework is derived from a set of equations representative of the average regional labour market within a given economy, and entails the estimation of the following three reduced-form equations:

$$\begin{aligned}
\Delta n_{it} &= \lambda_{10} + \lambda_{11}(L)\,\Delta n_{it-1} + \lambda_{12}(L)\,u_{it-1} + \lambda_{13}(L)\,pr_{it-1} + \varepsilon_{int}, && (3.1)\\
u_{it} &= \lambda_{20} + \lambda_{21}(L)\,\Delta n_{it} + \lambda_{22}(L)\,u_{it-1} + \lambda_{23}(L)\,pr_{it-1} + \varepsilon_{iut}, && (3.2)\\
pr_{it} &= \lambda_{30} + \lambda_{31}(L)\,\Delta n_{it} + \lambda_{32}(L)\,u_{it-1} + \lambda_{33}(L)\,pr_{it-1} + \varepsilon_{ipt}, && (3.3)
\end{aligned}$$

where n is relative employment (in logs), u is the relative unemployment rate (in %), and pr is the relative participation rate (in logs); L is the lag operator; i stands for region, t for period; the $\lambda's$ are parameters, and the $\varepsilon's$ are residuals.

The term *relative* affecting n, u, and pr indicates the beta-difference of these variables in region i with respect to the national average. More precisely, these three variables are defined as the residuals of the following equations:

$$\begin{aligned}
\Delta\log(N_{it}) &= \alpha_{1i} + \beta_i\Delta\log(N_t) + \mu_{1it}, && (3.4)\\
U_{it} &= \alpha_{2i} + \gamma_i U_t + \mu_{2it}, && (3.5)\\
\log(PR_{it}) &= \alpha_{3i} + \delta_i\log(PR_t) + \mu_{3it}, && (3.6)
\end{aligned}$$

where N, U and PR denote employment, the unemployment rate and the participation rate (all in absolute values); and β_i, γ_i, δ_i account for the regional sensitivity of these three variables with respect to changes in their national counterpart. The $\alpha's$ are the regional constants, whereas the $\mu's$ are the corresponding residuals.

In computing the values of the relative variables n, u, and pr as beta-differences, we follow the same methodology than Decressin and Fatás (1995) and Broersma and van Dijk (2002), who also use the residuals from equations (3.4)-(3.6). It is important to note that, by creating relative variables through beta-differenciation, we isolate the part of the variation in the regional variable that is not due to national changes. Each region is thus allowed to respond differently to a national shock. This is a situation empirically relevant whenever the beta coefficient $-\beta_i$, γ_i, δ_i in equations (3.4)-(3.6)– are significantly different from unity.

This procedure, however, is not the only possibility at hand, and a decision needs to be taken on the basis of the intended analysis to be performed. The alternative followed by Blanchard and Katz (1992) and Jimeno and Bentolila (1998) is to create regional specific variables as the log-difference between the regional and the national ones.[8] In this case, relative variables are not just considering regional responses to asymmetric shocks, but also regional asymmetric responses to common shocks.[9]

Another useful piece of information delivered by equations (3.4)-(3.6) is their adjusted R^2's. Taking as example equation (3.4), note that the value of this coefficient indicates the extent to which the pattern of regional employment growth fits the pattern of employment growth at the national level. For example, low values of the adjusted R^2's show that most movements in regional employment are not driven by national changes in labour demand. For the whole sample of our analysis (1996-2012), the average adjusted R^2 is 0.48 implying that just half of the regional employment growth variation can be explained, on average, by national trends.

Summing up, this methodology involves, first, the estimation of equations (3.4)-(3.6); and, second, the use of the residuals as relative variables to estimate the system of equations (3.1)-(3.3).

Once this process is completed, we shock the system with a one-off shift in the residual of equation (3.1). Such unexpected and temporary shocks on employment

[8]Although our main analysis is based on the results obtained when taking beta-differences, the Appendix provides the same set of results obtained on the basis of log-differencing the variables. It is important to note that both deliver the same qualitative picture, even though each sets of results requires its own assessment.

[9]See L'Angevin (2007) for a detailed explanation on the interpretation of different 'relative' regional variables.

growth are the ones evaluated in our analysis. It is important to note that, when we examine the dynamics of these shocks, we are not evaluating the persistence in the average value of the variable under scrutiny (no matter whether this is the employment growth or the unemployment and participation rates). What we are checking, rather, is the speed of convergence of this variable to the previous equilibrium, wherever this one is. This means that these shocks can be formally introduced as positive shocks in all cases, even if their effects are evaluated for a sample period just containing recessive years (as we do later on).

The interpretation of the effects of the shock is made under two common assumptions in the literature. The first one is that unexpected changes in regional relative employment within a year are due to changes in labour demand. This assumption is considered to be correct when most year-to-year unexpected movements in employment are caused by shifts in labour demand rather than shifts in labour supply. This is arguably the case in Spain in the last business cycle.

The second assumption concerns the identification of the shock. It states that employment growth is independent of current changes in the unemployment and participation rates, whereas these rates respond contemporaneously to changes in employment. Jimeno and Bentolila (1998) note that this assumption is more likely to hold when using quarterly data than when using annual data, which is precisely the case here.

When the shock takes place, it gives rise to three adjustment mechanisms, two of them directly arising from the estimated model, and a third one computed as a residual.[10] The first two arise from equations (3.2) and (3.3), where the reactions in terms of unemployment and participation rates take place. By construction of the model, whatever employment stimulus not absorbed by a decrease in unemployment or an increase in participation can be ascribed to spatial mobility. Mobility is thus the third adjustment mechanism, and is computed as the difference between the overall employment response to the shock and the unemployment and participation rates responses.

Blanchard and Katz (1992) consider an extension of the basic labour market model where price responses are also evaluated. This requires the addition, to the system of equations (3.1)-(3.3), of the following price equation:

$$w_{it} = \lambda_{40} + \lambda_{41}(L)\,\Delta n_{it} + \lambda_{42}(L)\,u_{it-1} + \lambda_{43}(L)\,pr_{it-1} + \lambda_{44}(L)w_{it-1} + \varepsilon_{iwt}, \quad (3.7)$$

[10] Alecke et al. (2010) augment Blanchard and Katz's (1992) model with a migration equation. Unfortunately, data on migrations is not available for Spain at quarterly frequencies.

where w_{it} represents two different price variables depending on the estimated model: nominal wages (in our case, the hourly total labour cost) and consumer prices (the standard CPI).[11] We consider this addition because it sheds new light on the price behaviour in the aftermath of average regional shocks in good and bad times.

We endeavour to examine price responses in Spain because of the conspicuous link between labour market prices and quantities. It is on this account that we incorporate nominal wages and a price deflator to the analysis. Jimeno and Bentolila (1998) had already dealt with this issue by examining the wage response to local economic conditions in 1983-1988. They found "a low responsiveness of wages to regional economic conditions" in which nominal wages were less flexible than prices. Here, we retake this issue, but sticking to Blanchard and Katz's (1992) methodology.

3.3 Data and estimation issues

As noted, we first need to estimate equations (3.4), (3.5) and (3.6), and then equations (3.1), (3.2) and (3.3). The first set of equations consist on a time-series estimation by Ordinary Least Squares (OLS), whereas the second set of equations calls for the use of Panel Vector AutoRegression (PVAR) techniques estimated by System OLS.

3.3.1 Data

Information on the labour market variables (employment, participation rate, and unemployment rate) is obtained from the Labour Force Survey (*Encuesta de Población Activa*, EPA, from its Spanish acronym). In turn, information on the labour and product prices variables comes, respectively, from the Quarterly Survey of Labour Costs (*Encuesta Trimestral de Coste Laboral*, ETCL) and the National Statistics Institute (*Instituto Nacional de Estadística*, INE). Table 3.1 presents the notation, sources and available sample periods for each of these variables.

The Labour Force Survey underwent a methodological change, in 2002, affecting the definitions of unemployment and participation rates. Since our study departs

[11] Further to the addition of equation (3.7), note that consideration of an aditional variable in the model implies the addition of the term $\lambda_4(L)w_{it-1}$ in each of the equations in model (3.1)-(3.3).

from 1996 (coinciding with the previous methodological change) to cover at length the last business cycle (1996 – 2012), we need to construct homogeneous series. We do that using the official link coefficients supplied by INE itself, and use the resulting series in our analysis.

TABLE 3.1: Definitions of variables

	Variables	Sources	Time Period
Labour market			
N_{it}	Employment	EPA	$1996q1 - 2012q4$
PR_{it}	Participation rate	EPA	$1996q1 - 2012q4$
U_{it}	Unemployment rate	EPA	$1996q1 - 2012q4$
Prices			
TC_{it}	Hourly total labour costs	ETCL	$2000q1 - 2012q3$
CPI_{it}	Consumer prices index	INE	$1996q1 - 2012q4$

From the ETCL we obtain the effective hourly total labour cost TC, which we think it is the relevant variable when examining how sensitive factor prices are to labour market shocks. The TC is the gross cost paid by the employer taking into account any other cost beyond the wage. Note that this variable has a shorter sample size starting in $2000q1$ and finishing in $2012q3$. The variable for prices, CPI, is the standard consumer price index.

All our variables are disaggregated regionally, have quarterly frequencies, and are seasonally adjusted by using the US X12 Census Bureau process.

3.3.2 Estimation methodology

The PVAR econometric model to be estimated takes the following reduced form in matrix notation:[12]

$$\mathbf{y}_{i,t} = \Gamma_0 + \Gamma_1(L)\mathbf{y}_{i,t-1} + \varepsilon_{it}, \tag{3.8}$$

where $y_{i,t}$ is the vector of endogenous variables (in our case $y_{i,t} = \Delta n_{it}, u_{ir}, pr_{it}$), $\Gamma_1(L)$ is a matrix of the reduced form coefficients relating past variable values to current values, Γ_0 is a vector of constants, and ε_{it} is a vector of idiosyncratic errors.

Since our relative variables are created by orthogonalising the regional variables with respect to the national average, there is no reason for fixed effects to be

[12] We estimate a PVAR(2). The lag order of the PVAR is chosen to use the maximum sample period available without neglecting the relevance of dynamics. For robustness, the PVAR has been estimated using different lag orders. The results remain roughly the same and are available upon request.

introduced in the estimated equations. This alleviates our estimates from the well-known dynamic panel data estimation problems, and allows us to proceed with System OLS, rather than System GMM, estimation.[13]

Once the estimation is performed, we compute impulse-response functions (IRFs) describing the reaction of the dependent variables to changes in the innovation of one particular variable in the estimated system. Following the model of Blanchard and Katz this variable is employment growth. We will thus evaluate the dynamics of the labour market to one-off shocks in regional employment.[14]

3.3.3 Spatial dependence

Regional variables may be liable to spatial correlation even though, in our study, the potential incidence of this problem should be lessened by the fact that national averages are subtracted from the regional values. Nevertheless, in order to discard that the regional distribution of each variable in the model is not random, and thus causes biased and inconsistent results, we run Moran's I test (Moran, 1950).

We construct a binary contiguity weighting matrix $\mathbf{W}$ in which the i, j elements (corresponding to the relative position of region i with respect to region j) take value 1, $\bar{w}_{ij} = 1$, if the involved regions share their borders, at least partially; and take value 0, $\bar{w}_{ij} = 0$, otherwise. We then standardise $\mathbf{W}$ so that the rows add up to unity and regions with a small number of borders do not have excessive weights.

For each quarter in our sample, we conduct the two tail version of the test so that the null hypotheses of randomness (i.e., no spatial dependence) is contrasted against the alternative of no randomness (or spatial correlation). Table 3.2 reports the results we obtain for all variables of interest at the 1% and 5% critical values. The information shown is the number of periods for which the null of randomness is rejected and the corresponding percentage over the total number of quarters in each subsample period.

[13]System GMM estimates of the PVAR models with fixed effects are provided in Sala and Trivín (2013). However, System GMM is a more inefficient estimation method in our context. We owe this point to an anonymous referee who made us note that fixed effects were already introduced when defining the relative variables and, thus, were no longer required when estimating. It is worth noting, also, that our findings are robust across econometric methodologies.

[14]All PVAR models in this paper are estimated by using the package provided by Ryan Decker, which is an update of the original package developed by Inessa Love and used in Love and Zicchino (2006).

TABLE 3.2: Results on Moran's I test

	1% critical value				5% critical value			
	1996-2007		2008-2012		1996-2007		2008-2012	
	# periods	%	# periods	%	# periods	%	# periods	%
Δn_{it}	1/47	2.1%	0/19	0.0%	5/47	10.6%	0/19	0.0%
u_{it}	2/48	4.2%	0/20	0.0%	10/48	20.8%	0/20	0.0%
pr_{it}	1/48	2.1%	0/20	0.0%	5/48	10.4%	2/20	10.0%
tc_{it}	1/32	3.1%	0/19	0.0%	3/32	9.4%	2/19	10.5%
cpi_{it}	6/48	12.5%	1/20	5.0%	14/48	29.2%	1/20	5.0%

Notes: Detailed test results are available from the authors upon request.

Regarding the labour market variables, there is no indication of serious spatial dependence. In the worst case, the unemployment rate variable examined at the 5% critical value in 1996-2007, we cannot reject the null of randomness in 38 out of 48 quarters. Then, with respect to prices, we can safely discard regional correlation at the 1% critical value. It is important to note that these results are consistent with the reported findings in Bande and Karanassou (2014) for labour market variables using annual data between 1980 and 2000.

3.3.4 Panel unit root tests

Another important issue is the potential presence of unit roots in the variables. Hence, to check the validity of our estimation we have to prove that stationary panel data techniques are appropriate given the integration order of our variables.

To do that, we conduct a series of panel unit root tests. Although it is well-known that the popular individual unit root tests –Dickey-Fuller (DF), Augmented Dickey-Fuller (ADF), and Phillips-Perron (PP) tests– have limited power in distinguishing the null of a unit root from stationary alternatives with highly persistent deviations from equilibrium, it is also generally accepted that the use of pooled cross-section time series data can generate more powerful unit root tests (Levin et al., 2002).

Taking this into account, we conduct a series of panel unit tests to check if the use of stationary panel data estimation techniques is appropriate in our context. We thus carry out the statistic test proposed by Maddala and Wu (1999), which is an exact nonparametric test based on Fisher (1928):

$$\zeta = -2\sum_{i=1}^{N} \ln \pi_i \sim \chi^2(2N),$$

where π_i is the probability value of the ADF unit root test for the ith unit (region in our case).

This test has the following attractive characteristics: (i) it does not restrict the autoregressive parameter to be homogeneous across i under the alternative of stationarity; and (ii) the choice of the lag length and of the inclusion of a time trend in the individual ADF regressions can be determined separately for each region.

TABLE 3.3: Panel unit root tests

	$\zeta(\Delta n_{it})$	$\zeta(u_{it})$	$\zeta(pr_{it})$	$\zeta(cpi_{it})$	$\zeta(tc_{it})$
96-07	557.23	104.72	122.08	63.88	
08-12	175.21	73.81	67.49	86.65	
00-07	319.66	100.10	111.71		237.93
08-12*	202.44	70.98	58.73		176.03

Notes: $\zeta(.)$ is the test proposed by Maddala and Wu (1999). It follows a chi-square distribution whose 5% critical value is 48.6. Lower case letters denote relative variables. * The data here runs from 2008q1 to 2012q3 due to data limitations in tc_{it}.

Table 3.3 shows the results of Maddala and Wu's (1999) unit root tests for our 5 variables of interest. The test statistic, ζ, follows a chi-squared distribution, which in our case has a 5% critical value of approximately 49. It is easy to see that all the panel unit root test statistics are greater than the critical value, and the null of a unit root can therefore be rejected at the 5% significance level. It is thus safe to proceed with the analysis by applying stationary panel data techniques.[15]

3.4 Aggregate results

We start by presenting an aggregate picture along the lines of previous literature. It comprises the whole sample period, does not yet consider prices, and all regions are taken into account. These three issues –splitting the sample, considering prices, and grouping the regions– will be faced in subsequent sections.

Table 3.4 shows the estimated β, γ and δs, together with the corresponding adjusted R^2s for the regional regressions of equations (3.4), (3.5), and (3.6) corresponding to years 1996-2012.

[15]Note that we only present a selection of test results, which correspond to some specific periods and do not cover entirely the estimated models in sections 3.4, 3.5, 3.6 and 3.7. Results for the whole sample period and all regions, and for the whole sample period and the high and low unemployment groups of regions, follow the same pattern. They are available upon request.

TABLE 3.4: Summary of estimates. 1996-2012

	Equation (3.4)		**Equation (3.5)**		**Equation (3.6)**	
	$\bar{R}^2$	$\hat{\beta}$	$\bar{R}^2$	$\hat{\gamma}$	$\bar{R}^2$	$\hat{\delta}$
AND	0.72	1.15	0.94	1.27*	0.98	1.08*
ARA	0.44	0.72*	0.92	0.84*	0.97	1.17*
AST	0.31	0.90	0.87	0.73*	0.95	1.30*
BAL	0.29	0.88	0.78	1.01	0.96	0.89*
CAN	0.46	1.12	0.83	1.35*	0.97	0.96*
CNT	0.34	0.84	0.74	0.73*	0.96	1.17*
CLE	0.65	0.76*	0.94	0.71*	0.98	1.01
CMA	0.64	1.04	0.91	1.11*	0.98	1.31*
CAT	0.81	1.15*	0.98	0.99	0.98	0.71*
VAL	0.74	1.17*	0.96	1.21*	0.98	0.87*
EXT	0.27	0.93	0.89	0.96	0.96	1.08*
GAL	0.29	0.62*	0.81	0.62*	0.95	0.69*
MAD	0.60	0.95	0.97	0.85*	0.98	1.27*
MUR	0.42	1.16	0.97	1.26*	0.98	1.11*
NAV	0.36	0.77	0.92	0.64*	0.95	0.87*
PVA	0.46	0.81	0.54	0.55*	0.95	0.68*
LRJ	0.28	0.92	0.83	0.82*	0.96	1.32*

Notes: * indicates $\hat{\beta}$, $\hat{\gamma}$ and $\hat{\delta}$ are different from unity at a 5% critical value; for the list of regions, see the notes below Figure 3.2.

We find the estimates of β to be significantly different from unity (at a 5% critical value) in 5 out of 17 regions. In contrast, the estimates of γ and δ are significantly different from unity in 14 and 16 regions respectively. In view of these results, we create the regional specific variables allowing regions to respond differently to common shocks.

The values of the adjusted R^2s corresponding to equation (3.4) provide a measure of the relevance of the regional shocks vis-à-vis the national ones. The fact that this value is below 0.50 in 11 out of 17 territories is an indication that regional-shocks are very relevant for the understanding of the Spanish labour market behaviour.

The residuals from these estimated equations are used to examine the aggregate impulse responses to a regional labour demand shock. These are plotted in Figure 3.4. Figure 3.4a shows the reaction of the average Spanish region to this shock in terms of employment, and the participation and unemployment rates. Employment converges to 0.273 indicating that 27.3% of the one-off shock is translated into a larger, long-run, relative regional employment level, which is covered by an increase in population (spatial mobility). The rest of the shock is absorbed by the growing participation rates and falling unemployment rates.

It should be noted that Figure 3.4a provides normalised responses to the shock, while Figures 3.4b to 3.4d deliver the original impulse-response functions, together with their standard errors. The reason for normalising is that the IRFs are calculated on one standard deviation shocks and may deliver small divergences from unity. We normalise to ensure comparability across results in next sections. The standard errors are calculated using Monte Carlo simulations with 500 replications.[16]

FIGURE 3.4: Aggregate IRFs to a regional employment shock. 1996-2012

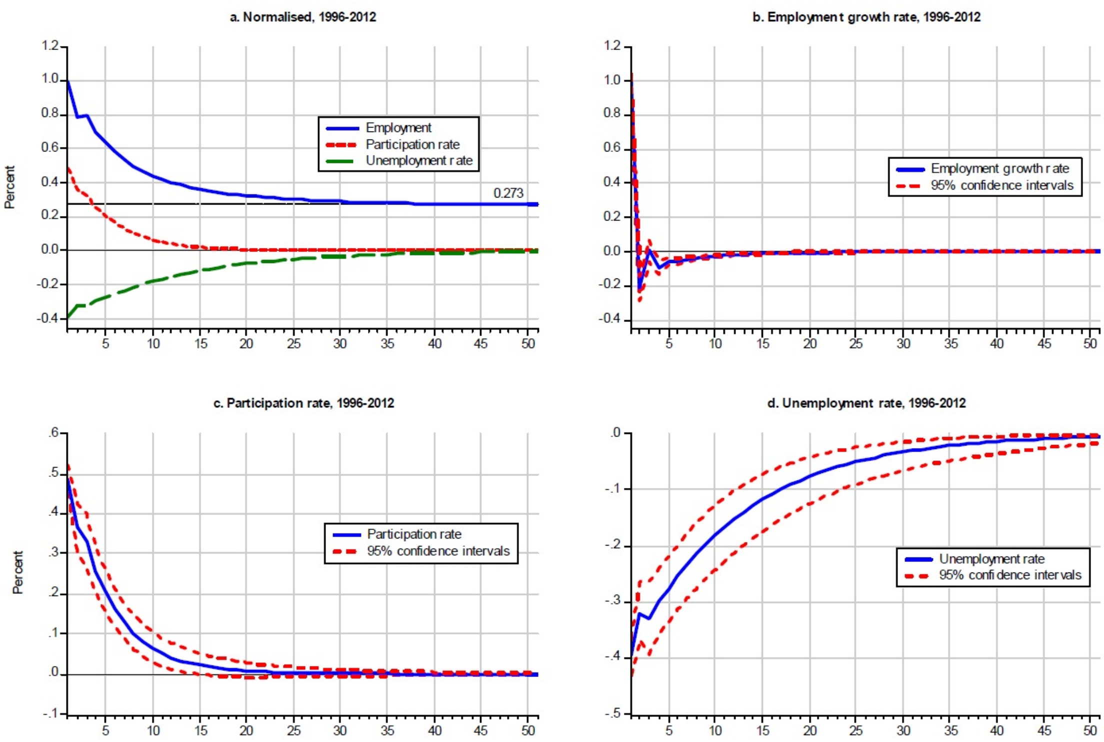


Figure 3.4a is directly comparable to Figure 2 in Jimeno and Bentolila (1998), and yields a similar picture both in terms of the labour market dynamics generated by the shock, and in terms of its employment effect. They place this effect at 0.40, while Decressin and Fatás (1995) place it at 0.60 for the European Union.

Table 3.5 provides a detailed comparison of the way a stylised region responds to a regional employment shock today (1996-2012) with its reaction in the past (1976-1994). The main differences observed are the following. First, adjustments via changes in participation rates are much more relevant today than in the past, especially in the short-run (43% versus 23% in the first year). However, there is a much lower persistence today causing the adjustment in the participation rate

[16] To conserve space, for the rest of the analysis we only show our results in terms of the normalised impulse-response functions.

TABLE 3.5: IRFs decomposition to a 1% regional employment shock

	Spain (1976 – 94)*			Spain (1996 – 2012)		
	Participation	Unemployment	Migration	Participation	Unemployment	Migration
Year 1	23%	36%	41%	43%	41%	16%
Year 2	18%	39%	43%	26%	43%	31%
Year 3	18%	33%	49%	14%	40%	46%

Notes: * Results taken from Jimeno and Bentolila (1998, p. 33). Quarterly data aggregated to annual values and normalised by the employment response in the year.

to fall by 4 percentage points in 1996-2012 relative to years 1976-1994. Second, unemployment displays slightly larger values and persistence today than in the past. The larger incidence of the participation and unemployment rate channels leaves spatial mobility as a minor adjustment mechanism in the short-run. In the long-run, however, quick convergence in participation allows migration to stay as relevant, today, as it was in the past (49% vs. 46%).

3.5 Labour market dynamics during the 'wild-ride' and the 'steep-fall' periods

So far we have studied the dynamics of the average region for the whole last business cycle. Some authors, however, have come out with the idea that some Spanish regional unemployment features (e.g. regional disparities) are related with the different phases of the business cycle (Bande et al., 2008). Accordingly, this section aims at examining regional-specific dynamic adjustment mechanisms when a positive shock hits the economy in good and bad times.

Table 3.6 shows the estimated βs, γs, δs, and the corresponding adjusted R^2s for the regional regressions of equations (3.4), (3.5) and (3.6) conducted for two sub-sample periods: 1996-2007, corresponding to the 'wild-ride' of the Spanish economy during those years, and 2008-2012, corresponding to the subsequent steep fall. We acknowledge the fact that these estimates may be sensitive to the sample period length. However, they correspond to such specific and contrasted developments that it is worth examining them individually. Moreover, in order to discard biases, next section provides two robustness checks related to the shortening of the first sub-sample period (to 2000-2007 due to data availability in the total labour costs variable) and to the addition of different price variables to the system. As we will see, the results are robust to these changes.

TABLE 3.6: Key estimates for equations (3.4), (3.5) and (3.6)

	1996-2007						**2008-2012**					
	$\bar{R}^2$	$\hat{\beta}$	$\bar{R}^2$	$\hat{\gamma}$	$\bar{R}^2$	$\hat{\delta}$	$\bar{R}^2$	$\hat{\beta}$	$\bar{R}^2$	$\hat{\gamma}$	$\bar{R}^2$	$\hat{\delta}$
AND	0.23	1.39	0.94	1.51*	0.98	0.95*	0.51	0.84	0.99	1.23*	0.47	2.52*
ARA	-0.01	0.24*	0.88	0.62*	0.96	1.16*	0.07	0.46*	0.96	0.87*	-0.04	-0.49
AST	0.06	1.25	0.83	0.74*	0.91	1.17*	0.30	1.18	0.96	0.97	-0.06	-0.04
BAL	0.04	0.98	0.63	0.48*	0.95	0.91*	0.06	0.82	0.93	0.98	0.05	1.20
CAN	0.13	1.45	0.86	0.77*	0.96	0.98	0.28	1.25	0.98	1.16*	0.11	1.94
CNT	-0.003	0.55	0.97	1.18*	0.94	1.24*	0.02	0.35*	0.99	0.77*	-0.05	-0.28
CLE	0.22	0.99	0.98	0.88*	0.97	0.93*	0.41	0.62*	0.99	0.75*	0.24	1.40
CMA	0.03	0.54	0.96	0.71*	0.98	1.21*	0.33	0.92	0.98	1.21*	0.20	1.80
CAT	0.27	1.13	0.97	1.02	0.99	0.77*	0.80	1.46*	0.99	0.99	-0.05	0.20
VAL	0.14	0.90	0.96	1.002	0.99	0.94*	0.56	1.12	0.98	1.14*	0.01	-0.91*
EXT	0.04	1.05	0.89	0.93	0.95	1.04	0.04	0.88	0.90	1.23*	0.11	2.24
GAL	0.03	0.69	0.83	0.71*	0.92	0.65*	0.05	0.44	0.96	0.87*	0.26	1.37
MAD	0.11	0.90	0.95	0.91*	0.98	1.37*	0.38	1.03	0.98	0.76*	0.15	1.11
MUR	-0.002	0.61	0.96	1.15*	0.96	1.08*	0.12	0.87	0.97	1.15*	0.31	1.55
NAV	0.16	1.27	0.86	0.49*	0.94	0.97	0.03	0.59	0.95	0.67*	0.02	-0.72*
PVA	0.11	0.92	0.98	1.10*	0.98	0.79*	0.26	1.04	0.94	0.59*	-0.03	0.42
LRJ	0.07	1.27	0.73	0.46*	0.95	1.43*	-0.06	-0.01	0.91	0.91	-0.03	0.58

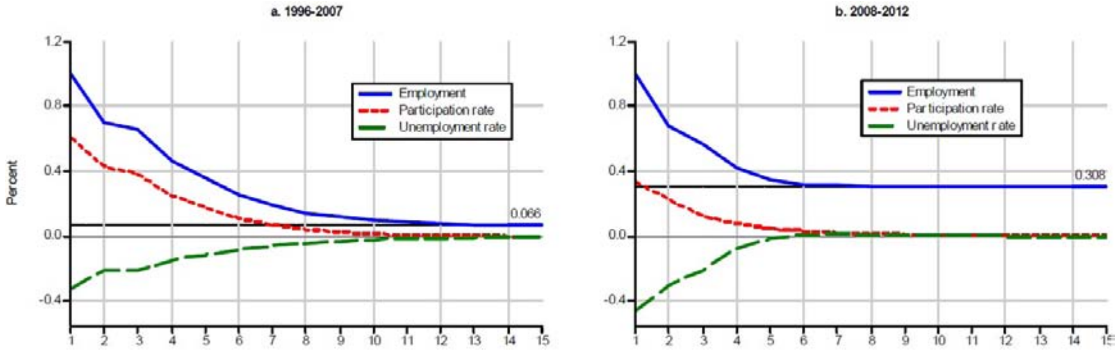
Notes: * indicates $\hat{\beta}$, $\hat{\gamma}$ and $\hat{\delta}$ are different from unity at a 5% critical value; for the list of regions, see the notes below Figure 3.2.

The first salient finding is the relatively large adjusted R^2s of equation (3.4) for years 2008-2012, in contrast to the low ones for 1996-2007. This is an indication that nation-wide shocks have become more relevant during the steep-fall years, and is consistent with the fact that Spain was fully caught by the Great Depression, which has driven all regions to an unprecedented slump. The few cases, not infrequent in the literature (see Decressin and Fatás, 1995), in which the adjusted R^2s fall around zero reveal the prominence of regional shocks.

A second noticeable result concerns the change in the participation rate behaviour. There is a stark contrast between the values of the adjusted R^2 in the first period (above 0.90 in all regressions) and the second one (in which they become low). To discard the possibility that this is due to the short sample period of 2008-2012, we shortened the first period to 2003-2007 to take the same length. Since the results remained essentially unchanged, we credit the hypothesis that regional participation rates have radically changed their behaviour. A behaviour that was mainly determined at the national level but has become regionally specific. This finding reinforces our strategy of disaggregating the analysis in subperiods.

As before, we shock the estimated systems and compute the resulting IRFs for each period of analysis. These are shown in Figure 3.5.

Figure 3.5: IRFs to a regional employment shock in 1996-2007 and 2008-2012



The long-run impact of the shock on the relative regional level of employment is smaller in 1996-2007 than in 2008-2012. This reflects the enhanced spatial mobility during the crisis, people in a region being more willing to migrate (relative to the national average) when there is a regional shock in bad times than when the shock impacts in good times. On top of these differences in the adjustment mechanisms across periods, persistence in the participation and unemployment rates responses is somewhat reduced in 2008-2012, when they take around 6/7 quarters, rather than 9, to converge to equilibrium. This result contrasts with the larger persistence of these variables (close to 20 and 35 quarters, respectively) when all years in the sample are examined together. This is mainly due to the analysis of much more homogeneous periods when the full sample is split in expansionary and recessive years, than when both subsamples are considered together (see Altissimo et al., 2006; and Altissimo et al., 2009).

Regarding the decomposition of the different components in terms of their influence on the adjustment process, Table 3.7 shows the large response of relative regional participation rates in 1996-2007. It explains around 60% of the adjustment in the first quarter, and still accounts for more than 48% of it in quarter 15. In contrast, the immediate response of unemployment dominates in 2008-2012 (more than 45% versus less than a third the participation rate response), although it vanishes progressively. The picture at the end is one of stark differences across periods. While adjustments during the wild ride are distributed more homogeneously with a dominant participation rate mechanism, during the crisis it is migration, by far, what leads the adjustment.

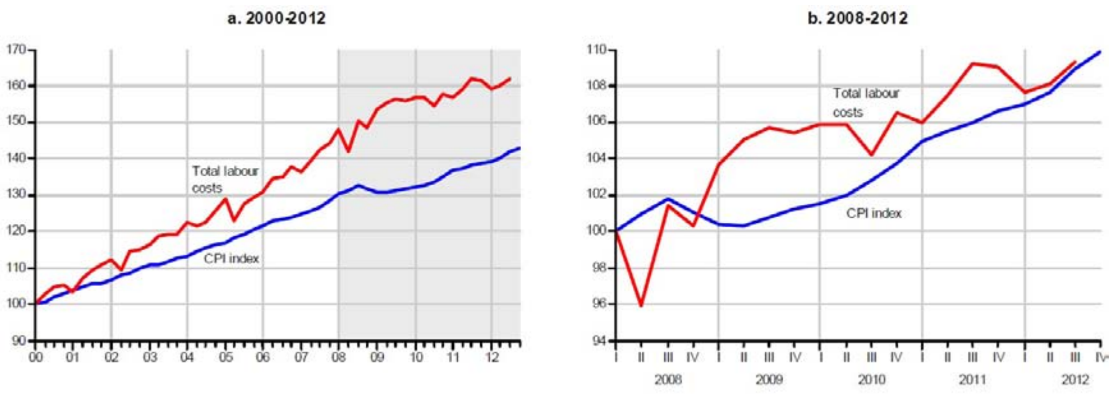
Table 3.7: IRFs decomposition to a 1% regional employment shock

	1996-2007	**2008-2012**
Final employment effect:	0.07%	0.31%
Adjustment in 1st quarter by:		
Participation	60.5%	33.4%
Unemployment	32.5%	46.3%
Migration	7.0%	20.3%
Cumulative adjustment in 15th quarter by:		
Participation	48.5%	14.8%
Unemployment	29.8%	16.8%
Migration	21.7%	68.4%

3.6 Price responses

Price rigidities are an important area of interest in macroeconomics and labour economics. In this section we enquire to which extent they are also relevant at the regional level. To outline the most recent developments, Figure 3.6 shows the average evolution of product and labour market prices in 2000-2012 and 2008-2012.[17]

Figure 3.6: Product and labour market prices. Index 100



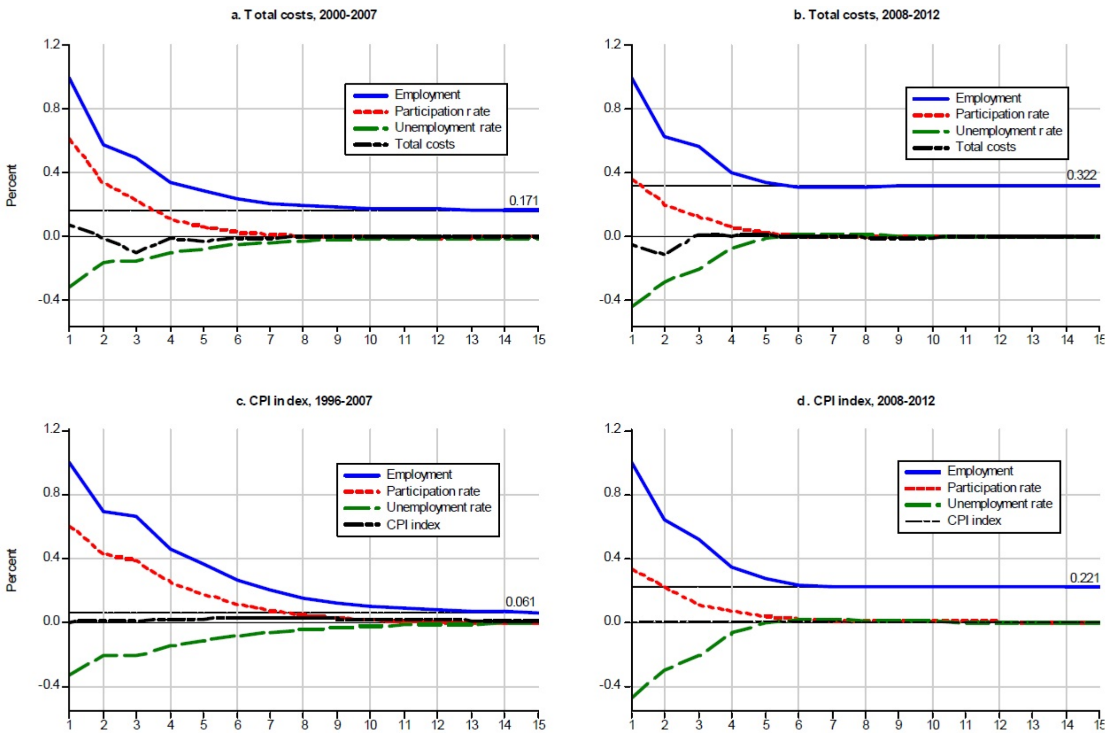
[17]It is worth noting that despite these price evolutions are presented at a national level, their patterns are alike across Spanish regions. The purpose of this figure, therefore, is to inform about the economic context in which the impact of the shock will be studied regarding its price effects. We start in 2000 since the total labour cost series is only available 2000 onwards, and we need full comparability across series.

The striking feature is the fact that nominal wages (as measured by the hourly total labour cost) have always grown above the CPI with hardly any exception, even during the most severe crisis years 2011-2012. Between 2000 and 2007, nominal wages grew by 40%, consumer prices by 20% and, as a result, real wages grew by around 20%. This figure, therefore, shows that nominal wages, consumer prices and, therefore, real wages, are highly insensitive to the business cycle.

Figure 3.7 displays the labour market IRFs when the system of equations (3.1)-(3.3) is augmented with equation (3.7) so that prices are included in the analysis.

Comparison of the responses by nominal wages (total labour costs) and consumer prices (CPI index) allows to infer the reaction of total compensation in constant terms through the dynamics of its two components. Due to restricted availability of data regarding the total labour costs, we are bound to reduce the system estimation for the first period to years 2000-2007, which implies loosing a third of the sample period (16 quarters of information). This is not the case when using the CPI index and we thus have a natural robustness check. More precisely, comparison of Figures 3.7a and 3.7c allows us to discard any significant bias resulting from the shortening of the sample period. This reassures us regarding the results for the second period, which is shorter by definition.

FIGURE 3.7: IRFs to a regional employment shock in total costs and product prices



The average regional labour market response across the four different models (Figure 3.7a to 3.7d) yields a substantive robust picture. Although consideration of total costs cause larger long-run relative regional employment effects, it is clear that the qualitative response in terms of the three adjustment mechanisms remains much alike across periods and price variables: general asymmetric responses, with participation rates acting as main adjustment channel in expansion, unemployment in recession, and a larger long-run employment level in 2008-2012 pointing to an enhanced relevance of the spatial mobility channel when the shock takes place in troublesome times.

Regarding prices, we confirm that price rigidities are not just present at the national level, but are also a regional matter. This was outlined in Jimeno and Bentolila (1998) for a sample period running from 1983 to 1988, and for real and nominal wages. Nominal wages have thus been considered as a main source of price rigidities in Spanish regions. Here, we further extend this result by disentangling the real wage in its two components. One of the contributions of our work, in this context, is the finding that price rigidities cannot be ascribed just to nominal wages, but also to the behaviour of prices.[18]

The implication of this finding is that the product and labour market reforms that have been passed since the mid 1990s have been essentially unsuccessful in increasing price flexibility. This result is complementary to the one claiming that labour adjustments in Spain are mainly achieved by adjusting temporary work in a labour market that is neatly segmented since the early 1990s (Dolado et al., 2002) in contrast to other European economies (Sala et al., 2012). And this result, in addition, helps to explain the virtual lack of labour market and product market price adjustments.

3.7 Cluster analysis and regional specificities

Beyond the differences along business cycle phases, a number of studies point out that Spanish regional unemployment could be regionally clustered into high and low unemployment regions –López-Bazo et al. (2005); Bande et al. (2008); Bande and Karanassou (2009, 2013, 2014); and López-Bazo and Motellón (2013). In this spirit, we next test for the existence of significant differences in the regional adjustment mechanisms in high and low unemployment regions when the analysis

[18]For a detailed discussion on how regional specific variables affect prices see Hubrich et al. (2011).

is performed on regional relative variables (as it is done here, in contrast to the mentioned studies, which focus on absolute regional variables).

3.7.1 Cluster analysis

We follow Bande et al. (2008), and subsequent articles by Bande and Karanassou (2009, 2013, 2014), and proceed in two steps. In the first one we use kernel density function analysis to uncover potential clusters in the Spanish regional relative unemployment rates.[19] In a second step, we perform a k-mean cluster analysis based on exogenous variables to rank each region in one of the groups.

Figure 3.8 depicts the estimated functions for several years. In 1996, most regional rates were slightly below the national unemployment rate, with two small groups in each extreme of the distribution. Then, as the analysis moves forward in time, two groups become clearly distinguishable. A "low unemployment group", where most of the regions take values around 0.8, and a "high unemployment group" with values around 1.25.

These results are similar to the ones obtained by Bande and Karanassou (2009, 2013, 2014). Like them, we cluster the Spanish regions in two groups, and proceed to assign each Spanish region to one of them. A tempting procedure would be to allocate the regions according to their regional unemployment rate, given that Figures 3.2 and 3.8 would provide immediate allocations. This, however, would generate a purely endogenous classification, which we want to avoid. Hence the second step.

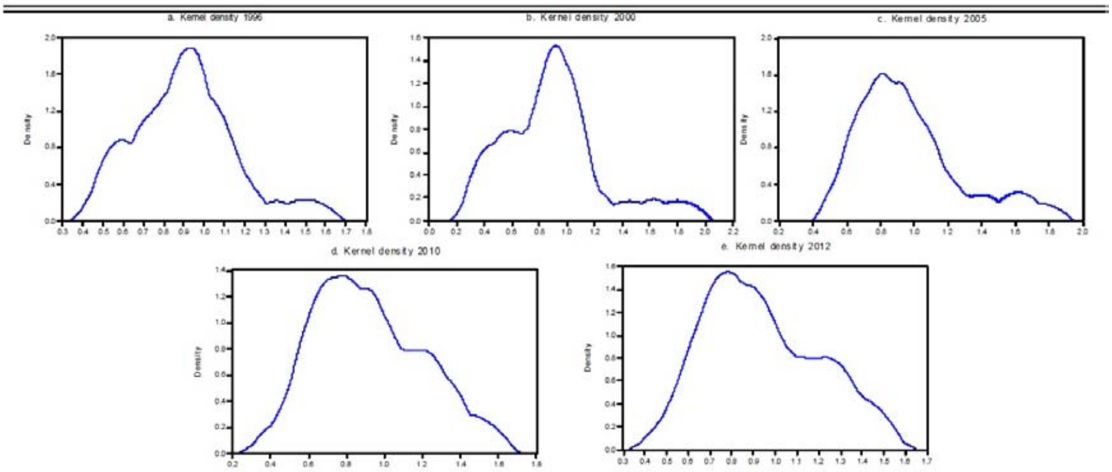
We conduct a cluster analysis following a k-means procedure.[20] This procedure is based on a selection of exogenous variables, which are used to allocate each region in a single group. We follow the literature and choose two variables closely related to regional social welfare: the participation rate and the relative total labour cost.[21] Table 3.8 shows the composition of the two groups and the group averages of the two variables examined.

[19]The regional relative unemployment rates are obtained by dividing each regional unemployment rate over the national one. A unit value, therefore, implies equal regional and national unemployment rates.

[20]For a detailed description of cluster analysis, see Everitt et al. (2001) and Bande et al. (2008).

[21]Although it is frequent to choose relative per capita income, our closest measure at hand to a quarterly and recent regional measure of social welfare is total compensation.

FIGURE 3.8: Kernel density functions: relative unemployment rates



Notes: Densities estimated with an Epanechnikov kernel. Results are robust to different kernel methods.
Source: Spanish Labour Force Survey (EPA).

Group 1 is formed by Andalusia, Aragon, Asturias, Balearic Islands, Canary Islands, Cantabria, Castile-Leon, Castile-La Mancha, the Valencian Community, Extremadura, Galicia, Murcia and La Rioja, whereas Group 2 is formed by just four regions: Catalonia, Madrid Community, Navarre, and the Basque Country. It is worth noting that Group 2 comprises regions with larger participation rates and compensations, and lower relative unemployment rates than the regions in Group 1. Thus, from now on we refer to Group 2 as the “low unemployment group” and to Group 1 as the “high unemployment group”.

TABLE 3.8: Composition of groups from the cluster analysis

Group 1			**Group 2**		
Andalusia	Castile-La Mancha		Catalonia		
Aragon	Valencian Community		Madrid Community		
Asturias	Extremadura		Navarre		
Balearic Islands	Galicia		Basque Country		
Canary Islands	Murcia				
Cantabria	La Rioja				
Castile-Leon					
	Mean	SD		Mean	SD
Activity rate	0.555	0.049		0.595	0.031
Relative total labour costs*	0.906	0.065		1.142	0.051
Relative unemployment rate	0.989	0.301		0.715	0.139

Notes: SD=standard deviation; data sources as explained in Table 3.1. * Analysis restricted to 2000q1-2012q3 due to data availability.

Once the two groups are identified, the next step is to create regional specific variables for each group in which the reference mean is the one of the group in which each region has been allocated. Then, in order to be consistent with our methodology, we follow Broersma and van Dijk (2002) and regress again equations (3.4) to (3.6), but this time considering two average regions: one representing those in Group 1 and another one representing those in Group 2.[22]

3.7.2 Regional specificities

Figure 3.9 shows the IRFs corresponding to the high and low unemployment rate groups.[23]

The first noticeable result is the similarity in the dynamic response of the participation rates and employment levels in both groups of regions to a regional employment shock (the employment response being slightly larger than for the high unemployment regions). The group of high unemployment regions (Group 1) responds in a similar way than the Spanish regions when considered all together, and achieves a long-run, relative regional employment level that is increased by 28% (quite close to the aggregate rise of 27.3%). In turn, the group of low unemployment regions (Group 2) responds in a slightly softer way with a rise of 20.0%.

In addition, even though both groups of regions deliver an almost identically short-run response to the shock, we find unemployment persistence to be larger in the long-run in the low employment regions, where convergence takes around 40 periods (instead of 35 in high unemployment regions).

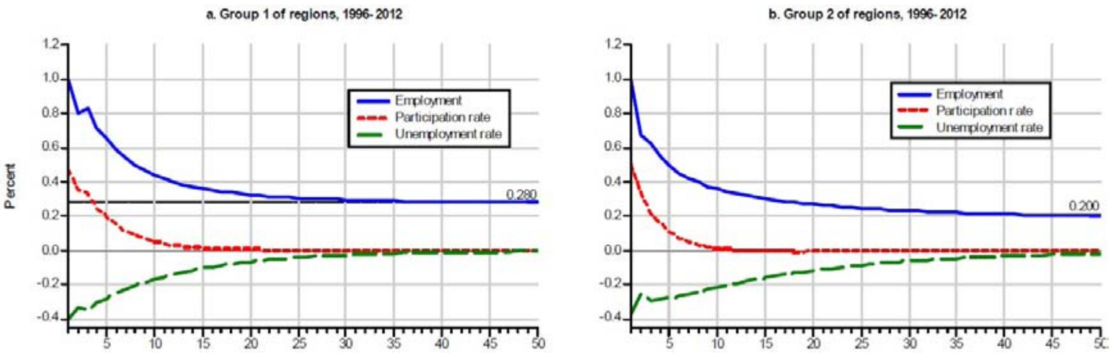
These results have important implications. The fact that both groups share similar dynamic adjustments in terms of participation rates and employment levels, but unemployment rates are more persistent in “low unemployment” regions, implies that spatial mobility is more relevant to explain relative adjustments in the “high unemployment” regions. Consequently, absolute differences between the regional unemployment rates of these groups should be ascribed to either regional structural differences, to nation-wide shocks producing different impacts, or to

[22]Note that we do not combine time and regional disaggregation simultaneously. The reason is the tight amount of degrees of freedom in which this analysis is conducted, which would be problematic in case of estimating equations (3.4) to (3.6) for Group 2 containing only four regions, for the restricted sample period 2008-2012.

[23]To conserve space, the results of the new estimation of equations (3.4) to (3.6) for each group are not reported.

both circumstances. Not, in any case, to different responses to one-off regional employment shocks.

FIGURE 3.9: IRFs to a regional employment shock in Groups 1 and 2 of regions



This result can be checked, numerically, in Table 3.9. While in regions with relative high unemployment rates there is a much stronger tendency to leave and there is large spatial mobility, low employment regions are more resilient, and the shock causes larger adjustments in the unemployment rates.

TABLE 3.9: IRFs decomposition to a 1% regional employment shock in high and low unemployment regions

	Group 1 $(1996-2012)$			Group 2 $(1996-2012)$		
	Participation	Unemployment	Migration	Participation	Unemployment	Migration
Year 1	42%	42%	16%	43%	42%	15%
Year 2	24%	42%	34%	14%	59%	27%
Year 3	11%	38%	52%	2%	60%	38%

Notes: quarterly data aggregated to annual values and normalised by the employment response in the year.

We find this result consistent with the one we have obtained in Section 3.5, when splitting the sample period to distinguish the wild-ride from the steep-fall years. We uncovered a stronger tendency to migrate in bad times which, according to these new set of results, may be connected to the fact that people face a relatively bad period (such as the 2008-2012 one), or to the fact that they live in a region with a relatively high unemployment rate.

3.8 Conclusions

How does the labour market of the average Spanish region respond to an employment shock? Is this response symmetric across business cycle phases? How do prices adjust in response to such shocks? Do regions react alike in spite of their unemployment rate differences? These are the central questions we have tried to answer.

Our aggregate analysis shows that persistence in the Spanish regional labour market is not substantially different today than in 1976-1994, as documented by Jimeno and Bentolila (1998). Although the Spanish labour market may be, on aggregate, much more flexible than it was, there has been little progress in terms of regional unemployment persistence (recall Figure 3.2b). The main difference lies in the larger adjustment, today, through changes in participation rates, although with a lower degree of persistence.

We also find evidence of different responses across business cycle phases. The main mechanism to adjust to a regional shock in expansion is the change in the participation rate, whereas unemployment and migration become the central ones in recession (in the short- and long-run, respectively).

Another finding is that the long-run relative employment level in the aftermath of the shock is higher in 2008-2012 than in 1996-2007. This is an outcome of the growing relevance, in the second period, of the migration adjustment mechanism to regional shocks. Although it is well-known that the Spanish labour market is not characterised by a high degree of regional mobility since the 1970's, our results provide support to the hypothesis that people in a particular region have become more willing to migrate (relative to the national average) when confronted to a shock that takes place in a recessive context than when confronted to an equivalent shock taking place in good times.

We also find that price stickiness is still very strong nowadays (1996-2012), a result that should come as no surprise since this was already documented by Jimeno and Bentolila (1998). Our contribution on this regard is that strong real wage rigidities arise both from nominal wages and consumer prices, and are present both in expansion and recession. This seems to indicate that the product and labour markets are still operating with a substantial degree of imperfect competition, in which case policy measures implemented to foster competition and a larger responsiveness of market prices to the changing (regional) economic environment have, to a large extent, failed to achieve their target.

Disaggregation of the Spanish regions by groups uncovers very similar responses in the short-run, and some divergence as time goes by. In the long-run, high unemployment regions are more reactive in terms of spatial mobility. This reveals a larger propensity to migrate from regions with high than from regions with low unemployment rates. It follows that differences in regional unemployment need to be explained by factors other than regional labour market dynamics. For example, by differences in regional amenities, or different responses to nation-wide shocks.

In any case, we need time to reduce our unemployment rate –participation rate is the main adjustment mechanism in good times–, but we are quick in rising it –unemployment becomes the most important adjustment channel in troublesome times. Given the current difficulties in reducing unemployment, this is an asymmetry that will surely receive attention in future work.

Technical Appendix and supplementary materials

The study of the impact of a region-specific shock requires disentangling the region-specific part of the shock from the common one. For this, variables need to be defined in relative terms, and there are different possibilities. In the Chapter, regional variables have been defined by orthogonalising the regional series with respect to national trends. As shown, this implies taking beta-differences. Although this choice is justified on the grounds that the slope coefficients in equations (3.4)-(3.6) are significantly different from 1 (as shown by Tables 3.4 and 3.6), for the sake of robustness it is worth checking how would the results look like under the alternative procedure of taking log-differences. In this case, relative terms are defined as log-differences between regional and national variables, implying that the impact of the shock needs to be intrepretted as capturing both the reaction to a regional idiosyncratic shock and the asymmetric response to a common shock.

Given this definition, estimation of equations (3.1), (3.2) and (3.3) is now conducted by System GMM, rather than by System OLS, and needs to control for fixed effects. Of course, given that we estimate a PVAR of order 2 where lagged dependent variables appear as explanatory variables, the use of a standard fixed-effects estimation could yield inconsistent coefficients even if the residuals were not serially correlated. To avoid this problem, we follow Love and Zicchino (2006) and use the "Helmert procedure", which consists in forward mean-differencing the variables to remove the fixed-effects (i.e. this procedure removes only the forward mean of each variable in each period, see Arellano and Bover (1995)). The advantage of this procedure is that it keeps the orthogonality between the transformed variables and the lagged regressors. This allows us to use lagged regressors as instruments to estimate the coefficients by System GMM.

Figures 3.A1 to 3.A4 are, respectively, the counterparts of Figures 3.4, 3.5, 3.7 and 3.9 in the article. In turn, Tables 3.A1 to 3.A3 are, respectively, the counterparts of Tables 3.5, 3.7, and 3.9 in the main text.

In spite of the methodological changes surrounding the definition of the relative variables and the estimation procedures, it is worth remarking that the qualitative results conveyed in the article hold and are robust to both definitions of relative terms. In particular, migration is the most important adjustment channel in the second period of analysis, and in the regions with the highest relative unemployment rate. Further, the responses in nominal wages and prices are characteristically low. In contrast, this new set of results reveal larger persistence differences across all regions not matter whether they are considered on aggregate or whether they are classified into high and low unemployment groups. This is mainly due to the fact that taking beta-differences involves controlling for business cycle oscillations whereas, under the log-difference strategy, underlying trends in the data may remain and cause delayed adjustments and larger relative long-run regional employment impacts.

A1. Aggregate IRFs and decomposition to a 1% regional employment shock for period 1996-2012

FIGURE 3.A1: Aggregate IRFs to a regional employment shock. 1996-2012

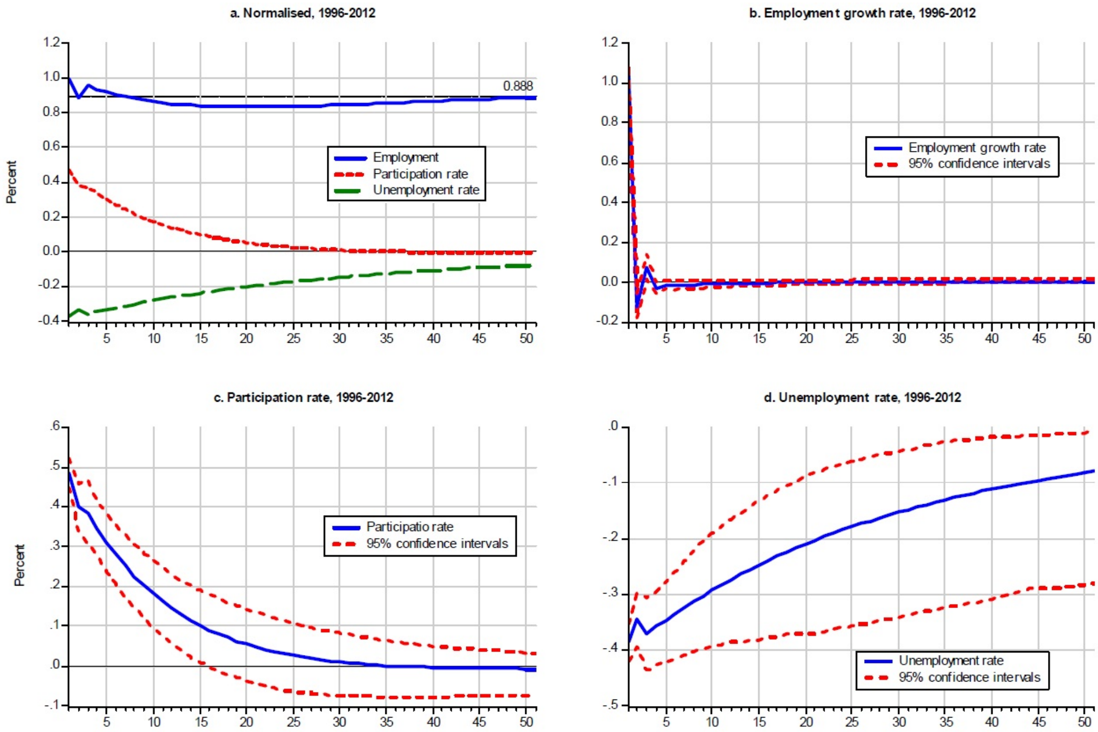


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TABLE 3.A1: IRFs decomposition to a 1% regional employment shock. 1996-2012

	Participation	Unemployment	Migration
Year 1	41%	37%	22%
Year 2	29%	35%	36%
Year 3	19%	32%	49%

Notes: quarterly data aggregated to annual values and normalised by the employment response in the year.

A2. Aggregate IRFs and decomposition to a 1% regional employment shock for periods 1996-2007 and 2008-2012.

FIGURE 3.A2: IRFs to a regional employment shock in 1996-2007 and 2008-2012

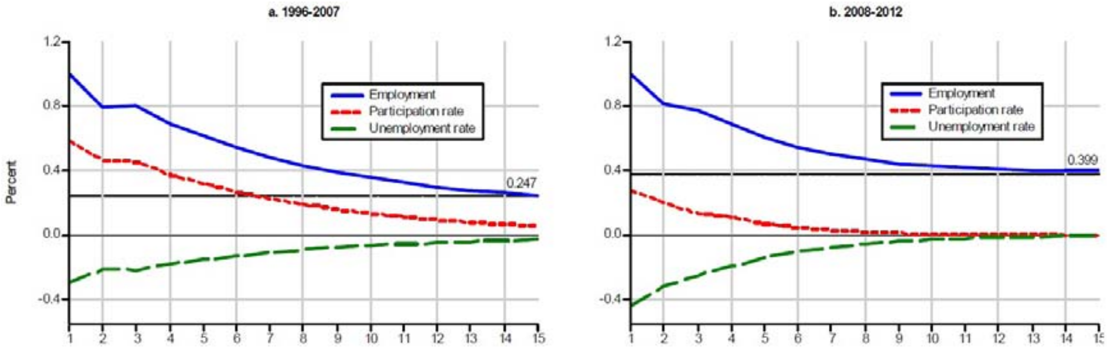
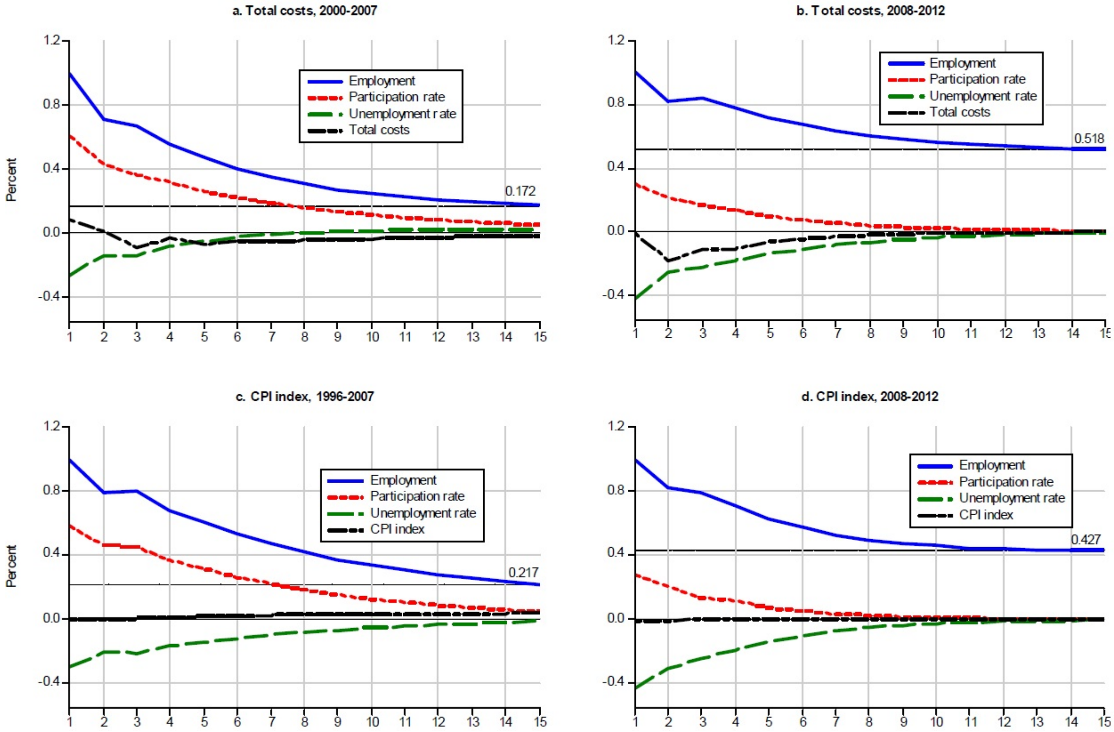


TABLE 3.A2: IRFs decomposition to a 1% regional employment shock

	1996-2007	**2008-2012**
Final employment effect:	0.25%	0.40%
Adjustment in 1st quarter by:		
Participation	58.3%	27.4%
Unemployment	29.9%	43.5%
Migration	11.8%	29.1%
Cumulative adjustment in 15th quarter by:		
Participation	47.4%	11.0%
Unemployment	23.1%	20.4%
Migration	29.5%	68.6%

A3. Aggregate IRFs and decomposition to a regional employment shock in total costs and product prices.

FIGURE 3.A3: IRFs to a regional employment shock in total costs and product prices



A4. Decomposition to a 1% regional employment shock for high and low unemployment regions.

FIGURE 3.A4: IRFs to a regional employment shock in Groups 1 and 2 of regions

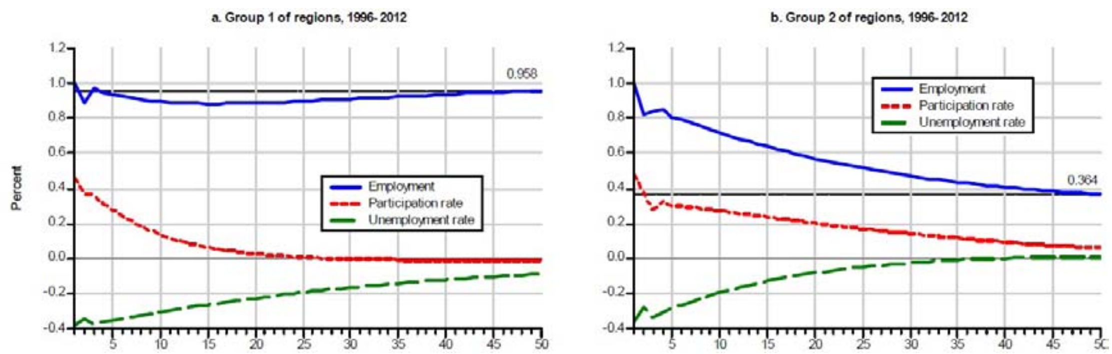


Table 3.A3: IRFs decomposition to a 1% regional employment shock in high and low unemployment regions. 1996-2012

	Group 1			**Group 2**		
	Participation	Unemployment	Migration	Participation	Unemployment	Migration
Year 1	40%	38%	22%	42%	36%	22%
Year 2	25%	37%	39%	37%	32%	30%
Year 3	14%	34%	52%	38%	26%	36%

Notes: quarterly data aggregated to annual values and normalised by the employment response in the year.

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