

# Structural changes in the Spanish labour demand: Does Rodrik's conjecture hold?\*

**Hector Sala**

Universitat Autònoma de Barcelona  
and IZA<sup>†</sup>

**Pedro Trivín**

Universitat Autònoma de Barcelona<sup>‡</sup>

17 February 2012

## Abstract

Rodrik's (1997) conjecture states that the labour demand elasticity with respect to wages should increase with globalization. We estimate this elasticity in Spain, and investigate whether the structural change experienced in the mid 1980s significantly altered the employment response to wage changes. This structural change is related to Spain's European integration, and the intensive internationalization and liberalization processes that followed. Hamermesh's (1993) framework is used to identify the total employment effect, which is made of the constant-output elasticity and the scale effect. This total effect is empirically quantified and its two components disentangled. We find that the total employment elasticity of a wage change rose from -0.7 in 1964-1984 to -1.3 in 1985-2007 in the new scenario of increased international exposure and progressive labour market liberalization. By components, the constant-output elasticity rose from -0.2 to -0.4, and the scale effect from -0.5 to -0.9. We also find that the elasticity of substitution between capital and labour moved from 0.9 to unit. Our results provide support to Rodrik's conjecture.

**JEL Codes:** F16, E24, J23.

**Keywords:** Globalization, Employment, Labour demand, Structural change.

---

\***Acknowledgments:** Hector Sala acknowledges financial support from the Spanish Ministry of Science and Innovation through grant ECO2009-07636.

<sup>†</sup>Departament d'Economia Aplicada, Universitat Autònoma de Barcelona, 08193 Bellaterra, Spain; tel.: + 34 93 5812779; email: hector.sala@uab.es.

<sup>‡</sup>Departament d'Economia Aplicada, Universitat Autònoma de Barcelona, 08193 Bellaterra, Spain; tel.: + 34 93 5811680; email: ptg66@msn.com.

# 1 Introduction

In the context of increasing globalization, labour outcomes and international trade patterns have become progressively intertwined. Labour market responses to policy measures have changed, and a growing body of literature has become interested in the interrelation between globalization, (un)employment, and labour market reforms.<sup>1</sup> One way of studying this interrelation is to focus on the labour demand elasticity with respect to wages and investigate whether structural changes in international exposure and legislation alter this elasticity. This is the route recently followed by Hijzen and Swaim (2010) who consider a large panel of 11 OECD countries and use industry-level data.

In this paper we follow this same route, but we specifically focus on a single country, Spain, which experienced a major structural change in the mid 1980s. This change is related to the most important labour market reform undertaken in the last four decades, which was passed in 1984 in view of accession to the European Economic Community (EEC) in January 1986. Our target is to evaluate how this structural change affected the elasticity of employment with respect to wages, and examine the resulting policy implications. On the road towards this target, we also provide estimates of the elasticity of substitution between capital and labour.

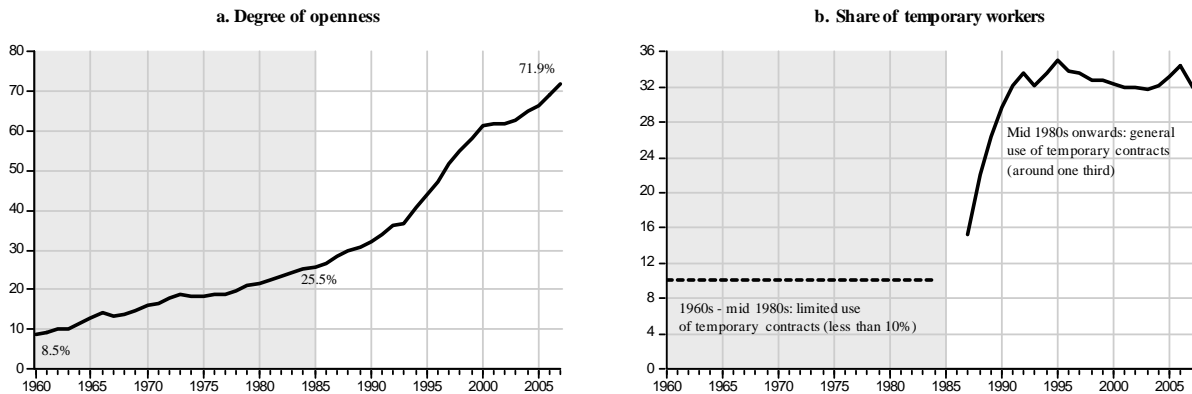
Spain provides a salient example of transition from a strictly closed economy to a wide-open and economically integrated country. The degree of openness –measured as the sum of exports and imports over output– evolved from around 10% in the early 1960s to 25% in the mid 1980s, and almost tripled afterwards, when it reached 72% in 2007 at the pick of last economic expansion. This rapid acceleration in the opening process was the consequence of a major institutional change brought by the 1986 accession to the EEC, and was followed (and enhanced) by further European commitments such as joining the European Monetary System (EMS) in 1989, achieving the Common Market in 1993, and becoming a member of the European Monetary Union (EMU) in 1999 (for details see Polo and Sancho, 1993; and Juselius and Toro, 2005). At the same time, Spain is a key example of structural change in the labour market. It also took place in the mid 1980s when a labour market reform was passed in 1984 to enhance the flexibility of the labour market. This enhanced flexibility was achieved by allowing a general use of temporary contracts, which at the time were essentially used in seasonal activities (mainly agriculture and tourism), and represented less than 10% of the total existing contracts.

---

<sup>1</sup>For example, Bentolila and Saint-Paul (1993) study how the globalisation process has affected the declining labour shares through the resulting changes in import prices and capital-augmenting technological progress. Senses (2010), in turn, studies how phenomena like offshoring and outsourcing are affecting employment, while other studies focus on the cyclical influence of foreign direct investment on the volatility of the labour market (Azariadis and Pissarides, 2007).

The result of this institutional change was a segmentation of the Spanish labour market that is well documented in the literature (see Dolado *et al.* 2002, among others). Figure 1 provides an illustration of these two crucial developments.

**Figure 1. Internationalization and labour market segmentation in Spain**



Source: Own calculation based on data from the OECD Economic Outlook and the Spanish Labour Force Survey (EPA).

Rodrik (1997) was the first one to conjecture on the labour demand consequences of the globalization process. He expected that this process would tend to increase the sensitivity of employment with respect to wages because of lower entry barriers in new markets, increased competition, and the emergence of new phenomena such as offshoring and outsourcing. He was even explicit about the likely channels through which globalization would directly affect the labour market. Among others, he referred to the amplified labour demand responses to global shocks, to shocks on non-labour costs (such as payroll taxes), and to the weakening of the workers' bargaining power.

The evidence on whether labour demand has become more elastic as globalization has deepened is not conclusive so far. The first study to explore this issue is Slaughter (2001). He considers skilled and non-skilled US manufacturing workers and shows, for different periods between 1961 and 1991, that the labour demand elasticities of the non-skilled have continuously increased. He finds difficulties, however, in ascribing these upward sensitivities to a concrete driving force (international trade among them). Krishna *et al.* (2001) focus on the dramatic trade liberalization process experienced by Turkey, but are unable to find significant effects on the labour demand elasticity. In the more general study of Bruno *et al.* (2004) evidence linking globalization and larger labour demand elasticities is only found for the UK out of seven OECD countries (Spain among them). Related evidence is provided in Molnar *et al.* (2007) by studying a panel of 11 OECD countries and disclosing a positive connection in the manufacturing sector between larger flows of foreign direct investment and a larger labour demand elasticity (this connection,

however, is much weaker or even negative in services). In turn, Hijzen and Swaim (2010) associate the rise in this elasticity in a large number of OECD countries to the growing use of offshoring practices, even though they find this positive relationship weaker the more strict the employment protection legislation is. This positive relationship is also verified in Senses (2010) for the US manufacturing sector.<sup>2</sup>

In this paper, we follow Slaughter (2001) and Hijzen and Swaim (2010), and resort to Hamermesh's (1993) model as theoretical framework for our empirical analysis. In terms of our work, the crucial feature of Hamermesh's (1993) model is the decomposition of the total employment effect of a wage change on a constant-output elasticity and a scale effect. Our contribution, therefore, is mainly empirical and has a twofold dimension.

First, we document a change in the elasticity of the demand for labour in response to the internationalization and liberalization of the Spanish economy. For this purpose we estimate two different panel data models covering periods 1964-1984 and 1985-2007, and accounting for the structural change documented in Figure 1. We find that the estimated long-run elasticities of labour demand (namely the total and the constant-output elasticities) have increased substantially in parallel with the globalization process. In particular, the total employment elasticity of a wage change rose from -0.7 in 1964-1984 to -1.3 in 1985-2007, while the constant-output elasticity increased from -0.2 to -0.4.

A second main contribution of this paper is the empirical computation of the scale effect and the evaluation of its change between our two reference periods. First of all, we find the magnitude of the scale effect to be at least twice the size of the constant-output elasticity no matter the period examined. In addition, we find that the scale effect has also shifted with the Spanish opening process, from -0.5 in 1964-1984 to -0.9 in 1985-2007. Since the scale effect reflects the net influence (of opposite signs) of product-supply and product-demand influences, we argue that not only the firms' sensitivity with respect to factor prices is larger than the consumers' sensitivity, but also that it has tended to grow across periods. These findings give empirical support to Rodrik's conjecture as the growing international exposure of the Spanish economy has clearly enhanced the sensitivity of employment with respect to wages.

The rest of the paper is structured as follows. Section 2 provides simple theoretical

---

<sup>2</sup>There is also a vast literature focusing on the elasticity of labour demand *per se*. Drazen *et al.* (1984), for example, find evidence that it varies with product demand, while Lawrence and Lawrence (1985) relate the small labour demand elasticity in the US steel industry to the fall of this industry on account of the high wage claims of its workers. For the whole US manufacturing industry, Revenga (1992) finds larger adjustments in quantities (i.e., employment) than in prices (i.e., wages) in response to shocks reflecting higher international competition. Along the same lines, Borjas and Ramey (1995) find that foreign competition reduces firm product-market power and thus labour rents. Benito and Hernando (2003), in turn, examine the Spanish labour market in 1985-2001 and find a larger labour demand elasticity for temporary than for permanent workers.

fundamentals to ensure a correct interpretation of the estimated elasticities. Section 3 discusses the data and the empirical strategy. Section 4 presents the empirical results. Section 5 concludes.

## 2 The elasticity of demand for labour

### 2.1 Theoretical underpinnings

As explained in Hamermesh (1993), the labour demand elasticity with respect to real wages ( $\eta_{LL}$ ) can be decomposed into a constant-output elasticity,  $-(1 - s_L)\sigma$ , and a scale effect,  $s_L\eta$ , so that:

$$\eta_{LL} = -(1 - s_L)\sigma - s_L\eta, \tag{1}$$

where  $s_L$  is the labour share over total income,  $\sigma$  is the elasticity of substitution between capital and labour, and  $\eta$  is the elasticity of employment with respect to output. The substitution effect reflects the extent to which a firm substitutes away from labour when faced with an increase in its price. In turn, the scale effect represents the reduction in employment due to the reduction in output holding production technology constant. Therefore, the short-run response to a wage change is fully captured by the scale effect (since the production function cannot be altered), whereas in the long-run the constant-output elasticity adds to the scale effect so that the total effect becomes the relevant measure of the impact.

A very important issue for our analysis is the fact that the relative role played by the constant-output elasticity and the scale effect depends on the a-priori modelling assumptions on how real wages move. This caused a hot debate between Dowrick and Wells (2004) and Lewis and McDonald (2002, 2004).

Dowrick and Wells (2004) develop a model in which prices are assumed to automatically and fully reflect all changes in unit labour costs resulting from variations in nominal wages. This is the mechanism by which real wages change (since changes in unit labour costs are not the exact counterpart of changes in nominal wages). This assumption is criticized by Lewis and McDonald's (2004) because of its arbitrariness. In Lewis and McDonald's model (2002, 2004) the equivalent assumption is that prices are exogenous (i.e., constant) as in perfect competitive markets. This is consistent with Hamermesh's (1993) framework and implies that the relevant variable is the real wage as a whole (and not the interplay between nominal wages and prices). In this setting, real wages may change in response to factors other than prices such as, for example, labour supply shifts or institutional variables like social security benefits (Lewis and McDonald, 2004).

What may be misleading is that Dowrick and Wells' (2004) framework yields a different

total effect on employment of changes in real wages, now denoted as  $\eta'_{LL}$ :

$$\eta'_{LL} = -\sigma - \eta \frac{s_L}{1 - s_L}. \quad (2)$$

Note that equations (1) and (2) clearly show the different quantitative role that the constant-output elasticity and the scale effect may play. Here we stick to Hamermesh (1993) analytical framework and drive our efforts to the empirical quantification of the total effect components.

Another important issue (to which we will return below) is the fact that the scale effect is the net effect of demand-side and supply-side product market influences. To preview this issue, it is useful to recall that, as shown in microfounded models behind the labour demand (see Karanassou *et al.*, 2007, for example), the labour demand curve is obtained by equating marginal costs and revenues (that is, by taking first order conditions from a standard profit maximization or cost minimization problem) *at those points where product-supply and product-demand intersect*. This is the reason why the scale effect should not be interpreted as a pure supply-side outcome, but as the net effect of demand-side and supply-side forces. As noted, this point will deserve further discussion below.

## 2.2 Empirical implementation

The empirical estimation of the constant-output elasticity and the scale effect has followed different routes. Some studies –Russell and Tease (1991), Lewis and MacDonald (2002), Bruno *et al.* (2004)– focus on a single equation in which the central estimates are the substitution ( $\sigma$ ) and the scale ( $\eta s_L$ ) effects. The standard equation in this first group of studies is:

$$n_t = \beta_0 + \beta_1 y_t + \beta_2 w_t + \beta_3 t + \varepsilon_{1t}, \quad (3)$$

and allows to estimate the constant-output elasticity. The total effect is then computed using equation (1).

A second group –Slaughter (2001), Hijzen and Swaim (2010)– estimate two different equations aiming at the individual identification of the total and the substitution effects. Therefore, on top of equation (3), the following equation yielding a direct estimate of the total effect is also considered:

$$n_t = \alpha_0 + \alpha_1 k_t + \alpha_2 w_t + \alpha_3 t + \varepsilon_{2t}. \quad (4)$$

One of the contributions of this paper lies in the methodology to empirically approach the scale effect not as the first group of works, but as the difference between the total

and the constant-output elasticity, which are obtained, respectively, from the estimation of equations (3) and (4).<sup>3</sup> Although Slaughter (2001) and Hijzen and Swaim (2010) also consider and estimate both equations, they do not focus on the scale effect. Here, in contrast, we are interested in empirically disentangling the two components of the total effect. As we will see, the scale effect will turn out to be quantitatively dominant.

Before explaining in detail our empirical procedure, let us review some simple theoretical endorsement to these equations.

### 2.2.1 Simple background for equation (3)

In his analysis of the US production function, Antràs (2004) departs from a CES production function of the form:

$$Y_t = A_t \left[ \delta K_t^{\frac{\sigma-1}{\sigma}} + (1 - \delta) N_t^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}},$$

where  $N$  is employment,  $K$  is capital stock,  $A$  accounts for Hicks-neutral technological change,  $\delta$  is a distribution parameter, and  $\sigma$  is the elasticity of substitution between capital and labour. Solving the first order condition with respect to labour and taking logs yields the following labour demand function (equation 2 in Antràs, 2004, p. 5):

$$\log(F_t/N_t) = b + \sigma \log W_t, \tag{5}$$

where  $F_t \equiv Y_t/A_t$  is the aggregate input function. Under the assumption of Hicks-neutral technological change  $F_t$  is independent of  $A_t$  (see Berndt, 1976) and allows the construction of an aggregate input index.<sup>4</sup> Antràs (2004), however, dismisses Berndt's (1976) procedure to obtain an empirical proxy of  $F_t$  because it cannot be computed in the presence of biased technological change. This leads him to replace  $F_t$  by real output  $Y_t$ . Therefore, the effective empirical counterpart of equation (5) is  $\log(Y_t/N_t) = b + \sigma \log W_t$  which, extended with an error term  $\varepsilon_{1t}$ , can be rewritten as equation (3).

### 2.2.2 Simple background for equation (4)

The Cobb-Douglas production function is  $Y = AN^{1-\alpha}K^\alpha$ , where  $\alpha$  ( $0 < \alpha < 1$ ) is a parameter accounting for the relative influence of capital and employment on output.

---

<sup>3</sup>Note that there is a direct relationship between the parameters in Hamermesh (1993) and the coefficients in equations (3) and (4). The elasticity of substitution is  $\sigma = \alpha_1 \approx -\beta_2 = \sigma$ ; the constant-output elasticity is  $-(1 - s_L)\sigma = \alpha_1(1 - s_L) \approx -\beta_2(1 - s_L) = -(1 - s_L)\sigma$ ; and the scale effect is  $\eta s_L = \beta_1 s_L$ . (The sign " $\approx$ " is used to make explicit that this equivalence is an empirical issue, to which we return below).

<sup>4</sup>This index is defined as  $(r_t K_t + W_t N_t)/P_t$ , where  $P_t$  is a Tornqvist price index of the rental prices of capital and labour (see Antràs, 2004, p. 9).

The first-order condition of the maximization problem with respect to labour,  $\frac{\partial \Pi}{\partial N} = 0$ , gives

$$N_t = [(1 - \alpha) A_t]^{1/\alpha} K_t W_t^{-1/\alpha}. \quad (6)$$

Assuming, as standard (see, for example, Antràs, 2004), that technological change increases at a constant growth rate so that  $A_t = A_0 e^{\lambda t}$  (where  $\lambda$  is the growth rate of technological change), and taking logarithms, we can rewrite equation (6) as:

$$\ln N_t = \frac{1}{\alpha} \ln(1 - \alpha) + \frac{1}{\alpha} (\ln A_0 + \lambda t) + \ln K_t - \frac{1}{\alpha} \ln W_t. \quad (7)$$

By writing  $n = \ln N$ ,  $k = \ln K$ ,  $w = \ln W$ ,  $\alpha_0 = 1/\alpha \ln(1 - \alpha + a_0)$  where  $a_0 = \ln A_0$ ,  $\alpha_1 = 1$ ,  $\alpha_2 = 1/\alpha$ , and  $\alpha_3 = \lambda$ , and adding  $\varepsilon_{2t}$  as the residual we obtain equation (4) as the empirical counterpart of equation (7).

Some authors consider extensions of equations (3) and (4) by adding price controls such as, for example, the price of capital. Clark and Freeman (1980), however, showed that labour demand elasticities tend to be biased upwards when the prices of other production factors –which are more subject to errors of measurement than the corresponding quantities– are considered in these type of equations. They suggested not imposing such extra price controls and this has been the route followed by the literature in the field. We also follow Clark and Freeman (1980) and do not consider the addition of other price variables. However, given our objective of testing the extent to which the globalization process has affected the labour demand elasticities, we find it critical to control for the growing degree of economic openness. This is the only extension of equations (3) and (4) we will consider in our empirical analysis.

### 2.3 Interpreting the crucial elasticities

When interpreting the resulting elasticities, the crucial coefficients are  $\alpha_2$  in equation (4) and  $\beta_2$  in equation (3). The coefficient  $\alpha_2$  measures the total elasticity of employment with respect to the real wage. It measures the overall effect of a change in wages on the level of employment, which is the sum of two components: (i) the constant-output elasticity; and (ii) the scale effect (in turn, the scale effect is the outcome of a product-supply effect and a product-demand effect). The coefficient  $\beta_2$  measures the substitution effect resulting from the change in the relative factor prices. When multiplied by the capital income share  $(1 - s_L)$ , the substitution effect becomes the constant-output elasticity:  $-(1 - s_L)\beta_2$ .

When wages fall, the relative price of labour *vis-à-vis* the price of capital is reduced and there is an incentive to substitute capital by labour. The extent to which this is feasible (which depends on the technology) is measured by  $\beta_2$ . In other words,  $\beta_2$  is



the substitution effect because it measures the employment response to a wage change. When it is multiplied by the capital share, it still measures the employment response to a wage change but *holding output constant*. The intuition behind this definition is that the firm can only substitute the existing amount of labour, hence the need to weight the substitution effect by the negative of the labour share (or capital share).

As noted, the scale effect has two components. On the one hand, the product-supply effect results from the fact that lower costs (in our example wages have fallen) allow an increase in production. This is a direct consequence of the fact that the labour demand derives from the production function. *Ceteris paribus*, the less costly becomes a production factor, the more output will be supplied, and the more labour needed. On the other hand, there is also a product-demand effect arising from the fact that wages are directly connected with consumption and, thus, with aggregate demand in the product market. All things constant, lower wages will diminish consumption and will cause an indirect fall in employment on account of the lower production required in the economy.

Our empirical analysis allows us to identify the total and the substitution effects ( $\alpha_2$  and  $\beta_2$ , respectively). The substitution effect can then be used to obtain the constant-output elasticity, and the difference between the total effect and the constant-output elasticity yields an empirical measure of the scale effect.

$$\text{Total effect} = \text{Constant-output elasticity} + \text{Scale effect} . \quad (8)$$

$$\underset{[\alpha_2]}{\quad} \quad \underset{[(1-s_L)\beta_2]}{\quad} \quad \underset{[=\alpha_2-(1-s_L)\beta_2]}{\quad}$$

The scale effect is by definition (recall equation (1) from in Hamermesh's (1993) model) the product of the labour share and labour demand elasticity with respect to output, we could expect the scale effect to decrease in response to labour share declines (as witnessed in recent decades), or to increase in response to rises in the elasticity of the demand for labour with respect to output arising, for example, from the globalization process. Because of the opposite direction exerted by these influences, their net impact is undetermined. To open this black-box empirically is one of the contributions of this paper.

Another crucial piece of information when interpreting the elasticities is the expected consistency between the estimates of  $\beta_2$  and  $\alpha_1$ . Recall the  $\beta_2$  is the elasticity of substitution between capital and labour, and observe that  $\alpha_1$  is the elasticity of employment with respect to capital. Although both estimates will be empirically different, we should expect them to be close from one another, of course with the opposite sign. This would be an indication of faithful results given that the selected empirical specifications of equations (3) and (4) could be the source of potentially significant differences.

The interpretation of the estimated coefficients  $\alpha_3$  and  $\beta_3$  is standard and reflects the influence of constant technological change.

### 3 Data and empirical modelling

#### 3.1 Data

Our main source database is the BD-MORES database supplied by the Spanish Ministry of Economy and Competitiveness. Our data is taken from its last version, made available in December 2011, with data going up to 2007 which is the last year before the Great Depression.<sup>5</sup> From this source we obtain data on our main set of variables: output, net capital stock, total employment, and average real wages. It is important to state that average real wages are computed as total worker’s compensation –a variable directly supplied by the database that includes self-employment compensation– over total employment (i.e., including dependant and self-employees). This implies that our constructed measure of the labour share is adjusted for self-employment income. To measure the degree of openness we take data from the OECD Economic Outlook.

Overall, we use annual data covering years 1964-2007. Definitions of the variables used are provided in Table 1.

**Table 1. Definitions of variables.**

	Variables	Sources	Subindices
$y_{ijt}$	Real GDP at market prices	BD Mores	$i = 1, \dots, 14$ sectors
$k_{ijt}$	Net real capital stock	BD Mores	$j = 1, \dots, 17$ regions
$n_{ijt}$	Total employment	BD Mores	$t = 1, \dots, 44$ years
$w_{ijt}$	Average real wage	BD Mores	
$ls_{it}$	Labour share $\left( = \frac{w_{ijt} * n_{ijt}}{y_{ijt}} \right)$		
$op_t$	Openness $\left( \frac{\text{exports} + \text{imports}}{\text{output}} \right)$	OECD Economic Outlook	
$t$	Linear time-trend	Constructed	

Note: These variables (but  $t$ ) are expressed in logs in the estimation process.

The main advantage of using BD-MORES data is its extensive disaggregation across sectors (14 non-agricultural sectors are considered)<sup>6</sup> and regions (17, because the two autonomous cities in Africa, Ceuta and Melilla, are not considered). Hence, we have homogeneous long time-series covering more than four decades of the Spanish recent economic history widely disaggregated allowing us to work with thousands of observations.

<sup>5</sup>A detailed description of this database is available in Dabán *et al.* (2002).

<sup>6</sup>(1) Mining and quarrying; (2) Food products; (3) Textiles; (4) Paper products; (5) Chemical products; (6) Non-metallic mineral products; (7) Metal products and machinery; (8) Transport equipment; (9) Other manufactured products; (10) Construction; (11) Wholesale and retail trade; (12) Financial intermediation; (13) Other market services (16) Non-market services.

In this way we are able use panel data techniques and estimate the relevant equations for our two periods of interest without worrying about the degrees of freedom. In particular, we work with two panel models corresponding to periods 1964-1984 and 1985-2007. We have a three-dimensional panel made of 14 sectors, 17 regions, and, respectively, 20 and 24 years per period. Therefore, once organized to conduct the estimation, the structure of the resulting bi-dimensional panel yields, for period 1, a time dimension of  $T = 340$  observations and a cross-section dimension of  $N = 14$ ; and, for period 2, a time dimension of  $T = 408$  observations and a cross-section dimension of  $N = 14$ .

Table 2 provides descriptive information on some crucial variables of interest. This information corresponds to aggregate averages of the Spanish economy for the relevant periods of analysis: the closed-economy period of 1964-1984, and the liberalization/economic integration period of 1985-2007.

**Table 2. Macroeconomic scenario in Spain, 1964-2007.**

	$\Delta Y$	$\Delta K$	$\Delta N$	$\Delta \frac{K}{Y}$	$\Delta \frac{Y}{N}$	$\Delta W$	$W/(Y/N)$
1964-2007	3.44	4.32	1.30	0.86	2.14	2.79	69.6
1964-1984	3.60	4.95	-0.14	1.33	3.73	5.02	76.6
1985-2007	3.32	3.77	2.56	0.45	0.75	0.76	62.7

Note:  $\Delta$  is the difference operator and indicates average growth rates. All variables are expressed in percent.

Real economic growth in Spain was around 3.5% on average since the mid 1960s until 2007, with no significant differences between periods. The sources of growth, however, differed markedly. The 1964-1984 years were mainly characterized by a shift from an agricultural to an industrialized-based economy followed, in 1977-1984, by a deep and prolonged crisis mainly affecting those industrial sectors that had previously led the expansionary period. The industrialization process, and subsequent crisis, resulted in an overall absence of net job creation (-0.14% on average), but very high growth rates of capital stock (close to 5%) and labour productivity (3.7%). Note that wage growth (5%) surpassed labour productivity growth resulting in large values of the labour income share of around 77% (recall that self-employment rents are considered as labour rents). This situation of growing labour share or, equivalently, rising unit labour costs, is consistent with the absence of net employment growth. In contrast, in the context of growing openness and the enhanced market competition that characterized the subsequent period, the labour market reform of 1984 and the resulting boom in temporary jobs prevented the unit labour cost to deteriorate, and allowed employment creation. Thus, in 1985-2007 employment rose more than 2.5% on average, in parallel with a mild increase in real wages

and labour productivity (around 0.75%). In turn, there was a deceleration in capital accumulation (3.8%), and the growth rate in the capital deepening ratio  $K/Y$  only reached a third than before.

## 3.2 Empirical strategy

Given the panel structure of our database, we will use panel data techniques applied to the estimation of dynamic one-way fixed-effect models (also known as Least Squares Dummy Variables, LSDV, models). Such techniques involve a number of interesting and important issues related to the use of stationary panel data (in case of estimating dynamic models), the use of homogeneous (i.e., having different intercept but the same slope for the different cross-section unit) or heterogeneous panels, and the treatment of endogenous variables. We next deal with these issues.

### 3.2.1 Dynamic panel data estimation

Due to the relevance of adjustment costs in labour demand decisions, equations (3) and (4) will be estimated as autoregressive distributed lag (ARDL) models of the following general form:

$$n_{ijt} = \alpha_i + \lambda \sum_1^S n_{ijt-s} + \beta \mathbf{X}_{ijt} + \beta \sum_1^V \mathbf{X}_{ijt-v} + u_{ijt}, \quad u_{ijt} \sim i.i.d.N(0, \sigma_{ij}^2), \quad (9)$$

where  $i$  denotes sector,  $j$  denotes region,  $t$  is the time index,  $s$  and  $v$  reflect the dynamic structure of the model,  $n$  is the dependent variable,  $\alpha$  is a sectoral cross-section intercept,  $\lambda$  is the persistence coefficient,  $\beta$  is a set of parameters reflecting the influence of the explanatory variables contained in vector  $\mathbf{X}$ , and  $u_{ijt}$  is the residual.

The choice of this specification is due to the availability of a large number of observations, which ensures enough sectoral and regional variation, and allows a robust estimation of equations (3) and (4). We will thus estimate specification (9) as a dynamic homogeneous fixed-effect model.

Regarding the dynamic structure of our estimated equations, the first issue to deal with is the well-known potential correlation problem between the lagged dependent variable and the error term, which is likely to bias the estimates. This is known as the Nickell bias (Nickell, 1981) and, provided  $T < N$ , results in inconsistent OLS estimates even when the error term is not serially correlated. The condition that  $T < N$  was the general case when the existing databases were still short in their time dimension. However, whenever  $N \rightarrow \infty$ ,  $T \rightarrow \infty$ , and  $T$  grows sufficiently fast relative to  $N$ , the OLS estimates will be consistent. This was shown by Álvarez and Arellano (2003) and it is relevant for us

because we work with a panel database with large  $T$ , large  $N$ , and  $T > N$ . This enables us to estimate a dynamic one-way fixed-effects model.

In a fixed-effect model estimation of equations (3) and (4), wages and capital stock (or output) are considered exogenous. Thus, for comparison purposes we estimate the GMM counterpart of these equations where  $w$  and  $k$  (or  $y$ ) are still treated as exogenous. This allows us to see whether there are significant differences in the estimated coefficients. Given that  $w$  and  $k$  (or  $y$ ) are treated equally, such differences can be ascribed to differences in the econometric methodology. The results presented in tables 4 to 7 reveal that this is indeed the case. Moreover, we experimented with different sets of instruments and found the elasticity of employment with respect to wages sensitive both in equations (3) and (4). Sensitivity to the choice of instruments is a well-known problem in GMM estimation (see, among others, Arellano, 2003). Thus, given that it is likely that  $w$  and  $k$  (or  $y$ ) could give rise to endogeneity problems in our estimation, we decided to instrument  $n$ ,  $w$  and  $k$  (or  $y$ ) and estimate the fixed-effects model using 3 Stages Least Squares. The third column in Tables 4 and 6 show these results which we take as the reference ones because they combine characteristics of dynamic panel data estimation and endogeneity control.

### 3.2.2 Unit Roots Test

One of the challenges of estimating dynamic panel data models is a correct specification of the long-run relationships between the variables. In order to check if it is appropriate to use stationary panel data estimation techniques, we conduct a series of unit root test.

These tests are different depending on the type of variables to be dealt with. Therefore, we use the KPSS unit root test<sup>7</sup> to test for the order of integration of the variable that is common across sectors and regions ( $op_t$ ), while for the variables that are sectoral and regional specific we use the simple statistic test proposed by Maddala and Wu (1999),<sup>8</sup> which is an exact nonparametric test based on Fisher (1932):<sup>9</sup>

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

where  $p_i$  is the probability value of the ADF unit root test for the  $i$ th unit (sector). This test has the following attractive characteristics: (i) it does not restrict the autoregressive

---

<sup>7</sup>See Kwiatkowski, Phillips, Schmidt and Shin (1992) for details.

<sup>8</sup>The use of pooled data can generate more powerful unit root tests than the popular Dickey-Fuller (DF), Augmented DF and Phillips-Perron (PP) tests.

<sup>9</sup>Maddala and Wu (1999), using Monte Carlo simulations, conclude that the Fisher test outperforms both the Levin and Lin (1993) and the Im, Pesaran and Shin (2003) tests.

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 sector.

**Table 3. Panel Unit Root Tests.**

---



---


$$\zeta(n_{it}) = 80.22 \quad \zeta(w_{it}) = 468.22 \quad \zeta(k_{it}) = 165.59 \quad \zeta(y_{it}) = 183.81 \quad \xi(op_t) = 0.15$$


---

Notes:  $\zeta(\cdot)$  is the test proposed by Maddala and Wu (1999). Its 5% critical value is approximately 51. In turn,  $\xi(\cdot)$  is the KPSS <sub>$c,t$</sub>  test using intercept and trend. Its 5% critical value is approximately 0.15.

The results of this test, displayed in Table 3, indicate that we can indeed proceed with stationary panel data estimation techniques.

### 3.2.3 Homogeneous versus heterogeneous models

The estimation of a homogeneous panel allows specific intercepts but restricts the different cross-section units to have common slopes. This restriction may be quite demanding when dealing with different countries –or sectors– and presumably different elasticities. To explain how would the potential bias from sectoral heterogeneity affect our estimates, let us assume (for simplicity and without loss of generality) that our general form specification (9) follows an ARDL(1,0) such as:

$$n_{ijt} = \alpha_i + \lambda n_{ijt-1} + \beta \mathbf{x}_{ijt} + u_{ijt}.$$

As explained by Pesaran and Smith (1995) and Pesaran *et al.* (1996) the heterogeneity bias may arise when a heterogeneous panel, as ours, is estimated as a homogeneous one. To the extent that our variables tend to a degree of stationarity close to  $I(1)$  without fully reaching a degree of integration 1, the probability limits of the fixed-effects estimator will be

$$P \lim_{p \rightarrow 1} \overset{\wedge}{\lambda}_{FE} = 1 \quad \text{and} \quad P \lim_{p \rightarrow 1} \overset{\wedge}{\beta}_{FE} = 0.$$

At a first glance, a situation in which the persistence coefficient is biased towards 1, and the role of the explanatory variables is biased towards zero may be too discouraging to undertake any sort of sensible analysis.<sup>10</sup> However, to the extent that this bias affects both set of estimated coefficients in a similar way, the resulting long-run elasticities are

---

<sup>10</sup>Furthermore, the heterogeneity bias cannot be dealt with by traditional instrumental variable estimators so that, even under the estimation by GMM or 3SLS, our estimation will still be potentially subject to this bias (Smith and Fuertes, 2010).

relatively save from the heterogeneity bias. As noted in Smith and Fuertes (2010, p. 32) “the asymptotic bias in the estimator of the long-run coefficient  $\hat{\theta} = \hat{\beta}_{FE} / (1 - \hat{\lambda}_{FE})$  is not as severe, because the biases in the numerator and denominator tend to cancel out”.

Furthermore, Baltagi and Griffin (1997) compare a large number of panel data estimators and find that standard homogenous estimators perform better than their heterogenous counterparts because “the efficiency gains from pooling appear to more than offset the biases due to intercountry heterogeneities” [Baltagi and Griffin (1997), p. 317].

Given these arguments, and the fact that our focus of analysis is on the long-run elasticities, our estimations are conducted using standard homogeneous panel data techniques.

## 4 Estimates

Equations (3) and (4) are estimated as a one-way fixed-effects (FE) model, by GMM, and by 3SLS. The GMM estimation considers wages and capital (or output) as exogenous variables to ensure comparability with respect to the fixed-effects model. Given the panel characteristics of our database, we should expect the FE and these GMM estimates to be similar, as they turn out to be (see Tables 4 and 6). However, given our exposure to the heterogeneity bias, the unavoidable differences between both sets of estimates, which is due to the way in which the GMM estimator makes use of the instruments, yield substantial differences in the estimated long-run elasticities (see Tables 5 and 7). This is the reason why we resort to the 3SLS estimation as a way of endogenizing wages and capital (or output) and as a way, at the same time, of avoiding the well-known sensitivity problem of the GMM estimates to the choice of instruments.

### 4.1 1964-1984

Table 4 presents the results for our first period of analysis. The estimated specifications are similar across models –equations (3) and (4)– and methodologies –FE, GMM, 3SLS–, and all explanatory variables are highly significant and take the expected signs. The only exception is the openness control in the 3SLS estimation of equation (3).

**Table 4. Estimated labour demand. Spain, 1964-1984.**

Dependent variable: $\Delta n_{ijt}$						
	FE		GMM		3SLS	
	<i>k</i>	<i>y</i>	<i>k</i>	<i>y</i>	<i>k</i>	<i>y</i>
<i>c</i>	0.118 [0.018]	0.311 [0.000]				
$n_{ijt-1}$	-0.010 [0.000]	-0.017 [0.000]	-0.008 [0.000]	-0.017 [0.000]	-0.011 [0.000]	-0.025 [0.000]
$\Delta n_{ijt-1}$	0.173 [0.000]	0.147 [0.000]	0.223 [0.000]	0.204 [0.000]	0.199 [0.000]	0.164 [0.000]
$\Delta n_{ijt-2}$	0.010 [0.028]	0.094 [0.026]				
$k_{ijt}$	0.008 [0.000]		0.006 [0.000]		0.009 [0.000]	
$\Delta k_{ijt}$	0.083 [0.000]		0.078 [0.000]		0.113 [0.000]	
$y_{ijt}$		0.015 [0.000]		0.016 [0.000]		0.022 [0.000]
$\Delta y_{ijt}$		0.170 [0.063]		0.214 [0.000]		0.280 [0.000]
$w_{ijt}$	-0.007 [0.043]	-0.015 [0.000]	-0.004 [0.008]	-0.013 [0.000]	-0.008 [0.030]	-0.023 [0.000]
$\Delta w_{ijt}$	-0.190 [0.000]	-0.249 [0.000]	-0.113 [0.000]	-0.217 [0.000]	-0.162 [0.000]	-0.254 [0.000]
$\Delta w_{ijt-1}$			0.040 [0.000]	0.041 [0.000]	0.035 [0.045]	0.036 [0.031]
<i>t</i>	-0.004 [0.000]	-0.006 [0.000]	-0.004 [0.000]	-0.004 [0.000]	-0.004 [0.000]	-0.002 [0.094]
$op_{ijt}$	0.046 [0.000]	0.130 [0.000]	0.054 [0.000]	0.048 [0.000]	0.050 [0.021]	0.021 [0.514]
$\Delta op_{ijt-1}$		-0.084 [0.000]			-0.031 [0.113]	
$\Delta op_{ijt-2}$		-0.077 [0.000]	-0.042 [0.000]	-0.032 [0.003]	-0.049 [0.005]	-0.206 [0.455]
<i>Obs.</i>	4, 284	4, 284	4, 284	4, 284	4, 284	4, 284

FE: Fixed-effects; GMM: Generalized Method of Moments.; 3SLS: Three Stages Least Squares

p-values in brackets. Instruments for GMM:  $n_{ijt-1}, n_{ijt-2}, k_{ijt}, k_{ijt-1}, w_{ijt}, w_{ijt-1}, w_{ijt-2},$

$t, op_{ijt}, \Delta op_{ijt-2}$  [ $y_{ijt}, y_{ijt-1}$  substitute  $k_{ijt}, k_{ijt-1}$  in model with  $y$ ];

Instruments for 3SLS:  $n_{ijt-1}, n_{ijt-2}, k_{ijt-1}, k_{ijt-2}, w_{ijt-1}, w_{ijt-2}, w_{ijt-3},$

$t, op_{ijt}, \Delta op_{ijt-1}, \Delta op_{ijt-2}$  [ $y_{ijt-1}, y_{ijt-2}$  substitute  $k_{ijt-1}, k_{ijt-2}$  in model with  $y$ ].



Table 5 presents the implied long-run elasticities arising from the base-run estimates displayed in Table 4. Recall that the total effect is obtained from equation (4), while our reference estimate of the substitution effect is the one obtained from equation (3). To clearly show the consistency across estimated models, as a robustness check we also present the estimate of the negative of the substitution effect obtained from equation (4). Then we use equation (8) to compute the scale effect as the total effect minus the constant-output elasticity (which is obtained from the estimated substitution effect).

**Table 5. Long-run elasticities of the demand for labour. 1964-1984.**

	FE	GMM	3SLS
<b>Substitution and total effects:</b>			
Estimated substitution effect	-0.91	-0.72	-0.92
Robustness check	0.82	0.85	0.86
Estimated total effect ( $A$ )	-0.75	-0.49	-0.71
<b>Total effect decomposition:</b>			
Computed constant-output elasticity ( $B$ )	-0.21	-0.17	-0.22
Computed scale effect ( $= A - B$ )	-0.54	-0.32	-0.49

Notes: total and substitution effects obtained, respectively, from estimated equations (4) and (3); constant-output elasticity = substitution effect\*( $1 - s_L$ ).

Qualitatively, the three estimated procedures deliver a similar picture. Quantitatively, the FE and the 3SLS procedures provide very similar estimates, while the GMM estimates are also quantitatively very close in terms of the substitution effect and the resulting constant-output elasticity. The main quantitative difference is in the estimated total effect with consequences on the computed scale effect.

Given these results and our preference for the 3SLS, we conclude that the substitution effect can safely be placed around -0.9, and the total effect around -0.7. That is, a 1% increase in real wages, will cause a 0.7% reduction in employment. This reduction can be attributed to less than one third (-0.2 of -0.7 percentage points) to the constant-output elasticity, and in more than two thirds (-0.5 of -0.7 percentage points) to the scale effect.

## 4.2 1985-2007

For the second period we estimate the same model specifications than for period one. The results, displayed in Table 6, are similar across estimation procedures and could be a reflection of the absence of serious endogeneity problems.

**Table 6. Estimated labour demand. Spain, 1985-2007.**

Dependent variable:  $\Delta n_{ijt}$

	FE		GMM		3SLS	
	<i>k</i>	<i>y</i>	<i>k</i>	<i>y</i>	<i>k</i>	<i>y</i>
<i>c</i>	-0.207 [0.066]	0.006 [0.934]				
$n_{ijt-1}$	-0.024 [0.000]	-0.086 [0.000]	-0.024 [0.000]	-0.087 [0.000]	-0.026 [0.000]	-0.073 [0.000]
$\Delta n_{ijt-1}$		-0.026 [0.202]	-0.028 [0.003]	-0.020 [0.017]	-0.046 [0.001]	0.028 [0.017]
$k_{ijt}$	0.023 [0.000]		0.024 [0.000]		0.026 [0.000]	
$\Delta k_{ijt}$	0.171 [0.000]		0.195 [0.000]		0.237 [0.000]	
$y_{ijt}$		0.085 [0.000]		0.086 [0.000]		0.072 [0.000]
$\Delta y_{ijt}$		0.422 [0.000]		0.408 [0.000]		0.208 [0.000]
$w_{ijt}$	-0.024 [0.006]	-0.078 [0.000]	-0.034 [0.000]	-0.084 [0.000]	-0.033 [0.000]	-0.075 [0.000]
$\Delta w_{ijt}$	-0.337 [0.000]	-0.384 [0.000]	-0.292 [0.000]	-0.333 [0.000]	-0.390 [0.000]	-0.403 [0.000]
$\Delta w_{ijt-1}$	-0.057 [0.002]	-0.039 [0.023]	-0.068 [0.000]	-0.037 [0.003]	-0.074 [0.000]	-0.053 [0.000]
<i>tr</i>	-0.004 [0.000]	-0.002 [0.000]	-0.007 [0.000]	-0.003 [0.000]	-0.004 [0.001]	-0.003 [0.002]
$op_{ijt}$	0.077 [0.000]	0.034 [0.004]	0.126 [0.000]	0.052 [0.000]	0.064 [0.004]	0.046 [0.012]
$\Delta op_{ijt-2}$	-0.159 [0.001]		-0.153 [0.000]	0.015 [0.408]	-0.159 [0.001]	-0.027 [0.477]
<i>Obs.</i>	5,474	5,474	5,474	5,474	5,474	5,474

FE: Fixed-effects; GMM: Generalized Method of Moments.; 3SLS: Three Stages Least Squares

p-values in brackets. Instruments for GMM:  $n_{ijt-1}, n_{ijt-2}, k_{ijt}, k_{ijt-1}, w_{ijt}, w_{ijt-1}, w_{ijt-2}, t, op_{ijt}, \Delta op_{ijt-2}$  [ $y_{ijt}, y_{ijt-1}$  substitute  $k_{ijt}, k_{ijt-1}$  in model with  $y$ ];

Instruments for 3SLS:  $n_{ijt-1}, n_{ijt-2}, k_{ijt-1}, k_{ijt-2}, w_{ijt-1}, w_{ijt-2}, w_{ijt-3},$

$t, op_{ijt}, \Delta op_{ijt-1}, \Delta op_{ijt-2}$  [ $y_{ijt-1}, y_{ijt-2}$  substitute  $k_{ijt-1}, k_{ijt-2}$  in model with  $y$ ].

Table 7 presents the long-run elasticities corresponding to the estimates displayed in Table 6.

**Table 7. Long-run elasticities of the demand for labour. 1985-2007.**

	FE	GMM	3SLS
<b>Substitution and total effects:</b>			
Estimated substitution effect	-0.90	-0.97	-1.03
Robustness check	0.95	0.98	0.98
Estimated total effect ( $A$ )	-0.99	-1.40	-1.27
<b>Total effect decomposition:</b>			
Computed constant-output elasticity ( $B$ )	-0.34	-0.36	-0.38
Computed scale effect ( $= A - B$ )	-0.65	-1.04	-0.89

Notes: total and substitution effects obtained, respectively, from estimated equations (4) and (3); constant-output elasticity = substitution effect $\cdot(1 - s_L)$ .

The qualitative picture given by these elasticities is consistent across methodologies, while the quantitative picture is even more consistent than for the previous period. For example, the estimated substitution effect ranges from -0.90 to -1.03, irrespective of the equation from which this estimate is obtained and irrespective of the econometric procedure in use. The resulting constant-output elasticity is placed in the narrow range between -0.34 and -0.38. In turn, the estimated total effect is larger under the 3SLS estimation than using fixed-effects, but the GMM value is even larger.

Overall, our reference values for this period are a unitary substitution effect and a total effect close to -1.3. This total effect is the result of a constant-output elasticity accounting for -0.4 percentage points (around a third of the total effect), and a scale effect amounting to -0.9 percentage points (almost 70% of the total effect).

### 4.3 Assessment

Comparison of our results for periods 1964-1984 and 1985-2007 yield two salient findings.

The first one is the increase in the substitution effect, which rises from -0.9 to -1. This implies that the growing exposure to international trade and the different waves of labour market reforms have contributed to enhance the sensibility of employment to wages. This is consistent with Rodrik's conjecture, according to which a critical consequence of the

globalization process is to facilitate labour-capital substitution since all economies have easier access to new markets and new avenues of productive specialization.

The second main finding of our analysis is the increase in the magnitude of the total effect, which jumps from -0.7 to -1.3. This jump of 0.6 percentage points, is the consequence of a rise in the two components of the total effect.<sup>11</sup>

The constant-output elasticity rises from -0.2 to -0.4, and accounts for one third of the increase in the total effect. Knowing that half of this increase is due to the rise in the substitution effect, the rest can be ascribed to the fall in the labour share elasticity. Economically, the increase in the sensitivity of employment with respect to wages even when holding output constant implies that firms have enhanced their internal flexibility to respond to price changes. This may be reflecting the enhanced possibilities brought by the new technologies and the growing pressure to which firms are subject to compete in the international arena.

In turn, the scale effect almost doubles its size. It augments by 0.4 percentage points, and accounts for two thirds of the increase in the total effect. The scale effect accounts for the net effect of the interplay between a product-supply effect (resulting from the fact that lower costs allow an increase in production) and a product-demand effect (arising from the fact that wages are directly connected with consumption and, thus, with aggregate demand in the product market). Our finding of a significant increase in the scale effect reflects the growing relative relevance of the product-supply effect. In a globalized context, a rise in the labour cost ends up reducing output and, thereby, employment. The increase in labour costs may be transferred on to prices and generate a fall in production larger than the product-demand increase resulting from the higher labour costs. In a context of large international exposure, such as the Spanish one in the aftermath of the EEC accession in 1986, it is appropriate to expect an increase in the product-supply effect because it is easier to circumvent labour costs increases by delocalizing part of the productive activity. In turn, it is also likely that the product-demand effect has diminished its influence because the consequences of wage rises have become more diluted. With increasing openness, wage rises affect not only domestic demand, but also the imports of goods and services, which in Spain have massively grown in 1985-2007.

---

<sup>11</sup>An interesting feature of the ARDL methodology is that it yields detailed information on the channels behind the rise in the long-run elasticities. Our results show that this increase is due to falls both in the persistence coefficient and in the short-run labour demand sensitivity with respect to wages.

## 5 Conclusions

Does the sensitivity of the demand for labour increase with globalization? Following Rodrik's (1997) conjecture a 'yes' response should be the expected one. However, and despite the growing attention that this issue has received in recent years, no consensus has yet been reached.

It is in this context that we aimed at determining whether the structural change caused by the internationalization process of the Spanish economy since the mid 1980s has significantly increased its labour demand elasticity. Our results indicate a substantial increase in the total employment sensitivity with respect to wages and give support to Rodrik's conjecture. Beyond that, we find that this increase is the outcome of both a larger constant-output elasticity and a larger scale effect. This means that growing international exposure stimulates the labour demand sensitivity through both potential channels.

Although we are not able to disentangle the relative size of the two scale effect components, we argue that both evolved in opposite direction (higher relevance of the product-supply effect relative to the product-demand effect) and jointly reinforced (given their opposite sign) the growing impact of the scale effect on the total employment response to wage changes.

The growing relevance of the scale effect in Spain indicates that firms have benefited more than households from the enhanced flexibility brought by globalization. Along with the internationalization process, firms have managed to circumvent increases in labour costs by increasing their mobility and the use of new technologies. This explains phenomena like offshoring and outsourcing, firm delocalization, and, more generally, the growing flows of outward foreign direct investment. In contrast, in a situation of real wage moderation as the Spanish one since the mid 1980s, even though households have also augmented their product-demand sensitivity to wage changes, their restricted mobility and use of new technologies (at least relative to firms) have prevented this sensitivity to increase further.<sup>12</sup>

The other side of the globalization coin is the downward trend followed by the labour share, in particular from the mid 1980s onwards. If firms have benefited more than households, it is licit to warn about the distributional perils of pursuing the cost-reducing strategy that has mainly been followed to face the increasing global market competition (recall that Spanish wages have grown less than 1% on average between 1985 and 2007).<sup>13</sup>

---

<sup>12</sup>Households augment their product-demand sensitivity to wage changes because they can substitute national consumption by a progressively cheaper international provision of goods and services. This explains why wage moderation diminishes the demand of national products more than before (i.e., when the economy was closed).

<sup>13</sup>Recall that by equation (1) movements in the labour income share have consequences (in opposite directions) in terms of both the constant-output elasticity and the scale effect. This implies an ambiguous

Rather, it would seem advisable to prioritize labour productivity growth so as to be able to allow some wage increase without harming competitiveness and without eroding job creation.

The database we have used contains rich information in terms of sectoral disaggregation. Further research should aim at carefully assessing how the Spanish transformation affected the labour demand elasticity across productive sectors. Our hypothesis is that the sectoral response is connected to the degree of exposure of each sector to international trade.

## References

- [1] Álvarez, J. and Arellano, M. (2003): “The time series and cross-section asymptotics of dynamic panel data estimators”, *Econometrica*, **71** (4), pp. 1121-1159.
- [2] Antràs, P. (2004): “Is the U.S. Aggregate Production Function Cobb-Douglas? New Estimates of the Elasticity of Substitution”, *Contributions to Macroeconomics*, **4**(1), Article 4.
- [3] Arellano, M. (2003): *Underidentification?*, Presidential Address, XXVIII Simposio de Análisis Económico, Sevilla (a lecture from joint work with L. P. Hansen & E. Sentana).
- [4] Azariadis, C. and Pissarides, C. (2007): “Unemployment dynamics with international capital mobility”, *European Economic Review*, **51**(1), pp. 27–48.
- [5] Baltagi, B.H. and Griffin, J.M. (1997): “Pooled Estimators versus their Heterogeneous Counterparts in the context of dynamic demand for gasoline”, *Journal of Econometrics*, **77**, 303-327.
- [6] Benito, A. and Hernando, I. (2003): “Labour demand , flexible contracts and financial factors: New evidence from Spain”, *Banco de España Working Papers* 0312. Bank of Spain.
- [7] Bentolila, S. and Saint-Paul, G. (2003): “Explaining movements in the labor share”, *Contributions to Macroeconomics*, Berkeley Electronic Press, **3**, article 9.
- [8] Berndt, E. (1976): “Reconciling Alternative Estimates of the Elasticity of Substitution”, *Review of Economics and Statistics*, **58** (1), pp. 59-68.
- [9] Borjas, G.J. and Ramey, V.A. (1995): “Foreign competition, market power, and wage inequality”, *Quarterly Journal of Economics*, **110**(4), pp. 1075-1110.
- [10] Bruno, G.S.F., Falzoni, A.M. and Helg, R. (2004): “Measuring the effect of globalization on labour demand elasticity: an empirical application to OECD countries”, *Working paper* No. 153.

---

net effect in terms of the total labour demand elasticity. Recall, also, that the scale effect is by definition the product of the labour income share and the output elasticity with respect to employment. This implies, in turn, that, other things equal, the fall in the labour share should push the scale effect down. The reason why this is not verified in Spain is the increase in the output elasticity with respect to employment, which is a reflection of the increased flexibility of firms when they have had to face labour adjustments.

- [11] Clark, K.B. and Freeman, R.B. (1980): “How elastic is the demand for labor?”, *The Review of Economics and Statistics*, **62**(4), pp. 509-520.
- [12] Dabán, T., Díaz, A., Escribá, F.J., and Murgui, M.J. (2002): “La base de datos BDMORES”, *Revista de Economía Aplicada*, **10**(30), pp. 165-186.
- [13] Dolado, J.J., García-Serrano, C., and Jimeno, J.F. (2002): “Drawing lessons from the boom of temporary jobs in Spain”, *The Economic Journal*, **112** (480), pp. F270-F295.
- [14] Dowrick, S. and G.M. Wells (2004): “Modelling aggregate demand for labour: a critique of Lewis and MacDonald”, *The Economic Record*, **80**, pp. 436-440.
- [15] Drazen, A., Hamermesh, D.S., and Obstm N.P. (1984): “The variable employment elasticity hypothesis: Theory and evidence”, *Research in Labor Economics*, **6**, pp. 287-309.
- [16] Fisher, R.A. (1932): *Statistical Methods for Research Workers*, Edinburgh: Oliver & Boyd, **4th** edition.
- [17] Hamermesh, D.S. (1993): *Labor Demand*, Princeton University Press, Princeton.
- [18] Hijzen, A. and P. Swaim (2010): “Offshoring, labour market institutions and the elasticity of labour demand”, *European Economic Review*, **54** (8), pp. 1016-1034.
- [19] Im, K.S., Pesaran, M.H., and Shin, Y. (2003): “Testing for unit roots in heterogeneous panels”, *Journal of Econometrics*, **115**, pp. 53-74.
- [20] Juselius, K. and Toro (2005): “Monetary transmission mechanism in Spain: The effect of monetization, financial deregulation and the EMS”, *Journal of International Money and Finance*, **24** (3), pp. 509-531.
- [21] Karanassou, M., Sala, H. and Snower, D.J. (2010): “Phillips curves and unemployment dynamics: A critique and a holistic perspective”, *Journal of Economic Surveys*, **24**, pp. 1-51.
- [22] Krishna, P., Mitra, D. and Chinoy, S. (2001): “Trade liberalization and labor demand elasticities: evidence from Turkey”, *Journal of International Economics*, **55**, pp. 391-401.
- [23] Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., and Shin, Y. (1992) “Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series are non stationary?”, *Journal of Econometrics*, **54**, pp. 159-178.
- [24] Lawrence, C. and Lawrence, R.Z. (1985): “Manufacturing wage dispersion: An end game interpretation”, *Brooking Papers on Economy Activity*, **1**, pp. 47-116.
- [25] Levin, A. and C.F. Lin (1993): “Unit Root Tests in Panel Data: New Results”, Department of Economics, *University of California at San Diego*, *Discussion Paper* No. 93-56.
- [26] Lewis, P.E.T. and G. MacDonald (2002): “The elasticity of demand for labour in Australia”, *The Economic Record*, **78**, pp. 18-30.
- [27] Lewis, P.E.T. and G. MacDonald (2004): “Modelling aggregate demand for labour: a reply to Dowrick and Wells”, *The Economic Record*, **80**, pp. 441-444.
- [28] Maddala and Wu (1999): “A comparative study of unit root tests with panel data and new simple test”, *Oxford Bulletin of Economics and Statistics*, **61**, pp. 631-652.

- [29] Molnar, M., Pain, N. and Taglioni, D. (2007): “The internationalization of production, international outsourcing and employment in the OECD”, *OECD Economics Department Working Paper No. 561*, Paris.
- [30] Pesaran, M.H. and Smith, R.P (1995): “Estimating Long-run relationships from Dynamic Heterogenous Panels”, *Journal of Econometrics*, **68**, pp. 79-113
- [31] Pesaran, M.H., Smith, R.P and Im, K.S. (1996): “Dynamic linear models for Heterogenous Panels” in L. Matyas and P. Sevestre (Eds.), *The Econometrics of Panel Data*, 2nd Ed., pp. 145-195. Dordrecht: Kluwer Academic Publishers.
- [32] Polo, C. and F. Sancho (1993): “An analysis of the Spain’s integration in the EEC”, *Journal of Policy Modelling*, **15**, pp. 157-168.
- [33] Nickell, S.J. (1981): “Biases in dynamic models with fixed-effect”, *Econometrica*, **49 (6)**, pp. 1417-1426.
- [34] Revenga, A.L. (1992): “Exporting jobs? The impact of import competition on employment and wages in US manufacturing”, *Quarterly Journal of Economics*, pp. 255-284.
- [35] Rodrik, D. (1997): *Has Globalization Gone Too Far?*, Institute for International Economics, Washington, DC.
- [36] Russell, B. and Tease, W. (1991): “Employment, output and real wages”, *The Economic Record*, **67**, pp. 34-45.
- [37] Senses, M.Z. (2010): “The effects of offshoring on the elasticity of labor demand”, *Journal of International Economics*, **81**, pp. 391-401
- [38] Slaughter, M. (2001): “International trade and labor-demand elasticities”, *Journal of International Economics*, **54**, pp. 27-56
- [39] Webster, E. (2003): “The Effects of Wages on Aggregate Employment: A Brief Summary of Empirical Studies”, *The Australian Economic Review*, **36**, pp. 134-142.