Labour Demand Effects of Internationalization in the Colombian Manufacturing Industry, 1974-2015*

Sonia A. Agudelo**

Recibido: 14 de marzo de 2022 • Aprobado: 09 de agosto de 2022
https://doi.org/10.22395/seec.v25n58a5

ABSTRACT

Colombia experienced a major structural change in the early nineties. This structural change is related to major trade and labour market reforms undertaken around 1992, which is considered the inflection year. This paper focuses on evaluating whether the structural change in international trade exposure significantly altered the employment response to wage changes in the manufacturing industry. Hamermesh's (1993) framework is used to explain the total employment effect, which is made of the substitution and scale effects. This total effect is empirically quantified, and its two components disentangled. We find that the employment elasticity of a wage change rose from -1.05 in 1974-1991 to -1.56 in 1992-2015. By components, the substitution effect rose from -0.68 to -1.19, and the scale effect remained stable at -0.38. These findings suggest that trade liberalization has had negative consequences on workers' welfare. A larger sensitivity of labour demand has been associated with a larger workers' tax burden and a greater instability in labour market outcomes (Rodrik, 1997). Therefore, the increases in payroll taxation experienced during the nineties may have led to job destruction and an amplified workers' tax burden, while capital-labour substitution processes may have accelerated on account of the larger employment sensitivity.

KEYWORDS

Labour demand elasticity; substitution elasticity; trade openness; Colombia.

JEL CLASSIFICATION

J23, F41, F16.

CONTENT

* This article is a research product of the thesis entitled: *Three essays on the Colombian labour market: A macroeconomic perspective* (Agudelo, 2019). It was submitted to the Autonomous University of Barcelona in partial fulfilment of the requirements for the degree of Ph.D. Program in Applied Economics, Department of Applied Economics, Advisor: Hector Sala Lorda, November 2019. Sonia A. Agudelo acknowledges helpful comments from Omar Arias, which have led to significant improvements in the paper. Sonia is also grateful for financial support received from Minciencias.

** Economist, Universidad de Medellín, Medellin, Colombia. Master of Research in Applied Economics and Ph.D. in Applied Economics, Autonomous University of Barcelona, Barcelona, Spain. Professor and researcher, Faculty of Economic and Administrative Sciences, Universidad de Medellín, Medellin, Colombia. Professor, Department of Basic Sciences and Common Areas. Tecnológico de Antioquia, Medellin, Colombia. E-mails: soagudelo@udemedellin.edu.co, sonia.agudelo@tdea.edu.co, soniaagudelo3@hotmail.com. Orcid: https://orcid.org/0000-0001-7499-9169.

Efectos de la internacionalización sobre la demanda de trabajo en la industria manufacturera colombiana, 1974-2015

RESUMEN

Colombia experimentó un importante cambio estructural a principios de los años noventa. Este cambio estructural está relacionado con las grandes reformas del comercio y del mercado laboral emprendidas en torno a 1992, que se considera el año de inflexión. Este trabajo se centra en evaluar si el cambio estructural en la exposición al comercio internacional alteró significativamente la respuesta del empleo a los cambios salariales en la industria manufacturera. Se utiliza el marco de Hamermesh (1993) para explicar el efecto total sobre el empleo, que se compone de los efectos de sustitución y de escala. Este efecto total se cuantifica empíricamente y se desentrañan sus dos componentes. Encontramos que la elasticidad del empleo de un cambio salarial pasó de -1.05 en 1974-1991 a -1.56 en 1992-2015. Por componentes, el efecto de sustitución aumentó de -0.68 a -1.19. y el efecto de escala se mantuvo estable en -0.38. Estos resultados sugieren que la liberalización del comercio ha tenido consecuencias negativas en el bienestar de los trabajadores. Una mayor sensibilidad de la demanda de mano de obra se ha asociado a una mayor carga fiscal de los trabajadores y a una mayor inestabilidad de los resultados del mercado laboral (Rodrik, 1997). Por lo tanto, los aumentos de la imposición sobre las nóminas experimentados durante los años noventa pueden haber provocado la destrucción de puestos de trabajo y un aumento de la carga fiscal de los trabajadores, mientras que los procesos de sustitución de capital por trabajo pueden haberse acelerado debido a la mayor sensibilidad del empleo.

PALABRAS CLAVE

Elasticidad de la demanda de trabajo; elasticidad de sustitución; apertura comercial; Colombia.

CLASIFICACIÓN JEL

J23, F41, F16.

CONTENIDO

Introducción; 1. Hechos estilizados; 2. Efectos del comercio internacional sobre la demanda de trabajo: Marco analítico; 3. Aplicación empírica; 4. Datos y modelización empírica; 5. Resultados; 6. Observaciones finales; Referencias; Apéndices.

Efeitos da internacionalização sobre a procura de trabalho na indústria manufatureira colombiana – 1974 a 2015

RESUMO

Colômbia experimentou uma significativa mudança estrutural no começo dos anos noventa. Esta mudança estrutural está relacionada com as grandes reformas do comércio e do mercado de trabalho empreendidas no ano de 1992, ano este considerado como a da mudança. Este trabalho centra-se em avaliar se a transformação estrutural na exposição ao comércio internacional afetou significativamente a resposta do emprego face as alterações salariais na indústria manufatureira. É utilizado o marco de Hamermesh (1993) para explicar o efeito geral sobre o emprego, que é composto dos efeitos da substituição e de escala. Este efeito geral quantifica-se empiricamente e se desentranham seus dois componentes. Encontramos que a elasticidade do trabalho de uma mudança salarial passou de -1.05 em 1974 a -1.56 em 1992-2015. Pelos componentes, o efeito de substituição aumentou de -0.68 a -1.19, e o efeito de escala se manteve estável em -0.38. Estes resultados sugerem que a liberalização do comércio tem tido consequências negativas no bem estar dos trabalhadores. Uma maior sensibilidade na procura de mão de obra tem sido associada a uma maior carga tributária dos trabalhadores e uma maior instabilidade dos resultados do mercado laboral (Rodrik, 1997). Por conseguinte, os aumentos da imposição sobre as folhas de pagamento experimentados durante os anos noventa podem ter provocado a diminuição dos postos de trabalho e um aumento tributário dos trabalhadores, enquanto que os processos de substituição de capital pelo trabalho podem ter sido acelerado devido a maior sensibilidade do emprego.

PALAVRAS-CHAVE

Elasticidade da procura de trabalho; elasticidade de substituição; abertura comercial; Colombia.

CLASIFICA CAO JEL

J23. F41. F16.

CONTEÚDO

Introdução. 1. Feitos estilizados; 2. Efeitos do comércio internacional sobre a procura de trabalho: marco analítico; 3. Aplicação empírica; 4. Dados e modelização empírica; 5. Resultado; 6. Observações finais; Referências; Apêndices.

INTRODUCTION

In a context of increasing globalization, changes in international trade patterns and labour outcomes have become gradually interlinked. Labour market responses to policy measures have transformed and a growing body of literature has become concerned about the link between globalization, employment, and labour market reforms. (e.g., Selwaness and Zaki, 2019; Malgouyres, 2017; Acemoglu *et al*., 2016; and Autor *et al*., 2013). A large part of the literature has focused on analyzing the impact of globalization on the labour demand. Nevertheless, the empirical evidence is far from reaching a consensus. Even more, the empirical literature is still scarce for developing countries.

On this account, this paper aims to study the labour demand effects of international trade in the Colombian manufacturing industry in 1974-2015. As most Latin American countries, Colombia experienced a major structural change in the early nineties. This structural change is related to the most important trade and labour market reforms undertaken in the early nineties. Trade regulations and a relatively tight labour market legislation in the seventies and eighties were superseded by trade liberalization and more flexible labour market institutions in the nineties.

Colombia provides a salient case to explore. It is one of the considered successful economies in Latin America which, in the eighties and nineties, embarked in an extensive liberalization program. As described by Agudelo and Sala (2016) this program was carried out in three stages. The first step in this process took place unilaterally in 1990, when the political authorities increased the Colombian exposure to international trade by reducing, simultaneously, import controls and import tariffs. The second step took place between 1992 and 2004, when Colombia enjoyed a new system of preferential tariffs to export to the US. Finally, this system was superseded in 2004 by Free Trade Agreements (FTAs) between Colombia and a few relevant trade partners such as the US, the European Union, Canada, Mexico, Korea, Chile, Salvador, Guatemala, and Honduras. The main consequence of this trade reform was a significant structural change in the degree of exposure to international trade. The degree of openness —measured as the sum of exports and imports over GDP— increased from below 25 % in 1991 to more than 38 % in 2015.

In this context, the manufacturing industry has been the most exposed economic activity in Colombia, with a steady increase in the imports share that attained around 90 % of total Colombian imports in 2015, and a share of exports above 60 % of total exports. Still more important, trade openness in this sector —measured as the ratio of industrial exports and imports over total industrial output— doubled from an average of 30.9 % in 1974-1991, to one of 68.2 % in 1992-2015 (Figure 3a). This is one of the reasons why our analysis places specific attention to the manufacturing industry.

In parallel to the external liberalization process, Colombia also embarked in a structural reform of the labour market. It took place in the early 1990s when Law 50 was passed in 1990 to enhance the flexibility of the labour market. This enhanced flexibility was achieved by reducing firing, training, and recruitment costs, and allowing a general use of temporary contracts. The main result of these institutional changes was a segmentation of the Colombian labour market. As shown in Figure 4a, between 1974 and 1991, the share of agency workers over total employment was less than 8 % in the manufacturing industry. This share, however, sharply increased due to the boom of outsourcing jobs, reaching a peak of 27 % in 2007.

Rodrik (1997) was the first one to conjecture on the labour demand consequences of the globalization process. He expected that this process would affect labour markets through two channels. One channel is the rise in the elasticity of labour demand with respect to wage changes —the wage elasticity effect—. The second channel is the reduction on the demand for low-skilled labour which results in an inward shift in the demand curve for low-skilled labour —the level effect—. In connection with Rodrick's conjecture (1997) of a higher labour demand elasticity, Hamermesh's (1993) model becomes useful to explain the driving forces of such increase. According to this author, the sensitivity of employment to wage changes is driven by the substitution and scale effects. The first 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 captures the reduction in employment due to the reduction in output holding production technology constant. On this account, a larger exposure to international trade increases the total employment sensitivity because, on one side, the lower entry barriers in new markets increases competition —enhancing the scale effect—. On the other side, the emergence of new phenomena such as outsourcing and offshoring enable firms to access to a larger variety of intermediate inputs and capital equipment —reinforcing the substitution effect—.

The wage elasticity and level effects described by Rodrik (1997) have been widely reviewed by the empirical literature, but separately. Although the impact of globalization on the elasticity of labour demand has received growing attention in the last decades, the empirical evidence is not conclusive so far. Most of studies on developed countries —mainly on OECD (2016) countries and the US— give support to Rodrik's conjecture of a more elastic labour demand. In contrast, the literature for developing countries provides mixed support.

Hijzen and Swaim (2010), for instance, associate the rise in wage labour demand elasticities in many OECD countries to the growing use of offshoring practices, even though they find that this positive relationship is weaker the stricter the employment protection legislation is. In addition, Seo *et al*. (2015) find that financial market liberalization is also an important factor driving the rise in wage-elasticities in OECD countries. In the case of US, the positive relationship between labour demand and trade was also verified by Senses (2010) and Slaughter (2001).

In contrast, for the developing economies, Krishna *et al*. (2001) did not find significant effects of the trade liberalization process in Turkey. Similar results emerge in the work by Fajnzylber and Maloney (2005) for manufacturing establishments in Chile, Colombia, and Mexico. Hasan *et al*. (2007), nevertheless, showed that trade made demand for labour more elastic in India.

In this context, our objective is twofold. First, we investigate whether the structural change in international exposure has affected the sensitivity of employment to wages in the Colombian manufacturing industry. Beyond that, we are interested in performing the empirical decomposition of this sensitivity between the substitution and scale effects. Second, we evaluate the possibility that the growing exposure to international trade has also caused a "level effect" on the labour demand. Although this level effect has often been associated with the hypothesis of skilled-biased technological change, which posits that a larger exposure to international trade tends to shift the relative demand for low/high-skilled labour, we focus on estimating the aggregate shift in the total labour demand.

We, therefore, contribute to the literature in a threefold dimension. First, the main contribution of this paper is the empirical decomposition of the total labour demand elasticity in a substitution and a scale effect. The second contribution is the empirical computation of the scale effect and the evaluation of its change between the two periods of analysis. The third contribution is the simultaneous evaluation of the wage elasticity and the level effect. Thence, in contrast to previous literature, in our analysis the way through which international trade may affect the labour demand is not constrained to a single channel.

To conduct the empirical analysis, we estimate two standard employment equations with different specifications and interpretation of the parameters (see section 4). These equations include, among other variables, three alternative measures of trade openness —a trade openness index, an import penetration ratio, and an export ratio—. We use a panel data base which covers a long period (1974-2015) and 16 manufacturing sectors. This data base allows to perform a comprehensive analysis of the labour determinants, and to account for the structural change in the early nineties, where the relevant periods of analysis are 1974-1991 and 1992-2015.

We show that the estimated long-run labour elasticity substantially increased, rising from -1.05 in 1974-1991 to -1.56 in 1992-2015. This result is consistent with Rodrik's conjecture (1997) on the existence of a wage elasticity effect of globalization. Beyond that, we find that the higher elasticity for the labour demand is the outcome of a larger substitution effect, which almost doubles its size, increasing from -0.68 in the first period to -1.19 in the second one.

Regarding the scale effect, we find that it can be placed around -0.38 in both periods of analysis. In Hamermesh's (1993) framework, this stability can be interpreted as a net consequence of the interplay between the decline in the labour income share and the increase in the price elasticity of the product demand.

Finally, our findings show that, although international trade enhances the labour demand elasticity, the level effect of a change in the degree of trade openness is scant and has remained stable across the two periods of interest, with a negative impact of exports and no significant impact of imports. These results are interpreted as evidence that progressive export orientation of the Colombian manufacturing industry has not been accompanied by improvements in technical efficiency.

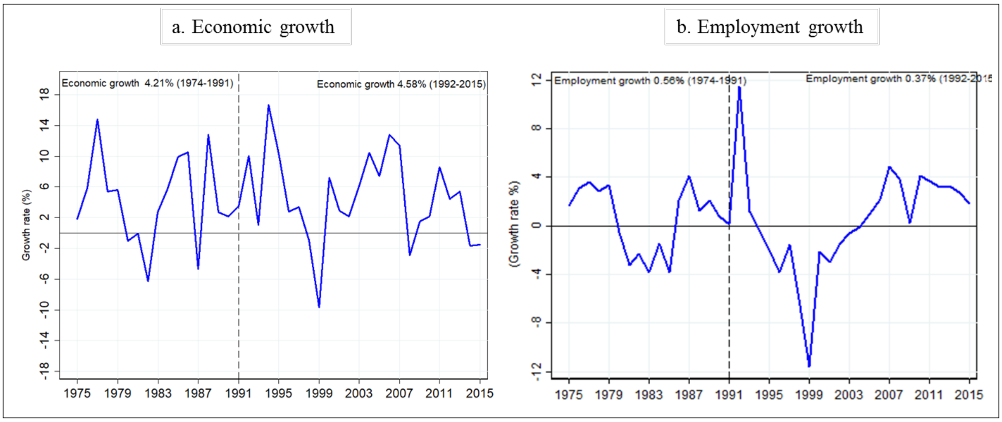
Overall, these findings suggest that trade liberalization has negative consequences on workers. As discussed in Rodrik (1997), higher labour demand elasticity has two important consequences, a greater workers' tax burden and higher employment volatility. It follows the increases in payroll taxation experienced during the nineties in Colombia may have led to a job destruction and a higher tax burden in the manufacturing industry. Moreover, the capital-labour substitution processes may have accelerated on account of the larger employment sensitivity. It implies on one side that if wages in the industrial sector have grown above productivity, there will have been substitution of workers; in contrast, if wages have grown below productivity, then the industry may have become more labour-intensive. As a result of the process of trade liberalization and institutional reforms, workers are placed under high pressure in the new open and deregulated environment.

The rest of the paper is structured as follows. Section 1 describes some macro developments in the Colombian manufacturing industry. Section 2 discusses the analytical framework on the labour demand effects of international trade Section 3 describe the empirical implementation. Section 4 discusses the data and explains the econometric methodology. Section 5 deals with the estimates. Section 6 provides an assessment of the main results.

1. STYLIZED FACTS

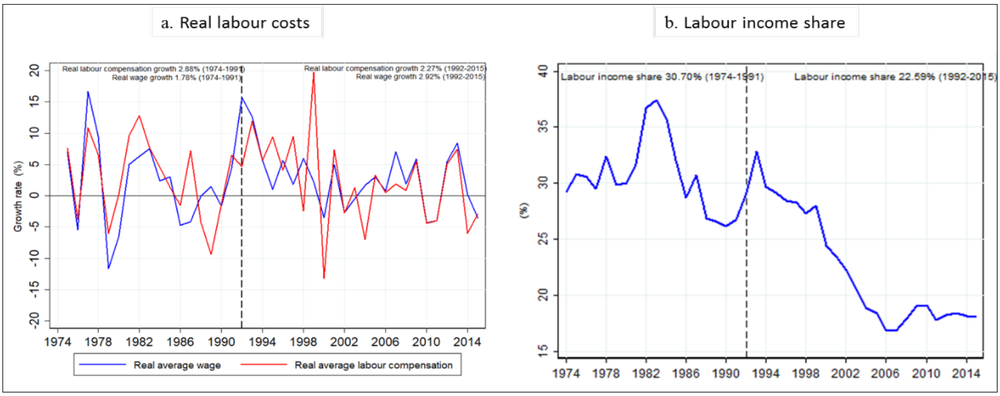
In this section, we provide descriptive information on some macroeconomic variables of the Colombian manufacturing industry across Figures 1 to 4. All data is supplied for the two relevant periods of analysis —the slow transition between import substitution and trade liberalization in 1974-991, and the trade and institutional reforms years of 1992-2015—.

Figure 1. Production and employment dynamics in the Colombian manufacturing industry 1974-2015.



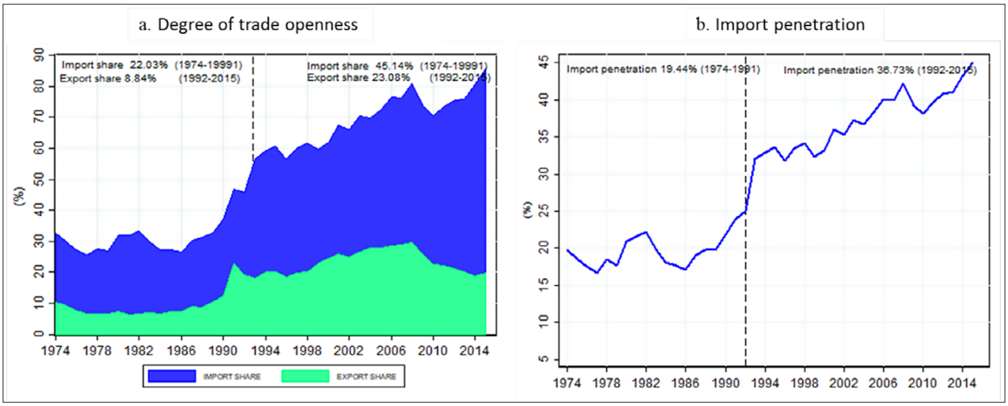
Source: Encuesta Anual Manufacturera (EAM).

Figure 2. Labour costs dynamics in the Colombian manufacturing industry 1974-2015.



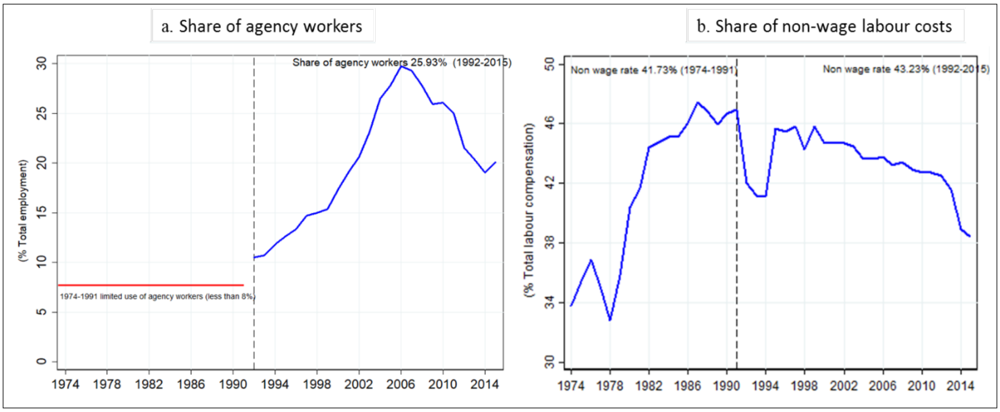
Source: Encuesta Anual Manufacturera (EAM).

Figure 3. Internationalization in the Colombian manufacturing industry 1974-2015.



Source: National Planning Department (DNP in its Spanish acronym).

Figure 4. Labour market segmentation and flexibilization in the Colombian manufacturing industry 1974-2015.



Source: Encuesta Anual Manufacturera (EAM).

Real economic growth of the Colombian manufacturing industry was around 4.4 % on average, since the mid-seventies up to 2015, with non-significant differences between periods (Figure 1a). Despite these growth rates throughout, the employment growth rates did not reach 1 % on average (Figure 1b). The industrial standstill in the seventies and the debt crisis in the eighties resulted in low employment growth rates (0.6 % on average in 1974-1991) and, thus, in high labour productivity rates (3.7 % on average). In contrast, real wages rose at average rates of 1.8 % (Figure 2a), which is almost equivalent to half of the progress in labour productivity growth.

Real labour compensation rose at average rates of 2.9 %, 1 percentage point above real wages growth. The share of non-wage costs over total labour compensation sharply increased. From a value of 34 % in 1975, this share rose by 12 percentage points up 1991 (Figure 4b).

The nineties were characterized not only for a growing exposure to trade, a larger payroll taxation and enhanced labour market flexibility, but also for a deindustrialization and growing expansion of the services sector. They were followed, though, by the recovery of the manufacturing industry in 2000-2009 driven by a rising capital accumulation (boosted both by domestic and foreign investment) and the substitution of domestic by imported raw materials. Altogether, these developments resulted in a scant net job creation (0.4 % on average), a segmentation of the labour market —the share of agency workers over total employment reach a peak of 27 % in 2007 (Figure 3a)— and a sharp acceleration in net capital accumulation in the manufacturing industry (11 % on average, 9 percentage points larger than in 1974-1991). The trade openness index doubled from an average of 30.9 % in 1974-1991, to 68.2 % in 1992-2015 (Figure 3a). The share of non-wage labour costs declined by almost 10 percentage points from a value of 46 % in 1991, to one of 36 % in 1975 (Figure 4b). In turn, real wages rose at average rates of 2.9 % and non-wage labour costs rose at average rates of 2.2 %. Regarding the labour income share, it declined by 8 percentage points in 1992-2015 (Figure 2b). This is the outcome of a permanent wage growth significantly lagging behind labour productivity growth.

Overall, the Colombian industry sector moved from a scenario of a closed economy, with a relatively tight labour market, to a situation in which the economy is widely exposed to international shocks because of the trade liberalization and labour market deregulation processes. This new scenario of the Colombian economy leads us to question whether the larger exposure to international trade has significantly affected the employment dynamics in the manufacturing industry. Next section discusses the labour demand effects of international trade from a theoretical perspective.

2. LABOUR DEMAND EFFECTS OF INTERNATIONAL TRADE: ANALYTICAL FRAMEWORK

Rodrik (1997) conjectured that the globalization process would affect labour markets through two channels. One channel is the rise in the elasticity of labour demand with respect to wage changes —the wage elasticity effect—. The other one is the reduction on the demand for low-skilled labour which results in an inward shift in the demand curve for low-skilled labour —the level effect—. We next overview in detail both transmission mechanisms.

2.1. International Trade Effects on the Elasticity for Labour Demand

As explained in Hamermesh (1993), in a competitive setting, the long-run labour demand elasticity with respect to real wages (η_{LL}) is determined by the weighted average of two components: (i) the substitution elasticity between capital and labour (σ) and (ii) the product demand elasticity with respect to product price (in absolute terms) (η), so that:

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

Where s_L reflects the labour share over total output and acts as the weighting factor.

The first term on the right-hand side captures the constant-output elasticity (or substitution effect), reflecting the extent to which a firm substitutes away from labour when faced with an increase in its price, holding the level of output constant. The extent to which this is feasible depends on whether firms are labour or capital intensive. Note that capital intensive firms could more easily substitute capital for labour.

The second term of equation (1) captures the scale effect (or output effect), which represents the fall in labour demand due to output reduction holding production technology constant. Output may fall on account of increases in labour costs, which lead to higher output prices and therefore to lower sales.

For a given labour share (s_L), the substitution and scale effects are both negative. The smaller the labour share (s_L), the greater the relative importance of the substitution effect in determining the total labour demand elasticity (η_{LL}).

As suggested by Rodrik (1997) and further elaborated by Slaughter (2001), a larger exposure to international trade is expected to increase the elasticity of labour demand with respect to wage (η_{LL}). The labour demand elasticity increases due to a rise in the substitution elasticity between capital and labour (σ) and in the price elasticity of demand for products (η). On one side, the emergence of new phenomena such as outsourcing and offshoring enable firms to access to a larger variety of intermediate inputs and capital equipment, produced both domestically and abroad. This expands the set of productive factors with which a firm can substitute away from workers when faced an increase in domestic wages. This is the way in which free trade tends to facilitate the substitution between capital and labour. On the other side, openness to trade makes a country's product market more competitive. It is well known that trade policy liberalization –lower entry barriers– may force domestic firms to face heightened foreign competition, and thus increase the price elasticity of demand for products[1].

Extensive empirical evidence highlights a critical consequence of the globalization process, which is the downward impact it exerts on the labour income share (s_L), (see, for example, Petra, 2017; Elsby *et al.*, 2013; Judzik and Sala, 2013; and Böckerman and Marilanta, 2012). It is not as clear; however, how international trade affects the labour share. This effect probably depends, among other factors, on the capital substitution and the production technology. But even if the direction of change in s_L were known, the effect on η_{LL} would be determined by the relatives' sizes of σ and η (see equation (1)). For instance, Hijzen and Swaim (2010) have conjectured that there would be a fall in the labour income share (s_L), as consequence of offshoring. According to these authors, when an economy opens to international trade, the most labour-intensive activities offshore first, particularly in developed economies which are relatively well endowed in capital and skilled labour. As a result, offshoring is expected to lead to a reduction in the labour share. In this context, assuming that $\Delta\sigma > 0$ and $\Delta s_L < 0$, implies that free trade reinforces the substitution effect and tends to make labour demand more elastic. On the contrary, assuming that $\Delta\eta > 0$ and $\Delta s_L < 0$, free trade may weaken the scale effect and thus the extent to which higher wages pass through into higher prices. Since the net impact on the labour demand elasticity is the sum of offsetting substitution and scale effects; its sign is theoretically indeterminate. As consequence, the validity or Rodrik's conjecture of a more elastic labour demand must be determined empirically.

Lastly, we turn our attention to the existence of a causal relationship between the labour share and the labour demand elasticity in a context of a larger exposure to the international trade. On one side, Hamermesh's (1993) framework posits that the sensitivity of employment to wage changes (η_{LL}) is determined by a constant labour share (s_L). Furthermore, the standard macroeconomic analysis has predicted that if an economy opened to international trade, the labour share would remain constant. On the other side, some studies have shown that labour share has fallen as globalization has made the labour demand more elastic (Petra, 2017 and Elsby et al., 2013). This finding can be interpreted as empirical evidence uncovering the reverse causality between s_L and η_{LL}. In other words, it indicates that endogeneity problems might arise in the estimation process. Even though these empirical issues clearly deserve to be dealt with, they lie beyond the scope of this particular research.

[1] There is also extensive literature (e.g., Ferguson and Maurice, 1973 and Krishna et al., 2001) in which it is argued that trade liberalization processes often do not occur in a competitive context. Hence, the trade effects often cannot be analysed through Hamermesh's expression (1). For example, Ferguson and Maurice (1973) show that under monopoly and with unspecified cost structures, the relation between η_{LL} and σ still goes through, while the impact of increases in η on η_{LL} is theoretically ambiguous.

This is the reason why we follow the empirical strategy of previous literature (see, for example, Lewis and MacDonald, 2002 and Bruno et al., 2004), where the labour share (s_L) is assumed as a given parameter and not as an endogenous variable. Still, the fact that we divide our analysis in two crucial periods allows introducing some flexibility in the value of the labour share. Next section provides a detailed description of our empirical approach.

2.2. International Trade Effects on the Level of the Labour Demand

Rodrik (1997) in line with Hercksher-Olin-Samuelson factor endowments model predicted a reduction on the relative demand for low-skilled labour. This is the standard expected result for developed economies whose trade partners (mainly developing countries) have abundant unskilled labour. Developing countries will export low-skill-intensive products to the developed market and import high-skill-intensive goods in return, as long as exports in developing countries replace some domestic production in the developed country. This will result in a fall in the demand for low-skilled workers.

What is the counterpart for the relative demand for skilled-unskilled labour in developing countries? Arbache *et al*. (2004) pointed out that one of the consequences of increasing trade openness in developing economies is a rapid inflow of foreign technology as result of both direct investment and increased imports. In-flowing technology is mainly designed in industrialized economies which are relative skill intensive. Thus, the acquisition of new technologies from developing countries is normally accompanied by a greater demand for skilled labour. Although this hypothesis of skill-biased technological change and the subsequent increase in skilled-labour demand in developing countries has been strongly supported by growing empirical studies (e.g., Caselli, 2014, Conte and Vivarelli, 2011 and Gonzaga *et al*., 2006), no consensus has yet been reached on the net impact of technology on the total or aggregate demand for labour.

Autor *et al*. (2003) and Manning (2004), for instance, have argued that the hypothesis of skill-biased technological change provides a "too simplistic description" of the impact of technology on the labour demand. According to Autor *et al*. (2003), capital and technology -mainly machines- may substitute for human labour in tasks that can be routinized (automated). The critical point is that these tasks correspond mostly to semi-skilled jobs. In contrast, technology may be complementary to tasks in which cognitive and interactive skills are widely used. Therefore, the demand for semi-skilled labour would decrease while the demand for high and low skill labour would increase.

On this account, one of the contributions of this paper is to estimate the level effect of international trade on the total demand for labour. To conduct the analysis, we take as reference the model proposed by Greenaway *et al*. (1999), which has become a popular reference in recent aggregate data analysis (e.g., Seo *et al*., 2015 and Njikam, 2016). The main hypothesis of interest in Greenaway *et al*. (1999) is that international trade might work as a channel of technological spill-over effects through: (i) import goods embodying foreign knowledge, (ii) foreign direct investment and (iii) the acquisition of useful information –international trade provides channels of cross-border communication that facilitates learning of production and organizational methods and market conditions–. All these factors may contribute to heighten the technical efficiency of production, thus resulting in a reduction of employment. Seo *et al*. (2015), however, pointed out that in case that an excessive dependence on imported parts and components attenuate complementarity among domestic firms, productivity may deteriorate as the trade liberalization process deepens.

Overall, trade liberalization has been associated both with job creation and job destruction. The acquisition of foreign technology is not the unique driving factor. From a theoretically perspective, exports are expected to have a positive effect on employment. Firms produce more due to the higher levels of exports, thus increasing the demand for labour. In contrast, sectors exposed to a higher competition, in case they are not competitive enough, reduce their labour demand. In conclusion, the net impact of international trade on labour demand is theoretically ambiguous and depends on factors such as the level of development of the partner (developed/developing), the skill structure of jobs (low/high), the sectorial specialization (more or less high-tech industries) and the nature of imported and exported goods (final/intermediate).

3. EMPIRICAL IMPLEMENTATION

This section provides a detailed description of our empirical procedure. First, we present the two benchmark equations that will be used in the empirical analysis. Since these equations are well known and widely used in the literature, we just provide brief theoretical underpinnings to clarify their interpretation. The novelty of our analysis lies in the empirical methodology we use to decompose the labour demand elasticity. Second, we explain how we use these equations to empirically assess the labour demand effects of international trade, and disclose our strategy to approach quantitatively, in particular, the level effect of international trade.

3.1 Baseline Equations

To conduct our empirical analysis, we estimate an extended version of the two following standard employment equations:

$$n_t = \alpha_0 + \alpha_1 w_t + \alpha_2 k_t + u_t \tag{2}$$

$$n_t = \beta_0 + \beta_1 w_t + \beta_2 y_t + e_t \tag{3}$$

These specifications have two attractive features. First, the coefficients can be interpreted as elasticities. Second and foremost, they have a direct connection with Hamermesh's equation (1). To be specific, equation (2) allows us to obtain a direct estimate of the labour demand elasticity with respect to wages (η_{LL} in Hamermesh's expression 1) while equation (3) allows us to estimate the substitution elasticity between capital and labour (σ in Hamermesh's expression 1). Hence, as discussed by several studies, the decomposition of the labour demand elasticity in a substitution and a scale effect can be performed by exploiting the connection between Hamermesh's expression (1) and equations (2) and (3) –see, for example, Slaughter, 2001; Lewis and MacDonald, 2002; Hijzen and Swaim, 2010; and Sala and Trivín, 2012–. Before explaining in detail our empirical approach, let us review some simple theoretical underpinnings to these equations. This will allow us to convey an accurate interpretation of all coefficients.

3.1.1. Theoretical Background for Equation (2)

Karanassou *et al*. (2007) assume a competitive labour market containing a fixed number f of identical firms with symmetric production and cost conditions, and monopoly power in the product market. In this context, the i'th firm has a Cobb Douglas production function $q_{it}^s = A n_{it}^{\alpha} \overline{k}_{it}^{(1-\alpha)}$, where q_{it}^s is output supplied, n_{it} is employment, $\overline{k_{it}}$ is capital stock, α $(0 < \alpha < 1)$ is a parameter accounting for relative influence of capital and employment, and A is a positive constant which captures the technological change[2]. Each firm faces a product demand function $q_{it}^D = \left(\frac{P_{it}}{P_t}\right)^{-\varepsilon} \frac{y_t}{f}$, where P_{it} is the price charged by firm i, P_t is the aggregate price level, $\varepsilon > 0$ is the price elasticity of product demand, and y_t stands for aggregate output.

As each firm chooses its employment at the profit maximizing level (for a given capital stock). Then, the following aggregate labour demand can be obtained by solving the first order condition and aggregating across the firms.

[2] Karanassou *et al*. (2007) assume that the technological change grows at constant rates λ, thus it can be expressed as $A = A_0{}^{e^{\lambda}}$.

$$N_t = \left[\alpha A\left(1-\frac{1}{|\varepsilon|}\right)\right]^{\frac{1}{(1-\alpha)}}\left[\frac{W_t}{P_t}\right]^{\frac{1}{(1-\alpha)}} K_t \tag{4}$$

Where $N_t = n_{it} f$ is aggregate employment, $\frac{W_t}{P_t}$ is real wage, and $K_t = k_{it} f$ is aggregate capital. Taking natural logarithms, introducing a white noise error term $u_t \sim i.i.N\left(0,\sigma^2\right)$to capture supply and demand shocks, and rearranging the terms as follows:

$$n_t = \ln\left(N_t\right);\ w_t = \ln\left(\frac{W_t}{P_t}\right);\ k_t = \ln\left(K_t\right);\ \alpha_0 = \frac{1}{(1-\alpha)} ln\left[\alpha A\left(1-\frac{1}{|\varepsilon|}\right)\right];\ \alpha_1 = \frac{-1}{(1-\alpha)};$$
$$\alpha_2 = 1$$

We obtain equation (2) as the empirical counterpart of equation (4) –usually called unconditional or capital constrained labour demand–.

Note that all coefficients in equation (2) will have to be interpreted as delivering elasticities, where α_1 measures the sensitivity of employment to respect real wages and α_2 quantifies the impact of capital stock on labour demand.

3.1.2. Theoretical Background for Equation (3)

Lewis and MacDonald (2002) assume, as standard, that the economy is described by a CES production function of the form:

$$Y_t = A\left[\theta N_t^{\frac{\sigma-1}{\sigma}} + (1-\theta) K_t^{\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{(\sigma-1)}} \tag{5}$$

Where Y_t is output, A is a positive constant which captures the technological change, N_t is aggregate employment, K_t is aggregate capital stock, θ can be interpreted as the share parameter and σ as the elasticity of substitution between capital and labour. Profit maximization by firms in a competitive framework implies the following first-order condition with respect to labour (equation 3 in Lewis and MacDonald, 2002, p. 20):

$$A^{\frac{\sigma-1}{\sigma}}\theta\left[\frac{Y_t}{N_t}\right]^{\frac{1}{\sigma}} = \frac{W_t}{P_t} \tag{6}$$

Where $\frac{W_t}{P_t}$ is the real wage. This first order condition can be rewritten as:

$$N_t = A^{\sigma-1}\theta^{\sigma}\left[\frac{W_t}{P_t}\right]^{-\sigma} Y_t \tag{7}$$

Taking natural logarithms, adding a noise error term $e_t \sim i.i.N\left(0,\sigma^2\right)$ to capture supply and demand shocks, and making the following rearrangements:

$$n_t = \ln\left(N_t\right);\ w_t = \ln\left(\frac{W_t}{P_t}\right);\ y_t = \ln\left(Y_t\right);\ \beta_0 = \sigma\ln\left(A\theta\right)-\ln\left(A\right);\ \beta_1 = -\sigma;\ \beta_2 = 1$$

We obtain equation (3) as the empirical counterpart of equation (7).

Note that all coefficients in equation (3) will have to be interpreted as delivering elasticities, where β_1 must be interpreted as the substitution elasticity between capital and labour, while β_2 captures the influence of output on employment.

3.2. Decomposition of the Elasticity for Labour Demand

We use the empirical estimation of models (2) and (3) to perform a quantitative decomposition of the labour demand elasticity. The crucial coefficients are α_1 in model (2) and β_1 in model (3). The coefficient α_1 measures the total elasticity of employment with respect to real wage (η_{LL} in Hamermesh's expression 1). It thus measures the overall effect of a change in wages on the level of employment, which is the sum of the constant-output elasticity (or substitution effect) and the scale effect (or output effect). The coefficient β_1 captures the substitution effect resulting from the change in the relative factor prices (σ in Hamermesh's expression 1). When it is multiplied by the capital income share ($1-s_L$), the substitution effect becomes the constant-output elasticity: $(1-s_L)\,\beta_1$ ($-(1-s_L)\sigma$ in Hamermesh's expression 1).

As well explained by Sala and Trivín (2012), when wages rise, the relative price of labour is increased and there is an incentive to substitute labour by capital. The extent to which this is feasible (which depends on the technology) is measured by β_1 In other words, β_1 is the substitution effect because it measures the employment response to a wage change. When it is multiplied by the capital income share, it still measures the employment response to a wage change but *holding output constant*. These authors have pointed that the intuition behind this definition is that a 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).

The scale effect arises from the fact that higher costs (in our example wage has increased) yield a reduction in production. This is a direct consequence of the fact the labour demand derives from a production function. H*olding the production* *technology constant*, the more costly becomes a production factor, the less output will be supplied, and the less amount of labour will be needed.

As our empirical analysis allows us to obtain a direct estimate of the labour demand elasticity (α_1) and the substitution elasticity between capital and labour (β_1), the constant-output elasticity can be computed with ease. In addition to the estimated value of β_1, we use a constructed measure of the labour income share (s_L). In turn, as suggested by Sala and Trivín (2012), the difference between the total effect and the constant-output elasticity yields an empirical measure of the scale effect. The following equation (8) provides a simple illustration of our empirical approach.

$$\underset{[\alpha_1]}{\text{Total effect}} = \underset{[(1-s_L)\,\beta_1]}{\text{Constant-output elasticity}} + \underset{[\alpha_1-(1-s_L)\,\beta_1]}{\text{Scale effect}} \tag{8}$$

It is worth highlighting that one of the contributions of this paper lies in the methodology to empirically approach the scale effect. Most previous studies have not estimated the scale effect as the difference between the total and the substitution effects, but as the product of the labour share and the price elasticity of demand for product (see, for example, Russell and Tease, 1991; Lewis and MacDonald, 2002; and Bruno *et al.*, 2004). Although, from a theoretical perspective, we have taken as reference the same definition of the scale effect –recall that this effect is the product of the labour share and the price elasticity of demand for product, $s_L\,\eta$ in Hamermesh's expression (1)–, our empirical approach has two outstanding advantages over previous literature.

First, we have not assumed, as Lewis and MacDonald (2002), that the price elasticity of the demand for product η is unitary. This route has been quite criticized given that the product elasticity can take a wide range of values, especially across countries.

Second, since we calculate the scale effect as the difference between the total and the substitution effects, our empirical measure is not constrained to be determined by the term $s_L\,\eta$. On this account, our empirical approach is also consistent with other analytical frameworks, for example with those by Ferguson and Maurice (1973) and Krishna *et al.* (2001). These authors show that under monopoly, and with unspecified cost structures, the scale effect cannot be explicitly stated in terms of η.

Finally, our empirical approach has an extra advantage, which is related to the empirical computation of the substitution effect. In contrast to several studies, we do not tend to interpret the coefficient β_1 as the constant-output elasticity. Lewis and MacDonald (2002) have remarked that this interpretation is a common mistake in a large part of the literature in the field (for example in Slaughter, 2001). As shown above, the model (3) is not a demand for labour function but a marginal productivity condition. Therefore, the coefficient β_1 represents the elasticity of substitution between capital and labour. Only when it is multiplied by the capital income share $(1-s_L)$, the substitution elasticity becomes the constant-output elasticity $(1-s_L)\,\beta_1$.

3.3. Capturing the International Trade Effects on the Level of the Labour Demand

As discussed before, the empirical decomposition of the total labour demand elasticity is the main contribution of this paper. The second contribution is the estimation of the aggregate shift in the total labour demand as consequence of a larger exposure to international trade –the level effect–. Next, we focus on describing the empirical strategy to compute this level effect. To be specific, following Greenaway *et al*. (1999), Seo *et al*. (2015) and Njikam (2016), we extend the equations (2) and (3) by adding a measure of trade openness, open (in logs).

The hypothesis of these authors is that international trade might work as a channel of technological spill-over effects through the import of goods embodying foreign knowledge, foreign direct investment, and the acquisition of useful information. These three factors may contribute to heighten the technical efficiency of the production process, thereby causing a change in the demand for labour.

Furthermore, as shown by Greenaway *et al*. (1999), if technological change is assumed to be correlated with trade changes in the firm's profit maximization problem, the labour demand function can be obtained, as standard, by solving the first order conditions. In this setting, however, the labour demand function includes a specific term that captures the trade effect on technical efficiency and acts as a labour demand shifter.

Therefore, if we assume, in the theoretical equations (4) and (7), that the efficiency parameter A , among other elements, is determined by international trade changes, then we can obtain the following extended equations (see Appendix 1 for details)

$$n_t = \alpha_0 + \alpha_1 w_t + \alpha_2 k_t + \alpha_3 open_t + u_t \tag{9}$$

$$n_t = \beta_0 + \beta_1 w_t + \beta_2 y_t + \beta_3 open_t + e_t \tag{10}$$

The coefficient α_3 in equation (9) can be interpreted as the international trade effect on the level of the labour demand. In contrast, the coefficient β_3 in equation (10) only captures a part of the shift in the technical efficiency of the production process. The critical point is that this equation is not a demand for labour function but a marginal productivity condition. Thereby, the coefficient β_3 cannot be interpreted as the international trade effect on the level of labour demand. It is affected by the substitution elasticity between capital and labour (see Appendix 1, extensions for equations (2) and (3), for details). In any case, if exposure to international trade competition improves technical efficiency, the coefficients α_3 and β_3 are expected to be positive. On the contrary, if excessive dependence on imported parts and components tend to attenuate the complementarity among domestic firms, the productivity is likely to deteriorate as the globalization process deepens. In that case, the coefficients α_3 and β_3 are expected to be negative.

3.4. Additional Considerations

To analyse the Colombian case, equations (9) and (10) are extended in two directions.

First, due to the relevance of adjustment costs in labour demand decisions, we consider the addition of the first lag of the explanatory variable. This enables us to perform a dynamic analysis and to compute short and long-run effects of each explanatory variable on the labour demand.

Second, because the nineties were characterized by a deindustrialization and a growing expansion of the service sector, we include a control variable which captures the growth dynamics of the commercial and services sector (Δser). These empirical adjustments leave us with the two following extended equations:

$$n_t = \alpha_0 + \lambda n_{t\text{-}1} + \alpha_1 w_t + \alpha_2 k_t + \alpha_3 open_t + \alpha_4 \Delta ser_t + u_t \tag{11}$$

$$n_t = \beta_0 + \lambda n_{t\text{-}1} + \beta_1 w_t + \beta_2 y_t + \beta_3 open_t + \beta_4 \Delta ser_t + e_t \tag{12}$$

Where λ represents the inertial or persistence coefficient, while α_4 and β_4 capture the influence of the expansion of the services sector on the dependent variable n_t. In this way, we control for sectoral composition effects on total employment.

4. DATA AND EMPIRICAL MODELLING

4.1. Data

We use a panel database with a cross-section dimension of sectors and a time dimension of $T = 42$ years covering the period 1974-2015[3]. Table 1 lists the variables and the corresponding sources.

[3] The detailed list of sectors is provided in List 1, in the Appendix 2.

Information on employment, labour compensation, net capital stock, and value added is taken from the Annual Manufacturing Survey (Encuesta Anual Manufacturera, EAM), which is produced by the National Administrative Department of Statistics (Departamento Administrativo Nacional de Estadística, DANE). It is worth noting that labour compensation is computed as total labour compensation in sector i (wages and salaries before taxes and employers' social security contributions) over total employment in that sector, where total employment only includes workers directly paid by the firm, either permanent or temporary workers (i.e., agency workers are excluded). In turn, our constructed measure of the labour income share is computed as the ratio of total labour compensation in sector i over real value added in that sector. All nominal variables are deflated with the manufacturing price index.

Table 1. Definitions of variables

	Variables	*Sources*	*Subindexes*
η_{it}	Employment	(1)	i = 1,…,16 sectors
ω_{it}	Average real labour compensation	(1)	t = 1,…,42 years
k_{it}	Net real capital stock	(1)	
y_{it}	Real gross value added	(1)	
s_{it}	Labour share $\left[\frac{w_{it} * \eta_{it}}{y_{it}}\right]$	(1)	
op_{it}	Trade openness $\left[\frac{(exports_{it} + imports_{it})}{output_{it}}\right]$	(2)	
m_{it}	Import penetration $\left[\frac{imports_{it}}{output_{it} - exports_{it} + imports_{it}}\right]$	(2)	
x_{it}	Export ratio $\left[\frac{exports_{it}}{output_{it}}\right]$	(2)	
ser_t	Service sector share	(3)	
d_{92}	Dummy: value 1 1992 onwards		
d_{8083}	Dummy: value 1 in 1980-1983		
d_{9700}	Dummy: value 1 in 1997-200		
d_{0809}	Dummy: value 1 in 2008-2009		

Note: All nominal variables are deflated by the manufacturing price index (base: June 1999).
All variables are expressed in logs in the estimation process. (1) EAM; (2) DNP;.(3) DANE.
Source: own elaboration.

To capture the degree of international trade openness, we use three alternative measures: the trade openness index, the import penetration ratio, and the export ratio. Information on these variables is obtained from the National Planning Department (D*epartamento* N*acional de* P*laneación*, DNP in its Spanish initials). The trade openness index is calculated as the ratio of total trade (the sum of exports plus imports) to the output in that sector. The import penetration ratio is computed as the value of imports divided by the value of apparent consumption or domestic output (output minus exports plus imports). The export ratio, on the other hand, is the export share over output.

To control for the expansion of the services sector, we use the ratio of value added in the service sector over total value added (in differences), this information is taken from DANE.

As shown in Table 1 all the variables have available data for all 16 sectors, except for the services sector share. Thus, our dataset provides detailed homogeneous time series information of the Colombian manufacturing industry across sectors.

Finally, we include a set of time dummies to control mainly for macroeconomic shocks that may affect all sectors. In this way, d_{92} help us to capture potential structural breaks in both the elasticity for labour demand and the elasticity of substitution arising from the trade and institutional reform process; d_{8083} accounts for the impact of the international debt crisis experienced by Latin America in the early eighties; d_{9700} and d_{0809} checks whether the international financial crisis at the end of the nineties and the Great Recession did also affect employment in Colombia.

4.2. Econometric Methodology

As we work with a two-dimensional panel data, we add sectorial fixed effects to control for unobserved heterogeneity among sectors. Thus, we will estimate one-way fixed effects models instead of two-way fixed effects. We do not include time fixed effects to control for temporal shocks that may affect all sectors, because the data for the control variable (ser_t) is common across sectors. Nonetheless, the main macroeconomic shocks that could affect employment are, indeed being captured through the time dummies discussed in data section.

Given the dynamic nature of the extended models, equations (11) and (12) will be estimated as partial adjustment models taking the following general form:

$$n_{it} = \gamma_i + \lambda n_{it-1} + \theta Z_{it} + v_{it} \qquad v_{it} \sim i.i.N\left(0, \sigma^2\right) \tag{13}$$

Where the subscripts i and t are sector and time indices, respectively; γ_i is a sectorial cross-section intercept; $n_{it\text{-}1}$ is the lagged dependent variable with λ as inertial (or persistence) coefficient; Z is a vector of explanatory and control

Semestre Económico, 25 (58) • enero-junio de 2022 • pp. 1-48 • ISSN (en línea): 2248-4345

variables with θ as the set of estimated parameters capturing their influence on the dependent variable n_{it}; and v_{it} is a stochastic perturbation.

4.2.1. Stationarity and Unit Root Tests

As we deal with a dynamic panel, we must ensure that a long-run equilibrium relationship exists among the variables considered. This implies testing that all variables are stationary I(0) which, by definition, yields a long-run cointegrating vector.

To check the order of integration of the variables, we carry out a set of stationery and unit root tests depending on the type of the variables to be dealt with. We use the test proposed by Maddala and Wu (1999) –MW henceforth– for the variables that are sector-specific (most of them). The MW test is a panel unit root test based on Fisher's (1932) results and it assumes, under the null hypothesis, that all panels contain unit roots, against the alternative that at least one panel is stationary. For the control variable (Δser_t) which is common across sectors, we use three standard tests: the Kwiatkowski et al. (1992) stationarity test –KPSS–; the Augmented Dickey-Fuller (1979) unit-root test –ADF–; and the Phillips-Perron (1988) unit-root test –PP–.

It is worth noting that we conduct the MW test, because, in general, panel unit root tests have higher power than unit root tests when applied to individual time series. Moreover, this test has two attractive characteristics. First, it does not restrict the autoregressive parameter to be homogeneous across sectors under the alternative of stationarity. Second, the choice of the lag length and the inclusion of a time trend in individual ADF test regressions can be determined separately for each sector. An important limitation of the MW test is the assumption of error cross-sectional independence. This assumption is quite restrictive, as the errors in macro panels often exhibit significant correlation among the different cross-section units[4]. To mitigate the impact of this error cross-sectional dependence, we follow the procedure suggested by Levin, *et al*. (2002) in which before performing the test, the mean of the series across panels is subtracted from the series.

We conduct the ADF and PP tests in order to test the null hypothesis of a unit root series as done in MW test. However, as discussed by Jafari *et al*. (2012) the PP and ADF unit root tests have a low power to reject the null hypothesis. Thus, the authors suggest using the KPSS test to deal with this problem. This is the reason why we also conduct the KPSS stationarity test. In case have mixed results, we will rely on the KPSS results.

[4] Using Monte Carlo simulations, Maddala and Wu (1999) conclude that this problem is less severe with the Fisher (1932) test than with other panel unit root tests such as the Levin and Lin test (1993) –LL– or the Im, Pesaran and Shin test (2003) –IPS–.

Table 2 shows the results of MW tests for the seven variables which are sector specific. It is straightforward to note that, the null hypothesis of an unit root can be rejected at the 1 % significance level for all variables (i.e. in all variables, at least one panel is stationary I(0)). In turn, Table 3 displays the results of the KPSS, ADF and PP test for the control variable (Δser_t). In the KPSS test, the null hypothesis of a stationary time series cannot be rejected at the 5 % significance level, while in both ADF and PP test the null hypothesis of a unit root can be rejected at the 5 % significance level.

The overall conclusion drawn from all these tests is that all variables are stationary. Hence, we have enough statistical evidence in favour of proceeding with stationary panel data estimation techniques.

Table 2. Panel Unit Root Test, 1974-2015

	η_{it}	ω_{it}	k_{it}	y_{it}	op_{it}	m_{it}	x_{it}
MW	122.89	184.99	117.55	122.89	104.85	96.54	119.86
Result	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
	$I(0)$	$I(0)$	$I(0)$	$I(0)$	$I(0)$	$I(0)$	$I(0)$

Notes: All variables are expressed in logs; MW tests computed using drift and removing cross-sectional means P-values in brackets.

Source: own elaboration.

Table 3. Stationary and Unit Root Tests, 1974-2015

	Δser_t		Δser_t		Δser_t
KPSS	0.11	ADF	-5.26	PP	-5.26
Result	[0.146]	Result	[-3.648]	Result	[-3.648]
	$I(0)$		$I(0)$		$I(0)$

Notes: variable is expressed in logs; DF and PP tests computed using drifts. 5 % Critical values in brackets.

Source: own elaboration.

4.2.2. Econometric Method

Models (11) and (12) are estimated by applying Ordinary Least Square (OLS) and one-way Fixed Effects (FE) (see Tables 4 and 5). In doing so, we need to take care of potential endogeneity problems caused by the introduction of lagged dependent variables in the set of regressors, as well as the well-known simultaneity between employment, real wages, value added and the net capital stock.

As explained by Nickell (1981), when lags of the dependent variable are included as regressors, as we do in models (11) and (12); the OLS estimator of the persistence coefficient will be upward biased, while the fixed effects estimator will be downward biased. However, Nickell (1981) pointed out that when T is large and N is small ($T>N$), the bias of the fixed effect estimator is likely to be insignificant. In contrast, Judson and Owen (1999) found that even with a time dimension as large as 3.0, this estimator would be biased downwards and inconsistent even in the absence of serial correlation in the error term. Although this may not be a critical problem in our analysis (we have $T = 42$ and $N = 16$), we cannot fully exclude the possibility existence of a bias in the persistence coefficients. We have hence estimated different versions of the Least Squares Dummy Variables Corrected (LSDVC) using Bruno's (2005) approximation to correct for the finite-sample bias. The corresponding results, presented in the Appendix 3 (Table 8), show that these sets of estimates do not differ significantly from the FE estimates. This allows us to conclude that the FE estimator is potent in our case[5].

On the other hand, to deal with the potential endogeneity of real wages, output, and net capital stock, we estimate FE by two stages least squares (FE-TSLS) using the real minimum wage and its first lag as instruments for real labour costs, and the energy consumption as instruments for output and net capital[6]. To confirm the appropriateness of these instruments in the different specifications of the (11) and (12), we rely on the performance of three tests. First, an LM test checking for under-identification (i.e., that the excluded instruments are not relevant, meaning non correlated with the endogenous regressors). This is denoted as U in Table 4 and 5, and the null hypothesis is that the equation is underidentified (against the alternative that the model is identified). Second, an F test checking for weak instruments (denoted as W), where the null hypothesis is that the instruments are correlated with the endogenous regressors, but only weakly[7]. Third, the Hansen (2001) test of overidentifying restrictions (denoted as H), in which the joint null hypothesis is, on one side, that the considered instruments are valid (i.e., uncorrelated with the error

[5] Nickell (1981) stand out that in small macro-panels (e.g., T around 30 and around 20) the System GMM estimator (Blundell and Bond, 1998) would not yield dramatic consistency gains over the FE estimator. The reason is that the consistency of this estimator depends on the fact that $N \rightarrow \infty$ grows sufficiently fast relative to T.

[6] It is important to note that we have estimated a wide set of specifications using different combinations of instruments depending on: (i) Whether we consider capital stock and value added either as weakly exogenous variables or endogenous; (ii) whether we use the ratio of non-wage labour costs over total labour compensation as instrument for real wages instead of using the first lag of the real minimum wage; (iii) whether we use the two first lags of value added and capital stock instead of the energy consumption as instruments for both capital and value added. Nevertheless, we only present those specifications with the best performance in the instrumental tests.

[7] Note that for the F-test, we show the standard critical values of Stock and Yogo (2005) at 10 % maximal IV relative bias.

term) and, on the other side, that the excluded instruments are correctly dropped from the estimated equation.

5. RESULTS

This section presents the empirical results of our study in three subsections. First, we present and discuss the estimates of models (11) and (12). Then we present and analyse the estimated value of the long-run labour demand elasticity with respect to wages and place attention on its stability, or not, between 1974-1991 and 1992-2015. We also compute the two channels through which the Colombian reform process could have affected this elasticity and compute the substitution and scale effects. Finally, we disclose the level effect of a higher exposure to international trade on the demand for labour. To capture this effect, we use the three alternative measures for the degree of trade openness discussed in data section –a trade openness index, an import penetration ratio, and an export ratio–.

5.1. Estimates

Tables 4 and 5 display, respectively, the estimated labour demand models (11) and (12) obtained through the various estimation methods just discussed: OLS in Columns (1), (4) and (7); FE in Columns (2), (5) and (8); and FE-TSLS in Columns (3), (6) and (9). Note that in both tables the information is classified into three blocks depending on which measure for the degree of international trade exposure is used in the estimation process, the one on the left-hand side corresponding to the trade openness index, the one on the intermediate side to the import penetration ratio, and the one on the left-hand side to the export ratio.

When examining the econometric analysis, if we have to favour some particular specifications, we would choose those obtained through the FE-TSLS estimator (Columns (3), (6) and (9) in Tables 4 and 5 for a three-fold reason: (i) instrumental variables are used to deal with potential endogeneity problems; (ii) the performance of the instrumental variable tests confirm the appropriateness of the instruments in all cases (since they are simultaneously exogenous and highly correlated with the endogenous regressors); and (iii) most explanatory variables are highly significant and take the expected sign according to the underlying theoretical relationships.

In any case, the six sets of estimates (i.e., those presented in Tables 4 and 5, Columns (3), (6) and (9)) provide a similar picture: high employment persistence, as widely tested by previous Colombian literature (e.g., Arango and Rojas, 2004; and Cárdenas and Bernal, 2004); significant but low short-run effects of real wages, output, and net capital stock; and large and highly significant effects of changes in the services sector. In contrast, the role of international trade exposure –measured either as trade openness or import penetration– is found to be non-significant in models (11) and (12), while the evidence on the effect of the export share is mixed. The export share is found to be significant in model (11) but non-significant in model (12).

It is worth noting that in these empirical models we have interacted the dummy variable d_{92} with the lagged employment and wage variables. These interactions allow us to test whether the elasticity for labour demand and the substitution elasticity have increased since 1992. The critical point is that, in the long-run, both elasticities are the joint outcome of the short-run effect of wage changes and the persistence coefficient. Our first hypothesis is that the liberalization process and the widespread use of more flexible employment methods such as short-term contracts and temporary workers, may have increased the short-run employment effects of wage changes. The second one is that the use of more flexible employment methods and the lower firing, training and recruitment costs could have made the demand for labour more flexible and allow faster adjustments. This should be reflected in a reduction in the persistence coefficient.

The results displayed in Tables 4 and 5, Columns (3), (6) and (9), however, show larger short-run sensitivities but also larger employment persistence[8]. We have a twofold explanation for this larger persistence, which is to be associated with the increases in social security revenues in the early nineties and the larger degree of labour market segmentation.

Although the objective of increasing payroll taxation was to expand the coverage of health and pension services, this measure may have reduced the speed of adjustment of labour demand (i.e., a larger persistence coefficient). This is one of the expected collateral employment effects of increasing non-wage costs. Our hypothesis is therefore that the increase in the estimated persistence coefficients captures the offsetting effect of reducing the firing, training and recruitment costs and simultaneously increasing payroll taxes.

On the other hand, as shown in Figure 4a, since 1974 up to 1991, the share of agency workers over total employment was less than 8 % and this share sharply increased due to the boom of outsourcing jobs, reaching a peak of 29 % in 2007. Our conjecture is that firms began to demand labour through outsourcing schemes as a mechanism to skip from the high non-wage labour costs (the share of non-wage labour costs was 46 % in 1992 and declined to 37 % in 2015, see Figure 4b), and as a mechanism to

[8] There is national literature which provides empirical support that the industrial manufacture employment in Colombia is still highly persistent after the nineties. (e.g., Medina *et al.*, 2012; and Cárdenas and Bernal, 2004)

deal with unexpected changes in factor prices. Hence, the demand for agency workers expanded and became more flexible while the demand for workers directly paid by the firm reduced and became more rigid. Since models (11) and (12) are estimated by using only workers directly paid by firms, we interpretate the significant increase in the persistence coefficient as an empirical evidence of a greater rigidity. In any case, we interpret this finding as a shortcoming of the measures undertaken during the nineties to make Colombian labour market more flexible.

Regarding the macroeconomic shocks controlled by the additive dummies (d_{8083}, d_{9700} and d_{0809}), it is found that the debt crisis at the early eighties had a mildly significant negative influence on employment. Likewise, the impact of the international financial crisis at the end of the nineties is also found to be negative but highly significant, while the results for the Great Recession crisis suggest that the last crisis did not have a significant impact on the labour demand.

Table 4. Estimates of model 11

Dependent variable: η_{it}									
	op_{it}			m_{it}			x_{it}		
	OLS (1)	FE (2)	FE-TLSLS (3)	OLS (4)	FE (5)	FE-TLSLS (6)	OLS (7)	FE (8)	FE-TLSLS (9)
η_{it-1}	0.99	0.92	0.87	0.99	0.91	0.87	0.99	0.91	0.84
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
ω_{it}	-0.03	-0.07	-0.14	-0.03	-0.07	-0.14	-0.03	-0.07	-0.17
	[0.058]	[0.030]	[0.005]	[0.061]	[0.038]	[0.006]	[0.042]	[0.012]	[0.002]
k_{it}	0.01	0.02	0.08	0.01	0.02	0.08	0.01	0.02	0.10
	[0.072]	[0.029]	[0.003]	[0.062]	[0.035]	[0.004]	[0.072]	[0.029]	[0.003]
$open_{it}$	0.00	-0.01	-0.01	0.00	-0.01	-0.01	-0.01	-0.01	-0.02
	[0.604]	[0.308]	[0.214]	[0.726]	[0.259]	[0.143]	[0.147]	[0.048]	[0.003]
Δser_t	-2.39	-1.98	-2.05	-2.38	-1.96	-2.04	-2.38	-1.99	-2.03
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
$\eta_{it-1}{*}d_{92}$	0.01	0.01	0.02	0.01	0.01	0.02	0.01	0.01	0.02
	[0.493]	[0.007]	[0.025]	[0.600]	[0.001]	[0.030]	[0.470]	[0.030]	[0.008]
$w_{it}{*}d_{92}$	-0.01	-0.01	-0.03	-0.01	-0.01	-0.03	-0.01	-0.01	-0.04
	[0.442]	[0.005]	[0.015]	[0.453]	[0.002]	[0.019]	[0.492]	[0.027]	[0.006]
d_{8083}	-0.03	-0.03	-0.02	-0.03	-0.03	-0.02	-0.03	-0.03	-0.02
	[0.001]	[0.004]	[0.060]	[0.001]	[0.004]	[0.070]	[0.001]	[0.001]	[0.043]
d_{9700}	-0.05	-0.06	-0.05	-0.05	-0.06	-0.05	-0.05	-0.05	-0.05
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
d_{0809}	0.02	0.02	0.02	0.02	0.02	0.02	0.03	0.02	0.02
	[0.117]	[0.168]	[0.262]	[0.115]	[0.165]	[0.267]	[0.092]	[0.145]	[0.190]
c	0.16	1.15		0.17	1.17		0.20	1.18	
	[0.076]	[0.003]		[0.071]	[0.001]		[0.038]	[0.003]	

Dependent variable: η_{it}									
	op_{it}			m_{it}			x_{it}		
	OLS (1)	FE (2)	FE-TLSLS (3)	OLS (4)	FE (5)	FE-TLSLS (6)	OLS (7)	FE (8)	FE-TLSLS (9)
Obs.	654	654	654	654	654	654	654	654	654
$Adj.R^2$	0.99	0.89	0.87	0.99	0.89	0.87	0.99	0.89	0.86
U			22.37 [0.000]			22.99 [0.000]			16.45 [0.001]
W			9.31 (6.61)			9.30 (6.61)			8.65 (6.61)
H			0.04 [0.998]			0.09 [0.996]			0.68 [0.714]

Notes: All variables are expressed in logs. P-values in brackets. OLS, Ordinary Least Square. FE, Fixed effects. FE-TSLS, Fixed effects using Two Step Least Squares. U, Under identification test. W, Weak identification test. Stock and Yogo (2005) weak ID test critical value at 10 % maximal IV relative bias in parentheses. H, Hansen (2001) test.

Source: own elaboration.

Table 5. Estimates of model 12

Dependent variable: η_{it}									
	op_{it}			m_{it}			x_{it}		
	OLS (1)	FE (2)	FE-TLSLS (3)	OLS (4)	FE (5)	FE-TLSLS (6)	OLS (7)	FE (8)	FE-TLSLS (9)
$\eta_{it\text{-}1}$	0.95 [0.000]	0.90 [0.000]	0.85 [0.000]	0.96 [0.000]	0.90 [0.000]	0.84 [0.000]	0.96 [0.000]	0.89 [0.000]	0.83 [0.000]
ω_{it}	-0.06 [0.000]	-0.08 [0.000]	-0.17 [0.002]	-0.06 [0.000]	-0.08 [0.000]	-0.17 [0.003]	-0.06 [0.000]	-0.08 [0.000]	-0.18 [0.001]
y_{it}	0.04 [0.000]	0.05 [0.004]	0.11 [0.001]	0.04 [0.000]	0.05 [0.003]	0.12 [0.002]	0.04 [0.000]	0.05 [0.009]	0.12 [0.001]
$open_{it}$	0.01 [0.105]	0.00 [0.909]	0.01 [0.216]	0.00 [0.225]	-0.00 [0.582]	0.00 [0.546]	-0.00 [0.278]	-0.01 [0.254]	-0.00 [0.372]
Δser_t	-2.14 [0.000]	-1.84 [0.000]	-1.54 [0.000]	-2.14 [0.000]	-1.83 [0.000]	-1.52 [0.001]	-2.16 [0.000]	-1.84 [0.000]	-1.47 [0.001]
$\eta_{it\text{-}1}{*}d_{92}$	0.01 [0.057]	0.02 [0.000]	0.03 [0.010]	0.01 [0.066]	0.02 [0.000]	0.03 [0.012]	0.01 [0.057]	0.02 [0.000]	0.04 [0.003]
$w_{it}{*}d_{92}$	-0.01 [0.056]	-0.02 [0.000]	-0.03 [0.012]	-0.01 [0.067]	-0.02 [0.000]	-0.04 [0.015]	-0.01 [0.075]	-0.02 [0.000]	-0.04 [0.005]
d_{8083}	-0.02 [0.003]	-0.02 [0.002]	-0.02 [0.094]	-0.03 [0.002]	-0.02 [0.001]	-0.02 [0.122]	-0.03 [0.002]	-0.02 [0.000]	-0.01 [0.187]
d_{9700}	-0.04 [0.000]	-0.04 [0.000]	-0.03 [0.013]	-0.04 [0.000]	-0.04 [0.000]	-0.03 [0.019]	-0.04 [0.000]	-0.04 [0.000]	-0.03 [0.023]
d_{0809}	0.02 [0.208]	0.02 [0.272	0.01 [0.522]	0.02 [0.199]	0.02 [0.250]	0.01 [0.508]	0.02 [0.164]	0.02 [0.225]	0.01 [0.490]

Dependent variable: η_{it}									
	op_{it}			m_{it}			x_{it}		
	OLS (1)	FE (2)	FE-TLSLS (3)	OLS (4)	FE (5)	FE-TLSLS (6)	OLS (7)	FE (8)	FE-TLSLS (9)
c	0.19 [0.031]	0.88 [0.021]		0.21 [0.018]	0.92 [0.013]		0.23 [0.010]	0.94 [0.021]	
Obs.	654	654	654	654	654	654	654	654	654
$Adj.R^2$	0.99	0.90	0.88	0.99	0.90	0.88	0.99	0.90	0.87
U			21.13 [0.000]			16.52 [0.001]			19.78 [0.000]
W			8.42 (6.61)			7.90 (6.61)			8.51 (6.61)
H			0.06 [0.970]			0.03 [0.937]			0.07 [0.964]

Notes: All variables are expressed in logs. P-values in brackets. OLS, Ordinary Least Square. FE, Fixed effects. FE-TSLS, Fixed effects using Two Step Least Squares. U, Under identification test. W, Weak identification test. Stock and Yogo (2005) weak ID test critical value at 10% maximal IV relative bias in parentheses. H, Hansen (2001) test.

Source: own elaboration.

5.2. International trade effects on the elasticity for labour demand

Table 6 presents the implied long-run elasticities arising from the base-run estimates displayed in Tables 4 and 5. Columns (3), (6) and (9). The information is classified in two blocks, the one on the left-hand side corresponds to the estimation for the first period (1974-1991) and the one on the right-hand side to the second period (1992-2015). The three different specifications in each block correspond to the inclusion of the three alternative measures used to control for the exposure to international trade, so that the estimates that incorporate the index of trade openness are presented in Columns (1) and (4); those that use the import penetration ratio, in Columns (2) and (5); and estimates that introduce the export ratio, in Columns (3) and (6). We next overview the findings for the two relevant periods of analysis, 1974-1991 and 1992-2015.

Looking at Table 6, Columns (1), (2) and (3), the estimated value for the labour demand elasticity during the first period (1974-1991) is around -1.05 (quite robust across specifications), and is tightly close to the substitution elasticity between capital and labour whose estimated value is around -1.08. It is worth highlighting that our result of -1.05 for the labour demand elasticity is consistent with a large part of the empirical literature for Colombia, which places this long-run elasticity for the manufacturing industry, between 1974 and 1991, in the range of -0.42 and -2.77 (see, for a detailed description, Isaza and Meza, 2004).

Given our results of -1.05 for the labour demand elasticity and -1.08 for the substitution elasticity, and taking into account that, for the 16 sectors in which the Colombian industry is disaggregated, the labour income share is 36.87 % on average; we conclude that the resulting constant-output elasticity can safely be placed around -0.68, while the scale effect can be placed around -0.38. That is, a 1 % increase in the real labour cost will cause a 1.05 % reduction in employment. This reduction can be attributed to less than two thirds (-0.68 of -1.05) to the substitution effect and in more than one third (-0.38 of -1.05) to the scale effect.

For the second period (1992-2015), the results are displayed in Columns (4), (5) and (6). All the estimated values are also very similar across specifications. For example, the estimated total effect ranges from -1.51 to -1.59 and the estimated substitution elasticity lies in the narrow interval between -1.64 and -1.68. In addition, the substitution effect is placed around -1.18 while the scale effect is placed around -0.37.

Table 6. Long-run wage elasticities for labour demand

	1974-1991			1992-2015		
	(1)	(2)	(3)	(4)	(5)	(6)
	Total effect and substitution elasticity					
η_{LL}	-1.07	-1.03	-1.06	-1.58	-1.51	-1.59
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
σ	-1.11	-1.08	-1.04	-1.68	-1.68	-1.64
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
	Total effect decomposition					
s_L	36.87	36.87	36.87	29.03	29.03	29.03
$-(1-s_L)\sigma$	-0.70	-0.68	-0.65	-1.19	-1.20	-1.16
$-s_L\eta$	-0.37	-0.35	-0.41	-0.38	-0.32	-0.42
	Robustness check					
	1974-1991			1992-2015		
	(1)	(2)	(3)	(4)	(5)	(6)
	Total effect and substitution elasticity					
η_{LL}	-1.20	-1.18	-1.16	-1.69	-1.69	-1.69
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
σ	-1.30	-1.32	-1.20	-1.94	-2.06	-1.88
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
	Total effect decomposition					
s_L	39.82	39.82	39.82	36.04	36.04	36.04
$-(1-s_L)\sigma$	-0.78	-0.79	-0.72	-1.24	-1.32	-1.21
$-s_L\eta$	-0.41	-0.39	-0.44	-0.45	-0.37	-0.48

Notes: P-values in brackets.
Source: own elaboration.

The comparison of our results for both periods 1974-1991 and 1992-2015 yields two salient findings. The first main finding is a rise in the substitution elasticity between capital and labour, which rises from –1.08 to –1.67. This finding is consistent with Rodrik's conjecture (1997) according to which a main consequence of the globalization process is the greater ease with which domestic workers can be substituted by capital, either through outsourcing, offshoring or foreign investment. Nevertheless, we interpret that the growing exposure to international trade in Colombia is not the only one factor driving this structural change. The widespread use of temporary contracts and the lower firing, training and recruitment costs may also have extended the ability of firms to substitute work for capital (and vice versa). Our interpretation is based on the results of previous literature, which give support to the hypothesis that relaxing hiring and firing regulations facilitates employment substitution possibilities in response to changes in factor prices (see Hasan *et al.*, 2007, and Saha *et al.*, 2013, for the Indian case and Hijzen and Swaim, 2010, for OECD countries).

The second salient finding is an increase in the magnitude of the elasticity of the demand for labour (or total effect) ranging from -1.05 to -1.56, on average. This increase of 0.51 percentage points in the labour demand elasticity is a consequence only of a rise in the substitution effect. The substitution effect almost doubles its size, rising from -0.68 to -1.19 and now accounts for three quarters of the total effect. The scale effect, in turn, remains stable around -0.38, and accounts for a quarter. Note that the rise in the substitution effect arises from the fall of 7.84 percentage points in the labour income share and from the increase in the substitution elasticity. In turn, the stability of the scale effect may be interpreted as a net effect of the interplay between the decline in the labour income share and the potential increase in the price elasticity of the demand for products. If we use Hamermesh's expression (1), the corresponding product demand elasticities can be straightforwardly computed as the scale effect over the labour income share. In doing so, we confirm that the product demand elasticity increased, going from 1.03 to 1.28. This change is probably a consequence of the heightening foreign competition that firms have faced since 1992, when free trade began to consolidate.

5.2.1. Robustness Check

An important issue for our empirical analysis is the role played by the labour income share in determining the magnitudes of the scale and substitution effects. The smaller the labour share, the greater the relative importance of the substitution effect in determining the total labour demand elasticity. Although our results show that this effect turns out to be quantitatively dominant, at least for the Colombian manufacturing industry, we need to take care of potential measurement errors when calculating the labour income share and estimating models (11) and (12). This concern requires a detailed discussion as follows.

First, as described above, we have computed an average labour income share of 36.87 % and 29.03 % for the two relevant periods of analysis, 1974-1991 and 1992-2015, respectively. These magnitudes, however, are relatively low compared to labour shares around 61.36 % and 59.78 % for advanced economies whose industries have more capital-intensive sectors (e.g., 61.35 % and 59.16 % for United States, as shown in IMF, 2017 and ILO 2015). These unexpected larger differences are mainly due to data limitations. As it is well-known, Latin American countries have data constraints which do not allow computing a "true" or adjusted share of labour income. The key problem is that the labour share measures do not encompass the labour compensation of informal employment (including self-employment), which accounts for the major fraction of total employment in developing countries (around 50 % for Colombia). The difficulty arises from the fact that most prevailing forms of self-employment in developing countries take place in micro and small enterprises whose economic activities are difficult to capture. Hence, the labour income shares cannot be adjusted for self-employment as suggested by international institutions such as the IMF or the ILO. Indeed, this is our case. As our database does not include information about informal employment or self-employment, we have not been able to compute the adjusted labour income share. This is the main reason why we have used the unadjusted labour income share.

The second crucial point is that, in the previous decomposition exercise, agency workers are omitted due to data constraints. To be specific, the available time series (up to 1991) do not consider such workers as part of employment. The reason is simple. Since 1974 up to 1991 there was a limited use of outsourcing jobs, and the share of agency workers over total employment was less than 8 % in the manufacturing industry. Nevertheless, since 1992 this data began to be collected because with the regulation of the temporary work agencies in 1990, the use of outsourcing jobs became standard. The share of agency workers sharply increased from 11 % in 1991 and reached a peak of 29 % in 2007.

To mitigate the impact of the source of measurement errors on our estimates, we calculate a new, or adjusted, labour income share considering a broader definition of employment. We include not only workers directly paid by the firm (either permanent or temporary workers), but also agency workers. To conduct this computation, we assume that all workers have the same average labour compensation and that the share of agency workers remained stable from 1974 to 1991, at 8 %. In this way, our alternative values for the labour income share became 39.82 % in 1974-1991 and 36.04 % in 1992-2015, indicating that the law 50 of 1990 and the resulting boom in outsourcing jobs prevented the labour share to deteriorate at the pace reported with the unadjusted calculation.

To check the robustness of the estimated values for the substitution and scale effects, we re-estimate models (8) and (9) using the new definition of employment. The corresponding results are displayed in Appendix 3 (Tables 9 and 10). As this addition affects mainly the employment series since 1992 onwards, this procedure allows us to test whether the boom of outsourcing jobs in the nineties might or not significantly affect the magnitude of the structural breaks just discussed.

As shown in Table 6, we find that the magnitudes of the new estimates have greater values, no matter the period examined, even though, they do not differ significantly from those obtained when using the original data set.

For the first period (1975-1991), the new estimated values for the labour demand elasticity are around -1.18 instead of -1.05, and the new estimates of the substitution elasticity between capital and labour are around -1.27 instead of -1.08. Given that, the labour income share is 39.82 % on average, we can conclude that the size of the substitution effect is placed around -0.76 which still accounts for less than two thirds of the total effect. These results are not surprising given that the labour income share did not change significantly during these years.

In turn, for the second period of analysis (1992-2015), the new estimated values for the labour demand elasticity, are placed around -1.69 (quite robust across specifications) instead of -1.56. In turn, with a labour share of 36.04 % on average, the substitution and scale effect are now computed around -1.25 and -0.41, respectively. The comparison between the new and previous estimates for both periods of analysis yield two salient findings. First, the increased in 0.51 percentage points in the labour demand elasticity has been strongly confirmed. And second, we have also found evidence that the greater sensitivity of the employment to wage changes is the outcome of a larger substitution effect.

5.3. International Trade Effects on the Level of the Labour Demand

We now turn the attention towards the level effect of international trade on employment. Table 7 displays the implied long-run effects on employment of the trade openness index (ε_{op}), the import penetration ratio (ε_m), and the export ratio (ε_x), all of them arising from the base-run estimates presented in Tables 4 and 5. This information is classified in two blocks, the one on the left-hand side corresponding to the estimation for the first period (1974-1991) and the one of the right-hand side to the second period (1992-2015). Estimates reported in Columns (1) and (3) corresponds to model (11), and Columns (2) and (4) to model (12).

As shown at the top of Table 7, the estimates for models (11) and (12) provide a similar picture with non-significant effects on employment of the trade openness index or the import penetration ratio, no matter the period of analysis. In contrast, in the case of the export ratio, the results differ between models. Estimates for model (11) suggest a significant and negative influence of exports on the labour demand, while the estimates for model (12) point to the irrelevance of exports on employment determination. On this account, if we must favour a particular model, we will choose model (11) –Columns (1) and (3)–. From a theoretical point of view, as described in section 4, model (11) makes up a capital constrained labour demand function. Therefore, the long run coefficients of the three measures of the degree of trade openness –including the export ratio– represent the trade effects on technology efficiency and act as demand shifters. In contrast, model (12) is not a demand for labour function but a marginal productivity condition. Thereby, the respective coefficients cannot be interpreted as the international trade effects on the level of the labour demand. These coefficients capture partially the changes in the technical efficiency of the production process since they are affected by the substitution elasticity between capital and labour (see Appendix 1, extensions for equations (2) and (3), for details).

Therefore, for the first period of analysis, results of model (11) reflect a scant influence exerted by the degree of international trade on employment adjustments. As mentioned before, the labour demand elasticities with respect to the trade openness index and to the import penetration ratio are found to be non-significant. Additionally, although the elasticity with respect to the export ratio is found to be significant, the volume of exports only exerts a low influence on the demand for labour. Specifically, the value of this sensitivity is estimated at -0.11 and implies that 1 % increase in the export ratio causes a reduction in the demand for labour by 0.11 %. These results are not surprising given that in the 1970s and 1980s Colombia was a closed economy, and the industry was mainly based on manufactures that had a low international exposure. The trade openness index was 30.87 %, on average. The export and import shares were 8.84 % and 22.03 %, respectively. And the import penetration was 19.44 %, on average.

For the second period, we find similar results. All long-run elasticities remain virtually unchanged. The long-run effect exerted by the export ratio continues to be negative and significant and the trade volume and the import penetration continue to be non-significant. Thus, these results are reflecting an absence of a structural break on the long-run sensitivity of employment with respect to trade openness.

In the light of these results, this paper allows us to draw two main conclusions. First, our findings do not provide empirical support to the hypothesis of the skilled-bias technological change, according to which a larger exposure to international trade in developing countries tends to raise the demand for high-skilled labour due to the acquisition of foreign technology. On the contrary, our results suggest the demand for labour in the Colombian manufacturing industry has not been affected by growing import penetration, at least at aggregate level. Second, our results indicate that export orientation in the Colombian manufacturing industry has not been accompanied by improvements in technical efficiency. This conclusion can be drawn from the fact that our findings show that the volume of exports has a negative effect on the level of the labour demand and has even remained stable across the two periods of interest –pre and post the liberalization program–.

Finally, at the bottom of Table 7, we provide additional estimates which correspond to those described in the robustness check subsection. As we rely on the results of model (11), the negative significant effects of exports are confirmed as well as their stability across periods. Their magnitudes, however, turn to be lower. They are equivalent to half of the estimated value in the original dataset.

Table 7. Long-run effects

	1974-1991		1992-2015	
	(1)	(2)	(3)	(4)
ε_{op}	-0.06	0.09	-0.08	0.12
	[0.221]	[0.140]	[0.222]	[0.150]
ε_m	-0.06	0.03	-0.07	0.04
	[0.148]	[0.499]	[0.152]	[0.502]
ε_x	-0.11	-0.03	-0.13	-0.03
	[0.001]	[0.428]	[0.001]	[0.417]
Robustness check				
ε_{op}	0.02	0.17	0.02	0.22
	[0.581]	[0.000]	[0.580]	[0.000]
ε_m	0.00	0.09	0.00	0.12
	[0.876]	[0.001]	[0.876]	[0.002]
ε_x	-0.06	0.02	-0.07	0.02
	[0.000]	[0.430]	[0.000]	[0.439]

Notes: P-values in brackets.

Source: own elaboration.

6. CONCLUDING REMARKS

We have studied the potential channels through which the internationalization process of the Colombian economy has affected employment in the manufacturing industry.

On one side, we investigated whether the structural trade changes experimented during the nineties have altered the sensitivity of employment with respect to wages. Our findings indicate that the demand for labour has become more responsive to wage changes. Specifically, the employment elasticity of a wage change increased from -1.05 in 1974-1991 to -1.56 in 1992-2015. This result is shown to be robust to three control variables of trade openness and is consistent with Rodrik's conjecture (1997) of a wage elasticity effect of globalization. Beyond that, we find that the increase of 0.51 percentage points in the labour demand elasticity is the outcome of a larger substitution effect between capital and labour, which almost doubles its size, rising from -0.68 in the first period to -1.19 in the second one. The rise in the elasticity of labour demand with respect to wages even when holding output constant (the substitution effect) implies that Colombian firms have enhanced their internal flexibility to react 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 context.

On the other side, we assessed whether the exposure to international trade has a level effect on the labour demand as predicted by Rodrik (1997). Our results show that this level effect is scant in the manufacturing industry. The long-run labour demand effects exerted by the openness index and the import penetration ratio are found to be non-significant. This implies the absence of empirical support to the hypothesis of the skilled-bias technological change which predicts an increase in the relative skilled demand for labour in developing countries. On the other hand, we found an unexpected negative employment effect of the export share, indicating that export orientation in the Colombian manufacturing industry is not accompanied by improvements in technical efficiency.

How do we assess these results? Our findings suggest that trade liberalization had negative consequences on workers' welfare. Under Rodrik's (1997) logic, the increase in the labour demand elasticity may have triggered more volatile responses of employment and wages to global shocks. In this context, increases in the payroll taxation, as those experimented in the Colombian labour market during the nineties, may have led to job destruction and a higher workers' tax burden. Moreover, the capital-labour substitution processes may have accelerated on account of the larger employment sensitivity. In particular, the manufacturing industry may have become more labour-intensive due to wages in this sector have grown below labour productivity. As a result of the process of trade liberalization and institutional reforms, workers are placed under high pressure in the new open and deregulated environment.

In this context, we call for a joint assessment of the international trade effects on the Colombian labour market. That is, the effect of the internationalization process on the labour demand should be analysed along with the impacts on the labour productivity and wages. It would allow identifying the net effect of international trade on workers' welfare. On this account, there are studies that have found evidence that international trade can contribute to enhance labour productivity and thus increase wages (see for example, Hansen, 2001). In such cases, workers may benefit from the process of trade liberalization even under increases in payroll taxation. Of course, it would be possible if wages are attached to labour productivity. We thus conclude that in the Colombian context, potential positive international trade effects on labour productivity stress the need for wages to become tied to labour productivity. This should be a critical policy target to offset in some extent the negative high pressures to which workers have been placed because of the trade liberalization and labour market deregulation processes.

Our analysis can be refined in a variety of directions. Further research should control for types of employment given that both institutional and trade reforms have different effects by type of worker. Another research avenue is to aim at an individual assessment of how these reforms affected employment in each productive sector. In that case, the starting hypothesis would be that each sector's response is connected to its specific production technology and degree of exposure to international trade.

REFERENCES

Acemoglu, D., Autor, D., Dorn, D., Hanson, G. H., & Price, B. (2016). Import competition and the great US employment sag of the 2000s. *Journal of Labor Economics*, *34*(1) Part 2, S141-S198. https://doi.org/10.1086/682384.

Agudelo, S. A. (2019). *Three essays on the Colombian labour market: a macroeconomic perspective* [Ph.D.Thesis, Universitat Autònoma de Barcelona]. http://hdl.handle.net/10803/669487.

Agudelo, S. A., & Sala, H. (2016). Wage setting in the Colombian manufacturing industry. *Economics of Transition*, *24*(1), 99-134. https://doi.org/10.1111/ecot.12085.

Anderson, T. W., & Hsiao, C. (1982). Formulation and estimation of dynamic models using panel data. *Journal of Econometrics*, *18*(1), 47-82. https://doi.org/10.1016/0304-4076(82)90095-1.

Arango, C. A., & Rojas, A.M. (2004). Demanda laboral y reforma comercial en el sector manufacturero colombiano: 1977-1999. *Ensayos sobre Política Económica*, *21*(44), 96-154. https://doi.org/10.32468/Espe.4403.

Arbache, J., Dickerson, A., & Green, F. (2004). Trade liberalisation and wages in developing countries. *The Economic Journal*, *114*(493), F73-F96. https://doi.org/10.1111/j.0013-0133.2004.00188.

Arellano, M., & Bond, S. (1991). Some test of specification for panel data: Montecarlo evidence and an application to employment equations. *Review of Economics Studies*, *58*(2), 277-297. https://doi.org/10.2307/2297968.

Autor, D. H., F. Levy and R. J. Murnane (2003). The skill content of recent technological change: an empirical exploration. *The Quarterly Journal of Economics*, *118*(4), 1279-1333. https://doi.org/10.1162/003355303322552801.

Autor, D. H., D. Dorn, and G.H. Hanson (2013). The China syndrome: Local labor market effects of import competition in the United States. *The American Economic Review*, *103*(6), 2121-2168. https://doi.org/10.1257/aer.103.6.2121.

Blundell, R., & Bond, S. (1998). Initial conditions and moment restrictions in dynamic panel data models. *Journal of Econometrics*, *87*(1), 115-143. https://doi.org/10.1016/S0304-4076(98)00009-8.

Böckerman, P., & Marilanta, M. (2012). Globalization, creative destruction, and labour share change: Evidence on the determinants and mechanisms from longitudinal plant-level data. *Oxford Economic Papers New Series*, *64*(2), 259-280. https://doi.org/10.1093/oep/gpr03.1.

Bruno, G. S. F. (2005). Approximating the bias of the LSDV estimator for dynamic unbalanced panel data models. *Economics Letters*, *87*(3), 361-366. https://doi.org/10.1016/j.econlet.2005.01.005.

Bruno, G. S. F., Falzoni, A. M., & Helg, R. (2004). Measuring the effect of globalization on labour demand elasticity: An empirical application to OECD countries. Working paper No. 153, KITeS, Centre for Knowledge, Internationalization and Technology Studies, Università Commerciale "Luigi Bocconi", 1-29.

Cárdenas, M., & Bernal, R. (2004). Determinants of labor demand in Colombia. 1976-1996. NBER Chapter 4, 229-272. En J. Heckman, J.J. & Pagés, C. Law and employment: Lessons from Latin American and the Caribbean, University of Chicago Press. https://www.nber.org/books-and-chapters/law-and-employment-lessons-latin-america-and-caribbean/determinants-labor-demand-colombia-1976-1996.

Caselli, M. (2014). Trade, skill-biased technical change and wages in Mexican manufacturing. *Applied Economics*, *46*(3), 336-348. https://doi.org/10.1080/00036846.2013.848033.

Conte, A., & Vivarelli, M. (2011). Imported skill-biased technological change in developing countries. *The Developing Economics*, *49*(1), 36-65. http://dx.doi.org/10.2139/ssrn.993480.

Dickey, D. A., & Fuller, W. A. (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association*, *74*(366), 427-431. https://doi.org/10.1080/01621459.1979.10482531.

Elsby, M. W. L., Hobjin, B., & Sahin, A. (2013). The decline of the US labor share. *Brokings Papers on Economic Activity*, *47*(2), 1-63.

Fajnzylber, P., & Maloney, W. F. (2005). Labor and trade reform in Latin America. *Journal of International Economics*, *66*(2), 423-446. https://doi.org/10.1016/j.jinteco.2004.08.002.

Ferguson, C. E., & Maurice, S. C. (1973). Factor demand elasticity under monopoly and monopsony. *Economica*, *40*(158), 180-186. https://doi.org/10.2307/2551779.

Fisher, R. A. (1932). Statistical Methods for Research Workers. 4th Edition, Edinburgh: Oliver & Boyd.

Gonzaga, G., Menezes, N., & Terra, C. (2006). Trade liberalization and the evolution of skill earnings differentials in Brazil. *Journal of International Economics*, *68*(2), 345-367. https://doi.org/10.1016/j.jinteco.2005.07.008.

Greenaway, D., Morgan, W. & Wright, P. (1999). Exports, export composition and growth. *The Journal of International Trade & Economic Development*, *8*(1), 41-51. https://doi.org/10.1080/09638199900000004.

Hamermesh, D. (1993). Labor demand. Princeton University Press, Princeton.

Hasan, R., Mitra, D., & Ramaswamy, K. V. (2007). Trade reforms, labor regulations, and labor demand elasticities: Empirical evidence from India. *The Review of Economics and Statistics*, *89*(3), 466-481.

Hansen, B. E. (2001). The new econometrics of structural change: Dating breaks in U.S. labour productivity. *Journal of Economic Perspectives*, *15*(4), 117-128. https://www.jstor.org/stable/2696520.

Hijzen, A., & Swaim, P. (2010). Offshoring, labour market institutions and the elasticity of labour demand. *European Economic Review*, *54*(8), 1016-1034. http://dx.doi.org/10.2139/ssrn.1104698.

Im, K. S., Pesaran, M. H., & Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics*, *115*(1), 53-74. https://doi.org/10.1016/S0304-4076(03)00092-7.

International Labour Organization -ILO- (2015). The Labour Share in G20 Economies.

International Monetary Fund -IMF- (2017). World Economic Outlook: Gaining Momentum? Washington, April, 1-238.

Isaza, J. G., y Meza, C.A. (2004). La demanda de trabajo: teoría y evidencia empírica para el caso colombiano. *Equidad y Desarrollo*, *1*(2), 35-64. https://doi.org/10.19052/ed.392.

Jafari, Y., Othman, J., & Nor, A. H. S. M. (2012). Energy consumption, economic growth and environmental pollutants in Indonesia. *Journal of Policy Modelling*, *34*(6), 879-889. https://doi.org/10.1016/j.jpolmod.2012.05.020.

Judson, R. A. & Owen, A. L. (1999). Estimating dynamic panel data models: A guide for macroeconomists. *Economic Letters*, *65*(1), 9-15. https://doi.org/10.1016/S0165-1765(99)00130-5.

Judzik, D. & Sala, H. (2013). Productivity, deunionization and trade: Wage effects and labour share implications. *International Labour Review*, *152*(2), 205-236. https://doi.org/10.1111/j.1564-913X.2013.00178.x.

Karanassou, M., Sala, H. & Snower, D.J. (2007). The macroeconomics of the labor market: Three fundamental views. *Portuguese Economic Journal*, *6*(3), 151-180. http://dx.doi.org/10.2139/ssrn.950947.

Krishna, P., Mitra, D., & Chinoy, S. (2001). Trade liberalization and labour demand elasticities: Evidence from Turkey. *Journal of International Economics*, *55*(2), 391-409. https://doi.org/10.1016/S0022-1996(01)00089-7.

Kwiatkowski, D., Phillips, P., Smith, P., & Shin, Y. (1992). Testing the null hypothesis of stationary against the alternative of unit root: How sure are we that economic time series have a unit root? *Journal of Econometrics*, *54*(1-3), 159-178. https://doi.org/10.1016/0304-4076(92)90104-Y.

Levin, A., & Lin, C. (1993). Unit root tests in panel data: New results. Discussion paper No. 93-56, Department of Economics, University California at San Diego.

Levin, A., Lin, C., & Chu, C. (2002). Unit root tests in panel data: Asymptotic and finite sample properties. *Journal of Econometrics*, *108*(1), 1-24. https://doi.org/10.1016/S0304-4076(01)00098-7.

Lewis, P.E.T., & MacDonald, G. (2002). The elasticity of demand for labour in Australia. *Economic Record*, *78*(240), 18-30. https://doi.org/10.1111/1475-4932.00036.

Maddala, G. S., & Wu, S. (1999). A comparative study of unit root tests with panel data and new simple test. *Oxford Bulletin of Economics and Statistics*, *61*(4), 603-605. https://doi.org/10.1111/1468-0084.0610s1631.

Manning, A. (2004). We can work it out: The impact of technological change on the demand for low-skill workers. *Scottish Journal of Political Economy*, *51*(5), 581-608. https://doi.org/10.1111/j.0036-9292.2004.00322.x.

Malgouyres, C. (2017). The impact of Chinese import competition on the local structure of employment and wages evidence from France. *Journal of Regional Science*, *57*(3), 411-441. https://doi.org/10.1111/jors.12303.

Medina, C., Posso, C. M., Tamayo, J. A., & Monsalve, E. (2012). Dinámica de la demanda laboral en la industria manufacturera colombiana 1993-2009: una estimación Panel VAR. *Borradores de Economía*, *694*. https://www.banrep.gov.co/es/borrador-694.

Nickell, S. (1981). Biases in dynamic models with fixed effects. *Econometrica*, *49*(6), 1417-1426. https://doi.org/10.2307/1911408.

Njikam, O. (2016). Trade liberalization, labor market regulations and labor demand in Cameroon. *International Review of Economics & Finance*, *43*. 525-541. https://doi.org/10.1016/j.iref.2016.02.001.

OECD (2016). Reviews of labour market and social policies: Colombia. January, 1-36.

Petra, D. (2017). Determinants of labour´s income share in the era of financialisation. *Cambridge Journal of Economics*, *41*(1), 283-306. https://doi.org/10.1093/cje/bew023.

Phillips, P. C. B., & Perron, P. (1988). Testing for a unit root in time series regression. *Biometrika*, *75*(2), 335-346. https://doi.org/10.2307/2336182.

Rodrik, D. (1997). Has globalization gone too far? Institute for International Economics, Washington.

Russell, B., & Tease, W. (1991). Employment, output and real wages. *Economic Record*, *67*(196), 34-45.

Saha, B., Sen, K. & Maiti, D. (2013). Trade openness, labour institutions and flexibilisation: Theory and evidence from India. *Labour Economics*, 24, 180-195. https://doi.org/10.1016/j.labeco.2013.08.008.

Sala, H., & Trivín, P. (2012). Structural changes in Spanish labour demand: Does Rodrik.s conjecture hold? Article presented at the IV Congreso Nacional sobre mercado de trabajo y relaciones laborales (Universidad de Valladolid, April 2012).

Selwaness, I., & Zaki, C. (2019). On the interaction between exports and labor market regulation: Evidence from the MENA countries. The Quarterly Review of Economics and Finance, 73, 24-33. https://doi.org/10.1016/j.qref.2018.05.011.

Senses, M. Z. (2010). The effects of offshoring on the elasticity of labor demand. *Journal of International Economics*, *81*(1), 89-98. https://doi.org/10.1016/j.jinteco.2010.02.001.

Seo, H. J., Kim, H. S., & Lee, Y. S. (2015). Globalization and labour demand elasticities: Empirical evidence from nine OECD countries. *The Korean Economic Review*, *31*(2), 413-439.

Slaughter, M. J. (2001). International trade and labor demand elasticities. *Journal of International Economics*, *54* (1), 27-56. https://doi.org/10.1016/S0022-1996(00)00057-X.

Stock, J. H. & Yogo, M. (2005). Testing for weak instruments in linear IV regression. In: Stock, J. H., Andrews, D. W. K., (Eds.), Identification and Inference for Econometric Models: A Festschrift in Honor of Thomas Rothenberg. Cambridge University Press, Cambridge, 80-108. (Ch. 5)

APPENDIX 1

Extension of Equation (4): Theoretical Background

Departing from equation (4)

$$N_t = \left[\alpha A \left(1 - \frac{1}{|\varepsilon|}\right)\right]^{\frac{1}{(1-\alpha)}} \left[\frac{W_t}{P_t}\right]^{\frac{1}{(1-\alpha)}} K_t$$

Where $N_t = n_{it} f$ is aggregate employment, A is an efficiency parameter, $\frac{W_t}{P_t}$ is real wage and $K_t = k_{it} f$ is aggregate capital.

If we assume that the efficiency parameter A is determined by international trade changes as $A = A_0 OPEN^{\delta}$, where $OPEN$ is a variable capturing the degree of exposure to international competition, then δ would stand for the effect of international trade on the technical efficiency of production.

If exposure to international trade competition improves technical efficiency, then $\delta > 0$.

If excessive dependence on imported parts and components attenuate complementarity among domestic firms so that productivity could deteriorate as the globalization process deepens, then $\delta < 0$.

Therefore, taking natural logarithms, introducing a white noise error term $u_t \sim i.i.N(0,\sigma^2)$ to capture supply and demand shocks, and rearranging the terms as follows:

$$n_t = \ln(N_t)\,;\ w_t = \ln\left(\frac{W_t}{P_t}\right);\ k_t = \ln(K_t)\,;\ open_t = \ln(OPEN_t)\,;\ \alpha_0 = \frac{1}{(1-\alpha)} ln\left[\alpha A_0 \left(1 - \frac{1}{|\varepsilon|}\right)\right];$$

$$\alpha_1 = \frac{-1}{(1-\alpha)}\,;\ \alpha_2 = 1\,;\ \alpha_3 = \frac{\delta}{(1-\alpha)}$$

We can obtain the labour demand (9) which incorporates the level effect of international trade α_3.

$$n_t = \alpha_0 + \alpha_1 w_t + \alpha_2 k_t + \alpha_3 open_t + u_t$$

Extension of Equation (8): Theoretical Background

Departing from equation (8)

$$N_t = A^{\sigma-1}\theta^{\sigma}\left[\frac{W_t}{P_t}\right]^{-\sigma} Y_t$$

Where $N_t = n_{it} f$ is aggregate employment, A is an efficiency parameter, $\frac{W_t}{P_t}$ is real wage and $K_t = k_{it} f$ is aggregate capital.

If we assume that the efficiency parameter A is determined by international trade changes as $A = A_0 OPEN^{\delta}$, where OPEN is a variable capturing the degree of exposure to international competition, then δ would stand for the effect of international trade on the technical efficiency of production.

If exposure to international trade competition improves technical efficiency, then $\delta > 0$.

If excessive dependence on imported parts and components attenuate complementarity among domestic firms so that productivity could deteriorate as the globalization process deepens, then $\delta < 0$.

Therefore, taking natural logarithms, introducing a white noise error term $e_t \sim i.i.N(0,\sigma^2)$ to capture supply and demand shocks, and rearranging the terms as follows:

$$n_t = \ln(N_t);\ w_t = \ln\left(\frac{W_t}{P_t}\right);\ y_t = \ln(Y_t);\ open_t = \ln(OPEN_t);\ \beta_0 = \sigma\ln(\theta) + (\sigma-1)\ln(A_0);$$

$$\beta_1 = -\sigma;\ \beta_2 = 1;\ \beta_3 = \delta(\sigma-1)$$

We can obtain the marginal condition (10) which incorporates a partial effect of international trade on technical efficiency through β_3.

$$n_t = \beta_0 + \beta_1 w_t + \beta_2 y_t + \beta_3 open_t + e_t$$

APPENDIX 2

List 1. Sector Classification

Classification based on International Standard Industrial Classification, Rev. 4.

Sectors: Manufacture of food products and beverages (S1); Manufacture of textiles (S2); Manufacture of wearing apparel; dressing and dyeing of fur (S3); Tanning and dressing of leather; manufacture of luggage, handbags, saddlery, harness and footwear (S4); Manufacture of wood and of products of wood and cork, except furniture; manufacture of articles of straw and plaiting materials (S5); Manufacture of paper and paper products (S6); Manufacture of coke, refined petroleum products and nuclear fuel (S7); Manufacture of chemicals and chemical products (S8); Manufacture of rubber and plastics products (S9); Manufacture of other non-metallic mineral products (S10); Manufacture of basic metals (S11); Manufacture of fabricated metal products, except machinery and equipment (S12); Manufacture of machinery and equipment n.e.c.; and manufacture of office, accounting and computing machinery (S13); Manufacture of electrical machinery and apparatus n.e.c.; manufacture of radio, television and communication equipment and apparatus; and Manufacture of medical, precision and optical instruments, watches and clocks (S14); Manufacture of motor vehicles, trailers and semi-trailers, and other transport equipment (S15); Manufacture of furniture; manufacturing n.e.c. (S16).

Note: Manufacture of tobacco products and publishing, printing and reproduction of recorded media are excluded due to data constraints.

Table 8: Bias-corrected LSDVC estimators.

		Model (11)			Model (12)		
Bias order	Estimator	op_{it}	m_{it}	x_{it}	op_{it}	m_{it}	x_{it}
$O(1/T)$	AH	0.92	0.91	0.91	0.92	0.92	0.91
$O(1/NT)$	AH	0.92	0.91	0.91	0.92	0.92	0.92
$O(1/T)$	AB	0.95	0.95	0.94	0.93	0.92	0.92
$O(1/NT)$	AB	0.95	0.95	0.95	0.93	0.93	0.92
$O(1/T)$	BB	0.97	0.96	0.96	0.94	0.93	0.94
$O(1/NT)$	BB	0.97	0.96	0.96	0.94	0.93	0.94

Note: This table only displays persistence coefficients. Column estimator provides the consistent estimator chosen to initialize the bias correction. AH = Anderson and Hsiao (1982); AB = Arellano and Bond (1991); BB = Blundell and Bond (1998).

Source: Own elaboration

APPENDIX 3

Table 9: Estimates of model 11. Robustness check

Dependent variable: η_{it}	op_{it}			m_{it}			x_{it}		
	OLS (1)	FE (2)	FE-TLSLS (3)	OLS (4)	FE (5)	FE-TLSLS (6)	OLS (7)	FE (8)	FE-TLSLS (9)
η_{it-1}	0.98 [0.000]	0.86 [0.000]	0.53 [0.000]	0.98 [0.000]	0.86 [0.000]	0.52 [0.001]	0.98 [0.000]	0.86 [0.000]	0.41 [0.019]
ω_{it}	-0.04 [0.047]	-0.11 [0.000]	-0.57 [0.008]	-0.04 [0.054]	-0.11 [0.000]	-0.57 [0.008]	-0.04 [0.036]	-0.11 [0.000]	-0.68 [0.003]
k_{it}	0.02 [0.099]	0.02 [0.000]	0.31 [0.011]	0.02 [0.093]	0.02 [0.000]	-0.31 [0.013]	0.01 [0.083]	0.02 [0.000]	0.40 [0.004]
$open_{it}$	0.00 [0.956]	-0.01 [0.353]	0.01 [0.602]	0.00 [0.965]	-0.02 [0.147]	-0.01 [0.874]	-0.01 [0.142]	-0.01 [0.216]	-0.04 [0.030]
Δser_t	-3.28 [0.000]	-2.57 [0.000]	-1.91 [0.006]	-3.28 [0.000]	-2.55 [0.000]	-1.92 [0.006]	-3.28 [0.000]	-2.59 [0.000]	-1.83 [0.024]
$\eta_{it-1^*}\, d_{92}$	0.00 [0.488]	0.02 [0.005]	0.02 [0.064]	0.01 [0.466]	0.03 [0.001]	0.02 [0.020]	0.01 [0.452]	0.02 [0.015]	0.02 [0.006]
$w_{it^*}\, d_{92}$	-0.01 [0.415]	-0.02 [0.020]	-0.10 [0.013]	-0.01 [0.414]	-0.02 [0.009]	-0.11 [0.015]	-0.01 [0.479]	-0.02 [0.046]	-0.14 [0.006]
d_{8083}	-0.03 [0.006]	-0.02 [0.011]	0.01 [0.639]	-0.02 [0.006]	-0.02 [0.011]	0.01 [0.594]	-0.03 [0.004]	-0.02 [0.003]	0.02 [0.567]
d_{9700}	-0.04 [0.000]	-0.06 [0.000]	-0.06 [0.000]	-0.04 [0.000]	-0.06 [0.000]	-0.06 [0.000]	-0.04 [0.000]	-0.06 [0.000]	-0.07 [0.002]
d_{0809}	-0.01 [0.561]	0.00 [0.967]	0.01 [0.651]	-0.01 [0.562]	0.00 [0.928]	0.01 [0.631]	-0.01 [0.675]	0.00 [0.322]	0.03 [0.384]
c	0.33 [0.022]	1.97 [0.000]		0.33 [0.021]	1.99 [0.000]		0.37 [0.018]	1.99 [0.000]	
Obs.	654	654	654	654	654	654	654	654	654
$Adj.R^2$	0.99	0.90	0.62	0.99	0.90	0.61	0.99	0.90	0.44
U			18.80 [0.000]			18.00 [0.000]			14.38 [0.002]
W			7.76 (6.61)			7.66 (6.61)			7.21 (6.61)
H			1.75 [0.416]			1.82 [0.404]			2.70 [0.259]

Note: All variables are expressed in logs. P-values in brackets. OLS, Ordinary Least Square. FE, Fixed effects. FE-TSLS, Fixed effects using Two Step Least Squares. U, Under identification test. W, Weak identification test. Stock and Yogo (2005) weak ID test critical value at 10 % maximal IV relative bias in parentheses. H, Hansen (2001) test.

Source: Own elaboration

Table 10: Estimates of model 12. Robustness check

Dependent variable: η_{it}	op_{it}			m_{it}			x_{it}		
	OLS (1)	FE (2)	FE-TLSLS (3)	OLS (4)	FE (5)	FE-TLSLS (6)	OLS (7)	FE (8)	FE-TLSLS (9)
η_{it-1}	0.97 [0.000]	0.85 [0.000]	0.61 [0.000]	0.97 [0.000]	0.85 [0.000]	0.56 [0.000]	0.94 [0.000]	0.85 [0.000]	0.64 [0.000]
ω_{it}	-0.03 [0.313]	-0.13 [0.000]	-0.51 [0.002]	-0.01 [0.703]	-0.12 [0.000]	-0.58 [0.003]	-0.09 [0.000]	-0.12 [0.000]	-0.43 [0.001]
y_{it}	0.03 [0.067]	0.05 [0.022]	0.31 [0.003]	0.03 [0.161]	0.05 [0.020]	0.38 [0.006]	0.05 [0.000]	0.05 [0.031]	0.29 [0.002]
$open_{it}$	0.01 [0.696]	-0.00 [0.967]	0.07 [0.016]	-0.01 [0.002]	-0.01 [0.381]	0.04 [0.047]	-0.01 [0.234]	-0.01 [0.504]	0.07 [0.453]
Δser_t	-2.52 [0.001]	-2.43 [0.000]	-0.78 [0.309]	-2.55 [0.001]	-2.42 [0.000]	-0.52 [0.566]	-2.97 [0.000]	-2.44 [0.000]	-1.07 [0.100]
$\eta_{it-1} * d_{92}$	-0.02 [0.037]	0.03 [0.000]	0.08 [0.000]	-0.02 [0.060]	0.03 [0.000]	0.11 [0.008]	-0.02 [0.013]	0.03 [0.000]	0.09 [0.002]
$w_{it} * d_{92}$	0.01 [0.913]	-0.03 [0.001]	-0.09 [0.009]	-0.01 [0.952]	-0.03 [0.001]	-0.11 [0.014]	-0.02 [0.039]	-0.02 [0.001]	-0.09 [0.004]
d_{8083}	-0.02 [0.008]	-0.02 [0.011]	0.01 [0.712]	-0.02 [0.017]	-0.02 [0.010]	0.01 [0.602]	-0.02 [0.012]	-0.02 [0.004]	0.01 [0.575]
d_{9700}	0.02 [0.429]	-0.05 [0.000]	0.01 [0.815]	0.02 [0.505]	-0.05 [0.000]	0.02 [0.567]	-0.04 [0.001]	-0.05 [0.000]	-0.01 [0.815]
d_{0809}	-0.10 [0.001]	-0.00 [0.714]	-0.01 [0.640]	-0.09 [0.001]	-0.00 [0.714]	-0.01 [0.738]	-0.01 [0.623]	-0.00 [0.872]	-0.01 [0.739]
c	-0.10 [0.703]	1.66 [0.000]		-0.10 [0.703]	1.66 [0.000]		-0.10 [0.703]	1.66 [0.000]	
Obs.	654	654	654	654	654	654	654	654	654
$Adj.R^2$	0.98	0.90	0.73	0.98	0.90	0.63	0.99	0.90	0.76
U			18.07 [0.000]			13.93 [0.003]			19.90 [0.000]
W			6.46 (6.61)			5.90 (6.61)			6.93 (6.61)
H			2.08 [0.353]			2.12 [0.347]			1.95 [0.377]

Notes: All variables are expressed in logs. P-values in brackets. OLS, Ordinary Least Square. FE, Fixed effects. FE-TSLS, Fixed effects using Two Step Least Squares. U: Under identification test. W: Weak identification test. Stock and Yogo (2005) weak ID test critical value at 10 % maximal IV relative bias in parentheses. H, Hansen (2001) test.

Source: Own elaboration

Table 11: Bias-corrected LSDVC estimators. Robustness check

		Model (11)			Model (12)		
Bias order	Estimator	op_{it}	m_{it}	x_{it}	op_{it}	m_{it}	x_{it}
$O(1/T)$	AH	0.87	0.86	0.86	0.85	0.85	0.87
$O(1/NT)$	AH	0.87	0.86	0.86	0.85	0.86	0.88
$O(1/T)$	AB	0.89	0.89	0.89	0.87	0.86	0.87
$O(1/NT)$	AB	0.90	0.89	0.89	0.88	0.87	0.88
$O(1/T)$	BB	0.92	0.92	0.92	0.88	0.89	0.89
$O(1/NT)$	BB	0.92	0.92	0.92	0.89	0.89	0.89

Notes: This table only displays persistence coefficients. Column estimator provides the consistent estimator chosen to initialize the bias correction. AH = Anderson and Hsiao (1982); AB = Arellano and Bond (1991); BB = Blundell and Bond (1998).

Source: Own elaboration