Universitat Autònoma de Barcelona

Three essays on the Colombian labou market: a macroeconomic perspective

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by

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Dedication

To my beloved husband Rubén, whose immense passion and constant dedication to the academy have been my greatest source of motivation and inspiration throughout this doctoral period. I will never be able to thank him enough for his unconditional support and unbroken trust in me. This doctoral dissertation belongs to him as much as me.

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Introduction

This Ph.D. dissertation consists of three essays on the Colombian labour market. It aims to contribute to the academic discussion and bring new insights into the configuration of the labour markets in developing countries. The essays are presented across chapters one, two and three. Each one of these chapters has a research paper structure: introduction and motivations, theoretical background, empirical issues, discussion of the results, and finally, conclusions and policy implications.

This introductory section presents a general motivation for our study together with the main findings and policy implications, which are fully developed along the thesis.

Motivation

This Ph.D. dissertation is motivated from the scientific urge to understand more about the functioning and dynamics of the labour markets in developing countries. On this account, Colombia provides an excellent experience to examine.

On one side, the Colombian economy represents a faithful example of the institutional reforms and the trade liberalization processes that were undertaken by most Latin American governments during the eighties and nineties. In the early nineties, in particular, the Colombian political authorities carried out a series of economic policy measures which led to a major structural change of the economy. Nevertheless, the effects of such measures on the labour market have not yet been widely and deeply assessed. Next we describe briefly the main policy measures that characterized the structural change.

First of all, relatively tight labour market legislation and trade regulations in the seventies and eighties were superseded by more flexible labour market institutions and free international trade in the early nineties. The main expectation was that such structural changes would to contribute to render formal employment more attractive and boost job creation. Second, payroll taxes were increased significantly aiming to expand the coverage of health and pension services. Third, in 1991, a new constitution conferred greater independence upon the board of directors of the Central Bank (Banco de la República) and defined that its principal function was to maintain the value of the national currency, which led to shift the monetary policy to the inflation targeting scheme. This policy, however, started to be implemented in 2000. Finally, in 1999, the minimum wage, which operates as a wage floor, started to be indexed on a yearly basis following the previous year CPI inflation rate. The objective of this measure was to protect the purchasing power of the minimum wage.

Although each one of these measures may have significantly affected the labour market through different channels –labour demand, wage setting and labour supply–, the empirical literature for Colombia is still scarce and, thus, far from reaching a consensus on the economic effects of such set of measures. Even more disturbing is the fact that the common wisdom might have been critically shaping policy decisions during the last decade. In this context, this thesis aims to shed light on the wage and labour demand effects of the institutional and trade changes of the nineties.

On another side, nowadays, Colombia is considered one of the successful economies in Latin America and the Caribbean. Its economic growth rate since 2000 has evolved around 4% on average, the inflation rate has been steadily reduced to stabilize at 3%, and the informal employment rate fell by 5 percentage points. Nonetheless, this positive macroeconomic performance has also been accompanied by a large deterioration of the international competitiveness, since the trade deficit sharply increased from 0.2% of the GDP in 2000 to 12.4% in 2014. Even more, the unemployment rate persists and remains stubbornly high around 10%. Indeed, Colombia has traditionally had one of the highest unemployment rates in Latin America –its rates have even surpassed by more than 4 percentage points the regional average–[1].

Why this high and persistent unemployment rate? Although the common wisdom has assumed that wages are downward rigid and that the job-creating sectors –for example the manufacturing industry– have not hardly developed, no empirical evidence is provided in the literature. This thesis tries to fill this void: First, by checking the extent to which

[1] This difference has diminished significantly after the commodity crisis of 2014.

wage rigidities are relevant in Colombia and, second, by analyzing the demand for labour focusing on the manufacturing industry.

To sum up, this Ph.D. dissertation aims to shed new light on the evolution of two relevant variables of the Colombian labour market –real wages and employment–. On one side, the discussion focuses on understanding the wage setting system in Colombia and assessing whether the changing institutional framework and the growing exposure to internationalization have altered the wage dynamics (Chapter 1). Furthermore, this dissertation posits special attention on testing the extent to which wage rigidities are relevant in Colombia. We identify its causes and discuss the policy implications (Chapter 2). On the other side, this research provides an extensive analysis of the labour demand effects of international trade (Chapter 3). Our period of analysis is from 1974 to 2015. Thus, the topics of international trade and labour market institutions become transversal issues across Chapters 1 and 3. Within this wide period since the seventies, the Colombia wage setting system and the demand for labour are assessed in two different periods –the slow transition between import substitution and trade liberalization in 1974-1991–; and –the trade and institutional reforms years of 1992-2015–. It is also important to highlight that the analysis performed in these two chapters focus on studying the manufacturing industry. The reason is twofold. It has been the most exposed sector to international trade. And it is the sector that has available the largest and most reliable time series in Colombia.

From a theoretical perspective, the wage setting system and the labour demand are important driving forces, not only of job creation and unemployment dynamics, but also of international competitiveness.

Standard wage setting models –collective bargaining and efficiency wage models– show that unemployment may arise from frictions among prevailing hiring decisions, labour demand and wage setting mechanisms. In this context, the wage level is important. It captures the main incentive to hire workers (labour demand) and to participate in the market providing a work effort (labour supply). In turn, the labour demand determines the level of employment. Therefore, the impact of any exogenous change in the factor prices or policy (e.g. payroll tax changes, trade liberalization processes, wage subsidies, etc.) on employment will depend on the structure of the labour demand. This research places special attention to the wage elasticity of labour demand which allows inferring the effects of changes in the wage rate on the amount of workers that employers are willing to hire.

All in all, this research pretends to contribute to the scientific knowledge about how labour markets work in the developing world. To achieve this purpose, we analyze the Colombian case and compare our results with the theoretical and empirical literature. In doing so, this research allows us to question the validity of some standard theoretical predictions about labour outcomes, which often apply for advanced economies. Next we develop in further detail the main results and the policy implications related with this research.

Chapter 1: "Wage setting in the Colombian manufacturing industry"

In this chapter, we analyze the wage determination in the Colombian manufacturing industry. In particular we identify the main drivers of net real wages and assess whether the institutional and trade reforms carried out in Colombia during the nineties have been translated into an efficient mechanism of wage determination.

Overall, our results show that wage setting system in Colombia is not efficient, at least at the industry level since the workers' compensation does not faithfully reflect labour's efficiency. That is, the net real wages are not fundamentally driven by labour productivity in contrast to the standard theoretical prediction. The reason why this theoretical prediction is far from holding in the Colombian manufacturing sector is twofold. On one side, adjustments of nominal wages are highly tied to the cost of living; in particular, nominal wage adjustments are indexed to the previous year CPI inflation rate, so that wage setting has a backward looking orientation. On the other side, payroll taxes and other non-wage costs make up a significant part of the total worker's compensation –they have been around 40%, on average, since 1980–. Both features weaken the sensitivity of net real wages to changes in labour productivity.

So, what have we learned from the institutional and the trade reforms of the nineties?

The main lesson we draw from this study is that the institutional measures undertaken by the government –such as the enhanced indexation of nominal wages to CPI inflation and the increasing in payroll taxes– have not been translated into an efficient wage determination. These policy reforms have not lead to enhance the scarce connection between workers' compensation and labour productivity. On the contrary, the relationship between these two variables has become weaker during the nineties and noughties. This can be considered one of the collateral pitfalls of the institutional and trade reform process since it is tempting to associate this finding to the increasing trade deficit experienced by Colombia in the last decades.

The second lesson is related to the identification of a fall in the ability of firms to shift the tax burden on workers through lower nominal net wages or higher prices. This fall is the joint outcome of the enhanced indexation of nominal wages to CPI inflation and the asymmetric downward pressures on manufacturing and consumption prices resulting from the liberalization process. The consequence of this lower tax shifting is again a loss in firms' cost competitiveness. In this context, firms probably contracted the demand for labour paid by firms and started to use outsourcing schemes as a mechanism to offset the lost competitiveness arising from the increases in payroll taxes experienced in the nineties.

The third lesson is that the wage setting has become slightly more persistent. That is, the speed of adjustment of real wages has been reduced. This implies that firms have a lower ability to face external shocks through wage and price adjustments. This result is also connected to the full indexation of nominal wages and the growing liberalization of the labour market.

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Chapter 2: "Wage rigidities in Colombia: Measurement, causes, and policy implications"

In this chapter, we explore the potential existence of a downward real wage rigidity –DRWR henceforth– in Colombia. We also evaluate its causes and discuss the policy implications of such rigidity.

Overall, our findings show that real wages are downward rigid in Colombia. Our measure of DRWR is computed as a fraction of the real wage cuts prevented in the process of wage setting in Colombia. Specifically, we estimate a deficit of real wage cuts of 12.09%, indicating that about 12 out of 100 potential expected real wage cuts do not result in actually observed wage cuts. The average of 12% is relatively high, when compared to values of 3.7% found for European economies; but relatively low with respect to the average of 15% for Latin America and the Caribbean countries.

Over time, this DRWR has followed a declining path. In particular, we estimate a value of DRWR above 20% in 2002, a sharp decline in subsequent years, and stabilization around 10% between 2011 and 2014. At the sectorial level, the differences are large since DRWR ranges between 0% and 40.43%. Although these large differences may be associated to disparities in trade union density rates across sectors, we show that wage rigidity in Colombia is not fundamentally connected to the wage bargaining system. This is in contrast to empirical evidence on advanced economies. In Colombia, the main driving forces of DRWR are: economic growth, inflation, real minimum wage growth and labour informality.

An outstanding characteristic of our results is that the falling path of DRWR has taken place simultaneously to the falling trend in price inflation. However, reductions in DRWR are usually associated with situations in which inflation grows. At least, this is the standard theoretical prediction and the common result for advanced economies. We attribute the positive influence of inflation on DRWR to the particular Colombian wage setting system, in which nominal wages are indexed to past CPI inflation. On this account, the chapter provides an illustrative and extensive discussion on the three channels by which inflationary pressures may have enhanced real wage rigidities: (i) the minimum wage anchor, (ii) backward looking wage setting and wage price feedback, and (iii) inflation persistence.

Policywise, our results imply that the feasible mechanism to fight rigidities is to envisage a far-reaching reform of the current wage setting process. Although boosting economic growth appears as an effective mechanism to reduce DRWR, the cyclical nature of economic growth is beyond any particular labour market policy. Other potential instruments may be increasing labour informality and reducing inflationary pressures. The problem with these options is that the first one is undesired and the second one has a narrow margin to operate due to the inflation targeting approach adopted by the Central Bank of Colombia, which in principle provides a guarantee that price inflation will remain under control.

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Chapter 3: "Labour demand effects of internationalization in the Colombian manufacturing industry"

In this chapter, we explore the potential channels through which the internationalization process of the Colombian economy might have affected employment in the manufacturing industry. Our objective is twofold. First, we examine whether the structural change in international exposure experimented during the nineties has altered the elasticity of labour demand with respect to wage changes. The novelty of our analysis lies in the empirical decomposition of this elasticity in a substitution and a scale effect. Second, we evaluate the possibility that the growing exposure to international trade has also caused a "level effect" on the labour demand.

Rodrik (1997) conjectured that a larger exposure to international trade would tend to increase the wage elasticity of the labour demand. The reason is twofold. 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 –reinforcing the substitution effect between capital and labour–. On the other side, the resulting lower entry barriers of the trade policy liberalization lead domestic firms to face heightened foreign competition –enhancing the scale effect—

In our research, we indeed uncover a substantial increase in the total employment sensitivity with respect to wages in response to the internationalization of the Colombian economy. This rise, however, is only the outcome of a larger substitution effect, which nearly doubles its size. This implies that the wide liberalization program has facilitated the substitution between capital and labour in the manufacturing industry since the Colombian economy has had easier access to new markets and new avenues of productive specialization.

Regarding the international trade effect on the level of the labour demand ("level effect"), it has often been associated with the hypothesis of skilled-biased 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. Although this hypothesis has been strongly supported by several 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. This is the reason why we expect to contribute to the literature by estimating the level effect on the total demand for labour.

Our findings indicate that this level effect is scant in the Colombian manufacturing industry, with no significant impact of imports and a negative impact of exports. These results are interpreted as evidence that the broad liberalization process in Colombia has not led to enhance the technical efficiency of the production process in the manufacturing industry. On the contrary, the growing export orientation of the industry has been accompanied by deterioration in the efficient use of the productive factors.

Overall our results suggest that trade liberalization is likely to have had negative consequences on workers. The increase in the labour demand elasticity may have resulted in larger employment volatility. As a consequence, it seems plausible that the increases in payroll taxation experimented during the nineties may have led to job destruction and a higher tax burden in the manufacturing industry. We thus conclude that workers were placed under pressure as a result of institutional changes and trade liberalization.

Conclusions and policy implications

All in all, the objective of this dissertation is to contribute to the scientific debate about the Colombian labour market dynamics from a macroeconomic perspective. This thesis provides new empirical evidence and insightful discussions about the labour market effects of trade and institutional reforms, and wage rigidities. Consequently, relevant policy implications arise.

If a single policy was to be outlined, it would be probably consist in a policy agenda in which a "true" wage bargaining system is brought in. Wage indexation needs to be reduced. Wages should become attached to labour productivity and start being fixed over expected prices. The arguments to sustain such strong statement are discussed below.

- *Wage setting system is a key determinant of international competitiveness.*

The wage setting system should endorse to economic growth and job creation in the long run. Nevertheless, the fact that wage growth rate is aligned with balanced economic growth depend largely on the underlying objective in the wage formation process. For a small and relatively opened economy as Colombia, achieving and maintaining cost-competitiveness should be a key target, as it lays the foundations for the success of the economy's tradable sector and, thus, production and employment growth. If cost-competitiveness becomes the starting point for the wage setting process, it must rely on labour's efficiency progress mainly. The former should be implemented, of course, without neglecting the role played by non wage costs, prices of other production factors, and the labour costs and prices of international competitors.

On the contrary, we show that the wage setting system in Colombia is not efficient. The connection between wages and labour productivity is scarce and has even been worsened during the nineties and noughties. This implies that in a context of growing exposure to international trade, the current wage-setting system will generate distortions in the process of achieving competitiveness, especially in the manufacturing industry. Thus, if the Colombian economy continues to globalize at the same pace than in the last decades, the link between wages and labour productivity needs to be further reinforced. This should be a critical policy target to solve the competitive problem of the Colombian industry, and help in this way rebalancing the trade deficit attained with the liberalization program.

- *Wage adjusments determine how the economy responds to global shocks.*

The wage and price setting mechanisms are crucial factors in determining the shape and extent to which an economy can face adverse shocks. Overall, it is expected that when wages and prices are rigid, demand and supply shocks lead both to greater cyclical fluctuations in GDP and unemployment than when wages and prices are completely flexible. Tobin (1972), for instance, was the first one to state the potential adverse macroeconomic effects of downward real wage rigidity –low inflation generally leads to higher unemployment rates, especially in periods of economic downturn–. On this account, there is ample evidence of downward real wage rigidity in developed economies (see for example, Dickenes *et al.*, 2007; Mesina *et al.*, 2014; and Holden and Wuslfsberg, 2009 and 2014). Even further, there is a growing number of international studies showing that firms typically respond to a weak cyclical situation by reducing the number of persons employed or hours worked, whereas wage cuts are less frequently applied (see for example, Fabiani *et al.*, 2010; Fabiani *et al.*, 2015; and Izquierdo *et al.*, 2017).

In this thesis we provide the first evidence that in Colombia real wages are also downward rigid. Although inflation is usually claimed to prevent negative effects of wage rigidities on unemployment, in the context of Colombia, inflation does not appear as a feasible mechanism to reduce wage rigidities. On the contrary, higher inflation rates deepen the downward rigidity of real wages since the current Colombian wage setting process has a backward looking orientation. Even further, although within this context it would be tempting to suggest a significant reduction of inflation, the Colombian monetary authority has already been carrying out a price stabilization program to keep inflation under control and to try to converge to the low values of the developed economies.

On the other hand, as well-known, under price stabilization programs, backward looking wage indexation has several undesirable effects. It makes inflation more persistent, increases the output loss caused by the program, causes a greater volatility of real wages, and may lead to a worsening of the external current account. Instead, making wage indexation forward looking –so that it starts being fixed over expected prices– would reduce the persistence of inflation and the costs of reducing it, while also lead to real wage stability. In other words, forward looking indexation would improve the efficiency and feasibility of the targeting inflation approach. In addition, it should be politically feasible to introduce due to the fact that it should lead to more stable real wages.

In line with the former discussion, our main policy advice is that the Colombian wage setting system should be reformed so that wages start being fixed over expected prices. Having a forward-looking wage setting system might become an effective tool not only to reach low inflation rates but also to attenuate tensions from the current wage-price spiral, thereby breaking the direct connection between inflation and wage rigidities.

- *Wage setting system is a key factor of the workers' welfare.*

According to Rodrik (1997), increasing international trade makes it more costly for workers to achieve a high level of labour standards and benefits. The cost of improving working conditions –generally non-wage costs– in relatively open economies cannot be shared with the employers with the same ease as in closed economies since the labour demand becomes more sensitive to wage changes as a result of the internationalization processes. The more elastic labour demand is, the greater the share of the non-wage cost increase that workers must bear. This implies that increases in non-wage costs cause a reduction in net wages. This reduction, however, will be higher for industries with a larger exposure to international trade. Likewise, Rodrik (1997) pointed out that there will be a larger reduction of employment as a consequence of increasing such costs.

In our dissertation, we provide empirical evidence that manufacturing employment in Colombia has become more sensitive to wage changes. Openness to international trade is an important factor driving this structural change. Therefore, in line with Rodrik's tax incidence analysis, our findings imply that the increases in payroll taxation experimented during the nineties in Colombia may have led to job destruction and a higher tax burden on workers in the manufacturing industry. In such case, potential positive international trade effects on labour productivity stress the need for wages become attached to labour productivity. This should be a critical policy target to alleviate to some extent the negative high pressures to which workers have been place in the new open and deregulated environment.

Chapter 1

Wage setting in the Colombian Manufacturing industry

Abstract[1]

We show that wage setting in the Colombian manufacturing industry is not fundamentally driven by labour productivity in contrast to the standard theoretical prediction. On the contrary, internal institutional arrangements —payroll taxation, the minimum wage or the price wedge between manufacturing and consumption prices— together with a higher exposure to international trade —connected to the increasing globalization of the Colombian economy— appear as the crucial drivers. These findings lead us to question the political strategy followed to attain cost competitiveness in a context of growing exposure to international trade. Implementation of a true wage bargaining system is suggested as a critical policy target to prevent the disruptive economic consequences of the current wage setting mechanism and help rebalance the trade deficit.

JEL Classification: J30, F16, J30.

Keywords: Wage setting, Labour productivity, Trade openness, Payroll taxes, Minimum wages, Price wedge

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1.1 Introduction

Two structural and worldwide phenomena are critically shaping policy decisions. The first one is globalization. The second one, going in parallel, is a deindustrialization process taking place in many countries. Globalization implies a growing exposure to international trade that needs to be strategically handled bearing in mind the specific strengths and weaknesses of the economy in change. The manufacturing industry is the most exposed economic activity, hence the need of a careful design of the policies that will allow industrial activities to cope.

On this account, Colombia provides an excellent experience to examine. It is one of the considered successful economies in Latin America which, in recent decades, embarked in an extensive liberalization program. The issue we assess here is whether this extensive program has been translated into an efficient mechanism of wage determination, at least at the industrial level. Our conclusion is that it has not. And the reason is twofold. First, because of the low connection between wages and productivity; and, second, because of the failure of the policy reforms to set up a *true* wage bargaining system and enhance the connection between wages and productivity.

We study wage setting in the Colombian manufacturing industry in a period in which relatively tight labour market and trade regulations in the seventies and eighties have been superseded by more flexible labour market institutions and trade liberalization. We investigate, in particular, which are the crucial drivers of net real wages with explicit focus on the role played by labour productivity, the changing institutional framework, and the growing exposure to international trade. This analysis is conducted through the estimation of a standard wage setting model using a panel database which covers a long time period (1974-2009) and 19 manufacturing sectors. This estimation involves the consideration of several type of interactions to disclose whether the institutional and trade reform process has affected the way wages respond to its driving forces.

In this context, a first salient result is the permanent low sensitivity of net real wages to changes in labour productivity. In other words, labour productivity is not the fundamental driver of the net real wage, is in sharp contrast with the standard theoretical prediction of a one-to-one long-run relationship between wages and labour productivity (see Judzik and Sala, 2013, for the simplest analytical case). The reason why this theoretical relationship is far from holding in the Colombian manufacturing industry is twofold. First, the adjustments of nominal wages are highly tied to the cost of living. Second, payroll taxes and other non-wage costs make up a significant part of total compensation. Both features weaken the sensitivity of net real wages to changes in labour productivity. A sensitivity that has even decreased after the institutional and trade reform process.

In contrast to the scarce influence of productivity, we find wages to be mainly driven by internal institutional settings (such as payroll taxation and minimum wages), other determinants related to the indexation of wages with respect to prices (such as the price wedge between manufacturing and consumer prices), and openness to trade. As we explain later, the real minimum wage and the price wedge can be considered two sides of the same coin capturing the close attachment of wages to the evolution of prices (with real minimum wages growing according to prices).

This leads us to specify the empirical model under two alternative settings, one with the real minimum wage and another one with the price wedge, revealing a high sensitivity of net real wages to changes in both the real minimum wage and the price wedge (depending on the specification). This result provides empirical evidence that the adjustments of nominal wages are highly tied to the cost of living, which is critically driving the growth of net real wages in the manufacturing industry. A specific feature of this result is the absence of a significant change in the wage reaction to these driving forces, with similar elasticities in 1976-1991 and 1992-2009.

The third main finding is the significant negative impact of payroll taxation on net wages. Two are the reasons by which payroll taxes may cause a fall in net real wages. The first one is to avoid an increase in total labour costs: firms try to lower net nominal wages to compensate the increase in payroll taxation and, therefore, there is a change in the composition of labour costs in response to the higher tax rates. The second mechanism consists in shifting the increase in payroll taxation to selling prices to try to keep the profitability rates unchanged. This, however, was not feasible in Colombia in a situation of growing exposure to international trade.

The impact of payroll taxes on wages has been lower after 1992 in spite that most of the increases in these taxes (paid by firms) took place in this period. That is, although the payroll tax shifting continued to be significant, firms started bearing a larger share of the payroll tax cost. This result provides evidence that some of the changes in the institutional environment have harmed the cost competitiveness of the industrial sector.

In this context, the explanation of the lower sensitivity of real wages to payroll taxation is twofold. First, the enhanced indexation of nominal wages to Consumer Price Index (CPI) inflation caused net real wages to be more responsive to the price wedge (with severe downward pressures on manufacturing prices resulting from the liberalization process) and less to payroll taxes. Second, it was gross wages, rather than net wages, who absorbed the bulk of the impact of the payroll tax rise (note that the gap between the growth rates of gross and net wages tripled from 0.21 to 0.65 percentage points, as shown in Table 1.2).

The last finding is the significant positive impact of globalization on wages in 1976-1991 and 1992-2009, and the slightly greater influence they have exerted in the second period. This positive and greater influence is interpreted along the lines of Arbache *et al.* (2004), who point to the skill-biased nature of in-flowing technology (through higher foreign direct investment and growing imports) to explain the greater demand of skilled labour, and the resulting pressure on relative wages.

In the light of these results, we conduct a complementary exercise in which we examine the marginal effects of real minimum wages and the price wedge on net real wages, conditional on the different values taken by payroll taxation and the degree of trade openness. In this way, we check whether the reform process (that is, the increases in payroll taxation and the exposure to international trade) have affected the way wages react to these key determinants. We find no difference in the extent of the wage impact of real minimum wages across periods, which remained stable in spite of the liberalization process; this does not preclude, however, that more open sectors tend to be less responsive to changes in the minimum wage. In turn, the sensitivity of real wages to the price wedge becomes larger along with the degree of trade openness in a scenario of high payroll taxation such as the one in 1992-2009. This response is clearly different than the one in 1976-1991, and calls for future research on the interaction between wages, payroll taxes and trade exposure.

The rest of the paper is structured as follows. Section 1.2 reviews some key features of the Colombian economy regarding the institutional setting and the trade liberalization process. Section 1.3 discusses the theoretical background on wage setting models, and their empirical implementation in the paper. Section 1.4 presents the data and explains the econometric methodology. Sections 1.5 and 1.6 deal with the results. Section 1.7 concludes.

1.2 Labour market institutions and trade liberalization

1.2.1 Labour market reform

The Colombian labour market is segmented. There is a massive informal sector accounting for around 50% of employment, and a formal sector accounting for the remaining 50%.[2] Within this context, the manufacturing industry is an economic activity with a relatively low incidence of informality close to 20%.[3]

Since 1990, a structural reform process has been developed in view, on one side, to enhance labour market flexibility and boost (formal) employment and expand, on the other side, the coverage of health and pension services.

To achieve the first target, Law 50 was passed in 1990 to lower firing, training and recruitment costs, and promote non-regular forms of employment. This was followed by Law 789 in 2002, which lowered the regulated costs of non-standard employment (for example,

[2] Value for thirteen metropolitan areas in year 2013 computed from data provided by the Gran Encuesta Integrada de Hogares (GEIH).

[3] It was 18.7% in 2007, in contrast to close to 75.0% in trade and services (values based on data for thirteen metropolitan areas supplied by the GEIH).

on weekend, night, and holiday working hours). The expectation was that such measures would contribute to render formal employment more attractive and boost job creation.

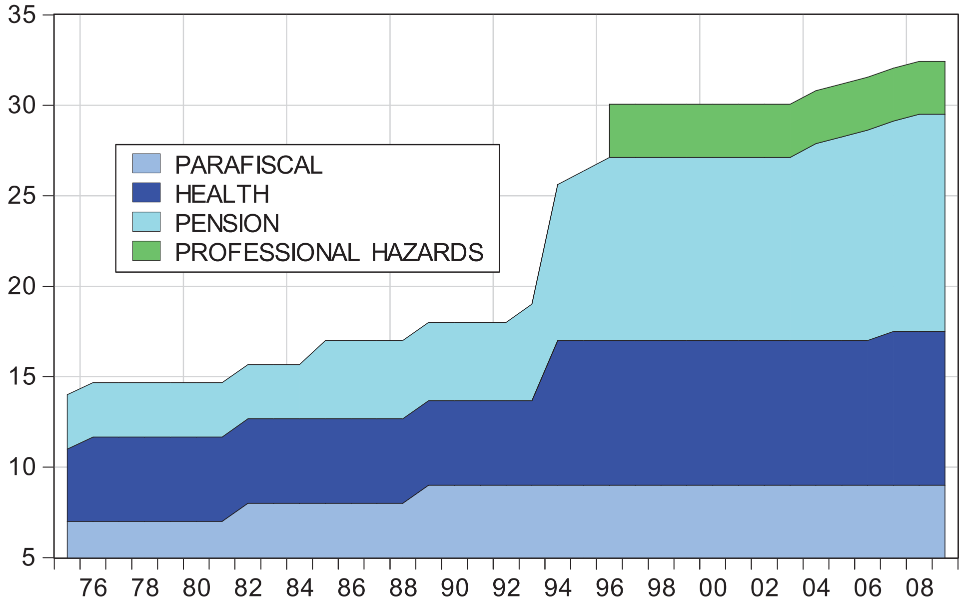
To expand the coverage of health and pension services, Laws 100 in 1993 and 797 in 2003 were passed to increase the Social Security revenues needed to expand the fledgling welfare state. The first of these laws increased the payroll tax rates covering the health and pension schemes, while the second one only focused on pensions. In this way, the total payroll tax rate paid by firms moved from the existing 17% before these laws, to more than 30% since the mid nineties, and beyond afterwards. Figure 1.1 shows this evolution by distinguishing the four components in which payroll taxes are classified in Colombia: parafiscal, health, pension and professional hazards.[4]

There is ample literature on institutions referring to the supposedly ‘labour-unfriendly’ impact of payroll tax rate increases. This common view may be one of the reasons why, without hardly any assessment on the estimated consequences of the previous measures, the government decided to lower first (by Law 1429 of Formalization and employment generation in 2010), and then eliminate, the parafiscal and health contributions paid by firms (by Law 1607 in 2012).

One exception in the virtual lack of assessment of the consequences of the rise in payroll taxes in Colombia is the work by Kugler and Kugler (2009). Their results show that a 10% increase in payroll taxes reduces wages of production workers by 1.46%, wages of nonproduction workers by 2.75%, production employment by 5.14%, and nonproduction

[4]The parafiscal contributions are payroll taxes which are only paid by firms. They have been used to finance the Family Compensation System, the National Service of Learning (SENA by its acronym in Spanish) and the Colombian Family Welfare Institute (ICBF by its acronym in Spanish). The health and pension payroll taxes correspond to contributions that firms and employees must shell out in a traditional social security system. The professional hazards are payroll taxes paid by firms that have been used to fund the General System of Occupational Risks, whose aim is to protect and assist workers from the effects of diseases and accidents that may befall them during or as a result of their work. In this study we take into account all payroll taxes paid by firms.

Figure 1-1: Payroll taxes paid by firms. 1975-2009.



Source: Own calculation based on data from LEGIS' official annual reports.

Cartilla laboral, years 1989-2009; and Cartilla de la Seguridad Social, years 1990-2009.

employment by 4.38%.[5]

Another critical labour market institution is the minimum wage. Although the government had the power to set a minimum wage since 1945 (by Law 6), it was in 1949 when it was effectively implemented for the first time. The minimum wage is indexed on a yearly basis following the previous year CPI inflation rate. Its growth cannot be inferior to this rate,[6] which is the same irrespective of the economic sector (see Hofstetter, 2005, for details).

In practice, it turns out that the real minimum wage operates as a sort of reservation wage which not only is a floor wage (in levels), but also a reference for wage increases in formal activities. Hence their strong correlation, as shown in Figure 1.2.

Finally, although the Colombian labour legislation recognizes unions as a part of the labour relations system, its role in wage-setting matters is nowadays minimal and essentially restricted to collective bargaining at the firm-level. Union density in Colombia is around 4%, while the coverage of collective agreements is less than 2% (data from ENS).

1.2.2 Trade reforms

In parallel to the labour market reform process, Colombia also embarked in a process of external liberalization in the 1990s. 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. In this way, free imports (i.e., custom non-controlled imports) rose from 14.8% in 1985 to 96.7% in 1990, while the average customs duty rate was reduced from 38.9% in 1990 to 12.0% in 1995.

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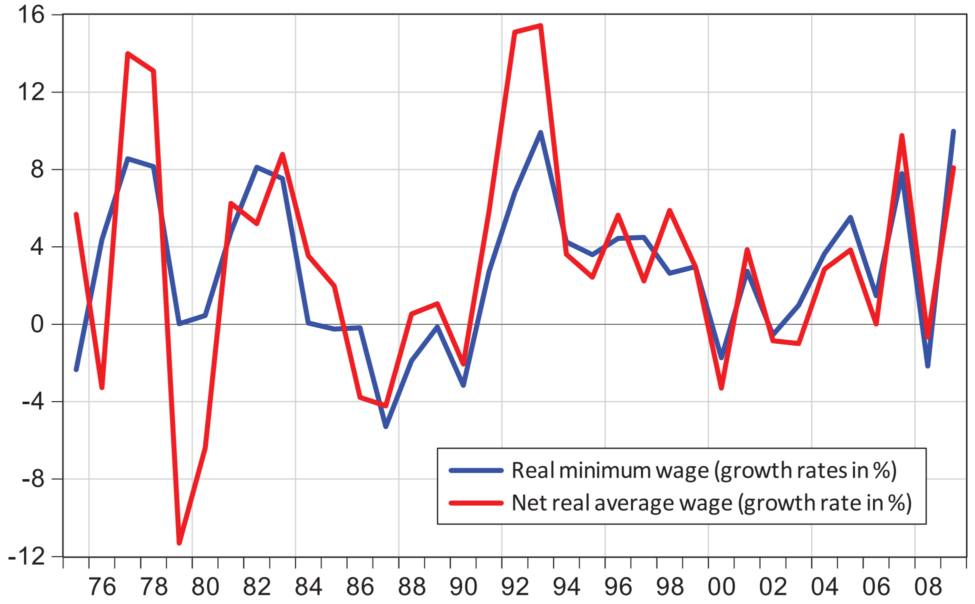
[5] Production workers (or employment) include workers in tasks strictly related to productive activities (usually called blue-collar), while nonproduction workers (or employment) include workers in administrative tasks (usually called white-collar). This is the distinction made by Kugler and Kugler (2009) because of the source data they work with.

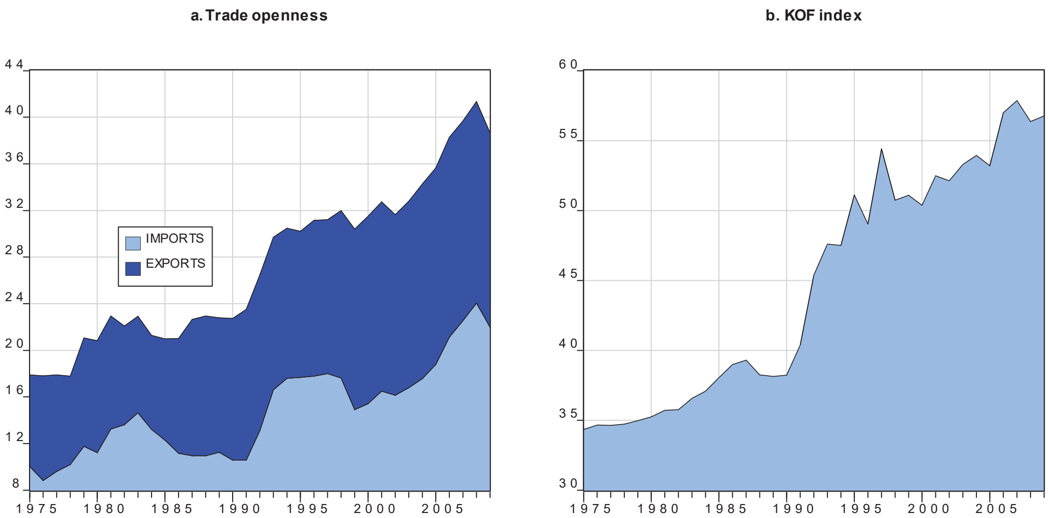
[6] This was started by judgment C-815 of the Constitution Court in 1999 as a consequence of the loss of purchasing power of the minimum wage in the eighties and early nineties.

Figure 1-2: Real wages of the manufacturing industry 1975-2009.



Source: EAM for the net real wage and legislation published annually for the minimum wage.

Figure 1-3: Globalization in Colombia. 1974-2009.



Source: DANE for Trade openness; Dreher *et al.* (2008) for the KOF index.

In addition, between 1992 and 2004, Colombia enjoyed a new system of preferential tariffs to export to the US.[7] This system was superseded in 2004 by Free Trade Agreements (FTAs) between Colombia and a number of relevant trade partners such as the US, the European Union, Canada, Mexico, Korea, Chile, Salvador, Guatemala and Honduras.

As a consequence, trade openness (i.e., the ratio of exports plus imports over GDP) increased from below 25% in 1990 to more than 40% today (Figure 1.3a). Therefore, although Colombia is still considered a relatively closed economy, in the last two decades it experienced a significant change in its overall degree of globalization as measured, for example, by the KOF index (Figure 1.3b). From a value of 34.3% in 1975, this index only increased by 4 percentage points up to 1990 (34.2%). The steep slope thereafter reflects the liberalization process of the Colombian economy, leading the KOF index to reach 57% in 2009.

In this context, the manufacturing industry has been the most affected sector in Colombia, with a steady increase in the imports share that attained 90% of total Colombian imports in 2009 (see Figure 1.6a in the Appendix 1), and a share of exports above 60% of total exports, in contrast to a share of less than 15% in terms of GDP. Still more important, trade openness in this sector (i.e., the ratio of industrial exports and imports over total industrial output) doubled from an average of 48.2% in 1974-1991, to one of 96.8% in 1992-2009.[8]

[7] The APTA was signed in December 1991. Through this agreement around 5,600 products were granted free access to the US market in exchange of a renewed focus on economic activities and jobs aiming at replacing the cocaine industry. This agreement was subsequently prolonged and extended, 2002 onwards, through the APTDEA.

[8] Detailed information by sectors is provided in the Appendix. All industries, but one (publishing and printing recorded media, just accounting for around 3% of total industrial GDP) were subject to this process, and there are no significant composition effects within the manufacturing industry in Colombia: the GDP share of the largest sector (on food products and beverages) remained stable at 28-29%, while the five largest sectors (S1, S3, S12, S10 and S9 in Table 1.6 in the Appendix 1) accounted for 62.2% the industrial GDP in the first period, and 65.9% in the second one.

A related critical issue is that the leading role played by the manufacturing sector in the opening process of Colombia has been accompanied by a large deterioration of its international competitiveness. This is illustrated in Figure 1.6b in the Appendix 1, which shows a structural trade deficit that became much larger on averages since the nineties.

1.3 Wage setting

1.3.1 Theoretical background

Standard wage setting models have been developed both in perfect and imperfect competition contexts.

A relevant example of a perfect competition wage setting model can be found in Kugler and Kugler (2009), where the market-clearing wage and employment levels are set to equate the demand and supply of labour. Firms choose employment by equating the gross marginal labour costs –net wage plus payroll taxes– with the marginal revenue of producing an extra unit of output. In turn, workers set their labour supply as a function of net wages and their prospective social security benefits received in exchange of the payroll tax paid by firms. This gives form to the tax/benefit linkage developed in Summers (1989). The outcome of this model is that wages are set as a function of the labour demand and labour supply elasticities with respect to wages, and also with respect to the payroll tax (paid by firms) and the valuation made by workers over the benefit received.

In a context of imperfect competition, the traditional classification distinguishes efficiency wage from collective bargaining models such as the insider-outsider or union models –Lindbeck and Snower (2001), Booth (2014). Within this vast strand of literature, the work by Podrecca (2011) provides an encompassing model in which a large number of identical unions and firms set real wages in a Nash bargaining framework. Empirically, her model states that the real wage is related to the following six fundamental variables: labour productivity (Y/N), the unemployment rate (u), the replacement ratio (BRR), union bargaining strength (η), payroll taxes/non-wage costs (τ^p), and income taxes $\left(\tau^i\right)$.

By solving the theoretical model in a Cobb-Douglas framework and log-linearizing the resulting expression, Prodecca (2011) obtains her core empirical equation:

$$\log(W_t) = a_0 + a_1 \log\left(Y_{t/}/N_t\right) - a_2 \log(u_t) + a_3 \log(BRR_t) + a_4 \log(\eta_t) - a_5 \log(\tau_t^p) + a_6 \log\left(\tau_t^i\right) + \varepsilon_t \tag{1.1}$$

To study the Colombian case, this benchmark expression needs to be adapted at least for a twofold reason. First, because some institutional features that are generally considered as crucial in developed economies, are less relevant in Colombia. Consequently, we lack times series information to be included in the analysis. Second, because the benchmark model we have considered does not account for the role of the external sector which, as we have discussed, is likely to have affected wage determination in the manufacturing industry. On top of these reasons, some data is simply not available.

Therefore, we next discuss the adjustments and extensions we introduce to the benchmark equation (1.1).

1.3.2 Empirical implementation

The first adjustment relates to the replacement ratio, which is essentially nonexistent in Colombia as a source of income for unemployed. There is a related in-kind benefit which was to a large extent improperly used (not as a true unemployment benefit) before the 2013 reform. This affects the term $a_3 \log(BRR_t)$, which is dropped from the analysis.[9]

The second adjustment refers to union power in whatever form (measures of trade union density, number of strikes, or days lost due to labour conflicts). Although there is data since the 1990s, any measure related to union power was not relevant as explanatory variable in subsample regressions conducted for 1992-2009. This is the reason why the term $a_4 \log(\eta_t)$

[9] At this point, it is worth noting that although a system of unemployment benefits exists in Colombia since 2002 (by Law 789), it had to be restructured in 2013 to enforce its use in correspondence to its nature. This implies, in contrast to the emphasis placed by the literature on the wage impact of such benefits, that the unemployment benefit system in Colombia has been irrelevant as a determinant of wage setting.

is omitted.

A final adjustment is the exclusion of $a_6 \log\left(\tau_t^i\right)$ due to the lack of information on income taxes in consistent time series. One problem is that effective tax rates are not available. The other problem is that using the statutory rates would not be accurate given the features of our panel data. The reason is that the income tax scheme sets differentiated rates by income group, with low incomes subject to tax exemptions. Therefore, as we only have data on average wages across sectors, the use of statutory tax rates could distort the true effect of income taxation on wages.

The central institutional extension is the inclusion of the real minimum wage, which acts as a reference for real wages. This implies considering the new term $a_2 \log\left(W_t^{\min}\right)$, where $W_t^{\min}$ denotes the real minimum wage. A positive sign on $\hat{a}_2$ is expected.

This extension is consistent with the findings in Iregui *et al.* (2012), who show that nominal minimum wage increases and past inflation are considered as important factors by firms when adjusting their nominal wages. Their analysis also reveals that most firms adjust nominal wages annually at rates that are roughly equivalent to the observed rate of CPI inflation, and none of them cut wages.

A second extension is related to relative prices and seeks to capture the role exerted by the price wedge between manufacturing and total prices in wage setting. The variable we consider is π_t, the ratio of manufacturing prices (p_t^m) over consumption prices (p_t^c), and we expect a negative influence over net real wages. The reason is the following. Over time, manufacturing prices tend to grow less than consumption prices for a twofold reason. First, it is a capital-intensive sector subject to quick technological change and efficiency gains that are translated into lower prices (or higher quality at equal prices). Second, it is the most exposed sector to global competition and, as such, bears the most important downward pressure on prices. This forces firms to use labour compensation as a key adjustment mechanism to ensure competitiveness. In this context, given that nominal wages in Colombia are indexed to the CPI, the larger is the wedge between the two, the bigger is the tension in terms of the wage setting mechanism: while workers grade their net wages having as reference consumption prices, firms appraise their expected benefits as a function of manufacturing prices. Here is the wedge and the tension pushing real net wages down when the wedge increases.

An important remark here is that the real minimum wage and the price wedge can be considered two sides of the same coin capturing the close attachment of wages to the evolution of prices. The reason is the following. If nominal minimum wages are closely tied to consumption prices, then the ratio $(W_t^{\min}/p_t^m)$ will not differ that much in its evolution from the price ratio p_t^c/p_t^m, which is the inverse of our price wedge variable p_t^m/p_t^c. Having confirmed this empirically, our empirical models will take two alternative forms to avoid multicollinearity problems, one with the real minimum wages as explanatory variable, as in equation (1.2) below; and a second one considering the price wedge instead: equation (1.3).

A final extension is related to the external sector and consists on the inclusion of a measure of trade openness (op) through the addition of the new aggregate term $a_4 \log(op)$. The expected incidence of international trade, or globalization, on net wages is not clear. On one side, the literature is still far from reaching consensus on the causal relationship between exports and productivity (and thus wages). Some relevant studies argue that exports cause efficiency gains (and would thus boost wages), while some other claim that efficiency progress is what allows exports to increase.

In the case of Colombia, we have two pieces of information that make us look at the imports' side. First, exports have not led the opening process (see Figure 1.3a). Second, as shown by Figure 1.6a in the Appendix 1, the share of manufacturing exports has remained stable between 1965 and 2010. Hence, if the industrial imports of goods are the catalyst variable, then we should expect a positive influence on net wages for a twofold reason.[10]

First, on account of the positive relationship between wages and the import of interme-

[10] Relevant studies claiming a negative relationship between imports and local wages point to massive imports of consumption goods in developed economies. For example, for the United States, Autor *et al.* (2013, 2014) show that wages/cumulative earnings were lower after the spectacular rising of the Chinese imports, mainly of consumption goods.

diate goods, as demonstrated in Amiti and Davis (2011). This hypothesis is endorsed by the fact that two thirds of total industrial imports in Colombia consist on intermediate inputs. And second because, as pointed out by Arbache *et al.* (2004), one of the consequences of increasing trade openness is a rapid inflow of foreign technology as a result both of foreign direct investment and increased imports. In-flowing technology is skill-biased because it is mainly designed in industrialized economies which are relatively skill intensive. Thus, the acquisition of new technologies from developing countries is normally accompanied by a greater demand of skilled labour. On the positive side this causes an upward pressure on relative wages (which is the effect we are capturing), although there is also a negative consequence in terms of increased wage inequality.[11]

Overall, these empirical adjustments and extensions leave us with two versions of the empirical model. In the first version, we have the real minimum wages:

$$\log\left(W_t\right) = a_0 + a_1 \log\left(Y_{t/}/N_t\right) + a_2 \log\left(u_t\right) + a_3 \log\left(W_t^{\min}\right) + a_4 \log(\tau_t^p) + a_5 \log\left(op_t\right) + \varepsilon_t, \tag{1.2}$$

while in the second version we have the ratio of manufacturing prices over consumption prices:

$$\log\left(W_t\right) = a_0 + a_1 \log\left(Y_{t/}/N_t\right) + a_2 \log\left(u_t\right) + a_3 \log\left(\pi_t\right) + a_4 \log(\tau_t^p) + a_5 \log\left(op_t\right) + \varepsilon_t. \tag{1.3}$$

To conclude, note that all coefficients will have to be interpreted as delivering elasticities. It is in this context that we treat payroll taxes as the tax wedge between gross and net wages.

[11] Growing empirical evidence on the effects of globalization gives support to a positive relationship between trade liberalization and wage inequality in developing countries –see, for example, Caselli (2014), Meschi and Vivarelli (2009), and Attanasio *et al.* (2004).

1.4 Empirical issues

1.4.1 Data

We use a panel database with a cross-section dimension of $N = 19$ sectors and a time dimension of $T = 34$ years covering the period 1976-2009.[12] Table 1.1 presents the variables and the corresponding sources.

Net average real wages per worker are obtained 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 calculated as the real wage bill before taxes in sector i over paid employment in that sector. Labour productivity is computed as the real value added in sector i over total employment in that sector, where total employment includes paid and unpaid workers. It is also obtained from the EAM. As explained above, trade openness is computed as exports plus imports in sector i over output in that sector. It is also obtained from the EAM. Note that these are the three variables for which detailed homogeneous time series information across sectors exists.

As stated, informal employment in Colombia is massive. This goes together with very limited statistics covering labour issues in non-urban areas. Consequently, the unemployment rate series available for the whole sample period is based on urban unemployment. Although it does not exactly fit the industrial sector, this is the only one we can use.

Data on nominal minimum wages is collected from public information/legislation published annually. Similarly, the statutory payroll tax rates (including health, pension, professional hazards, and parafiscal contributions) are computed from information supplied by LEGIS in its official annual reports. Note, therefore, that we use statutory rates instead of any sort of effective tax rate. In this way, we avoid dealing with econometric problems related to the use of effective tax rates, namely the simultaneity in the determination of wages and payroll tax rates, and spurious variability in payroll tax rates.

[12] The detailed list of sectors is provided in Table 1.6, in the Appendix 1.

Table 1.1: Definitions of variables.

Variables		Sources	Other variables and subindices	
W_{it}	Net real wage	(1)	d^{92}	Dummy: value 1 1992 onwards
Y_{it}	Real GDP	(1)	d^{8083}	Dummy: value 1 in 1980-1983
N_{it}	Total employment	(1)	d^{9700}	Dummy: value 1 in 1997-2000
$\frac{Y_{it}}{N_{it}}$	labour productivity	(1)	d^{0809}	Dummy: value 1 in 2008-2009
op_{it}	Trade openness $\left[\frac{(exports+imports)_{it}}{output_{it}}\right]$	(1)	i	$= 1, ..., 19$ sectors
u_t	Urban unemployment	(2)	t	$= 1, ..., 34$ years
$W_t^{\min}$	Real minimum wage	(3)		
τ_t^p	Statutory payroll tax rates	(4)		
p_t^m	Manufacturing prices	(2)		
p_t^c	CPI deflator	(2)		
π_t	Relative prices $\left[\frac{P^m}{P^c}\right]$	(2)		

Notes: All nominal variables are deflated by the manufacturing price index (base: June 1999).

(1) EAM; (2) DANE; (3) Legislation published annually; (4) LEGIS.

Manufacturing prices and the CPI index are also taken from the DANE. These two indices are used to compute the ratio of relative prices.

We include a set of time dummies to control for macroeconomic shocks that may affect all sectors. In this way, d^{92} help us to capture potential structural breaks in the wage elasticities arising from the institutional and trade 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 wages in Colombia. All nominal variables are deflated with the manufacturing price index.

1.4.2 Stylized facts

Table 1.2 provides descriptive information on some crucial macroeconomic variables of interest. Subscript i denotes information corresponding to the average of the 19 sectors in which the Colombian industry is disaggregated. Detailed industry information is not available for the real minimum wage, manufacturing prices, the CPI index, and the payroll tax rate. All data is supplied for the two relevant periods of analysis –the slow transition between import substitution and trade liberalization in 1976-1991; and the institutional and trade reforms years of 1992-2009.

Since the mid-seventies until 2009, real economic growth of the Colombian manufacturing sectors was around 4.2% on average, with no significant differences between periods. In spite of these positive growth rates throughout, the industrial standstill in the seventies and the debt crisis in the eighties resulted in low employment growth rates (0.3 % on average in 1976-1991) and, thus, in high labour productivity growth rates (3.8% on average). In contrast, gross wages (i.e., the addition of net wages and statutory payroll taxes) grew at 2.1%, thus prompting an annual fall in the unit labour costs of 1.7% on average. Note also, that net real wages (payroll taxes excluded) and real minimum wages grew in line (at around 1.9%) reflecting highly tied adjustments to the cost of living. These years were also characterized by high inflation rates, amounting to 22.5% if measured by the CPI,

Table 1.2: Macro developments in the Colombian manufacturing industry.

	Y_{it}	N_{it}	Y_{it}/N_{it}	W_{it}	τ_t^p	$\left[W_{it}(1+\tau_t^p)\right]$
1976-1991	4.16	0.31	3.85	1.94	16.1	2.13
1992-2009	4.30	-0.30	4.60	3.85	28.9	4.50
1976-2009	4.23	0.70	3.53	2.99	23.1	3.42
	W_t^{min}	p_t^m	p_t^c	π_t	op_{it}	$\left[\frac{W_{it}(1+\tau_t^p)}{Y_{it}/N_{it}}\right]$
1976-1991	1.86	21.5	22.5	-0.97	48.4	-1.72
1992-2009	3.59	9.0	11.2	-2.23	96.8	-0.10
1976-2009	2.80	14.9	16.5	-1.64	74.0	-0.11

Notes: denotes the growth rate corresponding to the annual average of the sectorial growth rates. All variables are expressed in percent.

and 21.5% by manufacturing prices. As a consequence, the price wedge went down by 1 percentage point annually, on average.

The nineties were years of labour reform and trade liberalization processes, but also of deindustrialization and growing expansion of the services sector. They were followed, though, by the recovery of the manufacturing sector in 2000-2009 driven by a rising capital accumulation (boosted both by domestic and foreign direct investment)[13] and the substitution of domestic by imported raw materials. Altogether, these developments caused employment to fall (by 0.30% annually, on average) and resulted in the acceleration of the growth rate of labour productivity (4.6%), which was 0.8 percentage points larger than in 1974-1991. In turn, gross real wages grew in line with productivity (4.5% on average), thus denying further progress in real unit labour costs. Beyond productivity gains, it should be

[13] In the seventies and eighties, FDI represented 0.3% of the GDP on average, while between 1997 and 2007 it represented 3.8%. Source: DANE.

noted that upward pressures on gross real wages arose from increases in payroll taxes paid by firms, and from the acceleration in the increase of the price wedge, reflecting the speedier deceleration of inflation in manufacturing prices than in the cost of living.

These contrasted periods, both in terms of economic developments and policy agenda (as discussed in Section 1.2), lead us to the estimation of equations (1.2) and (1.3) taking into account potential changes in the elasticities brought by the new institutional framework. As we explain next, this estimation is conducted by considering several types of dummies and interactions among selected explanatory variables.

1.4.3 Econometric methodology

Empirical equations (1.2) and (1.3) are extended in two directions. First, due to the relevance of adjustment costs in wage setting, 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 net real wage. Second, as we work with a two-dimensional panel data, we add fixed effects in order to control for unobserved heterogeneity among sectors.

Given the dynamic nature of the extended models, equations (1.2) and (1.3) will be estimated as partial adjustment models taking the following general form:

$$\ln(W_{it}) = \alpha_i + \gamma \ln W_{it-1} + \beta \ln X_{it} + u_{it}, \qquad u_{it} \sim i.i.N\ \left(0, \sigma^2\right) \tag{1.4}$$

where the subscripts i and t are sector and time indices, respectively; α_i is a sectorial cross-section intercept; $W_{i,t-1}$ is the lagged dependent variable with as inertial (or persistence) coefficient; X is a vector of explanatory and control variables with β as the set of estimated parameters capturing their influence on the dependent variable W_{it}; and u_{it} is the error term.

Unit root test

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

Table 1.3: Unit Roots Test, 1974-2009.

Different variables across sectors				Common variables across sectors				
	W_{it}	Y_{it}/N_{it}	op_{it}		W_t^{min}	π_t	τ_t^p	u_t
Null hypothesis: individual unit root process				Null hypothesis: variable is stationary				
MW	134.43 [0.000]	111.74 [0.000]	50.55 [0.084]	KPSS	0.08 [0.146]	0.08 [0.146]	0.09 [0.146]	0.07 [0.146]
Result	I(0)	I(0)	I(0)	Result	I(0)	I(0)	I(0)	I(0)

Notes: All variables are expressed in logs; MW and KPSS tests computed using intercept and trend; for the MW test, p-values in brackets; for the KPSS test, 5% critical values in brackets.

$I(0)$ which, by definition, yields a long-run cointegrating vector.

In order to check the order of integration of the variables, we perform a set of stationary and panel unit root tests depending on the type of variables to be dealt with. In particular, we use the Kwiatkowski *et al.* (1992) stationary-test –KPSS henceforth– for the variables that are common across sectors. For the variables that are sector-specific, we use the test proposed by Maddala and Wu (1999) –MW–, which is a panel unit root test based on Fisher's (1932) results.[14] The MW test assumes, under the null hypothesis, that all series are non-stationary, against the alternative that at least one series in the panel is stationary. We use the KPSS test because it allows direct testing of the null hypothesis of stationarity.

On the other hand, we conduct the MW test because, in general, panel-based 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

[14] Using Monte Carlo simulations, Maddala and Wu (1999) conclude that the Fisher test is a better test than the Levin and Lin (1993) and the Im *et al.* (2003) tests. They also highlight that the Fisher test is simple and straightforward to use.

test regressions can be determined separately for each sector.

As noted in the last row of Table 1.3, the overall conclusion drawn from these tests is that all variables are stationary $I(0)$. Hence, we can proceed with stationary panel data estimation techniques.

Estimation method

Given the panel structure of our database, models (1.2) and (1.3) are estimated by applying Ordinary Least Square (OLS) and Fixed Effects (FE). 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; by the well-known simultaneity between net real wages, labour productivity, and unemployment; and by the potential correlation between relative prices and the error term, on account of the simultaneity of wage and price setting.

Regarding the first potential problem, it is well known that OLS estimates gives rise to a "dynamic panel bias" (Nickell, 1981) causing the persistence coefficient to be overestimated. The reason is that the estimated coefficient on lagged wages will be inflated by attributing some predictive power to it that actually belongs to the sector's fixed effect. A potential response would then be to apply the within-groups (or fixed effects) estimator, but this does not fully offset the dynamic panel bias. This was explained by Nickell (1981), who pointed out, also, that when T is large and N is small ($T > N$), this bias is likely to be insignificant. In contrast, Judson and Owen (1999) found that even with a time dimension as large as 30, this estimator would be downward biased 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 $T = 34$ and $N = 19$), we can not fully exclude the existence of a dynamic panel bias. We have thus 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 1 (Table 1.7), show that these sets of estimates do not differ and allow us to conclude that the FE estimator is potent in our case.

We have not compute the System GMM estimator (Blundell and Bond, 1998), which is a common option to deal with dynamic panel biases. The reason is that the consistency of this estimator depends on the fact that $N \rightarrow \infty$ grows sufficiently fast relative to T. Since this is not the case here, the estimation by System GMM would not yield dramatic consistency gains over the FE estimator; not to mention the instrument proliferation problem, which emerges even we limit and collapse the massive amount of instruments due to the number of potentially endogenous regressors in our empirical models.

Thus, to deal with the potential endogeneity of labour productivity, the price wedge and unemployment, we estimate FE by Two Stage Least Squares (FE-2SLS), where the instruments are lags and differences of the explanatory variables. To select the most appropriate set of instruments we rely on the performance of the following sets of tests. First, an LM test checking for underidentification (i.e., that the excluded instruments are not relevant, meaning non correlated with the endogenous regressors). This is denoted as U in Tables 1.5 and 1.8 (Appendix 1), and the null hypothesis is that the equation is under-identified (against the alternative that the model is identified). Second, the Hansen 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 term) and, on the other side, that the excluded instruments are correctly dropped from the estimated equation. Third, an F test checking for weak identification (denoted as W), where the null hypothesis is that the instruments are correlated with the endogenous regressors, but only weakly.[15]

1.5 Results

Tables 1.4 and 1.5 present, respectively, the estimated short- and long-run elasticities obtained through the various estimation methods just discussed: OLS (in columns 1 and 5)

[15] Note that for this test we do not show the critical values (since the standard ones of Stock and Yogo, 2005, are not available), and we follow "the rule of thumb" of suggested by Staiger and Stock (1997) in such cases: an F-statistic near or above 10.

displayed as a reference to check whether the persistence coefficients obtained by the other methods are lower, as expected; FE (in columns 2 and 6); and FE-2SLS (in columns 3, 4, 7 and 8).

In the specification presented in column 3, we consider labour productivity and the respective interaction as weakly exogenous. By so doing, we use as instruments lags and differences (the first lag and the difference of the third lag for labour productivity, and just the first lag for the interaction term). In the specification of column 4, we also consider unemployment and their interaction with the dummy variable as weakly exogenous, and use the first lag in both cases. Regarding the specifications presented in columns 7 and 8, we keep these assumptions and further consider the price wedge and its interaction as endogenous, having their second lag as instrument.[16]

All this information is classified in two blocks, the one on the left-hand-side corresponding to the estimation of equation (1.2) and the one on the right-hand-side to equation (1.3).

To these benchmark empirical models we add interactions between dummy and explanatory variables so as to capture potential structural changes related to the Colombian institutional reform process. Accordingly, we next overview the results, and place some attention on the stability, or not, of the estimated elasticities between 1976-1991 and 1992-2009.

1.5.1 Wage setting before the reform process (1976-1991)

When examining the econometric analysis, if we have to favor some particular specification, we would choose the results in columns 4 and 8 for a three-fold reason: (i) instrumental variables are used to deal with potential endogeneity problems; (ii) the performance of the

[16] It is worth noting that we have estimated a wide set of specifications using different combinations of instruments depending on whether we consider labour productivity, unemployment and the price wedge either as weakly exogenous variables or endogenous. However, we only present those specifications with the best performance in the instrumental tests.

Table 1.4: Estimated wage equations.

Dependent variable: w_{it}

	Equation (1.2)				Equation (1.3)			
	OLS	FE	FE-2SLS	FE-2SLS	OLS	FE	FE-2SLS	FE-2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
W_{it-1}	0.79 [0.000]	0.42 [0.001]	0.25 [0.094]	0.27 [0.080]	0.82 [0.000]	0.54 [0.000]	0.45 [0.002]	0.46 [0.002]
Y_{it}/N_{it}	0.08 [0.000]	0.03 [0.363]	0.33 [0.020]	0.32 [0.015]	0.07 [0.000]	0.03 [0.382]	0.27 [0.041]	0.23 [0.067]
u_t	−0.12 [0.025]	−0.15 [0.000]	−0.05 [0.215]	−0.15 [0.318]	−0.01 [0.696]	0.04 [0.257]	0.11 [0.035]	0.18 [0.001]
W_t^{min}	0.38 [0.000]	0.69 [0.000]	0.78 [0.000]	0.95 [0.000]				
π_t					−1.37 [0.000]	−1.46 [0.000]	−2.37 [0.000]	−3.12 [0.000]
τ_t^p	−0.40 [0.001]	−0.24 [0.000]	−0.90 [0.026]	−0.94 [0.012]	−0.40 [0.000]	−0.13 [0.038]	−0.74 [0.034]	−0.82 [0.010]
op_{it}	0.01 [0.015]	0.00 [0.703]	0.08 [0.112]	0.08 [0.117]	0.01 [0.038]	0.00 [0.598]	0.06 [0.152]	0.05 [0.155]
d_{92}	−0.40 [0.509]	0.16 [0.717]	0.41 [0.272]	1.09 [0.332]	0.26 [0.534]	0.61 [0.034]	−1.54 [0.106]	−1.61 [0.009]
$W_{it-1}*d_{92}$	0.13 [0.005]	0.14 [0.027]	0.48 [0.000]	0.46 [0.000]	0.11 [0.019]	0.10 [0.103]	0.35 [0.002]	0.32 [0.012]
$Y_{it}/N_{it}*d_{92}$	−0.05 [0.032]	−0.04 [0.342]	−0.25 [0.000]	−0.24 [0.000]	−0.04 [0.072]	−0.02 [0.492]	−0.20 [0.004]	−0.17 [0.024]
u_t*d_{92}	0.05 [0.345]	0.12 [0.003]	−0.01 [0.862]	0.05 [0.769]	0.03 [0.347]	0.03 [0.417]	−0.11 [0.203]	−0.17 [0.053]
$W_t^{min}*d_{92}$	−0.05 [0.695]	−0.13 [0.123]	−0.46 [0.003]	−0.59 [0.034]				
π_t*d_{92}					0.11 [0.650]	−0.01 [0.970]	1.30 [0.002]	1.79 [0.011]
$\tau_t^p*d_{92}$	0.08 [0.555]	0.10 [0.223]	0.66 [0.057]	0.72 [0.033]	−0.29 [0.028]	−0.41 [0.000]	0.16 [0.627]	0.17 [0.423]
$op_{it}*d_{92}$	−0.01 [0.049]	−0.02 [0.028]	−0.04 [0.047]	−0.04 [0.045]	−0.01 [0.108]	−0.01 [0.108]	−0.03 [0.093]	−0.03 [0.093]
d_{8083}	−0.03 [0.172]	−0.06 [0.000]	−0.06 [0.000]	−0.08 [0.001]	−0.02 [0.195]	−0.02 [0.092]	−0.02 [0.145]	−0.02 [0.105]
d_{9700}	0.02 [0.047]	0.01 [0.062]	0.02 [0.051]	0.03 [0.028]	−0.03 [0.000]	−0.05 [0.000]	−0.02 [0.363]	−0.03 [0.068]
d_{0809}	−0.04 [0.029]	−0.02 [0.259]	−0.02 [0.177]	−0.03 [0.076]	−0.06 [0.000]	−0.04 [0.103]	−0.05 [0.047]	−0.06 [0.074]
c	−0.58 [0.195]	−0.46 [0.341]			2.49 [0.029]	4.37 [0.259]		
$Obvs.$	646	646	627	627	646	646	627	627
$Adj.R^2$	0.97	0.95	0.92	0.92	0.97	0.96	0.93	0.93
U			6.23 [0.044]	6.61 [0.037]			5.53 [0.063]	7.75 [0.021]
H			0.49 [0.486]	1.11 [0.291]			0.29 [0.588]	0.21 [0.643]
W			28.32	14.80			15.31	7.39

Notes: All variables are expressed in logs. P-values in brackets. OLS: Ordinary Least Squares. FE: Fixed Effects. FE-2SLS: Fixed effects using Two Step Least Squares.
U : Under identification test. H : Hansen test. W : Weak identification test.

instrumental variable tests confirm the appropriateness of the instruments in most cases (since they are simultaneously exogenous and highly correlated with the endogenous regressors); and (iii) all variables have the expected sign according to the underlying theoretical relationships.

In any case, the four sets of estimates (i.e., those presented in columns 3 and 4 in the left block, and those in columns 7 and 8 in the right block) provide a similar picture with significant effects on net wages of labour productivity, the real minimum wage, the price wedge and payroll taxes, all displaying the expected sign. In contrast, the role of trade openness is found mildly significant and not very strong, while the evidence on the unemployment rate is mixed.

Regarding the first model, results in column 4 display a short-run elasticity of wages with respect to labour productivity of 0.32 (Table 1.4) increasing to 0.44 in the long-run (Table 1.5) but still far away from unity, which is the theoretical benchmark value. Since only 44% of productivity gains are translated into a higher net compensation of workers (32% initially), wage setting in the Colombian manufactory industry is not fundamentally driven by labour productivity. This leaves space for other potentially more relevant determinants related to institutional arrangements (such as payroll taxation and minimum wages) and other determinants related to the indexation of wages with respect to prices (again the minimum wage or, alternatively, the price wedge), which will turn out to be crucial wage setting drivers.

The short-run elasticity of wages with respect to the real minimum wage is estimated at 0.95, increasing to 1.30 in the long-run. This captures the upward pressure that minimum wages exert on wages and implies that a 1% increase in the real minimum wage causes net real wages to grow by 1.30% (0.95% initially). The counterpart of this result is provided in column 8, in which the real minimum wage is replaced by the price wedge. We find that the short-run elasticity of wages with respect to relative prices is -3.12, increasing to -5.80 in the long-run. This implies that a 1% increase in the ratio of manufacturing prices over consumption prices cause wages to decrease by 5.80% overall (3.12% initially).

Table 1.5: Long-run elasticities.

	Equation (1.2)				Equation (1.3)			
1976-1991								
	OLS	FE	FE-2SLS	FE-2SLS	OLS	FE	FE-2SLS	FE-2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\xi_{(Y/N)}$	0.41	0.05	0.44	0.44	0.40	0.07	0.50	0.43
ξ_u	-0.56	-0.25	-0.07	-0.20	-0.07	0.08	0.20	0.34
$\xi_{W^{min}}$	1.83	1.18	1.05	1.30				
ξ_π					-7.48	-3.19	-4.34	-5.80
$\xi\tau^p$	-1.93	-0.41	-1.21	-1.30	-2.25	-0.34	-1.35	-1.52
ξ_{op}	0.07	-0.01	0.11	0.11	0.06	-0.01	0.11	0.10
1992-2009								
	Equation (1.2)				Equation (1.3)			
	OLS	FE	FE-2SLS	FE-2SLS	OLS	FE	FE-2SLS	FE-2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\xi_{(Y/N)}$	0.40	-0.01	0.29	0.29	0.41	0.02	0.40	0.28
ξ_u	-0.89	-0.07	-0.24	-0.35	0.80	0.04	-0.44	0.06
$\xi_{W^{min}}$	4.54	1.25	1.20	1.33				
ξ_π					-17.8	-4.05	-5.53	-6.04
ξ_{τ^p}	-4.44	-0.32	-0.90	-0.86	-9.84	-1.51	-3.02	-2.95
ξ_{op}	0.00	-0.05	0.15	0.15	0.00	-0.05	0.17	0.12

In other words, from this alternative specification, we are able to gauge to what extent manufacturing firms are forced to reduce wages to regain competitiveness when prices of manufacture goods become relatively more expensive (and vice-versa).

Regarding payroll taxation, the short-run coefficient of -0.94 increases to -1.30 in the long-run. This implies that 94% percentage of the tax burden is immediately shifted to workers in the form of lower net wages, which further decrease in the long-run. This is the outcome of the direct and indirect effects of payroll taxes on labour market outcomes: they have a negative impact on the labour demand (on account of the extra labour cost they represent for employers), but they also decrease net real wages and generate a compensation effect. Which effect dominates crucially depends on wage persistence. In case of significant wage persistence (and hence a large long-run effect), wage compression may overcome the direct negative effects on employment generated by payroll taxation.

In any case, whether the increase in the payroll taxes caused significant job cuts, or not, in Colombia is something we cannot answer properly in this study. It is generally expected that the larger the extent of payroll tax shifting on net wages, the lower the negative consequences on employment. However, there is not yet a consensus in the literature on the fact that cutting payroll taxes increases employment (and vice-versa). Rather, the empirical evidence suggests that payroll taxation might have asymmetric effects on wages and employment. On one side, there are studies suggesting that payroll tax increases have negative effects on net wages and employment –Kugler and Kugler (2009), Beach and Baulfour (1983), and Hamermesh (1979); while, on the other side, there are even more studies showing that payroll tax rate cuts do not generate significant effects on employment, even though they have sizeable positive effects on net real wages –Cruces *et al.* (2010), Bennmarker *et al.* (2009), Bauer and Riphahn (2002), and Gruber (1997).[17]

[17] There is also an issue related to the consequences of changes in social protection systems (and the corresponding taxation changes) on formal and informal employment. For Colombia this is studied in Camacho *et al.* (2014), who show that informal employment increased by 4 percentage points as a consequence of the government's expansion of social programs in the early 1990s.

In the case of the unemployment rate, the results differ between models. Estimates for model (1.2) point to the irrelevance of unemployment on wage determination (see the non-significant coefficients in columns 3 and 4), while the estimates for model (1.3) suggest a significant and positive influence, at odds the standard theoretical prediction (columns 7 and 8). On this account, we have verified a negative correlation of 0.59 between the unemployment rate and the price wedge, which seems to be affecting the estimated coefficient on unemployment (the one on the price wedge is robust to the absence of unemployment as explanatory variable, but the contrary does not hold). This is the reason why we credit the results from model (1.2). In this way, the irrelevance of urban unemployment as determinant of net real wages may be a reflection that it is a poor proxy of the dynamics of industrial unemployment or, being a reasonable proxy after all, may be a proof of the findings in Iregui *et al.* (2012), who show that Colombian firms consider some aggregate factors, the unemployment rate among them, as having minor relevance in determining nominal wage increases.

The last result is the scant influence exerted by the degree of international trade openness on wage setting. The short-run elasticity, which is 0.08 and increases to 0.11 in the long-run, indicates that a 1% increase in the degree of trade openness causes net real wages to grow by 0.11%. This result is not surprising given that in the seventies and eighties Colombia was a closed economy, and the industry was mainly based on manufactures that had a low international exposure. Of course, some branches had already high rates of trade openness (for example, the manufacture of machinery and equipment; and medical instruments), but they represented a minimal share of the industry (2.4%).

In any case, the positive influence of trade openness on wage setting should be interpreted in connection to the works by Amiti and Davis (2011) and Arbache *et al.* (2004). That is, our result is on one side connected to the positive relationship between wages and the import of intermediate goods, which was large in the Colombian industry. And, on the other side, to the hypothesis of skill-biased technological change and the subsequent increase in skilled-labour wages, which is likely to apply to those manufacturing sectors more (and growingly) exposed to international trade.

Finally, the macroeconomic shocks controlled by the additive dummies (d^{8083}, d^{9700}, and d^{0809}) have had, in general, a significant influence on wages (although there are some exceptions depending on the specification): clearly negative during the debt and Great Recession crises, but not conclusive during the financial bubble and burst of the second half of the nineties.

1.5.2 Wage setting with enhanced labour flexibility, payroll taxation, and trade exposure (1992-2009)

The analysis of the interactions allows us to check for the existence of structural changes in the wage setting effects of the various determinants.

A first salient result (robust across estimation methodologies and model specification) is an increase in wage persistence ranging from 0.32-0.35 in model (1.3), to 0.46-0.48 in model (1.2). This increase runs in parallel to the lower volatility of real net wages in the Colombian industry in 1992-2009 (specifically in 1995-2005), which echoes a much reduced volatility in the rate of inflation. In addition, the larger persistence is accompanied by lower short-run coefficients, so that it is the relative magnitude of these changes what determines the variation in the long-run response of wages to its various determinants.[18]

Regarding the relationship between wages and labour productivity, we find a fall in the long-run elasticity, going from 0.44 to 0.29 in model (1.2), and from 0.43 to 0.28 in model (1.3) (see Table 1.5, columns 4 and 8). This implies that the relationship between these two variables has become weaker. This can be considered one of the collateral pitfalls of the institutional and trade reform process since it is tempting to associate this finding to the increasing current account deficit experienced by Colombia in the last decades. The counterpart is, of course, that wage setting has become more persistent.

In contrast with productivity, the long-run elasticity of wages with respect to minimum

[18] Note that, as wage persistence increases, the long-run coefficients have a tendency to become larger. However, this increase might be offset, or even overtaken, by the falls in the short-run wage elasticities.

wages remained unaltered around 1.30 in 1992-2009. This implies that the institutional and trade reforms have not crucially affected the adjustments of net real wages vis-à-vis the net real minimum wage. This stability is confirmed when looking at the role played by the price wedge in model (1.3) which remains quite stable (with a slight increase in the long-run elasticity going from -5.80 to -6.04 in 1992-2009. In any case, this mildly larger sensitivity can be associated to the fact that adjustments in nominal wages became more tied to CPI inflation since 1991, and also to the increased exposure to international markets and the resulting dramatic downward pressures on producer and consumer prices. As noted before, the asymmetric price response to these pressures caused a fall in the price wedge. Thus, in the growing liberalized environment of 1992-2009, this increased sensitivity to the price wedge should come as no surprise.

Together with the result on productivity, another key novelty in this period is that net real wages became less sensitive to changes in payroll taxes with a long-run elasticity falling from -1.30 to -0.86 in model (1.2). Although, the long-run coefficients of payroll taxes in model (1.3) increase when interacted with d^{92}, it is important to note their lack of significance (column 8 in Table 1.4). According to this result, and taking into account the weaker instruments of this specification, we credit the findings obtained through model (1.2).

This lower sensitivity of wages with respect to payroll taxes takes place in the context of their significant rise in the 1990s. This result is consistent with the previous ones along the following lines. Recall that firms can shift the tax burden on workers through lower nominal net wages or higher prices. In a growing liberalized labour market, wage cuts become more feasible than rising prices (in an otherwise growing open environment), hence the preferred use of the first channel. However, the new situation of full indexation of nominal wages to the cost of living in the nineties was effectively introducing wage floors to firms. These wage rigidities, which prevented firms to shift the tax burden to the workers in the form of lower net wages, help to explain why the sensitivity of wages with respect to payroll taxes becomes lower in this period.

The last finding is related to the increase in the long-run elasticity of wages with respect to the degree of trade openness, which rises from 0.11 to 0.15 in model (1.2), and from 0.10 to 0.12 in model (1.3). This increase is consistent with the predicted behavior of wages according to the work by Arbache *et al.* (2004). Following a wide liberalization process, the relative demand for skilled labour in developing countries is expected to rise, thus leading to wage increases in this group of workers.

1.6 Multiplicative interactions and marginal effects

As a complementary exercise, in this Section we present the crucial findings from the estimation of equations (1.2) and (1.3) as multiplicative interaction models (see Berry *et al.*, 2012, for details). Although this procedure is not suitable to investigate changes in the form of structural breaks, it gives us further insights on the relationship between wages and their main determinants when they are assumed to be cross-dependent.

In particular, we look at the marginal effects of real minimum wages and the price wedge on net real wages, conditional on the different values taken by payroll taxation and the degree of trade openness in 1976-1991 and 1992-2009. In this way, we check whether the reform process (increases in the exposure to international trade and payroll taxation) have affected the way wages react to these key determinants.

Thus, the new models considered take the following general form:

$$W_{it} = \alpha_i + \quad W_{it-1} + \beta_X X_{it} + \beta_{X\tau}(X_{it} * \tau_t^p) + \beta_{op}(X_{it} * op_{it}) + \beta_{\tau op}(X_{it} * \tau_t^p * op_{it}) + \lambda \mathbf{Z}_{it} + e_{it}, \tag{1.5}$$

where X is a scalar comprising our two variables of interest, $W_t^{\min}$ and π_t; and $\mathbf{Z}_{it}$ is a vector comprising the rest of the explanatory variables $-Y_{it}/N_{it}$, u_t , τ_t^p and $op_{it}-$ and various controls.[19]

[19] We also conducted the analysis having labour productivity interacted with payroll taxes and the degree of trade openness, but it was not successful. Still, the results are available upon request.

The marginal effects of our variables of interest will thus be given by:[20]

$$\frac{\partial w_{it}}{\partial X_{it}} = \alpha_i + \beta_X + \beta_{X\tau}\tau_t^p + \beta_{op}op_{it} + \beta_{\tau op}(\tau_t^p * op_{it}). \tag{1.6}$$

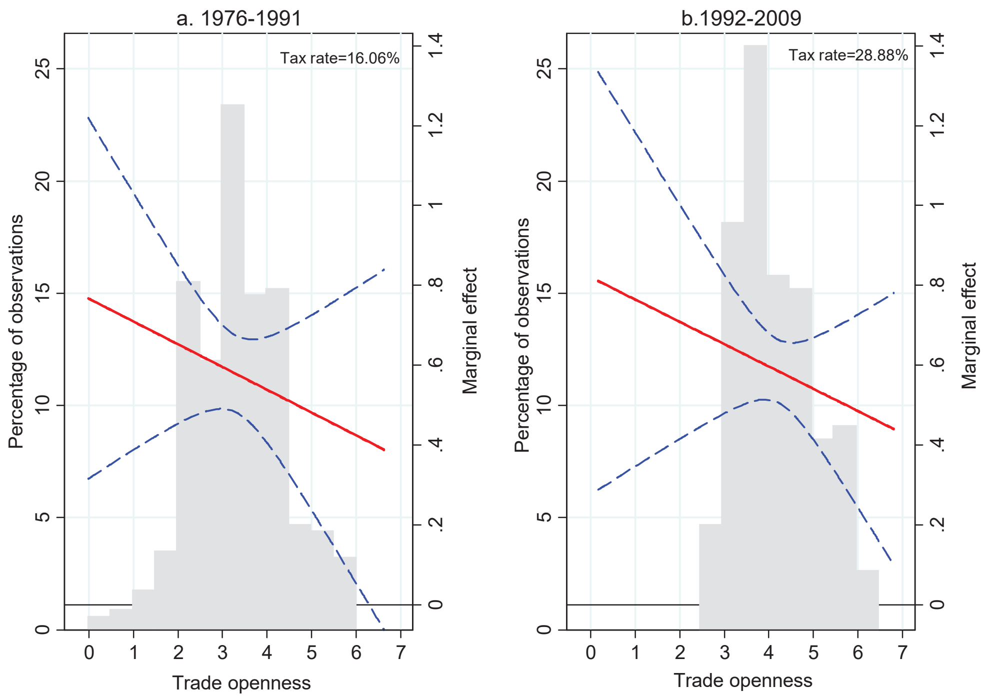
These marginal effects are presented in figures 1.4 and 1.5. Different values of trade openness (in logs) are listed in the horizontal axis, while the left vertical axis shows the histogram of trade openness corresponding to its distribution in 1976-1991 (in panel a) and 1992-2009 (in panel b). In the turn, the values in the right vertical axis indicate the magnitude of the marginal effect. The marginal effects are evaluated at the average payroll tax rate in each period. Note, also, that Figure 1.4 corresponds to the model specification with real minimum wages (column 4 in Table 1.8 in the Appendix 1), while Figure 1.5 corresponds to the model specification with the price wedge (column 8 in Table 1.8 in the Appendix 1).

The marginal effects of real minimum wages range between 0.4 and 0.8 depending on the degree of trade openness. Looking at the histogram, this range can be further narrowed to the interval 0.5-0.7 for the relevant values of trade openness in 1976-1991 (those between 7% and 55% corresponding to 2 and 4 in logs, as shown in Figure 1.4a) and to 0.55-0.65 in 1992-2009 (Figure 1.4b). Colombia was a relatively more open economy in 1992-2009, this is why the histogram shifts to the right. However, when the marginal effects are evaluated at the relevant trade openness and payroll tax values, there is no difference on the extent of the wage impact of real minimum wages. This result is consistent with the lack of structural change in this elasticity when evaluated through the analysis of multiplicative dummies as in Section 1.5.

Of course, this conclusion does not preclude the fact that larger values of trade exposure tend to reduce the sensitivity of wages with respect to the minimum wage. The point here is that this changing sensitivity has not shifted, on aggregate, in the new scenario of labour

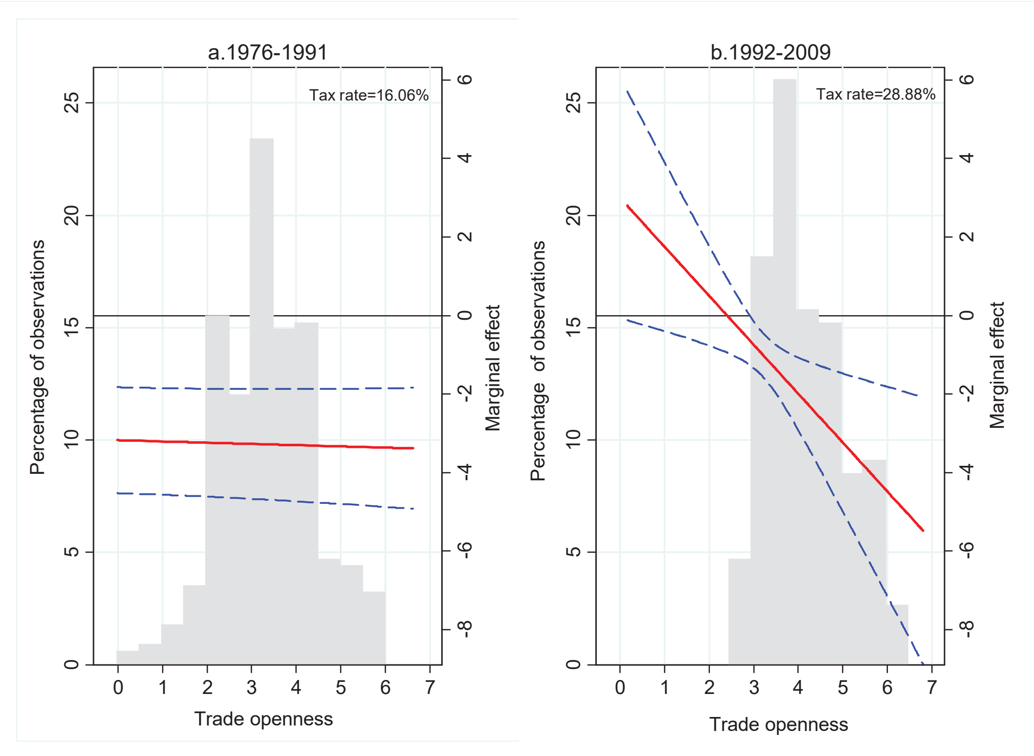
[20] The corresponding estimated equations are presented Table 1.8 in the Appendix 1. Here we only focus on the marginal effects derived from those estimations (more precisely, those presented in columns 4 and 8 in consistence with Section 1.5.

Figure 1-4: Marginal effects of real minimum wages on net real wages.



Notes: Dash lines represent 95% confidence interval. Trade openness in log.

Figure 1-5: Marginal effects of the price wedge on net real wages.



Notes: Dash lines represent 95% confidence interval. Trade openness in log.

market flexibility and enhanced trade liberalization. The implicit result at the sectorial level is that more open sectors will tend to have less responsive wages to changes in the minimum wage.

Figure 1.5 shows the marginal effects of the price wedge. The first remark on these results is that the marginal effects are negative and consistent with the analysis in Section 1.5. The second remark points to the flat slope in the first period, when payroll taxation was low on average (Figure 1.5a), which indicates that the sensitivity of wages with respect to the price wedge is not affected by the degree of trade openness.

In contrast, in a scenario of high payroll taxation such as the one in 1992-2009, the marginal effects are negatively sloped (Figure 1.5b). This implies that the sensitivity of real wages to the price wedge gets larger along with the degree of trade openness (cases of extreme openness, such as the ones in the right hand side of the histogram, correspond to the few very much open manufacturing sectors reported in the Appendix 1). This larger sensitivity goes in line with Bems (2014), who points out that increasing input imports leads to a greater response of the relative price to a given external adjustment.

1.7 Concluding remarks

We have studied wage determination in the Colombian manufacturing industry and paid specific attention to the consequences of the institutional and trade reforms carried out in Colombia during the nineties.

The first salient result is that productivity is not fundamentally driven by labour productivity in contrast to the standard theoretical prediction. In addition, the long-run elasticity of wages with respect to productivity has decreased to 0.30 in the nineties and noughties, from a value close to 0.45.

This suggests that the institutional measures undertaken by the government did not improve the scarce connection between workers' compensation and labour's efficiency progress. On the contrary, this connection has been worsened. Taking the benchmark one-to-one relationship between wages and productivity, the failure to reduce this detachment (or, even worse, increasing it) in a growing globalized economy should be given priority from policy makers, as it generates distortions in the process of achieving competitiveness in the manufacturing sector. This is the main lesson that can be drawn from our study in terms of policy measures.

If the Colombian economy continues to globalize at the same pace than in the last two decades, this call to strengthen the link between wages and productivity should be further reinforced.

Another main result is that wage progress in Colombia is largely tied to the cost of living. This conclusion is obtained from the important role played by the real minimum wage, or alternatively by the price wedge between manufacturing and consumer prices, as wage drivers. Here, however, we find no evidence of structural changes in the corresponding long-run relationships. It seems, therefore, that the purchasing power of workers has remained relatively isolated from the new institutional and trade scenario.

In terms of payroll taxation (paid by firms), we find the expected negative impact on net real wages, although we are unable to evaluate the net employment consequences of the increase in payroll taxes in the nineties. Our contribution here is the identification of a fall in the wage elasticity with respect to these taxes (from 1.30 to 0.86), which is the joint outcome of changes in the institutional setting such as the enhanced indexation of nominal wages to CPI inflation, and the asymmetric downward pressures on manufacturing and consumption prices resulting from the liberalization process. The consequence of this lower tax shifting is a loss in firms' cost competitiveness which, although we have not dealt with it, may use employment as an instrument to offset this loss.

What is, then, the counterpart of the lower explanatory power of productivity?

On one side, wage persistence has increased. This is connected to the lower capacity of firms to shift taxes. Although firms can shift the tax burden on workers through lower nominal net wages or higher prices, in a growing liberalized labour market wage cuts become more feasible than rising prices. Hence the preferred use of wage cuts. However, the new situation of full indexation of nominal wages to the cost of living in Colombia in the nineties was effectively introducing wage floors to firms. This explains the increase in wage persistence.

On the other side, the long-run elasticity of wages with respect to the degree of trade openness did also increase. This suggests that, following the liberalization process, the relative demand for skilled labour went up and pushed the compensation of this type of workers.

As a complementary exercise, we have also examined the marginal effects of real minimum wages and the price wedge on net real wages, conditional on the different values taken by payroll taxation and the degree of trade openness in 1976-1991 and 1992-2009. In this way, we have checked, from a different perspective, whether the impact of these wage determinants has changed along with the reform process. We have confirmed the stability of the relationship between net wages and the minimum wage, but we have identified a different sensitivity with respect to the price wedge depending on the level (low or high) of payroll taxation. This requires further research.

Overall, our findings call for a policy agenda in which wage indexation is reduced and a true wage bargaining system is brought in. This would allow workers' compensation to reflect more faithfully efficiency progress which, by all means, should be a critical target to solve the competitive problem of the Colombian industry and help, in this way, rebalancing the huge trade deficit attained with the liberalization program.

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, may have different effects on wages by type of worker. For example, regarding payroll taxes, international evidence suggests that there is less tax shifting for blue-collar than for white-collar workers. Beyond that, for developing countries there is growing empirical evidence showing that trade liberalization exerts a positive effect on high-skill labour wages, while there is no effect on low-skill labour wages.

Another research avenue is to aim at an individual assessment of how these reforms affected wage setting in each productive sector. In that case, the starting hypothesis would be that each sector's response is connected to its degree of exposure to international trade.

Appendix 1

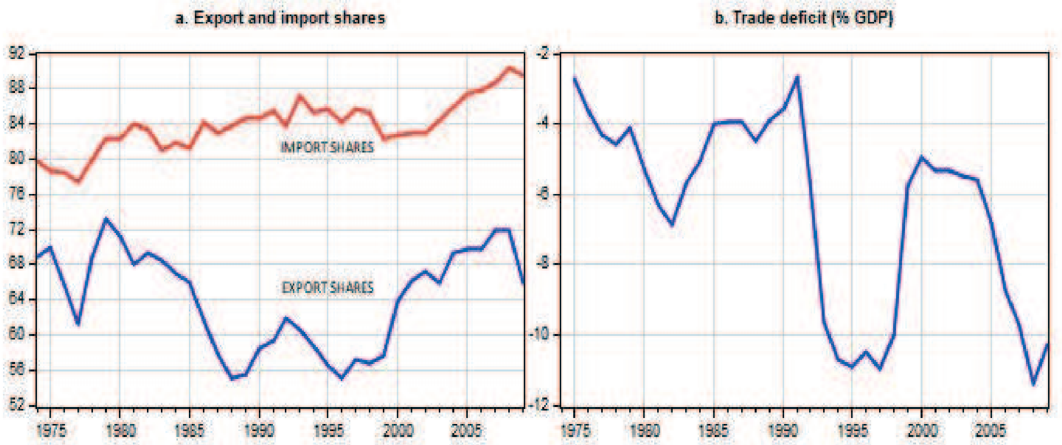
Table 1.6 informs on the homogeneity of the opening process across sectors. Within this relatively homogeneous process, we acknowledge that some branches experienced particularly intensive opening processes. Two of them were already globalized in 1974-1991 (S15 on manufacture of machinery and equipment including manufacture of offi ce, accounting and computing machinery, 232%; and S17 on manufacture of medical, precision and optical instruments, watches and clocks, 156%), but reached degrees of trade openness of 387% and 226%, respectively, in 1992-2009. The three other ones departed from values much below, but also became fully opened industries in 1992-2009: S16 on manufacture of electrical machinery, radio, television and communication equipment, 256%; S18 on manufacture of motor vehicles, trailers and semi-trailers, and other transport equipment, 151%; and S19 on manufacture of furniture, 125%. This particularly intensive transformations are taken into account in our empirical analysis.

Table 1.6: Trade openness and output shares by industries.

Sector	**Trade openness**		**Share in total industry GDP**	
	1974-1991	1992-2009	1974-1991	1992-2009
S2	5.1	29.2	2.6	0.7
S11	8.3	42.3	4.2	5.0
S1	9.7	28.2	28.4	28.9
S3	11.6	68.8	10.9	5.3
S12	13.3	22.1	5.8	7.3
S6	19.2	36.8	0.8	0.6
S7	23.1	39.5	3.7	3.8
S5	23.4	69.6	1.7	1.1
S14	24.5	60.3	3.7	2.8
S8	25.7	23.7	2.8	3.6
S4	29.2	39.2	3.0	3.6
S19	39.9	125	1.6	1.2
S10	47.6	79.8	12.8	15.4
S9	52.2	44.5	4.3	9.0
S18	59.7	151	4.1	3.2
S16	64.2	256	3.2	2.2
S13	71.9	111	4.0	3.9
S17	156	226	0.5	0.7
S15	232	387	1.9	1.7
Total	48.2	96.8	100	100

Notes: Classification based on International Standard Industrial Classification, Rev. 3. All variables are expressed in percent. Sectors: Manufacture of food products and beverages (S1); Manufacture of tobacco products (S2); Manufacture of textiles (S3); Manufacture of wearing apparel; dressing and dyeing of fur (S4); Tanning and dressing of leather; manufacture of luggage, handbags, saddlery, harness and footwear (S5); Manufacture of wood and of products of wood and cork, except furniture; manufacture of articles of straw and plaiting materials (S6); Manufacture of paper and paper products (S7); Publishing, printing and reproduction of recorded media (S8); Manufacture of coke, refined petroleum products and nuclear fuel (S9); Manufacture of chemicals and chemical products (S10); Manufacture of rubber and plastics products (S11); Manufacture of other non-metallic mineral products (S12); Manufacture of basic metals (S13); Manufacture of fabricated metal products, except machinery and equipment (S14); Manufacture of machinery and equipment n.e.c.; and manufacture of office, accounting and computing machinery (S15); Manufacture of electrical machinery and apparatus n.e.c.; and manufacture of radio, television and communication equipment and apparatus (S16); Manufacture of medical, precision and optical instruments, watches and clocks (S17); Manufacture of motor vehicles, trailers and semi-trailers, and other transport equipment (S18); Manufacture of furniture; manufacturing n.e.c. (S19).

Figure 1-6: Export and import shares, and trade deficit of the manufacturing industry, 1975-2010.



Source: DANE.

Table 1.7: Bias-corrected LSDVC estimators.

Equation (1.2)

	MODEL I			Equation (1.3)		
Bias order	AH	AB	BB	AH	AB	BB
O $(1/T)$	0.42	0.45	0.46	0.57	0.57	0.60
O $(1/NT)$	0.42	0.46	0.47	0.58	0.59	0.61
O $(1/N^{-1}T^{-2})$	0.42	0.47	0.52	0.58	0.60	0.69

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

Table 1.8: Multiplicative interaction models.

Dependent variable: w_{it}

	Equation (1.2)				Equation (1.3)			
	OLS	FE	FE-2SLS	FE-2SLS	OLS	FE	FE-2SLS	FE-2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
W_{it-1}	0.88 [0.000]	0.53 [0.000]	0.54 [0.000]	0.54 [0.000]	0.87 [0.000]	0.59 [0.000]	0.61 [0.000]	0.58 [0.000]
Y_{it}/N_{it}	0.05 [0.000]	0.02 [0.456]	−0.01 [0.312]	−0.01 [0.350]	0.05 [0.000]	0.03 [0.310]	0.01 [0.647]	0.00 [0.904]
u_t	−0.12 [0.000]	−0.10 [0.000]	−0.10 [0.000]	−0.11 [0.000]	0.00 [0.753]	0.05 [0.007]	0.11 [0.000]	0.21 [0.000]
W_t^{min}	0.21 [0.447]	0.50 [0.017]	0.50 [0.002]	0.50 [0.001]				
π_t					−2.69 [0.019]	−3.09 [0.004]	−18.96 [0.000]	−32.79 [0.002]
τ_t^p	−0.50 [0.492]	−0.60 [0.504]	−0.65 [0.484]	−0.68 [0.457]	−0.67 [0.000]	−0.50 [0.000]	−1.04 [0.000]	−1.55 [0.001]
op_{it}	0.06 [0.857]	0.39 [0.231]	0.40 [0.441]	0.41 [0.425]	0.00 [0.530]	−0.02 [0.165]	−0.01 [0.542]	0.00 [0.944]
$W_t^{min} * \tau_t^p$	0.05 [0.573]	0.08 [0.439]	0.09 [0.374]	0.09 [0.352]				
$W_t^{min} * op_{it}$	−0.01 [0.916]	−0.06 [0.257]	−0.06 [0.470]	−0.06 [0.450]				
$W_t^{min} * \tau_t^p * op_{it}$	−0.00 [0.933]	0.00 [0.551]	0.00 [0.721]	0.00 [0.687]				
$\pi_t * \tau_t^p$					0.47 [0.246]	0.53 [0.177]	5.96 [0.000]	10.68 [0.003]
$\pi_t * op_{it}$					0.01 [0.969]	0.02 [0.888]	3.13 [0.001]	5.80 [0.007]
$\pi_t * \tau_t^p * op_{it}$					−0.00 [0.650]	0.00 [0.970]	−1.13 [0.002]	−2.10 [0.002]
d_{8083}	−0.02 [0.137]	−0.04 [0.000]	−0.04 [0.000]	−0.04 [0.000]	−0.02 [0.071]	−0.02 [0.007]	−0.00 [0.780]	0.01 [0.515]
d_{9700}	0.03 [0.002]	0.03 [0.000]	0.02 [0.000]	0.02 [0.000]	−0.02 [0.001]	−0.04 [0.000]	−0.05 [0.000]	−0.07 [0.000]
d_{0809}	−0.06 [0.001]	−0.05 [0.016]	−0.05 [0.010]	−0.05 [0.005]	−0.04 [0.014]	−0.01 [0.769]	−0.01 [0.689]	−0.00 [0.996]
c	−0.36 [0.857]	−0.30 [0.820]			2.98 [0.000]	5.03 [0.000]		
$Obvs.$	646	646	627	627	646	646	627	627
$Adj.R^2$	0.97	0.95	0.94	0.94	0.97	0.96	0.93	0.87
U			8.29 [0.016]	16.10 [0.007]			8.22 [0.016]	7.01 [0.030]
H			1.91 [0.167]	9.76 [0.045]			0.29 [0.589]	0.00 [0.968]
W			407.8	160.4			2.78	2.04

Notes: All variables are expressed in logs. P-values in brackets. OLS: Ordinary Least Squares. FE: Fixed Effects. FE-2SLS: Fixed effects using Two Step Least Squares.
U : Under identification test. H : Hansen test. W : Weak identification test.

Chapter 2

Wage rigidities in Colombia: Measurement, causes, and policy implications

Abstract[1]

This paper evaluates the extent of wage rigidities in Colombia over a period, 2002-2014, in which the fall in unemployment was relatively slow with respect to sustained economic growth. Following Holden and Wulfsberg (2009), we compute a measure of downward real wage rigidity (DRWR) of 12.09%, four times bigger than their aggregate estimate for the OECD economies. Moreover, in contrast to the evidence for the advanced economies, the determinants of such rigidities show no connection to the wage bargaining system. Amid the absence of effective labour market institutions to make rigidities less prevalent, economic growth appears as the most powerful mechanism to ward them off. Under this light, we provide a stylized description of the wage setting rule in Colombia, compare it with the common one in the advanced economies, and call for a far-reaching reform of the Colombian wage bargaining setup.

JEL Classification: E24, J3, J48, J58.

Keywords: Downward Real Wage Rigidity, Wage bargaining, Minimum Wage, Informality, Unemployment.

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2.1 Introduction

Colombia is one of the successful Latin American economies. Its growth rate since 2000 has evolved around 5% on average, while inflation has been consistently reduced to stabilize at 3%. It is in this context that unemployment persists and remains stubbornly high at around 10%.

Why this persistence? Although common wisdom has implicitly assumed that wages are rigid, no empirical evidence is provided in the literature. This paper tries to fill this void by checking the extent to which wage rigidities are relevant in Colombia.

We contribute to the literature in two main respects. First, we apply to a very different case a methodology normally used to examine the labour market of advanced economies. Indeed, the wage setting environment in Colombia differs significantly from the one in most developed countries (nonexistent unemployment benefits, no modern welfare state, weak trade unions, and a large informal or underground economy). As a consequence, although we use a standard technique, its empirical implementation has had to be adapted to take into account both these features and the associated data constraints.

The second main contribution is that not only we estimate the extent of downward real wage rigidities (DRWR, henceforth), but we also study its causes and read them in terms of the ensuing crucial policy implications. This is relevant because it connects our labour market analysis to monetary policy issues. Common wisdom states that wage rigidities become progressively important in slumps (this dates back to Tobin (1972) who was the first one to claim that inflation is helpful to prevent negative effects of wages rigidities on unemployment). However, in the context of a developing economy, recent efforts to keep inflation under control, and to try to converge to the low values of the developed economies, have implied a sustained effort to avoid price increases.[2] This means that inflation cannot be seen as a conjunctural phenomenon (mainly hurtful during slumps), but rather as a

[2] Inflation targeting is one mechanism by which developing economies, Colombia among them, have enforced a downward trend in inflation. Note that the literature dealing with this important matter is beyond the focus of this paper and it is thus not covered here.

structural one. It is thus pertinent to enquire which kind of role does inflation play in Colombia as a determinant of DRWR, and to study how its trajectory has affected the extent of wage rigidities and, hence, unemployment. Furthermore, investigating the determinants of such rigidities will allow us to suggest policy recommendations that otherwise (i.e., in case of just quantifying DRWR) would not be possible to propose.

To conduct the analysis, we follow Holden and Wulfsberg (2009) and explore the potential existence of DRWR using aggregate data for the Colombian labour market. To be precise, we provide an sectorial-level analysis based on data from 59 sectors and 13 metropolitan areas over the period 2002-2014. This large panel data (large, at least, for the standards of an emerging economy) is used to compute and compare the observed real wage-change distributions with constructed counterfactual rigid-free distributions (which are called notional distributions). The difference between the two is used to quantify deficits in real wage cuts, which are then interpreted as reflecting the extent of wage rigidities in the labour market.

On this respect, the deficit of real wage cuts in Colombia is estimated at 12.09%, indicating that about 12 out of 100 notional real wage cuts do not result in actually observed wage cuts. In addition, a salient feature of this aggregate estimate is the declining path it has followed. We document a value of DRWR above 20% in 2002, a sharp decline in subsequent years, and a stabilization around 10% in recent years.

The average of 12% is relatively high, when compared to values of 3.7% found for major European areas in Holden and Wulfsberg (2009); relatively low with respect to the average of 15% found by Holden and Messina (2012, as cited in World Bank, 2012) for Latin America and the Caribbean countries; and similar to the case of Uruguay as reported in Messina and Sanz-de-Galdeano (2014).

Regarding the determinants of DRWR, we use the growth rate of GDP instead of the unemployment rate, which is the standard variable in the literature. Data on the later is highly imperfect, while data on the former is trustful. Given than Okun's Law connects the two, we work with economic growth and find it to be the most influential driver. As a consequence, the best mechanism to avoid binding wage rigidities is growth. The bad news here is that real economic growth in Colombia comes in parallel with inflation; in turn, the good news is that inflation has been on a steady downward trend in recent years.

Hence, a salient characteristic of our results is that the falling path of DRWR has taken place simultaneously to that of price inflation. Inflation targeting started to be implemented in 2000, and has certainly been crucial to explain the falling trend in price inflation. In turn, although reductions in DRWR are usually associated with situations in which inflation grows (for example by allowing firms to scape nominal wage cuts thanks to the inflationary scenario), this is not the case in Colombia.

This specific feature of the Colombian economy, which we discuss at length in Section 2.4, gives rise to a positive influence of inflation on wage rigidities. Moreover, because of the fact that real wage growth is based on observed (and not expected) inflation, we show that this positive influence is similar, but with the opposite sign, to the one exerted by real minimum wages. That is, as we document below, the more real minimum wages grow, the smaller wage rigidities are.

Beyond minimum wages, the impact of two other labour market institutions is also assessed. One is union power, which is small due to little trade union density and low coverage of collective agreements. It is thus unsurprising that we find trade union density irrelevant to explain wage rigidities. The second one is the degree of informality (as a share of total employment), which appears as another important mechanism to alleviate wage rigidities. The problem with this mechanism is that it goes against the pursuit of better quality jobs.

Thus, inflation being under control, union power largely irrelevant with regard to wage setting, and informality essentially unwanted, higher growth (or lower unemployment) is the only way out to prevent wage rigidities to be binding. Unless, of course, that the wage setting system is reformed so that wages become attached to productivity and start being fixed over expected prices. This is the main policy implication of our analysis, and the object of a detailed discussion in the last part of the article.

The rest of the paper is structured as follows. Section 2.2 measures the extent of downward real wage rigidity in Colombia. Empirical methods to evaluate its causes are discussed in Section 2.3. Section 2.4 presents the results and deals with some of the implications of such rigidities in a quite particular wage setting environment such as the Colombian one. Section 2.5 concludes.

2.2 Real wage rigidity in Colombia

We follow the empirical approach of Holden and Wulfsberg (2009), whose use has become popular for aggregate data analysis. Following their method, we explore the potential existence of DRWR by comparing observed real-wage-change distributions with constructed counterfactual or notional distributions in which wages are assumed to be flexible (i.e. free of wage rigidity). As explained below in detail, the difference between the two is a measure of the extent of real wage rigidities.

Since we work with macro data, our observation unit is the change in average hourly earnings of all workers.[3] This makes our data liable to compositional changes, for example because new workers' wages differ from the wages of those who retire (or quit), or because the share of low-skilled workers tends to decrease during recessions. This has two consequences.

First, the existence of spikes in the wage distribution is not indicative of DRWR, in contrast to the standard analysis based on microeconomic evidence.[4] Second, even if there is significant DRWR at the individual level, average wage growth at the industry level might be negative if those retiring (or quitting) have higher wages than the stayers.

This entails that any statistical evidence on DRWR implies that microdata wage rigidities are not fully offset by compositional and related changes. The existence of rigidities at

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[3] And not that of job stayers, as it is common in microeconomic studies such as Nickell and Quintini (2003), Bauer *et al.* (2007), Elsby (2009), or Stüber and Beissinger (2012).

[4] One implication of these differences, is that prevalent DRWR among individual workers, apparent in micro data by a spike in the wage change distribution at zero, would not imply that the change in average wage is zero. Hence, we cannot test for DRWR by looking for spikes and (as explained in Holden and Wulfsberg, 2009) we rather have to look for deficits of wage cuts.

the macro level is the object of our analysis.

2.2.1 Data

We use an unbalanced panel of sector level data for the annual percentage growth of hourly wages. This panel covers information for all workers, the period 2002-2014, 13 metropolitan areas, and 59 sectors classified according to the two-digits ISIC classification. Thus, we work with 9,156 observations distributed across 767 sector-year samples.

The areas considered are: (1) Bogotá; (2) Medellín and Valle de Aburrá; (3) Cali and Yumbo; (4) Barranquilla and Soledad; (5) Cartagena; (6) Manizales and Villamaría; (7) Monteria; (8) Villavisencio; (9) Pasto; (10) Cúcuta, Villa del Rosario, Los Patios and El Zulia; (11) Pereira, Dosquebradas and La Virginia; (12) Bucaramanga, Florida Blanca, Giron and Piedecuesta; and (13) Ibague.

The 59 sectors are distributed across 9 aggregate sectors: (1) Agriculture, Cattle Ranch, Forestry, Hunting and Fishing; (2) Mine and Quarry Exploitation; (3) Manufacturing Industry; (4) Electricity, Gas and Water; (5) Construction; (6) Commerce, Repairing, Restaurants and Hotels; (7) Transport, Storage and Communication; (8) Financial and Insurance; and (9) Social, Communal and Personal Services.

The data is taken from the Continuous Household Survey (Encuesta Continua de Hogares, ECH) for the period 2001-2006, and from the Household Integrated Survey (Gran Encuesta Integrada de Hogares, GEIH), which replaced the former for 2007-2014.

This information is used to construct sector-year specific estimates of DRWR by looking for a deficit of wage cuts in the observed wage changes. The notional distributions are derived from sector-year samples with high median nominal and real wage growth rates, which are assumed to be unaffected by rigidities. We assume that, in the absence of any rigidity, the notional real wage growth in area j, sector i, and year t is stochastic with an unknown distribution G, which is parameterized by the median real wage growth μ_{it}^{N}, and dispersion σ_{it}^{N}: $G(\mu_{it}^{N}, \sigma_{it}^{N})$.

This implies that we allow the location and dispersion of the notional (or feee of rigidities) wage growth to vary across sectors and years. We could have chosen to allow this variation across metropolitan areas (instead of sectors). However, as shown in Appendix 2, Table 2.6, variations in wage setting and other relevant variables are larger across sectors than across areas. This justifies our choice.

2.2.2 Notional distributions

The notional distributions are constructed in two steps.

In the first step, we compute an underlying distribution based on a subset H (where H denotes high) of the sample S, with $S^H = 592$ observations from 44 sector-year samples. These observations are selected on the basis that both the median nominal wage growth and the median real wage growth in the sector-year are in their respective upper quartiles over all sector-years. Specifically, the median nominal wage growth is above 20.39% (which is the value at which the 4th quartile starts), and the median real wage growth is above 15.71% (which is the value at which the 4th quartile starts for the real case). The selection of such subsample aims at ensuring that the constructed notional distributions are free of wage rigidities (as shown by Holden and Wulfsberg, 2009, the higher the growth rates, the lower the probability of wage rigidities is).

The underlying distribution is constructed by normalizing these 592 empirical observations as follows:

$$x_s \equiv \frac{(\Delta w_{jit} - \mu_{it})}{(P75_{it} - P35_{it})}, \quad \forall j, i, t \in H \text{ } and \text{ } s = 1, ..., S^H, \tag{2.1}$$

where Δw_{jit} is the observed wage growth in area j, sector i, and year t; μ_{it} is the corresponding observed sector-year specific median; and $(P75_{it} - P35_{it})$ is the inter-percentile range between the 75th and 35th percentiles, which is used as a measure of dispersion. Subscript s runs over j, i and t in the 44 samples. The resulting calculated x_s should thus be thought as an observation of the stochastic variable X from the underlying distribution $X \sim G(0, 1)$.

Figure 2.1 compares the underlying distribution of wages (solid line) with the standard normal distribution (dashed line). Note that the underlying distribution is slightly skewed left, with a skewness coefficient of -0.64, and has a larger peak than the normal distribution, with a kurtosis of 4.9.[5]

In the second step, for each of the 767 sector-years in the full sample, we compute the sector-year specific distributions of notional wage changes. This is done by adjusting as follows the underlying wage change distribution for the sector-specific observed median and inter-percentile range:

$$Z_{it} \equiv X(P75_{it} - P35_{it}) + \mu_{it}, \quad \forall i, t, \tag{2.2}$$

where Z_{it} is a set of 767 notional sector-year distributions $Z_{it} \sim G(\mu_{it}, P75_{it} - P35_{it})$, each consisting of $S^H = 592$ wage-change observations. By construction, these notional distributions have the same shapes across all notional sector-years. However, their median and inter-percentile range are the same as their empirical sector-year counterparts.

Once provided with the sector-year specific distributions, we explore the extent of DRWR by comparing the corresponding notional and empirical distributions. In particular, we compare the incidence rate of notional wage cuts $\widetilde{q}(k)_{it}$ below a specific floor k (this is the counterfactual), with the corresponding empirical incidence rate, $q(k)_{it}$.

For all sector-year samples, the notional incidence rate is defined as:

$$\widetilde{q}(k)_{it} \equiv \frac{\# Z^s_{it} < k}{S^H}, \quad s = 1, ..., S^H, \tag{2.3}$$

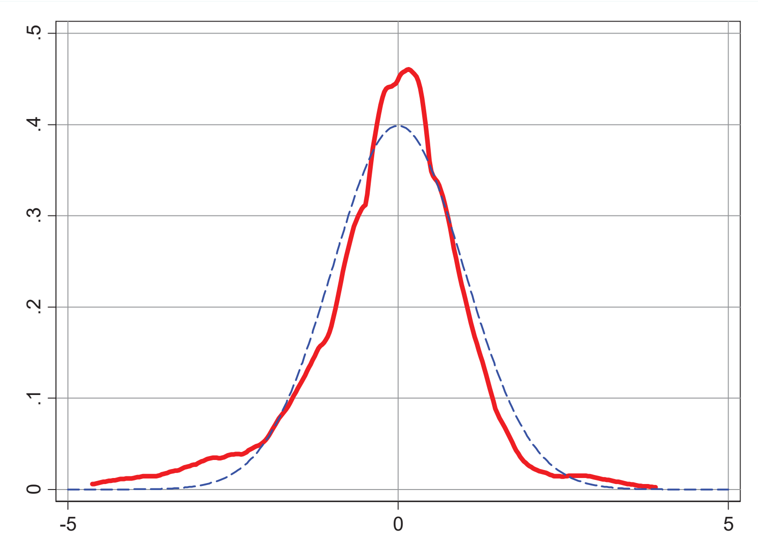
where $\# Z^s_{it} < k$ is the number of notional wage cuts below k, and S^H is the number of observations in the underlying distribution. Likewise, the empirical incidence rate is

$$q(k)_{it} \equiv \frac{\# \Delta w_{jit} < k}{S_{it}}, \quad \forall j, \tag{2.4}$$

where $\# \Delta w_{it} < k$ is the number of observed wage cuts below k, and S_{it} is the number of observed areas in sector-year it.

[5] A normal distribution would have a skewness of 0 and a kurtosis of 3, therefore our distribution is leptokurtic.

Figure 2-1: Underlying distribution.



Notes: Kernel density (solid line) of the normalized underlying distribution of wage changes compared to the normal density (dashed line). The ten extreme observations are omitted.

The deficit of observed wage cuts below floor k relative to the notional distribution is called the fraction of wage cuts prevented, $FWCP(k)$, and constitutes the measure of DRWR. This is calculated as:

$$FWCP(k)_{it}= 1 - \frac{q(k)_{it}}{\widetilde{q}(k)_{it}}. \qquad (2.5)$$

The estimates of $FWCP$ in single sector-years will be imprecise because of the small number of metropolitan areas in each sector-year sample (13 or less). Thus, we present estimates of the $FWCP$ for aggregate industries, aggregate periods, and the full sample, which are likely to be more precise.[6]

Finally, we use the simulation method of Holden and Wulfsberg (2008) to test for the statistical significance of our DRWR estimates. This method is considered very powerful for detecting statistical differences between the observed and notional (or free of rigidities) wage change distributions. The null hypothesis is that there is no DRWR, which means that observed and notional wage changes have the same distribution, while the alternative hypothesis is that the observed number of real wage cuts is smaller than the expected one from the notional distributions. Thus, if the null hypothesis is rejected, we conclude that wages are rigid downwards. In this context, the simulation method of Holden and Wulfsberg (2008) allows us to compute the respective $p-value$ associated with the null hypothesis of no DRWR.

2.2.3 Estimates

Table 2.1 presents estimates of DRWR computed when the real wage change is below zero, $k = 0$. The following information is reported: the notional incidence rate, $\widetilde{q}$ (that is, the percent of wage cuts that would take place in the absence of rigidities); the empirical incidence rate, q; the estimated fraction of wage changes prevented, $FWCP$ (which measures the extent of DRWR); the $p-values$ (which are used to test whether our estimates of $FWCP$ are statistically significant); and the number of observations, S.

[6] The sector-year estimates of $FWCP(k = 0)$ for the 9 sectors are reported in Appendix 2, Table 2.6. The number of observations in each sector-year sample is 13 or more.

For the full sample (last row in Table 2.1), we provide three important measurements. First, the percent of real wage cuts that would take place in a free wage rigidity system would be 49.60%. Second, 43.60% of the observed wage changes actually correspond to real wage cuts. As a consequence, we obtain a measurement of the $FWCP$ which is 12.09%. This means that about 12 out of 100 notional real wage cuts in the overall sample do not result in observed wage cuts due to DRWR.

At the sectorial level, differences are large. On the one hand, DRWR is not significant in three economic activities such as electricity, gas and water (S4); construction (S5); and commerce, repairing, restaurants, and hotels (S6). For S5 and S6, the absence of rigidities (or high wage flexibility) may be associated to their high degrees of informality (73.54% and 72.41%, respectively) and low levels of union density (0.72% and 1.70%, as reported in Table 2.3). This is in contrast to S4, which is characterized by the lowest level of informality (8.54%), and the highest level of union density (29.20%). However, S4 is also and by far the most productive sector and, as such, it is the one where wage rigidities are less likely to be binding. On the other hand, the estimated $FWCP$ is the highest in the agriculture and mining sectors (S1 and S2), with 32.10% and 40.43% respectively. Coincidentally, S2 shows a low share of informal employment and union density.

Regarding the evolution of DRWR over time, it has followed a downward path departing from above 20% in 2002, quickly falling to 7.51% in 2005, and subsequently stabilizing around 10%. This general evolution is marked by several peaks in 2002, 2007 and 2010 which, remarkably, coincide with low inflation periods in Colombia (2002 and 2010 on account of the effects of the financial crises previously experienced). It could thus be tempting to associate the existence of wage rigidities to periods of low inflation. Next we deal with the drivers of this evolution.

Table 2.1: Estimates of DRWR for k=0.

	$\widetilde{q}$	q	$FWCP$	$p-value$	S
Sectors					
S1	51.18	34.75	32.10	0.00	446
S2	52.04	31.00	40.43	0.00	500
S3	49.15	43.08	12.34	0.00	3498
S4	48.76	48.22	1.10	0.44	338
S5	51.72	49.11	5.03	0.27	169
S6	44.61	44.67	-0.15	0.53	676
S7	49.31	44.06	10.65	0.00	758
S8	50.60	44.01	13.02	0.00	668
S9	50.59	47.50	6.11	0.00	2103
Years					
2002	48.68	37.73	22.50	0.00	721
2003	58.43	50.71	13.21	0.00	704
2004	49.46	44.77	9.47	0.01	708
2005	45.32	40.54	10.55	0.01	708
2006	52.95	48.97	7.51	0.02	680
2007	38.87	31.90	17.93	0.00	699
2008	51.20	45.99	10.18	0.00	711
2009	56.28	50.93	9.52	0.00	701
2010	50.21	41.74	16.88	0.00	702
2011	49.63	44.64	10.06	0.00	699
2012	53.01	48.18	9.10	0.01	716
2013	49.83	44.16	11.37	0.00	702
2014	40.96	36.74	10.31	0.01	705
All observations	49.60	43.60	12.09	0.00	9,156

Notes: Data in percent. (S1) Agriculture, Cattle Ranch, Forestry, Hunting and Fishing; (S2) Mine and Quarry Exploitation; (S3) Manufacturing Industry; (S4) Electricity, Gas and Water; (S5) Construction; (S6) Commerce, Repairing, Restaurants and Hotels; (S7) Transport, Storage and Communication; (S8) Financial and Insurance; (S9) Social, Communal and Personal Services.

2.3 The determinants of DRWR

2.3.1 Empirical implementation

The theoretical model of Holden and Wulfsberg (2009) points out that the prevalence of DRWR can be explained by differences in economic and institutional variables. Economic variables such as the unemployment rate (u) and price inflation ($\triangle p$); and institutional variables such as the strictness of employment protection legislation (EPL), and union bargaining strength (θ) in whatever form (e.g. measures of trade union density, number of strikes, or days lost due to labour conflicts). In this way, a benchmark empirical version of their model may be expressed through the following general form:[7]

$$FWCR_{it} = f \ (\ u_{it}, \ \Delta p_{it}, \ EPL_{it}, \ \theta_{it} \ ) \, . \tag{2.6}$$

Previous empirical evidence on developed countries confirms the relevance of these explanatory variables (for example, Dias *et al.*, 2015; Druant *et al.*, 2012; and Holden and Wulfsberg, 2008, 2009 and 2014). However, in view of the different wage setting environment we are dealing with (weak union power, large informality, and a minimum wage system which acts as a crucial driver of the wage setting process), and also because data on some specific variables have limited availability (for example, unemployment, and EPL), some adjustments are needed for the case of Colombia.

The first adjustment on benchmark Equation (2.6) relates to unemployment. As no series are available for sectorial unemployment, one possibility is to proxy them by the number of unemployed in sector i (taking as reference the sector of the last job) as a ratio over the number of employees in that sector. The problem with this measure is that it does not correspond to the unemployment rate definition, and may not fit the actual dynamics of sectorial unemployment.[8] This is the reason why we rather use sectorial real economic

[7] Potential no-linearities may also be considered. However, Holden and Wulsfberg (2014) find they are not significant and stick to a linear representation. Since our attempts to include non-linearities were also unsuccessful, we follow the route of Holden and Wulsfberg (2009).

[8] We estimate regressions using this proxy variable, which turns out to be not significant.

growth (ΔY), which is fully available and, given Okun's Law, acts as a counterpart of the unemployment rate and should allow us to capture the effect of labour market tightness on wage rigidities. We thus expect a negative sign on the coefficient of ΔY.

The second adjustment relates to the absence of EPL simply due to the lack of data. The counterpart here is the inclusion of labour informality, as it is an important mechanism by which firms achieve flexibility. Although consideration of informality is a must in Colombia, given its relevance, the connection it has with DRWR has already been treated in the literature. For example, Ahmed *et al.* (2014) and Batini *et al.* (2010) highlight that wages in informal sectors tend to be flexible or, at least, tend to be significantly more flexible that those in formal sectors. In a macroeconomic context such as ours, in which the average wage growth is the observation unit, this flexibility should contribute to reduce the extent of DRWR both in the formal and informal sectors (the reason being that this extra flexibility is actually at the disposal of all industries, no matter their actual use of informal work).

Given these changes, benchmark Equation (2.6) becomes:

$$FWCR_{it} = f\ (\ \Delta Y_{it},\ \Delta p_{it},\ \iota_{it},\ \theta_{it}\ )\,, \tag{2.7}$$

where ι denotes the degree of informality defined as the proportion of informal workers over total workers.

The specificities of the wage setting system in Colombia (see Agudelo and Sala, 2016, and the discussion of our results below) lead us to consider two alternative specifications where price inflation is substituted, respectively, by the real minimum wage growth ($\Delta w^{\min}$) and growth in ratio between producer and consumer prices ($\Delta\pi$):

$$FWCR_{it} = f\ \left(\ \Delta Y_{it},\ \Delta w_{it}^{\min},\ \iota_{it},\ \theta_{it}\ \right), \tag{2.8}$$

$$FWCR_{it} = f\ (\ \Delta Y_{it},\ \Delta\pi_{it},\ \iota_{it},\ \theta_{it}\ ) \tag{2.9}$$

In specification (2.8), ΔP_{it} is substituted by $\Delta w_{it}^{\min}$ because of the intimate relationship between both variables. Wage indexation was set by law in 1999,[9] and made the minimum wage operate as a sort of reservation or floor wage irrespective of the economic sector (see Hofstetter, 2005, for details). As shown by Iregui *et al.* (2012), most firms adjust nominal wages annually at rates that are roughly equivalent to the observed (no the expected) rate of CPI inflation. This is why price inflation and the minimum wage growth appear as substitutable variables in determining DRWR: they both carry the effect of past prices on current outcomes, and hence specification (2.8).

This substitutability can be pushed forward by looking at it from another side. The fact that the nominal minimum wage is deflated by the GDP deflator implies that the growth rate of the real minimum wage essentially captures the changing wedge between consumption and production prices (this is so to the extent that nominal wage growth reflects CPI inflation). Hence specification (2.9), where the change in the ratio between the prices faced by producers (at the sectorial level) and those faced by consumers ($\Delta \pi_{it}$) substitutes $\Delta w_{it}^{\min}$.

Although similar alternative specifications were used in Agudelo and Sala (2016) to explain the wage setting mechanism in the Colombian industry, it is the empirical analysis which will ultimately endorse or reject them as plausible alternatives to explain wage rigidities.

2.3.2 Data

The database used to compute the $FWCP$, we now add a set of potential explanatory variables also having a cross-section dimension of $N = 9$ sectors, and a time dimension of $T = 13$ years covering the period 2002-2014. Table 2.2 lists the variables and the corresponding sources.

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[9] Judgment C-815 of the Constitution Court passed in 1999 to avoid further loss in the purchasing power of the minimum wage, which had deteriorated in the 1980s and 1990s.

Table 2.2: Definitions of variables.

Variables		Sources	Subindices
$FWCP_{it}$	Fraction of wage cuts prevented	(1)	$i = 1, ..., 9$ sectors
Y_{it}	Real GDP	(2)	$t = 1, ..., 13$ years
ΔY_{it}	Real GDP growth	(2)	
Sh_{it}	Sector's value added as % of GDP	(2)	
p_{it}	GDP deflator	(2)	
Δp_{it}	Sectorial inflation	(2)	
θ_{it}	Trade union density	(3-4)	
N_{it}	Total employment	(5)	
N_{it}^{in}	Informal employment	(5)	
ι_{it}	Informal employment share $\left[\frac{N_{it}^{in}}{N_{it}}\right]$	(5)	
$w_{it}^{\min}$	Real minimum wage	(2)	
p_{it}^{c}	CPI Index	(4)	
π_{it}	Relative prices $\left[\frac{p_{it}}{p_{it}^{c}}\right]$	(2-4)	

Notes: All nominal variables are deflated by the GDP deflator (base: December 2005). (1) Own estimates; (2) Banco de la República; (3) ENS; (4) DANE; (5) Own calculations based on ECH-GEIH.

Information on output per sector, total output and the corresponding deflators is taken from the Colombian Central Bank (Banco de la República). This allows us to compute real GDP (Y), the share of each sector's value added over total output (sh), and sectorial inflation (Δp). The same source is used to collect data on nominal minimum wages, which are also deflated by the GDP deflator (in each sector) to obtain the real minimum wage ($w^{\min}$).

The trade union density rate (θ) is computed as the number of unionized workers reported by the national trade union institution (Escuela Sindical Nacional, ENS) over total employment reported by the National Administrative Department of Statistics (Departamento Administrativo Nacional de Estadística, DANE).

Following the definition of the DANE and the International labour Organization (OIT), we consider informal employees those that are not affiliated to the social security system. Then, to compute the share of informal employment (ι) we use micro data from the Continuous Household Survey (Encuesta Continua de Hogares, ECH) for 2002-2006, and the Household Integrated Survey (Gran Encuesta Integrada de Hogares, GEIH) for 2007-2014.

Finally, the CPI index (obtained from the DANE) is used to compute the ratio of relative prices between sectorial and consumption prices (p^c).

Descriptive information on these variables at the sectorial level is provided in Table 2.3.

As noted when first presented in Table 2.1, wage rigidities across sectors are heterogeneous. They may be as high as 30% and 40% as in S1 and S2; but they may also be non-significant as in S4, S5 and S6, were wages are virtually flexible. This diversity leads us to use weighting factors in the estimation of equations (2.7) to (2.9).

Note, however, that in all sectors but S1 and S2, DRWR is below 13.5%. Given the low weight of these sectors (with a joint share of GDP of 15.28% on average during our sample period), DRWR at the aggregate level is 12.09% reflecting the general low sectorial values.

Looking at Table 2.3, this relatively low incidence of wage rigidities goes in parallel with a sustained period of quick real economic growth (with rates above 4% in most sectors which have coexisted with similar, or even larger, rates of price inflation); with a high degree of

Table 2.3: Macro developments in the Colombian sectors, 2002-2014.

Sector	$FWCP_i$	ΔY_i	Δp_i	θ_i	ι_i	$\Delta w_i^{\min}$	$\Delta \pi_i$	Sh_i
S1	31.39	2.28	3.86	2.25	57.93	1.89	-0.34	7.81
S2	40.79	5.15	9.44	6.35	28.56	-3.69	5.24	7.47
S3	12.28	2.85	4.68	4.06	52.38	1.07	0.49	14.51
S4	4.36	3.23	5.39	29.20	8.54	0.36	1.19	4.22
S5	8.45	7.65	7.10	0.72	73.54	-1.35	2.89	6.53
S6	2.35	4.84	3.86	1.70	72.41	1.89	-0.34	13.12
S7	11.40	5.57	2.89	5.06	59.18	2.86	-1.30	7.77
S8	13.45	4.81	4.04	2.74	22.46	1.71	-1.59	21.29
S9	6.21	3.98	5.14	12.42	47.71	0.61	0.94	17.28
Total	12.09	4.57	4.98	4.83	55.81	0.92	0.75	100

Notes: Data in percent. Definitions on sectors (S1 to S9) provided in notes to Table 2.1. Information on $FWCP$ per sector is computed as the average of the estimates presented in Appendix 2, Table 2.6.

informality (which is widespread across sectors with the exception of S4); and with a low power of the trade unions (reflected in very low affiliation rates, again with the salient exception of S4).

We should also note that the real minimum wage and the price ratio grow at similar rates, 0.92% and 0.75% per year. In addition, when looking at the different industries, we see that most increases/reductions in real minimum wage go in parallel with reductions/increases in the ratio of prices. This shows that the adjustments in the real minimum wage are highly tied to the wedge between sectorial and consumption prices.

2.3.3 Econometric methodology

Just as initial benchmark, we start by estimating equations (2.7), (2.8) and (2.9) as linear regression models. The corresponding results are presented in Table 2.7 in Appendix 2. This is not, however, the appropriate estimation technique in the context of our analysis. The reason is that our dependent variable, $FWCP_{it}$, takes the form of count data as it is obtained by dividing the number of observed wage changes below floor k in each sector-year sample, $Y(k)_{it}$, over the number of simulated wage cuts in each sector-year sample $\widehat{Y}(k)_{it}$.

Since we are looking at the number of times an event is observed, we are actually using count data. In this context, the Poisson regression methodology is the appropriate one, because it allows to take into account the limited number of possible values taken by the response variable. This is shown in Holden and Wulfsberg (2014) by means of the following expression:

$$[1 - FWCP(k)] \; = \frac{Y(k)_{it}}{\widehat{Y}(k)_{it}} = e^{x'_{it}\beta + \varepsilon_{it}}, \quad if \;\; \widehat{Y}(k)_{it} > 0, \tag{2.10}$$

where $[1 - FWCP(k)]$ captures the fraction of real wage cuts realized, β is the parameter vector, x'_{it} is a set of candidate explanatory variables, and ε_{it} is an error term. The results for equations (2.7), (2.8) and (2.9) obtained from their estimation as Poisson regressions are presented in Appendix 2 for reference (in Table 2.8), but they are not yet our selected estimates.

The reason is that the Poisson distribution may be subject to an important limitation, called overdispersion, which takes place when the conditional variance is larger than the conditional mean. The existence of overdispersion violates the equidispersion property of a Poisson distribution, and is particularly likely to occur in cases of unobserved heterogeneity such as ours. In such context, a reasonable alternative is to use a Negative Binomial regression, which is a generalized version of the Poisson regression that allows the variance to differ from the mean.

To check for the presence of overdispersion in our Poisson regressions, we use the Log likelihood ratio test of alpha=0, where alpha represents a specific parameter indicating that the variance differs from the mean. If the test rejects the null hypothesis that the errors do not exhibit overdispersion (alpha=0), then the Poisson regression model is rejected in favor of the Negative Binomial regression model.

The results on this test for each equation (denoted as $LL-alpha$) are reported in Table 2.4. The null hypothesis that the errors do not exhibit overdispersion is clearly rejected in all cases, and lead us to focus on the results from the estimation of Negative Binomial regressions.[10] This estimation is conducted using Maximum Log-Likelihood methods.[11]

2.4 Results and discussion

2.4.1 Results

The results for the Negative Binomial estimates are presented in Table 2.4. Columns (1), (3) and (5) show the pooled regressions, while columns (2), (4), and (6) include fixed effects across sectors. All regressions are estimated using the share of each sector's value added over total output as weighting factors. The Z-test is computed with robust standard errors

[10] The low negative values of the Log-Likelihood (LL) in the linear and Poisson regressions with respect to those in the Negative Binomial regressions, further endorse the selection of the Negative Binomial estimates.

[11] The combination of Negative Binomial regressions and instrumental variable methods (such as GMM) has not yet been used in a context such as ours. Even more, although these methods are available, the marginal effects cannot be computed as we do in our analysis (see Table 2.5).

clustered by sectors. In addition, although it is not common in the literature to provide information on the average marginal effects, we believe they are important to give the estimates full economic content. Hence, the computed average marginal effects are reported in Table 2.5 (moreover, the average marginal effects obtained from the Poisson estimation are reported in Table 2.9 in the Appendix 2 for the sake of comparison).

Although all sets of estimates provide a similar picture, our reference estimates are those obtained through the fixed effects estimation.

Economic growth reduces DRWR as expected. We find a robust result across econometric specifications according to which a 1 percentage point increase in sectorial GDP growth generates a fall in DRWR equivalent to -0.61 percentage points. To have a sense of this effect, recall that the DRWR is on average around 12.0%; in such context, an increase in the rate of economic growth, say from 3% to 6%, would reduce DRWR from 12% to around 10.0%. This negative impact is consistent with the literature and the standard finding that the higher unemployment is, the more binding the wage rigidities are (recall that u is the counterpart of ΔY).

In turn, sectorial price inflation increases DRWR. A 1 percentage point increase in sectorial prices generates a rise in DRWR equivalent to 0.23 percentage points. Hence, the magnitude of this impact is around a third of the economic growth effect.

The sign of the coefficient is in contrast to the common finding for the developed economies. It is standard in the literature to find that the larger the inflation rate, the lower the DRWR is, on the grounds that quicker progress in prices tend to make labour cost adjustments less binding for firms. In contrast, the positive influence of inflation on DRWR in Colombia is the outcome of a very particular system of wage setting (examined, for example, in Agudelo and Sala, 2016; and Iregui *et al.*, 2012). This critical result is discussed at length in the next section.

We find trade union density to have no significant influence on DRWR. This result is in contrast to the standard positive coefficient found in developed economies, but it is not surprising given the low and falling trade union density (with rates between 4% and 5%,

Table 2.4: Estimates. Negative Binomial regressions.

Dependent variable: $FWCP_{it}$

	Equation (2.7)		Equation (2.8)		Equation (2.9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.06** [-2.29]	-0.05* [-1.73]	-0.06** [-2.35]	-0.05* [-1.78]	-0.06** [-2.34]	-0.05** [-1.71]
Δp_{it}	0.03** [2.23]	0.02** [2.05]				
θ_{it}	-0.05*** [-3.18]	0.06 [1.38]	-0.05*** [-3.10]	0.06 [1.47]	-0.05*** [-3.06]	0.06 [1.46]
ι_{it-1}	-0.02*** [-3.75]	-0.02*** [-2.91]	-0.02*** [-3.71]	-0.02*** [-2.83]	-0.02*** [-3.68]	-0.02*** [-2.79]
$\Delta w_{it}^{\min}$			-0.03** [-2.06]	-0.02* [-1.90]		
$\Delta \pi_{it}$					0.03** [2.01]	0.01 [1.35]
c	3.78*** [12.89]	4.38*** [11.61]	3.94*** [13.71]	4.46*** [11.21]	3.90*** [13.51]	4.43*** [11.20]
$Obvs.$	117	117	117	117	117	117
LL	-45.90	-42.44	-45.91	-42.45	-45.93	-42.46
$LL-alpha$	103.6 (0.00)	28.68 (0.00)	102.6 (0.00)	29.78 (0.00)	103.6 (0.00)	30.00 (0.00)

Notes: FE, sectorial fixed effects. *** Significant estimates at 1%; **, at 5%; *, at 10%.

Z-test in brackets; P-values in parentheses. LL, Log-Likelihood;

$LL-alpha$, Log-Likelihood ratio test of alpha=0.

as shown in Figure 2.7 in the Appendix 2). Beyond this low density, collective bargaining coverage is also very limited (6.2% of all employees in 2014).[12]

Further to violations of trade unions rights which disincentive affiliation (OECD, 2016), the low trade union density is to be associated, also, to a deep labour market segmentation between formal and informal workers.[13] This is what makes the share of informal employment ι a relevant variable.

Ahmed *et al.* (2014) and Batini *et al.* (2010) show that wages in informal sectors tend to be significantly more flexible that those in formal sectors. Our result for equation (2.7) aligns with these studies: we find that labour informality exerts a significant negative influence on DRWR which, in addition, is of similar magnitude than the one exerted by inflation.

It is also remarkable that the average marginal effect $\Delta w_{it}^{\min}$ in specification (2.8) is almost identical to the alternative one of Δp in specification (2.7). Of course with the opposite sign. To be precise, we find that a 1 percentage point increase in real minimum wages generates a fall in DRWR equivalent to 0.21 percentage points, just below (in absolute value) to the one of Δp (+0.23).

The fact that $\Delta w_{it}^{\min}$ has grown in recent years is the outcome of nominal minimum wages growing more than sectoral price inflation. However, the gap between these two

[12] One characteristic of the Colombian labour market is the absence of a framework allowing for sectorial bargaining. To sign collective agreements, the Colombian labor code allows centralized levels of trade union organization (taking the form of industrial unions, federations, and cofederations). Sectorial or regional bargaining hardly occurs in practice, as it is not regulated by the Ministry of Labor (for instance, the Labor code only supplies rules affecting company-level wage bargaining). For further details see OECD (2016).

[13] Informality in Colombia is much larger than in most other emerging economies (OECD, 2016). This massive informal sector is largely characterized by an easy entry, lack of stable employer-employee relationships, a small scale of operations, and skills gained outside the formal education system. Within this context, one recent and important development is the steady fall in the degree of informality, which has evolved from around 67% in 2002 to 51% in 2014. This is shown in Figure 2.7 in the Appendix 2.

Table 2.5: Average marginal effects. Negative Binomial regressions.

	Equation (2.7)		Equation (2.8)		Equation (2.9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.76** [−2.29]	-0.61* [−1.74]	-0.78** [−2.35]	-0.63* [−1.79]	-0.78** [−2.33]	-0.61** [−1.72]
Δp_{it}	0.37** [2.27]	0.23** [2.07]				
θ_{it}	-0.72*** [−2.90]	0.77 [1.38]	-0.71*** [−2.83]	0.83 [1.48]	-0.71*** [−2.79]	0.83 [1.46]
ι_{it-1}	-0.25*** [−3.31]	-0.23*** [−2.92]	-0.25*** [−3.28]	-0.22*** [−2.83]	-0.25*** [−3.26]	-0.22*** [−2.80]
$\Delta w_{it}^{\min}$			-0.35** [−2.09]	-0.21* [−1.93]		
$\Delta \pi_{it}$					0.35** [2.01]	0.19 [1.36]

Notes: FE, sectorial fixed effects. *** Significant estimates at 1%; **, at 5%; *, at 10%.

Z-test in brackets.

growth rates has tended to fall, thereby causing a deceleration in the rate of growth of the real minimum wage. It is this deceleration what causes this negative influence: since the minimum wage is a key determinant of wage adjustments in Colombia, when the nominal minimum wage growth and sectoral inflation tend to converge, real wage freezes are more likely than real wage cuts, and DRWR becomes larger.

Equation (2.9) shows that $\Delta\pi$ exerts a positive, but far from significant, effect on DRWR (at standard critical values). It is positive, as expected, to the extent that the ratio of producer to consumer prices acts as a proxy of the inverse of the real minimum wage.[14] However, the fact that it is not fully significant –as in contrast $\Delta w_{it}^{\min}$ turns out to be in specification (2.8)– indicates that it is an inferior proxy.

[14] As noted before, nominal wage growth may be substituted by price inflation (since they move together). We can then take the inverse to place price inflation in the denominator and the growth rate of producer prices in the numerator.

Overall, the robustness of some crucial estimates is also noteworthy. For example, for a given econometric methodology, the estimated marginal effects of ΔY on DRWR are extremely stable across model specifications. Even across methodologies, the effects remain remarkably close. The estimates of informality are also notably robust across specifications. The same holds for the impact of θ.

The latter implies that consideration of the alternative specifications (2.8) and (2.9) are not distorting the analysis, but rather provide complementary views. In particular, the results on specification (2.8) show that $\Delta w^{\min}$ truly acts as a counterpart of ΔP, since there is virtually no change with respect to the marginal effects obtained via equation (2.7). On the contrary, $\Delta\pi$ in specification (2.9) is not such a precise counterpart, since it is not significant in our reference fixed effects estimation. Still, however, the estimates on ΔY and θ remain unchanged while, as noted before, the marginal effect of ι increases.

The complementary views obtained from models (2.7) and (2.8) point to the close association between price inflation and the wage setting system in Colombia with regard to its influence on wage or labour market rigidities. This important issue, and its policy implications, is discussed next.

2.4.2 Price, inflation, wage setting, and labour market rigidities

A simplistic, but illustrative view on the wage setting system in Colombia may be summarized through the following economic relationships:

$$\Delta NW_t = f\left(\ NW_t^{\min}\ \right), \tag{2.11}$$

$$\Delta NW_t^{\min} = f\ (\ \Delta P_{t-1}\ ), \tag{2.12}$$

where ΔNW denotes the rise in the nominal wage and $\Delta NW^{\min}$ denotes the rise in the nominal minimum wage, which in turn depends on past inflation (ΔP_{t-1}).

The way wages are fixed in Colombia provide three main channels by which inflationary pressures may enhance real wage rigidities. The first two act through the downward nominal wage rigidities ($DNWR$), while the third one operates via price inflation (ΔP).[15]

First channel: the minimum wage anchor. Wage negotiations in Colombia are based on the setup of the legal minimum wage, which is the first element to be determined and provides a floor for wage increases (note that this justifies the necessity to incorporate minimum wages in any model trying to explain wage rigidities in Colombia). It is on the basis of this arranged minimum wage that collective agreements are negotiated. Recent years have witnessed the difficulties to reach agreements between firms and unions, in which case it is the Ministry of Labour who sets the wage increase by law. The important point is that, irrespective of whether there is, or not, an agreement for the upcoming year, the reference increase is always given by the rise in the minimum wage (hence expression (2.11)).

Second channel: backward looking wage setting, and wage-price feedback. Minimum wages in Colombia are set as a function of past inflation (hence expression (2.12)). To illustrate this situation, Figure 2.2 plots the evolution of the minimum wage together with lagged inflation to uncover the close paths they follow. This explains the crucial role exerted by minimum wages in creating a wage-price spiral, since they act as a transmission channel from past inflation to the current rise in nominal wages.

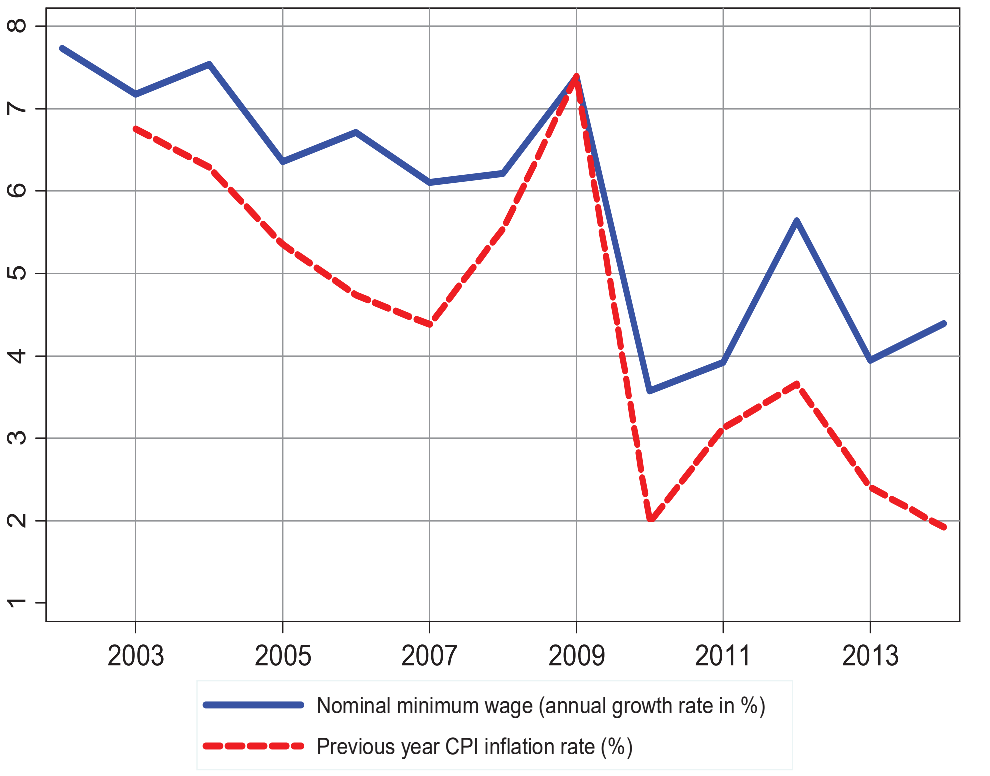
In contrast, most wage setting systems in developed economies are characterized by wage indexation over expected price inflation (ΔP^e), so that wage setting has a forward looking orientation. Given that wage progress is also tied to productivity growth (Δpr_t), such systems may be represented by the following expression:[16]

$$\Delta NW_t = f\left(\Delta P_t^e, \Delta pr_t\right). \tag{2.13}$$

[15] Again in a simplistic but illustrative view, it may be useful to see $DRWR$ as depending on $DNWR$ and ΔP^s, so that $DRWR = f(DNWR, \Delta P^s)$ on account of the approximation $\Delta RW \simeq \Delta NW - \Delta P^S$.

[16] It is well-known that forward-looking wage setting –on the basis of credible downward inflation expectations– was an effective tool to cut the wage-spiral problem that took place in many developed economies in the aftermath of the oil price shocks.

Figure 2-2: Indexation of nominal minimum wages, 2001-2014.



Source: Own calculations based on data from DANE and ENS.

Third channel: inflation persistence. Inflation persistence is positively related to DRWR. For example, Christoffel and Linzert (2012) show that larger degrees of DRWR tend to foster inflation persistence. In a backward-looking system like the Colombian one, reverse causality on this relationship may exist on account of the path dependency on past prices. Hence, if inflation persistence influences DRWR, it is plausible to expect a positive influence of inflation on DRWR in Colombia. This in contrast to some of the results obtained for developed economies, where the wage setting mechanism is based on price expectations and where productivity is the main driver of wage progress (as reflected in expression (2.13)).

Seen form the opposite side, what would be the incentive for Colombian firms, given the wage setting mechanism described above, to embark into costly wage cuts, in a context in which the wage-price feedback ensures inflation persistence? The absence of such incentive is what leads inflation to push DRWR.

To break this connection, the current wage setting system, as summarized by equations (2.11) and (2.12), should evolve towards a system closer to the one described by equation (*2.13*). This would contribute to attenuate the minimum wage anchor, tensions from the wage-price spiral, and pressures from inflation persistence. Next we close our analysis by calling for a wage setting system reform along these lines.

2.5 Conclusions and policy implications

We have explored the potential existence of DRWR in Colombia and evaluated its causes. The analysis is provided at the sectoral level, and is based on data for 59 sectors in 13 metropolitan areas over the period 2002-2014.

For the full sample, we estimate a deficit of real wage cuts of 12.09%, indicating that about 12 out of 100 notional real wage cuts do not result in an observed wage cut due to DRWR. In addition, the evolution of the DRWR is characterized by a declining path starting from values above 20% in 2002, showing a sharp decline in subsequent years, and then stabilizing around 10%.

At the sectoral level, differences are large. They may be as high as 30% and 40% as in

Agriculture and Mining; but they may also be non-significant as in Energy, Construction and Commerce, where wages are virtually flexible. Although these large differences may be associated to disparities in trade union density rates across sectors, we show that wage rigidity in Colombia is not fundamentally connected to the wage bargaining system. This is in contrast to empirical evidence on developed countries provided by Druant *et al.* (2012), Babecký et al. (2010) and Holden and Wulfsberg (2009).

On the contrary, economic growth, inflation, real minimum wage growth and labour informality appear as the crucial drivers. We find that a 1 percentage point increase in sectoral GDP growth generates a fall in DRWR equivalent to 0.61 percentage points. In turn, the effects on DRWR of inflation, the real minimum wage growth and labour informality have similar magnitudes, around a third of the estimated one for economic growth. In particular, a 1 percentage point increase in sectoral prices generates a rise in DRWR equivalent to 0.23 percentage points, while the same increase in real minimum wages or in labour informality generates a fall in DRWR of around 0.20 percentage points.

Policywise, our results imply that Colombia has two main mechanisms to fight rigidities. The most effective one, twice the alternatives, is to boost economic growth. The more the economy grows, the less wage cuts will be prevented, and the less relevant wage rigidities will be. The problem with this solution is the cyclical nature of economic growth, which responds to all sorts of stimulus beyond the specific ones related to the labour market. The consequence is that this mechanism is beyond any particular labour market policy.

The second mechanism is to embrace labour market institutional reforms. Our findings point to two alternatives of similar power. One is to lower inflation or, equivalently, to rise its counterpart in our analysis, the real minimum wage. This increase, however, can only be achieved by keeping the nominal minimum wage growth above inflation which, under the current wage setting system, will put immediate pressure on labour costs. The second alternative is to allow informality to stay and even grow, but this is problematic too, if the aspiration is to evolve towards better quality jobs.

To this situation of seeming no way out, we need to add two quasi-certainties and a fact.

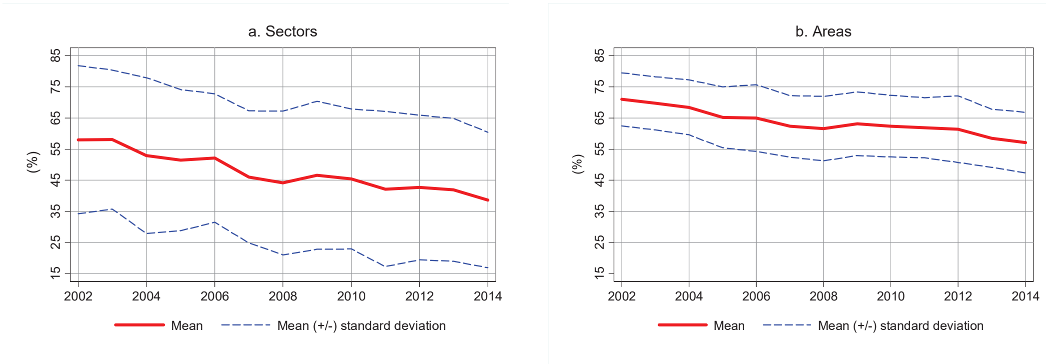
First, the inflation targeting approach adopted by the Central Bank of Colombia seems a guarantee that price inflation will remain under control. Second, informality has been on a downward trend since the beginning of the century. Being still at a 50% share over total employment, we should expect policy measures aiming at the continuation of this downward path. Third, the fact is that there is a very low trade union density rate (and a very weak system of collective bargaining) not even relevant to explain the wage behavior (Agudelo and Sala, 2016).

It is in this context that policy makers should envisage a far-reaching reform of the wage setting process. This reform should aim at stablishing a new system of collective bargaining in which wages, on one side, would become much more attached to productivity growth and, on the other side, would start being fixed over expected prices. This would connect labour compensation to economic performance and help in controlling inflation rates. Under such framework, wage rigidities would become less binding (since wages become aligned with productivity, and Δp is more easily controlled) without the need to reduce minimum wages (which constitutes one of the scarce mechanisms preventing work conditions to deteriorate). As a further outcome, such system would possibly allow for a larger degree of manouvre to reduce informality without firms facing unbearable labour costs and workers unbearable working conditions. A solid and well-designed system of collective bargaining is certainly crucial on all these grounds, and would help minimizing the undesired wage and unemployment costs of the current wage setting system. Murtin *et al.* (2014) provide critical and, in our view, notably helpful complementary evidence for 15 OECD economies on the consequences of the collective agreements system design.

Our study should be considered as a stepping stone towards a better understanding of real wage rigidities in the Colombian labour market. Whenever the data allows further research to be undertaken, the skill level of workers should be brought into the analysis (beyond the formal/informal distinction), and new controls related to non-wage costs and type of contracts should be considered.

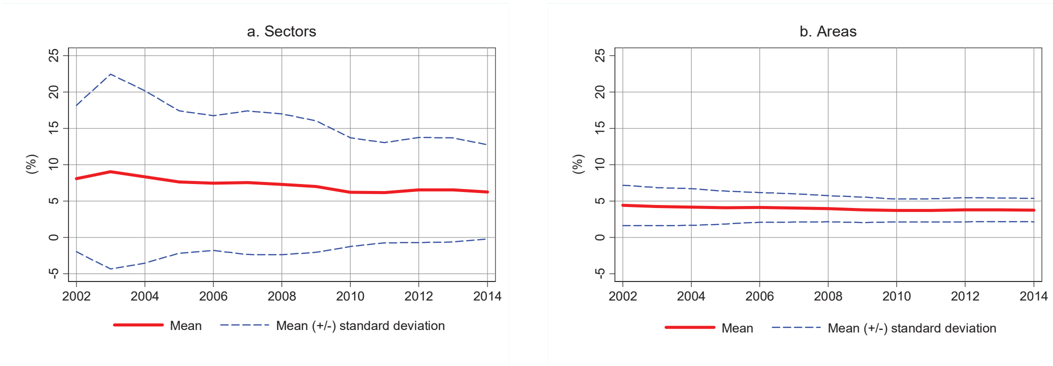
Appendix 2

Figure 2-3: Labour informality, 2002-2014.



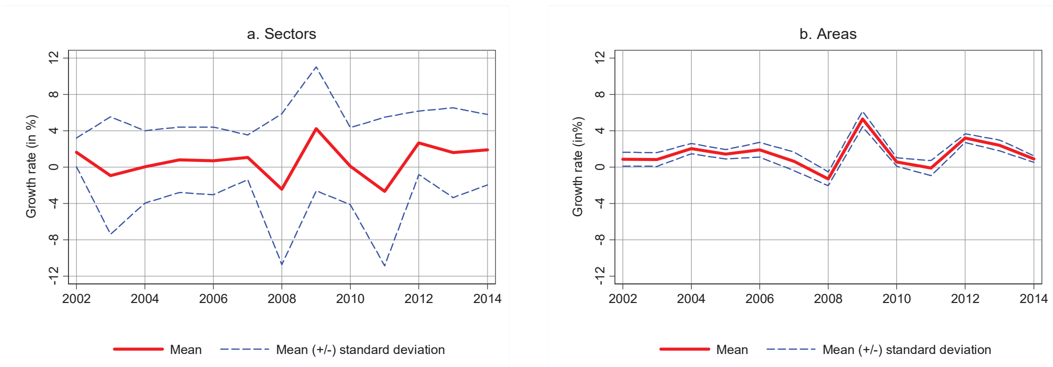
Source: Own calculations based on data from DANE.

Figure 2-4: Trade union density, 2002-2014.



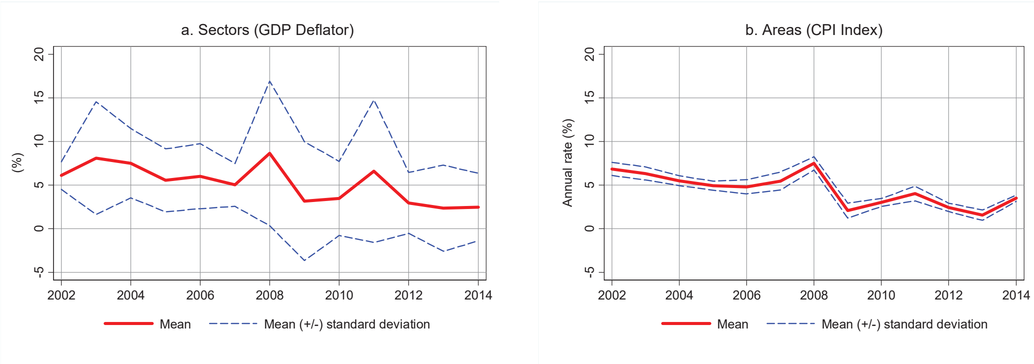
Source: Own calculations based on data from DANE.

Figure 2-5: Real minimum wage, 2002-2014.



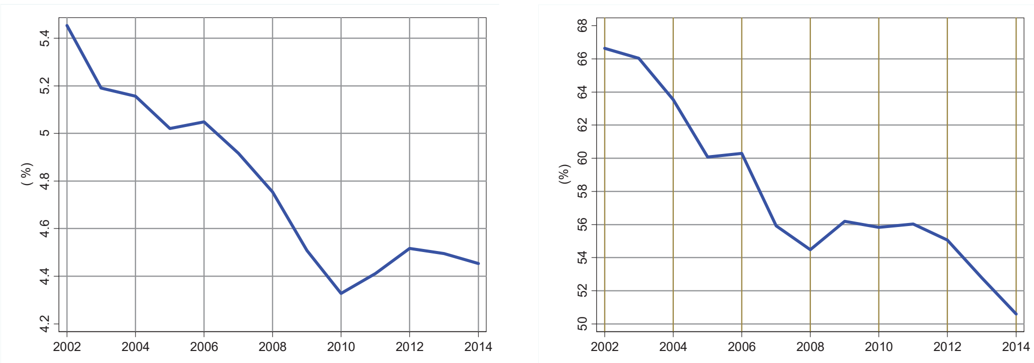
Source: Own calculations based on data from DANE.

Figure 2-6: Inflation, 2002- 2014.



Source: Own calculations based on data from DANE.

Figure 2-7: Trade union density and labour informality, 2002-2014.



Source: Own calculations based on data from DANE and ENS.

Table 2.6: Sectoral estimates.

Sector	Year	$\widetilde{q}$	q	$FWCP$	$p-value$	S
Agriculture & Fishing	2002	52.93	28.57	46.02	0.005	35
	2003	71.62	42.86	40.16	0.000	35
	2004	62.73	32.35	48.42	0.001	34
	2005	36.09	25.00	30.73	0.109	36
	2006	49.66	40.63	18.20	0.199	32
	2007	56.25	31.43	44.13	0.003	35
	2008	38.51	29.41	23.63	0.186	34
	2009	57.32	40.00	30.22	0.030	35
	2010	48.03	36.11	24.81	0.100	36
	2011	54.34	48.48	10.77	0.304	33
	2012	48.48	34.21	29.43	0.059	38
	2013	51.58	40.63	21.23	0.146	32
	2014	37.84	22.58	40.32	0.058	31
Mine & Quarry Exploitation	2002	56.12	23.68	57.80	0.000	38
	2003	73.52	43.24	41.18	0.000	37
	2004	41.72	29.27	29.85	0.067	41
	2005	47.00	27.50	41.49	0.012	40
	2006	64.02	37.14	41.98	0.001	35
	2007	27.20	8.33	69.36	0.006	36
	2008	70.27	44.74	36.34	0.000	38
	2009	40.37	27.78	31.19	0.079	36
	2010	62.58	33.33	46.74	0.000	39
	2011	63.22	44.44	29.70	0.015	36
	2012	41.26	28.57	30.75	0.064	42
	2013	45.95	26.83	41.61	0.012	41
	2014	43.24	29.27	32.32	0.048	41
Manufacturing Industry	2002	48.97	37.05	24.34	0.000	278
	2003	48.32	45.11	6.63	0.161	266
	2004	46.24	42.01	9.15	0.087	269
	2005	48.34	41.11	14.95	0.010	270
	2006	53.52	52.38	2.12	0.381	252
	2007	44.09	34.20	22.43	0.001	269
	2008	48.66	44.12	9.34	0.067	272
	2009	54.31	49.07	9.65	0.048	269
	2010	49.18	40.45	17.75	0.003	267
	2011	55.64	47.58	14.48	0.005	269
	2012	57.95	51.09	11.84	0.013	276
	2013	45.17	40.96	9.32	0.091	271
	2014	38.53	35.56	7.72	0.174	270

Continuation of Table 2.6.

Sector	Year	$\widetilde{q}$	q	$FWCP$	$p-value$	S
Electricity, Gas & Water	2002	54.48	53.85	1.16	0.544	26
	2003	75.34	57.69	23.42	0.035	26
	2004	46.37	46.15	0.46	0.563	26
	2005	31.00	42.31	-36.49	0.929	26
	2006	51.18	50.00	2.31	0.532	26
	2007	35.90	30.77	14.28	0.368	26
	2008	58.95	61.54	-4.39	0.672	26
	2009	30.41	38.46	-26.50	0.864	26
	2010	45.52	46.15	-1.38	0.604	26
	2011	33.28	38.46	-15.58	0.785	26
	2012	46.37	46.15	0.46	0.571	26
	2013	55.57	53.85	3.11	0.497	26
	2014	69.51	61.54	11.47	0.248	26
Construction	2002	41.72	46.15	-10.62	0.728	13
	2003	57.60	53.85	6.52	0.502	13
	2004	56.08	61.54	-9.73	0.747	13
	2005	32.43	46.15	-42.31	0.914	13
	2006	54.90	53.85	1.92	0.594	13
	2007	41.89	38.46	8.19	0.522	13
	2008	41.22	38.46	6.68	0.550	13
	2009	89.70	84.62	5.66	0.393	13
	2010	20.95	15.38	26.55	0.464	13
	2011	40.03	38.46	3.93	0.573	13
	2012	73.65	53.85	26.89	0.098	13
	2013	62.16	53.85	13.38	0.367	13
	2014	59.97	53.85	10.21	0.427	13
Commerce & Hotels	2002	36.06	32.69	9.35	0.372	52
	2003	57.90	57.69	0.35	0.543	52
	2004	39.61	38.46	2.90	0.505	52
	2005	31.59	30.77	2.59	0.523	52
	2006	42.44	40.38	4.85	0.447	52
	2007	33.99	36.54	-7.48	0.702	52
	2008	36.66	38.46	-4.93	0.671	52
	2009	69.51	63.46	8.70	0.220	52
	2010	41.72	44.23	-6.01	0.694	52
	2011	55.57	57.69	-3.81	0.682	52
	2012	49.45	51.92	-5.00	0.695	52
	2013	50.89	50.00	1.74	0.507	52
	2014	34.50	38.46	-11.48	0.778	52

Continuation of Table 2.6.

Sector	Year	$\widetilde{q}$	q	$FWCP$	$p-value$	S
Transport & Communication	2002	53.14	43.55	18.05	0.084	62
	2003	64.76	57.63	11.02	0.149	59
	2004	55.24	55.93	-1.26	0.597	59
	2005	41.49	46.43	-11.91	0.821	56
	2006	59.49	46.67	21.56	0.028	60
	2007	41.93	38.60	7.94	0.349	57
	2008	59.83	50.00	16.43	0.082	60
	2009	61.82	51.79	16.24	0.084	56
	2010	40.37	32.76	18.86	0.151	58
	2011	35.81	34.48	3.71	0.462	58
	2012	22.20	24.56	-10.66	0.737	57
	2013	70.24	64.91	7.58	0.224	57
	2014	34.76	25.42	26.87	0.082	59
Financial & Insurance	2002	41.13	30.77	25.19	0.089	52
	2003	58.99	51.92	11.99	0.178	52
	2004	57.56	46.15	19.81	0.063	52
	2005	53.00	49.02	7.51	0.341	51
	2006	44.59	49.02	-9.92	0.773	51
	2007	41.17	27.45	33.33	0.034	51
	2008	51.60	44.23	14.29	0.172	52
	2009	52.49	48.00	8.56	0.311	50
	2010	69.26	57.69	16.70	0.052	52
	2011	20.90	21.57	-3.18	0.627	51
	2012	69.13	59.62	13.76	0.087	52
	2013	50.93	46.00	9.68	0.286	50
	2014	47.00	40.38	14.08	0.213	52
Social Services	2002	49.05	42.42	13.51	0.057	165
	2003	63.66	56.71	10.92	0.038	164
	2004	52.74	51.85	1.69	0.451	162
	2005	48.12	43.90	8.76	0.163	164
	2006	52.69	50.94	3.31	0.362	159
	2007	29.08	30.63	-5.31	0.703	160
	2008	53.37	52.44	1.74	0.438	164
	2009	60.81	57.32	5.75	0.190	164
	2010	52.25	45.91	12.13	0.069	159
	2011	49.30	47.20	4.25	0.322	161
	2012	56.62	55.00	2.85	0.366	160
	2013	48.40	42.50	12.18	0.078	160
	2014	41.94	40.37	3.75	0.382	161

Notes: Data in percent. Negative values for $FWCP$ are replaced by zeros for estimating equations (7), (8) and (9).

Table 2.7: Estimates. Linear regressions.

Dependent variable: $FWCP_{it}$						
	Equation (2.7)		Equation (2.8)		Equation (2.9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.84** [-1.99]	-0.61* [-1.87]	-0.86** [-2.03]	-0.61* [-1.88]	-0.82** [-2.00]	-0.59** [-1.83]
Δp_{it}	0.58** [2.20]	0.08 [0.46]				
θ_{it}	-0.87*** [-4.68]	0.39 [0.93]	-0.86*** [-4.51]	0.41 [0.96]	-0.85*** [-4.44]	0.41 [0.94]
ι_{it-1}	-0.29*** [-4.03]	-0.09 [-0.94]	-0.28** [-3.90]	-0.09 [-0.92]	-0.28*** [-3.92]	-0.08 [-0.83]
$\Delta w_{it}^{\min}$			-0.55** [-2.01]	-0.07 [-0.39]		
$\Delta \pi_{it}$					0.54* [1.94]	0.00 [0.02]
c	35.2*** [6.79]	37.1*** [5.90]	38.3*** [7.94]	37.4*** [5.60]	37.3*** [7.55]	36.8*** [5.56]
$Obvs.$	117	117	117	117	117	117
LL	-465.5	-408.0	-465.8	-408.0	-466.3	-408.1

Notes: FE, sectorial fixed effects. *** Significant estimates at 1%; **, at 5%; *, at 10%. Z-test in brackets. P-values in parentheses. LL, Log-Likelihood.

Table 2.8: Estimates. Poisson regressions.

Dependent variable $FWCP_{it}$						
	Equation (2.7)		Equation (2.8)		Equation (2.9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.06** $[-2.21]$	-0.03* $[-1.92]$	-0.07** $[-2.25]$	-0.03* $[-1.93]$	-0.06** $[-2.08]$	-0.03* $[-1.83]$
Δp_{it}	0.04** [2.52]	0.01* [1.82]				
θ_{it}	-0.05*** $[-3.65]$	0.06 [1.45]	-0.05*** $[-3.58]$	0.06 [1.43]	-0.05*** $[-3.49]$	0.06 [1.30]
ι_{it-1}	-0.01*** $[-3.83]$	-0.01** $[-2.06]$	-0.01*** $[-3.71]$	-0.01* $[-1.91]$	-0.01*** $[-3.64]$	-0.01* $[-1.76]$
$\Delta w_{it}^{\min}$			-0.04** $[-2.22]$	-0.01 $[-1.56]$		
$\Delta \pi_{it}$					0.04* [1.87]	0.01 [1.03]
c	3.54*** [15.45]	3.92*** [13.44]	3.97*** [12.36]	4.46*** [11.21]	3.72*** [16.09]	3.93*** [12.18]
$Obvs.$	117	117	117	117	117	117
LL	-89.37	-55.66	-89.87	-55.78	-90.66	-55.94
D_{gof}	132.**3** (0.09)	64.85 (0.99)	133.3 (0.08)	65.07 (1.00)	134.8 (0.07)	65.39 (1.00)
P_{gof}	147.5 (0.01)	59.38 (1.00)	148.3 (0.01)	59.49 (1.00)	147.4 (0.01)	59.71 (1.00)

Notes: FE, sectorial fixed effects. *** Significant estimates at 1%; **, at 5%; *, at 10%. Z-test in brackets. P-values in parentheses. LL, Log-Likelihood. D_{gof}, deviance goodness of fit. P_{gof}, Pearson goodness of fit.

Table 2.9: Average marginal effects. Poisson regressions.

	Equation (2.7)		Equation (2.8)		Equation (2.9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.83** [−2.27]	-0.42** [−1.97]	-0.88** [−2.32]	-0.43** [−1.98]	-0.79** [−2.11]	-0.41** [−1.88]
Δp_{it}	0.56** [2.**66**]	0.16* [2.07]				
θ_{it}	-0.68*** [−2.90]	0.84 [1.46]	-0.67*** [−3.27]	0.85 [1.44]	-0.65** [−3.16]	0.81 [1.31]
ι_{it-1}	-0.20*** [−3.61]	-0.13*** [−2.12]	-0.20*** [−3.48]	-0.13** [−1.96]	-0.20*** [−3.40]	-0.12** [−1.80]
$\Delta w_{it}^{\min}$			-0.54** [−2.32]	-0.14 [−1.60]		
$\Delta \pi_{it}$					0.48* [1.92]	0.11 [1.05]

Notes: FE, sectorial fixed effects. *** Significant estimates at 1%; **, at 5%; *, at 10%. Z-test in brackets.

Chapter 3

Labour demand effects of internationalization in the Colombian manufacturing industry

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 –the most exposed economic activity–. 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.

JEL Classification: J23, F41, F16.

Keywords: Labour Demand Elasticity, Substitution Elasticity, Trade openness.

3.1 Introduction

In the context of increasing globalization, labour outcomes and international trade patterns have become progressively interlinked. Labour market responses to policy measures have changed and a growing body of literature has become interested in the interrelation between globalization, employment and labour market reforms (e.g. Selwaness and Zaki, 2018; Malgouyres, 2017; Acemoglu *et al.*, 2016; and Autor, 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 an excellent experience to examine. It is one of the considered successful economies in Latin America which, in recent decades, 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. In addition, between 1992 and 2004, Colombia enjoyed a new system of preferential tariffs to export to the US. This system was superseded in 2004 by Free Trade Agreements (FTAs) between Colombia and a number of 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 3.3a). 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 3.4a, 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 increase 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 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 a large number of 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 Manoley (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 3.4). 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 time 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 3.2 describes some macro developments in the Colombian manufacturing industry. Section 3.3 discusses the analytical framework on the labour demand effects of international trade Section 3.4 describe the empirical implementation. Section 3.5 discusses the data and explains the econometric methodology. Section 3.6 deals with the estimates. Section 3.7 provides an assessment of the main results.

3.2 Stylized facts

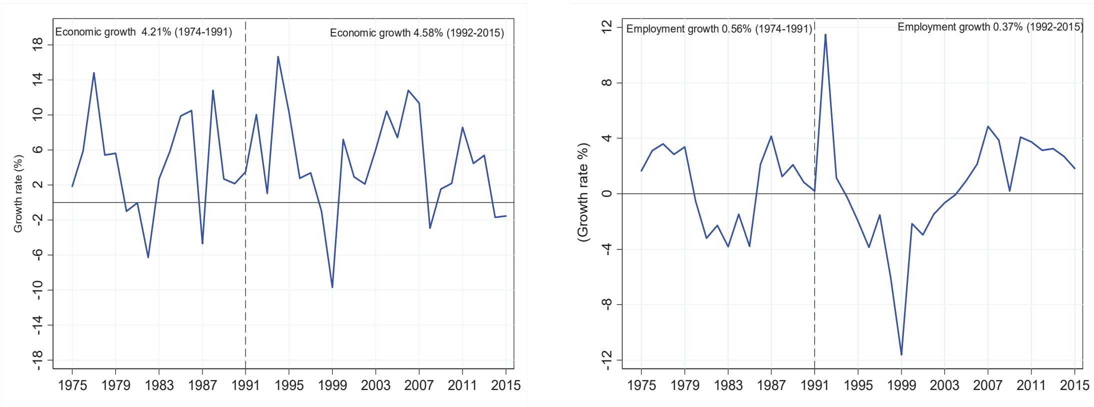
In this section, we provide descriptive information on some macroeconomic variables of the Colombian manufacturing industry across Figures 3.1 to 3.4. All data is supplied for the two relevant periods of analysis –the slow transition between import substitution and trade liberalization in 1974-1991; and the trade and institutional reforms years of 1992-2015–.

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 3.1a). In spite of these positive growth rates throughout, the employment growth rates did not reach 1% on average (Figure 3.1b). 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 3.2a), 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 3.4b).

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 3.3a)– 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 3.3a). The share of non wage labour costs

Figure 3-1: Production and employment dynamics in the Colombian manufacturing industry 1974-2015.

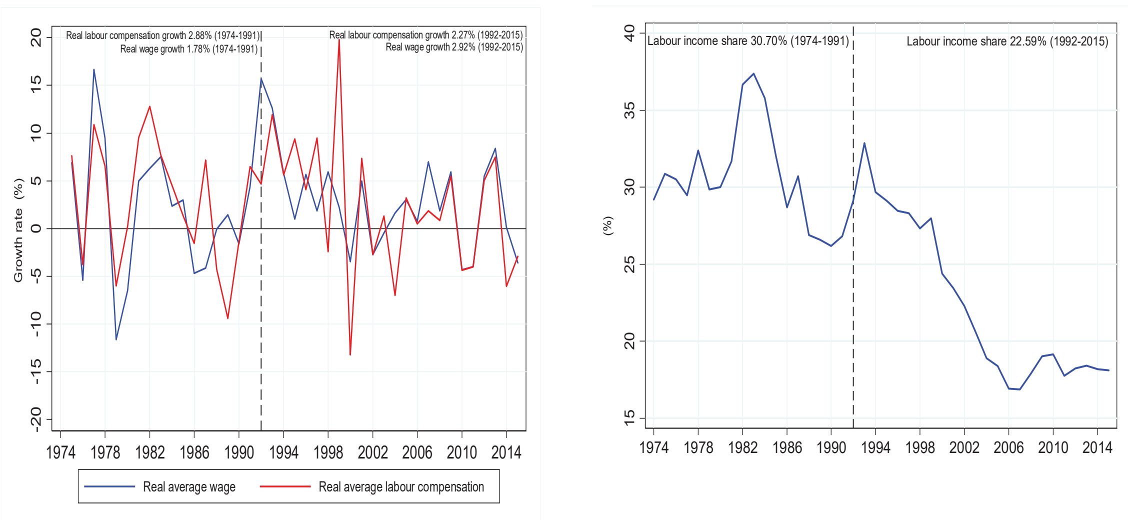
a. Economic growth **b. Employment growth**



Source: Encuesta Anual Manufacturera (EAM)

Figure 3-2: Labour costs dynamics in the Colombian manufacturing industry 1974-2015.

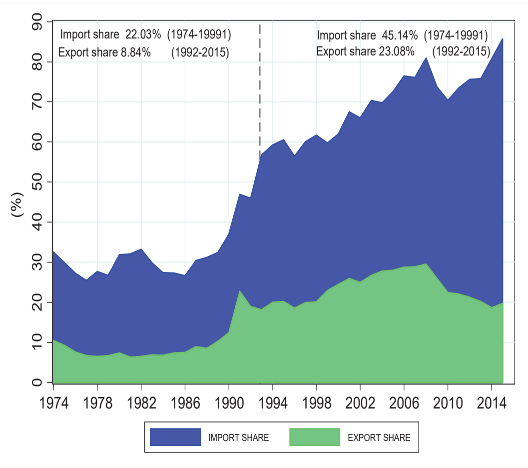
a. Real labour costs **b. Labour income share**



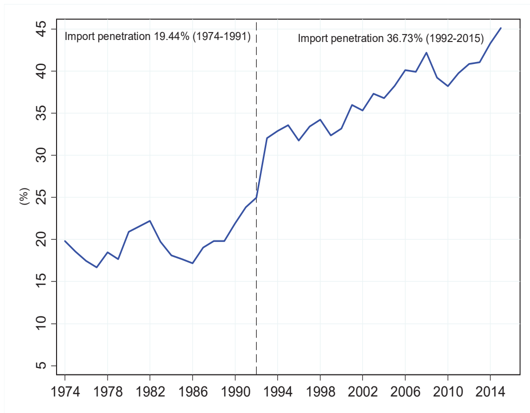
Source: Encuesta Anual Manufacturera (EAM)

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

a. Degree of trade openness



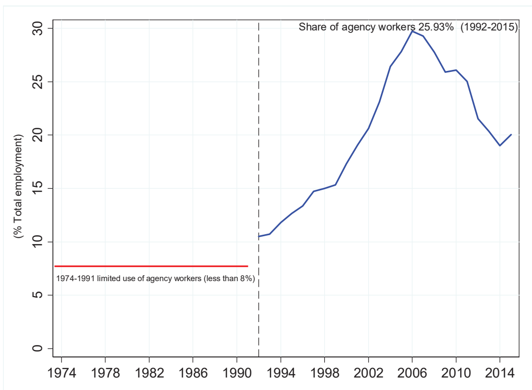
b. Import penetration



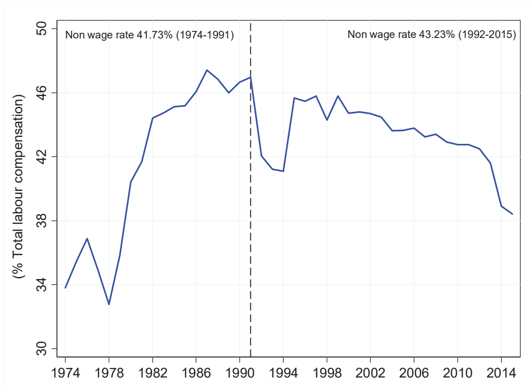
Source:Departamento Nacional de Planeación (DNP)

Figure 3-4: Labour market segmentation and flexibilization in the Colombian manufacturing industry 1974-2015.

a. Share of agency workers



b. Share of non wage labour costs



Source: Encuesta Anual Manufacturera (EAM) declined by almost 10 percentage points from a value of 46% in 1991, to one of 36% in 1975 (Figure 3.4b). 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 3.2b). 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 as a result 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.

3.3 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.

3.3.1 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:

$$\eta_{LL} = -(1 - s_L)\sigma - s_L\eta \tag{3.1}$$

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 (3.1) 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 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]

[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

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 Bockerman 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 (3.1)). For instance, Hijzen and Swaim (2010) have conjectured that there would be a fall in the labour income share (s_L) as a consequence of offshoring. According to these authors, when an economy opens up 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 has to 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 opens to international trade, the labour share would remain constant. On the other side, some studies

effects often cannot be analyzed through Hamermesh's expression (3.1). 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.

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. 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.

3.3.2 International trade effects on the level of the labour demand

Rodrik (1997) in line with Hercksher-Olin-Samuleson 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 a 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 spillover 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).

3.4 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 asses the labour demand effects of international trade, and disclose our strategy to approach quantitatively, in particular, the level effect of international trade.

3.4.1 Baseline equations

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

$$n_t = \alpha_0 + \alpha_1 w_t + \alpha_2 k_t + u_t, \tag{3.2}$$

$$n_t = \beta_0 + \beta_1 w_t + \beta_2 y_t + e_t. \tag{3.3}$$

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 (3.1). To be specific, equation (3.2) allows us to obtain a direct estimate of the labour demand elasticity with respect to wages (η_{LL}in Hamermesh's expression 3.1) while equation (3.3) allows us to estimate the substitution elasticity between capital and labour (σ in Hamermesh's expression 3.1). 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 (3.1) and equations (3.2) and (3.3) –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.

Theoretical background for equation (3.2)

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 $q_{it}^s = An_{it}^{\alpha}\overline{k_{it}}^{(1-\alpha)}$, 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.

$$N_t = \left[\alpha A\left(1 - \frac{1}{|\varepsilon|}\right)\right]^{\frac{1}{(1-\alpha)}}\left[\frac{W_t}{P_t}\right]^{\frac{-1}{(1-\alpha)}} K_t \ , \tag{3.4}$$

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 $u_t\tilde{}i.i.N(0, \sigma^2)$ to

[2] Karanassou *et al.* (2007) assume that the technological change grows at constant rates λ, thus it can be expressed as $A = A_0e^{\lambda}$.

capture supply and demand shocks, and rearranging the terms as follows:

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

we obtain equation (3.2) as the empirical counterpart of equation (3.4) –usually called unconditional or capital constrained labour demand–.

Note that all coefficients in equation (3.2) 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.

Theoretical background for equation (3.3)

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

$$Y_t = A\left[\theta N_t^{\frac{\sigma-1}{\sigma}} + (1-\theta)K_t^{\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{\sigma-1}}, \tag{3.5}$$

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):

$$A^{\frac{\sigma-1}{\sigma}}\theta\left[\frac{Y_t}{N_t}\right]^{\frac{1}{\sigma}} = \frac{W_t}{P_t}, \tag{3.6}$$

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

$$N_t = A^{\sigma-1}\theta^{\sigma}\left[\frac{W_t}{P_t}\right]^{-\sigma} Y_t \tag{3.7}$$

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

$$n_t = ln(N_t);\ w_t = ln(\tfrac{W_t}{P_t});\ y_t = ln(Y_t);\ \beta_o = \sigma ln(A\theta) - ln(A);\ \beta_1 = -\sigma;\ \beta_2 = 1;$$

we obtain equation (3.3) as the empirical counterpart of equation (3.7).

Note that all coefficients in equation (3.3) 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.

3.4.2 Decomposition of the elasticity for labour demand

We use the empirical estimation of models (3.2) and (3.3) to perform a quantitative decomposition of the labour demand elasticity. The crucial coefficients are α_1 in model (3.2) and β_1 in model (3.3). The coefficient α_1 measures the total elasticity of employment with respect to real wage (η_{LL} in Hamermesh's expression 3.1). 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 3.1). 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 3.1).

When wages rise, the relative price of labour *vis-à-vis* the relative price of capital 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*. Sala and Trivín (2012) 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. *Holding 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 (3.8) provides a simple illustration of our empirical approach.

$$\boxed{\begin{array}{cccc} \text{Total effect} = & \text{Constant-output elasticity} & + & \text{Scale effect} \\ [\alpha_1] & [(1-s_L)\beta_1] & & [\alpha_1 - (1-s_L)\beta_1] \end{array}} \tag{3.8}$$

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, Rusell 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 (3.1)–, 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 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 (3.3) 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 $(1 - s_L)\beta_1$.

3.4.3 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 a 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 (3.2) and (3.3) 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 spillover 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 (3.4) and (3.7), that the efficiency parameter A, among other elements, is determined by international trade changes, then we can obtain the following extended equations (see Appendix 3 for details)

$$n_t = \alpha_0 + \alpha_1 w_t + \alpha_2 k_t + \alpha_3 open_t + u_t. \tag{3.9}$$

$$n_t = \beta_0 + \beta_1 w_t + \beta_2 y_t + \beta_3 open_t + e_t. \tag{3.10}$$

The coefficient α_3 in equation (3.9) can be interpreted as the international trade effect on the level of the labour demand. In contrast, the coefficient β_3 in equation (3.10) only captures a part of the changes 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 3, extensions for equations (3.2) and (3.3), 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.

3.4.4 Additional considerations

To analyze the Colombian case, equations (3.9) and (3.10) 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, due to the fact that 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:

$$n_t = \alpha_0 + \lambda n_{t-1} + \alpha_1 w_t + \alpha_2 k_t + \alpha_3 open_t + \alpha_4 \Delta ser_t + u_t. \tag{3.11}$$

$$n_t = \beta_0 + \lambda n_{t-1} + \beta_1 w_t + \beta_2 y_t + \beta_3 open_t + \beta_4 \Delta ser_t + e_t. \tag{3.12}$$

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.

3.5 Data and empirical modeling

3.5.1 Data

We use a panel database with a cross-section dimension of $N = 16$ sectors and a time dimension of $T = 42$ years covering the period 1974-2015.[3]. Table 3.1 lists the variables and the corresponding sources.

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

[3] The detailed list of sectors is provided in List 3.1, in the Appendix 3.

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.

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 (Departamento Nacional de Planeación, DNP). 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 3.1 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.

Table 3.1: Definitions of variables.

Variables		Sources	Subindices	
n_{it}	Employment	(1)	i	$= 1, ..., 16$ sectors
w_{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 $\left[\frac{w_{it} * n_{it}}{y_{it}}\right]$	(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	Services sector share	(3)		
d^{92}	Dummy: value 1 1992 onwards			
d^{8083}	Dummy: value 1 in 1980-1983			
d^{9700}	Dummy: value 1 in 1997-2000			
d^{0809}	Dummy: value 1 in 2008-2009			

Notes: 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.

3.5.2 Econometric methodology

As we work with a two-dimensional panel data, we add sectorial fixed effects in order 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 (3.11) and (3.12) will be estimated as partial adjustment models taking the following general form:

$$n_{it} = \gamma_i + \lambda n_{it-1} + \theta Z_{it} + v_{it}, \qquad v_{it} \sim i.i.N\left(0, \sigma^2\right) \tag{3.13}$$

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

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.

In order to check the order of integration of the variables, we carry out a set of stationary and unit root tests depending on the type of the variables to be dealt with. In particular, 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 to use the KPSS test in order 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.

Table 3.2 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 3.3 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

[4] Using Monte Carlo simulations, Maddala and Wu (1999) conclude that this problem is less severe with the Fisher 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 3.2: Panel Unit Root Test, 1974-2015.

	n_{it}	w_{it}	k_{it}	y_{it}	op_{it}	m_{it}	x_{it}
MW	122.89 [0.000]	184.99 [0.000]	117.55 [0.000]	122.89 [0.000]	104.85 [0.000]	96.54 [0.000]	119.86 [0.000]
Result	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.

Table 3.3: Stationary and Unit Root Tests, 1974-2015.

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

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

be rejected at the 5% significance level, while in both ADF and PP test the null hypothesis of an 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 favor of proceeding with stationary panel data estimation techniques.

Econometric method

Models (3.11) and (3.12) are estimated by applying Ordinary Least Square (OLS) and one-way Fixed Effects (FE) (see Tables 3.4 and 3.5). 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 (3.11) and (3.12); 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 30, 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 can not 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 3.8), 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 stage 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 (3.11) and (3.12), we rely on the performance of three tests. First, an LM test checking for underidentication (i.e., that the excluded instruments are not relevant, meaning non correlated with the endogenous regressors). This is denoted as U in Table 3.4 and 3.5, and the null hypothesis is that the equation is under-

[5] Nickell (1981) stand out that in small macro-panels (e.g. T around 30 and N 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 $N \to \infty$ 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.

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

3.6 Results

This section presents the empirical results of our study in three subsections. First, we present and discuss the estimates of models (3.11) and (3.12). Then we present and analyze 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–.

3.6.1 Estimates

Tables 3.4 and 3.5 display, respectively, the estimated labour demand models (3.11) and (3.12) 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,

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

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 favor some particular specifications, we would choose those obtained through the FE-TSLS estimator (Columns (3), (6) and (9) in Tables 3.4 and 3.5 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 3.4 and 3.5, Columns (3), (6) and (9)) provide a similar picture: high employment persistence; 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 (3.11) and (3.12), while the evidence on the effect of the export share is mixed. The export share is found to be significant in model (3.11) but non-significant in model (3.12).

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 3.4 and 3.5, Columns (3), (6) and (9), however, show larger short-run sensitivities but also larger employment persistence. 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 3.4a, 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 3.4b), and also as a mechanism to 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 (3.11) and (3.12) are estimated by using only workers directly paid by firms, we interpretate the significant increase in the persistence coefficient as a 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

Table 3.4: Estimates of model 3.11.

Dependent variable: n_{it}	op_{it}			m_{it}			x_{it}		
	OLS	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
n_{it-1}	0.99 [0.000]	0.92 [0.000]	0.87 [0.000]	0.99 [0.000]	0.91 [0.000]	0.87 [0.000]	0.99 [0.000]	0.91 [0.000]	0.84 [0.000]
w_{it}	−0.03 [0.058]	−0.07 [0.030]	−0.14 [0.005]	−0.03 [0.061]	−0.07 [0.038]	−0.14 [0.006]	−0.03 [0.042]	−0.07 [0.012]	−0.17 [0.002]
k_{it}	0.01 [0.072]	0.02 [0.029]	0.08 [0.003]	0.01 [0.062]	0.02 [0.035]	0.08 [0.004]	0.01 [0.068]	0.02 [0.008]	0.10 [0.001]
$open_{it}$	0.00 [0.604]	−0.01 [0.308]	−0.01 [0.214]	0.00 [0.726]	−0.01 [0.259]	−0.01 [0.143]	−0.01 [0.147]	−0.01 [0.048]	−0.02 [0.003]
Δser_t	−2.39 [0.000]	−1.98 [0.000]	−2.05 [0.000]	−2.38 [0.000]	−1.96 [0.000]	−2.04 [0.000]	−2.38 [0.000]	−1.99 [0.000]	−2.03 [0.000]
$n_{it-1}*d_{92}$	0.01 [0.493]	0.01 [0.007]	0.02 [0.025]	0.01 [0.600]	0.02 [0.001]	0.02 [0.030]	0.01 [0.470]	0.01 [0.030]	0.03 [0.008]
$w_{it}*d_{92}$	−0.01 [0.442]	−0.01 [0.005]	−0.03 [0.015]	−0.01 [0.453]	−0.01 [0.002]	−0.03 [0.019]	−0.01 [0.492]	−0.01 [0.027]	−0.04 [0.006]
d_{8083}	−0.03 [0.001]	−0.03 [0.004]	−0.02 [0.060]	−0.03 [0.001]	−0.02 [0.004]	−0.02 [0.070]	−0.03 [0.001]	−0.03 [0.001]	−0.02 [0.043]
d_{9700}	−0.05 [0.000]	−0.06 [0.000]	−0.05 [0.000]	−0.05 [0.000]	−0.06 [0.000]	−0.05 [0.000]	−0.05 [0.000]	−0.05 [0.000]	−0.05 [0.000]
d_{0809}	0.02 [0.117]	0.02 [0.168]	0.02 [0.262]	0.02 [0.115]	0.02 [0.165]	0.02 [0.267]	0.03 [0.092]	0.02 [0.145]	0.02 [0.190]
c	0.16 [0.076]	1.15 [0.003]		0.17 [0.071]	1.17 [0.001]		0.20 [0.038]	1.18 [0.003]	
$Obvs.$	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 test.

Table 3.5: Estimates of model 3.12.

Dependent variable: n_{it}	op_{it}			m_{it}			x_{it}		
	OLS	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
n_{it-1}	0.95	0.90	0.85	0.96	0.90	0.84	0.96	0.89	0.83
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
w_{it}	−0.06	−0.08	−0.17	−0.06	−0.08	−0.17	−0.06	−0.08	−0.18
	[0.000]	[0.000]	[0.002]	[0.000]	[0.000]	[0.003]	[0.000]	[0.000]	[0.001]
y_{it}	0.04	0.05	0.11	0.04	0.05	0.12	0.04	0.05	0.12
	[0.000]	[0.004]	[0.001]	[0.000]	[0.003]	[0.002]	[0.000]	[0.009]	[0.001]
$open_{it}$	0.01	0.00	0.01	0.00	−0.00	0.00	−0.00	−0.01	−0.00
	[0.105]	[0.909]	[0.216]	[0.225]	[0.582]	[0.546]	[0.278]	[0.254]	[0.372]
Δser_t	−2.14	−1.84	−1.54	−2.14	−1.83	−1.52	−2.16	−1.84	−1.47
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.001]	[0.000]	[0.000]	[0.001]
$n_{it-1} * d_{92}$	0.01	0.02	0.03	0.01	0.02	0.03	0.01	0.02	0.04
	[0.057]	[0.000]	[0.010]	[0.066]	[0.000]	[0.012]	[0.057]	[0.000]	[0.003]
$w_{it} * d_{92}$	−0.01	−0.02	−0.03	−0.01	−0.02	−0.04	−0.01	−0.02	−0.04
	[0.056]	[0.000]	[0.012]	[0.067]	[0.000]	[0.015]	[0.075]	[0.000]	[0.005]
d_{8083}	−0.02	−0.02	−0.02	−0.03	−0.02	−0.02	−0.03	−0.02	−0.01
	[0.003]	[0.002]	[0.094]	[0.002]	[0.001]	[0.122]	[0.002]	[0.000]	[0.187]
d_{9700}	−0.04	−0.04	−0.03	−0.04	−0.04	−0.03	−0.04	−0.04	−0.03
	[0.000]	[0.000]	[0.013]	[0.000]	[0.000]	[0.019]	[0.000]	[0.000]	[0.023]
d_{0809}	0.02	0.02	0.01	0.02	0.02	0.01	0.02	0.02	0.01
	[0.208]	[0.272]	[0.522]	[0.199]	[0.250]	[0.508]	[0.164]	[0.225]	[0.490]
c	0.19	0.88		0.21	0.92		0.23	0.94	
	[0.031]	[0.021]		[0.018]	[0.013]		[0.010]	[0.021]	
$Obvs.$	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			16.52			19.78
			[0.000]			[0.001]			[0.000]
W			8.42			7.90			8.51
			(6.61)			(6.61)			(6.61)
H			0.06			0.03			0.07
			[0.970]			[0.987]			[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 test.

on the labour demand.

3.6.2 International trade effects on the elasticity for labour demand

Table 3.6 presents the implied long-run elasticities arising from the base-run estimates displayed in Tables 3.4 and 3.5, 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 3.6, 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. Given these results 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 3.6: 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 [0.000]	-1.03 [0.000]	-1.06 [0.000]	-1.58 [0.000]	-1.51 [0.000]	-1.59 [0.000]
σ	-1.11 [0.000]	-1.08 [0.000]	-1.04 [0.000]	-1.68 [0.000]	-1.68 [0.000]	-1.64 [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

	Robutsness check					
	1974-1991			1992-2015		
	(1)	(2)	(3)	(4)	(5)	(6)
Total effect and substitution elasticity						
η_{LL}	-1.20 [0.000]	-1.18 [0.000]	-1.16 [0.000]	-1.69 [0.000]	-1.69 [0.000]	-1.69 [0.000]
σ	-1.30 [0.000]	-1.32 [0.000]	-1.20 [0.000]	-1.94 [0.000]	-2.06 [0.000]	-1.88 [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.

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 and Maiti, 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 up 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 also 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 (3.1), 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.

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 the 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 (3.11) and (3.12). 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 taking into account 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 (3.8) and (3.9) using the new definition of employment. The corresponding results are displayed in Appendix 3 (Tables 3.9 and 3.10). 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 affect significantly the magnitude of the structural breaks just discussed.

As shown in Table 3.6, 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 changed 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.

3.6.3 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 3.7 displays the implied long-run effects on employment of the trade openness index (ξ_{op}), the import penetration ratio (ξ_m) and the export ratio (ξ_x), all of them arising from the base-run estimates presented in Tables 3.4 and 3.5. 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 (3.11), and Columns (2) and (4) to model (3.12).

As shown at the top of Table 3.7, the estimates for models (3.11) and (3.12) 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 (3.11) suggest a significant and negative influence of exports on the labour demand, while the estimates for model (3.12) point to the irrelevance of exports on employment determination. On this account, if we have to favor a particular model, we would choose model (3.11) –Columns (1) and (3)–. From a theoretical point of view, as described in section 3.4, model (3.11) 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 (3.12) 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 3, extensions for equations (3.2) and (3.3), for details).

Therefore, for the first period of analysis, results of model (3.11) 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. Adittionally, 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.

Table 3.7: Long-run effects

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

Notes: P-values in brackets.

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 3.7, we provide additional estimates which correspond to those described in the robustness check subsection. As we rely on the results of model (3.11), 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.

3.7 Concluding remarks

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

One 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. Economically, the increase in the sensitivity of employment with respect to wages even when holding output constant (the substitution effect) implies that Colombian firms have enhanced their internal flexibility to respond to price changes. This may be reflecting the enhanced possibilities brought by the new technologies and the growing pressure to which firms are subject to compete in the international arena.

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 analyzed 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, 2002). In such cases, workers may benefit from the process of trade liberalization even under increases in payroll taxation. Of course, it would be possible as long as 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 as a result 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.

Appendix 3

Extension of equation (3.4): Theoretical background

Departing from equation (3.4)

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

where $N_t = n_{it}f$ is aggregate employment, A is an efficiency parameter, $\frac{W_t}{P_t}$ 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 $u_t\tilde{}i.i.N(0, \sigma^2)$ to capture supply and demand shocks, and rearranging the terms as follows:

$n_t = ln(N_t); w_t = ln(\frac{W_t}{P_t}); k_t = ln(K_t); open_t = ln(OPEN_t); \alpha_0 = \frac{1}{(1-\alpha)} ln\left[A_0\left(1 - \frac{1}{|\varepsilon|}\right)\right];$ $\alpha_1 = \frac{-1}{(1-\alpha)}; \alpha_2 = 1; \alpha_3 = \frac{\delta}{(1-\alpha)}$.

We can obtain the labour demand (3.9) 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 (3.8): Theoretical background

Departing from equation (3.8)

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

where $N_t = n_{it}f$ is aggregate employment, A is an efficiency parameter, $\frac{W_t}{P_t}$ 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, adding a noise error term $e_t \tilde{} i.i.N(0, \sigma^2)$ to capture supply and demand shocks, and making the following rearrangements:

$n_t = ln(N_t)$; $w_t = ln(\frac{W_t}{P_t})$; $y_t = ln(Y_t)$; $open_t = ln(OPEN_t)$; $\beta_o = \sigma ln(\theta) + (\sigma - 1)ln(A_0)$; $\beta_1 = -\sigma$; $\beta_2 = 1; \beta_3 = \delta(\sigma - 1)$;

We can obtain the marginal condition (3.10) which incorporates a partial effect of international trade on technical efficiency through β_3.

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

List 3.1. 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 3.8: Bias-corrected LSDVC estimators.

		Model (3.11)			Model (3.12)		
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

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).

Table 3.9: Estimates of model 3.11. Robutsness check

Dependent variable: n_{it}	op_{it}			m_{it}			x_{it}		
	OLS	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
n_{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]
w_{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 [0.035]	−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 [0.0147]	−0.01 [0.874]	−0.01 [0.142]	−0.01 [0.215]	−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]
n_{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 [0.418]	−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 [0.822]	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]	
Obvs.	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]

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 test .

Table 3.10: Estimates of model 3.12. Robustness check

Dependent variable: n_{it}	op_{it}			m_{it}			x_{it}		
	OLS	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**	**OLS**	**FE**	**FE-TSLS**
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
n_{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]
w_{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 [0566]	−2.97 [0.000]	−2.44 [0.000]	−1.07 [0.100]
$n_{it-1} * d_{92}$	−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]	−0.0 [0.872]	−0.01 [0739]
c	−0.1 [0703]	1.66 [0.000]		−0.1 [0703]	1.66 [0.000]		−0.1 [0703]	1.66 [0.000]	
$Obvs.$	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	076
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 test .

Table 3.11: Bias-corrected LSDVC estimators. Robutsness check

		Model (3.11)			Model (3.12)		
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).

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