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ABSTRACT

Wage Rigidities in Colombia: Measurement, Causes, and Policy Implications*

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 labor 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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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 labor 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 labor 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 wage 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.[1] This means that inflation cannot be seen as a conjunctural phenomenon (mainly hurtful during slumps), but rather as a 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 labor market. To be precise, we provide a sectoral-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

[1]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.

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 labor 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 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 labor 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 measures the extent of downward real wage rigidity in Colombia. Empirical methods to evaluate its causes are discussed in Section 3. Section 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 5 concludes.

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.[2] 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.[3] 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

[2] 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).

[3] 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.

at the macro level is the object of our analysis.

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 industries 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 industries 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, variations in wage setting and other relevant variables are larger across sectors than across areas. This justifies our choice.

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:

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

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 sthochastic variable X from the underlying distribution $X \sim G(0, 1)$.

Figure 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.[4]

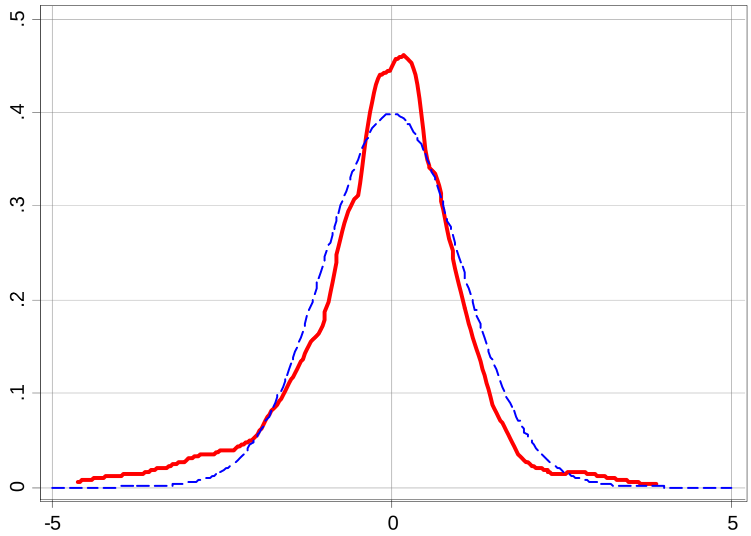
In the second step, for each of the 767 sector-years in the full sample, we compute the sector-year specific distributions of notional wage changes. This is done by adjusting as follows the underlying wage change distribution for the sector-specific observed median and inter-percentile range:

$$Z_{it} \equiv X(P75_{it} - P35_{it}) + \mu_{it}, \quad \forall i, t, \tag{2}$$

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.

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

Figure 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.

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:

$$\widetilde{q}(k)_{it} \equiv \frac{\#Z_{it}^{s} < k}{S^{H}}, \quad s = 1, ..., S^{H}, \tag{3}$$

where $\#Z_{it}^{s} < k$ 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

$$q(k)_{it} \equiv \frac{\#\Delta w_{jit} < k}{S_{it}}, \quad \forall j, \tag{4}$$

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.

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:

$$FWCP(k)_{it} = 1 - \frac{q(k)_{it}}{\widetilde{q}(k)_{it}}. \tag{5}$$

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.[5]

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.

2.3 Estimates

Table 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.

For the full sample (last row in Table 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 sectoral 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 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.

[5] The sector-year estimates of $\mathbf{FWCP(k=0)}$ for the 9 sectors are reported in Appendix 2. The number of observations in each sector-year sample is 13 or more.

Coincidentally, S2 shows a low share of informal employment and union density.

Table 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.

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.

3 The determinants of DRWR

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 labor conflicts). In this way, a benchmark empirical version of their model may be expressed through the following general form:[6]

$$FWCR_{it} = f\ \left(\ u_{it},\ \Delta p_{it},\ EPL_{it},\ \theta_{it}\ \right). \tag{6}$$

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 (6) relates to unemployment. As no series are available for sectoral 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 sectoral unemployment.[7] This is the reason why we rather use sectoral real economic 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 labor market tightness on wage rigidities. We thus expect a negative sign on the coefficient of ΔY.

[6] 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).

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

The second adjustment relates to the absence of EPL simply due to the lack of data. The counterpart here is the inclusion of labor 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 (6) becomes:

$$FWCR_{it} = f \ \left(\ \Delta Y_{it}, \ \Delta p_{it}, \ \iota_{it}, \ \theta_{it} \ \right), \tag{7}$$

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 ($\Delta \pi$):

$$\begin{aligned} FWCR_{it} &= f \ \left(\ \Delta Y_{it}, \ \Delta w_{it}^{\min}, \ \iota_{it}, \ \theta_{it} \ \right), & (8) \\ FWCR_{it} &= f \ \left(\ \Delta Y_{it}, \ \Delta \pi_{it}, \ \iota_{it}, \ \theta_{it} \ \right). & (9) \end{aligned}$$

In specification (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,[8] 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 (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

[8] 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.

reflects CPI inflation). Hence specification (9), where the change in the ratio between the prices faced by producers (at the sectoral 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.

3.2 Data

To 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 lists the variables and the corresponding sources.

Table 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}	Sectoral 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 sectoral 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 labor 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 sectoral and consumption prices (p^c).

Descriptive information on these variables at the sectoral level is provided in Table 3.

As noted when first presented in Table 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 (7) to (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 sectoral values.

Table 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 1. Information on $FWCP$ per sector is computed as the average of the estimates presented in Appendix 2.

Looking at Table 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 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 sectoral and consumption prices.

3.3 Econometric methodology

Just as initial benchmark, we start by estimating equations (7), (8) and (9) as linear regression models. The corresponding results are presented in Table A2 in Appendix 3. 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:

$$[1 - FWCP(k)] \; = \frac{Y(k)_{it}}{\widehat{Y}(k)_{it}} = e^{x'_{it}\beta + \varepsilon_{it}}, \quad if \;\; \widehat{Y}(k)_{it} > 0, \tag{10}$$

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 (7), (8) and (9) obtained from their estimation as Poisson regressions are presented in Appendix 3 for reference (in Table A3), 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 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.[9] This estimation is conducted using Maximum Log-Likelihood methods.[10]

4 Results and discussion

4.1 Results

The results for the Negative Binomial estimates are presented in Table 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 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 5 (moreover, the average marginal effects obtained from the Poisson estimation are reported in Table A4 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 sectoral 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).

[9] 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.

[10] 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 5).

In turn, sectoral price inflation increases DRWR. A 1 percentage point increase in sectoral 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 labor 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.*, 2005). This critical result is discussed at length in the next section.

Table 4. Estimates. Negative Binomial regressions.

Dependent variable: $FWCP_{it}$	Equation (7)		Equation (8)		Equation (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.

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%, as shown in Figure A5). Beyond this low density, collective bargaining coverage is also very limited (6.2% of all employees in 2014).[11]

[11] One characteristic of the Colombian labor market is the absence of a framework allowing for sectoral bargaining. To sign collective agreements, the Colombian labor code allows centralized levels of trade

Table 5. Average marginal effects. Negative Binomial regressions.

	Equation (7)		Equation (8)		Equation (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.

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 labor market segmentation between formal and informal workers.[12] This is what makes the share of informal employment ι a relevant variable.

Ahmed *et al.* (2014) and Batini and Levine (2010) show that wages in informal sectors tend to be significantly more flexible that those in formal sectors. Our result for equation (7) aligns with these studies: we find that labor 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 (8) is almost identical to the alternative one of Δp in specification (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 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

union organization (taking the form of industrial unions, federations, and cofederations). Sectoral 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).

[12] 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 A5.

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 (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.[13] However, the fact that it is not fully significant –as in contrast $\Delta w_{it}^{\min}$ turns out to be in specification (8)– indicates that it is an inferior proxy.

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 (8) and (9) are not distorting the analysis, but rather provide complementary views. In particular, the results on specification (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 (7). On the contrary, $\Delta\pi$ in specification (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 (7) and (8) point to the close association between price inflation and the wage setting system in Colombia with regard to its influence on wage or labor market rigidities. This important issue, and its policy implications, is discussed next.

4.2 Price inflation, wage setting, and labor market rigidities

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

$$\Delta NW_t = f\left(NW_t^{\min} \right), \tag{11}$$

$$\Delta NW_t^{\min} = f\left(\Delta P_{t-1} \right), \tag{12}$$

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

[13] 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.

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).[14]

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 Labor 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 (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 (12)). To illustrate this situation, Figure 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:[15]

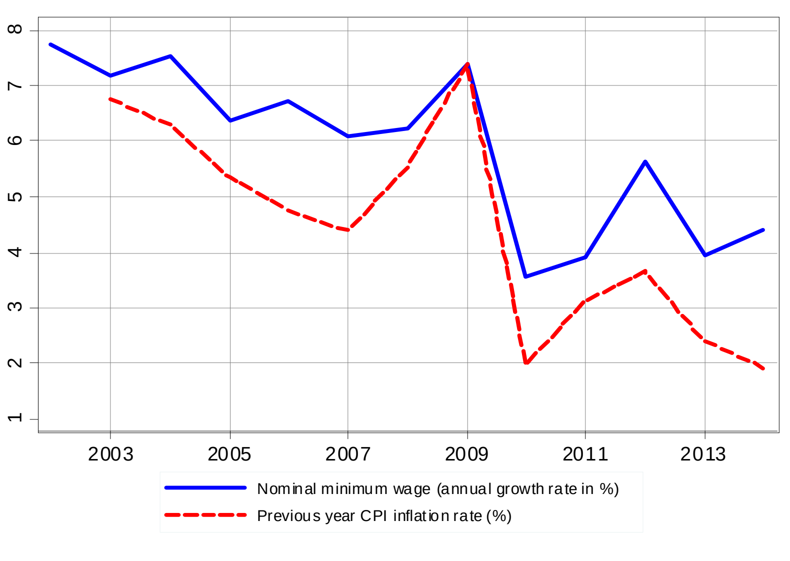
$$\Delta NW_t = f\left(\Delta P_t^e, \Delta pr_t\right). \tag{13}$$

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 (13)).

[14] 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 $\Delta RW \simeq \Delta NW - \Delta P^S$.

[15] 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. Indexation of nominal minimum wages. 2001-2014.



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

To break the connection between inflationary pressures and real wage rigidities, the current wage setting system, as summarized by equations (11) and (12), should evolve towards a system closer to the one described by equation (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.

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 labor 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 labor 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 labor 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 stimulous beyond the specific ones related to the labor market. The consequence is that this mechanism is beyond any particular labor market policy.

The second mechanism is to embrace labor 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 labor 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 labor 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 labor 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 labor 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.

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Appendix 1

Figure A1. Labor informality. 2002-2014.

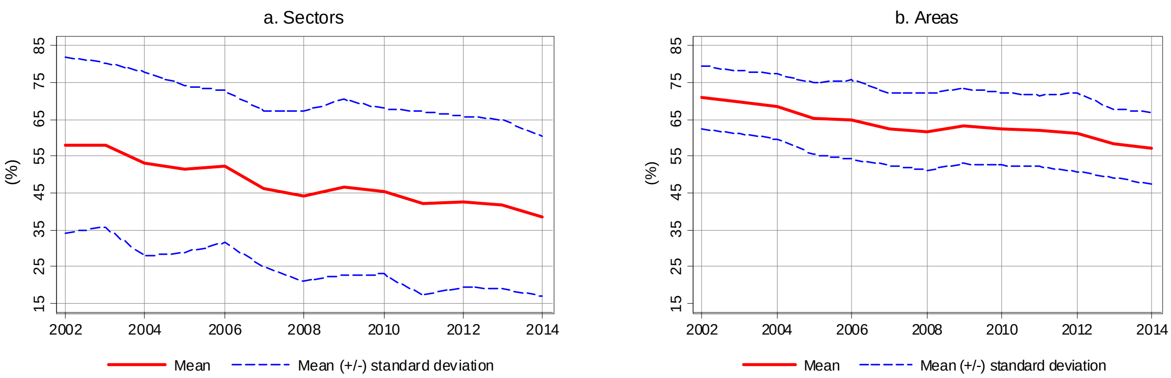


Figure A2. Trade union density. 2002-2014.

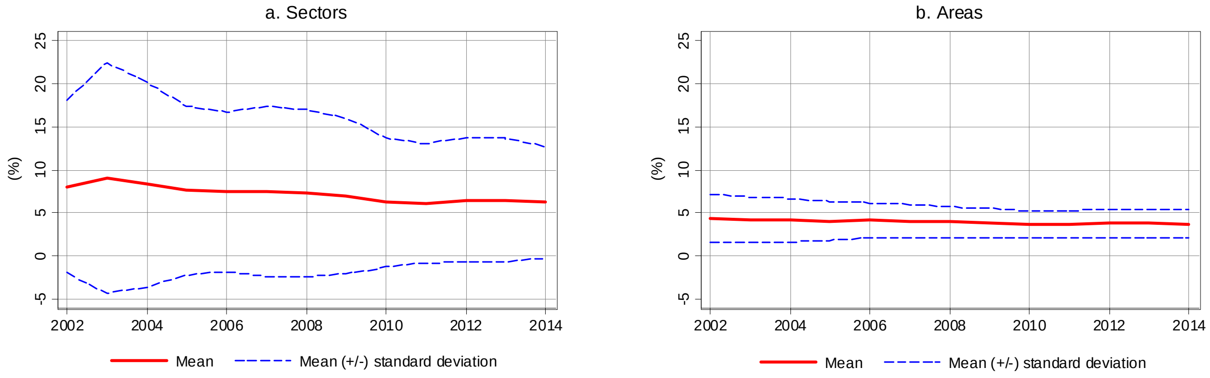


Figure A3. Real minimum wage. 2002-2014.

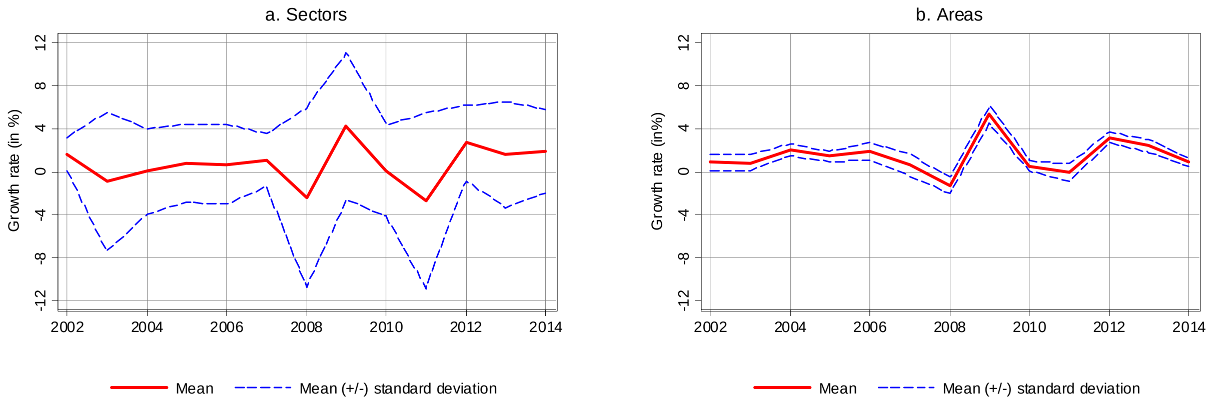
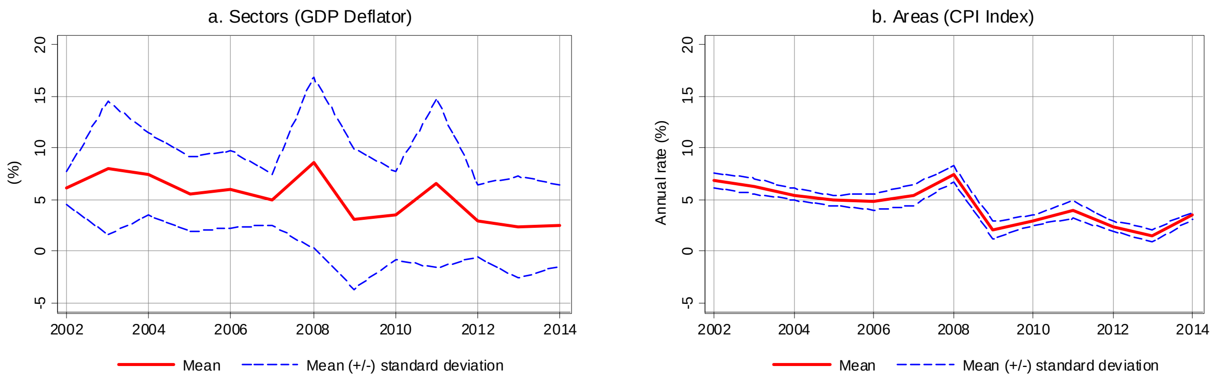


Figure A4. Inflation. 2002-2014.



Appendix 2

Table A1. Sector-year estimates.

Sector	Year	$\widetilde{q}$	q	$FWCP$	$p-value$	S
	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
Agriculture & Fishing	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
	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
Mine & Quarry Exploitation	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
	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
Manufacturing Industry	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 A1						
Sector	Year	$\widetilde{q}$	q	$FWCP$	$p-value$	S
	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
Electricity, Gas & Water	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
	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
Construction	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
	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
Commerce & Hotels	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 A1**						
Sector	Year	$\widetilde{q}$	q	$FWCP$	$p-value$	S
	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
Transport & Communication	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
	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
Financial & Insurance	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
	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
Social Services	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	5**6**.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).

Appendix 3

Table A2. Estimates. Linear regressions.

Dependent variable: $FWCP_{it}$

	Equation (7)		Equation (8)		Equation (9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.84**	-0.61*	-0.86**	-0.61*	-0.82**	-0.59**
	[−1.99]	[−1.87]	[−2.03]	[−1.88]	[−2.00]	[−1.83]
Δp_{it}	0.58**	0.08				
	[2.20]	[0.46]				
θ_{it}	-0.87***	0.39	-0.**86*****	0.4**1**	-0.85***	0.41
	[−4.68]	[0.93]	[−4.51]	[0.96]	[−4.44]	[0.94]
ι_{it-1}	-0.29***	-0.09	-0.28**	-0.09	-0.28***	-0.08
	[−4.03]	[−0.94]	[−3.90]	[−**0.92**]	[−3.92]	[−0.83]
$\Delta w_{it}^{\min}$			-0.55**	-0.07		
			[−2.01]	[−0.39]		
$\Delta \pi_{it}$					0.54*	0.00
					[1.94]	[0.02]
c	35.2***	37.1***	38.3***	37.4***	37.3***	36.8***
	[6.79]	[5.90]	[7.94]	[5.60]	[7.55]	[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 A3. Estimates. Poisson regressions.

Dependent variable $FWCP_{it}$

	Equation (7)		Equation (8)		Equation (9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.06**	-0.03*	-0.07**	-0.03*	-0.06**	-0.03*
	[−2.21]	[−1.92]	[−2.25]	[−1.93]	[−2.08]	[−1.83]
Δp_{it}	0.04**	0.01*				
	[2.52]	[1.82]				
θ_{it}	-0.05***	0.06	-0.05***	0.06	-0.05***	0.06
	[−3.65]	[1.45]	[−3.58]	[1.43]	[−3.49]	[1.30]
ι_{it-1}	-0.01***	-0.01**	-0.01***	-0.01*	-0.01***	-0.01*
	[−3.83]	[−2.06]	[−3.71]	[−1.91]	[−3.64]	[−1.76]
$\Delta w_{it}^{\min}$			-0.04**	-0.01		
			[−2.22]	[−1.56]		
$\Delta \pi_{it}$					0.04*	0.01
					[1.87]	[1.03]
c	3.54***	3.92***	3.97***	4.46***	3.72***	3.93***
	[15.45]	[13.44]	[12.36]	[11.21]	[16.09]	[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**	64.85	133.3	65.07	134.8	65.39
	(0.09)	(0.99)	(0.08)	(1.00)	(0.07)	(1.00)
P_{gof}	147.**5**	59.38	148.3	59.49	147.4	59.71
	(0.01)	(1.00)	(0.01)	(1.00)	(0.01)	(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.

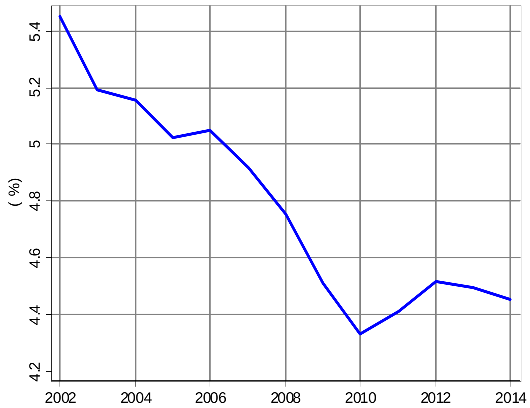
	Equation (7)		Equation (8)		Equation (9)	
	Pooled	FE	Pooled	FE	Pooled	FE
	(1)	(2)	(3)	(4)	(5)	(6)
ΔY_{it}	-0.83**	-0.42**	-0.88**	-0.43**	-0.79**	-0.41**
	[−2.27]	[−1.97]	[−2.32]	[−1.98]	[−2.11]	[−1.88]
Δp_{it}	0.56**	0.16*				
	[2.**66**]	[2.07]				
θ_{it}	-0.68***	0.84	-0.67***	0.85	-0.65**	0.81
	[−2.90]	[1.46]	[−3.27]	[1.44]	[−3.16]	[1.31]
ι_{it-1}	-0.20***	-0.13**	-0.20***	-0.13**	-0.20***	-0.12**
	[−3.61]	[−2.12]	[−3.48]	[−1.96]	[−3.40]	[−1.80]
$\Delta w_{it}^{\min}$			-0.54**	-0.14		
			[−2.32]	[−1.60]		
$\Delta \pi_{it}$					0.48*	0.11
					[1.9**2**]	[1.05]

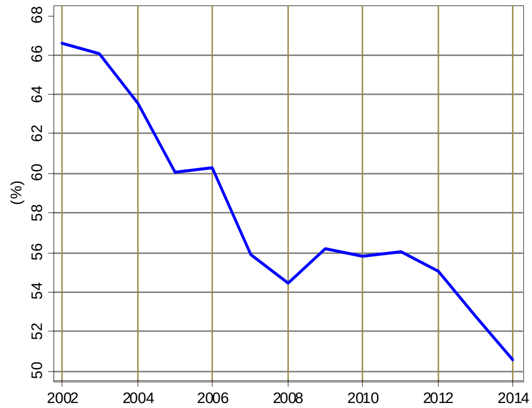
Table A4. Average marginal effects. Poisson regressions.

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

Appendix 4

Figure A5. Trade union density and labor informality. 2002-2014.





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