Understanding sectoral differences in downward real wage rigidity: workforce composition, competition, technology and institutions

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Abstract

This paper examines whether differences in wage rigidity across sectors can be explained by differences in workforce composition, competition, technology and bargaining institutions. We adopt the measure of downward real wage rigidity (DRWR) developed by Dickens and Goette (2006). The estimates are based on a large administrative matched employer-employee dataset for Belgium over the period 1990-2002. Our results indicate that DRWR is significantly higher for white-collar workers and sectors with a larger proportion of white-collar workers. DRWR is lower for older workers and for workers with higher earnings and bonuses. Sectors with larger dispersion of earnings and larger firms have lower DRWR. In addition, we find that DRWR is related to competition, technology and bargaining institutions: wages are more rigid in more competitive sectors, as well as in labour-intensive sectors and in sectors with predominant centralised wage setting at the sector level as opposed to firm-level wage agreements.

Keywords: wage rigidity, matched employer-employee data, wage bargaining institutions
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1. Introduction

Over the last decades, substantial effort has been devoted to the measurement of wage rigidity and to the understanding of its macroeconomic implications. Macroeconomic theories have put it forward as a source of resistance to wage moderation and therefore as a cause of high and persistent unemployment (see e.g. Jackman et al. (1991)). Moreover, it was suggested that rigid wages can be a cause of less frequent changes in prices of products with a high labour share (see Altissimo et al. (2006), Alvarez et al. (2006), Dhyne et al. (2006), Vermeulen et al. (2007)). In turn, price stickiness leads to higher output volatility in response to shocks, which requires stronger interest rate changes to affect inflation (see e.g. Altissimo et al. (2006)). For example, the New-Keynesian model of Blanchard and Gali (2006) shows that under wage and price rigidity, the optimal monetary policy has no longer the form of inflation targeting, and that it should rather aim to reduce but not eliminate the volatility of both inflation and unemployment. On the empirical side, there is an extensive literature measuring wage rigidity with macroeconomic data (see, for instance, Campbell (1997) and Layard et al. (1991)) and a growing volume of studies using microeconomic data. As a recent example, Dickens et al. (2006, 2007) present estimates of downward wage rigidity for a large number of countries, based on individual data.

This paper investigates the sources of wage rigidity using a large matched employer-employee dataset of individual earnings complemented with additional firm-level and sector-level data. Labour market rigidities can differ substantially across segments of the economy, i.e. there can be significant differences between groups of workers and between sectors of economic activity. Analysing differences across sectors is a natural way towards finding relevant factors of wage rigidity. Wage bargaining in Belgium is organised primarily at the sector level. Other relevant variables such as labour intensity or competition have a clear sector dimension.

Previous investigations of the factors underlying wage rigidity are rather scarce. Using a cross-country analysis, Dickens et al. (2006) point to national labour market institutions as factors explaining differences in the level of downward wage rigidity measured at the microeconomic level. More specifically, they show that unionisation and collective bargaining coverage at the country level are positively related to wage rigidity. Clar et al. (2007) examine the relations between national labour market features and macroeconomic estimates of the response of real wages to unemployment for OECD countries. They find that union density, centralisation of wage bargaining and employment protection legislation are negatively related to real wage flexibility. Coordination of wage bargaining, which allows for internalisation of the extern effects of wage changes on the economy, makes wages more responsive to labour market conditions and therefore increases real wage flexibility.

Differences in wage rigidity according to worker types were pointed out by Campbell (1997). He finds that wage flexibility, defined as the responsiveness of occupational wages to aggregate unemployment, is higher for blue-collar workers than for white-collar workers. Du Caju et al. (2007) provide estimates of downward wage rigidity using microeconomic data for Belgium and the methodology developed by the International Wage Flexibility Project (IWFP), see Dickens and

Goette (2006). They highlight differences across occupation, age, wage level, as well as firm size, but provide no formal statistical tests of these differences.

The literature on wage rigidity involving a sectoral dimension is rather limited. Using macroeconomic industry-level data of OECD countries, Holden and Wulfsberg (2007) find that downward nominal wage rigidity is higher in cases where employment protection legislation is stricter, union density is higher and unemployment is lower. Asking professional wage setters about the reasons for wage rigidity, Agell and Bennmarker (2007) find that the determinants of wages differ across segments of the labour market Their results suggest that the effects of firms' profits on wages are important in manufacturing and skilled services, and less important in unskilled services and in the public sector. They interpret this as an indication of incumbent workers' bargaining power and therefore as a possible source of rigidity. Campbell (1989, 1991) provides measures of wage flexibility for the United States, Canada and France, based on the response of sector-level wages to the aggregate unemployment rate and to sector product demand. Among others, he finds that sectors with a larger percentage of blue-collar workers are characterised by a higher degree of wage flexibility. His results for the United States also indicate that wage flexibility is lower in more capital intensive sectors. Finally, he finds no robust evidence that unionisation reduces wage flexibility.

In sum, the literature identifies several variables driving wage rigidity, such as those related to workers (e.g. occupation), firm's characteristics (size, sector), production technology (capital intensity), or labour market institutions (for example unionisation and wage bargaining). However, none of the studies mentioned above provides statistical tests of differences between the categories after controlling for the impact of labour force composition. The composition effects might be relevant especially at the sector level, as some sectors demand very specific labour with respect to skills. For instance, the construction sector employs a disproportional number of blue-collar workers.

The aim of this paper is to determine the relevant factors explaining differences in wage rigidity across sectors. We evaluate the importance of the labour force composition, sector-specific characteristics such as firm size, capital intensity and competition, and sector-specific institutional features related to wage bargaining. We rely on a large microeconomic data set on individual earnings from administrative sources for Belgium over the period 1990-2002. Du Caju et al. (2007) use the same data set and show that downward nominal rigidity (DNWR) is nearly absent during this period in Belgium, a country with full automatic indexation of wages. For this reason, we focus on downward real wage rigidity (DRWR) which we estimate using the procedure described by Dickens and Goette (2006).

The paper is organised as follows. Section 2 describes the dataset, relevant institutional features of the Belgian labour market and sector characteristics, as well as the methodology. Results are reported in Section 3. First, we highlight differences across worker types and shed light on the importance of labour force composition effects. Next, we investigate additional factors explaining differences in downward real wage rigidity between sectors. Section 4 concludes.

Appendix A defines the variables used in the paper while Appendix B provides robustness tests with respect to outliers and alternative definitions of variables.

2. Institutional background, data, and methodology

2.1. Institutional background

Some important institutional features of the labour market affect individual wages in Belgium, such as indexation and sector-level collective agreements, which can possibly be supplemented with agreements concluded at the firm level. These features explain why Belgium is characterised as a country with high real wage rigidity (see Dickens et al. (2006) or Du Caju et al. (2007)). We briefly describe these characteristics of the Belgian labour market. Firstly, practically all employees' gross wages are linked to an index of consumer prices through an automatic indexation mechanism.

Secondly, as in many countries, including Belgium, wages are to an important extent determined at the sector level. The level of gross wages is mainly determined through agreements concluded in joint committees organised per sector of economic activity.[1] The outcome of these sector-specific negotiations cannot undercut the legally determined guaranteed minimum wage.[2] In many sectors, pay scales are determined for blue-collar and white-collar workers separately. In the joint committees for blue-collar workers, pay scales are primarily defined in relation to the job description. Variations depending on age or length of service are not common. For white-collar workers, the pay scale usually varies not only by category, but also depending on age or tenure.[3] The joint committees at the sector level are also the main bargaining unit for the negotiations on collective wage increases. Quite often, these collective wage increases are defined as a rise in absolute terms of the (sometimes only minimum) pay scales, meaning that employees with wages above the scale can obtain a lower percentage collective wage increase. For the period under review, a lot of employees receive automatic wage increases, negotiated in sector-level collective agreements. These are a function of age and, to a lesser extent, tenure.

In addition, firm-level agreements can complement sector-level agreements. According to the favourability principle in hierarchical wage bargaining, the negotiated wages in these firm-level agreements cannot be below the sectoral agreements.[4] Decentralised wage setting through single-employer (SE) wage agreements is very common in the chemicals and transport equipment industries, and very rare in the construction and business services sectors. Also firm-level agreements are more common in large firms, with stronger union representation, than in smaller

1 They are called joint committees ('commissions paritaires'), because employers and employees share an equal representation in them. As the notion of economic sector is sometimes very narrowly defined, the number of joint committees exceeds 100.

2 The actual minimum wage in almost every joint committee exceeds the legally guaranteed minimum, with some exceptions for workers less than 21 years old.

3 During the period under review, age-related pay scales were not against European anti-discrimination rules and were applicable to the majority of Belgian white-collar workers.

4 Opt-out clauses are possible but are very rare.

firms. Note that union representation is compulsory in firms with 50 employees or more. SE collective wage agreements allow to better account for firm-specific features in the wage setting. Typically in Belgium, firms that do not have a firm-level agreement stick to the sector agreement. Firms with an SE agreement generally pay more and have a more dispersed earnings structure. This provides them with a wage cushion above the sector minima, creating room of manoeuvre for wage adjustment. From the individual data of the Belgian Structure of Earnings Survey (SES), it turns out that firms with SE agreements for blue-collar and white-collar workers pay earnings 12% higher and bonuses 53% higher. Further, the standard deviation of earnings is 2% larger, and that of bonuses is 16% larger, in firms with SE agreements compared to firms with no SE agreements.

In light of these institutional features, and consistent with labour market theories, Du Caju et al. (2007) find that DRWR is more important for white-collar workers than for blue-collar workers, and that wage rigidity decreases with age and with the wage level except for the lowest wages. Further, if institutions matter for wage setting practices, we expect to find differences in downward real wage rigidity across sectors, according to firm size and the importance of firm-level agreements.

2.2. Data

To measure downward real wage rigidity, we rely on an administrative database on individual labour earnings for Belgium, collected by the social security system. The data contain information on annual gross earnings (including bonuses and compensation for overtime hours), annual working days, age, sex and occupation category (blue-collar or white-collar). The data set contains a sample of around one-third of workers in the private sector and covers the period 1990-2002. It includes all persons that were born between the 5th and the 15th day of any month, except those employed by firms with less than 5 employees or by natural persons. The data set covers all sectors of activity including services. We focus on in firms active in branches with NACE codes from D to K, i.e. we exclude agriculture, extraction and non commercial services.

We restrict the sample to workers above the legal minimum age of obligatory schooling and below the retirement age, i.e. men between 18 and 64 and women between 18 and 59. We also exclude earnings below the legal minimum wage and we drop the same number of observations from the upper tail of the distribution. Finally, we restrict the sample to full-time permanent job-stayers. These are defined as working at least 11 months for the same employer over two consecutive years. Thereby we allow permanent workers to have at most one month of sick leave (or other "abnormal" days off) per year. We refer to Du Caju et al. (2007) for more details on the data.

It is important to note that annual earnings include variable compensation components, such as bonuses, premia, and overtime hours. Not all of these are subject to automatic increases such as indexation and collectively bargained increases. Therefore, annual earnings may be more flexible than the base wage. Further, because the importance of extra wage components varies across workers, firms and sectors, these may explain differences of wage rigidity across sectors.

The individual earnings data are complemented with information from firms' balance sheets.[5] Also, we use individual data from the Belgian Structure of Earnings Survey (SES), for the 1999, 2000, 2001 and 2002 waves.

Individual annual earnings data are used to estimate downward real wage rigidity by occupation, age category and sector. These rigidity measures are then related to three types of variables. The first set consists of variables defined at the category level, i.e. by occupation, age and sector. This is the case of the median level of earnings, computed from the individual earnings dataset, and the median level of bonuses, as reported in the four SES waves between 1999 and 2002. Note that this variable includes compensation for overtime hours. We also consider the earnings spread, defined as the ratio of the 80th to 20th percentiles for each branch, occupation category and year, computed from individual earnings data.

The second set of variables are sector-specific. From firms' balance sheets, we define the median firm size, as the number of employees, and the capital-labour ratio, as the median within the sector of firm-specific capital-labour ratios. Moreover, we estimate a measure of competition recently proposed by Boone et al. (2007), i.e. the elasticity of a firm's profits with respect to its marginal costs (profit elasticity). The intuition behind profit elasticity is that firms in less competitive sectors are not pure price takers and hence a given percentage increase in costs can be accommodated by a price increase, and therefore leads to a smaller fall in profits. The profit elasticity is thus larger for more competitive firms. Using firm level data for each branch, we regress log profits on log marginal costs (for more details on theoretical derivation as well as its relation to other measures of competition, see Appendix A). As a robustness test, we also consider two alternative measures of competition: the Herfindahl index which measures concentration within the sector, and sector-specific estimates of the price cost margin defined by Christopoulou and Vermeulen (2007). As argued in Boone et al. (2007), the three measures would correctly capture strengthened competition resulting from a fall in entry costs and a consequent increase in the number of firms. However, the Herfindahl index fails to capture rises in competition that cause inefficient firms to close. In such a case, the concentration of the industry increases, as well as the Herfindahl index. However, it would be misleading to interpret this as a fall in competition. Further, they argue that empirical measures of the price-cost margin, such as the ratio of profits to sales, may be less suited in highly concentrated markets. The estimates of Christopoulou and Vermeulen (2007) rely on the estimation of structural equations, but they are time-invariant. Because the profit-elasticity overcomes the drawback of the other measures and is time varying, it is our preferred measure of competition.

Finally, the third type of variable refers to sectoral wage bargaining practices, i.e. the coverage by collective wage agreements at the sector or firm level. For Europe in general and Belgium in particular this provides a much better indicator of union bargaining power than union membership for example. The reason is that, unlike in the US, wage agreements are negotiated between employer's representative and workers representatives, but apply to all workers, independent of

[5] Annual accounts data in Belgium is (nearly) exhaustive for credit institutions and is subject to substantial consistency checks.

whether they are unionised or not. As explained above, sector-level multi-employer agreements apply generally in Belgium. As an indicator of decentralised wage setting, we calculate the average proportion over time (1999-2002) of workers covered by a single-employer (SE) wage agreement, from the SES data set. Such agreements are expected to provide the firm with more flexibility as compared to the multi-employers (ME) agreements. Moreover, firms applying SE agreements are more likely to negotiate individually wage conditions that are more favourable than those of the collective agreement for part of their workforce. Appendix A gives more information on data sources and definitions.

Tables 1A and 1B provide information on sectoral differences in the variables of interest including the estimated DRWR per sector. The tables highlight important differences across sectors. For example, the proportion of blue-collar workers is very large in the construction and other manufacturing sectors, and very low in financial and business services. The level of earnings, bonuses, as well as earnings dispersion is particularly high in the chemical industry and in financial services. At the other extreme, earnings, bonuses and earnings dispersion are the lowest in the construction, and hotels and restaurants sectors. The textile industry is characterised by low earnings and bonuses but by a high earnings spread.

Table 1A - DRWR, labour force composition and wages by sector - averages over time

	DRWR[a]	Pct of white-collars	Average Age	Median earnings[b]	Average bonus[c]	Earning spread
food	0.526	41.66	36.15	72.41	2671	1.661
textile	0.600	33.62	36.75	58.35	968	1.850
wood and paper	0.648	43.54	36.84	79.10	2326	1.721
chemicals	0.467	60.27	37.87	101.34	4122	1.930
non metal	0.483	34.39	38.86	78.24	2292	1.581
metal	0.553	35.08	38.44	81.63	2509	1.649
machinery and equip.	0.618	47.51	37.46	81.11	2930	1.676
transport equipment	0.517	28.27	37.55	89.84	2428	1.498
other manufacturing	0.681	22.93	36.81	60.73	1327	1.422
construction	0.801	21.29	36.77	68.86	875	1.356
trade	0.648	72.58	36.20	71.02	3073	1.868
hotels and restaurants	0.590	36.87	34.12	53.67	354	1.517
transport and storage	0.354	46.14	37.41	74.29	1772	1.619
financial services	0.627	97.79	38.44	104.34	6043	1.969
business services	0.668	83.13	35.13	83.64	3354	2.031
Mean	0.585	47.01	36.99	77.24	2469	1.69
Standard deviation	0.107	22.25	1.26	14.34	1415	0.20

Notes: [a] DRWR estimated per branch using the IWFP procedure
[b] Gross total daily earnings in euros.
[c] Annual bonuses in euros.

The figures reported in Table 1B show that chemicals, non-metal manufacturing, transport storage and business services are capital intensive sectors while construction is the most labour intensive one. Firms are larger in chemicals, textile and transport equipment industries, and smaller in services. According to the profit elasticity competition is fiercer in other manufacturing, transport equipment and trade and the low in businessand financial services and in transport and storage. Note that alternative indicators of competition are not always consistent with the results using the profit elasticity. We therefore evaluate the robustness of our results with respect to the choice of competition indicator in Appendix. Finally, note that decentralised bargaining through SE agreements is much more widespread in the chemical industry and is essentially absent in construction and business services, i.e. in sectors with centralised bargaining.

Table 1B - Data by sector

	Median Firm size[a]	Median K/L[b]	Profit elasticity[c]	SE coverage of blue-collars[d]
food	6	17.8	7.957	35.50
textile	11	11.1	9.514	11.34
wood and paper	4	18.5	7.436	27.51
chemicals	13	20.7	7.809	54.80
non metal	7	20.5	8.514	37.49
metal	7	13.8	8.119	38.65
machinery and equip.	7	10.6	8.642	30.54
transport equipment	11	11.4	9.877	42.89
other manufacturing	5	12.3	9.159	14.44
construction	4	10.0	8.432	2.38
trade	3	14.5	9.821	15.10
hotels and restaurants	3	11.8	8.189	10.15
transport and storage	5	22.4	5.784	20.58
financial services	2	16.2	5.473	34.02
business services	2	19.5	6.007	3.61
Mean	6.0	15.4	8.049	25.27
Standard deviation	3.4	4.20	1.388	15.54

Notes: [a] Number of employees.
[b] Measured in thousands of euros.
[c] Values calculated for each branch and year. The table reports median over years.
[d] Percentage of blue-collar workers employed in firms with single-employer agreement.

2.3. Methodology

Measures of downward wage rigidity rest on the idea of an asymmetric behaviour of wage changes in case of wage increases versus the case of wage falls. Using microeconomic data, two strands of the literature have implemented this concept. The first relies on the distribution of wage changes. It rests on the assumption that in the absence of wage rigidity, the distribution of wage changes should be symmetric. One approach, developed by Kahn (1997) consists in testing whether the size of the bins, at a given distance from the median, is smaller when they correspond to wage decreases, than when they refer to wage increases. Another approach builds on the assumption that when downward wage rigidity binds, the distribution exhibits asymmetry and a spike at the reference point (zero in case of nominal rigidity, expected inflation in case of real rigidity). For example, Card and Hyslop (1997) and more recently Dickens and Goette (2006) develop methods based on these features.

The second strand of the literature is based on asymmetry in the sensitivity of wages to relevant variables or shocks (see for example Altonji and Devereux (1999) or Biscourp et al. (1995)). The method is appealing because it takes into account the motives to cut wages. The major drawback is that this method is very demanding in terms of data as it requires data on relevant workers' and firms' characteristics. The methodology of Dickens and Goette (2006) requires only information on wage changes. However, because it is based on the estimation of the distribution of wage changes, it demands datasets with a large cross-section dimension.

Another distinctive advantage of the methodology of Dickens and Goette (2006) is that the procedure jointly estimates nominal and real downward wage rigidity, together with the average reference point for real wage rigidity (expected inflation or bargaining focal point). In addition the factual distribution of wage changes is corrected for measurement errors. The major drawback is that identification of DRWR and DNWR becomes an issue in years with very low inflation, where the reference point for DRWR comes very close to zero, i.e. the reference point for DNWR. Lastly, because the analysis is performed at the individual-level, earnings changes can only be sensibly computed for job stayers.[6]

As in Du Caju et al. (2007), we adopt the approach of Dickens and Goette (2006). The measure of DRWR attempts to capture the fraction of workers who would not receive a real wage cut when they were scheduled for one, no matter what the reason for the wage cut. The estimation consists of two steps. We start with the empirical distribution of wage changes after cleaning the data set. Next, the procedure corrects for measurement error and the corrected distribution is dubbed "true distribution". In the second step, we fit a model of wage changes to the corrected distribution – called "theoretical distribution" - and estimate jointly downward nominal and real rigidity and the parameters of the distribution of reference points across individuals in each year using GMM. It is assumed that the distribution of wage changes under flexibility (called "notional distribution") follows a symmetric two-sided Weibull distribution. Each person's notional wage change is a draw from the notional distribution. A fraction of the population is potentially subject to DRWR. If their notional wage change falls below their reference point, π^e, they will receive a wage change equal to π^e instead of the notional wage change. Another portion of the population is assumed to be potentially subject to DNWR. Such individuals with negative notional wage change, who are not subject to DRWR, will receive a wage freeze instead of wage cut.[7] As an illustration, figures A1 to A3 in the Appendix A show the different distributions of earnings changes for three sectors in 2002.

With respect to the method based on the sensitivity of wages to given variables, note that the resistance to wage cuts following negative changes in relevant variables, implies that the distribution of wage changes is asymmetric in case negative shocks occur. And if one observes a resistance to wage cuts whatever the reason for the wage cut (so following the definition of the distribution-based approach) one should *a fortiori* find that wages do not respond to negative signals. The difference with respect to estimates of the sensitivity of wages to shocks at a more

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6 Analysing the wage change of job movers does not inform about wage rigidity because workers may move from a high-pay firm to a low-pay firm; but it is unclear whether it is related to a particular economic event the firms face. In addition, Du Caju et al. (2007) and Fuss (2007) find no evidence that firms use entrants to adjust their average wage cost. They report that, on average, entrants do not earn less than incumbents.

7 We make the same choice of specification and parameters as in Du Caju et al. (2007). More specifically, we allow the mean of π^e to be unrestricted in the 0-4 percent band and its variance to range from 4E-06 to 3.6E-05.

aggregated level is larger. Consider for example an adverse shock. Firms may reduce their average wage bill by changing the composition of their workforce, *even if* the wages of job stayers do not fall. In this case the distribution of earnings changes will exhibit no wage cuts, while the estimation of wage sensitivity will suggest that the average wage has fallen.

We estimate DRWR year by year for 90 categories defined as the combination of 15 branches, 2 occupation categories and 3 age categories. We consider two occupational categories (blue-collar and white-collar workers) and three age categories (18-24 years, 25-44 years, and 45 years or more). Branches have been defined as follows: (1) food (food products, beverages and tobacco), (2) textile (textiles, textile products, leather and footwear), (3) wood and paper (wood and products of wood and cork, and pulp, paper, paper products, printing and publishing), (4) chemicals (chemical, rubber, plastics and fuel products), (5) non metal (other non metallic mineral products), (6) metal (basic metals and fabricated metals products), (7) machinery and equipment, (8) transport equipment, (9) other manufacturing (manufacturing n.e.c., recycling), (10) construction, (11) trade (wholesale and retail trade, repair), (12) hotels and restaurants, (13) transport and storage, (14) financial services (financial intermediation), (15) business services (real estate, renting and business activities). We exclude categories for which the number of observations of earnings changes is below 2000. Also, we do not consider energy (electricity, gas and water supply) and transport and communication (post and telecommunication) because either the estimates of DRWR are not reliable or the observations appear to be outliers in the regressions estimated below.[8] We also exclude estimates of DRWR that hit the bound of zero or one as unreliable (142 occurrences).

Du Caju et al. (2007) provide evidence on differences in DRWR across categories of workers (occupation, age and wage level) and firm types (size and quit rate), but they simply report estimates of DRWR. In order to test more formally for differences across these categories, we first estimate DRWR for each branch, occupation and age category, then regress them on dummies for age and occupation and examine their significance in order to test differences across worker types. Second, in order to understand sector differences, we introduce additional variables, controlling for age and occupation. Using a finer classification, for instance considering each branch, occupation, age category, wage category and firm size is not feasible because this would leave too few observations per category to estimate DRWR. Rather we test the significance of the median wage level per occupation, age and sector, or the median firm size by branch. Finally, we augment the regression with a set of branch-specific variables like capital intensity, competition, and collective wage agreement coverage. So, we examine the role of branch-specific variables controlling for worker composition effects. The motivation for this exercise is twofold. First, we test the importance of worker heterogeneity in explaining differences in DRWR. Second, we examine differences across branches, after controlling for composition effects.

8 For energy, the highest standard deviation of the estimates of DRWR over time reaches 0.33, while the average standard deviation per sector equals to 0.15 (excluding energy).

3. Results

In Section 3.1, we test the importance of composition effects for DRWR. For this purpose, we first fit a simple OLS model with wage rigidity estimates over sector, age and occupation, $DRWR_{kajt}$, as dependent variable and dummy variables for each occupation, age and year as independent variables. We denote occupation by *k* (blue-collar or white-collar), age by *a*, sector by *j* and year by *t*. We are essentially interested in explaining the differences across sectors and therefore we include time dummies but no sector-specific fixed effects. Adding branch dummies to this model allows us to test whether there remain sectoral differences in wage rigidity beyond those due to the composition of the labour force. In Section 3.2, we include explanatory variables that can be suggested as determinants of DRWR. Finally, we combine the variables with the highest explanatory power into a single model and evaluate the contribution of each variable to DRWR in each sector.

3.1. Labour force composition effects

Regressing DRWR on occupation, age and year dummies indicates that white-collar workers have more rigid earnings than blue-collar workers, as shown in Model 1 in Table 2. More specifically, DRWR of white-collar workers is *ceteris paribus* higher by 0.07 and the difference is highly significant. This result is consistent with the shirking model of Shapiro and Stiglitz (1984) and with the turnover model of Stiglitz (1974). The idea from these theoretical models is that firms may be less inclined to cut wages of white-collar workers because they are more difficult to replace and to monitor, and therefore are more likely to shirk. In addition, white-collar workers in Belgium obtain automatic wage increases with age or tenure, while it is rarely the case for blue-collar workers. This makes white-collar workers less likely to experience real wage decreases. Campbell (1997) for the US and Du Caju et al. (2007) for Belgium also find higher wage rigidity for white-collar than for blue-collar workers.

Model 1 in Table 2 further highlights that real wage rigidity is the highest for workers less than 25 years old. DRWR is lower by 0.05 and 0.03 for workers in the age categories 25-44 years and above 44 years, respectively. Results based on the F-test suggest that DRWR for the two age categories are statistically identical.[9] Due to lower precision of the estimate, the difference between the old and the middle age category is not statistically significant in Model 1. Du Caju et al. (2007) also find that real rigidity is the highest for workers less than 25 years old. The result may be explained by the shirking model and the adverse selection model of Weiss (1980) applied to quits. It predicts that younger workers are more likely to quit or shirk, when their earnings increases are below their expected bargaining reference point, because the cost of job loss is smaller for them than for older workers, i.e. finding a job is more difficult for older workers. Furthermore, automatic

9 The F-statistics of the test for equality of the two coefficients is 0.91 with p-value of 0.34.

tenure and age related wage increases are more prominent for younger workers, while extra wage components are smaller leading to less flexible earnings.

Table 2: OLS estimates of DRWR per year, occupation, age category and sector

	Model 1 Dep. var. $DRWR_{kajt}$	Model 2 Dep. var. $DRWR_{kajt}$	Model 3 Dep. var. $DRWR_{kajt}$	Model 2 rank of sector dummies	$DRWR_{jt}$ rank of sectors
D white-collar	0.070***	0.084***	0.084***		
	(4.05)	(5.11)	(5.11)		
D age:25-44	-0.051**	-0.059***	-0.059***		
	(-2.29)	(-2.74)	(-2.74)		
D age:45+	-0.032	-0.047**	-0.047**		
	(-1.44)	(-2.20)	(-2.20)		
D Food		0.429***	-0.050	13	11
		(9.22)	(-1.14)		
D Textile		0.479***		8	8
		(10.98)			
D Wood and paper		0.619***	0.140***	3	4
		(14.40)	(3.22)		
D Chemicals		0.404***	-0.075*	14	14
		(9.71)	(-1.85)		
D Non-metal		0.533***	0.054	5	13
		(11.75)	(1.25)		
D Metal		0.447***	-0.031	11	10
		(10.18)	(-0.75)		
D Machinery & equipment		0.459***	-0.020	9	7
		(10.33)	(-0.48)		
D Transport equipment		0.529***	0.051	6	12
		(11.38)	(1.15)		
D Other manu-facturing		0.656***	0.177***	2	2
		(13.23)	(3.80)		
D Construction		0.695***	0.216***	1	1
		(15.86)	(5.16)		
D Trade		0.453***	-0.026	10	5
		(11.03)	(-0.64)		
D Hotels and restaurants		0.574***	0.095**	4	9
		(12.40)	(2.06)		
D Transport & storage		0.342***	-0.137***	15	15
		(8.43)	(-3.42)		
D Financial services		0.443***	-0.036	12	6
		(8.64)	(-0.71)		
D Business services		0.484***	0.006	7	3
		(11.63)	(0.13)		
Constant	0.488***		0.479***		
	(13.88)		(10.98)		
Year dummies	yes	yes	yes		
R^2_{adj}	0.056	0.199	0.199		
Number of obs.	758	758	758		
F test for sector dummies [p-val.]		10.51 [0.000]	11.30 [0.000]		

Note: The last column provides the ranking of $DRWR_{jt}$ obtained directly from the values estimated by the IWFP procedure per sector.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
t-statistics in brackets.

To obtain estimates of DRWR for sectors after controlling for workforce composition effects, we add sector dummies to Model 1, see Model 2 in Table 2. Once branch dummies are included, the adjusted R^2 increases from 0.056 to 0.199, suggesting that sector specific factors contribute to explaining DRWR beyond the effects of occupation and age. In model 3 we re-estimate the equation with sector dummies, taking textiles as the reference sector. The textiles industry was chosen because it has the median level of DRWR. The F-test for no significance of the differences across branches clearly rejects the null hypothesis.[10] In construction, other manufacturing and wood and paper industries, DRWR is significantly higher than in textiles; DRWR is significantly lower in chemicals and transport and storage. Additional factors may therefore explain sector differences in DRWR, as we turn to them in Section 3.2. Table 1 shows that the standard deviation of DRWR across sectors is 0.11, around a mean of 0.64. In the next section we attempt to explain this heterogeneity across sectors after controlling for age and occupation (and year dummies). The standard deviation of sector dummies in Table 2 evaluates this at 0.15.

The last two columns in Table 2 provide a ranking of relative DRWR for each industry with rank 1 assigned to the industry with the highest DRWR. First the ranking of DRWR after controlling for age and occupation is reported; then the last column displays the ranking of average DRWR directly obtained from the IWFP procedure (i.e. without controlling for composition effects). The highest DRWR is obtained unanimously for the following sectors: construction, other manufacturing and wood and paper. Sectors with the lowest degree of DRWR are transport and storage and chemicals. More strikingly, comparison of the last two columns of Table 2 documents that occupation can explain differences in DRWR across sectors. Indeed, non-metal manufacturing, transport equipment, hotels and restaurants, are ranked as sectors with much lower DRWR after controlling for occupation and age. Conversely, trade, financial services and business services are ranked at the upper end of the DRWR range after controlling for worker characteristics. According to the figures reported in Table 1, this is due to a very high proportion of blue-collar workers in the first set of sectors and a very high proportion of white-collar workers in the second set. Because blue-collar workers have a lower DRWR, rigidity after controlling for occupation increases relatively in sectors with a high proportion of blue-collar workers and vice versa.

In Model 2, both age dummies are statistically significant. Because we control for sector-specific fixed effects through the sector dummies, the coefficients capture solely the effect of an increasing age of an individual within any sector and occupation. Model 1 combines both the individual effect and the variation in age across sectors. Using the same groups of industries as above, we can see that differences in age do not always help to explain the difference in the rankings. Non-metal and transport equipment are characterised by older workers on average. Because DRWR decreases with age, this explains why DRWR increases after controlling for age. However, this is also the case for financial services, for which DRWR after controlling for age and occupation decreases. Conversely, trade and business services are characterised by a large fraction of young workers. Again this explains why DRWR after controlling for age and occupation decreases. However, the opposite is true for financial services, which employ relatively older

10 Considering other sectors as reference lead to the same conclusion. For example, taking food, construction and transport and storage as references, the value of the F-test is 10.94, 7.49 and 8.57, respectively.

workers but show a decrease in DRWR after the control. In sum, age explains differences in DRWR across workers whatever the sector of economic activity in which they work, rather than across sectors. Indeed, as shown in Table 1, there is much less heterogeneity in the age structure across sectors than in the occupational structure.

All in all, the results of this section suggest that age and occupation explain differences across worker types and sectors of economic activity. In addition to occupation, other sector-specific factors may explain differences in DRWR across branches, a question we address in the next section.

3.2. Explaining sector differences

In this section, we augment Model 1 introduced in Table 2 with candidate variables that may explain sector differences in DRWR, and therefore we drop sector dummies. Also some of the variables, such as the capital-labour ratio, bonuses and the firm-level collective agreement coverage, are time invariant.

First, we extend Model 1 with the median earnings per day calculated by branch, year, occupation and age category. The results in column (1) of Table 3 indicate that DRWR is negatively related to earnings after controlling for age and occupational status. A one standard deviation increase in median daily earnings (equivalent to an increase of 22.5 euros) reduces DRWR by 0.04. Du Caju et al. (2007) also report decreasing DRWR with earnings level, except for very low earnings. One explanation is that as income grows, a larger fraction of earnings is attributable to extra wage components, such as bonuses and premiums. Because these can be easily cut, higher earnings are less rigid. Another explanation is that low wages are close to the institutional minimum wage or to sectoral pay scales and therefore cannot be reduced freely. Lastly, wage levels are typically higher in firms with SE wage agreements as opposed to sector-level or multi-employer (ME) agreements.[11] As argued by Cardoso and Portugal (2005), a higher average wage and more wage dispersion in firms with SE agreements provides employers with a flexible wage cushion above the sectoral minima. This yields more flexibility in wage adjustment, hence lowers downward real wage rigidity.

In order to investigate the argument that variable components of earnings provide flexibility in earnings, we consider average bonuses (including compensation for overtime hours) per sector as an additional explanatory variable. Column (2) in Table 3 shows that the coefficient is significant and negative, confirming that sectors with higher bonuses and overtime compensation experience on average less downward rigidity of earnings.

In the same vein, higher earnings dispersion is related to lower DRWR, as shown in column (3) in Table 3. The coefficient on the earnings spread is negative and significant. Based on this result, DRWR in construction (sector with the lowest earnings spread, 1.36, see Table 1A) should be 0.09

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[11] See also evidence in Card and de la Rica (2006), Cardoso and Portugal (2005), Dell' Aringa and Lucifora (1994), Gerlach and Stephan (2006), Palenzuela and Jimeno (1996) and Rycx (2003) for Belgium. See also the numbers reported in the Section on institutional features.

higher than the average (the average earnings spread being 1.69). As argued above, the intuition behind the result may again be related to the fact that larger wage dispersion is a sign of firms applying a more flexible wage-setting policy. Again, wage dispersion may be higher in sectors with larger proportion of firms applying SE agreements, because these can better take into account firm-specific characteristics (e.g. rents) in the determination of wages.[12]

Table 3: Additional factors affecting DRWR

Dep. v. $DRWR_{kajt}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
D $white\text{-}collar_{kajt}$	0.109***	0.133***	0.085***	0.067***	0.080***	0.079***	0.071***
	(4.98)	(5.12)	(5.02)	(3.87)	(4.72)	(4.60)	(4.15)
D $age{:}25\text{-}44_{kajt}$	-0.022	-0.021	-0.068***	-0.044**	-0.062***	-0.061***	-0.046**
	(-0.91)	(-0.89)	(-3.10)	(-1.99)	(-2.84)	(-2.75)	(-2.07)
D $age{:}45+_{kajt}$	0.018	0.006	-0.052**	-0.029	-0.049**	-0.042*	-0.034
	(0.62)	(0.25)	(-2.37)	(-1.31)	(-2.23)	(-1.89)	(-1.53)
Median $earnings_{kajt}^{\dagger}$	-1.790***						
	(-2.88)						
Median $bonus_{kaj}^{\dagger}$		-0.030***					
		(-3.23)					
Earnings $spread_{jt}$			-0.271***				
			(-6.71)				
Median firm $size_{jt}^{\ddagger}$				-6.400***			
				(-2.66)			
Median capital-labour $ratio_{j}^{\dagger}$					-0.012***		
					(-5.96)		
Profit $elasticity_{jt}$						0.024***	
						(3.89)	
SE agreement $coverage_{kj}$							-0.003***
							(-4.75)
Constant	0.562***	0.493***	0.948***	0.522***	0.691***	0.330***	0.550***
	(12.94)	(14.11)	(12.38)	(14.01)	(14.28)	(6.15)	(14.85)
Year dummies	yes	yes	yes	yes	yes	yes	yes
Sector dummies	no	no	no	no	no	no	no
R^2_{adj}	0.070	0.068	0.109	0.063	0.098	0.073	0.082
Observations	758	758	758	758	758	758	758

Notes: † measured in thousands of euros.
‡ measured in thousands of employees.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
t-statistics in parentheses.

Next, we examine whether the median firm size within the sector affects wage rigidity. It has often been argued that larger firms have a different pay policy. Among others, they may develop a more complex wage structure, offer higher but also more dispersed wages. This is also the case in our sample.[13] Earnings decreases in smaller firms are more likely to be bound by minima collectively agreed outside the firm. On the contrary, in larger firms the wage cushion (above the sector-level agreement) provides some margin for earnings cuts. Finally, higher earnings may

12 This is supported by the findings of Card en de la Rica (2006), Cardoso and Portugal (2005), Hibbs and Lock (1996) and Rycx (2003) for Belgium.

13 For example, the average earnings in firms with less than 25 employees are 30 percent lower than those in firms with more than 500 employees, as is the standard deviation of earnings. More importantly, the mean and standard deviation of earnings changes are 15 percent lower for smaller firms compared to larger firms.

include a larger amount of extra wage components which are not always subject to indexation and collective agreements, and are thereby more flexible. Also, union representation is guaranteed in firms with 50 employees and more. This may ease negotiations of wage concessions in adverse times. Lastly, larger firms are more likely to negotiate SE wage agreements, which allow for a more flexible wage policy than the sector agreements do. Column (4) in Table 3 confirms these arguments. Larger firms show lower downward real wage rigidity. If the number of employees goes up from 5 to 25, DRWR would be according to our estimates lower by 0.128.

We also study whether production technology and market competition are related to DRWR. First, we introduce the median capital-labour ratio for each sector in column (5) of Table 3. Our estimates indicate that labour intensive sectors have higher DRWR. A one standard deviation increase in the capital-labour ratio reduces DRWR by 0.72. The increase is roughly equivalent to the difference in capital-labour ratio between construction (1.00) and financial services (1.62). Our result is consistent with the evidence provided by Fuss (2008). She finds that wage cuts in adverse times are essentially absent in the construction sector (the most labour intensive) as compared to the manufacturing and services sectors. Note that labour intensive sectors such as construction, textiles and transport equipment for example, are also characterised by a larger proportion of blue collar workers (see Table 1A), whose wages are less rigid. Table 3 shows that capital intensity is negatively related to DRWR after controlling for labour force composition. Our results contrast with the findings of Campbell (1991) who reports a negative correlation between sector-level wage flexibility and the capital-labour ratio in the US. However, our results support the view expressed by the Eurosystem Inflation Persistence Network (IPN) that the higher degree of price stickiness observed in more labour-intensive sectors results from wage rigidity, see Altissimo et al. (2006), Álvarez et al. (2006) and Vermeulen et al. (2006). In addition Dhyne et al. (2006) report evidence of downward consumer price rigidity in the services sector.[14]

In column (6), we report results for competition measured through the profit elasticity proposed by Boone et al. (2007) and estimated at the sector level. In appendix B, We report results with two other measures of competition, the Herfindahl index and the price-cost margin. Our result should be taken with some caution because estimates with these alternative measures predict the same direction of the impact of competition on DRWR but the coefficient is not statistically significant, as shown in Appendix B. Controlling for age and occupation, our estimates indicate that sectors with stronger competition experience higher DRWR. One explanation might be that fiercer competition reduces profits. As there is less rent to share, labour compensation is lower. Therefore, wages are closer to the collectively agreed minima and extra-wage components may be lower.[15] An alternative explanation is related to wage bargaining practices. SE wage agreements are more

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[14] Our measure of DRWR is negatively related to the sector-specific frequency of monthly producer price changes in the manufacturing sector, computed as in Cornille and Dossche (2006), suggesting that sectors with higher DRWR also experience higher price rigidity. The correlation coefficient between DRWR and the frequency of producer price change reaches -0.67. We thank M. Dossche for providing us with the estimates of the frequency of producer price change.

[15] Indeed, the correlation between the profit elasticity and average earnings or average bonuses, as reported in Table 1A, is 0.42.

common in sectors where firms are large and have more market power, and where firm-level unions try to appropriate the rents. SE wage agreements are far less common in sectors with competing small firms.[16] In this case, the main objective of unions is rather egalitarian as they are trying to avoid a wage race to the bottom; they organise themselves essentially at the sector level in order to negotiate equalised wages within the sector.

Finally, we examine whether differences in wage bargaining across sectors influence DRWR. In the literature, wage bargaining institutions have been cited as a cause of differences in downward wage rigidity across countries. Dickens et al. (2006) and Holden and Wulfsberg (2007) relate higher wage rigidity to higher union density and/or bargaining coverage. In the context of our paper, we examine whether sectoral differences in the wage bargaining mechanism are related to sectoral differences in wage rigidity. As already mentioned above, inter-sectoral co-ordination practices and indexation mechanisms are largely determined at the national level. These are common to all sectors and may explain the high level of DRWR in Belgium compared to other countries. Beyond this, other bargaining characteristics, such as the proportion of firms with SE agreements, vary across sectors. As already mentioned above, firm-level or single-employer (SE) agreements lead to higher wages on average, as well as increased wage dispersion across firms, because such agreements can better take into account firm-specific characteristics in the determination of wages[17]. In addition, according to Cardoso and Portugal (2005), a higher average wage and higher wage dispersion within firms provide employers with a flexible wage cushion above the sectoral minima, leaving these firms with a wider range of options in their wage-setting policy, i.e. it allows a more important role for workers' and firms’ characteristics in compensation. This in turn is expected to decrease downward real wage rigidity. Therefore, because firms with SE agreements have higher average earnings and wider earnings dispersion, they should be characterised by lower DRWR. This prediction is confirmed by the model in column (7) in Table 3. The higher the proportion of workers covered by an additional firm-level wage agreement in the sector, the lower the degree of downward real wage rigidity.

Appendix B provides a set of robustness tests with respect to outliers and alternative definitions of the explanatory variables. The estimates in Table B1 show that the results presented in Table 3 are robust to outliers and alternative definitions of variables. The coefficients are significant and of the same order of magnitude.

Up to now we investigated differences in both worker types and sectors. In order to provide a clearer view of the driving factors of DRWR across sectors, we perform similar regressions but focusing on sector differences alone. We estimate DRWR year by year for each sector, rather than for each occupation, age and sector category. We then regress DRWR by sector on the same variables as above, except that age and occupation dummies are replaced by the percentage of blue collars and the average age of workers in the sector. Results are presented in Table 4.

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16 However, the correlation between the profit elasticity and SE agreement as reported in Table 1B is only 0.01.

17 See Card and de la Rica (2006), Cardoso and Portugal (2005), Dell'Aringa and Lucifora (1994), Gerlach and Stephan (2006), Hibbs and Lock (1996), Palenzuela and Jimeno (1996) and Rycx (2003) for Belgium.

Table 4 - Explaining sectoral differences

Dep. var. $DRWR_{jt}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Average age_{jt}	-0.005	-0.019	-0.020*	-0.006	-0.012	-0.020*	0.016
	(-0.33)	(-1.17)	(-1.82)	(-0.52)	(-1.11)	(-1.84)	(1.14)
Percent. of $blue\text{-}collars_{jt}$	-0.001	-0.000	-0.003***	0.001	-0.002**	-0.001*	-0.000
	(-1.34)	(-0.32)	(-2.64)	(0.79)	(-2.32)	(-1.71)	(-0.74)
Median $earning_{jt}^{\dagger}$	-2.460						
	(-1.50)						
Average $bonus_{j}^{\dagger}$		-0.004					
		(-0.16)					
Earnings $spread_{jt}$			-0.353***				
			(-2.84)				
Median $size_{jt}^{\ddagger}$				-13.90***			
				(-2.91)			
Median capital-labour $ratio_{j}^{\dagger}$					-0.017***		
					(-4.91)		
Profit $elasticity_{jt}$						0.029**	
						(2.42)	
SE agr. coverage of $blue\text{-}collars_{j}$							-0.005***
							(-4.17)
Constant	0.890*	1.217**	1.937***	0.762*	1.252***	1.080***	0.055
	(1.87)	(2.41)	(4.18)	(1.76)	(3.29)	(2.65)	(0.11)
Year dummies	yes	yes	yes	yes	yes	yes	yes
R^2_{adj}	0.092	0.080	0.124	0.126	0.201	0.112	0.171
Number of obs.	173	173	173	173	173	173	173

Notes: † measured in thousands of euros.
‡ measured in thousands of employees.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
T-statistics in brackets.

Compared to Table 3, the coefficients have the same signs. However, median earnings and average bonus per sector in columns (1) and (2) in Table 4 are not statistically significant, contrary to median wage and bonus per worker type (age and occupation) and sector in Table 3. This suggests that earnings level and bonus mostly explain variation in DRWR across worker types, but not so much differences across sectors. We obtained the same conclusion for age already in Section 3.1 and Table 4 provides confirmation of the finding. Thus, being older and having higher wages or bonuses leads to more flexible earnings than younger and low-wage workers would experience, whatever the sector of economic activity.

Taken individually, the coefficients for the remaining variables in columns (3) to (7) are significant, have the expected sign, and are of the same order of magnitude as in Table 3, except that the size of the coefficients for average firm size and SE agreement coverage are larger.

In Table 5 we combines the explanatory variables discussed up to now into a single model. It was already suggested that the results might be plagued by multicollinearity between the explanatory variables giving rise to imprecise estimates of the coefficients. For example, older workers are likely to earn higher wages. Larger firms typically offer higher and more dispersed wages. And companies with firm level agreement are in general larger, pay higher wages and have a more dispersed wage structure. We therefore omit from the model the earnings level, bonuses and the earnings spread.[18] One may argue that these variables are the outcome of the firm's compensation policy, as is DRWR, and are therefore potentially endogenous. We prefer to maintain variables that are independent of the firm's pay policy, such as competition indices, capital intensity or SE agreement coverage.

Table 5 - Explaining sectoral differences, composite models

Dep. var. $DRWR_{jt}$	Model S1	Model S2	Model S3
Est. method	OLS	OLS	2SLS
Average	0.017		
age_{jt}	(1.24)		
Percent. of	-0.001*	-0.001*	-0.002**
$blue\text{-}collars_{jt}$	(-1.85)	(-1.73)	(-2.14)
Median	-11.012*	-10.764*	-14.900**
$size_{jt}^{‡}$	(-1.88)	(-1.83)	(-2.28)
Median capital-	-0.011***	-0.012***	-0.072
labour $ratio_j^{†}$	(-2.62)	(-2.77)	(-1.40)
Profit	0.023	0.021	0.049**
$elasticity_{jt}$	(1.65)	(1.49)	(2.13)
SE agr. coverage	-0.002*	-0.002	-0.001
of $blue\text{-}collars_j$	(-1.74)	(-1.28)	(-1.18)
Constant	0.091	0.685***	0.465**
	(0.18)	(4.95)	(2.33)
Year dummies	yes	yes	yes
R^2_{adj}	0.254	0.252	0.23
Number of obs.	173	173	173
F test for excluded instruments equal to 0			31.831
[p-val.]			[0.000]
Sargan's χ^2 test			2.875
[p-val.]			[0.238]

Notes: In Model S3, profit elasticity is treated as an endogenous variable. It is instrumented with the following excluded exogenous variables: Herfindahl index, number of firms per branch (and year) and the relative net increase in the number of firms in each branch and year.
† measured in thousands of euros.
‡ measured in thousands of employees.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
T-statistics in brackets.

In Model S1 in Table 5, there are two variables that are insignificant at the 10 percent level. One of them is profit elasticity, which is however only marginally insignificant (with p-value of 0.12). The other insignificant variable is age. Since we concluded that age does not explain the variation in DRWR across sectors, we estimate Model S2 that excludes the average age. As in Campbell (1989, 1991) we find that sectors with a higher proportion of blue-collar workers have higher wage rigidity. Table B3 in Appendix B3 shows that the qualitative results remain the same if we replace profit elasticity by the Herfindahl index.

18 The wage level and bonuses are highly collinear with the other variables included in the equation. Regressing the earnings level on the remaining explanatory variables excluding bonuses yields a coefficient of determination of 0.67; while the R^2 for the regression of bonuses on the other variables without the earnings level reaches 0.81.

One may argue that the profit elasticity in Model S2 is endogenous. Profits depend on wages and therefore, measures of profitability such as the profits-to-capital ratio, or empirical measures of the price-cost margin such as the profit-to-sales ratio will suffer from a simultaneity bias in an equation *for wages*. However, the simultaneity issue in an equation relating DRWR to the profit elasticity is much less clear. Under imperfect competition, an increase in costs, say raw materials, leads to a rise in prices and a smaller reduction in profits than for price-takers. Under DRWR, the real wage does not fall following the price increase; this reduces profits further. Therefore one observes a higher elasticity of profits with respect to marginal costs when DRWR is larger. This may bias the coefficient of profit elasticity upwards. Note however that this reasoning holds conditional on a *positive* cost shock only.

The last column in Table 5 reports estimates of Model S3 with instrumental variables for the profit elasticity in order to avoid potential simultaneity bias. We use the following instruments for profit elasticity: Herfindahl index, number of firms per branch (and year) and the relative net increase in the number of firms in each branch and year. The model is estimated by two-stage least squares.[19] In Model S3, the coefficient on profit elasticity increases and becomes significant when compared to Model S2. On the other hand, median capital-labour looses its significance. The estimates of the remaining coefficients are broadly in line with the results in Model S2.

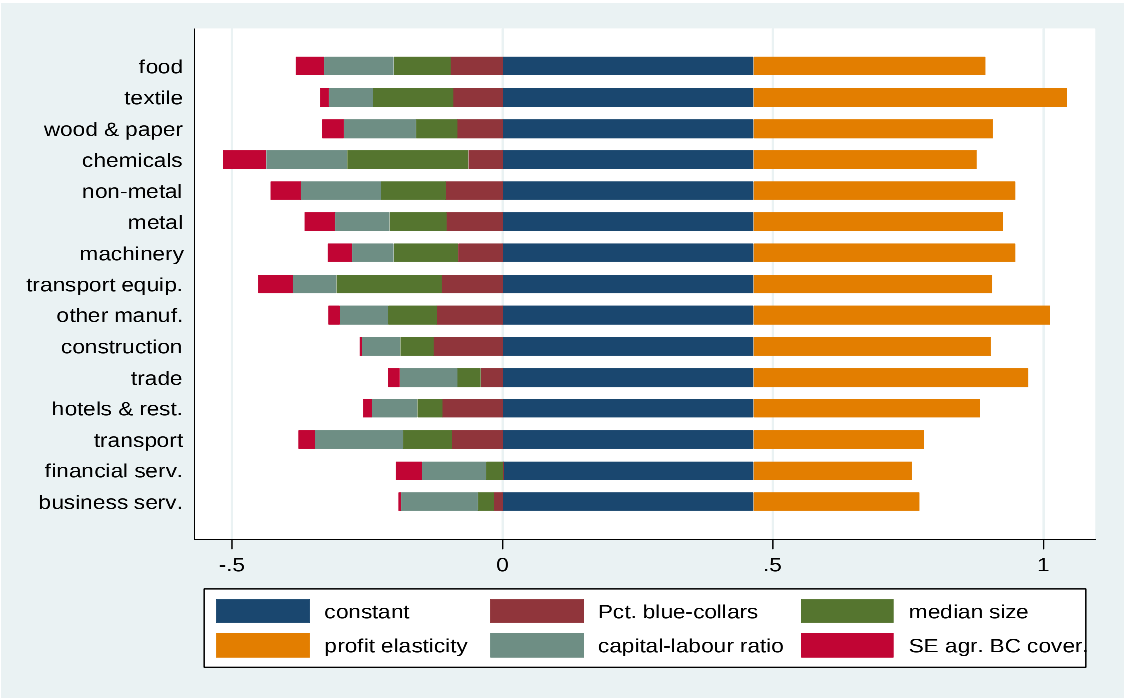

Figure 1 - Decomposition of DRWR based on Model S3 in Table 5 for year 2002

19 We run several tests of validity of our set of instrumental variables. First, the F-test for joint significance of the excluded exogenous variables in the first stage equation confirms that the instrumental variables are partially correlated with profit elasticity, the endogenous variable (see Table 5). Second, the insignificant test statistic of the Sargan's test of overidentifying restrictions states that the instruments are uncorrelated with the structural error term.

In order to highlight the importance of each variable to explaining DRWR across sectors, Figure 1 reports the contribution of each variable to sector-specific DRWR in the year 2002, based on Model S3 in Table 5.

Figure 1 shows that the constant picks up a large part of the level of DRWR in most sectors. This is in line with the observations made in Section 2. The standard deviation of DRWR across sectors is 0.11, around the mean of 0.64 (see Table 1A). This is consistent with the view that national institutions drive most of the observed wage rigidity. In particular, full automatic indexation and collectively agreed real wage increases are the main factors that explain wage rigidity.[20]

Leaving the constant aside, Figure 1 highlights the role of the different variables in explaining variation in DRWR across sectors. Let us first compare construction and chemicals, the sectors with the highest and one of the lowest values of DRWR in Table 1A (0.801 and 0.467, respectively). This may be attributed essentially to large difference in capital intensity, firm size and SE agreement coverage. As shown in Table 1B, chemical industry is characterised by one of the highest capital-labour ratios, the highest median firm size and the highest SE agreement coverage, while the opposite holds true for construction. As suggested by our estimates in Table 5, the higher the values of these variables, the lower is DRWR. These factors explain why DRWR is much higher in construction than in chemicals, despite the fact that construction has a disproportionately high percentage of blue-collar workers, while chemicals has a median proportion of blue-collar workers in its workforce.

The difference in DRWR between transport and storage (a service sector with low DRWR,0.354) and wood and paper (a manufacturing sector with medium-high DRWR, 0.648) is attributable mainly to competition. Indeed, the profit elasticity is 28 percent larger in wood and paper than in transport and storage, while the value of other variables are of the same order of magnitude.

In the services sector, financial and business services have approximately the same level of DRWR(0.627 and 0.668 respectively). They both employ a disproportionate number of white-collar workers. In spite of these similarities, they differ in that business services are characterised by higher capital intensity and financial services by higher SE agreement coverage.

4. Conclusion

Wage rigidity has been shown to have important implications both at the microeconomic and macroeconomic level. When wages are rigid, they no longer evolve hand in hand with productivity developments. Downward wage rigidity has been argued to be one of the causes of unemployment and price stickiness. It also bears implications for the design and effectiveness of monetary policy: These findings have led to a large empirical literature on the evaluation of wage rigidity, based on macroeconomic, sector-level or, more recently, microeconomic data.

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20 This statement cannot be tested in the framework of this study. The role of national institutions may be better evaluated by a cross-country analysis, as in Dickens et al (2006) or the meta-analysis of Clar et al. (2007).

This paper examines whether differences in wage rigidity across sectors can be explained by differences in workforce composition, competition, technology and bargaining institutions. Given the institutional features of the Belgian labour market, in particular full automatic indexation and previous results by Du Caju et al. (2007), we focus on downward rigidity of real wages. We adopt the measure of downward real wage rigidity developed by Dickens and Goette (2006). The estimates are based on a large administrative matched employer-employee dataset for Belgium over the period 1990-2002. We also use sector-level information derived from firms annual accounts over the same period and the 1999, 2000, 2001 and 2002 waves of the Belgian Structure of Earnings Survey (SES).

Our results are derived from regressions in two baseline specifications. First, we estimate downward real wage rigidity for different sectors and categories of workers (defined according to their occupation and age category). Using those estimates as dependent variable and different sets of sector and worker-type dummies as independent variables, we investigate the effects of differences in workforce composition on individual (worker-type) DRWR. The second specification focuses solely on the variation in DRWR across sectors and as such the number of observations is limited. Along with variables related to worforce composition we include explanatory variables that can be suggested as determinants of DRWR and finally we combine the variables with the highest explanatory power into a single model.

Even though several studies found that white-collar workers and younger workers have more rigid wages, the final piece of evidence showing statistical significance of the difference was missing. Our results indicate that DRWR is significantly higher for white-collar workers and younger workers. While investigating whether workforce composition helps to explain the differences in DRWR across sectors, we found that the larger the proportion of white-collar workers in a given sector, the larger the DRWR of the sector. On the other hand, differences in average age across sectors do not contribute to explanation of sectoral differences in DRWR. Similar conclusions hold for earnings level and bonuses: individuals with higher earnings and bonuses have lower DRWR, however, sectors with higher median wages and bonuses do not exhibit significantly lower degree of DRWR.

The paper confirms that sectoral variation in DRWR is significantly related to competition, technology and bargaining institutions. We find that wages are more rigid in more competitive sectors (measured through profit elasticity) and in labour-intensive sectors. Last but not least, we consider the impact of labour market institutions, through a measure of decentralisation of wage bargaining, i.e. the firm-level agreement coverage. Our findings suggest that the more centralised is wage formation, i.e. the lower is firm-level agreement coverage, the higher is DRWR.

References

Agell, J. and H. Benmarker (2007), "Wage incentives and wage rigidity: a representative view from within", Labour Economics, 14(3), 347-369.

Altissimo F., L. Bilke, A. Levin, T. Mathä and B. Mojon (2006), "Sectoral and aggregate inflation dynamics in the euro area", Journal of the European Economic Association, 4(2-3), 585-593..

Altonji J.G. and P.J. Devereux (1999), *Extent and Consequences of Downward Nominal Wage Rigidity*, in S.W. Polachek, Worker Well Being, Research in Labor Economics, 19, Elsevier.

Álvarez L., E. Dhyne, M. Hoeberichts, C. Kwapil, H. Le Bihan, P. Lünnemann, F. Martins, R. Sabbatini, H. Stahl, P. Vermeulen and J. Vilmunen (2006), "Sticky prices in the euro area: a summary of new micro- evidence", Journal of the European Economic Association, 4(2-3), 575-584.

Biscourp, P. O. Dessy, and N. Fourcade (2005), "Les salaires sont-ils rigides? Le cas de la France à la fin des années 90", Economie et Statistique, 386, 59-79.

Blanchard, O. and J. Gali (2006), "A New Keynesian model with unemployment", MIT Working Paper 06-22.

Boone, J. (2000), "Competitive pressure: the effects on investments in product and process innovation ", Rand Journal of Economics, 31(3), 549-569.

Boone, J., J.C. van Ours and H. van der Wiel (2007), "How (not) to measure competition", TILEC Discussion Paper 2007-014

Campbell, C. (1989), "Sectoral wage rigidity in the Canadian and French economies", European Economic Review, 66, 1727-1749.

Campbell, C. (1991), "Tests of efficiency wage theory and contract theory with disaggregated data from the US ", Weltwirtschaftliches Archiv, 127(1), 98-118.

Campbell, C. (1997), "The variation in wage rigidity by occupation and union status in the US", Oxford Bulletin of Economics and Statistics, 59 (1), 133-147.

Card D. and D. Hyslop (1997), "Does inflation grease the wheels of the labor market?", in *Reducing Inflation: Motivation and Strategy*, C. D. Romer and D. H. Romer (eds), NBER Studies in Business Cycles, v30.

Card, D. and S. de la Rica (2006), "Firm-level contracting and the structure of wages", Industrial and Labour Relations Review, 59(4), 573-592.

Cardoso, A and P. Portugal (2005), "Contractual wages and the wage cushion under different bargaining settings", Journal of Labor Economics, 23(4), 875-902.

Clar, M., C. Dreger and R. Ramos (2007), "Wage flexibility and labour market institutions: a meta-analysis", Kyklos, 60(2), 145-163.

Cornille, D. and M. Dossche (2006), "The patterns and determinants of price setting in the Belgian industry". NBB Working Paper, N°82.

Dell' Aringa, C. and C. Lucifora (1994), "Collective bargaining and relative earnings in Italy", European Journal of Political Economy, 10, 727-747.

Dhyne E., L. J. Ãlvarez, H. Le Bihan, G. Veronese, D. Dias, J. Hoffmann, N. Jonker, P. Lünnemann, F. Rumler, J. Vilmunen (2006), "Price Changes in the Euro Area and the United States: Some facts from Individual Consumer Price Data", Journal of Economic Perspectives, 20, pp. 171-192.

Dickens W.T., L. Goette, E.L. Groshen, S. Holden, J. Messina, M.E. Schweitzer, J. Turunen, and M.E. Ward (2007), "How wages change: micro evidence from the International Wage Flexibility Project", Journal of Economic Perspectives, 21(2), 195-214.

Dickens W.T., L. Goette, E.L. Groshen, S. Holden, J. Messina, M.E. Schweitzer, J. Turunen, and M.E. Ward (2006), "The Interaction of Labor Markets and Inflation: Analysis of Micro Data from the International Wage Flexibility Project", mimeo, Brookings Institution, November 2006.

Dickens, W.T., L. Goette (2006), "Estimating Wage Rigidity for the International Wage Flexibility Project". mimeo, Brookings Institution, October.

Du Caju, Ph., C. Fuss and L. Wintr (2007), " Downward wage rigidity for different workers and firms: an evaluation for Belgium using the IWFP procedure", ECB Working Paper N° 840.

Fuss, C. (2008), "How do firms adjust their wage bill in Belgium : A decomposition along the intensive and extensive margins", ECB Working Paper, forthcoming.

Gerlach, K. and G. Stephan (2005), "Wage distribution by wage-setting regime", IAB Discussion Paper, N°9/2005.

Kahn S. (1997), *"Evidence of Nominal Wage Stickiness From Micro Data"*, American Economic Review, 87(5), 993-1008.

Hibbs, H.A. and H. Locking (1996), "Wage compression, wage drift and wage inflation in Sweden", Labour Economics, 3, 109-141.

Holden S. and F. Wulfsberg (2007), "Downward nominal wage rigidity in the OECD", CESifo Working Paper N° 2009.

Layard R., S. Nickell and R. Jackman (1991), *Unemployment, macroeconomic performance and the labour market*, Oxford University Press.

Li, G. (1985), "Robust Regression", In *Exploring Data, Tables, Trends, and Shapes,* ed. D. C. Hoaglin, F. Mosteller, and J. W. Tukey, 281-340. New York: Wiley.

Palenzuela, D.R. and J.F. Jimeno (1996), "Wage drift and collective bargaining at the firm level: evidence from Spain", Annales d'Economie et Statistique, 41(42), 187-206.

Rycx, F. (2003), "Industry wage differentials and the bargaining regime in a corporatist country", International Journal of Manpower, 24(4), 347-366.

Shapiro, C. and J. E. Stiglitz (1984), “Equilibrium Unemployment as a Worker Discipline Device”, American Economic Review, 74, 433–44.

Stiglitz, J. E. (1974), “Alternative Theories of Wage Determination and Unemployment in L.D.C.’s: The Labor Turnover Model”, Quarterly Journal of Economics, 88, 194–227.

Vermeulen Ph., and R. Christopoulou (2007), "Markups in the Euro Area and the US over the period 1981-2004: a comparison of 50 sectors", ECB Working Paper, forthcoming.

Vermeulen P., M. Dias, M. Dossche, E. Gautier, I. Hernando, R. Sabbatini and H. Stahl (2007), "Price setting in the euro area: some stylised facts from individual producer price data and producer surveys", ECB Working Paper N° 727.

Weiss, A. (1980), “Job Queues and Layoffs in Labor Markets with Flexible Wages”, Journal of Political Economy, 88, 526–38.

Appendix A: Data

This appendix discusses the definitions of variables employed in this paper. The subscripts used in the paper in connection with variables refer to the particular group of observations over which the variable varies. The shortcuts for branch, occupation, age category and year are j, k, a and t, respectively. Unless explicitly stated, the variables come from the same administrative database on individual labour earnings for Belgium collected by the social security system used in Du Caju et al. (2007). Our trimming procedure of annual gross earnings is explained in section 2 of the paper. In addition, we use information from firms' balance sheets. Annual accounts data in Belgium is (nearly) exhaustive, on behalf of credit institutions, and is subject to substantial consistency checks. Finally, we also rely on individual data from the Belgian Structure of Earnings Survey (SES), for the 1999, 2000, 2001 and 2002 waves.

Variables related to compensation

Variables related to firms'pay policy are the level of earnings, the level of bonuses and the earnings dispersion. Earnings refers to daily gross earnings of an employee and are defined as annual gross earnings in euros divided by annual work days. The variable and the sample is identical to the one used in Du Caju et al. (2007). Table A3 on page 36 in Du Caju et al. (2007) provides descriptive statistics on the earnings distribution. Earnings spread is defined as the ratio of the 80^{th} and 20^{th} percentile of the daily earnings distribution defined above. Bonuses are obtained from the SES and refer to annual bonuses of an employee in euros. We do not make use of the variation of bonuses over years in order to retain as many observations as possible (our main data set covers the period 1991-2002, the SES 1999-2002 only).

To illustrate the data and our measure of downward real wage rigidity, Figures A1 to A3 below report histograms of the distribution of earnings changes for three sectors in 2002: chemicals, which according to our estimates has one of the lowest DRWR (0.490 in 2002), textiles having a median DRWR (0.617 in 2002), and construction with the highest DRWR (0.852 in 2002).

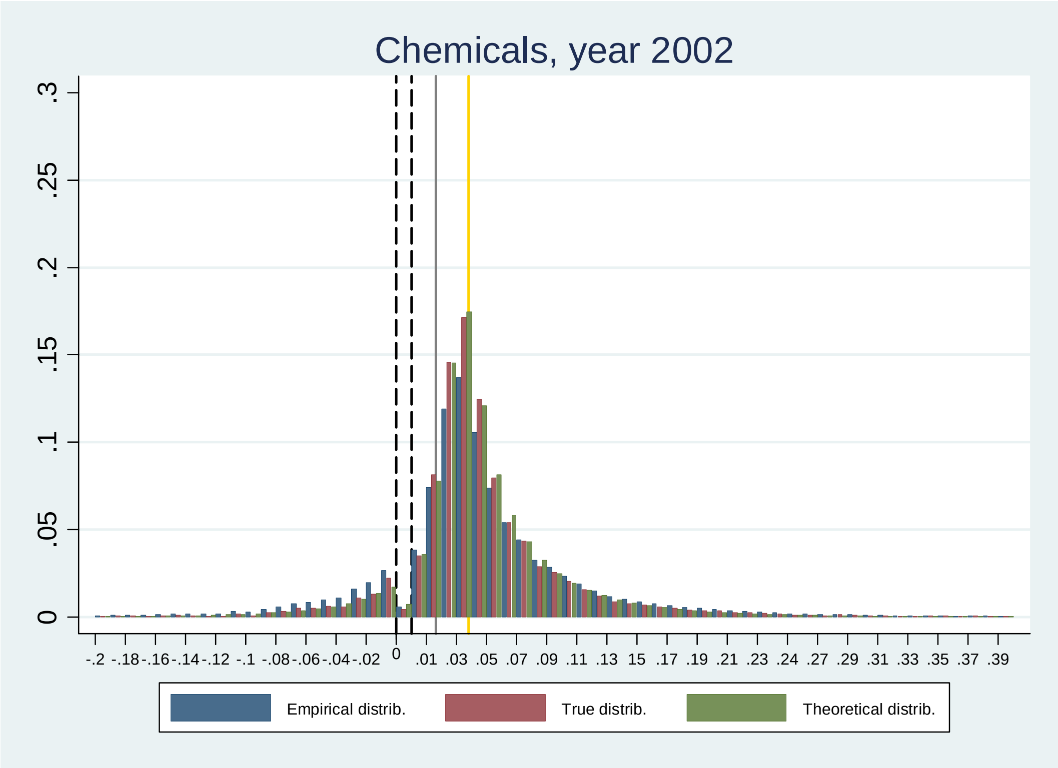

Figure A1 - Histograms of earnings changes for the chemicals sector in year 2002

Note: gray solid line shows the economy-wide CPI inflation while the yellow line is the economy-wide total collective wage increase.

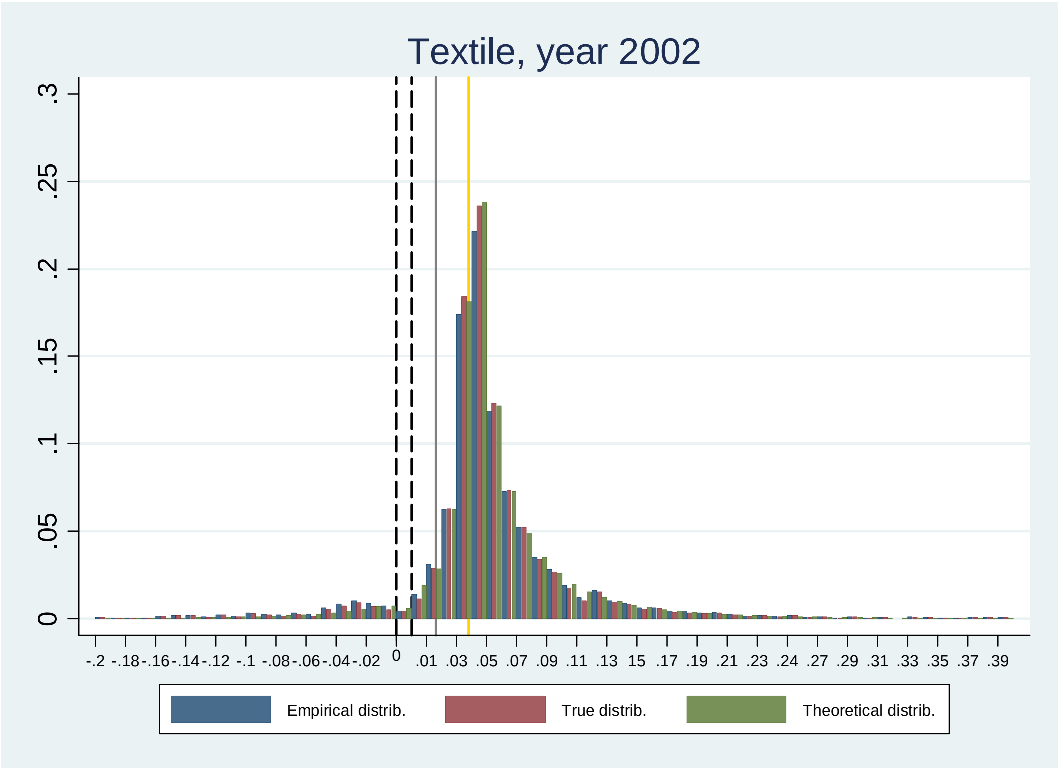

Figure A2 - Histograms of earnings changes for the textile industry in year 2002

Note: gray solid line shows the economy-wide CPI inflation while the yellow line is the economy-wide total collective wage increase.

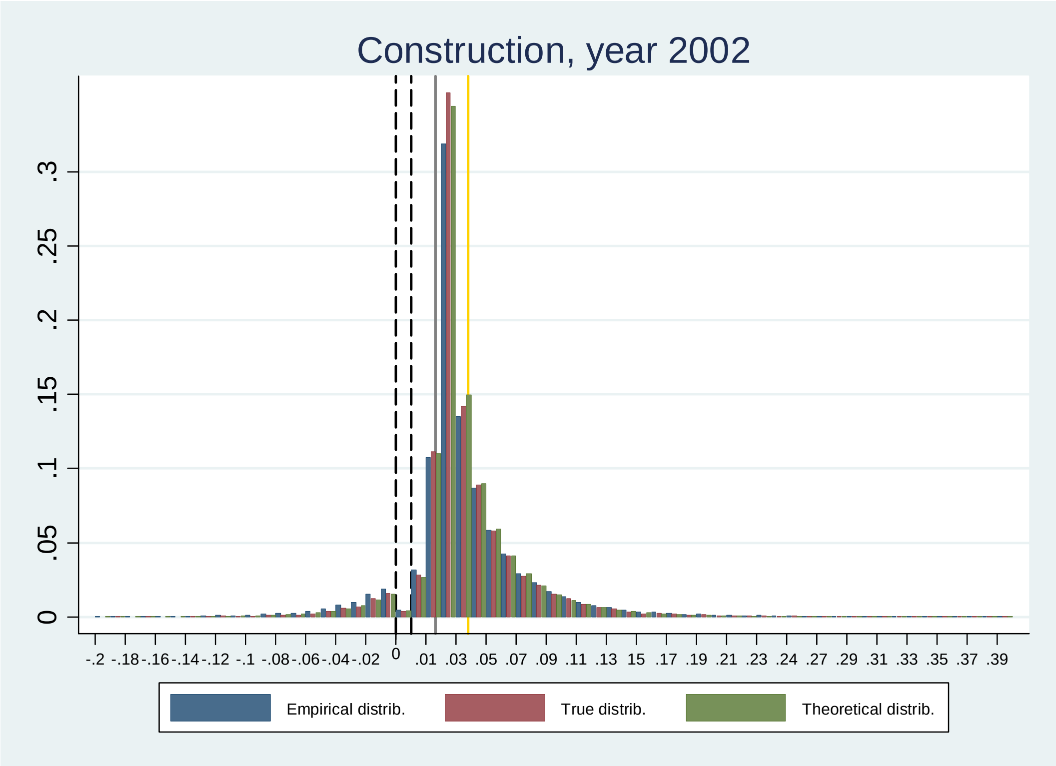

Figure A3 - Histograms of earnings changes for the construction sector in year 2002

Note: gray solid line shows the economy-wide CPI inflation while the yellow line is the economy-wide total collective wage increase.

Firm size

We measure the firm size by the number of employees. The definition of the number of employees in the balance sheet has changed over the period under examination. Since 1996 (and in some case 1997), firms report the total number of employees at the end of the year. Before 1996, only information on the average number of employees per year is available. Variables denoted as "Employees < 25_{jt}" and "Employees > 499_{jt}" refer to the percentage of firms in each branch and year that employ less than 25 and more than 499 employees, respectively.

Capital-labour ratio

Firm level capital-labour ratios are computed for all firms, both public companies and of non-profit associations, which publish an annual account for an entire year from January to December. The capital stock is computed based on the perpetual inventory method:

$$P^I_{st}K_{it} = (1-\delta_i)P^I_{st-1}K_{it-1}(P^I_{st}/P^I_{st-1}) + P^I_{st}I_{it}$$

with K_{it} representing the real capital stock, P^I_{st} the sector-specific deflator on gross capital formation and δ_i the firm-specific depreciation rate. The initial nominal capital stock is given by the book accounting value of the capital stock, plus revaluation gains, minus depreciation and amounts written down, all at the end of the preceding period. The firm-specific depreciation rate was estimated as the median depreciation expenditures on capital, over the years in which the firm exists. Labour is defined as the number of employees, as above. Capital labour ratios are taken as the median or average within each sector over the entire period 1991-2002 and are measured in euros. We do not consider the capital labour ratio by year, because in the short-run changes in the ratio can reflect employment changes rather than a change in the production technology.

As a robustness test, we calculate the capital-labour ratio also from the national accounts data. We take the net capital stock per sector at the end of the year divided by the yearly average of the number of workers in the branch. In national accounts, the gross capital stock is estimated according to the perpetual inventory method, based on historical series of gross fixed capital formation, the average service life of fixed assets and survival functions, from which the cumulative value of consumption of fixed capital is deducted to obtain the net stock. We deflate the net capital stock using the sector-specific deflator of gross capital formation at the A31 level.

Competition measures

We use a relatively new measure of competition, proposed by Boone et al. (2007): the elasticity of firms' profits with respect to marginal costs. A higher value of this profit elasticity suggests more intense competition. The intuition is as follows. An increase in costs lowers the firm's profits. As firms in less competitive markets have some market power over their prices, a given percentage increase in costs will lead to a smaller decline in profits in a less competitive market.

As shown in their paper, under monopoly the profit elasticity is a positive linear function of the demand elasticity. Consider a monopoly facing a constant elasticity demand function $x=p^{-\varepsilon}$. where x is output, p the output price and $-\varepsilon$ the price elasticity of demand. With c, the marginal cost, the firm's profit is given by $\Pi=x^{(\varepsilon-1)/\varepsilon}$-cx. Profit maximisation with respect to output yields the following expressions for output, $x=(\varepsilon-1)^{\varepsilon}/(\varepsilon c)^{\varepsilon}$, and profits, $\Pi=((\varepsilon-1)^{\varepsilon-1}/\varepsilon^{\varepsilon})c^{-(\varepsilon-1)}$. Taking the expression in logs yields $\ln\Pi=\ln((\varepsilon-1)^{\varepsilon-1}/\varepsilon^{\varepsilon})-(\varepsilon-1)\ln c$. The profit elasticity is equal to $-(\varepsilon-1)$.

Profit elasticity might also depend on marginal costs. To illustrate this, we consider the standard Cournot model with two firms and linear demand curve in the form $p(x_1,x_2) = a - b_1x_1 - b_2x_2$, where x_i is the output produced by firm *i*, the size of the market is captured by *a*, and b_i is each firm's own elasticity of demand. The cost function of each firm has the form of $C(x_i) = c_i x_i$. Firm *i* chooses its output x_i so that it maximizes its profits:

$$\max_{x_i}\{(a-b_1x_1-b_2x_2)x_i-c_ix_i\} \qquad i=\{1,2\} \tag{1}$$

In addition, we assume that $a > c_i > 0$ and that $b_i > 0$. Solving (1) yields reaction function of firm *i* to the output of firm *j*

$$x_i(x_j)=\frac{a-b_jx_j-c_i}{2b_i} \tag{2}$$

In Nash equilibrium, both firms are choosing best response to their competitor's output choice. After substituting x_j in equation (2) by $x_j(x_i)$, we obtain the equilibrium output of each firm

$$x_i^*=\frac{a+c_j-2c_i}{3b_i}$$

This can be substituted into the profit function to obtain the equilibrium profits

$$\Pi_i^* = \frac{\left(a + c_j - 2c_i\right)^2}{9b_i}$$

and we can show that each firm's profit elasticity is a non-linear function of marginal costs

$$\frac{\partial \Pi_i^*}{\partial c_i} \frac{c_i}{\Pi_i} = -\frac{4c_i}{a + c_j - 2c_i} \tag{3}$$

If DRWR has an impact on wages, it also influences profit elasticity through equation (3), causing simultaneity in our regression models. In Model S3 in Table 5 we use instrumental variables to account for potential simultaneity bias. See also Boone (2000) and Boone et al. (2007) for other examples and more general model specifications.

To estimate the profit elasticity, we follow Boone et al. (2007). Specifically, we regress the log of profits on the log of marginal costs. Marginal variable costs are defined as variable costs over turnover and are denoted as "mc" in what follows. We use information on all profit-maximizing firms that fill out annual accounts for the entire year from January to December. We identify outliers as firms with variable costs over turnover below or above the 5^{th} and 95^{th} percentile of the distribution and we use the same criterion for profits over total assets. For each branch, we estimate the following regression with firm-specific fixed effects for the period 1991-2002:

$\ln \pi_{it} = \alpha_i + \gamma_t - \beta_t \ln mc_{it} + \varepsilon_{it}$,

where β_t is the profit elasticity and π stands for profits.

As a robustness check, we also use the Herfindahl index as a simple measure of competition within each branch and year. We start by calculating the sum of the squares of the market shares of each individual firm within the branch. Market share is defined as the proportion of a firm's value added in the total value added of the branch. Finally, we normalize the index to range from 0 to 1. Formally, it is computed as

$$H_b = \frac{\sum_{i=1}^{N_b} \left(VA_i \Big/ \sum_{i=1}^{N_b} VA_i\right)^2 - 1/N_b}{1 - 1/N_b},$$

where VA_i is the value added of firm i in branch b, and N_b is the number of firms in branch b. As a measure of concentration, a small value of the Herfindahl index indicates a competitive industry with no dominant players.

We also consider the sector-level estimates of price-cost margins constructed by Christopoulou and Vermeulen (2007) for NACE 2 sectors for the US and several EU countries, including Belgium. The estimates are time-invariant. In cases where our sector definition is more aggregated than theirs, we consider the simple average of their estimated mark-ups.

Table A1 reports the three measures of competition as well as correlations with median firm size of the sector, the number of firms within the industry, the rate of firms entry the rate of firms exit.

Table A1: Competition measures

	Profit elasticity	Herfindahl index	Price Cost Margin
food	7.957	0.014	1.08
textile	9.514	0.009	1.09
wood and paper	7.436	0.007	1.13
chemicals	7.809	0.029	1.16
non metal	8.514	0.023	1.15
metal	8.119	0.036	1.13
machinery and equip.	8.642	0.023	1.20
transport equipment	9.877	0.074	1.06
other manufacturing	9.159	0.005	1.08
construction	8.432	0.001	1.17
trade	9.821	0.003	1.22
hotels and restaurants	8.189	0.013	1.23
transport and storage	5.784	0.010	1.26
financial services	5.473	0.012	1.44
business services	6.007	0.003	1.78
Mean	8.049	0.018	1.21
Standard deviation	1.388	0.019	0.18
number of firms	-0.46	-0.22	0.70
firm entry rate	-0.08	-0.42	0.64
firm exit rate	-0.07	-0.42	0.63
firm size	0.46	0.63	-0.55

Note: Correlation coefficients are calculated for sectors included in the data set.

Single-employer (SE) agreement coverage

The information on SE agreements is obtained from the Structure of Earnings Survey (SES) for Belgium covering the period 1999-2002 with annual frequency. The data set contains separate indicators of SE agreements for blue-collar and white-collar workers on the firm level and this allows us to match it with the occupational categories in our paper. For instance, for blue-collar workers, the SE agreement coverage refers to the percentage of blue-collar workers in each sector that work in firms with SE agreements covering blue-collar workers.

Appendix B: Robustness

In this appendix we consider the robustness of our results along two dimensions. Section B1 studies the robustness of our results with respect to outliers by re-estimating the models presented in Table 3 using the procedure suggested by Li (1985). Section B2 considers the sensitivity of our estimates to alternative definitions of the variables under study.

B1 Robustness with respect to outliers

Table B1 tests the robustness of the results presented in Table 3 with respect to outliers. We estimate identical models with the robust regression method suggested by Li (1985), as it is implemented in Stata. The procedure first eliminates large outliers (based on Cook's distance larger than 1). Afterwards, it iteratively runs regressions, calculates case weights based on absolute residuals, and regresses again using those weights (Huber weights and biweights). For more details, see Li (1985) or Stata reference manuals.

Table B1: Firm-level wage bargaining and DRWR, robust regressions

Dep. v. $DRWR_{kajt}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
D white-collar$_{kajt}$	0.115***	0.169***	0.090***	0.069***	0.087***	0.085***	0.073***
	(4.97)	(5.62)	(5.05)	(3.78)	(4.92)	(4.67)	(4.08)
D age:25-44$_{kajt}$	-0.018	-0.005	-0.068***	-0.042*	-0.065***	-0.061***	-0.041*
	(-0.70)	(-0.19)	(-2.96)	(-1.78)	(-2.86)	(-2.60)	(-1.76)
D age:45+$_{kajt}$	0.032	0.046	-0.044*	-0.019	-0.046**	-0.033	-0.023
	(1.09)	(1.61)	(-1.92)	(-0.80)	(-2.00)	(-1.42)	(-0.98)
Median earning$_{kajt}$ †	-1.935***						
	(-2.96)						
Average bonus$_{kaj}$ †		-0.036***					
		(-4.08)					
Earnings spread$_{jt}$			-0.287***				
			(-6.79)				
Median firm size$_{jt}$ ‡				-7.079***			
				(-2.79)			
Median capital-labour ratio$_{j}$ †					-0.013***		
					(-6.24)		
Profit elasticity$_{jt}$						0.023***	
						(3.61)	
SE agreement coverage$_{kj}$							-0.003***
							(-5.13)
Constant	0.554***	0.481***	0.968***	0.512***	0.703***	0.320***	0.544***
	(12.10)	(13.15)	(12.07)	(13.08)	(13.84)	(5.65)	(13.96)
Year dummies	yes	yes	yes	yes	yes	yes	yes
Sector dummies	no	no	no	no	no	no	no
R^2_{adj}	0.069	0.078	0.113	0.068	0.106	0.075	0.090
Observations	758	758	758	758	758	758	758

Notes: † measured in thousands of euros.
‡ measured in thousands of employees.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
T-statistics in parentheses, p-values in square brackets.

The coefficients of interest obtained from robust regressions are very similar to those reported in Table 3. They always have the same sign and also their significance level remains broadly similar. The largest difference between the robust and standard coefficients appears for median firm size, but even there the difference is only approximately 10 percent.

B2 Robustness with respect to definitions of the variables

In Table B2 we study the robustness of the results presented in Table 3 with respect to alternative definitions of the variables under consideration.

If we replace median wage or median bonus in Table 3 by average wage or average bonus, the coefficients and their significance level hardly change.

If we switch from median number of employees per firm to average, the coefficient looses its significance. We consider median number of employees as a better proxy for firm size because it does not get influenced by one (or few) extremely large firms in the sector. Du Caju et al. (2007) document large differences in DRWR between small and large firms (with less than 25 and more than 499 employees, respectively). For each branch and year, we calculate the percentages of firms that have less than 25 employees and more than 499 employees. The results show that the larger the percentage of small firms in a branch, the higher the rigidity even after controlling for age and occupational composition effect. On the other hand, branches with a higher proportion of firms with more than 499 employees do not show significantly lower degrees of DRWR. The result might stem from the fact that we are comparing branches with more or less firms in the large firm category, while Du Caju et al. (2007) compare large and small firms of all sectors.

Capital intensive sectors experience significantly lower DRWR, no matter we use our measure of capital intensity based on firm data or the one reported by national accounts (for details on their derivation, see Appendix A). The effect of capital-labour ratio obtained from the national accounts is approximately 60 times larger than the effect observed in Table 3 for the median capital-labour ratio that we calculated. This can be related to the fact that the average value of our median capital-labour ratio is more than 10 times lower than the one obtained from national accounts. Comparing average capital-labour ratio calculated from firm level data with the measure from national accounts leads to broadly comparable values and correlation coefficient of 97%.

We consider two alternative measures of competition, i.e. the structural estimates of price-cost margin of Christopoulou and Vermeulen (2007) and the more rudimentary Herfindahl index. The Herfindahl index is normalized to range from 0 to 1 with small value indicating competitive industry with no dominant players. The negative coefficient suggests that more competitive industries have higher DRWR, all else equal. The qualitative conclusion is the same as the one we obtained for profit elasticity, except that the coefficients are statistically insignificant. Following Boone et al. (2007), we consider profit elasticity to be a better measure of competitiveness in situations when increased competition leads to exit of inefficient firms, hence leading to an increase in Herfindahl's concentration index.

Under perfect competition, prices are equal to marginal costs and as a result the price-cost margin is zero. If competition becomes less fierce, prices increase above marginal costs and price-cost margin increases. Hence, then negative coefficient on price-cost margin in Table B2 implies that more competitive sectors exhibit higher degree of downward real wage rigidity, however, the coefficient is not statistically significant. The qualitative result is thus identical to the conclusion that we obtained for Herfindahl index and it is in line with the findings for profit elasticity discussed in the main text.

Table B2: Firm-level wage bargaining and DRWR, alternative definitions of variables

Dep. v. $DRWR_{kajt}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
D white-$collar_{kajt}$	0.114***	0.159***	0.069***	0.073***	0.070***	0.074***	0.072***	0.068***
	(4.82)	(5.55)	(3.97)	(4.26)	(4.02)	(4.31)	(4.13)	(3.95)
D age:25-44_{kajt}	-0.022	-0.012	-0.048**	-0.054**	-0.051**	-0.057**	-0.054**	-0.048**
	(-0.89)	(-0.49)	(-2.17)	(-2.44)	(-2.26)	(-2.56)	(-2.39)	(-2.16)
D age:45+$_{kajt}$	0.018	0.029	-0.031	-0.037*	-0.032	-0.039*	-0.034	-0.030
	(0.63)	(1.05)	(-1.40)	(-1.68)	(-1.43)	(-1.76)	(-1.52)	(-1.36)
Average $earnings_{kajt}^{\dagger}$	-1.536***							
	(-2.71)							
Average $bonus_{kaj}^{\dagger}$		-0.032***						
		(-3.87)						
Average firm $size_{jt}^{\ddagger}$			-0.285					
			(-1.33)					
Employees < 25_{jt}				0.002**				
				(2.51)				
Employees > $499_{jt}^{\ddagger}$					0.012			
					(0.02)			
NA capital-labour $ratio_{j}^{a}$						-0.019***		
						(-3.56)		
Price-cost $margin_{j}$							-0.048	
							(-0.94)	
Herfindahl $index_{jt}$								-0.667
								(-1.42)
Constant	0.549***	0.493***	0.496***	0.326***	0.488***	0.539***	0.547***	0.498***
	(13.19)	(14.15)	(13.91)	(4.45)	(13.07)	(14.30)	(7.64)	(13.90)
Year dummies	yes	yes	yes	yes	yes	yes	yes	yes
Sector dummies	no	no	no	no	no	no	no	no
R^2_{adj}	0.064	0.073	0.057	0.062	0.054	0.070	0.056	0.057
Observations	758	758	758	758	758	758	758	758

Notes: † measured in thousands of euros.
‡ measured in thousands of employees.
a capital-labour ratio obtained from national accounts (NA), measured in hundred-thousands of euros.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
T-statistics in brackets.

In table B3 we consider robustness of the regression models S1 and S2 presented in Table 5 with respect to the measure of competitiveness, i.e. we replace profit elasticity by Herfindahl index. In Models S1B and S2B in Table B3, the coefficient on Herfindahl index is negative implying that more concentrated industries have lower degree of DRWR. The coefficient is insignificant in Model SB1 but it is significant in Model SB2 at 10 percent level. The predicted effect has the same directions as our findings based on profit elasticity. i.e. more competitive sectors have higher DRWR. The remaining coefficients in Table B3 have the same sign as in Table 5, even though some of them loose their significance.

Table B3 - Robustness of models S1 and S2 in Table 5 with respect to Herfindahl index

Dep. var. $DRWR_{jt}$	Model S1B	Model S2B
Est. method	OLS	OLS
Average	0.010	
age_{jt}	(0.74)	
Percent. of	-0.001	-0.001
blue-$collars_{jt}$	(-1.19)	(-1.14)
Median	-7.036	-7.061
$size_{jt}^{‡}$	(-1.28)	(-1.29)
Median capital-	-0.177***	-0.180***
labour $ratio_j^{†}$	(-4.40)	(-4.52)
Herfindahl	-1.574	-1.708*
$index_{jt}$	(-1.51)	(-1.67)
SE agr. coverage	-0.001	-0.000
of blue-$collars_j$	(-0.43)	(-0.07)
Constant	0.533	0.879***
	(1.12)	(10.36)
Year dummies	yes	yes
R^2_{adj}	0.252	0.254
Number of obs.	173	173

Notes: † measured in thousands of euros.
‡ measured in thousands of employees.
*/**/*** indicate significance at the 0.10, 0.05 and 0.01 level, respectively.
T-statistics in parentheses, p-values in square brackets.