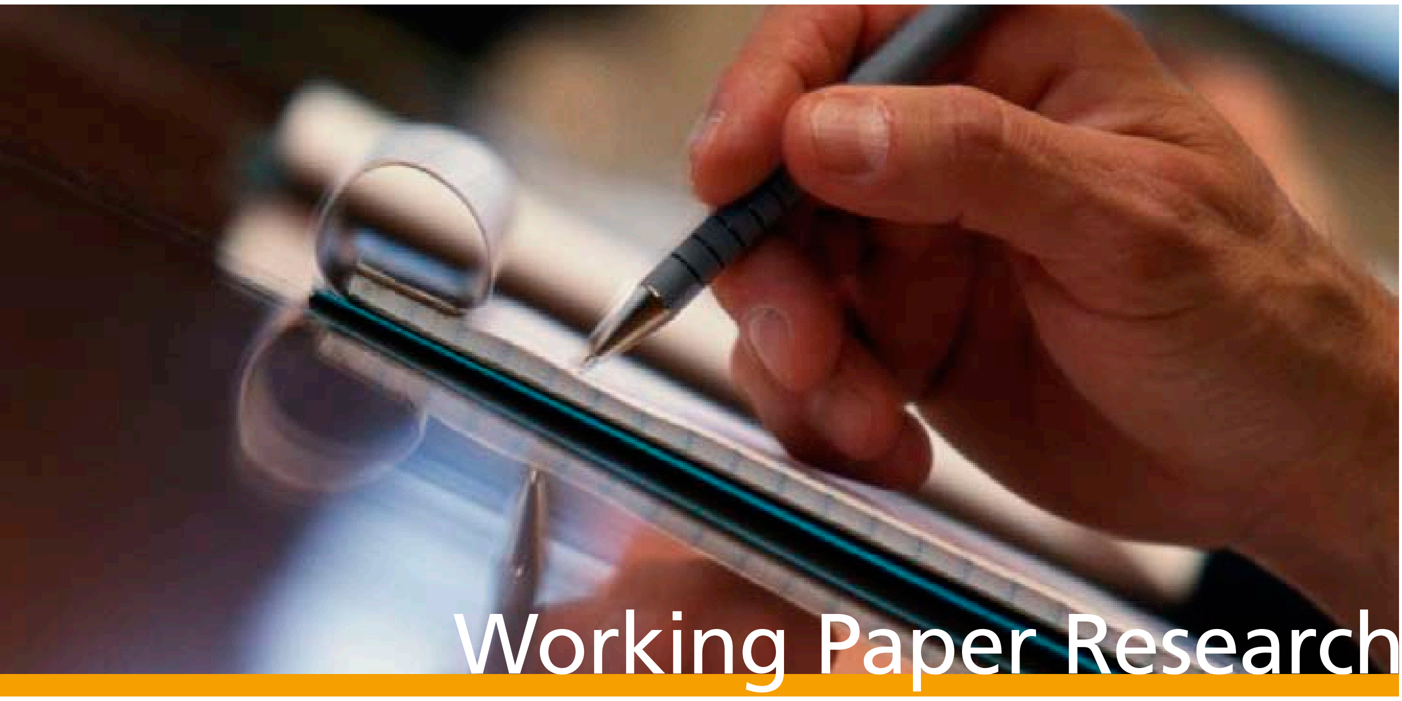
Rigid labour compensation and flexible employment? Firm-level evidence with regard to productivity for Belgium



by Catherine Fuss and Ladislav Wintr

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Abstract

Using firm-level data for Belgium over the period 1997-2005, we evaluate the elasticity of firms' labour and real average labour compensation to microeconomic total factor productivity (TFP). Our results may be summarised as follows. First, we find that the elasticity of average labour compensation to firm-level TFP is very low contrary to that of labour, consistent with real wage rigidity. Second, while the elasticity of average labour compensation to idiosyncratic firm-level TFP is close to zero, the elasticity with respect to aggregate sector-level TFP is high. We argue that average labour compensation adjustment mainly occur at the sector level through sectoral collective bargaining, which leaves little room for firm-level adjustment to firm-specific shocks. Third, we report evidence of a positive relationship between hours and idiosyncratic TFP, as well as aggregate TFP within the year.

Key-words: labour compensation, employment, hours, Total Factor Productivity.
JEL-code: J30, J60.

Corresponding authors:

Catherine Fuss, NBB, Research Department e-mail: catherine.fuss@nbb.be.
Ladislav Wintr, Central Bank of Luxembourg, Economics and Research Department,
e-mail: ladislav.wintr@bcl.lu.

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1. Introduction

This paper examines the response of the components of firms' total labour compensation to Total Factor Productivity (TFP) at the microeconomic level. More specifically, we investigate the impact of TFP on firms' real wage bill per employee (average labour compensation), firms' employment, total hours worked, as well as real labour compensation per hour (hourly labour compensation) and hours worked per employee. We distinguish between idiosyncratic firm-level TFP and aggregated sector-level TFP and evaluate and compare the elasticity of average labour compensation and labour to both firm-specific TFP and to sector-level TFP. Because average labour compensation vary with changes in the workforce composition as well as with variations in wages, our empirical specification includes control variables related to workforce composition. Therefore, the response of average labour compensation to TFP, after controlling for workforce composition, is mainly driven by changes in wages.

We find low response of average labour compensation and a high elasticity of labour to firm-specific TFP. This is consistent with the existence of real wage rigidity. Wage rigidity implies that wages have a smaller and more sluggish response to economic shocks than flexible wages. There are several reasons for wage rigidity and it can manifest itself in various ways, as the following examples show. Multi-period contracts may imply that wages do not respond to contemporaneous shocks. The resistance to wage cuts may imply reduced sensitivity to adverse shocks. This so-called downward wage rigidity may even reduce the sensitivity of wages to favourable shocks (Elsby, 2006). A sluggish response of wages may also be the outcome of wage bargaining between risk-averse workers and risk-neutral firms, leading to "wage insurance" (Azariadis (1975)).

Recent microeconomic research highlights the existence of various forms of wage rigidity. Firstly, Guiso et al. (2005), Cardoso and Portela (2005) and Katay (2007) provide microeconomic evidence that firms do insure workers against temporary firm-specific shocks to productivity. Note that firms can afford wage insurance only against temporary rather than permanent shocks. Second, recent evidence on downward wage rigidity in Belgium can be found in Dickens et al. (2006, 2007), Du Caju et al. (2007) and Knoppik and Beissinger (2005). These papers point to high downward real wage rigidity in Belgium. This is attributable mainly to the full automatic indexation of base wages. We therefore focus on adjustment of real labour compensation.

Models with real wage rigidity typically find wider variability of employment in response to productivity shock, as compared to the flexible wage scenario (see for example Blanchard and Galí (2007, 2008), Boldrin and Horvath (1995) or Hall (2005)). In addition to wage rigidities, other frictions alter the functioning of labour markets. Hiring and firing costs together with training expenses may generate considerable employment adjustment costs that impede labour adjustment. Which of the two constraints - wage rigidity or employment adjustment costs - is more binding has to be determined on empirical grounds.

This question is also relevant for inflation dynamics and monetary policy, as shown in the most recent strand of New Keynesian models. Examining jointly real wage rigidity and employment protection, Christoffel and Linzert (2005) show that, on the one hand, real wage rigidity increases the adjustment via the employment margin and explains inflation persistence. On the other hand, employment protection tends to smooth out labour flows, raise the volatility of wages following a monetary policy shock, increase the response of inflation, and thereby lower the persistence of inflation. Further, in the absence of wage rigidity, these models predict that the central bank should fully stabilise inflation at all times and at any cost (Goodfriend and King (1997)). On the contrary, price lumpiness (Christiano et al. (2005)), and real wage rigidity (Blanchard and Galí (2007, 2008)) generate inflation inertia and persistence of fluctuations in hours and output. Therefore, following an adverse economic shock, the monetary authority must decide whether to accommodate a higher level of inflation or, instead, keep inflation constant but allow for a larger decline in the output gap and employment. Pure inflation targeting is no longer the optimal monetary policy, which should rather aim at reducing, but not eliminating, the volatility of both inflation and unemployment.

In addition the degree of price stickiness may influence the employment adjustment mechanisms. First, under monopolistic competition conditions and flexible prices, a positive technology shock leads to a price reduction, an increase in demand and thereby a rise in output and labour demand. As discussed above, real wage rigidity may exacerbate the response of output and labour. Second, in a situation of price stickiness, prices and therefore demand remain unchanged. Following a positive productivity shock, the same volume of output is then produced using a smaller amount of labour (provided the shock is not offset by expansionary monetary policy action for instance, or that firms do not store unsold goods in expectation of future price change and increase in demand). Lastly, a negative relationship between technology shocks and labour may be also explained by a low elasticity of demand, high market power or a short-run negative impact on production due to necessary adaptation of stock and/or quality of labour and capital. For example, Francis and Ramey (2005) and Smets and Wouters (2007) point out that habit formation in consumption and adjustment costs in investment may induce a negative relationship between hours worked and technology.

The aim of this paper is to estimate the sensitivity of firm-level average real labour compensation per worker (referred to as average labour compensation hereafter), employment and hours to changes in firm-specific TFP, as well as in sector-level TFP. The main variables are obtained from companies' annual accounts and social balance sheet recorded in Belgium over the period 1997-2005. We estimate dynamic equations for average labour compensation, employment, total hours worked, as well as hours per worker and hourly labour compensation. The equations include TFP measures, as well as variables related to the workforce composition, firm size, and the capital stock, together with variables that capture sector-level fluctuations. TFP is measured through the growth accounting framework of Ackerberg et al. (2006) and corrected for fluctuations in hours per worker to account for variable utilisation of production factors (Basu and Kimball (1997)). Note that we use firm-level information on average labour compensation rather than individual earnings data. The advantage of using individual earnings data is that changes in wages cannot be confused with changes in the workforce composition. The main drawback is that wage changes can only be constructed for job stayers, while our approach also takes into account new entrants and workers that leave the firm together with permanent job-stayers when measuring average labour compensation changes. We believe that from the point of view of a firm, the relevant adjustment variable following a TFP shock is the average labour compensation rather than workers' individual wages, although changes in the composition of the labour force might have an impact on firm's productivity. We acknowledge this point by including control variables for labour force composition in our models.

First, we examine the relative sensitivity of average labour compensation and labour to firm-specific TFP changes. The finding of a small elasticity of average labour compensation and a large elasticity of labour to TFP supports the hypothesis that real wages are rigid. Two caveats should be mentioned here. One is that there is no theoretical reference point for the relative elasticity of wages and labour in cases of perfect wage flexibility. The other is that TFP measures may capture technology as well as demand shocks and input price shocks (see Klette and Griliches (1996), Katayama et al. (2003) and Foster et al. (2008)). Demand shocks will tend to induce a positive correlation with labour and a smaller correlation with wages (except to the extent that demand shocks raise profits and wages through rent-sharing mechanisms). Few recent papers look at these questions at the microeconomic level. Duhautois and Kramarz (2006) and Fuss (2008) examine the relative importance of average wage and employment flows in wage bill adjustment but do not estimate the elasticity of labour compensation and employment.

Second, we compare the elasticity of average labour compensation with respect to firm-specific and sector-level TFP, and explore the role of sector-level collective wage agreements in shaping the response of average labour compensation. To our knowledge there is no paper investigating differentiated behaviour of labour compensation and employment at idiosyncratic as opposed to aggregate level. The finding of a higher elasticity of average labour compensation to aggregate sector-level TFP than to idiosyncratic TFP highlights the role of centralisation and coordination of wage bargaining in facilitating wage adjustment.[1] It also stresses the fact that the importance of sector-level collective bargaining for wage-setting in Belgium may strongly limit the scope for firm-level adjustment. Lastly, it is consistent with the view that firms compete on the labour market to hire and keep workers, which makes them reluctant to undertake individual wage cuts, for effciiency wage or shirking considerations. Also, it may arise under a competitive product market environment in which firms may not afford wage increases because they will not be able to raise their prices if their competitors do not.

Third, by providing microeconomic estimates of the elasticity of hours to TFP, we contribute to the debate on the sign of the relationship between hours worked and technological change. Previous firm-level analyses of the impact of TFP on growth in hours suggest that a current TFP shock has a negative impact on hours, although the effect of lagged TFP is positive and compensates for the initial negative effect (Marchetti and Nucci (2005, 2007), and Carlsson and Smedsaas (2007)). This supports the hypothesis of sticky prices. One caveat of our exercise, as

1 This argument can be traced back to Bruno and Sachs (1985) who find that countries with more centralised wage bargaining find it easier to adjust real wages to adverse macroeconomic shocks.

well as those cited above, is again that TFP measures may capture demand shocks together with technological changes.

Our results may be summarised as follows: (1) the elasticity of average labour compensation with respect to TFP is close to zero, while the elasticity of labour is high. This is in line with the existence of real wage rigidity. Compared with microeconomic evidence for other European countries, our estimates of the average labour compensation sensitivity are in the lower range; (2) the elasticity of average labour compensation with respect to sector-level TFP is much larger than that of firm-level TFP. We relate this finding to the fact that the wage dynamics in Belgium is mostly driven by sector-level collective agreements; (3) we provide microeconomic evidence that hours respond positively to technological changes within the year.

The paper is organised as follows. Section 2 provides a brief overview of the Belgian labour market institutions, introduces the data and describes the methodology. Section 3 presents our main results. Robustness tests with respect to alternative measures of TFP and specifications are discussed in Section 4, while Section 5 concludes. Technical details on the construction of the dataset and measurement of TFP shocks are included in Appendix A.

2. Institutions, data and methodology

2.1 Institutional features of the Belgian labour market

In this section, we briefly introduce the main features of the Belgian labour market that are relevant for the interpretation of our results. Notable characteristics of the wage formation process in Belgium include the minimum wage, automatic indexation, a cap on average wage increases, and sectoral collective bargaining. As far as employment is concerned, strict employment protection may be eased by early retirement, temporary unemployment, as well as overtime.

Sector-level collective wage bargaining between trade unions and employers' representatives plays a major role in the wage formation process and concerns the vast majority of firms.[2] Wage setting in Belgium may be described as the outcome of three mechanisms. First, a prominent feature of the Belgian labour market is full automatic indexation of nominal gross wages to the so-called health index, which is the consumer price index excluding alcoholic beverages, tobacco and motor fuels. This impedes real wage reductions of job stayers through the pace of inflation. Second, the so-called wage norm, set at the national level, is a recommendation for a maximum nominal hourly labour compensation increase. It is set by an interprofessional agreement for two years and takes into account, among others, the predicted indexation and evolution of labour costs of Belgium's main trading partners (namely Germany, France and the Netherlands). Third, sector-level agreements, typically organised separately for white-collar workers and blue-collar workers,

[2] According to the Belgian WDN survey (Druant et al. (2008)), bargaining at sectoral level concern 98% of all firms. This collective bargaining determine various aspects of compensation, such as pay scales and real wage increases, as well as other aspects, such as training or mobility. Pay scales set a minimum wage by sector and occupation and vary with age or tenure for white-collar workers and some blue-collar workers. As a result of the EU's anti-discrimination rules, the relationship with age is less common today.

specify real wage increases, which often consist of an absolute rise in the minimum pay scale. On top of this, some companies have developed firm-level wage bargaining. These individual agreements are not common in Belgium and typically lead to higher earnings.[3] Note that union representation and involvment within the firm is compulsory for companies employing 50 workers or more, and they also have to have among others works councils.[4] Union participation is stronger and more structured in firms employing 100 and more people.[5]

These features explain why Belgium has substantial real wage rigidity. However, it should be noted that labour compensation involves extra-wage components such as bonuses, premiums and overtime hours, which make total compensation more flexible than the base wage.

Employment developments over the last decade have been characterised by changes in the labour force composition. The trends include a smaller proportion of blue-collar workers in private sector employment (from 54% in 1990 to 49% in 1997 and 46% in 2005 according to Social Security statistics), an increasing fraction of part-time workers (accounting for 13.5% of employment in 1990, 16.3% in 1997 and 18.1% in 2005 (OECD (2002, 2004, 2006)), fewer hours worked per employee (the annual number of hours worked per employee fell from 1,546 in 1999 to 1,534 in 2005 (OECD (2004, 2006))) and a slightly higher number of employees with fixed-term contracts. Fixed-term contracts represent only a small proportion of wage earners in Belgium, 6.3% in 1997 and 8.8% in 2005, in comparison with EU average of 12% in 1997 and 14% in 2005 (Eurostat New Cronos).

Among the OECD member states, Belgium has a slightly higher level of employment protection legislation than the OECD average. This results from below average protection of regular employment and above average protection of temporary jobs and specific requirements for collective dismissals (see OECD (2004)). On the other hand, flexibility of the labour market is enhanced by early retirement and temporary unemployment. For firms in distress or in the process of restructuring, early retirement is possible under specific conditions for workers aged 50 and over. For short periods, temporary unemployment allows firms to temporarily interrupt, but not breach, labour contracts. Workers then receive unemployment benefit for a defined period and are later re-employed by the same firm under the initial contract terms. Together with changes in the number of hours (e.g. due to overtime hours), temporary unemployment makes it possible to reduce the number of hours worked, and avoid costly layoffs, as does early retirement.

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3 From the Structure of Earnings Survey data for the years 1999 to 2005, in the manufacturing, construction and business service sectors, 16% of companies have a firm-level agreement (either for blue-collar workers, or for white-collar workers, or for both). Companies with firm-level collective wage agreement tend to pay 15% higher wages than firms with no firm-level collective wage agreement. This figure rises to 20% if one includes irregular payments such as bonuses and premiums. In addition, the standard deviation in individual earnings is 5% to 10% higher in companies with firm-level agreement, but the coefficient of variation is 8% lower.

4 The works council is jointly composed of employees representatives and management staff. Its aim is to provide a forum for consultation and negotiation between employers' and employees' representatives.

5 In firms with a workforce of more than 100, employees' representatives in the works council have to be elected every four years; in smaller firms the representatives mandate is simply renewed.

2.2 Data

The main variables of interest related to labour compensation (total wage bill, number of employees, total hours worked) are taken from firms' annual accounts. Almost all firms in Belgium have to file their annual accounts with the Central Balance Sheet Office. However, we focus on the manufacturing, construction and market services sectors, and we consider only firms with at least 50 employees. This accounts for 87% of total employment of our cleaned sample and represent the vast majority of jobs. We prefer to disregard smaller firms because they may have different employment dynamics.[6] We perform a range of consistency checks to identify possible data issues and exclude extreme observations as outliers. Technical details are discussed in Appendix A.1. In our analysis, we estimate equations of employment, labour compensation, hours etc. by System GMM. To make sure that sufficient history is available to build lagged instruments, we consider only spells with at least 6 consecutive observations per firm and the variables in levels. Last, we exclude sectors with either too few observations to estimate the production function, from which our measure of TFP is derived, or sectors with production function coefficients substantially different from the income shares. Altogether the data set contains 10.771 firm-year observations on 1.518 firms with more than 50 employees over the period 1997-2005. Table A1 in Appendix A provides more details on the composition of the dataset across the sectors considered in the paper. Basic descriptive statistics on the variables described below are given in Table A.2.

The real wage bill of firm *i* at time *t* is denoted as WB_{it} and includes total remuneration and direct social benefits deflated by sector-specific value added prices. Employment, abbreviated as L_{it}, is measured as the average number of employees in full-time equivalent positions over the year. Average labour compensation per firm (W_{it}) is simply calculated as the ratio of the total real wage bill to the average number of employees over the year in full-time equivalent. Total hours worked over the year for each firm are denoted as H_{it}. Value added per sector (VA_{st}) was obtained from national accounts statistics. Variables related to workforce composition, like the percentage of blue-collar workers ($\%BLUE_{it}$), the percentage of women ($\%WOMEN_{it}$) and the proportion of workers with fixed-length contract ($\%TEMP_{it}$), are provided in the social balance sheet, which has formed part of firms' annual accounts since 1996. This restricts the sample available to our study to the period 1997-2005.[7] The construction of capital stock (K_{it}) is based on the perpetual inventory method (see Appendix A.2 for details). In what follows, variables in lower case designate log transformation.

We measure average labour compensation per firm as its total labour compensation divided by the number of employees in full-time equivalent positions. This contrasts with empirical papers based on individual wages (such as Cardoso and Portela (2005), Guiso et al. (2005)). These studies focus on job stayers. Such analyses may underestimate the sensitivity of wages if the

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6 For example, Kaiser and Pfeiffer (2001) find that in Germany employment flexibility is lower for smaller firms, due to lesser use of recruitment dismissals, temporary employment contracts and overtime (short time) work. Our own estimates, although not reported, indicate that the elasticity of labour to TFP is higher for larger firms. This is also supported by the results of the WDN survey for Belgium (Druant et al. 2008).

7 We disregard the information for year 1996 due to data issues.

wages of job stayers are less flexible than those of new hires, for example because they are (partly) set by multi-period contracts. One advantage of our measure is that it also includes employees whose wages might be more easily adjusted than permanent job stayers, such as new entrants or workers on fixed-term contracts. Evidence that the wage of entrants or movers is more flexible than that of job stayers is provided by Fehr and Goette (2005) and Haefke et al. (2008). A potential disadvantage of our measure of average labour compensation is that it may vary with changes in the composition of the labour force. We account for this by including control variables related to workforce composition, namely the percentage of blue-collar workers, women, and workers under fixed-term contracts, in our equations. Note also that our measure of average labour compensation, i.e. the firm's average labour compensation per employee, may be more flexible than the base wage because it includes extra-wage components such as overtime hours, bonuses and premiums. Because fluctuations in hours per worker, reported in the social accounts, imply variations in labour compensation, that capture changes in compensation due to overtime hours or temporary unemployment in addition to the reaction of the wage, we also estimate an equation for hourly labour compensation defined as total labour compensation over total hours worked.

We attempt to capture the impact of sector-level collective agreements on each firm's average labour compensation. This is motivated by the considerable importance of sector-level collective agreements in the wage-setting process in Belgium and our estimates confirm their relevance for firms' average labour compensation. The variables are constructed as follows. The nominal index of collectively agreed nominal wage increases at the sector-level for blue-collar workers and white-collar workers, respectively, is published by the Ministry of Labour.[8] We deflate these by the corresponding sector-level value added deflator to obtain the real measure. We use the logarithm of the real index of collectively agreed wage increases for blue-collar workers and white-collar workers, I^{B}_{st} and I^{W}_{st}, respectively, and multiply these by the percentage of blue-collar workers and white-collar workers for each firm. The measure is not perfect because collectively agreed wage increases are set at a more detailed level (in terms of sectors, but also occupation and age or tenure).[9] Discrepancies with respect to the average labour compensation may capture the firm-specific pay policy but also reflect the fact that collective agreements do not apply to more flexible components of labour compensation which include bonuses, premiums and overtime hours paid.

Since the aim of the paper is to evaluate the response of labour compensation and employment to TFP, it is crucial to construct unbiased and consistent measures of productivity and avoid spurious correlation with labour. We estimate TFP through the method recently proposed by Ackerberg et al. (2006), who improve on several grounds the estimation procedures used by Olley and Pakes (1996) and Levinsohn and Petrin (2003). We take into account two important problems

8 Federal Public Service Employment Labour and Social Dialogue (FPS ELSD).

9 Note that collectively agreed nominal wage increases in Belgium are the result of two mechanisms: indexation and collective agreements concerning real wage increases. We do not attempt to estimate the latter, i.e. we do not try to discriminate between indexation and real wage increases negotiated within sectoral collective agreements. Rather, we evaluate the impact on the firm's labour compensation of wage increases triggered by the sector-level collective agreement that is decided outside the firm. From the point of view of the company, these costs have to be compared to the firm's real output prices. Therefore, we deflate the collectively agreed nominal wage increases by the value added deflator.

related to measures of TFP based on the residual of a production function. The first is a simultaneity bias arising from the fact that productivity shocks are likely to affect factor demand (Ackerberg et al. (2006)). The second is that, in addition to their impact on the demand for factors, productivity shocks may affect the rate of utilisation of production factors (Basu and Kimball (1997)).

Olley and Pakes (1996), Levinsohn and Petrin (2003) and Ackerberg et al. (2006) correct for the simultaneity bias by augmenting the production function equation with a proxy of technological shocks (based on capital and either investment or intermediate inputs). The procedures by Olley and Pakes (1996) and Levinson and Petrin (2003) are based on a two-step estimation. In the first step, the production function is estimated including the proxy for unobserved productivity to solve the simultaneity problem. Because capital appears in the proxy for productivity and as a production function factor, it is not identified. However, the equation provides an estimate of the labour coefficient. In the second step, the coefficient on capital is estimated, given the first-step estimate of the labour coefficient. The identification is based on the assumption that the current capital stock was built in the previous period and is independent of current productivity innovations.

Ackerberg et al. (2006) point out that when intermediate inputs are used to proxy unobserved productivity, as in the Levinsohn and Petrin (2003) methodology, the labour production coefficient cannot be identified in the first step if labour and intermediate inputs decisions are taken simultaneously. The problem is similar but less severe when investment is used to proxy unobserved productivity, as in Olley and Pakes (1996). Ackerberg et al. (2006) then propose an alternative estimation procedure in which all production function parameters are estimated in the second stage. Identification of the capital parameter is the same as in the Olley-Pakes and Levinsohn-Petrin procedures. Identification of the labour parameter is achieved under the assumption that lagged labour does not respond to current productivity shocks, contrary to current labour.

In this paper we adopt the Ackerberg et al. (2006) procedure. In addition, we correct the obtained measure of TFP for variable capacity utilisation. In order to deal with this problem, Basu and Kimball (1997) develop a structural model in which the rate of utilisation of labour can be proxied by hours per worker. Furthermore, we decompose TFP into a firm-specific or idiosyncratic TFP component, TFP_{it}, and a sector-level or aggregate TFP component, TFP_{st}. In short we regress the Ackerberg et al. (2006) measure of TFP on hours per worker and a full set of interactive sector and year dummies. The firm-level TFP corrected for variable utilisation rate, TFP_{it}, is obtained as the residual of this equation, and sector-level TFP, TFP_{st}, as the estimated values of the sector-specific time dummies. A thorough discussion, technical details and production function estimates are included in Appendix A.3.

2.3 Specification

We adopt a dynamic specification. This is standard in employment equations due for instance to adjustment costs (see Arellano and Bond (1991), Nickell and Nicolitsas (1999), Nickell and

Wadhwani (1991)). In the labour compensation equations case, the inclusion of lags of the endogenous variables may be motivated by multi-period contracts, wage smoothing, wage rigidity, or reference norms, for example. Dynamic wage equations have been used by Katay (2007) for the average wage bill per worker and by Guiso et al. (2005) for individual wages. In addition, from an empirical point of view, omitting lags of endogenous variables lead to serially correlated residuals. Equation (1) shows the baseline model that we estimate in Section 3:

$$y_{it} = \rho_1 y_{it-1} + \rho_2 y_{it-2} + \beta_1 tfp_{it} + \beta_2 tfp_{it-1} + \beta_3 k_{it} + \beta_4 \%BLUE_{it} + \beta_5 \%TEMP_{it} + \beta_6 \%WOMEN_{it} + \beta_7 L{>}100_{it} + \delta_i + \delta_{st} + \varepsilon_{it} \quad (1)$$

Variables in lower case are measured in logs and β_js are the coefficients to be estimated. A vector of dummy variables for the 14 sectors considered in the paper is denoted as δ_s, year dummies as δ_t and interactive year and sector dummies as δ_{st}. Firm fixed effects, δ_i, capture unobserved firm characteristics.

In equation (1) y_{it} denotes the dependent variable, which can be any of the following variables: average labour compensation (w_{it}), employment (l_{it}), hours (h_{it}), hours per worker ($h\text{-}l_{it}$), and hourly labour compensation ($wb\text{-}h_{it}$). Hence, we estimate dynamic equations for each component of the wage bill and we also provide estimates for total hours worked and the wage bill per hour worked. We include the same set of variables in all equations, which may therefore be viewed as reduced-form equations. We allow for sector-specific year dummies to capture aggregate sector-level conditions, such as aggregate demand or prices.

We control for the composition of the labour force by including the percentage of blue-collar workers, $\%BLUE_{it}$, the percentage of women, $\%WOMEN_{it}$, and the percentage of workers with fixed-term contracts, $\%TEMP_{it}$. In order to take into account the impact of firm size in our regressions, we include a dummy that is equal to one for firms with more than 100 employees, "$L>100_{it}$". This threshold is close to the median firm size in our sample. We favour this specification over one that would directly include firm size for two reasons. First, including the number of employees as a proxy for firm size would not be feasible in the employment equations for obvious perfect colinearity reasons, and would generate complex endogeneity problems in equations for average labour compensation. Second, the threshold can be motivated by the fact that union participation may be considered as more structured in firms with 100 employees and more (see section 2.1). Installed capital also enter the equation because it appears in labour demand equations under various sets of theoretical assumptions.[10]

[10] Assuming that firms are output constrained, employment depends on expected output and the relative capital-labour ratio. If firms are not output constrained and the capital is pre-determined, then employment depends on the pre-determined stock of capital and real wages.

In our alternative specification, the role of sector-specific variables is examined by replacing the sector-specific time dummies, δ_{st}, by a set of year dummies, δ_t, sector dummies, δ_s, and sector-level variables. These include sector-level TFP, sector-specific value added,[11] and weighted indices of wage increases for blue-collar workers and white-collar workers determined by sector-level collective agreements to capture the impact of sector-level collective agreement on firm' average labour compensation:

$$
\begin{aligned}
y_{it} = {} & \rho_1\, y_{it-1} + \rho_2\, y_{it-2} + \beta_1\, tfp_{it} + \beta_2\, tfp_{it-1} + \beta_3\, tfp_{st} + \beta_4\, tfp_{st-1} + \beta_5\, k_{it} + \beta_6\, va_{st} \\
& + \beta_7\, \%BLUE_{it} + \beta_8\, \%TEMP_{it} + \beta_9\, \%WOMEN_{it} + \beta_{10}\, L{>}100_{it} \\
& + \beta_{11}\, (\%WHITE_{it} * i^{W}_{st}) + \beta_{12}\, (\%BLUE_{it} * i^{B}_{st}) + \delta_i + \delta_s + \delta_t + \varepsilon_{it} \qquad (2)
\end{aligned}
$$

Equations (1) and (2) include firm-specific fixed effects, as is common in the literature. This implies that instrumental variables should be used to take into account endogeneity of the lagged dependent variable. The dynamic panel equations are estimated by the System GMM procedure proposed by Arellano and Bover (1995) and Blundell and Bond (1998). We report the two-step estimates with standard errors corrected by the Windmeijer (2004) procedure. We assume that TFP, firm size, labour force composition, sector-specific value added, and the impact of sector-level collective agreements on firms' wages are exogenous. Lags of the endogenous variable, capital stock and profits per worker are used as instruments.

We assume that TFP is exogenous. One may nevertheless argue that it is endogenous. For example, a demand shock might be correlated with our measure of TFP and with labour demand. However, lags of TFP, intermediate inputs and capital stock proved to be very poor instruments effectively making the TFP coefficient insignificant.[12] Therefore, we favour our specification that assumes TFP exogenous, although we are aware that TFP might be endogenous and reflect demand or factor price shocks together with productivity.

Note also that we do not take into account at this stage the fact that TFP is a generated regressor. Because of the uncertainty surrounding the estimates of production function coefficients used to construct TFP series, the standard errors of TFP reported in the tables below may be considered as a lower bound.[13]

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[11] Unemployment is often used as a determinant of wages. Because unemployment rates are not available at the sector level, we use a proxy for sector-level business conditions. Labour demand also depends on the sector output price (or sector demand) under monopolistic competition conditions. For both reasons we include the log of sector value added in equation (2).

[12] A least squares regression of changes in TFP on lags 4 and 5 of TFP, intermediate inputs and capital stock yield a R^2 of only 0.01. Consequently when estimating average labour compensation and labour equations by SGMM with TFP considered as endogenous, the TFP coefficient becomes insignificant. Results are available on request.

[13] We expect that estimating standard errors by bootstrap will not change our conclusions. The estimated effect of firm-specific TFP on labour compensation may turn insignificant (as is already the case in Table 1 below). But it is unlikely that the coefficients of sector-level TFP in the labour compensation equations, and those of firm-specific and sector-level TFP in the labour equations become insignificant because this would imply that the bootstrap standard errors are 3 times larger (or more) than the standard errors reported in the text.

3. Results

3.1 Estimating the elasticity of average labour compensation and labour to firm-specific TFP

In this section we compare the elasticity of labour compensation and labour to firm-specific TFP. We estimate equation (1) for average labour compensation, employment, hours worked, hourly labour compensation, as well as hours per worker. The results are reported in Table 1. The coefficients on control variables have the expected sign. Firms with a higher percentage of blue-collar workers and women have significantly lower average labour compensation, all else equal. Also, firms with a higher percentage of workers under fixed-term contracts have *ceteris paribus* lower average labour compensation. The capital stock has a positive coefficient in the employment equation, suggesting complementarities between the two production factors, capital and labour.

Our estimates indicate that the contemporaneous elasticity of average labour compensation to TFP is not significantly different from zero, while the elasticity of employment is large, 0.23. The sum of the coefficients on current and lagged TFP is 0.03 for average labour compensation and hourly labour compensation, which is four times smaller than the sum for employment (0.11) and total hours (0.10).

The elasticity of total hours worked, which accounts both for changes in hours per worker and changes in the number of employees, is of the same order of magnitude as that of employment. In the presence of adjustment costs in the short run, firms may adjust hours worked more easily than the number of employees, for example through overtime hours and temporary unemployment.[14] However, adjustment of hours per worker is very rare. Indeed, the contemporaneous elasticity of hours per worker to TFP is not significantly different from zero. This means that firms adjust labour to firm-specific productivity developments mainly through the extensive margin, rather than the intensive margin.

These results are consistent with the survey evidence in Druant et al. (2008). This indicates that, when reducing costs following an adverse shock, 60% of Belgian firms declare that they reduce employment, while only 14% of the companies adjust pay (and only do so through the variable components), while a very small proportion of enterprises actually reduce working time.[15]

Note that comparing the very low elasticity of average labour compensation to the substantial elasticity of labour to firm-level TFP supports the hypothesis of real wage rigidity in Belgium. In general, models with wage rigidity typically find wider variability of employment in response to productivity shock, as compared to the flexible wage scenario (see for example Hall (2005) and Blanchard and Galí (2007, 2008)) so that labour productivity can match the real wage. However, our estimates do not provide a test or a measure of real wage rigidity. There is no theoretical reference value for the average labour compensation elasticity and labour elasticity under the

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[14] See Fuss (2008) for evidence that variations in hours per worker and the number of days worked are significantly lower in cases of falling sales and wage bill contractions.

[15] For comparison, Bertola et al. (2008) report that on average over 15 European countries, around 30% of firms declare that they reduce employment, 11% of the firms reduce pay, and up to 7% cut working time.

flexible wage case. In the model of Blanchard and Galí (2008) without labour market frictions, labour does not respond to TFP under perfect wage and price flexibility. But this results from the fact that, in their model, income and substitution effects cancel each other out.

Our estimates of the elasticity of average labour compensation to TFP are partly related to the investigation of wage insurance. Following Guiso et al. (2005) for Italy, a number of authors decompose the sensitivity of wages to productivity into that due to permanent productivity changes and the other due to transitory productivity changes (see Cardoso and Portela (2005) for Portugal and Katay (2007) for Hungary). We estimate the average effect of TFP, while these papers go one step further by decomposing the average effect into the permanent and transitory part.[16] A zero response of wages to transitory changes in productivity is interpreted as evidence that risk-neutral firms insure risk-averse workers against wage fluctuations because wage insurance may apply to transitory but not to permanent shocks. Our finding of a zero elasticity of average labour compensation *a fortiori* suggests that permanent (if any) and transitory idiosyncratic TFP changes are insured. This interpretation translates into firm average wages provided that we appropriately control for changes in the composition of the labour force. Further, our estimates of the elasticity of labour is consistent with the following interpretation. In a contract model where firms insure workers against income fluctuations, Boldrin and Horvath (1995) show that the response of wages to shocks is smoother, and the volatility of hours worked is higher than in a situation without wage insurance.

Lastly, the low response of firms' wages to firm-specific shocks may be explained by labour market competition, efficiency wage considerations and product market competition. For example, in a tight labour market it may not be desirable for a company to reduce wages following a negative productivity shock, because it makes other companies more attractive for its workers. Further, this may generate adverse selection problems. This argument would explain why the firms' average labour compensation response to firm-level shocks is low, after controlling for workforce composition. Further, in a competitive product market, firms are price-taker and may not afford wage increases because they will not be able to raise their prices if their competitors do not.

[16] We do not attempt to decompose the effect of TFP into transitory and permanent effects because in our case the average (over permanent and transitory) effect is zero.

Table 1 - SGMM estimates of equation (1)

	w_{it}	l_{it}	h_{it}	$wb\text{-}h_{it}$	$h\text{-}l_{it}$
dep. var_{it-1}	0.70***	1.17***	1.07***	0.63***	0.28**
	(0.11)	(0.13)	(0.15)	(0.09)	(0.12)
dep. var_{it-2}	0.09	-0.28**	-0.18	0.24***	0.41***
	(0.09)	(0.12)	(0.13)	(0.08)	(0.08)
tfp_{it}	0.02	0.23***	0.18***	0.03*	-0.01
	(0.02)	(0.04)	(0.04)	(0.01)	(0.01)
tfp_{it-1}	0.01	-0.12***	-0.08***	0.00	0.01
	(0.02)	(0.03)	(0.03)	(0.01)	(0.01)
k_{it}	0.01	0.02***	0.02***	0.00	0.00
	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)
$\%BLUE_{it}$	-0.15***	0.03**	0.01	-0.08***	-0.04***
	(0.06)	(0.01)	(0.01)	(0.03)	(0.01)
$\%TEMP_{it}$	-0.07***	0.27***	0.27***	-0.07***	0.06**
	(0.03)	(0.06)	(0.06)	(0.02)	(0.03)
$\%WOMEN_{it}$	-0.08**	0.03*	0.01	-0.03	-0.04***
	(0.04)	(0.01)	(0.01)	(0.02)	(0.01)
$L>100_{it}$	-0.01	0.09***	0.09***	0.00	0.00
	(0.01)	(0.03)	(0.02)	(0.00)	(0.00)
Sargan	84.62	78.04	76.33	56.71	64.69
p-value	(0.02)	(0.05)	(0.06)	(0.56)	(0.29)
AR(1)	-4.59	-4.66	-3.71	-3.95	-2.08
p-value	(0.00)	(0.00)	(0.00)	(0.00)	(0.04)
AR(2)	0.20	-0.41	-0.20	-1.49	-4.15
p-value	(0.85)	(0.68)	(0.84)	(0.14)	(0.00)

Note: Firms with at least 50 employees and 6 consecutive annual accounts. 1.518 firms and 6.217 observations. Two-step System GMM estimates are reported with standard errors in parentheses following the correction proposed by Windmeijer (2004). The lagged dependent variable (denoted as dep. var_{it-1}) and the capital stock are treated as endogenous and instrumented with the Arellano-Bond instrument matrix with lags t-4 and earlier, as well as profit per worker. The remaining regressors are treated as exogenous. All equations include interactive sector and year dummies but their coefficients are not reported. AR displays the test for serial correlation in the first-differenced residuals. Lower case variables are in log. The remaining variables are defined in the text. * indicates significance at the 10% level, ** at the 5% level, *** at the 1% level.

These arguments may explain why the sensitivity of average labour compensation to TFP is low. But it does not explain differences across countries,[17] unless there are large differences in the degree of wage insurance, wage rigidity, or competition on the labour market. Compared to previous estimates, Belgium has the lowest elasticity of average labour compensation to TFP. In contrast, there is evidence of insurance against transitory shocks in Italy and Portugal, but the response to permanant shock is above zero. In the case of Hungary, the response of wages to

[17] Although previous microeconomic estimates are not fully comparable due to different approaches, data and sample definitions, they still provide a basis for comparison. Below, we consider the main differences with the results of Guiso et al. (2005), Cardoso and Portela (2005) and Katay (2007). First, Guiso et al. (2005) focus only on the manufacturing sector. Our own preliminary estimates indicate that the elasticity of wages to productivity is lower when services are included in the sample. Second, papers based on individual wage data restrict their sample to job stayers. However Fehr and Goette (2005) and Haefke et al. (2007) report that entrants' wages are more flexible than incumbents' wages. Hence, focusing on job stayers may bias the estimated elasticity downwards. Third, considering earnings per worker or hourly earnings may yield different results when hours per worker are adjusted to changes in productivity.

productivity is the highest but also there is no evidence of full insurance against temporary fluctuations in productivity.[18]

There are several reasons for differences between the results for Hungary and other Western European countries (see Katay (2007)). Among others, according to the OECD (2004), Hungary is among the countries with the most flexible labour market. Company and plant-level agreements are the dominant form of wage bargaining, with no coordination by upper-level associations and no centralisation. Further, the coverage of collective agreements is very low.[19] Therefore, firms have more freedom to change wages in response to TFP shocks. In contrast, sector-level agreements are common in the remaining European countries, combined with firm-level agreements (in France and Italy), or central agreements (in Portugal and Belgium), and a medium or high defree of coordination with high coverage of collective agreements.

Our results may be summarised as follows. First, labour has substantially high elasticity to firm-specific TFP; while conversely average labour compensation are insensitive to idiosyncratic TFP.[20] Although the elasticity of labour under perfect wage and price flexibility has no natural reference value, this finding is consistent with significant real wage rigidity in Belgium that shifts the burden of adjustment towards employment. It is also consistent with the importance of sector-level collective agreements in the wage formation process, as will be discussed below. Second, following a TFP change, firms adjust labour mainly on the extensive margin. Third, our estimates point to a positive relation between hours and TFP. The following two subsections examine in more detail whether the finding translates into sector-level TFP, the role of collective wage agreements in shaping the response of average labour compensation, and also discuss our finding of a positive elasticity of hours to TFP.

3.2. Differences in the elasticity to firm-specific TFP and sector-level TFP

Here we compare the response of labour compensation and labour to firm-specific TFP and to sector-level TFP. While the previous section highlights a close-to-zero and non-significant response of average labour compensation to firm-specific TFP, the same may not hold with respect to sector-level TFP. First, one of the reasons not to adjust wages downwards is the fear that the best workers leave for better-paying companies. In the Belgian WDN survey by Druant et al. (2008), more than 80% of firms with more than 50 employees report that this is a motive not to cut wages. However, the argument does not hold when all firms undertake wage contractions at the same time, as opposed to a single company doing so unilaterally. Also, if competition on the product market is strong, i.e. if markups are small or if firms are price takers, they cannot undertake

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[18] The estimated permanent effect of productivity is 0.11 for Hungary against 0.07 for Italy and 0.09 for Portugal. The estimated transitory effect of productivity is 0.05 for Hungary and not significant in other countries.

[19] According to Traxler and Behrens (2002), the percentage of employees covered by collective agreements (i.e. including not only union members but also other employees to whom the union-negotiated contract applies) is around 90% in Belgium, France, Italy and Portugal and 36% in Hungary (data for 2001).

[20] As shown in Section 4, this result is robust to alternative definitions of TFP (controlling or not for variable utilisation rate and considering TFP shocks rather than TFP in level).

isolated wage increases because they cannot raise their prices without incurring losses unless their competitors follow the suit. Lastly, it has been argued that centralisation of wage bargaining may ease wage contraction (see Bruno and Sachs (1985)), thanks to coordination of decisions and internalisation of the externalities of individual actions. We examine this last issue in the next sub-section.

Estimates including firm-level as well as sector-level TFP are reported in Table 2 below. We therefore estimate equation (2) with additive year and sector dummies, instead of sector-specific year effects. To account for other fluctuations at the sector level we also include value added per sector.

Table 2 - SGMM estimates of equation (2)

	w_{it}	l_{it}	h_{it}	wh_{it}	h/L_{it}
dep. var_{it-1}	0.84***	1.32***	1.09***	0.90***	0.33***
	(0.12)	(0.13)	(0.15)	(0.16)	(0.12)
dep. var_{it-2}	-0.07	-0.43***	-0.19	0.00	0.40***
	(0.09)	(0.11)	(0.13)	(0.13)	(0.07)
tfp_{it}	0.03**	0.23***	0.18***	0.04***	-0.01
	(0.02)	(0.05)	(0.04)	(0.02)	(0.01)
tfp_{it-1}	0.00	-0.13***	-0.09***	-0.02	0.01
	(0.02)	(0.03)	(0.03)	(0.02)	(0.01)
tfp_{st}	0.54***	0.13***	0.17***	0.49***	0.05**
	(0.04)	(0.05)	(0.05)	(0.04)	(0.02)
tfp_{st-1}	-0.30***	-0.06	-0.10**	-0.27***	-0.06**
	(0.06)	(0.05)	(0.05)	(0.08)	(0.03)
k_{it}	0.00	0.02**	0.02***	0.00	0.00
	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)
y_{st}	-0.12***	0.01	0.02	-0.07*	0.00
	(0.04)	(0.04)	(0.04)	(0.04)	(0.02)
$\%BLUE_{it}$	-0.16***	0.02**	0.01	-0.06**	-0.04***
	(0.06)	(0.01)	(0.01)	(0.03)	(0.01)
$\%TEMP_{it}$	-0.08***	0.29***	0.25***	-0.06***	0.05*
	(0.03)	(0.06)	(0.05)	(0.02)	(0.03)
$\%WOMEN_{it}$	-0.09**	0.03*	0.01	-0.02	-0.03***
	(0.04)	(0.02)	(0.01)	(0.02)	(0.01)
$L>100_{it}$	-0.01	0.09***	0.08***	0.00	0.00
	(0.01)	(0.03)	(0.02)	(0.00)	(0.00)
Sargan	98.62	82.26	76.22	70.84	58.48
p-value	(0.00)	(0.02)	(0.07)	(0.14)	(0.50)
AR(1)	-5.29	-5.39	-3.72	-3.54	-2.13
p-value	(0.00)	(0.00)	(0.00)	(0.00)	(0.03)
AR(2)	2.07	0.25	-0.24	0.87	-4.10
p-value	(0.04)	(0.80)	(0.81)	(0.39)	(0.00)

Note: Firms with at least 50 employees and 6 consecutive annual accounts. 1.518 firms and 6.217 observations. Two-step System GMM estimates are reported with standard errors in parentheses following the correction proposed by Windmeijer (2004). The lagged dependent variable (denoted as dep. var_{it-1}) and the capital stock are treated as endogenous and instrumented with the Arellano-Bond instrument matrix with lags t-4 and earlier, as well as profit per worker. The remaining regressors are treated as exogenous. All equations include additive sector and year dummies but their coefficients are not reported. AR displays the test for serial correlation in the first-differenced residuals. Lower case variables are in log. The remaining variables are defined in the text. * indicates significance at the 10% level, ** at the 5% level, *** at the 1% level.

The most striking result is that the elasticity of average labour compensation to sector-level TFP is very large compared to that with respect to firm-specific TFP. The sum of the coefficients on current and lagged sector-level TFP for labour compensation is much larger (0.22 to 0.24) than that for firm-specific TFP (0.02 to 0.03). This holds for both unit or hourly labour compensation. By contrast, the sum of the coefficients on current and lagged aggregate TFP for the number of employees and hours worked equations is slightly smaller (0.07) than that for idiosyncratic TFP (0.09 to 0.10).

The picture that emerges from these results is one of sluggish average labour compensation and large employment fluctuations in response to idiosyncratic TFP but more flexible average labour compensation and less sensitive labour in response to aggregate fluctuations. One interpretation of these results is that firms in Belgium are bound by sector-level collective wage agreements and tend not to deviate too much from them. This hypotheses is examined in the next section.

3.3. The role of sector-level collective wage bargaining

Collective bargaining plays a dominant role for wage-setting in Belgium. First, an indicative norm for maximum nominal hourly labour cost increases is set at the national level. Then, given expected indexation, sector-level agreements decide on real increases of the base wages or minimum pay scale. In order to illustrate the role of sector-level collective bargaining in shaping the response of average labour compensation to TFP, Table 3 below reports three sets of estimates for average labour compensation. Column (1) is directly taken from Table 2. Column (2) includes the impact of sector-level collective wage agreements on firms' average labour compensation, but omits sector-level TFP (we do not include the two together for colinearity reasons explained below).

The impact of sector-level collectively agreed wage increases at the firm level is positive and significant. The point estimates imply that a one percent increase in the collectively agreed wage induces firms to raise average labour compensation on average by 0.65 and 0.68 percent for blue-collar and white-collar workers, respectively. Note that once sector-specific time dummies are allowed for (column (3)), the coefficients for sector-level collectively agreed wage increase, $\%WHITE_{it}.i^{W}_{st}$ and $\%BLUE_{it}.i^{B}_{st}$, become smaller and insignificant. This suggests that the impact of sector-level bargaining on firms' average labour compensation ($\%WHITE_{it}.i^{W}_{st}$ and $\%BLUE_{it}.i^{B}_{st}$) is driven mostly by sector-level agreements, i^{W}_{st} and i^{B}_{st}, rather than by the firm workforce composition.[21]

[21] The result is driven by the fact that wage increases set by sector-level collective agreement follow the same trend for blue-collar and white-collar workers from the same sector but different patterns across sectors.

Table 3 - SGMM estimates of equation (2) for average labour compensation

	(1)	(2)	(3)
w_{it-1}	0.84***	0.70***	0.61***
	(0.12)	(0.18)	(0.18)
w_{it-2}	-0.07	0.15	0.21
	(0.09)	(0.15)	(0.15)
tfp_{it}	0.03**	0.02	0.02
	(0.02)	(0.02)	(0.02)
tfp_{it-1}	0.00	0.00	0.01
	(0.02)	(0.02)	(0.02)
tfp_{st}	0.54***		
	(0.04)		
tfp_{st-1}	-0.30***		
	(0.06)		
$\%WHITE_{it}.i^{W}_{st}$		0.68***	-0.81
		(0.12)	(0.59)
$\%BLUE_{it}.i^{B}_{st}$		0.65***	-0.80
		(0.09)	(0.59)
k_{it}	0.00	0.00	0.00
	(0.01)	(0.01)	(0.01)
y_{st}	-0.12***	-0.09	
	(0.04)	(0.06)	
$\%BLUE_{it}$	-0.16***	-0.12**	-0.11**
	(0.06)	(0.06)	(0.05)
$\%TEMP_{it}$	-0.08***	-0.09***	-0.09***
	(0.03)	(0.03)	(0.03)
$\%WOMEN_{it}$	-0.09**	-0.04	-0.04
	(0.04)	(0.03)	(0.03)
$L>100_{it}$	-0.01	0.00	0.00
	(0.01)	(0.01)	(0.01)
year and sector dummies	δ_s and δ_t	δ_s and δ_t	δ_{st}
Sargan	98.62	51.26	54.39
p-value	(0.00)	(0.11)	(0.06)
AR(1)	-5.29	-2.87	-2.51
p-value	(0.00)	(0.00)	(0.01)
AR(2)	2.07	-0.22	-0.64
p-value	(0.04)	(0.83)	(0.52)

Note: Firms with at least 50 employees and 6 consecutive annual accounts. 1.518 firms and 6.217 observations. Two-step System GMM estimates are reported with standard errors in parentheses following the correction proposed by Windmeijer (2004). The lagged dependent variable (denoted as dep. var_{it-1}) and the capital stock are treated as endogenous and instrumented with the Arellano-Bond instrument matrix with lags t-4 and earlier, as well as profit per worker. The remaining regressors are treated as exogenous. All equations include additive sector and year dummies but their coefficients are not reported. AR displays the test for serial correlation in the first-differenced residuals. Lower case variables are in log. The remaining variables are defined in the text. * indicates significance at the 10% level, ** at the 5% level, *** at the 1% level.

We then perform a simple back-of-the-envelope calculation of the impact of TFP through sector-level collective agreements. OLS regressions of changes in the log of the indices of sector-level collectively agreed wage increases (deflated by value-added prices) on changes in sector-level TFP per sector confirm that there is a highly significant and positive relationship between TFP and collective wage increases at the sector level - the coefficient is equal to 0.38 for white-collar workers and 0.39 for blue-collar workers. This confirms the conjecture that productivity developments are taken into account in wage-setting practice in Belgium. Importantly, this suggests that the impact of TFP on average labour compensation is not zero, but labour compensation in Belgium adjusts mainly through collectively agreed wage increases, which take into account sector-level common productivity evolutions rather than idiosyncratic or firm-specific TFP changes. A simple calculation suggests that the impact of sector-level TFP on firm-level average real labour compensation per worker is 0.26.[22] This is close to the sum of the coefficient on sector-level TFP in column (1), 0.24. For the sake of comparison, in a structural VAR model for the US, Ravn and Simonelli (2007) find that the impact of neutral technology shocks on real wages is 0.15 after a year. The estimates of the elasticity of wages aggregated over all workers to labour productivity for the US over the period 1984-2006 obtained by Haefke et al. (2007) range from 0.17 to 0.37.

All in all, our results point to the fact that firms have little room for adjusting their average labour compensation to firm-specific developments but respond to sector-level TFP via sector-level collective bargaining. Firms may not deviate too far from sector-level collective agreements for workers already employed by the firm, i.e. job stayers. In addition, bonuses and premiums generally do not account for a substantial proportion of earnings in Belgium.[23] Lastly, one might argue that firms could adjust their average wage bill by applying a different pay scheme to entrants and workers under fixed-term contracts. However, the percentage of workers under fixed-term contracts in Belgium is below the average for Europe (see Section 2.1). In addition, minimum pay scale defined at the activity, occupation and tenure level, within sector-level collective agreements, provide a lower bound for new entrants' wages.

3.4 The hours-TFP relationship

After having examined the sensitivity of average labour compensation to TFP we now turn to the hours-TFP relationship. In the macroeconomic literature, there is some controversy about the sign of the relationship between hours and technology. Few papers investigate this issue at the microeconomic level. In this section, we first review the macroeconomic debate, then we describe previous microeconomic estimates. We then go on to discuss the differences between macroeconomic and microeconomic estimates. Lastly, we report our own estimates of the elasticity of hours with respect to TFP and compare with other microeconomic evidence.

As mentioned in the introduction, under monopolistic competition conditions in the product market, the short-run response of hours to technology shocks should be positive when prices are flexible, and negative in the case of sticky prices. Additional explanations for a negative relationship between hours and technology shocks include a short-run negative effect on production due to

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22 This is obtained as follows. Differentiating equation (2) for average labour compensation as a dependent variable with respect to the index of collectively agreed wage increases for blue-collar and white-collar workers leads to $\Delta w_{it} = \beta_{11}$ %WHITE$_{it}$ * Δi^{W}_{st} + β_{12} %BLUE$_{it}$ * Δi^{B}_{st}. An OLS regression of Δi^{W}_{st} (Δi^{B}_{st}) on Δtfp_{st} yields a coefficient of 0.38 for white-collar workers and 0.39 for blue-collar workers. Therefore, given the average percentage of blue-collar workers in our sample (0.38) and the coefficient estimates in Table 2, the impact of tfp_{st} on firm-level wages is equal to 0.26.

23 Data from the Belgian Structure of Earnings Survey (SES) indicates that bonuses form on average 8.4 percent of earnings. The proportion varies from 2.4 percent in hotels and restaurants to 13.3 percent in financial services.

adaptation to the new technology, habit formation in consumption and adjustment costs in investment.

The macroeconomic empirical evidence is not conclusive. Starting with Galí (1999), a number of structural VAR analyses point to a negative short-run impact of neutral technology shocks on hours worked. Francis and Ramey (2005) find that this result is robust to alternative VAR specifications and identification schemes. Smets and Wouters (2007) confirm this finding in a DSGE model with price and nominal wage stickiness. Basu at al. (2006) find that their growth accounting measure of technology has a negative impact on hours within the quarter, but a positive effect after one year. Using both VAR and growth accounting measures of technological change, Alexius and Carlsson (2007) find that technology shocks are positively correlated with output growth and negatively correlated with changes in hours worked. Galí's finding has been challenged on several grounds. It has been argued that the effect of productivity on hours turns positive if one assumes that hours are stationary rather than difference stationary (Christiano et al. (2003, 2004)), if one allows for an investment-specific technology shock (Ravn and Simonelli (2007)), or if one allows shocks other than technology changes to have a long-run effect on labour productivity (Dedola and Neri (2007)).

Few recent papers evaluate this question at the microeconomic level. Based on firm-level data, Marchetti and Nucci (2005, 2007) and Carlsson and Smedsaas (2007) evaluate the impact of growth accounting productivity measures on total hours worked. These papers suggest that current TFP shock has a negative impact on hours, although the effect of past TFP is positive and compensates for the initial negative effect. The main finding of Marchetti and Nucci (2005, 2007) is that the effect of a TFP shock on hours within the year is more negative for firms with stickier prices.

Note that there are several reasons why the elasticity might differ when macroeconomic productivity is considered. First, microeconomic exercises typically evaluate the impact of idiosyncratic productivity shocks, while macroeconomic or sector-level analyses consider the response to a common aggregate change. A new secret recipe or a patented innovation can serve as examples of idiosyncratic productivity shocks. The innovating firm will then have a productivity advantage over its competitors. The introduction of new software by Microsoft would be a common productivity shock. In principle all firms have access to this technological improvement. Firms might react differently to idiosyncratic and common shocks. An individual firm might have different incentives and varying ability to change its price when it is the only one facing the shock than if the shock is common to all firms.

Second, macroeconomic investigations rely on structural estimations of technology shocks, typically assumed to have long-run impact on output, while microeconomic studies rely on TFP growth accounting measures. One traditional caveat of these measures is that nominal variables are deflated using sector-level price indices rather than firm-level output prices, due to the lack of relevant data. As shown in Klette and Griliches (1996), this may bias the production function coefficient estimates. Further, TFP measures may capture demand shocks (Foster et al., 2008) or variation in factor prices (Katayama et al., 2003), together with technological changes. Demand shocks will induce a positive correlation between TFP and hours which is not related to technological change and may not be present in macroeconomic models.

Lastly, macroeconomic series of hours cannot be directly compared to the microeconomic data due to different weights and composition. Firstly, in panel data all firms have the same weight, while in a macroeconomic series firms are implicitly weighted by their size. Secondly, the sample composition may be different. Using aggregate productivity decompositions of the type developed by Baily et al. (1992) and Foster et al. (2001), several authors have found evidence that aggregate productivity fluctuations are due not only to productivity developments of continuing firms, but also to a substantial extent to reallocation of workers across establishments, including entries and exits of plants or firms. Following a positive common TFP shock, some companies may immediately invest in the new technology, increase output and total hours and expand their market share at the expense of some other enterprises that might be driven out of the market. In such cases, hours worked in these firms will of course drop. What is observed in our panel dataset are only surviving firms (with 50 and more employees, six consecutive years and active over the entire calendar year). On the contrary, macroeconomic series aggregate hours across all firms. This may drive the response of aggregate hours towards more negative values.

The debate on the sign of the hours-technology relationship focuses on the short-run response of hours. In the medium run, for example once prices adjust, hours increase following a productivity shock. With the above caveats in mind, note that our estimate of the within-the-year response of hours to aggregate sector-level TFP reported in Table 2 is positive.

Comparing our estimates of idiosyncratic elasticity to previous microeconomic studies, one important difference between our results and Marchetti and Nucci (2005, 2007) and Carlsson and Smedsaas (2007) is their finding that the effect of current TFP shocks on hours is negative, while our results point to a positive contemporaneous effect.

Several factors may explain the difference in the estimated sign of the hours-technology relationship. We first examine differences due to methodology or data issues. First, we consider the level of TFP, while Marchetti and Nucci (2005, 2007) consider TFP shocks estimated from an AR(2) regression on TFP. Replacing TFP level by AR(2) shocks leaves the estimated elasticity essentially unchanged, as shown in the robustness tests in part (2) of Table 4 in Section 4.

Second, the TFP measure used by Marchetti and Nucci (2005, 2007) and Carlsson and Smedsaas (2007) takes into account variable utilisation of production factors which can be approximated by hours per worker, following the structural approach of Basu and Kimball (1997). We adopt a more empirical approach and simply regress the Ackerberg et al. (2006) TFP measure of hours per worker to clean TFP for variable utilisation rate. When considering a TFP measure not corrected for hours per worker, we still obtain a positive elasticity with respect to TFP (see part (3) of Table 4 in Section 4). And the results in Marchetti and Nucci (2007) using Olley and Pakes' estimation procedure that also neglects variable utilisation still lead to a negative contemporaneous effect of TFP on hours.[24]

Third, we also examine potential omitted variable bias. The above-mentioned papers only include year and sector dummies as control variables. If we include only two lags of hours, current and lagged TFP in the hours equation with interactive or additive year and sector dummies, the coefficient on current TFP changes is smaller (equal to 0.10) but remains significant. The same result applies when estimating the hours equation with no lags of hours by OLS in difference.

Fourth, the difference in the results may be due to the measurement of labour. Carlsson and Smedsaas (2007) find a negative impact of TFP on average employment but a zero impact on employment measured at the end of the year. This is consistent with the result in Basu et al. (2006) that, following a TFP shock, hours diminish within the quarter but increase after a year. These results indicate that the probability of finding a negative relationship is lower when using annual data. Note that in our estimation, we consider the average employment over the year, so the argument does not apply.

All in all, we consider our finding that the response of hours to TFP is positive within the year to be robust. Below, we examine whether the different results can be due to differences across the countries. As already discussed above, real wage rigidity shifts the burden of adjustment towards labour following a productivity shock, so that labour productivity can match the real wage. On the other hand, sticky-price models, such as Galí's (1999), imply that hours respond negatively to TFP shocks because demand and therefore output remain constant under unchanged prices. In sum real wage rigidity tends to produce a positive relationship between labour and productivity, while price stickiness goes in the opposite direction. It is nevertheless not clear how this could alter the sign of the labour-technology relationship, given the evidence of both wage and price rigidity.

Firstly, Belgium would have to have much more flexible prices than Italy in order to explain the difference in the sign and size of the hours-technology relationship between our results and those of Marchetti and Nucci (2005, 2007). However, the recent findings of the Inflation Persistence Network summarised by Dhyne et al. (2006) report an average (weighted) frequency of consumer price changes equal to 0.15 for Belgium and 0.12 for Italy.[25] It is hard to believe that this alone can explain the difference in the sign of the hours-technology relationship between our estimates and those of Marchetti and Nucci (2005, 2007).

Secondly, our estimates support the hypothesis of real wage rigidity in Belgium. Our results point to a very small (and hardly significant) response of average labour compensation in Belgium, while the response of individual wages in Italy is not significantly different from zero in response to

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24 Note also that the Olley and Pakes (1996) methodology is closer to that used in this paper than Basu and Kimball's (1997) approach, in the sense that production function coefficients are estimated neglecting fixed effects. In the Basu and Kimball (1997) methodology, they are estimated by GMM, thus allowing for fixed effects.

25 Vermeulen et al. (2007) report a higher average (weighted) frequency of producer price changes for Belgium (0.24) than for Italy (0.15). However, part of the lower frequency in Italy can be explained by the absence of energy products that typically have the highest frequency of price changes. In general, one finds stronger heterogeneity in price adjustment across sectors than across countries.

transitory productivity shocks, but amounts to 0.07 in response to permanent productivity shocks.

In sum, this section present robust evidence that at the firm-level hours respond positively to TFP. To judge whether this conclusion translates to macroeconomic aggregates would require a different exercise.

4. Robustness tests

Our base measure of TFP is obtained as the firm-specific component of the production function residual with production function coefficients estimated according to the Ackerberg et al. (2006) methodology (see Appendix A.3) and regressed on hours per worker to control for variable utilisation of production factors. We continue to focus on firm-level response to firm-level productivity developments corrected for hours per worker. Results from Table 1 are reported in the first section of Table 4.

First, we consider TFP shocks instead of the level of TFP. Indeed, the estimated coefficient on the level of TFP may be a mix of the dynamic response to current and lagged TFP shocks. Following Marchetti and Nucci (2005, 2007), we construct the shocks as the residuals from an AR(2) model on TFP with sector-specific intercept and slopes. The estimates of the current elasticity to TFP hardly change but the lagged impact is now positive.

Table 4 - SGMM estimates - alternative definitions of TFP

	w_{it}	l_{it}	h_{it}	$wb\text{-}h_{it}$	$h\text{-}l_{it}$
(1) TFP corrected for hours per worker (as in Table 1)					
tfp_{it}	0.02	0.23***	0.18***	0.03*	-0.01
	(0.02)	(0.04)	(0.04)	(0.01)	(0.01)
tfp_{it-1}	0.01	-0.12***	-0.08***	0.00	0.01
	(0.02)	(0.03)	(0.03)	(0.01)	(0.01)
(2) AR(2) TFP shock corrected for hours per worker					
tfp_{it}	0.03	0.25***	0.23***	0.01	0.00
	(0.02)	(0.04)	(0.04)	(0.02)	(0.01)
tfp_{it-1}	-0.00	0.12***	0.13***	-0.01	-0.00
	(0.01)	(0.04)	(0.04)	(0.01)	(0.01)
(3) TFP not corrected for hours per worker					
tfp_{it}	0.06***	0.25***	0.24***	0.01	0.02*
	(0.02)	(0.05)	(0.04)	(0.01)	(0.01)
tfp_{it-1}	-0.03	-0.11***	-0.13***	-0.00	-0.02*
	(0.02)	(0.04)	(0.03)	(0.01)	(0.01)
(4) AR(2) shock on TFP not corrected for hours per worker					
tfp_{it}	0.04*	0.24***	0.24***	0.00	0.02
	(0.02)	(0.04)	(0.04)	(0.02)	(0.02)
tfp_{it-1}	-0.00	0.12***	0.13***	-0.01	0.01
	(0.01)	(0.04)	(0.04)	(0.01)	(0.01)

Notes: The table presents only results for current and lagged TFP. In addition, each equation includes in addition the same control variables as Table 1. For details on GMM estimation, see note under Table 1. Standard errors in parentheses. * indicates significance at the 10% level, ** at the 5% level, *** at the 1% level.

Second, we examine the Ackerberg et al. (2006) estimate of TFP that is not corrected for variation in hours per worker, as is the case for Olley and Pakes (1996) and Levinsohn and Petrin

(2003) measures. As discussed above, our measure of average labour compensation, i.e. firm wage bill per employee in full time equivalent employee, includes compensation for overtime hours. Therefore, using a measure of TFP that does not take into account variable utilisation may induce a spurious positive correlation with our measure of wage and hours. The results in Table 4 suggest that such an upward bias may indeed be present. The coefficient on TFP in the average labour compensation equation increases and turns significant. In addition, the elasticity of hours per workers is significantly positive, although of a small magnitude.

Lastly, for comparison with the results in Marchetti and Nucci (2005) based on the Olley and Pakes (1996) methodology, we report estimates based on the shocks to TFP not corrected for variable utilisation rate in part (4) of Table 4.

5. Conclusion

In this paper we estimate the sensitivity of average labour compensation, employment, hours, hourly labour compensation and hours worked per employee to firm-specific and sector-level Total Factor Productivity. The sign and size of these elasticities may be affected by the presence of wage rigidity, employment adjustment costs, as well as price stickiness. On the one hand, real wage rigidity reduces the sensitivity of wages to shocks and shifts the burden of adjustment towards labour (Boldrin and Horvath (1995), Hall (2005)). On the other hand, hiring and firing costs may restrict adjustment through employment. In addition, price stickiness may induce a negative response of hours worked to productivity changes in the short or medium-run (see, for instance, Gali (1999)).

We compare the response of average labour compensation and labour to firm-level TFP with the response to sector-level TFP. When firms compete for workers on the labour market, they may implicitly coordinate their pay policies and refrain from isolated wage adjustment. When they compete on the product market they may not be able to transfer wage increases to prices. Lastly, it has been argued that collective wage bargaining may ease wage adjustment, especially in adverse times. Given the prominent role of sector-level wage bargaining in Belgium, this argument would again translate into a larger elasticity of average labour compensation with respect to sector-level TFP than with respect to firm-specific TFP.

We rely on a dataset obtained from firms' annual accounts and social balance sheets in Belgium over the period 1997-2005. Belgium is typically pointed out as a country with substantial real wage rigidity, due in part to its system of full automatic indexation of base wages. In addition, wage developments are largely driven by sector-level collective wage agreements. This makes Belgium a relevant case to study the role of real wage rigidity in alternative adjustment margins, and the role of centralised collective agreements in wage dynamics.

Our models are dynamic regression equations for each of the variables mentioned above and including TFP, variables that control for the workforce composition, capital stock, a firm size dummy and proxies of sector-level conditions. Our results can be summarised as follows. Focusing on the response to firm-level TFP, our estimates of the elasticity of average labour compensation to TFP is close to zero, while the elasticity of labour is high. Even though studies for other European countries surveyed in the paper also find low elasticity of wages to firm-level productivity, our estimate for Belgium is amongst the lowest. Although our analysis does not provide a test or evaluation of the extent of real wage rigidity, our finding of a low sensitivity of average labour compensation and large volatility of labour in response to firm-specific TFP is consistent with the hypothesis of real wage rigidity with respect to idiosyncratic productivity developments.

In contrast to the response to idiosyncratic TFP, the elasticity of average labour compensation to aggregate sector-level TFP is large, while that of labour is smaller. This is consistent with the fact that aggregate shocks cannot be insured, contrary to idiosyncratic ones. It is also consistent with the view that firms compete on the labour market to hire and keep workers, which makes them reluctant to undertake individual wage cuts. With respect to this argument, our results support the view that the high importance of centralised and coordinated wage bargaining at the sector level in Belgium may ease wage adjustment to aggregate changes. Indeed, the response of average labour compensation to sector-level TFP is large. Additional estimates suggest that (a large part of) sector-specific TFP developments are transmitted to average labour compensation changes through the sector-level collective wage agreements.

Lastly, we provide microeconomic evidence that hours worked respond positively to TFP within the year in our dataset. This is in contrast to the findings of Marchetti and Nucci (2005, 2007) for Italy and Carlsson and Smedsaas (2007) for Sweden. We run several robustness tests and we find our results robust to the specifications considered.

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Appendix A: Data

A.1 Dataset

Section 2.2 provides a brief description of the data used in the paper. Here we describe the technical details involved in construction of the dataset. We start with all firms that report their annual accounts in Belgium. This represents nearly all active firms. In this study, we consider only firms with 50 and more employees in manufacturing, construction and market services (NACE codes DA-KK). Additional data trimming concerns the legal situation of the firm. Foreign and public companies, as well as non-profit associations, are excluded from the sample. Only annual accounts covering the period from January to December are considered in order to ensure consistency between firms.

Outliers are removed by excluding observations below the 1st percentile or above the 99th percentile (defined year by year) of the following variables: employment growth (defined with respect to the number of employees and full-time equivalents), level and growth of the average wage per worker and average wage per hour, growth in firm-specific value added (in nominal and real terms), hours per worker, and the investment-capital ratio.

We consider only spells with at least 6 consecutive observations to make sure that sufficient history is available to build lagged instruments in the GMM procedure. We exclude observations with a missing value in any of the variables and in intermediate inputs. Last, we do not consider sectors with either too few observations to obtain a reliable estimate of the production function, or sectors where the production function coefficients were too far away from those obtained using alternative estimation procedures and the income shares (see Appendix A.3). Table A1 shows the composition of the dataset across the selected sectors.

Table A1 - Composition of the data set

Sector	NACE	firms	obs.
food	DA	143	1038
textile	DB	85	611
wood	DD	17	124
paper	DE	96	687
rubber	DH	67	480
metals	DJ	131	943
machinery and equipment	DK	54	387
electrical equipment	DL	60	424
other manufacturing	DN	47	339
construction	FF	249	1796
trade	GG	333	2325
hotels and restaurants	HH	29	203
financial services	JJ	23	164
real estate	KK	184	1250
total		1518	10771

A.2 Definitions of variables

The nominal wage bill is defined as remuneration and direct social benefits. The real wage bill, denoted by WB_{it}, is obtained as deflated nominal wage bill, the deflator being the sector-specific value added prices (31 branches).

Employment, abbreviated as L_{it}, is measured as the average number of employees in full-time equivalent positions over the year.

Hours are denoted as H_{it} and comprise the total number of hours worked at a firm in the particular year. These are reported in the firms' annual accounts.

We measure average labour compensation per firm (W_{it}) by the total real wage bill divided by the average number of full-time equivalent employees over the year. Hourly labour compensation per firm is given by the total real wage bill divided by the total number of hours worked over the year.

Variables related to workforce composition, like the percentage of blue-collar workers ($\%BLUE_{it}$), the percentage of women ($\%WOMEN_{it}$) and the proportion of workers with fixed-length contracts ($\%TEMP_{it}$), are provided in the social balance sheet, which has formed part of firms' annual accounts since 1996.

Nominal value added for large firms is equal to operating income minus operating subsidies and compensatory amounts received from public authorities, and the following operating charges: raw materials consumables and services and other goods. For small firms, valued added is proxied by the gross operating margin.

Profits are measured after financial charges, depreciation, amortisation and taxes. The unit of account for profits in our paper are millions of euro.

Real values are obtained using the sectoral value added deflator.

The index of collectively agreed wage increases is constructed as follows. The nominal indices of the collectively agreed nominal wage increases, defined at the sector-level separately for blue-collar and white-collar workers, are published by the Ministry of Labour (more precisely, the Federal Public Service Employment Labour and Social Dialogue) and we deflate it by the corresponding sector-level value added deflator to obtain the real measure. We then take the logarithm of the real indices for blue-collar workers and white-collar workers, I^B_{st} and I^W_{st}, and multiply them by the percentage of blue-collar workers, $\%BLUE_{it}$, and white-collar workers, $\%WHITE_{it}$ in each firm. The measure is not perfect because collectively agreed wage increases are defined at a more detailed level (in terms of sectors but also occupation and age or tenure). Furthermore, discrepancies with the average wage bill changes may be due to the fact that collective agreements do not apply to more flexible components of labour compensation such as bonuses or premiums paid in addition to wages. Deviations from these collective agreements are more frequent in large firms, and are very rare in the construction and business services sectors.

The firm-level capital stock is constructed using 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, from the earliest available annual account for the firm. We use the full history of annual accounts, since 1985, to determine the initial capital stock. The firm-specific depreciation rate is estimated as the median depreciation expenditure on capital, over the years in which the firm is in business.

Value added per sector (VA_{st}) was obtained from national accounts statistics.

Table A.2 below reports descriptive statistics on the variable used, including the difference of logarithm of our measure of firm-level TFP that we describe in the next section.

Table A2 - Descriptive statistics

Variable	obs.	mean	st dev.	P5	median	P95
W_{it}	10771	27904	8424	17542	26008	44539
L_{it}	10771	265.7	811.2	55.5	109.3	709.2
H/L_{it}	10771	1554	150	1301	1562	1777
WB/H_{it}	10771	17.91	4.86	11.86	16.84	27.4
Δtfp_{it}	10771	0.000	0.090	-0.14	0.00	0.12
$I_{it}/K_{it}^{(b)}$	10749	0.820	3.410	0.05	0.49	2.09
$\%BLUE_{it}$	10771	0.570	0.310	0.00	0.69	0.91
$\%TEMP_{it}$	10771	0.040	0.090	0.00	0.01	0.14
$\%WOMEN_{it}$	10771	0.260	0.220	0.02	0.19	0.70
$\%L>100_{it}$	10771	0.560	0.500	0.00	1.00	1.00
Δw_{it}	10771	0.020	0.080	-0.10	0.01	0.14
Δl_{it}	10771	0.010	0.110	-0.13	0.01	0.18
$\Delta(h\text{-}l)_{it}$	10771	-0.010	0.060	-0.09	0.00	0.07
$\Delta(wb\text{-}h)_{it}$	10771	0.020	0.080	-0.09	0.02	0.15
Δva_{st}	104	0.015	0.052	-0.07	0.02	0.09

Notes: Descriptive statistics for firms with more than 50 employees and 6 consecutive annual accounts over the years 1999-2005. P5 and P95 refer to the 5th and 95th percentile. Lowercase variables are in log. Δ stands for the difference operator.

W_{it} : Real wage bill per average number of employees in euro.
L_{it} : Average number of employees over the year
H/L_{it} : Total hours worker over the average number of employees
WB/H_{it} : Real wage bill over the total number of hours worked
Δtfp_{it} : difference log of firm-level TFP

I_{it}/K_{it} : Investment-capital ratio.

$\%BLUE_{it}$: percentage of blue collar workers

$\%TEMP_{it}$: percentage of employees under fixed-term contract

$\%WOMEN_{it}$: percentage of women

$\%L>100_{it}$: dummy equals to one when the firm employs 100 workers or more

Δva_{st} : difference log of real value added at sector-level

A.3 Productivity estimates

Two important issues related to the estimation of productivity shocks have been raised in the literature recently. First, factor demand is likely to be correlated with productivity shocks. Olley and Pakes (1996) correct for this simultaneity bias by augmenting the production function equation with a proxy of technological shocks based on investment and capital. Levinsohn and Petrin (2003) use intermediate inputs and capital as proxy.

Second, estimated productivity shocks may capture variations in the rate of utilisation of production factors. Indeed, with adjustment costs to input changes, firms may increase (decrease) effort and hours worked after a productivity shock rather than undertake a costly increase (decrease) in labour or capital. If this issue is ignored, the estimated residual will capture variations in the rate of utilisation of production factors along with the "true" productivity shock. Basu and Kimball (1997) develop a model where hours per worker are used to correct for this.

In this paper, we adopt the Ackerberg et al. (2006) estimation procedure that improves the techniques developed by Olley-Pakes (1996) and Levinsohn and Petrin (2003) in several directions. All three methods take into account the simultaneity bias due to the fact that factor demand is correlated with productivity shocks, but not variable utilisation of production factors. Ackerberg et al. (2006) correct for a colinearity issue that is present in the Levinsohn and Petrin (2003) and, to a lesser extent, Olley and Pakes (1996) procedures. Below, we first briefly describe the Olley and Pakes (1996) and Levinsohn and Petrin (2003) approaches, and then explain the Ackerberg et al. (2006) methodology.

Consider the following production function:

$$y_{it} = \beta_L l_{it} + \beta_K k_{it} + z_{it} + \eta_{it} ,$$

where the residual, $z_{it} + \eta_{it}$ is decomposed into one component observable to the firm when making its input decision, z_{it}, the productivity shock, and another component not observed by the firm at that time, η_{it}, which can be associated with unexpected productivity changes as well as measurement error. Because factor demand depends on productivity shocks, estimation of the above equation suffers from a simultaneity problem since the error term, $z_{it} + \eta_{it}$, is correlated with labour, l_{it}, and capital, k_{it}.

The estimation proceeds in two steps. First, a proxy for unobserved productivity, $\phi_t(.)$, is included in the equation to solve the simultaneity problem:

$$y_{it} = \beta_L l_{it} + \beta_K k_{it} + \phi_t(.) + \eta_{it} .$$

As shown in Olley and Pakes (1996), under some standard assumptions, the productivity shock can be expressed as a polynomial function of investment and capital, $\phi_t(i_{it},k_{it})$. Levinsohn and Petrin (2003) propose to use intermediate inputs instead of investment to solve the simultaneity bias, $\phi_t(m_{it},k_{it})$.[26] Because k_{it} also appears in the equation through $\phi_t(.,k_{it})$, the parameter β_K is not identified. However, the equation provides an estimate of β_L.

In the second stage, the coefficient on capital is estimated. For this purpose, expressing productivity as the sum of expected productivity, $E[z_{it}|z_{it-1}]$, and productivity innovations, ω_{it}, yields

$$y_{it} - \hat{\beta}_L l_{it} = \beta_K k_{it} + E[z_{it}|z_{it-1}] + \omega_{it} + \eta_{it} .$$

Assuming that the capital stock in year *t* was built in period *t-1* through i_{it-1}, k_{it} is independent of productivity innovations, ω_{it}. Therefore the proxy for expected productivity introduced in the equation enables β_K to be identified.

Finally, productivity growth is computed as $d\hat{z}_{it} = dy_{it} - \hat{\beta}_L dl_{it} - \hat{\beta}_K dk_{it}$.

Ackerberg et al. (2006) point out a fundamental colinearity problem that invalidates the estimation of the labour coefficient, and consequently the capital coefficient in the above

[26] The use of intermediate inputs was essentially motivated by data availability issues. In addition, as argued by Levinsohn and Petrin (2003), intermediate inputs may provide a better proxy of productivity growth if they adjust more easily to productivity shocks than capital.

procedures. In Levinsohn and Petrin (2003) methodology, if labour decisions and intermediate input decisions are taken simultaneously, i.e. if one assumes that both are flexible production factors, labour is collinear with the non-parametric function that proxies for productivity, $\phi_t(m_{it},k_{it})$. This leaves the labour production coefficient unidentified in the first step. Ackerberg et al. (2006) discuss in detail alternative assumptions concerning the timing of decisions on production factors and productivity data-generating processes but find no realistic set of assumptions that can solve the problem. The issue is less severe with the Olley and Pakes (1996) procedure, which remains valid if one assumes that the labour input decision is taken under incomplete information about productivity, while investment decisions are made after the productivity outcome is fully known.

Ackerberg et al. (2006) then propose an alternative estimation procedure that makes it possible to circumvent these issues, and is robust to a less restrictive set of assumptions.[27] The main feature of the estimation procedure is that the first stage equation is no longer used to estimate the labour production coefficient. But it serves to isolate production (and therefore TFP estimates) from the noise η_{it}, the unanticipated shocks at time *t* and measurement errors. All production function parameters are estimated in the second-stage equation. Identification of the capital parameter is again based on the assumption that the capital stock in period *t* was built through investment in period *t-1* and is therefore orthogonal to the productivity shock in *t*. Identification of the labour parameter uses the fact that lagged labour does not react to current productivity shocks, contrary to current labour. This defines a set of moment conditions used to estimate production function parameters through the method of moments. Alternative identification set-ups can be used according to the assumption made on labour decisions.

Our estimates of the production function and TFP are based on the method proposed by Ackerberg et al. (2006). Estimates using the Olley and Pakes (1996) and Levinsohn and Petrin (2003), not reported in the paper for the sake of brevity, yield less plausible production function coefficients than the Ackerberg et al. (2006) procedure. Data availability pleads in favour of an approach based on value added rather than gross output, seldom available for smaller firms. The dataset differs slightly from the one described in Appendix A.1. To maximise the number of observations on which our estimates are based, we include all firms disregarding their size and remove the restriction on 6 consecutive observations. Estimation of production function parameters is done separately for each sector. Three sectors were excluded because of insufficient number of observations: electricity, gas and water (EE); coke, refined petroleum and nuclear fuel (DF); leather and footwear (DC). We remove four additional sectors because the estimated coefficients do not appear to be plausible when compared to the income shares of labour and capital, see Table A3 below.

[27] Their procedure is consistent with the assumption that labour decisions may have an impact on future output and profits. Because this is what may result from important features of labour market rigidities, such as hiring, firing and training costs, it may be an important advantage of this procedure. When the Ackerberg et al. (2006) procedure is applied with intermediate inputs to invert productivity, the procedure is also consistent with unobservable and serially correlated variables affecting labour and capital decisions.

Table A3 - Estimated production function coefficients and income shares

Sectors in the data set	NACE	Prod. function coefficients: capital	Prod. function coefficients: labour	Income shares: capital	Income shares: labour
food	DA	0.24	0.56	0.20	0.59
textile	DB	0.12	0.50	0.25	0.69
wood	DD	0.15	0.91	0.24	0.62
paper	DE	0.11	0.52	0.17	0.63
rubber	DH	0.17	0.85	0.20	0.65
metals	DJ	0.28	0.44	0.17	0.74
machinery & equipment	DK	0.19	0.46	0.14	0.70
electrical equipment	DL	0.14	0.48	0.17	0.74
other manufacturing	DN	0.36	0.34	0.23	0.67
construction	FF	0.17	0.81	0.10	0.58
trade	GG	0.13	0.87	0.13	0.56
hotels and restaurants	HH	0.09	0.25	0.19	0.56
financial services	JJ	0.11	0.30	0.16	0.60
real estate	KK	0.26	0.66	0.22	0.29
mean		0.18	0.57	0.18	0.61
Sectors with unreliable coefficients					
chemicals	DG	0.11	1.15	0.13	0.57
non-metallic products	DI	-0.02	1.19	0.19	0.65
transport equipment	DM	-0.01	1.12	0.14	0.77
post and telecoms	II	0.08	1.13	0.30	0.60

Note: The following sectors were excluded owing to an insufficient number of observations: electricity, gas and water; coke, refined petroleum and nuclear fuel; leather and footwear.

We also account for another problem related to traditional production function estimates put forward by Basu and Kimball (1997). The assumption of constant rate of utilisation of production factors may not be consistent with adjustment costs in labour and capital accumulation. When the estimated productivity growth is not corrected for variable utilisation of production factors, it may be correlated with variations in hours per worker. This in turn may induce a spurious correlation between estimated productivity shocks and hours per worker. Basu and Kimball (1997) develop a structural model of imperfect competition on the product market, perfect competition on factor markets, increasing returns to scale and convex adjustment costs in both capital and number of employees. In their model, the rate of utilisation of labour can be proxied by hours per worker. Estimating TFP based on this methodology involves much more data and is beyond the scope of this paper. However, we adopt a simple method that corrects for variation in hours per worker. We simply regress our measure of TFP on hours per worker, and consider the residual from such regression as the "corrected" TFP measure.

Lastly, we decompose TFP into an idiosyncratic firm-level component and an aggregate sector-level component. The measure of total TFP obtained from the Ackerberg et al. (2006) procedure is estimated as $z_{it} = y_{it} - \hat{\eta}_{it} - \hat{\beta}_L l_{it} - \hat{\beta}_K k_{it}$. We regress this measure on hours per worker in order to account for the variable utilisation rate and on a set of sector-specific time dummies in order to remove the sector-specific component from z_{it}. Firm-level TFP is given by the residual of this equation and sector-level TFP is obtained as the estimated sector-specific time dummies.

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