Abstract. Using multivariate cointegration tests based on Johansen's approach, this paper tests for the existence of long-run relationships between real wages and other explaining variables in Belgium. We retain two cointegrating vectors which can refer to wage and price setting behaviours.

1. Introduction

The empirical estimation of wage determinants is an expanding research area, because of theoretical as well as technical considerations.² In a Phillips curve context, empirical studies relate wages to a set of variables including productivity, price and unemployment. Bogaert et al. (1991), Hall (1986, 1989), Graafland and Huizinga (1988) have used cointegration tests in this framework.²

A very promising theoretical approach has been more recently proposed and estimated by Layard, Nickell and Jackman (1991, further referred as LNJ). It specifically focuses on labour market
behaviour, where two types of economic agents meet: wage-setters ask for a target wage, while price-setters accept a feasible real wage considering their profit maximizing behaviour. Two wage equations are therefore assumed in this kind of framework. Equilibrium is reached when these behaviours are compatible: no inflationary pressure remains when both feasible and target wages are equal. And unemployment rate is the variable that can reconcile the two kinds of behaviour: the unemployment value that leads to equilibrium is called the Non-Accelerating Inflation Rate of Unemployment (NAIRU).

In this paper, we want to test the number of long-run relationships explaining real wages with respect to variables related to the NAIRU approach. On technical grounds, the presence of cointegrating vectors is the only case in which non-spurious relationships exist between non-stationary time series. We specifically use the maximum likelihood estimator (Johansen, 1988; Johansen and Juselius, 1990) in order to try to identify two expected cointegration relations for wage-setting (WS) and price-setting (PS), as Collard and Hénin (1993) noticed.

We proceed in two steps:

- we first present the theoretical model, available variables and unit root tests;
- we then estimate the model. We test for the number of cointegrating vectors, show how to restrain the cointegrating space in order to identify and interpret the structural form. This is the main technical point of this paper, allowing the estimation of the structural long-run model. We test for weak exogeneity for the long-run parameters. Introducing other wage-pressure variables, we finally test for the robustness of our estimated model. Throughout this technical approach, we comment on economic insights from our estimations.

2. Wage- and price-setting: Theoretical background

2.1. General framework

LNJ analyse wage- and price-setting behaviours. They consider two structural relationships explaining real wage and unemployment rate.

A. In the first one, wage-setters ask for a nominal “target” wage ($t_w$), represented by the yearly net wage earned by a full-time
worker. This wage, deflated by the consumption price index ($cp^4$), is the target real wage. The nominal target wage is related to the level of consumption price workers expect to pay (we assume full indexation in this analysis of the Belgian situation) and to the following wage-pressure variables:

- The unemployment rate ($un$), second dependent variable in this framework, influences competition on the labour market and thereby the target wage (Phillips effect) through different channels related to micro-founded explanations: lower insider power, higher outsider effectiveness, higher mismatch or lower incentives for firms to pay efficiency wages.

- These microfoundations also justify why other independent variables may influence the wage-setting behaviour and have to be integrated in our macroeconomic schedule:
  - wage-bargaining between insiders and firms (Mac Donald and Solow, 1981; Lindbeck and Snower, 1989), where insiders try to catch a mark-up on wages paid in non-unionized firms. This process leads to equilibria where negotiated wages are determined by variables that can affect the objectives of both negotiators and their relative power. In this context, variables related to insider power are union representation rate, size of wages with respect to total output, firms’ competitiveness or profits and human capital value. The degree of bargaining centralization also matters: outsider interests are more represented the more centralized the bargaining, as outside opportunities are then fewer for insiders if they are laid off. Replacement ratio can also influence wage pressure by insiders as they also condition outside opportunities. Among these potential variables, our data set allows for the following ones: union representation rate ($ur$), labour productivity ($lp$), reserved profits ($rp$), duration of work ($cd$) and a competitiveness index measured as the ratio between Belgian goods and services export prices and their world competitors’ export prices ($ci$);
  - together with insider power, higher search effectiveness from outsiders also affects the target wage as it strengthens firms capabilities to replace quitting workers. Aside from the fact that the total unemployment rate has already been considered to estimate this capability, we also include the ratio between long-term unemployed individuals and total labour force ($lu$) as another potential explaining variable.
Higher long-term unemployment rate should lower the expected effect of unemployment on the target wage, if longer duration leads to decreasing search effectiveness;

- besides the factors related to the bargaining process, increasing mismatch among labour micro-markets, either industrial, geographical or educational, can also raise the overall wage pressure. Even if aggregate labour supply constrains demand and wage-pressure should be low, some micro-markets can still be constrained on the labour demand side. As a consequence, overall wage pressure can be higher in case of bigger disparities between micro-markets. We have built an industrial index for mismatch (mm);

- another reason for wage-pressure, refers to firms’ behaviour. They can efficiently pay higher wages than those related to classical profit maximizing behaviour, when higher wages in turn increase productivity to a bigger extent. Workers can first work harder if they expect to receive a gift from firms (Akerlof, 1984). Higher wages can also serve as a screening device (Weiss, 1991), prevent workers from shirking (Shapiro-Stiglitz, 1984), limit union development (Dickens, 1986) or reduce voluntary quits and their turnover costs (Salop, 1979). Unfortunately, it is not possible to define an adequate empirical proxy for efficiency wages so far.

B. To define the other relationship between real wages and unemployment rate, price-setters are supposed to maximize profits in a non-competitive world. They decide a mark-up between product prices ($p$) and expected wage-costs ($w$). The feasible real wage is the ratio between wage-costs and product prices. We consider yearly wage costs for a full-time worker and estimate production prices from the GDP deflator.

Under this PS behaviour, higher labour productivity ($lp$) raises feasible real wage. Unemployment proxies product demand: the higher the unemployment rate, the higher the feasible real wage. Note that higher unemployment could proxy lower labour demand that also allows for higher feasible wages.

C. This WS–PS system assumes therefore two relations between wages, unemployment and independent variables. PS specifies a feasible real wage, WS a target real wage. To be able to specify the same (feasible) real wage ($w-p$) for both PS and WS, we consider

the wage wedge \((ww)\) between wages and introduce it in the WS specification. This wedge results from a tax wedge paid by employers and employees \((ta)\) and a price-wedge between consumption and GDP prices \((pw)\).

2.2. The model

Considering the previous theoretical background, we have to test for the existence of the two following relations:

\[
\text{PS: } (w-p)_t = \gamma_1 x_{1t} + u_{1t} \tag{1}
\]

\[
\text{WS: } (w-p)_t = \delta_1 x_{1t} + \delta_2 x_{2t} + \delta_3 x_{3t} + u_{2t} \tag{2}
\]

- \((w-p)\) is the feasible real wage;
- \(x_{1t}\) contains variables common to price and wage setting, i.e. unemployment rate \((un)\) and labour productivity \((lp)\);
- \(x_{2t}\) includes wage pressure variables, i.e. union representation rate \((ur)\), reserved profits \((rp)\), competitiveness index \((ci)\), long-term unemployment rate \((lu)\) and mismatch \((mm)\);
- \(x_{3t}\) refers to variables that measure the wage-wedge between feasible and target real wages, i.e. the tax-wedge \((ta)\) and the price-wedge \((pw)\).

2.3. Data and unit root tests

Testing for the presence of a unit root in a time series has become a starting point in empirical studies. Indeed, whether an economic variable is stationary has important consequences for the interpretation of economic models and data [see the seminal work of Nelson and Plosser (1982) \textit{inter alia}].

Unlike in other countries like France, Germany, Japan or the United Kingdom, quarterly data for Belgian national accounts are not available. Furthermore, given that annual data have been available since 1953, we finally consider the period 1953–90 to estimate our wage equations. These 38 observations are not too constraining as we observe a rather long period that matches pretty well the long-run concepts we consider, referring to unit root tests and cointegration. Perron (1989, 1991), Shiller and Perron (1985), Hakiao and Rush (1991) or Pierse and Snell (1993) show that increasing data periodicity does not improve the power of the tests. It is therefore better to analyse the longest possible period which is...
economically consistent, like the post-Second World War period looks to be in terms of wage bargaining.

Before estimating the model, we apply unit root tests to our variables using the Augmented Dickey–Fuller test (1979, 1981). The procedure consists of testing the null hypothesis of a unit root $H_0: \rho = 1$ against the alternative of a stationary process $H_A: \rho < 1$ in the following model:

$$ y_t = \rho y_{t-1} + \mu [1 + \beta t] + \sum_{i=1}^{k} \delta_i \Delta y_{t-i} + \epsilon_t \tag{3} $$

The most usual way is to use the standard $t$ statistic $(\hat{\rho} - 1)/\sigma(\hat{\rho})$ to test the unit root hypothesis $H_0: \rho - 1 = 0$. But this test is not asymptotically standardly distributed under the null, so Dickey and Fuller tabulated critical values for different parametrizations of equation (3). For variables given in absolute terms or in first difference, Table 1 gives the values of the unit root tests, alternatively with a constant or a constant and a trend. This last test has been introduced to control for the constant. An asterisk means that we can reject the null hypothesis at a 5 percent significance level. For each variable, the number of lags $k$ required to make the residuals white noises in equation (3) are presented in brackets.

### Table 1. Augmented Dickey–Fuller unit root test results

<table>
<thead>
<tr>
<th>Variables</th>
<th>Variables in level</th>
<th>Variables in first difference</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Constant</td>
<td>Constant and trend</td>
</tr>
<tr>
<td>$w$</td>
<td>-0.947 (1)</td>
<td>-2.283 (1)</td>
</tr>
<tr>
<td>$w$</td>
<td>-0.715 (1)</td>
<td>-2.164 (10)</td>
</tr>
<tr>
<td>$w-p$</td>
<td>-1.363 (1)</td>
<td>-0.647 (1)</td>
</tr>
<tr>
<td>$w-cp$</td>
<td>-1.688 (1)</td>
<td>-0.768 (1)</td>
</tr>
<tr>
<td>$l$</td>
<td>-2.646 (0)</td>
<td>0.969 (0)</td>
</tr>
<tr>
<td>$l$</td>
<td>-0.974 (1)</td>
<td>-2.661 (1)</td>
</tr>
<tr>
<td>$l$</td>
<td>-0.907 (1)</td>
<td>-3.340 (1)</td>
</tr>
<tr>
<td>$u$</td>
<td>-1.528 (0)</td>
<td>-0.416 (0)</td>
</tr>
<tr>
<td>$p$</td>
<td>-2.124 (1)</td>
<td>-2.368 (1)</td>
</tr>
<tr>
<td>$t$</td>
<td>-0.246 (0)</td>
<td>-2.285 (0)</td>
</tr>
<tr>
<td>$w$</td>
<td>0.180 (0)</td>
<td>-2.277 (0)</td>
</tr>
<tr>
<td>$c$</td>
<td>-1.277 (1)</td>
<td>-1.302 (1)</td>
</tr>
<tr>
<td>$m$</td>
<td>-1.630 (0)</td>
<td>-1.305 (0)</td>
</tr>
<tr>
<td>$mm$</td>
<td>-1.922 (2)</td>
<td>-2.358 (2)</td>
</tr>
</tbody>
</table>
We see that all the variables follow $I(1)$ processes, except nominal wages that are $I(2)$. So, taking real wages we can estimate long-run relations between these variables using the cointegration technique.

At this stage, we have presented the economic framework, the model and the data. To estimate it, we first determine the number of long-run relationships in our system. We then test for restrictions on the cointegrating vectors and on loadings. We finally comment the robustness of our estimated model.

3. Estimating the PS–WS model

3.1. Estimation technique: Insights

We use cointegration estimation technique to avoid spurious correlation problems encountered using other methods [see Granger and Newbold (1974), Banerjee et al. (1993) inter alia]. Estimating the long-run target wage using standard least squares regression techniques, Jackman and Leroy (1995) note that “[…] these results are not satisfactory as no statistics are provided about cointegration and the $t$-statistics are not appropriate to test the significance of non-stationary variables.” In economic terms, cointegration refers to a system of one or more long-run equilibrium relationships between relevant variables.

On statistical grounds, it tests and estimates the existence of stationary relationships between non-stationary variables. In a multivariate system, the Engle and Granger (1987) procedure does not allow to consider more than one cointegrating vector between different variables. As we want to test for the two possible relationships assumed in the preceding economic model, we use test statistics defined in Johansen (1988), Johansen and Juselius (1990, 1992). This maximum likelihood (ML) procedure can be summarized in the following way.5

We assume a $p$-dimensional non-restricted VAR process $\{Z_t\}$, referring to $p$ variables that have to be either $I(0)$ or $I(1)$:

$$Z_t = \mu + \sum_{i=1}^{k} A_i Z_{t-i} + u_t$$  \hspace{1cm} (4)

$\mu$ refers to the constant term, while $u_1 \ldots u_t$ are identically, independently and normally distributed $N_p(0, \Sigma)$. We do not consider seasonal components as they need not be included in our
estimations. The VAR process expressed in absolute levels (4) can further be transformed in first difference:

\[
\Delta Z_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} - \Pi Z_{t-k} + u_t
\]

(5)

where \( \Gamma_i = -(I_k - A_1 - \ldots - A_i), \ \Pi = I_p - A_1 - \ldots - A_k \)

\( (i = 1 \ldots k - 1). \)

Johansen (1988) suggests analysing the rank of the \( p \times p \) square matrix \( \Pi \) to determine the existence and the number \( r \) of cointegrating vectors. Three situations may occur:

1° if \( \text{rank}(\Pi) = p \), all the \( p \) variables are stationary;
2° if \( \text{rank}(\Pi) = 0 \), there is no cointegrating vector and analysis can be applied to a VAR process in first difference;
3° if \( \text{rank}(\Pi) = r \) with \( r < p \), \( \Pi \) can be transformed as \( \Pi = \alpha \beta' \), where \( \alpha \) and \( \beta \) are two \( p \times r \) matrices. The rows of \( \beta \) form the cointegrating vectors while the matrix \( \alpha \) contains the factor loadings, i.e. the weights of the cointegrating vectors in the various equations. The \( r \) cointegrating vectors of order \( p \), \( \beta_i \), are linearly independent \( (i = 1 \ldots r \leq p) \) and such as \( \beta_i Z_t \) are stationary. Johansen and Juselius propose two tests (trace and \( \lambda_{\text{max}} \)) to show the existence of cointegrating relations.

3.2. Estimating the most "robust" WS–PS model

We first test for the existence and the number of cointegrating vectors. Estimating coefficients of the long-run vectors, we then try to constrain the cointegrating space. We then test for weak exogeneity for the long-run parameters. It is indeed better to test for restrictions on the cointegrating space, because the hypothesis on weights (error correction coefficients) depends on a normalization and requires to identify (constrain) and adequately normalize the cointegrating space (Johansen and Juselius, 1992).

3.2.1. Determining the number of cointegrating vectors

The theoretical approach from LNJ allows for normalizing the cointegrating space \textit{ex post}, choosing the feasible real wage as a dependent variable for the two relations: and there has to be a positive correlation between real wage and unemployment in the PS relation and a negative in the WS.
Our data set does not allow to determine and identify WS and PS with all the variables presented in the theoretical framework. Therefore, we consider a more simple, apparently more robust model whose variables are the real wage, labour productivity, unemployment rate, long-term unemployment rate and competitiveness.

Rank tests for the ML procedure are presented in Table 2. Choosing the number of lags in the unrestricted VAR process and modelling deterministic variables is crucial in the Johansen procedure (Johansen, 1992). We have tested the overall strategies for deterministic components. We retain the model with an unconstrained constant in the short-run and a constrained trend in the long-run (see Juselius and Hargreaves, 1992). The Hannan–Quin statistic determines a multivariate lag process which follows a VAR of third order. Following Reimers (1993), it is better to use HQ criteria rather than the other ones in choosing the lag order into the Johansen’s procedure. With this specification residuals do not present evidence of serial correlation (LM test) or non-normality (BJ test).

Table 2 indicates that we cannot reject the null hypothesis of two cointegrating vectors, considering either the $\lambda_{\max}$ or the trace tests.

Table 3 details the coefficients of the two cointegrating vectors normalized by real wages.

### Table 2. Estimating the number of cointegrating vectors

<table>
<thead>
<tr>
<th>Null hypothesis</th>
<th>Alternative hypothesis</th>
<th>$\lambda_{\max}$ (95%)</th>
<th>Trace (95%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$r = 0$</td>
<td>$r = 1$</td>
<td>60.42*</td>
<td>37.52</td>
</tr>
<tr>
<td>$r \leq 1$</td>
<td>$r = 2$</td>
<td>43.27*</td>
<td>31.46</td>
</tr>
<tr>
<td>$r \leq 2$</td>
<td>$r = 3$</td>
<td>20.80</td>
<td>25.54</td>
</tr>
<tr>
<td>$r \leq 3$</td>
<td>$r = 4$</td>
<td>11.66</td>
<td>18.96</td>
</tr>
<tr>
<td>$r \leq 4$</td>
<td>$r = 5$</td>
<td>7.76</td>
<td>12.25</td>
</tr>
</tbody>
</table>

Table 3. Long-run coefficients of the two cointegrating vectors

<table>
<thead>
<tr>
<th>$w-p$</th>
<th>$un$</th>
<th>$lp$</th>
<th>$lu$</th>
<th>$ci$</th>
<th>Trend</th>
</tr>
</thead>
<tbody>
<tr>
<td>$-1$</td>
<td>0.031</td>
<td>0.665</td>
<td>0.083</td>
<td>0.242</td>
<td>0.00566</td>
</tr>
<tr>
<td>$-1$</td>
<td>-1.791</td>
<td>-6.316</td>
<td>1.267</td>
<td>3.488</td>
<td>0.218</td>
</tr>
</tbody>
</table>
When wage responsiveness to unemployment rate is positive (negative), we infer that the long-run relationship refers to price-wage-setting.

3.2.2. Restricting the cointegrating space

These two estimated relations are unidentifiable, as no a priori restriction is imposed and all the variables appear in both relations (Johnston, 1984).

Before restricting one or more coefficients to zero in one relation, we must ensure that they are non-zero in both equations to avoid that the system remains unidentified. This condition is tested imposing a zero coefficient on the two cointegrating vectors. Tests for non-significance of variables in both cointegrating relations are performed in Table 4.

Referring to our economical background, wage-pressure variables like long-term unemployment and union power (proxied by competitiveness for instance) only appear in the WS relation. Therefore, we can jointly test the restriction of a zero coefficient for long-term unemployment and competitiveness in the PS relation. We calculate a value of 0.013 for a test which is asymptotically distributed under the null as a $\chi^2(1)$ and accept the restriction. Results for constrained vectors then appear in Table 5.

This zero restriction is often used (Darby and Wren-Levis, 1993). But the cointegrating space remains statistically unidentified: using the order condition for identification, we note that the PS relation is overidentified, while the WS remains unidentified. To solve this problem, we have to put a restriction on WS for a variable that has

| Table 4. Long-run coefficients of the two cointegrating vectors |
|------------------|-------|-------|-------|-------|-------|
|                  | $w-p$ | $un$  | $lp$  | $lu$  | $ci$  |
| Test             | 24.25 | 20.63 | 13.78 | 16.63 | 8.66  | 11.06 |
| Distribution     | $\chi^2(2)$ | $\chi^2(2)$ | $\chi^2(2)$ | $\chi^2(2)$ | $\chi^2(2)$ | $\chi^2(2)$ |

| Table 5. Long-run restricted coefficients |
|------------------|-------|-------|-------|-------|-------|
|                  | $w-p$ | $un$  | $lp$  | $lu$  | $ci$  |
| $\mu$            | 0.158 | 1.168 | 0     | 0     | 0.328 |
| $\mu_1$          | 0.345 | -0.808| 0.948 | 0.0504|

a non-zero effect in the PS relation. We follow Manning (1993), rejecting that labour productivity has an effect on the WS relation. He suggests that observing a positive relation between real wages and productivity merely reflects a change in the conditions that determine the maximizing profit behaviour of the price-setters, not in the conditions explaining wage pressure.⁸

We finally impose $lu = 0$ in PS and $lp = 0$ in WS, estimating the two equations from the structural form without simultaneity bias. Table 6 summarizes the values of the tests for restrictions imposed on the two cointegrating vectors, together with a test for trend non-nullity in PS.

Results show that both WS and PS can be restricted ($lp = 0$ for WS, $lu = 0$ for PS). We then obtain two structural equations where we normalize PS by the unemployment rate (calculated standard-errors figure in brackets):⁹

$$
\text{PS: } un = 6.532(w-p) - 5.895lp \\
(1.01606) (1.04876)
$$

$$
\text{WS: } w-p = -0.211un + 0.245lu + 0.713ci + 0.025 \text{ trend} \\
(0.09883) (0.06086) (0.13491) (0.00207)
$$

These two structural equations do not allow to reject the model that determines wages and unemployment rate in the long-run.¹⁰

- Referring to the assumed price-setting relation, we note that elasticity between target wage and labour productivity is estimated to be close to unity (5.895/6.532), although a bit smaller. The latter phenomenon could probably be related to the fact that unions seem to have accepted higher manning ratios during the 1980s in Belgium.
- On the other hand, we observe significant and close coefficients in absolute terms and opposite in signs for total unemployment and long-term unemployment rates in WS. These estimates are consistent with the assumption that

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**Table 6.** Restriction on the cointegrating space

<table>
<thead>
<tr>
<th>Constraining PS</th>
<th>Constraining WS</th>
<th>Test</th>
<th>Distribution</th>
</tr>
</thead>
<tbody>
<tr>
<td>$lu = ci = 0$</td>
<td>$lp = 0$</td>
<td>0.013</td>
<td>$\chi^2$ (1)</td>
</tr>
<tr>
<td>$lu = ci = \text{trend} = 0$</td>
<td>$lp = 0$</td>
<td>1.971</td>
<td>$\chi^2$ (2)</td>
</tr>
</tbody>
</table>
outsiders’ effectiveness on the labour market strongly decreases as average unemployment duration grows. Unemployment duration probably first serves as a screening device on the Belgian labour market where inflows to employment are severely constrained. But we cannot estimate whether this decreasing effectiveness has to be explained by lower search intensity and/or lower expected productivity. However, our estimations do not mean that we can subtract long-term unemployment from total unemployment and consider that only short-term unemployment could influence the target wages in WS (Bogaert et al., 1991; Hénin, 1993), as our variables are not in absolute terms but represent logistic transformations.

● The competitiveness effect is quite amazing. The index of competitiveness is built as the ratio between Belgian export prices and world competitors’ prices. The fact that the target wage is positively related to this index could be interpreted as follows: higher Belgian prices could represent higher profitability for price-making firms and hence higher bargaining power for insiders. Alternatively, we could also assume that causality between the target wage and competitiveness index could be reversed. In the Belgian small open economy context, higher negotiated wages can also (partly) provoke higher prices for some price-making firms. But this assumption that merely reflects price-setting conditions seems not to be consistent with the 3 restriction that we were able to accept for the competitiveness effect in the PS relation.

● Interpreting the deterministic trend is more difficult. First, one or more of the variables could be stationary around this deterministic trend. We have tested this assumption for variables whose behaviour seems to follow this trend, i.e. real wage and labour productivity. Our multivariate schedule could question the results of ADF unit root tests that seem to indicate stochastic rather than deterministic trends. In fact, Table 7 confirms that we can reject the assumption of stationarity around a deterministic trend for these two variables. On the other hand, maintaining the three constraints in PS and only constraining the trend to zero in WS (leaving the productivity variable free), we obtain a value of 12.32 for a test distributed as $\chi^2(2)$ and reject the hypothesis that the trend is a proxy for productivity.
From the two previous arguments, we can infer that the trend does not proxy labour productivity in WS. It could alternatively proxy increasing insider power or bigger propensity to pay efficiency wages given that industrial mismatch does not seem to increase on the overall period.

3.2.3. Testing the matrix of loadings

Table 8 details the matrix of factor loadings, i.e. the weights of the cointegrating vectors in the various equations of the VAR. Given the fact that normalization has been imposed, we can test for weak exogeneity of one or several variables for the parameters of interest, i.e. the long-run parameters (Urbain, 1993). Such a procedure allows to estimate the speed of short-run adjustments when variables deviate from their long-run equilibrium values. Adjustments from the five variables to a disequilibrium in either PS or WS are reported in the first two columns on Table 8. We can first notice in column 2 that wages do not seem to adjust quickly to WS long-run equilibrium, reflecting a (short-run) rigidity that could partly explain Belgian persistence of unemployment encountered in the late 1980s. On the other hand, unemployment adjustment to PS disequilibrium is much quicker. This result could also mean that Belgian labour demand does not encounter that much difficulties to adjust to

<table>
<thead>
<tr>
<th>Null hypothesis</th>
<th>Test</th>
<th>Distribution</th>
</tr>
</thead>
<tbody>
<tr>
<td>((w-p) - \text{deterministic trend})</td>
<td>12.81</td>
<td>(\chi^2 (3))</td>
</tr>
<tr>
<td>(lp - \text{deterministic trend})</td>
<td>19.21</td>
<td>(\chi^2 (3))</td>
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<table>
<thead>
<tr>
<th>Test for stationarity around a deterministic linear trend</th>
</tr>
</thead>
<tbody>
<tr>
<td>Null hypothesis Test Distribution</td>
</tr>
<tr>
<td>((w-p) - \text{deterministic trend}) 12.81 (\chi^2 (3))</td>
</tr>
<tr>
<td>(lp - \text{deterministic trend}) 19.21 (\chi^2 (3))</td>
</tr>
</tbody>
</table>

Table 8. Loading matrix and exogeneity tests

<table>
<thead>
<tr>
<th></th>
<th>PS</th>
<th>WS</th>
<th>(z_1(\text{PS}) = 0)</th>
<th>(z_1(\text{WS}) = 0)</th>
<th>(z_1 \text{ and } z_2 = 0)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(w-p)</td>
<td>-0.0625</td>
<td>-0.0140</td>
<td>11.075</td>
<td>2.015</td>
<td>22.109</td>
</tr>
<tr>
<td>(un)</td>
<td>-0.8558</td>
<td>2.4365</td>
<td>13.286</td>
<td>15.570</td>
<td>32.808</td>
</tr>
<tr>
<td>(lp)</td>
<td>0.0226</td>
<td>-0.3183</td>
<td>2.665</td>
<td>23.255</td>
<td>29.676</td>
</tr>
<tr>
<td>(lu)</td>
<td>-0.3041</td>
<td>0.5769</td>
<td>9.922</td>
<td>3.302</td>
<td>10.057</td>
</tr>
<tr>
<td>(ci)</td>
<td>0.0389</td>
<td>-0.5091</td>
<td>2.589</td>
<td>9.517</td>
<td>9.66</td>
</tr>
</tbody>
</table>

disequilibria. And the difference between Belgian observed unemployment rate and NAIRU could therefore be partly related to both wage stickiness and huge capability for labour demand to adjust to labour productivity/cost ratio.

In columns 3 and 4, we test for null adjustment in the two relations considered apart and together in column 5. We also simultaneously test for the null that all zero coefficients equal zero. We get a value of 3.86 for the test distributed as $\chi^2(4)$. Under these zero restrictions on the loading matrix, we estimate two new WS and PS relations (calculated standard errors appear in brackets) and get the following results.

\[
\text{PS: } un = 6.939 (w-p) - 6.314 lp \\
\quad (0.84050) \quad (0.87203)
\]

\[
\text{WS: } w-p = -0.131 un + 0.202 lu + 0.603 ci + 0.025 \text{ trend} \\
\quad (0.03446) \quad (0.03391) \quad (0.07642) \quad (0.00158)
\]

Estimated coefficients are quite similar to the ones observed in the first estimation, while standard errors are much smaller. Same qualifications apply therefore to explain the estimated relations between variables. Note, however, that the estimated effect of unemployment on wages is lower, though the confidence interval is narrower and presents fewer probabilities to contain the zero value.

These estimations also support LNJ theoretical approach. On the PS, the feasible wage is positively influenced by unemployment rate and labour productivity. While the target wage is a negative function of unemployment rate and is positively related to long-term unemployment rate, competitiveness and a deterministic trend proxying either bigger insider power or propensity to pay efficiency wages.

Table 9 details the new matrix of loadings to the two long-run equilibria (calculated standard errors appear in brackets).

**Table 9. Loading matrix of the restricted model**

<table>
<thead>
<tr>
<th></th>
<th>PS</th>
<th>WS</th>
</tr>
</thead>
<tbody>
<tr>
<td>(w-p)</td>
<td>$-0.0649 \ (0.08083)$</td>
<td>0</td>
</tr>
<tr>
<td>(un)</td>
<td>$-0.6967 \ (0.10197)$</td>
<td>2.4319 \ (0.68573)</td>
</tr>
<tr>
<td>(lp)</td>
<td>0</td>
<td>$-0.3563 \ (0.05748)$</td>
</tr>
<tr>
<td>(lu)</td>
<td>$-0.2446 \ (0.0838)$</td>
<td>0</td>
</tr>
<tr>
<td>(ci)</td>
<td>0</td>
<td>$-0.4966 \ (0.15511)$</td>
</tr>
</tbody>
</table>

We also confirm previous observations where we partly explained Belgian observed unemployment rate by wage stickiness on the WS side and labour demand quick adjustments on the PS side.

The residuals of both relations are shown in the Graphs 1 and 2. Cointegrating vectors are multiplied by variables in absolute terms corrected for short-run evolutions.

Graph 1. First cointegrating vector (price-setting)

[Graph showing data from 1956 to 1990]

Graph 2. Second cointegrating vector (wage-setting)

[Graph showing data from 1956 to 1990]
3.2.4. Robustness of the results

Estimating the existence of two cointegrating relations for wages, Darby and Wren-Lewis (1993) raise the problem of a lack of robustness in their estimations. They focus on three explanations:

- inadequacy of this theoretical background to explain the empirical situation;
- difficulty to measure adequately concepts like mismatch, insider power, efficiency wages or search effectiveness;
- applying cointegrating techniques needs to consider long periods together with huge datasets.

We also observe a lack of robustness in our results. Coefficient values or signs are sensitive to different factors like the number of lags we retain for the VAR process, the introduction of additional explaining variables or the period considered.

Estimating the model with additional variables seems promising. But it is econometrically difficult to do because of an induced lack of degrees of freedom. When we add variables to our model (union representation rate, wedge and mismatch), we have to estimate 25 coefficients for each relation. In such an estimation, we are unable to reject the existence of eight cointegrating vectors! We could infer that some variables are weakly exogeneous and condition our model under such an information. But weak exogeneity tests could not exert their asymptotic properties with five or six degrees of freedoms. So we should probably choose which variable to retain in order to condition the system.

At this stage, we decide to follow an exploratory walk and test for the significance of one or two variables we added in our model. Assuming the existence of two cointegrating vectors, we can exclude the duration of work and the wage-wedge for both relations.

Union representation rate presents serious difficulties. It cannot be excluded from both equations but has a negative effect on wages and does not permit to reject the significance of all cointegrating vectors. Furthermore, this indicator does probably not proxy insider power in a convenient way. While the decrease of Belgian wage-profit share during the 1980s seems to indicate decreasing union power, union representation rate does not decline during the same period.

Mismatch cannot be excluded from the wage-setting relation but has a negative effect on wages. The index we built to measure this variable could reveal another explaining phenomenon than mismatch, proxying for example the unemployment rate. Increasing
mismatch could also be related to another type of relation than wage-pressure, i.e. economic activity. Economic growth could provoke wage pressure together with either decreasing unemployment and mismatch. The estimated relation between the later variables do then not reflect causality from mismatch to wage pressure.

Finally, we cannot analyse the stability of rank tests and coefficient values, because of an insufficient number of degrees of freedom.

4. Conclusion

Referring to the theoretical NAIRU schedule, we have tested for the existence of two long-run relations representing wage- and price-setting. Wage-setters claim a target wage which is a function of unemployment rate and wage-pressure variables. Price-setters accept a “feasible” wage related to unemployment rate (as a proxy for the size of products demand) and labour productivity.

Using the maximum likelihood estimation technique, we are not able to reject the existence of two cointegrating vectors. Testing and imposing identification conditions to the structural form of our model, we retain two relations that can be referred to as price- and wage-setting. In the first relation, feasible wages are positively related to unemployment rate and labour productivity in a way that is consistent with profit maximizing behaviours. The elasticity between wages and labour productivity is estimated to be close to unity. On the wage-setting side, target wage is negatively related to unemployment rate and is a positive function of competitiveness (in the sense that it could proxy higher insider power) and long-term unemployment rate (as a proxy for outsiders’ search effectiveness). So insider–outsider explanations seem relevant to explain Belgian equilibrium wages. But neither union representation rate (whose quality to proxy insider power can be questioned), mismatch nor duration of work are estimated to have a significant effect in both relations.

We finally would like to qualify these considerations. Our results are sensitive to additional explaining variables and to a modification of the number of lags in the VAR process. We can probably partly explain this lack of robustness by technical reasons like a small number of degrees of freedom, a lack of appropriate data to measure some relevant explaining factors like efficiency wages or to
proxy appropriately mismatch or union bargaining power. Nevertheless, we believe that the way we test the wage determinants and especially restrict the long-run relationships can open the way to new empirical findings.

Appendix. Representing the data (absolute levels)

Graph A1. Feasible (——) and target (—–) real wages

Graph A2. Feasible (——) and target (—–) real wages: growth rates

Graph A3. Labour productivity

Graph A4. Total (Utot), short term (Uct) and long-term (Ult) unemployment rate

Notes

1 Estimating a wage equation, Sargan has introduced error correction models. Comparing the two following papers from Hall (1986, 1989) where he estimates wages in the United Kingdom, one can figure out how cointegration techniques develop, from Engle and Granger (1987) to Johansen (1988) maximum likelihood estimators.

2 Using Engle and Granger (1987) estimator, these authors test for only one wage equation.

3 […] "Using error correction models, the true wage equation should include two cointegration relations, in terms of supply and demand" (Hénin, 1993).

4 All variables are expressed in logarithms.

5 This technique has already been used to estimate wage equations, for example in Hall (1989), Darby and Wren-Lewis (1993) or Nymoen (1992).

6 The computations were carried out using a procedure written in RATS 4 by Mosconi.

7 Note that these tests are independent of an ex post normalization of cointegrating vectors.

8 […] In neither case do productivity variables play an important role in the structural wage equation. […] this does not mean that productivity growth does not cause the growth in real wages over time. But suppose that the labour market was competitive and that labour supply was totally inelastic. Then productivity growth leads to real wage growth, but it would be strange to argue from this that one should include productivity directly in the estimate of the structural labour supply curve.

9 The standard errors were calculated within a non-linear least squares system framework using the procedure Nisystem in RATS 4. More precisely, the unrestricted VAR is rewritten as in (5). We add, however, in place of all the
variables in level at \( t - 3 \) in the matrix \( \Pi \), the two cointegrating vectors we obtained. One must let all the coefficients free but constrain the long-run parameters of the two relationships to be equal in the five ECM equations. The estimates give numerically the long-run coefficients obtained in the Johansen’s procedure. If the model is correctly over-identified, these calculated standard errors are the same as the estimator of the variance–covariance matrix developed in Johansen (1991).

Standard error of unemployment rate is important in WS: the value of the \( t \)-test is 2.14, facing a theoretical value of 1.64. We will see further that this standard error decreases as we introduce restrictions in the adjustment matrix. Note that one will accept the same restrictions if the short-term unemployment rate instead of the total one (un) was used.

References


