Econ Change Restruct (2012) 45:121–133
DOI 10.1007/s10644-011-9112-1

Price-wage nexus and the role of a tax system

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Received: 15 September 2009 / Accepted: 15 May 2010 / Published online: 25 May 2011

Abstract The presented study was aimed to test empirically major economic hypotheses dealing with long-term relationships between wages, producer prices, prices of consumer goods and services, the consumer price index, productivity of labour, unemployment and payroll expenses other than wages themselves. It is particularly important for this approach to distinguish between net wages shaping employees' decisions and gross wages driving employers' decisions. Because the variables are generated by non-stationary stochastic processes integrated of order 1 and 2, the analytical tool applied was a vector equilibrium correction model, VEqCM. The findings demonstrate that prices and payroll expenses are the major sources of shocks in the system in question. Wages and prices (particularly producer prices) are the most sensitive to this type of stochastic trends. In the Polish economy prices are integrated of order two, so they can be effectively influenced by anti-inflationary policy.

Keywords Inflation modelling · Cointegration · I(2) analysis

JEL Classification C32 · E24 · E31

1 Introduction

It is usually assumed that wages are formed in the course of a bargaining process and that an unemployment rate does not accelerate inflation (the NAIRU hypothesis, see Wallis 2004). Most aggregate models, however, do not distinguish between total

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costs of labour (i.e. gross payroll plus all related expenses) and net wages, notwithstanding the fact that employers' decisions are determined by the former, while employees decide to take jobs in response to the latter. Therefore, because of the time-varying dynamics of the charges related to wages (payroll), i.e. taxes and social insurance contributions, taking account of them is crucial for explaining inflationary phenomena correctly.

Another simplification that can be found in many models is that the producer prices (PPI) and the consumer price index (CPI) are treated as equivalent to each other. The former is determined by (unit) costs that mainly consist of labour cost, whereas CPI is shaped not only by producer prices, but also by the prices of services, of imports and also taxes which show variability in time that is not only strong but also essentially different from the variability of producer prices.

The presented study aimed to build a model taking account of both payroll expenses and their effects on inflationary processes and to analyse the factors that make CPI grow faster than PPI.

The study employed deseasonalized monthly data spanning the period January 1993–December 2008, i.e. covering also the first years of Poland's membership in the EU. At the same time, the final months of the sample mark the outset of a global financial crisis that emerged in the fall of 2008.

Because the variables are generated by nonstationary stochastic processes integrated of order one and two, the analytical tool used was a vector (equilibrium) correction model, VECM (VEqCM).

2 The model

The starting point for the model is producer prices that depend on unit labour costs and import prices; assuming that the import-output ratio is constant in time this can be written as:

$$pr = \delta_1^r pm^I + \left(1 - \delta_1^r\right)(wc - z) \tag{1}$$

where pm^I—prices of imported raw materials and intermediate goods, wc—unit labour costs (nominal), z—productivity of labour, δ_1^r—the parameter. The variables denoted by lower case are expressed in natural logarithms.

According to the above equation, wages must grow faster than productivity of labour to produce a pro-inflation effect (see Tobin 1972). It should be noted that this (reduced) cost function has to meet a homogeneity condition.

Index of consumer goods and the producer price index show different variability for the following reasons. Firstly, the former is additionally influenced by the prices of imported final goods. Secondly, VAT is charged on both domestic and imported goods and its (effective) rate varies in time. As a result, a correctly constructed deflator of consumer goods pc should be a (weighted) sum of the deflator of imported final goods pm^F, the producer price index pr and the VAT rate tv:

$$pc = \left(1 - \delta_1^c\right) pm^F + \delta_1^c pr + tv. \tag{2}$$

Besides the prices of goods the consumer price index p (CPI) also includes the prices of consumer services ps:

$$p = \delta_1^p pc + \left(1 - \delta_1^p\right) ps. \tag{3}$$

Regarding real wages, it can be assumed that they depend on productivity of labour z and labour force pressures of intensity related to the rate of unemployment U:

$$w - p = \delta_1^w z - \delta_2^w U. \tag{4}$$

The acceptance of the hypothesis about a bargaining model of wages (see Nickell 1984; Layard et al. 1991) induces a unit (long run) elasticity of wages with respect to productivity of labour, $\delta_1^w = 1$, that corresponds to the standard wage function (see Tobin 1995).

When the impact of indirect and direct taxes and of social insurance contributions is added to the model, then the definition of total (real) payroll expenses (employment costs) is as follows:

$$wc - p = w + tw - p \tag{5}$$

where tw denotes payroll taxes paid by employers, social insurance contributions and other elements of wages and w stands for gross wages that diminished by the direct taxes represent net wages:

$$w - td - p = wc - tx - p \tag{6}$$

where td is direct taxes and tx total costs of payroll expenses ($tx = tw + td$).

Altogether, the model consists of five equations:

– producer prices:

$$pr = \delta_1^r pm^I + \left(1 - \delta_1^r\right)(wc - z), \tag{7a}$$

– the deflator of consumer goods:

$$pc = \left(1 - \delta_1^c\right) pm^F + \delta_1^c pr + tv, \tag{7b}$$

– the deflator of services:

$$ps = \delta_1^s pc + \delta_2^s h, \tag{7c}$$

– the consumer price index:

$$p = \delta_1^p pc + \left(1 - \delta_1^p\right) ps, \tag{7d}$$

– net wages:

$$wc - tx - p = z - \delta_2^w U \tag{7e}$$

where h stands for the share of services in total consumption, whose expansion generates upward market pressure on prices. The structure and degree of disaggregation make the above model close to the classic macromodels constructed for national economies (see Klein et al. 1999).

Let us note that the consumer price index can also be written as:

$$p = \left[\delta_1^p\left(1-\delta_1^s\right)+\delta_1^s\right](pr+tv+tm)+\left(1-\delta_1^p\right)\delta_2^s h \tag{8}$$

where $tm = \left(1-\delta_1^c\right)\left(pm^F - pr\right)$ can be interpreted as "a tax" resulting from (high) import prices (see Wallis 2004). The long-run dynamics of the prices of services does not diverge from the dynamics of consumer goods prices, so $\delta_1^s = 1$. Therefore, Eq. 8 comes down to a weighted sum of the services share in total consumption and the PPI index augmented with tax rates:

$$p = (pr+tv+tm)+\left(1-\delta_1^p\right)\delta_2^s h. \tag{9}$$

According to model 7a–7e, an exchange rate contributes to inflationary processes through the import prices of raw materials and intermediate goods (consumed during production) pm^I and the import prices of consumer goods pm^F, because both these categories are by definition a product of the import price index and the exchange rate. It must be emphasised, though, that the model treats pm^I and pm^F, and consequently the exchange rate, as exogenous.

3 Analysis in the I(2) environment

The variability of M cointegrated variables generated by stochastic processes I(1) and I(2) can be described as

$$\Delta^2\mathbf{Y}_t = \Pi\mathbf{Y}_{t-1} + \Gamma\Delta\mathbf{Y}_{t-1} + \sum_{s=1}^{S-2}\Psi_s\Delta^2\mathbf{Y}_{t-s} + \Sigma_t, \tag{10}$$

which is equivalent to

$$\Delta\mathbf{Y_t} = \Pi\mathbf{Y}_{t-1} + \sum_{s=1}^{S-1}\Gamma_s\Delta\mathbf{Y}_{t-s} + \Sigma_t \tag{11}$$

where $\Gamma = \sum_{s=1}^{S-1}\Gamma_s - \mathbf{I}$, $\Psi_s = -\sum_{j=s+1}^{S-1}\Gamma_j$ and $\mathbf{Y}_t$ are matrixes M of variables, Π—a matrix of long-run multipliers, Γ—a matrix of medium-run multipliers, Ψ_s—a matrix of short-run reactions, $s = 1,\ldots,S$.

If the hypothesis about cointegration rank $r = R < M$ is not rejected, then matrix Π of the long-run multipliers can be decomposed into matrixes $\mathbf{A}$ and $\mathbf{B}$ of the $M \times R$ dimension, $\Pi = \mathbf{AB}^T$. Matrix $\mathbf{B}$ consists of $R > 0$ base cointegration vectors. It is purposeful then to further decompose the variable space into an R-dimensional cointegration space and an M – R—dimensional space of common stochastic trends. Two cases can be distinguished here.

One is when $r\left(\mathbf{A}_\perp^T\Gamma\mathbf{B}_\perp\right) = M - R$. This means that common stochastic trends I(2) do not exist and the procedure is reduced to a standard cointegration analysis I(1).

The other takes place when $r\left(\mathbf{A}_\perp^T\Gamma\mathbf{B}_\perp\right) < M - R$, which means that both direct equilibrium relations (CI(2,2)) that are observable in the long and medium run and cointegration relations of the (CI(2,1)) type can be found between the variables. Because the latter become effective after a very long time, the adjustments they induce are slower. There are also short-run relations or exclusively medium-run relations that according to Juselius (2004) can be linked to stochastic cycles.

To estimate the direct cointegration relations CI(2,2) and the cointegration of the CI(2,1) type, matrix **B** has to be projected onto respective subspaces (see Haldrup 1999):

$$\mathbf{B}_1 = \mathbf{B}\Lambda^T \tag{12}$$

$$\mathbf{B}_0 = \mathbf{B}\Lambda_{\perp}^T \tag{13}$$

where:

$\mathbf{B}_0$—$M \times R_0$ matrix consisting of R_0 base cointegration vectors of the CI (2,2) type,

$\mathbf{B}_1$—$M \times R_1$ matrix consisting of R_1 base cointegration vectors of the CI (2,1) type,

$\Lambda_{\perp}^T - R \times R_0$ projection matrix into directly stationary cointegration space CI(2,2),

Λ^T—projection matrix into I(1) cointegration space CI(2,1),

$$\Lambda^T = \left(\mathbf{A}^T\mathbf{A}\right)^{-1}\mathbf{A}^T\Gamma\mathbf{B}_{2\perp}\left(\mathbf{B}_{2\perp}^T\mathbf{B}_{2\perp}\right)^{-1} \tag{14}$$

$$R_0 + R_1 = R \tag{15}$$

Decomposing the space of the common stochastic trends we obtain:

$$\mathbf{A}_{1\perp} = \mathbf{A}_{\perp}(\mathbf{A}_{\perp}^T\mathbf{A}_{\perp})^{-1}\Xi \tag{16}$$

$$\mathbf{B}_{1\perp} = \mathbf{B}_{\perp}(\mathbf{B}_{\perp}^T\mathbf{B}_{\perp})^{-1}\mathrm{N} \tag{17}$$

$$\mathbf{A}_{2\perp} = \mathbf{A}_{\perp}\Xi_{\perp} \tag{18}$$

$$\mathbf{B}_{2\perp} = \mathbf{B}_{\perp}\mathrm{N}_{\perp} \tag{19}$$

where:

$\mathbf{A}_{1\perp}$—$M \times P_1$ matrix of coefficients defining P_1 base common stochastic trends I(1),

$\mathbf{A}_{2\perp}$—$M \times P_2$ matrix of coefficients defining P_2 base common stochastic trends I(2),

$P_1 + P_2 = M - R$,

$\mathbf{B}_{1\perp}$ and $\mathbf{B}_{2\perp}$—matrixes $M \times P_1$ and $M \times P_2$, respectively,

Ξ, N—matrices $(M - R) \times P_1$ $(P_1 < M - R)$ such that $\mathbf{A}_{\perp}^T\left(\sum_{s=1}^{S-1}\Gamma_s - \mathbf{I}\right)\mathbf{B}_{\perp} = \Xi\mathrm{N}^T$.

The above procedure starts with a VAR model and tests the hypothesis about order Π being reduced and then about $r(\mathbf{A}_{\perp}^T\Gamma\mathbf{B}_{\perp}) < M - R$. It is assumed therefore that $I(2) \subset I(1) \subset I(0)$.

An alternative strategy takes as a starting point the VAR model solution for a model with common stochastic trends I(2) (see Johansen 1996, p. 35):

$$\mathbf{Y}_t = C_1\sum_{i=1}^{t}\Sigma_i + C_2\sum_{j=1}^{t}\sum_{i=1}^{j}\Sigma_i + C(L)\Sigma_t, \tag{20}$$

where C_1 measures medium-run shocks I(1) and $C(L)$ stationary (short-run) shocks. If matrix C_2 related to the long-run shocks is non-zero, then it can be decomposed into $C_2 = \tilde{\mathbf{B}}_{2\perp}\mathbf{A}_{2\perp}^T$, where $\tilde{\mathbf{B}}_{2\perp} = \mathbf{B}_{2\perp}\left(\mathbf{A}_{2\perp}^T\left(\Gamma\mathbf{B}\left(\mathbf{B}^T\mathbf{B}\right)^{-1}\left(\mathbf{A}^T\mathbf{A}\right)^{-1}\mathbf{A}^T\Gamma - \sum_{s=1}^{S-2}\Psi_s\right)\mathbf{B}_{2\perp}^T\right)^{-1}$ is a matrix of adjustments to common stochastic trends I(2). If $C_2 = \mathbf{0}$, then (20) can be simplified to a model of common stochastic trends I(1):

$$\mathbf{Y}_t = \tilde{\mathbf{B}}_{\perp}\mathbf{A}_{\perp}^T\sum_{i=1}^{t}\Sigma_i + C(L)\Sigma_t \tag{21}$$

where $\tilde{\mathbf{B}}_{\perp} = \mathbf{B}_{\perp}\left(\mathbf{A}_{\perp}^T\left(\sum_{s=1}^{S-1}\Gamma_s - \mathbf{I}\right)\mathbf{B}_{\perp}\right)^{-1}$ is a matrix of adjustments to common stochastic trends I(1). If $\tilde{\mathbf{B}}_{\perp} = \mathbf{0}$, then (21) reduces to VMA for stationary variables. Therefore, this strategy assumes that $I(2) \supset I(1) \supset I(0)$.

The omission of the stochastic trends I(2) produces serious consequences. Although the cointegration matrix estimator is super-super-consistent, the dimension of matrix $\mathbf{B}$ may be incorrectly determined, because the results of the classic tests for cointegration order are based on the assumption that $r\left(\mathbf{A}_{\perp}^T\Gamma\mathbf{B}_{\perp}\right) = M - R$, which is not fulfilled then. As a result, the dimension of the common stochastic trends space may be overestimated and hence some cointegration relations will not be identified.

Because the common stochastic trends I(2) are present, matrix $\mathbf{B}$ containing both stationary relations CI(2,2) and non-stationary combinations CI(2,1) cannot have a single interpretation. Decomposition of (12)–(13) offers the following advantages.

The relations inside matrix $\mathbf{B}$ are always long run, but some of them can be observed in the medium run too. The speed of adjustments to the relations is also different. Matrix $\mathbf{B}_1$ contains relations CI(2,1) that take place in the long run but not necessarily in the medium run and the adjustments they bring about are quantified by the elements of matrix $\mathbf{A}_1$. The fast adjustment, i.e. towards the relation CI(2,2), are expressed by the elements of matrix $\mathbf{A}_0$.

Matrices $\mathbf{A}_{1\perp}$ and $\mathbf{A}_{2\perp}$ have similar interpretations. Matrix $\mathbf{A}_{1\perp}$ defines the medium-run stochastic trends I(1), while $\mathbf{A}_{2\perp}$—the long-run ones. Therefore, testing the restrictions imposed on this matrix could be an alternative to the standard ADF tests, especially when economic arguments exist that the variable in question may be integrated of order 2 (see Juselius 2006).

4 Empirical results

The investigation used deseasonalized monthly data on Poland spanning the period January 1993–December 2008, so the sample had 192 observations. Labour productivity was measured with value added in industry (constant prices) divided by the number of employees.

The nominal stock data (prices in particular) are suspected to be generated by stochastic processes integrated of order two which is supported by economic analysis (see Juselius 2004). The results of the ADF and KPSS tests presented in Table 1 do not support this hypothesis, perhaps because they underestimate the order of integration (see the discussion in: Kębłowski and Welfe 2004). Nevertheless, the results of a system test using restrictions imposed on the elements of matrixes $\tilde{\mathbf{B}}_{2\perp}$ and $\mathbf{A}_{2\perp}$ suggest that stochastic trends I(2) are present.

The results of a joint test for cointegration rank (see Rahbek et al. 1999) shown in Table 2 allow concluding that the system has five linearly independent cointegration vectors, but no common stochastic trends I(2) (see Haldrup 1999; Paruolo 1996). It seems that the absence of the I(2) trends should be attributed to the fact that in a system thus defined they undergo annihilation and become stationary as soon as they form a direct cointegration relation CI(2,2). Consequently, no polynomial cointegration relations exist and the number of direct cointegration relations is *R*. Because in the analysed system only this type of cointegration relations was identified, this means that they are long-lived and that an adjustment to the equilibrium path follows a possible shock very quickly.

At the next step, a standard VEqCM model was considered, where the initially assumed order of lags was five. Then the order of lags was reduced to three, according to the standard SBC criterion (see Table 3).

The model's structure points to the following group of potential (weakly) exogenous variables: the price index of imported raw materials and intermediate goods, the price index of imported consumer goods, labour productivity in industry, the rate of unemployment, value added tax (VAT), the rates of payroll taxes and the share of services in total consumption (constant prices). The results of a likelihood ratio test *LR* (see Table 4) show that except for the services' share in total consumption and payroll taxes all other variables are weakly exogenous in the long run (for a structurization strategy see Greenslade et al. 2000).

Exogeneity was tested again, assuming this time that the five earlier identified variables were (weakly) exogenous. The test results (see Table 5) allowed adding the two isolated variables into the set of weakly exogenous variables.

Table 1 Testing variables for integration order

Variable	ADF	KPSS	$\tilde{\mathbf{B}}_{2\perp}$	$\mathbf{A}_{2\perp}$
wc	I(0)	I(2)	I(2)	I(2)
pm^F	I(2)	I(1)	I(2)	I(2)
pm^I	I(1)	I(1)	I(2)	I(2)
z	I(1)	I(1)	I(2)/I(1)	I(2)
tv	I(1)	I(1)	I(2)	I(2)
U	I(1)	I(1)	I(2)/I(1)	I(2)
ps	I(0)	I(2)	I(2)	I(2)
pc	I(0)	I(2)	I(2)	I(2)
p	I(0)	I(2)	I(2)	I(2)
pr	I(1)	I(1)	I(2)	I(2)
h	I(1)	I(1)	I(1)	I(2)
tx	I(1)	I(1)	I(1)	I(2)

Table 2 Results of the joint test for cointegration order (the numbers in the brackets are p values)

$M-R/P_2$	7	6	5	4	3	2	1	0
7	1342.129 (0.000)	1100.536 (0.000)	939.581 (0.000)	814.554 (0.000)	740.625 (0.000)	690.040 (0.000)	651.160 (0.000)	637.305 (0.000)
6		930.512 (0.000)	756.077 (0.000)	639.033 (0.000)	550.043 (0.000)	476.391 (0.000)	441.354 (0.000)	425.486 (0.000)
5			600.349 (0.000)	473.495 (0.000)	386.312 (0.000)	328.067 (0.000)	294.706 (0.000)	268.762 (0.000)
4				366.932 (0.000)	293.449 (0.000)	233.096 (0.000)	195.520 (0.000)	164.941 (0.000)
3					211.659 (0.000)	163.451 (0.000)	128.158 (0.000)	95.718 (0.002)
2						113.651 (0.000)	74.969 (0.000)	54.392 (0.022)
1							34.303 (0.000)	24.625 (0.076)

Table 3 Testing for order of lags

Order of lags	SBC
1	−60.34
2	−67.22
3	**−70.46**
4	−69.09
5	−68.58

Bold points lag order

Table 4 Results of testing for weak exogeneity (the critical value is 14.067, because RV = 7)

Variable	$\chi 2(7)$
lh	84.37
lp	72.19
lpc	68.02
lps	57.21
ltx	43.23
lpr	31.98
lwc	19.87
**lpm^I	10.94
***ltv*	9.34
***U*	9.30
***lz*	7.81
**lpm^F	7.34

** Denotes variables identified as weakly exogenous

The order of cointegration was found using standard tests: a trace test λ_{trace} and a maximum eigenvalue test $\lambda_{\max}$, although the first one is believed to be more reliable (see Lüutkepohl et al. 2001). Because the results are sensitive to system marginalization, inference was performed assuming (weak) exeogeneity of the variables (see Table 6). The asymptotic critical values (for a 5% level of significance, see Pesaran et al. 2000) were used, because they are relevant to inference based on a well-identified conditional model. According to the results, the system has five independent cointegration vectors.

The long-run structure was identified by imposing the exclusion restrictions and homogeneity conditions on the appropriate elements of matrix **B** (discussion of the homogeneity conditions in: Grabowski and Welfe 2010) Twenty-five restrictions resulting from Eq. 7a–7e and five normalization conditions produced 30 testable restrictions for the long-run parameters. Because $\chi^2 = 16.92 < 43.773$, there were no grounds for rejecting the restrictions (see Table 7). The constant term appeared in both the cointegration space and beyond it, according to the identity making use of the orthogonal complements of matrix **A** (see Johansen 1996). Moreover, all major restrictions were tested one by one and none of them was rejected.

The random terms in particular equations are normal and uncorrelated. The ARCH effect was not found (see Table 8).

Table 5 Results of testing for weak exogeneity—assuming system marginalization (the critical value is 11.07, because RV = 5)

Variable	$\chi 2(5)$
***lh*	9.17
lp	42.59
lpc	38.57
lps	40.32
**ltx*	10.32
lpr	27.88
lwc	15.59

** Denotes variables identified as weakly exogenous

Table 6 Results of testing for cointegration order (limited trend, unlimited constant term, assumed (weak) exogeneity of variables)

Order	Eigenvalue	TRACE	*p* value	MAX	*p* value
0	0.50053	368.87	0	131.21	0
1	0.42237	237.67	0	103.73	0
2	0.19980	133.94	0	42.126	0.0138
3	0.18013	91.815	0	37.538	0.0072
4	0.15197	54.277	0.002	31.154	0.0066
5	0.11392	23.124	0.1055	22.860	0.0124
6	0.00139	0.2634	1	0.263	1

Table 7 Results of restriction tests

Restriction	LR value	Critical value
$wc - z$	2.32	3.841
$pc - tv$	1.77	3.841
$ps = pc$	2.14	3.841
$pc + ps$	1.16	3.841
$pm^F + pr$	1.19	3.841
$wc - tx - p - z$	3.72	7.815
Total "homogeneity"	6.23	15.507
All restrictions	16.92	43.773

The following parameter estimates of the long-run equilibrium equations were obtained (*t*-statistics are in the brackets below estimates):

$$pr = 0.456 + \underset{(8.33)}{0.167}\, pm^I + 0.823(wc - z) \tag{22a}$$

$$pc - tv = -0.082 + \underset{(3.45)}{0.152}\, pm^F + \underset{(6.19)}{0.848}\, pr \tag{22b}$$

$$ps = 1.084 + pc + \underset{(7.25)}{2.183}\, h \tag{22c}$$

Table 8 Specification tests

Equation	Autocorrelation test LM $\chi2(2)$	Normality test $\chi2(2)$	ARCH(3) $\chi2(3)$	R^2
pr_t	2.251	1.012	7.675	0.898
pc_t	1.372	4.987	6.198	0.934
ps_t	4.298	4.227	5.986	0.975
p_t	2.126	5.023	6.129	0.942
wc_t	5.125	1.048	3.098	0.981

$$p = 0.003 + \underset{(23.47)}{0.695}\, pc + \underset{(49.92)}{0.305}\, ps \tag{22d}$$

$$wc - tx - p - z = 0.401 - \underset{(-14.33)}{0.037}\ U \tag{22e}$$

The results, fully interpretable in economic terms, allow identifying the mechanisms through wage pressures stimulate inflation, the impact of consumer prices on wages, and the role played by payroll taxes.

The elasticities of the producer price index towards payroll expenses and the price index of imported raw materials and intermediate goods are 0.823 and 0.167, respectively (Eq. 22a). They are similar to the categories shares in total production costs. The growing significance of payroll expenses is also notable (the same elasticity calculated for the period preceding Poland's EU membership was 0.742, see Kębłowski et al. 2008).

The estimates of the consumer price equation do not stir any reservations, because the respective elasticities are approximately equal to the shares of imported and domestic goods in personal consumption. In the case of the CPI equation, the elasticity estimates are also similar to the shares of goods and services in total household consumption.

The estimate of net real wages elasticity with respect to the rate of unemployment is negative (−0.037) and not different from the results obtained for other European economies. The assumed unit elasticity of net real wages with respect to productivity of labour guarantees that improving productivity, *ceteris paribus*, will neither allow entrepreneurs to earn unlimited profits in the long term (if the elasticity were lower than one), nor will it cause persistent market disequilibrium (if the elasticity were greater than one). On the other hand, the unit elasticity of net wages with respect to prices is based on the assumption that long-term inflation, *ceteris paribus*, should not bring real wages down, thus reducing households incomes.

A non-growing share of services in total consumption effectively defuses the upward pressure on services prices (see Eq. 22c). Rising direct taxes (VAT) and payroll expenses accelerate inflation, as this factor operates for a long time (the appropriate elasticities are close to one).

5 Conclusions

All I(2) shocks in the wage-price nexus model including taxes and payroll expenses are directly cointegrated and form stationary relations (of the CI(2,2) type), so applying the polynomial cointegration is not necessary. In other words, a cost-push inflation model can be analysed with a standard model of the I(1) variables without the risk that any meaningful statistical and economic information will be lost, although the I(2) variables are also present in the model.

A simple integration analysis as well as a more complex examination involving the VEqCM model provided solid grounds for the conclusion that prices in the Polish economy are I(2), so they can be shaped by effective anti-inflationary policy.

A notable fact is that the results are resistant to Poland's entry into the European Union and to the global crisis in the financial markets, as shown by their significant similarity to those obtained in the study where the sample ended in 2003.

Acknowledgments Financial assistance of KBN grant no N111 032 32/3723 is gratefully acknowledged.

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