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by

Gaetano D'Adamo

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Departamento de Estructura Económica (Economía Aplicada II), Facultad de Economía, Universitat de València
Avda dels Tarongers, s/n, 46022 – Valencia, Spain
Phone +34 96 382 83 49 / Fax +34 96 382 83 54
http://www.estructuraeconomica.es
e-mail: investigacion.estructura@uv.es

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Wage spillovers across sectors in Eastern Europe

Gaetano D'Adamo*

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Abstract

This paper studies the interactions between wages in the public sector, the traded private sector and the closed sector in ten EU Transition Countries during the period 2000-2010. The theoretical literature on wage spillovers, as well as the Balassa-Samuelson hypothesis, suggest that the internationally traded sector should be the leader in wage setting, with sheltered and public sector wages adjusting. Using a Cointegrated VAR approach we show that a large heterogeneity across countries is present, and non-traded and public sector wages are often leaders in wage determination or at least affect traded sector wages in the short run. In some countries, public sector wages are weakly exogenous, with the private sectors adjusting. This result is relevant from a policy perspective since wage spillovers, leading to costs growing faster than productivity, may affect the international cost competitiveness of the traded sector and thus the catching-up process may be accompanied by accumulation of large international imbalances.

JEL Classification: C32; E62; J31
Keywords: Cointegrated VAR, wage setting, public sector.

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Correspondance address: Departamento d'Estructura Economica – Facultat d'Economia - Av.da dels Tarongers s/n, Valencia. E-mail address: gaetano.dadamo@uv.es.

1. Introduction

The aim of this paper is to study the interactions and spillovers in wage determination across different macro sectors. In particular, we analyze this interplay in three broad sectors: the open (internationally traded) sector, the closed or “sheltered” sector, which can also be called the market non-traded sector, and the public sector, also called the non-market, non-traded sector. The focus of this work is on Central and Eastern European Countries (CEEC) that have joined the European Union during the last decade.

This issue is relevant for several reasons. First, public sector employment is large: in the OECD, around 25% of the work force is employed in the public sector. Second, wage spillovers across sectors may lead to wage costs growing faster than productivity and this may affect the international cost competitiveness of the country's tradable goods sector. In particular, theoretical models generally assume that the traded sector is leader in wage determination and there is free mobility of labor across sectors which, in turn, ensures wage equalization. This is assumed, for example, by the Froot and Rogoff [1995] model of the Balassa-Samuelson (henceforth B-S) effect, and the so-called Scandinavian Model of wage determination (Aukrust [1970], see Section 3). If these assumptions hold, during the process of catching-up excess inflation will be witnessed and the real exchange rate will appreciate, but, on the other hand, only wages in the non-traded sector should grow faster than productivity, and therefore competitiveness should not be harmed. However, it can also happen that wage bargaining in the non-traded sector – which is not subject to international competition – and/or in the public sector – which is influenced by political rather than productivity considerations – may lead to higher outcomes and, in turn, push traded sector wages up. A large strand of empirical literature that has tried to measure the B-S effect has found that it can only account for a small part of the excess inflation observed in Central and Eastern European Countries[1], while other factors like the switch to consumption of higher quality goods might be at play[2]. This paper, by rigorously testing whether the hypothesis of wage leadership of the traded sector is valid, studies an additional potential source of fallacy of the B-S hypothesis[3].

Further, testing wage leadership in the case of Central and Eastern European transition Countries is important because they are in the process of catching up: entry in the EU, which also fostered international labor mobility, may have influenced wage determination in these countries, so that prices and wages converge to the western European level faster than productivity, thus leaving room for competitiveness loss. As long as the leader in wage setting is the traded sector, the problem is not relevant. But if this is not the case, increased labor mobility together with union pressures might lead to a decoupling of real and nominal convergence, i.e. price convergence occurring faster than productivity convergence. From a policy perspective, this would be worrying since we would expect catching-up to occur at the cost of large international imbalances.

The contribution of this paper is twofold: first, while there is limited existing literature on public/private wage spillovers, no empirical work to date studies the issue from the three-sector perspective that is adopted here. Second, this is the first work on the topic focusing on transition countries.

The paper is organized as follows. Section 2 reviews the literature on wage setting and wage spillovers across the public and private sector. Section 3 presents three alternative models of wage determination and causality relationships. Section 4 outlines the empirical model, a Cointegrated VAR; section 5 describes the data set and presents the results of the empirical analysis. Section 6 concludes.

2. Related literature on wage spillovers

While this is not a paper on wage setting models, the literature on wage setting is indeed relevant for the topic. In particular, since we are dealing with the issue of externalities in wage determination, we are interested in models studying the impact of centralization of wage setting and its impact on the overall macroeconomic performance. In principle, we can imagine different degrees of centralization in wage setting, ranging from firm-level bargaining (no centralization) to

[1] See, for example, Mihaljek and Klau [2003], Egert [2003 and 2007].

[2] See Egert [2010].

[3] Empirical models testing the size of the B-S effect in CEECs often estimate a regression equation having CPI inflation differential as dependent variable and dual productivity growth differential (productivity growth in tradeables vs. non-tradeables) as explanatory variable. This is the result of assuming that wages in T grow in step with productivity.

national bargaining. The outcome of the bargaining process will therefore depend, among other things, on how much wage setting is centralized. According to the Calmfors – Driffill hypothesis (Calmfors and Driffil [1988]) both firm-level bargaining and national bargaining are likely to produce wage moderation, while industry-level bargaining tends to produce higher real wages. In other words, the relationship between the real wage and the extent of centralization is a hump-shaped curve (see Figure 1). The reason for this is that in the case of no centralization, with perfect competition, firms will not increase wages above productivity because they are price takers, and thus cannot increase prices when the real consumption wage is raised by the firm and not elsewhere in the sector. However, such wage moderation is restrained if firms have market power. On the other hand, with national bargaining, firms and unions internalize the negative externalities produced by aggregate real wage increases[4] and thus the outcome will be a lower wage level. Other things equal, however, when the economy is open the employment and profit losses due to real wage increases are dampened in centralized systems (which therefore tend to have higher wages), but this will in any case depend on the degree of competition between domestic and foreign producers (Calmfors [1993]).

The literature on inter-sectoral wage spillovers, in particular private and public wages, is quite scarce. Theoretical models generally assume that public wages are exogenous or follow the same pattern as private wages (Quadrini and Trigari [2007], Ardagna [2007]). Demekas and Kontolemis [2000], instead, in a static model show that public wages can affect private sector wages through the labor supply channel: when public wages increase, workers move to that sector, and private firms are forced to increase wages too. As far as the direction of causality is concerned, the main theoretical reference is the so-called Scandinavian model of wage determination (Aukrust [1970]). This model assumes that the sector that is open to international competition is the leader in wage setting, since productivity should increase faster in the traded goods sector and firms there cannot increase wages above productivity, because they would otherwise become less competitive[5]. In this sense, also public sector wages should be led by private sector wages. Evidence in favor the Scandinavian Model was found by Aukrust for Norway, the U.S. and France, while Bemmels and Zaidi [1990] successfully applied it to Canada.

However, this model has been found to be at odds with more recent data. Ultimately, the results seem to be country-dependent. Demekas and Kontolemis [2000] find weak exogeneity of government wages over private wages. Jacobsson and Ohlsson [1994], in a Vector Error Correction Model for Sweden, find long-run wage leadership of the private sector, thus confirming the predictions of the Scandinavian model, but Friberg [2007], using a broader sectoral decomposition[6] does not find evidence of the Scandinavian model for Sweden.

Christou et al. [2007] show a bidirectional causality relationship between private and public sector wages in Romania using monthly wages over the period 1993-2006.

Lamo et al. [2008] used several empirical methods to study the co-movement and causality relationship between private and public wages using annual data for 18 OECD countries (plus the Euro Area as a whole), finding that private and public wages generally do not decouple and the former seem to exert a stronger influence on the latter than the reverse; moreover, they find that prices seem to play an important role in the transmission of wage leadership.

None of the cited works is applied to transition countries (with the exception of Christou et al. [2007]), and, more importantly, none decomposes the private sector into a traded and non-traded sector, which is relevant when we want to understand the role of wage spillovers in affecting competitiveness. In fact, as we will see in the next section, not only a leading role of public sector wages but also of non-traded goods sector wages can lead to traded sector wages growing faster than productivity and thus harm the country's international competitiveness.

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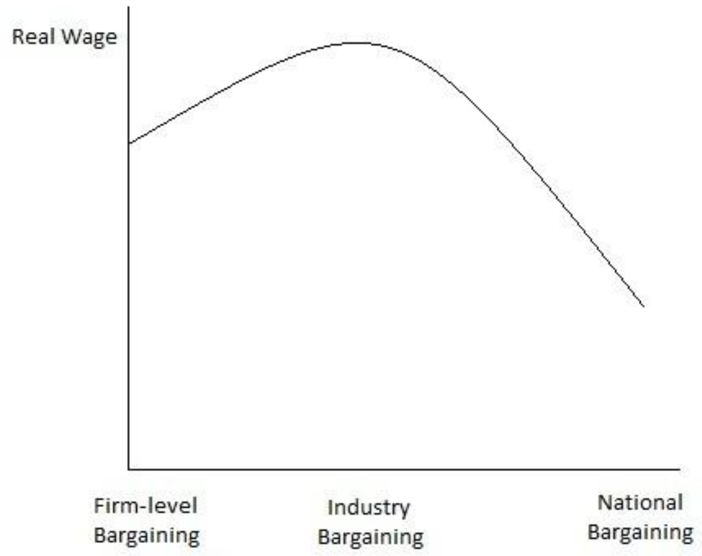
[4] There are several kinds of externalities in this sense: Consumer price externalities, since prices will increase (Layard et a. [1991]); Input price externalities, since the price of inputs will go up; fiscal externalities due to higher unemployment benefits that have to be paid since higher wages reduce employment (Blanchard and Summers [1987], Calmfors and Driffill [1988]; envy externalities, if the welfare of an agent depends on the wage of other agents, and thus we have forms of social comparisons (Oswald [1979] and others). For a broader discussion, see Calmfors [1993].

[5] An outline of this model is presented in the next section.

[6] In particular, he distinguishes between private sector, manufacturing sector, construction, wholesale and retail trade, financial sector, central government and county/municipal government.

Figure 1. The relationship between centralization and the aggregate real wage (Calmfors [1993])



3. Three competing theories of wage determination

When dealing with inter-sectoral wage spillovers, theoretical models generally assume that the leader in wage determination is the sector that is exposed to international competition, i.e. the traded sector. However, as it was outlined in the previous section, the results obtained by the empirical literature are often at odds with this hypothesis. Alternative models of inter-sectoral wage linkages can therefore be imagined.

In this section we will outline three alternative theories of wage determination and inter-sectoral wage spillovers. This will allow us to come up with testable hypotheses that will be taken to the data. The first model is the main theoretical reference as far as inter-sectoral wage linkages and expected causality is concerned, and it is the so-called "Scandinavian Model" of inflation, which was first developed by Aukrust [1970][7]. The Scandinavian Model rests on three fundamental assumptions: (1) the different sectors in the economy can be classified as either traded (exposed) sectors or non-traded (sheltered) sectors; (2) wage increases in the traded sectors can be expected to be transmitted to wage increases in the non-traded sectors of the economy, and therefore wage decisions are not taken simultaneously; (3) the exchange rate is fixed. In the original model we have two sectors (traded, T, and nontraded, N) and two countries (home, H, and foreign, F). We add the public sector as a non-traded, non-market sector and label it P. The timing structure of the model is pictured in figure 2, panel (a).

In this highly stylized model, nominal wages in the traded sector are determined by the productivity in that sector and the prices obtainable internationally for the output of those firms[8]:

$$w_{T,t} = p_{T,t} + c_{T,t} + \eta_{T.1,t} \qquad (3.1)$$

where $c_{T,t}$ is productivity, p_T are prices and η_T represents stationary deviations from this long-run equilibrium relation. Labor is mobile across sectors; for this reason wage equalization occurs:

$$w_{N,t} = \mu_{N.1} + \theta_{N.1} w_{T,t} + \eta_{N.1,t} \qquad (3.2)$$
$$w_{P,t} = \mu_{P.1} + \theta_{P.1} w_{T,t} + \eta_{P.1,t} \qquad (3.3)$$

[7] Aukrust [1970] first developed this model to describe price dynamics in Norway, and tested it on that country. However, Aukrust's model was applied to other countries' data to test for inter-sectoral wage linkages: France, USA and Australia (Aukrust [1977]), Canada (Bemmels and Zaidi [1990]), Sweden (Friberg et al. [2004], Friberg [2007], Jacobson and Ohlsson [1994]).

[8] The model outlined here heavily draws on Aukrust [1977].

where $\theta_{i.1} \geq 0$, i = N,P are parameters describing the degree of wage adaptability across sectors. The Scandinavian model concludes that not only should wages in different sectors co-move, but the transmission of wage shocks should be one to one, a testable hypothesis that we will term "full wage adaptability". In other words, according to the Scandinavian Model the deviations from equilibrium, η_j, j = N,P,T should be stationary, $\mu_{N.1} = \mu_{P.1} = 0$ and $\theta_{N.1} = \theta_{P.1} = 1$.
Finally, firms in the non-traded sector set the prices for their goods and services accordingly, in order to avoid losses[9], so that

$$p_{N,t} = w_{N,t} - c_{N,t} + \eta_{pn.t} \qquad (3.4)$$

As a consequence, non-traded goods prices will increase more than traded goods prices, a result which is in line with the so-called "internal version" of the Balassa-Samuelson hypothesis.
A peculiar argument may be relevant for transition countries that have entered the European Union. The catching up process of transition countries, in terms of productivity and price level, has been documented in many studies on the so-called Balassa-Samuelson effect[10]. However, when a country joins the European Union, the liberalization of labor mobility across countries could itself foster wage convergence with the Union. This, in turn, can be due to convergence in productivity, but also to migration of labor to countries with higher wage levels or increased union pressures for wage raises. In the latter two cases, wage increases can harm the competitiveness of the country if they are above productivity growth. In the B-S hypothesis, instead, labor is assumed to be internationally immobile.

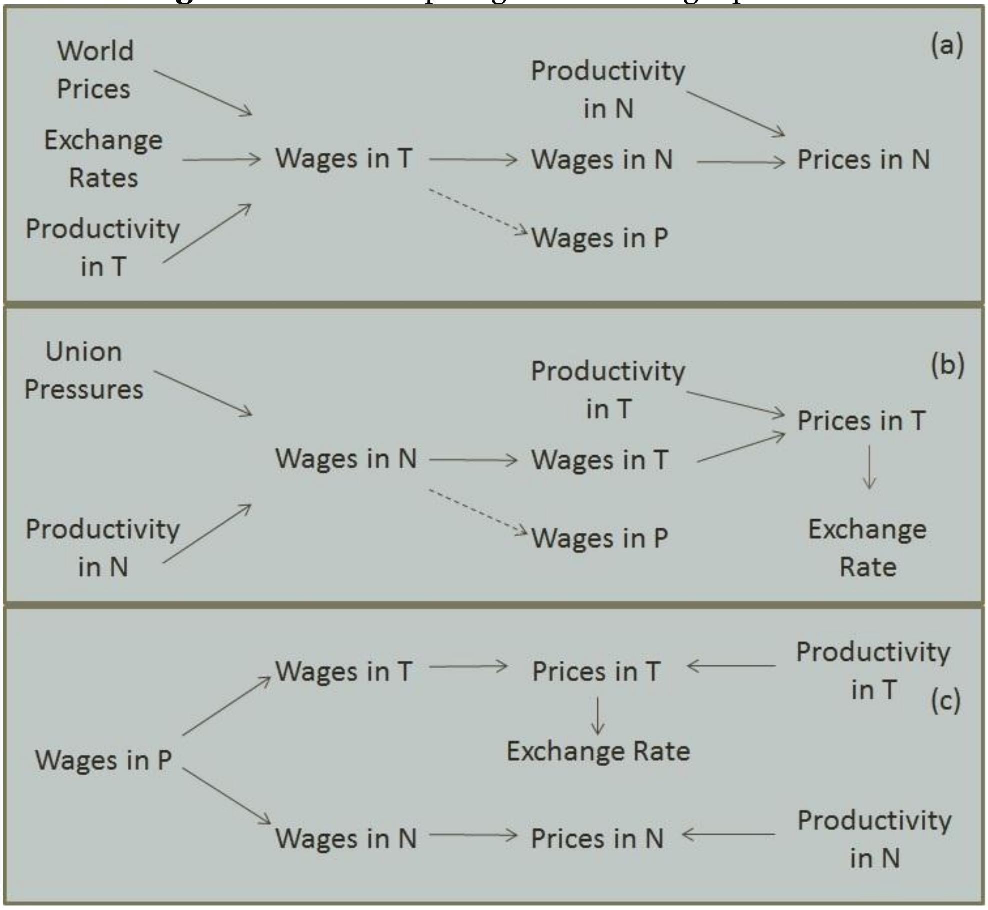
As long as the traded sector is the leader in wage determination, pressures due to international competition will avoid wage increases in excess of productivity. However, alternative models of inter-sectoral wage spillovers might be in place, and the empirical analysis in the next section will prove that this is indeed the case. As it was pointed out by Friberg [2007], non-traded sector firms operate in a less competitive environment, since they are not subject to international competition: therefore, wage bargaining in non-traded sectors may lead to higher outcomes, *ceteris paribus*. This is not the only issue, since in the present paper we further distinguish, within non-traded sector wages, market non-traded and from non-market non-traded (public) sector wages. Theoretical models of public sector wage setting generally assume that wages for public employees are set exogenously or, as in Demekas and Kontolemis [2000], that the government maximizes an objective function in public goods provision and public wages (a form of political patronage). In this sense, also wage bargaining in the public sector may lead to higher outcomes, depending on the political pressure that public employees are able to exert on the government. If the mobility of labor across sectors is high and therefore wages tend to equalize, a leading role in wage setting of either the public sector or the non-traded sector can harm international competitiveness.
We can sketch models of wage determination which are parallel to the Scandinavian model for these two cases.

A model with the non-traded sector leading is pictured in figure 1, panel (b). For practical reasons, we will call it the "wage mark-up model". Nominal wages in the non-traded sector are set according to productivity in that sector, the prices that firms can obtain on the internal market, and the mark-up on productivity that unions are able to extract from employers, m:

$$w_{N,t} = p_{N,t} + c_{N,t} + m_{N,t} + \eta_{N,t} \qquad (3.5)$$

[9] One additional assumption that we have left implicit is that productivity growth is higher in the traded sector than in the non-traded sector. This assumption, which is empirically sound, is also at the basis of the so-called Baumol-Bowen effect that the price wedge between non-tradables and tradables should be co-moving with productivity. If this assumption did not hold, then we could either have p_N growing more than p_T (if firms in N set prices under mark-up pricing) or less than p_T (if there is pricing-to-market in N).
[10] A small subset includes Mihaljek and Klau [2001], Egert [2002a, 2002b, 2002c, 2007], Fischer [2004], Dobrinsky [2006] and Staher [2010].

Figure 2. Three competing models of wage spillovers



Pressures for wage equalization across sectors due to free intersectoral labor mobility, as in the previous case, will make sure that wages in T and P adjust to w_N:

$$w_{T,t} = \mu_{T.2} + \theta_{T.2} w_{N,t} + \eta_{T.2,t} \tag{3.6}$$

$$w_{P,t} = \mu_{P.2} + \theta_{P.2} w_{N,t} + \eta_{P.2,t} \tag{3.7}$$

where again $\mu_{T.2} = \mu_{P.2} = 0$ and $\theta_{T.2} = \theta_{P.2} = 1$ if full wage adaptability holds and η_j, j = N,P,T are stationary i.i.d. disturbances.

The firms in the traded sector, then, will set the prices of their goods according to the level of wages and the productivity in the sector:

$$p_{T,t} = w_{T,t} - c_{T,t} + \eta_{pT.t}{}^{11} \tag{3.8}$$

and this in turn causes real appreciation. Thus, unless a nominal exchange rate depreciation occurs, it will lead to the country's firms becoming less competitive on the international markets[12]. In this case, the presence of low wage adaptability ($\theta_{T.2} < 1$) might offset (at least partially) the negative effect of public sector wages on competitiveness.

Finally, Figure 1, panel (c) pictures the case of the public sector leading the other two. We will call this the "envy-effect model" (Friberg [2007], Strom [1995]).In this case, we assume (following Ardagna [2007] and Quadrini and Trigari [2007]) that the public sector wages are predetermined:

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[11] The implicit assumption here is that $c_{N,t} + m_{N,t} > c_{T,t}$.

[12] In order to avoid losing market share, firms in T might keep prices unchanged for some time, but this strategy would not be sustainable since it will generate losses. Alternatively, they might reduce employment and/or try to push productivity up, for example by eliminating previous slack in the work process (see, for example, Juselius and Ordonez [2009]). However, in order to be able to account for this we would need a more general theoretical and empirical model which is beyond the scope of this paper.

$$w_{P,t} = \overline{w} \tag{3.9}$$

where $\overline{w} > c_{i,t}$, i= N,T, i.e. public wages lead wage determination and they also grow faster than productivity in the other sectors. In fact, while in the private sector the argument of competition between different production units (either different firms in the same sector, or domestic vs. foreign firms) producing wage restraint holds, this is not true for the public sector. Fiscal discipline itself may not be sufficient to produce wage restraint in the public sector, in particular under centralized wage bargaining, because in that case the (central) government would be bargaining with a large share of the electorate (Calmfors et al., [1985], [1988]). As in the previous cases, wage equalization occurs due to free inter-sectoral labor mobility and/or envy externalities (see fn. 4):

$$w_{T,t} = \mu_{T.3} + \theta_{T.3} w_{P,t} + \eta_{T.3,t} \tag{3.10}$$

$$w_{N,t} = \mu_{N.3} + \theta_{N.3} w_{P,t} + \eta_{N.3,t} \tag{3.11}$$

Again, full wage adaptability holds if $\theta_{i.3}$=1, i=N,T. As in the previous case, the consequences for the domestic firms' international competitiveness might be serious due to traded goods prices growing faster than productivity. Moreover, since wages in both market sectors (N and T) will be growing ahead of productivity, in this latter case the impact on CPI inflation will be larger than in the previous cases. The envy-effect framework, thus, introduces an additional negative fiscal externality of wage increases: when the public sector is the leader in wage setting, the overall level of the real wage will be higher, resulting in higher prices and/or lower employment.

Albeit highly stylized, these models give us an idea of the scenarios we are likely to find in an empirical analysis of wage spillovers and leave us with clear testable hypotheses to find out which model fits best in describing the spillovers in wage determination in a specific country.

4. The Empirical Model

The ideal empirical model to test the assumptions embedded in the theoretical models sketched in the previous section, for reasons that will be clarified shortly, is a Cointegrated VAR (Johansen [1988, 1995]). The main advantage of the CVAR is that it is a data-driven approach, and can therefore challenge – as it has been proved in several recent works[13] – theoretical macro models which have often been proved to fail to explain many empirical facts. Instead of pre-specifying the correct economic model from the outset, by using the CVAR we "let the data speak".

One drawback of VAR models is that they are quite data-demanding, which can be a problem when we deal with transition countries, given the limited data availability. However, in this specific case, the dimension of the VAR (three variables) and the characteristics of the series (see next section) make this problem less relevant.

Suppose we have a vector of p variables in time series. When the p variables, in levels, are nonstationary and integrated of order 1, I(1), their first difference will be I(0). Moreover, if two or more variables are I(1) but their linear combination is I(0), they are said to be cointegrated (or CI(1,0)). The CVAR, as we will see, has the variables in first differences as dependent variables but is simply a reparametrization of the VAR (see Juselius [2006], Chp.2). Therefore, the value of the likelihood is the same and there is no information loss in moving from the VAR to the Cointegrated (or Vector Error Correction) counterpart.

A Cointegrated VAR model of order n with p variables is defined as:

$$\Delta x_t = \Pi \tilde{x}_{t-1} + \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_n \Delta x_{t-n} + \phi D_t + \mu + \varepsilon_t; \;\; \varepsilon_t \sim Niid(0, \Omega) \tag{4.1}$$

where D is a vector of dummy variables (seasonals and other unrestricted dummies), μ is a vector of constant terms[14] and Ω is the (pxp) covariance matrix of (white noise) residuals. Γ_j, j=1,...,n are

[13] See, for example, Juselius and Franchi [2007].

[14] Notice that an unrestricted constant in the differences is equivalent to a linear trend in the levels of the variables.

the matrices of the short-run coefficients. When variables are I(1) and cointegrated, the matrix Π will be of rank r < p, where r is the number of (long-run) cointegration relations. Therefore, the Π matrix can be decomposed as:

$$\Pi = \alpha\tilde{\beta}'; \ \tilde{\beta} = [\beta \quad \mu_0 \quad \beta_d \quad \beta_t]; \qquad \beta = [\beta_T \quad \beta_N \quad \beta_P]$$

$$\widetilde{x'}_{t-1} = [x_{t-1} \quad 1 \quad D_s \quad t]; x'_{t-1} = [w_{T,t-1} \quad w_{N,t-1} \quad w_{P,t-1}]$$

where the elements in $\tilde{\beta}$ are r x 1 vectors, 1 represents a constant restricted to lie in the cointegration relation, D_s are dummies capturing shocks that do not cancel in the cointegration relations and t is a linear trend. The $\tilde{\beta}$'s are called the *cointegration relations* while the α are the *loadings*. In other terms, the long run (stationary) relations that characterize the variables are the $\tilde{\beta}$'s, while the α's show how each variable adjusts to disequilibria in the corresponding long-run relation. When we decompose Π, the coefficients in the α and β matrices will not be, in general, identified (i.e. we have an infinite number of matrices α and β that, if multiplied, are equal to Π); therefore, in order to achieve identification, we will need to impose restrictions on β: a necessary (but not sufficient) condition for a Cointegrated VAR model to be empirically identified is that *r*-1 restrictions are imposed on each long-run relation. If the number of restrictions is larger, the model is *overidentified* and such restrictions are testable.
Including a linear term (the time trend t) in the cointegration relations is useful when variables in the system are trending and trends do not cancel in the cointegrating relations; the constant, instead, allows for a non-zero mean in the relations.
The CVAR (4.1), in sum, says that the changes in the variables in each period are explained by adjustments to equilibrium relations and the effect of past changes in the variables of the system, plus some deterministics; all contemporaneous effects are in the covariance matrix, i.e. it is a reduced form model.
Moreover, the CVAR classifies the variables in *r* long-run relations by which the system is pulled and *p*-*r* common stochastic trends, by which the system is pushed. In other words, when the rank is *r*, the shocks in the equations of the VAR can be rewritten as *r* transitory shocks (i.e. shocks that cause disequilibria which are gradually absorbed through adjustment to the long-run relations) and $p - r$ permanent shocks. We can therefore re-write the CVAR in the so-called common trends representation:

$$x_t = C\sum_{i=1}^{t}(\epsilon_i + \mu_0 + \phi_1 Dp_i) + C^{\circ}(L)(\nu_t + \mu_0 + \mu_1 t) + \tilde{X}_0 \qquad (4.2)$$

where Dp are permanent (shift) dummies, ν are the transitory shocks, $\tilde{X}_0$ the initial values and

$$C = \beta_{\perp}(\alpha_{\perp}\Gamma\beta_{\perp})^{-1}\alpha'_{\perp} = \tilde{\beta}_{\perp}\alpha'_{\perp} \qquad (4.3)$$

where $\alpha_{\perp}$ and $\beta_{\perp}$ are the p x (p − r) orthogonal complements of α and β describing the common stochastic trends, $\alpha_{\perp}\sum\epsilon$, and their loadings.
If a variable is found to be weakly exogenous, i.e. it does not adjust to any cointegrating relation[15], then shocks to that variable are identified as one of the common stochastic trends of the system. In other words, the weakly exogenous variable "causes" movements in the other variables.

It should now be clear why the Cointegrated VAR model is a natural candidate to test empirically the predictions of the theoretical model sketched in the previous section. Theoretical models of wage determination make precise statements on the long-run relations as well as the causal links between wages in different sectors.
The Scandinavian Model and the Froot and Rogoff [1995] model of the Balassa-Samuelson effect state that the open sector should be the driving force of the system, since on one hand productivity grows faster in this sector (as it has been empirically observed), and on the other hand it faces international competitiveness and therefore wages should increase in step with productivity to

[15] This amounts to testing that the corresponding row in the α matrix is zero.

make sure that prices to not grow and there is no competitiveness loss. Free mobility of labor across sectors, then, makes sure that wages are equalized.
The hypothesis of wage equalization (or, more precisely, constancy of the wage ratio)[16] implies that in a model including wages in the three sectors (industry, services and public administration) we should be able to find two cointegrating (long-run) relations and one common trend[17] and the coefficients in the cointegration relations should satisfy long-run homogeneity, (i.e. we should find one-to-one long-run relations and the constant in the cointegration vectors should be zero). Finally, in the Scandinavian Model, the common trend should be identified with shocks to industry wages, i.e. industry wages should be weakly exogenous. On the other hand, if the hypothesis of one-to-one relations (coefficients in the β – vector equal to 1) is rejected, this is interpreted as low wage adaptability between the sectors considered (see Friberg [2007]). If less than two cointegration relations are found, this means that there is more than one common trend affecting the wages: in other words, a weaker form of our three theoretical models holds. For example, if both N and T are found to be weakly exogenous, with public sector wages adjusting, this would imply that the (market) traded and non-traded sector wages are subject to different kinds of shocks (i.e. bargaining processes), and also a lower mobility of labour across sectors is present. By classifying the model into *pulling* and *pushing* forces, we will be able to identify how the wage shocks load into the three sectors. If the weak exogeneity of traded sector wages does not hold, the empirical model suggests that the Scandinavian model is not a good approximation of wage dynamics in the country under analysis, and the hypotheses at the basis of the Balassa-Samuelson effect do not hold.
Put it another way, absence of weak exogeneity of w_T would imply that the traded sector has not been a driving force in wage determination, and rather it has been influenced by shocks to the other sectors, leaving room for competitiveness loss if wages have grown more than productivity.
Equation (4.1), where we omitted, for simplicity, the trend and the constant from the cointegration relations, before any restrictions are imposed and making the matrices explicit, can be written as:

$$\begin{bmatrix}\Delta w_{T,t}\\ \Delta w_{N,t}\\ \Delta w_{P,t}\end{bmatrix} = \begin{bmatrix}\alpha_{1T} & \alpha_{2T} & \alpha_{3T}\\ \alpha_{1N} & \alpha_{2N} & \alpha_{3N}\\ \alpha_{1P} & \alpha_{2P} & \alpha_{3P}\end{bmatrix}\begin{bmatrix}\beta_{1T} & \beta_{1N} & \beta_{1P}\\ \beta_{2T} & \beta_{2N} & \beta_{2P}\\ \beta_{3T} & \beta_{3N} & \beta_{3P}\end{bmatrix}\begin{bmatrix}w_{T,t-1}\\ w_{N,t-1}\\ w_{P,t-1}\end{bmatrix} + \sum_{i=1}^{n}\Gamma_i \Delta x_{t-i} + \phi D_t + \mu + \varepsilon_t \quad (4.4)$$

Table 1 summarizes the hypotheses embedded in the theoretical models outlined in Section 3 and the corresponding testable restrictions within the Cointegrated VAR framework.

5. The Data and empirical results

Our dataset contains quarterly data from 2000Q1 until 2010Q3 for ten Central and Eastern European Countries that have joined the European Union in either of the two waves of 2004 and 2007 (Bulgaria, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovak Republic, Slovenia). The time sample was chosen in order to make sure we had data for all countries, as well as excluding data from the very first years after the fall of Communism which might be less reliable or, for what concerns the consumer price index, depend heavily on price administration.

There are three wage series for each country: Industry (proxying the traded sector), Services (proxying the market non-traded sector) and Public Administration (the non-market, non-traded sector)[18].
While the definition of the Public Sector (wages in public administration, defence and compulsory social security) is not really debatable, some assumptions had to be done in order to define the traded and non-traded sector.

[16] Since we have wage *indexes*, when testing the hypothesis of wage adaptability we will actually test the constancy of the wage *ratio*, rather than equality of wages, which a weaker hypothesis but it is more realistic and, on the other hand, it is the assumption actually made by the empirical literature on the Balassa-Samuelson effect.
[17] Moreover, the model should be specified with an unrestricted constant and and no trend, since the presence of a trend in the cointegrating relation would imply that one sector has been gaining purchasing power with respect to the others.
[18] The exact definition of the series and the sources is given in Appendix 1.

Table 1. Testable hypothesis on wage spillovers

Model	Cointegration Rank	Leader	Identifying Restrictions[19]	Weak Exogeneity	Full Wage adaptability
1. Scandinavian Model	2	w_T	$\begin{bmatrix}\alpha_{1T} & \alpha_{2T}\\ \alpha_{1N} & \alpha_{2N}\\ \alpha_{1P} & \alpha_{2P}\end{bmatrix}\begin{bmatrix}\beta_{1T} & 1 & 0\\ \beta_{2T} & 0 & 1\end{bmatrix}$	$\alpha_{1T} = \alpha_{2T} = 0$	$\beta_{1T} = \beta_{2T} = -1$
2. Wage mark-up model	2	w_N	$\begin{bmatrix}\alpha_{1T} & \alpha_{2T}\\ \alpha_{1N} & \alpha_{2N}\\ \alpha_{1P} & \alpha_{2P}\end{bmatrix}\begin{bmatrix}1 & \beta_{1N} & 0\\ 0 & \beta_{2N} & 1\end{bmatrix}$	$\alpha_{1N} = \alpha_{2N} = 0$	$\beta_{1N} = \beta_{2N} = -1$
3. Envy-effect model	2	w_P	$\begin{bmatrix}\alpha_{1T} & \alpha_{2T}\\ \alpha_{1N} & \alpha_{2N}\\ \alpha_{1P} & \alpha_{2P}\end{bmatrix}\begin{bmatrix}1 & 0 & \beta_{1P}\\ 0 & 1 & \beta_{1P}\end{bmatrix}$	$\alpha_{1P} = \alpha_{2P} = 0$	$\beta_{1P} = \beta_{2P} = -1$

To this end, we have followed a common practice in the literature on the Balassa-Samuelson effect by identifying the traded sector with industry (excluding construction, which is not unanimously treated in the literature) and the market non-traded sector with "Services of the business economy"[20]. Agriculture is excluded from the sample as agricultural prices heavily rely on state support and price administration, and are influenced by policies as the CAP in the EU.

Finally, the wage indexes are deflated using the Harmonized CPI index. This is done in order to concentrate on the relationship between *real* wages, excluding price shocks from the model.

Most of the empirical works on this issue define wages as compensation per employee in the corresponding sector, calculated as compensation of employees divided by the number of workers in that sector; however, these data are only available on an annual basis and the available time series for transition countries are too short to use annual data[21]. The dimension of the VARs estimated in this paper makes the use of quarterly data suitable; moreover, unlike other long-run equilibrium relations, deviations from a long-run equilibrium of wage series should be less persistent, and therefore 11 years represent a long enough time span to detect cointegration. Finally, we prefer using quarterly data to annual data because the latter can be misleading – or at least less informative – if inter-sectoral wage spillovers occur within the year, i.e. with a frequency which is higher than that of the sample. As we will show through impulse response analysis, indeed, disequilibria in the wage relations are generally absorbed within the year.

As already stated, the empirical model is a Vector Error Correction Model in three variables: (real) wages in Industry, w_T , Services, w_N, and Public Administration, w_P. Since the wage series, from a graphical inspection, seem to exhibit a deterministic trend, we also include a linear term in the baseline VAR. As a result, the corresponding VECM can either have an unrestricted constant and no trend (if the trend cancels out in the cointegrating relations), or an unrestricted constant and a restricted trend, in the case the latter does not cancel out in the cointegration relations[22,23].

Table 2 shows the results of the cointegration rank tests. As it is common in the literature, we do not limit ourselves to looking at the results of the trace test but combine it with information on the largest unrestricted root of the characteristic polynomial, the significance of the adjustment coefficients in the corresponding row of the α-matrix, and a graphical inspection of the cointegration relations. We do this in order to make sure that, by excluding a cointegration relation, albeit persistent, we are not throwing away potentially important information. These criteria need not (and in general, do not) suggest the same rank, as it is the case in Table 2, and in this sense some room for judgment is left. Table 2 reports the results of the trace test and the modulus of the largest unrestricted root of the characteristic polynomial. As a rule of thumb, when

[19] The restriction to 1 in this column is just a normalization, and thus it is not an identifying restriction, but is included here for clarity. Each vector is normalized on the variable that is significantly adjusting to that cointegration relation.

[20] The literature generally identifies the open sector with industry and the sheltered sector with services or "all the rest". For example, Egert [2002, 2003a, 2003b] and Golinelli and Orsi [2002] define Industry as the Open sector and the rest as the closed sector. Halpern and Wyplosz [2001] include respectively Industry and Services, and exclude construction from the latter. Agriculture is excluded in Coricelli and Jazbec [2004]. Nenovsky and Dimitrova [2002] include also construction in the open sector.

[21] For example, Lamo et al. [2008] and Demekas and Kontolemis [2000]. Friberg [2007] uses average monthly wages.

[22] See Juselius [2006], Chp. 6. A linear term in the cointegration vector here would imply that wages one sector have been gaining purchasing power with respect to wages in the other sector(s).

[23] The exact specification of the deterministic part and the order of the underlying VARs is given in Appendix 2.

using quarterly data the chosen rank should not leave out roots larger than 0.83-0.84 (Juselius [2006], chp. 8).
The trace test suggested no cointegration in the case of Lithuania and Slovenia, but this was not supported by the other criteria we have listed above, and therefore we opted for $r = 1$.
Only for three countries out of ten (Estonia, Romania and Slovakia) we could not reject the null of $r = 2$. In the other cases, the rank is 1. This might suggest low intersectoral mobility across sectors, so that the wage series do not share a single common trend but two different trends.

We can now move to testing restrictions on the cointegration vectors, i.e. the β matrices, in order to find out the degree of wage adaptability; on the other hand, by testing restrictions on the α matrices (i.e. testing for weak exogeneity) we will be able to classify the variables in the system in pushing and pulling forces, i.e. variables that drove the system out of equilibrium and variables which brought the system back to equilibrium, adjusting to the cointegration relations. Then, we will employ Granger non-causality tests (Granger [1969]) to test for wage leadership. The process of adjustment of the variables after a shock to other sectors will be investigated by means of impulse response analysis.
Tables 3-5 report the results of the empirical model. Table 3 provides the null hypothesis, the imposed restrictions and normalization on the β-vectors and the likelihood ratio test results.
If we have, for example, two cointegration relations and the null hypothesis of one-to-one relations (i.e. long-run homogeneity) can be rejected at the 10% level, then we interpret this as a signal of low wage adaptability (see Friberg [2007]). In order to achieve economic identification, i.e. making the cointegration vectors interpretable as equilibrium relations, the vectors in Table 3 were normalized to the variable which is significantly adjusting to the corresponding cointegration relation.

Table 4 reports the results of the weak exogeneity tests and Table 5 shows the common trends representation of the model, where the coefficients to the stochastic trends of the system are reported. We can now comment on the results by country.
Bulgaria shows only one cointegration relation and Industry and Services are weakly exogenous. The joint hypothesis of weak exogeneity could not be rejected with a fairly high p-value, and since there is only one cointegration vector, the model is identified without having to impose restrictions on β. Looking at the common trends representation, we see that shocks to industry wages had a significant permanent effect only on the industry sector, while shocks to the services sector significantly affected the services and public sector. In other words, we can label the two permanent shocks a “traded sector wage shock” and a “sheltered sector shock”. Overall, a weak version of the Scandinavian Model therefore holds for Bulgaria.
The Czech Republic exhibits one cointegration vector; the two closed sector wages, w_N and w_P, are found to be weakly exogenous and the joint weak exogeneity could not be rejected with a p-value of 0.170. Wages in T, therefore, have been adjusting, and table 4 shows that the sector that has been leading on T was the public sector. Wages in N were found to be long-run excludable, and full wage adaptability between T and P was rejected. This structure of wage spillovers is consistent with a weak version of our “Envy-effect model” and shows a potential for wage costs increasing more than productivity and competitiveness loss in the Czech Republic.
Estonia has two cointegration relations; however, there is not full wage adaptability since the hypothesis of long-run homogeneity is rejected at all significance levels. On the other hand, industry wages are weakly exogenous (though only at 1%) and shocks to this sectors load significantly on the other sectors.

Table 2. Cointegration rank.

Country	Null	Trace test p-value	Largest unrestricted characteristic root		Choice
Bulgaria	r=0	0.030	r=1	**0.748**	
	r≤1	**0.433**	r=2	0.911	r=1
	r≤2	0.898	r=3	1.045	
Czech Republic	r=0	0.021	r=1	**0.428**	
	r≤1	**0.060**	r=2	0.935	r=1
	r≤2	0.204	r=3	0.916	
Estonia	r=0	0.025	r=1	0.712	
	r≤1	**0.061**	r=2	**0.789**	r=2
	r≤2	0.201	r=3	0.959	
Hungary	r=0	0.001	r=1	**0.689**	
	r≤1	0.027	r=2	0.956	r=1
	r≤2	**0.767**	r=3	0.960	
Latvia	r=0	0.005	r=1	**0.654**	
	r≤1	**0.066**	r=2	0.847	r=1
	r≤2	0.077	r=3	1.005	
Lithuania	r=0	**0.278**	r=1	**0.844**	
	r≤1	0.497	r=2	0.919	r=1
	r≤2	0.177	r=3	0.971	
Poland	r=0	0.022	r=1	**0.720**	
	r≤1	**0.496**	r=2	0.973	r=1
	r≤2	0.916	r=3	0.985	
Romania	r=0	0.017	r=1	0.436	
	r≤1	**0.535**	r=2	**0.735**	r=2
	r≤2	0.890	r=3	0.822	
Slovak Republic	r=0	0.001	r=1	0.326	
	r≤1	0.018	r=2	**0.650**	r=2
	r≤2	**0.159**	r=3	0.954	
Slovenia	r=0	**0.665**	r=1	0.705	
	r≤1	0.692	r=2	**0.728**	r=1
	r≤2	0.933	r=3	0.997	

Note: the rank choice suggested by each method is indicated in boldface.

In the case of Hungary only one cointegration vector was found; w_P and w_N are weakly exogenous and w_T is adjusting to a cointegration relation with non-traded sector wages. In other terms, w_P is long-run excludable and therefore public wages have been following a separate pattern. The wage mark-up model is therefore a good approximation of inter-sectoral wage spillovers in Hungary.
Latvia has one cointegration relation with the Traded sector adjusting, while the Public and Non-traded sector are weakly exogenous. Moreover, public sector wages are long-run excludable but shocks to w_P are the only ones that significantly affected the three variables in the system in the long run.
Lithuania presents one cointegration relation; w_P and w_N are weakly exogenous and w_T is adjusting. Moreover, from the common trends representation we can see that shocks to N had a significant long-run effect on that sector and T, while shocks to the P significantly affected only public sector wages in the long run.
Poland shows one cointegration relation; wages in T and P are weakly exogenous but only shocks to w_T significantly affect w_N in the long run.
Romania presents two cointegration vectors with wages P being weakly exogenous. Full wage adaptability was rejected.
The Slovak Republic is the only country for which all hypotheses in the Scandinavian model and were found to hold. There are two cointegration vectors, full wage adaptability and w_T is weakly exogenous and significantly affects the other sectors in the long run.

Table 3. Cointegration Vectors and wage adaptability.

Country	# of Coint. Vectors	Restriction/Normalization	LR-Statistic	p-value
Bulgaria	1	$\hat{\beta}' = [\beta_{1T} \quad \beta_{1N} \quad 1]'$	N.A.	N.A.
Czech Republic	1	$\hat{\beta}' = [1 \quad 0 \quad \beta_{1P}]'$	5.874	[0.118]
		$\hat{\beta}' = [1 \quad 0 \quad -1]'$	16.409	[0.003]
Estonia	2	$\hat{\beta}' = \begin{bmatrix} 1 & -1 & 0 \\ 0 & 1 & -1 \end{bmatrix}'$	19.259	[0.000]
		$\hat{\beta}' = \begin{bmatrix} \beta_{1T} & 1 & 0 \\ 0 & \beta_{2N} & 1 \end{bmatrix}'$	N.A.	N.A.
Hungary	1	$\hat{\beta}' = [1 \quad -1 \quad 0]'$	12.897	[0.005]
		$\hat{\beta}' = [\beta_{1T} \quad 1 \quad 0]'$	2.896	[0.089]
Latvia	1	$\hat{\beta} = [1 \quad -1 \quad 0]$	13.513	[0.001]
		$\hat{\beta}' = [1 \quad \beta_{1N} \quad 0]'$	0.243	[0.622]
Lithuania	1	$\hat{\beta}' = [1 \quad \beta_{1N} \quad \beta_{1P}]$	N.A.	N.A.
Poland	1	$\hat{\beta}' = [\hat{\beta}_{1T} \quad 1 \quad \hat{\beta}_{1P}]'$	N.A.	N.A.
Romania	2	$\hat{\beta}' = \begin{bmatrix} 0 & 1 & -1 \\ 1 & 0 & -1 \end{bmatrix}'$	32.071	[0.000]
		$\hat{\beta}' = \begin{bmatrix} 0 & 1 & \beta_{1P} \\ 1 & 0 & \beta_{1P} \end{bmatrix}'$	4.465	[0.347]
Slovak Republic	2	$\hat{\beta}' = \begin{bmatrix} -1 & 1 & 0 \\ 0 & -1 & 1 \end{bmatrix}'$	7.538	[0.110]
Slovenia	1	$\hat{\beta}' = [\beta_{1T} \quad 1-\beta_{1T} \quad 1]'$	1.811	[0.612]

Note: the vectors are normalized on the variable which was found to be significantly adjusting. N.A. = not applicable.

Table 4. Weak exogeneity.

Country	Industry	Services	Public	Joint weak exogeneity test
Bulgaria	0.014 [0.906]	0.489 [0.484]	38.788 [0.000]	w_T and w_N : 1.166 [0.558]
Czech Republic	1.510 [0.219]	0.339 [0.560]	0.159 [0.690]	w_N and w_P: 3.539 [0.170] w_N and w_T: 3.116 [0.213] w_T and w_P: 12.469 [0.002]
Estonia	5.991 [0.030]	7.306 [0.026]	13.394 [0.001]	
Hungary	6.056 [0.014]	0.612 [0.434]	1.765 [0.184]	w_N and w_P: 0.711 [0.339]
Latvia	2.606 [0.106]	0.006 [0.939]	3.119 [0.077]	w_T and w_N: 12.194 [0.002] w_T and w_P: 3.359 [0.186] w_N and w_P: 1.745 [0.418]
Lithuania	2.524 [0.112]	0.388 [0.533]	2.038 [0.153]	w_T and w_N: 7.012 [0.030] w_T and w_P: 4.471 [0.107] w_P and w_N: 2.862 [0.239]
Poland	0.068 [0.795]	12.877 [0.000]	2.937 [0.087]	w_T and w_P: 4.712 [0.095]
Romania	30.275 [0.000]	15.914 [0.000]	3.439 [0.179]	
Slovak Republic	2.664 [0.264]	6.451 [0.040]	12.555 [0.002]	
Slovenia	1.367 [0.242]	0.066 [0.798]	4.192 [0.041]	w_T and w_N: 3.946 [0.139]

LR tests, distributed as χ^2_r , where r = number of cointegrating relations. P-values in brackets.

Table 5. Common Trends representation: estimated elements of the *C* matrix

Country	Equation	Shocks			
		ε_I	ε_S	ε_P	Det. Trend
Bulgaria	w_{Tt}	**1.124 [2.717]**	0.370 [0.697]	0.000 [N.A.]	0.009
	w_{Nt}	0.066 [0.155]	**1.360 [2.504]**	0.000 [N.A.]	0.013
	w_{Pt}	-0.456 [-1.068]	**1.750 [3.191]**	0.000 [N.A.]	0.014
Czech Republic	w_{Tt}	0.000 [N.A.]	0.048 [0.337]	**0.645 [2.073]**	0.012
	w_{Nt}	0.000 [N.A.]	**0.894 [4.060]**	0.251 [0.517]	0.011
	w_{Pt}	0.000 [N.A.]	0.099 [0.337]	**1.346 [2.073]**	0.011
Estonia	w_{Tt}	**1.316 [1.874]**	0.000 [N.A.]	0.000 [N.A.]	0.014
	w_{Nt}	**1.316 [1.874]**	0.000 [N.A.]	0.000 [N.A.]	0.011
	w_{Pt}	**1.316 [1.874]**	0.000 [N.A.]	0.000 [N.A.]	0.013
Hungary	w_{Tt}	0.000 [N.A.]	**1.557 [2.314]**	-0.000 [-0.004]	0.011
	w_{Nt}	0.000 [N.A.]	**1.750 [2.314]**	-0.000 [-0.004]	0.012
	w_{Pt}	0.000 [N.A.]	2.055 [1.272]	**0.783 [5.636]**	0.010
Latvia	w_{Tt}	0.000 [N.A.]	0.378 [0.920]	**0.412 [2.408]**	0.012
	w_{Nt}	0.000 [N.A.]	0.445 [0.920]	**0.484 [2.408]**	0.014
	w_{Pt}	0.000 [N.A.]	-1.202 [-0.980]	**1.961 [3.840]**	0.005
Lithuania	w_{Tt}	0.000 [N.A.]	**1.881 [2.054]**	-0.467 [-1.848]	0.005
	w_{Nt}	0.000 [N.A.]	**1.915 [2.083]**	-0.261 [-1.029]	0.005
	w_{Pt}	0.000 [N.A.]	1.080 [1.274]	**0.447 [1.909]**	0.004
Poland	w_{Tt}	**0.827 [2.419]**	0.000 [N.A.]	-0.049 [-0.179]	0.008
	w_{Nt}	**0.961 [2.623]**	0.000 [N.A.]	-0.305 [-1.043]	0.007
	w_{Pt}	0.009 [0.032]	0.000 [N.A.]	**1.025 [4.801]**	0.010
Romania	w_{Tt}	0.000 [N.A.]	0.000 [N.A.]	**0.304 [2.489]**	0.022
	w_{Nt}	0.000 [N.A.]	0.000 [N.A.]	**0.505 [2.489]**	0.021
	w_{Pt}	0.000 [N.A.]	0.000 [N.A.]	**0.860 [2.489]**	0.036
Slovak Republic	w_{Tt}	**0.810 [3.086]**	0.000 [N.A.]	0.000 [N.A.]	0.008
	w_{Nt}	**0.810 [3.086]**	0.000 [N.A.]	0.000 [N.A.]	0.008
	w_{Pt}	**0.810 [3.086]**	0.000 [N.A.]	0.000 [N.A.]	0.008
Slovenia	w_{Tt}	**0.431 [2.067]**	0.299 [0.777]	0.000 [N.A.]	0.008
	w_{Nt}	-0.199 [-0.963]	**0.938 [2.463]**	0.000 [N.A.]	0.005
	w_{Pt}	**-0.760 [-2.943]**	**1.508 [3.164]**	0.000 [N.A.]	0.003

T-values in parenthesis. Significant long-run effects of the shocks on the dependent variable are given in boldface.

Finally, Slovenia has one cointegration relation, where the public sector wages adjust to a linear combination of Industry and the Services, and the latter two sectors being jointly weakly exogenous and shocks to these variables significantly affecting public sector wages.

In order to test for wage leadership, we follow Tagtström [2000] and Friberg [2007] and employ Granger non-causality tests on our VECM model (4.4). Testing for wage leadership in (4.4) with two cointegration vectors therefore implies the following null hypotheses:

$$\alpha_{1T}\beta_{1P} = \alpha_{2T}\beta_{2P} = \Gamma_{iT} = 0 \ : \ w_P \overset{not}{\longrightarrow} w_T \tag{5.1}$$

$$\alpha_{1T}\beta_{1N} = \alpha_{2T}\beta_{2N} = \Gamma_{iT} = 0 \ : \ w_N \overset{not}{\longrightarrow} w_T \tag{5.2}$$

$$\alpha_{1N}\beta_{1P} = \alpha_{2N}\beta_{2P} = \Gamma_{iN} = 0 \ : \ w_P \overset{not}{\longrightarrow} w_N \tag{5.3}$$

$$\alpha_{1N}\beta_{1T} = \alpha_{2N}\beta_{2T} = \Gamma_{iN} = 0 \ : \ w_T \overset{not}{\longrightarrow} w_N \tag{5.4}$$

$$\alpha_{1P}\beta_{1T} = \alpha_{2P}\beta_{2T} = \Gamma_{iP} = 0 \ : \ w_T \overset{not}{\longrightarrow} w_P \tag{5.5}$$

$$\alpha_{1P}\beta_{1N} = \alpha_{2P}\beta_{2N} = \Gamma_{iP} = 0 \ : \ w_N \overset{not}{\longrightarrow} w_P \tag{5.6}$$

Table 6. Granger Causality test results.

	Null Hypothesis	Wald test (p-value)		Null Hypothesis	Wald test (p-value)
Bulgaria			Lithuania		
	$w_N \overset{not}{\longrightarrow} w_T$	3.31 [0.050]		$w_N \overset{not}{\longrightarrow} w_T$	0.77 [0.520]
	$w_P \overset{not}{\longrightarrow} w_T$	4.27 [0.024]		$w_P \overset{not}{\longrightarrow} w_T$	1.40 [0.264]
Czech Republic			Poland		
	$w_N \overset{not}{\longrightarrow} w_T$	3.54 [0.068]		$w_N \overset{not}{\longrightarrow} w_T$	1.42 [0.257]
	$w_P \overset{not}{\longrightarrow} w_T$	4.63 [0.039]		$w_P \overset{not}{\longrightarrow} w_T$	0.77 [0.469]
Estonia			Romania		
	$w_N \overset{not}{\longrightarrow} w_T$	3.67 [0.037]		$w_N \overset{not}{\longrightarrow} w_T$	2.45 [0.103]
	$w_P \overset{not}{\longrightarrow} w_T$	2.02 [0.150]		$w_P \overset{not}{\longrightarrow} w_T$	5.19 [0.011]
Hungary			Slovak Republic		
	$w_N \overset{not}{\longrightarrow} w_T$	1.09 [0.348]		$w_N \overset{not}{\longrightarrow} w_T$	0.90 [0.349]
	$w_P \overset{not}{\longrightarrow} w_T$	2.47 [0.100]		$w_P \overset{not}{\longrightarrow} w_T$	0.05 [0.830]
Latvia			Slovenia		
	$w_N \overset{not}{\longrightarrow} w_T$	1.57 [0.225]		$w_N \overset{not}{\longrightarrow} w_T$	2.04 [0.162]
	$w_P \overset{not}{\longrightarrow} w_T$	5.61 [0.008]		$w_P \overset{not}{\longrightarrow} w_T$	2.96 [0.094]

Table 7. Models of wage spillovers

	Bulgaria	Czech Rep.	Estonia	Hungary	Latvia	Lithuania	Poland	Romania	Slovak Rep.	Slovenia
Model	1	3	1	2	3	2	1	3	1	1

Note: 1. Scandinavian Model; 2. Wage mark-up model; 3. Envy-effect model.

Where Γ_{ij}, i = 1, ... ,n ; j = P,N,T is the row of the Γ matrix (the matrix of coefficients of the lagged differences) corresponding to sector j for lag i. In other words, if (5.1) and (5.2) cannot be rejected, then the traded sector is wage leader and the Scandinavian model holds. A similar reasoning can be done for the non-traded and public sector.

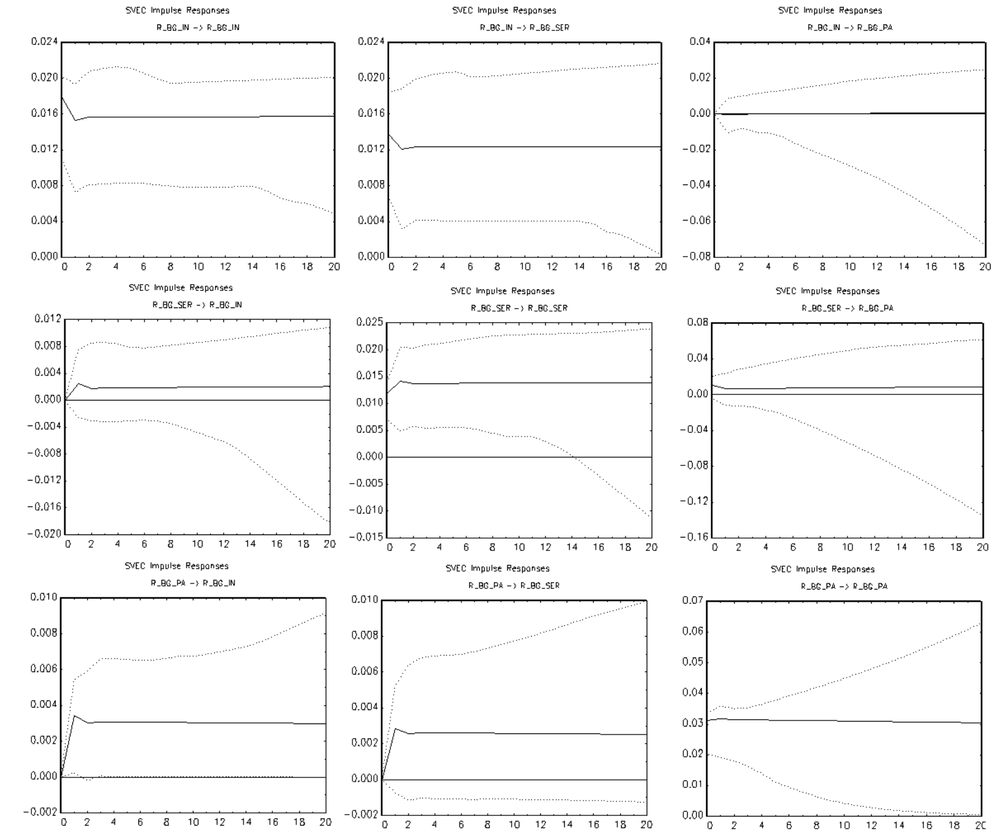
We have ten countries here, and thus the number of tests to be performed is quite high; however, since we are mainly interested in discussing the role of the Balassa-Samuelson assumption of traded sector wage leadership in the process of convergence as well as for brevity, we only report the results of (5.1) and (5.2)[24]. The leadership of the Traded sector wages vis à vis the Public sector (i.e. null hypotheses (5.1)) is rejected at 5% significance in the case of Bulgaria, Czech Republic, Latvia and Romania and at 10% for Slovenia. The null (5.2) is rejected at 5% for Estonia and Bulgaria, and 10% for Hungary. Therefore, in Bulgaria and Estonia, while w_T was found to be weakly exogenous, the traded sector was not wage leader, meaning that w_P and w_N affected it at least in the short run. In other words, there was bidirectional (Granger) causality.

In figure 3, I plotted the Impulse Response functions for each country. The graphs of the responses up to 20 quarters ahead confirm the results reported in Table 5; moreover, we notice that, in general, the adjustment to the new long-run value after a permanent shock is pretty quick, taking in general at most one year and a half. In the case of Poland, Bulgaria and Estonia the adjustment requires less than one year. This is interesting since, if most of the adjustment occurs within the year, previous works on wage spillovers that use yearly data (for example, Lamo et al. [2008]) might not be able to capture these feedback effects.

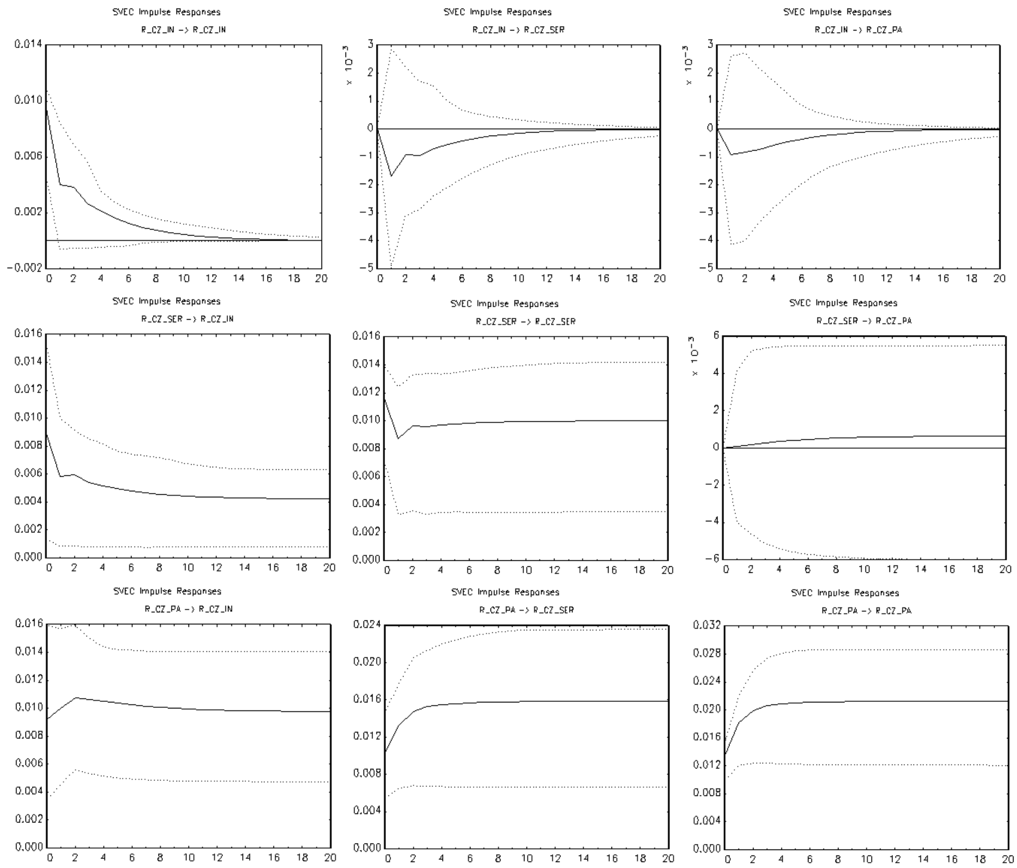
[24] Test results on (5.3) – (5.6) are available on request.

Figure 3. Impulse Response Analysis

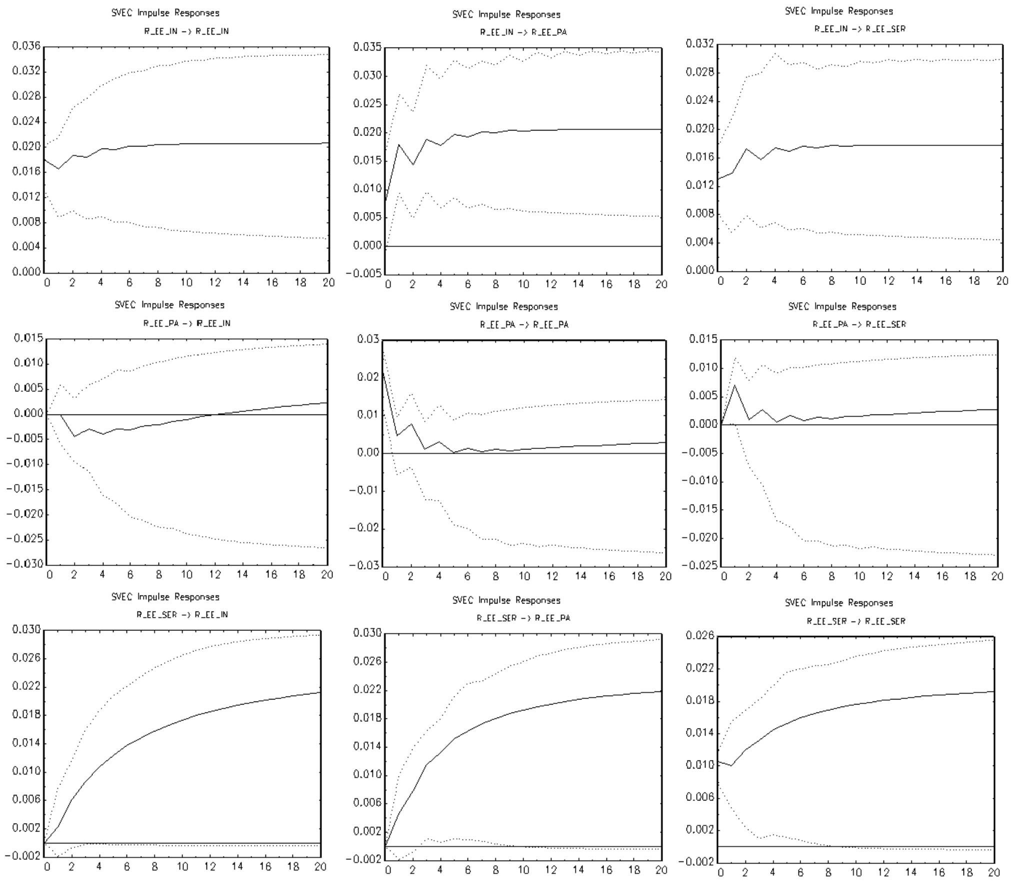
Bulgaria



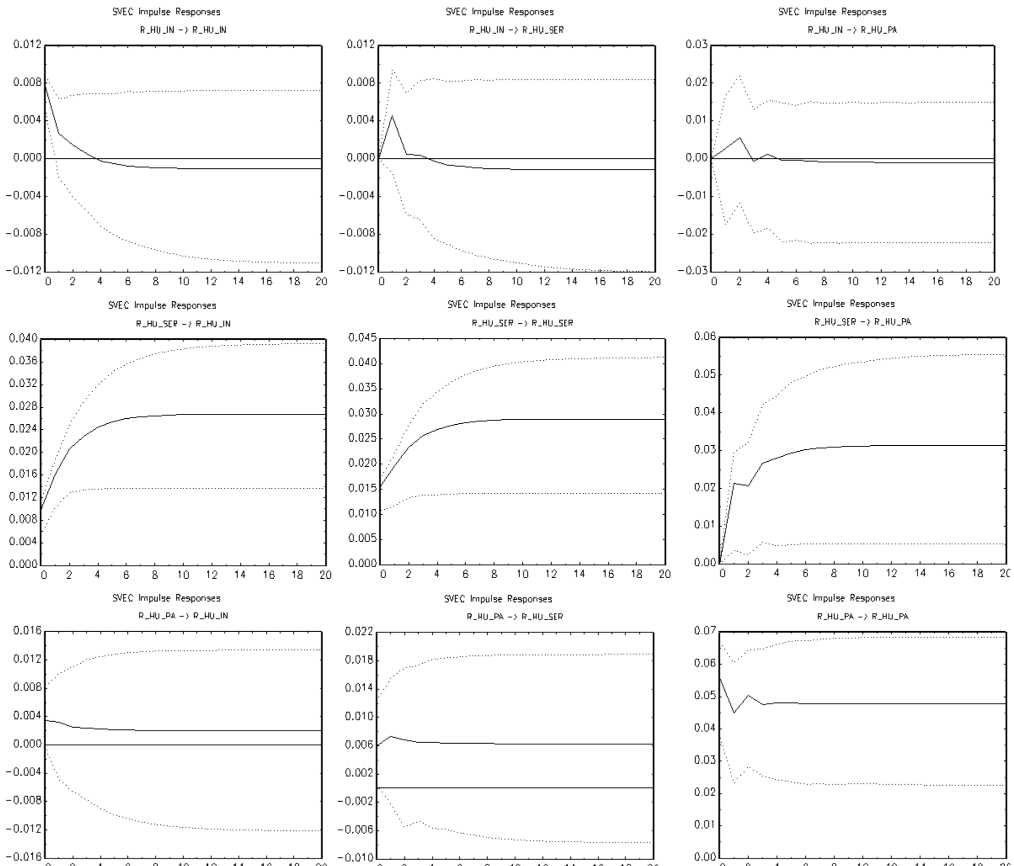
Czech Republic



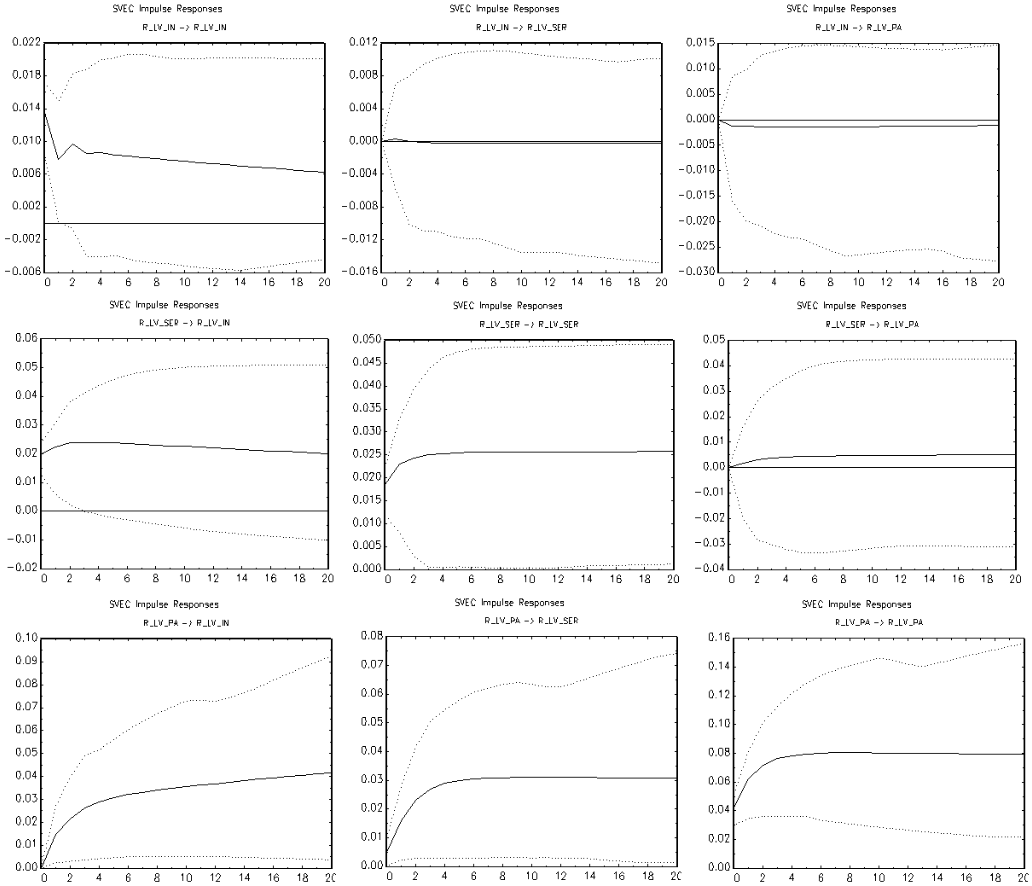
Estonia



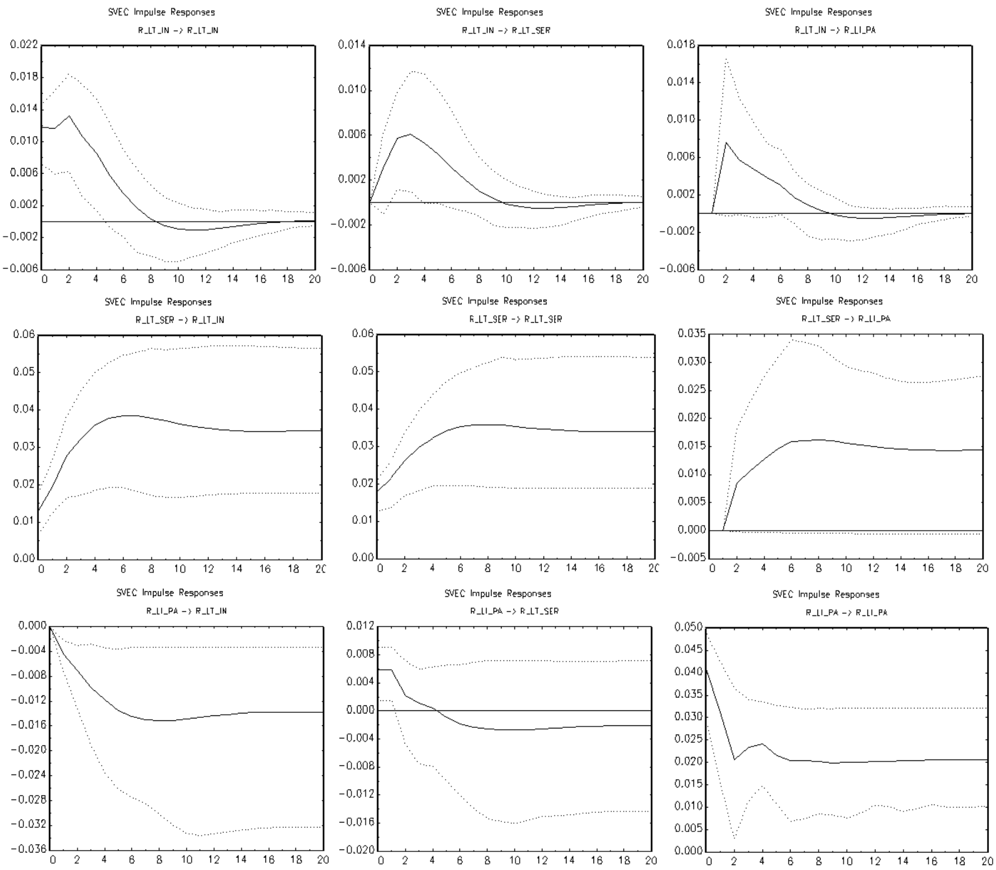
Hungary



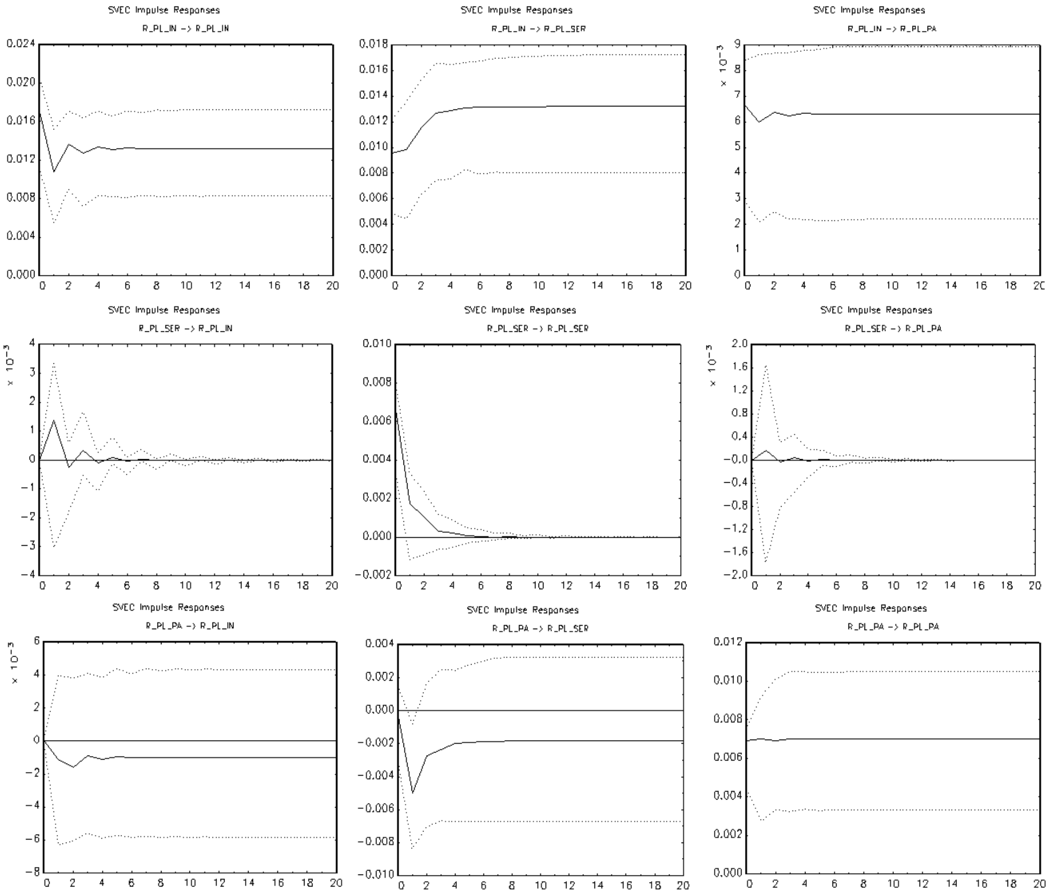
Latvia



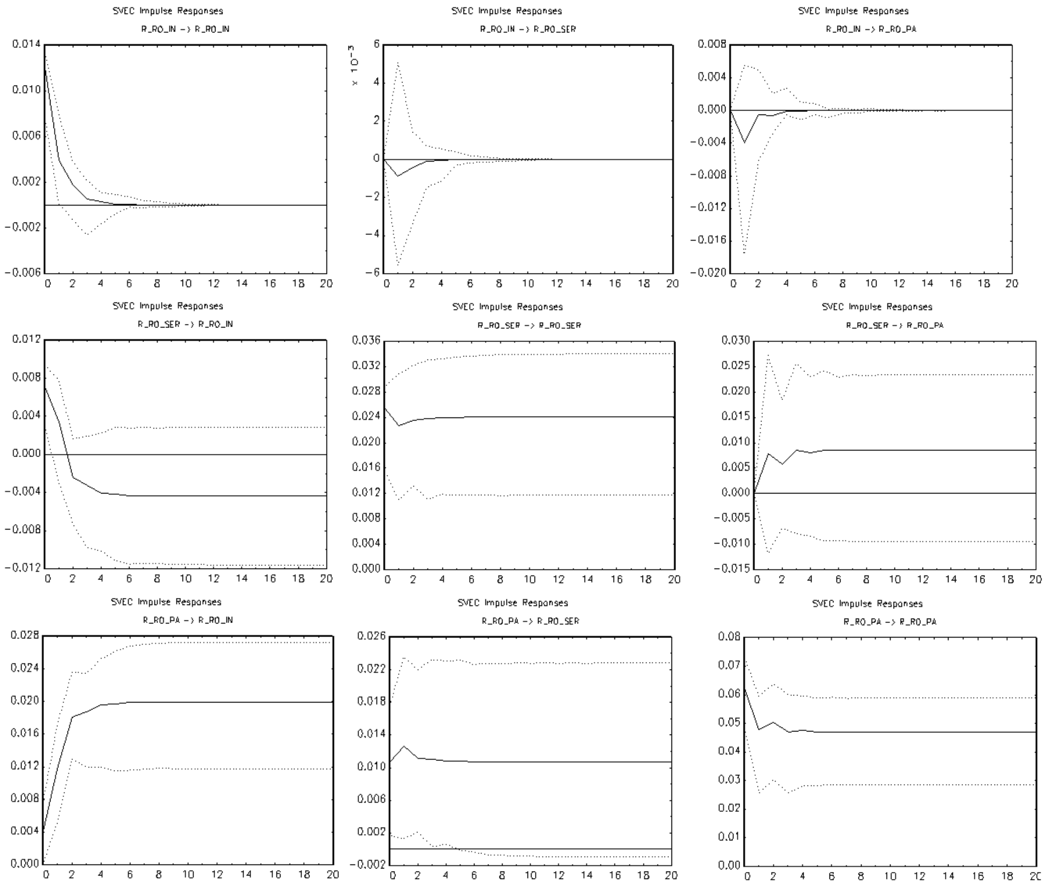
Lithuania



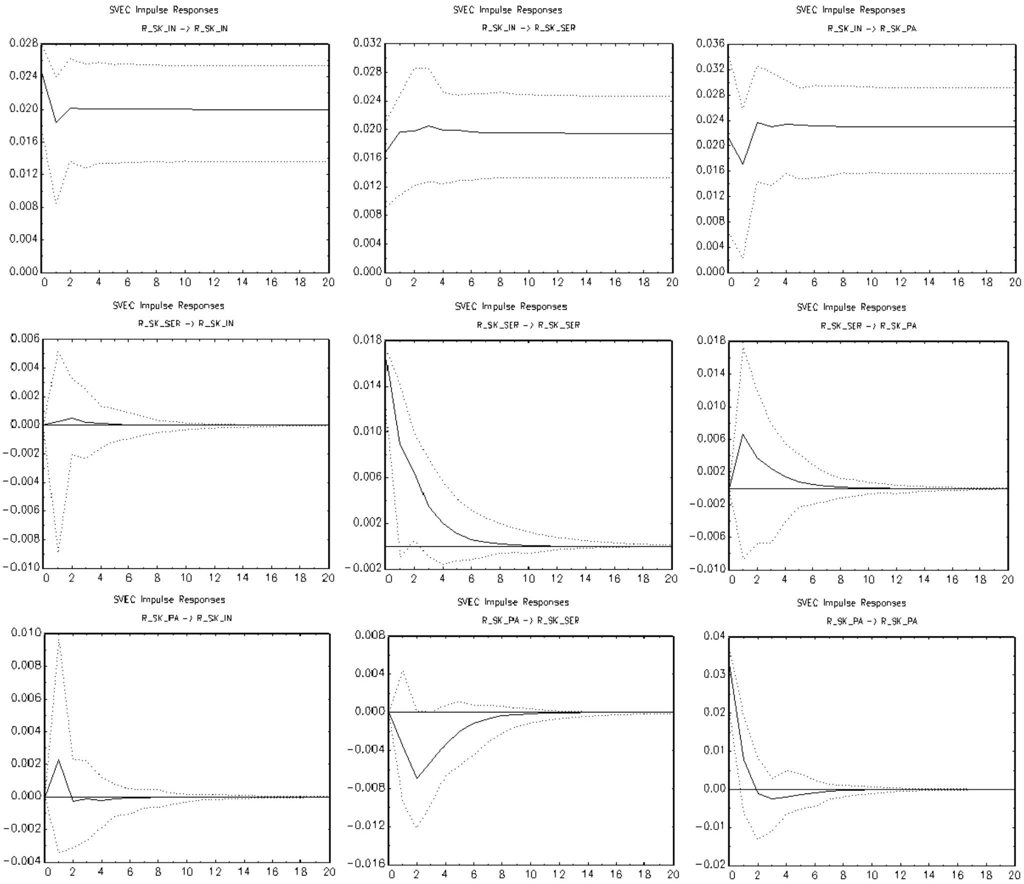
Poland



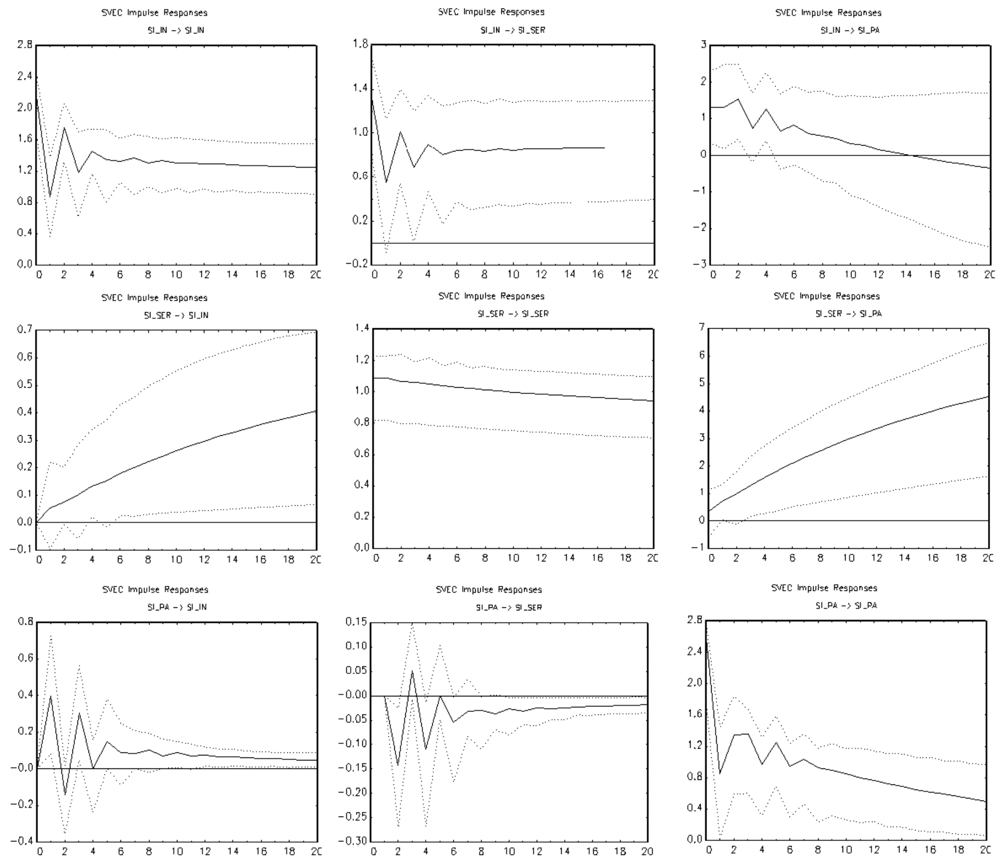
Romania



Slovak Republic



Slovenia



6. Conclusions

This paper analyzed spillovers in wage determination across macro sectors in ten european transition countries that joined the European Union within the last decade. Several previous studies have shown, either with a descriptive approach or using econometric techniques, that wages in different sectors generally equalize or, more generally, that the wage ratio tends to be constant.

However, as it was discussed in Section 3, different frameworks of leads and lags in wage determination may have different impacts as regards the international cost competitiveness of a country's firms.

When wages in sectors that are not exposed to international competition lead the process of wage setting, inter-sectoral labor mobility as well as union pressure may cause the traded sector wages to grow more than productivity, thus harming competitiveness. In this case, the catching up process will be characterized by wages converging faster than productivity.

The main results of the paper are the following: first of all, in transition countries (with the exception of Slovakia) we do not find support for full wage adaptability, and in most cases two common trends are present, meaning that shocks to wages in the sectors that are found to be leaders in wage determination are not fully transmitted to the other sectors. Thus, pushes to wage equalization coming from inter-sectoral labor mobility might have been low. Moreover, the fact that we have two common trends may imply separate bargaining processes. Second, the "right" model to describe wage interactions is different across countries. In Section 3, we have outlined three alternative models of wage spillovers and derived the hypotheses they implied. Table 7 shows which model best describes wage setting in each of the sample countries; in general, we found support for a weaker form of those models, with the exception of the Slovak Republic, for which the strict version of the Scandinavian Model holds. In Romania, the Czech Republic and Latvia public sector wages were found to be weakly exogenous, "causing" movements in traded and non-traded sector wages. In the case of Hungary and Lithuania, the non-traded sector has been leading wage determination. Thus, for these five countries, the process of convergence may be accompanied by the creation of large international imbalances due to competitiveness loss.

As far as the remaining countries are concerned, while traded sector wages were weakly exogenous, with the exception of Estonia we found shocks to wages in N to significantly affect public sector wages in the long run. Thus, we could find what we called a traded sector wage shock and a non-traded sector wage shock. In the case of Estonia, Bulgaria and (only at 10% significance) Slovenia, while model 1 appears to hold, the traded sector was not wage leader and actually we could observe bidirectional causality.

Finally, Granger "causality" tests and Impulse Response analysis have shown that, in some cases, even when the traded sector is weakly exogenous, there is bidirectional causality (Bulgaria and Estonia); adjustment after a shock to wages in the leading sector occurs quite quickly, taking in general less than six quarters.

As it was stated in the introduction, in principle, when wage setting in sectors that are not subject to international competition leads that in the traded sector, there is room for competitiveness loss since wage costs will grow faster than productivity. At a micro level, firms in the export sector will either have to increase prices or, to avoid losing market share, reduce profits. At a country level, excess inflation may be balanced by a depreciation of the nominal exchange rate to re-establish (relative) purchasing power parity. When the nominal exchange rate is fixed, however, due to an exchange rate arrangement or entry in a monetary union, nominal depreciation is not an option. For this reason, in principle, empirical support for the wage mark-up or the envy-effect model in central and eastern European countries, as it was found in the present paper, may suggest the potential for accumulating large current account imbalances when these countries will finally have to adopt the euro, and this may be the subject of further research.

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Appendix 1. Data Sources

Variable	**Source and Definition**
Industry Wages	Eurostat, LCI (Labor Cost Index) – Wages and Salaries. Seasonally adjusted and adjusted data by working days – B-E ("Industry except construction"). Nominal value.
Services Wages	Eurostat, LCI (Labor Cost Index) – Wages and Salaries. Seasonally adjusted and adjusted data by working days – G-N ("Services of the business economy"). Nominal value.
Public Wages	Eurostat, LCI (Labor Cost Index) – Wages and Salaries. Seasonally adjusted and adjusted data by working days – O ("Public administration and Defence, compulsory social security"). Nominal value.
Price Adjustment	Eurostat, HICP Overall Index – Seasonally and Working day adjusted

Appendix 2. Specification of the empirical model by country.

As it was outlined in section 4, che CVAR can be specified with a constant and a linear trend which can be restricted to lie in the cointegration space or left unrestricted.
In the former case, a restricted constant accounts for a mean of the cointegration relation which is different from zero; a restricted trend is included whenever some of the variables in the system exhibit a linear trend while others do not or have a trend with a different slope, so that possibly the cointegration relation would be trending as well. In the present empirical model, if the hypothesis of wage equalization holds, a trend should not be present in the long-run relations. However, as it is clear from table A2, we had to include a linear trend in the cointegration relation in some countries and for some specific time periods to account for cases when wages in some sectors were indeed growing at higher rates than others. While it is not realistic that a trend is present over longer periods, we think this is still a good local approximation, while excluding cointegration because of the presence of a trend would be a potentially important information loss.

Table A2. VECM specification.

Country	n	Model Specification
Bulgaria	2	Unrestricted constant, no trend, transitory dummies (2001Q4, 2009Q4)
Czech Rep.	1	Unrestricted Constant, restricted trend, shift in 2004Q3.
Estonia	2	Unrestricted constant, no trend
Hungary	2	Unrestricted constant, no trend
Latvia	2	Unrestricted constant, no trend, shift at 2006Q4.
Lithuania	3	Unrestricted constant, no trend
Poland	2	Unrestricted constant, no trend, blip dummy D0402P restricted in the CV.
Romania	2	Unrestricted constant, restricted trend. Trend break at 2007Q1, which might account for entrance in the E.U.
Slovak Rep.	1	Unrestricted constant, no trend
Slovenia	1	Unrestricted constant, no trend