

How flexible are wages in EU accession countries?*

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Abstract

The transition to a market economy and increased economic integration have fostered regional disparities in Central and Eastern European countries. This paper investigates whether and to what extent wages could act as an equilibrating mechanism in these countries by adjusting to local market conditions. Using regional data for the 1990s, we estimate static and dynamic wage curve models for Bulgaria, Hungary, Poland and Romania. We find empirical evidence indicating that regional average earnings adjusted to local unemployment rates in Bulgaria, Hungary and Poland. This result suggests that in these countries wages could help equilibrate labour markets following demand shocks. In the case of Romania, the unemployment elasticity of pay was not significantly different from zero.

Keywords: Wage flexibility, Panel data, EU accession countries

JEL classification: C23, J30, J60

* This paper uses a data set generated in the framework of the project “Regional Labour Market Adjustment in EU Accession Candidate Countries”. The data are available upon request from the authors.

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1 Introduction

The transition to a market economy in Central and Eastern European countries (CEECs) and increasing integration with Western economies have resulted in significant labour demand changes across both sectors and regions, leading to rising unemployment and falling employment and participation rates (EBRD, 2000). Furthermore, there is growing evidence of a strong regional dimension of the restructuring process, with regional disparities increasing in most CEECs (Boeri and Scarpetta, 1996, Petrakos, 1996 and 2000). In particular, there are increasing regional differentials in labour market performance, which raises the question about possible equilibrating mechanisms such as interregional labour mobility and regional wage flexibility. That these mechanisms function well will gain importance with the upcoming accession of CEECs to the EU and later to the Economic and Monetary Union. Without flexible nominal exchange rates and with low interregional mobility (Boeri and Scarpetta, 1996, Fidrmuc, 2003), wage flexibility could play an important role in helping labour markets adjust to supply and demand shocks.

In this paper, we assess whether and to what extent wages represent an equilibrating mechanism in CEECs. In particular, we investigate the responsiveness of regional average earnings to local labour market conditions over the last decade in Bulgaria, Hungary, Poland and Romania. We contribute to the literature on labour markets in EU accession countries in two ways. First, we provide empirical evidence on the responsiveness of wages to local market conditions in the above-mentioned countries using a new data set. Second, we go beyond the standard static models and address three critical concerns raised in the related literature: potential endogeneity in the relationship between regional wages and unemployment, bias in dynamic panel models and spatial dependence in relationships across regions.

To uncover the responsiveness of wages to local labour market conditions, we follow the literature flowing from Blanchflower and Oswald (1994) in estimating wage curve models using panel data for these countries for the last decade. The wage curve model relates wage levels to local unemployment rates. Blanchflower and Oswald (1994) argue that there seems to be a high responsiveness of wages to local market conditions and, moreover, that countries have similar unemployment elasticity of pay, around -0.10 , despite their different institutions. In contrast to the macroeconomic Phillips curve that describes the aggregate relationship between changes in money wages and unemployment, the wage curve uncovers a mechanism for local labour market equilibration.

We first estimate standard static models with fixed regional and time effects, allowing for comparisons with results from existing literature. In addition, we account for potentially different effects of local unemployment rates on regional average earnings in the early and late transition periods. Given the simultaneous determination of wages and unemployment on the one hand, and the possibility of wage inertia, on the other (see for example, Büttner, 1999a, and Baltagi and Blien, 1998), the results of the static panel models are likely to be biased and inconsistent (Baltagi, 2001). Taking this fact into account, we further estimate dynamic panel models suggested by Arellano and Bond (1991).

As pointed out by a growing literature, regions are likely to be interdependent due to production and trade linkages, as well as due to technology spillovers (Fingleton, 1999, Quah, 1996). Büttner (1999a), Elhorst et al. (2002) and Longhi et al. (2002) demonstrate that neglecting spatial dependence can lead to an underestimation of the effect of local unemployment on wages. We therefore correct for spatial dependence using a filtering procedure based on Getis and Ord (1992) and Getis (1995). The results of our first estimations of static models suggest that, during the 1990s, regional average earnings adjusted

to local labour market conditions in Bulgaria and Poland. The local unemployment rate effect on regional average earnings was stronger in the early transition period (1992 -1994) in Bulgaria, while in Poland it was stronger in the late transition period (1995-1998). Furthermore, the results we obtained with dynamic specifications confirm this labour market adjustment process in the cases of Bulgaria and Poland. At -0.12 , unemployment elasticity of pay in Bulgaria was highest in our sample and close to that found in advanced economies (-0.10). In the case of Poland, the unemployment elasticity of pay was only half of the standard result, -0.05 . Spatial dependence was important in Hungary but had no effect in the other countries. In Hungary, only the dynamic specification with spatially filtered variables revealed the wage curve. In the case of Romania, we find no empirical support for a wage curve.

This paper is organised as follows. Section 2 presents previous results on wage curves in transition countries and selected Western economies. In section 3 we outline our econometric strategy and discuss the related model specifications. The obtained estimates for the responsiveness of regional average earnings to local market conditions are discussed in section 4. Section 5 concludes.

2 Stylized facts

Conventional economic theory holds that regional wages are positively related to regional unemployment rates. This result was formalised by Harris and Todaro (1970) and supported during the 1970s and 1980s by empirical evidence from both individual and aggregated regional data (Hall, 1970 and 1972; Reza, 1978; Adams, 1985; Marston, 1985).

The consensus on a positive relationship between regional wages and unemployment rates was challenged by empirical work uncovering a negative relationship between these variables

in the late 1980s and the 1990s. The new work, including contributions by Blackaby and Manning (1987), Freeman (1988), and Card (1990), uses regional data and controls for regional fixed effects. Blanchflower and Oswald (1994) called this negative relationship between regional wages and local unemployment rates a genuine “empirical law in economics”, the wage curve. They brought a considerable amount of empirical evidence from large numbers of individuals in the US, UK and other developed countries supporting not only the negative unemployment elasticity of pay but also that this elasticity is the same in all cases, around -0.10 . This result implies that a doubling of the unemployment rate reduces contemporaneous regional wages by ten percent. The 1994 publication of their book, “The Wage Curve”, generated a large amount of research on the wage curve for different countries, including developing and transition economies. We discuss next the main stylized facts coming out from existing studies on transition countries and selected Western countries for comparison.

A number of existing studies on transition countries use individual micro data and estimate standard static models with regional and time fixed effects, similar to Blanchflower and Oswald (1994). Micro data sets have the advantage of allowing the use of control variables specific to standard wage equations *à la* Mincer, such as gender, education, and experience. On the other hand, using micro data has disadvantages as well. Micro data usually exclude specific groups such as those with high earnings (see Partridge and Rickman, 1997, for the case of the US, and Büttner, 1999a, for the case of Germany). An alternative option is using aggregate regional data. In this latter case, however, changes in the composition of the labour pool and of the unemployed cannot be controlled for.

In many cases, wages are proxied by earnings. As pointed out by Card (1995), the elasticity of earnings with respect to unemployment rates is determined by the unemployment elasticity of hourly wages and the unemployment elasticity of hours worked. However, most studies do

not control for the numbers of hours worked. If the hours worked are correlated with local unemployment rates, this omission implies that the unemployment elasticity of wages is overestimated (Card, 1995, Büttner, 1999b). Failing to control for the variation of hours worked does not bias however the results if there is no evidence that hours worked are correlated with local unemployment rates (Partridge and Rickman, 1997).

A number of existing studies estimated wage curves in transition countries during the 1990s and found unemployment elasticities of pay close to the standard result of -0.10 . For example, Kertesi and Köllô (1997 and 1999) found unemployment elasticities of pay in Hungary in the range of -0.09 to -0.11 using individual micro data matched with data from 170 labour office districts. In the case of Poland, Duffy and Walsh (2001) used individual data from labour force surveys and data for 49 regions and found unemployment elasticities of pay in the range of -0.08 to -0.11 . In the case of Eastern Germany, Elhorst et al. (2002) obtained an unemployment elasticity of pay of -0.112 using individual data for 114 districts. Kállai and Traistaru (2001) use aggregate regional data from 41 regions in Romania and found an unemployment elasticity of pay of -0.09 .

Furthermore, Blanchflower (2001) estimates standard wage curves for a number of 15 transition countries, including nine EU accession countries and six successors of the former Soviet Union, using both individual micro data and aggregate regional data sets. He finds unemployment elasticities of pay ranging from -0.02 to -0.46 in regressions without fixed effects, and 0.003 to -0.52 in regressions with fixed effects. These results imply that controlling for unobserved time invariant regional characteristics gives a higher responsiveness of earnings to unemployment rates. This conclusion is supported by the findings of Kállai and Traistaru (2001) and Pannenberg and Schwarze (1998a).

For Western European countries, typically lower unemployment elasticities of pay than for transition countries have been found, ranging from -0.01 to -0.07 (see for example, Winter-Ebmer, 1996, for the case of Austria, Baltagi and Blien, 1998, Büttner, 1999a, Longhi et al., 2002, for the case of Western Germany, Jimeno and Bentolila, 1998, and Bajo et al., 1999, for the case of Spain, and Bell et al. 2002 for Great Britain). In contrast, a recent paper by Montuenga et al. (2003) finds higher unemployment elasticities of pay for the United Kingdom, France and Spain, at -0.24 , -0.29 and -0.30 , respectively.

A frequent criticism of the wage curve estimations centers on the potential endogeneity of the unemployment rates (see for example, Baltagi and Blien, 1998, Longhi et al., 2002, Jimeno and Bentolila, 1998, and Montuenga et al., 2003). To address this problem, a number of studies use lagged unemployment rates as instruments for unemployment rates (see for example Duffy and Walsh, 2001, and Pannenberg and Schwarze, 1998a). However, instrumenting unemployment rates by own lagged values yields inconsistent and biased results from panel estimators. Consequently, other authors use estimation techniques robust to the lack of strict exogeneity of unemployment rates, such as the Arellano-Bond Generalized Method of Moments (GMM) procedure (Jimeno and Bentolila, 1998) and first-differenced two-stage least squares (FD-2SLS) (Baltagi and Blien, 1998).

A growing literature also points to the need to control and correct for spatial dependence in regressions using regional data. Neglecting spatial correlations between labour market characteristics of neighbouring regions could result in biased estimates. For example, Longhi et al. (2002) address this concern and find that, in the case of Western Germany, correcting for spatial dependence gives a higher unemployment elasticity of pay in comparison to the standard estimation.

Our paper fills a gap in the literature by providing wage curve estimates for four transition countries during the 1990s correcting for potential endogeneity of unemployment rates and spatial dependence in regional data. We exploit a specially created data set for Bulgaria, Hungary, Poland and Romania that allows us to correct for both endogeneity and spatial dependence in a dynamic panel model.

3 Model specifications and estimation issues

Most existing studies on wage curves in transition countries typically estimate the unemployment elasticity of pay using a standard static panel model which includes regional and time fixed effects (the Least Square Dummy Variables, LSDV estimator). One advantage of this approach is the possibility to investigate the responsiveness of wages to local unemployment over periods corresponding to distinct phases of the transition to a market economy. However, this estimator has a number of shortcomings.

Static panel models may fail to capture characteristics specific to the relationship between wages and unemployment for a number of reasons. First, regional unemployment rates and wages may be simultaneously determined.¹ This calls for a panel estimation methodology which is robust to the endogeneity of regressors. Second, the possibility of wage inertia needs to be allowed for (see for example Büttner, 1999a), which requires a dynamic model. As Nickell (1981) and Kiviet (1995) point out, the LSDV estimator is biased and inconsistent in the case of dynamic panels (see also Baltagi, 2001). While the bias may be not too large in

¹ Blanchflower and Oswald (1994) argue that static wage curves do not suffer from simultaneity bias. However, for example, Baltagi and Blien (1998) find evidence against the strict exogeneity of unemployment rates with respect to wages in Western Germany in the 1980s.

very large samples, it is a significant problem in small samples.² Third, wages may react to unemployment with delay, or unemployment hysteresis may be present in the wage curve (see for example Jimeno and Bentolila, 1998, for the Spanish wage curve exhibiting hysteresis) which implies that lagged values of the unemployment rate variable should be included in the regression. To appropriately address these concerns, we estimate the wage curve in an autoregressive distributed lag model framework, using the estimator proposed by Arellano and Bond (1991).

Previous studies also fail to account for spatial dependence. As mentioned above, spatial dependence may arise from correlations in labour market characteristics of neighbouring regions. As pointed out by Büttner (1999a), one can distinguish between three types of spatial dependence. In the first type, unobserved regional characteristics, such as labour market accessibility, may be spatially correlated (see also Elhorst et al., 2002). A second type of spatial dependence arises from common shocks to contiguous regions, causing error autoregression. Finally, spatial dependence might exist in the dependent variable or the regressors resulting from the similarity of employment conditions in neighbouring districts. For example, Longhi et al. (2002) point out that one region's wage levels may raise because of higher alternative wages in surrounding regions. With respect to wage curves, Büttner (1999a), Longhi et al. (2002) and Elhorst et al. (2002) found that neglecting spatial effects leads to an underestimation of the unemployment elasticity of pay.

In this paper, we estimate wage curves for Bulgaria, Hungary, Romania and Poland using first a standard static fixed effects model and then a dynamic fixed effects model. We account next

² For example, Judson and Owen (1999) find that even when $T = 30$, the size of the bias could be around 20 per cent of the true value of the estimated coefficient.

for spatial dependence and re-estimate the dynamic panel model with spatially filtered variables.

In order to allow for comparability with previous studies and the assessment of the bias from neglecting the dynamic nature of the relationship between unemployment and earnings, we first estimate the following standard static fixed effects model:

$$\ln w_{r,t} = c + \beta \ln U_{r,t} + \gamma X'_{r,t} + \mu_r + \lambda_t + \varepsilon_{r,t} \quad (1)$$

where $w_{r,t}$ is the average monthly earnings in region r at time t , deflated with the national inflation index (the consumer price index), c is a constant, $U_{r,t}$ is the unemployment rate in region r at time t , $X'_{r,t}$ is a vector of variables controlling for the regional economic structure,³ μ_r is a time invariant region-specific effect, $\mu_r \sim \text{i.i.d. } N(0, \sigma^2_\mu)$, λ_t is a region-invariant time specific effect, $\lambda_t \sim \text{i.i.d. } N(0, \sigma^2_\lambda)$, and $\varepsilon_{r,t}$ is the stochastic error term, $\varepsilon_{r,t} \sim \text{i.i.d. } N(0, \sigma^2_\varepsilon)$.

The literature on transition countries highlights two phases in the transition to a market economy in CEECs. In the first years of transition (up to 1994), market institutions were put in place while the second phase of transition consists of structural reforms. One can expect that the relationship between earnings and unemployment rates was different in the first transition phase than in the second phase. To take account of this fact, we estimate separate wage curves for each transition phase in each country.

Next, we estimate a dynamic panel model with fixed effects using the GMM procedure suggested by Arellano and Bond (1991). The estimated dynamic model has the following form:

³ As elements of X we use the shares of employment in industry and services in the cases of Bulgaria, Poland and Romania, and the shares of employment in agriculture and industry in the case of Hungary (see the Appendix). In each case, the selected sectoral shares had the lowest correlation.

$$\ln w_{r,t} = c + \sum_{k=1}^K \alpha_k \ln w_{r,t-k} + \sum_{l=0}^L \beta_l \ln U_{r,t-l} + \sum_{m=0}^M \gamma_m X'_{r,t-m} + \mu_r + \lambda_t + \varepsilon_{r,t} \quad (2).$$

The Arellano-Bond GMM procedure includes the following estimation steps. The model is first-differenced in order to remove the fixed effects. The differenced equation is then estimated using instrumental variables. As instruments, for each year, all available lags of the variables in levels are used. Since these are correlated with differenced variables, but uncorrelated with differenced error terms (unless the error terms in levels display serial correlation), they provide a set of valid instruments. While first order autocorrelation in the first-differenced residuals complies with the estimator's consistency requirements, it is necessary that the differenced error terms are free of second order autocorrelation (Arellano and Bond, 1991).

We choose the most appropriate specification of the dynamic wage curve model for each country by the following procedure. We start with a model specification where each variable is included with up to its third lag among the regressors.⁴ When the third year lagged variables are not significant we start with the two years lagged specification. In order to decide whether the unemployment rate is exogenous or predetermined we use the Sargan test statistics. Then, in the chosen model, we gradually reduce the number of regressors by dropping insignificant lagged variables. For each of these models, we report the one-step GMM estimator with robust standard errors. Since the standard errors from the two-step GMM are frequently found downward biased (Arellano and Bond 1991), for inference on single variables' coefficients we rely on the one-step estimator. For the choice between specifications, however, we use the Sargan test of over-identifying restrictions after the

⁴ Due to the low number of time periods available for our data, more lags would substantially reduce the quality of statistical inference from our estimations. Therefore, we do not consider the possibility of further lags.

corresponding two-step GMM estimator.⁵ Since consistency of the estimator requires the absence of second-order autocorrelation in the differenced residuals, we consider only specifications that fulfill this criterion. This is checked by the respective tests developed by Arellano and Bond (1991).

The dynamic wage curve model with one lag corresponds to the Phillips curve. This model is often used in contrast to the wage curve with the aim to understand whether the wage equation is a relationship between unemployment and wage levels, as suggested by the wage curve, or wage changes, as suggested by the Phillips curve (see Bell et al., 2002). To assess whether a Phillips curve interpretation of the regional labour market dynamics in the four countries included in our study rather holds, we re-estimate our dynamic wage curve model with the restriction that the dependent variable enters with the first lag only (for more details, see Büttner, 1999b). The estimated model is the following:

$$\ln w_{r,t} = c + \alpha \ln w_{r,t-1} + \sum_{l=0}^L \beta_l \ln U_{r,t-l} + \sum_{m=0}^M \gamma_m X'_{r,t-m} + \mu_r + \lambda_t + \varepsilon_{r,t} \quad (3).$$

This specification includes a test of the Phillips curve nested into the wage curve model. In particular, if $\alpha=1$, one gets the familiar result that the wage change is determined by unemployment, whereas $\alpha=0$ indicates a static wage curve. Intermediate values point to the presence of both an error correction mechanism and nominal wage inertia (Pannenberg and Schwarze, 1998b).

As discussed above, the fixed effects included in the wage curve models are likely to show spatial autocorrelation due to regional interaction and the presence of spillover effects (Longhi et al, 2002). We first check for spatial autocorrelation using the Lagrange multiplier

⁵ No robust Sargan test using one-step residuals is available.

(LM) calculated on the basis of Moran's I statistics. The Moran's I coefficients are calculated as follows:

$$I = \frac{(x - \mu)' W (x - \mu)}{(x - \mu)' (x - \mu)} \quad (4),$$

where x is the variable to be checked for spatial autocorrelation, μ is its mean, and W is a row-standardized weights matrix. The elements of the weights matrix represent the inverse distances between pairs of regions' capitals (in km on public roads). The LM statistics is asymptotically χ^2 -distributed with one degree of freedom,⁶ and it is obtained as follows:

$$LM = \frac{(NI)^2}{tr(W'W + W^2)} \quad (5),$$

where N is the number of observations and W is the spatial weights matrix as described above.

As pointed out by Badinger et al. (2002), an estimation procedure for a spatial dynamic panel model incorporating spatially lagged regressors or an error process with spatial autocorrelation is not yet available. Therefore, in order to control for spatial effects, they use a two-step procedure: first, spatial autocorrelation is removed from the variables by a filter based on a spatial association measure put forward by Getis and Ord (1992) and Getis (1995). Then, the model is re-estimated with standard techniques using the filtered variables.⁷ The filtering methodology is defined as follows:

$$x_i^F = x_i \frac{\sum_j w_{ij}(\delta)}{(N-1)G_i(\delta)} \quad (6)$$

⁶ For details on this methodology, see Longhi et al. (2002).

⁷ The Getis-Ord filter is also used by Badinger and Url (2002) to estimate determinants of regional unemployment in Austria in 1991.

with

$$G_i(\delta) = \frac{\sum_j w_{ij}(\delta)x_j}{\sum_j x_j}, \quad i \neq j \quad (7),$$

where w_{ij} are elements of the spatial weights matrix W , and δ is a distance parameter indicating the extent to which further distant observations are downweighted. Following the approach of Badinger et al. (2002), we repeat the estimation procedure described above with variables from which spatial correlation is eliminated by filtering. Here, we again use the above mentioned spatial weights matrix without assigning over-proportionally decreasing importance to farther distant observations, i.e. we assume $w_{ij}(\delta)=(d_{ij})^{-\delta}$ with $\delta=1$, where d_{ij} denotes the road distance between county capitals.

4 Empirical results

In this paper we use a specially created data set including annual regional labour market data at NUTS 3 level for Bulgaria, Hungary, Poland and Romania for the period from 1992 to 1999.⁸ The data were collected from national statistical offices. More detailed information about our data set and summary statistics are given in the Appendix.

Over the last decade, unemployment rates were high, in particular in Bulgaria and Poland. Furthermore, regional disparities of unemployment rates were high and, except Romania, increasing (see Figure 1 in the Appendix). In Hungary and Poland unemployment rates were more persistent compared to Bulgaria and Romania (Figure 2 in the Appendix). In contrast,

⁸ This data set was generated in the framework of the research project "Regional Labour Market Adjustment in the Accession Candidate Countries" undertaken with financial support from the European Commission's RTD

over the same period, regional differentials of average earnings were low but more persistent than in the case of unemployment rates in all countries considered (see Figures 3 and 4 in the Appendix). We next discuss the estimation results of the model specifications explained in section 3.⁹

Table 1 shows the results of the standard static wage curve model with time and region fixed effects following the model specification described by Equation (1).

Insert Table 1 about here

We find that, over the whole period, average earnings were negatively and significantly associated with regional unemployment rates in Bulgaria and Poland as suggested by the wage curve literature. The unemployment elasticity of pay was around -0.05 in Bulgaria and -0.06 in Poland. These results are close to the findings of Blanchflower (2001) for the case of Bulgaria and Duffy and Walsh (2001) for the case of Poland. In the cases of Hungary and Romania, the coefficients of regional unemployment rates are negative but not significant, suggesting no clear pattern in the relationship between regional real earnings and unemployment rates.

As discussed above, we expect a structural difference between early and late phases of transition. Consequently, we re-estimated the static wage curve model for two sub-periods, namely, 1992-1994 and 1995-1999. The estimation results are shown in Table 2.

Insert Table 2 about here

5th Framework Programme. The data set includes: 28 regions ('oblast') in Bulgaria, 20 regions ('megye') in Hungary, 49 regions ('województwa') in Poland, and 41 regions ('judet') in Romania.

⁹ Estimations were obtained using the STATA version 7 software.

In the case of Bulgaria we find a significant and negative relationship between average real regional earnings and regional unemployment rates in the early transition period, with an unemployment elasticity of pay of -0.07 , but no clear pattern of this relationship in the later transition phase. In contrast, Poland exhibits a wage curve in the second sub-period, with an unemployment elasticity of pay similar in magnitude to that of Bulgaria, -0.07 . In the cases of Hungary and Romania, unemployment elasticities of pay were not significantly different from zero. The F statistics indicate that the hypothesis of equal coefficients for the two sub-periods can be rejected with the exception of the case of Hungary.

As mentioned above, the standard static LSDV estimator does not capture the dynamic relationship between wages and unemployment. In addition, wages and unemployment are likely to be endogenous. To address these concerns, we estimate the wage curve using a dynamic panel model as suggested by Arellano and Bond (1991). The model specification is described by Equation (2). The estimation results are shown in Table 3.

Insert Table 3 about here

The Sargan test indicates that, with the exception of Poland, unemployment rates are predetermined. We find that in Bulgaria and Poland, regional unemployment rates are negatively and significantly related to average real regional earnings. While in Bulgaria the responsiveness of earnings is high, -0.12 , and contemporaneous, in the case of Poland real earnings adjust to changes in regional unemployment rates with a one year delay, and the elasticity is lower, -0.04 .

We further calculate the long-run effect of changing regional unemployment rates on regional real earnings captured by the long-run multiplier obtained from the coefficients for the lagged regional average real earnings and unemployment rates. In the case of Bulgaria, the size of the calculated long-run multiplier is -0.18 , suggesting that, in the long run, a doubling of local

unemployment rates results in a declining of real regional earnings by 18 percent. In Poland, such a doubling would only reduce earnings by 3 percent. The long term relationship between regional unemployment rates and regional average real earnings appears positive in Hungary, with a long-run multiplier amounting to 0.06. The long-run multiplier is significant at the 5 percent level in the cases of Bulgaria and Poland and the 10 percent level in the case of Hungary.

In the next step, we compare our results obtained from the unrestricted dynamic wage curve model with the Phillips curve. Table 4 shows the results of the corresponding model specification given in Equation (3).

Insert Table 4 about here

The obtained short-term unemployment elasticities of pay in the Phillips curve specification are close to the previous results obtained with the dynamic wage curve model with the exception of Hungary. In this latter case, we find a negative and significant coefficient for the one year lagged regional unemployment rate suggesting that average real regional earnings adjust with a one year delay to a change in local market conditions. The unemployment elasticity of pay is -0.05 .

We further check for spatial dependence using the LM test given in (5). The results of the LM statistics on spatial autocorrelation are shown in Table 5.

Insert Table 5 about here

We find no evidence of spatial autocorrelation in the case of regional real earnings. In contrast, our results suggest that, with the exception of Bulgaria, regional unemployment rates were affected by spatial autocorrelation in specific years: over the period 1994-1998, in the case of Hungary; 1992-1993, in Poland; and 1992, 1995, and 1996, in the case of Romania.

The control variables were spatially autocorrelated only in Poland. Taking these results into account we apply the spatial filtering procedure explained in the previous section and re-estimate the dynamic wage curve model with the resulting spatially filtered variables.

Table 6 shows the results of the estimated dynamic model with spatially filtered variables.

Insert Table 6 about here

The estimated unemployment elasticities of pay are close to those obtained with the non-filtered variables with the exception of Hungary. After correcting for spatial dependence we find that regional real earnings are negatively and significantly related to the two year lagged local unemployment rates. A doubling of the unemployment rate results in a decline of regional real earnings by 5 percent two years later.

The calculated long run effect of unemployment rates on regional real earnings is -0.20 in Bulgaria and -0.04 in Poland and Hungary. For the latter, however, the long run effect is not significant. Consequently, we can conclude that in Hungary, earnings respond to unemployment in the short run only, while in the long run, this effect is annihilated by the dynamics of the adjustment process.

5 Conclusion

If wages are responsive to unemployment at the regional level, regional wage adjustment can allow markets characterized by low interregional migration and inflexible exchange rates to adjust to labour demand shocks. This is important in transition countries given the growing regional disparities in labour market performance and the need to adjust to potential external shocks following their accession to the European Union and later to the Economic and Monetary Union.

Previous studies found that in many transition countries regional wages seem to respond to local labour market conditions. The estimated unemployment elasticity of pay is typically close to the standard result of the literature on the wage curve, -0.10 . However, most of these studies use static estimators and do not account for potential endogeneity and spatial dependence.

Using improved econometric techniques we bring new empirical evidence about the relationship between regional wages and unemployment rates in Bulgaria, Hungary, Poland and Romania. We first estimated a standard static fixed effects model allowing for comparisons with results from existing literature. We accounted for potentially different effects of unemployment on regional earnings in the early and late transition periods. Further, we corrected for endogeneity in a dynamic wage curve model. In addition, we checked and corrected for the presence of spatial dependence in the regional variables.

The estimations of static models uncover that, over the last decade, regional average real earnings adjusted to local market conditions in Bulgaria and Poland. This adjustment was stronger in the early transition phase (1992-1994) in Bulgaria, while in Poland it was stronger in the late transition phase (1995-1998). Further, the estimations we obtained with the dynamic specifications confirm this adjustment process during the 1990s in Bulgaria and Poland. The unemployment elasticity of pay was the highest in Bulgaria, -0.12 , while in Poland it was lower, -0.04 . While in Bulgaria the regional earnings adjustment to local labour market conditions took place contemporaneously, in Poland this adjustment took place with a one year delay. The spatial effects played an important role in Hungary. The estimations obtained with spatially filtered variables uncovered a wage curve in Hungary. In this latter case, the adjustment of regional average pay to local unemployment took place with a two

years delay. In the case of Romania, we find no evidence for the adjustment of regional earnings to local labour market conditions.

Our results indicate that wages could act as an adjustment mechanism in equilibrating regional labour markets in the accession EU member states. This adjustment is likely to take place however with a certain delay, which implies that labour market disequilibria might persist.

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Appendix: Data and summary statistics

1. Data set characteristics

Bulgaria: 28 regions, data from 1991 to 1999, 252 observations

Hungary: 20 regions, data from 1992 to 1999, 160 observations

Poland: 49 regions, data from 1992 to 1998, 343 observations

Romania: 41 regions, data from 1992 to 1999, 328 observations

2. Definitions of variables

Regional average earnings: average gross monthly earnings (Bulgaria: average net monthly earnings) in national currency, 1995 prices (CPI-deflated)

Regional unemployment rate: registered unemployment rates, end-year

Regional sectoral employment shares:

- s_empl1: Bulgaria, Poland, Romania: the share of regional employment in industry (mining, manufacturing, electricity, gas, water, construction) in total regional employment; Hungary: the share of regional employment in agriculture (agriculture, hunting, forestry, fishing) in total regional employment

- s_empl2: Bulgaria, Poland, Romania: the share of regional employment in services in total regional employment; Hungary: the share of regional employment in industry (mining, manufacturing, electricity, gas, water, construction) in total regional employment

Figure 1 Regional unemployment rates: Averages and regional variation

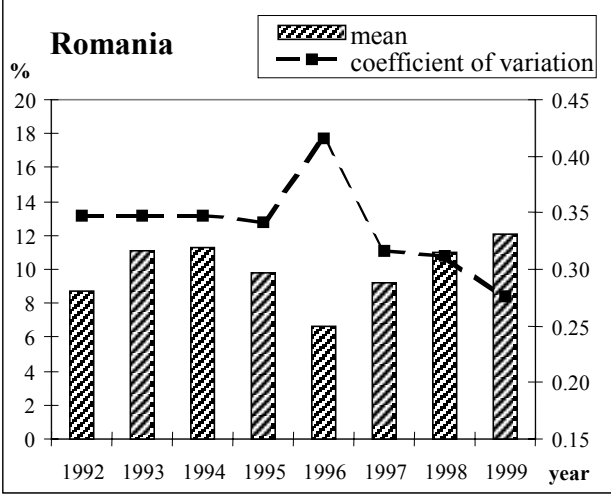
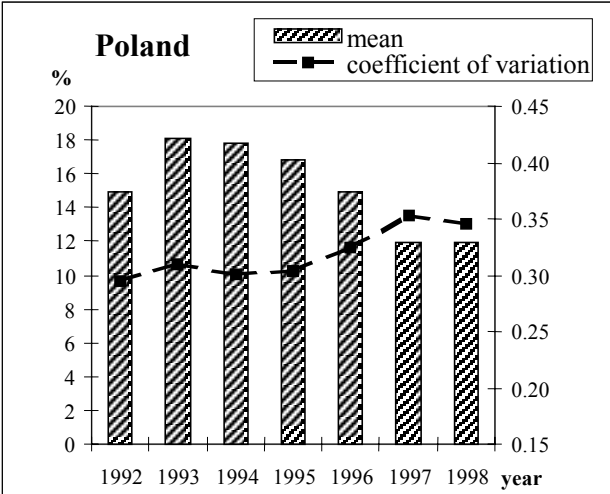
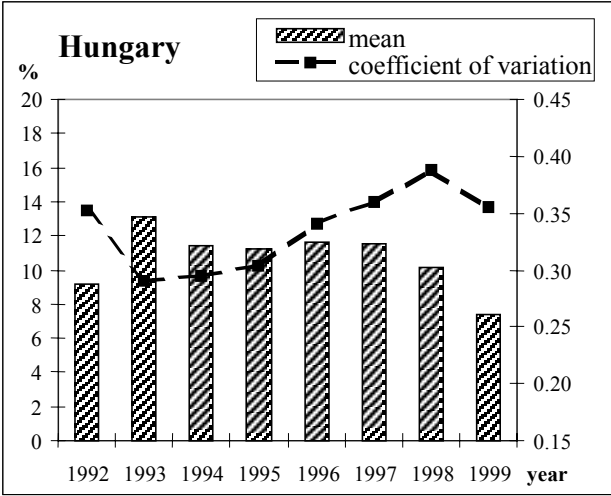
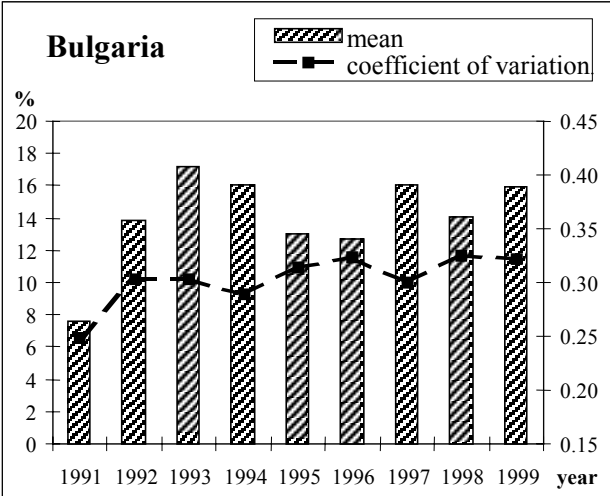


Figure 2 Regional unemployment rates: Persistence

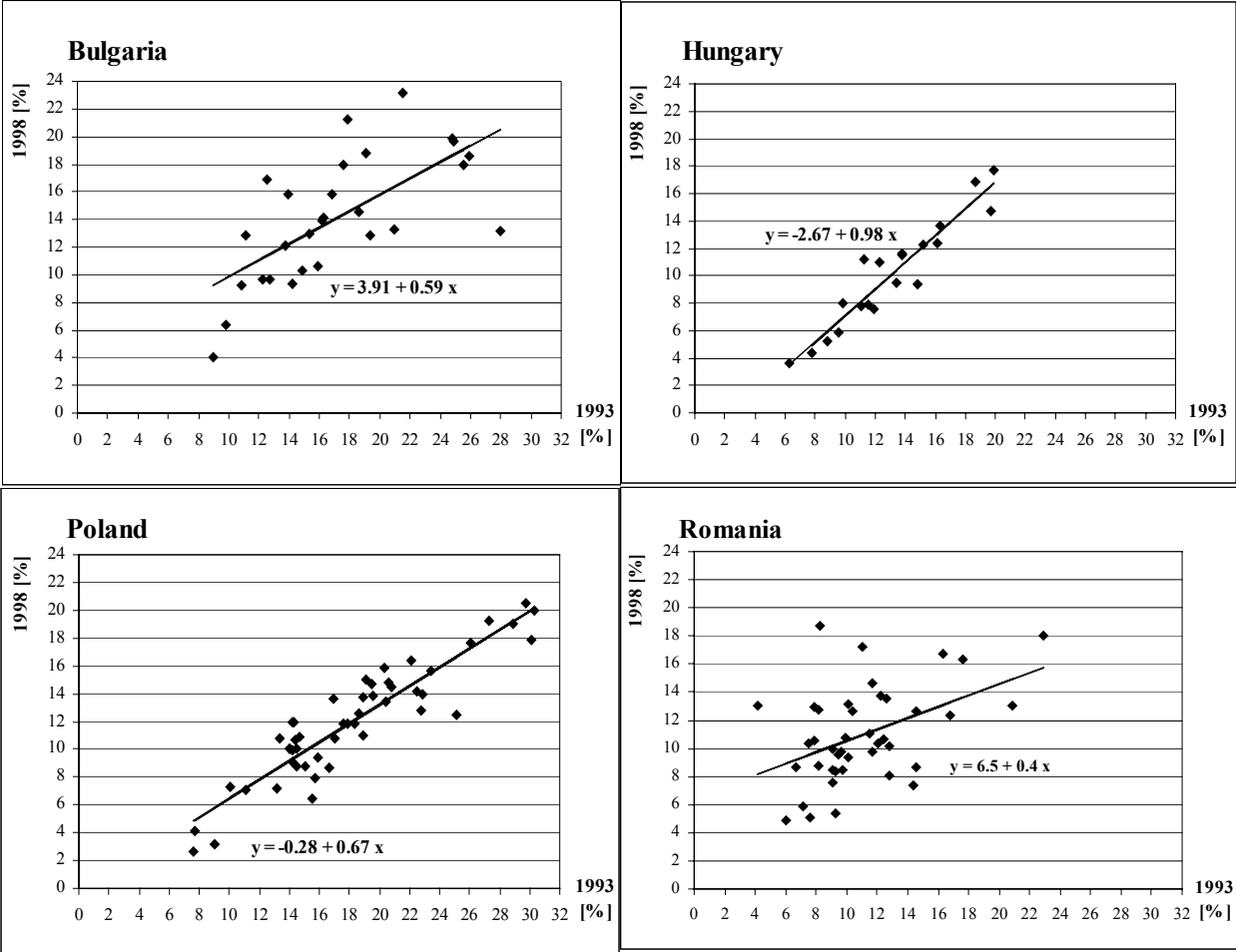


Figure 3 Regional average monthly earnings: Averages and regional variation

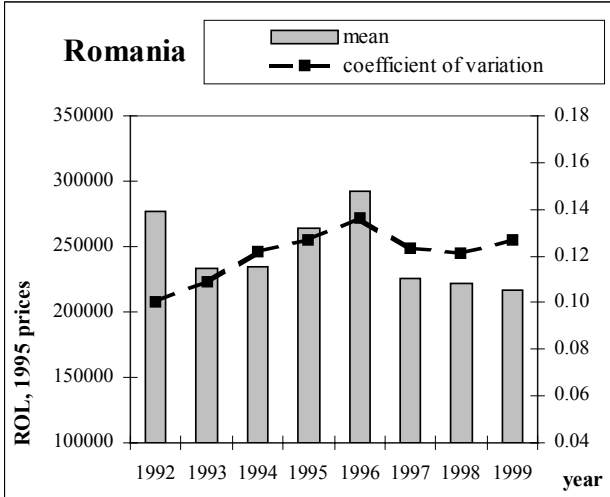
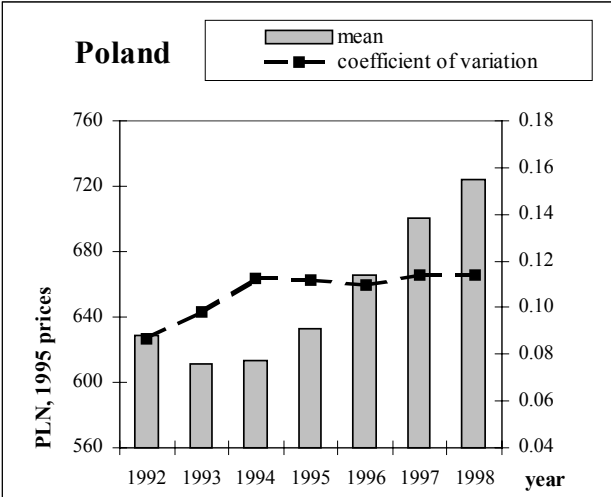
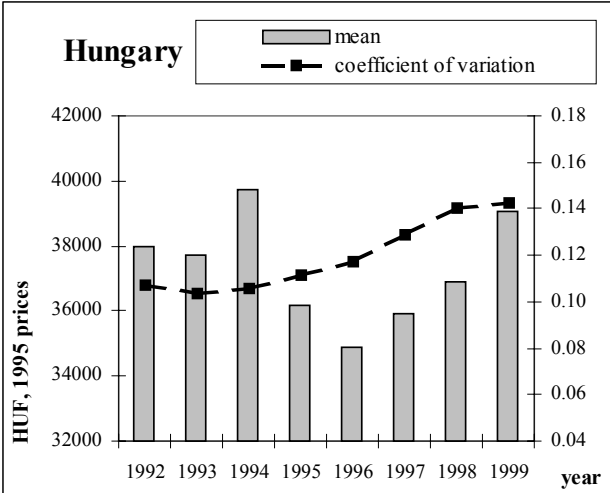
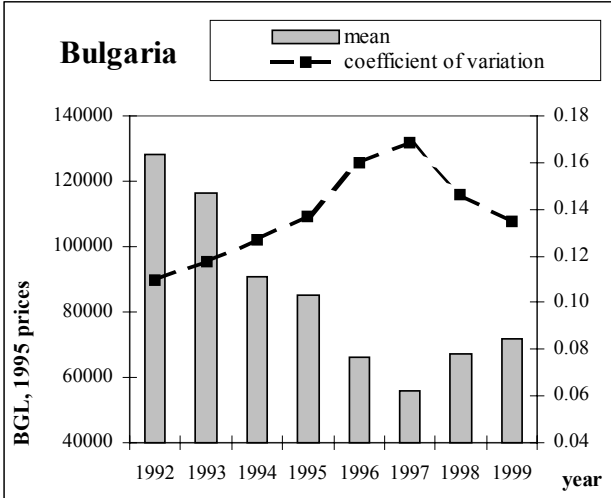
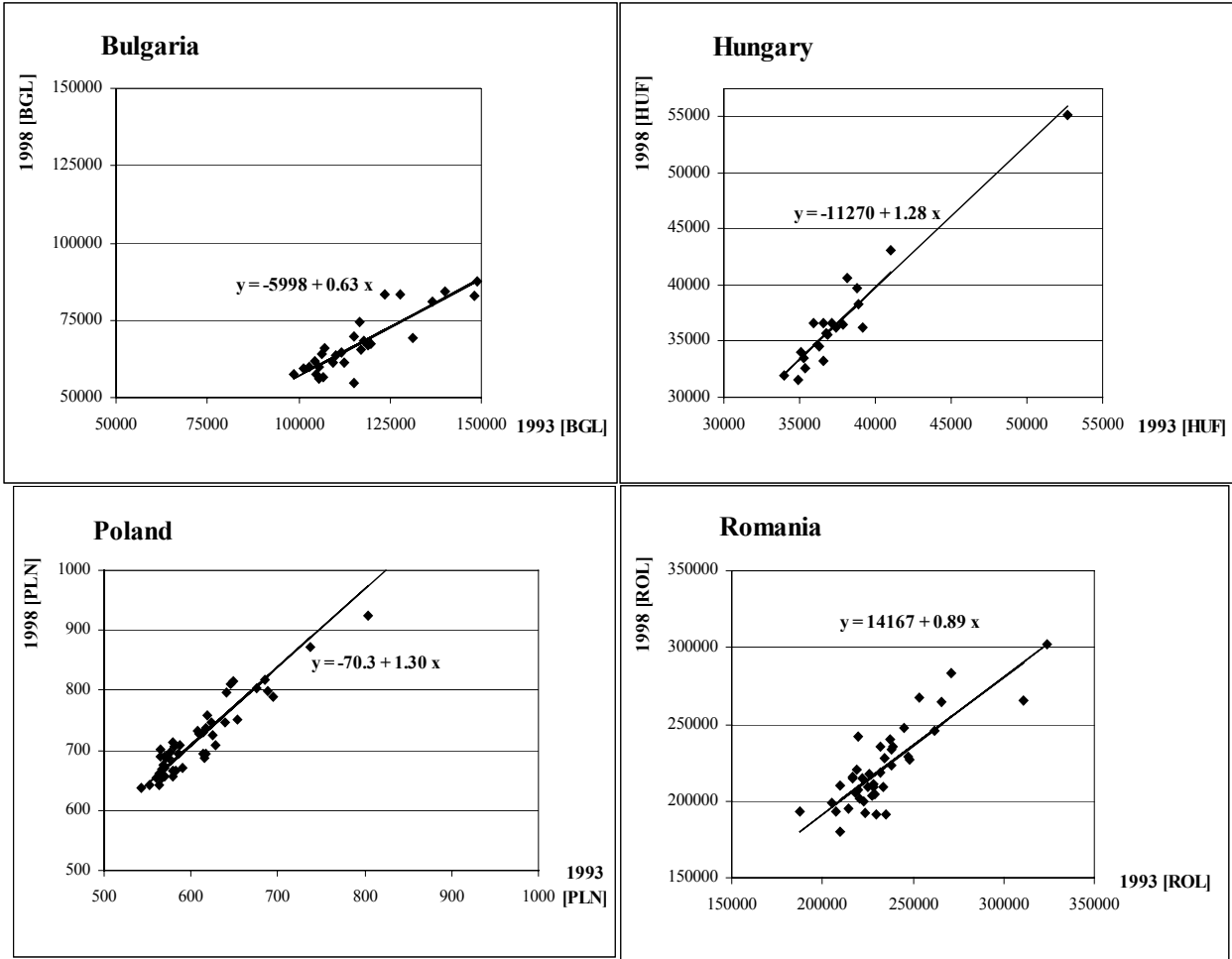


Figure 4 Regional average monthly earnings: Persistence



References

- Adams, J.D., 1985. Permanent differences in unemployment and permanent wage differentials. *Quarterly Journal of Economics* 100, 29-56.
- Arellano, M. and S. Bond, 1991. Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations. *Review of Economic Studies* 58, 277-297.
- Badinger, H., W. Müller and G. Tondl, 2002. Regional convergence in the European Union (1985-1999) : A spatial dynamic panel analysis. Research Institute for European Affairs (IEF) Working Paper 47, forthcoming in *Regional Studies*.
- Badinger, H. and T. Url, 2002. Determinants of regional unemployment: Some evidence from Austria. *Regional Studies* 36, 977-988.
- Bajo, O., I. Rabadán, and R. Salas, 1999. Regional wage flexibility in Spain, 1989-1992. Instituto de Estudios Fiscales del Ministerio de Hacienda, working paper.
- Baltagi, B., 2001. *Econometric analysis of panel data*. John Wiley, West Sussex.
- Baltagi, B. and U. Blien, 1998. The German wage curve: Evidence from the IAB employment sample. *Economics Letters* 61, 135-142.
- Bell, B., S. Nickell and G. Quintini, 2002. Wage equations, wage curves and all that. *Labour Economics* 9, 341-360.
- Blanchflower, D.G. and A.J. Oswald, 1994. *The wage curve*. The MIT Press, Cambridge, MA.

- Blanchflower, D.G., 2001. Unemployment, well-being and wage curves in Eastern and Central Europe. Mimeo.
- Blackaby, D.H. and D.N. Manning, 1987. Earnings, unemployment and the regional employment structure in Britain. *Regional Studies* 24, 529-535.
- Boeri, T. and S. Scarpetta, 1996. Regional mismatch and the transition to a market economy. *Labour Economics* 3, 233-254.
- Büttner, T., 1999a. The effect of unemployment, aggregate wages, and spatial contiguity on local wages: An investigation with German district level data. *Papers in Regional Science* 78, 47-67.
- Büttner, T., 1999b. Agglomeration, growth, and adjustment: A theoretical and empirical study of regional labor markets in Germany. Physica-Verlag, Heidelberg, New York.
- Card, D., 1990. Unexpected inflation, real wages and employment determination in union contracts. *American Economic Review* 80, 669-688.
- Card, D., 1995. The wage curve: A review. *Journal of Economic Literature* 33, 785-799.
- Duffy, F. and P. Walsh, 2001. Individual pay and outside options: Evidence from the Polish Labour Force Survey. IZA Discussion Paper 295.
- EBRD (European Bank for Reconstruction and Development), 2000. Transition report 2000: Employment, skills and transition. London.
- Elhorst, J.P., U. Blien and K. Wolf, 2002. A spatial panel approach to the East German wage curve. Paper presented at the 42nd congress of the European Regional Science Association, August 27-31, Dortmund/Germany.

- Fidrmuc, J., 2003. Migration and regional adjustment to asymmetric shocks in transition economies. CEPR Discussion Paper 3798.
- Fingleton, B., 1999. Estimates of time to economic convergence: An analysis of regions of the European Union. *International Regional Science Review* 22, 5-34.
- Freeman, R.B., 1988. Evaluating the European view that the United States has no unemployment problem. *American Economic Review Papers and Proceedings* 78, 294-299.
- Getis, A. and J.K. Ord, 1992. The analysis of spatial association by use of distance statistics. *Geographical Analysis* 24, 189-206.
- Getis, A., 1995. Spatial filtering in a regression framework: Examples using data on urban crime, regional inequality, and government expenditures. In: A. Anselin and R.J. Florax (Eds.), *New directions in spatial econometrics*. Springer, Berlin, pp. 172-185.
- Hall, R.E., 1970. Why is the unemployment rate so high at full employment? *Brookings Papers on Economic Activity*, 369-402.
- Hall, R.E., 1972. Turnover in the labour force. *Brookings Papers on Economic Activity*, 707-757.
- Harris, J.R. and M.P. Todaro, 1970. Migration, unemployment and development: A two-sector analysis. *American Economic Review* 60, 126-142.
- Jimeno, J. and S. Bentolila, 1998. Regional unemployment persistence (Spain, 1976-1994). *Labour Economics* 5, 25-51.
- Judson, R.A. and A.L. Owen, 1999. Estimating dynamic panel data models: A guide for macroeconomists. *Economics Letters* 65, 9-15.

- Kállai, E. and I. Traistaru, 2001. Characteristics and trends of regional labour markets in transition economies: Empirical evidence from Romania. Mimeo.
- Kertesi, G. and J. Köllô, 1997. The birth of the 'wage curve' in Hungary, 1989-95. WDI Working Paper 104.
- Kertesi, G. and J. Köllô, 1999. Unemployment, wage push and the labour cost competitiveness of regions: The case of Hungary, 1986-1996. Budapest Working Papers on the Labor Market 1999/5.
- Kiviet, J.F., 1995. On bias, inconsistency and efficiency of various estimators in dynamic panel data models. *Journal of Econometrics* 68, 53-78.
- Longhi, S., P. Nijkamp and J. Poot, 2002. Spatial contiguity and the wage curve revisited. Paper presented at the 42nd congress of the European Regional Science Association, August 27-31, Dortmund/Germany.
- Marston, S.T., 1985. Two views of the geographic distribution of unemployment. *Quarterly Journal of Economics* 100, 57-79.
- Montuenga, V., I. García, and M. Fernández, 2003. Wage flexibility: Evidence from five EU countries based on the wage curve. *Economics Letters* 78, 169-174.
- Nickell, S.J., 1981. Biases in dynamic models with fixed effects. *Econometrica* 49, 1417-1426.
- Partridge, M.D., and D.S. Rickman, 1997. Has the wage curve nullified the Harris-Todaro model? Further US evidence. *Economics Letters* 52, 277-282.
- Pannenberg, M. and J. Schwarze, 1998a. Labor market slack and the wage curve. *Economics Letters* 58, 351-354.

- Pannenberg, M. and J. Schwarze, 1998b. Phillips curve or wage curve: Is there really a puzzle? Empirical evidence for West Germany. Economic Discussion Papers 83, Bamberg University.
- Petrakos, G., 1996. The regional dimension of transition in Eastern and Central European Countries: An assessment. *Eastern European Economics* 34, 5-38.
- Petrakos, G., 2000. The spatial impact of East-West integration in Europe. In: G. Petrakos, G. Maier and G. Gorzelak (Eds.), *Integration and transition in Europe: The economic geography of interaction*. Routledge, London, 38-68.
- Quah, D., 1996. Regional convergence clusters across Europe. *European Economic Review* 40, 951-958.
- Reza, A.M., 1978. Geographical differences in earnings and unemployment rates. *Review of Economics and Statistics* 60, 201-208.
- Winter-Ebmer, R., 1996. Wage curve, unemployment duration and compensating differentials. *Labour Economics* 3, 425-434.

Tables

Table 1 Estimation results: Static fixed effects model

| dependent variable: ln $w_{r,t}$ | Bulgaria 1992-99 | Hungary 1992-99 | Poland 1992-98 | Romania 1992-99 |
|-------------------------------------|---------------------------|--------------------------|---------------------------|---------------------------|
| ln $U_{r,t}$ | -0.05*** (0.02) | -0.01 (0.03) | -0.06*** (0.02) | -0.003 (0.01) |
| s_empl1 $_{r,t}$ | 0.60*** (0.22) | -0.06 (0.24) | 0.10 (0.08) | -0.33 (0.21) |
| s_empl2 $_{r,t}$ | 0.13 (0.20) | 0.23*** (0.08) | -0.09** (0.04) | -0.89*** (0.17) |
| N obs. | 784 | 160 | 343 | 328 |
| overall R² | 0.86 | 0.15 | 0.38 | 0.14 |

Notes:

Regressors include time dummies and fixed effects.

Robust standard errors are reported in parentheses.

***, **, * indicate significance at 1, 5, 10 percent level respectively.

Table 2 Estimation results: Static fixed effects model, two sub-periods

| dependent variable: ln $w_{r,t}$ | Bulgaria | | Hungary | | Poland | | Romania | |
|------------------------------------|---------------------------|-----------------|------------------|-------------------------|------------------|---------------------------|-----------------|---------------------------|
| | 1992-94 | 1995-99 | 1992-94 | 1995-99 | 1992-94 | 1995-98 | 1992-94 | 1995-99 |
| ln $U_{r,t}$ | -0.07*** (0.02) | -0.03 (0.05) | -0.003 (0.02) | -0.01 (0.03) | -0.002 (0.03) | -0.07*** (0.02) | -0.01 (0.02) | -0.003 (0.02) |
| s_empl1 $_{r,t}$ | 0.17 (0.17) | -0.08 (0.28) | -0.24 (0.24) | -0.94* (0.52) | 0.17 (0.20) | 0.11 (0.09) | 0.23 (0.21) | -0.53 (0.34) |
| s_empl2 $_{r,t}$ | 0.47*** (0.20) | -0.43 (0.29) | 0.30 (0.24) | 0.10 (0.08) | -0.02 (0.05) | -0.23*** (0.07) | -0.09 (0.25) | -1.12*** (0.27) |
| N obs. | 140 | 84 | 60 | 100 | 147 | 196 | 123 | 205 |
| overall R² | 0.78 | 0.40 | 0.03 | 0.25 | 0.18 | 0.24 | 0.48 | 0.10 |
| F - ln $U_{r,t}$ | 0.47 | | 0.04 | | 4.02 | | 0.03 | |
| Pr>F | 0.49 | | 0.84 | | 0.05 | | 0.86 | |
| F - all variables | 2.44 | | 0.56 | | 3.29 | | 2.62 | |
| Pr>F | 0.06 | | 0.64 | | 0.02 | | 0.05 | |

Notes:

Regressors include time dummies and fixed effects.

Robust standard errors are reported in parentheses.

***, **, * indicate significance at 1, 5, 10 percent level respectively.

F test on equality of coefficients β_1 and β_2 in two subperiods – $H_0: \beta_1 - \beta_2 = 0$.

Table 3 Estimation results: One-step GMM

| dependent variable: ln w _{r,t} | Bulgaria 1991-99 | Hungary 1992-99 | Poland 1992-98 | Romania 1992-99 |
|--|---------------------------|---------------------------|---------------------------|---------------------------|
| ln w _{r,t-1} | 0.75*** (0.10) | 0.51*** (0.09) | 0.31 (0.20) | 0.07 (0.13) |
| ln w _{r,t-2} | -0.19*** (0.06) | 0.13* (0.07) | -0.25*** (0.08) | -0.16** (0.07) |
| ln U _{r,t} | -0.12*** (0.04) | -0.01 (0.01) | | 0.001 (0.02) |
| ln U _{r,t-1} | 0.04 (0.04) | -0.03 (0.03) | -0.04** (0.02) | |
| ln U _{r,t-2} | 0.04* (0.02) | -0.04 (0.03) | | |
| ln U _{r,t-3} | | 0.02* (0.01) | | |
| s_empl1 _{r,t} | | 0.43** (0.21) | | -0.69** (0.32) |
| s_empl1 _{r,t-1} | | | 0.14** (0.06) | |
| s_empl1 _{r,t-2} | | 0.33** (0.18) | | |
| s_empl2 _{r,t} | | 0.19*** (0.06) | -0.14*** (0.04) | -1.18*** (0.20) |
| s_empl2 _{r,t-1} | | -0.12*** (0.04) | | |
| ln U _r long-run multiplier | -0.18 | 0.06 | -0.03 | -- |
| Wald χ^2 Pr> χ^2 | 4.39 0.04 | 2.92 0.09 | 4.50 0.03 | |
| predet. | ln U _r | ln U _r | -- | ln U _r |
| No. of obs. | 168 | 80 | 196 | 164 |
| Wald χ^2 | 4078.45 | 15701.86 | 109.21 | 5851.65 |
| AR1 errors z Pr>z | -3.45 0.001 | -2.96 0.003 | -1.26 0.21 | -1.78 0.07 |
| AR2 errors z Pr>z | 0.67 0.51 | 0.40 0.69 | 1.31 0.19 | 0.30 0.76 |
| Sargan χ^2 Pr> χ^2 | 14.97 1.00 | 6.07 1.00 | 1.02 0.31 | 33.27 0.69 |

Notes:

Variables are in first differences. Regressors include time dummies.

Robust standard errors are reported in parentheses.

***, **, * indicate significance at 1, 5, 10 percent level respectively.

Tests:

- Wald test on significance of long-run multiplier – H₀: The long-run multiplier calculated from the individually significant coefficients is insignificant.
- Arellano-Bond test on average order 1 autocovariance in residuals (AR1 errors) – H₀: The residuals are not autocorrelated.
- Arellano-Bond test on average order 2 autocovariance in residuals (AR2 errors) – H₀: The residuals are not autocorrelated.
- Sargan test of over-identifying restrictions (results from 2-step GMM with standard errors not corrected for heteroskedasticity) - H₀: The over-identifying restrictions are valid.

Table 4 Estimation results: One-step GMM, Phillips curve specification

| dependent variable: ln $w_{r,t}$ | Bulgaria 1991-99 | Hungary 1992-99 | Poland 1992-98 | Romania 1992-99 |
|-------------------------------------|---------------------------|--------------------------|---------------------------|---------------------------|
| ln $w_{r,t-1}$ | 0.65*** (0.08) | 0.55*** (0.07) | 0.25* (0.14) | 0.11 (0.12) |
| ln $U_{r,t}$ | -0.12*** (0.04) | -0.005 (0.02) | | 0.007 (0.02) |
| ln $U_{r,t-1}$ | 0.05 (0.04) | -0.05* (0.03) | -0.03** (0.02) | |
| ln $U_{r,t-2}$ | 0.05** (0.02) | -0.03 (0.03) | | |
| ln $U_{r,t-3}$ | | 0.02* (0.01) | | |
| s_empl1 $_{r,t}$ | | 0.40 (0.25) | | -0.59** (0.27) |
| s_empl1 $_{r,t-1}$ | | | 0.13*** (0.04) | |
| s_empl1 $_{r,t-2}$ | | 0.32* (0.18) | | |
| s_empl2 $_{r,t}$ | | 0.18*** (0.56) | -0.09*** (0.03) | -1.14*** (0.19) |
| s_empl2 $_{r,t-1}$ | | -0.13*** (0.04) | | |
| ln U_r long-run multiplier | -0.19 | -0.06 | -0.05 | -- |
| Wald χ^2 Pr> χ^2 | 2.78 0.09 | 0.73 0.39 | 3.87 0.05 | |
| predet. | ln U_r | ln U_r | -- | ln U_r |
| No. of obs. | 168 | 80 | 245 | 246 |
| Wald χ^2 | 2388.84 | 11774.65 | 347.63 | 3534.92 |
| AR1 errors z Pr>z | -3.05 0.002 | -2.61 0.01 | -1.49 0.14 | -2.67 0.01 |
| AR2 errors z Pr>z | -0.84 0.40 | 1.20 0.23 | -2.12 0.03 | -1.39 0.16 |
| Sargan χ^2 Pr> χ^2 | 18.75 1.00 | 7.27 1.00 | 9.10 0.82 | 31.99 0.81 |

Notes:

Variables are in first differences.

Regressors include time dummies. Robust standard errors are reported in parentheses.

***, **, * indicate significance at 1, 5, 10 percent level respectively.

On the tests reported, see Notes to Table 3.

Table 5 Spatial autocorrelation in the variables

| | ln w _r | | ln U _r | | s_empl1 _r | | s_empl2 _r | |
|-----------------|-------------------|-------|-------------------|-------------|----------------------|-------------|----------------------|-------------|
| | LM | Pr>LM | LM | Pr>LM | LM | Pr>LM | LM | Pr>LM |
| Bulgaria | | | | | | | | |
| 1992 | 0.0050 | 0.94 | 1.1804 | 0.28 | 1.9672 | 0.16 | 0.9915 | 0.32 |
| 1993 | 0.0340 | 0.85 | 0.1528 | 0.70 | 1.3050 | 0.25 | 1.8240 | 0.18 |
| 1994 | 0.0670 | 0.80 | 0.0528 | 0.82 | 1.2424 | 0.27 | 0.6963 | 0.40 |
| 1995 | 0.0003 | 0.99 | 0.1948 | 0.66 | 0.8206 | 0.36 | 1.5263 | 0.22 |
| 1996 | 0.1715 | 0.68 | 0.5922 | 0.44 | 0.9193 | 0.34 | 0.9314 | 0.33 |
| 1997 | 0.1437 | 0.70 | 0.7642 | 0.38 | 0.3040 | 0.58 | 1.5703 | 0.21 |
| 1998 | 0.0005 | 0.98 | 0.9337 | 0.33 | 0.1344 | 0.71 | 2.4336 | 0.12 |
| 1999 | 0.0303 | 0.86 | 1.3618 | 0.24 | 0.4670 | 0.49 | 1.7752 | 0.18 |
| Hungary | | | | | | | | |
| 1992 | 0.0592 | 0.81 | 1.6607 | 0.20 | 0.0000 | 1.00 | 1.1253 | 0.29 |
| 1993 | 0.0008 | 0.98 | 2.2632 | 0.13 | 0.0876 | 0.77 | 0.4656 | 0.50 |
| 1994 | 0.2266 | 0.63 | 2.8646 | 0.09 | 0.2312 | 0.63 | 0.3640 | 0.55 |
| 1995 | 0.1388 | 0.71 | 2.7132 | 0.10 | 0.1715 | 0.68 | 0.5156 | 0.47 |
| 1996 | 0.3404 | 0.56 | 3.1808 | 0.07 | 0.3818 | 0.54 | 0.0934 | 0.76 |
| 1997 | 0.8689 | 0.35 | 2.7409 | 0.10 | 0.2651 | 0.61 | 0.0324 | 0.86 |
| 1998 | 0.6365 | 0.42 | 2.8495 | 0.09 | 0.4688 | 0.49 | 0.0003 | 0.99 |
| 1999 | 0.8371 | 0.36 | 1.9154 | 0.17 | 0.4222 | 0.52 | 0.0978 | 0.75 |
| Poland | | | | | | | | |
| 1992 | 0.1646 | 0.68 | 3.6663 | 0.06 | 11.3508 | 0.00 | 5.8671 | 0.02 |
| 1993 | 0.2687 | 0.60 | 4.2243 | 0.04 | 10.5313 | 0.00 | 2.0134 | 0.16 |
| 1994 | 0.4166 | 0.52 | 1.9126 | 0.17 | 10.9388 | 0.00 | 3.4046 | 0.07 |
| 1995 | 0.5720 | 0.45 | 1.2581 | 0.26 | 9.5231 | 0.00 | 1.1651 | 0.28 |
| 1996 | 0.6188 | 0.43 | 0.9060 | 0.34 | 14.7458 | 0.00 | 5.9518 | 0.01 |
| 1997 | 0.8126 | 0.37 | 0.1499 | 0.70 | 15.0584 | 0.00 | 7.1536 | 0.01 |
| 1998 | 0.6637 | 0.42 | 0.0088 | 0.93 | 15.1339 | 0.00 | 6.7609 | 0.01 |
| Romania | | | | | | | | |
| 1992 | 0.0226 | 0.88 | 3.1051 | 0.08 | 0.0004 | 0.98 | 0.5739 | 0.45 |
| 1993 | 0.0653 | 0.80 | 0.9988 | 0.32 | 0.0567 | 0.81 | 0.3662 | 0.55 |
| 1994 | 0.2751 | 0.60 | 2.0767 | 0.15 | 0.0197 | 0.89 | 0.6251 | 0.43 |
| 1995 | 0.0881 | 0.77 | 3.3467 | 0.07 | 0.0047 | 0.95 | 0.5631 | 0.45 |
| 1996 | 0.0448 | 0.83 | 3.5424 | 0.06 | 0.0133 | 0.91 | 0.4134 | 0.52 |
| 1997 | 0.0075 | 0.93 | 2.2205 | 0.14 | 0.0486 | 0.83 | 0.2788 | 0.60 |
| 1998 | 0.6295 | 0.43 | 1.2585 | 0.26 | 0.0086 | 0.93 | 0.6523 | 0.42 |
| 1999 | 0.4778 | 0.49 | 1.0421 | 0.31 | 0.0111 | 0.92 | 0.6694 | 0.41 |

Note:

LM test on spatial autocorrelation in the variables – H₀: The variable is not spatially autocorrelated.

Table 6 Estimation results: One-step GMM, spatially filtered variables

| dependent variable: ln $w_{r,t}$ | Bulgaria 1991-99 | Hungary 1992-99 | Poland 1992-98 | Romania 1992-99 |
|-------------------------------------|---------------------------|---------------------------|---------------------------|---------------------------|
| ln $w_{r,t-1}$ | 0.67*** (0.13) | 0.48*** (0.08) | 0.35* (0.20) | 0.08 (0.14) |
| ln $w_{r,t-2}$ | -0.27*** (0.07) | | -0.25*** (0.08) | -0.15** (0.07) |
| ln $U_{r,t}$ | -0.12** (0.06) | -0.003 (0.01) | | 0.01 (0.02) |
| ln $U_{r,t-1}$ | | -0.04 (0.03) | -0.04** (0.02) | |
| ln $U_{r,t-2}$ | | -0.05* (0.03) | | |
| ln $U_{r,t-3}$ | | 0.03** (0.02) | | |
| s_empt1 $_{r,t}$ | | | | -0.63* (0.34) |
| s_empt1 $_{r,t-1}$ | | | 0.15*** (0.07) | |
| s_empt1 $_{r,t-2}$ | | 0.28* (0.16) | | |
| s_empt2 $_{r,t}$ | | 0.16*** (0.06) | -0.13*** (0.04) | -1.17*** (0.21) |
| s_empt2 $_{r,t-1}$ | | -0.14*** (0.04) | | |
| ln U_r long-run multiplier | -0.20 | -0.04 | -0.04 | -- |
| Wald χ^2 | 4.73 | 0.50 | 5.91 | |
| Pr> χ^2 | 0.03 | 0.48 | 0.02 | |
| predet. | ln U_r | ln U_r | -- | ln U_r |
| N obs. | 168 | 80 | 196 | 164 |
| Wald χ^2 | 4203.17 | 12193.81 | 104.28 | 6420.19 |
| AR1 errors z | -3.26 | -2.61 | -1.43 | -1.67 |
| Pr>z | 0.001 | 0.01 | 0.15 | 0.10 |
| AR2 errors z | 0.62 | 1.29 | 1.40 | -0.27 |
| Pr>z | 0.54 | 0.20 | 0.16 | 0.78 |
| Sargan χ^2 | 20.70 | 9.31 | 7.30 | |
| Pr> χ^2 | 1.00 | 1.00 | 0.64 | |

Notes:

Variables are in first differences. Regressors include time dummies.

Robust standard errors are reported in parentheses.

***, **, * indicate significance at 1, 5, 10 percent level respectively.

On the tests reported, see Notes to Table 3.