How Flexible are Wages in EU Accession Countries?*

Anna Iara[a**] and Iulia Traistaru[b]

[a] Center for European Integration Studies (ZEI), University of Bonn

[b]Center for European Integration Studies (ZEI), University of Bonn

Abstract

The transition to a market economy and increased economic integration has 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 for a wage curve in Bulgaria, Hungary and Poland suggesting that wages could help equilibrate labour markets following labour demand shocks. In the case of Romania, the elasticity of pay to unemployment is not significantly different from zero.

Keywords: wage flexibility, panel data, EU accession countries

JEL classification: C23, J30, J60

---

* This research was undertaken with financial support from the European Community's 5th Framework Programme. We are grateful for helpful comments and suggestions to Eric Crampton, Mihails Hazans, Peter Huber, Simonetta Longhi, Ron Moomaw, Gabriele Tondl, and Guntram Wolff. We also benefited from discussions with the participants in the seminars and workshops held at the Center for European Integration Studies, University of Bonn and University of Budapest. 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.

** Corresponding author. Center for European Integration Studies, Walter-Flex-Str. 3, D-53113 Bonn, Germany; Tel. +49-228-731831; Fax +49-228-731809; E-mail: anna.iara@uni-bonn.de.

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 (European Bank for Reconstruction and Development, 2000). Furthermore, there is growing evidence of a strong regional dimension to 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 inter-regional 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, wage flexibility could play an important role in helping labour markets adjust to labour 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 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 wage responsiveness to local market conditions in the above mentioned countries using a unique 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. 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 to technology spillovers (Fingleton, 1999, Quah, 1996). Büttner (1999a), Elhorst et al. (2002) and Longhi et al. (2002) demonstrate that neglecting spatial dependence can bias downwards the unemployment rate effect on wages. We therefore correct for spatial dependence using a filtering procedure based on Getis and Ord (1992) and Getis (1995). We find empirical support for the wage curve in Bulgaria, Hungary 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). Hungarian and Polish elasticities, -0.05 and –0.04 respectively, were only half of advanced economy elasticities.

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 discusses previous results on wage curves in transition countries and selected Western economies. In section 3 we present our data followed by a brief discussion of summary statistics of regional unemployment rates and average earnings in Bulgaria, Hungary, Poland and Romania. Section 4 outlines our estimation strategy. The estimates for the regional earnings' responsiveness to local market conditions are discussed in section 5. Section 6 concludes.

2 STYLIZED FACTS

Conventional economic theory going back to Adam Smith 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 stylised 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 datasets have the advantage of allowing the use of control variables specific to standard wage equations *a 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). One way to solve this problem is to aggregate individual data at regional level. Another 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 to unemployment rates is determined by the elasticity of hourly wages to unemployment and the elasticity of hours worked with respect to unemployment. However, most studies do not control for the numbers of hours worked. This implies that the magnitude of the unemployment elasticity of wages is overestimated.

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 the nine EU accession countries and six successors of 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 is related to 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, Rabadán and Salas, 1999, for the case of Spain, and Bell et al. 2002 for Great Britain). In contrast, a recent paper by Montuenga et al. (2003)[1] 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

---

[1] The cross-country comparability of these results is impeded by the different size of the regions. While spatial dependence may induce bias especially in the case of small regions that may not be exactly separable into distinct labour market areas, aggregation to larger regions may hide cross-sectional variation in the data. Kertesi and Köllô (1999) find that higher aggregation of a single dataset lowers the estimated elasticity of pay by 0.02.

results from panel estimators. Consequently, other authors use estimation techniques robust to the lack of strict exogeneity of unemployment rates, such as Arellano-Bond GMM (Jimeno and Bentolila, 1998) and 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. In contrast, using a different spatial model, Büttner (1999a) finds that controlling for spatial correlation in unemployment rates of neighbouring regions lowers estimates of unemployment elasticity of pay.

Our paper fills a gap in the literature by providing transition country wage curve estimates that correct for potential endogeneity of unemployment rates and spatial dependence in regional data. We exploit a unique regional dataset for Bulgaria, Hungary, Poland and Romania that allows us to correct for both endogeneity and spatial dependence in a dynamic panel model.

3 THE DATA AND SUMMARY STATISTICS

In this paper we use a unique data set[2] that includes annual regional labour market data at NUTS 3 level for Bulgaria, Hungary, Poland and Romania for the period 1992-1999.[3] The average size of regions varies in the four countries as shown in Table 1, with average numbers of inhabitants ranging from 300,000 (in Bulgaria) to 790,000 (in Poland).

Insert Table 1 about here

The data were collected from national statistical offices. The regional data used in our analysis include:

- regional average monthly earnings in 1995 prices in the respective national currency
- regional unemployment rates based on end-year numbers of registered unemployed
- sectoral employment shares including agriculture, industry and services

The latter are used as control variables. Given differences in data collection and availability, the datasets are not fully comparable across countries. More detailed information on the variables is given in Table 2.

Insert Table 2 about here

Compared to wage curve studies using micro individual data, this data set does not allow us to control for hours worked or for composition effects in terms of individual characteristics of

---

[2] This data set was generated in the framework of the research project ”Regional Labour Market Adjustment in the Accession Candidate Countriues” undertaken with financial support from the European Commission’s RTD 5th Framework Programme.

workers and the unemployed. In addition, regional variation in real earnings data is not fully reflected by our data since regional price indices could not be used. For these reasons, our estimates of the unemployment elasticity of pay may be biased upward[4].

Our panel exhibits substantial variation in unemployment rates, both cross-sectionally and over time (Figures 1a-d). In the 1990s, unemployment was relatively low in Hungary and Romania, with unemployment rates between 7 to 12 percent, while Poland and Bulgaria experienced rather high unemployment, with average regional unemployment rates ranging from 8 to 18 percent. The coefficients of variation of unemployment rates suggest that regional disparities with respect to unemployment rates were high in Hungary but less important in Bulgaria, Romania and Poland.

Insert Figures 1a-d about here

As shown in Figures 2a-d, average unemployment rates in 1998 were lower than the comparable 1993 figures in all countries investigated.

Insert Figures 2a-d about here

Regional disparities with respect to earnings were the highest in Hungary and the lowest in Romania, as shown in Figures 3a-d.

Insert Figures 3a-d about here

---

[3] The respective spatial units in the countries considered are: oblast (Bulgaria), megye (Hungary), województwa (Poland), and judet (Romania). For Poland, due to administrative reform, appropriate data are no longer available for 1999, while for Bulgaria, our dataset also includes data from 1991.

[4] Blanchflower and Oswald (1995) found that the wage curve estimates for the United Kingdom were robust to the use of regional price indices. Although this may not directly translate to the countries included into our study, bias from using a national price deflator can be considered small.

Real average monthly earnings increased by about 30 percent in Hungary and Poland over the period from 1993 to 1998, but decreased slightly in Romania and decreased strongly in Bulgaria over the same period (see Figures 4a-d).

Insert Figures 4a-d about here

4 MODEL SPECIFICATIONS

The existing literature on transition economies' wage curves 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). 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.[5] 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 e.g. 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 very large samples, it is a significant problem in small samples[6]. 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).

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

6 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.

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 to regions' wage levels raising 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 then account for spatial dependence and re-estimate the dynamic panel model with spatially filtered variables.[7]

In order to allow 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:

$$\log w_{rt} = \beta \log U_{rt} + \gamma X'_{rt} + \mu_r + \lambda_t + \varepsilon_{rt} \qquad (3)$$

where

[7] As concerns the data on Bulgaria which cover the years 1991-1999, for cross-country comparability the 1991 observations are left out in the estimations of the static model, while we made use of them in the dynamic panel

w_{rt} is the monthly average of earnings from work in region r at time t,

deflated with the national inflation index (the consumer price index),

U_{rt} is the unemployment rate in region r at time t,

X'_{rt} is a vector of variables controlling for the regional economic structure[8],

μ_r is a time invariant region-specific effect, $\mu_r \sim$ i.i.d. $N(0, \sigma^2_\mu)$,

λ_t is a region-invariant time specific effect, $\lambda_t \sim$ i.i.d. $N(0, \sigma^2_\lambda)$,

ε_{rt} is the remainder stochastic error term, $\varepsilon_{rt} \sim$ i.i.d. $N(0, \sigma^2_\varepsilon)$.

The transition economics literature points to two phases in CEEC transition to market economies. 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 as suggested by Arellano and Bond (1991). The estimated dynamic model has the following form:

$$\log w_{rt} = \Sigma_k \alpha_k \log w_{r,t-k} + \Sigma_l \beta_l \log U_{r,t-l} + \Sigma_m \gamma_m X'_{r,t-m} + \mu_r + \lambda_t + \varepsilon_{r,t} \quad (4).$$

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

---

data estimations to be able to allow for a higher-order lag structure. On the other hand, for Poland no data are available for 1999.

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 correlation (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.[9] 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.[10] For the choice between specifications, however, we use the Sargan test of over-identifying restrictions after the corresponding two-step GMM estimator.[11] Since consistency of the estimator requires the absence of second-order autocorrelation in the differenced residuals, we consider only

---

[8] As elements of X we employ 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 Table 2). In each case the sector pair selected exhibited the lowest correlation with each other.

[9] 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.

[10] Arellano and Bond (1991) recommend the one-step GMM estimator for inference on coefficients' significance, since according to their findings, standard errors from the two-step estimator tend to contain substantial downward bias in small samples.

[11] No robust Sargan test using one-step residuals is available.

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:

$$\log w_{rt} = \alpha \log w_{r,t-1} + \Sigma_l \beta_l \log U_{r,t-l} + \Sigma_m \gamma_m X'_{r,t-m} + \mu_r + \lambda_t + \varepsilon_{r,t} \quad (5).$$

This specification includes a test of the Phillips curve nested into the wage curve model. In particular, if α=1, one gets the familiar result that the wage change is determined by unemployment, whereas α=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 (LM) calculated on the basis of Moran's I statistics. The Moran's I coefficients are calculated as follows:

$$I = [(x-\mu)' W (x-\mu)] / [(x-\mu)'(x-\mu)] \quad (6),$$

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,[12] and it is obtained as follows:

$$LM = (N\ I)^2 / [tr\ (W'W + W^2)\ ] \quad (7),$$

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[13]. The filtering methodology is defined as follows:

$$x^F_i = x_i\ [\Sigma_j\ w_{ij}\ (\delta) / (N-1)\ ] / G_i\ (\delta), \quad (8)$$

with

$$G_i\ (\delta) = \Sigma_j\ w_{ij}\ (\delta)\ x_j / \ \Sigma_j\ x_j, \quad i \neq j\ . \quad (9),$$

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

[12] For details on this methodology, see Longhi et al. (2002).

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

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.

5 EMPIRICAL RESULTS

Results of the wage curve estimations for Bulgaria, Hungary, Poland and Romania are described below.[14]

Table 3 shows the results of the standard wage curve static model with time and region fixed effects following the model specification described by (3).

Insert Table 3 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 percent 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-estimate the static wage curve model for two sub-periods, namely, 1992-1994 and 1995-1999. The estimation results are displayed in Table 4.

Insert Table 4 about here

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 (4). The estimation results are shown in Table 5.

Insert Table 5 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 low, -0.04.

We further calculate the long-run effect of changing regional unemployment rates on regional real earnings captured by the long-run multiplier calculated on the basis of the obtained coefficients for the lagged regional 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

[14] Estimations were obtained using the STATA version 7 software.

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 real earnings appears positive in Hungary, with a long-run multiplier amounting to 0.06. The long-run multiplier is significant for all the three countries at the 5 percent (Bulgaria, Poland) and 10 percent (Hungary) level, respectively.

In the next step, we compare our results obtained from the unrestricted dynamic wage curve model with the Phillips curve. Table 6 shows the results of the corresponding model specification given in (5).

Insert Table 6 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 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 (6). The results of the LM statistics on spatial autocorrelation are shown in Table 7.

Insert Table 7 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 are 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 are 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 8 shows the results of the estimated dynamic model with spatially filtered variables.

Insert Table 8 about here

As shown above, 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.

6 CONCLUSIONS

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 estimate a standard static fixed effects model allowing for comparisons with results from existing literature. Further, we account for endogeneity in a dynamic wage curve model. In addition, we check and corrected for the presence of spatial dependence in the regional variables.

We find evidence on the adjustment of average regional real earnings over the past decade in Bulgaria, Hungary and Poland. The unemployment elasticity of pay was the highest in Bulgaria, -0.12, while in Hungary and Poland it was lower, -0.05 and –0.04 respectively. While in Bulgaria the regional earnings adjustment to local labour market conditions took place contemporaneously, in Hungary and Poland this adjustment took place with a two years and one year delay, respectively. The spatial effects played an important role in Hungary. 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 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.

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 No. 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, 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. and A.J. Oswald, 1995, An introduction to the wage curve, Journal of Economic Perspectives 9, 153-167.

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 No. 295.

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.

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), 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, William Davidson Institute Working Paper No. 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 No. 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 No. 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).

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 AND FIGURES

Table 1: Data Set Characteristics

Country	No. of regions	Average size, sqkm	Average population 1996, 1000s	Data available	No. of obs.
Bulgaria	28	3965	298	1991-99	252
Hungary	20	4651	511	1992-99	160
Poland	49	6381	789	1992-98	343
Romania	41	5876	551	1992-99	328

Table 2: Variable Definitions

Country	**Earnings in 1995 prices (CPI), national currency** (in logarithms)	**Unemploy-ment rates** (in logarithms)	**Employment shares by sectors:** Persons employed in sectors … as share of total employment	
Variable labels	lrwage	lurate	s_empl1	s_empl2
Bulgaria	Average net monthly earnings of persons employed	Official unemployment rates	industry (mining, manu-facturing, electricity, gas, water, and construction)	services
Hungary	Average gross monthly earnings of employees	Official unemployment rates, end-year	agriculture (incl. hunting, forestry, and fishing)	industry (mining, manu-facturing, electricity, gas, water, and construction)
Poland	Average gross monthly wages and salaries	Official unemployment rates, end-year	industry (mining, manu-facturing, electricity, gas, water, and construction)	services
Romania	Average gross monthly salary of employees	Official unemployment rates, end-year	industry (mining, manu-facturing, electricity, gas, water, and construction)	services

Figures 1a-d: Regional Unemployment Rates: Average and Coefficient of Variation

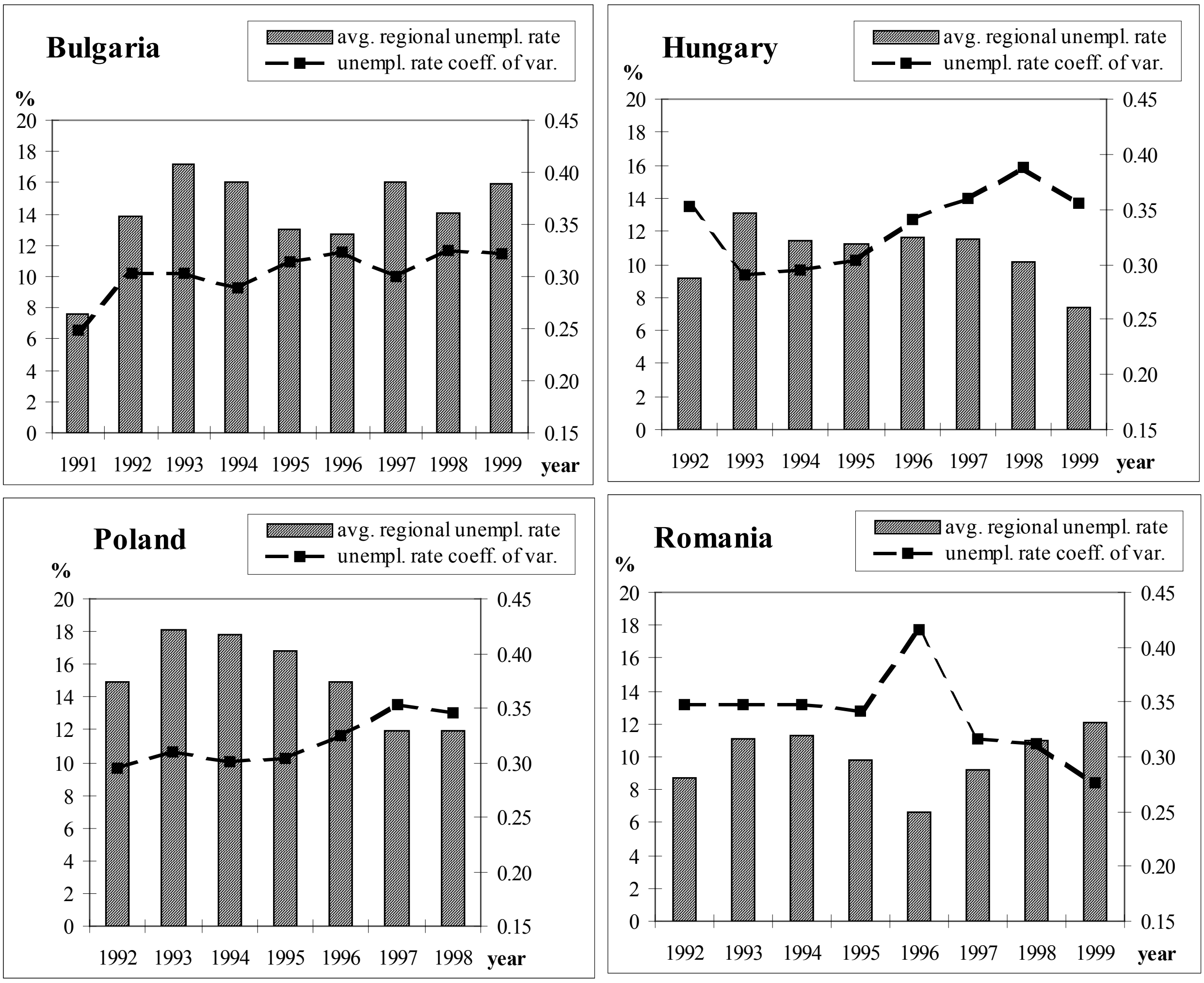


Figure 2a: Regional Unemployment Rates 1993 and 1998, Bulgaria

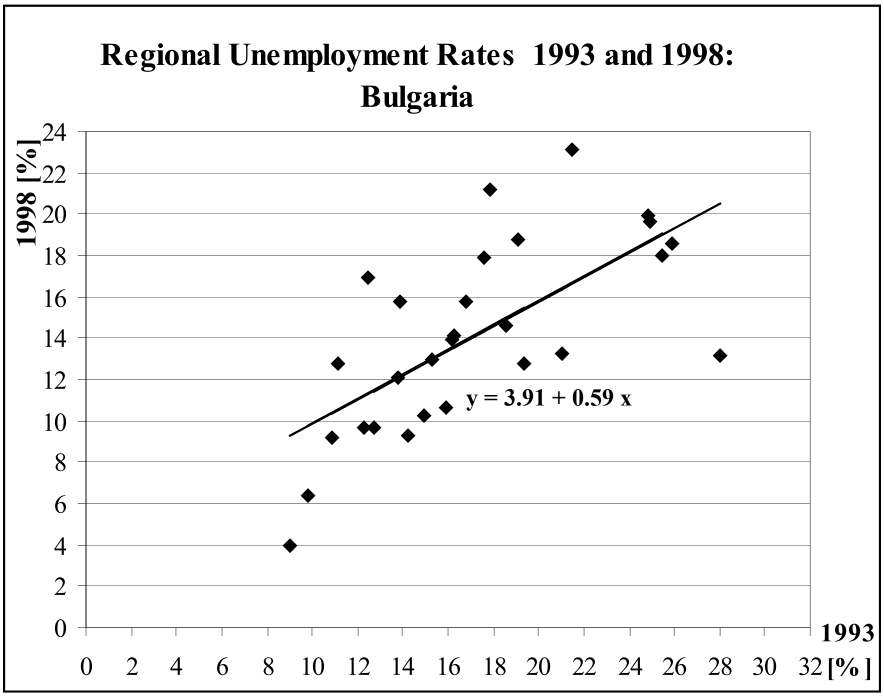


Figure 2b: Regional Unemployment Rates 1993 and 1998, Hungary

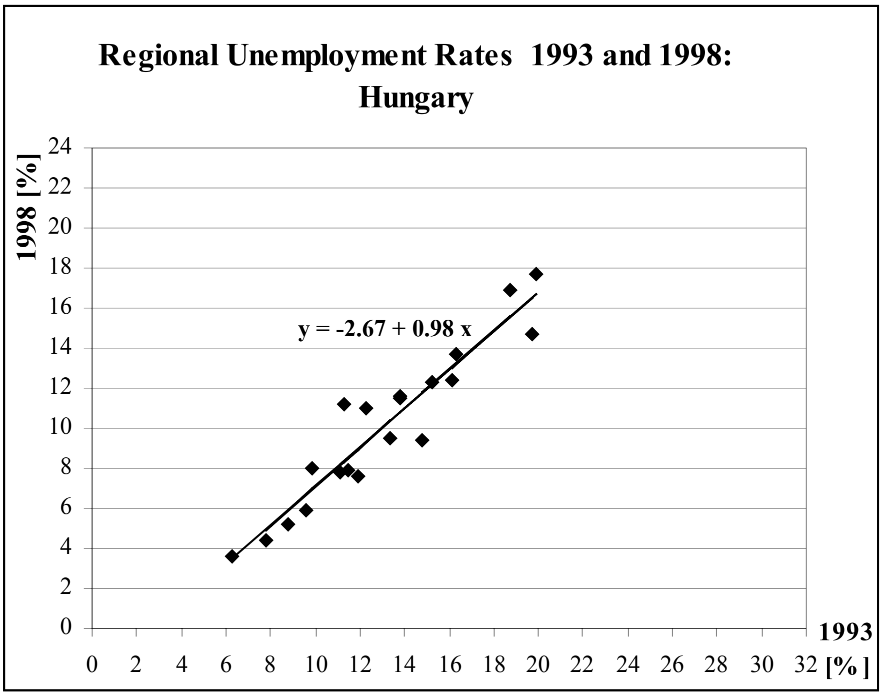


Figure 2c: Regional Unemployment Rates 1993 and 1998, Poland

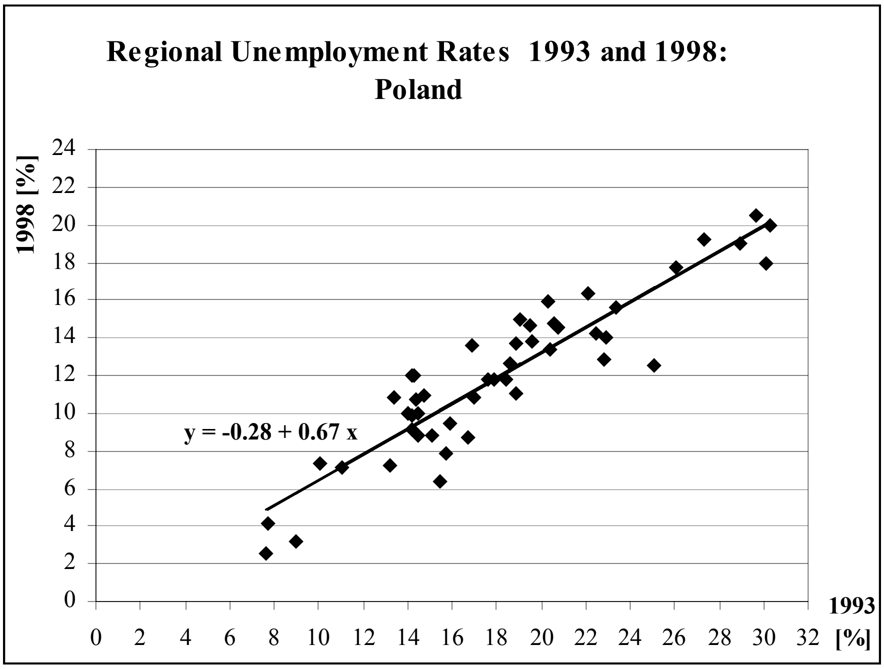
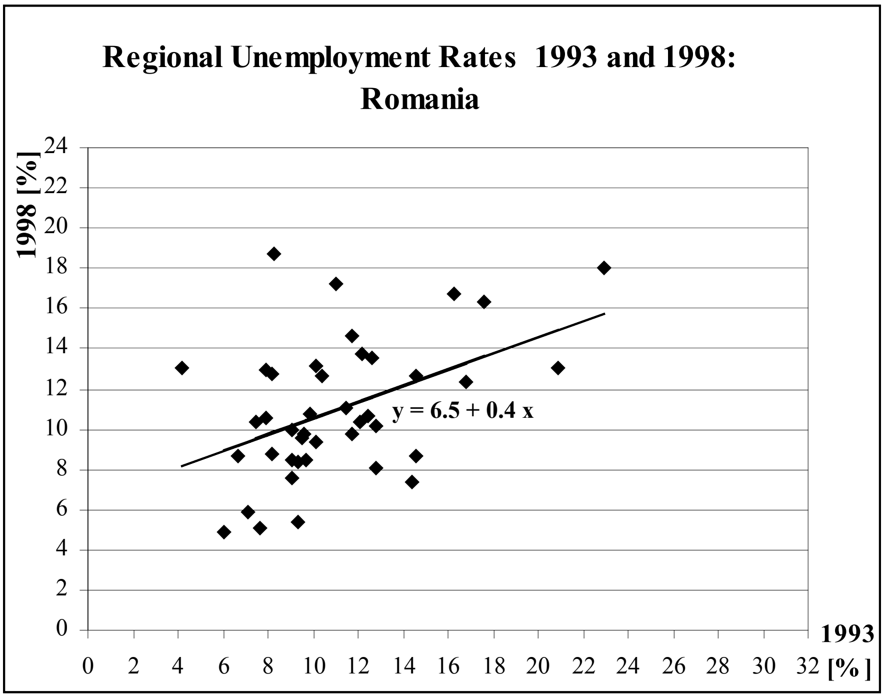


Figure 2d: Regional Unemployment Rates 1993 and 1998, Romania



Figures 3a-d: Regional Monthly Average Earnings: Average and Coefficient of Variation

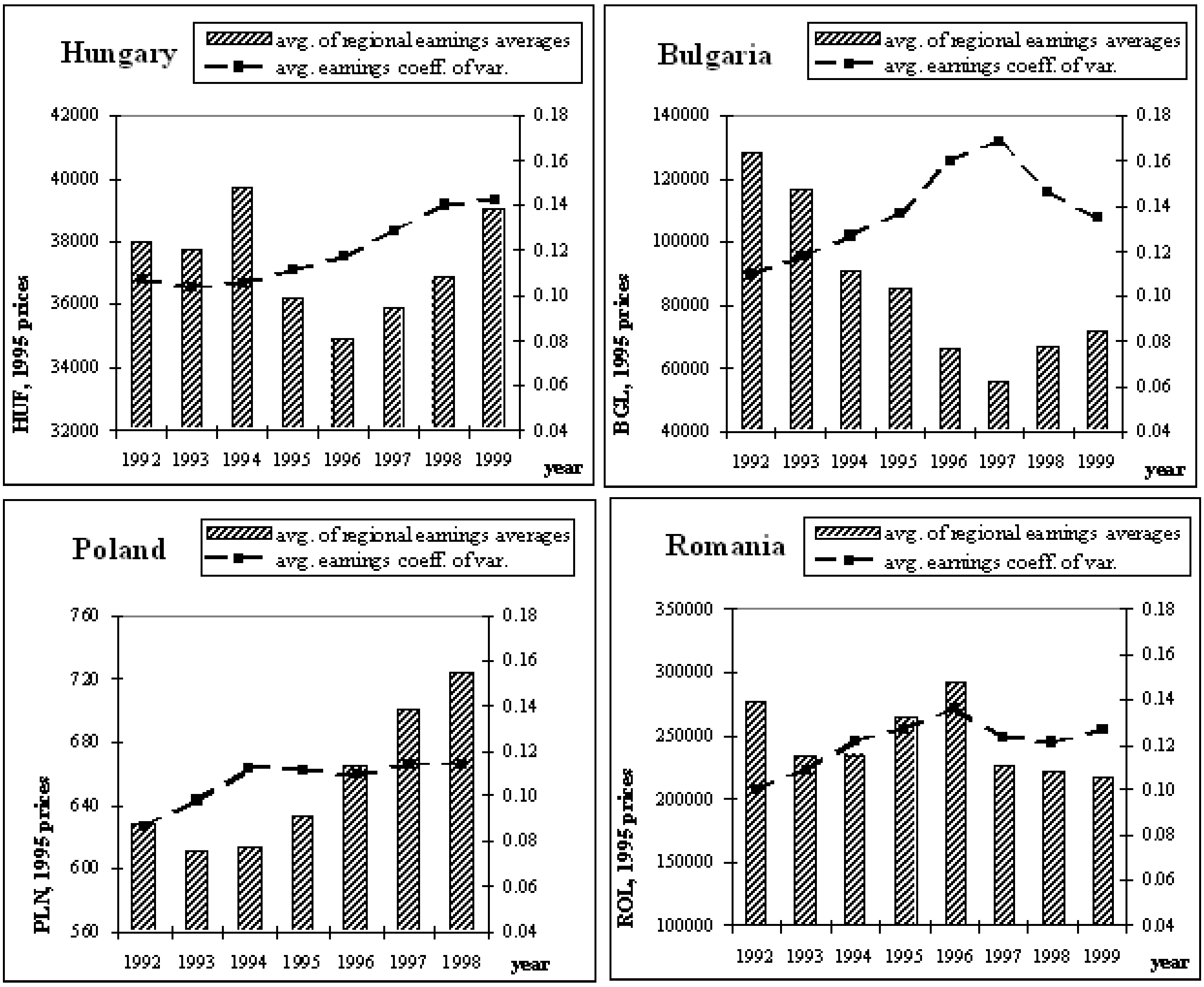


Figure 4a: Regional Average Earnings 1993 and 1998, Bulgaria

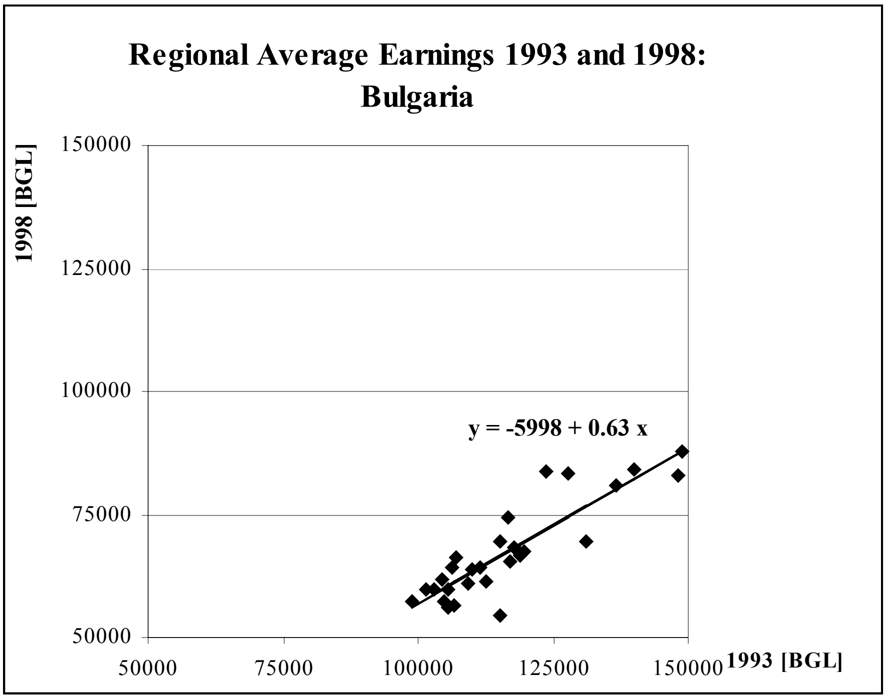


Figure 4b: Regional Average Earnings 1993 and 1998, Hungary

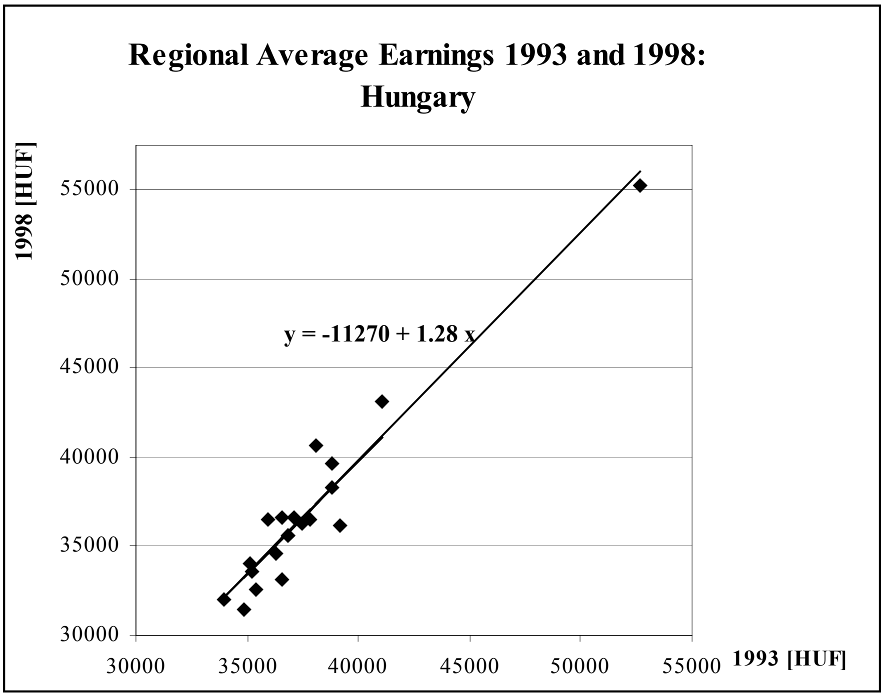


Figure 4c: Regional Average Earnings 1993 and 1998, Poland

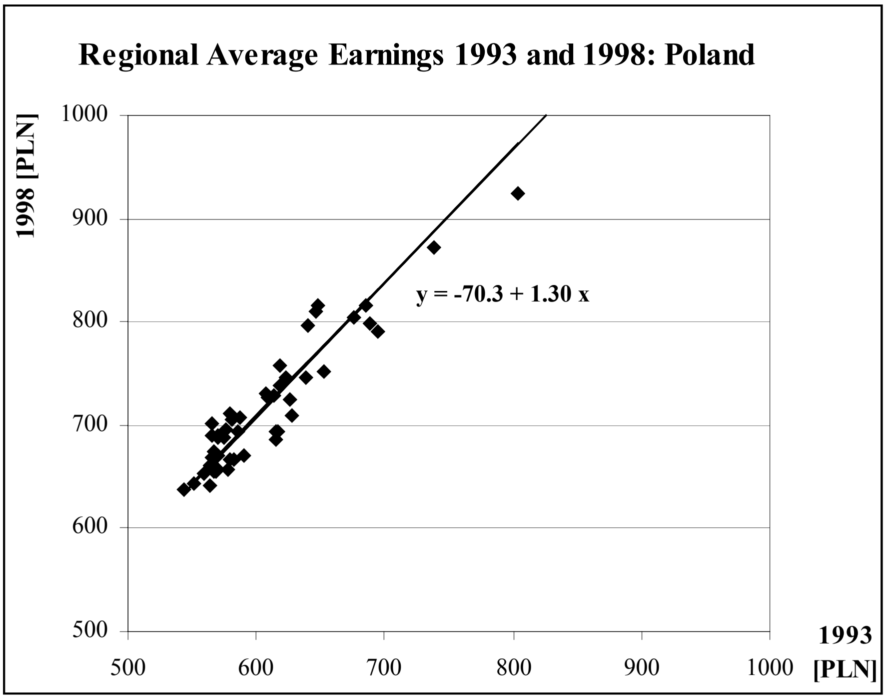


Figure 4d: Regional Average Earnings 1993 and 1998, Romania

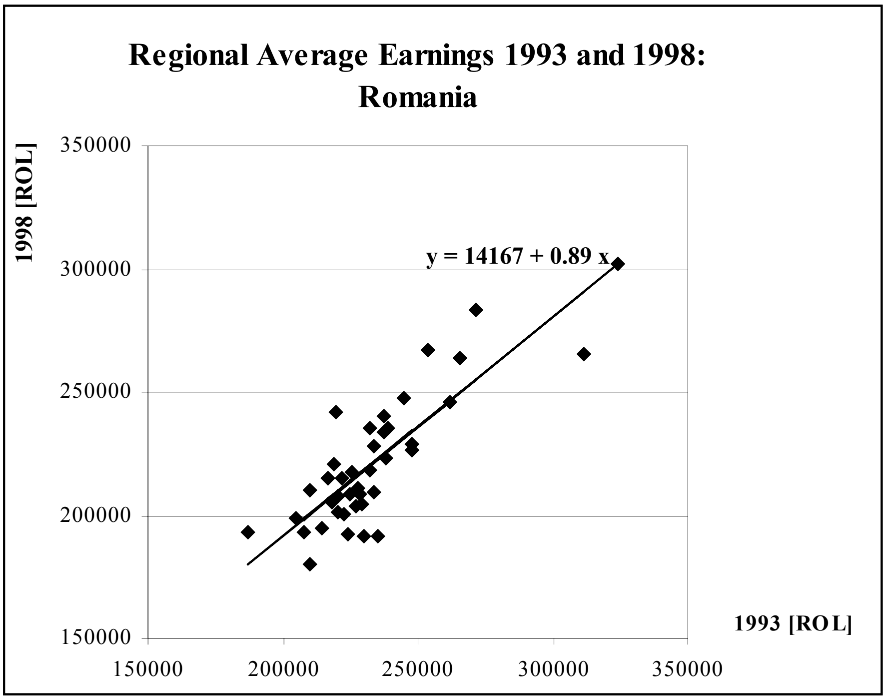


Table 3: Estimation Results: Static Fixed Effects Model

dep. variable: lrwage	Bulgaria 1992-99	Hungary 1992-99	Poland 1992-98	Romania 1992-99
lurate	**-0.05***** (0.02)	-0.01 (0.03)	**-0.06***** (0.02)	-0.003 (0.01)
s_empl1	**0.60***** (0.22)	-0.06 (0.24)	0.10 (0.08)	-0.33 (0.21)
s_empl2	0.13 (0.20)	**0.23***** (0.08)	**-0.09**** (0.04)	**-0.89***** (0.17)
N obs.	784	160	343	328
R^2	0.86	0.15	0.38	0.14

Note: Regressors include time dummies and fixed effects. Robust standard errors are reported in parentheses. ***, **, * indicate significance at 1, 5, 10 percent level respectively.

Table 4: Estimation Results: Static Fixed Effects Model, Two Subperiods

dep. variable: lrwage	Bulgaria		Hungary		Poland		Romania	
	1992-94	**1995-99**	**1992-94**	**1995-99**	**1992-94**	**1995-98**	**1992-94**	**1995-99**
lurate	**-0.07***** (0.022)	-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	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	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 R^2	0.78	0.40	0.03	0.25	0.18	0.24	0.48	0.10
lurates F Pr>F	0.47 0.49		0.04 0.84		4.02 0.05		0.03 0.86	
all variables F / Pr>F	2.44 0.06		0.56 0.64		3.29 0.02		2.62 0.05	

Note: 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: β_1-β_2=0.

Table 5: Estimation Results: 1-step GMM

dependent variable: lrwage	Bulgaria 1991-99	Hungary 1992-99	Poland 1992-98	Romania 1992-99
$lrwage_{t-1}$	**0.75***** (0.10)	**0.51***** (0.09)	0.31 (0.20)	0.07 (0.13)
$lrwage_{t-2}$	**-0.19***** (0.06)	**0.13*** (0.07)	**-0.25***** (0.08)	**-0.16**** (0.07)
$lurate_t$	**-0.12***** (0.04)	-0.01 (0.01)		0.001 (0.02)
$lurate_{t-1}$	0.04 (0.04)	-0.03 (0.03)	**-0.04**** (0.02)	
$lurate_{t-2}$	**0.04*** (0.02)	-0.04 (0.03)		
$lurate_{t-3}$		**0.02*** (0.01)		
s_empl1_t		**0.43**** (0.21)		**-0.69**** (0.32)
s_empl1_{t-1}			**0.14**** (0.06)	
s_empl1_{t-2}		**0.33**** (0.18)		
s_empl2_t		**0.19***** (0.06)	**-0.14***** (0.04)	**-1.18***** (0.20)
s_empl2_{t-1}		**-0.12***** (0.04)		
lurate long-run multiplier	-0.18	0.06	-0.03	--
Wald χ^2 **$Pr>\chi^2$**	4.39 0.04	2.92 0.09	4.50 0.03	
predet.	lurate	lurate	--	lurate
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 **$Pr>\chi^2$**	14.97 1.00	6.07 1.00	1.02 0.31	33.27 0.69

Note: 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 – H_0: The long-run multiplier calculated from the individually significant coefficients is insignificant. Arellano-Bond test on average order 1 autocovariance in residuals (AR1 errors) – H_0: The residuals are not autocorrelated. Arellano-Bond test on avareage order 1 autocovariance in residuals (AR1 errors) – H_0: The residuals are not autocorrelated. Sargan test of over-identifying restrictions (results from 2-step GMM with standard errors not corrected for heteroskedasticity) - $H_{0:}$ The over-identifying restrictions are valid.

Table 6: Estimation Results: 1-step GMM, Phillips Curve Specification

dependent variable: lrwage	Bulgaria 1991-99	Hungary 1992-99	Poland 1992-98	Romania 1992-99
$lrwage_{t-1}$	**0.65***** (0.08)	**0.55***** (0.07)	**0.25*** (0.14)	0.11 (0.12)
$lurate_t$	**-0.12***** (0.04)	-0.005 (0.02)		0.007 (0.02)
$lurate_{t-1}$	0.05 (0.04)	**-0.05*** (0.03)	**-0.03**** (0.02)	
$lurate_{t-2}$	**0.05**** (0.02)	-0.03 (0.03)		
$lurate_{t-3}$		**0.02*** (0.01)		
s_empl1_t		0.40 (0.25)		**-0.59**** (0.27)
s_empl1_{t-1}			**0.13***** (0.04)	
s_empl1_{t-2}		**0.32*** (0.18)		
s_empl2_t		**0.18***** (0.56)	**-0.09***** (0.03)	**-1.14***** (0.19)
s_empl2_{t-1}		**-0.13***** (0.04)		
lurate 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.	lurate	lurate	--	lurate
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

Note: 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 5.

Table 7: Spatial Autocorrelation in the Variables

	lrwage		lurate		s_empl1		s_empl2	
	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

LM test on spatial autocorrelation in the variables – H_0: The variable is not spatially autocorrelated.

Table 8: Estimation Results: 1-step GMM, Spatially Filtered Variables

dependent variable: lrwage	Bulgaria 1991-99	Hungary 1992-99	Poland 1992-98	Romania 1992-99
lrwage$_{t-1}$	**0.67***** (0.13)	**0.48***** (0.08)	**0.35*** (0.20)	0.08 (0.14)
lrwage$_{t-2}$	**-0.27***** (0.07)		**-0.25***** (0.08)	**-0.15**** (0.07)
lurate$_t$	**-0.12**** (0.06)	-0.003 (0.01)		0.01 (0.02)
lurate$_{t-1}$		-0.04 (0.03)	**-0.04**** (0.02)	
lurate$_{t-2}$		**-0.05*** (0.03)		
lurate$_{t-3}$		**0.03**** (0.02)		
s_empl1$_t$				**-0.63*** (0.34)
s_empl1$_{t-1}$			**0.15***** (0.07)	
s_empl1$_{t-2}$		**0.28*** (0.16)		
s_empl2$_t$		**0.16***** (0.06)	**-0.13***** (0.04)	**-1.17***** (0.21)
s_empl2$_{t-1}$		**-0.14***** (0.04)		
lurate long-run multiplier	-0.20	-0.04	-0.04	--
Wald χ^2 **Pr>χ^2**	4.73 0.03	0.50 0.48	5.91 0.02	
predet.	lurate	lurate	--	lurate
N obs.	168	80	196	164
Wald χ^2	4203.17	12193.81	104.28	6420.19
AR1 errors z **Pr>z**	-3.26 0.001	-2.61 0.01	-1.43 0.15	-1.67 0.10
AR2 errors z Pr>z	0.62 0.54	1.29 0.20	1.40 0.16	-0.27 0.78
Sargan χ^2 **Pr>χ^2**	20.70 1.00	9.31 1.00	7.30 0.64	

Note: 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 5.