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Regional Wage Adjustments and Unemployment: Estimating the Time-Varying Wage Curve

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1. Introduction

The relationship between unemployment and wages is one of the key equilibrating channels in the economy. The strength of this relationship is determined, among other things, by the legal framework of bargaining between employees and employers, the institutional setting of the welfare system and the mobility of labour. Blanchflower and Oswald (1994) enriched the academic discussion on the relationship between unemployment and wages by empirically analysing the effect of the regional unemployment rate on the level of individual or region-average wages. They found that the local unemployment elasticity of real wages is about –0.1 for a number of developed and developing countries. This evidence, called the wage curve, is widely used as an indicator of real wage flexibility at the regional level.[1]

The wage curve evidence is less straightforward in the economies of Central Europe (*Table 1*). The regional elasticity of wages in the Czech Republic differs markedly from Slovakia and Hungary. Existing studies, such as Blanchflower (2001) and Huitfeldt (2001), find the elasticity in the Czech Republic to be between 0.00 and –0.04. A low elasticity estimate is also reported for Austria by Winter-Ebmer (1996). Given that estimation methods are appropriate in these studies, this evidence points to drawbacks emanating from the bargaining between employers or employees or from the welfare system. In particular, the wage formation might not reflect local labour market conditions, or it might be the welfare scheme that restrains wage adjustments. Galuščák and Münich (2003) provide evidence that the problem of low elasticity estimates is rather in the estimation procedure itself and in the heterogeneity of the wage curve relationship. Although the baseline elasticity estimate is –0.03, the estimate is greater in absolute value after controlling for the endogeneity of unemployment and in particular groups of districts. In any case, all these estimates are still below the benchmark specified in Blanchflower and Oswald (1994).

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[1] See, for example, (Nickell, 1997).

TABLE 1 Previous Wage Curve Estimates

Country	Period	Coefficient log(u)[a]	Data level	Estimation method	Number of regions	Authors
Czech Republic	1992, 1994–1997	–0.21 (4.35)	individual	OLS	8	Blanchflower (2001)
Czech Republic	1992, 1994–1997	0.03 (0.41)	individual	OLS, regional dummies	8	Blanchflower (2001)
Czech Republic	1992–1998	–0.013 (1.86)	district	OLS	73	Huitfeldt (2001)
Czech Republic	1992–1998	–0.042 (5.25)	district	OLS, district fixed effects	73	Huitfeldt (2001)
Czech Republic	1996–2001	–0.032 (2.0)	district	OLS, district fixed effects, instruments	74	Galuščák, Münich (2003)
Czech Republic	1996–2001	–0.079 (2.2)	district	OLS, district fixed effects, instruments	45[b]	Galuščák, Münich (2003)
Slovakia	1995	–0.049 (3.85)	individual	OLS	42	Blanchflower (2001)
Slovakia	1992–1998	–0.128 (9.1)	district	OLS	29	Huitfeldt (2001)
Slovakia	1992–1998	–0.109 (7.3)	district	OLS, district fixed effects	29	Huitfeldt (2001)
Hungary	1990–1997	–0.052 (2.07)	individual	OLS	20	Blanchflower (2001)
Hungary	1990–1997	0.042 (1.23)	individual	OLS, regional dummies	20	Blanchflower (2001)
Hungary	1994–1995	–0.11	individual	OLS, regional dummies	14	Kertesi, Köllö (1997)
Poland	1991–1997	–0.153 (3.59)	individual	OLS	8	Blanchflower (2001)
Poland	1991–1997	–0.127 (2.28)	individual	OLS, regional dummies	8	Blanchflower (2001)
Austria	1983	–0.029 (2.7)	individual	OLS, regional dummies	99	Winter-Ebmer (1996)

Notes: [a] *t*-statistics in parentheses
[b] without the districts that experienced the largest rise in unemployment rate between 1996 and 2001

Do the previous estimates of the regional elasticity of wages mean that Czech real wages are inflexible with respect to regional unemployment? Or alternatively, are these estimates low because of inappropriate estimation techniques? In this chapter, we adopt the methodology described in (Galuščák – Münich, 2003) with the aim of assessing the degree of real wage flexibility at the regional level by estimating the unemployment elasticity of wages. The panel of district-level data on unemployment and average wages covers the period 1993–2001, allowing us to assess how the regional flexibility has changed in this period. We improve the estimation employed in (Galuščák – Münich, 2003) by using contemporaneous instruments in order to utilise the full length of the panel. Because local structural conditions are important determinants of wage flexibility at the district level, we allow for different functional forms of the wage curve across districts.

Our prime goal is to analyse how the regional flexibility of wages changed during 1993–2001. A potential deterioration in the flexibility would have an adverse impact on the economy during periods of adverse shocks. In addition, the regional flexibility will be important as regards the prospect of losing independent monetary policy upon EMU entry.

This paper is set out as follows. The next section summarises the theory and previous evidence on the wage curve. Section 3 is devoted to the estimation method, while Sections 4 and 5 describe the data and results. The last section summarises the paper.

2. Theoretical Background

Real wage flexibility at the regional level may be investigated using equations of the following type:

$$w_{rt} = \alpha_r + \beta u_{rt} + \delta_t \qquad (1)$$

where w_{rt} is the logarithm of the average nominal wage in region r at time t and u_{rt} is the logarithm of the regional unemployment rate.[2] The region-specific (α_r) fixed effects represent differences in prices across regions, which are given by different industrial and demographic structures. On the other hand, the time-specific (δ_t) fixed effects capture differences in prices over time and contemporary aggregate shocks affecting all regions. The time-specific fixed effects contain, for example, potential effects of the economy-wide unemployment rate and the economy-wide previous wage level – see (Bell *et al.*, 2002).[3] The coefficient β is therefore the local unemployment elasticity of the real wage level. It is a measure of the regional real wage flexibility. Controlling for fixed effects, the estimation is not affected by the spurious effects observed in the case of cross-sectional estimates.

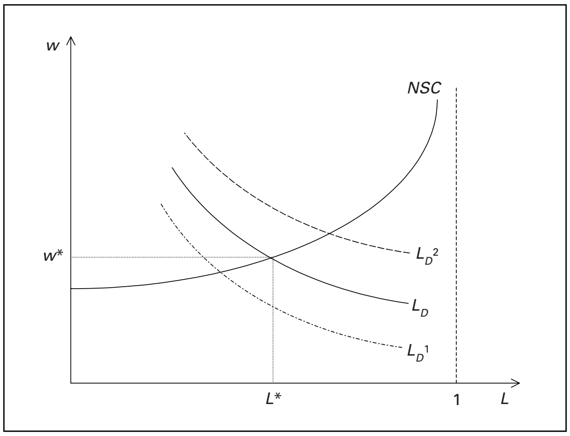
Based on the estimation of individual or regional-average wage data, the literature agrees that the estimates of β in equation (1) are negative at about –0.1. Estimates for the countries of Central and Eastern Europe have been published recently. Blanchflower (2001) estimated wage curves in 23 transition countries from Eastern and Central Europe for the period 1990–1997. For some of these countries, the estimates of β are larger in absolute terms than in developed economies, while they are lower for Hungary and the Czech Republic (Table 1). The wage curve disappears for these two countries when region-specific dummies are included.[4] Huitfeldt (2001) finds a significant wage curve using district-level panel data from the Czech Republic between 1992 and 1998. His estimates of β for district average wages are greater in absolute terms after allowing for district fixed effects, but still lower than the results obtained for Slovakia and other countries.

[2] Blanchflower and Oswald (1994) used the log-specification based on the best empirical fit.

[3] While the national rate of unemployment reflects some wages set nationally, the previous economy-wide wage level provides comparability factors in the contemporaneous wage bargaining.

[4] On the other hand, Kertesi and Köllö (1997) estimated a significant wage curve in Hungary in 1992–1995.

FIGURE 1 The No-shirking Condition and Aggregate Labour Demand



In our earlier work (Galuščák – Münich, 2003) we found a similarly low elasticity of wages in regressions pooling all districts. The elasticity is, however, greater in absolute terms for particular groups of districts, indicating that the functional form might differ across districts.[5] These results all indicate that real wage flexibility with respect to local unemployment might be low in the Czech Republic.

What size of the regional unemployment elasticity of wages is expected in a typical transitional economy? What should be expected in terms of its changes during the transition? The wage curve should be observed in an economy where at least some sectors work on the principles of supply and demand. As the proportion of market sectors increases during the transition, we might expect to observe an emerging wage curve. In other words, regional wage flexibility might increase in the course of the economic transition.

In order to interpret changes in the elasticity, we refer to the framework of efficiency wages described in (Shapiro – Stiglitz, 1984). In their stylised model, an upward-sloping no-shirking condition (NSC) is derived describing the minimum wage level needed to induce effort and prevent shirking of employees at a given rate of employment (*Figure 1*). Blanchflower and Oswald (1994) show that the wage curve is identical with the NSC and that the NSC is the same in all regions if institutional factors do not differ across regions. According to these authors, regions exhibit different unemployment-wage combinations due to non-pecuniary benefits. A region with a non-pecuniary benefit greater than in another region pays lower wages and suffers from higher unemployment in order to meet the zero migration

[5] Card (1995) pointed out that the wage curve might be different for different groups of workers. He found that younger, less educated, less unionised, male workers are more likely to have a significant unemployment elasticity of wages. Examining earnings and unemployment in ten regions of the UK between 1972 and 1995, Cameron and Muellbauer (2001) estimated a significant negative unemployment elasticity of pay in the case of manual men.

condition in the long run. In each region, the equilibrium is reached at the intersection of the NSC and a negatively sloped demand function. The labour demand curve emanates from the usual relationship of the value marginal product of labour.

The convex profile of the NSC implies that the slope of the wage curve may differ across regions if regional differences in the rate of unemployment are high. Using equation (1) to model the wage curve, one implicitly assumes that the slope of the wage curve is the same for all regions. This is depicted by the aggregate labour demand curve L_D in Figure 1.

The equation for the NSC is given as:

$$w \geq e + B + \frac{e}{q}\left(\frac{b}{1-L} + r\right) \tag{2}$$

where e is the effort of employees ($e = 0$ if shirking), B is the level of unemployment benefits, q is the probability of being fired if caught shirking, b is the exogenous risk of job termination, L is the employment rate ($1 - L = u$ is the unemployment rate), and r is the interest rate. The first two derivatives of (2) with respect to L are positive, implying that the slope of the wage curve is lower during recessionary periods resulting in a decline of the aggregate demand for labour. On the other hand, the wage curve becomes steeper during periods of high aggregate labour demand (depicted as $L_D{}^1$ and $L_D{}^2$ in Figure 1).[6]

In order to estimate the change in the flexibility over time, we estimate the wage curve by adopting a static equation (1) and using contemporaneous instruments in the estimation. We estimate the wage curve using the full sample (1993–2001) and in the early and late transition separately (1993–1997 and 1998–2001). We focus on how the shape of the wage curve has changed during the economic transition by estimating the time-varying wage curve. We expect the elasticity to have increased during the economic transition to a market economy. On the contrary, the recession of 1997–1999 might have led to a temporary deterioration in regional wage flexibility.

Wage flexibility might be heterogeneous across regions or specific groups of workers. Hence, we allow for different functional forms of the wage curve. According to (Galuščák – Münich, 2003), the wage curve estimates might be biased by including districts with the highest rise in unemployment rate between 1996 and 2001 in the sample. In these districts, the unemployment rate was prevalently low until 1996, while it increased sharply during the 1997–1999 recession to the level observed in other districts. Unemployment increased to the level observed in other districts for a given wage level, indicating that these districts experienced a "delayed" transition. To account for this, we also provide estimates of the elasticity using the sample excluding these districts.[7] We expect some improvement in the overall elasticity due to the delayed transition observed in these districts. Finally,

[6] Increases in B or r shift the NSC upward, implying a lower slope in the wage curve for a particular labour demand curve. Effects of changes in other parameters of the NSC on the slope of the wage curve are ambiguous.

we estimate the regional elasticity of wages in districts excluding Prague, Prague-East and Prague-West.[8]

3. Estimation

In the empirical analysis, we extend our deterministic specification (1) by including a component ε_{rt} that represents the variation in observed wages not explained by unemployment:

$$w_{rt} = \alpha_r + \beta u_{rt} + \delta_t + \varepsilon_{rt} \tag{3}$$

In line with equation (1), we assume that regional fixed effects capture price differences across regions and that prices grow at the same rate in all regions.[9] Fixed effects represent district- or time-specific unobserved factors contributing to the variance in observed wages. Not controlling for fixed effects, while these effects are correlated with the explanatory variables through unobserved relationships, the parameter estimates based on the standard ordinary least squares procedure would be biased.[10]

The standard approach to control for the presence of fixed effects is to transform the observations into cross-sectional unit specific mean deviations, subtracting the unit specific mean values:

$$w_{rt} - \overline{w_r} = \beta(u_{rt} - \overline{u_r}\,) + \delta_t' + \varepsilon_{rt} - \overline{\varepsilon_r} \tag{4}$$

where bars denote mean values across time periods.

District-specific fixed effects may also be removed by first differences. In particular, transforming equation (3) into first differences yields:

$$w_{rt} - w_{r,t-1} = \beta(u_{rt} - u_{r,t-1}) + \delta_t'' + \varepsilon_{rt} - \varepsilon_{r,t-1} \tag{5}$$

In the estimation of equation (4) and equation (5) we assume that unemployment is exogenous. This may not be the case, for two reasons.

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[7] The selection of districts is arbitrary. In (Galuščák – Münich, 2003) we calculated the ratio of unemployment rates in 2001 and 1996 and sorted the districts by that ratio. Using K-means cluster analysis, we created four groups of districts. Then we merged these groups into two groups of districts: a group of districts that experienced the lowest increase in unemployment rates between 1996 and 2001 (45 districts), and another group of districts with the highest increase in unemployment (32 districts, with Prague at the top). In fact, we discard almost half the sample. The aim is to demonstrate that the wage curve might be weaker in problematic regions.

[8] As capital cities have specific local labour markets. We exclude the districts of Prague-East and Prague-West as many residents in these districts commute to Prague. While they affect the unemployment statistics in their home districts, their wages are counted in the average wage in Prague.

[9] It seems implausible to assume that regional price differences are constant over the period used in the estimation. Potential changes in regional differences transform to the measurement error of the left-hand side variable in equation (3), leading to an efficiency loss of the estimates. In addition, the Czech Statistical Office does not provide district inflation rates.

[10] Not controlling for district fixed effects, we might observe a positive long-run relationship between unemployment and real wages. Such evidence would be in line with the theory of compensating differentials (Harris – Todaro, 1970). In particular, higher wages compensate for high unemployment in order to achieve the zero migration condition.

First, suppose that there is an unobservable time-varying variable that is correlated with both the unemployment rate and the wage level. For example, migrants moving to regions with low unemployment and high wages might represent such a variable. Migration into a region increases the rate of unemployment and decreases the average wage at the same time. Not controlling for migration, the error term is correlated with unemployment, leading to biased estimates of elasticity. A second source of violation of the assumption that the unemployment rate is exogenous is the time aggregation. Using annually aggregated data, annual wage rates are likely to contain the accumulated effect of unemployment. To overcome the problem of endogeneity, we have to instrument for the unemployment rate.

Appropriate instruments for unemployment are the lagged values of the unemployment rates. However, using the lagged values as instruments shortens the time dimension of the sample. In order to avoid this problem, we look for instruments among contemporaneous variables. We use the average unemployment rate in neighbouring districts as instruments. In the notation of equation (5), the difference in unemployment rates $u_{rt} - u_{r,t-1}$ might be instrumented using $u^s_{rt} - u^s_{r,t-1}$. In addition, unemployment rates are at least partly determined by inflows into unemployment. Given that i_{rt} is the logarithm of the number of inflows into unemployment, we can use the differences in inflows in neighbouring districts $i^s_{rt} - i^s_{r,t-1}$ as instruments for $u_{rt} - u_{r,t-1}$. However, we do not use both differences in unemployment rates and differences in inflows, since unemployment rates are highly correlated with lagged inflows.

Both the mean-deviation and the first-order transformation also limit the scope for possible heteroscedasticity because the transformation removes all scale effects. For example, suppose that district size is a source of heteroscedasticity. Assuming that the district labour force does not change over time, the difference of the log-variables has the same effect as multiplying the variables by the district labour force, removing the source of heteroscedasticity. It should also be noted that the difference multiplies the impact of errors in the variables.[11]

4. Data

Aggregate wage data are published regularly by the Czech Statistical Office. They come from regular reports that economic units are obliged to fill in by law. In this paper, we use monthly district averages of the annual data covering the period 1993–2001. The sample covers all employees in the public sector, but is restricted to firms with more than 25 employees in the private sector until 1996, and more than 20 employees since 1997. However, in 1995 and 1996, the sample covers firms with more than 100 employees in industry, construction and hotels and restaurants. The data refer to the location of the workplace. The first row in *Table 2* shows the average monthly earnings statistics.

[11] Note that $var(\varepsilon_{rt} - \varepsilon_{r,t-1}) = 2\ var(\varepsilon_{rt})$.

TABLE 2 Data Statistics

	1993	1994	1995	1996	1997	1998	1999	2000	2001
Average monthly earnings (CZK)	5518.1 (406.6)	6455.5 (493.3)	7602.2 (561.8)	8961.5 (674.7)	9860.8 (714.3)	10684.5 (845.5)	11484.0 (909.7)	12196.6 (931.3)	13144.2 (1009.3)
Unemployment rate (in %)	3.4 (1.6)	3.7 (1.8)	3.4 (1.7)	3.6 (1.8)	4.8 (2.1)	6.6 (2.5)	9.1 (3.2)	9.3 (3.8)	8.8 (3.9)
Short-term unempl. rate (less than 12 months, in %)	3.0 (1.3)	3.0 (1.3)	2.6 (1.2)	2.8 (1.3)	4.0 (1.5)	5.3 (1.7)	6.8 (1.8)	5.9 (1.9)	5.5 (1.7)
Long-term unempl. rate (more than 12 months, in %)	0.5 (0.3)	0.7 (0.5)	0.8 (0.6)	0.8 (0.6)	1.0 (0.7)	1.5 (1.0)	2.5 (1.5)	3.4 (2.0)	3.3 (2.2)
Number of districts	73	73	73	74	74	74	74	74	74

Note: Mean values across districts for the year, standard deviations in parentheses.
Sources: Czech Statistical Office; Ministry of Labour and Social Affairs

The district-level data on registered unemployment come from the registers of 77 district labour offices in the Czech Republic (76 until 1995) and represent detailed and standardised monthly sources of information collected for the Ministry of Labour and Social Affairs. The data include end-of-month values of stock variables and period-cumulative values of gross flows of unemployment and vacancies. Table 2 shows the mean values and standard deviations of the unemployment rate and the short-term and long-term unemployment rate defined as the number of the unemployed divided by the labour force. Short-term unemployment covers persons seeking a job for less than 12 months, while long-term unemployment includes persons registered with labour offices for more than 12 months.

The registry unemployment data are likely to underestimate the actual number of unemployed. Some people do not register with a labour office when they change jobs. Under-reporting is more likely in urban areas, where other channels of job search are used. The under-reporting is consequently likely to be uneven across districts. Assuming that the differences in the under-reporting of unemployment across districts are time-invariant, this effect is removed by the differences used in this chapter. In contrast to this problem, the registry unemployment might be over-reported since some people register with a labour office in order to be eligible for social security benefits. Again, we assume that this effect is to a great extent removed by mean-specific and first-specific differences.

5. Results

In order to illustrate the effect of different estimation techniques on the results, we first estimate the wage curve by the ordinary least squares applied to the pooled sample. The results are shown in columns 1 and 5 of *Table 3*. The elasticity is insignificant on the pooled sample and even positive (+0.03) when the districts with the highest rise in the rate of unem-

TABLE 3 Looking for the Wage Curve, 1993–2001

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
β	–0.001 (0.007)	–0.016*** (0.004)	–0.018*** (0.005)	–0.086*** (0.022)	0.033*** (0.009)	–0.032*** (0.006)	–0.030*** (0.008)	–0.094*** (0.027)
N	74	74	74	74	45	45	45	45
NT	663	663	589	589	402	402	357	357
R^2 (overall)	0.94	0.94	0.89		0.95	0.95	0.90	
R^2 (within)		1.00				1.00		
F statistics				3104***				1943***
Hausman test				17.4***				8.6***
Low unemployment rise districts					x	x	x	x
OLS	x				x			
Fixed effects		x				x		
First differences			x				x	
First differences, IV				x				x

Notes: * significant at 10 %, ** at 5 %, *** at 1 %; robust standard errors in parentheses; time dummies not reported.

Hausman test: $H0$: instrumented variable (log u) is exogenous (F statistics reported).

Estimates for all districts and for the districts with the lowest rise in unemployment rate between 1996 and 2001 (columns 5–8).

ployment between 1996 and 2001 are excluded from the sample. Not controlling for district fixed effects, we estimate a positive long-run relationship between unemployment and real wages.[12]

Accounting for district fixed effects yields estimates that are similar to the results published by other authors for the Czech Republic. We do so by estimating equations (4) and (5) allowing for the mean-specific and first--specific deviations. The elasticity is –0.02 for the full sample and –0.03 for the sample without the districts with the highest rise in unemployment (columns 2, 3, 6 and 7 in Table 3). The size of the elasticity is, however, still notably smaller than found in most other countries.

Allowing for endogeneity of the unemployment rate raises the elasticity estimate dramatically to –0.09 (see columns 4 and 8 in Table 3).[13] The results of the Hausman test do not allow us to reject the hypothesis that the unemployment rate is endogenous, implying that the instrumenting is

[12] For example, heavy industry districts exhibit higher wages due to wage differentials compensating for less favourable work conditions. These districts also exhibit persistently higher unemployment rates due to a lower educational level of the labour force and occasionally due to lasting restructuring. Not controlling for district-specific effects, the estimate of β is biased upward. Given that the actual effect is negative, we underestimate the actual wage-curve effect. A strong bias even leads to a positive coefficient β.

[13] We instrument the unemployment rate using average unemployment rates in neighbouring districts. Alternatively, using the inflow rate in neighbouring districts defined as the ratio of inflows and labour force as instruments for the unemployment rate yields the estimates –0.10 and –0.11. In order to account for possible effects associated with different wage methodologies in 1995 and 1996 (see Section 4), we repeated the estimation with the sample excluding the years 1995 and 1996. The results are similar to those in Table 3.

TABLE 4 2SLS Estimates of the Wage Curve, 1993–2001

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
β	–0.086*** (0.022)	–0.090*** (0.017)	–0.094*** (0.027)	–0.105*** (0.026)	–0.099*** (0.034)		
β*d		0.017 (0.041)		0.033*** (0.010)	0.021 (0.052)		
β (short-term u)						–0.068 (0.061)	–0.135* (0.080)
β (long-term u)						–0.024 (0.053)	0.020 (0.070)
N	74	74	45	74	45	74	45
NT	589	589	357	589	357	589	357
F statistics	3104***	2738***	1943***	2776***	1705***	2693***	1364***
Hausman test	17.4***	8.9***	8.6***	8.6***	4.5**	8.8***	5.1***
Low unemployment rise districts			x		x		x
d = 1 if high unemployment rise districts				x			
d = 1 if 1998–2001		x			x		

Notes: Columns 1 and 3 are from Table 3 for comparison.
* significant at 10 %, ** at 5 %, *** at 1 %; robust standard errors in parentheses; time dummies not reported.
Hausman test: $H0$: instrumented variables are exogenous (F statistics of the joint test reported).
Estimates for all districts and for the districts with the lowest rise in unemployment rate between 1996 and 2001 (columns 3, 5, 7).

appropriate. This also implies that the findings so far reported by other studies might be biased due to inappropriate estimation techniques or data imperfections. Our results indicate that, contrary to previous studies, the elasticity of real wages might be at the level observed in other countries and reported by Blanchflower and Oswald (1994).

We explore whether excluding the districts with the highest increase in the unemployment rate between 1996 and 2001 has a significant effect on the elasticity estimate. For this purpose, we estimate the first-specific deviation equation (5) with an additional term on the right-hand side representing the unemployment rate interacted with a dummy variable. The dummy equals one for the districts with the highest rise in unemployment and zero otherwise. While column 3 in *Table 4* is the same as the last column in Table 3, column 4 in Table 4 reports estimates of equation (5) with the added term. The estimate of the added term is significant and positive, indicating that the wage curve does not have an identical functional form across all districts.

Although all the results in Table 3 indicate that excluding the districts with the highest rise in the unemployment rate between 1996 and 2001 from the sample might improve the results, the main difference is between the estimation techniques. In particular, one has to treat the unemployment rate in the wage curve relationship as endogenous for yearly district--level data.

We now turn our attention to estimating whether the regional flexibility of real wages has improved, deteriorated or remained the same during 1993–2001. Using our preferred method, we estimate the first-specific deviation equation (5) with an additional term on the right-hand side (β^*d) representing the unemployment rate interacted with a period-specific dummy variable (denoting 1998–2001 for the late transition). The regression results are shown in column 2 of Table 4 for the full sample and in column 5 for the sample with only the low unemployment rise districts. The estimates of β^*d are insignificant, suggesting that the flexibility has been roughly the same throughout the period 1993–2001. The positive estimate of the coefficient β^*d, although insignificant, is an indication of a possible deterioration in the flexibility during 1993–2001.[14] This result is striking, as we could expect that ongoing restructuring should increase the flexibility.

We explore the possible sources of weakening flexibility. We estimate equation (5) with separate terms representing the short-term and long-term unemployment rates. The estimates are shown in columns 6 and 7 of Table 4. For the districts with the lowest rise in unemployment, the elasticity of the short-term unemployed is –0.14, while it is insignificant (+0.02) for the long-term unemployed. The wage curve is observed for the short-term unemployed, while the long-term unemployed do not affect wage formation.[15] Although the short-term unemployed have a high elasticity of real wages, the elasticity is lower in absolute value at the economy-wide level due to the incidence of long-term unemployment.

In order to investigate further how the shape of the wage curve has changed during the economic transition, we split the sample into early and late transition and repeat the estimation. Splitting the sample allows the district-specific fixed effects to differ between the two periods. The results are reported in *Table 5*. The elasticity has worsened between the early and late transition (compare columns 1–2 and 3–4) by the same size as the estimates of (β^*d) in Table 4. Hence, allowing district-specific fixed effects to be different between the early and late transition does not change the results. Estimating the effects of short-term and long-term unemployment rates indicates that while the elasticity for the short-term unemployed might have improved between the early and late transition, the elasticity for the long-term unemployed deteriorated during the same period (columns 5–6 and 7–8 in Table 5).

In order to disentangle the effects of the 1997–1999 recession on the wage curve, we estimate equation (5) separately by 2-year intervals. The estimates of the time-varying wage curve are shown in *Table 6*. The elasticity is –0.11 in 1994–1995 and –0.13 in 1995–1996. It decreased markedly to

[14] This result seems to be robust to the choice of the break point as well as to the length of the sample.

[15] This evidence confirms our previous results (Galuščák – Münich, 2003). High incidences of long-term unemployment are observed in districts with high unemployment. These districts are, for example, mining or heavy industry districts with both high unemployment and high wages. Wages do not adjust downward in these districts because the welfare system leads to higher reservation wages.

TABLE 5 The Wage Curve in the Early and Late Transition

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
β	–0.090*** (0.027)	-0.073** (0.030)	–0.099*** (0.034)	–0.079*** (0.039)				
β (short-term u)					–0.042 (0.083)	–0.167 (0.154)	–0.139 (0.180)	–0.144** (0.070)
β (long-term u)					–0.048 (0.064)	0.092 (0.167)	0.011 (0.123)	0.086 (0.103)
N	74	74	45	45	74	74	45	45
NT	293	296	177	180	293	296	177	180
F statistics	4148***	1328***	2392***	1018***	3096***	673***	1544***	523***
Hausman test	14.8***	3.0*	8.3***	0.6	7.3***	2.6*	5.2***	2.5*
Low unemployment rise districts			x	x			x	x
Early transition, 1994–1997	x		x		x		x	
Late transition, 1998–2001		x		x		x		x

Notes: 2SLS estimates
* significant at 10 %, ** at 5 %, *** at 1 %; robust standard errors in parentheses; time dummies not reported.
Hausman test: $H0$: instrumented variables are exogenous (F statistics of the joint test reported).
Estimates for all districts and for the districts with the lowest rise in unemployment rate between 1996 and 2001 (columns 3–4, 7–8).

TABLE 6 The Time-varying Wage Curve

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
β	–0.105*** (0.033)	–0.125** (0.043)	–0.052 (0.047)	–0.017 (0.033)	–0.058 (0.060)	–0.114*** (0.040)	–0.081*** (0.032)	–0.112*** (0.033)
$\beta^* d_1$								0.095** (0.047)
$\beta^* d_2$								0.029 (0.048)
N	73	73	74	74	74	74	74	74
NT	146	146	147	148	148	148	148	589
F statistics	3085***	3443***	4551***	1741***	1149***	874***	1166***	2327***
1994–2001								x
1994–1995	x							
1995–1996		x						
1996–1997			x					
1997–1998				x				
1998–1999					x			
1999–2000						x		
2000–2001							x	

Notes: 2SLS estimates
d_1 = 1 if 1997–1998
d_2 = 1 if 1999–2001
* significant at 10 %, ** at 5 %, *** at 1 %; robust standard errors in parentheses; time dummies not reported.

–0.02 during the recession (1997–1998), but fell back after the recession (columns 6 and 7), although probably not to the level observed before the recession. In order to estimate how the recession affected the elasticity, we repeated the estimation using the whole sample with two dummy variables denoting the recession (1997–1998) and the after-recession period (1999–2001). The results in the last column of Table 6 indicate that the elasticity was –0.11 in 1994–1996 and that the deterioration was significant during the recession. The elasticity is insignificantly higher in the period after than before the recession.

The results indicate that the regional flexibility was high between 1994 and 1996 at the level observed in most developed and developing countries.[16] The significant deterioration in the flexibility during the recession of 1997–1999 is consistent with the predictions of the efficiency wage model. After 1999, the flexibility might not have returned to the level observed before the recession, probably due to hysteresis effects in the Czech labour market. The sharp rise in long-term unemployment at the end of the 1990s (see Table 2) is, therefore, the prime suspect for why the regional real wage flexibility might not have recovered to the pre-recession level.

6. Conclusion

In this paper we estimate the degree of regional real wage flexibility. We estimate the elasticity of wages using a static version of the relationship between district-level unemployment rates and district-level wages. We build on the methodology described in our earlier work (Galuščák – Münich, 2003). We estimate the wage curve using estimation methods accounting for endogeneities and show that previous estimates could have been low due to inappropriate estimation techniques. The elasticity is about –0.1, just at the level reported by Blanchflower and Oswald (1994) for a number of developed and developing countries. In contrast to previous studies, this result indicates that Czech real wages are flexible with respect to local unemployment rates. Furthermore, we show that the wage flexibility might not be homogenous across districts. We show that changes in the shape of the wage curve observed during the 1997–1999 recession are in line with the standard efficiency wage model.

In accordance with Galuščák and Münich (2003), the elasticity is significantly greater in absolute value after excluding the districts that experienced delayed restructuring. Prior to the recession, the wage elasticity was high in all districts. Our results show that the degree of wage flexibility has not changed significantly between the early and late transition. A significant deterioration in flexibility is observed during the recession of 1997–1999, which we explain using the efficiency wage model. The degree of regional flexibility might not have returned to its pre-recession level. We associate that observation with hysteresis effects in the Czech labour market.

[16] It seems that the elasticity was at the same level in all districts. Repeating the estimation in column 8 of Table 6 for the districts which experienced the highest rise in unemployment, the results indicate that the elasticity was –0.117 prior to the recession. Consequently, these districts were more heavily affected by the recession.

We provide some evidence that the sharp rise in long-term unemployment observed at the end of the 1990s might have weakened the regional flexibility of real wages. Given that long-term unemployment will probably continue rising, the flexibility of real wages at the regional level will deteriorate, meaning the loss of an important equilibrating channel in the economy when facing negative shocks, particularly after EMU entry. This is an important issue to be incorporated into labour market policies still relying only on information on growth of aggregate wages and unemployment. As a result, this chapter delivers additional evidence on deteriorating labour market performance in the Czech Republic, a message which is consistent with the findings of other papers included in this issue.

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SUMMARY

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Regional Wage Adjustments and Unemployment: Estimating the Time-Varying Wage Curve

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This paper investigates the flexibility of real wages at the regional level by estimating the wage curve, the relationship between regional unemployment, and the regional level of wages. For this purpose the authors use a sample of annual district-level unemployment and wage data in the Czech Republic from 1993 to 2001. Previous estimates of the wage curve for the Czech Republic suggested that the regional flexibility of real wages is extraordinarily low. Taking into account the endogeneity of unemployment, the results indicate that regional real wages are flexible at the level observed in most developed and developing economies. The temporary deterioration in the regional flexibility observed during the 1997–99 Czech recession is explained by the standard efficiency wage model. Some indication of weakening elasticity since the end of the 1990s is probably associated with the sharp rise in the incidence of long-term unemployment. As this trend is expected to continue, it could further attenuate the elasticity and complicate adjustment processes if adverse shocks appear in the future, particularly after the Czech Republic's anticipated entry into the European Monetary Union.