2003

WORKING PAPER SERIES /1/

CNB CZECH NATIONAL BANK

WORKING PAPER SERIES

Microfoundations of the Wage Inflation in the Czech Republic

Kamil Galuščák
Daniel Münich

1/2003

Reviewed by: Michaela Erbenová (Czech National Bank)
Randall Filer (CERGE-EI, Prague, and City University of New York)
Miroslav Singer (PricewaterhouseCoopers)

Project Coordinator: Vladislav Flek

Microfoundations of the Wage Inflation in the Czech Republic

Kamil Galuščák and Daniel Münich*

Abstract

We investigate whether microfoundations might increase the predictive power of macroeconomic models of wage inflation. By comparing past predictions to observed values, we find that the Phillips curve with the average unemployment rate in districts with prevalently low unemployment rates delivers more accurate predictions of aggregated wage inflation than the Phillips curve with the overall unemployment rate. The identification of specific groups of districts is based on our estimates of the wage curve at the regional level, i.e. the relationship between the regional level of wages and regional unemployment. Real wages adjust to changes in local unemployment in districts with low unemployment rates, a low share of public sector employment, and for the short-term unemployed. On the other hand, the welfare system might represent a floor preventing downward wage adjustments in districts with high unemployment rates and for the long-term unemployed. In the public sector, wages are negotiated at the economy-wide level, while the variance in regional unemployment does not play a role.

JEL Codes: E24, J64, J31, C23.
Keywords: Panel data, partial adjustment model, Phillips curve, unemployment, wage curve.

* Kamil Galuščák, Czech National Bank, Na příkopě 28, 115 03 Praha, Czech Republic, e-mail: kamil.galuscak@cnb.cz. Daniel Münich, CERGE-EI, POB 882, Politických vězňů 7, Praha, Czech Republic, e-mail: daniel.munich@cerge-ei.cz. CERGE-EI is a joint workplace of the Center for Economic Research and Graduate Education, Charles University, and the Economics Institute of the Academy of Sciences of the Czech Republic.

The paper was written within the framework of Czech National Bank Research Project No. 18/2001: "Using the Wage Curve Approach in the Analysis of Wage Inflation." The research of Daniel Münich was partly supported by grant No. P98-1129-R from the Phare ACE Program. We thank Michaela Erbenová, Randall Filer and Miroslav Singer for valuable comments. We also benefited from discussions with Vladislav Flek and Štěpán Jurajda. Our thanks go to the Department of Analysis and Statistics at the Ministry of Labour and Social Affairs for providing us the data on registered unemployment. We are responsible for all remaining errors and omissions.

Nontechnical summary

In this paper we estimate the regional unemployment elasticity of wages (the wage curve) for particular groups of districts and for specific unemployment rates in order to analyse the microfoundations of the relationship between unemployment and wages. The prime question is whether the microfoundations of wage inflation increase the predictive power of the macroeconomic models of wage inflation developed in the Czech National Bank. For this purpose, we propose a partial adjustment model to decompose the short-term and long-term effects in the relationship between regional unemployment and wages. We estimate the unemployment elasticity of wages using the average monthly earnings and registered unemployment rates in Czech districts in 1996–2001.

The results indicate that the short-term elasticity is negative and significant. The size of the elasticity is close to the estimates reported in the literature for other countries, but differs from what has been published for the Czech Republic. It seems that the wage curve does not have the same shape in all districts. In particular, one has to exclude the districts that experienced the highest increase in unemployment rates during the recession of 1997–1999 in order to obtain a significant estimate of the unemployment elasticity of wages. Between 1996 and 2001, the most pronounced rise in unemployment rates was observed in districts with prevalently low unemployment in 1996. In these districts, unemployment rates increased to the level observed in other districts. While wages also rose in these districts during the same period, this leads to bias in the elasticity estimates when all districts are included in the sample.

Our results indicate that the wage curve does not exist, or is weak, in districts with more long-term unemployed persons and in districts with high unemployment. Wages do not adjust to changes in local unemployment in these districts, owing, for example, to a relatively generous welfare system. In addition, increasing welfare benefits might increase the group of districts where wages do not adjust to changes in local unemployment. The unemployment elasticity of wages might also depend on the share of public sector employment and on the cumulative value of foreign direct investment.

We investigate whether these findings might increase the predictive power of macroeconomic models of wage inflation. For this purpose, we estimate the aggregated Phillips curve within a shorter sample. Using coefficient estimates, we compute out-of-the sample predictions and compare these predictions to observed values. We find that the average unemployment rates calculated for districts with prevalently low unemployment deliver more accurate predictions of aggregated wage inflation than the overall unemployment rate.

1. Introduction

Wage inflation is linked with unemployment in the concept of the Phillips curve. This relationship has robust implications for macroeconomic analysis and policy. It allows one to predict wage dynamics from unemployment. The relationship between wage growth and unemployment represents an important equilibrating channel in the economy. The scale and scope of the relationship emanates primarily from the micro-level, including, among other things, the form of the welfare system, the legal framework of bargaining between employees and employers, and the mobility of labour. Understanding the strength and the dynamics of this relationship is an important component of the overall empirical background necessary to build mutually consistent policies at the macro and micro level.

The relevance and validity of the macroeconomic policy implications of the conventionally used Phillips curve can be tested at the regional level. The standard approach for the analysis of unemployment and wages at the micro-level is called the wage curve. This is the relationship between the level of individual wages and regional unemployment, investigated, for example, in Blanchflower and Oswald (1994). The regional aggregation of individual wages allows them to examine wage inflation at the regional level by including a lagged wage level among the explanatory variables. The estimate of the lagged wage coefficient is the test of the Phillips curve at the regional level. Blanchflower and Oswald (1994) interpret their virtually zero estimate of the lagged wage as proof that the Phillips curve is an incorrectly specified wage curve. Their evidence has been a source of many disputes in the economic literature (Card, 1995, Blanchard and Katz, 1997, 1999). On the other hand, the literature agrees on the robustness of the size and statistical significance of the unemployment elasticity of wages across countries.

The time dimension of the panel data is short in the transition countries of Central and Eastern Europe, including the Czech Republic, indicating that the test of the Phillips curve is not robust in this case. On the contrary, the estimates of the unemployment elasticity of wages have implications for wage adjustments over the business cycle. In particular, the size of the elasticity is interpreted as the degree of wage flexibility with respect to changes in regional unemployment.

The aim of this paper is to estimate the regional unemployment elasticity of wages for particular groups of districts and for specific unemployment rates in order to analyse the microfoundations of the relationship between unemployment and wages. The prime question is whether the microfoundations of wage inflation increase the predictive power of the macroeconomic models of wage inflation developed in the Czech National Bank.

We propose a partial adjustment model to decompose the short-term and long-term effects in the relationship between regional unemployment and wages. We control for fixed effects by cross-sectional specific first differences. We use the average monthly earnings and registered unemployment rates in 1996–2001 to estimate the unemployment elasticity of wages using two-stage least squares. Using past predictions computed using coefficients of the macroeconomic model of wage inflation estimated within a shorter sample and comparing these predictions to the observed values, we test the predictive power of the model. The model explains the economy-wide wage inflation using specific unemployment rates.

The paper is organised as follows. The next section outlines the basic specifications and summarises the literature related to this topic. Section 3 introduces the partial adjustment model. The estimation method and the data are discussed in Sections 4 and 5. Section 6 reports the results, while the last two sections discuss the policy implications and give concluding remarks.

2. Background

The traditional Phillips curve describes the relationship between wages and unemployment at the macroeconomic level. In particular, the expected growth rate of real wages is estimated against the rate of unemployment using equations such that

$$w_t - p_t^e = \alpha + \beta u_t + w_{t-1} - p_{t-1}, \tag{1}$$

where w_t and p_t are logarithms of the wage and price levels and u_t is the unemployment rate. This equation describes deviations from the equilibrium represented by a non-accelerating wage rate of unemployment (NAWRU). If the unemployment rate equals the NAWRU, there is zero growth in real wages. Real wages increase if the unemployment rate is lower than the NAWRU, while they decrease if the rate of unemployment is greater than the NAWRU.

Assuming that the expected inflation in this period is equal to the last period's inflation, $p_t^e - p_{t-1} = p_{t-1} - p_{t-2}$, equation (1) transforms to

$$\Delta w_t = \alpha + \beta u_t + \Delta p_{t-1}. \tag{2}$$

Coefficient estimates of equation (2) have robust implications for macroeconomic analysis and policy. Using particular values of unemployment rates, they allow one to calculate forecasts of wage dynamics, and, given the assumption on the mark-up between wage growth and inflation, price inflation. Assuming that $\Delta w_t - \Delta p_{t-1} = 0$ in equation (2), the NAWRU is derived as $u^* = -\alpha/\beta$.

The validity of the aggregated Phillips curve of equation (2) can be tested at the microeconomic level. In particular, is it appropriate to say that the expected growth rate of regional or individuals' real wages depends on the regional unemployment rate? If not, then the concept of the aggregated Phillips curve is put in question. The real wage dynamics at the regional level can be explored using equations of the following type:

$$w_{rt} = \alpha_r + \beta u_{rt} + \gamma w_{r,t-1} + \delta_t, \tag{3}$$

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.[1] The region-specific (α_r) and time-specific (δ_t) fixed effects capture differences in prices across regions and changes in prices over time. Region-specific price differentials represent persistent differentials given by different industrial and demographic structures that change very little over time. Time-specific effects capture

[1] Blanchflower and Oswald (1994) used the log-specification based on the best empirical fit. Other studies estimate equations of both types with the unemployment rate and the log unemployment rate to provide a comparison. In this paper, we argue that the log-specification is more appropriate in the estimation of equation (3).

contemporary aggregate shocks affecting all regions in the same way. Controlling for fixed effects, we are able to estimate the actual elasticity of real wages and the real wage dynamics. The estimation is not affected by the spurious effects observed in the case of simple cross-sectional estimates. In particular, the coefficient β is the local unemployment elasticity of the real wage level, while the estimate of γ is the test of the Phillips curve at the regional level. If the estimate of γ is near zero, the Phillips curve does not exist at the regional level and the concept of the aggregated Phillips curve might be incorrect.

Blanchflower and Oswald (1994) found that the unemployment elasticity of pay is about –0.1 for a number of developed and developing economies. The empirical evidence is called the wage curve. Its estimation is based on repeated cross-sections of individual data matched with regional rates of unemployment using regressions of the following form:

$$w_{irt} = \alpha_r + \beta u_{rt} + \rho X_{irt} + \delta_t, \tag{4}$$

where w_{irt} is the logarithm of the nominal monthly earnings of individual i living in region r at time t, and X_{irt} is a vector controlling for personal characteristics.

Blanchflower and Oswald's estimates of γ in equation (3) were all less than 0.3. Based on these results, they concluded that the macroeconomic Phillips curve of the form (1) is an incorrectly specified aggregate wage curve. This means that the macroeconomic implications based on estimates of equation (2) are incorrect.

Many authors argue that Blanchflower and Oswald's estimates of γ are invalid (Card, 1995, Blanchard and Katz, 1997, 1999). In particular, the estimates based on monthly earnings mix the information on changes in wage rates with changes in hours of work. Based on individual-level hourly wage regressions, Blanchard and Katz (1997) report estimates of γ that are close to unity.[2] These results are interpreted as a good empirical fit of the Phillips curve.

Blanchflower and Oswald's estimates rely on time fixed effects to capture aggregated variables. Blanchard and Katz (1999) argue that this assumption cannot give us a reliable estimate of γ. In equation (3), no inter-regional mobility is assumed. Blanchard and Katz argue that if one relaxes this assumption, the regional wage depends not only on the lagged regional wage, but also on the aggregate wage. The lagged aggregate wage effect is hidden in the time fixed effects in equation (3), leading to a downward bias in the estimates of γ. Although inter-regional mobility is high in the United States, it is persistently low in Europe and particularly in the Czech Republic.

The literature agrees on the negative estimates of β in equation (3) and (4). Blackaby and Hunt (1992) test whether there is a differential effect of the long-term unemployed in the wage curve in the UK. Their results indicate that there is a wage curve for the short-term unemployed. This is in accordance with other literature examining the role of the long-term unemployed on wage determination, but not with Blanchflower and Oswald (1994), who found that the addition of the long-term unemployment rate explains nothing in the wage curve.

[2] Blanchard and Katz estimate an individual hourly wage equation using regional dummies for each year. The regional dummy coefficients represent demographically adjusted wage rates that are regressed on their lagged levels and the regional unemployment rate. This two-step estimation is also applied in Jurajda (2002).

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 pay. Cameron and Muellbauer (2001) examine earnings and unemployment in ten regions of Great Britain between 1972 and 1995. They estimate a significant negative unemployment elasticity of pay for manual men.

While wages adjust to changes in local unemployment, there are regional differences in unemployment observed in many countries. These differences are determined primarily by wage rigidities with respect to local unemployment and low labour mobility. For example, McCormick (1997) analyses the relationship between regional unemployment and labour mobility in the UK. He finds that the regional unemployment differences in the UK are largely determined in the manual labour market, while little variation in unemployment across regions is observed among non-manual workers. There is little evidence that manual workers move to regions with lower unemployment rates. This provides support to the results of Decressin and Fatas (1995) that region-specific shocks lead to changes in participation in the EU region, while more migration is observed in the US.

Regional differences in unemployment might change during the business cycle. According to McCormick (1997), the range of unemployment rates increased significantly in the 1982 recession. During the recovery of 1982–1993, the unemployment differences increased further, while they declined during the 1989–1993 recession. As McCormick points out, the more prosperous regions in the South and East have become more cyclical relative to other regions. The financial liberalisation in the 1980s enabled households to increase borrowing against house equity. This phenomenon was observed particularly in the South, with its larger owner-occupied sector and higher house prices. The financial liberalisation contributed to increased relative sensitivity of demand in the South to interest rate fluctuations.

The different unemployment elasticities of wages for specific groups of workers and specific unemployment rates might be explained by labour mobility or changes in the composition of jobs due to increased supply of more educated workers or technological change. Acemoglu (1999) presents a model where firms decide what type of jobs to create and then search for workers. He shows that an increase in the proportion of skilled workers (due to increased supply of more educated workers, for example) or skill-based technical change may create a change in the composition of jobs. It is more profitable for firms to create jobs for skilled and unskilled workers separately and to screen among applicants. The demand for skilled labour increases, while demand for unskilled labour declines. The unemployment rates of both groups are higher and wage inequality rises. The approach of his paper is consistent with the empirical evidence. In particular, the wage inequality across education groups increased in the US between 1979 and 1987, while at the same time the unemployment rates of all education groups also increased. Acemoglu shows that the driving source of this evidence was a change in the composition of jobs in the US since the late 1970s. Given that both unemployment and wages increase for a particular group of workers, it leads to a positive unemployment elasticity of wages.

Wage curve 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 in equation (4) are larger in absolute terms than in developed economies, while for Hungary and Czech Republic they are lower. For these two countries, the wage curve disappears when region-specific dummies are included.[3] Blanchflower's results indicate that wage flexibility is low in Hungary and the Czech Republic. His results are not overly convincing in the case of the Czech Republic, as he used unemployment data disaggregated to only eight regions.

Huitfeld (2001) provides clear evidence on the wage curve using panel data from 77 districts in 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.

Jurajda (2002) analyses the Czech wage curve using repeated cross-sections of matched employer-employee data for 1998–2001 from firms employing more than 100 workers. His district-fixed-effect estimates of the unemployment elasticity of wages are insignificant for the most part, which he blames on short time series. On the other hand, he obtains a negative wage effect of occupational unemployment rates.

The estimation of the wage curve for the countries of Central and Eastern Europe focuses on parameter β. On the other hand, the estimates of γ are not robust since the time dimension of the data is short in transition economies. The results of this paper should be interpreted within these limits.

This paper is intended not to provide additional estimates of the unemployment elasticity of wages in line what has been published elsewhere for the Czech Republic, but to investigate first why the other estimates are lower in absolute value, or even positive. The wage curve might have a different shape in different districts. The elasticity will be estimated for specific groups of districts and for specific unemployment rates. Given that a significant wage curve exists in some districts, the average unemployment rate in these districts might be a good predictor of wage inflation in macroeconomic models.

3. Model

In this part, we develop a partial adjustment model that decomposes the relationship between unemployment and wages into short-term and long-term effects. With w_{rt} denoting the logarithm of the real wage in region r and time t and u_{rt} the logarithm of the regional unemployment rate,[4] the basic specification is

$$w_{rt} = \alpha_r + \beta u_{rt} + \gamma w_{r,t-1}. \tag{5}$$

[3] On the contrary, Kertesi and Köllö (1997) found a significant wage curve for Hungary between 1992 and 1995.

[4] A log-specification provides a better fit of the observed values. In particular, a one-percentage-point increase in the unemployment rate represents a greater change to the stock of the unemployed in a region where unemployment is low than in a region of prevalently high rates of unemployment. Accordingly, it is more appropriate to investigate ratios, rather than differences in unemployment rates, where the regional differences in unemployment are large. Blanchflower and Oswald (1994) refer to the same point by using a logarithmic specification. Furthermore, a log-specification simplifies the interpretation of the coefficients as elasticities.

In the long-run steady state, the wage reaches its equilibrium:

$$w_r^* = \frac{\alpha_r + \beta u_r^*}{1-\gamma}. \tag{6}$$

While β is the short-term elasticity of wages reflecting the contemporary impact of unemployment growth on wages, the long-term unemployment elasticity of wages is $\beta/(1-\gamma)$. If $0<\gamma<1$, the long-term impact of unemployment exceeds the short-term one.

Equation (5) has a plausible interpretation. Subtracting and adding $w_{r,t-1}$ and $\beta u_{r,t-1}$ in equation (5) and rearranging the terms yields

$$\Delta w_{rt} = \beta \Delta u_{rt} + (1-\gamma)(\frac{\alpha_r + \beta u_{r,t-1}}{1-\gamma} - w_{r,t-1}). \tag{7}$$

Noting that the first term in the parentheses represents the long-term steady-state wage level in equation (6), the whole term in the parentheses corresponds to the departure of actual wages from their long-term steady state. Equation (7) can be simplified in notation so that

$$\Delta w_{rt} = \beta \Delta u_{rt} + (1-\gamma)(w_{r,t-1}^* - w_{r,t-1}). \tag{8}$$

Equation (8) describes the partial adjustment of wages to a new equilibrium as a response to a contemporary change in unemployment. The immediate impact of unemployment appears through the short-term elasticity β, while the consecutive adjustment of wages to the new equilibrium proceeds with rate $(1-\gamma)$. The speed of adjustment is proportional to the departure of actual wages from the steady-state level.

There is no non-accelerating wage rate of unemployment (NAWRU) in equation (5). The wage changes at the same rate for a particular change in the unemployment rate. At the same time, equation (6) suggests that there is a steady state for every wage level. The NAWRU is equal to any level of the unemployment rate.[5]

The wage-curve effect might not be the same for different levels of unemployment, reflecting the specific consequences of long-term unemployment such as deterioration of skills and a willingness to accept jobs at given wages among the long-term unemployed. In order to account for these effects, some studies of the wage curve incorporate higher polynomials of unemployment rates in equation (5). In this paper, we do not use non-linearity in unemployment. Instead, we test the effects of short-term and long-term unemployment.

[5] The concept of the NAWRU could be incorporated into the model by adding a non-linear unemployment term in equation (5).

4. Estimation

In order to evaluate the effects discussed in the previous section, we need to estimate β and γ from equation (5). With w_{rt} denoting the logarithm of the nominal wage, which is equal to the logarithm of the real wage plus the logarithm of the price level p_{rt}, equation (5) becomes

$$w_{rt} = \alpha_r + \beta u_{rt} + \gamma w_{r,t-1} + p_{rt} - \gamma p_{r,t-1} + \varepsilon_{rt}. \qquad (9)$$

For the purposes of our empirical analysis, we extend our deterministic model to include a component ε_{rt} representing the variation in observed wages unexplained by unemployment. Assuming further that regional price differences are captured by regional fixed effects and that prices grow at the same rate in all regions,[6] we introduce time fixed effects δ_t:

$$w_{rt} = \alpha_r + \beta u_{rt} + \gamma w_{r,t-1} + \delta_t + \varepsilon_{rt}. \qquad (10)$$

Fixed effects represent district- or time-specific unobserved factors contributing to the variance in observed wages. Not controlling for fixed effects, these effects being correlated with the explanatory variables through unobserved relationships, the parameter estimates based on the standard ordinary least squares procedure would be biased.[7]

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. However, the mean difference transformation is not appropriate in the partial adjustment model with a lagged dependent variable as in equation (10). The reason is that the mean deviation introduces a correlation between the transformed lagged dependent variable and the transformed error term. This makes the estimated parameters inconsistent. The problem can be shown simply by subtracting the mean values in equation (10):

$$w_{rt} - \overline{w_r} = \beta(u_{rt} - \overline{u_r}) + \gamma(w_{r,t-1} - \overline{w_r}) + \delta_t^{'} + \varepsilon_{rt} - \overline{\varepsilon_r}, \qquad (11)$$

where bars denote mean values across time periods. A closer look at equation (11) shows that the mean of the error terms consists of ε_{rt} for all time periods t. The lagged wage $w_{r,t-1}$ contains the error term ε_{rt-1} by definition. Therefore, the transformed lagged wage and the error term are correlated.[8]

[6] It seems implausible to assume that regional differences in prices are constant. Potential changes in regional differences transform to the measurement error of the left-hand side variable in equation (10), leading to an efficiency loss of the estimates. The length of the panel used in the paper is short enough so that the problem is mitigated. Furthermore, the Czech Statistical Office does not provide district inflation rates.

[7] 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 would be biased upward. Given that the actual effect is negative, we would underestimate the actual wage-curve effect. A strong bias could even lead to a positive coefficient β.

[8] Balestra and Nerlove (1967), Maddala (1983), Nickell (1981) and Arellano and Bond (1991) provide detailed evidence on the size and sign of the biases of both estimates of γ and β. For γ>0, the bias in estimated γ is negative and decreases with the number of available time periods. The bias does not approach zero when γ=0, and it is larger if other right-hand-side variables are included. Moreover, the bias in β is positive if other right-hand-side variables are positively related to the lagged wage.

A possible way of dealing with the correlation between the explanatory variables and the error term is to instrument the endogenous explanatory variables with proper instruments. By proper instruments we mean such variables which predict the instrumented explanatory variables and which are not correlated with the error term. The mean deviation illustrated in equation (11) does not leave the proper instruments among the set of lagged explanatory variables. A solution to this problem is to transform equation (10) into first differences:

$$w_{rt} - w_{r,t-1} = \beta(u_{rt} - u_{r,t-1}) + \gamma(w_{r,t-1} - w_{r,t-2}) + \delta_t^{"} + \varepsilon_{rt} - \varepsilon_{r,t-1}. \qquad (12)$$

The transformed wage on the right-hand side of equation (12) is still correlated with the transformed error term, but if there is no serial correlation in the error term, further lags of $w_{r,t-1}$–$w_{r,t-2}$ are independent of the error term.

The first difference leaves further lags of wages, $w_{r,t-2}$–$w_{r,t-3}$, etc. and also more exogenous variables available as proper instruments. For example, a change in vacancies reflects the dynamics of the demand for labour. Since changes in labour demand are reflected in the wage dynamics, wages are at least partly determined by vacancies.[9] Therefore, lagged vacancies $v_{r,t-1}$–$v_{r,t-2}$, $v_{r,t-2}$–$v_{r,t-3}$, etc. of the logarithm of the number of vacancies v_{rt} could be considered instruments for $w_{r,t-1}$–$w_{r,t-2}$ in equation (12).

There are other important issues to be reflected in the empirical analysis. We may consider the case where the unemployment rate is not exogenous in equation (12). Suppose that there is an unobservable time-varying variable that is correlated with both the unemployment rate and the wage level. This variable might be represented by migrants moving to regions of low unemployment rates and high wages. Migration into a region increases the rate of unemployment and decreases the average wage. Another source of violation of the assumption that the unemployment rate is exogenous in equation (12) is the time aggregation. If we use annually aggregated data in the estimation of equation (12), wage rates are likely to affect unemployment. Under such circumstances, we have to find instruments for the rate of unemployment which are not correlated with unexplained wage components.

The natural instruments for unemployment are the lagged values of the unemployment rates. In the language of equation (12), the difference in unemployment rates $u_{r,t}$–$u_{r,t-1}$ might be instrumented using $u_{r,t-1}$–$u_{r,t-2}$, $u_{r,t-2}$–$u_{r,t-3}$, etc. In addition, unemployment rates are likely to be partly predetermined by inflows into unemployment. Given that i_{rt} is the logarithm of the number of inflows into unemployment, we can construct instruments such as $i_{r,t}$–$i_{r,t-1}$, $i_{r,t-1}$–$i_{r,t-2}$, etc. However, we do not use both lagged differences in unemployment rates and lagged differences in inflows, since unemployment rates are highly correlated with lagged inflows. Given that lagged unemployment rates might also have an effect on wages, differences in inflows might be better instruments than lagged unemployment rates.

The first-order transformation limits the scope for possible heteroskedasticity because the transformation removes scale effects. Suppose, for example, that the district size is a source of heteroskedasticity. Assuming that the district labour force does not change over time, the first difference of the log-variables has the same effect as multiplying the variables by the district

[9] See also Jackman (1990).

labour force, removing the source of heteroskedasticity. It should also be noted that the first difference multiplies the impact of errors in the variables, if there are any.[10] This results in higher standard errors in the estimated parameters compared to the results from ordinary least squares.

5. Data

The aggregate wage data published regularly by the Czech Statistical Office 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 1996–2001. The sample covers all employees in the public sector, but is restricted to firms with more than 20 employees in the private sector. The first row in Table 1 provides the average monthly earnings statistics.

Table 1: Data statistics

	1996	1997	1998	1999	2000	2001
Average monthly earnings (CZK)	9031.7 (790.2)	9954.2 (879.5)	10801.2 (1059.3)	11625.1 (1183.2)	12360.0 (1275.9)	13062.7 (1297.7)
Unemployment rate (%)	3.44 (1.87)	4.67 (2.22)	6.42 (2.69)	8.82 (3.33)	9.08 (3.88)	8.56 (3.95)
Short-term unemployment rate (%)						
less than 6 months	2.17 (1.01)	3.02 (1.18)	3.93 (1.23)	4.62 (1.19)	3.94 (1.16)	3.61 (1.05)
less than 12 months	2.73 (1.33)	3.86 (1.60)	5.17 (1.81)	6.60 (1.96)	5.74 (1.94)	5.19 (1.81)
Long-term unemployment rate (%)						
more than 6 months	1.33 (0.93)	1.79 (1.16)	2.71 (1.59)	4.37 (2.27)	5.06 (2.84)	4.79 (2.96)
more than 12 months	0.77 (0.60)	0.95 (0.73)	1.47 (1.01)	2.39 (1.51)	3.25 (2.05)	3.21 (2.21)
Inflows into unemployment (% of the labour force)	0.66 (0.27)	0.88 (0.31)	1.09 (0.28)	1.21 (0.26)	1.11 (0.26)	1.09 (0.25)
Vacancies (% of the labour force)	2.01 (0.70)	1.63 (0.66)	1.17 (0.46)	0.77 (0.36)	1.00 (0.46)	1.22 (0.58)
Public sector employment share (%)	22.5 (3.4)				21.3 (3.4)	

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

The district-level data on registered unemployment come from the registers of all 77 district labour offices in the Czech Republic and represent detailed and standardised monthly sources of information collected for the Ministry of Labour. The data include end-of-month values of stock variables and period-cumulative values of gross flows of unemployment and vacancies. Table 1 shows the mean values and standard deviations of the unemployment rate, the short-term and

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

long-term unemployment rates, inflows into unemployment, and the vacancy rate in the years 1996 to 2001.[11]

The characteristics of Czech districts have been published by the Czech Statistical Office every year since 1994. The data include basic socio-economic district-level characteristics and information on housing, employment, industries, the environment, and crime statistics. The last row of Table 1 displays the shares of employment in the public sector in 1996 and 2000.

The district average monthly wages published by the Czech Statistical Office are reported for a sample of firms and branches of firms located in each particular district between 1996 and 2000. However, in 2001, the earnings in branches are reported only in districts where headquarters are located. For example, if a firm based in Prague has local branches in other districts, earnings in these branches contribute to the average earnings in Prague in 2001. This effect might influence the average wage in cities with many corporate headquarters as well as the average wage in other districts where branches are located. Table 2 shows growth rates of average monthly earnings in Prague, Prague-East, Prague-West, and other cities (the former regional capitals). In 1997–2000, wages grew faster in the districts of Prague than in the other districts reported in Table 2. In 2001, the wage growth declined sharply in Prague, while this is not the case for the other cities. This indicates that firms located in Prague pay wages in other districts that are lower than the average wage in Prague. The change in reporting earnings reduced the wage growth in Prague in 2001, while this effect might be neglected in the other cities. The effect might be less significant in districts other than Prague and could be reduced by excluding Prague from the sample. This is supported by the results in Table 3. The variation coefficient of the average earnings reflecting earnings differentials across all districts increased between 1996 and 2000, while it declined in 2001. Excluding Prague or Prague and other districts from the sample yields increasing earnings differentials in the whole period 1996–2001.

Table 2: Average monthly earnings growth rates (y-o-y in %)

District name	**1997**	**1998**	**1999**	**2000**	**2001**
Prague	12.2	12.8	9.8	8.2	-2.4
Prague-East	13.6	9.0	11.8	11.9	1.7
Prague-West	16.4	12.8	9.8	5.5	1.9
Ceske Budejovice	10.2	11.0	7.8	6.2	3.2
Plzen	7.1	8.6	7.9	5.0	8.1
Usti n.L.	8.7	6.8	8.1	4.7	5.6
Hradec Kralove	9.1	7.9	8.4	8.2	5.8
Brno	11.7	8.9	7.9	5.9	0.7
Ostrava	8.8	8.8	5.3	4.9	8.7

Source: *Czech Statistical Office.*

[11] Although employers are compelled to notify labour offices of job openings, the vacancy data are likely to underestimate the actual number of job openings. Underreporting seems more likely in urban areas, where wider use of other channels of the job search is possible. The underreporting is consequently likely to be uneven across districts. Assuming that the differences in underreporting across districts are time-invariant, the effect of underreporting is removed by the first difference used in this paper.

Table 3: Variation coefficient of average monthly earnings across districts

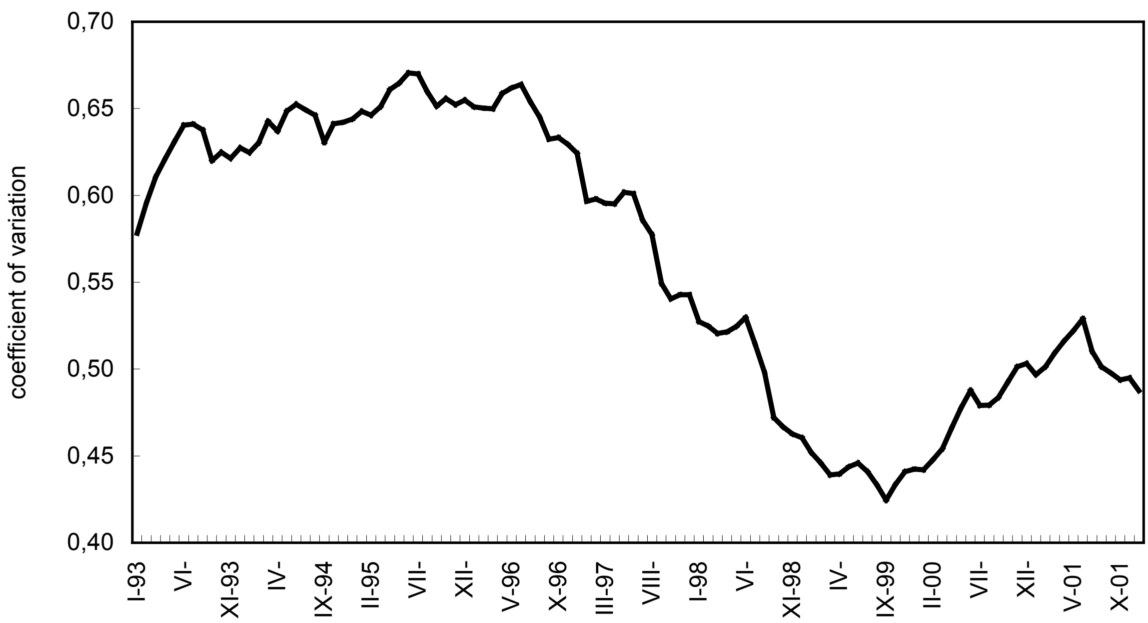
	1996	1997	1998	1999	2000	2001
All districts	0.087	0.088	0.098	0.101	0.103	0.099
Districts excluding Prague*	0.081	0.079	0.085	0.086	0.084	0.088
Districts excluding former regional capitals**	0.068	0.069	0.078	0.081	0.081	0.086

Source: *Czech Statistical Office.*
Note: *The variation coefficient is the standard deviation divided by the mean value, weighted by the district labour force.*
* *District excluded: Prague.*
** *Districts excluded: Prague, Ceske Budejovice, Plzen, Usti n.L., Hradec Kralove, Brno, Ostrava.*

6. Results

Before commenting on the regression results, we turn our attention to unemployment and wage patterns across districts during the years 1996–2001. This will help us understand why previous wage curve estimates are lower in absolute value or insignificant for the Czech Republic. The differences in unemployment rates across districts are illustrated using the coefficient of variation in Figure 1.[12]

Figure 1: Unemployment differentials, 1993–2001



Source: *Ministry of Labour and Social Affairs.*

[12] The variation coefficient is the standard deviation weighted by the district labour force divided by the mean unemployment rate.

During the 1997–1999 recession, the differences declined sharply. This was followed by a temporary upward movement in 2000. The differences started to decline again in the middle of 2001.[13] Despite the remarkable changes in unemployment differences, the differences in wages across districts did not change significantly during 1996–2001. The last two columns of Table 4 show that unlike the variation coefficient of unemployment rates, the variation coefficient of district average wages increased only slightly between 1996 and 2001. Unemployment rates are flexible, while wage differences across districts are rigid during 1996–2001, indicating that the unemployment elasticity of wages might be low in absolute value or insignificant.

Although the unemployment differences declined during the 1997–1999 recession, the unemployment rates increased in all districts. This follows from Table 4, where for each year we have sorted all the districts except Prague by unemployment rate and created two groups of districts. Low-unemployment districts are districts with a cumulative labour force of 25% of the total labour force among the districts with the lowest unemployment rates. High-unemployment districts are defined in the same way for the districts with the highest unemployment rates. The results in Table 4 show that average unemployment rates increased in both groups of districts, but are much lower in low-unemployment districts than in high-unemployment districts during the period 1996–2001. On the other hand, the other columns in Table 4 suggest that average wages are similar in these groups of districts. Except for 1998, wages grew more slowly in high-unemployment districts than in districts with low unemployment. This indicates that the wage growth in districts might be correlated with district unemployment and that the regional Phillips curve might exist.

Table 4: Unemployment and wage differentials (Prague excluded)

	Low-unemployment districts			**High-unemployment districts**			**Coefficient of variation**	
	Average unempl. rate (%)	Average wage (CZK)	Wage growth, y-o-y (%)	Average unempl. rate (%)	Average wage (CZK)	Wage growth, y-o-y (%)	Unempl. rates	Wages
1996	1.5	9364		6.1	9305		0.52	0.081
1997	2.3	10405	11.1	7.8	10211	9.7	0.45	0.079
1998	3.7	11090	6.6	10.4	11303	10.7	0.40	0.085
1999	5.6	11883	7.1	14.0	11955	5.8	0.36	0.086
2000	5.4	12635	6.3	15.5	12644	5.8	0.41	0.084
2001	4.8	13396	6.0	15.0	13142	3.9	0.44	0.088

Source: *Ministry of Labour and Social Affairs, Czech Statistical Office.*
Note: *Low- (high-) unemployment districts are the groups of districts with the lowest (highest) unemployment rates. For every year, the groups are defined by the cumulative labour force of 25% of the total labour force among districts sorted by unemployment rates. Average wages are weighted by the district labour force. Coefficients of variation are calculated across all districts except Prague, weighted by the district labour force.*

[13] The pattern of the variation coefficient observed in Figure 1 follows the downward and upward trends of the business cycle. The seasonally adjusted variation coefficient of the district unemployment rates might be used as an indicator of the business cycle. This point deserves further investigation.

Although unemployment rates increased in all districts during the 1997–1999 recession, some districts experienced a more pronounced rise in unemployment than others. According to the results in Table 5, the districts that experienced the highest rise in unemployment between 1996 and 2001 are the districts with prevalently low unemployment rates in 1996.[14] Unemployment in these districts increased to the level observed in other districts. In particular, the gap in unemployment rates between the two groups narrowed from 1.77 in 1996 to 1.18 in 2001. The other columns of Table 5 indicate that wages in the two groups of districts grew almost at the same rate between 1996 and 2001, while the gap between average wages did not change at all (see the last column of Table 5). In the districts that experienced the most pronounced rise in unemployment, wages grew at the same rate as in the other districts, while unemployment increased rapidly between 1996 and 2001, indicating that the estimates of the unemployment elasticity of wages might be biased if all districts are included in the sample.[15] The bias might be reduced if the districts that experienced the highest increase in unemployment rates between 1996 and 2001 are excluded from the sample.

Table 5: Unemployment and wages in the districts that experienced the lowest and highest increase in unemployment rates between 1996 and 2001 (Prague excluded)

	The lowest increase in unemployment rates*			**The highest increase in unemployment rates****				
	Average unempl. rate (%)	Average wage (CZK)	Wage growth, y-o-y (%)	Average unempl. rate (%)	Average wage (CZK)	Wage growth, y-o-y (%)		
	(1)	**(2)**	**(3)**	**(5)**	**(6)**	**(7)**	**(1)/(5)**	**(2)/(6)**
1996	4.5	8961		2.5	9381		1.77	0.96
1997	5.9	9851	9.9	3.6	10365	10.5	1.62	0.95
1998	7.7	10680	8.4	5.6	11255	8.6	1.39	0.95
1999	10.3	11462	7.3	8.3	12057	7.1	1.24	0.95
2000	10.5	12169	6.2	8.8	12788	6.1	1.20	0.95
2001	10.0	12839	5.5	8.4	13495	5.5	1.18	0.95

Source: *Ministry of Labour and Social Affairs, Czech Statistical Office.*
Note: *Groups of districts are defined across districts sorted by the ratio of unemployment rates in 2001 and 1996 using K-means cluster analysis. Average wages are weighted by the district labour force.*
* *45 districts*
** *31 districts*

Our finding that rates of unemployment do not rise as rapidly in districts with already high unemployment in the recession might be misleading. In particular, registry unemployment rates might decrease to some extent in high-unemployment districts, as recipients have less incentive to register after exhausting benefits. In such case, the survey unemployment provided by the Labour

[14] 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).

[15] Further research will be directed at explaining the driving forces of these changes in unemployment.

Force Survey would be a better measure than the registry unemployment used in this paper. However, there is no difference between these measures for those who are unemployed for between 6 and 12 months.[16] At the economy-wide level, long-term survey unemployment exceeds registry unemployment, while short-term survey unemployment rates are lower than registry rates of unemployment. Although regional differences are lower for survey than for registry unemployment rates, they follow the same trend as differences in the registry measure of unemployment. In addition, survey unemployment rates are not provided at the district level.

The elasticity should be estimated in districts excluding Prague, Prague-East, and Prague-West. While Prague is excluded in order to tackle the effect associated with the wage measures in 2001 (see Section 5), the other districts are neighbours of Prague.[17] In addition, the elasticity will be estimated using the districts that experienced the lowest increase in unemployment between 1996 and 2001, as defined in Table 5. The regression results of the two-stage least-square estimation of equation (12) are reported in Table 6.

Table 6: 2SLS regression results

	(1)	(2)	(3)	(4)
logw(-1)-logw(-2)	0.087	0.251	0.943	1.019
	(0.725)	(0.645)	(0.661)	(0.645)
logu-logu(-1)	-0.032**	-0.014	-0.059**	-0.079**
	(0.016)	(0.019)	(0.027)	(0.036)
dummy 1999	0.076	0.057	0.015	0.015
	(0.059)	(0.051)	(0.051)	(0.047)
dummy 2000	0.054	0.042	-0.008	-0.013
	(0.052)	(0.047)	(0.048)	(0.047)
dummy 2001	0.050	0.041	-0.007	-0.013
	(0.045)	(0.041)	(0.041)	(0.042)
Observations	222	222	135	135
Number of districts	74	74	45	45
F statistics	702.5***	628.5***	243.5***	237.0***
Hausman test	0.01	0.84	2.85*	1.93

Note: *(1), (2) Districts of Prague, Prague-West and Prague-East excluded.*
(3), (4) Districts with the largest increase in unemployment rates between 1996 and 2001 excluded.
(1), (3) logw(-1)-logw(-2) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), dummies.
(2), (4) logw(-1)-logw(-2), logu-logu(-1) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), logi-logi(-1), logi(-1)-logi(-2), logi(-2)-logi(-3), dummies.
Dependent variable: logw-logw(-1), robust standard errors in parentheses.
* *significant at 10%,* ** *significant at 5%,* *** *significant at 1%.*
Hausman test: H0: instrumented variable(s) exogenous (F statistics reported).

[16] The duration of unemployment benefits is six months, while twelve months is the internationally recognised threshold used to define long-term unemployment.

[17] We can also argue that country capitals have specific local labour markets. We exclude Prague-East and Prague-West because many residents in these districts commute to Prague. They affect the wage statistics in Prague and the unemployment statistics in Prague-East and Prague-West.

Columns (1) and (2) in Table 6 show the results for districts excluding Prague, Prague-East and Prague-West. The unemployment elasticity of wages is –0.032, while it is insignificant when the unemployment rate is instrumented (column 2). The elasticity is greater in absolute value and significant when the districts with the highest increase in unemployment rates between 1996 and 2001 are excluded from the sample (columns 3 and 4). The elasticity is significant and its size is close to the values reported for other countries in the literature. It seems that the wage curve does not have an identical shape in all districts. Including the districts that experienced the largest increase in unemployment rates between 1996 and 2001 in the sample leads to bias in the elasticity estimates.[18] The effects of short-term and long-term unemployment are described in Table 7.

Table 7: 2SLS results for short-term and long-term unemployment

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
logw(-1)-logw(-2)	0.190	0.716	0.206	0.637	1.037	1.169	1.056	1.399
	(0.723)	(0.689)	(0.700)	(0.653)	(0.823)	(0.743)	(0.856)	(0.883)
logst6u-logst6u(-1)	-0.004	-0.062**			-0.042	-0.046		
	(0.022)	(0.029)			(0.050)	(0.055)		
loglt6u-loglt6u(-1)	-0.017	0.054			-0.018	-0.034		
	(0.019)	(0.034)			(0.038)	(0.070)		
logst12u-logst12u(-1)			-0.016	-0.046**			-0.044	-0.078*
			(0.016)	(0.019)			(0.038)	(0.044)
loglt12u-loglt12u(-1)			-0.008	0.054**			-0.017	0.023
			(0.015)	(0.025)			(0.038)	(0.095)
dummy 1999	0.066	-0.002	0.064	0.005	0.004	0.001	0.006	-0.031
	(0.060)	(0.060)	(0.057)	(0.054)	(0.070)	(0.067)	(0.073)	(0.090)
dummy 2000	0.048	-0.009	0.045	-0.009	-0.020	-0.029	-0.019	-0.059
	(0.056)	(0.053)	(0.053)	(0.049)	(0.070)	(0.063)	(0.074)	(0.085)
dummy 2001	0.044	0.013	0.043	0.017	-0.013	-0.023	-0.015	-0.037
	(0.046)	(0.044)	(0.044)	(0.042)	(0.052)	(0.047)	(0.054)	(0.055)
Observations	222	222	222	222	135	135	135	135
Number of districts	74	74	74	74	45	45	45	45
F statistics	559.6***	329.4***	553.7***	359.5***	187.0***	174.6***	184.0***	140.1***
Hausman test	0.04	2.86**	0.05	3.63**	2.72	1.45	2.20	1.82

Note: *(1)–(4) Districts of Prague, Prague-West and Prague-East excluded.*
(5)–(8) Districts with the largest increase in unemployment rates between 1996 and 2001 excluded.
(1), (3), (5), (7) logw(-1)-logw(-2) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), dummies.
(2), (4), (6), (8) logw(-1)-logw(-2), logu-logu(-1) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), logi-logi(-1), logi(-1)-logi(-2), logi(-2)-logi(-3), dummies.
Dependent variable: logw-logw(-1), robust standard errors in parentheses.
* *significant at 10%,* ** *significant at 5%,* *** *significant at 1%..*
Hausman test: H0: instrumented variable(s) exogenous (F statistics reported).

[18] We calculated the F-test of the difference between groups of districts. The difference is significant between the districts of Prague and other districts, while it is insignificant for the districts that experienced the highest increase in unemployment rates between 1996 and 2001 with respect to other districts. The choice of district groups is arbitrary in Table 6. The results of the F-test would be different for another choice of district groups.

Two measures of short-term and long-term unemployment are applied in Table 7. In columns (1), (2), (5), and (6), short-term unemployment is defined as the unemployed being out of work for less than 6 months, while the long-term unemployed are idle for more than 6 months. The other columns in Table 7 show the results for unemployment rates less than and more than 12 months.

The results in Table 7 indicate that the unemployment elasticity of wages is negative and significant for the short-term unemployed, while it is lower in absolute value or even positive for the long-term unemployed. A higher level of long-term unemployment is observed in high unemployment districts. These districts are, for example, heavy industry or mining districts exhibiting both higher wages and higher unemployment. The prevalently high wages there do not adjust downward with increasing unemployment. A tentative explanation is that the welfare system represents a floor preventing further downward wage adjustments in these districts.

Table 8: 2SLS results for districts with low and high unemployment rates in 2001

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
logw(-1)-logw(-2)	0.302	0.315	0.424	0.223	0.398	0.135	2.549	0.413
	(0.414)	(0.373)	(0.480)	(0.310)	(1.230)	(0.942)	(3.074)	(0.886)
logu-logu(-1)	-0.039**	-0.024	-0.076**	-0.070*	-0.018	0.005	0.006	0.033
	(0.020)	(0.022)	(0.035)	(0.040)	(0.049)	(0.055)	(0.088)	(0.081)
dummy 1999	0.063*	0.057*	0.061*	0.075***	0.043	0.057	-0.126	0.032
	(0.034)	(0.030)	(0.036)	(0.022)	(0.098)	(0.084)	(0.234)	(0.086)
dummy 2000	0.040	0.040	0.031	0.046*	0.029	0.046	-0.131	0.023
	(0.031)	(0.028)	(0.036)	(0.024)	(0.085)	(0.067)	(0.223)	(0.068)
dummy 2001	0.034	0.035	0.022	0.036	0.035	0.051	-0.088	0.030
	(0.027)	(0.025)	(0.033)	(0.022)	(0.072)	(0.056)	(0.168)	(0.052)
Observations	135	135	75	75	87	87	60	60
Number of districts	45	45	25	25	29	29	20	20
F statistics	389.1***	378.9***	219.2***	266.8***	253.7***	278.3***	43.2***	244.8***
Hausman test	1.04	1.12	1.47	0.93	0.06	0.09	0.96	0.10

Note: *(1)–(4) Districts with low unemployment rates in 2001.*
(5)–(8 Districts with high unemployment rates in 2001.
(1), (2), (5), (6) Districts of Prague, Prague-West and Prague-East excluded.
(3), (4), (7), (8) Districts with the largest increase in unemployment rates between 1996 and 2001 excluded.
(1), (3), (5), (7) logw(-1)-logw(-2) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), dummies.
(2), (4), (6), (8) logw(-1)-logw(-2), logu-logu(-1) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), logi-logi(-1), logi(-1)-logi(-2), logi(-2)-logi(-3), dummies.
Dependent variable: logw-logw(-1), robust standard errors in parentheses.
* *significant at 10%,* ** *significant at 5%,* *** *significant at 1%..*
Hausman test: H0: instrumented variable(s) exogenous (F statistics reported).

The results in Table 8 show that the wage curve is observed in districts with low unemployment rates in 2001, while the wage curve does not exist in districts with high unemployment.[19] The

[19] We separated the districts into two groups according to the unemployment rate in 2001 using cluster analysis. There are 48 districts with low unemployment and 29 districts with high unemployment in 2001.

reason why the elasticity is not significant in high-unemployment districts might be the same as in the case of long-term unemployment. In particular, workers do not accept lower wages when unemployment rises because they prefer being unemployed, owing to the relatively generous welfare system.

Table 9: 2SLS results for districts with a low and high share of public sector employment in 1996

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
logw(-1)-logw(-2)	-0.200	-0.085	0.888	0.894	0.377	1.366	0.568	0.978
	(0.741)	(0.633)	(0.694)	(0.667)	(1.548)	(0.869)	(1.042)	(0.803)
logu-logu(-1)	-0.041**	-0.008	-0.082**	-0.080*	-0.009	0.014	-0.021	-0.037
	(0.021)	(0.026)	(0.038)	(0.044)	(0.053)	(0.044)	(0.042)	(0.051)
dummy 1999	0.102	0.081	0.024	0.023	0.049	-0.031	0.037	0.010
	(0.062)	(0.053)	(0.054)	(0.050)	(0.128)	(0.070)	(0.085)	(0.062)
dummy 2000	0.075	0.067	-0.001	-0.002	0.032	-0.042	0.017	-0.014
	(0.053)	(0.045)	(0.050)	(0.048)	(0.115)	(0.065)	(0.078)	(0.060)
dummy 2001	0.068	0.063	-0.007	-0.008	0.032	-0.025	0.022	-0.003
	(0.048)	(0.042)	(0.047)	(0.046)	(0.088)	(0.051)	(0.058)	(0.047)
Observations	147	147	81	81	75	75	54	54
Number of districts	49	49	27	27	25	25	18	18
F statistics	401.2***	380.0***	143.8***	159.0***	366.2***	124.7***	208.2***	127.4***
Hausman test	0.17	1.07	1.71	0.88	0.16	4.07**	0.81	2.07

Note: *(1)–(4) Districts with low public sector employment share in 1996.*
(5)–(8) Districts with high public sector employment share in 1996.
(1), (2), (5), (6) Districts of Prague, Prague-West and Prague-East excluded.
(3), (4), (7), (8) Districts with the largest increase in unemployment rates between 1996 and 2001 excluded.
(1), (3), (5), (7) logw(-1)-logw(-2) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), dummies.
(2), (4), (6), (8) logw(-1)-logw(-2), logu-logu(-1) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), logi-logi(-1), logi(-1)-logi(-2), logi(-2)-logi(-3), dummies.
Dependent variable: logw-logw(-1), robust standard errors in parentheses.
** significant at 10%; ** significant at 5%; *** significant at 1%.*
Hausman test: H0: instrumented variable(s) exogenous (F statistics reported).

The wage curve exists in districts where the share of public sector employment is low. The elasticity is lower in absolute value and insignificant in districts with higher employment in the public sector (Table 9).[20] It should be recognised here that public sector wages are negotiated at the economy-wide level, while the variance in regional unemployment does not play a role.[21]

Wage adjustments to exogenous changes in the local unemployment rate could depend on the existing industrial or corporate structures. In particular, firms owned by foreign investors might

[20] Applying K-means cluster analysis, there are 28 districts with a high share of public sector employment in 1996 and 49 districts with a prevalently low share of persons employed in the public sector in 1996.

[21] We could argue that districts with a high share of public sector employment tend to have low unemployment. However, a high share of public sector employment is also observed in those high-unemployment districts which are regional centres with many public services.

introduce more efficient utilisation of inputs, leading to higher labour productivity and therefore larger wage adjustments. We estimate the wage curve in two groups of districts according to the cumulative amount of foreign direct investment at the end of 1998.[22] The results in Table 10 indicate that districts with high foreign direct investment exhibit a stronger wage curve, as indicated by more negative elasticity. The coefficient estimates are significant after exclusion of the districts with the largest increase in unemployment rates between 1996 and 2001. However, the results in Table 10 should be interpreted with caution because the volume of foreign direct investment flowing into the districts might not be exogenous with respect to the wage level.

Table 10: 2SLS results for districts with low and high foreign direct investment (end of 1998)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
logw(-1)-logw(-2)	0.482	0.553	1.003	0.918	-1.000	-0.452	0.167	0.460
	(0.485)	(0.470)	(0.765)	(0.686)	(1.547)	(0.885)	(0.817)	(0.895)
logu-logu(-1)	-0.022	-0.011	-0.059*	-0.059*	-0.050	-0.056	-0.121**	-0.184***
	(0.019)	(0.022)	(0.030)	(0.034)	(0.062)	(0.073)	(0.045)	(0.062)
dummy 1999	0.042	0.033	0.011	0.018	0.180	0.132	0.094	0.085
	(0.040)	(0.040)	(0.058)	(0.050)	(0.131)	(0.081)	(0.067)	(0.067)
dummy 2000	0.026	0.021	-0.010	-0.004	0.132	0.092	0.045	0.026
	(0.035)	(0.034)	(0.056)	(0.050)	(0.111)	(0.064)	(0.061)	(0.066)
dummy 2001	0.024	0.020	-0.014	-0.009	0.123	0.091	0.053	0.036
	(0.030)	(0.030)	(0.050)	(0.046)	(0.092)	(0.056)	(0.034)	(0.038)
Observations	180	180	114	114	42	42	21	21
Number of districts	60	60	38	38	14	14	7	7
F statistics	412.6***	401.9***	184.7***	210.0***	102.7***	165.8***	416.7***	222.7***
Hausman test	1.37	1.01	1.68	0.9	1.45	0.22	2.27	1.33

Note: *(1)–(4) Districts with low FDI.*
(5)–(8) Districts with high FDI.
(1), (2), (5), (6) Districts of Prague, Prague-West and Prague-East excluded.
(3), (4), (7), (8) Districts with the largest increase in unemployment rates between 1996 and 2001 excluded.
(1), (3), (5), (7) logw(-1)-logw(-2) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), dummies.
(2), (4), (6), (8) logw(-1)-logw(-2), logu-logu(-1) instrumented using logw(-2)-logw(-3), logv(-1)-logv(-2), logv(-2)-logv(-3), logi-logi(-1), logi(-1)-logi(-2), logi(-2)-logi(-3), dummies.
Dependent variable: logw-logw(-1), robust standard errors in parentheses.
** significant at 10%, ** significant at 5%, *** significant at 1%,*
Hausman test: H0: instrumented variable(s) exogenous (F statistics reported).

[22] Using K-means cluster analysis, there are 60 districts with low and 17 districts with high foreign direct investment per capita at the end of 1998.

7. Policy Implications

The prime conclusion of this paper is that local unemployment affects the level of wages. This finding is robust after the districts of Prague or Prague and other districts that experienced the highest increase in unemployment between 1996 and 2001 are excluded from the sample. We observe the wage curve for specific unemployment rates and in particular groups of districts. In particular, the elasticity is negative and significant for the short-term unemployed, indicating that short-term unemployment rates have robust implications for wage dynamics. The intensive job search of the short-term unemployed enhances competitive forces on the labour supply side, generating downward pressure on wages. The long-term unemployed are mostly discouraged from the job search and do not affect labor supply and wage levels.

The wage curve has a different shape in specific groups of districts. The elasticity is negative and significant in districts with prevalently lower unemployment rates. The effect of unemployment on wages is strong in this case, indicating that wage inflation is determined primarily by the unemployment rates observed in these districts.

A significant elasticity is observed in districts with a low share of public sector employment in 1996. This is in line with the interpretation that in the public sector, wages are determined at the economy-wide level, while local unemployment does not have a significant role. This finding indicates that the unemployment rate in districts with a low share of public sector employment might be a good predictor of wages in the business sector.

The wage curve is weak, or does not exist, in districts with prevalently high unemployment rates or for the long-term unemployed. In these cases, wages do not adjust to changes in unemployment, owing, for example, to the welfare system and the diminishing effects of long-term unemployment. Increasing welfare benefits might extend the group of districts where the wage curve does not work. This implies weaker equilibrating forces in high unemployment districts.

We are able to investigate whether the microfoundations of wage inflation increase the predictive power of the macroeconomic models of wage inflation. In particular, suppose that wages are estimated using a modification of the Phillips curve defined in equation (2):

$$\Delta w_t = \alpha + \beta_1 u_t^{-1} + \beta_2 u_t^{-2} + \beta_3 \Delta u_t + \eta \Delta p_t + \varepsilon_t . \tag{13}$$

In equation (13), w_t and p_t are logarithms of the wage and price levels and u_t is the unemployment rate. The nominal wage growth with respect to the same period in the previous year is regressed against a function of unemployment rates and inflation, defined as the consumer price growth with respect to the same period in the previous year. Equation (13) is estimated for the economy-wide data by ordinary least squares using a quarterly sample of wages in the business sector, inflation, and unemployment rates over the period from the first quarter of 1995 to the first quarter of 2002. The unemployment rate is calculated for all districts. The results of Section 6 indicate that the wage curve is observed in districts with a low share of public sector employment and in districts with prevalently low unemployment rates. We therefore replicate the estimation of equation (13) using average unemployment rates calculated for these groups of districts. In all cases, we exclude the districts of Prague or the districts that experienced the highest increase in unemployment rates between 1996 and 2001 from the sample.

Table 11 presents estimates of equation (13). For the unemployment rate defined in each group of districts, the results are better than for the overall unemployment rate. In particular, the unemployment rates in districts with a low share of public sector employment (Group 1) and in districts with low unemployment rates excluding the districts that experienced the highest increase in unemployment rates between 1996 and 2001 (Group 4) provide significantly better estimates of equation (13) than the results for the unemployment rate in all districts.

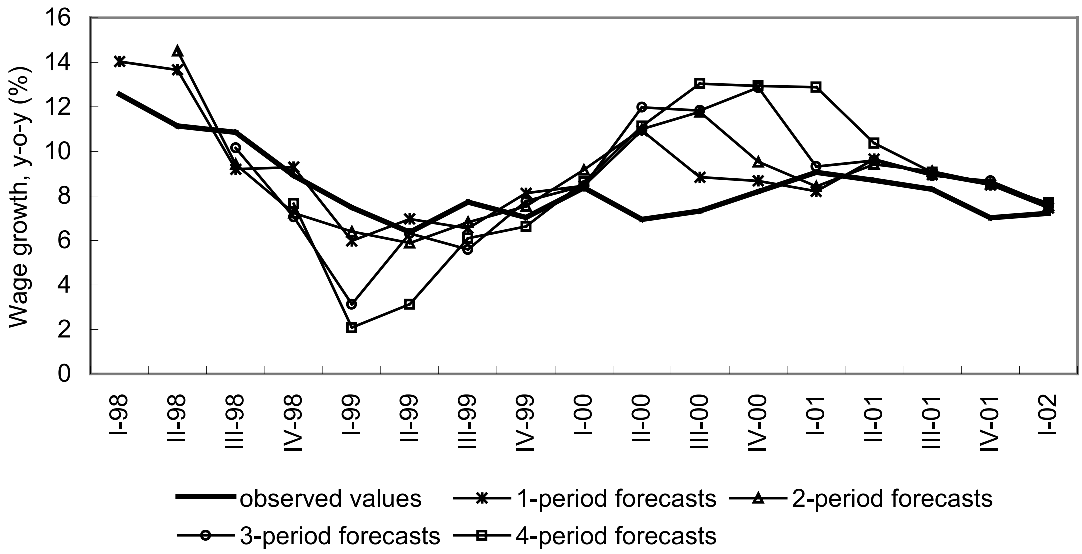
Table 11: Least square estimates of the Phillips curve, 1995:1–2002:1

	All districts	**Group 1**	**Group 2**	**Group 3**	**Group 4**
constant	-2.90	-0.18	-1.06	-2.98	-0.34
	(2.34)	(1.70)	(1.97)	(2.78)	(1.39)
1/u	83.49***	75.26***	95.35***	132.27***	58.35***
	(24.65)	(21.94)	(28.33)	(45.33)	(13.44)
1/(u*u)	-82.99*	-77.21	-119.31*	-200.62	-48.37**
	(48.39)	(47.36)	(69.73)	(137.36)	(21.29)
u-u(-4)	0.04	-0.32*	-0.33*	-0.30*	-0.12
	(0.19)	(0.18)	(0.17)	(0.17)	(0.17)
inflation	0.17*	0.14	0.11	0.10	0.16*
	(0.09)	(0.09)	(0.09)	(0.10)	(0.08)
R-squared	0.957	0.966	0.963	0.960	0.966
Akaike info criterion	2.91	2.69	2.78	2.84	2.67
Durbin-Watson stat.	2.17	2.13	2.02	2.13	2.22
Observations	29	29	29	29	29
Number of districts	77	49	27	45	25

Note: *Coefficient estimates, standard errors in parentheses.*
** significant at 10%, ** significant at 5%, *** significant at 1%.*
All districts: Average unemployment rate across all districts.
Group 1: Unemployment rate for districts with low public sector employment share in 1996, districts of Prague, Prague-West, Prague-East excluded.
Group 2: Unemployment rate for districts with low public sector employment share in 1996, districts that experienced the highest increase in unemployment rates between 1996 and 2001 excluded.
Group 3: Unemployment rate for districts with low unemployment districts in 2001, districts of Prague, Prague-West, Prague-East excluded.
Group 4: Unemployment rate for districts with low unemployment districts in 2001, districts that experienced the highest increase in unemployment rates between 1996 and 2001 excluded.

The predictive power of equation (13) estimated using the unemployment rates for particular groups of districts could be investigated by comparing past predictions to the observed values. We compute past predictions using the coefficients of equation (13) estimated within a shorter sample and compare these predictions to the observed values. In particular, we estimate equation (13) in the period from the first quarter of 1995 to every period between the fourth quarter of 1997 and the fourth quarter of 2001. In each case, we calculate 1-period to 4-period forecasts. Figure 2 shows the observed values and forecasts in the case of the overall unemployment rate. While the 1-period and 2-period forecasts are relatively accurate, the 3-period and 4-period forecasts deviate significantly from the observed values at the beginning of 1999 and in 2000.

Figure 2: Past predictions of average wage growth and observed values



Note: *Unemployment rate for all districts, quarterly data from 1995:1 to 2002:1.*

Table 12 provides the mean values and standard deviations of the differences between the forecasts and the observed values. The mean values show a systematic difference, while the standard errors might be used to calculate the confidence intervals for the forecasts, provided that the deviations are normally distributed. The results in Table 12 indicate that the systematic difference between the forecasted and observed values is lower in the case of the groups of districts than for the overall unemployment rate. The standard deviations suggest that the predictive power of the 1-period and 2-period forecasts is significantly better when the rate of unemployment is calculated for districts with low unemployment rates excluding the districts that experienced the highest increase in unemployment rates between 1996 and 2001 (Group 4). The seasonally adjusted time series of the unemployment rate calculated for this group of districts has robust implications for the wage inflation trend in these districts and at the economy-wide level.

Table 12: Past predictions of average wage growth and observed values

	1-period forecasts	2-period forecasts	3-period forecasts	4-period forecasts	Number of districts
Forecasts from 1998:1 to 2001:4					
All districts	0.60 (1.40)	0.73 (1.82)	0.65 (2.49)	0.80 (3.03)	77
Group 1	0.41 (1.43)	0.48 (2.08)	0.37 (3.24)	0.57 (4.03)	49
Group 2	0.40 (1.51)	0.43 (2.23)	0.24 (3.46)	0.36 (4.31)	27
Group 3	0.40 (1.51)	0.39 (2.13)	0.12 (3.25)	0.14 (4.05)	45
Group 4	0.43 (1.34)	0.51 (1.83)	0.41 (2.62)	0.57 (3.18)	25
Forecasts from 1999:1 to 2001:4					
All districts	0.58 (1.35)	0.88 (1.64)	0.95 (2.54)	0.96 (3.08)	77
Group 1	0.56 (1.27)	0.89 (1.69)	0.80 (3.27)	0.81 (4.09)	49
Group 2	0.59 (1.29)	0.91 (1.74)	0.75 (3.44)	0.64 (4.35)	27
Group 3	0.49 (1.37)	0.75 (1.77)	0.56 (3.26)	0.37 (4.11)	45
Group 4	0.45 (1.22)	0.72 (1.57)	0.71 (2.69)	0.74 (3.24)	25
Forecasts from 2000:1 to 2001:4					
All districts	0.95 (1.28)	1.49 (1.59)	2.01 (1.98)	2.57 (1.94)	77
Group 1	0.84 (1.19)	1.47 (1.62)	2.15 (2.17)	2.90 (2.35)	49
Group 2	0.91 (1.23)	1.53 (1.66)	2.19 (2.15)	2.90 (2.29)	27
Group 3	0.90 (1.27)	1.45 (1.64)	2.00 (2.05)	2.59 (2.08)	45
Group 4	0.74 (1.16)	1.27 (1.53)	1.82 (2.00)	2.40 (2.06)	25

Note: *Quarterly data from 1995:1 to 2002:1. Means of forecasts minus observed values, standard deviations in parentheses.*

Group 1: Unemployment rate for districts with low public sector employment share in 1996, districts of Prague, Prague-West, Prague-East excluded.

Group 2: Unemployment rate for districts with low public sector employment share in 1996, districts that experienced the highest increase in unemployment rates between 1996 and 2001 excluded.

Group 3: Unemployment rate for districts with low unemployment in 2001, districts of Prague, Prague-West, Prague-East excluded.

Group 4: Unemployment rate for districts with low unemployment in 2001, districts that experienced the highest increase in unemployment rates between 1996 and 2001 excluded.

8. Concluding Remarks

We propose a partial adjustment model that decomposes the short-term and long-term effects of regional unemployment on wages. In order to control for fixed effects, we use cross-sectional specific first differences in the basic specification. We apply the two-stage least-square method to estimate the unemployment elasticity of wages.

The results indicate that the short-term elasticity is negative and significant, while it is even greater in absolute value when the districts that experienced the largest increase in unemployment rates between 1996 and 2001 are excluded from the sample. The elasticity is negative and significant for the short-term unemployed and in districts with low unemployment rates in 2001.

At about –0.08, the elasticity is close to the estimates reported in the literature for other countries, but differs from what has been published for the Czech Republic. It seems that one has to exclude the districts that experienced the highest increase in unemployment rates between 1996 and 2001 in order to obtain an elasticity estimate that is close to the values observed in other countries. The results of this paper indicate that districts with prevalently low unemployment in 1996 exhibited the largest rise in unemployment rates between 1996 and 2001 to the level observed in other districts. While wages have also been rising in these districts during the same period, this leads to bias in the elasticity estimates.

The wage curve does not exist, or is weak, in districts with more long-term unemployed and in districts with high unemployment. Wages do not adjust to changes in local unemployment in these districts, owing, for example, to a relatively generous welfare system. In addition, increasing welfare benefits might increase the group of districts where wages do not adjust to changes in local unemployment.

The elasticity is negative and significant in districts with a low share of public sector employment, while it is insignificant in districts where the share of employment in the public sector is high. In the public sector, wages are determined at the economy-wide level and are not affected by regional unemployment.

The elasticity is greater in absolute value in districts with higher foreign direct investment. Were FDI purely exogenous with respect to district-level wage rates, this finding would indicate that firms owned by foreign investors introduce more efficient utilisation of inputs so that wage adjustment might be more flexible according to the level of the local rate of unemployment. However, foreign direct investment might not be exogenous across districts and these results should be interpreted with caution.

The model allows us to test whether the Phillips curve exists at the regional level. However, the panel is short (1996–2001), indicating that the estimates of the long-term unemployment elasticity of wages are not robust. This paper is, therefore, focused on the estimation of the short-term elasticity of wages.

The estimates of the aggregated Phillips curve indicate that the unemployment rates calculated for particular groups of districts explain wage dynamics better than the overall unemployment rate. The predictive power examined by comparing past predictions to observed values was higher for unemployment rates calculated for districts with low unemployment rates excluding the districts that experienced the highest increase in unemployment rates between 1996 and 2001 than for unemployment rates calculated for all districts.

References:

ACEMOGLU, D. (1999): "Changes in Unemployment and Wage Inequality: An Alternative Theory and Some Evidence". *The American Economic Review* 89(5), pp. 1,259–1,278.

ARELLANO, M. AND BOND, S. (1991): "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations". *Review of Economic Studies* 58, pp. 277–297.

BALESTRA, P. AND NERLOVE, M. (1967): "Pooling Cross Section and Time Series Data in the Estimation of a Dynamic Model: The Demand for Natural Gas". *Econometrica* 34(3), pp. 585–612.

BLACKABY, D.H. AND HUNT, L.C. (1992): "The 'Wage Curve' and Long-Term Unemployment: A Cautionary Note". *The Manchester School* 45(4*),* pp. 419–428.

BLANCHARD, O. AND KATZ, L.F. (1997): "What We Know and Do Not Know About the Natural Rate of Unemployment". *Journal of Economic Perspectives* 11(1), pp. 51–72.

BLANCHARD, O.J. AND KATZ, L. (1999): "Wage Dynamics: Reconciling Theory and Evidence". NBER Working Paper No. 6924.

BLANCHFLOWER, D.G. (2001): "Unemployment, Well-Being and Wage Curves in Eastern and Central Europe". *Journal of the Japanese and International Economies* 15(4), pp. 364–402.

BLANCHFLOWER, D.G. AND OSWALD, A.J. (1994): *The Wage Curve*. Cambridge: MIT Press.

CAMERON, G. AND MUELLBAUER, J. (2001): "Earnings, Unemployment, and Housing in Britain". *Journal of Applied Econometrics* 16, pp. 203–220.

CARD D. (1995): "The Wage Curve: A Review". *Journal of Economic Literature* 33, pp. 785–799.

DECRESSIN, J. AND FATAS, A. (1995): "Regional Labor Market Dynamics in Europe". *European Economic Review* 39, pp. 1,627–1,655.

HUITFELD, H. (2001): "Unemployment, Labour Market Programmes and Wage Determination: Evidence From the Czech and Slovak Republics". *CERGE-EI Discussion Paper No. 54.*

JACKMAN, R., PISSARIDES, CH. AND SAVOURI, S. (1990): "Unemployment Policies". *Economic Policy*, October, pp. 449–491.

JURAJDA, Š. (2002): "Estimating Czech Wage Curves within Firms". mimeo, CERGE-EI.

KERTESI, G. AND KÖLLÖ, J. (1997): "The Birth of the 'Wage Curve' in Hungary, 1989–95". Working Paper 104*,* The William Davidson Institute.

MADDALA, G.S. (1983): *Limited-Dependent and Qualitative Variables in Econometrics*. Cambridge: Cambridge University Press.

MCCORMICK, B. (1997): "Regional Unemployment and Labour Mobility in the UK". *European Economic Review* 41, pp. 581–589.

NICKELL, S. (1981): "Biases in Dynamic Models with Fixed Effects". *Econometrica* 49(6), pp. 1,417–1,426.

CNB Working Papers Series

1/2003	Kamil Galuščák: Daniel Münich	*Microfoundations of the wage inflation in the Czech Republic*
4/2002	Vladislav Flek: Lenka Marková Jiří Podpiera	*Sectoral productivity and real exchange rate appreciation: Much ado about nothing?*
3/2002	Kateřina Šmídková: Ray Barrell Dawn Holland	*Estimates of fundamental real exchange rates for the five EU pre-accession countries*
2/2002	Martin Hlušek:	*Estimating market probabilities of future interest rate changes*
1/2002	Viktor Kotlán:	*Monetary policy and the term spread in a macro model of a small open economy*

Czech National Bank
Economic Research Department
Na Příkopě 28, 115 03 Praha 1
Czech Republic
phone: +420 2 244 12 321
fax: +420 2 244 14 278
e-mail: eva.grenarova@cnb.cz