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THE WAGE CURVE: A PANEL DATA VIEW OF LABOUR MARKET SEGMENTS

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The wage curve: A panel data view of labour market segments

The case of Slovakia.

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Pavel Gertler[1]

Abstract

The paper studies the relationship between the local unemployment rate and wage level – commonly referred to as the wage curve. Using a panel data setup for annual enterprise-level microdata, we confirm previous findings that wages in Slovakia are, on the whole, relatively flexible – with a rise in the local unemployment rate of 10 percent being associated with a drop in wages by 0.8%. We find, however, that these elasticities differ considerably across sectors, regions and, in particular, skills. Our results indicate that overall wage flexibility in the Slovak labour market is driven more by the wage flexibility of higher-skilled employees, and their broader opportunities for employment, than by the institutional arrangements of the labour market.

JEL classification: "E 21, E 27, C 53"
Key words: wage curve, panel data, unemployment elasticity of wages, wage flexibility, Slovakia, Phillips curve, microdata

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INTRODUCTION

Flexible wage formation is at the core of labour market flexibility and, in the contemporary environment of negative demand shocks, it is becoming ever more important to the real economy. The concept of estimating unemployment elasticity of wages goes directly to the heart of the wage formation process. In a world of flexible wages, adverse changes in demand and the consequential rise of the unemployment rate are accommodated by contraction in employee wage bills. Wage flexibility is therefore increasingly important in the absence of an exchange rate channel to support the competitiveness of enterprises.

The aim of this study is to identify those labour market segments where wages are rigid to changes in the rate of unemployment. Identifying these segments and assessing their degree of rigidity allows us to see potential risks and hence facilitates labour market policy-making.

During the course of Slovakia's economic transition, many aspects of the economy underwent considerable changes in their structure. Therefore, average country estimates across time, sectors or occupation types do not provide a consistent picture. By using a large microdata database of companies, we are able to examine these structures and reveal how elasticities differ across them.

The outline of the paper is as follows. In the first part, we set the scene by elaborating on the theoretical background, which is based on the argument of a negative relationship between wages and unemployment. In addition, we pinpoint the techniques that are being used to measure wage rigidities. The second part briefly describes the microdata we used for the estimations and the data classification referred to later in the paper. In the third part, we test the validity of the Phillips curve and/or wage curve relationship on our dataset. This is necessary because both approaches model a negative relationship between wages and unemployment, although the assumptions in both cases differ. Then we propose specifications of the wage curve, which we use for the estimation across various classifications. The results are then discussed and interpreted in more detail in the fourth part. In general, they confirm the overall flexibility of wages in Slovakia, but also reveal a rather large heterogeneity across sectors, occupations and regions.

1 THEORETICAL BACKGROUND

From the extensive literature, we may distinguish two main approaches to measuring wage flexibility. The first is based on the long-run relationship between the level of wages and the level of local unemployment[2] (this concept, known also as "the wage curve", was introduced by Blanchflower and Oswald (1990) and has since been challenged by many other authors).

[2] Blanchflower and Oswald argue for the use of microeconomic data and document the existence of a logarithmic curve linking the level of pay to the unemployment rate in the local area. In their 1994 book, they draw this conclusion from an analysis of 16 nations including the US; in their 2005 publication, they expand their sample to 40 countries from around the world.

The other approach is based on aversion to wage cuts (also called "downward wage rigidity").

The wage curve approach is the one that we will apply in this study. There is an important divide in the concept of the unemployment/wage relationship. Traditionally, economic theory uses the Phillips curve, which models the trade-off relationship between the variation of wages and the level of unemployment. This is well established in a great number of specifications as a tool for economic policy or for assessing the functioning of the labour market. However, the wage curve departs from the Phillips curve approach by assuming a long-run relationship between the level of wages and the level of local unemployment.[3] Several methods to validate this relationship have been developed (more details in part 3). From the point of view of economic theory, three main approaches to wage formation are usually referred to when rationalising the long-term relationship.

The first is a demand-supply framework where both parties – labour demand and labour supply – have a certain position in the collective bargaining market (Diamond, 1982, Pissarides, 1990). In this stream of theory, an employee's position in the bargaining process is measured by the reservation wage, i.e. the wage level that makes an employee neutral between accepting a wage offer and being unemployed. The level of the reservation wage depends mainly on factors such as utility of leisure, unemployment benefits, and opportunities to obtain another form of income (or access to loans), the scale of accessible illegal employment, and others. The reservation wage will therefore rise during economic upswings and decrease when the rate of unemployment is rising. Since the reservation wage cannot be measured directly, it is in practice derived from flows of labour. Based on relevant legislation and specific conditions, these may include flows within the labour market, flows between activity and inactivity, or geographical flows of labour.

The second approach builds upon the reservation wage. Assuming that the wage level and wage growth affects productivity growth, employers rather pay in excess of the reservation wage in order to ensure the retention and motivation of key employees (Katz, 1986). This higher pay is often called the aspiration wage, i.e. what an employee assumes to be a fair pay, usually on the basis of his previous wage level. In the so-called "shirking model" of Shapiro and Stiglitz (1984), the aspiration wage serves as a reference value below which the employee will tend to underperform his employment duties. Similarly to the demand-supply model, increased demand and the consequent higher outflow from unemployment lifts the aspiration wage, i.e. the wage level required to avoid shirking.

The third approach is mainly applied to specific groups of employees (usually low-skilled and paid close to the reservation wage) where the standard demand-supply relationship does not apply. Since their chances of getting a well-paid job are worse in comparison with the rest of

[3] Blanchard and Katz (1992) conversely find a consistent positive long-run relationship between the level of wages and the rate of unemployment. They argue that the negative relationship holds only while the benefits from migration are smaller than its costs. These conditions apply in relation to the shock and to the subsequent assessment of whether it is transitory or permanent. This leads them on to argue that the wage curve is a short-run relationship. In the case of Slovakia, however, we presume that the wage curve is a long-term relationship, since the costs of migration are permanently higher than the benefits arising from wage differences, mainly due to the relatively lower nation-wide level of income and faster acceleration of real estate prices.

the workforce, they are also in a weaker bargaining position. Under this approach, the main factors of wage formation are therefore flows from and to unemployment and to inactivity and existing social assistance schemes (Card and Riddell, 1993).

No matter which approach prevails in a given labour market, the long-run relationship between the level of wages and the rate of unemployment is easily subjected to empirical observation. Based on a number of estimates, Blanchflower and Oswald (1994) claimed to have found an empirical law that the regional rate of unemployment elasticity of wages is around -0.1. They included lagged wages in their model, and finding them to be statistically insignificant, they argued that the Philips curve was no longer valid in this case. The issue, however, turned out to be not so clear-cut. Contrary evidence from US data (e.g. in Bell et al., 2002) indicated a failure to examine the autoregressive nature of hourly wages – a point that Blanchflower and Oswald admitted in their subsequent work (2005).

This contradiction between the diverse evidence has given rise to various ways of getting around it. The first suggestion comes directly from Blanchflower and Oswald (hereinafter "B and O"), who test for the autoregressive nature of wages on aggregated cell means of the individuals observed.[4] This is done because individual agents may share a common component of variance, entirely attributable neither to the measured characteristics nor to the local rate of unemployment. In such case, the common group effect may cause an error term to be positively correlated across agents from the same local labour market.

As a response to this, Card in his review of the B and O paper suggests a modified approach of testing. He argues that due to the possible technical problem associated with "the presence of both a lagged dependent variable and a regional fixed effect in the augmented 1994 version of the B and O wage curve and to possible serial correlation in the error term"[5], it is better to use the first differenced version of the equation, while testing for off-setting behaviour of the two slope coefficients of the rate of unemployment vis-à-vis the wage change.

In addition to these tests, Whelan (1997) suggests a compromise in which neither approach is to be rejected in any case. He argues that taking into account the mark-up price setting mechanism and the way in which expectations are formed, an aggregate Phillips curve may be obtained from a microeconomic specification of the wage curve without making any assumptions about the autoregressive nature of the wage level.

That wage dynamics may play an important role is proven by the different evidence of autoregression in wages in the US and other countries and by some of the limitations in testing for it. Recently, therefore, the focus of the literature has shifted overwhelmingly from static wage curve modelling (e.g. as in Baltagi and Blien, 1998) to dynamic panel data models (e.g. as in Baltagi, Blien and Wolf, 2008).

Bearing these issues in mind, we will first test for the nature of autoregression in wages according to the approaches suggested earlier. Since we find this to be far from unity, we

[4] The cell means approach in this paper relates to a single observation being the gap between the wage level observation and its annual mean.
[5] Card, D.: The Wage Curve, *A Review*, Journal of Economic Literature, American Economic Association, vol. 33(2), pages 285-299, June.

then examine the static wage curve according to Baltagi, Blien and Wolf (2008), using the two-stage method of Bell et al. (2002).

So far, to our knowledge, there have been three studies that used Slovak data to estimate some form of wage rigidity. The B and O paper (2001) uses microdata for one year (1995) and finds an elasticity of wages of -0,049. Huitfeldt (2001), examining regional data for effects of unemployment and labour market policies on real wages in the Czech Republic and Slovakia in 1992-1998, finds significantly less wage rigidity in Slovakia (elasticity under -0.1) compared to both the previous study and to the Czech Republic. Finally, Babetskii (2006) confirms the elevated level of wage flexibility before 2000 on a basic Phillips curve specification (referring mainly, however, to the fixed exchange rate regime period), while showing evidence of its deterioration thereafter.[6]

The first two of these studies, as well as the vast majority of other papers estimating wage curves for individual economies, present overall estimates of unemployment elasticities of wages as evidence for the wage curve. The variable structure of emerging economies may however reveal many new findings that could put economy-wide estimates into a different perspective. This is because the structure of an emerging economy is often a subject of more frequent and intense changes, usually driven by investments in certain sectors and certain regions. Therefore, any averaging across the economy may lead to distortions. The aim of this paper is therefore to exploit the wage curve approach at a disaggregated level in the estimates in order to examine more closely the relationship between wages and unemployment as it appears in different sectors, regions or types of occupations.

2 DATA

We use enterprise-level microdata collected by the Statistical Office of Slovak Republic in three different enterprise surveys[7] and merge them into one concise database with annual data. The database is filtered to retain only those observations where the identifier matches the relevant wage observations in two consecutive years.[8] The number of overall

[6] For wage curve specifications, so-called hysteresis represents the presence of positive or insignificant values of wage elasticity of unemployment, which indicate an absence of wage flexibility. This finding is confirmed e.g. by Galuščák and Münich (2005) and Babetski (2006) for some central and eastern European countries after 2000. They identified two main reasons for hysteresis: (i) an upswing of long-term unemployment, which causes exclusion of an ever larger part of the labour supply from the labour market (see Winter-Ebmer, 1993) and therefore has a lessening impact on wages, (ii) a transition to a low-inflation environment, where low nominal wage changes imply higher nominal and real wage rigidities. This mechanism of grease and sand effects caused by a different level of inflation determines the relationship between nominal and real wage rigidities and is further described in Groshen and Schweitzer (1996).

[7] The Annual Questionnaire on Labour (Praca 3-01), the Annual Questionnaire on Business Statistics by Establishments (ROC Zav 1-01) and by Enterprises (ROC 1-01).

[8] In addition, it is filtered for missing entries, illogical entries (e.g. negative value added), and entries that are very likely to be wrong.

observations is thus reduced by more than one quarter – to 32.944 observations[9] covering the time period from 2001 to 2007 (one lag included).

Overall, this filtered database contains 7903 separate establishments (firms).[10] We scaled down the sample to include only those firms for which there are at least five consecutive annual observations. Filtering for this condition leaves us with 3498 firms, for which there are 22814 annual observations. This further filter is designed to avoid unnecessary noise caused by keeping excessively short time series in the panel data sample.[11]

Table 1: Filtered data statistics: frequency of observations across businesses

Frequency	Percent	Cumul.	Pattern in year						
			2001	2002	2003	2004	2005	2006	2007
2439	69.73	69.73	1	1	1	1	1	1	1
339	9.69	79.42			1	1	1	1	1
241	6.89	86.31	1	1	1	1	1		
237	6.78	93.08		1	1	1	1	1	1
209	5.97	99.06	1	1	1	1	1	1	
33	0.94	100.00		1	1	1	1	1	
3498	100.00								

Note: Value 1 stands for the presence of data for the respective year. Thus, the first row of the table represents the statistics of businesses for which there is a complete dataset for all the observed years.

Table 1 shows the structure of the dataset analysed in this paper. For cross-checking robustness, we make calculations and several estimates for the former structure as well as for the subsample of 2439 firms with complete datasets.[12]

We use the concept of a wage bill per hour of work in a company, which is calculated as compensation paid to the employees of a company divided by the number of hours worked. Our extensive database allows us to measure wages as wages paid per physical employee or normative employee (working 40 hours per week). While the results are similar, we stick to

[9] Observations include an exhaustive list of enterprises that have more than 20 employees and are obliged to report to the Statistical Office of the Slovak Republic. These are corporate sector enterprises covered by the NACE Rev.1.1 classification from A to K. The categories of Financial intermediation (class J), public administration (L) and education (M) are not included.

[10] An establishment is an organisational unit of a company which maintains separate accounts and operates in a different region. There are 7357 enterprises in the basic sample (For simplicity, we also call them firms in this paper).

[11] In general, records are eliminated in three cases: (i) where the firm started reporting after 2003, (ii) where the firm stopped reporting before 2005, and (iii) where the firm did not, for any reason, report in at least one of the years of its existence (gaps in annual reporting). In the third case, a firm whose records include observations in 2001, 2002, 2003, 2005 and 2006 would be eliminated as it would yield three non-consecutive annual changes, whereas a firm with the same number of annual observations, but with observations present in the years 2001 to 2005 would be included also in the reduced filtered sample.

[12] Since coefficients in these cases do not deviate significantly from the estimates of the Table 1 structure, we do not report them in the paper.

the former definition so that we can capture the potential shift towards more flexible employment contracts, especially at the end of the period under review.

Since the wage curve is usually estimated on individual employee data, the use of company microdata may yield some risk of information loss in aggregation to enterprise-level data. To a certain degree, this may be balanced by the broad variety of attributes in the database. Thus, we can identify a relationship between a company's characteristics and its general approach to wage setting, assuming equal wage-related treatment within the company.

The overall interpretation will also be altered by having different units of observation. The interpretation of slope coefficients of the rate of unemployment with a sample of individual employees lies at the core of the employer-employee relationship and wage bargaining, while the enterprise-level sample sheds light on productivity-related wage setting.

The rate of unemployment data are drawn from Labour Office Statistics. We use monthly rates of unemployment per district (NUTS 4 geographical unit) averaged to annual frequency. Since the methodology for reporting the rate of unemployment in Slovakia was changed in 2004, we calculate monthly rates of unemployment before this date based on its current definition (unemployed available for work[13] in a given month as a share of the average labour force for the previous year).

To sum up, we will work with an unbalanced panel of cross-section company-level data over seven annual observations. The database also includes variables and attributes assigned to each company (or, alternatively, a geographically localised or activity-based firms), which we use to build up a set of individual characteristics of firms. These include: share of foreign capital, number of hours worked per employee, and occupational structure as presented in Table 2.

The structures in which the wage curve will be analysed are defined as follows.

Occupation type – value 1 is assigned to companies where the share of a certain type of occupation exceeds the given threshold; value 0 is assigned in other cases. The methodology for setting the variable is shown in Table 2.

Table 2: Occupational breakdown: definition and statistics

Occupation type	Condition (% of empl.)	Observations (full)	ISCO-08 major groups
White-collar high-skilled	ISCO 1 to 3 > 1/3	5984	1 Managers
			2 Professionals
			3 Technicians and associate professionals
White-collar low-skilled	ISCO 4 and 5 > 1/2	4768	4 Clerical support workers
			5 Service and sales workers
Blue-collar and low-skilled	ISCO 7 to 9 > 3/4	9314	7 Craft and related trades workers
			8 Plant and machine operators + assemblers
			9 Elementary occupations

[13] The Statistical Office of the Slovak Republic refers to this group as "disposable unemployed" on the English-language version of its website.

NACE classification – the twofold definition comprises, firstly, primary sectors and, secondly, the two digit NACE classification with some adjustments[14] made for the number of observations in the database.

Geographical classification – this is defined according to NUTS 3 units, i.e. by regions, although unemployment is measured by NUTS 4 in order to better capture local labour market conditions and relatively large inter-regional differences.

3 TESTING THE VALIDITY OF THE WAGE CURVE HYPOTHESIS AND THE MODEL

3.1 ORIGINS

We depart from B and O's (1994) standard formulation of the wage curve, which lets the individual wage level be a function of individual characteristics and the regional rate of unemployment:

$$w_{irt} = \alpha + \beta u_{rt} + \lambda X_{irt} + \mu_r + \varepsilon_{irt} \qquad (1)$$

where w_{irt} denotes the wage of an individual *i* in region *r* at time *t,* and where u_{rt} is the rate of unemployment in the region; X_{irt} is a vector of factors specific to the individual *i* at time *t* that may influence wages (while each individual belongs to some regional group *r*); μ_r is a fixed regional effect which captures structural features of the region;[15] ε_{irt} denotes the idiosyncratic error term, and β and λ are the parameters to be estimated.

In its perception of a single observation, the dataset used in this paper differs from the standard wage curve concept. Consequently, the terms are reformulated as follows: w_{irt} will denote an average hourly wage paid to an employee in a firm *i* in region *r* at time *t*, and X_{irt} is a vector of factors specific to the firm *i* at time *t*.

In the first stage, we focus on testing the scope of dynamics in wages. This is to validate the wage curve hypothesis for our dataset. To do this, we follow the three above-mentioned approaches

[14] The manufacture of textiles (DB), leather (DC) and wood (DD) products are merged into one class. The same is done for the manufacture of chemicals (DG), rubber (DH) and other non-metallic (DI) products. On the other hand, wholesale and retail trade is separated into the sale, maintenance and repair of motor vehicles (G0), wholesale trade (G1) and retail trade (G2). Similarly, transport of any kind (I0) is separated from post and telecommunications (I4) and real estate and renting (K0) from computer, research and other business activities (K4). This partly reflects the new concept of the NACE Rev.2 classification.

[15] This term represents permanent features of the environment in the region, while the rate of unemployment should be affected by transitory aspects of the relationship between wage level and unemployment.

3.2 TESTS APPLIED

a) Departing from B and O (1994), a simple procedure would include lagged nominal wages. The reduced version to cell means, which ensures that group effects of individual companies in specific regions are avoided, is set as:

$$w_{rt} = \alpha + \rho w_{rt-1} + \beta_1 u_{rt} + \beta_2 q_{rt} + \lambda^T X_{rt} + \mu_r + \varepsilon_{rt} \qquad (2)$$

where ρ is the tested parameter of wage dynamics.[16] This parameter should measure the speed of wage adjustment to the unemployment rate. If this value is significantly different from one, it may be argued, according to B and O, that the Phillips curve can be rejected.[17] If it is not significantly different from one, the validity of the Phillips curve should not be rejected. Intermediate cases of 0<ρ<1 require examination by dynamic wage curve specification, where the impact of changes in the local rate of unemployment lasts for longer than one period.

b) We then apply a differenced type of testing suggested by Card (1995). Among other things, he criticised specification (2) as yielding inaccurate coefficient estimates due to the presence of both the regional fixed effect and lagged dependent variable and to a possible serial correlation in the regional market error term ε_{rt}. He suggests avoiding the dynamics component by first differencing the equation (dropping the regional fixed effect term) and estimating as follows:

$$\Delta w_{rt} = \alpha_1 u_{rt} + \alpha_2 u_{rt-1} + \beta_1 q_{rt} + \beta_2 q_{rt-1} + \lambda_1 X_{rt} + \lambda_2 X_{rt-1} + g_t + \Delta\varepsilon_{rt} \qquad (3)$$

where the wage curve hypothesis implies that $\alpha_1 = -\alpha_2$, while the Phillips curve hypothesis implies that $\alpha_2 = 0$.

c) For the third step, we apply the two-stage testing proposed by Solon et al. (1994) and used e.g. by Blanchard and Katz (1997). In the first stage, wages are regressed on company characteristics and regional effects. In the second stage, these regional dummies are then used as a measure of average regional wages, and are regressed on the fixed-time effects, the unemployment, the value added and lagged wages.

1st stage $$w_{irt} = a + b^T X_{irt} + \gamma_r^T \delta_r + \gamma_t^T \tau_t + \varepsilon_{irt} \qquad (4a)$$

2nd stage $$\gamma_r = a + \rho w_{rt-1} + \tau_t + \beta u_{rt} + \varepsilon_{rt} \qquad (4b)$$

[16] The term q_{rt} denotes labour productivity, which may also be a part of the matrix X_{rt}.

[17] Parameter value ρ=1 yields a traditional Phillips curve $(w_{ir} - p)_t = (w_{ir} - p)_{t-1} + \alpha_i - \beta u_{rt}$, while ρ=0 yields the simplest form of the wage curve $w_{irt} - p_t = \alpha_i - \beta u_{rt}$. No consensus, however, has been reached on what it means that ρ is sufficiently distant from unity.

3.3 TESTS RESULTS

Since the lagged dependent variable is present in equation (2), it implies dynamics in the panel data. The estimation of fixed effects or random effects cannot then be used, since the time dimension is too short and the estimator of ρ would be biased (Anderson and Hsiao, 1981). We therefore use the Arelano and Bond (1991) estimation technique, which yields $\rho = 0.09$. This is far enough distant from one to rule out a Phillips curve relationship, although it is not possible to rule out autocorrelation in regional market errors,[18] as suggested by Card-based test results on this dataset.

The second test does not reject either of the two theoretical hypotheses as being of some relevance. The estimated coefficients $\alpha_1 = 0.03, \alpha_2 = -0.03$ fit well the test of $\alpha_1 = -\alpha_2$, although they are too close to zero to draw a conclusive answer.

Finally, the two-stage testing yields $\rho = 0.27$, this being the only test that would conclusively confirm the validity of the wage curve relationship. All in all, none of the tests that we undertook implies that the wage curve relationship is not applicable to our dataset (see Table 3 for results of these tests). On this basis, we will consider the wage curve hypothesis to be relevant to the dataset in question (see appendix for test reports).

3.4 THE MODEL

The rationale behind the model is to estimate the slope coefficients of regional unemployment on the hourly average wage paid in the companies.

We depart from the basic model of one-way error component regression:

$$w_{it} = \alpha + \beta u_{rt} + \chi q_{it} + \lambda X_{it} + \mu_R + \varepsilon_{it} \qquad (5)$$

where q_{it} is log of company value added per hour of employees' work, u_{rt} is log of regional rate of unemployment, μ_R is a regional fixed effect,[19] ε_{it} is an idiosyncratic component, and X_{it} is composed of the following characteristics:

hrse number of hours worked by one employee (in logs)
cafa share of foreign capital in fixed assets
flex ratio of the use of flexible hours (employees count per normalised hours count)
size number of employees (in logs)
blue dummy for firms with a high proportion of blue-collar and low-skilled employees (see Table 2)

[18] Based on the Arelano-bond test, the hypothesis of zero autocorrelation was rejected with 99% significance.
[19] Regional breakdown *r* corresponds to the NUTS 4 classification, while regional breakdown *R* corresponds to NUTS 3.

whil dummy for firms with a high proportion of clerical, sales and service employees (see Table 2)

whih dummy for firms with a high proportion of high-skilled white-collar employees (see Table 2)

The structure of the panel we work with is unbalanced and largely in favour of a cross-section dimension (N~4700 and 5≤T≤7).[20]

It is almost impossible to employ a standard-type fixed-effects model with individual heterogeneity term μ_i, firstly because of the enormous loss of degrees of freedom with large N, and, secondly, since such a term would cancel out the cross-section explanatory power of the individual characteristics variables X_{it}. This would then be in contradiction with the prime objective to estimate the local unemployment elasticity of firms' wage setting.

Keeping the dependent variable and explanatory variables in nominal terms does not rule out that trends on both sides of equation are equivalent. We therefore have a control for time effect. The simplest way of controlling for this effect is to include time dummies in the former specification, so that:

$$w_{it} = \alpha + \beta u_{rt} + \chi q_{it} + \lambda X_{it} + \mu_R + \partial D_t + \varepsilon_{it} \qquad (6)$$

The validation of the use of fixed effects is obvious from the panel setup. On one hand, very large N would clog the model with (N-1) cross-section dummy variables, resulting in an enormous loss of degrees of freedom and aggravating the problem of mulitcollinearity among the regressors. On the other hand, N being large in this sample F-test suggests that OLS would report lower standard errors and thus be biased. Applying the Hausman test to further discriminate between the fixed-effects and random-effects model is for the full sample robustly tilted to the fixed effect specification. The Breush-Pagan test also confirms the inappropriateness of random effects in this case. Therefore a fixed effects model is the preferred method of estimation.

Since the very limited time dimension of the model may entail a certain undesirable trend, we can remove it entirely it by mean differencing both sides of the equation in each observed year. In this way, we arrive at a stacked cross-section model (7) that may be further cleaned for possible nominal inconsistencies of wage and productivity trends by employing annual dummies (8):

$$(w_{it} - \overline{w}_t) = \alpha + \beta(u_{rt} - \overline{u}_t) + \chi(q_{it} - \overline{q}_t) + \lambda X_{it} + \mu_R + \left[\partial D_t\right] + \varepsilon_{it} \qquad (7) \text{ and } [8]$$

For the full sample, we report the above four model specifications (5-8). In general, the fixed effects model (5-6) is compared with the stacked cross-section model (7-8), while both specifications are reported with and without yearly dummy variables. The preferred

[20] The numbers of companies in a specific year in our sample range from 4535 to 4885. Some estimations were also made with a balanced panel, achieved by retaining in the sample those establishments that reported to the Statistical Office for each year of observation (2439 establishments x 7 years). Since, however, these estimates yield very similar coefficients as the unbalanced dataset, we do not report them.

specification is then used to examine the individual segments of the labour markets and to draw conclusions on the specifics related to the occupational structure of firms, sectors and geographical regions[21].

4 RESULTS

Results (see Table 3) are reported in columns as follows:

(1) Fixed effects model (eq. 5)
(2) Fixed effects model with time dummies (eq. 6)
(3) Stacked demeaned model (eq. 7)
(4) Stacked demeaned model with time dummies (eq. 8)

Estimating the fixed time-invariant model of the specification (7) proves to be broadly equivalent to specification (6), which implies that time trend was indeed an issue in (5). By contrast, applying the same annual dummies in (8) does not yield any further added value, i.e. mean differencing is quite sufficient for detrending of the sample.

Table 3: Overall sample estimates

w (wdif)	(5)	(6)	(7)	(8)
u	-0.0974***	-0.0794***		
q	0.2868***	0.2865***		
udif			-0.0798***	-0.0794***
qdif			0.2868***	0.2865***
hrse	-0.2509***	-0.2415***	-0.2330***	-0.2415***
cafa	0.1680***	0.1674***	0.1665***	0.1674***
flex	-0.0814***	-0.0799***	-0.0798***	-0.0799***
blue	-0.0966***	-0.0962***	-0.0957***	-0.0962***
size	-0.0046*	-0.0044*	-0.0041*	-0.0044*
whil	-0.1217***	-0.1199***	-0.1201***	-0.1199***
whih	0.1909***	0.1933***	0.1930***	0.1933***
rok1		-0.0528***		0.0246***
rok2		-0.0281***		0.0106
rok3		-0.0394***		0.0052
rok4		-0.0504***		0.0087
rok5		-0.0369***		0.0068
rok6		-0.0278***		0.0032
rok7		(omitted)		(omitted)
cons	5.1601***	5.0826***	1.8219***	1.8780***
N	22623	22623	22623	22623
N_g	8	8	8	8
r2_w	0.5587	0.5599	0.5518	0.5521

Note: udif is the annually demeaned local unemployment rate u, and qdif is the annually demeaned value added per hour worked. Specifications (7) and (8) that use these explanatory variables have demeaned wages per hour worked as a dependent variable. rok1-rok7 are annual dummies for 2001-2007. N represents the number of observations; N_g relates to μ_R *and r2_w is an R2-within fit.*

[21] The variables used throughout this study are nominal rather than real (deflating with consumer price inflation) for two reasons: (1) a distorting trend would still be present in our short time dimension as only the boom period of the business cycle is captured, and (2) wage curve literature has historically proposed an inverse relation between the level of nominal wages and the unemployment rate in local labour markets.

The marginal loss on the fit is traded off by circumventing the possible distortion stemming from the time-variant effect in the estimated coefficients. Therefore, specification (7) further serves in this paper as the preferred specification.

In general, overall estimates of the wage curve confirm the existence of a wage curve in Slovakia. The resulting overall elasticities are considerably close to the golden rule promoted by B and O. The model is quite robust, reflecting also the size of the sample we use.

The wage curve is, however, far from being constant across different market segments. To assess the variability of unemployment elasticity of wages across market segments, we estimate the subsample for each one of them. There is a somewhat higher slope (-0.113) where the only companies accounted for are those employing a high proportion of "white-collar professionals", while the wage curve ceases to exist (at -0.012 and non-significant) where the only companies are those employing a high proportion of "white-collar clerical" employees. The large variation between these subgroups provides prima facie evidence that wages of specific high profile professions are extensively more flexible than wages of the clerical and medium-skilled labour supply. This pattern is also clearly present in nearly all sample subgroups.

Low-skilled clerical, sales and service employees, and to some extent blue-collar workers, are very likely to be destined to work only in the domestic market and are also quite unlikely to migrate to other occupational segments within the labour market. This could suggest that these groups of employees are more vulnerable and weaker in wage bargaining. As is clear from the results in Table 4, they do not enjoy the wage increases that unemployment elasticity of wages endows employees in high-skilled occupational segments (since our focus is mainly on the period of economic upturn).

Table 4: Preferred specification estimates by occupational breakdown

wdif	(7)	(7) wc-hs	(7) wc-ls	(7) bc
udif	-0.0798***	-0.1133***	-0.0114	-0.0725***
qdif	0.2868***	0.3336***	0.3588***	0.3258***
hrse	-0.2330***	-0.3793***	-0.2794***	-0.0697*
cafa	0.1665***	0.1666***	0.2754***	0.0625***
flex	-0.0798***	-0.4912***	-0.0442**	-0.4770***
blue	-0.0957***			
size	-0.0041*	-0.0195***	-0.0357***	0.0235***
whil	-0.1201***			
whih	0.1930***			
cons	1.8219***	3.5683***	2.1541***	0.8008**
N	22623	4109	3146	6111
N_g	8	8	8	8
r2_w	0.5518	0.4546	0.5571	0.4612

Note: (7) represents the specification (7) estimate, wc-hs in the white-collar high-skilled segment, wc-ls in the white-collar low-skilled segment (clerical, sales and service workforce), and bc in the blue-collar segment (workers); see Table 2.

We check the vulnerability assumption by using dummy variables[22] to control for wage cuts and decreases in the local unemployment rate. The results suggest that estimated elasticities remain almost intact, except in the case of clerical, sales and service employees, for which a moderately sloped wage curve is found only in a pseudo-bust environment (i.e. dummies for observed wage cuts).

This finding suggests two things: firstly, that the measured wage flexibility is relatively symmetric upwards and downwards and, secondly, that high-skilled workers enjoy wage increases in a tightened labour market, but also considerable wage decreases when the labour market conditions loosen. However, this finding must be considered only in respect of the wage measure used in our model, i.e. *de facto* compensation per hour worked.[23]

We also found some differences across the main sectors analysed (agriculture, industry, construction, business services and governmental services). The least flexible wages are reported in agriculture (-0.0495), although with considerable variability between urban and rural regions. In the largest segment of the economy (represented by industry and business services - 69% of the sample) unemployment elasticity of wages is close to the nation-wide elasticity.

Table 5: Preferred specification estimates by sector

wdif	(7)	(7) agr	(7) ind	(7) bus	(7) con	(7) gov
udif	-0.0798***	-0.0495***	-0.0729***	-0.0696***	-0.0973***	-0.1604***
qdif	0.2868***	0.1177***	0.3081***	0.3759***	0.3491***	0.1749***
hrse	-0.2330***	-0.5270***	-0.0940**	-0.3252***	-0.1598**	-0.6571***
cafa	0.1665***	-0.0171	0.0763***	0.1932***	0.0670*	0.1012**
flex	-0.0798***	-0.2611***	-0.6264***	-0.0598***	-0.1162	-0.9342***
blue	-0.0957***	-0.0598***	-0.0903***	-0.1202***	-0.0391***	-0.0051
size	-0.0041*	-0.0028	0.0222***	-0.0032	0.0097	-0.0612***
whil	-0.1201***	0.0411	-0.0795***	-0.0763***	0.1761**	-0.1436***
whih	0.1930***	0.2177***	0.0976***	0.2353***	0.0997***	0.1892***
cons	1.8219***	4.0820***	1.2266***	2.4641***	1.2528**	5.9860***
N	22623	3954	8611	7049	2269	740
N_g	8	8	8			8
r2_w	0.5518	0.4609	0.5742	0.6193	0.4447	0.446

Note: (7) agr represents the specification (7) estimate in the agriculture segment, ind in the industry segment, bus in business services, con in construction, and gov in the government segment.

The wage curve can be found in all the years analysed and its coefficient is stable throughout the time, ranging between -0.075 and -0.095. One finding, however, stands out: labour productivity (q) was reflected in wages at a decreasing rate across time. Observing the contradictory increasing effect of the foreign capital share in wages (*cafa*), while assuming that foreign-capital-intensive firms have higher labour productivity, it must be concluded that the distribution of gains to the workforce is ever less generous among locally owned firms.

[22] A new set of variables was defined for cases of negative annual wage change and the corresponding local rate of unemployment increase over the year.
[23] Since bonuses constitute a larger share of the compensation of high-skilled employees, the freezing of them is reflected in a more pronounced drop in the company wage bill.

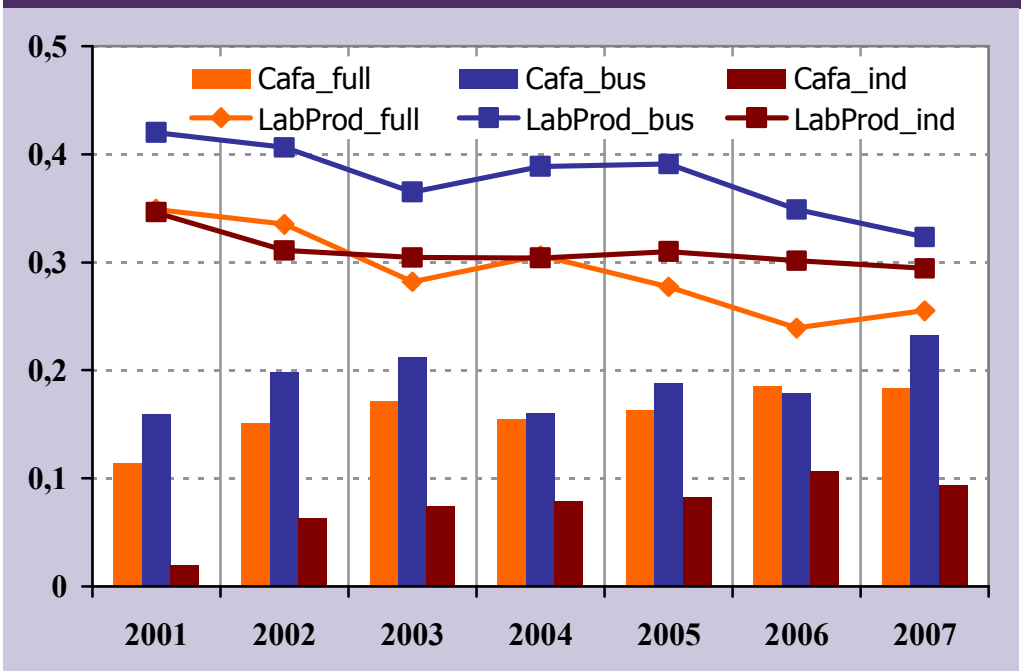
Table 6: Cross-section estimates in individual years

w	(7) 01	(7) 02	(7) 03	(7) 04	(7) 05	(7) 06	(7) 07
U	-0.082***	-0.092***	-0.077***	-0.092***	-0.080***	-0.075***	-0.092***
Q	0.349***	0.335***	0.282***	0.306***	0.277***	0.239***	0.255***
hrse	-0.338***	-0.269***	-0.188***	-0.302***	-0.128*	-0.250***	-0.336***
cafa	0.114***	0.151***	0.172***	0.155***	0.164***	0.185***	0.184***
flex	-0.581***	-0.336**	-0.381**	-0.054**	-0.057**	-0.405***	-0.758***
blue	-0.062***	-0.076***	-0.092***	-0.090***	-0.103***	-0.131***	-0.111***
size	-0.012**	-0.009	-0.004	-0.007	0.001	0.004	0.004
whil	-0.089***	-0.107***	-0.125***	-0.124***	-0.124***	-0.140***	-0.120***
whih	0.163***	0.171***	0.180***	0.197***	0.198***	0.228***	0.199***
cons	5.977***	5.320***	4.967***	5.399***	4.204***	5.649***	6.621***
N	2859	3129	3467	3470	3465	3233	3000
N_g	8	8	8	8	8	8	8
r2_w	0.61	0.589	0.553	0.565	0.551	0.52	0.535

Note: Suffixes 01 to 07 relate to the breakdown into the annual observations of 2001 to 2007.

This is especially valid in the business services sector, but less so in industry as reported in Chart 1.

Chart 1: Gains redistribution hypothesis for IND, BUS and overall sample



As regards separate estimates in NUTS 3 regions, the resulting elasticities were by far the most surprising. No wage curve was observed in two regions of central Slovakia (Žilina and Trenčín), and very low elasticities were observed in the other two. By contrast, the far western and far eastern regions in Slovakia are reported with elasticities of over -0.1. This suggests some of the following effects are likely to be in place:

1. Labour mobility (from the southwest to abroad and from east to west and to abroad) exploits regional and cross-border wage differentials, and this in turn affects employers' wage setting (with factors including infrastructure, tradition, educational community formation, or the proximity of wealthier regions across the border).

2. A return to education is starting to appear. Higher elasticities align closely with the presence of major universities in respective regions.

3. A discrepancy between lower-skilled industry job offers in central Slovakia and more services-oriented job offers in the west and east of Slovakia.

Table 7: Preferred specification estimates by region

wdif	(7) BA	(7) TT	(7) TN	(7) NR
udif	-0.1009***	-0.1634***	-0.0118	-0.0687***
qdif	0.3207***	0.2755***	0.2994***	0.2950***
hrse	-0.3040***	-0.2046***	-0.2705***	-0.2643***
cafa	0.2172***	0.1024***	0.0718***	0.1489***
flex	-0.0498**	-0.6664***	-0.8084***	-0.4382***
blue	-0.1033***	-0.0384***	-0.1128***	-0.0750***
size	-0.0495***	0.0247***	0.0132**	0.0013
whil	-0.0821***	-0.0281	-0.0928***	-0.0800***
whih	0.2616***	0.2143***	0.1256***	0.2098***
cons	2.5042***	2.0129***	2.7669***	2.3511***
N	4292	2764	2943	2904
r2	0.612	0.57	0.5641	0.5848
wdif	(7) ZA	(7) BB	(7) KE	(7) PO
udif	-0.0277	-0.0791***	-0.1265***	-0.1658***
qdif	0.2585***	0.2237***	0.2563***	0.2820***
hrse	-0.2766***	-0.3781***	-0.2163***	-0.2594***
cafa	0.1619***	0.1835***	0.1318***	0.1050***
flex	-0.8329***	-0.3320**	-0.5453***	-0.3923**
blue	-0.1271***	-0.1052***	-0.1360***	-0.0772***
size	0.0119*	0.0214***	-0.0007	0.0013
whil	-0.1422***	-0.1955***	-0.1435***	-0.1974***
whih	0.1683***	0.1089***	0.1506***	0.1223***
cons	2.8216***	3.0402***	2.1955***	2.4141***
N	2874	2131	2564	2151
r2	0.5514	0.5236	0.4969	0.5363

Note: The regional breakdown is based on the NUTS 3 classification; BA stands for Bratislava, TT for Trnava, TN for Trencin, NR for Nitra, ZA for Zilina, BB for Banska Bystrica, PO for Presov and KE for Kosice

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The validity of the above three points is also emphasised when the strong regional patterns are cross-checked with the occupational structure. The regional unemployment flexibility of wages is even higher in firms with a high proportion of high-skilled employees than in the overall sample documented in Table 8. At the same time, there is almost no difference in the wage curve in firms with a high proportion of lower-skilled white-collar (clerical) employees, and only some difference in firms with a high blue-collar workforce.

Table 8: Preferred specification wage curve coefficients across occupations and regions

	White collar - high	White collar – low	Blue collar
Regions BA,TT,PO and KE	- 0.183 ***	-0.090 ***	-0.086 ***
Regions NR,TN,ZA and BB	- 0.036 *	-0.073 ***	-0.040 **

These results reveal that higher-skilled jobs and higher-productivity activities tend to concentrate in the westernmost and easternmost regions of the country. Opportunities for labour mobility are generally greater for the higher-skilled workforce and therefore the wage curve coefficients found here are more negative. The opposite applies to the mostly rural and mountainous central regions of the country.

5 CONCLUSION

We have found clear evidence of the wage curve in the economy-wide enterprise-level microdata. The local rate of unemployment elasticity of wages in the full sample is revealed to be at -0.08 and is relatively stable over time. In other words, an increase in the rate of unemployment by 10% is found to be associated with an average decline in wages by 0.8%[24]. We found this elasticity to be rather symmetric in upturns and in cost-cutting business environments. Based on the extensive empirical literature, we conclude that wages possess a considerable degree of flexibility.

The wage curve is also present in most of the estimated segments of the labour market structure. It showed considerable stability over the history of our sample and also across major sectors, the only exception being agriculture, where the elasticity was revealed to be somewhat lower. We did, however, find more flexible wages associated with firms that have a more high-skilled workforce, which tend to be concentrated in the westernmost and easternmost regions of the country.

At the same time, estimates in multiple segments of the market provide evidence of higher wage flexibility in companies that employ more people on a part-time basis or on fewer hours. However, the effect of company size is not so clear and is often found to be insignificant.

The overall message about relatively flexible wages in Slovakia is broadly in line with previous findings in Gertler and Senaj (2008). However, a closer look at the individual segments based on labour market participants reveals that wage elasticity is quite substantially driven by the high-skilled segment of the labour market. On the other hand some segments have no wage curve or very low unemployment elasticity of wages (e.g. regions where firms with a high proportion of clerical, sales and service employees are more numerous, rural regions with a higher share of agriculture, etc.). Since these are usually also the most vulnerable groups/segments in economic distress, measures taken to make the labour market more flexible should keep this diversity in mind.

[24] Since variance of the rate of unemployment in Slovakia is very high across regions, we use a logarithmic scale also for the unemployment rate in all specifications. Therefore, 10% increase may in this setting be associated with an increase e.g. from 5 to 5.5% as well as from 15 to 16.5%.

6 ANNEX

Table A1: Test 1 – Blanchflower and Oswald

Fixed-effects (within) regression — Number of obs = 553
Group variable: REG — Number of groups = 79

R-sq: within = 0.9064 — Obs per group: min = 7
between = 0.9207 — avg = 7.0
overall = 0.9147 — max = 7

F(6,468) = 755.05
corr(u_i, Xb) = -0.1309 — Prob > F = 0.0000

w	Coef.	Std.err.	t	P>|t|	[95% Conf.	Interval]
w_lag	0.6612	0.0302	21.91	0.0000	0.6019	0.7205
u	-0.1423	0.0155	-9.16	0.0000	-0.1729	-0.1118
Q	0.0388	0.0063	6.12	0.0000	0.0263	0.0513
hrse	-0.3931	0.0985	-3.99	0.0000	-0.5867	-0.1996
cafa	0.0119	0.0203	0.59	0.5590	-0.0280	0.0518
flex	-0.4241	0.4648	-0.91	0.3620	-1.3375	0.4893
_cons	5.0992	0.9482	5.38	0.0000	3.2360	6.9624
sigma_u	0.0590					
sigma_e	0.0538					
		(fraction of variance due to u_i)				
		F(78, 468) = 1.79				

F-test that all u_i =0: Prob > F = 0.0001

Table A2: Test 2 – Card

Linear regression, absorbing indicators — Number of obs = 553
F(6, 540) = 13.87
Prob > F = 0.0000
R-squared = 0.1972
Adj R-squared = 0.1794
Root MSE = 0.0541

dw	Coef.	Std.err.	t	P>|t|	[95% Conf.	Interval]
u	-0.0333	0.0216	-1.54	0.125	-0.0757	0.0092
u_lag	0.0312	0.0228	1.37	0.173	-0.0137	0.0760
q	0.0400	0.0055	7.32	0.000	0.0293	0.0508
q_lag	-0.0370	0.0052	-7.09	0.000	-0.0473	-0.0268
hrse	-0.3673	0.1131	-3.25	0.001	-0.5895	-0.1452
hrse_lag	0.4578	0.1084	4.22	0.000	0.2448	0.6708
		F(6, 540) = 2.229		0.039	(7 categories)	

Table A3: Test 3 – Solon et al.

Linear regression, absorbing indicators — Number of obs = 546
F(2, 537) = 685.68
Prob > F = 0.0000
R-squared = 0.7186
Adj R-squared = 0.7144
Root MSE = 0.0517

REGdum	Coef.	Std.err.	t	P>|t|	[95% Conf.	Interval]
u	0.2742	0.0140	19.53	0.000	0.2466	0.3018
u_lag	-0.0546	0.0048	-11.48	0.000	-0.0639	-0.0452
		F(6, 537) = 76.777		0.000	(7 categories)	

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