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Regional labor demand and national labor market institutions in the EU15*

Helmut Herwartz[†] and Annekatriin Niebuhr[‡]

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Abstract

The labor market effects of the recent financial and economic crisis are rather heterogeneous across countries and regions. Such differences in labor market performance among industrialized countries are an issue of ongoing research. The objective of this paper is to analyse labor market disparities among European regions and to provide evidence on the factors behind these differences. Whereas previous research focused on effects of national labor market institutions, we also take structural characteristics of regions into account and investigate differences in labor demand responsiveness and their potential determinants. The data set covers the NUTS 2 regions in the EU15 for the period 1980 to 2008. We apply an error correction model that is combined with a spatial modeling approach in order to account for interaction among neighboring labor markets. Our findings point to substantially distinct labor demand responses to changes in output and wages among European countries and regions. Moreover, the rate of adjustment to disequilibrium is subject to a significant variation across units of observation. Whereas evidence on the significance of region specific variables as explanatory factors is weak, labor market institutions, especially regulations that affect the determination of wages, explain an important fraction of the disparities.

Keywords: Regional labor markets, labor demand, institutions, Europe, error correction model.

JEL Classification: C23, J23, R23.

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1 Introduction

The labor market effects of the recent financial and economic crisis are rather heterogeneous across countries and regions (see Eichhorst et al. 2010, Artha and de Haan 2011). Eichhorst et al. (2010) argue that the structure of the economy as well as labor market institutions likely influence the impact of the crisis. The pronounced disparities in labor market performance among industrialized countries and their potential causes are an issue of research for a long time. Differences in unemployment between European countries and the US are frequently attributed to more rigid labor market institutions in Europe (e.g. Nickell 1997). According to Decressin and Fatás (1995) there is an insufficient response of wages to shocks in Europe compared with the US. Blanchard and Wolfers (2000) and Bertola et al. (2002) have highlighted the role of institutions as potential determinants of these differences. In contrast, Solow (2000) stresses low output growth and a corresponding weakness of labor demand as primary factors behind the persistently high unemployment of several European economies. Evidence provided by Eichhorst et al. (2010) suggests, however, that the effects of GDP reductions on labor demand and unemployment during the recent economic crisis are marked by a considerable variation across countries.

Our analysis of labor market disparities among European regions draws on different strands of literature: studies that deal with demand for labor, research on the labor market effects of institutions and finally, investigations that provide evidence on regional labor market disparities in Europe. Previous research on demand for labor has focused on estimating labor demand elasticities based on industry level data.¹ Only a few studies consider the impact of labor market institutions on labor demand. Neumark and Wascher (2004) investigate the employment effects of minimum wages for a cross section of OECD countries. Buscher et al. (2005) examine the impact of different labor market institutions on aggregate labor demand for a sample of EU countries. Most studies that deal with the influence of labor market institutions focus on their significance with respect to the level of unemployment or long-run changes in unemployment (e.g. Nickell et al. 2005, Blanchard and Wolfers 2000) whereas research on regional labor market differences deals primarily with the spatial pattern and persistence of unemployment disparities (e.g. Overman and Puga 2002). Decressin and Fatás (1995) investigate regional labor market dynamics in Europe and the US. They provide empirical evidence on adjustment mechanisms to region specific shocks. Baddeley et al. (2000) analyse wage flexibility across EU regions and US states. and explore if regional differences in wage flexibility are associated with structural characteristics of labor markets. Herwartz and Niebuhr (2011) consider regional differences in Okun's law and provide evidence on the main regional and national triggers of these disparities. Elhorst (2003) notes that there is a lack of corresponding studies that integrate research on national and regional factors for European countries. Furthermore, evidence in Kosfeld and Dreger (2006) suggests that an analysis at the regional level has to account for spatial dependence since regions are linked by labor mobility, aggregate demand and other forms of interaction. Summing up, evidence on the impact of institutions and characteristics of regional labor markets on labor demand is scarce.

¹See Hamermesh (1986) for a survey.

Responding to this lack of comprehensive findings this analysis attempts to integrate the different strands of literature and investigates several aspects of regional labor market performance. The objective of this paper is to analyse labor market disparities among European regions and to provide evidence on both national and regional factors behind these differences. Regional labor markets in Europe exhibit considerable diversity and, thus, provide a good basis to uncover potential factors behind detected heterogeneity of labor market performance. The focus is on regional labor demand that is specified by means of an error correction model (ECM) coupled with a so-called spatial error specification (SEM, Anselin 1988) in order to account for spatial interaction among neighboring labor markets. We refrain from imposing strong homogeneity restrictions on the models' slope parameters and investigate if marginal responses of employment with respect to output and wages differ significantly among regions and countries in Europe. Moreover, the adjustment of labor demand in response to violations of the long run equilibrium is subjected to regional and national differentiation. Intuitively one may imagine that measurable cross sectional features impact on the responsiveness of employment to output and factor prices as well as on the speed of adjustment of labor demand. To uncover potential triggers of distinguished labor market performance we follow a two step approach: After estimating region specific adjustment coefficients and labor demand responses to output and factor prices these estimates are subjected to surface regressions conditioning on exogenous national and regional characteristics.

Our findings point to a considerable variation in the marginal responses of labor demand to output and wages across and within EU countries. Moreover, the rate of adjustment of labor demand to deviations from the long run equilibrium spreads across regional labor markets. The results of surface regressions suggest that different labor market institution matter for regional labor demand. Especially regulations that influence the determination of wages seem to play a significant role. In contrast, characteristics of regional labor markets appear to be of minor importance for differences in labor demand.

The next Section provides a brief view at the empirical literature on the links between labor market performance, institutions and structural characteristics of regional labor markets and motivates our choice of potential triggers of labor market performance. Section 3 introduces the error correction model with spatial error distribution (ECM/SEM) and the strategy employed to uncover potential factors behind the region specific model specification. The data is introduced in Section 4. Section 5 provides the empirical results. Concluding remarks are given in Section 6.

2 Triggers of region specific labor market performance

An extensive literature has examined the effects of labor market institutions on labor market outcomes.² Altogether, this literature has documented that rigid institutions tend to increase unemployment. If a specific regulation exerts an adverse or beneficial effect on market performance depends, however, on the type of the considered institution. Most studies point to significant effects of the unemployment benefit system. According to

²See Herwartz and Niebuhr (2011) for a discussion with respect to Okun's Law.

Nickell and Layard (1999) generous and long-lasting benefit entitlements generate higher unemployment. Unemployment insurance reduces the search effort and increases reservation wages, thereby resulting in fewer matches between employers and workers and in fewer offers being accepted. Checchi and García-Peñalosa (2008) argue that high labor tax rates have similar effects because labor taxation increases labor costs and thereby reduces labor demand. Furthermore, strong trade unions are expected to raise unemployment. Evidence in Eichhorst et al. (2010) suggests that countries with a low collective bargaining coverage tend to be characterized by a relatively high wage flexibility. The adverse effects of strong trade unions might, however, be offset if wage setting is characterized by highly coordinated bargaining (Nickell and Layard 1999).

Findings are less clear-cut for other institutions. There is no unambiguous evidence that stricter labor standards and employment protection legislation (EPL) result in higher unemployment. Since EPL reduces the risk of job loss and shifts associated costs from workers to employers the latter might refrain from firing in downturns but also from hiring in booms. Thus, the overall effect of EPL on employment is ambiguous. Bertola et al. (2002) detect a significantly positive impact of EPL on unemployment, however, Nickel et al. (2005) argue that this correlation mainly operates via the effect of EPL on unemployment persistence. There is also no clear indication that minimum wages affect employment growth or unemployment. Eichhorst et al. (2010) also point to the capacity of labor market institutions to absorb shocks. The benefit system, for instance, might act as an automatic stabilizer during recessions. Agell (1999) concludes that although it seems likely that certain institutions adversely affect labor market performance, others might give rise to beneficial effects.³

Apart from the institutional settings other factors are most likely to impact on regional labor market performance. Firstly, the sectoral structure of the economy possibly affects the responsiveness of labor demand to output growth. Economies specialized in services tend to be characterized by a relatively high employment intensity of growth since labor productivity is low in this sector. The same argument applies to construction. In the light of productivity differences across industries, one might expect stronger marginal responses of labor demand to output in regions that are characterized by high shares of technologically less developed branches (see e.g. Mourre (2004)).

Secondly, referring to the reallocative aspect of growth emphasized in Aghion and Howitt (1994) structural change may influence the labor market effects of output variations. The implementation of new technologies requires labor reallocation, i.e. permanent shifts of labor demand between sectors. Thus, the intensity of structural change might matter since - with given labor market flexibility - regions characterized by more pronounced reallocation of jobs between industries face higher adjustment burdens. Matching frictions can arise because of industry-specific skills. Skill requirements in expanding branches may not coincide with skills possessed by workers laid off in declining industries (Petrongolo and Pissarides 2001). Therefore the marginal effect of output changes on labor demand might be smaller in economies characterized by fast structural change. However, this effect could be offset if structural change goes along with a relative strengthening of

³Checchi and García-Peñalosa (2008) note that some institutions might also have conflicting labor market effects.

employment intensive service industries.

Thirdly, agglomeration economies most likely affect the matching process in regional labor markets. The likelihood of a match possibly improves when more agents try to match. In case of increasing returns to scale a proportional increase in the number of job seekers and vacancies results in a more than proportional increase in job matches. More vacant jobs and job seekers reduce search frictions on local labor markets and the proportion of unemployed workers (Duranton and Puga 2004). Benefits of a matching function that exhibits increasing returns to scale accrue in dense urban labor markets. Therefore, one might expect that output growth results in more pronounced increases of employment in highly agglomerated regions in comparison with rural labor markets.

Moreover, drawing upon suggestions in Elhorst (2003) the age structure of the labor force, labor market participation and the educational attainment of the population as potential factors could contribute to the distinguished performance of regional markets. A relatively young and well-educated work force might allow a flexible and more rapid adjustment to shocks. Similarly, regional patterns of long-term unemployment could affect employment responses to changes in output and wages due to matching frictions and insider power in wage setting.

In this paper, we analyse regional patterns of the speed of adjustment of labor demand to deviations from the long run equilibrium and the long run effects of wages and aggregate demand on employment. With regard to potential triggers of (regional) dynamics of labor demand, we draw upon the quoted literature since labor market institutions, the sectoral structure and other characteristics of regional labor markets are natural candidates for a conditional description of region specific employment patterns. We investigate if the rate of adjustment of employment and the labor market responsiveness to output and wage fluctuations depend on labor market institutions (unemployment benefit system, the system of wage determination and EPL) and on structural characteristics such as the sectoral composition of the economy, the region type or the intensity of structural change.⁴

3 The spatial regression model

3.1 Model representation

The empirical analysis rests on a neoclassical framework where cost minimization of firms subject to a production constraint gives rise to an expression for labor demand as a function of planned output and factor prices. Addison and Teixeira (2005) note that the ECM has become the standard approach to investigate the dynamic characteristics of labor demand. In the regression analysis, we use employment as a proxy for labor demand. Let l_{it} denote the log of employment in region i and year t . For a cross section comprising $N = 192$ European regions the following ECM is applied to investigate regional labor

⁴A detailed list of considered factor variables and their definitions is given in the Appendix

demand

$$\begin{aligned}\Delta l_{it} &= \mu_i + \alpha_i(l_{it-1} - \beta_{1i}q_{it-1} - \beta_{2i}w_{it-1} - \beta_{3i}r_{it-1}) \\ &+ \gamma_{1i}\Delta q_{it} + \gamma_{2i}\Delta w_{it} + \gamma_{3i}\Delta r_{it} + e_{it}, \quad i = 1, \dots, N, t = 1, \dots, T.\end{aligned}\quad (1)$$

In (1) q_{it} is the log real output, w_{it} and r_{it} are the log of the real prices of labor and capital, and Δ is short for the first difference operator, e.g. $\Delta l_{it} = l_{it} - l_{i,t-1}$. Presample values are available by assumption. The model (1) formalizes labor demand adjustment in response to lagged violations of an equilibrium relationship linking labor, output and factor prices conditional on contemporaneous growth of output and factor prices, i.e. the latter are assumed to be weakly exogenous.⁵ The parameter α_i , $-2 < \alpha_i < 0$, governs the degree of correcting recent violations of market equilibrium. In case of $\alpha_i = -1$ the long run equilibrium is reestablished within one period. Thus, the larger is $\alpha_i^* = |\alpha_i + 1|$ the more persistent are deviations from the equilibrium relation. Respecting the cross sectional dimension the model fits into the framework of seemingly unrelated regressions (Zellner 1962) if some parsimonious representation of contemporaneous error correlation is available.

Making use of vector (e.g. $\Delta l_t = (\Delta l_{1t}, \Delta l_{2t}, \dots, \Delta l_{Nt})'$) and stacked vector (e.g. $\Delta l = (\Delta l'_1, \Delta l'_2, \dots, \Delta l'_T)'$) representation the model in (1) reads as

$$\begin{aligned}\Delta l_t &= \underset{(N \times 1)}{\boldsymbol{\mu}} + \underset{(N \times 1)}{l_{t-1}} \odot \underset{(N \times 1)}{\boldsymbol{\alpha}} + \underset{(N \times 1)}{q_{t-1}} \odot \underset{(N \times 1)}{(\boldsymbol{\beta}_1 \odot \boldsymbol{\alpha})} \\ &+ \underset{(N \times 1)}{w_{t-1}} \odot \underset{(N \times 1)}{(\boldsymbol{\beta}_2 \odot \boldsymbol{\alpha})} + \underset{(N \times 1)}{r_{t-1}} \odot \underset{(N \times 1)}{(\boldsymbol{\beta}_3 \odot \boldsymbol{\alpha})} \\ &+ \underset{(N \times 1)}{\Delta q_t} \odot \underset{(N \times 1)}{\boldsymbol{\gamma}_1} + \underset{(N \times 1)}{\Delta w_t} \odot \underset{(N \times 1)}{\boldsymbol{\gamma}_2} + \underset{(N \times 1)}{\Delta r_t} \odot \underset{(N \times 1)}{\boldsymbol{\gamma}_3} + \underset{(N \times 1)}{e_t},\end{aligned}\quad (2)$$

$$\begin{aligned}\Delta l &= \underset{(TN \times 1)}{(j_T \otimes I_N)} \underset{(TN \times N)}{\boldsymbol{\mu}} + \underset{(TN \times N)}{\mathbf{l}_-} \underset{(N \times 1)}{\boldsymbol{\alpha}} + \underset{(TN \times N)}{\mathbf{q}_-} \underset{(N \times 1)}{(\boldsymbol{\beta}_1 \odot \boldsymbol{\alpha})} + \underset{(TN \times N)}{\mathbf{w}_-} \underset{(N \times 1)}{(\boldsymbol{\beta}_2 \odot \boldsymbol{\alpha})} \\ &+ \underset{(TN \times N)}{\mathbf{r}_-} \underset{(N \times 1)}{(\boldsymbol{\beta}_3 \odot \boldsymbol{\alpha})} + \underset{(TN \times N)}{\Delta \mathbf{q}} \underset{(N \times 1)}{\boldsymbol{\gamma}_1} + \underset{(TN \times N)}{\Delta \mathbf{w}} \underset{(N \times 1)}{\boldsymbol{\gamma}_2} \\ &+ \underset{(TN \times N)}{\Delta \mathbf{r}} \underset{(N \times 1)}{\boldsymbol{\gamma}_3} + \underset{(TN \times 1)}{\mathbf{e}}, \quad \mathbf{e} = (e'_1, e'_2, \dots, e'_T)',\end{aligned}\quad (3)$$

$$= \underset{(TN \times 8N)}{\mathbf{Z}} \underset{(8N \times 1)}{\boldsymbol{\phi}} + \underset{(TN \times 1)}{\mathbf{e}},\quad (4)$$

where

$$\begin{aligned}\mathbf{Z} &= [(j_T \otimes I_N), \mathbf{l}_-, \mathbf{q}_-, \mathbf{w}_-, \mathbf{r}_-, \Delta \mathbf{q}, \Delta \mathbf{w}, \Delta \mathbf{r}], \\ &\text{and } \boldsymbol{\phi} = (\boldsymbol{\mu}', \boldsymbol{\alpha}', (\boldsymbol{\beta}_1 \odot \boldsymbol{\alpha})', (\boldsymbol{\beta}_2 \odot \boldsymbol{\alpha})', (\boldsymbol{\beta}_3 \odot \boldsymbol{\alpha})', \boldsymbol{\gamma}_1', \boldsymbol{\gamma}_2', \boldsymbol{\gamma}_3')'.\end{aligned}$$

In (2) ' \odot ' indicates element-by-element vector multiplication, and vectors $\boldsymbol{\mu}$, $\boldsymbol{\alpha}$, $\boldsymbol{\beta}_j$, $\boldsymbol{\gamma}_j$, $j = 1, \dots, 3$, collect cross section specific model parameter, e.g., $\boldsymbol{\beta}_j = (\beta_{1j}, \beta_{2j}, \dots, \beta_{Nj})'$. In (3), j_T is a $T \times 1$ vector of ones, Δl and \mathbf{e} are TN dimensional vectors and e.g. \mathbf{q}_-

⁵When discussing estimation results we will reconsider the issue of weak exogeneity

and Δq are 'blockdiagonal' $TN \times N$ dimensional matrices. Matrices like \mathbf{q}_- consist of observations that are lagged by one period, i.e. $\mathbf{q}_- = \text{blockdiag}(q_0, q_1, \dots, q_{T-1})$.

The introduction of cross section specific parameters $\mu_i, \beta_{ji}, \gamma_{ji}, j = 1, 2, 3$, or α_i may follow a deterministic or random coefficient approach. A fixed effect approach could be justified in light of the time dimension of available data covering a period of $T = 28$ annual observations. A key purpose of this study is to uncover cross sectional patterns explaining the marginal responses of labor demand to changes in output and wage level. We follow the idea that in particular the model parameters β_{1i}, β_{2i} and $\alpha_i^* = |\alpha_i + 1|$ depend in a systematic fashion on measurable national and regional characteristics. Given that we do not a-priori know the most effective factor variables governing labor market responsiveness, first step parameter estimates are subjected to surface regressions on exogenous variables. For efficiency of first step estimates we allow for potential spatial correlation characterizing error terms e_{it} . Distributional assumptions are addressed in the next Section.

3.2 Correlation pattern and feasible GLS

As in Herwartz and Niebuhr (2011) we assume that error vectors e_t in (2) are contemporaneously correlated, i.e.

$$E[e_t e_t'] = \Omega \quad (5)$$

($N \times N$)

In light of the large cross section dimension which exceeds the time dimension by a factor of more than 6 it is not possible to estimate Ω without structural assumptions. Single region regressions reveal that (estimated) residual variances differ markedly across regions. Moreover, geographic distance is likely governing the dependence of error terms. For both reasons, diagnosed heteroskedasticity and spatial correlation, we construct estimates of Ω from (unrestricted) cross sectionally heterogeneous variances combined with a correlation pattern built from a spatial weights matrix. To be explicit,

$$\Omega = \Sigma^{1/2} R \Sigma^{1/2}, \quad \Sigma = \text{diag}(\sigma_1^2, \sigma_2^2, \dots, \sigma_N^2), \quad \sigma_i^2 = E[e_{it}^2], \quad (6)$$

with R denoting the correlation matrix associated with Ω . To implement GLS estimation the a-priori presumed correlation pattern and variance estimates are, respectively,

$$\hat{R} = \text{Cor} \left((\hat{B}' \hat{B})^{-1} \right), \quad \hat{B} = I_N - \hat{\rho} W, \quad \text{and} \quad \hat{\sigma}_i^2 = \frac{1}{T - |Z|/N} \sum_{t=1}^T \hat{e}_{it}^2. \quad (7)$$

In (7) $|Z|$ is the column dimension of Z , W is a known spatial $N \times N$ weights matrix with zero diagonal elements and $\hat{\rho}$ is an estimated scalar spatial correlation parameter, $-1 < \rho < 1$. By construction the rows of W sum to unity. As characterized by (2) and (6) the ECM/SEM can be estimated by means of the following feasible GLS estimator:

$$\hat{\phi} = (Z' \hat{\Omega}^{-1} Z)^{-1} Z' \hat{\Omega}^{-1} \Delta l, \quad \hat{\Omega}^{-1} = I_T \otimes \hat{\Omega}^{-1}. \quad (8)$$

Under multivariate normality the log likelihood conditional on ρ is

$$\ln L = -\frac{NT}{2} \ln 2\pi - \frac{T}{2} \ln |\widehat{\Omega}| - \frac{1}{2} \sum_{t=1}^T \hat{e}'_t \widehat{\Omega}^{-1} \hat{e}_t. \quad (9)$$

We use an iterative feasible GLS approach. At each step of the iteration a grid search over parameters $\tilde{\rho} = 0.01r$, $r = -99, -98, \dots, 99$, (Cochrane-Orcutt 1949) is employed to determine the correlation pattern that maximizes the log likelihood in (9). Providing an outer loop for adjusting the cross sectional variances the feasible GLS estimator is obtained after convergence of the log likelihood. The log likelihood estimates offer guidance to determine the most preferable choice of the spatial weights matrix W in (7) from a set of alternative suggestions. The common OLS estimator corresponding to (8) is obtained with choosing $R = I_N$, or equivalently $W = 0$.

3.3 Surface regressions

From first step feasible GLS regressions we obtain N -dimensional vectors of coefficients $(\hat{\alpha}^*, \hat{\beta}_1, \hat{\beta}_2)$. In a second modeling step, these parameter estimates are conditioned on measurable cross sectional characteristics. A priori it is unlikely that all these conditioning variables govern cross sectional parameter variation jointly such that the determination of a suitable subset model becomes an essential modeling step. In light of the large dimension of the set of potential conditioning variables we employ a specific-to-general strategy building on Lagrange Multiplier (LM) specification tests (Godfrey 1988). Herwartz (2010) recommends the specific-to-general strategy especially for cases with relatively small sample sizes.

After initializing $S = \{j_N\}$ we pursue the following strategy to choose the 'most informative' set $S = \{j_N, s_1, s_2, \dots, s_M\}$ out of potential covariates $\tilde{S} = \{\tilde{s}_k, k = 1, \dots, K, K \geq M\}$ to explain measurements $\hat{\beta}_1 = (\hat{\beta}_{11}, \hat{\beta}_{12}, \dots, \hat{\beta}_{1N})'$, $\hat{\beta}_2 = (\hat{\beta}_{21}, \hat{\beta}_{22}, \dots, \hat{\beta}_{2N})'$ or $\hat{\alpha}^* = (\hat{\alpha}_1^*, \hat{\alpha}_2^*, \dots, \hat{\alpha}_N^*)'$. Referring to $\hat{\beta}_1$ the sequential selection includes the following steps:

1. Regress $\hat{\beta}_1$ on S , and obtain the implied residuals, $\hat{\varepsilon} = \hat{\beta}_1 - S\hat{q}$, $\hat{q} = (S'S)^{-1}S'\hat{\beta}_1$.
2. Estimate regressions of $\hat{\varepsilon}$ on sets of variables $S_k = \{S, \tilde{s}_k\}$, $k = 1, \dots, \tilde{K}$, where \tilde{K} is the column dimension of \tilde{S} . For all regressions compute the degree of explanation, R_k^2 , and an LM measure of the marginal explanatory content of \tilde{s}_k , i.e. $\lambda_k = NR_k^2$.
3. The set of explanatory variables in S is augmented with the covariate \tilde{s}_{k^*} that obtains the maximum LM-statistic if λ_{k^*} exceeds the $(1 - \delta)$ quantile of a $\chi^2(1)$ distribution. In this case the particular variable is removed from \tilde{S} .
4. Steps 1) to 3) are iterated until the largest LM statistic is insignificant at level δ . Then, the variables in S are regarded 'most effective' in explaining $\hat{\beta}_1$. Noting that sequential testing sacrifices control over the exact significance level we will refer to δ as the tuning level of the selection procedure.

4 Data

We investigate regional labor demand for a cross section of 192 EU15 regions (NUTS 2 level) based upon annual data on employment, real Gross Value Added (GVA) and compensation per employee for the period 1980 to 2008.⁶ The corresponding information is taken from the European regional database of Cambridge Econometrics (CE) which, in turn, draws upon the EUROSTAT Regio database and official data from national providers. The data on national interest rates are collected from the annual macroeconomic database (AMECO) of the European Commission's Directorate General for Economic and Financial Affairs (DG ECFIN). Missing data of Sweden, Portugal and Greece are replaced by data extracted from the International Financial Statistics (IFS) database of the International Monetary Fund (IMF).

As outlined in Section 2 numerous factors might affect the marginal response of labor demand to changes in output and wage level. In the surface regressions, we consider country specific and region specific influences. Out of the full set of 37 conditioning variables 9, respectively 10, factor measures are incomplete as quotes for regions in Greece, respectively Greece and Ireland, are not available. In principle, surface regressions become infeasible or involve considerable loss of information if for particular cross section members factor observations are not available. Since for most factor variables cross sectional data are complete we decide in favor of a simple nearest neighbor imputation technique as described in Herwartz and Niebuhr (2011). The factor variables allow a classification in two categories that are briefly described in turn.

Country specific factors At the country level we consider the OECD indicators of labor market institutions (OECD 2004) that refer to three main areas: the unemployment benefit system, wage determination and employment protection (Nickell et al. 2005). Measurable features of the benefit system include the unemployment benefit replacement ratio, the duration of entitlement and expenditures on active labor market policies. Variables referring to wage determination comprise collective bargaining coverage, union density, a coordination index that captures the extent to which bargaining is coordinated and information on extension laws. Moreover, the OECD index of the strictness of EPL, its variation in the 1980s, and the total labor tax rate are included in the factor set. We allow these factors to influence the marginal responses of labor demand to GDP and wages as well as the adjustment speed.

Region specific factors Potential regional determinants of labor demand responsiveness include the percentages of regional employment across 15 distinct industries. In the surface regressions, sector specific time means of annual employment shares enter as (potential) explanatory variables. As an indicator for the intensity of structural change we

⁶Exceptions to the NUTS 2 level include Denmark (3 former NUTS regions) and Germany (East German regions and Berlin excluded). Furthermore, Départements d'outre-Mer (France), Açores, Madeira (Portugal), Ceuta y Melilla, Canarias (Spain) are not considered because of data restrictions. Information on compensation per employee is available at NUTS 1 level only for Germany and Sweden. In these cases we assigned NUTS 1 wage data to corresponding NUTS 2 observations in the first stage regressions.

include the sum of absolute annual changes in employment shares between 1980 and 2008 across all industries. Employment data is taken from the CE database.

To address agglomeration economies we rely on a density indicator derived from population figures and a partition of EU regions into spatial categories. Conditional on population density and the size of regional centers three groups of regions are distinguished (agglomerated, urbanized and rural regions). Moreover, accessibility of a region is considered by means of population potential. The classification scheme and the accessibility measure are taken from the database established by the Study Program on European Spatial Planning (SPESP).⁷

Further factor variables are considered as proxies for structural characteristics of regional labor markets. These include the age structure (shares of young and old workers in total work force), the educational attainment (share of high- and low-skilled workers), participation rates and the share of long-term unemployment in total unemployment. The labor market indicators originate from the EUROSTAT Regio database.

We restrict effects of region specific factors primarily to the employment intensity of output changes. In contrast, most of the regional factors should be of minor importance for wage flexibility and speed of adjustment. In the corresponding surface regressions we only consider the share of long-term unemployment in total unemployment as a potential explanatory factor at the regional level as it might reflect insider power in wage setting.

5 Empirical results

5.1 Model selection and diagnostics

OLS diagnostics Cross sectional heteroskedasticity is diagnosed by means of a F -ratio of two cross section specific OLS variance estimates $f_{ij} = \hat{\sigma}_i^2 / \hat{\sigma}_j^2$, $i \neq j$. In total we perform $N(N - 1)/2 = 18336$ pairwise comparisons of estimated error variances. With 5% significance the null hypothesis of homoskedasticity is rejected if f_{ij} is too small or too large in comparison with the .025 or .975 quantile of the respective F -distribution. The actual rejection rate is 50.1%. Given the likelihood of cross section specific variances we decide in favor of the (generalized) ECM/SEM outlined in Section 3.2.

OLS regressions deliver 192 sequences of estimated error terms subjected to testing for first order serial correlation by means of an LM test (Godfrey 1978). The actual rejection rate for testing with 5% significance is 25.0%. Given the moderate time dimension, finding 25.0% significant test statistics does not mirror 'severe' model misspecification.⁸ Therefore we do not further discuss potential specification improvements for single region models.

Spatial correlation Feasible GLS estimation in (8) requires some presumption on the structure of spatial interaction. For choosing among alternative settings of W the (max-

⁷See SPESP indicator set: <http://www.bbr.bund.de/raumordnung/europa/espon.htm>.

⁸It is noteworthy that individual test decisions are not independent as a consequence of spatial error correlation. Therefore, even under the null hypothesis the empirical rejection frequencies do not necessarily correspond to the nominal level of the test.

imized) log-likelihood as given in (9) provides useful guidance. We consider three alternative 'raw' adjacency matrices \widetilde{W} in this analysis. A first and frequently applied specification is a binary spatial weights matrix such that $\tilde{w}_{ij} = 1$ if the regions i and j share a border and $\tilde{w}_{ij} = 0$ otherwise. Secondly, \tilde{w}_{ij} is set to the inverse of travel time between the capitals of regions i and j . Thirdly, we use the inverse of travel time for regions within the same country and set $\tilde{w}_{ij} = 0$ for regions located in distinct countries. From the 'raw' matrices \widetilde{W} spatial weights matrices W are derived by normalizing such that $\sum_j w_{ij} = 1, \forall i$.

The following display documents alternative log-likelihood estimates of the ECM/SEM and provides a comparison with the OLS regression:

OLS	W binary	W travel	W travel border
21059.64	21727.67	21865.58	21915.54

Apparently, spatial models offer a marked improvement of fitting accuracy and in particular, specifying the weight matrix with (inverse) travel times within countries offers the most accurate data description. The choice of the weights matrix 'travel border' is also in line with economic intuition. Spatial dependence of regional labor market performance is likely caused by migration and commuting. As both forms of labor mobility are influenced by frictional effects of geographical distance, a weighting scheme based on inverse distance should be a more appropriate specification than a simple binary weights matrix. However, labor mobility across national borders is still low in the EU15. This suggests to restrict spatial interaction to national labor markets. The differences in log-likelihood estimates between 'travel' and 'travel border' reflect this feature of EU15 labor markets.⁹

Weak exogeneity The model outset in (1) is a cross sectional collection of single equation ECMs formalized under the assumption of weak exogeneity of output and factor prices. To address the issue of potential endogeneity of Δq_{it} , Δw_{it} or Δr_{it} we perform ECM/SEM type regressions for these variables conditioning on the presumed equilibrium relation, a constant and one autoregressive lag. Interestingly all these ECM/SEM implementations are in favor of the weights matrix 'travel border' and support the SEM in comparison with OLS. Over the cross section of 192 adjustment parameters we obtain the median estimates and intervals covering 90% of the estimators, 0.206 [−0.219; 0.682] (output), 0.201 [−0.246; 0.548] (wages) and 0.003 [−0.004; 0.011] (capital cost). Although these parameter estimates are skewed to the right and thereby hint at potential endogeneity for single covariates or regions, they hardly call for a fully fledged multivariate ECM/SEM approach. Taking the respective statistics for the adjustment of labor, −0.392 [−0.799; −0.164], it is apparent that labor demand reacts most strongly and uniformly to violations of longer term relations between labor demand, output and factor prices.

With regard to biases invoked by incidental endogeneity it is noteworthy that adjustment parameters for labor change only mildly if the model in (1) is fully specified in

⁹Similar evidence with respect to Okun's law is provided in Kosfeld and Dreger (2006) as well as Herwartz and Niebuhr (2011).

terms of predetermined variables. Conditioning labor changes on the presumed equilibrium, lagged changes of output and factor prices, and an autoregressive component the median and 90% range of adjustment coefficient estimates is $-0.435 [-0.969; -0.074]$.

5.2 Distributional features of estimated first stage coefficients

The systematic differences in log-likelihood estimates between non-spatial and the spatial models point to the importance of spatial dependence as regards regional labor demand. Therefore we focus on the estimates of the spatial ECM for weights matrix 'travel border' in the following. Table 1 summarizes the distributional features of selected coefficients included in (1). We only display the adjustment coefficient and the long run effects of GDP and factor prices.¹⁰ At the mean group level of inference the error correction parameter is negative at common significance levels and, moreover, the documented quantiles for this coefficient strongly indicate that the ECM formalizes intuitive adjustment dynamics for the vast majority of cross sections.¹¹ The mean EC parameter is -0.392. This corresponds with evidence in Addison and Teixeira (2005). They estimate the speed of employment adjustment for Portugal (-0.342) and Germany (-0.108). The signs of the average elasticity parameters correspond with theoretical expectation. As indicated by the interquartile range or 95% coverage intervals the distribution of parameter estimates for GDP is concentrated to the right from zero. The same applies to the region specific wage coefficients with the majority of estimates being located left from zero. Moreover, the size of the effects in absolute terms is in line with some previous estimates of labor demand functions. The median coefficient estimates for output and wage level are 0.596 and -0.380 respectively. Estimates of demand responses to wages range between -0.04 and -1.09 in a survey by Hamermesh (1986). Falk and Koebel (2001) apply a dynamic labor demand model to German manufacturing and detect wage effects on labor demand between -0.10 and -0.21. Corresponding results in Buscher et al. (2005) vary between -0.08 and -0.99 for a cross section of European countries. Output coefficients in Buscher et al. (2005) exceed 0.5 (range from 0.53 to 1.64). Hamermesh (1993) reports long run output effects from 0.03 to 0.98. In summary, our estimated coefficients are well within the range that can be found in the related literature.

As shown in the Figures 1 to 3 the estimated coefficients $\hat{\alpha}_i$, $\hat{\beta}_{1i}$ and $\hat{\beta}_{2i}$ are marked by a substantial variation across both, regions and countries. Figure 1 displays the cross country and within country dispersion of the adjustment coefficient. There is both variation at the country and at the regional level. At the national level Luxembourg (-0.061) reveals a much slower rate of adjustment to disequilibrium than the UK (-0.559). Country specific medians of GDP coefficients vary between 0.250 for Spain and 0.863 for the Netherlands. As regards the impact of wages on employment we get a rather small effect for Luxembourg (-0.054), whereas wages seem to play a prominent role for regional labor

¹⁰Regression results for the short run coefficients are available upon request.

¹¹Interestingly, smallest EC parameters in absolute value are obtained for the two NUTS2 regions in Ireland. For these regions we fail to diagnose cointegration and, as a consequence, the model implied long run output and wage elasticities are rather large in absolute value. To keep graphical displays of estimation results at a reasonable scale, later, results for Irish regions are excluded from respective figures.

Table 1: Coefficient estimates for labor demand functions

	MG		quantiles								
	$\bar{\cdot}$	$t(\bar{\cdot})$.025	.05	.10	.25	.50	.75	.90	.95	.975
	ECM/SEM										
$\hat{\alpha}_i$	-0.421	-31.42	-0.894	-0.799	-0.677	-0.490	-0.392	-0.312	-0.246	-0.164	-0.092
$\hat{\beta}_{1i}$	0.594	22.01	-0.165	0.113	0.201	0.362	0.596	0.764	1.020	1.184	1.382
$\hat{\beta}_{2i}$	-0.422	-15.50	-1.170	-1.118	-0.785	-0.554	-0.380	-0.243	-0.152	-0.066	0.017
$\hat{\beta}_{3i}$	0.437	8.10	-0.114	-0.048	0.011	0.094	0.231	0.443	1.034	1.423	2.167

Notes: The table provides average estimated coefficients ($\bar{\cdot}$), a corresponding mean group (MG) t -ratio and selected quantiles of the unconditional distribution of cross sectional ECM/SEM estimates $\hat{\alpha}_i$ (error correction parameters) and cointegration parameters $\hat{\beta}_1$ (output), $\hat{\beta}_2$ (wages) and $\hat{\beta}_3$ (capital cost). The spatial weights matrix is 'travel border', $\hat{\rho} = 0.90$.

demand in the Netherlands (-0.809). Although there is some variation of wage coefficients within specific countries, intra-national dispersion seems to be small compared with cross country variance. This becomes even more apparent if we compare Figure 2 and Figure 3. Country specific output coefficients show considerable variation as well. However, relative to differences at the country level disparities within countries seem to be more pronounced when compared with intra- and international variance of wage coefficients. This might be interpreted as a first indication for the significance of national factors such as labor market institutions in explaining differences in wage coefficients. In contrast, the distributional features of the output coefficients suggests that labor market institutions alone cannot account for the entire variation of estimates since they cannot capture the distinct disparities between coefficients of regions located in the same country.

5.3 Determinants of regional labor demand responses

Having described the distributional properties of cross sectional labor market responsiveness we now characterize the cross sectional shape by means of surface regressions. When implementing surface regressions it turned out that the outcome of the model selection strategy depends on some outlying estimates that are also depicted in the Figures 1 to 3. In light of this dependence we decide to remove the C largest and smallest parameter estimates from the sample of cross sectional estimates. The number of removed tail estimates is chosen as $C = 5, 8, 10$. Consequently, the following results are 'representative' for a cross section of 171 up to 181 regions since Luxembourg is also excluded from the surface regressions owing to missing factor observations. In order to provide some evidence on the robustness of the surface regressions we discuss results for two values of C and the two applied tuning levels.

Table 2 documents the surface regression estimates \hat{q}_{fin} for the adjustment coefficients along with standard t -statistics and heteroskedasticity consistent t -ratios (White 1980). Furthermore the partial degrees of explanation achieved by singular covariates are given.

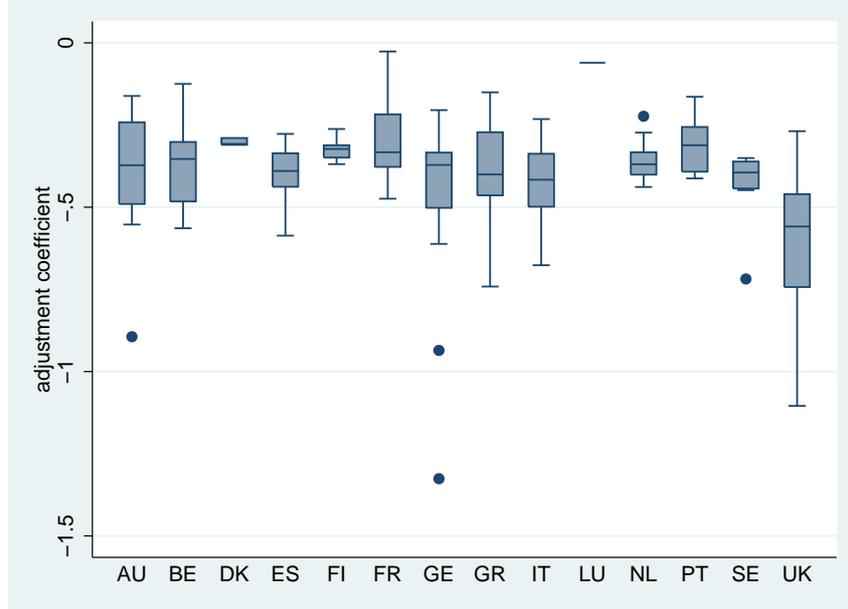


Figure 1: Distribution of estimated error correction parameters ($\hat{\alpha}_i$, ECM/SEM, 'travel border', $\hat{\rho} = 0.90$) within and across European countries. Austria AU (9 regions), Belgium BE (11), Germany GE (30), Denmark DK (3), Finland FI (5), France FR (22), Greece GR (13), Italy IT (20), Luxemburg LU (1), Netherlands NE (10), Portugal PT (5), Sweden SE (8), Spain ES (16), United Kingdom UK (37). The box represents the interquartile range of estimated adjustment coefficients and the horizontal line corresponds to the median. The whiskers mark the last observation within 1.5 times the interquartile range. Estimates outside this support are classified as outliers.

With regard to sample trimming we mostly discuss results for $C = 5$ since this choice obtains most selected factors. Although the sequential selection of explanatory variables forgives exact control over the statistical significance level, it is noteworthy that almost all robust t -ratios exceed absolute values of 2.3, say, underscoring significance at conventional levels. This also refers to the corresponding results of the wage and output coefficients (see Tables 3 and 4). However, it is not only interesting to consider the documented coefficient estimates but also to keep in mind that particular factor variables have been ruled out according to apparently low explanatory content.

The results of the surface regression for adjustment coefficients after trimming with $C = 5$ and $\delta = 0.01$ indicate that some labor market institutions indeed affect the speed of adjustment of labor demand in response to deviations from the long run equilibrium. Both selected factors reduce the speed of adjustment according to our results. The transformed adjustment coefficient α_i^* increases, on average, by 0.12 if the wage determination of the country is subject to extension laws. Considering that the average adjustment parameter estimate is $\bar{\alpha} = -0.42$, this is a rather strong effect. Moreover, the estimates for alternative

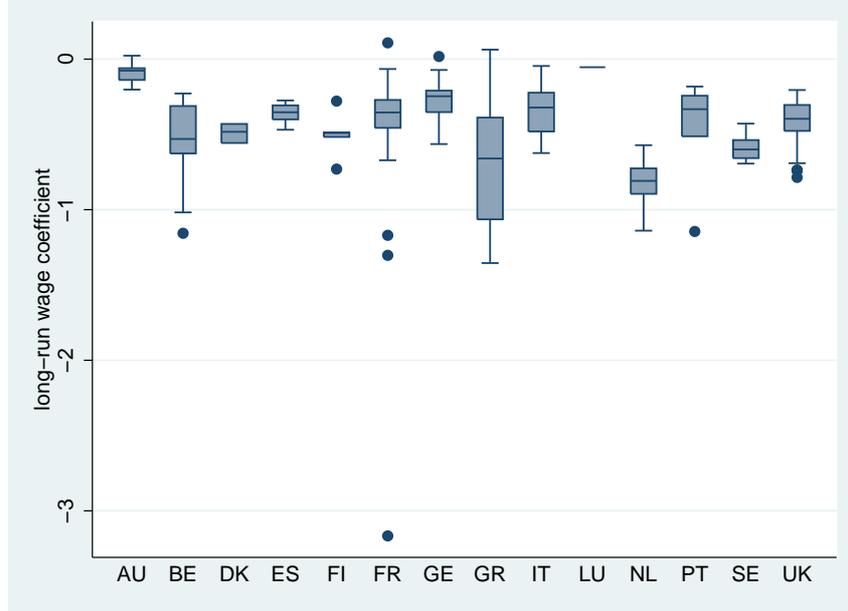


Figure 2: Distribution of estimated responses to wages ($\hat{\beta}_{2i}$, ECM/SEM, 'travel border', $\hat{\rho} = 0.90$) within and across European countries. For further notes see Figure 1.

combinations as regards the number of excluded outliers (C) and tuning levels (δ) indicate that this seems to form the lower limit of the impact. Altogether the influence of extension laws is fairly robust and the partial degree of explanation ($R_p^2 = .11$) suggests that it is an important factor - although the R_p^2 varies across the displayed combinations of C and δ .

The second institutional factor that impacts on the adjustment speed of labor demand is EPL. The result is in line with theoretical arguments since the positive and significant coefficient indicates that strict legislation tends to reduce the flexibility of labor demand. The OECD employment protection index ranges from 0.35 (UK) to 1.93 (Portugal). This variation in EPL translates into a difference in the adjustment coefficient of 0.05 if we apply the estimate for $C = 5$ and $\delta = 0.01$ and 0.17 for the combination $C = 5$ and $\delta = 0.05$. The R_p^2 of the factor varies between 6% and 10%. Finally, the selection of the unemployment benefit replacement ratio for two out of three constellations points to the role of the benefit system as regards a swift adjustment of labor demand to shocks. Again considering the size of the effect as implied by the cross country variation of the ratio (in 1999 from 0.17 in the UK to 0.74 in Sweden) we arrive at a difference in the rate of adjustment of 0.06. Moreover, the average replacement ratio increased by 5 percentage points between 1980 and 1999. This gives rise to an average increase of the transformed adjustment coefficient by merely 0.005 (0.9% of the average ECM/SEM estimate $\bar{\alpha}$). Thus, compared with the impact of the extension laws these effects are rather small. Moreover, the evidence on an important influence of extension laws is more robust than for the other

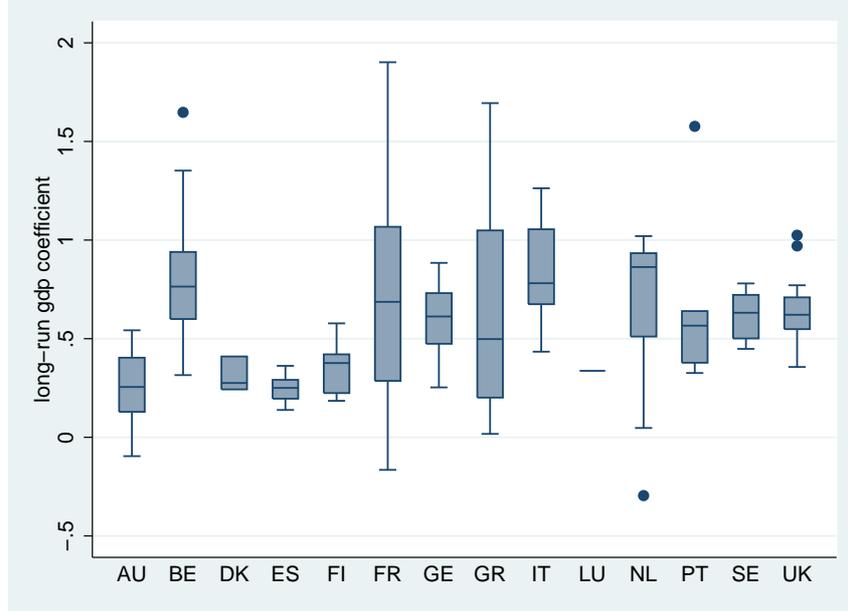


Figure 3: Distribution of estimated responses to output ($\hat{\beta}_{1i}$, ECM/SEM, 'travel border', $\hat{\rho} = 0.90$) within and across European countries. For further notes see Figure 1.

two selected factors.

In Table 3 the results of surface regressions for the wage coefficients $\hat{\beta}_{2i}$ are summarized. The factor selection for the region specific responses to wages seems to be even more robust with respect to a variation of C and δ than the corresponding findings for the adjustment coefficients. The group of selected factors does not change at all for the three combinations.¹² A comparison of partial degrees of explanation in Table 3 suggests that primarily active labor market policy matters for labor demand responses to wage changes. According to the partial R^2 more than 30% of the variation of region specific wage coefficients can be explained by cross country differences in the expenditure on active labor market policies. Thus, it is by far the most important factor as regards the marginal employment response to wages. The negative and highly significant coefficient of the factor indicates that countries that spend a relatively large share of GDP on such measures, e.g. Denmark (1.7% of GDP in 1998), tend to achieve more pronounced employment effects for a given change of the wage level than countries such as the UK or Austria (0.34% and 0.44% of GDP, respectively) where labor market policies play a much less prominent role. Applying the maximum spread of GDP shares we get a difference in the long run wage

¹²Only for $C = 8$ we detect slight changes with respect to the selected factors. We restrict the presentation of results in Table 3 to the combinations with $C = 5$ since deviations in estimates for $C = 10$ are marginal. The selection procedure identifies the same factor variables with only slight changes of their estimated impacts as a consequence of distinct trimming. The regression results for all other combinations of C and δ are available upon request.

Table 2: Surface regressions for region specific rate of adjustment

Var.	Description	\hat{q}_{fin}	t -rat	rob. t	R_p^2	p_{LM}
$C = 5, \delta = 0.01$						
-	Constant	0.426	19.30	16.16	0.629	0.00
6	Extension laws	0.118	4.674	4.767	0.108	0.00
9	EPL late 1980s	0.029	3.294	3.689	0.057	0.13
$C = 5, \delta = 0.05$						
-	Constant	0.391	14.34	12.03	0.504	0.00
6	Extension laws	0.137	2.156	2.418	0.025	0.00
9	EPL late 1980s	0.111	4.404	4.657	0.098	0.13
1	Unemployment benefit replacement rat.	0.021	2.212	2.725	0.027	3.14
$C = 10, \delta = 0.01$						
-	Constant	0.451	17.41	15.88	0.613	0.00
6	Extension laws	0.142	2.615	2.685	0.039	0.00
1	Unemployment benefit replacement rat.	0.099	4.603	4.399	0.111	0.97

Notes: Estimation results are coefficient estimates \hat{q}_{fin} , ' t -rat' and 'rob t ' are the common and heteroskedasticity consistent t -ratios (White 1980). R_p^2 is the partial degree of explanation associated to particular factor variables. The nominal level of the selection procedure is δ . The ordering of variables reflects the sequential outcome of the specific-to-general selection described in Section 3 and particular p -values ($\times 100$) of the LM-statistic are also given (p_{LM}). More detailed information on the factor variables is given in the Appendix.

coefficients of 0.51. This even exceeds the average wage coefficient (-0.42).

The other institutional factors selected at the second stage of our regression analysis adversely affect the marginal wage effect. We get a dampening effect of the total labor tax rate on the impact of wages on employment. This factor explains around 5% of the wage coefficient's variance. The tax rate is rather low in Ireland (mean rate in the period 1996-2000: 33%), whereas in Scandinavian countries and especially in Sweden taxation of labor is much higher. In the latter country the tax rate on average amounted to 77% between 1996 and 2000. A decline of the labor tax rate by 1 percentage point will, according to our estimates reduce the wage coefficient by 0.098. This correspond with a decrease of the average wage coefficient by more than 20%. Furthermore, the estimates suggest that extension laws affect the labor demand response to wage changes. The partial R^2 (19%) points to an important contribution of this factor to the explanation of cross country differences in the wage coefficient. However, extending bargained wages to non-union firms increases the marginal wage effect by only 0.012. This is less than 3% of the average wage coefficient.

Finally, Table 4 shows the results of the selection procedure for the region specific output coefficients. Although the number of selected factors is large compared with the surface estimates for the adjustment and the wage coefficient robust evidence on influential factors is scarce. Both national factors and region specific features are selected and seem

Table 3: Surface regressions for long run region specific labor demand responses to wage

Var.	Description	\hat{q}_{fin}	t -rat	rob. t	R_p^2	p_{LM}
	$C = 5, \delta = 0.01; C = 5, \delta = 0.05 (C = 10, \delta = 0.01)$					
-	Constant	-0.732	-8.625	-8.742	0.288	0.00
3	Active labor market policy	-0.383	-9.800	-9.751	0.349	0.00
8	Total labor tax rate	0.098	3.188	3.239	0.054	0.00
6	Extension laws	0.012	6.562	6.666	0.194	0.17

Notes: See Table 2.

to affect the size of labor demand responses to output variations. The estimates for the three combinations of C and δ are characterized by a considerable variation with respect to selected factors, signs of coefficients and size of effects. Evidence on region specific factors is altogether weak with only three corresponding indicators being selected. The estimated coefficient of fuels and chemicals ($C = 5, \delta = 0.05$) implies that a specialization in this branch reduces the impact of output changes on labor demand. This is consistent with the productivity argument put forth in Section 2. Moreover, the participation rate and the share of low skilled labor are chosen for some combinations of C and δ . However, the selection of all regional factors is not robust with respect to a variation of sample trimming (C) and tuning level (δ). Due to this lack of robustness we refrain from a detailed discussion of the corresponding results.

The results suggest that labor market institutions are the main drivers of cross regional variation of output coefficients since corresponding findings and especially evidence on the coordination of wage bargaining are fairly stable. The factor is selected in all combinations. There are no changes of the sign of the coefficient across distinct constellations and only minor changes of the size. The partial degree of explanation indicates that the coordination index of wage bargaining ($R_p^2 = .16$ for $C = 5, \delta = 0.01$) is an influential factor with respect to labor demand responses to output. Coordination of wage bargaining tends to increase the impact of GDP on employment. The positive effect of coordination corresponds with the role of coordinated bargaining discussed in the literature. Nickell and Layard (1999) argue that the extent to which bargaining is coordinated might matter for labor demand since highly coordinated bargaining can offset adverse effects of unionism. By implication, countries marked by a high degree of consensus between the actors in collective bargaining such as Germany and Austria should, *ceteris paribus*, realize stronger impacts of output changes on employment than Italy that shows only a medium degree of coordination.

There is more plausible evidence on significant effects of different labor market institutions. Extension laws and EPL are associated with a decline of the output coefficient. In line with theory and previous empirical evidence high union density tends to reduce labor demand effects of output variations. Since these factors enter the surface regression only for specific trimming (C) and tuning levels (δ) we do not put too much emphasis on them and refrain from a discussion of the size of corresponding effects. The same applies to the impact of the unemployment benefit replacement ratio. Although this factor is selected

for all combinations summarized in Table 4 a consistent interpretation is impeded by the changing sign of the coefficient.

Table 4: Surface regressions for long run region specific labor demand responses to output

Var.	Description	\hat{q}_{fin}	t -rat	rob. t	R_p^2	p_{LM}
$C = 5, \delta = 0.01$						
-	Constant	1.387	8.520	6.923	0.286	0.00
7	Coordination index	0.595	5.730	5.511	0.155	0.00
1	Unemployment benefit replacement ratio	-0.269	-9.082	-13.14	0.315	0.00
17	Participation rate	-0.009	-3.253	-2.864	0.056	0.14
$C = 5, \delta = 0.05$						
-	Constant	1.062	5.455	3.812	0.144	0.00
7	Coordination index	0.610	5.302	4.822	0.138	0.00
1	Unemployment benefit replacement rat.	0.004	3.351	3.557	0.060	0.00
17	Participation rate	0.160	2.444	2.031	0.033	0.14
10	Change EPL	-0.297	-7.786	-9.231	0.257	2.82
5	Union density	-0.506	-3.243	-3.331	0.057	1.04
6	Extension laws	-0.005	-1.484	-1.115	0.012	2.74
25	Fuels and Chemicals	-2.827	-2.072	-2.118	0.024	3.63
$C = 10, \delta = 0.01$						
-	Constant	1.200	15.84	16.87	0.593	0.00
7	Coordination index	0.414	4.471	4.055	0.107	0.00
1	Unemployment benefit replacement rat.	0.210	5.081	4.545	0.134	0.00
6	Extension laws	-0.474	-11.65	-13.37	0.448	0.18
9	EPL late 1980s	-0.616	-4.324	-4.097	0.101	0.11
15	Share low skilled workers	0.084	5.199	4.845	0.139	0.00

Notes: See Table 2

6 Conclusions

We detect a considerable variation in the marginal responses of labor demand to output and wages across and within EU countries. Moreover, the speed of adjustment of labor demand to deviations from the long run equilibrium spreads across regional labor markets. The dispersion within specific countries tends to be more pronounced for output coefficients than for wage effects and the rate of adjustment. The variance of region specific employment effects points to national as well as regional factors as potential causes. But evidence on significance of region specific factors is rather weak. There is some indication for the sectoral structure of regional labor markets and the participation rate might matter for employment effects of output. However, corresponding findings are not robust. The poor performance of region specific characteristics does not, however, imply that the regional dimension is irrelevant in this context. Our findings show that spatial dependence matters for regional labor demand. Taking into account interaction among neighboring regions significantly improves the efficiency of regression results. Thus, an analysis of regional labor demand has to consider the impact of labor mobility and demand linkages.

Whereas structural characteristics of regional labor markets seem to be of minor importance for the responsiveness of labor demand, there is much evidence on the significance of labor market institutions. We detect important effects on different indicators of labor market performance, more precisely on the marginal effects of output and wages on labor demand as well as on the speed of adjustment. This suggests that focusing on unemployment only as variable of interest might present a too narrow perspective of the issue. The findings also indicate that specific institutions seem to influence labor market outcomes via different channels. This applies especially to extension laws. Moreover, we discover distinct differences in the size of effects associated with various labor market institutions.

Our findings point to adverse and beneficial effects of different institutions on labor demand that are almost uniformly in line with theoretical expectations. This corresponds with the differentiated evidence on employment effects of institutions summarized in Blau and Kahn (1999). Institutions that influence the determination of wages, in particular extension laws and the coordination of the wage bargaining process, are important according to our results. The extension of bargained wages to non-union firms seems to affect the speed of adjustment of labor demand and the long run effects of wages and output on employment. The beneficial effect of coordination on the marginal response of employment to GDP corresponds with findings of previous studies. EPL impacts on the rate of adjustment of labor demand. This effect might rest on its influence on layoffs and hirings and conforms to the role of EPL discussed in the literature. Institutions likely influence short run adjustment in the labor market since the ability of firms to adjust production to changes of aggregate demand might be subject to institutional restrictions. Moreover, there seem to be significant effects of the unemployment benefit system. We identify adverse effects of the unemployment benefit replacement ratio on the speed of adjustment of labor demand. As regards the long run wage effects on employment the taxation of labor and active labor market policies seem to matter. The last result confirms arguments in Eichhorst et al. (2010) who argue that active labor market programmes mitigated adverse labor market effects of the sharp decline of GDP during the recent crisis.

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Appendix

A.1 Data

CE regional database (NUTS 2 level); annual data 1980 to 2008
Employment data by region and industry (in 1000 employees); see A.2 for details on classification)
Unemployment rate (number of unemployed as a percentage of the labor force)
Gross value added (10E06 EUROS, 1995 prices)
Population density (1000 inhabitants per km^2)
Compensation per employee (in EURO)
SPESP database (NUTS 2 level)
Classification scheme region types (agglomerated, urbanized, rural)
Population potential accessible by road 1996 (10E06 inhabitants)
Gross value added (10E06 EUROS, 1995 prices)
EUROSTAT Regio database (NUTS 2 level); annual data 1999 to 2002
Working population by age (15-24 years, 45-54 years, 55-64 years)
Working population by highest level of education attained according to International Standard Classification of Education (ISCED) 1997 (ISCED0 2: Pre-primary, primary and lower secondary education, ISCED3 4: Upper secondary and post-secondary non-tertiary education, ISCED5 6: Tertiary education)
Total and long term unemployed (12 months and more)
Participation rate (in percent)
OECD indicators on labor market institutions (see Nickell et al. 2005 for a detailed description of the data)
Benefit replacement rate (Benefit entitlement as a percentage of previous earnings, data refers to first year of unemployment)
Benefit duration index
Expenditure on active labor market policy (Expenditure as a percentage of GDP)
Collective bargaining coverage (Percentage of employed labor force whose pay is determined by collective agreements)
Extension laws (Dummy variable indicating that bargained wages are extended to non-union firms at the behest of one bargaining party)
Trade union density (Union members as percentage of employees)
Coordination index (Consensus of collective bargaining actors, from low (1) to high(3))
Employment protection index (Strictness of EPL).
Labor taxes (Payroll plus income plus consumption tax rates)

Notes: Due to differences in data availability and delineation of regions between the various data sets some information from SPESP and data on working population had to be adjusted for the following regions: Hovedstadsregionen, Ost for Størebælt, ex.Hovedst, Vest for Størebælt, Vlaams Brabant, Brabant Wallon, Oost-Nederland, Scotland and Wales. Weighted averages are assigned to corresponding NUTS 2 regions.

A.2 Employed

factors explaining estimates of labor market responsiveness ($\hat{\beta}_{ji}$, $j = 1, 2$ and $\alpha_i^* = |\hat{\alpha}_i + 1|$)

Var.		Description
1	x	Unemployment benefit replacement ratio (mean 1980-1999)
2	x	Unemployment benefit duration index (mean 1980-1999)
3	x	Expenditure on active labor market policies (mean 1985-1998)
4	x	Collective bargaining coverage (mean 1980-1994)
5	x	Union density (mean 1980-1998)
6	x	Dummy variable extension laws
7	x	Coordination index (mean 1980-1995)
8	x	Total labor tax rate (average percentage 1980-2000)
9	x	OECD employment protection index late 1980s
10	x	Change of EPL index 1980-1987
11	x	Share of long-term unemployed in total unemployment (mean 1999-2002)
12		Accessible population by road 1996 (in 10E06 inhabitants)
13-14		Share of age groups in working population (mean 1999-2002); 13: 15-24 years; 14: 45-64 years.
15-16		Shares of low/high skilled in working population (mean 1999-2002), 15: low, i.e. pre-primary, primary and lower secondary; 16: high, i.e. tertiary education
17		Participation rate (mean 1999-2002)
18-19		Regiontypes with agglomerated regions as reference, 18: urban; 19: rural
20-35		Mean employment shares 1980-2008, 20: Agriculture; 21: Construction; 22: Electronics; 23: Financial services; 24: Food, beverages and tobacco; 25: Fuels, chemicals, rubber and plastic products; 26: Hotels and restaurants; 27: Mining and energy supply; 28: Non-market services; 29: Other manufacturing; 30: Other market services; 31: Textiles and clothing; 32: Transport and communications; 33: Transport equipment; 34: Manufacturing; 35: Market services
36		Mean population density 1980-2008 (1000 inhabitants per km ²)
37		Structural change (mean sum of absolute changes in employment shares across all industries 1981 to 2008)

Notes: For factors 20 to 35: Employment shares for region i and industry j are given by $\frac{1}{T} \sum_{t=1}^T \frac{L_{jit}}{L_{.it}}$, where L_{jit} is employment in industry j in region i and time t and $L_{.it}$ is total employment in region i and time t . The structural change variable for region i (factor 37) is defined as $\frac{1}{T} \sum_{t=1}^T \sum_{j=1}^J \left| \frac{L_{jit}}{L_{.it}} - \frac{L_{jit-1}}{L_{.it-1}} \right|$. Factor quotes for regions in Greece are not available for factors 1-10. Moreover, factor 4 is not reported for regions in Ireland. The left hand side column distinguishes two sets of factors, the larger set consists of all 37 factors and has been used for the modeling of output response. The smaller set consists of those factors indicated with 'x' and has been used to model wage responsiveness and the intensity of error correction dynamics.

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