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ABSTRACT

Local Employment Growth in West Germany: A Dynamic Panel Approach*

In this paper we study the dynamics of local employment growth in West Germany from 1980 to 2001. Using dynamic panel techniques, we analyse the timing of the impact of diversity and specialisation, as well as of the human capital structure of local industries. Diversity has a positive effect on employment growth in the short run, which is stronger in manufacturing than in services. Concerning specialization we find evidence for mean reversion, which is inconsistent with the idea that growth emphasizes itself. But there is considerable inertia in this process. A positive effect of education is only found in manufacturing. Additionally, we look at the impact of firm size and regional wages on local employment growth.

JEL Classification: R11, O40

Keywords: regional labour markets, externalities, local employment growth, dynamic panel estimation, urbanization and localisation effects

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1) Introduction

What economic structure is conducive for the employment growth performance of different industries at the local level? The seminal paper of Glaeser et al. (1992) argues that a local industry thrives if it faces a diversified surrounding economic structure, whereas the study of Henderson et al. (1995) finds that own industry specialisation is the major employment growth engine. Understanding the impact of the local economic structure is a crucial issue for policymakers that try to tailor specific regional development programmes. If own industry specialisation increases job creation, policies appear promising that aim at promoting "regional clusters" with the intention of a self-sustained employment growth take-off due to local concentration. The same policies seem less appropriate if job creation is primarily fostered by a diversification of the regional production composition.

The empirical local growth literature has considerably grown in recent years, with mixed results on the relative importance of diversity versus specialisation for different countries (for a recent survey, see Combes and Overman 2004). Most of this literature, however, has neglected an issue that seems equally important: What is the *timing* of the impact? Is it the *current* local economic structure that matters for current employment growth, or rather some historical pattern that influences things like the "image", or the "business climate" of a specific location? If the former turns out to be the case, regional policies might become effective immediately. In the latter case the impact might be slower, but also longer lasting.

Conventionally this literature has used a pure cross-section approach, as long-run employment growth rates are regressed on control variables that reflect the regional industry composition in some base year.[1] It is thus assumed that a historical pattern from 10-30 years ago affects employment growth, but no real test is provided about the relevant time structure. To be able to do this, one needs data of local industries for many consecutive years in order to make full use of the three dimensions of the panel (location, industry, time period). An additional advantage of panel techniques is the possibility to control for time invariant fixed effects that can not be easily disentangled from the impact of the local economic structure in a cross-section analysis.

In the present paper we analyse both, the nature and the timing of the impact of the local economic structure for employment growth by means of dynamic panel estimation. We draw on a 22-year panel, covering the complete population of full-time employees in West Germany from 1980 to 2001, disaggregated into local (manufacturing and service) industries

[1] Both Glaeser et al. (1992) and Henderson et al. (1995) are cross-sections, as well as the influential study on France by Combes (2000). Among this literature is also the paper by Blien and Suedekum (2005) on Germany (1993-2001) that draws on a less comprehensive version of the same data set that is used in this paper.

at the level of NUTS3-regions. It turns out that timing is in fact a crucial issue. Rather than some historical pattern, it is the current and very recent economic structure that affects employment growth. Concerning the nature of the impact we find a positive effect of diversification, both in manufacturing and services. The evidence for the impact of specialisation is mixed. Employment growth rates in West Germany exhibit mean reversion, which is inconsistent with the idea that past growth feeds on itself due to local concentration/specialisation. But there is considerable inertia involved in this process.

Our paper is most closely related to two contributions that have also explicitly looked at the timing issue. Henderson (1997) was probably the first who went in this direction. He draws on a 14-year panel (1977-1990), covering employment in 5 manufacturing industries across US urban counties and finds evidence for a positive impact of both, specialisation and diversity. The relevant lag structure is 6-7 years in the former, an even longer time horizon in the latter case. The second important contribution comes from Combes, Magnac and Robin (2004), who study the employment growth of 36 different (manufacturing and service) industries in 341 French local areas between 1984 and 1993. The authors are able to decompose local industry employment into average plant size and the number of plants, thereby distinguishing between the growth of existing, and the creation of new plants. They find a positive effect of the current diversity of the surrounding economic environment.

In the present paper we can not observe employment at the plant level. Therefore our empirical model is closer to Henderson (1997), while taking a specification problem into account that has been spelled out by Combes (2000) and that leads to an overstatement of the impact of specialisation. Apart from providing novel evidence for Germany, Europe's largest economy, an additional contribution of the present study is that we look at two issues that have not yet been addressed in this literature. First, unlike previous studies we have information about workers' qualification structure in the different local industries, which does not depend on the local economic structure. Its inclusion in the empirical analysis allows observing whether the impact of diversity and specialisation remains robust. As it turns out, human capital has a positive short-run impact on employment growth (at least in manufacturing), but it leaves the conclusions with respect to the local economic structure unaffected. Second, we will analyse the impact of wages in a more elaborate way compared to previous studies. If at all, unsettled wages were directly plugged into the employment growth regression (e.g. Glaeser et al., 1992; Henderson et al., 1995). Instead of including unsettled wages, we use the methodology of Blien and Suedekum (2004) to develop a measure of the "neutralized" regional wage level, which is detached from various productivity influences.

In sum, our empirical analysis points at a robust positive short-run impact of industrial diversity on employment growth, which largely corroborates the findings by Combes, Magnac and Robin (2004) for the case of West Germany.

Note that we have avoided the terms Jacobs- and Marshall-Arrow-Romer (MAR)-externalities, which are commonly used in this literature. Knowledge spillovers are an important theoretical rationale why the local economic structure should have an effect on industry growth. It is known that external knowledge flows rapidly decay with distance (Jaffe el al., 1993; Audretsch and Feldman, 1996), hence local environments are good natural laboratories to study their precise nature (Lucas, 1988). Starting with Glaeser et al. (1992), researchers have taken a positive impact of diversity on employment growth as evidence for (inter-industry) Jacobs-externalities, whereas a positive impact of own industry specialisation was set equal with (intra-industry) MAR-externalities.[2] Cingano and Schivardi (2004) argue that this identification is problematic, however, because knowledge spillovers affect productivity, not employment directly. Cingano and Schivardi (2004), as well as Dekle (2002) and Henderson (2003) consequently use output and productivity instead of employment data to study the nature of urban knowledge spillovers. In particular, Cingano and Schivardi (2004) receive conflicting results in productivity and employment regressions using Italian data. With productivity growth as the dependent variable, only specialisation is found to matter. If employment growth is used instead, diversity comes to dominate, and the positive effect of specialisation vanishes. This leads them to conclude that employment growth regressions are ill-suited for studying the nature of externalities.

With respect to the employment growth regression presented in this paper, one has to be careful interpreting the results on diversity and specialisation as evidence for or against a particular theory of knowledge spillovers. Combes, Magnac and Robin (2004) derive the conditions under which a positive productivity shock leads to an increase in equilibrium employment. The elasticity of goods demand must be sufficiently elastic (see also Appelbaum and Schettkat, 2001 for a similar result), and the reaction is stronger the more elastic labour supply. Inference about externalities based on employment growth regressions is viable only to the extent that these conditions are met. More generally, however, our empirical results can also be seen as non-structural estimations on the sources of local employment growth related to the underlying local economic structure.

The rest of the paper is organized as follows. Section 2 describes the empirical model, data, and the specification of variables. Results are presented in section 3. Section 4 concludes.

[2] For an intensive discussion of urban knowledge spillovers, see Duranton and Puga (2004).

2) The model

2.1. The basic model

We rely on the following estimation equation that represents a dynamic panel setup.

$$emp_{z,s,t} = \alpha + \sum_{\ell=1}^{m} \rho_{\ell}\, emp_{z,s,t-\ell} + \sum_{\ell=0}^{m} \delta_{\ell}\, X_{z,s,t-\ell} + U_{z,s} + D_t + \varepsilon_{z,s,t} \qquad (1)$$

$emp_{z,s,t}$ is the log scale of industry s ($s = 1,\ldots,S$) in area z ($z = 1,\ldots,Z$) at time t ($t = 1,\ldots,T$). $emp_{z,s,t-\ell}$ ($l = 1,\ldots,m$) are the lagged dependent variables, $X_{z,s,t-\ell}$ are the (current or lagged) time variant characteristics (in logs) that are discussed at length below. $U_{z,s}$ is a fixed time invariant location and industry specific effect, and D_t is a general time effect. The standard error term is denoted $\varepsilon_{z,s,t}$.

As Nickell (1981) has shown, the standard within-group estimate is biased and inconsistent in the dynamic panel model, because of a correlation between the transformed error term $\varepsilon_{z,s,t} - \bar{\varepsilon}_{z,s}$ and the transformed endogenous variable $emp_{z,s,t-1} - \overline{emp}_{z,s}$ with $\bar{\varepsilon}_{z,s} = (1/T)\sum_{t=1}^{T} \varepsilon_{z,s,t}$ and $\overline{emp}_{z,s} = (1/T)\sum_{t=1}^{T} emp_{z,s,t-1}$. Following Arellano and Bond (1991) we use GMM method to get consistent estimates for the unknown coefficients. We take first differences to get rid of the time invariant effect $U_{z,s}$, so we obtain

$$\Delta emp_{z,s,t} = \sum_{\ell=1}^{m} \alpha_{\ell} \Delta emp_{z,s,t-\ell} + \sum_{\ell=0}^{m} \delta_{\ell}\, \Delta X_{z,s,t-\ell} + \Delta d_t + \Delta \varepsilon_{z,s,t} \qquad (2)$$

where $\Delta emp_{z,s,t-\ell} = emp_{z,s,t-\ell} - emp_{z,s,t-\ell-1}$.[3] This transformation allows us to use values of $emp_{z,s,t}$ (lagged twice or more) as instruments (Anderson and Hsiao, 1982; Arellano and Bond, 1991). Crucial for the validity of these instruments is the assumption about the order of autocorrelation of the error term. Under the assumption of serially uncorrelated $\varepsilon_{z,s,t}$, the first differenced error terms $\varepsilon_{z,s,t} - \varepsilon_{z,s,t-1}$ follow a MA(1) process, so $emp_{z,s,t-p}$ ($p = 2,3,\ldots$) are valid instruments for $\Delta emp_{z,s,t-1}$. Furthermore, we assume that the remaining right-hand side variables $X_{z,s,t}$ are strictly exogenous with respect to $\varepsilon_{z,s,t}$, i.e.

[3] Since $emp_{z,s,t}$ is measured in logs, the left-hand side of (2) is (approximately) the employment growth rate.

$$E(\Delta\varepsilon_{z,s,t} \mid \Delta X_{z,s,t'}) = 0 \quad \forall t,t' \in [1,\ldots,T].$$

Test statistics for these assumptions are presented below.

2.2. Data

The data for this study is provided by the German Federal Employment Agency (Bundesagentur fuer Arbeit). This highly reliable official information covers the entire territory of West Germany, and the complete population of full-time employment relationships subject to social security (i.e. excluding civil servants and self-employed individuals) between 1980 and 2001. Employment is observed in 326 NUTS3-districts ("Landkreise" and "kreisfreie Städte")[4]. Data refer to the workplace location, hence there are no upward biases in the income levels of metropolitan areas due to inward commuting. Furthermore, the data is not subject to any censoring. For every district-industry and every year we know

- the total employment level
- the employment shares in small (<20 workers), medium-sized (20-99) and large (>100) establishments
- the employment shares of three skill categories (without formal vocational qualifications, completed apprenticeship, higher education)
- the average age of the employees
- the fraction of men
- the average wage income per employee per calendar day, including all bonuses and extra payments subject to social security.

Two things should be noted with respect to the income data. Firstly, income levels that exceed the threshold for social security contributions are reported with this value. Our data therefore understates the true wage dispersion in West Germany. Secondly, although we deflate the wages and work with prices of 1977, we are restricted to use a common price deflator for all districts (the CPI for West Germany), because price level data and price indices are not available on a regional level.

[4] For obvious historical reasons, East Germany is not part of the analysis. Moreover, we also excluded West Berlin from the data.

We can distinguish 28 different industries in each district, but we perform our analysis only for 21 of them. Specifically, we do not analyse employment growth in agriculture, mining, the public sector and four basic service industries (e.g. household-related services, gastronomy). The reason is that these service sectors are strongly locally oriented, and specialisation in these non-tradable activities seems hardly possible. We end up looking at 15 manufacturing industries, and 6 advanced service sectors (like banking) that produce tradable goods, in which a region could reasonably specialise in, and for which final goods prices are determined on nationwide, or even international markets.[5] In the empirical analysis we could estimate industry by industry, but in order to restrict the number of results we lump all manufacturing and all advanced service industries together and only perform global regressions for the two broad classes of activities.

2.3. Specification of explanatory variables

In this subsection we discuss our collection of explanatory variables. Recall that the estimation equation is formulated in first differences. Therefore, all time invariant fixed effects are eliminated. Furthermore, because the exogenous variables are in logs, we analyse the impact of their growth rates on the growth rate of the local employment. Particular emphasis is put on those variables that are related to the local economic structure. The additional controls that capture the impact of firm size, qualification and neutralized wages are described in e-g.

a) Sector specific effects

To control for pure sector effects such as structural change at the national level, we include the total size of sector s across all $Z = 326$ districts without the own regional employment:

$$sect_{s,t} = \sum_{z'=1}^{Z} emp_{z',s,t} - emp_{z,s,t} \tag{3}$$

b) Total regional size

To capture total market size and agglomeration effects unrelated to the industrial structure, we include the total size of region z without the own sectoral employment to avoid endogeneity.

[5] The 15 manufacturing industries are Electronics, Chemical Industry, Synthetic Material, Nonmetallic Mineral Mining, Glass&Ceramics, Primary Metal Manufacturing, Machinery, Automobile, Office Supplies & IT, Musical Instruments&Jewellery, Wood-working, Paper & Printing, Leather&Apparel, Food&Tobacco and Building&Construction. The advanced services include the sectors Commerce, Finance&Insurance, Business-Related Services, Education, Leisure-Related Services and Social Services.

$$size_{z,t} = emp_{z,t} - emp_{z,s,t} = \sum_{s'=1}^{S} emp_{z,s',t} - emp_{z,s,t} \tag{4}$$

where $S = 28$ denotes the total number of local industries.[6] An alternative measure for addressing agglomeration effects that has often been used is employment density, which would be (4) divided by some appropriate measure of area size (see Ciccone, 2002 for an analysis that includes Germany). We prefer $size_{z,t}$, because there has been no variation in the territorial classification of districts over time. In either case, one must be careful how to measure the effects of local specialisation when both, the total area employment (or employment density) and the (lagged) employment level of the district-industry is included.

c) Specialisation

The usual measure for own industry specialisation is the local employment share. But the employment share of industry s in area z is perfectly collinear with the employment level $emp_{z,s,t}$ and total area employment $size_{z,t}$. Thus, as Combes (2000) has shown, the inclusion of all three elements at once leads to an overstatement of the effect of specialisation.

An alternative indicator is the (size of the) coefficient for the lagged dependent variable, used also by Combes, Magnac and Robin (2004). The auto-regressive parameter in (2) indicates whether a local industry grows faster in environments with strong past growth performance. Strictly speaking, there is only evidence for a positive effect of own industry specialisation if the estimated coefficient is larger than one, as this would imply an explosive growth path. A parameter between zero and one indicates mean reversion in the long run, which is inconsistent with the idea that local industry growth feeds on itself. However, there can be some inertia in the transition dynamics towards the long-run target. This inertia is stronger the larger (the closer to one) is the coefficient.

d) Diversity

Most of the former studies have used a variant of the commonly used Herfindahl-Hirshman index to capture the diversity of the surrounding industrial environment. For example, Combes (2000) has suggested the following specification that was later also used in Combes, Magnac and Robin (2004) and in Blien and Suedekum (2005).

[6] Note that this variable includes also those 7 industries for which we do not perform the regressions.

$$HHI_{z,s,t} = \left[\sum_{s'=1, s' \neq s}^{S} \left(emp_{z,s',t} / (emp_{z,t} - emp_{z,s,t}) \right)^2 \right]^{-1} \quad (5)$$

This measure increases with local diversity faced by sector s. It reaches a maximum when all surrounding industries account for an identical employment share. The coefficient measures the impact of diversity, as sector s faces a more balanced local industrial environment the higher $HHI_{z,s,t}$. But there is an identification problem when using $HHI_{z,s,t}$ in logs together with the total regional employment (4) as an additional explanatory variable.

$$\log(HHI_{z,s,t}) = -\log\left(\sum_{s'=1, s' \neq s}^{S} \frac{emp_{z,s',t}^2}{(emp_{z,t} - emp_{z,s,t})^2} \right) = -\log\left(\frac{\sum_{s'=1, s' \neq s}^{S} emp_{z,s',t}^2}{(emp_{z,t} - emp_{z,s,t})^2} \right)$$

$$= -\log \sum_{s'=1, s' \neq s}^{S} emp_{z,s',t}^2 + 2\log\left(emp_{z,t} - emp_{z,s,t} \right) \quad (6)$$

The diversification effect can not be identified from the total regional employment, as can be seen by comparing (6) with (4) in logs. Because of this, we prefer to use an alternative measure for diversification, namely a standard Krugman-diversification index that is defined in the following way

$$diversity_{z,s,t} = -\sum_{s'=1, s' \neq s}^{S} \left| \frac{emp_{z,s',t}}{emp_{z,t}} - \frac{emp_{s',t}}{emp_t} \right| \quad (7)$$

This index sums the absolute differences of the regional and the national employment shares of all sectors (except for the one under consideration). It takes on the value of zero if the surrounding local economic structure exactly mirrors the average national structure, and it is stronger negative the more idiosyncratic (and less diversified) the district z. It uses a different reference structure than the HHI, the national average structure instead of a setting with identical employment shares. Since the size of the sectors under consideration varies substantially, this is taken into account by the Krugman index. Additionally it is not flawed with comparable identification problems.

Apart from these measures, we will additionally include three further control variables.

e) Firm size

The first one relates to average firm size, which is a common feature of several papers in this literature. Specifically, we include the employment share in small firms

$$firmsize_{z,s,t} = emp\left[in\ \ firms < 20\, employees\right]_{z,s,t} \big/ emp_{z,s,t}\ . \tag{8}$$

Traditionally, inference about the impact of local product market competition was based on firm size measures (e.g., Glaeser et al., 1992).[7] As argued by Combes (2000), this is quite problematic, however, as firm sizes measure the effect of internal scale economies instead of competition.

f) Education

Most previous studies did not include information about education due to data limitations. In the present analysis we include the employment share of college educated workers,

$$education_{z,s,t} = high\ skilled_{\ z,s,t} \big/ emp_{z,s,t}\ \ . \tag{9}$$

This variable measures the human capital intensity of a local industry, which is not related to the local economic structure. A straightforward theoretical basis for including education in a growth regression are human capital spillovers. In his extensive overview, Moretti (2004) distinguishes two channels of human capital externalities, direct technological spillovers (as in the famous model of Lucas, 1988), and complementarities between different skill types. Although negative human capital spillovers are also conceivable, the most plausible prior is to expect a positive effect of a higher stock of human capital on productivity in a local industry, operating through the two channels.[8] To the extent that a positive productivity shock translates into higher equilibrium employment, which relates to the discussion about Jacobs- and MAR-externalities and the caveat pointed out by Cingano and Schivardi (2004), one would expect a positive effect of an increasing share of high skilled workers on employment growth.

[7] Small average firm sizes are frequently identified with strong competition on local product markets, whereas large firm are set equal with a high degree of monopolisation. The debate is then whether strong competition, or high market power is conducive to innovation and, ultimately, growth.

[8] The most fundamental methodological issue in this debate is how to separate the technological spillover from the complementarity in an empirical analysis, where typically either output or wages are used as dependent variable. In the present context we are only interested in the total effect of human capital on employment growth.

A particular interesting issue in the present context concerns the relation between the general effect of human capital versus the regional industry composition. As argued above, inter- and intra-sectoral knowledge spillovers can be an important rationale why diversity and specialisation influence the performance of a local industry. These externalities are typically thought of as having their origin in the communication of educated workers (Duranton and Puga, 2004). Controlling for the general human capital intensity independent of the local economic structure will thus serve as an important robustness check for the impact of diversity and specialisation.

g) Wages

Lastly, we will not include raw wages, but rather use a methodology (described at length in Blien and Suedekum, 2004) for constructing a "neutralized" regional wage level. To do so, we take a preceding step and regress the average (log) wage in every industry, region and year ($wage_{z,s,t}$) on a variety of explanatory variables and fixed effects. Period-by-period we estimate the following wage regression

$$\ln(wage_{z,s,t}) = a' + W_z' + \varsigma_s' + \beta X'_{z,s,t-1} + \varepsilon_{irt}' \tag{10}$$

where $X'_{z,s,t-1}$ is the matrix of observable characteristics of the respective local industry (firm size structure, qualification, age and gender), ς_s' is an industry fixed effect, and W_z' a location fixed effect. From this analysis we take the regional fixed effects W_z' and include them in main regressions (1) and (2) as our measure of the neutralized regional wage level. That is, a "high wage region" in our interpretation is not a region with high wages per se, but a region whose wages are higher than expected, given a variety of characteristics.

We estimate (10) subject to the restriction that all regional fixed effects, weighted by the aggregate employment share of regional over total national employment, must sum up to zero. This method, which is simply a normalization of coefficients that does not affect the other estimators, is useful since it allows interpreting the values of W_z' as percentage deviations from a national grand mean of zero, not in relation to some arbitrarily omitted reference category from the complete (and thus, perfectly multicollinear) set of regional dummies. "High wage" and a "low wage" regions are characterised by values of W_z' that are significantly higher (lower) than zero.

In our view, this is a more meaningful measure than the raw average regional wage, as used e.g. in Henderson (1997), since it controls for static productivity differences due to qualification, firm sizes etc. that influence wages in a different way than employment growth.

3) Results

We estimate our empirical model separately for the manufacturing sector (15 industries) and the advanced services sector (6 industries). We report the results of a parsimonious model specification in table 1a, where we include two lags of the dependent variable and the independent variables with up to two lags.[9] That is, we specify an autoregressive distributed lag model, ADL(2,2), as laid out e.g. in Davidson and MacKinnon (2004: 577). The results are robust with respect to different specification tests. Our baseline regression (I) reported in table 1a leaves out the control variables on education and neutralized wages, which are novel features of the present study. These two variables are included in a second estimation (II) to check the robustness of the results. For instrumenting the first difference of the lagged dependent variable we use higher order time lags of the dependent variable in levels (Arellano and Bond, 1991). In table 1b we provide the results of diagnostics tests for the validity of the used instruments. The Sargan-test can not be rejected at the 5% level for manufacturing and for services in both specifications.[10] The other assumption that is necessary for the validity of instruments, serial uncorrelated error terms $\varepsilon_{z,s,t}$, can not be rejected at conventional significance levels. We conclude that $\log(emp_{z,s,t-l})$ with $l \geq 2$ are valid instruments.

3.1. Manufacturing

In manufacturing, diversity is found to matter in the short run. Focussing on estimation (I) the contemporaneous coefficient is positive (0.184) and significantly different from zero. The impact of the variable is insignificant with a time lag of two or more periods. This agrees with the conclusion of Combes, Magnac and Robin (2004) for the case of France, who also find that current instead of historical changes in the degree of diversity have a positive effect on employment growth. It stands in contrast to the conclusion of Henderson (1997), who finds a time lag of six years or more to be relevant. Including education and wages as additional

[9] We have estimated also specifications with more than two lags. But the coefficients of the variables lagged three and more were not significant. Also all the estimated coefficients for the contemporaneous variables remain stable. Results are available on request from the authors.

[10] It is commonly known from Monte Carlo studies (e.g. Hansen et al., 1996) that the Sargan-test rejects the null hypothesis of valid instruments too easily. Hence, given the strong support we get from the autocorrelation test, we do not worry too much about the low p-value of the Sargan-test. Moreover, note that the p-values increase due to the inclusion of the two additional control variables, which are partly significant. This is so, because the Sargan-test also reacts sensitive to misspecifications.

control variables leaves the estimated coefficient nearly unchanged (0.180). The positive impact of diversity on employment growth in the short run thus seems to be robust result.

Table 1a: Results Dynamic Panel Estimation

Dep.variable: **y(t)** Time dummies: **YES**		**Manufacturing (I)**	**Manufacturing (II)**	**Services (I)**	**Services (II)**
number of groups		4717	4717	1955	1955
y	t-1	0,836***	0,840***	0,877***	0,869***
		(0,000)	(0,000)	(0,000)	(0,000)
	t-2	-0,084	-0,079	-0,025	-0,006
		(0,372)	(0,392)	(0,825)	(0,958)
sect	t	0,540***	0,539***	0,832***	0,836***
		(0,000)	(0,000)	(0,000)	(0,000)
	t-1	-0,320**	-0,321***	-0,732***	-0,742***
		(0,000)	(0,000)	(0,000)	(0,000)
	t-2	-0,049	0,058	0,069	0,038
		(0,537)	(0,450)	(0,580)	(0,761)
size	t	0,160***	0,155***	0,810**	0,079*
		(0,001)	(0,002)	(0,064)	(0,066)
	t-1	-0,093**	-0,096*	-0,028	-0,030
		(0,098)	(0,087)	(0,526)	(0,502)
	t-2	0,032	0,033	0,045	0,043
		(0,483)	(0,459)	(0,173)	(0,201)
diversity	t	0,184***	0,180***	0,096***	0,095***
		(0,000)	(0,000)	(0,001)	(0,002)
	t-1	-0,147***	-0,145***	-0,050**	-0,049*
		(0,000)	(0,000)	(0,067)	(0,077)
	t-2	-0,047	-0,043	-0,012	-0,0141
		(0,132)	(0,159)	(0,595)	(0,540)
firmsize	t	-0,050***	-0,049***	-0,325***	-0,322***
		(0,000)	(0,000)	(0,000)	(0,000)
	t-1	0,042***	0,042***	0,303***	0,312***
		(0,000)	(0,000)	(0,000)	(0,000)
	t-2	0,007	0,008	0,018	0,026
		(0,431)	(0,373)	(0,658)	(0,539)
education	t		0,020***		0,004
			(0,000)		(0,173)
	t-1		-0,014***		-0,012***
			(0,000)		(0,000)
	t-2		0,001		-0,007**
			(0,777)		(0,016)
wages	t		-0,067		-0,050
			(0,531)		(0,557)
	t-1		0,109		0,148*
			(0,227)		(0,055)
	t-2		-0,224**		0,004
			(0,011)		(0,959)

Table 1b: Test for validity of instruments

	Manufacturing (I)	Manufacturing (II)	Services (I)	Services (II)
Sargan test of over-identifying restrictions	chi2(17) = 27,12 Prob > chi2 = 0,056	chi2(17) = 25,74 Prob > chi2 = 0,079	chi2(17) = 27,36 Prob > chi2 = 0.053	chi2(17) = 24,85 Prob > chi2 = 0.098
AB-test that average auto-covariance of order 1 is 0. H0: no autocorrelation	z = -17,83 Pr > z = 0,0000	z = -18,09 Pr > z = 0,0000	z = -6,70 Pr > z = 0,0000	z = -6,63 Pr > z = 0,0000
AB-test that average auto-covariance of order 2 is 0 H0: no autocorrelation	z = 1,15 Pr > z = 0,2506	z = 1,13 Pr > z = 0,2577	z = 0,901 Pr > z = 0,3703	z = 0,81 Pr > z = 0,4200

The coefficient for the lagged dependent variable is very similar, but smaller than one in both estimations (0.836 and 0.840). We thus find no evidence for an explosive growth path. The coefficient seems reasonably close to one, which suggests that mean reversion exhibits considerable inertia. To have an impression of the size of the effect a numerical calculation is useful. Assume an initial exogenous growth impulse of 5 % percent. Due to the inertia, the next year's growth rate would be 3,8 % even without any further exogenous impulse. After ten years, the employment level will be approximately 18 % higher (not counting the initial impulse), which seems to be a substantial increase. Note, however, that we can not judge how much of this increase is potentially due to MAR-externalities, since we do not know how strong mean reversion would be in the absence of any effect.

The positive sign of the contemporaneous coefficient for total regional size (0.160, resp. 0.155) shows global agglomeration, or market size effects, which qualitatively agrees with the conclusion of Ciccone (2002). We also find that an increasing share of small firms reduces employment growth (-0.050, resp. -0.049), which is roughly consistent with the findings of Combes (2000). It stresses the role of internal scale economies, as a setting with many small firms does not seem to be a very growth friendly environment. Note that impact of both control variables, which is almost identical in both estimations, fades away quickly over time. Only the contemporaneous variable is significant, hence it seems to be the current rather than some historical setting that influences employment growth.

Concerning the education variable, we find a significantly positive impact (0.020) on employment growth in estimation (II). This finding is consistent with general human capital spillovers, as emphasised in the endogenous growth literature. Interestingly, these general effects of education do not interfere with the variables that reflect the local economic structure, as an elimination of the education leaves the other coefficients unchanged. This seems to be an important finding for the robustness of the positive impact of diversity.

Finally, the short run effect of the neutralized regional wages is negative, but not significant. There is a significantly negative effect with a time lag of two periods, however. Higher regional wages should, ceteris paribus, depress employment growth according to neoclassical arguments. On the other hand, they also point to a higher purchasing power of local consumers, which might have positive demand side effects on employment growth. We interpret the non-significant coefficient in table 1a such that cost push effects dominate over potential demand side effects, but that the latter moderate the former to an extent that renders the contemporaneous relationship between neutralized wages and employment growth insignificant.[11] In the longer run, as expected, cost push effects grow stronger whereas the mitigating influence of local demand disappears.

As argued above, the impact of diversity and most other control variables appears to be of a contemporaneous nature. However, we also check if there is a long-run impact. Given the ADL-specification, the long-run effects on employment growth can be determined by computing (for each independent variable) the following coefficient δ^*

$$\delta^* = \frac{\sum_{l=0}^{2} \delta_l}{1 - \sum_{p=1}^{2} \rho_p} \qquad (11)$$

where δ_i are the coefficients for the lagged independent variables $X_{z,s,t}$, and δ_p for the lagged dependent variable. The long-run results are reported in table 2, with p-values for the significance of the coefficients in parentheses.

There is a significantly positive and robust long-run impact of diversity, total regional size and education on employment growth. This finding supports the view that general agglomeration and urbanization effects are conducive for employment growth. Note, however, that due to our specification the "long run" is relatively short, compared e.g. to the more traditional cross-section analysis by Blien and Suedekum (2005). Concerning the long-run impact of the dependent variable, we can provide the result that the null hypothesis of a coefficient equal to one can not be rejected at any reasonable level of significance.

[11] This conclusion is in accordance with the findings of Suedekum and Blien (2004), who look at the relation between wages and employment growth in West Germany (1993-2001). There, the wage effect is broken down according to single industries. For a subset of industries it is insignificant, but for some it is significantly negative.

Table 2: Long run effects

	Manufacturing (I)	Manufacturing (II)	Services (I)	Services (II)
sect	0,691*** (0.000)	0.670*** (0.000)	1,142*** (0.000)	1,197*** (0,000)
size	0,401*** (0.002)	0.384*** (0.005)	0,658*** (0.009)	0.838* (0.080)
diversity	0,341*** (0.004)	0.324*** (0.007)	0,228 (0.135)	0.293 (0.211)
firmsize	-0,007 (0.929)	-0.001 (0.990)	-0,019 (0,938)	0.145 (0.757)
education		0.028*** (0.005)		-0.136 (0.315)
wages		-0.762 (0.300)		0.934 (0.457)

3.2. Services

Going over to the service sector, we also receive a positive short run impact of diversity that is robust across specifications and that dies out quickly over time (see table 1a). The estimated coefficient (0.095) is considerably smaller as compared to the manufacturing sector (0.180), and there is no positive impact of diversity in the longer run (see table 2). For the lagged dependent variable we obtain conclusions that are similar to manufacturing. The coefficient is close to one in the two specifications (0.877, resp. 0.869), but not larger than one. In the longer run, the null hypothesis of an impact equal to one can also not be rejected.

With respect to firm size and total regional size, the signs of the coefficients are the same as for manufacturing, but the magnitudes differ. A large share of small firms depresses employment growth in advanced services even stronger, although the effect is also not significant in the longer run. Global agglomeration effects also appear to be stronger, both in the short and longer run, at least for the baseline specification. Somewhat surprisingly, advanced service industries do not grow stronger the larger their share of college educated workers. If anything, it is even the opposite. Note, however, that this need not imply that human capital spillovers are absent in advanced services, because we test the impact on employment, not on productivity. Furthermore, the education variable is important insofar as it underpins again the robustness of the impact of diversity and specialisation, because the coefficients that address these two central issues are very similar in the two specifications. Lastly, we find no negative impact of neutralised regional wages on employment growth, neither in the short nor in the longer run. This suggests that the employment growth effects of are less adverse in services than in manufacturing, which agrees with the conclusions of Suedekum and Blien (2004).

4.) Conclusion

The local economic structure is an important determinant of the employment growth performance of different industries, but timing is a crucial issue. In West Germany, it is predominantly the current rather than some historical industry pattern that matters. For economic policy this can be good news, because structural interventions that influence the industry structure at the local level will have an immediate impact on employment growth. According to our estimations, policymakers do not have to wait for several years before results become visible. On the other hand, the effects of policy might also not be long-lasting. Our results show that employment growth of local industries benefits from a diversified and urbanized surrounding environment, whereas there is no clear evidence that local clustering of an industry leads to a take-off of employment growth rates.

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