Convergence in West German Regional Unemployment Rates

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ABSTRACT
Differences in regional unemployment rates are often used to describe regional economic inequality. This paper asks whether changes in regional unemployment differences in West Germany are persistent over time. Only if such changes are persistent, the differences are a sensible measure of inequality and only then can policies be effective that aim at lowering the dispersion of unemployment rates.
Our analysis follows a time-series approach to economic convergence and we test whether unemployment differences between regions are stationary or not. While univariate tests show that changes in unemployment differences are persistent, more powerful panel tests find them to be only transitory. However, these tests reveal only a moderate speed of convergence. Since there is a structural break following the second oil crisis, we also employ unit-root tests that allow for such break. Again we find strong evidence for convergence and now also the speed of convergence is found to be very high.
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1 Introduction

The large literature on economic convergence between countries and regions mostly focuses on per capita income or other related income and productivity measures. However, as Quah (1996, p. 1354) points out:

> "Certainly, understanding economic growth is important. But growth is only one of many different areas in economics where analyzing convergence sheds useful insight."

Therefore, this paper borrows techniques from the growth convergence literature to examine the evolution of regional disparities in unemployment rates within a country, a topic that has gained much attention since the seminal paper of Blanchard and Katz (1992). Unemployment disparities are often perceived as persistent. They are at the heart of the "regional problem" and in the focus of regional economic policy (Armstrong and Taylor, 2000). In particular the persistence of unemployment disparities has attracted much attention, see for example Decressin and Fatas (1995) or Obstfeld and Peri (1998). Regional disparities may be persistent because they reflect stable equilibrium differentials or because shocks to regional unemployment rates have long-lasting effects, as Martin (1997) points out. Only in the latter case policy interventions are likely to be effective. On the contrary, if the differentials reflect an equilibrium that has been stable over time, (short-term) policy interventions can hardly be expected to change this equilibrium.

It is thus interesting in particular in the German context to study whether regional unemployment rates converge to the national average over time and how fast this convergence happens, as the federal government in Germany aims to reduce the gap between unemployment rates in East and West Germany by granting subsidies and by spending on public infrastructure. We employ annual data from the "Mikrozensus" database on unemployment rates for the West German federal states during the period 1960-2002 and adopt the stochastic definition of convergence proposed by Bernard and Durlauf (1995, 1996) and others. On the basis of this definition of convergence, our study characterizes the evolution of the gap between the unemployment rate in a federal state and the unemployment rate in Germany as a whole.

For the US, Blanchard and Katz (1992, pp.12) analyzed the dynamics of regional employment and unemployment. While they do not explicitly find evidence for stationarity of regional unemployment rates, they attribute this to a power problem of the tests they apply. Indeed, Decressin and Fatas (1995) and Obstfeld and Peri (1998) provide some evidence that regional unemployment disparities are a more persistent phenomenon in Europe than in the US. However, these results have recently been questioned by Rowthorn and Glyn (2003) who do find substantial persistence also in US regional unemployment rates. For the UK by contrast, Martin (1997) finds that regional unemployment shocks are only short lived. Yet, he also finds that regional unemployment rates differ in the long run which reflects a stable equilibrium distribution around the national average.

For Germany, this study is to the best of our knowledge the first one analyzing the convergence of unemployment rates at the federal state level (see Section 3 for a data description). There are a number of studies which examine the related issue of hysteresis for West German unemployment rates: Belke (1996), Belke and Göcke (1996), Hansen (1991), Reutter (2000). However, these studies analyze the absolute level of aggregate or regional unemployment rates and not relative unemployment rates as we do. As a consequence, these papers cannot shed much light on convergence.

The main results of our study are the following. While univariate techniques, which do not account for structural breaks, do not show evidence for stochastic convergence in relative unemployment rates, more powerful panel-based methods do so but suggest that it is sluggish. The estimated half-life of a shock to regional unemployment is 5.5 years. A graphical analysis reveals that there is a structural break in the data. Therefore, we subsequently include the possibility of such a break in the analysis, which is specified as an endogenously determined single level shift in the mean of the series. After this extension we are able to reject the null hypothesis of a persistent shock to unemployment differences in favor of conditional convergence for most regions on the basis of their individual time series. Regional unemployment rates are found to converge up to a constant differential with the national average, but this differential is subject to a one-time permanent shift which occurred following the second oil crisis. Moreover, allowing for a structural break, the estimated speed of convergence increases substantially, so that the estimated half-life goes down from 5.5 to less than 2 years on average. Consequently, persistence in regional unemployment disparities reflects an equilibrium phenomenon to which the German economy adjusts quickly.

The remainder of this paper is organized as follows: Section 2 introduces the theoretical concepts. After describing the dataset in Section 3, we begin with a graphical analysis which serves as a guideline for the rest of the paper. Section 4 analyzes convergence on the basis of univariate and panel unit root tests which do not account for structural breaks. This analysis is extended to the possibility of a structural break in Section 5. Finally, Section 6 concludes.

2 Theoretical concepts

When labor markets adjust towards equilibrium in the long run, there will be convergence of regional unemployment rates, because unemployed workers take jobs in other areas or because capital flows into low-wage regions when producers take advantage of lower labor costs (for details see Blanchard and Katz, 1992). However, if the speed of adjustment is slow, unemployment disparities may arise during adjustment as a result of negative demand shocks affecting some regions more than others (Armstrong and Taylor, 2000).

We can test this theory of long-run convergence empirically by using Bernard and Durlauf's (1995, 1996) time series approach. This test focuses on the permanence of shocks to relative variables and uses a stochastic definition of convergence (Carlino and Mills 1993). The idea of Bernard and Durlauf's test (1996) for stochastic convergence can best be explained using income as an example. Let y_i and y_j be per-capita income of two countries i and j, respectively. In the starting period, country i has a larger income than country j, $y_{i0} > y_{j0}$. The output gap between the two regions is $y_{it} - y_{jt}$. Define I_t as the information set available at period t. Then, Definition 2 in Bernard and Durlauf (1996, p. 165) understands convergence as the equality of long-term forecasts at any fixed time. This means

$$\forall t : \lim_{k \to \infty} E(y_{i,t+k} - y_{j,t+k} | I_t) = 0 \qquad (1)$$

Stochastic convergence implies that differences between the economies will always be transitory in the sense that long-run forecasts of the difference between any pair of countries converge to zero as the forecast horizon grows (Oxley and Greasley, 1997). The important testable implication of long-run convergence is that stochastic convergence can only be present if shocks to the disparity are temporary. Hence, the disparities between economies should follow a stationary process, which means that y_i and y_j are cointegrated. Without stationarity, shocks to the relative variable lead to permanent differences. Carlino and Mills (1993) and Evans and Karras (1996) demonstrate that a test for stochastic convergence can be conducted as a Dickey-Fuller test for the presence of a unit root in the relative variable. If the series has a unit root, shocks are permanent and there will be no convergence. Moreover, long-run convergence also precludes a deterministic trend in the cross-country differences. Additionally, the series of gaps should have a zero mean.

However, the hypothesis of perfect convergence might be too strict. Consider as a simple example regional amenities that lead to wage differentials which compensate workers for differences in the quality of life and regional price levels. Now, additionally assume a national unemployment insurance pays a fixed unemployment benefit that is equal among regions. Because in regions rich of amenities the wages are lower, the equal unemployment benefit results in higher rates of voluntary unemployment in those amenity-rich regions. Or to put it differently, the voluntarily unemployed would move to the amenity-rich regions in this simplistic setting. Therefore, a stable difference between regional unemployment rates may simply reflect fundamental economic differences, such as differences in natural endowments. In such a setting, the regional economic policy could only aim at shifting the equilibrium by policy interventions. However, it is unlikely that such policy interventions are actually effective, if the equilibrium has been very stable over the past.[1]

To capture the notion of a stable long-run difference, we define conditional convergence as

$$\forall t : \lim_{k \to \infty} E(y_{i,t+k} - y_{j,t+k} | I_t) = constant. \tag{2}$$

This means y_i and y_j converge towards a (time-invariant) equilibrium differential. An empirical test for stochastic conditional convergence is again related to the time series properties. Conditional convergence implies that the series is level-stationary but it is not required that the series has an intercept of zero. For example, consider the series generated by the autoregressive model $y_t = \phi + \rho y_{t-1} + \varepsilon_t$. This series is stationary if $|\rho| < 1$ and the intercept ϕ controls the mean of y_t through the relationship $E(y_t) = \mu = \phi/(1-\rho)$. If y is relative unemployment, we find conditional convergence if $\rho < 1$ and unconditional convergence if additionally $\phi = 0$.

3 Data and graphical examination

3.1 Data

We use data from the German "Mikrozensus" database which is available from 1957 onwards. In this database the term "unemployed" refers to all people without employment contract who search for a job irrespective of whether they are registered as unemployed or not at the Federal Employment Agency. Therefore, the definition of unemployment in our data is somewhat different to the statistics of the Federal Employment Agency. Data on registered unemployment on the federal state level are available only since 1974 (depending on the federal state). Before 1974, data on registered unemployment are available only for the "Landesarbeitsamtsbezirke" but not for the federal states. Besides the advantage of longer time-series, the Mikrozensus definition of unemployment has also

[1] See Marston (1985) for a more elaborated theoretical underpinning of the equilibrium and disequilibrium perspective of regional unemployment disparities.

the advantage of being more similar to the definition of the unemployment rate used in other countries, in particular the US.

Since there was virtually no unemployment in Germany during the late 1950s, we restrict the data to the time-period 1960-2002. Moreover, West Berlin is excluded from the analysis because of its special status before unification.

The unemployment rate is defined as unemployment divided by the labor force (multiplied by 100). Labor force data were also derived from the Mikrozensus. According to the Mikrozensus definition, the labor force ("Erwerbspersonen") is the sum of the employed and the unemployed ("Erwerbstätige" and "Erwerbslose").

We denote the unemployment rate for federal state i by ur_i and the unemployment rate for Germany as a whole by ur_{Ger} (without West-Berlin). Time indices are suppressed for notational convenience. For the period after German reunification, 1991-2002, the unemployment rate for Germany, ur_{Ger}, is calculated on the basis of the data from West German federal states only.

As explained in the previous section, stochastic convergence requires that relative unemployment rates follow a stationary process. Unemployment rates are themselves relative numbers and bounded between 0 and 100 percent. Hence, in theory a linear model for the differences must be inconsistent if both unemployment-rates are close to the opposite extreme bounds. Yet, in practice these observed unemployment rates are never close to the upper bound, so that relative unemployment rates can be calculated as simple differences in the levels,[2] which means that the relative unemployment rate u_i for federal state i is computed as

$$u_i = ur_i - ur_{Ger} \tag{3}$$

The unemployment rate for Western Germany, ur_{Ger}, is selected as a reference. This reflects that unemployment rates for the different federal states do not evolve differently from the national average if they converge.

3.2 Graphical examination

To get a first impression of the time-series characteristics of u_i, we display the series graphically. Figure (1) plots relative unemployment rates during the period 1960-2002.

It can be seen that the dispersion of unemployment rates has sharply increased in times of recessions (1966/67 and at the beginning of the 80s) parallel to the increase in the aggregate unemployment rate. In the beginning of the 1960s unemployment was not

[2] Using differences in logs has the disadvantage that minor differences in unemployment rates and rounding errors get inflated by the low aggregate unemployment rates during the 1960s.

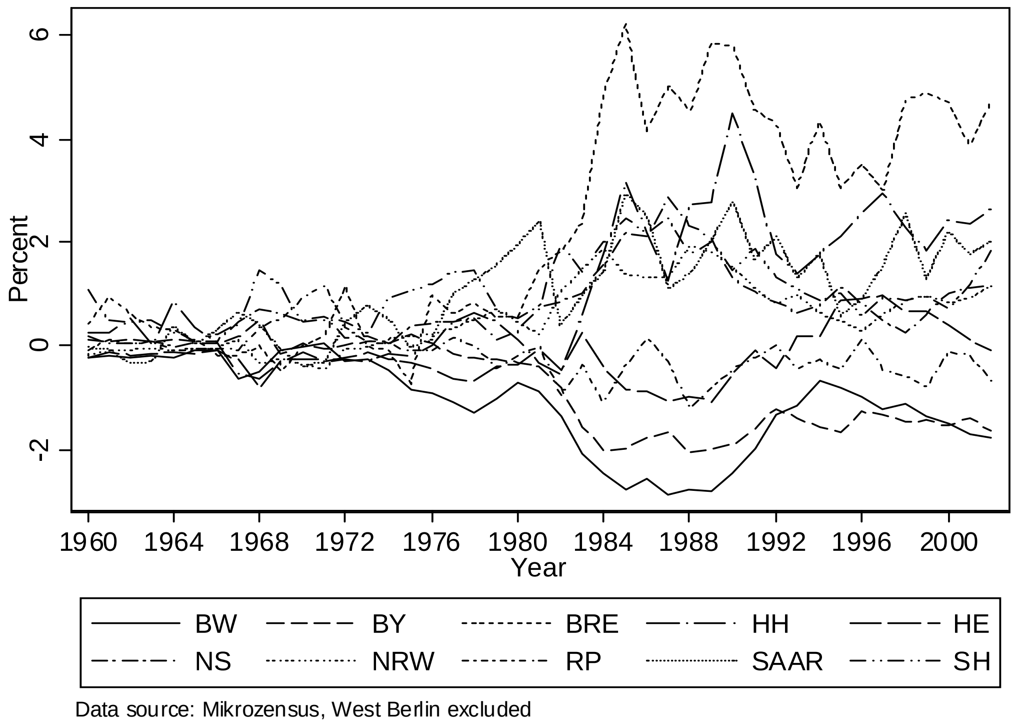

Figure 1: Relative unemployment rates in Western Germany, 1960-2002.

a problem in Germany, in fact there was rather a shortage of labor, similarly there is not much of a difference in unemployment rates across states. After 1980 the situation is dramatically different, the dispersion of unemployment rates sharply increases with the general rise in unemployment rates. Thereafter economic differences between the northern and southern part of Germany become apparent. Since the beginning of the 80s, the city states Bremen and Hamburg have the highest relative unemployment rates, while Bayern and Baden-Würtemberg have unemployment rates around 2 percentage points below the national average.

At first glance, most of the series look non-stationary. However, splitting the sample in the period before and after 1980 shows that the lack of stationarity might be just due to a single structural break that occurs in the early 80s after the second oil crisis. In order to illustrate this, figure 2 displays the data for both subperiods; one ranging from 1960-1979 and the second from 1980-2002 (Figure 2). The series look more stationary now. Additionally, the two graphs illustrate that the dispersion of relative unemployment

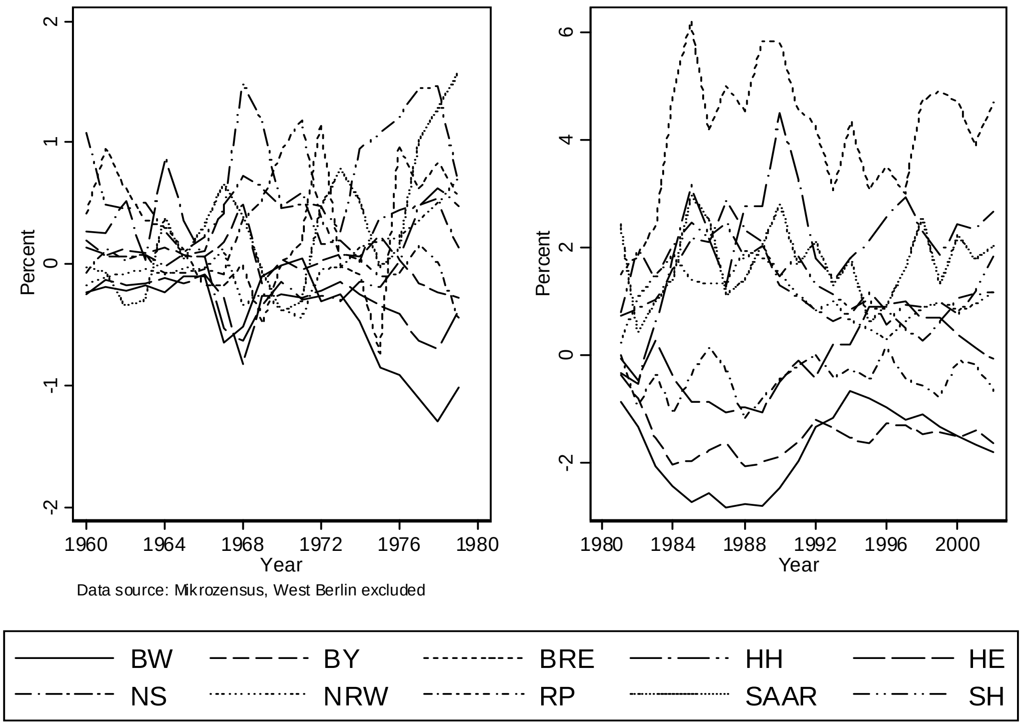

Figure 2: Relative unemployment rates in Western Germany, subperiods 1960-1979 and 1980-2002.

rates is bigger during the second subperiod than during the first. It seems as if the levels of the series have changed due to a structural break and that the series have only slowly reverted to the mean. Finally, note that the data do not display a deterministic trend.

4 Unit root tests without structural breaks

Having displayed the series graphically, we turn to a formal characterization of the stochastic behavior. The hypothesis being tested is that relative unemployment rates contain a unit root. To provide a benchmark for our later results, we first employ a univariate unit root test without structural breaks. Thereafter, we analyze convergence with more powerful panel-based unit root tests.

4.1 Univariate unit root tests

As explained before, tests of convergence can be conducted as Dickey-Fuller (1979) type tests based on the difference between the unemployment rate in federal state i and the

Table 1: ADF test for relative unemployment rates (without trend)

Federal State	ADF lags (k)	μ	α	p-value	Federal State	ADF lags (k)	μ	α	p-value
BW	1	-.117*	-.0828	0.402	NRW	5	.099	-.122	0.440
		(.066)	(.047)				(.061)	(.073)	
BY	5	-.087	-.066	0.658	RP	0	-.071	-.402	0.027**
		(.061)	(.054)				(.062)	(.130)	
BRE	0	.283	-.082	0.650	SAAR	3	.321*	-.189	0.547
		(.201)	(.065)				(.169)	(.128)	
HH	3	.176	-.081	0.791	SH	0	.278**	-.241	0.187
		(.144)	(.091)				(.133)	(.107)	
HE	2	-.044	-.219	0.233	ur_{Ger}	2	.302**	-.038	0.729
		(.052)	(.103)				(.191)	(.036)	
NS	0	.105	-.108	0.509					
		(.072)	(.070)						

standard errors in parentheses; *,**,*** significant at the 10, 5 and 1 percent levels

unemployment rate for Western Germany:

$$\Delta(u_{i,t} - u_{Ger,t}) = \mu + \alpha(u_{i,t-1} - u_{Ger,t-1}) + \sum_{n=1}^{k} \delta_n \Delta(u_{i,t-n} - u_{Ger,t-n}) + \varepsilon_t \qquad (4)$$

If the series contains a unit root ($\alpha = 0$), the proposition for both absolute and conditional convergence is violated. The alternative hypothesis is that $\alpha < 0$, which implies that the series is stationary. Moreover, unconditional convergence implies insignificance of the constant term, μ. Since a deterministic time trend is neither compatible with long-run convergence nor apparent in our data, we do not include a trend in the regressions.

The convergence tests for relative unemployment rates are reported Table 1, optimal lag lengths have been determined by sequential t-tests. It can be seen that there are considerable differences in the time series properties of relative unemployment rates across the federal states, but the most important result is that for nearly all federal states we cannot reject the null hypothesis of a unit root. The unit root is safely rejected only for Rheinland-Pfalz.

This means that the ADF tests provide no evidence of stochastic convergence during the period under study. Other studies of convergence often include a deterministic time trend in the ADF regressions. In our setting, the derived results do not depend on the absence or presence of a trend. When we allow for a time trend, the series for Rheinland-

Pfalz remains (trend) stationary and all other series remain non-stationary.[3] We also tried with Phillips and Perron and KPSS tests, but again the results did not change qualitatively.

4.2 Panel unit root tests

It is well known that unit root tests such as the ADF test have low power against stationary alternatives in small samples. Panel-based unit root test have been shown to be more powerful, since they exploit the cross-sectional dimension of the data. Most prominent tests include the Levin and Lin (2002), Im, Pesaran and Shin (2003) and the Maddala and Wu (1999) testing procedures.

The basic regression for these panel unit root tests is[4]

$$y_{it} = \rho_i y_{i,t-1} + z'_{i,t}\gamma + u_{i,t}$$
$$i = 1, ..., N; t = 1, ..., T$$

where z_{it} is the deterministic component and u_{it} is a stationary error. The set of exogenous regressors z_{it} could be empty, or include a one, fixed effects, μ_i, or fixed effects and a time trend, t.[5] The Levin and Lin (2002) test assumes that each individual unit in the panel shares the same AR(1) coefficient: $\rho_i = \rho$ for all i. Hence, the power of the single ADF tests is increased not only by pooling the data but also by exploiting a cross-equation parameter restriction on the unit root parameters.[6] The null hypothesis of the Levin and Lin test states that the relative unemployment series of each state contains a unit root against the alternative that all series are stationary.

The Levin and Lin test combines individual Dickey-Fuller regressions. In a first step, we test for conditional convergence of regional unemployment rates by including a constant term in the individual Dickey-Fuller regressions. Hence, we control for heterogeneity by allowing for unit-specific (fixed) effects. We do not include a deterministic time trend in the regressions. In our setting, the Levin and Lin test with constant terms can be interpreted as a test for the convergence of regional unemployment rates to a pattern of stable differences to the national average.

[3] We also tried the Dickey-Fuller GLS test proposed by Elliot, Rothenberg and Stock (1996). The qualitative results are the same as obtained with the conventional ADF tests.

[4] See Baltagi (2001) for an overview of non-stationary panels.

[5] Note that in the more general case, when the error disturbances u_{it} are serially correlated, the serial correlation can be corrected by including lagged terms similar to the ADF procedure.

[6] Note that the Levin and Lin test statistic converges more rapidly with respect to the time dimension T than with respect to the cross-section dimension N. Hence, the Levin and Lin test is well-suited for our dataset with $N = 10$ and $T = 43$.

Table 2: Levin-Lin and IPS tests for a unit root in relative unemployment rates

Levin-Lin test					IPS-test			
Lags	Obs.	$\rho - 1$	t^{*2}	$P > t^*$	Lags	Obs.	$W(\bar{t})^3$	$P > \bar{t}$
0	420	-0.116	-1.850	0.032**	0	420	-1.958	0.025**
1	410	-0.112	-1.643	0.050**	1	410	-1.506	0.066*
2	400	-0.117	-1.307	0.096*	2	400	-1.615	0.053*
3	390	-0.096	-0.013	0.495	3	390	-0.003	0.499
4	380	-0.101	-0.026	0.490	4	380	-0.099	0.461
mixed[1]	401	-0.117	-1.478	0.0697*	mixed[1]	401	-1.249	0.106

*,** significant at the 10, and 5 percent levels

[1]Average augmentation 1.9 lags

[2] t^* is distributed standard normal under the null

[3] $W(\bar{t})$ is distributed standard normal under the null

Table 2 summarizes the results of the Levin and Lin test. Qualitatively the results do not change when a trend is included. We can reject the null hypothesis of a unit root safely, if no or only one lag is included to allow for serial correlation in the errors. If a second lag is included, we can still reject the null at the 10 percent level. Moreover, the parameter estimate for the autoregressive coefficient does not change substantially across the different specifications. If three or more lags are included, we cannot reject the null anymore. Since the univariate ADF tests of the previous section suggest an average optimal lag length of roughly 2, we suppose that the model specification with two lags is most preferable. Alternatively, we include different lag lengths for all individual series according to the ADF specifications reported in Table 1. This test is reported in the last line of Table 2. Again, we can reject the null hypothesis of a unit root at the 10 percent level.

The parameter estimate for $(\rho - 1) = -0.117$ implies an autoregressive parameter of 0.883. This in turn means that the half-live of a shock to relative unemployment rates is 5.5 years, which is the number of years for the shock to decay by 50% and can be computed as $\ln(0.5 - \rho)$. This seems a moderate degree of persistence.

A restriction of the Levin and Lin test is that it requires ρ to be homogenous.[7] Im, Peasaran and Shin (2003) (IPS) propose an alternative testing procedure which allows for heterogenous ρ_i, which means a difference in the speed of convergence among regions. While the null hypothesis of the IPS test is the same as for the Levin and Lin test, the alternative hypothesis is more flexible. It states that at least one of the series is

[7]However, a panel data approach primilary deals with the problem of heterogeniety in intercepts and not with heterogeneities in the slopes. Therefore, the assumption of homogenous ρ need not be too restrictive.

Table 3: Levin-Lin test for a unit root, no constant

Lags	Obs.	$\rho - 1$	t^{*3}	$P > t^*$
mixed[1]	401	-0.035	-2.108	0.018**

** significant at the 5 percent level

[1]Average augmentation 1.9 lags;

[2] t^* is distributed standard normal under the null

Table 4: Pooled AR(1) estimation with fixed-effects (not reported)

Dependent variable: $(ur_i - ur_{ger})_t$		
constant	0.073	(2.74)***
$(ur_i - ur_{ger})_{t-1}$	0.880	(36.75)***
Obs.: 420, Years: 42, N : 10		
R^2 within: 0.79		
F(9, 409) = 2.38** (indiv. effect is zero)		

***,** significant at the 1 and 5 percent level respectively.

stationary. The results of the IPS tests are also reported in Table 2.

Again, the inclusion of a time trend does not alter our findings. We find a similar pattern as with the Levin and Lin test. However, our preferred specification with mixed lag lengths is marginally insignificant at the 10 percent level. Although the result obtained with the IPS test is less clear-cut than the results of the Levin and Lin test, the panel-based unit root tests are more in favor of the convergence hypothesis for regional unemployment rates than the univariate ADF tests in general.

In order to analyze whether our results are driven by the inclusion of intercepts we run the Levin and Lin without a constant term. This is equivalent to testing for absolute convergence conditional on the assumption that there are no deterministic long-run differences in regional unemployment rates to the aggregate level. A formal test of the conditional convergence hypothesis will be performed afterwards.

The results of the Levin and Lin test without constant terms are reported in the first line of Table 3. In accordance with the previous findings, the unit root can still be rejected, even at a higher level of significance.

In a last step, we try to discriminate between the conditional and unconditional convergence hypothesis. Having shown that the time series for relative unemployment rates are jointly stationary, we estimate a simple AR(1) fixed-effects model. This allows us to formally test for unconditional convergence by testing the joint significance of the fixed-effects. The fixed-effects estimation is reported in Table 4.

The $F-$test that all unit effects are zero is reported in the last row of the table.

Since we have to reject the hypothesis that all fixed-effects are insignificant, we find no evidence for unconditional convergence of regional unemployment rates. Therefore, the Levin and Lin test with fixed-effects reported in table 2 is the most appropriate one.

To sum up, the panel-based tests show some support for conditional convergence of relative unemployment rates during the period 1960-2002. However, the estimated speed of convergence is not fast. Moreover, differences in unemployment rates do not totally disappear since we only find conditional convergence. In other words, the panel-based tests suggest that there is a stable distribution of relative regional unemployment rates. Yet, the graphical examination of the time series for relative unemployment rates suggested that there might be a structural break in the means of the series. If there is indeed a structural break, the estimated degree of persistence will be biased upwards. The interesting question is whether the estimated speed of convergence changes substantially if one accounts for a structural break.

5 Unit root tests with structural breaks

5.1 Test procedure

As displayed in Figure 1, around the year 1980 the relative unemployment rates for the federal states seem to change permanently. After 1980, the northern regions, especially the city-states Bremen and Hamburg, exhibit a higher level of unemployment, while the southern states, e.g. Bayern and Baden-Würtemberg, have below average unemployment rates.

This observation calls for the inclusion of a structural break in the analysis. It also explains why relative unemployment rates can only be found to converge conditionally. Absolute convergence implies a zero mean of the relative series at all times, so that there cannot be a structural change. By contrast, conditional convergence implies an equilibrium relationship of regional unemployment rates and the stationarity of their distribution. If the equilibrium relation is non-unique, a one-time major shock may shift the economy from one equilibrium to the other and the relative unemployment rates are only regime-wise stationary. With this regime-wise stationarity, conditional convergence with a structural break implies on the one hand that in the absence of major shocks there is an equilibrium relationship between the unemployment rates of the various states, i.e. regional shocks have no persistent effect. On the other hand, a permanent change of the equilibrium relationship occurs when the regime shifts because of a one-time major shock. To put it simple, only very few regional shocks have persistent effects, most of them do not.

Although a theoretical explanation of an apparent level shift is interesting and important (Hansen, 2001), in this paper we only try to find the structural break and test for convergence. A theoretical explanation could for example base on: induced technological change, hysteresis effects, differences in regional specialization, differences in union density and bargaining power, see Martin (1997) for further examples.

Since we do not specify a structural model for the regime shift, we continue to follow the univariate time-series approach but extend the model to allow for a one-time level shift. The timing of the level shift, i.e. the structural break, is determined endogenously and data-dependent. Perron (1990) has shown that conventional ADF tests perform poorly when there is a structural break in the means of the series. As a consequence, a stationary series subject to a structural break can look like a non-stationary series if the break is not accounted for. The non-rejection of the univariate tests presented in Section 4 might be associated with the permanent change in the level of the series. Similarly, the moderate speed of convergence we find on the basis of the panel-based tests could also result from a structural break. However, the original test for a unit root in presence of a structural break proposed by Perron (1990) requires the break date to be known. Since we want to choose the breaking date data-dependent, we employ the Perron and Vogelsang (1992) test.

The literature on structural change in time series suggests two different models which differ in the way the transition from the old to the new level occurs. The first is called the "additive outlier model" (AO). Here, the transition after the break occurs instantaneously. The second is called the "innovational outlier model" (IO), where the break is assumed to follow the same path as the innovations and to occur slowly over time.

In our application, Figure 1 suggests that the adjustment which follows a level shift needs some years to take effect and does not occur instantaneously. Consequently, the IO model is more appropriate for our data. This finding is also in line with the general remark of Hansen (2001) who argues that a structural break is unlikely to be immediate.[8]

The IO model of the Perron and Vogelsang (1992) test can be described as follows: Denote T_b the date of the break with $1 < T_b < T$, where T is the sample size. The null hypothesis is specified as

$$y_t = y_{t-1} + \psi(L)(e_t + \theta D(TB)_t), \qquad t = 2, ..., T \tag{5}$$

where $\psi(L) = A^*(L)^{-1}B(L)$ defines the moving average representation of the noise

[8] We also tried the AO model, but as expected, its performance turned out to be inferior compared to the IO model, which means that the AO model rejects the null hypothesis in fewer cases.

function. The dummy variable $D(TB)_t$ is set to 1 if $t = T_b + 1$ and 0 otherwise. The dummy $D(TB)_t$ is a one-off impulse dummy which changes the level of the series after the break by θ under the null of a unit root. The long-run impact of the level change is given by $\psi(1)\theta$.

Under the alternative hypothesis of stationarity, the model is represented by

$$y_t = a + \phi(L)(e_t + \delta DU_t), \qquad t = 2, ..., T \tag{6}$$

where $\phi(L) = A(L)^{-1}B(L)$. The dummy variable DU_t is equal to 1 if $t > T_b$ and 0 otherwise. Hence, after the break the level becomes $(a + \delta)$. As suggested by Perron and Vogelsang (1992), models (5) and (6) can be nested and approximated by the finite-order autoregressive model

$$y_t = \mu + \delta DU_t + \theta D(TB)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \tag{7}$$

$$t = k + 2, ..., T$$

Similar to the augmented Dickey-Fuller regression, lags of first-differences of the dependent variable y are included on the right hand side of the equation. Model (7) can be estimated by OLS. Under the null hypothesis of a unit root, the autoregressive parameter α is equal to 1, which implies $\delta = \mu = 0$ because there is no trend. Since the appropriate value of T_b is unknown, there are two strategies to perform the unit root test. Under both options we first perform regression (7) for all possible breaking dates. Then, under the first option, the break date is chosen to minimize the t statistic on the autoregressive coefficient. In other words, this option selects the break date provide most evidence against the random walk hypothesis.

The alternative option identifies the breakpoint so that the t statistic (in absolute terms) on the coefficient associated with the change in the mean, δ, is maximized.

Asymptotic distributions and finite-sample critical values of the test statistics are derived in Perron and Vogelsang (1992). We generated critical values that correspond to our sample size with T=43 and $k_{\max}$=8 from 5000 replications of a Monte-Carlo experiment.

There are various procedures to select the appropriate order k of the estimated autoregressions. This in turn influences the critical values for the test statistics. As for the ADF regressions in Section 3, we choose k according to a significance test on the last included lag, given a pre-specified maximum of 8 years.

5.2 Test results

The results of the Perron and Vogelsang (1992) unit root tests obtained by minimizing the t-statistic on α over all possible breakpoints are summarized in Table 5. In seven out of ten cases we are able to reject the null hypothesis of a unit root in favor of regime-wise stationarity at least at the 10 percent level of significance. Recall that the univariate unit root tests without structural breaks rejected the random walk hypothesis only for one federal state. For three of the ten federal states we still cannot reject the null hypothesis of a random walk even after accounting for a structural break. These are Baden-Würtemberg, Niedersachsen and Schleswig-Holstein. However, with point estimates of 0.5 - 0.7 for α the non-rejection seems to be due to a lack of power. The weak power of the test can also be seen if we look at the opposite extreme cases. Although the estimates of α for Bremen and Hessen are virtually zero, the test rejects the hypothesis $\alpha = 1$ only at the 5 percent level of significance.

The data-dependent choice of the break date mostly coincides with the a priori assumption that the second oil crisis and the following recession had a huge and persistent impact on relative unemployment rates. For all but three series the chosen break date falls into the period of 1978-1982.

The three states for which the estimated break date is outside this period are Rheinland-Pfalz, Hessen and Schleswig-Holstein. For Rheinland-Pfalz, the ADF test without structural break already rejected the unit root. For Schleswig-Holstein, the estimated break date coincides with the first oil-crisis, but the unit root cannot be rejected. Only for Hessen, the break date is hard to interpret. It could be the German re-unification of 1989/90 that affects Hessen with a three year time lag in 1993.

The unemployment rate for Germany as a whole remains non-stationary even after accounting for a structural change in the level. This result is in line with the findings of Papell, Murray and Ghiblawi (2000)[9].

A comparison of the results for the two different methods to determine the break point reveals that our results are robust. Only in two cases, the two methods do estimate a different break point and /or a different number of augmentation lags. For Niedersachsen, the second method estimates the break point one year later without a change in the qualitative result of non-stationarity. For Rheinland-Pfalz the $\delta-$method yields a number of augmentation lags of 8 and cannot reject non-stationarity anymore.

To further test the robustness of our results, we also tried unit root tests which

[9] Papell, Murray and Ghiblawi (2000) analyze hysteresis in OECD unemployment rates. They adopt unit root tests with multiple structural breaks and show the West German unemployment rate to be non-stationary.

Table 5: Perron-Vogelsang (1992) unit root tests with level shift for nontrending data

Fed. State	T_b	k	α	δ	Fed. State	T_b	k	α	δ
BW	80	6	0.49	-0.69	NRW	80	6	0.28	0.78
			(-4.23)	(-3.76)				(-5.54)**	(5.05)
BY	81	5	0.30	-1.08	RP	70	0	0.38	-0.36
			(-5.03)**	(-4.81)				(-4.72)*	(-2.75)
BRE	82	5	0.05	3.75	SAAR	78	2	-0.41	2.06
			(-4.99)**	(4.78)				(-5.21)**	(4.56)
HH	82	1	0.27	1.75	SH	72	0	0.60	0.39
			(-5.22)**	(4.96)				(-3.36)	(2.37)
HE	93	4	0.05	0.75	ur_{Ger}	79	6	0.65	1.94
			(-4.60)*	(3.57)				(-3.67)	(3.53)
NS	78	2	0.69	0.27					
			(-3.51)	(2.30)					

k chosen according to a significance test on the last included lag, given a pre-specified maximum of $k = 8$; T_b, k, α, θ are obtained by minimizing the t-statistic on α;
*,** significant at the 10, and 5 percent levels, t–statistics in parenthesis

Critical Values	1%	2.5%	5%	10%
T_b chosen by min. $t_{\hat{\alpha}}$	-5.61	-5.25	-4.91	-4.53

obtained from the empirical distribution of 5000 replications
of a Monte Carlo experiment, T=43, max(k)=8,

allow for a break both in the intercept and the trend (Perron, 1997, Zivot and Andrews, 1992). Allowing for slope breaks provides little additional evidence against the unit root hypothesis. It cannot be rejected at a higher significance level because the power of the tests decline when unnecessary breaks are included.

5.3 Speed of convergence

One of the initial motivations for the structural break unit root tests was the moderate speed of convergence that we inferred from the panel-based unit root tests. Therefore, we analyze the half-life of a shock to relative unemployment on the basis of our results of the Perron and Vogelsang test. This of course only makes sense for those regions for which non-stationarity could be rejected. For those states where the unit root hypothesis cannot be rejected, shocks have a persistent effect and the implied half-life is infinite. From the Perron and Vogelsang regressions, we generated a moving average representation of the estimated autoregressive process for relative unemployment rates. This MA-representation is used to compute impulse-response functions that include the

Table 6: Half-lifes (in years) of shocks to relative unemployment rates, computed from impulse-response functions based on regression results as reported in Table 5

Federal State	BY	BRE	HH	HE	NRW	RP	SAAR
Half-life	2	1	2	1	3	1	1

Note: Three federal states are omitted, for which the relative unemployment series were found to be non-stationary.

augmentation lags. Thereafter, we define the half-life of a shock as the date at which the initial impulse has lost at least half of its effect for the first time.

The estimated half-lives are reported in Table 6 and vary between 1 and 3 years. In comparison, the implied half-life is 5.5 years when the Levin and Lin test regressions with unit-specific effects are used. These regressions do not include a structural break. Consequently, measured persistence is substantially upwards biased if the structural break is omitted.

6 Conclusions

The question of this paper was whether there are forces that lead to convergence in the levels of regional unemployment rates over time. We used German regional data on unemployment from the Mikrozensus covering the period 1960-2002 and performed univariate as well as panel unit root tests to examine the hypothesis of stochastic convergence. On the basis of univariate ADF tests the hypothesis of non-convergence cannot be rejected. But using more powerful panel unit root tests we found some evidence for conditional convergence in regional unemployment rates up to a stable equilibrium distribution. Yet, these tests imply a moderate speed of convergence. Since the graphical examination of the series suggested the presence of a shift in the equilibrium differential of regional unemployment rates after the second oil crisis, we extended the convergence tests to allow for such a shift. We employed the univariate unit root test of Perron and Vogelsang (1992) that includes a level shift in the series analyzed. In contrast to the univariate ADF test, the non-convergence hypothesis could be rejected for seven out of ten federal states. Moreover, in comparison to the results of the panel-based tests, the estimated speed of convergence increased substantially. Consequently, regional unemployment rates are found to converge quickly to a constant difference from the national average but this difference is not the same for the two regimes before and after the second oil crisis.

Both results, the presence of regime-wise conditional convergence in regional unemployment rates and fast equilibrium adjustment, have important implications for economic policy targeted at regional unemployment. On the one hand, small government interventions loose their effect quickly as unemployment rates adjust back to their equilibrium levels. On the other hand, large interventions might move the economy from one equilibrium of regional unemployment rates to the other. This means the policy intervention needs to take the form of a substantial regime shift. Most policies that aim at reducing relative unemployment differentials are unlikely to make permanent contributions to social welfare.

On the side of the econometric analysis, our paper, like many others, provides once more evidence of the low power of univariate tests in small samples. This problem is especially apparent in the setting with a structural break and we have dealt with it in two ways. Including the panel dimension and accounting for the structural break. An even more powerful approach would combine panel techniques and structural change. Following the general proposal of Madalla and Wu (1999), individual Perron and Vogelsang (1992) tests can be combined in a single test statistic. We leave this for further research.

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