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Do regional price levels converge?

Paneleconometric evidence based on German districts

by Christian Dreger[*] and Reinhold Kosfeld[**]

Abstract: We investigate price index convergence on the base of regional data for 439 German districts. Prices refer to the overall consumer price index as well as to the index without housing prices. To increase the efficiency of the testing framework, a panel unit root analysis is performed, where cross section dependencies are taken into account. The tests indicate a lack of regional price convergence. While the idiosyncratic component of price differentials is mostly stationary, their common component is driven by a unit root. The results are very similar for the overall price index and the index without housing prices, and for the Western and Eastern part of the German economy. Obviously the elimination of housing prices is not sufficient to obtain a price index where tradable products dominate. One rationale of our findings is the persistent west-east divide in consumer prices. A second argument is related to the persistence of the price gradient between urban and rural regions.

Keywords: Regional price differentials, price convergence, panel unit roots

JEL: E31, R10, C33

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1 Introduction

The often reported failure of the purchasing power parity (PPP) relationship in international datasets has led many researchers to study the condition on the intranational level, see Culver and Papell (1999), Cecchetti, Sonora and Mark (2000), and Chen and Devereux (2003), among others. In fact, there are numerous arguments in favour of this line of research: within the country borders, a higher degree of market integration and lower barriers to trade can be detected. Although transportation costs are still an obstacle to arbitrage, they are usually less important than in international markets. In addition, exchange rate fluctuations cannot distort PPP equilibrium, as the regions share the same currency. The composition of price indexes within a country is supposed to be more homogeneous than the one for different countries. On the other hand, however, a price gradient between urban and rural areas can be conspicuous. The price levels of comparable types of regions across countries may be even more homogenous as those of different types within the same country. Substantial deviations from PPP can as well occur even at the regional level, for example because of market segmentation or the presence of non traded goods.

Previous studies have been mostly able to find cointegration of regional price indexes, i.e. the relative price index of two regions seems to be stationary. However, the speed of adjustment is surprisingly slow. Often, rates of convergence are even lower than the rates in an international setup. Due to Cecchetti, Sonora and Mark (2000) annual inflation rates measured over 10 years intervals can differ by as much as 1.6 percentage points between US cities, see also Engel and Rogers (2001). The half life of convergence towards the PPP condition, i.e. the time span how long it takes for the impact of a shock to diminish by 50 percent is estimated to be around 9 years. Because the nominal interest rate is the same within the country, real interest rate differentials with a high degree of persistence would be implied and might have a large effect on the allocation of resources.

This paper adds to the literature in two ways. First, price index convergence is explored by means of a large regional dataset, where 439 German districts (*Kreise*) are considered. Because of the lack of official data, the series constructed by Kosfeld, Eckey and Lauridsen (2007) are used to proxy the regional price evolution. Second, panel unit root tests are applied to obtain a more efficient testing design of the null hypothesis of no convergence. In contrast to previous papers, cross section correlation is taken into account. In particular, a common factor approach is employed to capture the dependencies.

Overall, the evidence does not indicate price index convergence. While the idiosyncratic component of price differentials is mostly stationary, their common component is driven by a unit root. This result is rather robust and can be obtained for the whole index as well as for the one without housing prices, and for the Western and Eastern part of the economy. Obviously the elimination of housing prices is not sufficient to obtain a price index where tradable products dominate. Our findings can be traced to the persistent west-east divide in consumer prices and to the price gradient between urban and rural regions.

The paper is organized as follows: Section 2 introduces the theoretical model that is the workhorse in the study. Section 3 reviews the panel unit root tests, and section 4 describes the dataset used in the analysis. The results are discussed in section 5. Finally, section 6 concludes.

2 Convergence and unit root analysis

According to the PPP hypothesis, regional prices should be equalized in the long run. In other words, they could not permanently deviate from the cross country average. Under these conditions, the log of the relative price level,

$$(1) \qquad q_t = p_t - p_t^* .$$

is a stationary variable, where p is the regional price level and p^* denotes the numeraire, i.e. the regional average, both expressed in logrithms. As prices refer to the same currency, the relative price level corresponds with the real exchange rate. A simple structure to analyse price convergence is the AR(1) process,

$$(2) \qquad q_t = \alpha + \theta q_{t-1} + \varepsilon_t$$

where ε is a white noise innovation. Prices tend to converge to the same level if α=0 and θ<1, see Durlauf and Quah (1999). Then, the long run real exchange rate is expected to be equal to 0. A non zero constant captures the case of conditional convergence. Permanent, but constant deviations might be justified due to limitations of the arbitrage process, such as transportation costs. However, a slope parameter less than 1 is essential for the process of convergence to occur. Equation (2) might be rewritten in error correction form

$$(3) \qquad \Delta q_t = \alpha + \beta q_{t-1} + \varepsilon_t \quad , \quad \beta = \theta - 1$$

where the change in the real exchange rate depends negatively on its previous level, i.e. the gap between the regional price level and the cross country average. Thus, price convergence can be established by means of a unit root test, where the null hypothesis β=0 should be rejected in favour of the alternative, β<0. However, it has been widely acknowledged that standard unit root tests can have low power against stationary alternatives, see Campbell and Perron (1991). Panel unit root tests offer a promising alternative. Because the time series dimension is enhanced by the cross section, the results rely on a broader information set. Thus gains in power are expected, and more reliable evidence should be obtained (see Levin, Lin and Chu, 2002).

3 Panel unit root tests

Levin, Lin and Chu (LLC, 2002) and Im, Pesaran and Shin (IPS, 2003) have suggested tests of the null of a unit root in cross sectional independent panels. Both are generalizations of the ADF principle. Heterogeneity of panel members is allowed to some extent, and is shown in individual deterministic components (constants and time trends) and individual short run dynamics. The tests differ in the alternative considered. In the LLC approach, a homogeneous first order autoregressive parameter is assumed. The within-type statistic is built on the t-value of its estimator in a pooled regression. The IPS test is a between-type test and emerges as a standardized average z of individual ADF tests. If the null of a unit root is rejected, the series are stationary for at least one individual. Hence, the IPS test extends heterogeneity to the long run behaviour of the series. The test statistic

$$(4) \qquad z_t^* = \frac{z_t - \mu}{\sigma}$$

is asymptotically distributed as standard normal with a left hand side rejection area. Standardization factors μ and σ are obtained by simulation and depend on deterministic components involved in the regressions. Given that the panel members are independent, the Gaussian distribution can be justified by central limit arguments. In case of cross section dependencies, the tests are not valid. However, if the dependencies are caused by common time effects they can be eventually removed by subtracting cross-sectional means from the data.

Second generation panel unit root tests have relaxed the independency assumption, see Hurlin (2004) and Gengenbach, Palm and Urbain (2004) for recent surveys. In this paper, the tests suggested by Pesaran (2003) and Bai and Ng (2004) are applied. Both procedures capture the cross sectional correlation pattern by a common factor structure.

Pesaran (2007) has motivated a single factor approach. The common component is assumed to be stationary and embedded in the error process of the model. The procedure is a cross sectional extension of the ADF framework (CADF). The standard ADF regression is extended by cross sectional averages of lagged levels and differences of the series of interest. In the model

$$(5) \qquad \Delta y_{it} = a_{0i} + \alpha_{1i} y_{i,\,t-1} + \alpha_{2i} \bar{y}_{t-1} + \alpha_{3i} \Delta \bar{y}_{t-1} + \varepsilon_{it}$$

$$\bar{y}_t = n^{-1} \sum\nolimits_{i=1}^{n} y_{it}\ ,\ \ \Delta \bar{y}_t = n^{-1} \sum\nolimits_{i=1}^{n} \Delta y_{it}$$

the cross sectional average $\bar{y}_t$ of the series under consideration serves as a proxy for the single factor. Testing for the null of a unit root refers to the t-ratio of the first order autoregressive parameter. Equation (5) can be interpreted as an alternative to the ADF in a time series setting, where information from other individuals is allowed to enter through the common component. Due to this extension, the critical values exceed those in the standard ADF setting in absolute value. The panel version arises from a cross sectional extension of the IPS test (CIPS), where t-ratios are pooled across the individuals. The limiting distribution of the CIPS procedure is non-standard, whereas the extent of the deviations from normality depend on the deterministic terms included in the model. The critical values have been tabulated by Pesaran (2007).

Since the common component is stationary by assumption, nonstationarities can occur only because of idiosyncratic developments. In the PANIC approach (Panel Analysis of Nonstationarity in Idiosyncratic and Common components) suggested by Bai and Ng (2004), both common and idiosyncratic components are allowed to show nonstationary behaviour. The unit root hypothesis is tested separately for the two components. Therefore, further information into the possible sources of nonstationarity is provided.

In particular, the variable of interest Y_{it} is expressed as the sum of a deterministic component, a common component expressed by a dynamic factor structure, and an idiosyncratic component, which accounts for the error term. For the i-th panel member and time t, the decomposition

(6) $$Y_{it} = \alpha_i + \lambda_i ' F_t + \varepsilon_{it}$$

is applied, where α_i is a fixed effect, eventually including a linear time trend, F_t is the rx1 vector of common factors, λ_i is an rx1 vector of factor loadings and ε_{it} is the idiosyncratic component. The parameter r denotes the number of factors and is estimated by information criteria discussed in Bai and Ng (2002). The series Y_{it} includes unit roots if one or more of the common factors are nonstationary, or the idiosyncratic error is nonstationary, or both. Principal components are employed to estimate the common factors. However, as the components might be integrated, a suitable transformation is required in advance. Bai and Ng (2004) suggest to perform the principal component analysis on the differenced data, which are assumed to be stationary. Once the components have been estimated, they are cumulated again to match the integration properties of the original variables.

A unit root in the idiosyncratic component can be examined by ADF tests. However, they will have low power, as in the time series case. First generation panel unit root tests are more efficient. They can be used, as the defactored data are uncorrelated across the panel members. For the common component, the appropriate strategy depends on the number of factors of the series considered. If there is only a single factor, a standard ADF regression with a constant

(7) $$\Delta F_t = \alpha + \phi_0 F_{t-1} + \sum_{i=1}^{p} \phi_i \Delta F_{t-i} + u_t$$

is employed, and inference is based on the Dickey Fuller distribution. If there are multiple common factors, the Johansen trace test can be employed to determine the cointegration rank. Jang and Shin (2005) conclude from their small sample analysis that the PANIC approach has a better test performance than the procedure suggested by Pesaran (2003).

4 Construction of regional price data

Regional price levels are not available from official statistics. Therefore, the panel data reported by Kosfeld, Eckey and Lauridsen (2007) are used to study price convergence. Data are based on regional price models for the consumer price index without housing (*CPI-H*) and the housing rent index (*HRI*), the latter as a proxy for non tradable goods. The models are derived from utility maximization of representative consumers and are calibrated using fragmentary price data from two sources. In case of the *CPI-H*, a price survey of 50 German cities is available for 1993. Data on housing rents are published for 2004 by the Federal Office for Building and Regional Planning at the level of German districts.

Panel data for the *CPI-H* and *HRI* are obtained by using a two step approach, see Kosfeld, Eckey and Lauridsen (2007) for a detailed exposition of the procedure. Preliminary indices refer to the fit of regressions equations derived from a price model for consumer goods and housing. Afterwards the indices are adjusted by state wide inflation. The overall regional price level (*CPI*) is obtained through a linear combination of both subindices,

$$CPI = \begin{cases} 0.815 * CPI - H + 0.185 * HRI & , \quad \text{West German districts} \\ 0.879 * CPI - H + 0.121 * HRI & , \quad \text{East German districts} \end{cases}$$

for 1995-1999 and

$$CPI = 0.788 * CPI - H + 0.212 * HRI$$

for both parts of the country in the other half of the period (2000-2004). The coefficients reflect the weights of both types of commodities in the particular consumer baskets that are equal across the respective regions (Statistisches Bundesamt, 1998, 2003). Population-weighted price indices are normalized to be equal to 100 in 1995 for the entire economy.

In the regressions carried out at the level of districts, more than 90 percent of the *CPI-H* dispersion can be attributed to purchasing power, external demand, wages, agglomeration effects, and demographic variables. *HRIs* tend to be more difficult to explain. The coefficient of determination for rents decreases by 10 percentage points. Table 1 shows some measures on the three regional price indices for the sample period. Note that disparities in housing rents are quite larger. High *HRI* regions are not only located in the Western, but also in the Eastern part of the economy, particularly in the surrounding area of Berlin.

-Table 1 about here-

5 Empirical results

The analysis is based on panel data for 439 German districts covering the 1995-2004 period. The evidence is based on the IPS test, where common time effects are removed, the CIPS test and the PANIC approach. These measures are applied to the overall *CPI* and the *CPI* without housing (table 2).

-Table 2 about here-

The optimal lag length for IPS and CIPS is selected using a general-to-simple approach (Campbell and Perron, 1991). According to the alternative information criteria suggested by Bai and Ng (2002), the number of factors for the PANIC approach is not unique. Either one or two factors seem to be appropriate, and the analysis is conducted for both settings. However, the evidence is robust to this choice. The results are reported for the case of a single factor, but look very similar in the two factor model. All tests are specified with a constant, but no time trend.

Relative prices seem to be nonstationary in each case. This finding holds both for the overall index and the one without housing prices, and for the Western as well as for the Eastern part of the economy. The stochastic trend of regional price differentials has a unit root driven by their common component. This finding might be caused by a substantial part of non-tradables in the *CPI-H*. Putting it differently, the elimination of housing from the overall *CPI* is not sufficient to construct a price index, where tradable products dominate.

One explanation for the lack of price convergence across districts over the whole country can be found in the west-east divide. Table 3 shows that the consumer price indices in the Western and Eastern part differ significantly at the beginning and end of the sample period. Absolute and relative price differentials have been increased. The West-East indices differ also significantly when housing is excluded. While the absolute difference for *CPI-H* is unchanged, the relative difference has decreased slightly.

-Table 3 about here-

Another explanation for persistent price differentials comes from the downward price gradient from urban to rural areas. It does not only account for non-convergent price behaviour in the entire economy, but also for persistent disparities within the Western and Eastern part of the country. Price differentials between different types of regions within Germany may be even stronger than those between the same types of regions across countries. For both price indices, *CPI* and *CPI-H,* they have increased during the period of investigation.

6 Conclusion

In this paper we investigate price index convergence on the base of regional data for 439 German districts. Prices refer to the overall consumer price index as well as to the index without housing prices. To increase the efficiency of the testing framework, a panel unit root analysis is conducted, where cross section dependencies are taken into account. The tests indicate a lack of regional price convergence. The nonstationarity of price differentials is due to a unit root in their common component. For Germany as a whole, one explanation of this outcome is the persistent west-east divide in prices of consumer prices. A second explanation lies in the persistence of the price gradient between urban and rural regions. This explains why the results are similar for the Western and Eastern part of the economy.

The results are very similar for the overall price index and the index without housing prices, and for the Western and Eastern part of the German economy. While the idiosyncratic component of the price index without housing is always stationary, the performance is somewhis uniformly nonstationary for both indices. This finding might be due to a large share of non-tradables in the *CPI*. Therefore, the elimination of housing prices proves to be not sufficient to obtain a price index where tradable products dominate.

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Table 1: Descriptive price statistics

Variable	Year	Mean	Std deviation	Minimum	Maximum
CPI	1995	92.2	3.816	86.0	108.1
	1996	93.5	3.798	87.4	109.5
	1997	95.4	3.854	88.7	112.2
	1998	96.3	3.958	89.4	114.4
	1999	96.8	4.081	89.8	115.5
	2000	98.0	4.771	89.0	119.3
	2001	99.9	4.919	91.1	121.7
	2002	101.3	5.144	91.2	123.5
	2003	102.2	5.597	93.9	129.8
	2004	104.0	5.668	95.4	131.2
CPI-H	1995	93.3	2.377	88.4	99.7
	1996	94.3	2.338	89.4	100.8
	1997	96.0	2.253	91.0	102.5
	1998	96.9	2.254	91.7	103.6
	1999	97.4	2.343	91.9	104.3
	2000	99.0	2.490	93.3	106.3
	2001	101.2	2.469	95.4	108.7
	2002	102.6	2.577	96.2	110.6
	2003	103.8	2.628	96.7	111.9
	2004	105.8	2.719	97.8	114.5

Note: Data provided by Kosfeld, Eckey and Lauridsen (2007). Price data of 439 German districts: *CPI* Consumer price index, *CPI-H* Consumer price index without housing.

Table 2: Panel unit root tests of relative CPIs

Entire economy

	CPI	*CPI-H*
IPS	7.288	3.323
CIPS	-1.820	-1.233
PANIC	-1.639 | -4.774*	-2.009 | -3.655*

West Germany

	CPI	*CPI-H*
IPS	6.245	5.533
CIPS	-0.893	-1.082
PANIC	-1.863 | -5.474*	-1.910 | -7.974*

East Germany

	CPI	*CPI-H*
IPS	3.424	6.117
CIPS	-0.795	-1.121
PANIC	-1.564 | -1.193	0.488 | -6.002*

Note: Sample period 1995-2004, 439 German districts. IPS=Im, Pesaran and Shin (2003) test with cross sectional means, CIPS=Pesaran (2007) test, PANIC= Bai and Ng (2004) test, left entry common component, examined by ADF test, right entry idiosyncratic component, examined by IPS test. * indicates a rejection of the null hypothesis of nonstationarity at the 0.05 level of significance.

Table 3: Tests on differences in price level

	Region	1995	2004	Districts	1995	2004
CPI	West	93.4	105.6	Urban	93.9	106.9
	East	88.7	99.3	Rural	91.5	103.0
	t-value	13.38*	11.71*	*t*-value	6.08*	6.27*
CPI-H	West	94.4	107.0	Urban	94.6	107.5
	East	89.9	102.5	Rural	92.8	105.2
	t-value	30.16*	21.63*	*t*-value	7.42*	8.31*

Notes: Two-sample *t*-test for differences of population-weighted means of regional price indices; * indicates a rejection of the null hypothesis of equal average price levels at the 0.01 level of significance.