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CONVERGENCE OF PRICES AND RATES OF INFLATION

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Abstract

We consider how unit root and stationarity tests can be used to study the convergence properties of prices and rates of inflation. Special attention is paid to the issue of whether a mean should be extracted in carrying out unit root and stationarity tests and whether there is an advantage to adopting a new (Dickey-Fuller) unit root test based on deviations from the last observation. The asymptotic distribution of the new test statistic is given and Monte Carlo simulation experiments show that the test yields considerable power gains for highly persistent autoregressive processes with "relatively large" initial conditions, the case of primary interest for analysing convergence. We argue that the joint use of unit root and stationarity tests in levels and first differences allows the researcher to distinguish between series that are converging and series that have already converged, and we set out a strategy to establish whether convergence occurs in relative prices or just in rates of inflation. The tests are applied to the monthly series of the Consumer Price Index in the Italian regional capitals over the period 1970-2003. It is found that all pairwise contrasts of inflation rates have converged or are in the process of converging. Only 24% of price level contrasts appear to be converging, but a multivariate test provides strong evidence of overall convergence.

JEL classification: C22, C32.

Keywords: Dickey-Fuller test, Initial condition, Law of one price, Stationarity test.

Contents

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1. Introduction[1]

The issue of price and inflation convergence between countries belonging to a common currency or trade area or between regions in the same country has attracted considerable interest in the recent years. This is especially pertinent in Europe because of increased economic integration and the establishment of the European Monetary Union.

There are economic reasons why prices may not converge within countries belonging to a monetary union or within regions in the same country. A branch of recent economic literature (see, for example, Engel and Rogers, 1998, 2001; Parsley and Wei, 1996; Cecchetti *et al.* 2002) has pointed out that theories of market segmentation typically applied to the field of international economics can also explain permanent or temporary deviations from the law of one price within a currency or trade union, or within a single country. The dynamics of relative price levels can be influenced by transportation costs that impede the effective arbitrage across areas, by firms exercising local monopoly power and pricing to segmented markets, by the presence of non-traded goods in the price basket considered and by different speed in sticky price adjustment across areas.

Empirical evidence at the regional level is rather scant and refers mainly to the United States. Parsley and Wei (1996) analyse convergence to purchasing power parity across United States cities on the basis of price levels of individual goods and find that convergence rates are much higher than those found for cross-country data, although transport costs seem to account only for a small portion of the difference. Chen and Devereux (2003) observe a decline in the dispersion of tradable price levels across United States cities, hence supporting the convergence hypothesis. Cecchetti *et al.* (2002), on the basis of disaggregated consumer price indices, argue that deviations from the law of one price across cities in the United States can be mainly attributed to the distance between locations and to nominal price stickiness. Using the same type of data, but at the aggregate level, Engel and Rogers (2001) find that relative price levels mean revert but at a surprisingly slow rate.

[1] We would like to thank Roberto Sabbatini, Rob Taylor and participants in the *Frontiers in Time Series Analysis* conference, Olbia 2005, and the Bank of Italy's *Workshop on Inflation Convergence in Italy and the Euro Area*, Roma 2004, for insightful comments on earlier versions of the paper and Angela Gattulli for her valuable support with the data. The views expressed here are those of the authors, not the Bank of Italy. Andrew Harvey gratefully acknowledges the hospitality and financial support of the Research Department of the Bank of Italy.

Studies on regional price level patterns within European countries are even scarcer. Caruso, Sabbatini and Sestito (1993) focus on the time series properties of the Italian provincial consumer price indices and find, using univariate unit root tests, that the structure of relative prices is rather stable. On the other hand, in a study of provincial inflation and relative price shifts in Spain, Alberola and Marqués (1999) show that, while inflation differentials are rather small, deviations of relative prices from equilibrium can be large and very persistent.

In this paper we consider how unit root and stationarity tests can be used to study the convergence properties of price levels and inflation rates. We pay special attention to the issue of whether a mean should be subtracted when carrying out stationarity tests and whether there is an advantage to working in terms of deviations from the last observation when carrying out unit root tests for convergence. We derive the asymptotic distribution of a Dickey Fuller test statistic for data expressed as deviations from the last observation and evaluate its power properties by Monte Carlo simulation experiments. It is shown that this test allows considerable power gains for highly persistent autoregressive processes with "relatively large" initial conditions, the case of primary interest for analysing convergence.

Our work contributes to the existing literature in a number of ways. First, in focussing on regions within the same country, we indirectly examine the effect of real factors related to market segmentation in preventing a complete adjustment in relative price levels and hence in accounting for deviations from the law of one price, as opposed to other factors which might be more relevant in an international framework, such as tariff and exchange rate movements. In this light, the results of our analysis might provide some evidence for a tentative understanding of price and inflation convergence within the European Monetary Union. Second, we introduce a new Dickey-Fuller type-test and evaluate its properties. Third, we use econometric tests in a rather novel way that has relevance for other studies of convergence. The results of stationarity and unit root tests are combined to give information on whether inflation rates and prices have converged or whether they are in the process of converging. Furthermore, when we use multivariate tests we account for the cross-correlation between regions, rather than following most of the existing empirical studies in using panel unit root tests under the unlikely assumption of cross-sectional independence. Indeed, these panel techniques, while allowing considerable gains in terms of power of the tests, can also lead to serious size distortions by neglecting cross-sectional correlation, as demonstrated in O'Connell (1998). In such circumstances, when the number of units is not excessively large, a better strategy is to apply multivariate unit root tests that specifically account for such correlation, as in Abuaf and Jorion (1990), Taylor and Sarno (1998), Flôres et al (1999), Phillips and Sul (2002), Harvey and Bates (2003).

The unit root and stationarity tests are applied in this paper to the monthly series of the Consumer Price Index (CPI) in the twenty Italian regional capitals over the period 1970-2003. As the index is an aggregate measure built up from prices of individual goods and not an absolute price level, we investigate what might be labelled, following Engel and Rogers (2001), the "proportional law of one price", or, in other words, convergence in relative price levels across regions.

The paper proceeds as follows. Section 2 sets out the theoretical framework for testing the hypotheses of stability and convergence. The properties of the new Dickey-Fuller-type test on data expressed in terms of deviations from the last observation are compared with those of standard unit root tests. In section 3 it is shown that the joint use of unit root and stationarity tests in levels and first differences allows one to distinguish between series that are converging from series that have already converged, and a strategy for establishing whether convergence occurs in relative prices or just in rates of inflation is proposed. It is also shown that stationarity tests on first differences can be biased if the data in levels are highly persistent. The application to Italian regions is described in section 4 and section 5 concludes.

2. Stability and convergence

In the time series literature on convergence there is often some confusion on the role played by unit root and stationarity tests for detecting convergence. The two types of tests are in fact meant for different purposes. Unit root tests are more useful to establish whether two (or more) variables are in the process of converging, with large part of the gap between them depending on the initial conditions. Stationarity tests, on the other hand, are the more appropriate tool to verify whether the series have converged, i.e. whether the difference between them tends to remain stable. In other words, there is the need to distinguish between *convergence* and *stability*, as defined in the following subsections.

2.1 *Stability*

If the difference between two nonstationary time series, y_t, is a stationary process with finite non-zero spectrum at the origin, we will say they have a stable relationship. The null hypothesis of stability may be tested by a stationarity test. Such a test will reject for large values of

$$\xi_1 = \frac{\sum_{t=1}^{T}\left(\sum_{j=1}^{t} e_j\right)^2}{T^2\widehat{\omega}^2}, \tag{1}$$

where $e_t = y_t - \overline{y}$ and, following Kwiatkowski, Phillips, Schmidt and Shin (1992), hereafter KPSS, $\widehat{\omega}^2$ is a non-parametric estimator of the long run variance of y_t, that is

$$\widehat{\omega}^2 = \hat{\gamma}(0) + 2\sum_{\tau=1}^{m} w\left(\tau, m\right)\hat{\gamma}(\tau), \tag{2}$$

with $w\left(\tau, m\right)$ being a weight function, such as the Bartlett window, $w(\tau, m) = 1 - |\tau|/(m+1)$, and $\widehat{\gamma}(\tau)$ the sample autocovariance of y_t at lag τ. The bandwidth parameter m must be such that, as $T \to \infty$, $m \to \infty$ and $m^2/T \to 0$; see Stock (1994). The 10%, 5% and 1% critical values for the asymptotic distribution are 0.347, 0.461 and 0.743, respectively.

If the mean is known to be zero under the null, then y_j rather than e_j is used to construct the test statistic, now denoted[2] by ξ_0. Under the null hypothesis of zero-mean stationarity of y_t, the asymptotic distribution of ξ_0 is given by the integral of a squared Brownian motion process, rather than a Brownian bridge; see McNeill (1978) and Nyblom (1989). The 10%, 5% and 1% critical values are 1.196, 1.656 and 2.787, respectively.

The ξ_0 test will have power against a stationary process with a non-zero mean as well as against a non-stationary process. As shown in Busetti and Harvey (2002), another effective test can be based on the non-parametrically corrected 't-statistic' on the mean of y_t, that is $t_{\overline{y}} = \widehat{\omega}^{-1}T^{-\frac{1}{2}}\sum_{t=1}^{T} y_t$. Under the null hypothesis of zero mean stationarity $t_{\overline{y}}$ converges to a standard Gaussian distribution. Busetti and Harvey (2002) show that this t-test is nearly as powerful as ξ_0 against non-stationarity but is much more powerful against the alternative of a non-zero mean; they advise it be used when either alternative is of interest. Parametric versions of the tests are also possible.

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[2] Unlike the case when the mean is subtracted, the statistic is different when reverse partial sums are used; see Busetti and Harvey (2002). This is not of any practical importance in the present context.

2.2 *Convergence*

If y_t is stationary (with finite non-zero spectrum at the origin), the series have already converged. However, they may be in the process of converging, have just converged or have converged some time ago but with a large part of the series dependent on initial conditions. We therefore need a modelling framework that can capture the convergence process. A suitable model will be asymptotically stationary, satisfying the condition that

$$\lim_{\tau \to \infty} E(y_{t+\tau}|Y_t) = \alpha, \tag{3}$$

where Y_t denotes current and past observations. Convergence is said to be *absolute* if $\alpha = 0$, otherwise it is *relative* (or conditional); see, for example, Durlauf and Quah (1999). The simplest such convergence model is the AR(1) process

$$y_t - \alpha = \phi\,(y_{t-1} - \alpha) + \eta_t, \quad t = 2, ..., T, \tag{4}$$

where η_t's are martingale difference innovations and y_0 is a fixed initial condition. By rewriting (4) in error correction form as

$$\Delta y_t = \gamma + (\phi - 1)y_{t-1} + \eta_t, \tag{5}$$

where $\gamma = \alpha(1 - \phi)$, it can be seen that the expected growth rate in the current period is a negative fraction of the gap between the two series after allowing for a permanent difference, α. We can therefore test against convergence, that is $H_0 : \phi = 1$ against $H_1 : \phi < 1$, by a unit root test. The power of the test will depend on the initial conditions, that is how far y_0 is from α. If α is known to be zero, the test based on the Dickey-Fuller (DF) $t-$statistic with no constant, denoted τ_0, is known to perform well, with a high value of $|y_0|$ actually enhancing power; see Müller and Elliott (2003) and Harvey and Bates (2003). Although the τ_0 test is not invariant to y_0 it appears to be quite robust in that y_0 has little effect on its distribution under the null hypothesis.

What happens when testing for relative convergence? Including a constant in the DF regression and computing the $t-$statistic, denoted as τ_1, reduces power considerably. The test of Elliott, Rothenberg and Stock (1996), hereafter denoted ERS, also performs rather poorly as $|y_0 - \alpha|$ moves away from zero; again see Müller and Elliott (2003) and Harvey and Bates

(2003). A possible way of enhancing power in this situation is to argue that in view of (3) we should set α equal to y_T and then run the simple DF test (without constant) on the observations $y_t - y_T$, $t = 1, ...T - 1$. We will denote this test statistic as τ^*. When $\phi = 1$, the asymptotic distribution of τ^* is

$$\tau^* \rightarrow \frac{-(W(1)^2 + 1)}{2\left[\int_0^1 W(r)^2 dr\right]^{1/2}} = \frac{-(\chi_1^2 + 1)}{2\left[\int_0^1 W(r)^2 dr\right]^{1/2}} \tag{6}$$

where $W(r)$, is a standard Wiener process; see the appendix. Note that this differs from the asymptotic distribution of the τ_0 statistic in the sign attached to the one in the numerator. Simulated quantiles are shown in table 1 in the column labelled $N = 1$. The power properties of the τ^* test are evaluated in the next subsection by Monte Carlo simulation experiments. It turns out that it is considerably more powerful than τ_1 for series that start far apart.[3]

A possible objection to τ^* is that it introduces noise into the proceedings because of the variability in the last observation. This effect might be mitigated by estimating α by a weighted average of the most recent observations.[4] Some rationale for this may be obtained by considering the theory for the ERS test. This involves the estimation of α by

$$\widehat{\alpha}_c = \left[y_1 + (1 - \overline{\phi}) \sum_{t=2}^{T} (y_t - \overline{\phi} y_{t-1})\right] / [1 + (T-1)(1-\overline{\phi})^2] \tag{7}$$

where $\overline{\phi} = 1 + \overline{c}/T$. The recommended value of $\overline{c}$ is 7, as in Elliott, Rothenberg and Stock (1996). If $\overline{c} = 0$ we end up subtracting the first observation. The asymptotic distribution for the t-statistic formed from $y_t - \widehat{\alpha}_c$ is the standard one for τ_0. The de-meaning is based on GLS estimation, assuming that $\alpha = y_0$. If instead we set $\alpha = y_{T+1}$, then we find

$$\widehat{\alpha}_c^* = \left[\overline{\phi}^2 y_T + (1 - \overline{\phi}) \sum_{t=2}^{T} (y_t - \overline{\phi} y_{t-1})\right] / [\overline{\phi}^2 + (T-1)(1-\overline{\phi})^2] \tag{8}$$

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3 The test based on subtracting the last observation, τ^*, would also display some power against an explosive autoregression. However unreported simulations show that in such circumstances the upper tail Dickey-Fuller test with constant, τ_1, rejects the null much more frequently than τ^* does.

4 It is, however, worth noting that in the LM type test - what Stock (1994) calls the Sargan-Bhargava test - the test statistic is constructed from deviations from the first observation. (Subtracting the last observation instead makes little difference to power).

As in (7) the weights sum to unity. Denote the resulting test statistic as $\tau^* - GLS$. As $\overline{\phi}$ approaches one, all the weight goes on to y_T and we obtain τ^*. More generally, a higher order autoregression is used, that is

$$\Delta y_t = \gamma + (\phi - 1)\, y_{t-1} + \gamma_1 \Delta y_{t-1} + ... + \gamma_{p-1} \Delta y_{t-p+1} + \eta_t, \tag{9}$$

The Augmented Dickey-Fuller (ADF) test is based on such a regression. ERS recommend the use of (9), without the constant, having first subtracted $\widehat{\alpha}_c$ from y_{t-1}. An alternative would be to estimate α from (9) with ϕ set to $\overline{\phi}$. When $y_{T+1} = \alpha$ this leads to an estimator that places relatively more weight on the last p observations; see the appendix. Another possibility is to work within an unobserved components framework where the model is an AR(1) plus noise. In this case $\widehat{\alpha}_c^*$ is replaced by an estimator close to an exponentially weighted moving average (EWMA). The asymptotic distribution of all these modified ERS statistics under the null hypothesis is the same as τ^*.

The contrast between (log) price indices in Florence and Aosta shown in Figure 1 for seasonally unadjusted data provides an illustration. After seasonal adjustment, we use ADF-type regressions to compute the statistics τ_1 and τ^* with number of lags chosen according to the modified AIC criterion (MAIC) of Ng and Perron (2001). We obtain $\tau_1 = -2.53$ and $\tau^* = -2.95$, where α is estimated as the average of the last twelve months. Thus including a constant term implies a non-rejection of the null hypothesis even at 10% level of significance, while with τ^* we reject the null at 5% significance. Notice that in this example the series start quite far apart: the ratio of the initial condition to the residual standard deviation is about 26 in a sample of 408 observations.

2.3 *Monte Carlo evidence on the power of the τ^* test*

Here we report Monte Carlo simulation experiments designed to compare the power properties of τ_1 and τ^* for a near-unit root data generating process, for a range of initial conditions. Specifically we consider the AR(1) data generating process, $t = 1, 2, ..., T,$

$$\begin{aligned} y_t &= \alpha + u_t \\ u_t &= (1 - c/T)u_{t-1} + \eta_t, \qquad \eta_t \sim NIID(0, 1) \end{aligned}$$

with c taking on the values $0, 1, 2.5, 5, 10$ and $u_0 = \alpha + K$, with K varying among 0, 5, 10, 15, 20, 25, 30 and 50. The notation $NIID(a, b)$ indicates a Gaussian independent and identically distributed process with mean a and variance b. Thus y_t is a highly persistent process for $c > 0$ and a unit root process for $c = 0$. The τ^* test is simply based on $y_t - y_T$ without constant. Since the test statistics are invariant to α this is set equal to zero. K is the magnitude of the initial condition in units of the errors standard deviation.

Tables 2a,b contain the simulated rejection frequencies of these tests for $T = 100$ and 400 and a 5% significance level, which for τ^* is -2.69. For quarterly data, $T = 100$ might be most relevant. In this case $c = 5$ is quite plausible as it corresponds to $\phi = 0.95$; a smaller ϕ would mean unusually fast convergence. A value above 0.975 ($c = 2.5$) is quite slow. As can be seen, for $c = 2.5$ and 5, τ^* is considerably more powerful than the standard ADF test τ_1 when the initial condition is relatively large. In fact τ_1 is only better when K is 5 or zero and then the power is so low as to render the tests useless. The case of $T = 400$ is more relevant for monthly data. Here $c = 5$ corresponds to $\phi = 0.9875$ and this is a typical value. When $c = 1, 2.5, 5$ τ^* is more powerful than τ_1 for $K \geq 20$.

In this local-to-unity framework (with the autoregressive parameter depending on the sample size and the initial condition fixed), enlarging the sample results in lower power. It also implies that the power gains of τ^* for a large initial condition are lower the larger is the sample. On the other hand, if the autoregressive parameter is kept fixed (e.g. $c = 2.5$ with $T = 100$ versus $c = 10$ with $T = 400$) the power increases with the sample size for given initial condition.

2.4 *Multivariate tests*

Let $\mathbf{y}_t$ be the $N = n - 1$ vector of contrasts between each region and a benchmark. If the benchmark is the n-th region, then $\mathbf{y}_t = \left(y_t^{1,n}, y_t^{2,n}, \ldots, y_t^{n-1,n}\right)'$ where $y_t^{i,j} = \log p_{i,t} - \log p_{j,t}$. Most of the empirical literature on convergence across a group of regions is based on panel unit root tests, as in Evans and Karras (1996), Levin et al. (2002) and Im et al. (2003), and panel stationarity tests, as in Hadri (2000). However, these panel tests assume the contrasts to be mutually independent, a condition that is unlikely to be satisfied for most macroeconomic series. O'Connell (1998) and Bornhorst (2003) have investigated the size distortion and power loss of these tests under cross-sectional dependence and shown that it can be considerable. In the Appendix we describe a class of multivariate unit root and stationarity tests that take account of cross correlations among the series[5] and they are invariant to pre-multiplication of $\mathbf{y}_t$ by a nonsingular $N \times N$ matrix (thus, in our context, they are invariant to which region is chosen as a benchmark[6]).

3. Testing stability and convergence in levels and first differences

For data on prices it is of interest to test the hypotheses of stability and convergence in both levels and first differences, that is to analyze the dynamics of both relative prices and inflation differentials. Let $P_{i,t}$ denote some weighted average of prices in region i at time t. If information is available only for a price index, the observations are

$$p_{i,t} = P_{i,t}/P_{i,b}, \quad i = 1, ..., n, \quad t = 1, ..., T$$

where $b \in \{1, ..., T\}$ is the base year. The difference - or contrast - between (the log of) this price index and one in another region, say region j, denoted $y_t^{i,j}$, is

$$y_t^{i,j} = \log p_{i,t} - \log p_{j,t}, \quad t = 1, ..., T \tag{10}$$

where $y_b^{i,j} = 0$ by definition. This is the logarithm of the relative price between the two regions. The base can always be changed to a different point in time, τ, by subtracting y_τ from all the observations. It is not possible to discriminate between absolute and relative convergence with price indices; all that can be investigated is convergence to the proportional law of one price. The appropriate test for stability is ξ_1. Not subtracting the mean gives a test statistic, ξ_0, that is not invariant to the base and does not give the usual asymptotic distribution under the null hypothesis of a zero mean stationary process since treating the $y_t's$ as independent is incorrect. A test of convergence, on the other hand, can be based on a DF statistic,τ^*, formed by taking the base to be the last period.

The contrasts in the rate of inflation, or inflation differentials,

$$\Delta y_t^{i,j} = \Delta \log p_{i,t} - \Delta \log p_{j,t}, \quad t = 1, ..., T \tag{11}$$

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[5] Panel tests that relax the assumption of cross-sectional independence are described in the recent survey of Breitung and Pesaran (2005). See also Banerjee (1999).

[6] The tests are also invariant if the contrasts are formed by subtracting a weighted average of the series. However, this is not true if the weighted average is constructed before taking logarithms.

are invariant to the base year since this cancels out yielding $\Delta y_{i,t} = \Delta \log P_{i,t} - \Delta \log P_{j,t}$. A test of the null hypothesis that there are no permanent, or persistent, influences on an inflation rate contrast amounts to testing that Δy_t is stationary with a mean of zero. The appropriate tests are therefore ξ_0 and $t_{\overline{y}}$. Similarly the null hypothesis of no convergence in an inflation rate contrast against the alternative of absolute convergence can be tested using τ_0, the t-statistic obtained from an ADF regression without a constant.

3.1 *A testing strategy*

Taking account of the results of unit roots and stationarity tests allows the researcher to distinguish between regions that *have already converged* (characterized by rejection of unit root and non-rejection of stationarity test) and regions that *are in the process of converging* (rejection by both tests[7]). However, since both levels and first differences are of interest, the order of testing is also important: do we start the testing procedures with levels or first differences?

As regards convergence tests, Dickey and Pantula (1987), argue that it is best to test for a unit root in first differences and if this is rejected, to move on to test for a unit root in the levels.[8] On the other hand, stationarity of the levels implies that the spectrum of first differences is zero at the origin, thereby invalidating a (nonparametric) stationarity test on first differences. This suggests that the sequence of stability tests should be one in which the stationarity of Δy_t is tested only if stationarity of y_t has been rejected; see also Choi and Yu (1997).

Taking those arguments into account we end up with the strategy described in the chart in figure 2, with five possible outcomes A,B,C,D,E. The starting point is the unit root test on inflation differentials. If this doesn't reject we have the case of non-convergence (E), while a rejection will lead to testing the unit root hypothesis in relative prices. The result of the latter test will lead to a stationarity test in either levels or first differences. The final outcomes are as follows.

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[7] As shown in Muller (2005), a stationarity test will tend to reject the null hypothesis for highly persistent time series. In other words, it is difficult to control the size of stationarity tests in the presence of strong autocorrelation; see also KPSS.

[8] The results in Pantula (1989) indicate that the test of a unit root in inflation will tend to reject if the price level is stationary.

(A) Relative prices are converging: rejection of unit root in first differences and levels, rejection of levels stationarity test.

(B) Relative prices have converged: rejection of unit root in first differences and levels, non rejection of levels stationarity test.

(C) Inflation rates are converging: rejection of unit root in first differences but not in levels, rejection of first differences stationarity test.

(D) Inflation rates have converged: rejection of unit root in first differences but not in levels, non rejection of first differences stationarity test.

(E) Non convergence: non rejection of unit root in first differences.

The price and inflation contrasts between Florence and Aosta provide again an illustration. The null hypothesis of non convergence is rejected at 1% level by the ADF test on inflation differentials: the modified AIC lag selection criterion of Ng and Perron (2001) suggests 19 lags and resulting τ_0 statistic[9] is -3.21. The unit root in levels is also rejected, as was seen in sub-section 2.2, and a rejection also occurs for the level stationarity tests. Thus, the sequential testing procedure of figure 2, leads to the conclusion that relative prices are converging between Florence and Aosta, that is, case A. Further details are given in the row labelled AO-FI of table 4. One aspect of these results that might cause concern is the fact that, although prices seem to be converging, the stationarity test on inflation differentials rejects the null hypothesis. The next sub-section explains why this happens.

3.2 *First differences stationarity tests for highly persistent process in levels*

The properties of first differences stationarity tests when the DGP is a highly persistent process in the levels depend on whether the initial condition is small or large. In the former case the test is undersized, in the latter it is oversized with the degree of oversizing increasing with the magnitude of the initial condition. We present a small Monte-Carlo simulation experiment that illustrates the point. A theoretical analysis of this and related issues is beyond the scope of this paper.

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[9] Including fewer lags would imply even stronger evidence against the null.

We consider the AR(1) data generating process, $t = 1, 2, ..., T$,

$$y_t = (1 - c/T)y_{t-1} + \eta_t, \qquad \eta_t \sim NIID(0, \sigma^2_\eta) \tag{12}$$

for some given initial condition y_0. Thus, as in section 2.3, y_t is a highly persistent process for $c > 0$ and a unit root process for $c = 0$. Notice that a relatively small c and a large initial condition are associated with y_t converging to its long run value of zero.[10]

The validity of stationarity tests in first differences requires that $c = 0$ in (12). If this is not the case then the properties of the test depend on the magnitude of the initial condition y_0 relatively to the standard deviation of η_t. In particular, the test is undersized if y_0 is small and (often dramatically) oversized if y_0 is large. We take $\sigma^2_\eta = 1$, $c = 0, 1, 2.5, 5, 10$ and $y_0 = 0, 5, 10, 15, 20, 25, 30, 50$. Table 3a,b reports rejection frequencies for the stationarity tests ξ_0, ξ_1 computed on the first differenced data Δy_t, for $T = 100$ and 400, where the bandwidth parameter for spectral estimation is equal to $int(m(T/100)^{0.25})$ and $m = 0, 4, 8$.

For $c = 0$ the stationarity tests in first differences have (approximately) the correct size, while they are undersized when $c > 0$ and the initial condition is small. Oversizing occurs for a large initial condition, at least as large as 15 when $T = 100$ and 25 when $T = 400$. Notice that oversizing can be huge, with the probability of rejecting the null equal or close to 1 in many cases.

Intuitively, this oversizing problem can be explained if we think of a converging path in levels (starting from a large initial value): the first difference is the slope of the series which keeps changing mostly in the same direction in order to bring the level to its long run value. Notice that these large values of the initial conditions, for which oversizing occurs, are quite typical for converging series, as can be seen in the Florence-Aosta example and in other empirical results of next section.

4. Convergence properties of the CPI among Italian regions

In this section we provide evidence on the nature and features of inflation and relative price differentials between Italian regions. The data used are the monthly Istat series of the

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[10] Clearly, as in section 2.3, we could also specify a model that for $c > 0$ would converge to a nonzero long run equilibrium α. The simulation results will be unchanged as long as we interpret the initial conditions as deviations from α.

Consumers' Price Index in nineteen "regional capitals" for the period 1970M1-2003M12. Due to the presence of large outliers, Potenza was excluded from the analysis. The cities included are Ancona (AN), Aosta (AO), L'Aquila (AQ), Bari (BA), Bologna (BO), Cagliari (CA), Campobasso (CB), Firenze (FI), Genova (GE), Milano (MI), Napoli (NA), Palermo (PA), Perugia (PG), ReggioCalabria (RC), Trento (TN), Torino (TO), Trieste (TS), Venezia (VE), Roma (RM). As the original series refer to different base years, they have been rebased, taking 2003 as the base year. They have also been seasonally adjusted by removing a stochastic seasonal component using the STAMP package of Koopman *et al.* (2000).

Figure 3 shows the time pattern of the log of relative price levels, computed as the difference between each (log) regional price index and the average national one. As we have set 2003 as the base year the contrasts are constrained by construction to tend to zero near the end of the sample period. The picture seems consistent with high persistence in price differentials, either a unit root or a converging process. The dynamics of the cross-sectional standard deviation of regional inflation rates is depicted in figure 4. Despite the high variability of the data due to the monthly frequency of observation, we observe an overall reduction in the geographical dispersion of inflation since the beginning of the eighties. This reduction is partly correlated with a decrease in average inflation; see Caruso *et al.* (1993). [11]

The results of the battery of convergence and stability tests on inflation and price differentials on the 171 regional contrasts are reported in table 4. For inflation contrasts we report significance levels of rejections for the ADF test and the stationarity test, τ_0 and ξ_0 respectively (both computed without fitting a mean), and the number of lags in the ADF regression chosen according to the modified Akaike information criterion of Ng and Perron (2001). For price contrasts we report significance levels of rejections for the ADF test with a constant term, τ_1, the modified ADF test, τ^* (where the data are transformed by subtracting the average of the observations in the final year), and the KPSS stationarity test[12], ξ_1. We also report the number of lags in the ADF regression and the magnitude of the initial condition

[11] Changes in the variance of the series are likely to affect, to some extent, the properties of the statistical tests of convergence. In particular, the results of Kim et al. (2002) and Busetti and Taylor (2003b) would predict some degree of oversizing for both stationarity and unit root tests in the presence of a variance decrease.

[12] For computing the stationarity tests both in level and first differences the data have been additionally prefiltered by the seasonal sum operator in order to guard against "unattended" unit root and structural breaks at the seasonal frequencies; see Busetti and Taylor (2003a). The reported results refer to a bandwidth parameter $m = 15$ in the nonparametric long-run variance estimator. The conclusions however are quite robust for a wide range of values of m.

in units of residuals standard deviation. The last column of the table contains the summary results, coded A to E according to the framework described in figure 2.

In all cases the unit root test on inflation differentials easily rejects the null hypothesis, thus excluding case E of non-convergence. Out of 171 regional contrasts we obtained 89 cases of D, stability in the inflation rates, 41 C's, converging inflation, and 41 A's, converging relative prices. Among the largest cities, it turns out that inflation rates have been stable between Milano, Napoli and Torino, while relative prices are converging between Roma and Milano and Roma and Napoli. In six cases (namely AN-RM, AO-TS, CA-PG, PG-RC, RC-TN and TN-RM) we obtained the somewhat contradictory result that, for relative prices, both unit root and stationarity tests are unable to reject. These cases have been labelled as D, stable inflation rates, because of the non rejection of τ^*. However, given the low power of DF tests for small initial conditions - which is the case for all six pairs here - there is a strong argument for following the stationarity test and labelling them as B, stability of relative prices.

It is also interesting to observe that, as predicted by the simulation results of table 3b, there are many cases (denoted in italics in the column reporting the results of the stationarity tests for inflation contrasts) where a simultaneous rejection of the unit root and the stationarity hypothesis in the levels with a large initial condition is accompanied by a rejection of the stationarity test in first differences. This simply reflects the bias in the stationarity test for highly persistent processes, as described in section 3.2. Notice also that in most cases where the initial condition is large the τ^* test provides much stronger evidence against the null than does τ_1, as predicted by the power study reported in tables 2a and 2b.[13]

If the failure to reject the null hypothesis of a unit root in relative price levels is put down to the low power of unit root tests, it is worth considering the possibility of exploiting the higher power of a multivariate test. We therefore applied the MHDF test, both with and without constant, on all the regional constrasts computed with respect to Rome[14]. It turns out that $\tau^*(18)$ is less than -6.81 (the 10% critical value taken from table 1) for nearly all lag structures in the ADF regression, thus providing stronger evidence for convergence of relative

[13] To guard against possible biases induced by variance shifts in the data, the same empirical analisys has been carried out also for the shorter subsample 1985.1-2003.12. It has been found that overall the results do not change much, although there are cases where the final outcome of the tests (A,B,C,D) is switched among pairs of regions. Full details are available from the authors, on request.

[14] The properties of the test and its results are invariant to the region chosen as a benchmark.

prices. On the other hand, the MHDF test with constant, $\tau_1(18)$, never rejects the null even at the 10% significance level[15], confirming the loss in power from fitting a constant, even in the multivariate case.

5. Concluding remarks

In examining the behaviour of relative price time series between different regions it is important to distinguish between stability and convergence. Stability is assessed by stationarity tests, while convergence is determined by unit root tests. For pairwise contrasts of inflation rates, these tests are best carried out without removing a constant term. As an alternative to the stationarity test, a ‘t-test’ on the sample mean may be used. For price index contrasts, running a Dickey-Fuller unit root test with the base year at the end leads to power gains in testing for relative convergence. (We derive the asymptotic distribution of this test statistic and report critical values). We set out a sequential testing strategy to establish whether convergence occurs in relative prices or just in rates of inflation. This strategy is applied to the monthly series of the Consumer Price Index in the Italian regional capitals over the period 1970-2003. It is found that all 171 pairwise contrasts of inflation rates have converged or are in the process of converging. Only 24% of price level contrasts appear to be converging, but a multivariate test provides strong evidence of overall convergence.

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[15] The critical value with 18 degrees of freedom is -6.43, obtained by interpolation from Harvey and Bates (2003).

Appendix

Distribution of the DF statistic τ^* constructed from data with the last observation subtracted

Let $y_t^* = y_t - y_T,\ t = 1, ..., T-1$. Under the null hypothesis that $\phi = 1$ in (5) it follows from the standard argument used to derive the distribution of τ_0 - for example, Hamilton (1994, p476-7) - that

$$\sum_{t=2}^{T} y_{t-1}^* \eta_t = \frac{1}{2}(y_T^{*2} - y_1^{*2} - \sum_{t=2}^{T} \eta_t^2).$$

Now $y_t^* = -\sum_{t+1}^{T} \eta_j,\ t = 1, ..., T-1$ and so, since $y_T^* = 0$,

$$\frac{1}{\sigma^2 T} \sum_{t=2}^{T} y_{t-1}^* \eta_t \xrightarrow{d} \frac{1}{2}(-W(1)^2 - 1).$$

The distribution in (6) then follows by application of the continuous mapping theorem. If the test statistic is calculated by subtracting the first observation it is easy to see that the sign of $W(1)$ changes.

Derivation of the ERS-type statistic

Write down the likelihood for the observations from $t = p+1, ..., T+1$ and set $y_{T+1} = \alpha$ before differentiating with respect to α. This yields

$$\widehat{\alpha}_c^* = \frac{\sum_{j=1}^{p} \overline{\phi}_j \sum_{j=1}^{p} \overline{\phi}_j y_{T-j+1} + (1 - \sum_{j=1}^{p} \overline{\phi}_j) \sum_{t=p+1}^{T} (y_t - \sum_{j=1}^{p} \overline{\phi}_j y_{t-j})}{(\sum_{j=1}^{p} \overline{\phi}_j)^2 + (1 - \sum_{j=1}^{p} \overline{\phi}_j)^2 (T-p)}$$

with $\sum_{j=1}^{p} \overline{\phi}_j = 1 - c/T$. If c is set to zero, α is estimated from a weighted average of the last p observations, with the weights summing to one.

Multivariate tests

Let $\mathbf{y}_t$ be the $N = n-1$ vector of contrasts between each region and a benchmark. If the benchmark is the n-th region, then $\mathbf{y}_t = \left(y_t^{1,n},\ y_t^{2,n},\ \ldots,\ y_t^{n-1,n}\right)'$ where $y_t^{i,j} = \log p_{i,t} - \log p_{j,t}$. Multivariate stationarity tests applied to $\mathbf{y}_t$ can be used to test whether the series for the n regions are stable. In a simple multivariate random walk plus noise model, the Lagrange multiplier is easily constructed in a homogeneous model in which the covariance matrix of the random walk is proportional to the covariance matrix of the noise; see Nyblom and Harvey (2000). The more general statistic is now given by

$$\xi_0(N) = Trace\left(\widehat{\mathbf{\Omega}}^{-1}\mathbf{C}\right),$$

where $\mathbf{C} = \sum_{t=1}^{T}\left(\sum_{j=1}^{t}\mathbf{y}_j\right)\left(\sum_{j=1}^{t}\mathbf{y}_j\right)'$ and $\widehat{\mathbf{\Omega}}$ is a non-parametric estimator of the long run variance of $\mathbf{y}_t$, obtained by a straightforward multivariate extension of (2). Under the null hypothesis of zero mean stationarity, $\xi_0(N)$ converges in distribution to the sum of the integrals of the squares of independent Brownian motions; critical values are provided in Nyblom (1989) and Hobijn and Franses (2000). A multivariate Wald-type test on the mean of $\mathbf{y}_t$ can also be constructed by generalizing the nonparametric t- statistic to give $t_{\overline{y}}(N) = T\overline{\mathbf{y}}'\widehat{\mathbf{\Omega}}^{-1}\overline{\mathbf{y}}$. Under the null hypothesis of zero mean stationarity of $\mathbf{y}_t$, $t_{\overline{y}}(N) \xrightarrow{d} \chi^2(N)$.

The simplest multivariate convergence model is the zero-mean VAR(1) process

$$\mathbf{y}_t = \Phi\mathbf{y}_{t-1} + \boldsymbol{\eta}_t,$$

where $\mathbf{\Phi}$ is a $N \times N$ matrix and $\boldsymbol{\eta}_t$ is a N dimensional vector of martingale differences innovations with constant variance $\mathbf{\Sigma}_\eta$. The model is said to be *homogeneous* if $\mathbf{\Phi} = \phi\mathbf{I}_N$. Following Abuaf and Jorion (1990) and Flôres et al (1999), we use the *multivariate homogeneous Dickey-Fuller* (MHDF) statistic; this is given by the Wald statistic on $\rho = \phi - 1$, that is

$$\tau_0(N) = \frac{\sum_{t=2}^{T}\mathbf{y}_{t-1}'\widetilde{\mathbf{\Sigma}}_\eta^{-1}\Delta\mathbf{y}_t}{\left(\sum_{t=2}^{T}\mathbf{y}_{t-1}'\widetilde{\mathbf{\Sigma}}_\eta^{-1}\mathbf{y}_{t-1}\right)^{\frac{1}{2}}},$$

where $\widetilde{\mathbf{\Sigma}}_\eta$ is the ML estimator of $\mathbf{\Sigma}_\eta$. Critical values for the MHDF test are tabulated in Harvey and Bates (2003). One of the attractions of the MHDF test is that it is invariant to pre-multiplication of $\mathbf{y}_t$ by a nonsingular $N \times N$ matrix; in contrast, such invariance is lost if $\mathbf{\Phi}$ is assumed to be diagonal as in Taylor and Sarno (1998). Serial correlation in the innovations can be accounted for by the $VAR(p)$ process

$$\Delta\mathbf{y}_t = (\mathbf{\Phi} - \mathbf{I})\,\mathbf{y}_{t-1} + \mathbf{\Gamma}_1\Delta\mathbf{y}_{t-1} + ... + \mathbf{\Gamma}_{p-1}\Delta\mathbf{y}_{t-p+1} + \boldsymbol{\eta}_t,$$

written in error correction form. The analogue of the homogeneous model has $\mathbf{\Phi} = \phi\mathbf{I}_{n-1}$. In this case the test will be computed by the same statistic $\tau_0(N)$ where $\Delta\mathbf{y}_t$ and $\mathbf{y}_{t-1}$ are replaced by the OLS residuals from regressing each of them on $\Delta\mathbf{y}_{t-1}$, ..., $\Delta\mathbf{y}_{t-p+1}$. The same limiting distribution and critical values apply.

The distribution of the test statistic changes if it is constructed using the demeaned observations $\mathbf{y}_t - \overline{\mathbf{y}}$ in place of $\mathbf{y}_t$. As regards the multivariate τ^* test statistic, obtained by working with $\mathbf{y}_t - \mathbf{y}_T$, the asymptotic distribution under the null hypothesis is

$$\tau^*(N) \to \frac{-(1/2)\sum_{i=1}^{N}(W_i(1)^2+1)}{\left[\sum_{i=1}^{N}\int_0^1 W_i(r)^2 dr\right]^{1/2}} = \frac{-(1/2)(\chi_N^2+N)}{\left[\sum_{i=1}^{N}\int_0^1 W_i(r)^2 dr\right]^{1/2}}$$

where $W_i(r), i = 1, ..N$ are independent standard Wiener processes. The power of the $\tau^*(N)$ test relative to MHDF with mean subtracted, that is $\tau_1(N)$, will depend on the distribution of the initial conditions. Series with large initial conditions will tend to increase power.

Table 1. Limiting distribution of the MHDF test constructed after subtracting the last observation

Quantiles	N=1	N=2	N=3	N=4	N=5	N=6	N=7	N=8	N=9	N=10
0.01	-3.16	-3.50	-4.15	-4.45	-4.80	-5.15	-5.43	-5.69	-5.84	-5.92
0.05	-2.69	-3.27	-3.64	-4.03	-4.33	-4.59	-4.92	-5.11	-5.31	-5.59
0.10	-2.43	-3.03	-3.42	-3.75	-4.11	-4.34	-4.61	-4.85	-5.07	-5.33
0.90	-0.98	-1.51	-1.94	-2.28	-2.59	-2.87	-3.18	-3.42	-3.65	-3.84
0.95	-0.87	-1.38	-1.79	-2.15	-2.48	-2.76	-3.00	-3.20	-3.46	-3.70
0.99	-0.71	-1.11	-1.54	-1.90	-2.14	-2.43	-2.71	-2.95	-3.18	-3.40

Quantiles	N=11	N=12	N=13	N=14	N=15	N=16	N=17	N=18	N=19	N=20
0.01	-6.11	-6.33	-6.52	-6.79	-6.90	-7.11	-7.35	-7.50	-7.81	-7.96
0.05	-5.71	-5.92	-6.21	-6.39	-6.58	-6.77	-6.97	-7.07	-7.25	-7.41
0.10	-5.53	-5.72	-5.92	-6.15	-6.34	-6.52	-6.65	-6.81	-6.98	-7.20
0.90	-4.07	-4.28	-4.45	-4.63	-4.82	-5.00	-5.16	-5.36	-5.53	-5.68
0.95	-3.94	-4.15	-4.33	-4.49	-4.67	-4.85	-5.02	-5.22	-5.39	-5.49
0.99	-3.55	-3.81	-4.00	-4.13	-4.36	-4.51	-4.79	-4.91	-5.09	-5.20

N is the number of series.

Table 2a. Power comparison of convergence tests - T=100

T=100		Initial Condition							
		0	5	10	15	20	25	30	50
c=10	τ_1	0,28	0,37	0,64	0,92	1,00	1,00	1,00	1,00
	τ^*	0,15	0,23	0,55	0,89	0,98	1,00	1,00	1,00
c=5	τ_1	0,10	0,12	0,19	0,35	0,59	0,82	0,95	1,00
	τ^*	0,05	0,08	0,20	0,54	0,90	0,99	1,00	1,00
c=2.5	τ_1	0,06	0,06	0,07	0,09	0,13	0,19	0,27	0,74
	τ^*	0,03	0,04	0,08	0,19	0,44	0,76	0,95	1,00
c=1	τ_1	0,05	0,05	0,05	0,05	0,04	0,04	0,04	0,06
	τ^*	0,03	0,04	0,05	0,07	0,11	0,18	0,28	0,85
c=0	τ_1	0,04	0,04	0,04	0,04	0,04	0,04	0,04	0,04
	τ^*	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05

Table 2b. Power comparison of convergence tests - T=400

T=400		Initial Condition							
		0	5	10	15	20	25	30	50
c=10	τ_1	0,30	0,32	0,40	0,54	0,72	0,88	0,97	1,00
	τ^*	0,16	0,18	0,26	0,42	0,65	0,85	0,94	1,00
c=5	τ_1	0,11	0,12	0,13	0,16	0,21	0,29	0,39	0,87
	τ^*	0,06	0,07	0,09	0,13	0,22	0,39	0,61	0,99
c=2.5	τ_1	0,07	0,07	0,07	0,08	0,08	0,09	0,10	0,21
	τ^*	0,04	0,04	0,05	0,06	0,09	0,13	0,20	0,79
c=1	τ_1	0,06	0,06	0,06	0,06	0,05	0,05	0,05	0,05
	τ^*	0,04	0,04	0,04	0,04	0,05	0,06	0,07	0,18
c=0	τ_1	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
	τ^*	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05

Rejection frequencies of the DF test with constant, $\tau 1$, and the the test without constant τ^*.
The data generating process is
$y(t)=\alpha+u(t)$
$u(t)=(1-c/T)u(t-1)+e(t)$
$u(0)$ given initial condition
$e(t)$ NIID(0,1)

Table 3a. Rejection of first differences stationarity tests for highly persistent levels - T=100

T=100			Initial Condition							
		m	0	5	10	15	20	25	30	50
c=10										
		0	0,00	0,00	0,00	0,30	0,97	1,00	1,00	1,00
	ξ_0	4	0,00	0,00	0,01	0,26	0,83	0,99	1,00	1,00
		8	0,00	0,00	0,02	0,24	0,64	0,91	0,98	1,00
		0	0,00	0,00	0,03	0,28	0,80	0,99	1,00	1,00
	ξ_1	4	0,00	0,00	0,03	0,23	0,62	0,90	0,99	1,00
		8	0,00	0,00	0,04	0,18	0,43	0,70	0,87	1,00
c=5										
		0	0,00	0,00	0,01	0,36	0,93	1,00	1,00	1,00
	ξ_0	4	0,00	0,00	0,03	0,37	0,88	0,99	1,00	1,00
		8	0,00	0,00	0,05	0,39	0,82	0,98	1,00	1,00
		0	0,00	0,01	0,05	0,20	0,46	0,76	0,93	1,00
	ξ_1	4	0,00	0,01	0,05	0,18	0,39	0,67	0,87	1,00
		8	0,00	0,01	0,05	0,15	0,33	0,56	0,77	1,00
c=2.5										
		0	0,00	0,00	0,03	0,22	0,61	0,91	0,99	1,00
	ξ_0	4	0,00	0,00	0,04	0,25	0,63	0,90	0,99	1,00
		8	0,00	0,01	0,06	0,29	0,64	0,90	0,98	1,00
		0	0,02	0,02	0,04	0,08	0,15	0,25	0,36	0,84
	ξ_1	4	0,01	0,02	0,04	0,07	0,13	0,22	0,31	0,79
		8	0,01	0,02	0,03	0,07	0,12	0,18	0,27	0,73
c=1										
		0	0,01	0,01	0,03	0,09	0,18	0,33	0,49	0,96
	ξ_0	4	0,01	0,02	0,04	0,10	0,21	0,35	0,52	0,96
		8	0,01	0,02	0,05	0,12	0,24	0,38	0,55	0,96
		0	0,04	0,04	0,04	0,04	0,05	0,05	0,06	0,11
	ξ_1	4	0,03	0,03	0,03	0,04	0,04	0,05	0,05	0,10
		8	0,03	0,03	0,03	0,03	0,04	0,04	0,05	0,08
c=0										
		0	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
	ξ_0	4	0,07	0,07	0,07	0,07	0,07	0,07	0,07	0,07
		8	0,08	0,08	0,08	0,08	0,08	0,08	0,08	0,08
		0	0,04	0,04	0,04	0,04	0,04	0,04	0,04	0,04
	ξ_1	4	0,04	0,04	0,04	0,04	0,04	0,04	0,04	0,04
		8	0,03	0,03	0,03	0,03	0,03	0,03	0,03	0,03

ξ_0 is stationarity test without constant, with bandwidth equal to int(m(T/100)^.25)
$\xi 1$ is stationarity test, with constant with bandwidth equal to int(m(T/100)^.25)
The initial condition is in units of the error standard deviation

Table 3b. Rejection of first differences stationarity tests for highly persistent levels - T=400

T=400			Initial Condition							
		m	0	5	10	15	20	25	30	50
c=10										
		0	0,00	0,00	0,00	0,00	0,00	0,08	0,71	1,00
	ξ_0	4	0,00	0,00	0,00	0,00	0,00	0,10	0,63	1,00
		8	0,00	0,00	0,00	0,00	0,01	0,13	0,56	1,00
		0	0,00	0,00	0,00	0,01	0,04	0,19	0,50	1,00
	ξ_1	4	0,00	0,00	0,00	0,01	0,05	0,19	0,46	1,00
		8	0,00	0,00	0,00	0,01	0,05	0,18	0,42	1,00
c=5										
		0	0,00	0,00	0,00	0,00	0,01	0,10	0,45	1,00
	ξ_0	4	0,00	0,00	0,00	0,00	0,01	0,12	0,45	1,00
		8	0,00	0,00	0,00	0,00	0,02	0,14	0,45	1,00
		0	0,00	0,00	0,01	0,02	0,06	0,12	0,23	0,84
	ξ_1	4	0,00	0,00	0,01	0,02	0,06	0,12	0,23	0,82
		8	0,00	0,00	0,01	0,02	0,06	0,12	0,22	0,79
c=2.5										
		0	0,00	0,00	0,00	0,01	0,02	0,08	0,21	0,93
	ξ_0	4	0,00	0,00	0,00	0,01	0,03	0,09	0,22	0,93
		8	0,00	0,00	0,00	0,01	0,03	0,10	0,23	0,93
		0	0,02	0,02	0,02	0,03	0,04	0,06	0,09	0,27
	ξ_1	4	0,02	0,02	0,02	0,03	0,04	0,06	0,09	0,26
		8	0,02	0,02	0,02	0,03	0,04	0,06	0,09	0,25
c=1										
		0	0,00	0,00	0,01	0,02	0,03	0,05	0,07	0,30
	ξ_0	4	0,00	0,01	0,01	0,02	0,03	0,05	0,08	0,31
		8	0,01	0,01	0,01	0,02	0,03	0,05	0,09	0,33
		0	0,04	0,04	0,04	0,04	0,04	0,04	0,05	0,06
	ξ_1	4	0,04	0,04	0,04	0,04	0,04	0,04	0,04	0,06
		8	0,04	0,04	0,04	0,04	0,04	0,04	0,04	0,05
c=0										
		0	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
	ξ_0	4	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
		8	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
		0	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
	ξ_1	4	0,05	0,05	0,05	0,05	0,05	0,05	0,05	0,05
		8	0,04	0,04	0,04	0,04	0,04	0,04	0,04	0,04

ξ_0 is stationarity test without constant, with bandwidth equal to int(m(T/100)^.25)

Table 4. Results of the tests on the CPI in the Italian regional capitals

Cities	Inflation contrasts			Price contrasts					Summary			
	τ_0	n.lags	ξ_0	τ_1	τ^*	init. cond.	n.lags	ξ_1	A	B	C	D
AN-AO	1%	20	*1%*	1%	1%	31.6	2	1%	*			
AN-AQ	1%	20				-0.2	9	1%				*
AN-BA	1%	3	*10%*	1%	1%	24.6	1	1%	*			
AN-BO	1%	24	*1%*	10%	5%	29.5	1	1%	*			
AN-CA	1%	10				1.2	2	1%				*
AN-CB	1%	23				-10.9	11	1%				*
AN-FI	1%	13				11.9	9	1%				*
AN-GE	1%	1	*1%*	10%	5%	21.4	9	1%	*			
AN-MI	1%	2	*1%*	1%	1%	28.6	4	1%	*			
AN-NA	1%	1		1%	1%	21.0	1	1%	*			
AN-PA	1%	1				-8.5	4	1%				*
AN-PG	1%	19				-0.4	9	1%				*
AN-RC	1%	1				2.2	3	5%				*
AN-TN	1%	16				4.0	3	5%				*
AN-TO	1%	2	*5%*	10%	5%	23.7	2	1%	*			
AN-TS	1%	24	*1%*	1%	1%	44.1	1	1%	*			
AN-VE	1%	20	*1%*	5%	1%	44.3	6	1%	*			
AN-RM	1%	20				4.2	1					*
AO-AQ	1%	23	*1%*		10%	-34.3	2	1%	*			
AO-BA	1%	3	5%			-12.9	19	1%			*	
AO-BO	1%	7		10%		-7.6	2	1%				*
AO-CA	1%	24	*1%*		10%	-35.0	5	1%	*			
AO-CB	1%	24	5%			-40.1	22	1%			*	
AO-FI	1%	19	*5%*		5%	-26.4	2	1%	*			
AO-GE	1%	3		5%	5%	-18.0	2	1%	*			
AO-MI	1%	3	*10%*	10%	10%	-13.5	2	1%	*			
AO-NA	1%	1	5%			-12.4	5	1%			*	
AO-PA	1%	24	*1%*		10%	-39.7	2	1%	*			
AO-PG	1%	23	*1%*		5%	-34.4	2	1%	*			
AO-RC	1%	23	*1%*			-31.5	1	1%			*	
AO-TN	1%	3	*5%*		10%	-32.9	6	1%	*			
AO-TO	1%	23	10%			-15.2	2	1%			*	
AO-TS	1%	3				9.6	2					*
AO-VE	1%	1				5.2	3	5%				*
AO-RM	5%	24	*1%*	10%	10%	-35.0	22	1%	*			
AQ-BA	1%	1	10%			25.3	10	1%			*	
AQ-BO	1%	17	10%			29.8	1	1%			*	
AQ-CA	1%	1				1.6	1	5%				*
AQ-CB	1%	1				-10.7	3	1%				*
AQ-FI	1%	18				13.1	1	1%				*
AQ-GE	1%	21	5%			21.7	1	1%			*	
AQ-MI	1%	17	10%			27.6	1	1%			*	
AQ-NA	1%	1				21.0	1	1%				*
AQ-PA	1%	1				-9.3	24	1%				*
AQ-PG	1%	4		5%	5%	-0.2	2	5%	*			
AQ-RC	1%	1				2.6	3	5%				*
AQ-TN	1%	3		10%		4.4	5	1%				*
AQ-TO	1%	12				25.0	4	1%				*
AQ-TS	1%	8	*1%*		5%	47.2	1	1%	*			
AQ-VE	1%	23	1%			43.1	1	1%			*	
AQ-RM	1%	21				4.3	1	5%				*
BA-BO	1%	14				5.4	5	1%				*
BA-CA	1%	11				-24.3	1	1%				*
BA-CB	1%	23				-35.1	2	1%				*
BA-FI	1%	13			10%	-14.9	1	5%	*			
BA-GE	1%	1		10%	10%	-5.4	2	5%	*			
BA-MI	1%	3				-0.2	1	1%				*

BA-NA	1%	1				-2.1	1	5%			*
BA-PA	1%	23	*1%*		10%	-33.4	5	1%	*		
BA-PG	1%	2	10%			-26.5	2	1%		*	
BA-RC	1%	5				-22.5	1	1%			*
BA-TN	1%	3				-23.6	6	1%			*
BA-TO	1%	2		10%		-2.3	2				*
BA-TS	1%	1	5%			24.4	9	1%		*	
BA-VE	1%	1				18.9	1	1%			*
BA-RM	1%	13		10%	5%	-24.9	2	1%	*		
BO-CA	1%	24	10%			-29.9	1	1%		*	
BO-CB	1%	10	10%			-35.8	11	1%		*	
BO-FI	1%	12				-22.4	1	1%			*
BO-GE	1%	5				-12.2	7	1%			*
BO-MI	1%	24				-7.0	1	5%			*
BO-NA	1%	14				-7.2	5	1%			*
BO-PA	1%	24	5%			-33.5	5	1%		*	
BO-PG	1%	24	5%			-32.5	3	1%		*	
BO-RC	1%	23				-26.9	1	1%			*
BO-TN	1%	1	5%			-30.6	1	1%		*	
BO-TO	1%	3				-8.6	2	1%			*
BO-TS	1%	1		1%	1%	20.7	1	1%	*		
BO-VE	1%	12				15.4	2	5%			*
BO-RM	1%	24	5%			-32.7	20	1%		*	
CA-CB	1%	1				-12.0	1	1%			*
CA-FI	1%	12				12.2	1	1%			*
CA-GE	1%	24	5%			20.4	10	1%		*	
CA-MI	1%	23	10%			27.4	12	1%		*	
CA-NA	1%	1				23.4	16	1%			*
CA-PA	1%	1				-10.7	2	1%			*
CA-PG	1%	2				-1.8	19				*
CA-RC	1%	1				1.3	20	5%			*
CA-TN	1%	3				3.2	1	5%			*
CA-TO	1%	24				26.9	6	1%			*
CA-TS	1%	13	*1%*		5%	44.7	2	1%	*		
CA-VE	5%	22	1%			44.6	23	1%		*	
CA-RM	1%	24				3.0	10	5%			*
CB-FI	1%	24				23.4	4	1%			*
CB-GE	1%	14	10%			28.4	1	1%		*	
CB-MI	1%	24	10%			36.2	9	1%		*	
CB-NA	1%	1				29.3	3	1%			*
CB-PA	1%	1			10%	2.5	6	10%	*		
CB-PG	1%	1				10.7	10	1%			*
CB-RC	1%	8				12.5	1	1%			*
CB-TN	1%	1				15.4	4	1%			*
CB-TO	1%	13				31.9	1	1%			*
CB-TS	1%	24	*1%*		10%	50.9	2	1%	*		
CB-VE	1%	23	5%			44.5	1	1%		*	
CB-RM	1%	15				16.0	9	1%			*
FI-GE	1%	14				10.6	7	1%			*
FI-MI	1%	15				17.2	9	1%			*
FI-NA	1%	14				11.6	3	10%			*
FI-PA	1%	24				-22.4	6	1%			*
FI-PG	1%	2				-14.5	2	1%			*
FI-RC	1%	7				-10.6	20	1%			*
FI-TN	1%	17				-10.5	5	1%			*
FI-TO	1%	13				14.8	1	1%			*
FI-TS	1%	15	*1%*	10%	1%	41.1	1	1%	*		
FI-VE	1%	13	10%			38.6	3	1%		*	
FI-RM	1%	19				-11.4	13	1%			*
GE-MI	1%	1				6.7	2	5%			*
GE-NA	1%	1				3.1	1	1%			*
GE-PA	1%	1	1%			-28.4	1	1%		*	

GE-PG	1%	1	5%			-22.4	1	1%		*	
GE-RC	1%	1	10%			-17.6	3	1%		*	
GE-TN	1%	3				-18.9	5	1%			*
GE-TO	1%	1				4.3	3	5%			*
GE-TS	1%	1		10%	5%	32.1	11	1%	*		
GE-VE	5%	24				27.7	1	1%			*
GE-RM	1%	24	5%			-20.7	1	1%		*	
MI-NA	1%	1				-2.0	4	1%			*
MI-PA	1%	23	1%			-35.4	15	1%		*	
MI-PG	1%	24	5%			-28.0	2	1%		*	
MI-RC	1%	4	10%			-23.0	4	1%		*	
MI-TN	1%	17				-26.6	6	1%			*
MI-TO	1%	15				-2.5	3	1%			*
MI-TS	1%	1	*10%*	10%	5%	28.3	13	1%	*		
MI-VE	1%	1				22.0	11	1%			*
MI-RM	1%	24	*5%*		10%	-28.3	2	1%	*		
NA-PA	1%	1	5%			-29.2	3	1%		*	
NA-PG	1%	24				-22.9	1	1%			*
NA-RC	1%	1				-18.6	1	1%			*
NA-TN	1%	14				-19.6	5	1%			*
NA-TO	1%	10				0.2	5	1%			*
NA-TS	1%	1	1%			24.5	1	1%		*	
NA-VE	1%	14				20.1	11	1%			*
NA-RM	1%	1		5%	1%	-21.8	1	1%	*		
PA-PG	1%	23				8.9	13	1%			*
PA-RC	1%	1				11.1	11	1%			*
PA-TN	1%	3				13.3	5	1%			*
PA-TO	1%	23	5%			31.0	3	1%		*	
PA-TS	1%	24	*1%*		5%	50.6	7	1%	*		
PA-VE	1%	23	1%			45.8	1	1%		*	
PA-RM	1%	24				13.7	6	1%			*
PG-RC	1%	1				2.8	1				*
PG-TN	1%	1		10%		5.0	8	5%			*
PG-TO	1%	2	10%			25.1	2	1%		*	
PG-TS	1%	24	*1%*		5%	45.9	1	1%	*		
PG-VE	1%	24	*1%*		10%	44.2	2	1%	*		
PG-RM	1%	18				4.5	2	5%			*
RC-TN	1%	6				1.6	4				*
RC-TO	1%	23				21.8	1	1%			*
RC-TS	1%	4	*1%*		5%	43.3	1	1%	*		
RC-VE	1%	22	5%			40.9	9	1%		*	
RC-RM	1%	1				1.6	19	10%			*
TN-TO	1%	3				23.3	5	1%			*
TN-TS	1%	1	*1%*	10%	5%	44.6	1	1%	*		
TN-VE	1%	12	1%			41.4	1	1%		*	
TN-RM	1%	16				-0.3	12				*
TO-TS	1%	1	*5%*	5%	1%	28.6	1	1%	*		
TO-VE	1%	23	5%			24.0	2	1%		*	
TO-RM	1%	19	10%			-26.5	1	1%		*	
TS-VE	1%	1				-6.0	1	5%			*
TS-RM	1%	24	*1%*	5%	1%	-46.4	2	1%	*		
VE-RM	5%	22	*1%*		10%	-45.5	13	1%	*		

The figures in italics in the columns reporting stationarity test on inflation contrasts correspond to the case of rejection of unit root in the levels.

Figure 1 – Relative prices and inflation rates in Florence and Aosta

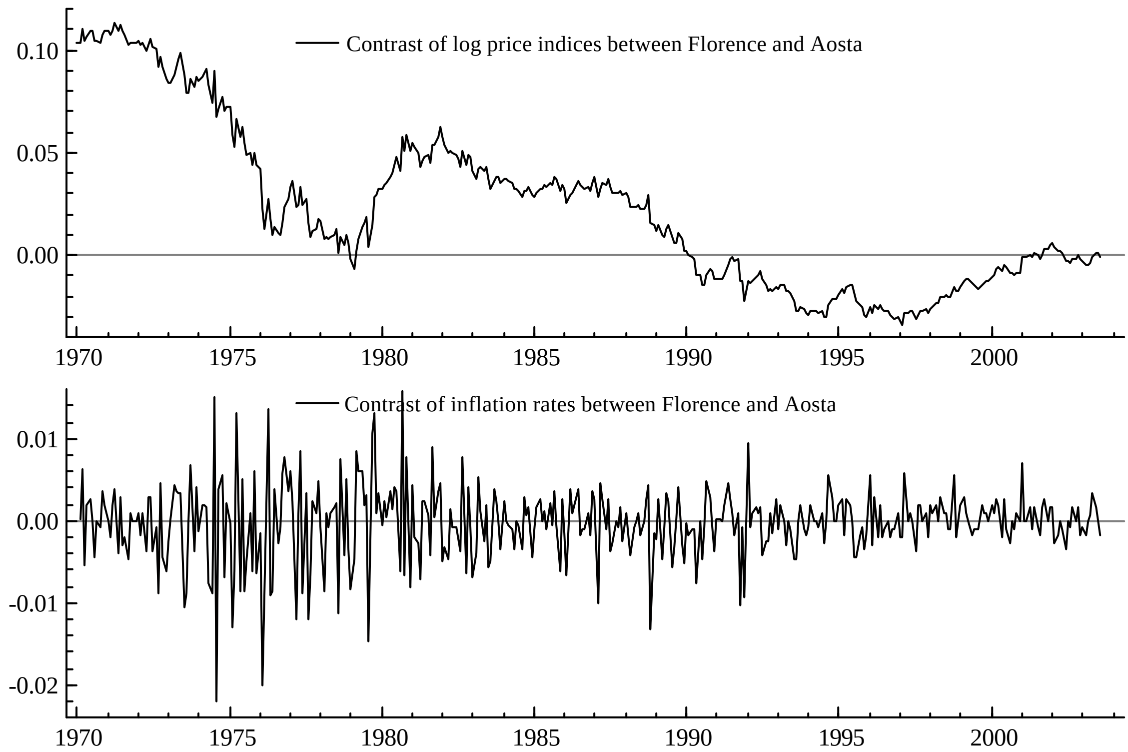


Figure 2 – Testing convergence in levels and first differences

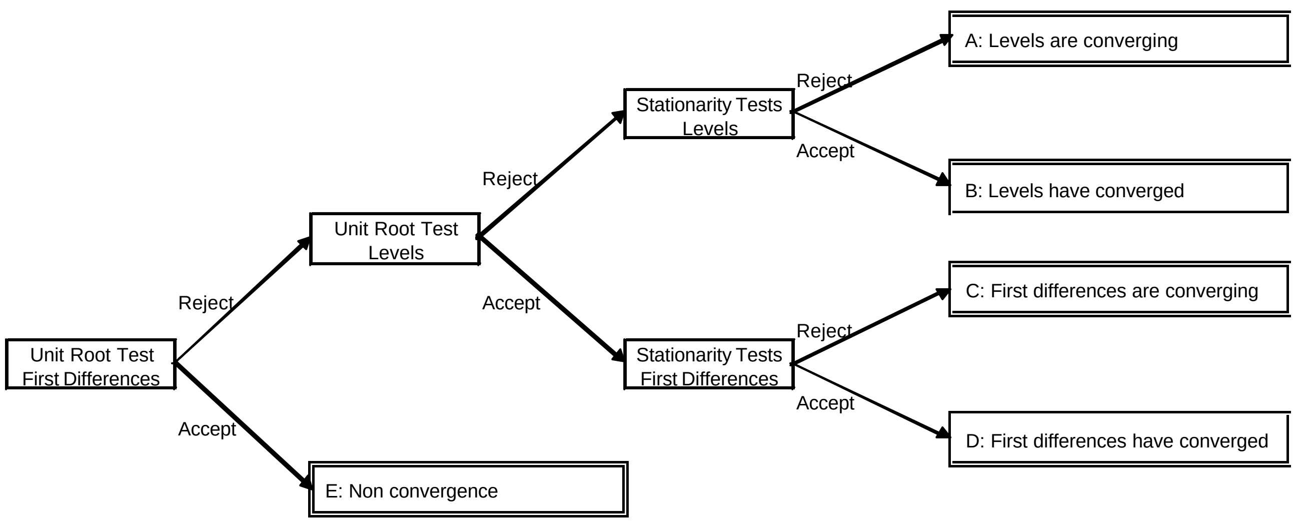


Figure 3 – Regional relative prices, base year=2003

(computed as differences with respect to the Italian average cost of living index)

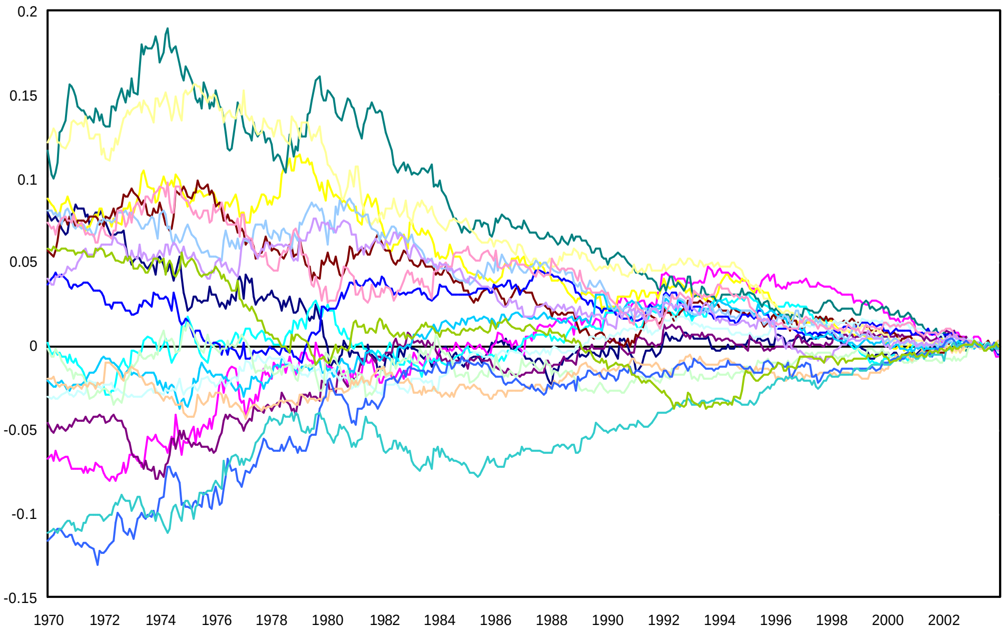
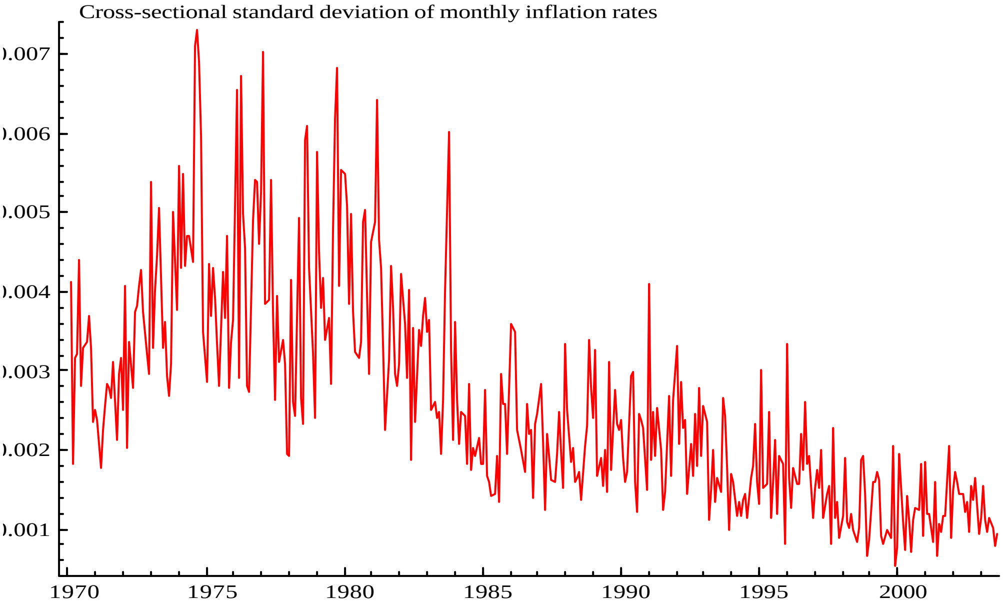


Figure 4 – Dispersion across regional inflation rates



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