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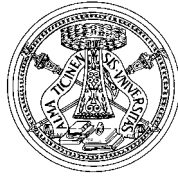
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## **Quaderni di Dipartimento**

### **A Cross-sectional Approach to Regional Long-run Inflation in Italy**

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# A Cross-sectional Approach to Regional Long-run Inflation in Italy

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# A Cross-sectional Approach to Regional Long-run Inflation in Italy

## Abstract

Considering a sample of about 70 Italian regions, this paper goes beyond the assumption that there exists a unique core inflationary process in a macroeconomy. We show that local long-run inflation rates can display remarkable variability. On the one hand they are negatively correlated with productivity growth, on the other the less competitive is the local retail sector and the higher is long-run inflation.

**Keywords:** purchasing power parity, long-run inflation, Balassa-Samuelson model, Kaldor-Verdoorn model.

**JEL codes:** R1, E31, F49.

## ***Introduction***

Testing for the purchasing power parity (PPP) hypothesis has been a classical research topic in economics (for a review see Rogoff, 1996). There exist two versions of the PPP hypothesis: the absolute and the relative ones. The former asserts that the real exchange rate is a constant or otherwise that the nominal exchange rate accommodates relative changes in the level of prices. “Relative PPP requires only that the rate of growth in the exchange rate offsets the differential between the rate of growth in home and foreign price indices” (Rogoff, 1996).

This paper focuses on the relative purchasing power parity hypothesis by exploiting a regional dataset, that is a dataset of small open economies belonging to a monetary union. When many economies share the same currency they will have a fixed nominal exchange rate, implying that if relative PPP holds, local inflation rates will converge to the same value in the long-run.

This study departs from standard ways of testing for PPP or for Balassa-Samuelson effects. Building on the literature on inflation persistence (Lünnemann and Mathä, 2004 and Vaona and Ascari, 2007) we first specify an autoregressive process for inflation and we compute the local long-run level of inflation as its unconditional mean. We find that it can display remarkable variation at the regional level. Long-run inflation also appears to be negatively correlated with productivity growth both for the whole local economy and in the non-traded sector. Furthermore, the less competitive is the local retail sector and the higher is long-run inflation. Testing for endogeneity does not point to the existence of sizeable biases.

In comparison with the various methods proposed in the literature to isolate long-run inflation (Taillon, 1997 and Stock and Watson, 1998), our approach might seem naïve, as it corresponds to a constant trend in the level of prices during the period of observation. However, ignoring time variation in long-run inflation does not hamper our analysis. Indeed, some of the determinants of long-run inflation, such as the degree of competitiveness of the local retail sector, change very slowly across time. As a consequence temporal variation might not always offer help in identifying the factors

underlying long-run inflation. Furthermore, data on real variables at the local level are not usually produced with the same frequency as inflation data. Finally, Vaona and Ascari (2007) show that the inflation generating process does not display major structural breaks in the sample here considered, so a constant trend in prices does not appear to be a too stringent assumption for the data we considered. On the other hand, focusing on cross-sectional regional variation we can go beyond the assumption “that there is a unique core inflationary process in a macroeconomy – across all sectors and all regions” – an assumption that “might seem improbable” (Quah and Vahey, 1995).

Finally, by focusing on Italy, we can overcome the lack of data characterizing the whole of Europe, which currently hampers the analysis of the long-run determinants of inflation differentials as admitted by Altissimo et al. (2005)<sup>1</sup>.

### ***Literature survey***

The PPP literature was subject in the past to a number of methodological shifts. At first, time-series data were used. One of the most common exercise was to run an Augmented-Dickey-Fuller test within the following model for the real exchange rate ( $q_t$ ):

$$\Delta q_t = \mu + \alpha q_{t-1} + \sum_{i=1}^k c_i \Delta q_{t-i} + \varepsilon_t$$

where  $\Delta$  is the first difference operator,  $\mu$  is a constant,  $\alpha$  and  $c_i$  are coefficients,  $k$  is the number of lags considered and  $\varepsilon$  is the stochastic error. If  $\alpha$  is significantly less than zero, there will be evidence of mean reversion and therefore of PPP.

In order to dispel doubts regarding the performance of tests for unit roots in small samples, researchers started to analyse very long time-series, which, however, have their own shortcomings, being prone to structural breaks. One further way out to this problem was to use panel data, as panel unit root tests have a better small sample performance than their time-series counterparts. On this ground, Imbs et al. (2005) showed that convergence towards PPP can be quite fast, though there exists some variability in the estimates they produced. Panel data estimates have been recently criticized because they ignore cross-unit

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<sup>1</sup> At page 21.

cointegration relationships leading to an excessive rejection of the unit root hypothesis (Banerjee et al., 2005).

The possible presence of factors hampering the adjustment of relative prices or inflation rates spurred researchers to move to consider regional datasets. Among the main factors hampering relative price adjustment it is possible to list: a) tariff barriers; b) non-tariff barriers; c) nominal exchange rates failing to adjust to relative price-level shocks; d) market imperfections allowing firms to apply different price policies in different countries; e) costs in adjusting prices; f) transportation costs hampering arbitrage between different countries; g) the presence of non traded goods, for which arbitrage is impossible (Cecchetti et al., 2002).

Other explanations that have been offered by the literature to explain price (inflation) differentials are: i) a positive correlation between the level of income and the level of prices, implying that catching up regions or economies should experience positive inflation differentials; ii) macro-economic disequilibria, whereby it is not said that all the regions within a country experience the same demand pressures; iii) even in presence of the same demand pressures there might be different market rigidities, implying stronger or weaker inflationary bottlenecks (Alberola, 2000).

One of the major studies of price differences within countries is Cecchetti et al. (2002). They analysed a dataset of the price indexes of 19 major US cities from 1918 to 1995 finding that relative price adjustment has a half-life of 8.5 years. Three explanations for such a slow convergence were proposed: distance – on the account that the price differential between two cities is larger the farther the two cities are -, different adjustment costs for small and large deviations and non traded goods. Remarkably, they did not manage to find any statistical support for these three explanations. They also could not test if the real wage or productivity differentials could affect their results due to data constraints.

Parsley and Wei (1996) analysed a quarterly data set including 51 tradable and non-tradable goods and services for 48 cities from 1975 to 1992. They found that distance, proxying for arbitrage costs, does affect the size of price differences and its convergence rate, therefore the more two cities are distant the more price differentials are variable and

wide and the longer they take to converge. A similar role for distance was found by Engel and Rogers (1996). Besides the role of distance, Parsley and Wei (1996) highlighted that prices of tradable goods converge faster than non-tradable ones, in contrast with the results found by Cecchetti et al. (2002).

Weber and Beck (2006) analysed a panel of 77 European regions from 1991 to 2002 using monthly data and a similar model to Cecchetti et al. (2002) but for inflation instead of the price level. They find that: i) regional inflation rates do not display smooth sigma convergence; ii) they do display a lot of internal volatility – whereby regions with a high inflation ranking in the present may have a low one in the future; iii) there is a positive relationship between regional inflation dispersion and mean which can allow central banks to decrease the average inflation down to 1% without pushing a sizeable percentage of regions into deflation; iv) mean-reversion takes place at a slow pace, that is the inflation half-life can be rather long, ranging from 0.5 to 75.1 years for different sub-samples.

Busetti et al. (2006) used a dataset of 19 Italian cities at a monthly frequency from 1970 to 2003 and they find evidence of convergence in both the level of prices and inflation rates by using unit roots and stationarity tests. However, Vaona (2007a) merged the PPP and the Phillips Curve literatures, applying Dynamic Panel Data methods on a sample of eighty one Italian provinces from the year 1986 to the year 1998 with an annual frequency. Inflation appeared to be characterized by a low degree of persistence and by conditional  $\beta$ -convergence across provinces. Similarly to previous contributions<sup>2</sup>, macroeconomic factors, such as the unemployment rate, can explain deviations from PPP.

The finding of conditional beta convergence is important because it is conceptually similar to the finding of Papell and Prodan (2006), that the real exchange rate might not revert to a constant mean, but rather to a constant trend determined by productivity growth differentials according to the Balassa-Samuelson hypothesis<sup>3</sup>. From the theoretical point of view, Obstfeld (1993) offered a model in which real exchange rates have a trend.

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<sup>2</sup> See Rogoff (1996), p. 663, where it is discussed the hypothesis that government spending might have an effect on PPP.

<sup>3</sup> One of the most active researchers in the field is David Papell. See for instance the papers quoted in Banerjee et al. (2005).



Bahmani-Oskooee and Nasir (2005) reviewed the literature on the Balassa-Samuelson effect, distinguishing between three groups of studies: “The first group includes studies that have used cross-sectional data. This group has provided mixed results. The second group which mostly supports the hypothesis includes studies that use time-series data. Finally, a third group has recently emerged and includes studies that use panel data and provide strong support for the hypothesis”. On the other hand, contrary to the Balassa-Samuelson model, the Kaldorian tradition postulates that productivity growth will decrease long-run inflation by reducing firm’s cost inflation (Dixon and Thirlwall, 1975).

It is worth recalling that the present study produces results relevant also to another strand of literature, given that long-run inflation has attracted considerable attention among economists in recent years (among others King and Wolman, 1996; Ascari, 2004). In particular, Altissimo et al. (2005) built a theoretical model showing that, within a monetary union, regional variations in productivity in non-tradables can be the primary cause of inflation differentials, whereby a faster productivity growth leads to a decrease in long-run inflation.

## ***Econometric Analysis***

### **Estimating Long-run Inflation**

The analysis here proposed builds on the results obtained by Vaona and Ascari (2007), who analysed a dataset of 71 Italian provinces (NUTS3 regions<sup>4</sup>) between 1996Q1 and 2006Q3. They found, first, that AR(1) models with seasonal dummies can well approximate the stochastic generating processes of local inflation rates in Italy. Second, the estimated degree of inflation persistence is very low and immune of structural breaks in the period considered, implying that it is possible to rely on standard methods to have reliable coefficient estimates. Finally, inflation persistence is statistically different across provinces

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<sup>4</sup> NUTS is the French acronym for Nomenclature of Territorial Units for Statistics used by Eurostat. In this nomenclature NUTS1 refers to European Community Regions and NUTS2 to Basic Administrative Units, with NUTS3 reflecting smaller spatial units most similar to counties in the US. Local inflation rates are computed on the basis of surveys conducted in the main cities of NUTS3 regions.

and this difference can be explained by the degree of competitiveness of the local retail sector.

As a consequence, let us suppose local inflation rates to be generated by different AR(1) processes:

$$\pi_{it} = \alpha_i + \beta_i \pi_{it-1} + \sum_{j=1}^3 \gamma_{ij} m_{ijt} + u_{it} \quad (1)$$

where  $\pi_{it}$  is the inflation rate in the main city of province  $i$  at time  $t$ ,  $m_{ijt}$  is a quarterly dummy accounting for the possible effects of seasonality,  $u_{it}$  is a stochastic error,  $\alpha$ ,  $\beta$  and  $\gamma$  are the parameters to be estimated<sup>5</sup>. This implies that the long run inflation rate in region  $i$ ,  $\bar{\pi}_i$ , will be different than those in the other provinces and it will assume the following form:

$$\bar{\pi}_i = \frac{\alpha_i - \log \left( \frac{4}{\sum_j e^{\gamma_{ij}}} \right)}{1 - \beta_i} \quad (2)$$

where  $\log \left( \frac{4}{\sum_j e^{\gamma_{ij}}} \right)$  is a normalization necessary to correct the fact that, inserting quarterly

dummies and dropping one of them to avoid the dummy trap, entails arbitrarily assuming that long-run inflation shows up in the quarter of the dropped dummy (Suits, 1984). The possibility to use (2) as a measure of long-run inflation hinges on the absence of major structural breaks in the underlying parameters which was successfully tested by Vaona and Ascari (2007).

Exploiting their dataset, we first estimated an AR(1) model for each one of the series of the local inflation rates. Then, we tested whether (2) is equal in all the provinces or not.

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<sup>5</sup> It is possible to consider the model also as an heterogeneous panel one. In that case, the seasonal dummies will account also for the possible effects of national common factors, though factor loadings have been restricted to be constant across time and let to vary across different spatial units. Common factors in regional inflation dynamics have been investigated by Beck et al. (2006) and they were not found to reduce the variability of idiosyncratic parameters.

Considering the whole of the inflation series as a heterogeneous panel dataset, this is a test of 70 nonlinear restrictions on the estimated parameters of the AR(1) processes. We followed Greene (2003, pp. 108 - 111) and we relied on a Wald test robust to heteroskedasticity, which is distributed as a  $\chi^2$  with 70 degrees of freedom. The test returns a value of 451.21 with a p-value of 0.00, strongly rejecting the null of equality of local long-run inflation rates.<sup>6</sup>

The fact that (2) is not equal across different locations implies that there is not full convergence of inflation in the long-run at the regional level, which is in contrast with the evidence produced by Busetti et al. (2006) using an Italian dataset. To reconcile these apparently conflicting results, it is necessary to suppose that regional inflation rates are characterized by conditional rather than absolute convergence, where the conditioning factors are real economic variables at the local level (Vaona, 2007a). This is consistent with the trend purchasing power parity hypothesis (Papell and Prodan, 2006).

Part A of Table 1 sets out some descriptive statistics of our data on inflation. Comparing the distribution of inflation with that of long-run inflation it is possible to see that they have a similar average, but the latter one has a smaller dispersion. Nonetheless, regional disparities in long-run inflation remain remarkable as its minimum and maximum annual values are about 1.6% and 2.8%.

(Table 1 about here)

Table 2 shows that the four Italian macro-regions display on average very similar values of trend inflation. On the other hand, its variability appears to be starker within macro-regions, as showed by columns 2 to 4 of Table 2. Considering an analysis of

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<sup>6</sup> We also tested if (2) is stable across time for each time series, by checking whether it displays a structural break before and after 2001Q1, a date in between the EMU kick-off and the physical introduction of Euro banknotes. Confirming the results of Vaona and Ascari (2007) for the 72% of the regions involved in the study the null of no structural break was accepted at a 5% level. Furthermore, a Li (1996) test could not reject the null that the regional distribution of (2) before and after 2001Q1 is the same, returning a value of -1.18. A Li (1996) test is distributed like a standardized normal.

variance of long-run inflation rates across the four Italian macro-regions leads to a very similar result as an F-test of the model returns a p-value of 0.88. This is remarkable because economic disparities among Italian macro-regions has been a prominent economic policy issue since the unification of the country in the nineteenth century (Brunello et al. 2001). Furthermore, Vaona and Ascari (2007) find that inflation persistence is higher in the South, than in the Centre or in the North. Therefore, it is important to look for a plurality of factors that might explain such a regional pattern.

(Table 2 about here)

Finally, long run inflation is spatially correlated as a Moran's I test on this variable returned a value of 4.064 with a p-value of 0.00. Therefore, consistently with previous studies, inflation rates tend to be more similar the closer are the location they belong to.

### **The determinants of long-run inflation**

We further investigate the possible sources of the regional dispersion of long-run inflation, merging our data on (2) with a dataset of economic indicators produced by the Italian statistical office regarding local labour market areas (LLMAs). LLMAs are functional regions defined on the basis of the commuting flows, so they have an economic nature. Their size is in between NUTS3 regions and municipalities and it is possible to consider them as "metropolitan areas". We consider only the LLMAs of the main cities of the NUTS3 provinces, as for the other LLMAs there exist no data about inflation. It is also worth noting that after the 2001 census the boundaries of the LLMAs have been redesigned so that there is no temporal continuity of the data about LLMAs after 2003. Furthermore, no data regarding the labour force and the unemployed has been produced for the year 2003.

In the end, our baseline model regresses (2) on the unemployment rate, the population density, the log of the resident population, labour productivity growth, and the percentage of retail firms with no more than two employees.

We consider the unemployment rate to capture its possible effect on the local long-run inflation rate following the Phillips curve tradition (Vaona, 2007a). A high population density might increase local aggregate demand and exacerbate inflationary bottlenecks. Therefore, a model trying to explain the long-run level of local inflation has to include this factor too.

There exists a number of reasons to consider also the resident population as explanatory variable. First it can capture possible agglomeration effects, whereby larger LLMAAs might be more efficient in the use of their resources (Duranton and Puga, 2004) and enjoy a lower rate of long-run inflation. An alternative reason to insert the log of the resident population is that larger metropolitan areas are specialized in different activities than smaller ones, and so this regressor might capture the effect of differences in the industrial specialization of LLMAAs (Camagni, 1993, chp. 4).

If a Balassa-Samuelson effect is at work, faster labour productivity growth will lead to more inflation<sup>7</sup>. On the other hand, the Kaldorian tradition postulates that faster labour productivity growth leads to a decrease in long-run inflation by reducing the impact of cost inflation on firms' pricing decisions (Dixon and Thirlwall, 1975). Therefore, using labour productivity growth as regressor can offer a test for these two different schools of thought. Our measure of productivity growth is the percentage change of value added per worker deflated by the local CPI.

Vaona and Ascari (2007) showed inflation persistence - measured by  $\beta_i$  in (1) - to depend on the percentage of retailers with no more than two employees, which is often considered as a proxy of the degree of competitiveness of the local retail sector (Boylaud and Nicoletti, 2001). The fact that differences in the local degree of competitiveness could cause deviations from PPP was theoretically explored in the pricing to market literature (Rogoff, 1996). Therefore it is natural to include also this indicator when trying to explain (2). Part B of Table 1 sets out some descriptive statistics of the proposed regressors.

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<sup>7</sup> Froot and Rogoff (1995) list among the key assumptions of the Balassa-Samuelson model the international mobility of capital and an exogenous world real interest rate. Under these respects, considering a regional dataset should in principle provide a favourable environment to test for the presence of Balassa-Samuelson effects.

Regression results are shown in Table 3 (Model 1). Labour productivity growth is negatively and significantly correlated with long-run inflation. On the contrary, the less competitive is the retail sector and the higher is long-run inflation. The other regressors appear to be less successful in explaining long-run inflation. Residuals are well-behaved: they have zero-mean and the assumption of a normal distribution could not be rejected by a battery of tests.

(Table 3 about here)

Remarkably, no spatial correlation was detected in the residuals. Therefore, the chosen regressors were able to explain the spatial pattern assumed by local inflation rates. In other words, the role of geographic distance in explaining long-run inflation is just proxying for the similarity of the economic structure and performance of nearby regions<sup>8</sup>.

For Model 1 in Table 3, we also computed a nonparametric specification test after Ellison and Ellison (2000)<sup>9</sup>. Its null hypothesis is that the model fits the data in terms of functional specification and absence of omitted variables. We used a multivariate normal distribution for the kernel function and Silverman's optimal smoothing bandwidth as window width. We computed a different bandwidth for each regressor. The test has an asymptotic standard normal distribution and it returned a p-value of 0.33 supporting the model.

We performed two different robustness checks as well. First we inserted dummies accounting for the macro-regions where LLMA are located. This step is taken to check if some regional specificities bias our results (Model 2 in Table 3). Second, we distinguished

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<sup>8</sup> As a further robustness check for this conclusion we took the difference of each variable from the value of the Rome LLMA and we inserted among the regressors the geographic distance of each LLMA from Rome. This new regressor did not turn out to be significant even at a 10% level. The same happened when considering the log of the distance from Rome.

<sup>9</sup> The Ellison and Ellison test proved to be more successful when compared to other nonparametric tests in detecting functional misspecification (see also Miles and Mora, 2003). Ellison and Ellison (2000) mention the possibility to use their test also to detect the absence of omitted variables.

between the average labour productivity growth in manufacturing and service activities<sup>10</sup>, to assess the possible effects of local industrial specialization. We also included among the explanatory variables a dummy accounting for the presence of an industrial district within each LLMA because industrial districts might lead to more economic efficiency and a lower inflation rate in the long run. Furthermore, given the temporal discontinuity in unemployment data after 2002, we changed the average unemployment rate between 1998 and 2005 with the average unemployment rate between 1998 and 2002 (Model 3 in Table 3).

The results obtained in Model 1 proved to be robust to our checks. Comparing Model 3 with Model 2 it is possible to see that the coefficient of productivity growth in the service activities is negative and significant, while that in manufacturing is not significantly different from zero. In both the cases, tests on the residuals support the model.

### **Testing for Endogeneity**

We further tested for endogeneity of productivity growth. The reason underlying this choice is that the cross-country empirical literature on the connection between economic growth and inflation often found that the latter one might have a significant impact on productivity growth (Vaona and Schiavo, 2007 and Temple, 2000). We used as instruments for labour productivity growth, the level of real value added per worker at the beginning of the period of observation and the average percentage change in the labour force between 1998 and 2002. Similar explanatory variables are customarily considered in the empirical studies addressing the issue of the connection between inflation and economic growth. The real valued added per head at the beginning of the period of observation captures convergence forces<sup>11</sup>, while the faster the labour force grows and the slower will be the growth of productivity.

Descriptive statistics of the instruments used are offered in Table 1 Part D.

We ran a preliminary regression to check that our candidate instruments are in fact correlated with the instrumented variable. As expected, real value added per worker had a

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<sup>10</sup> Descriptive statistics of these variables are showed in Table 1, Part C.

<sup>11</sup> Economic convergence within Italy is a rather well documented phenomenon (Vaona, 2007b and Aiello and Scoppa, 2005).

negative coefficient estimate equal to -0.05 and a t-statistic equal to -2.04, whereas the average change in the labour force had a coefficient estimate of -0.32 and a t-statistic of -2.14. The F-statistic testing for the hypothesis that both the coefficients are equal to zero returned a value of 5.64 with a p-value of 0.005. Therefore, our instruments passed customary preliminary tests and allowed us to proceed implementing a 2SLS estimator.

We tested for endogeneity of the financial indicators by means of a Durbin-Wu-Hausman test which compares the 2SLS estimator with the OLS one. Having two instruments for the average percentage labour productivity growth we have one overidentifying restriction. In order to assess its validity, we also computed the test statistic given by the product between the number of observations and the  $R^2$  of the regression of the residuals of the 2SLS estimator on the percentage of firms in the retail sector with less than 2 employees and the instruments (Wooldridge, 2001).

As shown in Table 4, estimates resulting from the OLS and 2SLS are rather close. In fact the Hausman-Durbin-Wu test could not reject the hypothesis that they are equal, excluding sizeable endogeneity biases. The test for the overidentifying restriction supported the model. In the end, we can exclude sizeable endogeneity biases<sup>12</sup>.

(Table 4 about here)

## **Conclusions**

To conclude, this paper shows that there can exist statistically significant variability in local long-run inflation even when considering a 146-years-old economic union as Italy. Differences in regional inflation rates can be explained by differences in productivity growth and the degree of competitiveness of the local retail sector.

Productivity growth appears to be negatively correlated with long-run inflation supporting more the Kaldorian tradition, than the Balassa-Samuelson one. Considering the service sector as the non-traded one, productivity growth in non-tradables appears to affect long-run inflation more than that in tradables in line with the theoretical results obtained by

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<sup>12</sup> The p-value of the Durbin-Wu-Hausman test increases to 0.99 inserting the non significant regressors of Model 1 in Table 3.



Altissimo et al. (2005). This can be explained on the ground that the “traded sector relies more than others on intermediate inputs produced by other sectors in the economy [...]. Movements in the prices of non-traded goods that enter in the production or transportation of traded goods can be an important source of price dispersion for traded goods at the consumer level.” (Altissimo et al., 2005, p. 17).

Regarding the degree of competitiveness of local economies, Dixon and Thirlwall (1975) already postulated that changes in local mark-ups could produce changes in long-run inflation. However, they ignored that even the level of competitiveness of the local economy could generate differences in regional long-run inflation rates. This happens because arbitrage is hampered by market distortions, so that monopolistic rents can appropriate the benefits arising from productivity growth. As a matter of consequence, lack of competitiveness in the product market might obstacle the virtuous cycle hypothesized by the Kaldorian tradition: faster productivity growth in lagging regions might not fully translate into lower inflation rates reducing one area’s competitive advantage and the speed of the convergence process.

Finally, in the present study geographic distance appeared to proxy for the similarity of nearby regions more than having a role in itself.

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**Table 1 – Descriptive statistics of the variables involved in the study**

<b>Variable</b>	<b>Mean</b>	<b>Standard Deviation</b>	<b>Minimum Value</b>	<b>Maximum Value</b>
<b><u>Part A: inflation variables</u></b>				
<b>Inflation rate<sup>1</sup></b>	0.0059	0.0046	-0.0051	0.0747
<b>Long-run inflation<sup>1</sup></b>	0.0052	0.0005	0.0043	0.0069
<b><u>Part B: candidate determinants of long-run inflation</u></b>				
<b>Average unemployment rate between 1998 and 2005<sup>1</sup></b>	0.0829	0.0611	0.0302	0.2525
<b>Average unemployment rate between 1998 and 2002<sup>1</sup></b>	0.0863	0.0660	0.0261	0.2721
<b>Population density (thousands of people per km<sup>2</sup>)</b>	0.4260	0.5083	0.0439	3.2128
<b>Average population between 1998 and 2002 (thousands of people)</b>	384.92	589.17	53.96	3287.19
<b>Average growth rate of real labour productivity between 1996 and 2003<sup>1</sup></b>	0.0032	0.0093	-0.0190	0.0208
<b>Percentage of firms with no more than two employees in the retail sector (2001 census)<sup>1</sup></b>	0.7947	0.0371	0.6892	0.8823
<b><u>Part C: further candidate determinants of long-run inflation</u></b>				
<b>Average growth rate of real labour productivity in manufacturing between 1996 and 2003<sup>1</sup></b>	0.0543	0.0165	0.0088	0.0989
<b>Average growth rate of real labour productivity in service activities between 1996 and 2003<sup>1</sup></b>	0.0042	0.0114	-0.0238	0.0275
<b><u>Part D: instruments used to check for endogeneity of productivity growth</u></b>				
<b>Real value added per worker in 1996<sup>2</sup></b>	0.4159	0.0377	0.3357	0.5670
<b>Average percentage change in the labour force between 1998 and 2002<sup>1</sup></b>	0.0088	0.0070	-0.0087	0.0232

<sup>1</sup>To obtain percentages multiply values by 100. Inflation data have a quarterly frequency. All the data are produced by ISTAT, the Italian national statistical office. <sup>2</sup>: real value added per worker is measured in hundred thousands of 1995 Euros.

**Table 2 – Descriptive statistics of long-run inflation by macro-region**

<b>Macro-region</b>	<b>Mean</b>	<b>Standard Deviation</b>	<b>Minimum</b>	<b>Maximum</b>	<b>Observations</b>
	<b>(1)</b>	<b>(2)</b>	<b>(3)</b>	<b>(4)</b>	<b>(5)</b>
<b>North-West</b>	0.0052	0.0006	0.0045	0.0069	18
<b>North-East</b>	0.0053	0.0005	0.0043	0.0060	18
<b>Centre</b>	0.0052	0.0004	0.0044	0.0063	17
<b>South and Islands</b>	0.0054	0.0006	0.0046	0.0070	18

To obtain percentages multiply values by 100. Inflation data have a quarterly frequency.

**Table 3 – The determinants of local long-run inflation. Regression results.**

Dependent variable: long-run inflation\*100. Estimation method: Least Squares with Robust Standard Errors

	<b>Model 1</b>	<b>Model 2</b>	<b>Model 3</b>
<b>Average unemployment rate between 1998 and 2005</b>	-0.13	-0.22	-
<b>t-statistics</b>	-1.27	-0.76	-
<b>Average unemployment rate between 1998 and 2002</b>	-	-	-0.08
<b>t-statistics</b>	-	-	-0.40
<b>Population density</b>	0.01	0.01	0.01
<b>t-statistics</b>	0.21	0.61	0.54
<b>Log. of average population between 1998 and 2002</b>	0.01	-0.01	-0.01
<b>t-statistics</b>	0.24	-0.13	-0.40
<b>Average percentage labour productivity growth between 1996 and 2003</b>	-1.84*	-1.76*	-
<b>t-statistics</b>	-2.78	-2.58	-
<b>Average percentage labour productivity growth between 1996 and 2003 in manufacturing</b>	-	-	0.20
<b>t-statistics</b>	-	-	0.51
<b>Average percentage labour productivity growth between 1996 and 2003 in service activities</b>	-	-	-1.80*
<b>t-statistics</b>	-	-	-2.99
<b>Percentage of firms with no more than two employees in the retail sector</b>	0.66*	0.68*	0.68*
<b>t-statistics</b>	9.41	8.70	9.82
<b>The local labour market area is located in North-East Italy<sup>1</sup></b>	-	0.02	0.01
<b>t-statistics</b>	-	0.78	0.65
<b>The local labour market area is located in Central Italy<sup>1</sup></b>	-	0.01	-0.01
<b>t-statistics</b>	-	0.35	0.01
<b>The local labour market area is located in the South of Italy<sup>1</sup></b>	-	0.01	-0.01
<b>t-statistics</b>	-	0.23	-0.23
<b>The local labour market area is located in the Italian Islands<sup>1</sup></b>	-	0.04	0.03
<b>t-statistics</b>	-	0.65	0.58
<b>The local labour market areas has an industrial district<sup>1</sup></b>	-	-	0.11
<b>t-statistics</b>	-	-	0.81
<b>Test for zero mean in the residuals (p-value)<sup>2</sup></b>	0.95	0.96	0.97
<b>Shapiro – Francia test (p-value)<sup>3</sup></b>	0.16	0.18	0.08
<b>Shapiro – Wilk test (p-value)<sup>3</sup></b>	0.18	0.27	0.07
<b>Skeweness – Kurtosis test (p-value)<sup>3</sup></b>	0.14	0.10	0.14
<b>Test for spatial correlation in the residuals (p-value)<sup>4</sup></b>	0.92	0.80	0.71
<b>Observations<sup>5</sup></b>	70	70	71

Following Eisenhauer (2003), the constant was dropped because it was not significantly different from zero at a 5% level. \*: significantly different from zero at the 1% level. 1: dummy variables. The control group is constituted by the LLMA in the North-West of Italy. 2: the null hypothesis is that residuals have zero mean. 3: the null hypothesis is that residuals are normally distributed. 4: the test for spatial correlation is the Moran's I statistic which is distributed as  $N(0,1)$  and whose 5% critical value is 1.96. The null hypothesis is absence of spatial correlation. For an introduction to this test see Anselin (1988). The spatial weight matrix was obtained setting to one the elements of a null matrix in correspondence to LLMA belonging to contiguous NUTS3 regions. 5: for model 3 we have 71 observations because the LLMA of Chieti was suppressed after 2002.

**Table 4 – The determinants of local long-run inflation. Endogeneity tests.**

Dependent variable: long-run inflation\*100.

	<b>OLS<sup>a</sup></b>	<b>2SLS</b>
<b>Percentage of firms with no more than two employees in the retail sector</b>	0.67**	0.68**
<b>t-statistics</b>	78.79	57.18
<b>Average percentage labour productivity growth between 1996 and 2003</b>	-2.05**	-3.96*
<b>t-statistics</b>	-2.96	-2.04
<b>Hausman-Durbin-Wu test (p-value)<sup>1</sup></b>		0.57
<b>Test for overidentifying restrictions (p-value)<sup>2</sup></b>		0.40
<b>Observations</b>	71	71

Notes: \*\*: significant at a 1% level. \*: significant at a 5% level. Instruments in the 2SLS regression in the third column include the real value added per worker in 1996 and the percentage change of the labour force between 1998 and 2002. a: preferred estimates. <sup>1</sup>: the null is no endogeneity in the comparison between the OLS and the 2SLS estimators. <sup>2</sup>: the null is that over-identifying restrictions are not rejected.