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Abstract

Considering a sample of 71 Italian metropolitan areas, 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[1]. 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).

One of the reasons why relative PPP might not hold is different productivity growth rates across countries. Balassa (1964) originally acknowledged that the “productivity bias”, as called by Bahmani-Oskooee and Nasir (2005), could have different directions. Suppose there exist two sectors in a given economy, a traded one and a not-traded one. Suppose further that productivity growth is stronger in the former than in the latter one and that on aggregate it outpaces the growth rate of wages. As a consequence, inflation will decrease, but this decline will be less marked in the non-traded sector. This possible negative relationship between inflation and productivity growth has been emphasized in regional economics by the Kaldorian tradition (Dixon and Thirlwall, 1975).

Alternatively, it might be, first, that wages in the tradable sector grow as fast as productivity does and, second, that wages in the non-tradable sector are pegged to those in the tradable one (possibly due to competition among labour groups). Under these circumstances, inflation will rise due to an acceleration of price growth in the non-tradable sector.

This paper focuses on the issues above by exploiting a dataset of metropolitan areas, that is a dataset of small open economies belonging to a monetary union. When many economies share the same currency they will have

[1] For reviews see Rogoff (1996) and Taylor and Taylor (2004).

a fixed nominal exchange rate, implying that if relative PPP holds, local inflation rates will converge to the same value in the long-run.

We depart from standard ways of testing for PPP or for Balassa-Samuelson effects. We build on the literature on inflation persistence (Lünnemann and Mathä, 2004 and Vaona and Ascari, 2007), but we move beyond the short-run. 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 local 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 intra-national 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).

Furthermore, our measure is similar, but superior to the measures of trend inflation used in the literature investigating the relationship between inflation and the slope of the Phillips curve. For instance, Ball, Mankiw and Romer (1988) used just the average inflation, Ball (1994) and Boschen and Weise (2001) a nine quarters moving average, Hofstetter (2008) a 10 years inflation average and Senda and Smith (2008) experimented with 2, 5 and 10 years inflation averages. By using a Schwartz criterion, we let as much as possible the data speak about the length of the time period over which average inflation should be computed.

Our research strategy is possible because we consider a dataset with a large cross-sectional dimension (71 metropolitan areas), much larger, for instance, than that of datasets concerning the 19 major U.S. cities studied by Cecchetti et al. (2002) and Chen and Deveraux (2003).

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)[2].

The rest of the paper is structured as follows. The next section offers a review of papers testing for PPP mainly in a regional/urban context. The second section shows some features of our dataset and our econometric results. The last section concludes.

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):

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

[2] At page 21.

where Δ is the first difference operator, μ is a constant, α and c_i are coefficients, k is the number of lags considered and ε is the stochastic error. If α 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, often being characterized by 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 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 an 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 traded and not-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 (2005) 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 a smooth decline in their dispersion; 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 (2007) 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 reversion to the mean, which resulted to be conditional on local unemployment rates. Therefore, similarly to previous contributions[3], macroeconomic factors, such as the unemployment rate, can explain deviations from PPP[4].

The finding of conditional mean reversion 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[5]. From the theoretical point of view, Obstfeld (1993) offered a model in which real exchange rates have a trend caused by differential productivity growth in tradable and non-tradable goods.

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

[3] See Rogoff (1996), p. 663, where it is discussed the hypothesis that government spending might have an effect on PPP.

[4] Intranational price convergence has recently become the topic of a number of different papers such as Fan and Wei (2006) for China, Ceglowski (2003) for Canada, Dan and Battacharya (2008) for India. Morshed (2007), instead, focused on Bangladeshi and Pakistani cities trying to understand if state borders have an impact on price convergence.

[5] One of the most active researchers in the field is David Papel. See for instance the papers quoted in Banerjee et al. (2005).

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, the Kaldorian tradition postulates that output growth is positively connected to labour productivity growth, which decreases long-run inflation by offsetting firm's cost inflation. The decrease in long-run inflation will in its turn lead to an increase in export and to more output growth, starting an economic virtuous cycle. However, empirical papers belonging to this stream of literature have been more concerned with testing the real part of the model (the connection of output growth, labour productivity growth and export) than the connection between labour productivity growth and inflation as we do here (see for instance Fingleton and McCombie, 1998 and McCombie, 1985).

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[6].

Econometric Analysis

The urban dispersion of inflation rates, unit root testing and estimating local long-run inflation

The analysis here proposed builds on the results of Vaona and Ascari (2007). There a dataset of 71 local Italian inflation rates between 1996Q1 and 2006Q3 was analysed within a short-run framework. AR models with seasonal dummies were fit to intra-national inflation time series and the estimated degree of inflation persistence resulted to be low and hardly affected by structural

[6] Where the long-run is identified by a low elasticity of substitution of labour inputs in the tradable and non-tradable sectors.

breaks in the period considered, implying that standard econometric methods provide reliable coefficient estimates. Furthermore, inflation persistence appeared to be statistically different across provinces and this difference could be explained by the degree of competitiveness of the local retail sector.

Analysing this dataset it is possible to find that inflation displays similar features to those emerged in the literature reviewed above. Figure 1 shows the cross-sectional coefficient of variation among Italian metropolitan areas across time.

(Figure 1 about here)

Furthermore, we run both for the level of the CPI and for inflation the panel unit root tests proposed by Levin, Lin and Chu (2002), Im, Pesaran and Shin (2003) and the Fisher-type tests using ADF and PP tests after Maddala and Wu (1999) and Choi (2001). Results are set out in Table 1.

(Table 1 about here)

While the null hypothesis of the presence of a stochastic trend was strongly accepted for the CPI, it was strongly rejected for inflation. So, similarly to Weber and Beck (2005) and Busetti et al. (2006), inflation rates appear to converge towards the mean across metropolitan areas in Italy, though their dispersion does not steadily decline. However, we do not stop here and we tackle the issue whether they converge to a common mean or there exists a "productivity bias" in our data.

In order to do so, let us suppose local inflation rates to be generated by different AR processes:

$$\pi_{it} = \alpha_i + \sum_{k_i=1}^{K_i} \beta_{ik_i} \pi_{it-k_i} + \sum_{j=1}^{3} \gamma_{ij} m_{ijt} + u_{it} \qquad (1)$$

where π_{it} is the inflation rate in the main city i at time t, m_{ijt} is a quarterly dummy accounting for the possible effects of seasonality, u_{it} is a stochastic error, α, β and γ are the parameters to be estimated, K_i is the maximum lag length chosen for city i[7]. This implies that the long run inflation rate in city i, $\bar{\pi}_i$, will be different from those in the other cities and it will assume the following form:

$$\bar{\pi}_i = \frac{\alpha_i - \log\left(\frac{4}{\sum_j e^{\gamma_{ij}}}\right)}{1 - \sum_{k_i}^{K_i} \beta_{ik_i}} \tag{2}$$

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)[8]. 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).

We estimate (1) for each one of the series of the local inflation rates, choosing K_i by means of a Schwartz criterion. Part A of Table 2 sets out some

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[7] 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.

[8] We used the normalization for a log-linear equation given that $\pi_{it} = \log\left(\frac{P_{it}}{P_{it-1}}\right)$

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, intra-national disparities in long-run inflation remain remarkable as its minimum and maximum annual values are about 1.7% and 2.9%[9].

(Table 2 about here)

Table 3 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 3. Considering an analysis of 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 0.52. 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 3 about here)

The determinants of long-run inflation

We further investigate the possible sources of the intra-national 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

[9] Long-run inflation rates by metropolitan areas are set out in detail in Table A1 in the Appendix.

the commuting flows, so they have an economic nature. Their size is in between NUTS3 regions[10] and municipalities and it is possible to consider them as metropolitan areas. We consider only the LLMAs of the main cities of the NUTS3 regions, 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 (U), the population density (PD), the resident population (RP), labour productivity growth (LPG), and the percentage of retail firms with no more than two employees (MP), the "mom-and-pop" stores as called by Boylaud and Nicoletti (2001).

$$\overline{\pi}_i = \alpha + \beta_0 U_i + \beta_1 PD_i + \beta_2 RP_i + \beta_3 LPG_i + \beta_4 MP_i + \xi_i \qquad (3)$$

where α and β_j with j=0,…,4 are coefficients, ξ is a stochastic error and i indicates the i-th metropolitan area.

We consider the unemployment rate to capture its possible effect on the local long-run inflation rate following the Phillips curve tradition (Vaona, 2007). 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 LLMAs might be more efficient in the use of their resources (Duranton and Puga, 2004) and enjoy a lower rate of long-run inflation. An

[10] 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.

alternative reason to insert 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 LLMAs (Camagni, 1993, chp. 4).

As explained in the introduction, the effect of labour productivity growth is one of the main issues of this paper. Our measure for it is the percentage change of value added per worker deflated by the local CPI.

Vaona and Ascari (2007) showed inflation persistence - measured by $\sum_{k_i}^{K_i} \beta_{ik_i}$ in (1) - to depend on the percentage of retailers with no more than two employees, which is often considered as a proxy for the degree of protection of the local retail sector (Boylaud and Nicoletti, 2001)[11]. 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 2 sets out some descriptive statistics of the proposed regressors.

Regression results are shown in Table 4 (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[12]. Residuals are well-behaved: they have zero-mean and the assumption of a normal distribution could not be rejected.

(Table 4 about here)

[11] The intuition being that small shops cannot stay on the market in presence of economies of scale, which are present in the retail sector (see for instance Betancourt and Malanoski, 1999).

[12] Though having the expected signs.

Remarkably, no spatial correlation was detected in the residuals, that is residuals of contiguous metropolitan areas are not more correlated than those of more distant ones. Therefore, the chosen regressors were able to explain the spatial pattern assumed by local inflation rates. In other words, geographic distance would not appear to explain long-run inflation differentials once inserting variables accounting for the economic structure and performance of metropolitan areas. 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.

For Model 1 in Table 4, a nonparametric specification test after Ellison and Ellison (2000) was also computed[13]. Its null hypothesis is that the model fits the data in terms of functional specification and absence of omitted variables. We used a quartic distribution for the kernel function and we set the smoothing bandwidth as in Miles and Mora (2003): we divided each regressor by its standard deviation and, then, we used $h=\lambda n^{-1/(d+4)}$ as bandwidth, where d is the number of regressors. The test supports the model.

We performed various robustness checks as well (Models 2 and 3 in Table 4). First we inserted dummies accounting for the macro-regions where LLMAs are located. This step is taken to check if some regional specificities bias our results[14]. Second, we distinguished between the average labour productivity

[13] 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.

[14] For instance we cannot control for money growth, which could be different in different cities possibly due either to credit market segmentation and credit rationing, or to different stages of development of the credit system or to different liquidity preferences of lenders and borrowers (Dow and Rodriguez-Fuentes, 1997). Data on credit for Italian LLMAs exist but they are based on the location of banks and not of borrowers. Inserting the average

growth in manufacturing and service activities[15], 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[16].

The results obtained in Model 1 proved to be robust to our checks. Comparing Model 1 with Model 2 it is possible to see that the coefficient of productivity growth in service activities is negative and significant, while that in manufacturing is not significantly different from zero. Tests on the residuals support the model.[17] Finally, once resorting to a different measure of long-run inflation, that is the log of the ratio between the CPI indexes at the beginning and at the end of the period of observation, our results would not change (Table 4, Model 3).

Testing for Endogeneity

We further tested for endogeneity of both productivity growth and percentage of retailers with no more than two employees.

One 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).

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growth rate of this variable between 1998 and 2002 into our regressions would not return a significant t-statistic. Further details are available from the author upon request.

[15] Descriptive statistics of these variables are showed in Table 2, Part C.

[16] Given the temporal discontinuity in unemployment data after 2002, we also tried to change the average unemployment rate between 1998 and 2005 with the average unemployment rate between 1998 and 2002 and results are robust.

[17] It would be possible to argue that our dependent variable is estimated in a first stage regression and that this might induce heteroskedasticity. For this reason we used robust regression analysis. However, following Lewis and Linzer (2005) we computed also a weighted least squares estimator and a feasible GLS one. Results are stable as showed in Table A2 in the Appendix.

The other reason is that, following Chirinko and Fazzari (2000), it would be possible to think that more inflation spurs consumers to look for better deals decreasing market power in the retail sector. To the extent that the percentage of retailers with no more than two employees captures distortions generated by regulations adopted by local authorities, coefficient estimates will not be biased by endogeneity. However, we prefer to take a conservative stance and to test for endogeneity by means of a Durbin-Wu-Hausman test, which compares a 2SLS estimator with an OLS one (Wooldridge, 2002).

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 worker at the beginning of the period of observation captures convergence forces, while the faster the labour force grows and the slower will be the growth of productivity.

To instrument the percentage of retailers with no more than two employees we built on the fact that a less protective regulation of the retail sector has been adopted in the most developed parts of the country (Argiolas and Ventura, 2002). So we used as instrument the level of real value added per worker at the beginning of the period of observation[18].

On the other hand, excluding our instruments from the model of inflation is consistent with the original reasoning of Balassa (1964)[19]. Regarding real value added per worker, Balassa (1964) argues that differences in the level of productivity affect the level of prices and that this might translate into inflation differentials because productivity growth is faster in countries with a lower initial level of productivity. Concerning the growth rate of the labour force,

[18] In this way the system is exactly identified, having two instruments for two instrumented variables.

[19] It is worth recalling that exclusion restrictions cannot be tested (see for instance, Hsiao, 1983).

Balassa (1964) assumes that wage setting is determined by the competition among different “labour groups”, alluding to a non-competitive structure of the labour market where the bargaining power of either unions or insiders’ groups might be a key factor. Furthermore, the growth rate of the labour force is not customarily inserted in models of inflation dynamics, such as Phillips curve ones where either the unemployment rate or the real unit labour cost or the output gap are used (Vaona, 2007).

Descriptive statistics of the instruments used are offered in Table 2 Part D.

We ran preliminary regressions to check that our candidate instruments are actually correlated with the instrumented variable. For both the equations the F statistic returned a p-value of 0.00. So our instruments passed customary preliminary checks. Furthermore, the value of the F-statistic was equal to 2504.45, which, being much greater than 10, alleviates possible concerns that our results are affected by a weak instruments problem after Staiger and Stock (1997).

As shown in Table 5, 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. So our preferred estimator is OLS. In the 2SLS estimator, productivity growth is not significantly different from zero. However, first, the Hausman-Durbin-Wu test does not support these estimates and, second, it is a well known fact that “2SLS standard errors have a tendency to be large” (Wooldridge, 2002, p. 102).

(Table 5 about here)

Conclusions

To conclude, this paper shows that there can exist significant variability in local long-run inflation even when considering a 146-years-old economic union as Italy. Differences in metropolitan 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. 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 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. Consistently with this assumption, we showed that the degree of competitiveness of the local economy can generate differences in intra-national 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, geographic distance did not appear to have a role in the present study as, once inserted in a regression equation, its coefficient was not significantly different from zero. Furthermore, spatial econometric testing could not detected any correlation in the residuals, so spatially closer observations did not appear to be more correlated than farther ones.

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(Appendix and Tables A1 and A2 here)

Table 1 – Unit root tests for CPI and Inflation in 71 Italian metropolitan areas, 1996Q1-2006Q3

CPI					**Inflation**				
Method	Statistic	Prob.**	sections	Obs	Method	Statistic	Prob.**	sections	Obs
Levin, Lin & Chu	1.096	0.86	71	2888	Levin, Lin & Chu	-33.67	0.00	71	2819
Im, Pesaran and Shin W-stat	14.23	1.00	71	2888	Im, Pesaran and Shin W-stat	-34.06	0.00	71	2819
ADF - Fisher Chi-square	27.95	1.00	71	2888	ADF - Fisher Chi-square	1300.90	0.00	71	2819
PP - Fisher Chi-square	40.29	1.00	71	2941	PP - Fisher Chi-square	1671.81	0.00	71	2870

Notes: ** Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. Exogenous variables: Individual effects. Maximum lags were automatically selected on the basis of the Schwartz criterion. The Newey-West bandwidth was selected using the Bartlett kernel. Null hypothesis: unit root.

Table 2 – Descriptive statistics of the variables involved in the study

Variable	Mean	Standard Deviation	Minimum Value	Maximum Value
Part A: inflation variables				
Inflation rate[1]	0.0059	0.0046	-0.0051	0.0747
Long-run inflation[1]	0.0054	0.0005	0.0044	0.0072
Part B: candidate determinants of long-run inflation				
Average unemployment rate between 1998 and 2002[1]	0.0863	0.0660	0.0261	0.2721
Population density (thousands of people per km^2)	0.4260	0.5083	0.0439	3.2128
Average population between 1998 and 2002 (thousands of people)	384.92	589.17	53.96	3287.19
Average growth rate of real labour productivity between 1996 and 2003[1]	0.0032	0.0093	-0.0190	0.0208
Percentage of firms with no more than two employees in the retail sector (2001 census)[1]	0.7947	0.0371	0.6892	0.8823
Part C: further candidate determinants of long-run inflation				
Average growth rate of real labour productivity in manufacturing between 1996 and 2003[1]	0.0543	0.0165	0.0088	0.0989
Average growth rate of real labour productivity in service activities between 1996 and 2003[1]	0.0042	0.0114	-0.0238	0.0275
Part D: instruments used to check for endogeneity of productivity growth				
Real value added per worker in 1996[2]	0.4159	0.0377	0.3357	0.5670
Average percentage change in the labour force between 1998 and 2002[1]	0.0088	0.0070	-0.0087	0.0232

Notes. [1]: 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. [2]: real value added per worker is measured in hundred thousands of 1995 Euros.

Table 3 – Descriptive statistics of long-run inflation by macro-region

Macro-region	Mean (1)	Standard Deviation (2)	Minimum (3)	Maximum (4)	Observations (5)
North-West	0.0052	0.0006	0.0044	0.0071	18
North-East	0.0053	0.0005	0.0045	0.0061	18
Centre	0.0053	0.0003	0.0046	0.0059	17
South and Islands	0.0055	0.0006	0.0049	0.0070	18

Notes: to obtain percentages multiply values by 100. Inflation data have a quarterly frequency.

Table 4 – The determinants of local long-run inflation. Regression results.

Dependent variable: in Models 1 and 2 $\bar{\pi}_i$*100, where $\bar{\pi}_i$ is defined in (2), in Model 3 $\ln P_T - \ln P_0$ where P is the price index, T is the last period of observation and 0 the first period of observation.. Estimation method: Least Squares with Robust Standard Errors

	Model 1	Model 2	Model 3
Average unemployment rate between 1998 and 2002	-0.11	-0.16	-0.02
t-statistics	-1.14	-0.71	-0.34
Population density	0.01	0.02	0.01
t-statistics	1.36	1.25	0.61
Average resident population between 1998 and 2002	-0.01	-0.01	-0.01
t-statistics	-1.30	-1.45	-0.80
Average percentage labour productivity growth between 1996 and 2003	-1.34*	-	-
t-statistics	-2.00	-	-
Average percentage labour productivity growth between 1996 and 2003 in manufacturing	-	0.61	0.25
t-statistics	-	1.71	1.42
Average percentage labour productivity growth between 1996 and 2003 in service activities	-	-1.47*	-0.61*
t-statistics	-	-2.52	-2.32
Percentage of firms with no more than two employees in the retail sector	0.68*	0.64*	0.25*
t-statistics	48.68	18.21	17.23
The local labour market area is located in North-East Italy[1]	-	0.02	0.01
t-statistics	-	1.01	1.29
The local labour market area is located in Central Italy[1]	-	0.01	-0.01
t-statistics	-	0.60	-0.09
The local labour market area is located in the South of Italy[1]	-	0.02	-0.01
t-statistics	-	0.59	-0.45
The local labour market area is located in the Italian Islands[1]	-	0.05	0.01
t-statistics	-	1.00	0.38
The local labour market areas has an industrial district[1]	-	0.01	0.01
t-statistics	-	0.34	0.82
Test for zero mean in the residuals (p-value)[2]	0.96	0.97	-
Shapiro – Francia test (p-value)[3]	0.24	0.08	-
Test for spatial correlation in the residuals (p-value)[4]	0.88	0.83	-
Ellison and Ellison test (p-value)[5]	0.83	0.34	-
Observations	71	71	71

Notes: 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 5% level. 1: dummy variables. The control group is constituted by the LLMAs 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 asymptotically distributed as N(0,1). 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 LLMAs belonging to contiguous NUTS3 regions. 5: the null is that the model is well specified in terms of functional form and absence of omitted variables.

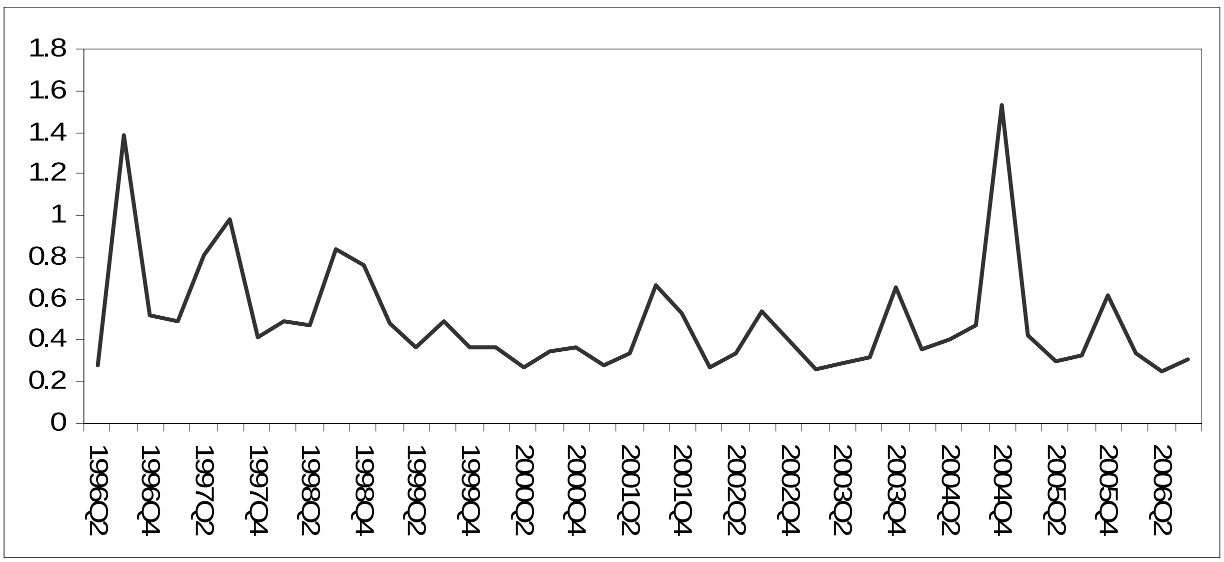
Table 5 – The determinants of local long-run inflation. Endogeneity tests.
Dependent variable: long-run inflation*100.

	OLS[a]	2SLS
Percentage of firms with no more than two employees in the retail sector	0.68***	0.69***
t-statistics	73.04	45.40
Average percentage labour productivity growth between 1996 and 2003	-1.52**	-4.65
t-statistics	-2.18	-1.47
Hausman-Durbin-Wu test (p-value)[1]	0.58	
Observations	71	71

Notes: ***: significant at a 1% level. **: significant at a 5% level.

Instruments in the 2SLS regression in the second 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. [1]: the null is no endogeneity in the comparison between the OLS and the 2SLS estimators.

Figure 1 – Inflation dispersion across metropolitan areas in Italy, 1996Q2-2006Q3



Note: Dispersion is measured by the coefficient of variation.

Appendix

Table A1 – Long-run inflation rates by metropolitan areas – Italy, 1996Q1-2006Q3

Metropolitan Area	Long-run Inflation	Standard deviation of long-run inflation	Number of lags in model (1)	Metropolitan Area	Long-run Inflation	Standard deviation of long-run inflation	Number of lags in model (1)
North-West				**Centre**			
Alessandria	0.005094	0.000436	2	Ancona	0.005314	0.000776	2
Aosta	0.004482	0.000447	3	Arezzo	0.005248	0.000971	3
Asti	0.004754	0.000384	2	Ascoli Piceno	0.005632	0.000698	4
Brescia	0.005604	0.000393	2	Firenze	0.004934	0.000414	2
Como	0.005145	0.000418	1	Grosseto	0.005815	0.000388	1
Cremona	0.004993	0.000465	1	Latina	0.00544	0.000919	2
Cuneo	0.005579	0.000392	3	Livorno	0.004776	0.000671	4
Genova	0.005056	0.000512	2	Lucca	0.005123	0.000403	1
Mantova	0.005899	0.000625	3	Macerata	0.005922	0.000331	1
Milano	0.00499	0.000431	2	Perugia	0.005119	0.000668	4
Novara	0.004571	0.000515	1	Pesaro	0.005689	0.000694	3
Pavia	0.005333	0.000421	1	Pisa	0.005488	0.000495	2
Savona	0.005469	0.000388	1	Pistoia	0.005117	0.000518	3
Sondrio	0.004357	0.000505	4	Roma	0.005479	0.000488	3
La Spezia	0.007167	0.000937	2	Siena	0.004633	0.000562	3
Torino	0.006122	0.000515	1	Terni	0.005395	0.000777	4
Varese	0.00516	0.000247	4	Viterbo	0.00548	0.000631	4
Vercelli	0.004478	0.002056	4	**South**			
North East				L'Aquila	0.005748	0.001272	4
Belluno	0.004983	0.000451	1	Bari	0.005188	0.000949	4
Bologna	0.005194	0.000345	4	Brindisi	0.005431	0.001121	4
Ferrara	0.005132	0.000447	3	Campobasso	0.004872	0.000412	1
Forlì	0.00471	0.000487	1	Chieti	0.005276	0.000294	3
Modena	0.005748	0.000859	3	Cosenza	0.006247	0.000867	2
Padova	0.004756	0.000544	2	Foggia	0.005901	0.000419	1
Parma	0.005588	0.000696	4	Napoli	0.005922	0.000471	1
Piacenza	0.004703	0.000729	4	Pescara	0.005444	0.001359	4
Pordenone	0.006005	0.000496	1	Potenza	0.004884	0.001136	3
Ravenna	0.006076	0.000402	1	Reggio Calabria	0.005101	0.000396	3
Reggio Emilia	0.0056	0.000391	1	Teramo	0.00699	0.000509	1
Rovigo	0.004457	0.000621	3	Cagliari	0.00499	0.000418	1
Trento	0.00541	0.000445	3	Catania	0.005326	0.000517	1
Treviso	0.005979	0.000441	1	Palermo	0.004913	0.000563	3
Trieste	0.005773	0.000533	4	Sassari	0.005532	0.000307	2
Udine	0.005111	0.00048	1	Siracusa	0.005925	0.000455	2
Venezia	0.005679	0.000439	1	Trapani	0.006011	0.000776	3
Verona	0.005108	0.000383	1				

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

Table A2 – The determinants of local long-run inflation. Regression results.
Dependent variable: long-run inflation*100. Estimation method: Least Squares with Robust Standard Errors

	Estimation Method	
	Weighted Least Squares°	**Feasible GLS°°**
Average unemployment rate between 1998 and 2002	-0.23	-0.18
t-statistics	-1.08	-0.83
Population density	0.02	0.02
t-statistics	1.02	0.95
Average resident population between 1998 and 2002	-0.01	-0.01
t-statistics	-0.81	-0.93
Average percentage labour productivity growth between 1996 and 2003	-	-
t-statistics	-	-
Average percentage labour productivity growth between 1996 and 2003 in manufacturing	-0.05	0.34
t-statistics	-0.12	0.87
Average percentage labour productivity growth between 1996 and 2003 in service activities	-1.20*	-1.39*
t-statistics	-2.34	-2.55
Percentage of firms with no more than two employees in the retail sector	0.67*	0.66*
t-statistics	20.32	20.02
The local labour market area is located in North-East Italy[1]	0.03	0.02
t-statistics	1.83	1.11
The local labour market area is located in Central Italy[1]	0.03	0.01
t-statistics	1.66	0.74
The local labour market area is located in the South of Italy[1]	0.02	0.02
t-statistics	0.83	0.70
The local labour market area is located in the Italian Islands[1]	0.04	0.04
t-statistics	1.04	1.00
The local labour market areas has an industrial district[1]	0.01	0.01
t-statistics	0.14	0.40
Error variance attributable to the sampling error in the dependent variable	-	0.01
Remaining error variance	-	0.01
Observations	71	71

Notes: 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 5% level. 1: dummy variables. The control group is constituted by the LLMAs in the North-West of Italy. °: the weights used are the standard deviations of inflation persistence, resulting from the estimation of equation (1). °°: weights have been estimated form the standard deviations of inflation persistence following Lewis and Linzer (2005), p. 353.