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Regional Inflation Persistence: Evidence from Italy

Abstract

Regional patterns of inflation persistence have received attention only at a very coarse level of territorial disaggregation, that of EMU member states. However economic disparities within EMU member states are an equally important policy issue. This paper considers a country with a large regional divide, i.e., Italy, at a fine level of territorial disaggregation (NUTS3). Our results show that economically backward regions display greater inflation persistence. Moreover, we show that higher persistence is linked to a lower degree of competitiveness in the retail sector. Finally, the inflation persistence at the national level does not present any geographical aggregation bias, because it equals the mean of inflation persistence of provincial data.

Keywords: inflation persistence, retail sector, regions
JEL Codes: E0, E30, R0, R10.

1. Introduction

Inflation persistence has become one of the central issues regarding the modelling of the inflation process. Indeed, the degree of inflation persistence is crucial for monetary policy, since if the inflation process is less persistent, the task of monetary policy is easier in terms both of sacrifice ratio and of controlling inflation fluctuations around a given target.

By now quite a large literature investigated the nature of the inflation process. First, some papers explore the nature of inflation persistence across countries (e.g., Benigno and Lopez-Salido, 2006), and also across sectors looking at different levels of disaggregation (see, e.g., Lünnemann and Mathä, 2004 and Cecchetti and Debelle, 2006). The main conclusions are: (i) once the estimation allows for shifts in the mean, inflation persistence is much lower than previously believed; (ii) inflation persistence differs across sectors.

Second, the ECB launched a big research project called the Inflation Persistence Network, in order to better understand the pricing behaviour of the single firms, and what is the impact of aggregating these different behaviours on the persistence of the aggregate inflation series. The main results are summarized in Altissimo et al. (2006) - see also Dhyne et al. (2005). Moreover, Cecchetti and Debelle, (2006) and Altissimo et al. (2007) underline the existence of an across sectors aggregation bias, because aggregate series inherit the properties of their most persistent component, thus justifying the importance to look at different sectoral prices. The same argument, obviously, should also hold when one aggregates geographically, that is, across regions within the same country rather than across sectors. Looking at the Euro data, Altissimo et al. (2007) stress that cross-country heterogeneity is less pronounced that cross-sector heterogeneity.

However, surprisingly enough, only few papers investigate the difference in the inflation process at a regional level within a single country. Cecchetti et al. (2002) analyse regional US price data to focus on deviations from PPP across the US and the dynamics of relative prices across regions. Beck and Weber (2005a,b) study inflation rate dispersion across US, Japan, Canada and EMU regions and investigate the issue of convergence of regional inflation rates using distribution dynamics methodology.

Busetti et al. (2006) consider the same issue on Italian regional data, but with a different methodology. Beck et al. (2006) use country specific factors as well as idiosyncratic regional components to examine the causes of the inflation dispersion across EMU regions.

None of the above papers, however, focuses on *regional inflation persistence*, that is, on the difference between *inflation persistence across regions*. The implications of regional, rather than sectoral, differences in inflation persistence for the functioning of monetary policy is rather limited as it stops at a very coarse level, that of EMU member states. However, it is obvious that the national indexes are built aggregating along the two dimensions: the sectoral and geographical ones.[1] While the former dimension has been extensively investigated (e.g., by the IPN in the Euro Area), the latter dimension, instead, has been somewhat neglected. This limitation should not be understated, since differences in regional inflation persistence could evidently be as much important as differences in sectoral inflation persistence. As wide regional disparities exist within EMU member states, a common monetary policy should take them into account. Indeed, Benigno (2004) showed that central banks should overweight, within their target index, regions with stickier price developments and underweight more flexible regions, in order to avoid that the former ones bear a disproportionate part of the adjustment process following a monetary shock. Suppose that more backward regions are also the more rigid ones, because, though having the same labour market institutions as more developed regions, they have less competitive product markets. If the central bank does not adjust its target inflation index, backward regions will be affected for a longer time by monetary shocks.

The novelty of this paper is to investigate regional inflation persistence, that is, the nature and causes of differences in inflation persistence across regions at a fine level of territorial dis-aggregation in Italy (NUTS 3 regions[2]). Like for the sectoral disaggregation in IPN types of studies, we want to know: (i) if and how much inflation persistence differs across regions; (ii) if so, what are the possible causes; (iii) if there is

[1] In the case of Italy, in computing the national consumer price index, the national statistical agency (ISTAT) builds first the provincial index, then the regional and, finally, the national one for each item, aggregating, thus, first geographically the indexes to obtain the national index for each item. It then aggregates across items to obtain the national consumer price index.

[2] 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.

a geographical aggregation bias. Finally, following the theoretical suggestion in Benigno (2004), it would be interesting to understand the policy implications for a common monetary policy of possible differences in geographical inflation persistence. In the end, the suggestion of Benigno (2004) to target the inflation index of the more rigid regions has to be judged empirically, by looking at whether the difference in regional inflation persistence is quantitatively relevant to justify such a strategy.

We focus on Italy, since the Italian regional divide is a well known problem, with Northern Italy being the most developed part of the country followed by the Centre and then by the South and Islands, also called "Mezzogiorno" (Brunello et al. 2001). Furthermore, the "Mezzogiorno" ghost has been evoked in trying to understand regional disparities within other European countries, notably Germany (Sinn and Westermann, 2001). Therefore, Italy provides a good example where studying how inflation persistence may interact with regional disparities.[3]

Moreover, from a methodological point of view, with respect to the previous literature: (i) we test if the differences in inflation persistence across regions are statistically significant, by means of a poolability test; (ii) we test the impact of possible structural breaks on the degree of persistence in inflation, employing a non parametric test on the estimated kernel density functions pre-break and post-break; (iii) we investigate the possible causes of these differences in inflation persistence across Italian regions using both non-parametric and parametric estimation methods.

The paper contains four main results. First, we do document that disparities in inflation persistence are statistically significant in Italy in our sample period. Moreover, it turns out that inflation persistence is higher in the relatively poorer part of the country. Second, there is no geographical aggregation bias, because the inflation persistence measured using the national index is roughly equal to the mean of the provincial inflation persistences. Third, we look at possible determinants of inflation persistence, and we find robust evidence that the degree of competitiveness in the retail sector is a key variable in explaining it. Finally, to investigate the policy implications of the previous analysis, and following Beningo (2004), we build a new CPI index

[3] The same analysis could also be interesting for the Euro Area, because inflation persistence differs across Euro area countries, or for other European countries with strong regional disparities (e.g., Spain, Germany). Such an analysis would complement the one in the present paper, also because the recent regional developments have differed considerably between Italian and most other European regions. This analysis is however outside the scope of the present paper and it is left to future research.

weighted by the persistence measure. This index, however, does not seem to behave very differently from the standard national CPI.

The rest of this paper is structured as follows. In the next section, we will deal with our estimates of inflation persistence. Then, we will illustrate its link with the local degree of competitiveness of the retail sector and we will consider a number of different robustness checks for our results. The last section concludes.

2. Estimating Persistence at the Regional Level

The Italian Statistical Office (ISTAT) has a long tradition in collecting data about prices in the main cities of NUTS3 regions. In this contribution we consider the data for the Consumer Price Index (CPI) from 1996Q1 to 2006Q3 at quarterly frequency for 70 out of 103 Italian provinces. The sample period is constrained only by electronic data availability and it is very similar to that studied by Lünnemann and Mathä (2004). We can distinguish 3 groups of provinces according to the number of observations. Five provinces have 36 observations[4], two provinces 40 observations[5] and all the rest 43 observations[6]. All the time series are continuous.

For each provincial index we estimate a univariate autoregressive process with a constant, as illustrated by the following equation

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

where π_{it} is the inflation rate in province i at time t, m_{ijt} is a quarterly dummy accounting for the possible effects of seasonality, u_{it} is a stochastic error, α_i, β_{ik_i} and γ_{ij} are the parameters to be estimated, K_i is the maximum lag length chosen for province i.

We choose the optimal lag length resorting to the Schwartz criterion, starting from a maximum lag length of 4. The selected optimal lag length is equal to one for 27 provinces, to 2 for 15 provinces, to 3 for 18 provinces and to 4 for 10 provinces.

Our indicator of inflation persistence is the sum of the autoregressive coefficients:

[4] Vercelli, Livorno, Latina, Teramo and Sassari.
[5] Sondrio and Macerata.
[6] Torino, Novara, Cuneo, Asti, Alessandria, Aosta, Savona, Genova, Spezia, Varese, Como, Milano, Pavia, Cremona, Mantova, Bolzano, Trento, Verona, Belluno, Treviso, Venezia, Padova, Rovigo, Udine, Trieste, Piacenza, Parma, Reggio Emilia, Modena, Bologna, Ferrara, Ravenna, Forlì, Pesaro, Ancona, Ascoli Piceno, Lucca, Pistoia, Firenze, Pisa, Arezzo, Siena, Grosseto, Perugia, Terni, Viterbo, Roma, Napoli, L'Aquila, Pescara, Chieti, Campobasso, Foggia, Bari, Brindisi, Potenza, Cosenza, Trapani, Palermo, Catania, Siracusa, Cagliari, Pordenone.

$$\rho_i = \sum_{k_i=1}^{K_i} \beta_{ik_i} \tag{2}$$

A number of measures of persistence have been offered in the literature and they usually return comparable results. It is worth stressing that one of the advantages of the sum of the autoregressive coefficients compared to other measures of persistence is that its confidence interval is very easy to compute. Thus, a poolability test is rather straightforward and, as in Lünnemann and Mathä (2004), we stick to this simple measure.

After estimating the model for the whole sample, we test - by means of Wald tests robust to heteroskedasticity - the restrictions of ρ being equal to zero or one, and for the presence of a structural break in the intercept and in ρ. For the last two tests, we take as possible reference times the EMU kick-off (1999Q1) and the introduction of Euro banknotes and coins (2002Q1). These possible breakpoints were chosen because a change in the monetary regime might affect the pricing behaviour of economic agents[7]. We also test, first, for a structural break in the variance by means of an F test, second, for the presence of serial correlation resorting to Durbin's *m* test[8], and third for omitted variables by means of a RESET test. Finally, we perform a poolability test[9] to control whether the estimates of ρ obtained for each province are statistically different. Indeed, different lag lengths in different provinces are a rather strong sign of geographic heterogeneity, however this does not per se imply that ρ is different across provinces.

In general, we find a rather low level of persistence: across different provinces ρ has an average of 0.25 and a standard deviation of 0.24[10]. Table 1 shows the results of

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[7] "[T]he launch of European Monetary Union and the establishment of a clearly defined nominal anchor [was] the defining event that changed the very nature of the inflationary process in the Euro area. This institutional break has eradicated the "intrinsic" component of the inflation formation mechanism, namely the extent to which economic agents – in resetting prices or negotiating wages – look at the past history of inflation, rather than into the eyes of the central bank" (Trichet, 2007, quoted in Benati, 2008).

[8] We always test for an order of serial correlation equal to the number of lags detected by the Schwartz criterion for the autoregressive model of inflation. So if, for instance, the Schwartz criterion points to an AR(3) model for inflation, we will test for third order serial correlation in the residuals.

[9] A Wald test robust to heteroskedasticity with null hypothesis that all the provinces have the same ρ.

[10] This low level of persistence is confirmed by unit root testing after 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). Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. The estimated model includes individual effects. Maximum lags were automatically selected on the basis of the Schwartz criterion. The Newey-West bandwidth was

the specification tests. It is evident that structural breaks do not appear to play a major role in our sample. These results are similar to those of Lünnemann and Mathä (2004) and consistent with those of Angeloni et al. (2006), who find a structural break in the inflation process in EMU member states in the mid-90s and a low, stable degree of inflation persistence thereafter. Given the low level of persistence, resorting to more sophisticated methods to avoid bias or unreliable inference does not seem necessary (Cecchetti and Debelle, 2006 and Hansen, 1999)[11]. Also serial correlation in the residuals does not affect our results and the autoregressive model of inflation does not appear to be plagued by the omission of relevant variables[12].

To further explore the impact of possible structural breaks on ρ, we estimate its kernel density function for the complete, pre-break and post-break estimates. We focus on the regional distribution of inflation persistence as its general picture might be more interesting to a central bank than its single local values. Figure 1 shows that structural breaks do not appear to have sizeable effects on the regional distribution of inflation persistence in our data, with the exception of the 2002 case where the number of provinces with a negative ρ increases somewhat. However, a Li (1996) test for equality between the regional distribution of ρ, before and after the structural break, could never reject the null returning a p-value of respectively 0.21 and 0.82 for the 1999 and the 2002 cases.

Finally, the poolability test is strongly rejected. Regional disparities in inflation persistence are quite remarkable. The average of ρ is rather similar in the North East and in the North West of Italy, where it is respectively equal to 0.22 and 0.21. Whereas in both the Centre and the South and Islands it is about 40% higher[13]. Therefore, inflation persistence appears to be greater in the lagging part of the country[14].

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selected using the Bartlett kernel. The null hypothesis of presence of unit root in all the series was rejected at a 1% level by all the tests.

[11] This is confirmed by our robustness checks below.

[12] Therefore the results here obtained are robust to the possible misspecification problems highlighted by Vaona (2007a) with reference to simple autoregressive models applied to long aggregate inflation time series for 19 countries.

[13] The few structural breaks detected were concentrated among Northern provinces more than Southern ones. For the 2002 tests, we found a structural break in the intercept in the provinces of Asti (North), Como (North), Livorno (Centre), Lucca (Centre), Reggio Emilia (North), Rovigo (North), Trapani (South), Varese (North). A structural break in ρ was found in the provinces of Alessandria (North), Ancona (Centre), Bari (South), Bologna (North), Bolzano (North), Como (North), Firenze (Centre), Latina (Centre), Lucca (Centre), Padova (North), Palermo (South), Pesaro (Centre), Pisa (Centre), Rovigo (North), Sassari (South), Siracusa (South), La Spezia (North). Similar results were found when testing for

Remarkably, once estimating an autoregressive model on national inflation data we obtain an estimate of ρ equal to 0.26, significantly different from both 0 and 1. No structural break, no serial correlation and no omitted variable is detected. The Schwartz criterion chooses an AR(1) model as the most suitable specification. The estimate of ρ is very close to the cross-section mean of provincial estimates, entailing that, contrary to the findings of Altissimo et al. (2007) for sectoral data, aggregate inflation series might not inherit the properties of the more persistent disaggregate regional inflation series.

It is tempting to derive some policy implications from the above analysis, regarding to what extent monetary policy should take into account regional inflation persistence dispersion. One immediate implication is that the cost of a disinflation should be larger in the South then in the North of the country, as higher inflation persistence means a larger sacrifice ratio. Interestly, Brunello et al. (2001) find that the unemployment rate increased more in the South than in the North during the disinflation of the nineties.

Moreover, our findings could suggest something about the type of inflation measure that a central bank should target. We find a generally low level of persistence across different provinces. Nonetheless, the difference is statistically significant, since the poolability test is rejected. According to Benigno (2004), in order to avoid welfare losses monetary policy should give an higher weight to the inflation in the region that features an higher degree of nominal rigidity, and hence of inflation persistence. Moreover, we also find that there is no geographical aggregation bias. Therefore, contrary to the sectoral dimension, targeting national inflation would not imply an indirect, and unintentional, bias on the more persistent region, since the national index does not inherit the properties of the most persistent provincial indexes. Benigno (2004, see Table 2, p. 316) suggests to compute a CPI index taking into account the relative

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structural breaks in 1999. It is therefore possible to exclude that the result of greater inflation persistence in the Centre and South is spurious, due to more structural breaks there, as structural breaks appeared more often in the North for the intercept, while for ρ they were concentrated in the North and in the Centre to a similar extent.

[14] Poolability has been at the centre of a number of different recent econometric contributions. Baltagi et al. (2003) and Baltagi et al. (2004) recommended adopting pooled estimators because they provide better forecasts and more plausible estimates. In our context we are not concerned with forecasting. Regarding the plausibility of estimates, we can distinguish poolability across space and across time. Inspecting Figure 1 it is possible to see that heterogeneity across space and homogeneity across time provide plausible estimates. Once allowing for structural breaks and therefore not pooling across time, we cannot rule out that inflation has an explosive behaviour in some provinces, which is contrary to the evidence on convergence of regional inflation rate in Italy provided by Vaona (2007b) and by Busetti et al. (2006).

size and the relative degree of nominal rigidity in the different regions. Then he shows that targeting such an inflation index would theoretically provide results very close to the fully optimal policy. We can try to get a flavour about the empirical importance of this theoretical argument. Following Beningo (2004), we build a CPI index weighted by the size of the province and the degree of inflation persistence. More precisely, l_i, the weight for each province i used to calculate our persistence-population weighted CPI (PPCPI), is given by

$$l_i = \frac{a_i \rho_i}{\sum_i a_i \rho_i},$$

where a_i is the relative size of the province i,[15] and ρ_i the inflation persistence in province i. Figure 2 shows the annualized inflation derived by using our PPCPI index and the one implied by the standard CPI. The two indexes basically overlaps, being only marginally different and comoving very closely[16]. This suggests that targeting a PPCPI inflation index would provide only marginal gains. As a result, our analysis implies that geographical dispersion of inflation persistence may not be an issue for monetary policy. In other words, when judged empirically, the suggestion in Benigno (2004) does not seem to be quantitatively relevant.[17]

Obviously, this policy conclusion should be taken with some caveats. First, the fact that there is no geographical aggregation bias in calculating inflation persistence may induce this result. Second, the same result may not hold for other countries or macroregions (as the Euro Area). Third, the inflation persistence in the provinces is different, but it is anyway overall quite modest, given our "great moderation" sample period. If it had been possible to get a longer sample, including periods of higher inflation persistence, also the registered geographical dispersion in inflation persistence, and hence the difference between the two indexes, might have been quantitatively larger.

To sum up, our results point towards a rather strong policy conclusion: opposite to sectoral heterogeneity, geographical dispersion in inflation persistence does not seem to be empirically relevant for monetary policy. Naturally, future work should check the

[15] We measure a_i as the average in our sample of the relative population in the province with respect to the national population.

[16] Their correlation is 0.96.

[17] Interestingly, Eusepi et al. (2008) perfom a similar exercise based on data on sectoral prices in a microfounded New Keynesian model and find that their index delivers substantial welfare gains.

robustness of this conclusion with respect to other sample periods and other countries or regions.

3. Regional Inflation Persistence and the Retail Sector

In order to shed further light on the possible sources of the geographical pattern above, we investigate the connection between regional inflation persistence and the structure of the retail sector. Recent empirical contributions have compared the persistence properties of consumer and producer price indexes in Italy (Sabbatini et al., 2004). These works show that the rate of change of the consumer price index displays more persistence than the producer price index, pointing to the degree of competitiveness of the retail sector as one of the possible sources of inflation persistence. In the present paper, we show that more persistent regions tend to have a less competitive retail sector.

The retail sector is often regulated at the local level and Italy makes no exception (Boylaud and Nicoletti, 2001). Furthermore, recent reforms, though originally aiming at liberalizing the sector, actually increased the regulatory power of local authorities, which often used it to inhibit competition, limiting the entry of large stores (Schivardi and Viviano, 2007).

The entry of large stores would generally be beneficial for competition, but only to the extent that they do not acquire local monopolistic power. On the other hand, if the effect of market regulation were to protect small shops, boosting their share in the total number of firms belonging to the sector, local monopoly power would tend to increase. Therefore, the connection between firm size and competitiveness in the retail sector is a rather complex one entailing possible non-linearities.

Regarding the connection between competitiveness and inflation persistence, a lower degree of competitiveness would induce higher inflation persistence through two channels. First, Leith and Mulley (2003) show that firms in a less competitive environment tend to adjust prices less frequently and are less likely to do so in a forward-looking manner. Second, the recent New Keynesian literature suggests that inflation persistence requires the existence of “rule of thumbers” among price setters (Galí and Gertler, 1999, and Altissimo et al., 2006). Intuitively, small shops are good candidates for “rule of thumb” price-setters, because they are unlikely to have the ability to forecast future inflation.

To this regard, Veronese et al. (2004) analyse the price dynamics of items of a specific brand sold in a specific outlet, for a total of 750,000 elementary price quotes in the Italian CPI basket. One finding is that traditional outlets tend to change the price significantly less frequently than large stores.

Following the results of the studies above, we want, therefore, to check in our database if there is a link between inflation persistence and the competitiveness of the retail sector, which we try to capture by means of two indicators. We consider both the large store floor space over resident population (*LS*) and, after Boylaud and Nicoletti (2001), the share of firms with no more than two employees in the retail sector, the “mom-and-pop stores” (*MP*)[18]. The former indicator is available for 1999 from Schivardi and Viviano (2007), while the latter one is available from the 2001 census. The South of Italy has a higher incidence of small shops than the North, where large stores gained ground in recent decades (Argiolas and Ventura, 2002).

However, inflation persistence might be driven by other factors than the degree of competitiveness of the local retail sector. In particular we would like to test two hypotheses. The first one is whether regional disparities in inflation persistence might also originate from local labour market characteristics, even though different regions of a country usually share the same labour market institutions. The second hypothesis is whether the industrial mix of a region has an impact on inflation persistence.

In order to test the first hypothesis we add the average of the local unemployment rate between 1998 and 2005 (*AU9805*) as one further explanatory variable. Unemployment might have different effects on inflation persistence via wage rigidity. Suppose that wage rigidity, induced by labour market institutions, produces a higher unemployment rate. If wage rigidity is also the source of a more pronounced inflation persistence, a higher unemployment rate will be connected to a greater inflation

[18] Following Chirinko and Fazzari (2000), it would also be possible to argue that inflation persistence affects the degree of competitiveness of the retail sector, rather than the opposite. Suppose that after a shock inflation takes more time to go back to its steady state value in a region with respect to an other, this will induce economic agents in the first region to spend more time shopping in search for better deals, decreasing monopoly rents in the retail sector. However, to the extent that - building on Schivardi and Viviano (2007) and on Boylaud and Nicoletti (2001) - regional variation in *LS* and *MP* can be attributed to local product market regulations, we can consider these variables as exogenous.

persistence. On the other hand, a higher unemployment rate might erode the bargaining power of insiders, reducing wage rigidity and thereby inflation persistence.

In order to test the second hypothesis we include in the model the average share of service activities in total local value added between 1996 and 2003 (*SS9603*). We also insert a set of regional dummies, to account for further factors that we might not directly capture (*D*). Descriptive statistics for the proposed explanatory variables are showed in Table 2.

We start investigating by a non-parametric estimator a possible non linear relationship between inflation persistence and the large store surface over resident population in 1999. Figure 3 shows that their relationship appears to be negative up to approximately 220 squared meters per 1000 inhabitants and positive thereafter. As a consequence we add to the explanatory variables above the square of large store surface over resident population in 1999. Regarding the share of small firms in the retail sector, non-linearities are far less marked (Fig. 4), therefore we specify a linear model for this variable, not neglecting, however, to test for possible omitted non-linearities in the residuals.

To sum up the estimated model is

$$\rho_i = \delta_0 + \delta_1 LS_i + \delta_2 LS_i^2 + \delta_3 MP_i + \delta_4 AU9805_i + \delta_5 SS9603_i + \delta_6 D_{NE} + \delta_7 D_{NW} + \delta_8 D_S + \varepsilon_i \quad (3)$$

where δ_k for k=0,…, 8 are coefficients, ε_i is a stochastic error, D_j for j=NE, NW, S are regional dummies for the North East (NE), the North West (NW) and the South (S) of Italy. The control group is therefore made by provinces belonging to the Centre of Italy.

Table 3 shows estimation results. The only regressors with a significant coefficient are those concerning the competitive structure of the retail sector. The higher is the share of "mom-and-pop" stores within the retail sector and the higher is inflation persistence. Non-parametric results concerning the ratio of large stores floor over resident population are confirmed. The presence of large stores decreases inflation persistence up to a threshold level, after which they increase it due to their nonlinear effect on competitiveness explained above.

The estimated model supports the view that regional inflation persistence is positively affected by the degree of monopoly power in the retail sector, as suggested by the New Keynesian theoretical literature. On the contrary, the local labour market conditions and the sectoral composition of the regional economy do not affect regional inflation persistence.

Moreover, as showed in Table 4, the residuals of the proposed model are very well behaved. Their normality could not be rejected by a battery of tests, as well as the fact that they have zero mean. A test for spatial correlation could not reject the absence of spatial correlation in the residuals[19]. A RESET test, obtained adding to the right hand side of (3) the square of the fitted values of the model (Davidson and MacKinnon, 1999), could not detect the presence of either omitted variables or of further nonlinearities. This result is also supported by a non-parametric test after Ellison and Ellison (2000) and by further RESET tests obtained by adding to the model in the first instance the squares and the cubes of the fitted values and, then, their cubes and fourth powers, that returned respectively a p-value of 0.31 and 0.39.

4. Robustness checks

In the present section we consider a number of different robustness checks for our results concerning the link between inflation persistence and the degree of competitiveness of the retail sector. In the first place we adopt a different estimator than OLS for (1) resorting to Andrews and Chen (1994). In the second place we insert a measure of money growth in (1). Thirdly, we use both a weighted least squares (WLS) and a feasible generalized least squares (FGLS) estimators to better account for heteroskedasticity in (3). Finally, we use a different measure of inflation persistence, the cumulative impulse response function (CIRF), as dependent variable in (3).

OLS estimates of the sum of autoregressive parameters in AR models are well known to be downward biased (see for example Orcutt and Winokur, 1969 or Quenouille, 1956). We show here that our results are robust to adopting the approximately median unbiased (AMU) estimator proposed by Andrews and Chen (1994) for autoregressive models.

The method devised by Andrews and Chen (1994) builds on Andrews (1993) for AR(1), which is as follows. Given the OLS estimator of ρ, $\hat{\rho}$, whose median function is $m(\cdot)$, a median unbiased estimator of ρ is:

[19] Testing for spatial correlation in the residuals is important because its presence might induce biased standard errors and unreliable statistical inference. We used as weight matrix a contiguity matrix whose elements are equal to one for bordering provinces and zero otherwise.

$$\tilde{\rho} = \begin{cases} 1, & if \;\; \hat{\rho} > m(1) \\ m^{-1}(\hat{\rho}), & if \;\; m(-1) < \hat{\rho} \le m(1) \\ -1, & otherwise \end{cases} \tag{4}$$

where $m^{-1}(\cdot)$ is the inverse of $m(\cdot)$ and $m(-1) = \lim_{\rho \to -1} m(\rho)$. The median of $\hat{\rho}$ usually is numerically evaluated on a fine grid of ρ values and interpolation is used to obtain both $m(\cdot)$ and $m^{-1}(\cdot)$. In a similar fashion it is possible to compute the 5th and the 95th quantiles of $\hat{\rho}$ and to build a confidence interval of $\tilde{\rho}$.

Andrews and Chen (1994) extended this estimator to AR(p) models, where p is the number of lags. Consider the Augmented Dickey-Fuller regression form of an AR(p) model for inflation in the province *i*:

$$\pi_{it} = \mu_i + \delta_i t + \rho_i \pi_{it-1} + \psi_{i1} \Delta\pi_{it-1} + \cdots + \psi_{ip-1} \Delta\pi_{it-p+1} + v_{it}$$

where *t* is a time trend, μ_i, δ_i, and $\psi_{i.}$ are parameters, Δ is the difference operator and v_{it} is a disturbance. It is possible to show that the distribution of $\hat{\rho}_i$ is independent from μ_i, δ_i and from the variance of v_{it} but not from $\psi_{i.}$. Therefore, an iterative procedure is suggested. First compute the OLS estimates of $(\mu_i, \delta_i, \rho_i, \psi_{i1}, \ldots, \psi_{ip-1})$, calling them $(\hat{\mu}_{LSi}, \hat{\delta}_{LSi}, \hat{\rho}_{LSi}, \hat{\psi}_{LSi1}, \ldots, \hat{\psi}_{LSip-1})$. Second, treat $(\hat{\psi}_{LSi1}, \ldots, \hat{\psi}_{LSip-1})$ as if they were the true values of $(\psi_{i1}, \ldots, \psi_{ip-1})$ and compute the bias corrected estimator of ρ, $\tilde{\rho}_{U1i}$, using (4) and switching in the simulations $\hat{\mu}_{LSi}$ and $\hat{\delta}_{LSi}$ to zero and the variance of v_{it} to 1. Third, treat $\tilde{\rho}_{U1i}$ as if it was the true value of ρ and compute a second round of estimates of $(\psi_{i1}, \ldots, \psi_{ip-1})$. Iterate between the estimates of ρ and $(\psi_{i1}, \ldots, \psi_{ip-1})$ until convergence or for a fixed number of times.

When implementing the Andrews and Chen (1994) estimator, we simulated the distribution of the OLS by generating 1000 random samples, we used a number of lags equal to that chosen by the Schwartz criterion for the OLS estimates and we fixed the maximum number of iterations to 10. We also included in our model a set of seasonal dummies.

Our results are set out in Tables 5 and 6. The first table shows the OLS and the AMU estimates of inflation persistence, together with the 95% confidence intervals of the latter ones. As it is possible to expect considering Table 2 in Andrews and Chen

(1994), the AMU estimates have a larger mean, as they are less downward biased than the OLS ones[20]. However, they have a greater dispersion too[21]. Once using the AMU estimates instead of the OLS ones in (3), econometric results do not change much (Table 6).

Table 7 sets out our results once inserting money growth and its lags into (1). This step was taken in order to account for possible common trends originating from monetary policy. We use as measure of money growth the quarter on quarter logarithmic growth rate of the sum of the liabilities of the monetary and financial institutions present in Italy and of all the items included in M3. The data can be downloaded from the website of the Bank of Italy. We start with an autoregressive distributed lag model with 4 lags in the inflation rate and 3 in the money growth rate, decreasing afterwards the number of lags and using a Schwartz criterion to select the most suitable specification. As it is possible to see our results are robust, as the sign and the significance of the explanatory variables included in (3) are not substantially altered.

In our third robustness check we further consider that ρ_i is an estimated dependent variable in (3), which might induce heteroskedasticity. Using robust standard errors is already a possible fix for this problem. However, we follow Lewis and Linzer (2005)[22] and we compute a WLS and a FGLS estimators as well. In the first case we use as weights the standard errors of the OLS estimates of ρ_i obtained in (1), in the second case the square roots of the linear projections of the squares of the estimated residuals of (3) on the variances of the OLS estimates of ρ_i. Our results are stable.

Finally, our results are not altered by using the CIRF as a measure of inflation persistence instead of $\hat{\rho}_i$[23], where $CIRF_i = \dfrac{1}{1-\hat{\rho}_i}$. The last column of Table 8 shows econometric results of (3) once using CIRF as dependent variable instead of $\hat{\rho}_i$ and adopting the WLS estimator illustrated above. In this case, *LS*, LS^2, *MP* and *AU9805* have all significant coefficients at the 95% level. So we find also some evidence that a high unemployment rate might have an impact on inflation persistence by eroding the bargaining power of insiders.

[20] The two means are respectively 0.27 and 0.25 for AMU and OLS estimates.
[21] The two standard deviations are respectively 0.26 and 0.24 for AMU and OLS estimates.
[22] p. 353.
[23] For a review of various measures of inflation persistence see Robalo Marques (2004).

5. Conclusions

In this contribution we assessed whether there exist statistically significant differences in inflation persistence at the regional level in a country with sizeable regional disparities like Italy. First, we document that inflation persistence is statistically different across Italian provinces, and that economically backward regions have greater inflation persistence. Nonetheless, the dispersion in inflation persistence across provinces does not show up in the aggregate inflation rate. In other words, there is not a geographical aggregation bias, since the inflation persistence in the national index is equal to the average of the inflation persistences in the provincial indexes. Moreover, we provide an empirical assessment of the suggestion of recent theoretical contributions (Benigno, 2004), about the need for monetary policy to take into consideration dispersion in regional inflation persistence, supplementing traditional inflation indicators, with one that weighs regional inflation time series according to their persistence. Our dataset does not support such policy prescription. Finally, another important result of our analysis is that different levels of regional inflation persistence are associated with different local degrees of competitiveness in the retail sector. Therefore, if a policymaker would like to decrease the persistence of inflation in a given province should try to decrease the market power of local retailers.

Contrary to sectoral inflation persistence, dispersion in regional inflation persistence is an issue that the literature has somewhat overlooked. We hope that the results in this paper push future research in this area.

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Table 1 - Model specification tests - based on regional inflation time series at quarterly frequency (Total number of series: 70)

Wald test on the sum of the autoregressive coefficients to be equal to zero[†] °	0.41	**Wald test on a structural change in the intercept (2002Q1)[‡] °**	0.89
Wald test on the sum of the autoregressive coefficients to be equal to one[†] °	0.03	**One sided F-test on a structural change in the standard deviation of inflation (1999Q1)[‡] °**	0.96
Wald test on a structural change in sum of the autoregressive coefficients (1999Q1)[‡] °	0.87	**One sided F-test on a structural change in the standard deviation of inflation (2002Q1)[‡] °**	0.91
Wald test on a structural change in the sum of the autoregressive coefficients (2002Q1)[‡] °	0.76	**Wald test for poolability of all the provinces (restricted sample) - p-value[♣] °**	0.00
Wald test on a structural change in the intercept (1999Q1)[‡] °	0.90	**Durbin's *m* test for serial correlation[‡ ‡]**	1
		RESET test for model misspecification[‡ ‡ ♣]	0.99

°: tests based on heteroskedasticity consistent standard errors.

[†]: frequency of accepting the null at the 5% level, in percent of the total number of time series

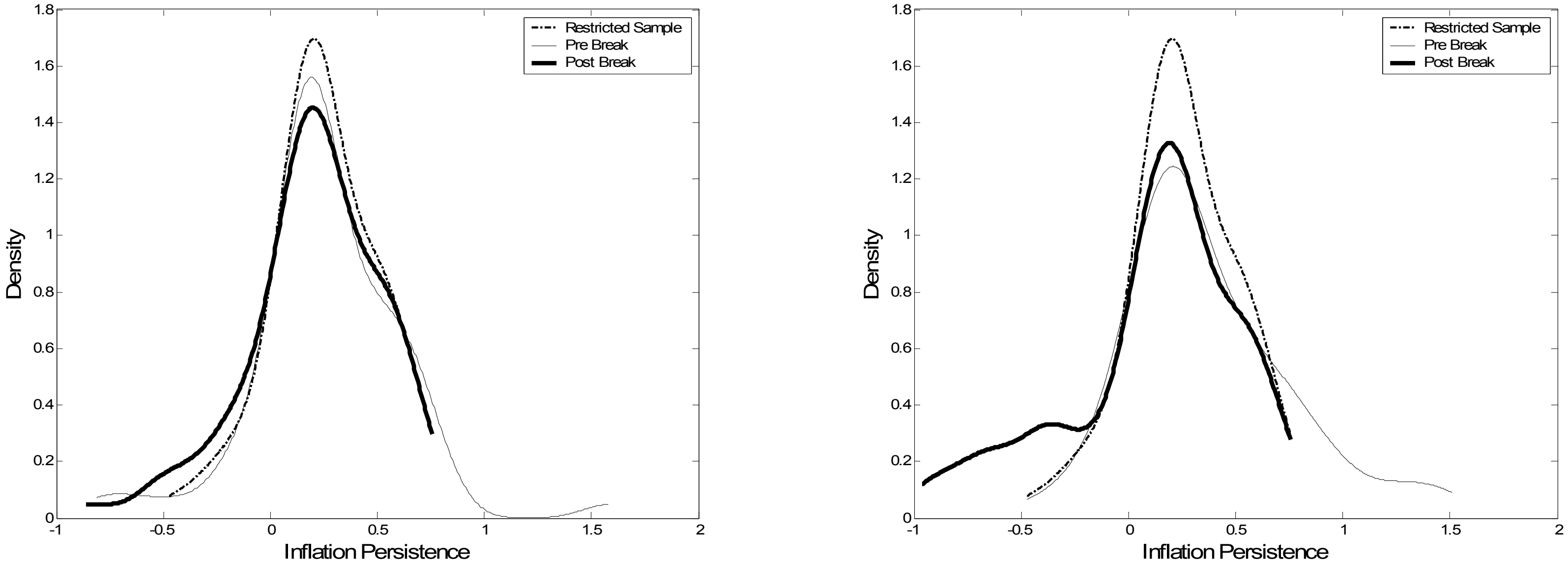
[‡]: frequency of accepting the null of no break at the 5% level, in percent of the total number of time series analysed.

[♣]: the test is distributed as a chi squared with 69 degrees of freedom. The null is that all the provinces can be pooled, namely that each province has the sum of autoregressive coefficients equal to that of Vercelli in the North-West of Italy (Piemonte region). The restricted sample is the one imposing absence of structural breaks.

[‡ ‡] : frequency of accepting the null at the 5% level of no serial correlation in the residuals, in percent of the total number of time series analysed.

[‡ ‡ ♣]: frequency of accepting the null at the 5% level of no model misspecification, in percent of the total number of series analysed. The test was obtained by adding to the model the squares and the cubes of the regressors and testing for the significance of their coefficients.

Figure 1 – Regional Distribution of Inflation Persistence – General Consumer Price Index



Break in 1999Q1 **Break in 2002Q1**

Note: kernel density estimators with Silverman's optimal smoothing bandwidth. The inflation persistence measure is the sum of the coefficients of autoregressive processes estimated for the Italian provinces included in the sample. The number of lags was selected thanks to a Schwartz criterion. For the reasons underlying the choice of this persistence measure see the body of the paper.

Figure 2. Annualized inflation using standard CPI and a persistence-population weighted CPI

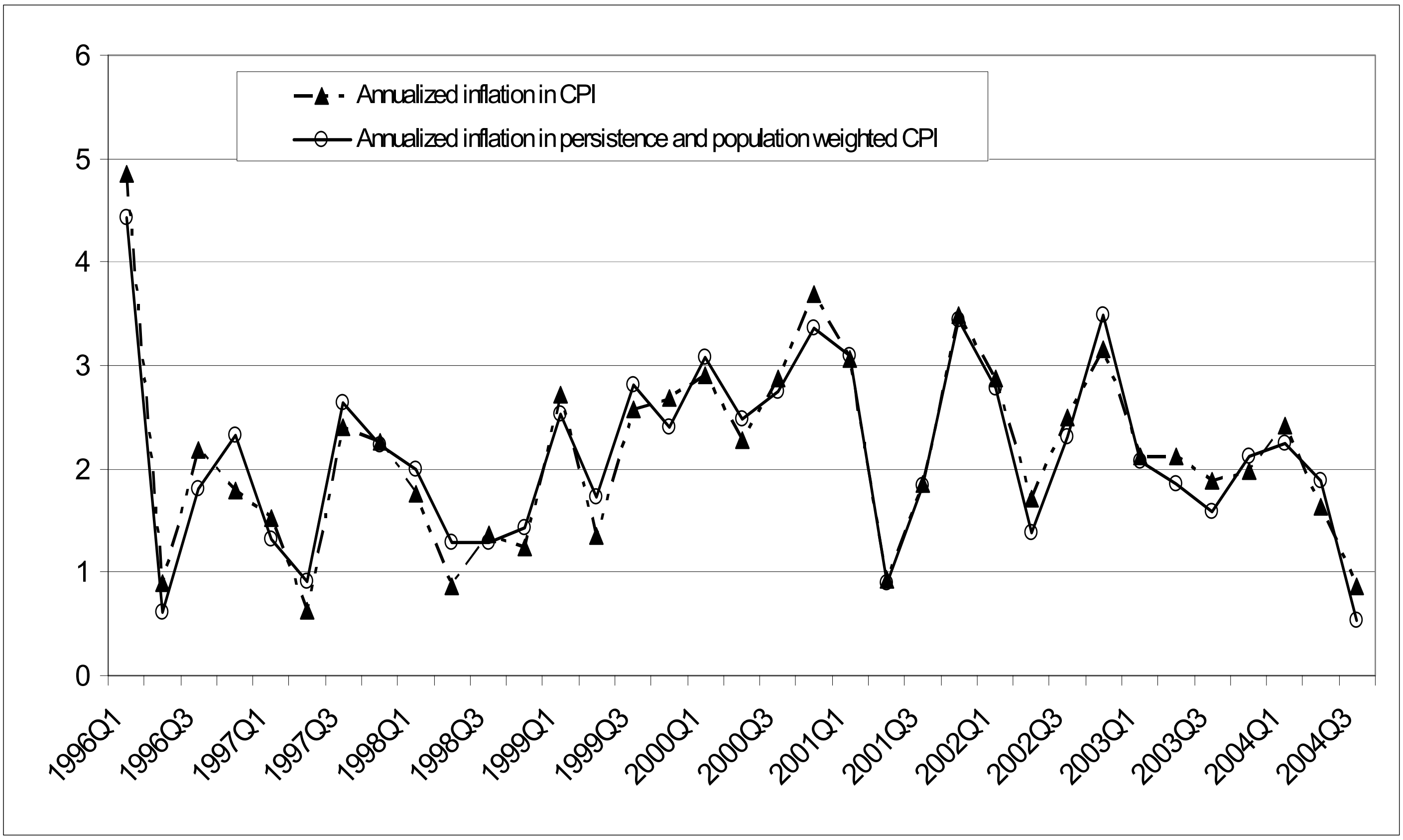
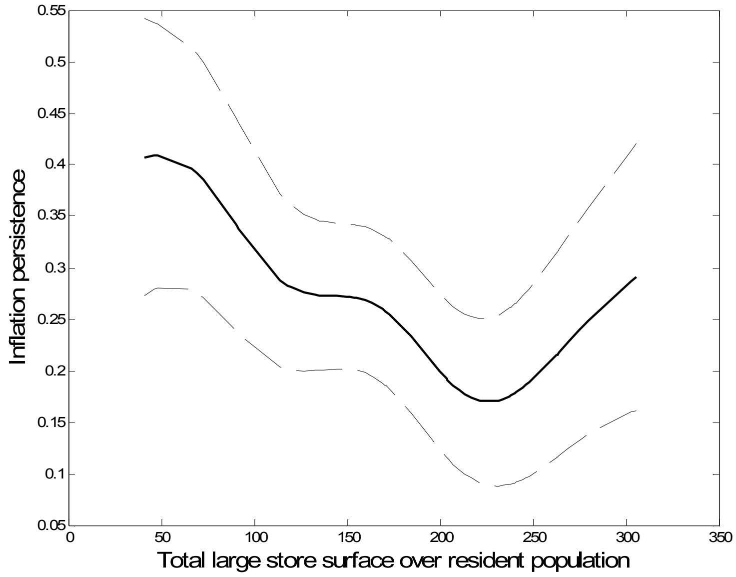
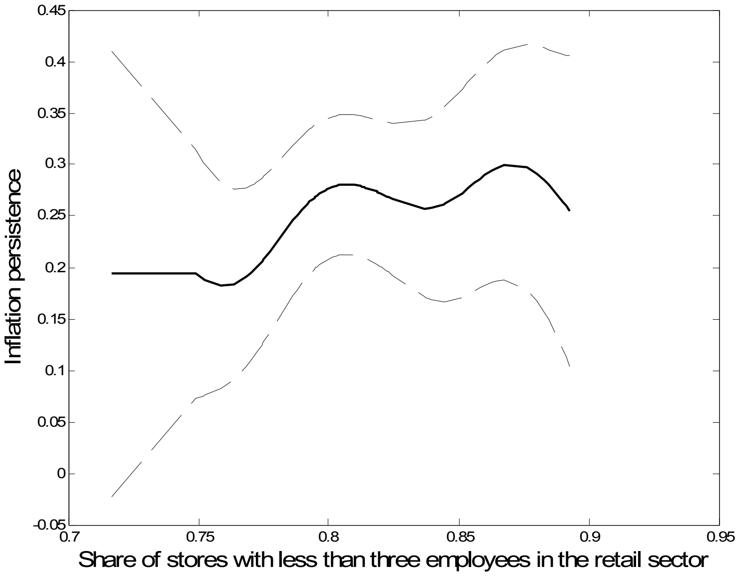


Figure 3 – Nonparametric estimator of the relationship between regional inflation persistence and total large store surface over resident population in 1999.



Note: for this picture we resorted to a Gaussian kernel estimator with Silverman's optimal smoothing bandwidth.

Figure 4 – Nonparametric estimator of the relationship between regional inflation persistence and the share of stores with less than 3 employees in 2001.



Note: for this picture we resorted to a Gaussian kernel estimator with Silverman's optimal smoothing bandwidth.

Table 2 - Descriptive statistics of the regressors				
	Mean	**Std. Dev.**	**Min.**	**Max.**
Total large store surface in 1999 (squared meters per thousands of inhabitants)	182.13	65.13	41.13	305.65
Share of retail stores with less than 3 employees in 2001	0.81	0.05	0.65	0.89
Average unemployment rate between 1998 and 2005	8.18	5.58	2.36	25.04
Share of service activities in total local value added between 1996 and 2003	0.71	0.08	0.55	0.90

Table 3 - The determinants of regional inflation persistence
Dependent variable: inflation persistence. Estimation method: Ordinary Least Squares with Robust Standard Errors. Observations: 70.

	Coefficients	**t-statistics**
Total large store surface over resident population in 1999	-0.005*	-2.45
(Total large store surface over resident population in 1999)^2	0.001*	2.37
Share of retail stores with less than 3 employees in 2001	0.905*	2.39
Average unemployment rate between 1998 and 2005	-0.009	-0.78
Share of service activities in total local value added between 1996 and 2003	0.202	0.57
North East Italy†	-0.070	-1.05
North West Italy†	-0.102	-1.18
South and Islands†	-0.039	-0.23

Following Eisenhauer (2003), the constant was dropped because it was not significantly different from zero at a 5% level.
†: dummy variables. The control group is constituted by provinces belonging to Central Italy.
*: significant at a 5% level.

Table 4 - Tests on the residuals of the model of the determinants of regional inflation persistence (Table 2)

Normality tests	
Shapiro - Francia (p-value)^{o}	0.42
Shapiro - Wilk (p-value)^{o}	0.54
Skeweness - Kurtosis (p-value)^{o}	0.59
RESET test for omitted variables (p-value)†	0.58
Test for zero mean residuals (p-value)††	0.97
Test for spatial correlation$^{\dagger\dagger o}$	0.84
Ellison and Ellison (p-value)$^{\dagger\dagger oo}$	0.81
Observations	70

o: the null is that residuals have a normal distribution.
†: the null is absence of omitted variables.
††: the null is that residuals have a zero mean.
††o: 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 is absence of spatial correlation. For an introduction to this test see Anselin (1988).
††oo: the null is that the model fits the data well in terms of functional specification and omitted variables. We used a quartic kernel function as in Miles and Mora (2003) and Silverman's optimal smoothing bandwidth as window width. The test has an asymptotic standard normal distribution.

Table 5 - OLS and approximately median unbiased estimates of inflation persistence

	Observations	Number of lags of AR model	OLS estimate of ρ	Approx. median unbiased estimates of ρ	95% confidence interval of the approx. median unbiased estimates of ρ	
Alessandria	43	2	0.219	0.232	-0.093	0.499
Ancona	43	2	0.471	0.505	0.200	0.707
Aosta	43	1	0.210	0.216	-0.044	0.433
Aquila	43	3	0.584	0.619	0.242	0.807
Arezzo	43	3	0.616	0.661	0.286	0.837
Ascoli Piceno	43	4	0.477	0.525	0.034	0.762
Asti	43	2	0.030	0.042	-0.302	0.325
Bari	43	3	0.673	0.707	0.402	0.861
Belluno	43	1	0.198	0.198	-0.043	0.426
Bologna	43	4	-0.065	-0.063	-0.561	0.337
Bolzano	43	3	0.424	0.449	0.024	0.706
Brindisi	43	4	0.655	0.708	0.245	0.890
Cagliari	43	1	0.129	0.130	-0.114	0.358
Campobasso	43	2	-0.162	-0.152	-0.521	0.177
Catania	43	1	0.279	0.290	0.027	0.508
Chieti	43	3	-0.270	-0.297	-0.853	0.168
Como	43	1	0.171	0.179	-0.068	0.406
Cosenza	43	1	0.515	0.530	0.260	0.700
Cremona	43	1	0.224	0.224	-0.044	0.453
Cuneo	43	3	0.087	0.085	-0.405	0.447
Ferrara	43	2	0.075	0.070	-0.269	0.348
Firenze	43	2	0.042	0.051	-0.303	0.347
Foggia	43	1	0.148	0.150	-0.103	0.385
Forli	43	1	0.281	0.294	0.038	0.509
Genova	43	3	0.446	0.451	0.013	0.704
Grosseto	43	1	0.112	0.112	-0.132	0.344
Latina	36	2	0.532	0.561	0.270	0.743
Livorno	36	2	0.199	0.213	-0.140	0.489
Lucca	43	1	0.100	0.099	-0.157	0.357
Macerata	40	1	-0.128	-0.133	-0.370	0.132
Mantova	43	3	0.460	0.512	0.046	0.745
Milano	43	2	0.062	0.087	-0.284	0.366
Modena	43	3	0.575	0.598	0.219	0.797

continues

continued

Napoli	43	1	0.236	0.248	-0.022	0.485
Novara	43	1	0.290	0.304	0.025	0.526
Padova	43	2	0.298	0.324	-0.003	0.566
Palermo	43	3	0.452	0.469	0.087	0.710
Parma	43	3	0.509	0.558	0.118	0.781
Pavia	43	2	-0.137	-0.127	-0.528	0.222
Perugia	43	1	0.273	0.284	0.034	0.503
Pesaro	43	3	0.400	0.414	-0.010	0.667
Pescara	43	4	0.704	0.804	0.286	0.967
Piacenza	43	1	0.031	0.030	-0.212	0.264
Pisa	43	2	0.262	0.278	-0.053	0.526
Pistoia	43	3	0.215	0.221	-0.256	0.539
Pordenone	43	1	0.263	0.276	0.015	0.485
Potenza	43	3	0.601	0.648	0.221	0.845
Ravenna	43	1	0.105	0.107	-0.149	0.333
Reggio Emilia	43	1	0.106	0.116	-0.136	0.351
Roma	43	3	0.199	0.200	-0.282	0.513
Rovigo	43	1	0.253	0.263	0.012	0.480
Sassari	36	2	-0.301	-0.307	-0.637	0.030
Savona	43	1	0.086	0.086	-0.156	0.342
Siena	43	3	0.351	0.378	-0.104	0.656
Siracusa	43	2	0.167	0.181	-0.126	0.444
Sondrio	40	4	0.204	0.207	-0.315	0.549
Spezia	43	2	0.621	0.669	0.341	0.826
Teramo	36	1	0.223	0.237	-0.039	0.489
Terni	43	4	0.482	0.510	0.079	0.747
Torino	43	1	0.297	0.299	0.022	0.503
Trapani	43	3	0.475	0.523	0.092	0.749
Trento	43	1	0.202	0.193	-0.069	0.434
Treviso	43	3	0.110	0.112	-0.355	0.458
Trieste	43	4	0.305	0.310	-0.117	0.580
Udine	43	1	0.234	0.246	-0.019	0.469
Varese	43	4	-0.475	-0.488	-1.084	0.017
Venezia	43	1	0.179	0.184	-0.071	0.397
Vercelli	36	4	0.757	0.869	0.273	1.000
Verona	43	1	0.074	0.070	-0.192	0.309
Viterbo	43	4	0.331	0.365	-0.140	0.655

Table 6 - The determinants of regional inflation persistence

Dependent variable: approximately median unbiased estimates of inflation persistence. Estimation method: Ordinary Least Squares with Robust Standard Errors. Observations: 70.

	Coefficients	t-statistics
Total large store surface over resident population in 1999	-0.005*	-2.41
(Total large store surface over resident population in 1999)^2	0.001*	2.33
Share of retail stores with less than 3 employees in 2001	0.968*	2.39
Average unemployment rate between 1998 and 2005	-0.011	-0.84
Share of service activities in total local value added between 1996 and 2003	0.209	0.55
North East Italy†	-0.081	-1.14
North West Italy†	-0.106	-1.14
South and Islands†	-0.027	-0.14

Following Eisenhauer (2003), the constant was dropped because it was not significantly different from zero at a 5% level.

†: dummy variables. The control group is constituted by provinces belonging to Central Italy.

*: significant at a 5% level.

Table 7 - The determinants of regional inflation persistence

Dependent variable: estimates of inflation persistence in an ARDL model including the money growth rate. Estimation method: Ordinary Least Squares with Robust Standard Errors. Observations: 70.

	Coefficients	t-statistics
Total large store surface over resident population in 1999	-0.005*	-2.40
(Total large store surface over resident population in 1999)^2	0.001*	2.26
Share of retail stores with less than 3 employees in 2001	0.882*	2.50
Average unemployment rate between 1998 and 2005	-0.007	-0.63
Share of service activities in total local value added between 1996 and 2003	0.119	0.37
North East Italy†	-0.011	-0.17
North West Italy†	0.019	0.28
South and Islands†	-0.024	-0.14

Following Eisenhauer (2003), the constant was dropped because it was not significantly different from zero at a 5% level.

†: dummy variables. The control group is constituted by provinces belonging to Central Italy.

*: significant at a 5% level.

Table 8 - WLS and FGLS estimates of the determinants of regional inflation persistence

Observations: 70

Dependent variable	Sum of AR coefficients				CIRF	
Estimation method	Weighted Least Squares°		FGLS°°		Weighted Least Squares°	
Explanatory variables	**Coefficients**	**t-statistics**	**Coefficients**	**t-statistics**	**Coefficients**	**t-statistics**
Total large store surface over resident population in 1999	-0.003*	-2.99	-0.004*	-2.14	-0.009*	-7.81
(Total large store surface over resident population in 1999)^2	0.001*	3.11	0.001*	2.01	0.001*	7.53
Share of retail stores with less than 3 employees in 2001	0.562*	3.61	0.714*	2.17	0.862*	3.38
Average unemployment rate between 1998 and 2005	-0.004	-0.83	-0.007	-0.79	-0.023*	-4.93
Share of service activities in total local value added between 1996 and 2003	0.175	1.09	0.245	0.75	0.300	1.45
North East Italy†	-0.061	-1.60	-0.068	-1.01	-0.221*	-5.68
North West Italy†	-0.064	-1.54	-0.079	-1.08	-0.209*	-4.83
South and Islands†	-0.026	-0.45	-0.025	-0.23	0.070	1.19
Error variance attributable to sampling error in the dependent variable			0.034			
Remaining error variance			0.013			

Following Eisenhauer (2003), the constant was dropped because it was not significantly different from zero at a 5% level, with the exception of the CIRF model whose estimated constant was equal to 1.63 with a t-statistic of 5.37. CIRF is the acronym for cumulative impulse response function, which is equal to the inverse of 1 minus the sum of the autoregressive coefficients in the AR model for inflation.

°: 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.

†: dummy variables. The control group is constituted by provinces belonging to Central Italy.

*: significant at the 5% level.