PRICING BEHAVIOR AND THE INTRODUCTION OF THE EURO: EVIDENCE FROM A PANEL OF RESTAURANTS

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Abstract

This paper assembles an original panel of data from 2,500 restaurants in Italy over 1998-2004, with the objective of studying whether the euro cash changeover had an impact on individual pricing behavior, as apparently perceived by consumers, and which economic mechanisms may explain it.

On the first point, the data show that only about a price increase of 3-4 percentage points can be attributed to the new currency; the changeover focussed the public attention over a medium-run trend, prompting the attribution of the whole increase to the introduction of the euro.

On the second point, we reach two conclusions. We find evidence consistent with the existence of "menu-costs": during the changeover the rise in the average meal price is mainly due to a larger fraction of agents who simultaneously revise their price. We also find that during the changeover more market power (proxied by a index of concentration on local markets) was associated with larger price increases; we propose a simple interpretation based on consumer behavior which may also explain why the effects of the cash changeover were especially pronounced in this industry as opposed to more competitive ones.

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1. Introduction

The introduction of the euro banknotes and coins (cash changeover), which occurred during the early months of 2002, was followed by a heated debate on the alleged inflationary effect that such a renomination of the unit of account supposedly exerted on the price level in

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several Euro-area countries, as well as on the reasons underlying such an effect.[1] Since 2002, the dynamics of consumers' "perceived inflation" (based on qualitative judgements) was systematically greater than that of actual inflation; in Italy, such a phenomenon was more marked than in other countries (Figure 1).

The Italian media and consumers' associations frequently reported about extraordinary price increases. Among those, the ones recorded by restaurants gave rise to great controversies. According to a survey conducted in an Italian region, one third of citizens blamed restaurants for excessive increases, a percentage second only to the share of respondents who blamed food prices.[2] Various newspaper articles reported anecdotal evidence allegedly showing that the cost a meal in a restaurant or a *pizzeria* had increased by 80-100% after the introduction of the euro.[3] It became almost common wisdom to argue that price setters took advantage of the psychological conversion of 1,000 lire into 1 euro, which would basically amount to a doubling of the price:[4] "everybody knows that one euro is now worth 1,000, not 2,000, lire",[5] an opinion sometimes also put forward by members of the administration.

After the introduction of the euro, in Italy and in most euro area countries, restaurant prices were indeed among the fastest growing items in the Harmonized Index of Consumer Prices (HICP); however, the size of the increase is in contrast with the perceptions voiced in the public discussions. Inflation in the "restaurant and cafe's" sector showed a blip in January 2002 (a month-on-month increase of 0.8 percent) and reached a 12-month growth rate of 5 percent at the end of the same year, compared to previous figures ranging between 2 and 3 percent (Figure 2).

Such facts raise two distinct questions. First, it is of interest to understand what happened at a more disaggregated level, which may shed light on why consumers' perceptions about inflation diverged so markedly from the official measures. Preliminary

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[1] See European Central Bank (2003a, 2003b) and Deutsche Bundesbank, (2004).

[2] EU.R.E.S., *Rapporto 2002 sullo stato delle provincie del Lazio*.

[3] E.g., *La Repubblica*, 23/8/2003 and 24/8/2003.

[4] The official conversion rate was 1936.21 lire per 1 euro.

[5] *La Repubblica*, 21/8/2003.

investigations into this problem have conjectured that consumers' inflation expectations may have been particularly affected by rise in the price of frequently purchased goods, and/or by the detection of a greater-than-usual fraction of rising prices, whose impact on the general price level would however be limited.[6]

The second question concerns the identification of the economic mechanism which may have favored increases of some prices in the face of an event, such as the change in the unit of account, which was widely expected to be "neutral". Two candidate explanations have been advanced so far. The first one emphasizes the presence of collusion and/or lack of competition. This is the prevailing interpretation in the popular debate and in the press; for restaurant prices, a more formal version of this view, based on collusive behavior and multiple equilibria is advanced by Adriani *et al.* (2003). The second set of explanations stresses the existence of sticky prices and small menu costs: the changeover is viewed as inducing more firms than usual to review their prices simultaneously, which may have short-run inflationary consequences. Hobjin *et al.* (2004) propose a model along these lines and use it to study the dynamics of restaurant prices after the changeover, based on HICP figures.

Our strategy is to address these issues by making extensive recourse to micro-data. We begin by assembling an original panel of data from 2,500 restaurants, *pizzerie* and the like in Italy over 1998-2004 using data from a leading guide to Italian restaurants. The sample includes well-known establishments and is not designed to be representative of the whole sector. Our interest stems from the possibility of gaining deeper insights from the individual data. In particular, the individual data allow us to study whether there are features of the distribution of individual price changes in the year 2002 (e. g. the frequency of price revisions or the number of 'exceptionally' large increases) which may contribute to explain the widespread perception of a large effect of the introduction of the euro on prices.[7]

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[6] See Del Giovane and Sabbatini (2004) for an analysis of the Italian case.

[7] Note that disaggregated individual data are essential to analyze these questions. A related study based on disaggregated data is offered by Adriani *et al.* (2003), who use data from six European countries for the years 2002 and 2003. These authors find a greater than average increase in the prices of the euro-area countries, interpreting such effect as a consequence of the changeover. Their analysis, however, only measures the increase in prices of a single year, for a limited number of restaurants (155 for Italy). We believe that our analysis goes a step further by significantly extending the number of restaurants (2500) and years (seven) included in the sample, which allows us to measure the effect of the changeover in comparison to a historical

The descriptive evidence drawn from these data opens the way to the analysis of the two above-mentioned mechanisms capable of explaining a changeover-related increase in prices. First, we study whether the observed price setting behavior provides any support to the "menu cost" hypothesis, which suggests that the year 2002 is a special one as the change in the unit of account invites all agents to revise their prices. We analyze the distribution of individual price increases, asking whether the average price increase and the frequency of price revisions in the year of the cash changeover were larger than "normal".

Next, we analyze whether the price increases recorded during the changeover relate to the local degree of competition, on the basis of a pricing equation that controls for several determinants of restaurant prices. We argue that if the changeover made prices less transparent to consumers, producers might have exploited such confusion to raise prices, at least temporarily. Such an incentive is greater the less elastic the demand faced by the consumer, as this ensures the producer a smaller fall in sales in case the consumer realizes the true value of the new price.

Both exercises provide a first quantification of the magnitude of, respectively, the "menu-cost" and the "market power" channel. The analysis is completed by the presentation of a formal model in which both effects appear simultaneously. The model is calibrated on historical data and then used to simulate the effect of the changeover on prices.

The paper is organized as follows. The next section describes the database: an open panel of about 2500 establishments (including restaurants, "*pizzerie*" and "*trattorie*") for the years 1998-2004, a total of about 17.000 observations which were purposely assembled for this paper. The main facts affecting demand and costs in this industry over the period and the key descriptive statistics on the distribution of individual price changes are presented in Section 3. Section 4 discusses the frequency and size of price revisions, a decomposition which is helpful in identifying the source of the price increase and the relevance of the "menu cost" assumption. An econometric analysis of the role of "market power" in the setting of prices is in Section 5. Section 6 formally analyzes the role of the "menu cost" and the "market power" hypotheses by means of a simple price setting model, attempting to

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benchmark.

quantify the relative importance of these two channels. A concluding section summarizes the main findings.

2. A description of the database

Restaurant prices are collected from the yearly publications *Guida dei ristoranti d'Italia* and *Roma*; both are leading guides to Italian restaurants, wine-bars, *trattorie* and *pizzerie*.[8]

The data report the price of a "full meal".[9] The price records are expressed in lire until 2001 (rounded to the nearest 5,000 lire, approximately 2.6 euro), in euro since 2002 (rounded to the nearest euro). For the sole "restaurant" category, the guide also assigns a score, on a 0-100 scale, concerning its quality; it includes information about each establishment, such as location or category. Data are annual, collected each year during the first semester; the guide is published towards the end of September. The data span the 1998-2004 period.

The panel is unbalanced, since the list of surveyed establishments is revised each year. On average, 85 per cent of the establishments surveyed in a given year also appear in the guide the following year. This subset allows us to compute price increases. We converted the lira prices in euro, rounding them to the nearest euro. In order to avoid the results to be affected by the rounding conventions used by the guide in each year, in computing price increases we set to zero all price changes smaller than 1 euro in absolute value.[10]

Tables A1-A5 in appendix A provide a detailed description of the sample composition. The number of establishments surveyed in each year hovers between 1,800 and 2,500 observations, for a total of 17,000 observations (Table A1). Due to sample turnover, the

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[8] The guides are edited by the company *Il Gambero Rosso*. The second guide is specifically about the Rome area and contains more detailed information, such as the location of the various restaurants in the city (e.g. Vatican, city centre vs. suburbs, etc).

[9] This includes appetizer, first and second course and a dessert (beverages are not included).

[10] This adjustment operates only on 2002, 2003 and 2004 data and it has negligible effects (a few basis points) on most results reported in the paper. The only exception is the percentage of unchanged prices in 2002, 2003 and 2004 which is discussed at greater length in section 5.

number of price changes computed in each year is smaller, between 1,600 and 2,300 per year, for a total of 12,300 observations.

The data cover the entire national territory. The distribution by geographical area (Table A2) reveals a slight under-representation of the establishments located in the South (23 per cent of the total, which compares with a 32 per cent computed from the registry of the Chamber of Commerce).[11]

As to the type of establishments, the majority is represented by restaurants (about 70 per cent; see Table A3); the remaining part consists of *trattorie*, *pizzerie* and wine bars. Overall, the surveyed establishments cover approximately 3.5 per cent of the total number of businesses recorded by the Chamber of Commerce (Table A4).

Assuming that the surveyed establishments are scarcely substitutable with lower-quality and less visible dining places, we constructed an indicator of the degree of market competition at the local (province) level. A first measure is given by the ratio between the number of quality dining places that are present in a given province, proxied by the number of establishments reported in the guide, and resident population. This indicator, whose averages are reported in the first column of Table A5 for the Italian *Regioni*,[12] suggests a greater degree of competition in the North-east regions, a smaller one in the South. By overlooking the presence of tourists, however, the indicator might overestimate competition in the provinces where tourism is highly relevant. To this end, a second measure is constructed as the ratio between the number of quality dining places and a weighted sum of local population and foreign and domestic tourist presence.[13] The indicator, reported in the second column of Table A5, broadly confirms the previous picture, with the notable exception of a few small regions which attract considerable tourism flows.

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[11] The regional distribution of establishments, for the year 1999, is drawn from *Rapporto sul turismo italiano 2001*, page 47, tav.7.

[12] In Italy there are 20 regions and 103 provinces.

[13] Statistics on tourist presence in hotels and rented apartments in each province are from Chambers of commerce data; we used data referring to the year 1998. Population and tourist presence are not directly comparable: the former is measured in heads, the latter in days of presence; also, tourists are likely to go to the restaurants more often than local resident. Assuming local population eat out once a month (an assumption broadly consistent with indications from the Istat survey on household consumption), we multiplied population by 12 to make the two magnitudes comparable.

Our sample was obviously not designed with the aim of being representative of the whole restoration industry and should not be considered as such. As discussed in the next section, the rise in the price index in the whole sector is smaller than in our high-quality establishments. One possible explanation for the divergence is due to the fact that these belong to a segment facing a less elastic demand (the implications of this hypothesis is discussed in detail in Sections 5 and 6). More generally, the demand for quality food increased significantly in Italy in the last few years, as witnessed by the growing diffusion of publications and guides on the subject, possibly commanding an increasing premium on quality restaurants.

Table A6 reports several descriptive statistics on the price of a meal from 1998 to 2004, sorted by year and establishment type. The average price of a meal in a restaurant varies from 33 euro in 1998 to 49 euro in 2004; in wine bars, *pizzerie* and *trattorie* it increases from 20 to 27 euro. The price dispersion (standard deviation) increases in time.

3. Demand and costs conditions and the dynamics of meal prices over 1999-2004

In the period under examination, both demand and supply conditions in the industry were subject to shocks that should be considered when examining the data. Between the second half of 2000 and the first half of 2001, a strong rise in the demand for meal services took place, related to the spike of tourism sparked by the Jubilaeum of the Catholic church in the year 2000. Expenses by foreign tourists (Figure 3) then decreased sharply following the terrorist attacks of September 11, 2001.[14] The growth of households' expenditure (at constant prices) in the "restaurant and hotel services" sector also reached a peak in 2000; employment and value added in this industry grew at around 8 per cent.

On the costs side, substantial increases in the price of unprocessed food took place in 2001-2003, due to exceptional events, as the spreading of first BSE ("mad cow") and then foot-and-mouth diseases in 2000-2001, or adverse climatic conditions at the beginning of 2002. According to Eurostat data (Figure 4), the agricultural production prices increased rapidly in 2001, at a rate above 5 percent; they slowed down, but still kept increasing, in the

[14] See a discussion in *"Rapporto sul Turismo Italiano" 2002*, p.3.

following years. Fresh food consumer prices followed a similar, although more persistent, pattern; in this case, growth was above 4 percent until 2003. Unit labour costs in the "hotels and restaurants" sector increased substantially in 2001 and, notably, 2002 and 2003. All in all, the annual growth in costs can be roughly estimated to have been between 4 and 6 percent in 2001-2003.

The main statistics on price increases in our sample are reported in Table 1. Two facts emerge from the data: in the year of the euro changeover the average increase is sizable (9.3 per cent), but it does not represent a peak (which is located in the year 2001, at 10,5 per cent). This holds irrespective of the type of establishments under investigation (Table A7 in Appendix A). A sample split by geographical areas (North, center and South) shows that the year of the changeover records a peak increase in prices only in the South (Table A8).

A similar picture emerges from other descriptive statistics, such as the median, or an indicator of the largest increases (the 95th percentile or the maximum increase). In all these cases, the increases recorded in 2002 and 2003 are smaller than in the previous year. The year of the changeover is characterized by price increases that, although sizable, are not out of line with those recorded in previous years.

Table 1: The rise in the price of a meal (12-month growth)

Year	**1999**	**2000**	**2001**	**2002**	**2003**	**2004**
Sample average	4.3	6.7	10.5	9.3	5.8	3.8
5th percentile	0.0	0.0	0.0	0.0	0.0	-0.1
Median	0.0	0.0	10.0	8.7	4.9	0.0
95th percentile	25.0	30.0	33.3	29.0	20.0	20.0
Standard deviation	10.4	12.3	12.5	11.3	8.6	9.2

A comparison of the average annual price increase in our dataset with those from other studies is shown in Table A9. Our data are basically identical to those obtained from a similar sample of quality restaurants by Adriani *et al.* (2003), although their sole observation concerns 2003. In contrast, notable differences emerge between our data and the increase of the average price level for the whole sector produced by the National statical offices.

Although the time profile of the changes is similar, the rise in the price index in the whole sector is much smaller than in our high-quality establishments. This holds for all years and it is not limited to 2002.

According to the facts presented above, price increases in 2000 and 2001 may be related to increases in demand and costs, which are likely to have affected prices with a lag. Such a lag may just reflect the way our data are constructed: the prices reported in the Guide are surveyed in the first half of the year.

The trends recorded in the city of Rome, most directly connected to the Jubilaeum, confirm the conjecture that tourism played a role in the evolution of prices. In Rome (Table 2) the largest increase occurred in the year 2000 (about 12,7 per cent); in that year, the average price growth in the establishments located in the *Vatican* area and the close-by tourist district of *Trastevere* are, respectively, of 23 and 29 per cent, much greater than in the rest of the city.

Table 2: The rise in the price of a meal - selected Rome districts (12-month growth)

	1999	**2000**	**2001**	**2002**	**2003**	**2004**
Average	*6,1*	*12,7*	*10,7*	*8,6*	*3,8*	*3,8*
City Center	5,4	12,0	12,9	7,1	4,8	3,8
Vatican	10,6	23,4	8,9	3,0	0,0	5,1
Trastevere	6,4	29,3	11,6	1,3	4,8	2,1

Considering the tails of the distribution, there is simply no evidence of any doubling of prices in 2002. Table 3 shows the number of price increases larger than a given threshold (respectively, 50, 75 or 90 per cent) in each year. As in the previous case, the years 2002 and 2003 stand out for the smaller number (virtually nil) of “large” increases. A few increases greater than 50 percent can be found in 2000-2002, but not later. The conjecture that high perceived inflation was caused by large increases which hit the imagination of the consumers finds little support in the data (such increases represent less than 2 per cent of the total sample observations).

Table 3: Number of "extreme" price increases

Year	**1999**	**2000**	**2001**	**2002**	**2003**	**2004**
Increases greater than 50%	11	20	30	17	5	6
Increases greater than 75%	4	4	4	4	0	0
Increases greater than 90%	2	1	2	1	0	0
Total # observations	2239	2292	2144	2023	2028	1552

A major factor affecting the public perception, however, may have been the substantial cumulate price increase which took place over the whole 1998-2002: around 40 per cent on average, or even 75 per cent for the 10 per cent of establishments which recorded the largest price rises. These observations help to form a conjecture on the factors that may have favored the "perception" of an exceptional rise in prices among consumers; such an increase, while it actually took place, occurred more gradually during the years than perceived.

4. The frequency of price revisions

The histograms in Figure 5 show the distribution of price changes in each year. The distributions are strongly asymmetric, with few negative records. The mode is at zero. This indicates that in most years a large mass of firms leave their price unchanged, or change it by a small amount. In the years 1999, 2000 and 2004 this mass amounts to about 60 per cent of the whole population (Table 4); on average, the percentage is around 40 per cent. These data correspond to an average duration of prices of respectively 23 and 13 months[15]. In both 2001 and 2002 the percentage of unchanged prices is much smaller, reducing to below 30 percent.[16]

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[15] Under the assumption that the decision whether to revise prices is taken with a monthly frequency, duration is computed as $[1-(1-a)^{1/12}]^{-1}$, where *a* is the annual share of revised prices; a similar result is obtained by assuming continuous decisions and applying the formula reported by Bils e Klenow (2002), $[-1/\log(1-a)]*12$. According to Bils e Klenow, the average price duration for items "lunch" or "dinner" is 11 months.

[16] As mentioned, in constructing the sample we set to zero all price variations smaller than one euro in absolute value. This was done for two reasons. First, the figures reported in table 4 would otherwise not be comparable through time: before 2002, any price change smaller than one euro in absolute value would not have been recorded by the Guide due to the rounding convention described in Section 2. Second, in 2002 many

The data are consistent with the "menu cost" hypothesis: small price revisions are prevented, in normal years, by an adjustment cost.[17] From this standpoint, the changeover, and the associated renomination of all prices in the new unit of account, can be though of as an aggregate shock which requires each agent to repost its price, creating an opportunity to revise it, consequently determining an increase in average inflation.[18]

Table 4: Number of constant and falling prices

Year	**1999**	**2000**	**2001**	**2002**	**2003**	**2004**
a) # unchanged prices	1432	1309	600	542	861	874
b) # price reductions	108	72	74	90	45	109
c) # observations	2239	2292	2144	2023	2028	1552
a/c	*0.64*	*0.57*	*0.28*	*0.27*	*0.42*	*0.56*

The size of the average (as well as median) price revision (Table 5) reinforces the impression that pricing polices in years 2002 and 2003, rather than representing a discontinuity, were quite moderate in relative terms. Average increases were among the smallest since 1999 (13 and 10 per cent, respectively, versus 15 and 16 per cent in the previous 2 years).

"small" variations in prices (equal to 1 euro) would be observed, even if prices were unchanged. This is due to the new rounding convention adopted by the Guide since 2002. For example, in 2001 a "true" price of 17,500 lire would have been rounded to 20,000 lire, corresponding to 10 euro; in 2002 the same price would be rounded to 9 euro. Had we not made these adjustments, the figures in the fourth row of Table 4, would have been equal to 5%, 30% and 50% in 2002, 2003 and 2004 respectively (there would have been no change in 1999-2001).

[17] Fabiani, Gattulli e Sabbatini (2003) study a sample of Italian firms and find that in the services sector the main reason given by the firms for not adjusting their price is fear of competitors not following. In this sector, according to their study, about 30% of firms kept prices unchanged in 2002.

[18] The fact that, according to mainstream "menu costs" models, at the moment of the changeover the revision of most prices should produce a blip in inflation is discussed in section 6. Also see Hobjin *et al.* (2004).

Table 5: Distribution of price revisions, unchanged prices excluded (12-month growth)

Year	**1999**	**2000**	**2001**	**2002**	**2003**	**2004**
Sample average	12.0	15.7	14.6	12.7	10.1	8.7
5th percentile	-12.5	-8.3	3.7	-3.8	3.1	-12.4
Median	11.1	13.3	12.5	11.1	8.3	8.3
95th percentile	33.3	42.8	36.3	32.2	25.0	27.3
Standard deviation	14.5	14.5	12.6	11.5	9.2	12.4

Altogether, it may be conjectured that the dynamics of prices reflects several shocks: an increase in demand and costs affecting both 2000 and 2001; and a "mandatory" rewriting of menus in 2002, coinciding with the cash changeover, which prompted most agents to post new prices.

To compare the features of the pricing policies during these two episodes, we decomposed the average price change as follows: we first computed the differential between the average price change in each year (*i*) and in a base period (*j*), and then decomposed it into two contributions: the change in the share of prices which were reset (α) and the change in the average increase of those prices which were reset (φ) [19]:

(1) $$\pi_i - \pi_j \equiv (\alpha_i - \alpha_j)\left[\frac{\varphi_i + \varphi_j}{2}\right] + (\varphi_i - \varphi_j)\left[\frac{\alpha_i + \alpha_j}{2}\right]$$

A model of infrequent state-dependent price revisions has implications for the decomposition (1). Assuming that the main effect of the introduction of the euro is the simultaneous revision of all menus, the first component in (1) should explain most of the change in inflation; in contrast, in case of a shock to demand, costs or profit margins both components in (1) should contribute to the increase in inflation[20].

[19] The first term is weighted with the average increase in prices, the second with the average share α. Eq. (1) is obtained by manipulating the identity: $\pi_i=\alpha_i\ \varphi_i$,, where π_i is the average price increase, α_i is the share of price revisions in year *i* and φ_i is the average increase of those prices that are revised.

[20] See section 6 below.

The results are reported in Table 6. The year 1999 is used as the benchmark period *j*.

Table 6: Decomposition of price increases

	2000	**2001**	**2002**	**2003**	**2004**
Price increase: differential with respect to 1999	2.4	6.2	5.0	1.5	-0.5
i) due to larger share of price revisions	*1.0*	*4.8*	*4.6*	*2.4*	*0.8*
ii) due to larger average increase of revised prices	*1.4*	*1.4*	*0.4*	*-0.9*	*-1.3*

In 2001, the acceleration of prices with respect to the benchmark year (more than 6 percentage points) is due to both factors: almost five percentage points are due to a larger number of firms who revised their prices; about 1.5 percentage points are due to larger increases made by those restaurants which changed their price. This evidence is consistent with a demand and costs shock, which may have prompted more agents than usual to revise their price, while also increasing them substantially (e. g., because marginal costs increased more rapidly).

In 2002, the share of unchanged prices is also particularly low. However, the considerable number of price revisions determines 4.6 percentage points of higher inflation, the bulk of the overall increase. Unlike 2001, the effect of the size of the individual revisions on inflation is negligible (0.4 per cent).

This evidence is consistent with the conjecture that in 2002 a price shock was due to a "mandatory" price revision, due to the cash changeover. Many prices changed, but the amount of individual increases was not particularly large. This followed two years of already sustained price growth, related to exogenous demand and cost factors. Perceptions of a large price increase may have been the cumulated effect of all these events

5. Market structure and the effect of the changeover

The previous section suggests that the "revision of the menus" was the main facrtor explaining the impact of the euro on prices. Still, market structure could also have contributed to the decision about how much to revise prices in the face of the changeover. A possible explanation of such an effect is articulated in section 6 below: in industries characterized by low competition, producers may, at least temporarily, take advantage of the consumers' confusion to raise their price. Such an incentive is greater the less elastic the demand faced by the consumer, as this ensures the producer a smaller fall in sales in case the consumer realizes the true value of the new price. An implication of this conjecture is that only in year 2002 - and possibly in 2003, since increases taking place in the second half of 2002 are recorded in the following year's data - the magnitude of price increases should be negatively correlated to the degree of local market competition, proxied by the indicator discussed in section 2 (and shown in table A 5).

Some descriptive evidence, consistent with this conjecture, is presented in Table 7. Firms operating in more competitive local markets recorded a smaller increase in prices in 2002, and vice-versa. The table compares the average price increase with those of the upper and lower tails of the distribution, ordered according to the competition indicator; the second measure discussed in section 2 is employed. The difference is notable: about two and a half percentage points between the top and the bottom fourth of firms.

Table 7: Distribution of price increases in 2002 by market competition

	market competition < 1st quartile	**Market competition < median**	**Whole sample**	**Market competition > median**	**market competition > 3rd quartile**
Mean	9.7	9.8	9.3	8.8	7.3
5th percentile	0.0	0.0	0.0	-3.2	-13.0
Median	8.7	8.7	8.7	7.7	7.4
95th percentile	29.0	29.6	29.0	28.6	28.2

To address the issue more formally, we estimate a model of the determination of price changes over the period 1999-2003. Since price increases are truncated at zero (with a negligible number of exceptions), we model the price change as a case of censored data. To control for selectivity bias, we implement the estimates with the maximum likelihood equivalent of Heckman's two step selection model.[21] The main results are quite robust to the use of alternative estimation methods, as Tobit, OLS, random effects or fixed effect estimators.

We estimate the equation:

$$\Delta p_i = \mathbf{X}_i \boldsymbol{\beta} + \varepsilon_i \tag{2}$$

where p_i is the individual price of firm i and $\mathbf{X}_i$ is the vector of explanatory variables.[22]

Among the variables in $\mathbf{X}_i$ (affecting both the likelihood to change the price and the size of the price increase) we include the lagged log-price level of the firm (p_{t-1}); a dummy taking value one for restaurants, zero otherwise (D^{rist}); and, for those establishments classified as restaurants, their lagged quality judgement (i.e. the *score*),[23] and the (lagged) "age" of the price (*age*), expressed as the number of years since the last price revision. We assume that the actual price revision is an increasing function of the distance of the lagged price from its equilibrium level, a function of the *score* variable. The (lagged) "age" of the price captures a similar effect: under positive inflation the time elapsed since the last change increases the probability that the price is revised again, and increases the expected size of the revision.

As a proxy for cyclical demand pressure, we include the annual inflows of foreign tourists in each province, normalised by the population in the province (*tourism*). As for domestic demand, we include the (lagged) per capita value added in each province (*vadd*). Lacking other measures of real expenditure by households which have sufficient cross-sectional variability, any remaining demand effect is captured by the time dummies included in the regression. The same applies to cost variables (like fresh food prices) who are not available at the disaggregate level.

Among the explanatory variables we include our measure of local market competition

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[21] See Johnston and DiNardo (1997), ch.13.

[22] We assume that the same explanatory variables enter in the "selection" equation underlying Heckman estimates.

[23] Since the score variable is only available for restaurants, the variable is interacted with dummy D^{rist} in the regression.

(*comp*) defined in Section 2, based on the ratio of the number of restaurants in the guide to a weighted sum of local population and tourist presence, interacted with the time dummies (T_{1999} to T_{2003}). Our conjecture is that the degree of competition affects price dynamics only after the changeover; hence, the variable is entered separately for each year.

The estimation results are reported in the first column of Table 8. Most variables are strongly significant and have the expected sign. In 2002 and in 2003, the degree of local market competition significantly enters the equation for price determination. In contrast, it does not significantly enter the model in the previous years.

Table 8- Determinants of price changes

(Heckman selection model)

	Coef.	P>|z|		Coef.	P>|z|	
p(t-1)	-0.13	*0.00*	(°°)	-0.13	*0.00*	(°°)
score*D_{rest}	0.44	*0.00*	(°°)	0.44	*0.00*	(°°)
D_{rest}	-1.81	*0.00*	(°°)	-1.81	*0.00*	(°°)
age	0.01	*0.00*	(°°)	0.01	*0.00*	(°°)
tourism	0.00	*0.05*	(°°)	0.00	*0.05*	(°°)
v.add.	0.04	*0.00*	(°°)	0.04	*0.00*	(°°)
comp*T1999	-0.30	*0.14*		-0.30	*0.14*	
comp*T2000	-0.06	*0.77*		-0.06	*0.78*	
comp*T2001	-0.29	*0.08*		-0.29	*0.08*	
comp*T2002	-0.76	*0.00*	(°°)	-0.76	*0.00*	(°°)
comp*T2003	-0.91	*0.00*	(°°)	-0.90	*0.00*	(°°)
T2000	0.04	*0.00*	(°°)	0.04	*0.00*	(°°)
T2001	0.09	*0.00*	(°°)	0.06	*0.00*	(°°)
T2002	0.08	*0.00*	(°°)	0.04	*0.00*	(°°)
T2003	0.04	*0.00*	(°°)	0.01	*0.15*	
Δ Input cost				1.00	*restr.*	
Observations	10525			10525		
Censored	4689			4689		

Heckman selection model. Intercept (not shown) included. (°°) indicates 1% rejection confidence level.

Market structure thus seems to have contributed to the impact of the changeover, therefore giving a possible rationale for its non-neutrality. The size of the effect is relevant: differences in market concentration can be shown to explain a few points' differences in the price increases across provinces. Considering that in the equation for Δp_t the estimated coefficient on *comp* (when interacted with either T_{2002} or T_{2003}) is around -0.8, a decrease in *comp* (i.e., less competition in the local market) equal to its standard deviation (which is 0.03) would determine a price increase of about 2.5 percentage points.

The coefficients of the time dummies T_{2000}-T_{2003} provide us with an estimate of the time effects. In each year there is an "unexplained" increase relative to the 1999 benchmark, which captures the effect of price determinants not included in our regression, e.g. input costs. An attempt to control for the latter is presented in the second column of Table 8, where a proxy for the growth rate of input costs is added to the price equation (with a coefficient restricted to one). Once the model dynamics are accounted for,[24] the "unexplained" growth in prices in the year 2000 and 2001 is, respectively, 4 and 6 percentage points. In the year of the changeover "unexplained" inflation is still positve but smaller, around 3 percentage points; it is slightly negative in 2003.

6. The euro, competition and sticky prices: a simple model

This section proposes an analytical framework to provide a rigorous formulation of the "menu cost" and the "market power" hypotheses and to quantify the effect implied by each one, thus providing an assessment of their relevance.

We begin by showing how the cash changeover may cause a larger price increase where the market is less competitive. Next, we show how this effect may interact with the presence of menu costs and the consequent clustering of adjustments at the moment of introduction of the euro. We use the model presented above to fit some characteristics of the micro data before the changeover and to study the implications of the changeover under

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[24] Defining the time-effect on prices in each period j as E_j, the model dynamics imply $E_j=\beta_{Tj} + [(1+\beta_1)E_{j-1} - \beta_{Tj-1}]$, where β_{Tj} is the coefficient of the *dummy* T_j pertaining to the same period and β_1 is the coefficient on the lagged price level *p(t-1)*.

three alternative set of assumptions, namely that the introduction of the euro determined *i)* a simultaneous revision of all prices; *ii)* a temporary increase in market power; *iii)* both.

A key assumption is that the introduction of a new currency increases the difficulty faced by the consumer in evaluating the relative price of the goods he or she is purchasing. Such a lack of price visibility, similar to what happens when engaging in transactions in a foreign currency, has the effect of enabling producers to increase their selling prices the more, the greater is their market power.

We consider a firm *i* producing a good with a constant unit cost of production, defined as W_i, under monopolistic competition, facing a demand curve $D_i = D(P_i \; , P_k)$, where P_i is firm *i*'s price and P_k is the price set by the firm's competitors (which firm *i* takes as given); price elasticity is constant, $\varepsilon > 1$. We omit the firm suffix in what follows. The first order condition from profit maximization ($\max_P DP - WD$) yields the price: $P = W(1 + 1/(\varepsilon - 1)) \equiv \overline{P}$. We define the quantity demanded at this equilibrium price $\overline{D} = D(\overline{P})$; we assume that before the cash changeover the market clears, with the firm producing and selling $\overline{D}$ at price $\overline{P}$.

The consumer is confused by the cash changeover. This means that the firm can increase the price without the consumer noticing. More precisely, we assume that the consumer is presented with a new price $P^{€}$, not necessarily corresponding to the previous one. We assume that the probability that the consumer notices that the price has been changed is γ. If the consumer realizes that the price has changed, he just follows its demand curve. If he does not realize it, he keeps demanding the previous quantity $\overline{D}$.

In this framework, the problem of the firm is:

(3) $$\max_P \gamma(P - W)D + (1 - \gamma)(P - W)\overline{D}$$

The F.O.C. s are:

(4) $$\gamma D + (1 - \gamma)\overline{D} + \gamma(P - W)\frac{\partial D}{\partial P} = 0 \, ,$$

which yield the new equilibrium price $P^{€} = W\left[1+\frac{1}{\gamma h\varepsilon - 1}\right]$, where the new (lower) demand elasticity is $\gamma h\varepsilon$; h is a positive term not larger than 1, increasing in γ.[25]

The percent price change with respect to the old level is:

$$(5) \qquad \pi^{€} \equiv \frac{P^{€} - \overline{P}}{\overline{P}} = \frac{1-\gamma h}{\gamma h\varepsilon - 1}.$$

Equation (5) shows that after the changeover the price is unchanged if $\gamma=1$ (trivially, in this case there is no incentive to move the price since the consumer is on the original demand curve with probability 1)[26] or if $\varepsilon \to \infty$ (perfect competition). The inflation induced by the euro-changeover is larger, the larger is the firm's market power (small ε) and, obviously, the larger the probability that the consumer does not perceive the price change (small γ).

The interpretation of the result is intuitive. The consumer's "confusion" caused by the cash changeover, captured by the variable γ, decreases the demand elasticity expected by the firm; the decrease is proportional to the original demand elasticity, ε. In the problem (3) faced by the firm, the marginal benefit of a higher price is the increase in the unit value of the old quantity sold, $\overline{D}$; the marginal disadvantage depends on the probability that the consumer realizes that the price has changed and on the consequent decrease in sales, which is function of the elasticity ε (in this case the consumer follows the demand curve $D(P)$) .

To capture the high observed degree of price stickyness, the model needs to be cast in a dynamic setting that accounts for costly price adjustment. We follow Dotsey, King and Wolman (1999) and assume that firms incur a random menu cost ξ_t if their price is revised. The decision that in period t the firm will not revise its price thus depends on the comparison between the adjustment cost and the increase in profits that would be obtained by changing

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[25] It is $h = \left(\gamma + (1-\gamma)\frac{\overline{D}}{D}\right)^{-1}$. Using a linear approximation of the demand function, it may be shown that in equilibrium it is $h = \frac{\gamma(\varepsilon-1)-(1-\gamma)\varepsilon}{\gamma(\varepsilon-1)-\gamma(1-\gamma)\varepsilon}$. From S.O.C.s one gets $\gamma(\varepsilon-1)-(1-\gamma)\varepsilon > 0$. hence $0<h<1$. Condition $\gamma h>1/\varepsilon$ is necessary to have an internal solution.

[26] Note that if $\gamma=1$, then $h=1$.

the price. Marginal costs and the demand curve $D(P_t)$ are assumed to be the same for all firms, so we omit the firm suffix in our notation, but introduce time suffixes. We define $\alpha_{t|t-j}$ as the probability that in period t a price originally set in period t-j is left unchanged; as in Dotsey, King and Wolman (1999), the probability depends on the distribution in the menu cost: firms balance the gains from adjusting the price with the menu cost (see Appendix B).

Firms must take into account the effect of the current price on future periods' profits, in case they will not revise the price again. The solution of the optimisation problem is derived in Appendix B. From the first order conditions, the optimally reset price in each period is equal to the present discounted value of future marginal costs divided by the present discounted value of the (inverse of) future mark-ups:

(7) $$P_t^* = \left[\sum_{i=0}^{\infty} \beta^i A_{t+i|t} \varphi_{t+i} \frac{1}{X_{t+i}} \frac{\varepsilon_{t+i}-1}{\varepsilon_{t+i}}\right]^{-1} \left[\sum_{i=0}^{\infty} \beta^i A_{t+i|t} \varphi_{t+i} \frac{W_{t+i}}{X_{t+i}}\right]$$

where β is the discount factor, $A_{t+i|t} \equiv \prod_{k=0}^{i} \alpha_{t+k|t}$, φ_t is the (possibly time varying) price derivative of the demand function (normalised, without loss of generality, as a ratio to a steady state value), ε_t is the (possibly time varying) demand elasticity and X_t is the general (whole economy) price level.[27] The sector's *aggregate* price level P_t is obtained as a weighted average of vintages of past prices P^*_t 's, where the share of vintage h prices in period t, is a function of the $\alpha_{i|j}$'s.

As assumed above, in $t0$ with probability (1-γ) consumers keep demanding the benchmark quantity $D(\bar{P}_{t0})$, where $\bar{P}_{t0}$ is the price that would have prevailed without the changeover. Considering expected demand $D_{t0} = \gamma D(P_{t0}) + (1-\gamma) D(\bar{P}_{t0})$, φ_{t0} and ε_{t0} in (7) can be derived.[28]

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[27] If $\varepsilon_t = \varepsilon$, (7) can be shown to be consistent with the result of Dotsey, King and Wolman (1999). If $\alpha_{t|k} = \alpha$, $\varphi_t = 1$, $\varepsilon_t = \varepsilon$, the standard result of Calvo (1983) obtains.

[28] It is $\varphi_{t0} = \gamma$. In a neighbourhood of equilibrium, $\varepsilon_{t0} = \varepsilon\gamma \left(1 + (1-\gamma)\varepsilon \frac{P_{t0} - \bar{P}_{t0}}{P_{t0}}\right)^{-1}$. This expression can be solved simultaneously with (7) to derive ε_{t0} and P_{t0}.

We calibrate (7) assuming the annual discount factor is β=0.96 and, based on the features of our sample in 1999, that the steady state marginal cost W_t grows at a constant annual rate of 4 percent. Using a relatively standard assumption, demand elasticity is set at $\varepsilon_{t|t\neq t0}$=11 and $\varphi_{t|t\neq t0}$=1, which corresponds to an equilibrium markup of 10%. The path of $\alpha_{t|t-j}$ is endogenously determined by positing a uniform distribution for adjustment costs, calibrated to reproduce the fact than on average about 60 percent of firms do not revise their price, as from our panel data.

The probability that the consumer does not notice that the price has changed is set at $(1-\gamma)$=0.3. This assumption matches the evidence of the European Commission's "eurobarometer"; according to this survey, at end 2003 36% of Italian consumers reckoned, in retrospective, that their attitude was "to buy more, because they did not realize how much they are spending" after the introduction of the euro.[29]

In the first scenario the changeover involves a simultaneous "forced" revision of most prices (the "menu cost" is nil for all firms). In the second we posit that the consumer is confused with probability 1-γ, which increases the firm's degree of monopolistic power. In the third scenario both shocks occur.

The results of the simulations are shown in the four panels of Figure 6, which include the rate of inflation, the size of the average price revision, the share of unrevised prices and the price level. In each panel, the solid line shows the responses to a "menu cost" shock, the dashed line the responses to a "market power" shock and the dotted line the responses to both shocks occurring simultaneously. All shocks are assumed to take place in *t0*.

The menu-cost shock scenario causes an increase in inflation, reaching 12.9 percent in *t0* (top-left panel). The increase mostly reflects the share of unrevised prices dropping to zero (bottom-left panel), thereby sustaining average inflation; in contrast, the average size of individual price revisions is only marginally larger than in steady-state (top-right panel). Before *t0*, inflation falls slightly: the agents incorporate less future expected marginal cost

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[29] European Commission "Eurobarometer", *The euro, two years later*, January 2004.

increases in their prices, knowing they will be soon resetting prices. Afterwards, the inflation rate gradually returns to its steady state value (4%) with small oscillations.[30]

In the second scenario, the "market power" shock has an impact on period-*t0* inflation, which increases to 6.3 percent.[31] The increase in inflation is the result both of a temporary increase in margins and of an (endogenously) larger number of price revisions: the share of prices which stay fixed drops to 53 percent, as more firms find it convenient to sustain the adjustment cost to exploit the increase in market power. The third scenario involving both shocks yields an higher inflation in *t0*, around 15 percent.

In all three cases, the deviation of the price level from its long-run trend is relatively short-lived (bottom-right panel). After temporarily rising above their steady state path, prices slow down and are back to baseline on the second year after the shock.

Table 9: Model-based decomposition of price increases

	Menu cost shock	**Market power shock**	**Both**
Price increase: differential with respect to steady state	8.9	2.3	11.0
i) due to larger share of price revisions	*8.1 (92%)*	*1.7 (74%)*	*8.8 (80%)*
ii) due to larger average increase of revised prices	*0.8 (8%)*	*0.6 (26%)*	*2.2 (20%)*

The decomposition of the price increase, along the lines described in Section 4, is quite different in the three cases (Table 9). After the "menu cost" shock, the increase in inflation is almost entirely (92%) due to the large number of price revisions. After the "market power"

[30] As discussed by Hobjin *et al.*, 2004, these oscillation are typical of models which include a maximum lag beyond which all prices are adjusted.

[31] The size of this effect, however, is strongly dependent on the exact shape assumed for the menu cost distribution, and it would be greater under parametrisations which increase the degree of state-dependency of the adjustment.

shock, the contribution to the increase in inflation is more evenly split among the two components (74 and 26 percent, respectively). When both shocks hit the economy, the decomposition of inflation lies somewhere in between.[32]

How does the model compare to the facts described in the previous sections? The profile of inflation and its decomposition are consistent with the idea that the "menu cost" shock may have determined the bulk of the impact of the introduction of the euro, with the "market power" shock possibly playing a complementary role. The overall size of the increase in inflation observed in the sample is fully consistent with what would have been produced by a mere revision of all menus, given the relatively high degree of price stickiness on this market. It is possible, however, that an increase in market power reinforced the effect of the "menu cost" shock.

According to the model the price level will gradually return to its steady-state path. Such a convergence takes the form of small price revisions and a relatively high share of sticky prices. This implication is qualitatively - albeit maybe not quantitatively - matched by the features observed in our sample in 2004; in that year, the average price revision and the number of firms adjusting prices are below their medium-run levels.

7. Conclusions

This paper analyzed a panel of restaurant prices to study price setting behavior at the firm level during the cash changeover. The unique database allows us to shed light on two controversial issues concerning the effects of the cash changeover: first, to document what happened to prices at the micro level, in the year of the changeover and in the neighbouring ones. This evidence is a first key step in understanding what caused the public perceptions to deviate so markedly from official inflation measures. Second, to begin a systematic analysis of the mechanisms that may have affected price setting during the cash changeover.

On the first question, our results show that a sizeable average price increase took place in 2002 (about 9 per cent). This increase is slightly smaller than the one recorded by the

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[32] These percentages are rather robust to different assumptions and calibrations for the menu cost distribution.

industry in the previous year (about 10 per cent). The evidence suggests that such increases are related to rising demand and costs conditions; this should induce caution in attributing a large inflationary effect *exclusively* to the introduction of the euro banknotes.

The evidence clearly dismisses the hypothesis that the euro changeover favored a "doubling of prices". Such a widespread perception might be ascribed to the substantial price increase which took place in this sector over a longer time period (between 1998 and 2003 the average price of a meal increases by 40 per cent; in the 10 per cent of restaurants recording the largest increases the rise is 75 per cent). The changeover might have focussed the public attention on the price level, prompting the attribution of the whole increase to the introduction of the euro.

On the second question, we found that both the "menu cost" and the "market power" hypotheses find support in the data. As to the former, the evidence shows that much of the aggregate price increase during the changeover results from a greater number of prices being revised, rather than from "large" individual price increases. This appears consistent with the idea that the new currency denomination made it "mandatory" for almost all firms to revise their prices in 2002. In normal years, a great proportion of restaurants do not adjust their price, probably due to the presence of small menu costs.

We also found that market structure affected price dynamics after the changeover. In 2002 and 2003 price increases were larger in the provinces which featured a smaller degree of competitiveness.

Both the "menu cost" and the "market power" hypothesis imply only a temporary effect on the price level. As a consequence, one may conjecture that price increases will be gradually reabsorbed, although it is too early to have decisive evidence on the latter conjecture. So far, in 2003 and in 2004 both the average price increase and the number of price changes were relatively moderate.

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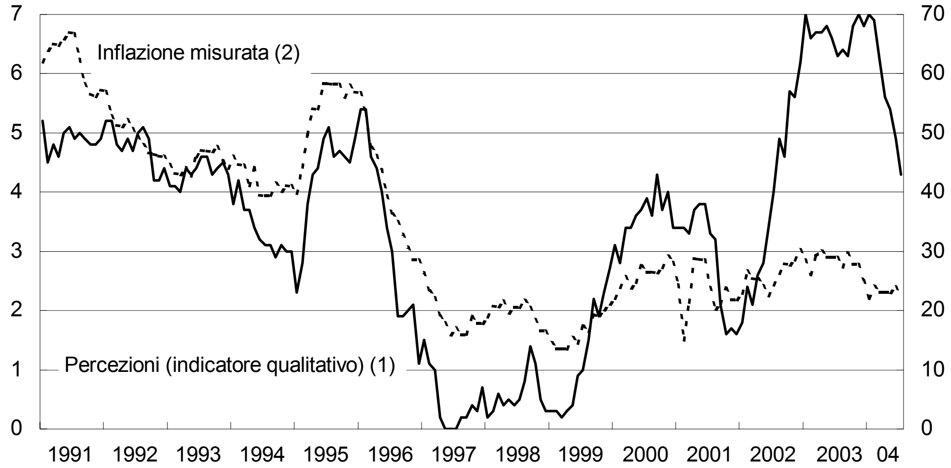
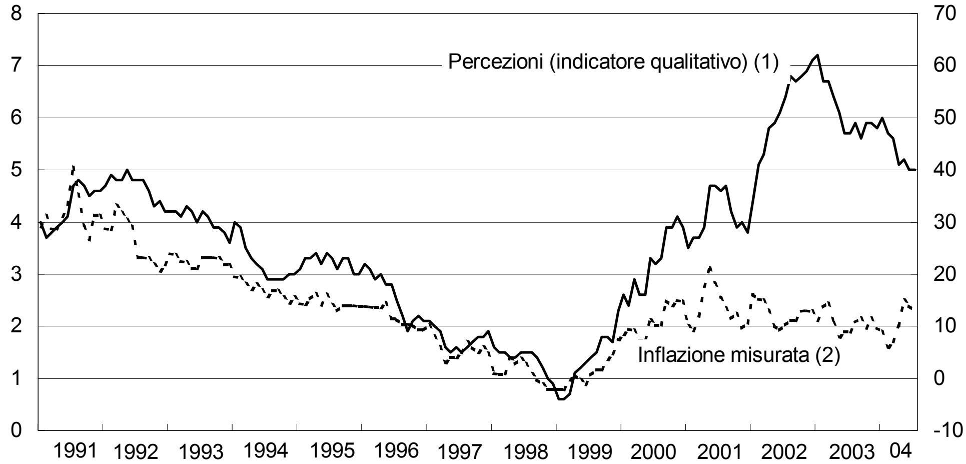
Figure 1 - Perceived inflation in Italy and in the euro area

(a) Italy



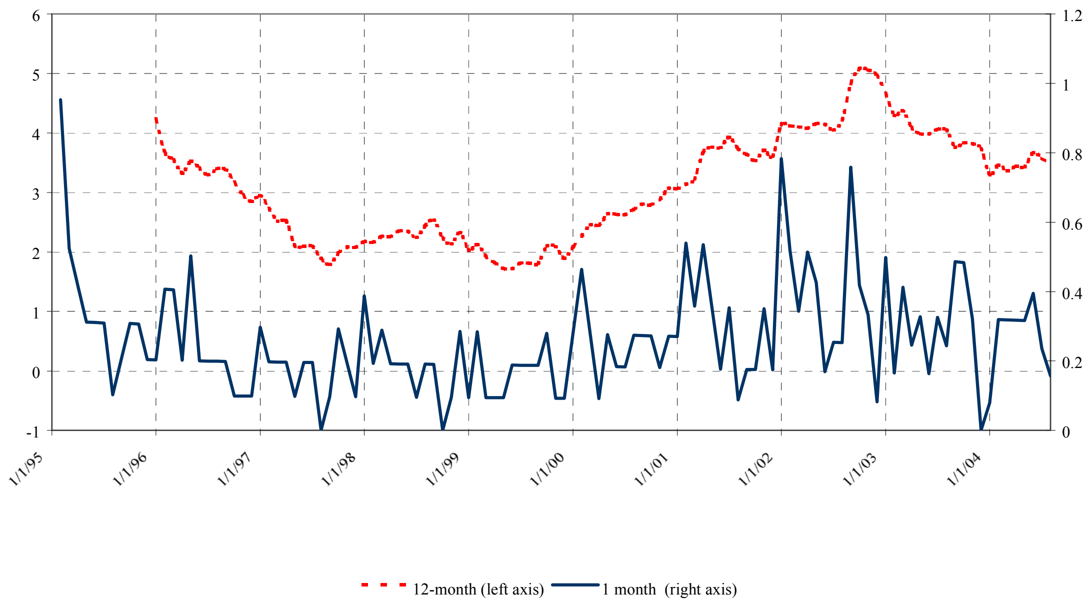
(b) Euro area



Source: European Commission, Eurostat. (1) Perceived inflation: based on surveys by ISAE and by the European Commission. Differences between the share of respondents reporting "strongly increased" or "moderately increased" prices and the share of respondents reporting "stable" or "decreased" prices (right axis). (2) HICP inflation: twelve-month growth rates

Figure 2 - Italy: HICP, category "restaurants and cafes"

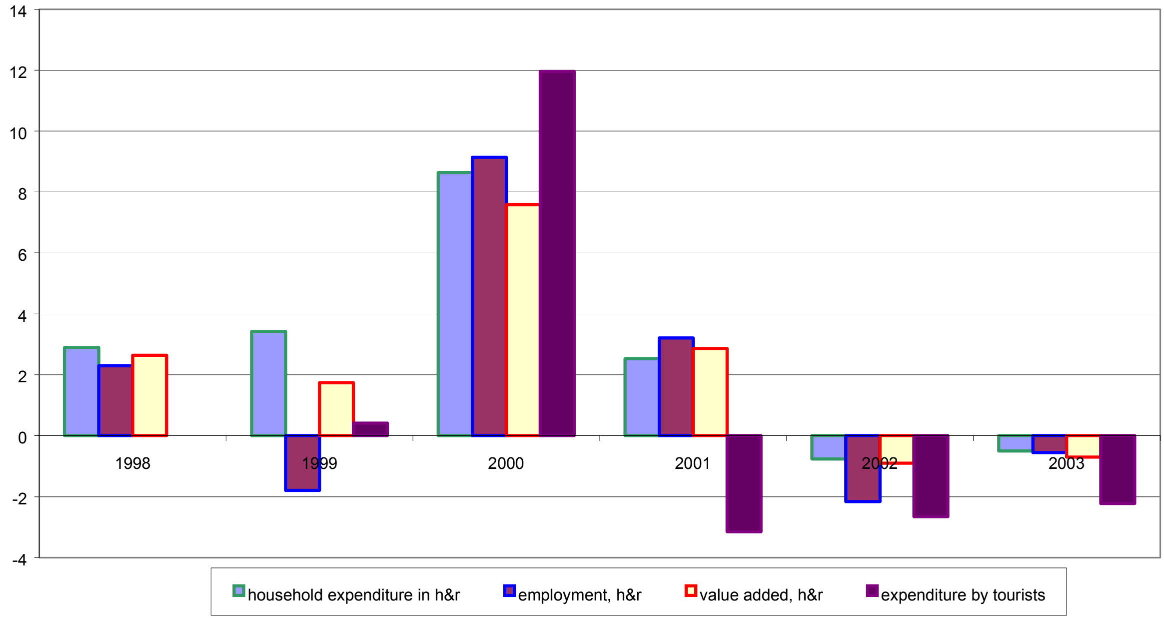
(percentage changes)



Source: Eurostat.

Figure 3 - Activity and demand in the sector "hotels & restaurants"

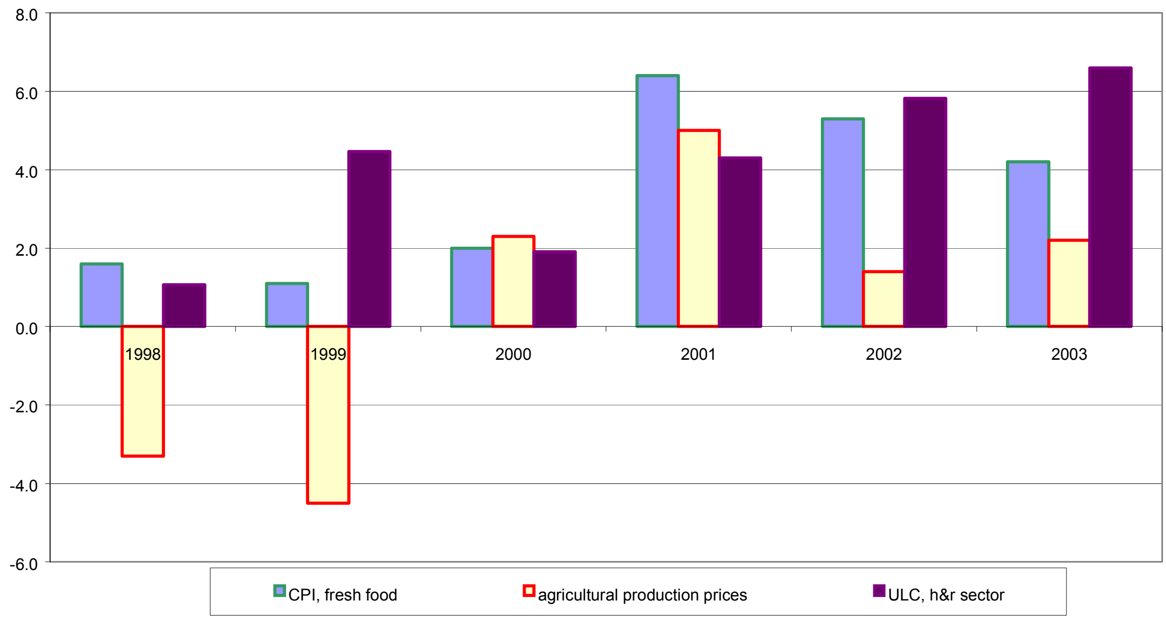
(annual percentage changes)



Source: Istat, UIC.

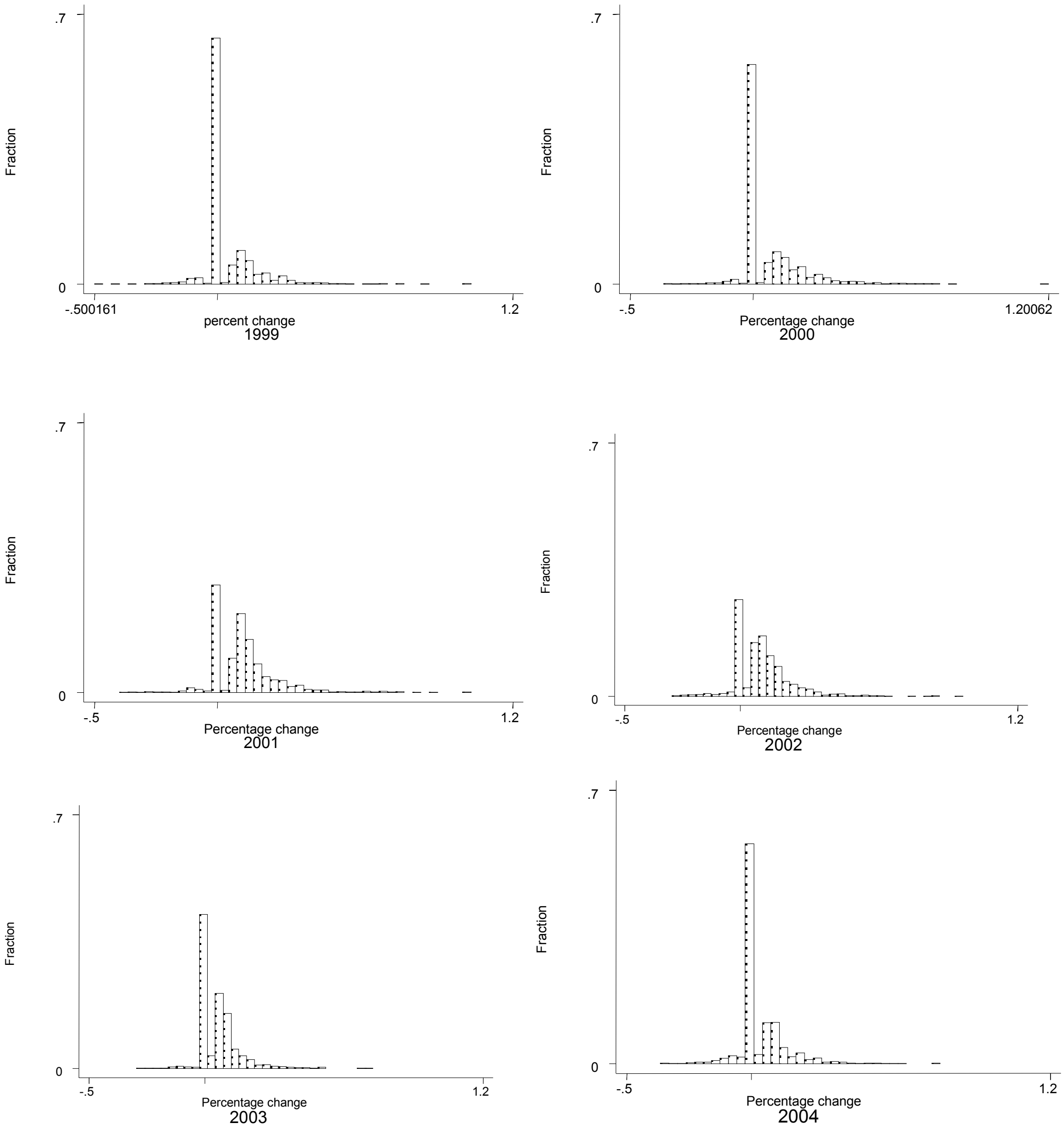
Figure 4 - Fresh food prices and ULC in the sector "hotels and restaurants"

(annual percentage changes)



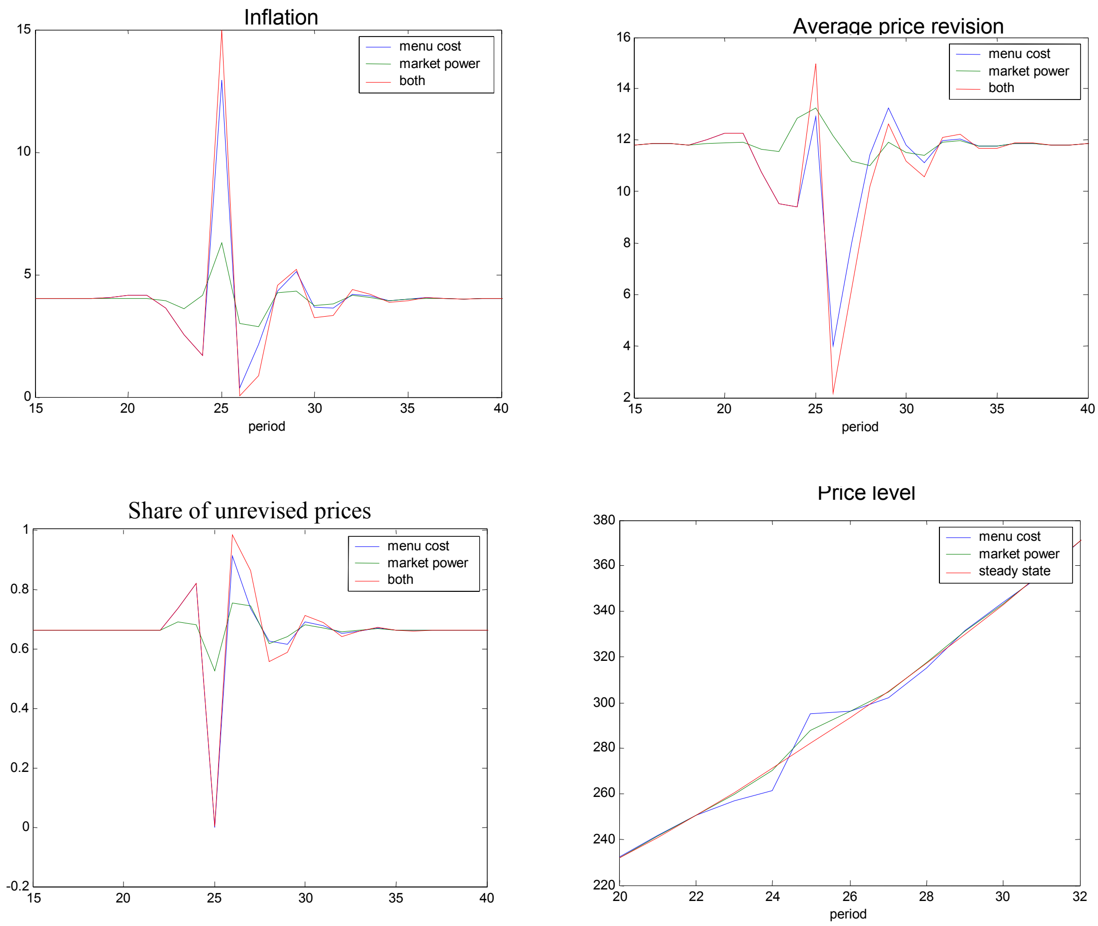
Source: Eurostat, Istat.

Figure 5 – Frequency distribution of price increases



Frequency distribution of price increases in each year. Range of percentage increases from -0.5 (-50%) to 1,2 (+120%) (horizontal axis). The tick marks 0. Each bar includes a range of price increases from *t* to *t+3,5* percentage points.

Figure 6 – Model based effect of the changeover



Based on the model in section 6.

Appendix A: descriptive statistics

Table A 1 - Sample size

(number of observations)

	price	price change
1998	2632	
1999	2739	2239
2000	2650	2292
2001	2569	2144
2002	2356	2023
2003	2321	2028
2004	1769	1552
Total	17036	12278

Source: authors' computation based on data from *Il Gambero Rosso.*

Table A 2 - Sample composition, by region

(number of observation and percentages)

region	observations	frequency	Italy: distribution of all restaurants, 1999
Abruzzo	336	2.7%	2.9%
Basilicata	151	1.2%	0.7%
Calabria	269	2.2%	3.4%
Campania	574	4.7%	9.7%
Emilia Romagna	1022	8.3%	6.5%
Friuli Venezia Giulia	426	3.5%	3.1%
Lazio	1471	12.0%	9.5%
Liguria	621	5.1%	4.2%
Lombardia	1608	13.1%	13.6%
Marche	396	3.2%	3.0%
Molise	73	0.6%	0.7%
Piemonte	984	8.0%	7.3%
Puglia	530	4.3%	6.0%
Sardegna	307	2.5%	3.2%
Sicilia	600	4.9%	5.4%
Toscana	1240	10.1%	7.5%
Trentino Alto Adige	445	3.6%	2.6%
Umbria	271	2.2%	1.7%
Val d'Aosta	113	0.9%	0.3%
Veneto	841	6.8%	8.9%
North	6060	49.4%	46.4%
Center	3378	27.5%	21.2%
South & islands	2840	23.1%	31.9%
Total	12278	100.0%	100.0%

Source: authors' computation based on data from *Il Gambero Rosso* and Chambers of Commerce.

(*) Restaurants registered at the Chambers of Commerce in 1999 (source: *Rapporto sul turismo italiano - X edizione).*

Table. A 3 - Sample composition, by type

(number of observations and percentages)

region	(a) restaurants	(b) others	(a) / (a+b)
Abruzzo	218	118	64.9%
Basilicata	113	38	74.8%
Calabria	170	99	63.2%
Campania	323	251	56.3%
Emilia Romagna	703	319	68.8%
Friuli Venezia Giulia	325	101	76.3%
Lazio	846	625	57.5%
Liguria	442	179	71.2%
Lombardia	1208	400	75.1%
Marche	230	166	58.1%
Molise	27	46	37.0%
Piemonte	861	123	87.5%
Puglia	328	202	61.9%
Sardegna	218	89	71.0%
Sicilia	404	196	67.3%
Toscana	837	403	67.5%
Trentino Alto Adige	366	79	82.2%
Umbria	198	73	73.1%
Val d'Aosta	76	37	67.3%
Veneto	644	197	76.6%
North	4625	1435	76.3%
Center	2111	1267	62.5%
South & islands	1801	1039	63.4%
Total	8537	3741	69.5%

Source: authors' computation based on data from *Il Gambero Rosso.*

Table A 4 - Sample representativeness

(number of observations and percentages)

region	(a) sample observations in 1999	(b) total establishments In Italy In 1999	(a) / (b)
Abruzzo	52	1914	2.72%
Basilicata	27	476	5.67%
Calabria	48	2196	2.19%
Campania	101	6272	1.61%
Emilia Romagna	209	4203	4.97%
Friuli Venezia Giulia	70	1985	3.53%
Lazio	247	6157	4.01%
Liguria	103	2701	3.81%
Lombardia	302	8834	3.42%
Marche	72	1933	3.72%
Molise	14	437	3.20%
Piemonte	154	4764	3.23%
Puglia	92	3889	2.37%
Sardegna	57	2053	2.78%
Sicilia	129	3483	3.70%
Toscana	252	4851	5.19%
Trentino Alto Adige	84	1672	5.02%
Umbria	54	1100	4.91%
Val d'Aosta	21	206	10.19%
Veneto	151	5758	2.62%
North	*1094*	*30123*	*3.63%*
Center	*625*	*14041*	*4.45%*
South & islands	*520*	*20720*	*2.51%*
Italy	*2239*	*64884*	*3.45%*

Source: authors' computation based on data from *Il Gambero Rosso* and Chambers of Commerce.

Table A 5 - Competitive pressure

(number of restaurants included in the guide per capita)

region	(a)	(b)
Abruzzo	0.56	0.39
Basilicata	0.58	0.50
Calabria	0.27	0.22
Campania	0.22	0.17
Emilia Romagna	0.63	0.44
Friuli Venezia Giulia	0.81	0.55
Lazio	0.70	0.51
Liguria	0.79	0.44
Lombardia	0.40	0.33
Marche	0.64	0.40
Molise	0.51	0.46
Piemonte	0.67	0.59
Puglia	0.29	0.26
Sardegna	0.46	0.32
Sicilia	0.35	0.28
Toscana	0.93	0.51
Trentino Alto Adige	1.01	0.26
Umbria	0.71	0.55
Val d'Aosta	2.07	0.62
Veneto	0.44	0.23

Averages of ratios computed at the province level. (a) Number of establishments included in the guide over total population in the province (tens of thousands); (b) Number of establishments included in the guide over (days of tourist presence/12 + local population) - see text. Source: Chambers of Commerce, Istat and authors' computation based on data from *Il Gambero Rosso*.

Table A 6 - The price of a meal

(euros)

Restaurants

	mean	min	1°quartile	median	4°quartile	max	stand. dev.	coeff. var.
1998	33.2	12.9	25.8	31.0	36.2	113.6	11.6	0.35
1999	34.3	12.9	25.8	31.0	38.7	129.1	12.4	0.36
2000	36.3	12.9	28.4	33.6	41.3	129.1	13.2	0.36
2001	39.3	15.5	31.0	36.2	43.9	147.2	14.8	0.38
2002	42.8	16.0	32.0	40.0	48.0	220.0	16.7	0.39
2003	45.5	20.0	35.0	40.0	50.0	220.0	17.6	0.39
2004	48.6	21.0	36.0	45.0	55.0	250.0	19.8	0.41

Others

	mean	min	1°quartile	median	4°quartile	max	stand. dev.	coeff. var.
1998	19.5	7.8	15.5	18.1	23.2	72.3	5.6	0.29
1999	19.8	7.8	15.5	20.7	23.2	56.8	4.9	0.25
2000	20.6	7.8	18.1	20.7	23.2	49.1	4.6	0.22
2001	22.5	8.8	18.1	23.2	25.8	46.5	5.3	0.23
2002	24.7	9.0	20.0	25.0	28.0	75.0	6.1	0.25
2003	28.0	11.0	23.0	27.0	30.0	105.0	9.1	0.32

Source: authors' computation based on data from *Il Gambero Rosso.*

Table A 7 - Price changes, by type

(percent changes)

Restaurants

	mean	min	5° percentile	median	95° percentile	max	std. dev.
1999	4.8	-50.0	-6.7	0.0	25.0	100.0	10.8
2000	7.1	-33.3	0.0	0.0	30.8	120.1	12.3
2001	10.3	-36.8	0.0	9.1	33.3	100.1	12.7
2002	9.1	-27.1	-3.0	7.7	29.0	83.3	11.2
2003	5.5	-27.0	0.0	4.3	20.0	71.9	8.5
2004	3.9	-36.4	-7.5	0.0	21.2	75.0	9.6

Others

	mean	min	5° percentile	median	95° percentile	max	std. dev.
1999	3.3	-42.9	0.0	0.0	19.9	75.0	9.5
2000	5.8	-33.3	0.0	0.0	28.5	80.0	12.2
2001	10.8	-33.3	0.0	11.1	33.4	80.0	12.1
2002	9.8	-27.8	0.0	8.7	29.0	94.4	11.6
2003	6.4	-29.4	0.0	6.7	20.0	66.7	8.7
2004	3.6	-21.4	0.0	0.0	18.5	60.0	8.2

Source: authors' computation based on data from *Il Gambero Rosso.*

Tav. A 8 - Price changes, by area

(percent changes)

North

	mean	min	5° percentile	median	95° percentile	max	std. dev.
1999	3.9	-27.3	-7.1	0.0	23.1	100.0	10.2
2000	6.1	-33.3	0.0	0.0	27.8	75.0	11.1
2001	10.9	-30.0	0.0	10.0	33.3	87.5	12.0
2002	8.9	-26.8	-3.8	7.7	30.4	84.6	11.8
2003	5.9	-24.3	0.0	5.0	20.0	66.7	8.1
2004	3.9	-31.8	-6.7	0.0	20.0	75.0	9.5

Center

	mean	min	5° percentile	median	95° percentile	max	std. dev.
1999	5.9	-23.1	0.0	0.0	25.0	99.9	10.6
2000	8.6	-27.3	0.0	0.0	37.5	120.1	14.3
2001	11.2	-36.8	0.0	10.0	37.5	100.0	13.4
2002	9.2	-26.9	0.0	7.7	29.0	80.6	11.1
2003	5.3	-29.4	0.0	0.0	20.7	50.0	8.8
2004	4.1	-36.2	0.0	0.0	20.5	57.1	9.3

South and islands

	mean	min	5° percentile	median	95° percentile	max	std. dev.
1999	3.5	-50.0	0.0	0.0	21.1	75.0	10.6
2000	5.8	-33.3	0.0	0.0	28.5	75.0	11.8
2001	8.8	-33.3	0.0	9.1	25.0	100.1	12.2
2002	10.2	-27.8	0.0	9.5	26.9	94.4	10.4
2003	5.9	-12.7	0.0	4.7	20.7	71.9	9.0
2004	3.1	-36.4	-7.1	0.0	19.4	46.9	8.3

Source: authors' computation based on data from *Il Gambero Rosso.*

Table A 9 - Restaurant prices in Italy: a comparison among datasets

(annual percentage changes)

	1999	2000	2001	2002	2003	2004
GR guide *(our panel)*	4.3	6.7	10.5	9.3	5.8	3.8
Michelin Red Guide *(Adriani et al., 2003)*	n.a.	n.a.	n.a.	n.a.	5.6	n.a.
HICP restaurant and cafes (*Hobjin et al., 2004)*	1.9	3.1	3.6	5.0	3.7	3.5
Istat ristoranti, pizzerie, pubblici esercizi	1.7	3.2	3.8	4.8	4.6	3.6

Source: authors' computation based on data from *Il Gambero Rosso,* Istat, Eurostat, Adriani et al. (2003).

Appendix B - Mathematical details on the model of Section 6

The problem analyzed in this appendix is a small variation of Dotsey, King and Wolman (1999), which allows for a time-varying elasticity of demand.[33]

The value of the firm in period t if it adjusts the price to the new optimal level is defined as $v_{t|j}$ and it is:

$$\text{(B1)}\qquad v_{t|t} = \max_{P^*_t} \left\{ \frac{(P^*_t - W_t)}{X_t} D_t(P^*_t, X_t) + \beta\alpha_{t+1|t} v_{t+1|t} + \beta(1-\alpha_{t+1|t}) v_{t+1|t+1} - \beta\,\Xi_{t+1|t} \right\}$$

where X_t is the general price level, W_t denotes the marginal cost, $\alpha_{t|j}$ the period t (endogenous) probability of not revising a price which was set in period j, and $\Xi_{t+1|t}$ is the present expected value of the adjustment costs to be sustained next period. The value of the firm in period t if it maintains its price at the current level, P^*_{t-j} (where j is the price "vintage") is given by:

$$\text{(B2)}\qquad v_{t|t-j} = \left\{ \frac{(P^*_{t-j} - W_t)}{X_t} D_t(P^*_{t-j}, X_t) + \beta\alpha_{t+1|t-j} v_{t+1|t-j} + \beta(1-\alpha_{t+1|t-j}) v_{t+1|t+1} - \beta\,\Xi_{t+1|t-j} \right\}$$

There is a fixed distribution of menu costs, which we assume for simplicity to be uniform over the interval (0, $1/k$). The marginal firm (i. e., the one indifferent about resetting the price) is the one for which the realized menu cost equates the gain in its value resulting from price adjustment ($v_{t|t}$ - $v_{t|t-j}$); hence, the share of firms adjusting their price is given by the c.d.f. of the menu cost, evaluated at the point ($v_{t|t}$- $v_{t|t-j}$).

$$\text{(B3)}\qquad 1-\alpha_{t+1|t-j} = k(v_{t+1|t+1} - v_{t+1|t-j})$$

The expected adjustment cost next period is:

$$\text{(B4)}\qquad \Xi_{t+1|t-j} = \int_0^{(1-\alpha_{t+1|i})/k} kx\,dx = W_t \frac{k}{2}(v_{t+1|t+1} - v_{t+1|t-j})^2$$

Taking the derivative of (B3) and (B4):

$$\text{(B5)} \qquad \frac{\partial \alpha_{t+1|t-j}}{\partial (v_{t+1|t+1} - v_{t+1|t-j})} = -k$$

$$\text{(B6)} \qquad \frac{\partial \Xi_{t+1|t-j}}{\partial (v_{t+1|t+1} - v_{t+1|t-j})} = k(v_{t+1|t+1} - v_{t+1|t-j})$$

The dynamic program (B1) implies the condition:

$$\text{(B7)} \qquad \begin{aligned} 0 &= (P^*_t - W_t)\frac{1}{X_t}\frac{\partial D_t}{\partial P_t^*} + D_t + \beta\alpha_{t+1|t}\frac{\partial v_{t+1|t}}{\partial P_t^*} - \beta k v_{t+1|t}\frac{\partial (v_{t+1|t+1} - v_{t+1|t})}{\partial P^*_t} - \\ &+ \beta k v_{t+1|t+1}\frac{\partial (v_{t+1|t+1} - v_{t+1|t})}{\partial P_t^*} - \beta k (v_{t+1|t+1} - v_{t+1|t})\frac{\partial (v_{t+1|t+1} - v_{t+1|t})}{\partial P_t^*} = \\ &= (P_t^* - W_t)\frac{1}{X_t}\frac{\partial D_t}{\partial P_t^*} + \frac{1}{X_t}D_t + \beta\alpha_{t+1|t}\frac{\partial v_{t+1|t}}{\partial P_t^*} \end{aligned}$$

Taking the derivative of (B2) and using (B5) and (B6):

$$\text{(B8)} \qquad \begin{aligned} \frac{\partial v_{t|t-j}}{\partial P^*_{t-j}} &= (P^*_t - W_t)\frac{1}{X_t}\frac{\partial D_t}{\partial P_{t-j}^*} + \frac{D_t}{X_t} + \beta\alpha_{t+1|t-j}\frac{\partial v_{t+1|t-j}}{\partial P^*_{t-j}} - \beta k v_{t+1|t-j}\frac{\partial (v_{t+1|t+1} - v_{t+1|t-j})}{\partial P^*_{t-j}} - \\ &+ \beta k v_{t+1|t+1}\frac{\partial (v_{t+1|t+1} - v_{t+1|t-j})}{\partial P^*_{t-j}} - \beta k (v_{t+1|t+1} - v_{t+1|t-j})\frac{\partial (v_{t+1|t+1} - v_{t+1|t-j})}{\partial P^*_{t-j}} = \\ &= (P^*_{t-j} - W_t)\frac{1}{X_t}\frac{\partial D_t}{\partial P_{t-j}^*} + \frac{1}{X_t}D_t + \beta\alpha_{t+1|t-j}\frac{\partial v_{t+1|t-j}}{\partial P^*_{t-j}} \end{aligned}$$

Substituting (B8) into (B7) and iterating yields:

$$\text{(B9)} \qquad 0 = \sum_{i=0}^{+\infty}\beta^i\left[\left(\prod_{j=0}^{i}\alpha_{t+j|t}\right)\left((P^*_t - W_{t+i})\frac{\partial D_{t+i}}{\partial P^*_t} + D_{t+i}\right)\frac{1}{X_{t+i}}\right]$$

Defining $\varphi_{t+i} \equiv \frac{\partial D_{t+i}}{\partial P_t^*}$ and $\varepsilon_{t+i} \equiv -\frac{\partial D_{t+i}}{\partial P_t^*}\frac{P_t^*}{D_{t+i}}$, yields the expression for the optimal reset price P_t^* (equation 7 in the main text).

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[33] We also assume consumers have access to complete markets so that the ratio of marginal utility affecting the stochastic discount factor in DKW is constant at 1.

Based on (B8), $v_{t|t}$ can be linearized in a neighbourhood of $[P_{t-j}^*, v_{t|t-j}]$ to obtain an expression for the impact of a price revision on firm's value:

(B10) $$v_{t|t} - v_{t|t-j} \cong \left(\frac{\partial v_{t|t-j}}{\partial P_{t-j}^*}\right)(P_t^* - P_{t-j}^*) = \sum_{i=0}^{+\infty} \beta \overline{A}_{t+i|t-j} \left((1-\varepsilon_{t+i}(1-\frac{W_{t+i}}{P_{t-j}^*}))D_{t+i}\right)\frac{1}{X_{t+i}}(P_t^* - P_{t-j}^*)$$

where $\overline{A}_{t+i|t-j} \equiv \prod_{k=0}^{i} \alpha_{t+1+k|t-j}$. Finally, the current price P_t is a weighted average of past P^*_t's :

(B11) $$P_t = \sum_{h=0}^{+\infty} \omega_{t|t-h} P^*_{t-h}$$

where the share of vintage h prices in the total of period t prices, $\omega_{t|t-h}$, is recursively given by:

(B12) $$\omega_{t|t-j} = \alpha_{t|t-j}\omega_{t-1|t-j}$$

The simulation follows these steps: *i)* a matrix of probabilities $\{\alpha_{i|j}\}$ is assumed as an initial guess; *ii)* the path for P_t^* is derived using (7); *iii)* the composition of firms by price vintage $\omega_{t,j}$ is derived using (B12), given beginning-of-period values; *iv)* the path for P_t is derived, as a weighted average of past P_t^*'s, using (B11); *v)* the increase in profits resulting from price adjustment is derived using (B10); *vi)* a new matrix $\{\alpha_{i|j}\}$ is derived using (B3); the process is iterated until convergence.

To avoid the curse of dimensionality, a steady state solution is first derived, assuming marginal cost grow at a constant rate and keeping constant φ_t and ε_t through time (without loss of generality, φ_t can be normalised to 1). Subsequently, we introduce in period *t0* the shocks discussed in the text. In the first simulation ("menu cost") we set $\alpha_{t0|t0-j}=0$ (for all *j*'s); in the second simulation ("market power") we set $\varphi_{t0}=\gamma$ and determine ε_{t0} according to the expression in footnote 28 above. In each simulation, we re-estimate the elements $\alpha_{t|t-j}$'s in a sufficiently large neighbourhood of *t0*, i.e. for $t \in (t0-c, t0+c)$,[34] while keeping the remaining elements in $\{\alpha_{i|j}\}$ at their steady state value.

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[34] We first try with c= 1 and then increase c until the estimates of $\alpha_{t|t-j}$ are no longer affected.