BANCA D'ITALIA

Temi di discussione

del Servizio Studi

Consumer price setting in Italy

by Silvia Fabiani, Angela Gattulli, Roberto Sabbatini and Giovanni Veronese

Number 556 - June 2005

The purpose of the Temi di discussione *series is to promote the circulation of working papers prepared within the Bank of Italy or presented in Bank seminars by outside economists with the aim of stimulating comments and suggestions.*

The views expressed in the articles are those of the authors and do not involve the responsibility of the Bank.

CONSUMER PRICE SETTING IN ITALY

Silvia Fabiani*, Angela Gattulli*, Roberto Sabbatini*and Giovanni Veronese*

Abstract

This paper investigates the microeconomic behaviour of consumer prices in Italy using the individual price records underlying the Italian CPI dataset collected by Istat. We discuss how to analyse price stickiness using such a detailed database and compute a quantitative measure of the unconditional degree of price rigidity in the Italian economy. The analysis focuses on the monthly frequency of price changes and on the duration of price spells, with a sectoral breakdown as well as with a classification by type of outlet.

Prices are in general found to be rather sticky, remaining unchanged on average for around 10 months; price spells last longer for non-energy industrial goods and services, much less for energy products. Prices are revised more frequently upwards than downwards, while the size of price changes is quite symmetric. Price stickiness is found to be less marked in large modern stores than in smaller traditional shops. Price changes display considerable synchronisation, in particular in the services sector. The average frequency of price changes and the probability of observing a price change over time and across items are positively related to headline inflation and increases in VAT rates and negatively related to the share of attractive prices. These findings are consistent with the ones reported in similar national studies for other countries of the euro area, which were conducted by the National Central Banks within the Eurosystem Inflation Persistence Network.

Keywords: consumer prices, nominal rigidity, frequency of price change.

JEL codes: D21, D40, E31.

* Bank of Italy, Economic Research Department.

Contents

1. Introduction[1]

Firms' price setting behaviour has long been investigated in the economic literature. The implications of nominal price stickiness for the short-term non-neutrality of monetary policy are widely acknowledged in macroeconomics; in particular, the dynamics of the response of output, inflation and employment to monetary policy shocks depend on both the degree and the characteristics of nominal price rigidities in the economy (see, for instance, Wolman, 1999).

The recent theoretical literature on optimal monetary policy has emphasised that monetary policy should aim to stabilise those prices that are adjusted less frequently, since they can be expected to become more misaligned in an environment that requires prices to move in either direction (Aoki, 2001; Benigno, 2004). In a framework with positive aggregate inflation in which not all prices are adjusted instantaneously, the consequent relative price distortions may lead to inefficient sectoral allocation of resources. Hence, as pointed out in Woodford (2003), a justification for an approach to monetary policy that aims at price stability is found when taking into account the delays of adjustment of wages and prices.

One of the fundamental assumptions underlying the so-called New-Keynesian literature (Goodfriend and King, 1997) is that prices are set in a staggered fashion. This is coupled with the assumption that individual suppliers can take pricing decisions as they have a certain degree of market power. In this setting a supplier failing to adjust its price in response to changed market conditions does not suffer an infinite loss.

Despite the recognised relevance of the issue of price stickiness, the related empirical evidence is rather scant and is based mainly on macroeconomic data (Gali and Gertler, 1999; Gali *et al*., 2001), which might, however, not capture important aspects of price behaviour due to their aggregate nature.[2] Very little evidence is, instead, available at the very micro level, that is, based on the direct investigation of elementary prices. Indeed, the number of empirical studies of firms' pricing behaviour based on quantitative micro data is limited, mainly because such data, typically collected by the national statistical institutes (NSIs) for the computation of price indices, are not

---

[1] This study is part of a project of the Eurosystem (*Inflation Persistence Network* – IPN). We are grateful to Istat for having allowed the use of disaggregated data, to Federico Polidoro and Roberto Monducci for their comments on a preliminary version of the paper, to Enzo Agnesse for his help with the dataset, to the participants at the meetings of the IPN, in particular Stephen Cecchetti, Hervé Le Bihan and Emmanuel Dhyne, for their helpful comments at different stages of this research. We are also indebted to an anonymous referee for his/her comments. The views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank of Italy.

[2] This is certainly a limitation of these macroeconomic models, which pay considerable attention to the micro-foundations and in formalising the optimising behavior of agents, but may still be hence unsuited for practical policy analysis (Woodford, 2003).

publicly available. For this reason, most of the available studies investigate pricing behaviour referring to a specific firm or market (Cecchetti, 1986; Kayshap, 1995; Genesove, 2003; Kackmeister, 2001; Konieczy and Skrzypacz, 2005; Lach and Tsiddon, 1992 and 1996; Ratfai, 2005).[3]

The recent papers by Bils and Klenow (2004) and Klenow and Kryvstov (2005) represent the first attempts to assess the degree of price stickiness in the United States on the basis of the micro data underlying the U.S. Consumer Price Index (CPI). Their findings stimulated an extensive project on micro data analysis also within the Eurosystem, which launched the Inflation Persistence Network (IPN) at the beginning of 2003 with the aim of studying the behaviour of inflation and prices in the euro area on the basis of a wide range of quantitative and qualitative evidence. Thanks to this project, most euro area countries (Austria, Belgium, Finland, France, Germany, Italy, Luxembourg, Netherlands, Portugal and Spain) have had the unique opportunity to access the micro consumer price data collected by NSIs and to produce evidence on price setting behaviour (Alvarez and Hernando, 2004; Aucremanne and Dhyne, 2004; Baudry *et al*., 2004; Dias *et al*., 2004; Jonker *et al.*, 2004; for a survey see Dyhne *et al*., 2004).

In this paper we present the analysis of consumer price behaviour in Italy, using for the first time a large fraction of the data underlying the Italian CPI.[4] The approach adopted is essentially statistical. We assess the degree of price stickiness computing the frequency of price changes (defined as the monthly percentage of price changes), with reference to different items and types of outlets. We produce a quantitative measure of the unconditional degree of nominal price rigidity in the Italian economy and investigate the extent of sectoral differences. The available micro data also allows investigating other key features of price setting (Taylor, 1999), such as the existence of downward price rigidities and the degree of synchronisation of price movements. The related findings contribute to discriminate between the main theories of pricing behaviour, in particular between time and state-dependent rules.

The paper is structured as follows. Section 2 describes our database. Section 3 defines the main statistics on which our empirical analysis of the characteristics of price behaviour is based; the

---

[3] For a survey of the existing empirical work on price stickiness, see Wolman (2000) and Taylor (1999).

[4] Additional evidence for Italy, complementary to that described here, is provided by two companion studies, which investigate price-setting from a different perspective. The first (Fabiani *et al*., 2003) analyses the statistical properties of the cross-sectional distribution of sectoral price changes, using official price *indices* and applying the methodological approach proposed by Ball and Mankiw (1994). The second (Fabiani *et al*., 2004) investigates firms' pricing practices on the basis of qualitative information collected through a survey of a sample of Italian firms (following the work of Blinder, 1994; Blinder *et al*., 1998; Hall *et al*., 2000).

issue can be approached by adopting either a frequency or a duration approach. Section 4 presents the empirical results. Section 5 concludes.

2. The dataset

We use the monthly data collected by the Italian statistical institute (Istat) to compute the Italian CPI. These are the prices actually observed by the surveyors when they visit stores. Overall, for each month the Istat dataset averages 300,000 price quotes for some 550 COICOP product categories, which are mainly collected by surveyors in 87 municipalities throughout Italy and, for a subset of products, directly by the Central Bureau of Istat.

2.1 Sample coverage

This paper is based on only a fraction of the Istat dataset, for the period January 1996-December 2003. Our database consists of 750,000 elementary price quotes referred to a specific item and brand, sold in a given outlet, for 48 items in the Italian CPI. We focus only on the 20 regional capitals and disregard data related to the rest of the country used in computing the official overall index. The sample accounts for almost 20% of the CPI basket in terms of the product weights in the *Classification of Individual Consumption by Purpose* (COICOP; see Table A1.1 in Appendix 1). The subset was selected in agreement with the other central banks of the euro area doing similar research within the IPN, to ensure some cross-country comparability (Dyhne *et al*., 2004). By restricting the analysis to a subset of the CPI, we lose in terms of coverage but gain in accuracy and more careful preliminary treatment of the data.

The representativeness of our sub-sample can be assessed in several ways. First, the comparison between the general CPI and the index referring only to the 48 goods and services we consider shows that inflation trends are well captured by our sub-sample, especially if re-scaled weights are used[5]; the correlation coefficient between the two series, computed over the years 1997-2003, is equal to 0.85.[6] Second, to check the robustness of our results with respect to the geographical selection, we compare the overall inflation rate with the inflation rate calculated by

5 In order to aggregate our elementary data we can rely on two different sets of weights. In one, each item has the same weight as in the overall CPI basket; in the second, these weights are re-scaled to sum up to the overall weight of all the items included in the same category of expenditure. For instance, if only 2 out of 11 unprocessed food products are considered in the sub-sample, under the first option the sub-index "unprocessed food" in our sub-sample will have a weight given by the sum of the two items, and this can differ substantially from its weight in the overall basket. Under the second weighting scheme, each weight is re-scaled so that the overall importance of unprocessed food in our sub-sample is equal to that in the overall index (i.e., the selected products are assumed to be fully representative of all unprocessed food goods).

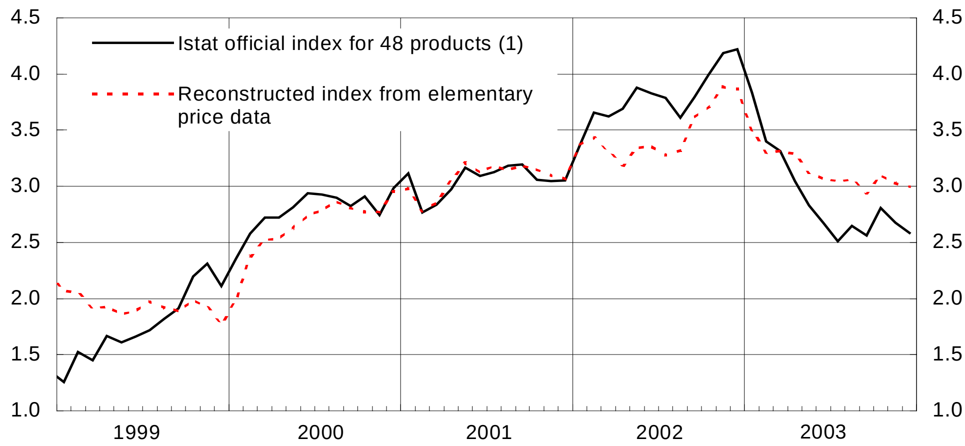
6 The differences are most pronounced in 1999, 2000 and 2002, largely because of developments in energy prices.

aggregating the CPIs for the 20 regional capitals.[7] We find that the latter represents national inflation trends quite well. Finally, we compute the monthly inflation rate on the basis of the elementary prices in our dataset. Figure 1 confirms that our "reconstructed CPI" closely mirrors that based on aggregation of the official Istat indices for the same 48 items; the correlation coefficient between the two series is equal to 0.92.

All in all, our sub-sample of elementary prices captures the pattern of those same items in the overall CPI and, more importantly, tracks headline inflation quite well.

The items in our dataset are classified according to two different criteria. One is the standard classification into the categories of expenditure typically used by statistical institutes (COICOP). The second is the five-category breakdown used by the Eurosystem: non-energy industrial goods, processed food, services (the so-called "core components"), unprocessed food and energy products.

Figure 1 – 48 items inflation – Official and reconstructed price index from elementary price data in our sub-sample
(percentage changes on the year-earlier period)



Source: Based on Istat data.
(1) Each of the 48 items has the same weight as in the overall CPI basket and the price index considered is the official one released by Istat.

The proportion of processed food products and, in particular, of non-energy industrial goods in our sub-sample is lower than in the overall CPI basket (Table 1), while that of services is higher. With reference to the COICOP classification, the percentage of items under "Food and non-alcoholic beverages" is lower than in the overall CPI basket, whereas that of "Restaurants and hotels" is higher. Finally, three important categories of expenditure – "Health", "Communication"

[7] Up to the end of the nineties Istat computed, for regional capitals, the index of consumer prices for worker and employee households, whose evolution mirrors that of the CPI.

and "Education" – are not represented in our sub-sample; these are the categories that show the greatest incidence of regulated prices, which we decided not to consider.

Table 1 – Classification of the elementary price quotes by type of items

	Number of obs.	Percentage	Percentage weight of the items included in our sample in the CPI basket (2002=100)		Percentage weight in the official CPI basket (2002=100)
			weights	weights re-scaled to 100	
COICOP category					
01 - Food and non-alcoholic beverages	195,452	26.3	1.7	8.5	15.9
02 - Alcoholic beverages and tobacco	98,819	13.3	0.8	4.1	2.7
03 - Clothing and footwear	87,851	11.8	1.5	7.2	10.5
04 - Housing, water, electricity, etc.	26,724	3.6	1.3	6.5	9.2
05 - Furnishings, household equipment, etc.	48,871	6.6	2.6	12.8	10.2
06 - Health	-	-	0.0	0.0	7.3
07 - Transport	63,349	8.5	4.1	20.1	13.2
08 - Communication	-	-	0.0	0.0	3.2
09 - Recreation and culture	83,385	11.2	0.5	2.6	8.4
10 - Education	-	-	0.0	0.0	1.1
11 - Restaurants and hotels	76,727	10.3	6.5	32.1	10.9
12 - Other goods and services	62,988	8.5	1.2	6.1	7.3
Type of product					
Unprocessed food	94,023	12.6	0.9	4.3	6.9
Energy	32,374	4.4	2.9	14.2	5.9
Processed food	204,376	27.5	1.7	8.2	11.7
Non-energy industrial goods	221,787	29.8	3.4	16.6	35.3
Services	191,606	25.8	11.5	56.7	40.1
Total	**744,166**	**100.0**	**20.2**	**100.0**	**100.0**

Source: Based on Istat data.

2.2 The information available in the dataset

Crucial to our study is characterising the elementary price records that underlie the empirical analysis. This hinges on the "non-price" information (metadata) collected together with the price quotes. The Istat dataset has a wealth of information pertaining not only to prices, but also to the type of product and of outlet as well as to the causes of possible replacements.

By product we denote each of the items included in the CPI classification and in our sub-sample. Take, as an example, "Coffee". For each product, several price quotes are available, each referring to a specific variety and brand, sold in a specific outlet. Accordingly, by elementary product we denote a specific variety and brand of the product, sold in a specific outlet, in a given town. Each elementary product has an elementary price quote at a particular time *t*. In terms of our example, by elementary price quote we mean the price of, say, a "1 kilogram package of coffee, of brand X, sold at time *t* in store Y, located in town Z".[8] In our analysis it is crucial to follow the exact same elementary price quote through time; this is possible using the information associated with each quote (see Table A1.2 in Appendix 1).

Our dataset includes 750,000 elementary price quotes (8,000 quotes each month). Most of them cover the entire period from January 1996 to December 2003.[9] Istat's sampling procedure aims at maintaining within a given base period[10] a balanced sample of product price quotes. For example, in the collection of price quotes for "packaged coffee" for the city of Turin, Istat may collect 30 different elementary quotes for every month in a given base period (e.g. between December 2001 and December 2002). This method requires the "forced" replacement of some products when, say, the brand in question happens to be unavailable on shelves.[11] Forced replacements are not to be confused with replacements due to the standard turnover of outlets and products that can occur at each base-year change and due to a rotation in the sample ("optional" replacements; see below).

The metadata at our disposal is sufficiently rich to detect product replacements, which occur for various reasons: the product is no longer sold in a given store (a new brand of coffee is now sold); the product is temporarily unavailable on shelves; in this case the previous month's price quote is imputed, but such imputation can only last for at most two consecutive months; the shop has closed; the periodical rotation of outlets that occurs when Istat re-bases the CPI.

The first three are "forced" replacements. In these cases the surveyor replenishes the sample by collecting the price of an elementary product with similar characteristics (for instance, another brand of coffee in the same shop); if a shop was closed, another outlet with similar characteristics in the same area is included. The fourth case instead leads to an "optional" replacement, since Istat deliberately introduces a change to maintain a representative of sample. From the metadata we know that a replacement has occurred, but we cannot tell whether it was forced or optional. As will

---

8 At this level of disaggregation, the elementary price quotes are not associated with specific weights in the basket. Istat, in computing the CPI, aggregates elementary price quotes for each good or service, collected in different stores in a given town, first in the so-called "elementary aggregate index"; namely, for each municipality and for each product, the geometric mean of the ratios between each current quote and the base-period quote is used to calculate the "municipal product index". Then, using a provincial weight structure, Istat applies weighted arithmetic averages to produce "national product indices". Finally, the CPI results from a weighted average of the national product indices, where the weights are derived from the yearly Survey of Final Household Consumption.

9 To compare prices over time with the euro cash changeover, we converted into euro all prices before January 2002. When comparing prices between January 2002 and December 2001, figures were rounded to the 2nd decimal place, as required by Eurostat regulations.

10 The base period for the CPI is annual, from December to December.

11 Istat signals replacements with various flags, according to the type of replacement enacted by the price collector (variety-brand-size-outlet). If the product package size changes (for instance from ½ to 1 kilogram) we consider the product to be the same and, accordingly, we use the collected price per unit instead of the actual package price.

become clear in the empirical section, this requires us to make some auxiliary assumptions when dealing with the problem of censored observations.[12]

Some of our products – essentially clothing and footwear – are subject to significant price swings due to temporary sales or other special offers. Since January 2002 Istat has been following the European Commission Regulation on the treatment of sales prices in the Harmonised Index of Consumer Prices (HICP), which requires collection also of the prices subject to temporary reductions. But for the national CPI Istat has continued to disregard reductions lasting less than a month. So after 2001 the observations for products on sale have included both full price and sales price for each month. For meaningful intertemporal comparisons, we have ignored sales prices. This may produce downward bias in the estimate of the frequency of price changes.

We selected products that would always be available throughout the year. Thus in categories with strong seasonal dependence, such as clothing or unprocessed food, we took only highly standardised items (e.g. men's shirts, fresh tomatoes), in stock all year long.

As for type of outlet, around 60% of our elementary price quotes come from traditional stores, such as corner shops and other small sized stores (Table 2); this share is in line with that of the whole CPI basket.

Table 2 – Classification of the elementary price quotes by type of outlet and frequency of price collection
(percentages)

Distribution channel	**Percentage of the elementary price quote in our sample**	
Modern	27.0	
Traditional	56.8	
Other	16.2	
Total	100.0	
Category of expenditure	**Frequency of price collection (%)**	
	Monthly	Quarterly
Unprocessed food	100	0
Energy	100	0
Processed food	100	0
Non-energy ind.goods	60.3	39.7
Services	46.8	53.3
Total	74.4	25.6

Source: Based on Istat data.

Finally: (i) none of our 48 items has regulated price, with the partial exception of "taxis", for which fares are usually set by the local municipality (hence, only one price per town is collected);

[12] The problem of censoring arises when the "true" time of beginning (ending) of the first (last) price spell does not correspond to the one observed in the dataset, since it occurs before (after) the first (last) price collection. A detailed discussion of censoring is in Section 4.1.

(ii) the periodicity of data collection can be either monthly, quarterly or twice a month (around 75% are monthly quotes; see Table 2). For prices observed quarterly, Istat leaves the price unchanged except in observation months (“carry forward” approach); for those collected twice a month (products with high price variability, such as some unprocessed food and energy products) an average monthly price is computed.

2.3 Basic definitions

We follow the notation introduced by Baudry *et al*. (2004) to provide the basic definitions underlying our empirical analysis. We define:

- Elementary price quote: $P_{jl,t}$, for product j (j =1,.., n_j, where n_j is the total number of products; in our analysis n_j=48), sold in a specific outlet l (l = 1,…, n_l) and observed at time t (t= 1,…T). An elementary product is therefore defined by the pair (j,l).

- Price spell: an uninterrupted sequence of unchanged price quotes associated with the elementary product (j,l), that is the sequence $P_{jl,t}$, $P_{jl,t+1}$,…, $P_{jl,t+k-1}$, with $P_{jl,t+s} = P_{jl,t}$ for $s=1,\ldots,t+k-1$. A price spell is therefore an episode of fixed price, which can be fully described by three elements: the date of the first quote (t),its price ($P_{jl,t}$) and the duration of the spell (k), that is $\{P_{jl,t}, t, k\}$.

- Price trajectory: a sequence of s successive price spells for the product (j,l), that is ($\{P_{jl,t}, t_1, k_1\}$,$\{P_{jl,t+k1}, t_2, k_2\}$, $\{P_{jl,t+k1+k2}, t_3, k_3\}$,….,$\{P_{jl,t+k1+\ldots+ks-1}, t_s, k_s\}$). By computing $L_{jl}= (k_1 +\ldots +k_s)$ we can define the trajectory length for the elementary product (j,l), which is just the sum of the individual price spells.

These definitions are represented in Figure 2. Trajectory 1 can be described as:

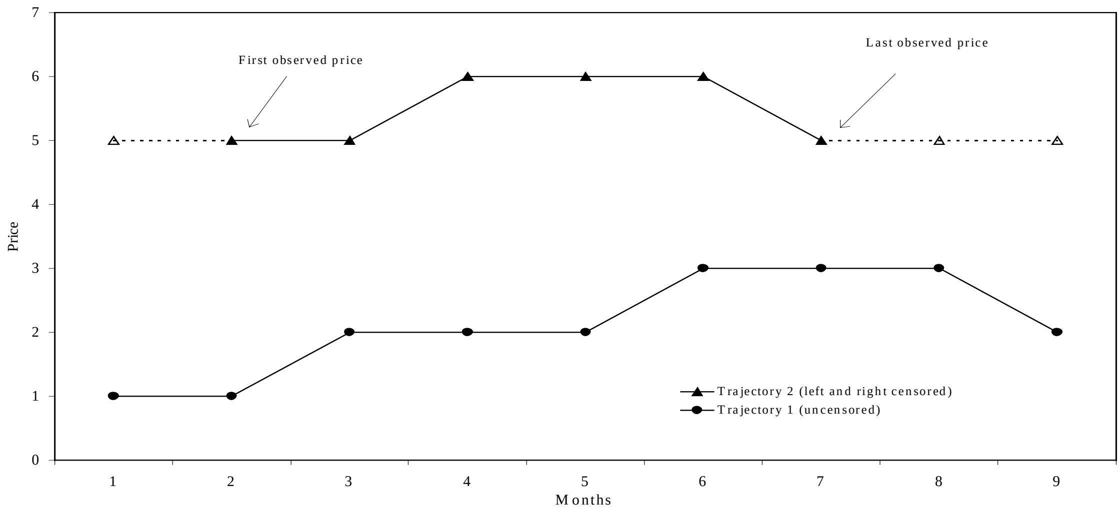
$$(\{P=1, t_0=1,k_1=2\};\{P=2, t_0=3, k_2=3\};\{P=3, t_0=6, k_3=3\};\{P=2, t_0=9, k_4=1\})$$

Similarly, Trajectory 2 can be described as

$$(\{P=5, t_0=1, k_1=3\};\{P=6, t_0=4, k_2=3\};\{P=5, t_0=7, k_3=3\})$$

Trajectory 1 has four price spells (with durations of 2, 3, 3 and 1 month), while Trajectory 2 has three price spells (all with a duration of 3 months).

Figure 2 -A typical price trajectory



3. Price duration and the frequency of price changes

The first aim of our analysis is to characterise price behaviour across products and outlets. Two methodological approaches can be used: the first is based on duration between two price changes, the second on the frequency of price changes. The former, which is denoted as “duration” approach, measures the duration of a price spell (the number of months in which a price remains unchanged) directly and derives an “implied” frequency of price changes indirectly, as the inverse of duration. The “frequency” approach first computes the frequency of price changes as the proportion of all quotes that change in a given period and then derives an indirect measure of “implied” duration of the spells. If the sample does not have censored price spells, the two approaches give the same results.

Under both approaches, in order to determine synthetic estimates of average duration and of average frequency of price changes we need to aggregate statistics on elementary products. As Baharad and Eden (2004) show the method of aggregating can change the results for the overall statistic substantially. Here we take a “bottom-up” approach (see below), assuming that products are more homogeneous within certain subgroups (for instance in a given price trajectory or in a given product category).

3.1 The duration approach

The duration approach is illustrated in Figure 2. The average duration of price spells in Trajectory 1 (the uncensored one) is the simple average of the spells durations (2, 3, 3 and 1). The same result is obtained by dividing the number of observations (9) by the number of price spells (4).

An important issue arises in aggregating price spells over different trajectories. Consider again Figure 3; it is clear that we can either ignore the information on the trajectory to which each spell belongs and then compute a simple average over the entire pool of spells, or else take the information on the trajectory into account. The latter approach entails first computing average spell duration within trajectories and then aggregating over trajectories. While the results in this example differ, they will be similar if the number of price spells is large enough to presume that duration processes are fairly homogenous.

We can compute average duration using a number of different formulas. If we pool all price spells, we obtain the unweighted average duration of all price spells (all spells have equal weight), that is:

$$\bar{d} = \frac{1}{N_{spells}} \sum_{j=1}^{n_j} \sum_{s=1}^{N_{sj}} d_{js} = \frac{N_{elementary\ price\ quotes}}{N_{spells}} \qquad (1)$$

where j refers to the product (n_j =48), s identifies a specific spell, d_{js} is the duration of spell s of product j, N_{sj} refers to the number of price spells for product j.

Obviously, products that exhibit more frequent price changes and thus more spells (typically, energy and unprocessed food) will have a greater weight in the formula for $\bar{d}$. This problem can be overcome by computing the unweighted average duration of price spells, averaged by product, obtained by first computing the average duration for each product j, $\bar{d}_j = \frac{N_{elementary\ price\ quotes\ for\ product\ j}}{N_{spells\ for\ product\ j}}$ (averaging over all its spells, across outlets and time), and then averaging it over products, that is:

$$\bar{\bar{d}} = \sum_{j=1}^{n_j} \frac{1}{n_j} \bar{d}_j \,. \qquad (2)$$

If CPI product weights (ω_j, $j=1,\ldots,n_j$) are used in calculating the previous statistics, we obtain the weighted average duration of price spells, averaged by product, defined as:

$$\bar{\bar{d}}^{\omega} = \sum_{j=1}^{n_j} \omega_j \sum_{j=1}^{N_s} \bar{d}_j \qquad (3)$$

A further possibility is first averaging price spells in a given trajectory and then averaging over products, obtaining the unweighted average duration of price spells, averaged by individual trajectory:

$$\overline{\overline{d}}^{traj} = \sum_{j=1}^{n_j} \frac{1}{n_j} \sum_{r=N_{traj;1}}^{N_{traj;j}} \frac{{}_j\overline{d}_r^{traj}}{N_{traj;j}} \qquad (4)$$

where $N_{traj;j}$ is the number of price trajectories for product j, and ${}_j\overline{d}_r^{traj}$ represents the average duration of the spells in the r^{th} trajectory of product j.

Similarly, the weighted average duration of price spells, averaged by individual trajectory is given by:

$$\overline{\overline{d}}^{traj,w} = \sum_{j=1}^{n_j} \omega_j \sum_{r=N_{traj,1}}^{N_{traj,j}} \frac{\overline{d}_r^{traj}}{N_{traj;j}} \qquad (5)$$

In our analysis we mostly aggregate the individual durations according to expression (3), but as a robustness check we also present those based on (5).

The main advantage of the duration approach consists in the possibility of reporting the full distribution of price durations in each period (in our analysis we focus on the median and the inter-quartile range of the distribution of durations). Moreover, only under this approach we can compute the hazard and survival functions. In the empirical section we present the standard non-parametric estimate of the hazard function based on the Kaplan-Meier product limit estimator. The estimator is based on the number of spells ending at period t, h_t ("failures"), divided by the number of spells still in the risk set (R_t, which at each time t consists of the number of spells lasting *at least* until t):

$$\hat{\lambda}(t) = \frac{h_t}{R_t} \qquad (6)$$

3.2 The frequency approach

The frequency approach, used by Bils and Klenow (2004) among others, defines the frequency of price changes, given for each product j as the number of price changes in each period (denoted by NUM_{jt}, for the j^{th} product at time t, where t=2,…T) over the total number of price quotes for that product in that period (DEN_{jt})[13], that is:

[13] See Appendix 2 for details.

$$F_j = \frac{\sum_{t=2}^{T} NUM_{jt}}{\sum_{t=2}^{T} DEN_{jt}} \tag{7}$$

Under certain conditions of stationarity of the process determining price spells both cross-sectionally and over time (see Lancaster, 1990)[14], the data on average spell duration for product *j* can be recovered as the inverse of the frequency of price changes (<u>implied average spell duration</u>):

$$\overline{T}_j = \frac{1}{F_j} \tag{8}$$

Equation 7 and consequently equation 8 are appropriate when retailers change their price only at discrete intervals. We can also assume, like Bils and Klenow (2004), that prices are changed in continuous time: in such an environment, and under a constant hazard model, the implied average duration can be computed by

$$\overline{T}_j = \frac{-1}{\ln(1-F_j)} \tag{9}$$

and the median time elapsing between two price changes is equal to:

$$T_j^{50} = \frac{\ln(0.5)}{\ln(1-F_j)} \tag{10}$$

As these formulas show, the frequency approach has some clear advantages. First it does not require a long span of data; in principle, the observation window may even be shorter than the average duration of a price spell. Second, if some months are deemed to be exceptional due to specific events (like a VAT change), in principle they could be easily be dropped in computing the average frequency of price changes. Finally, the approach also allows to use the maximum amount of information from the dataset, not discarding all the observations from censored price spells (as in the duration approach; see below) but only those observations that involve transitions from or to unobserved prices.

Nonetheless, as pointed out in Baudry *et al.* (2004), the frequency approach too requires sufficiently homogenous products (and sub-periods). In our analysis this means that in aggregating frequencies over products we first compute statistics at the product level and then average different goods and services; at this stage, in computing aggregate weighted statistics (i.e. averaging the F_j) we use CPI weights.

[14] The stationarity assumption regards the process determining the sample of price spells: this ensures that the probability of a price spell's starting in each period is stable over time. Lancaster (1990), in the context of the theory of renewal processes, provides a formal argument for the convergence of the inverse of the frequency of price changes with the mean spell duration.

The frequency approach can also be used to calculate the frequency of price increases (decreases) for each product, where the expression in the numerator of equation 7, NUM_{jt} is replaced by $NUMUP_{jt}$ ($NUMDW_{jt}$), indicating the number of positive (negative) price changes in each period *t*.[15]

4. The empirical analysis

4.1 How to deal with censoring

On these definitions we now illustrate two main problems faced in the empirical analysis - censoring and attrition - both arising in the process of price collection.

The issue of censoring can be introduced by turning again to the example in Figure 2. The first price spell in Trajectory 2 is left-censored because collection only begins at time t=2, even though the price was also available at time t=1; the last is right-censored, as collection ends in period t=7. Hence, the true time of beginning (ending) of the first (last) price spell does not correspond to that one observed in the dataset, as it comes before (after) the first (last) price observation.[16] It is worth remarking that if the start of the collection period coincided with the first time a price is set in the outlet (as in Trajectory 1), then the price trajectory would not be censored; this occurs, for instance, if a shop sells a product for the first time in the same month as that in which the statistical agency starts to observe it. As a consequence, any measurement of the average duration of the price spells in such a trajectory that ignored censoring in the first and in the last spell would be biased.[17]

In practice, the frequency of price changes can be computed on the basis of different sets of price quotes. Two "extreme" strategies are the following.

<u>Strategy 1: No censoring</u>: the statistics completely disregard the issue of censoring, therefore using all price spells.

<u>Strategy 2: Full censoring</u>: the first and last price spells within each price trajectory are considered as censored (thus obtaining full symmetry between the number of right- and left-censored price spells), and the statistics on price duration and on the frequency of price changes are computed only on the basis of the remaining uncensored observations. This approach, as is pointed

---

[15] See Appendix 2 for details.

[16] In terms of the example of product "Coffee", in the case of Trajectory 1 we are considering the price of, say, 1 kg. of coffee of a specific brand sold for the first time in period t=1 in a given store, or which is no longer sold in that shop from time t=9 onwards. In the case of Trajectory 2, this product was sold before t=1 and it will continue to be sold after t=9.

[17] For an extensive review of duration analysis and the various methods of dealing with censoring, see Kiefer (1988).

out in Baudry *et al*. (2004) and in Aucremanne and Dhyne (2004), has the disadvantage of excluding far too many spells, especially when considering products (such as clothing and footwear) subject to very few price changes in their lifecycle.

Intermediate strategies to overcome these difficulties can be devised. If the price of a specific product can always be observed in the same outlet and for the whole time interval, then at most the spells at the start and at the end of each trajectory are censored; our dataset is truncated on the left in January 1996 and on the right in December 2003. In practice, however, when the statistical institute re-bases the index (every December for the Italian CPI) generally drops and adds some products and outlets. This cause of censoring is motivated by statistical considerations and can be considered as an optional replacement (a choice by the statistical institute).

The price collector may also have to replace a product variety or an outlet in order to continue to follow a given product if the price observed to that point ceases to be available (on shelf) or the outlet has closed. This situation is also known as "attrition" and gives rise to a type of censoring that is considered as a forced replacement, since the statistical agency must undertake specific actions independently of its wishes.

In practice, on our information the distinction between forced and optional replacements is not possible, since we have only a flag signalling whether there was a product substitution, independent of sources. We shall soon see that the two different sources of censoring can affect how we estimate our statistics. We can distinguish between the two situations using two *auxiliary* assumptions:

Assumption 1 - Nature of the replacement. Since the CPI is re-based in December each year, we assume that the replacements effected in January are optional, part of the CPI revision and do not coincide with the start/end of the price spell. Replacements in other months are assumed to be forced.

On this basis we devise an intermediate strategy to compute the frequency and the size of price changes.

<u>Strategy 3: Intermediate censoring.</u> We consider as uncensored also the first and the last price spell of a price trajectory that starts/ends due to a forced replacement (i.e. it occurs in non-rebasing months). Hence, price spells are considered as left-censored only if they belong to a trajectory beginning in January, and right-censored when they are the last in a trajectory ending in December. The main advantage of Assumption 1 is that it increases the number of price spells characterised as uncensored. However, it also introduces an asymmetry in the number of left- and right-censored

spells and in some situations it can be too extreme. In our "coffee" example suppose that up to March a given retailer was selling 1 kg. of "Lavazza" brand coffee which was in the list of prices to be collected in that shop, and that in April he/she has decided to sell only 1 kg. of another brand. In this case the new price is not left-censored, so our assumption is appropriate. But if a shop where the price of 1 kg. of "Lavazza" is observed closes in March and is replaced by a similar shop that also sells 1 kg. of "Lavazza", we consider this new price spell not to be left-censored, though as a matter of fact we are facing a situation similar to that depicted by Trajectory 2 in Figure 2.

In computing the frequency of price changes, the consequences of replacement can be severe for some products, leading to very low estimated frequencies. This holds, in particular, for clothing and footwear items. New fashion models are typically introduced at the beginning of each season; they are sold at a price that remains unchanged until a slightly different model replaces the previous one on the market. Hence, under strategy 3 the price for that product never changes, since as long as the product is on the shelf its price is unchanged. To mitigate this effect, a specific assumption concerning products that are no longer available at some point in time is made:

Assumption 2 - Pseudo price-change: when a new variety of a given product (for instance a new pair of trousers) replaces a variety that is no longer sold and whose price has been collected through a given month, the price of the new variety is likely to be different, and the two prices cannot be directly compared. However, we can assume that the price of the old model would have changed in the same month in which it was replaced.

On this basis we can device a fourth intermediate strategy.

<u>Strategy 4: Intermediate censoring with pseudo price-change.</u> We still rely on Assumption 1 to distinguish between forced and optional replacements, but for the former we now assume that the implied price change (the difference between the prices for the old and the new models, collected in periods t-1 and t) is an actual price change (as if the two prices referred to the same item), which is taken into account when computing the frequency of price changes. Since we are actually referring to two different goods, however, we disregard this information in computing the average size of the price change, which thus remains the same as under strategy 3.

In conclusion, the treatment of censoring can be crucial for the results, though on the basis of the available metadata an optimal strategy does not exist. To assess robustness, we compute all the relevant frequency statistics under the four different strategies described above, focusing in particular on strategy 4.

4.2 Price trajectories and price spells

To allow for the impact of the euro cash changeover on the duration and frequency of price changes, we computed all the main statistics both for the time interval excluding the euro cash changeover (1996-2001; 19,000 observations) and for the entire one (1996-2003; 23,600 price trajectories). The analysis that follows focuses on the former, on the assumption that the euro cash changeover at most exerted a temporary impact on price-setting behaviour.[18]

Table 3 – Duration of trajectories – Period 1996-2001
(*unweighted statistics; months*)

	Number of obs.	Mean	Median	Standard deviation
Total sample	**18,858**	**36**	**30**	**28**
By nature:				
Unprocessed food	1,297	48	46	29
Processed food	6,304	33	25	26
Non energy industrial goods	6,815	30	24	24
Energy	717	43	36	27
Services	3,725	49	47	31
By COICOP:				
Food and non-alcoholic bev.	4,562	40	36	29
Alcoholic bev. and tobacco	3,039	29	25	23
Clothing and footwear	2,364	34	26	26
Housing, water, electricity, etc.	561	45	45	29
Furnishings, household equipment, etc.	1,218	38	32	29
Health				
Transport	1,369	44	37	30
Communication				
Recreation and culture	2,443	31	23	26
Education				
Restaurants and hotels	1,643	43	38	29
Other goods and services	1,659	35	30	27

Source: Based on Istat data.

In general, our elementary observations (trajectories) are quite long; in fact the unweighted mean length of price trajectories is 3 years (Table 3), ranging from an average of 30 months for non-energy industrial goods to 49 for services. As for individual items, some quotes are observed over the entire time horizon (96 months), others for just two months. The median trajectory has a duration of 30 months for the whole dataset, ranging from 24 months for non-energy industrial goods to 47 months for services. With reference to price spells (without any assumption on censoring), their numbers is about four times as numerous as that of price trajectories (Table 4).

Prices in Italy are rather sticky. The weighted average duration of all price spells (equation 1), with no censoring (strategy 1), is 10 months; it falls to 8 months under the assumption of intermediate censoring (strategy 3), and to 6 months under full censoring (strategy 2). Note that there is a substantial heterogeneity in average duration across products. In particular, as for the five sub-indices, average duration is much longer for non-energy industrial goods and services (14 and 15 months, respectively) and very short for energy products (2 months). The longer than expected duration for unprocessed foods (9 months) is entirely due to “meat” and “milk”, whereas “fruit and vegetables” prices have much shorter duration. These results suggest that the impact of censoring on estimated average duration is not dramatic. For the full sample, the difference between the no censoring and strategy 3 estimates is only 2 months. For the whole period 1996-2003, average duration falls to 8 months, suggesting that the euro cash changeover did in fact induce more price adjustments.[19]

Table 4 - Duration of price spells – Period 1996-2001
(weighted statistics; months)

	Number of obs.	Mean	Median	Standard deviation
1. No censoring				
Total sample	**69,308**	**10**	**5**	**13**
Unprocessed food	13,447	9	3	14
Processed food	19,689	9	5	11
Non energy industrial goods	13,505	14	10	13
Energy	14,845	2	1	1
Services	7,822	15	11	16
Averaged by individual trajectory (1)				
Total sample	**18,858**	**14**	**8**	**16**
Unprocessed food	1,297	12	4	18
Processed food	6,304	11	6	12
Non energy industrial goods	6,815	15	10	15
Energy	717	2	1	2
Services	3,725	23	14	22
2. Full censoring				
Total sample	**43,886**	**6**	**2**	**9**
3. Intermediate censoring				
Total sample	**58,397**	**8**	**4**	**11**

Source: Based on Istat data.
(1) See equation (5) in the text.

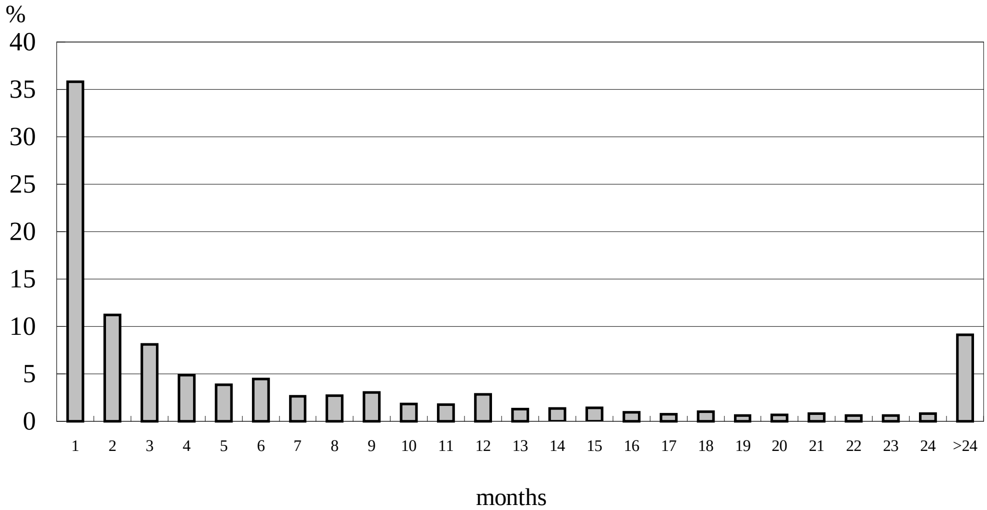
<u>Fact 1</u>*: the weighted average duration of price spells is between 6 and 10 months, depending on estimation strategy. It is longer for non-energy industrial goods and services and very short for energy products.*

The distribution around the mean duration is markedly asymmetrical, as is suggested by comparison with the median of 5 months (2 and 4 by the other two methods; Table 4). For around 35% of the spells the duration is 1 month (Figure 3), as some unprocessed food and especially

[18] We checked this by comparing the results for the two periods; the statistical evidence is available upon request from the authors.
[19] All sub-indices have shorter average duration including 2002-2003.

energy prices change very often indeed. The distribution of spell duration shows also a peak at 6, 9 and 12 months.

Figure 3 - Distribution of spells duration
(*unweighted statistics)*

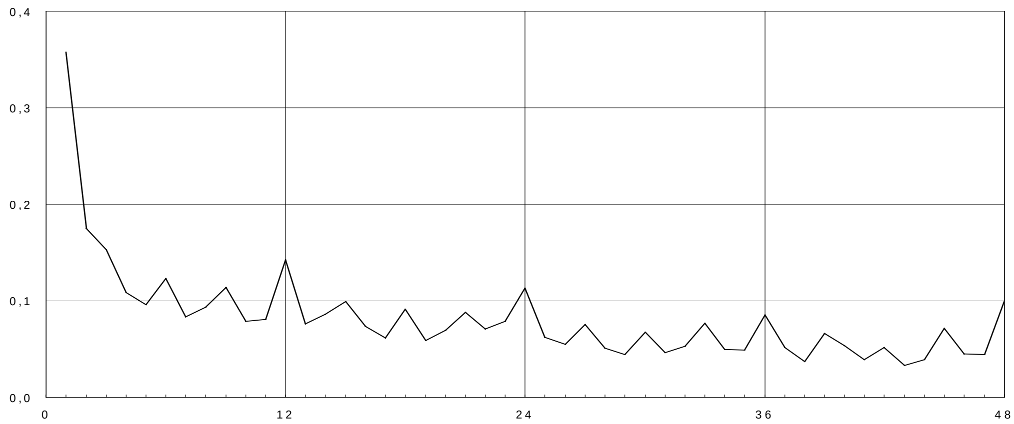


Source: Based on Istat data.

The overall average tends to overweight the products with very short price duration (quite significant in our database), since for given trajectory length more spells are observed. Therefore, we also compute the weighted average duration of price spells averaged by individual trajectories (equation 5 in section 3.1), which increases average duration to 14 months (13 for the whole period;) and the median to 8 months (Table 4).

Finally, we compute the hazard function of price spells for the overall sample (Figure 4), which describes the probability of a price spell to end (vertical axis) in a given month (horizontal axis). The hazard function is a decreasing function of time, as found in other country studies (for instance in Dias, 2004), and it is characterised by local modes at durations of 12 and 24 months, a strong indication that firms apply yearly pricing rules, a result in line with the survey evidence in Fabiani *et al*. (2004). Furthermore, when distinguishing items according to their nature (Appendix 4), we find that the shape of the hazard function is either moderately decreasing, or flat, for services and food products, while it displays an increasing shape for energy products after a 12 months period. Declining hazard functions are clearly at odds with the theoretical pricing rules most widely considered in macroeconomic models. For example in the model by Calvo (1983) firms adjust prices infrequently because the opportunities to adjust their prices are modelled as an exogenous Poisson process; this implies that the theoretical hazard function is constant. Nonetheless, the literature on duration suggests that downward sloping hazards may be due to unobserved heterogeneity in the sample.[20]

Figure 4 – Hazard function for the whole sample (1)



Source: Based on Istat data.
(1) Hazard functions are estimated using the Kaplan-Meier method.

4.3 Price adjustment in Italy: the frequency approach

In this section we describe some basic features of consumer price adjustment in Italy adopting the frequency approach. As already mentioned in the introduction, we investigate the frequency of price changes, their direction and size and their degree of synchronisation. The analysis relies on the statistics described in section 2.2.

The weighted average frequency of monthly price changes is 8.8% with no censoring; it rises to 10% if either of the two intermediate approaches to censoring is adopted (Table 5). The weighted average implied duration, which is approximated by the inverse of the average frequency, ranges from 9 to 11 months. This evidence support the view that prices in Italy are considerably stickier than in the US where, on the basis of the estimates provided by Bils and Klenow (2004), around 25% of consumer prices changes every month.[21] The estimate of the frequency of price changes for Italy is also lower than that for the euro area as a whole (around 15%; Dyhne *et al*., 2004); this suggests a higher degree of nominal price rigidity in our economy than in the rest of the euro area. The percentage for Italy rises marginally when the period covered is extended to include 2002-03, as expected on the basis of the results described in section 3.2.

---

[20] See for instance Heckman and Singer (1985) for a theoretical discussion on this issue.

[21] As stated by Dhyne *et al*. (2004) with reference to the whole set of euro area studies, the higher frequency of consumer price changes estimated for the US compared to the euro area is not due to differences in the consumption structure (Dhyne *et al*., 2004).

It is worth recalling that in the case of the Italian CPI (as in most other countries) the prices of some items are collected only quarterly.[22] This introduces a complication in our analysis, and a possibly downward biased estimate of the frequency of price change. Following Bils and Klenow (2004), we postulate that the monthly frequency of price changes is constant over time within each outlet: this is equivalent to ruling out duration dependence within each quarter, and allows us to measure monthly frequency of price changes from the corresponding observed quarterly statistics.

Fact 2: *On average, in Italy, 9% of prices are changed each month. The implied average duration of price spells is 9-11 months, depending on the estimation method. This result denotes a higher price stickiness in Italy than in the US and, marginally, than in the euro area as a whole.*

The analysis across different goods and services shows pronounced heterogeneity, which makes it difficult to provide a single answer to the question of how often prices change. The disaggregated analysis provides useful insights, however. Under our preferred intermediate strategy (censoring, with pseudo price-change; strategy 4) most energy prices change every month (Table 5). Unprocessed food prices also change quite often, with an implied average duration of 5 months.[23] The prices of other goods and services change less frequently. Processed food's and beverages change, on average, almost once a year, with modest differences between products. The dispersion of the implied average duration of non-energy industrial goods prices is also quite moderate (the average is 17 months). Services prices change very seldom, on average every 21 months, with some remaining fixed for more than 2 years.

Fact 3: *Results on the frequency of price changes are rather heterogeneous across the five product groups. In particular, price changes are very frequent for energy products (every month), while non-energy industrial goods and services prices last longest, about 17 and 21 months, respectively.*

---

[22] These are, in particular, services and some durable goods, around one third of the items included in our sample (see Table A1.1).

[23] The average implied duration of unprocessed food prices is higher than expected; this is due to the behaviour of meat prices, which are classified in the unprocessed food component and exhibit a much longer duration than fruit and vegetables prices.

Table 5 – Frequency and size of price changes and duration - Period 1996-2001

(weighted statistics)

	Frequency of price changes (%)			Duration implied by average frequency (months)			Frequency of price increases (%)			Frequency of price decreases (%)			Average price increase (%)		Average price decrease (%)	
	1. no cens.	3. interm. cens.	4. pseudo price-changes	1. no cens.	3. interm. cens.	4. pseudo price-changes	1. no cens.	3. interm. cens.	4. pseudo price-changes	1. no cens.	3. interm. cens.	4. pseudo price-changes	1. no cens.	3. interm. cens.	1. no cens.	3. interm. cens.
By COICOP:																
1. Food and non-alcoholic bev.	13.9	15.0	14.6	6.7	6.2	6.4	7.8	8.3	7.7	6.2	6.7	6.1	6.2	6.2	-6.4	-6.5
2. Alcoholic bev. and tobacco	7.7	8.9	10.0	12.5	10.7	9.5	4.9	5.4	4.8	2.8	3.5	2.7	7.2	7.0	-7.3	-7.2
3. Clothing and footwear	3.8	3.8	5.4	25.6	25.8	18.0	3.3	3.2	3.2	0.6	0.6	0.6	7.0	6.7	-7.2	-9.1
4. Housing, water, electricity, etc.	26.6	27.7	27.0	3.2	3.1	3.2	15.5	16.4	15.4	11.0	11.3	10.9	3.1	2.9	-2.6	-2.5
5. Furnishings, household equip.	3.1	3.0	4.5	31.6	32.4	21.8	2.8	2.7	2.7	0.3	0.3	0.3	7.2	6.3	-8.3	-8.6
7. Transport	23.8	24.4	24.5	3.7	3.6	3.6	13.2	13.6	13.1	10.6	10.8	10.4	3.6	3.4	-3.2	-2.9
9. Recreation and culture	5.2	5.9	7.7	18.8	16.4	12.5	2.5	2.7	2.3	2.7	3.2	2.4	7.5	8.0	-8.6	-9.4
11. Restaurants and hotels	4.8	6.2	5.7	20.4	15.6	17.0	3.9	4.6	3.8	0.9	1.6	0.9	9.0	9.1	-13.6	-12.9
12. Other goods and services	3.3	3.4	4.3	29.8	28.8	22.7	2.6	2.6	2.5	0.7	0.8	0.7	7.9	7.1	-7.8	-7.4
By Product type																
Unprocessed food	19.1	21.1	19.5	4.7	4.2	4.6	10.1	11.2	10.0	9.1	9.8	9.0	7.3	7.3	-7.8	-8.0
Processed food	7.8	8.4	9.4	12.4	11.4	10.1	5.0	5.2	4.9	2.8	3.2	2.7	6.1	5.9	-6.0	-5.9
Energy	60.5	61.9	60.8	1.1	1.0	1.1	32.9	34.0	32.6	27.5	27.9	27.1	1.8	1.8	-1.6	-1.7
Non-energy industrial goods	3.9	4.0	5.8	24.9	24.7	16.7	3.0	3.0	2.9	0.9	1.0	0.9	6.8	6.5	-7.3	-8.1
Services	3.9	4.8	4.6	25.2	20.5	21.2	3.2	3.7	3.2	0.7	1.1	0.7	9.1	8.9	-12.8	-11.9
Total - 48 products	**8.8**	**9.5**	**10.0**	**10.8**	**10.0**	**9.5**	**5.6**	**6.0**	**5.5**	**3.2**	**3.5**	**3.1**	**7.2**	**6.9**	**-7.9**	**-7.4**

Source: Based on Istat data.

Considering the direction of price changes, the weighted average frequency of price increases ranges from 5.5% to 6.0% depending on the strategy used (Table 5). On average, price reductions are less frequent than price increases, the average frequency ranging from 3.1% to 3.5% (depending on the approach).

When considering the sectoral dimension, under strategy 4 the frequency of price increases is 10% for unprocessed food, 32.6% for energy products and only around 3% for non-energy industrial goods and for services. The asymmetry in the direction of price changes is more pronounced for services and for non-energy industrial goods, for which price increases occur, respectively, four and three times more often than decreases.

Fact 4: Price changes tend to be asymmetric, as one expects given positive inflation; under strategy 4 the average frequency of price increases is 5.5%, compared with 3% for decreases.

The average size of price increase and decrease is about the same (7.2% and –7.9%, respectively; Table 5); this symmetry broadly holds for all components and it is in line with what found for the euro area as a whole (Dhyne *et al*., 2004). This result is robust to sample period and is thus not affected by the euro cash changeover.

Fact 5: The average percentage change is about the same, whether prices are adjusted upwards or downwards.

Table 6 – Monthly frequency of price changes of non-energy industrial goods and food products by type of outlet (strategy 4)
(weighted statistics, percentages)

Distribution channel	Frequency of price changes	Frequency of price increases	Frequency of price decreases
	Period 1996-2003		
Modern	14.5	7.1	5.3
Traditional	9.3	4.9	3.0
Other	9.7	4.2	3.8
	Period 1996-2001		
Modern	13.7	6.7	4.8
Traditional	8.7	4.5	2.8
Other	9.4	3.9	3.9

Source: Based on Istat data.

As for type of outlet, traditional outlets tend to change the price of non-energy industrial goods and food products significantly less often than large outlets, with frequencies of 8.7% and 13.7%, respectively (Table 6). This suggests that differences in the market structure of the retail sector may play an important role in explaining sectoral (as well as cross-country) differences in the observed degree of price stickiness.

Fact 6: *The monthly frequency of price changes differs significantly by type of outlet, being higher for larger stores.*

Table 7 - Synchronisation ratio of price changes, with intermediate censoring and pseudo price-change (strategy 4) - Period 1996-2001
(weighted statistics)

Product	Synchronisation ratio of price changes	Synchronisation ratio of price increases	Synchronisation ratio of price decreases	Average of the synchronisation ratio of price changes by cities *(weighted with the weight of the cities)*	Average of the synchronisation ratio of price changes by cities *(weighted with the number of observations)*
Unprocessed food					
Steak	12.0	11.2	14.3	27.6	29.0
1 fresh fish	14.1	17.5	15.9	42.7	41.5
Tomatoes	23.3	53.3	52.7	53.9	47.5
Banana	12.8	21.8	18.8	36.9	35.2
Energy					
Gasoline (heating)	40.3	56.6	44.4	67.6	69.2
Fuel type 1	52.5	75.6	71.2	70.5	70.0
Fuel type 2	59.8	77.3	73.0	75.4	74.6
Processed food				0.0	0.0
Milk	28.3	31.7	9.0	58.2	58.5
Sugar	13.0	10.2	16.5	30.1	29.9
Frozen spinach	18.9	20.8	7.2	34.7	36.2
Mineral water	34.8	11.5	9.7	41.6	43.3
Coffee	14.5	20.6	10.9	30.0	29.5
Whisky	13.2	15.9	5.7	27.9	28.7
Beer in a shop	11.6	13.0	7.7	29.4	29.2
Non-energy ind. goods					
Socks	14.1	14.7	6.8	35.4	34.3
Jeans	20.7	19.8	8.5	39.7	39.8
Sport shoes	18.6	16.5	13.3	43.5	42.5
Shirt (men)	18.4	15.7	6.6	34.8	35.8
Tiles	31.0	26.4	10.1	53.6	54.2
Toaster	38.4	31.5	20.3	62.4	59.2
Electric bulb	15.1	15.9	8.9	39.9	39.1
1 type of furniture	36.5	26.8	12.5	53.6	55.0
Towel	13.6	13.2	7.7	35.5	35.7
Car tyre	45.8	29.4	27.3	62.3	64.4
Television set	47.4	23.1	29.5	59.3	59.0
Dog food	39.0	27.0	21.3	53.9	50.6
Football	25.2	16.5	13.6	45.2	46.2
Construction game (Lego)	27.5	17.7	11.7	47.3	48.1
Toothpaste	8.3	8.3	6.9	27.5	26.5
Suitcase	33.0	22.9	17.2	53.6	54.5
Services					
Dry cleaning	23.8	20.9	15.2	40.0	44.1
Hourly rate of an electrician	28.6	26.2	12.2	57.3	58.6
Hourly rate of a plumber	29.0	27.3	12.7	61.0	61.0
Domestic services	33.2	32.7	15.9	84.8	70.5
Hourly rate in a garage	28.4	26.1	10.9	54.1	55.1
Car wash	26.1	22.0	13.6	54.1	55.8
Balancing of wheels	25.7	21.1	14.8	51.4	53.9
Taxi	23.9	23.2	-	67.4	70.6
Movie	45.2	28.3	56.3	84.4	84.9
Videotape hiring	19.8	16.2	16.4	46.0	41.2
Photo development	27.8	20.8	18.2	46.7	45.4
Hotel room	24.6	26.5	15.1	55.5	56.1
Glass of beer in a café	12.6	9.4	7.2	37.8	38.7
1 meal in a restaurant	10.5	10.2	11.4	44.0	41.6
Hot-dog	10.2	9.6	8.8	38.8	38.4
Cola based lemonade in a café	11.7	10.8	8.9	35.1	36.6
Haircut (men)	31.9	17.6	9.4	49.2	54.4
Hairdressing (ladies)	17.7	17.3	10.7	44.9	42.8

Source: Based on Istat data.

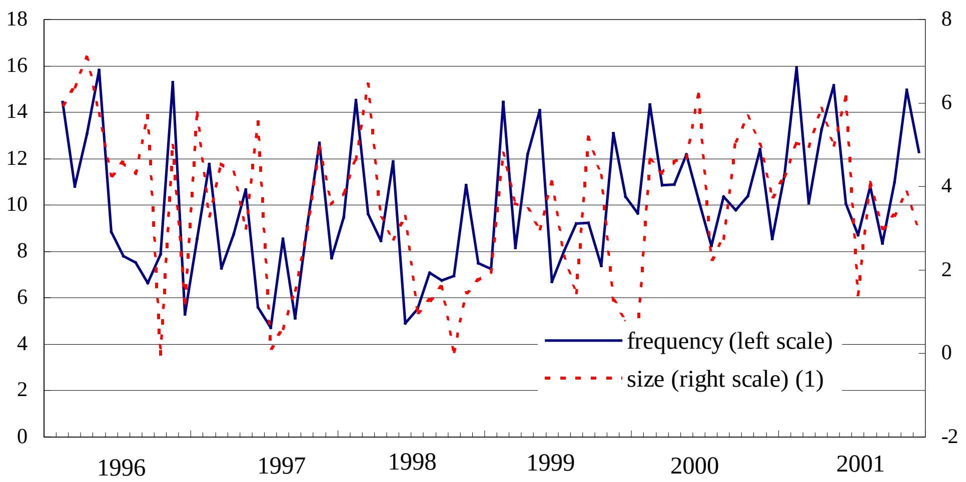
In general the degree of synchronisation of price changes, measured following Fisher and Konieczny (2000), which yields a value of 1 in case of perfect synchronisation, is rather low except for energy prices (Table 7; see Appendix 2 for the formula). However, it is worth remarking that for the other items, the lack of synchronisation may result from the fact that the geographical location of stores is not considered. In fact, price changes for individual products are likely to be more similar within each town than in a pool of observations from all towns. To adjust, we compute our statistics separately for each item and each town and then aggregate the results in an average for the whole country using two weighting schemes, one based on the weight of each town in the national CPI, the other on the price quotes for each item in a town as a percentage of all the observations for that product nation-wide. The results show that for a number of items price movements are actually quite highly synchronised, in particular for services.

Fact 7: Price movements of individual products are not highly synchronised at the national level, but they are much more so if geographical location is taken into account.

4.4 Factors affecting the frequency and the size of price changes

In this section we perform a regression analysis to assess the quantitative importance of the main factors determining variations in the frequency and size of price changes over time. As shown in Figure 5 there is a substantial variation over time in both the size (median) and the frequency of price changes.

Figure 5 – Frequency and size of price changes



Source: Based on Istat data.
(1) Median size of price changes in each month.

The explanatory factors considered in the analysis are average inflation, seasonal factors, the incidence of threshold prices and indirect tax changes. Unfortunately, due to the lack of detailed information, we cannot take into account crucial elements such as product characteristics, market structure, degree of competition within the sector, and so on.

The aim of the exercise is to shed some light on firms' pricing strategies, in particular concerning the adoption of time-dependent or state-dependent rules. Under both strategies firms hold their price stable for some period of time. However, following a significant shock (to costs, to demand, etc.) firms following a state-dependent rule will immediately adjust their price, while those following a time-dependent rule will wait to be in a position to reset prices.

The variables examined can help explain variations in the frequency and the size of price changes over time as follows.

Average inflation - From a firm's point of view, the cost of not changing its price is likely to rise with the rate of inflation. Under time-dependent rules, the duration of prices should rise as inflation falls, so we would expect a positive correlation between the frequency of price changes and the inflation rate. Under state-dependent rules what matters is the magnitude of the shock, i.e. whether it is large enough to warrant a change in price; effects of nominal variables (money, interest rates, etc.) on the real side of the economy are less likely to vary with the rate of inflation.

Seasonality - A strong seasonal pattern in the data could be interpreted as evidence in favour of time-dependent pricing. Such a pattern could have several causes: weather, the timing of sales, institutional factors, such as changes in regulated prices typically coming in specific periods of the year. In the case of the Italian CPI, part of the seasonal pattern could be related to the fact that some prices are collected only quarterly (remaining unchanged for the other two months)[24]; however Cubadda and Sabbatini (1997) show that the actual impact of the data collection method is only marginal.

Threshold prices - There is substantial empirical evidence for Italy and for other countries in the euro area that a large percentage of consumer prices are set at "attractive" levels.[25] Attractive prices are of three types: (i) "psychological" (e.g. €1.99 instead of €2.00), aiming at gaining additional customers[26]; (ii) "fractional" (e.g. €1.70 instead of €1.67), to simplify payments; and (iii) "exact" (typically for large amounts, e.g. €50), to avoid the use of coins and small banknotes. With attractive prices, a retailer might decide not to change the price until the adjustment brings a new attractive level. This results in longer price duration.

In our database, the share of attractive prices is indeed quite large, almost 80% in the years before the euro cash changeover.[27] There are significant differences by sector: attractive prices are most important for services and non-energy industrial goods (around 90%), a bit less for unprocessed food (80%) and significantly less (50%) for processed food. We expect that the higher this share, the lower the frequency and the larger the size of price change.

VAT changes – Changes in value added tax rates are likely to induce a sudden increase in the frequency (and size) of price changes, as firms pass the tax on to the consumer. Over the sample period there was only one major change in VAT rates, in October 1997.[28]

Empirically, we evaluate the importance of the above factors by regressing the average frequency (F_t) and size (S_t) of price changes, respectively, on the following variables:

- D_i *(i=1, 2, …, 12)* = seasonal dummies.
- *VAT* = dummy variable equal to 1 in October 1997 and 0 elsewhere, capturing the change in the VAT rate occurred in that period.
- $CPTOT_t$ = headline consumer price inflation (year-on-year rates) at month *t*.
- $ATTR_t$ = percentage of attractive prices in month *t*.

The following two linear regressions are run, over the period 1996-2001, with reference to the average statistics computed on the whole dataset and, separately, for the five sub-indices:

$$F_t = \alpha + \sum_{i=2}^{12} \beta_i D_i + \gamma VAT9710 + \delta\, Cptot_t + \phi\, attr_t + \varepsilon_t$$

$$S_t = \alpha + \sum_{i=2}^{12} \beta_i D_i + \gamma VAT9710 + \phi\, attr_t + \varepsilon_t$$

---

24 In our sample, the items whose prices are collected on a quarterly basis are mostly services (maintenance and repair of the dwelling, maintenance and repair of personal transport equipment, hairdressing, etc.) and non-energy industrial goods (TV, furniture and furnishings, electric household appliances, etc.).

25 See Mostacci and Sabbatini (2003).

26 This is explained by the fact that customers may unconsciously undervalue the last digit, so that €0.99 is "assimilated" to €0.90; for a discussion see Mostacci and Sabbatini (2003).

27 These prices have been identified by the same methodology of Mostacci and Sabbatini (2001, 2003). The share is similar to that computed in Mostacci and Sabbatini (2003) on monthly elementary CPI price quotes in 2002; their coverage of the index is much larger, though data are referred only to 2002.

28 In the years considered by our analysis other episodes of VAT changes concerned only a fraction of goods and services and their impact was very moderate.

Tables 8 and 9 report the estimated coefficients[29]. The main findings can be summarised as follows.

Headline inflation increases the frequency of price changes, in line with our expectations (Table 8). The low value of the coefficient might depend on the fact that the variability of price dynamics over our time horizon is fairly small. As for the share of attractive prices, in our sample a higher share is correlated with lower frequency of price changes. This is evidence that firms indeed change attractive prices less frequently, waiting until a new attractive level can be set. The increase in VAT rates in October 1997 induced a higher frequency of price changes. Under the "extreme" assumption that this VAT change was fully and immediately passed on to prices, its impact on overall average inflation for our subset of items would have amounted to around 0.1 percentage points.[30] Seasonal dummies help explain the frequency of price changes.

Table 8 – Determinants of the time-series variation of the frequency of price changes (1)

	All items		**Unprocessed food**		**Processed food**		**Energy**		**Non-energy industrial goods**		**Services**	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	0.664	*0.00*	0.277	*0.00*	0.000	*0.99*	0.528	*0.00*	0.195	*0.10*	-1.213	*0.02*
VAT	0.026	*0.02*	0.097	*0.00*	0.036	*0.00*	-0.091	*0.43*	0.019	*0.00*	0.000	*0.99*
CPTOT	0.012	*0.00*	0.006	*0.00*	0.098	*0.00*	0.018	*0.01*	0.005	*0.00*	0.000	*0.99*
ATTR	-0.842	*0.00*	-0.272	*0.00*	0.058	*0.63*	0.260	*0.58*	-0.213	*0.09*	1.322	*0.02*
Number of observations	71		71		71		71		71		71	
adj. R^2	0.75		0.34		0.49		0.09		0.91		0.57	
Wald joint significance test	113.20		138.28		45.28		24.00		944.05		45.52	
p-value	*0.00*		*0.00*		*0.00*		*0.00*		*0.00*		*0.00*	
Wald seasonality test (2)	32.59		4.76		3.35		1.60		119.35		37.05	
p-value	*0.00*		*0.00*		*0.00*		*0.12*		*0.00*		*0.00*	

Source: Based on Istat data.
(1) All the regression include monthly seasonal dummies; standard errors are calculated using a HAC robust covariance estimator. - (2) Test of the joint significance of the monthly seasonal dummies.

As for the various sub-indices, average sectoral inflation always has a positive impact on the frequency of price changes, except for services, where the coefficient is not significant (Table 8). The share of attractive prices affects frequency significantly and negatively for unprocessed food and non-energy industrial goods. The frequency of price changes of processed food, non-energy industrial goods and services is highly seasonal.

29 Results for upward and downward changes are not reported in the Tables but are available from the authors on request.
30 At least part of the adjustment is likely to have taken place gradually over the months following the VAT change, in a pattern that might be impossible to distinguish from the other factors in price developments; nevertheless, in October 1997 an impact should be recorded.

The results presented in Table 9 indicate that the size of price changes is not significantly affected by any of the above factors.

Table 9 – Determinants of the time-series variation of the size of price changes (1)

	All items		**Unprocessed food**		**Processed food**		**Energy**		**Non-energy industrial goods**		**Services**	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	-0.060	*0.74*	0.286	*0.00*	0.144	*0.01*	-0.002	*0.09*	0.130	*0.79*	-0.412	*0.66*
VAT	0.007	*0.10*	-0.066	*0.00*	0.011	*0.10*	0.011	*0.02*	-0.011	*0.22*	0.034	*0.29*
ATTR	0.120	*0.65*	-0.377	*0.01*	-0.250	*0.01*	0.019	*0.62*	-0.088	*0.86*	0.473	*0.62*
Number of observations	71		71		71		71		71		71	
adj. R^2	0.13		0.21		0.07		0.00		0.04		0.16	
Wald joint significance test	5.74		79.65		5.24		1.58		8.32		6.13	
p-value	*0.00*		*0.00*		*0.00*		*0.12*		*0.00*		*0.00*	
Wald seasonality test (2)	4.94		1.70		1.07		0.61		8.27		6.09	
p-value	*0.00*		*0.10*		*0.40*		*0.81*		*0.00*		*0.00*	

Source: Based on Istat data.
(1) All the regression include monthly seasonal dummies; standard errors are calculated using a HAC robust covariance estimator.
(2) Test of the joint significance of the monthly seasonal dummies.

4.5 Factors affecting the probability of a price change

As a final piece of evidence, we exploited both the time series and the cross-sectional dimension of our dataset in order to investigate the factors influencing the probability of a price change. We carried out a logit regression in which the dependent variable is a binary indicator indicating whether a price change is recorded in each month. The explanatory variables considered in the regression are:

- DUR_{it}: the *number* of months since the last price change, capturing the degree of duration dependence of price changes;
- D_i $(i=1, 2, \ldots, 12)$ = seasonal dummies;
- type of item = unprocessed food, processed food, non-energy industrial good, energy, services;
- ATTR_i = dummy variable indicating whether the price is an attractive one;
- $CPTOT_t$ = headline consumer price inflation (year-on-year rates) at month *t*.
- *VAT* = dummy variable equal to 1 in October 1997 and 0 elsewhere, capturing the change in the VAT rate occurred in that period.

The results are reported in Table 10. First, they point to duration dependence: the longer the price spell is, the lower is the probability of a subsequent price change, quite surprisingly since the unchanged price gradually becomes more misaligned from the ones prevailing in the corresponding market.

Furthermore, the change in VAT rate significantly and positively affected on impact the probability of observing a price change, suggesting a state-dependent component in firms' price setting policies. However, under a state-dependent model we would expect changes in prices to have occurred also before the actual VAT changes (front-loading), as the fiscal manoeuvre was announced well in advance by the Government.[31]

Table 10 – Logit regression – Determinants of the conditional probability of a price change (1)

	Price changes				**Price increases**				**Price decreases**			
	coeff.	p-value	marginal effect	p-value	coeff.	p-value	marginal effect	p-value	coeff.	p-value	marginal effect	p-value
Unprocessed food	1.886	*0.00*	0.339	*0.00*	1.293	*0.00*	0.149	*0.00*	2.405	*0.00*	0.226	*0.00*
Processed food	0.460	*0.00*	0.562	*0.00*	0.256	*0.00*	0.020	*0.00*	1.019	*0.00*	0.046	*0.00*
Non-energy ind. goods	0.279	*0.00*	0.034	*0.00*	0.173	*0.00*	0.013	*0.00*	0.623	*0.00*	0.027	*0.00*
Energy	2.809	*0.00*	0.563	*0.00*	1.990	*0.00*	0.292	*0.00*	2.773	*0.00*	0.316	*0.00*
DUR	-0.019	*0.00*	-0.002	*0.00*	-0.007	*0.00*	-	*0.00*	-0.048	*0.00*	-0.002	*0.00*
VAT	0.502	*0.00*	0.069	*0.00*	0.686	*0.00*	0.069	*0.00*	-0.134	*0.11*	-0.005	*0.09*
CPTOT	0.418	*0.00*	0.048	*0.00*	0.392	*0.00*	0.030	*0.00*	0.200	*0.00*	0.008	*0.00*
ATTR	-0.374	*0.00*	-0.045	*0.00*	-0.245	*0.04*	-0.019	*0.00*	-0.382	*0.00*	-0.015	*0.00*
Intercept	-3.350	*0.00*			-3.848	*0.00*			-4.018	*0.00*		
Number of observations	291048				291048				291048			
Pseudo R^2	0.18				0.10				0.17			
Wald joint significance test	41285				19784				20814			
p-value	*0.00*				*0.00*				*0.00*			

Source: Based on Istat data.
(1) All the regressions include monthly seasonal dummies.

Average inflation is found to positively and significantly affect the probability of a price change.[32] On the contrary, a high share of attractive prices lowers such a probability. Although the related coefficients are not shown in the Table, seasonal dummies are always significant. As mentioned in Section 4.4, this is only partly due to statistical effects related to the data collection procedure and could signal the presence of some form of time-dependency.

[31] This finding is also consistent with an interpretation based on theories of "consumer anger" (Rotemberg, 2002) in which producers delay the price adjustments to avoid sudden reactions in demand determined by the disappointment of consumers.

[32] Note that the coefficient of this variable in the regression for price decreases is wrongly signed.

Finally, the results confirm the sectoral differences, the probability of a price change being much higher for energy and unprocessed food products than for non-energy industrial goods and services. There are no major differences with respect to the direction of the price change.

5. Conclusions

We investigated the characteristics of consumer price changes in Italy by analysing price behaviour at the most disaggregated level, namely items of a specific brand sold in a specific outlet. The empirical analysis was conducted in co-ordination within the *Inflation Persistence Network* of the Eurosystem.

We adopted both a duration approach, directly tracking the number of months for which a price remains unchanged, and a frequency approach, computing the proportion of price quotes that change in a given period. We used a variety of empirical strategies to assess the robustness of the results with respect to censoring and attrition.

The average duration of price spells, which captures how long prices tend to remain unchanged, is between 6 and 10 months, depending on how censoring is treated. It is much longer for non-energy industrial goods and services (14 and 15 months, respectively) and very short for energy products (2 months). The average duration slightly falls including in the sample the years 2002-03, suggesting that the euro cash changeover did induce a larger number of price adjustments.

Turning to the frequency approach, every month 9% of prices are changed, on average, which implies an average duration of 9-11 months. Again, results are rather heterogeneous across products. At the two extremes, energy prices change every month and services prices last almost two years.

Price changes tend to be asymmetrical, as is only to be expected given inflation. Under our preferred strategy on censoring, the average percentage of price increases each month is 7%, that of decreases 4%. The asymmetry is most pronounced for services and non-energy industrial goods, for which increases are respectively three and two times as frequent as decreases. However, the size of the average increase and decrease is approximately the same, a symmetry that broadly holds for all components.

As for type of outlet, traditional outlets tend to change the price of non-energy industrial goods and food products significantly less frequently than large stores (7.1% *versus* 10.8%). Price changes tend to be more symmetrical in mass than in traditional retailing.

In general, synchronisation of price changes is rather low (except for energy prices). However, when the statistics measuring synchronisation are computed separately for each item and each town and then aggregated in a national average, there is a much greater synchronisation for a number of items (in particular in the service sector).

In a regression analysis we investigated, exploiting the time series dimension of our dataset, the main factors determining variations over time in the average frequency and size of price changes. We considered for this purpose headline inflation, seasonal factors, attractive pricing thresholds and indirect tax changes. Frequency is positively related to the inflation rate and to the increase in VAT rates in October 1997 and negatively related to the share of attractive prices; seasonal effects are present. As for the size of price changes, none of the considered factors seems to significantly affect it.

Finally, a logit regression aimed at assessing which elements affect the probability of observing a price change in our dataset, both over time and across items, shows that such a probability is positively correlated with average inflation and with the increase in VAT rate and negatively correlated with the share of attractive prices and with the duration of the price spell. Seasonal dummies are always significant. There is evidence of differences with respect to sectors but not to the direction of the price change.

All in all, the analysis offers relevant evidence on pricing behaviour in Italy, confirming and significantly extending the results found in previous studies based on different data and approaches (Fabiani *et al.*, 2003 and 2004).

Appendix 1: the database

Table A1.1 – Items included in our dataset

Item	Frequency of collection	Item	Frequency of collection
Steak	Monthly	Dog food	Quarterly
1 fresh fish	Twice a month	Football	Quarterly
Tomato	Twice a month	Construction game (Lego)	Quarterly
Banana	Twice a month	Toothpaste	Monthly
Fresh milk	Monthly	Suitcase	Monthly
Sugar	Monthly	Dry cleaning (suit)	Monthly
Frozen spinach	Monthly	Hourly rate of an electrician	Quarterly
Mineral water	Monthly	Hourly rate of a plumber	Quarterly
Coffee	Monthly	Domestic services	Quarterly
Whisky	Monthly	Hourly rate in a garage	Quarterly
Beer in shop	Monthly	Car wash	Quarterly
Gasoline (heating)	Monthly	Balancing of wheels	Quarterly
2 fuels	Twice a month	Taxi	Monthly
Socks	Monthly	Movie	Monthly
Jeans	Monthly	Videotape hiring	Quarterly
Sport shoes	Monthly	Photo development	Quarterly
Shirt (men)	Monthly	Hotel room	Monthly
Tiles	Monthly	Glass of beer in a café	Monthly
Iron	Quarterly	1 meal in a restaurant	Monthly
Electric bulb	Monthly	Hot dog	Monthly
Furniture	Quarterly	Cola based lemonade in a café	Monthly
Towel	Monthly	Haircut (men)	Quarterly
Car tyre	Quarterly	Hairdressing (ladies)	Quarterly
Television set	Quarterly		

Table A1.2: Information available for each elementary price quote (metadata)

Date	Reference year and month
Outlet code	Each outlet has a numeric code
City	Name of the city
Area of the outlet	Code describing whether the location of the outlet is in a residential area, city center, etc.
Type of Outlet	Code describing the type outlet (supermarket, hard discount, small retailer…)
Brand	Description of the brand
Product code	Coicop or Istat code
Istat trajectory code	Numeric code which identifies univocally each price trajectory: i.e. it changes with any of the following characteristics: brand, variety, outlet, quantity.
Variety	Description of the product
Collected price	Price of the item actually collected by the surveyor
Collected price per unit	Price for a reference quantity. E.g. the price of 1 egg, computed on the basis of the price collected for a pack of 6 eggs.
Sale price	Starting from 2002 both the sale and the full price are collected. Before 2002 only one of the two was collected and no information was given on the nature of the price.
Base period price	The price in the base period (e.g. in October 2003 the base price refers to December 2002)
Switch of brand	Flag variable =1 if the price collected is referred to a different brand, 0 otherwise.
Switch of variety	Flag variable =1 if the price collected is referred to a different variety, 0 otherwise.
Switch of quantity	Flag variable =1 if the price collected is referred to a different quantity, 0 otherwise.
Switch of outlet	Flag variable =1 if the price collected is referred to a different outlet, 0 otherwise.

Appendix 2: definitions and formulas

Variables. We define the following binary variables for each product *j*:

$$DEN_{jt} = \begin{cases} 1 & \text{if } \mathrm{P}_{jt} \text{ and } \mathrm{P}_{j,t-1} \text{ are observed} \\ 0 & \text{if } \mathrm{P}_{jt} \text{ exists but not } \mathrm{P}_{j,t-1} \end{cases} \tag{1}$$

$$NUM_{jt} = \begin{cases} 1 & \text{if } \mathrm{P}_{jt} \neq \mathrm{P}_{j,t-1} \\ 0 & \text{otherwise} \end{cases} \tag{2}$$

$$NUMUP_{jt} = \begin{cases} 1 & \text{if } \mathrm{P}_{jt} > \mathrm{P}_{j,t-1} \\ 0 & \text{otherwise} \end{cases} \tag{3}$$

$$NUMDW_{jt} = \begin{cases} 1 & \text{if } \mathrm{P}_{jt} < \mathrm{P}_{j,t-1} \\ 0 & \text{otherwise} \end{cases} \tag{4}$$

On the basis of the above variables we analyse the frequency of price changes.

Basic statistics for each product *j*:

frequency of price changes:

$$F_j = \frac{\sum_{t=2}^{\tau} NUM_{jt}}{\sum_{t=2}^{\tau} DEN_{jt}} \tag{5}$$

implied median price duration (continuous time):

$$T_j^{50} = \frac{\ln(0.5)}{\ln(1-F_j)} \tag{6}$$

implied mean price duration (continuos time):

$$\overline{T}_j = -\frac{1}{\ln(1-F_j)} \tag{7}$$

frequency of price increases:

$$F_j^{+} = \frac{\sum_{t=2}^{\tau} NUMUP_{jt}}{\sum_{t=2}^{\tau} DEN_{jt}} \tag{8}$$

average price increase in p.c.

$$\overline{\Delta P}_j^{+} = \frac{\sum_{t=2}^{\tau} NUMUP_{jt}\left(\ln P_{jt} - \ln P_{j,t-1}\right)}{\sum_{t=2}^{\tau} DEN_{jt}} \tag{9}$$

frequency of price decreases:

$$F_j^- = \frac{\sum_{t=2}^{\tau} NUMDW_{jt}}{\sum_{t=2}^{\tau} DEN_{jt}} \tag{10}$$

average price decrease in p.c.:

$$\overline{\Delta P_j^-} = \frac{\sum_{t=2}^{\tau} NUMDW_{jt}\left(\ln P_{j,t-1} - \ln P_{jt}\right)}{\sum_{t=2}^{\tau} DEN_{jt}} \tag{11}$$

Price synchronisation. We rely on the synchronisation ratio proposed by Konieczny and Skrzypacz (2005), which is based on the standard deviation of the proportion of price changes (increases and/or decreases) evaluated each month *t*. In case of perfect synchronisation, the proportion of price changes at time t is either equal to 1 or to 0. As the average frequency of price changes over the sample period is equal to *Fj*, it means, in the case of perfect synchronisation, that all firms change their price simultaneously in *Fj* per cent of the cases and do not change their price in *(1- Fj)* per cent of the cases. Using the probability of price changes, it is then possible to compute the theoretical value of the standard deviation of the proportion of price changes over time in case of perfect synchronisation. This standard deviation associated to perfect synchronisation is:

$$SDMAX_j = \sqrt{F_j\left(1-F_j\right)^2 + \left(1-F_j\right)\left(0-F_j\right)^2} = \sqrt{F_j\left(1-F_j\right)} \tag{12}$$

This theoretical value is an upper limit for the standard deviation of the proportion of price changes. Similarly, in the case of perfect staggering, a constant proportion F_j of firms changes its price each month and the standard deviation of the proportion of price changes over time is equal to 0. The true standard deviation of price changes for product classification *j* is:

$$SD_j = \sqrt{\frac{1}{\tau-1}\sum_{t=2}^{\tau}\left(F_{jt} - F_j\right)^2} \tag{13}$$

The synchronisation ratio of product classification j is then given by:

$$SYNC_j = \frac{SD_j}{SDMAX_j} \tag{14}$$

The closer the ratio is to one, the higher is the degree of synchronisation. Similar expressions can be derived for $SYNC^+_j$ and $SYNC^-_j$, the synchronisation ratio of price increases and price decreases.

Appendix 3: results for individual items

Table A3.1 - Frequency of price changes and duration, with intermediate censoring and pseudo price-change (strategy 4) - Period 1996-2001
(weighted statistics)

Product	Frequency of price changes	Implied median duration (months)	Implied average duration
Unprocessed food	**19.5**	**3.2**	**4.6**
Steak	7.7	8.7	12.5
1 fresh fish	51.2	1.0	1.4
Tomatoes	76.4	0.5	0.7
Banana	34.0	1.7	2.4
Energy	**60.8**	**0.7**	**1.1**
Gasoline (heating)	59.0	0.8	1.1
Fuel type 1	63.7	0.7	1.0
Fuel type 2	61.2	0.7	1.1
Processed food	**9.4**	**7.0**	**10.1**
Milk	7.0	9.6	13.8
Sugar	7.9	8.4	12.1
Frozen spinach	10.0	6.6	9.5
Mineral water	10.2	6.4	9.3
Coffee	13.9	4.6	6.7
Whisky	9.1	7.3	10.5
Beer in a shop	12.3	5.3	7.6
Non-energy ind. goods	**5.8**	**11.6**	**16.7**
Socks	4.5	15.0	21.6
Jeans	5.7	11.9	17.1
Sport shoes	5.9	11.5	16.5
Shirt (men)	4.9	13.7	19.8
Tiles	4.1	16.7	24.1
Iron	5.0	13.5	19.4
Electric bulb	5.6	12.0	17.3
1 type of furniture	5.0	13.6	19.7
Towel	4.4	15.3	22.1
Car tyre	8.6	7.7	11.2
Television set	9.7	6.8	9.8
Dog food	6.8	9.8	14.2
Football	4.1	16.7	24.2
Construction game (Lego)	5.5	12.2	17.6
Toothpaste	10.1	6.5	9.4
Suitcase	5.5	12.3	17.8
Services	**4.6**	**14.7**	**21.1**
Dry cleaning	2.4	28.3	40.8
Hourly rate of an electrician	3.5	19.4	28.0
Hourly rate of a plumber	3.5	19.2	27.7
Domestic services	3.6	18.7	27.0
Hourly rate in a garage	3.2	21.4	30.9
Car wash	2.8	24.8	35.8
Balancing of wheels	2.4	28.5	41.2
Taxi	2.2	30.8	44.5
Movie	7.4	9.0	13.0
Videotape hiring	1.5	45.8	66.1
Photo development	3.2	21.4	30.9
Hotel room	8.3	8.0	11.5
Glass of beer in a café	2.5	27.2	39.3
1 meal in a restaurant	4.2	16.2	23.4
Hot-dog	2.7	25.3	36.5
Cola based lemonade in a café	2.4	28.9	41.7
Haircut (men)	3.1	22.0	31.7
Hairdressing (ladies)	2.7	25.0	36.1
Total	**10.0**	**6.6**	**9.5**

Source: Based on Istat data.

Table A3.2 - Frequency and size of price increases (decreases), with intermediate censoring and pseudo price-change (strategy 4) - Period 1996-2001

(weighted statistics)

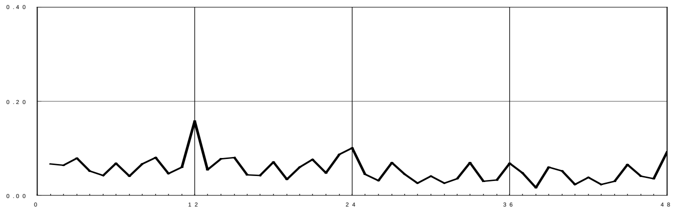
Product	Frequency of price increases	Average price increase (%)	Frequency of price decreases	Average price decrease (%)
Unprocessed food	**10.0**	**7.3**	**9.0**	**-7.8**
Steak	4.3	5.7	2.9	-6.3
1 fresh fish	24.4	10.6	26.8	-10.9
Tomatoes	38.4	15.5	37.9	-15.3
Banana	17.8	10.1	16.1	-10.5
Energy	**32.6**	**1.8**	**27.1**	**-1.6**
Gasoline (heating)	32.4	2.2	25.9	-2.1
Fuel type 1	33.1	1.9	29.5	-1.7
Fuel type 2	32.7	1.6	27.3	-1.5
Processed food	**4.9**	**6.1**	**2.7**	**-6.0**
Milk	5.4	4.7	0.9	-3.7
Sugar	2.9	4.7	4.1	-5.3
Frozen spinach	5.8	6.0	1.9	-6.6
Mineral water	4.2	7.8	3.2	-7.8
Coffee	5.8	5.2	6.3	-5.5
Whisky	5.3	6.3	1.5	-6.7
Beer in a shop	6.8	6.8	3.7	-6.8
Non-energy ind. goods	**2.9**	**6.8**	**0.9**	**-7.3**
Socks	2.9	6.9	0.3	-7.3
Jeans	3.7	6.8	0.4	-7.2
Sport shoes	2.9	7.4	1.0	-7.8
Shirt (men)	2.9	6.6	0.4	-7.7
Tiles	2.7	8.0	0.2	-3.0
Toaster	1.8	7.1	1.2	-6.3
Electric bulb	3.2	7.7	1.0	-9.1
1 type of furniture	2.5	5.8	0.3	-5.6
Towel	2.6	8.6	0.3	-8.6
Car tyre	3.3	6.2	3.0	-10.1
Television set	1.9	4.6	3.4	-8.5
Dog food	2.8	7.6	2.0	-9.8
Football	1.5	9.5	0.3	-12.3
Construction game (Lego)	1.8	6.8	0.5	-6.2
Toothpaste	5.0	6.6	2.7	-6.9
Suitcase	2.4	5.4	0.9	-8.1
Services	**3.2**	**9.1**	**0.7**	**-12.8**
Dry cleaning	1.6	7.9	0.6	-12.9
Hourly rate of an electrician	2.9	6.7	0.2	-8.3
Hourly rate of a plumber	2.9	7.1	0.1	-11.7
Domestic services	3.3	6.8	0.1	-4.9
Hourly rate in a garage	2.4	8.8	0.3	-8.2
Car wash	1.6	11.9	0.3	-15.5
Balancing of wheels	1.4	12.1	0.5	-9.6
Taxi	2.0	7.6	0.0	-
Movie	4.1	8.5	2.8	-7.1
Videotape hiring	1.0	18.4	0.4	-22.4
Photo development	1.8	6.2	1.0	-10.1
Hotel room	5.5	10.6	1.6	-15.3
Glass of beer in a café	1.7	12.2	0.1	-12.7
1 meal in a restaurant	2.8	7.7	0.5	-10.5
Hot-dog	2.0	13.0	0.1	-12.5
Cola based lemonade in a café	1.9	12.9	0.1	-10.6
Haircut (men)	1.8	9.4	0.2	-9.5
Hairdressing (ladies)	2.1	10.2	0.3	-11.5
Total	**5.5**	**7.2**	**3.1**	**-7.9**

Source: Based on Istat data.

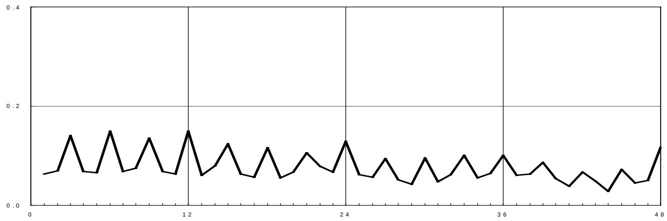
Appendix 4: hazard functions for sub-indices

Hazard functions presented below are estimated using the Kaplan-Meier method.

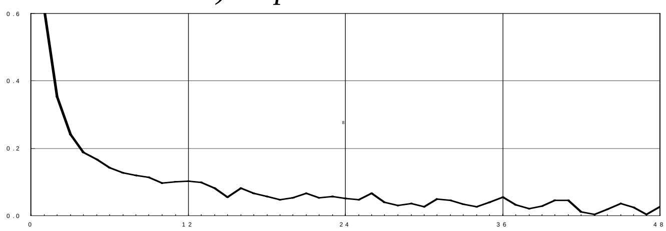
a) Services



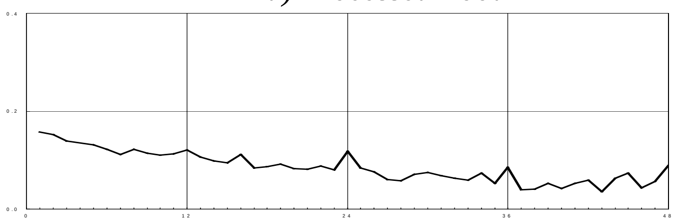
b) Non Energy Industrial Goods



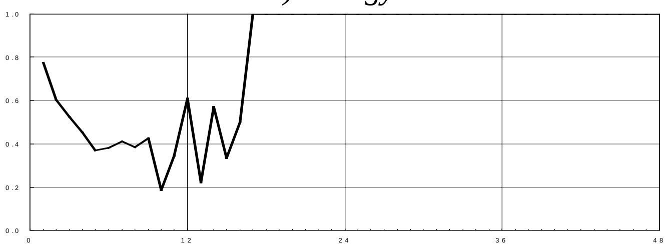
c) Unprocessed Food



d) Processed Food



e) Energy



Source: Based on Istat data.

References

Alvarez, L.J. and I. Hernando (2004), "Price setting behaviour in Spain: stylised facts using consumer price micro data", *forthcoming ECB Working Paper*.

Aoki, K. (2001), "Optimal Monetary Policy Responses to Relative Price Changes", *Journal of Monetary Economics*, vol. 48, pp. 55-80.

Aucremanne, L. and E. Dhyne (2004), "How frequently do price change? Empirical evidence based on the micro data underlying the Belgian CPI", *ECB Working Paper*, No. 331.

Baharad, E. and B. Eden (2004), "Price rigidity and price dispersion: evidence from micro data", *Review of Economic Dynamics*, vol. 7(2004), pp. 613-41.

Ball, L. and N.G. Mankiw (1994), "Asymmetric price adjustment and economic fluctuations", *The Economic Journal*, vol. 104(422), pp. 247-61.

Baudry, L., H. Le Bihan, P. Sevestre and S. Tarrieu (2004), "Price rigidity in France – evidence from consumer price micro-data", *ECB Working Paper*, No. 384.

Benigno, P. (2004), "Optimal Monetary Policy in a Currency Area", *Journal of International Economics*, vol. 63(2), pp. 293-320.

Blinder, A. (1994), "On sticky prices: academic theories meet the real world", in N.G. Mankiw (ed.), *Monetary Policy*, The University of Chicago Press, Chicago and London.

Blinder, A., E. Canetti, D. Lebow and J. Rudd (1998), *Asking about prices: A new approach to understand price stickiness*, Russel Sage Foundation, New York.

Bils, M. and P. Klenow (2004), "Some evidence on the importance of sticky prices", *Journal of Political Economy*, vol. 112, pp. 947-985.

Calvo, G.A. (1983), "Staggered Contracts in a Utility-Maximizing Framework", *Journal of Monetary Economics*, vol. 12(3), pp. 383-398.

Cecchetti, S.G. (1986), "The frequency of price adjustment: a study of newsstand prices of magazines", *Journal of Econometrics*, vol. 31, pp. 255-74

Cubadda, G. and R. Sabbatini (1997), "The seasonality of the Italian cost-of-living index", Bank of Italy, *Temi di Discussione*, No. 313.

Dias, M., D. Dias and P.D. Neves (2004), " Stylised features of price setting behaviour in Portugal: 1992-2001", *ECB Working Paper*, No. 332.

Dhyne, E., L. Alvarez, H. Le Bihan and G. Veronese (2004), " Price setting in the euro area: Some stylised facts from Individual Consumer Price Data", ECB mimeo.

Fabiani S., A. Gattulli and R. Sabbatini (2003), "Le rigidità dei prezzi in Italia", *Moneta e Credito*, No.233, pp. 325-358.

Fabiani, S., A. Gattulli and R. Sabbatini (2004), "The pricing behaviour of Italian firms: new survey evidence on price stickiness", *ECB Working Paper*, No. 333.

Fisher, T.C.G. and J.D. Konieczny (2000), "Synchronization of price changes by multiproduct firms: evidence from Canadian newspaper prices", *Economics Letters*, vol. 68, pp. 271-277.

Gali, J. and M. Gertler (1999) "Inflation Dynamics: a Structural Econometric Analysis," *Journal of Monetary Economics*, vol. 44, pp. 195-222.

Gali, J., M. Gertler and D. Lopez-Salido (2001), “European Inflation Dynamics”, *European Economic Review,* vol. 45, pp. 1237-1270.

Genesove, D. (2003), “The Nominal Rigidity of Apartment Rents”, *The Review of Economics and Statistics,* vol. 85(4), pp. 844-853.

Goodfriend, M. and R. King (1997), “The New Neoclassical Synthesis and the Role of Monetary Policy”, *NBER Macroeconomics Annual*, vol. 12, pp. 231-283.

Hall, S., A. Yates and M. Walsh (2000), “Are UK companies’ prices sticky?”, *Oxford Economic Papers*, vol. 52(3), pp. 425-446.

Heckman, J.J. and B. Singer (1985), “Social Science Duration Analysis”, in J.J. Heckman (ed.), *Longitudinal Analysis of Labor Market Data.* Cambridge: Cambridge University Press, pp. 39-110.

Jonker, N., C. Folkertsma and H. Blijenberg (2004), “An empirical analysis of price-setting behaviour in the Netherlands in the period 1998-2003 using micro data”, *forthcoming ECB Working Paper*.

Kackmeister, A. (2001), “Has retail price behavior changed since 1889? Evidence from microdata”, mimeo.

Kashyap, A. (1995), “Sticky Prices: New Evidence from Retail Catalogs”, *Quarterly Journal of Economics*, vol. 110, pp. 245-74.

Kiefer, N. (1988), “Economic duration data and hazard functions”, *Journal of Economic Literature,* vol. 26(2), pp. 646-722.

Klenow, P. and O. Kryvtsov (2005), “State-Dependent or Time-Dependent Pricing: Does It Matter for recent U.S. Inflation?”, mimeo.

Konieczny, J.D. and A. Skrzypacz (2005), “Inflation and price setting in a natural experiment”, *Journal of Monetary Economics*, forthcoming.

Lach, S.E. and P. Tsiddon (1992), “The behaviour of prices and inflation: an empirical analysis of disaggregated price data”, *Journal of Political Economy*, vol. 100(2), pp. 349-389.

Lach, S.E. and P. Tsiddon (1996), “Staggering and Syncronization in Price Setting: Evidence from Multiproduct Firms”, *American Economic Review,* vol. 86(5), pp. 1175-1196.

Lancaster, T. (1990), *The econometric analysis of transition data*, Cambridge University Press, Cambridge (UK).

Mostacci, F. e R. Sabbatini (2001), “Una stima ex ante dell'impatto del cash changeover sui prezzi al consumo in Italia”, *Contributi Istat*, No. 11.

Mostacci, F. and R. Sabbatini (2003), “L’euro ha creato inflazione? Changeover e arrotondamenti dei prezzi al consumo in Italia nel 2002”, *Moneta e Credito*, vol. 221, pp. 45-95.

Ratfai, A. (2005), “The Frequency and Size of Price Adjustment: Microeconomic Evidence”, *Managerial and Decision Economics,* forthcoming.

Rotemberg, J.J. (2002), “Costumer Anger at Price Increases, Time Variation in the Frequency of Price Changes and Monetary Policy”, *NBER Working Papers,* No. 9320.

Taylor, J.B. (ed.) (1999), *Monetary Policy Rules,* University of Chicago Press.

Wolman, A.L. (1999), “Sticky Prices, Marginal Cost, and the Behavior of Inflation”, *Federal Reserve Bank of Richmond Economic Quarterly*, vol. 85, pp. 29-48.

Wolman, A.L. (2000), "The Frequency and Costs of Individual Price Adjustment", *Federal Reserve Bank of Richmond Economic Quarterly*, vol. 86, pp. 1-22.

Woodford, M. (2003), *Interest and Prices: Foundations of a Theory of Monetary Policy*, Princeton University Press.

RECENTLY PUBLISHED "TEMI" (*).

N. 531 – *Cyclical asymmetry in fiscal policy, debt accumulation and the Treaty of Maastricht,* by F. BALASSONE and M. FRANCESE (December 2004).

N. 532 – *L'introduzione dell'euro e la divergenza tra inflazione rilevata e percepita,* by P. DEL GIOVANE and R. SABBATINI (December 2004).

N. 533 – *A micro simulation model of demographic development and households' economic behavior in Italy*, by A. ANDO and S. NICOLETTI ALTIMARI (December 2004).

N. 534 – *Aggregation bias in macro models: does it matter for the euro area?,* by L. MONTEFORTE (December 2004).

N. 535 – *Entry decisions and adverse selection: an empirical analysis of local credit markets*, by G. GOBBI and F. LOTTI (December 2004).

N. 536 – *An empirical investigation of the relationship between inequality and growth,* by P. PAGANO (December 2004).

N. 537 – *Monetary policy impulses, local output and the transmission mechanism*, by M. CARUSO (December 2004).

N. 538 – *An empirical micro matching model with an application to Italy and Spain*, by F. PERACCHI and E.VIVIANO (December 2004).

N. 539 – *La crescita dell'economia italiana negli anni novanta tra ritardo tecnologico e rallentamento della produttività,* by A. BASSANETTI, M. IOMMI, C. JONA-LASINIO and F. ZOLLINO (December 2004).

N. 540 – *Cyclical sensitivity of fiscal policies based on real-time data*, by L. FORNI and S. MOMIGLIANO (December 2004).

N. 541 – *L'introduzione dell'euro e le politiche di prezzo: analisi di un campione di dati individuali*, by E. GAIOTTI and F. LIPPI (February 2005).

N. 542 – *How do banks set interest rates?*, by L. GAMBACORTA (February 2005).

N. 543 – *Maxmin portfolio choice*, by M. TABOGA (February 2005).

N. 544 – *Forecasting output growth and inflation in the euro area: are financial spreads useful?*, by A. NOBILI (February 2005).

N. 545 – *Can option smiles forecast changes in interest rates? An application to the US, the UK and the Euro Area*, by M. PERICOLI (February 2005).

N. 546 – *The role of risk aversion in predicting individual behavior*, by L. Guiso and M. PAIELLA (February 2005).

N. 547 – *Prices, product differentiation and quality measurement: a comparison between hedonic and matched model methods,* by G. M. TOMAT (February 2005).

N. 548 – *The Basel Committee approach to risk weights and external ratings: what do we learn from bond spreads?*, by A. RESTI and A. SIRONI (February 2005).

N. 549 – *Firm size distribution: do financial constraints explain it all? Evidence from survey data,* by P. ANGELINI and A. GENERALE (June 2005).

N. 550 – *Proprietà, controllo e trasferimenti nelle imprese italiane. Cosa è cambiato nel decennio 1993-2003?,* by S. GIACOMELLI and S. TRENTO (June 2005).

N. 551 – *Quota dei Profitti e redditività del capitale in Italia: un tentativo di interpretazione*, by R. TORRINI (June 2005).

N. 552 – *Hiring incentives and labour force participation in Italy,* by P. CIPOLLONE, C. DI MARIA and A. GUELFI (June 2005).

N. 553 – *Trade credit as collateral* by M. OMICCIOLI (June 2005).

N. 554 – *Where do human capital externalities end up?,* by A. DALMAZZO and G. DE BLASIO (June 2005).

N. 555 – *Do capital gains affect consumption? Estimates of wealth effects from italian households' behavior,* by L. GUISO, M. PAIELLA and I. VISCO (June 2005).

(*) Requests for copies should be sent to:
Banca d'Italia – Servizio Studi – Divisione Biblioteca e pubblicazioni – Via Nazionale, 91 – 00184 Rome (fax 0039 06 47922059). They are available on the Internet www.bancaditalia.it.

"TEMI" LATER PUBLISHED ELSEWHERE

1999

L. GUISO and G. PARIGI, *Investment and demand uncertainty*, Quarterly Journal of Economics, Vol. 114 (1), pp. 185-228, **TD No. 289 (November 1996)**.

A. F. POZZOLO, *Gli effetti della liberalizzazione valutaria sulle transazioni finanziarie dell'Italia con l'estero*, Rivista di Politica Economica, Vol. 89 (3), pp. 45-76, **TD No. 296 (February 1997)**.

A. CUKIERMAN and F. LIPPI, *Central bank independence, centralization of wage bargaining, inflation and unemployment: theory and evidence*, European Economic Review, Vol. 43 (7), pp. 1395-1434, **TD No. 332 (April 1998)**.

P. CASELLI and R. RINALDI, *La politica fiscale nei paesi dell'Unione europea negli anni novanta,* Studi e note di economia, (1), pp. 71-109, **TD No. 334 (July 1998)**.

A. BRANDOLINI, *The distribution of personal income in post-war Italy: Source description, data quality, and the time pattern of income inequality*, Giornale degli economisti e Annali di economia, Vol. 58 (2), pp. 183-239, **TD No. 350 (April 1999)**.

L. GUISO, A. K. KASHYAP, F. PANETTA and D. TERLIZZESE, *Will a common European monetary policy have asymmetric effects?*, Economic Perspectives, Federal Reserve Bank of Chicago, Vol. 23 (4), pp. 56-75, **TD No. 384 (October 2000)**.

2000

P. ANGELINI, *Are banks risk-averse? Timing of the operations in the interbank market*, Journal of Money, Credit and Banking, Vol. 32 (1), pp. 54-73, **TD No. 266 (April 1996).**

F. DRUDI and R: GIORDANO, *Default Risk and optimal debt management*, Journal of Banking and Finance, Vol. 24 (6), pp. 861-892, **TD No. 278 (September 1996)**.

F. DRUDI and R. GIORDANO, *Wage indexation, employment and inflation,* Scandinavian Journal of Economics, Vol. 102 (4), pp. 645-668, **TD No. 292 (December 1996)**.

F. DRUDI and A. PRATI, *Signaling fiscal regime sustainability*, European Economic Review, Vol. 44 (10), pp. 1897-1930, **TD No. 335 (September 1998)**.

F. FORNARI and R. VIOLI, *The probability density function of interest rates implied in the price of options,* in: R. Violi, (ed.) , Mercati dei derivati, controllo monetario e stabilità finanziaria, Il Mulino, Bologna, **TD No. 339 (October 1998)**.

D. J. MARCHETTI and G. PARIGI, *Energy consumption, survey data and the prediction of industrial production in Italy,* Journal of Forecasting, Vol. 19 (5), pp. 419-440, **TD No. 342 (December 1998)**.

A. BAFFIGI, M. PAGNINI and F. QUINTILIANI, *Localismo bancario e distretti industriali: assetto dei mercati del credito e finanziamento degli investimenti,* in: L.F. Signorini (ed.), Lo sviluppo locale: un'indagine della Banca d'Italia sui distretti industriali, Donzelli, **TD No. 347** (**March 1999)**.

A. SCALIA and V. VACCA, *Does market transparency matter? A case study,* in: Market Liquidity: Research Findings and Selected Policy Implications, Basel, Bank for International Settlements, **TD No. 359 (October 1999)**.

F. SCHIVARDI, *Rigidità nel mercato del lavoro, disoccupazione e crescita*, Giornale degli economisti e Annali di economia, Vol. 59 (1), pp. 117-143, **TD No. 364 (December 1999)**.

G. BODO, R. GOLINELLI and G. PARIGI, *Forecasting industrial production in the euro area,* Empirical Economics, Vol. 25 (4), pp. 541-561, **TD No. 370 (March 2000)**.

F. ALTISSIMO, D. J. MARCHETTI and G. P. ONETO, *The Italian business cycle: Coincident and leading indicators and some stylized facts,* Giornale degli economisti e Annali di economia, Vol. 60 (2), pp. 147-220, **TD No. 377 (October 2000)**.

C. MICHELACCI and P. ZAFFARONI, *(Fractional) Beta convergence*, Journal of Monetary Economics, Vol. 45, pp. 129-153, **TD No. 383 (October 2000)**.

R. DE BONIS and A. FERRANDO, *The Italian banking structure in the nineties: testing the multimarket contact hypothesis*, Economic Notes, Vol. 29 (2), pp. 215-241, **TD No. 387 (October 2000)**.

2001

M. CARUSO, *Stock prices and money velocity: A multi-country analysis,* Empirical Economics, Vol. 26 (4), pp. 651-72, **TD No. 264 (February 1996)**.

P. CIPOLLONE and D. J. MARCHETTI, *Bottlenecks and limits to growth: A multisectoral analysis of Italian industry*, Journal of Policy Modeling, Vol. 23 (6), pp. 601-620, **TD No. 314 (August 1997)**.

P. CASELLI, *Fiscal consolidations under fixed exchange rates,* European Economic Review, Vol. 45 (3), pp. 425-450, **TD No. 336 (October 1998)**.

F. ALTISSIMO and G. L. VIOLANTE, *Nonlinear VAR: Some theory and an application to US GNP and unemployment*, Journal of Applied Econometrics, Vol. 16 (4), pp. 461-486, **TD No. 338 (October 1998)**.

F. NUCCI and A. F. POZZOLO, *Investment and the exchange rate*, European Economic Review, Vol. 45 (2), pp. 259-283, **TD No. 344 (December 1998)**.

L. GAMBACORTA, *On the institutional design of the European monetary union: Conservatism, stability pact and economic shocks,* Economic Notes, Vol. 30 (1), pp. 109-143, **TD No. 356 (June 1999)**.

P. FINALDI RUSSO and P. ROSSI, *Credit costraints in italian industrial districts,* Applied Economics, Vol. 33 (11), pp. 1469-1477, **TD No. 360 (December 1999)**.

A. CUKIERMAN and F. LIPPI, *Labor markets and monetary union: A strategic analysis,* Economic Journal, Vol. 111 (473), pp. 541-565, **TD No. 365 (February 2000)**.

G. PARIGI and S. SIVIERO, *An investment-function-based measure of capacity utilisation, potential output and utilised capacity in the Bank of Italy's quarterly model*, Economic Modelling, Vol. 18 (4), pp. 525-550, **TD No. 367 (February 2000)**.

F. BALASSONE and D. MONACELLI, *Emu fiscal rules: Is there a gap?*, in: M. Bordignon and D. Da Empoli (eds.), Politica fiscale, flessibilità dei mercati e crescita, Milano, Franco Angeli, **TD No. 375 (July 2000)**.

A. B. ATKINSON and A. BRANDOLINI, *Promise and pitfalls in the use of "secondary" data-sets: Income inequality in OECD countries*, Journal of Economic Literature, Vol. 39 (3), pp. 771-799, **TD No. 379 (October 2000)**.

D. FOCARELLI and A. F. POZZOLO, *The determinants of cross-border bank shareholdings: An analysis with bank-level data from OECD countries*, Journal of Banking and Finance, Vol. 25 (12), pp. 2305-2337, **TD No. 381 (October 2000)**.

M. SBRACIA and A. ZAGHINI, *Expectations and information in second generation currency crises models,* Economic Modelling, Vol. 18 (2), pp. 203-222, **TD No. 391 (December 2000)**.

F. FORNARI and A. MELE, *Recovering the probability density function of asset prices using GARCH as diffusion approximations,* Journal of Empirical Finance, Vol. 8 (1), pp. 83-110**, TD No. 396 (February 2001)**.

P. CIPOLLONE, *La convergenza dei salari manifatturieri in Europa,* Politica economica, Vol. 17 (1), pp. 97-125, **TD No. 398 (February 2001)**.

E. BONACCORSI DI PATTI and G. GOBBI, *The changing structure of local credit markets: Are small businesses special?*, Journal of Banking and Finance, Vol. 25 (12), pp. 2209-2237, **TD No. 404 (June 2001)**.

G. MESSINA, *Decentramento fiscale e perequazione regionale. Efficienza e redistribuzione nel nuovo sistema di finanziamento delle regioni a statuto ordinario*, Studi economici, Vol. 56 (73), pp. 131-148, **TD No. 416 (August 2001)**.

2002

R. CESARI and F. PANETTA, *Style, fees and performance of Italian equity funds*, Journal of Banking and Finance, Vol. 26 (1), **TD No. 325 (January 1998)**.

L. GAMBACORTA, *Asymmetric bank lending channels and ECB monetary policy*, Economic Modelling, Vol. 20 (1), pp. 25-46, **TD No. 340 (October 1998)**.

C. GIANNINI, *"Enemy of none but a common friend of all"? An international perspective on the lender-of-last-resort function*, Essay in International Finance, Vol. 214, Princeton, N. J., Princeton University Press, **TD No. 341 (December 1998)**.

A. ZAGHINI, *Fiscal adjustments and economic performing: A comparative study*, Applied Economics, Vol. 33 (5), pp. 613-624, **TD No. 355 (June 1999)**.

F. ALTISSIMO, S. SIVIERO and D. TERLIZZESE, *How deep are the deep parameters?,* Annales d'Economie et de Statistique,.(67/68), pp. 207-226, **TD No. 354 (June 1999)**.

F. FORNARI, C. MONTICELLI, M. PERICOLI and M. TIVEGNA, *The impact of news on the exchange rate of the lira and long-term interest rates*, Economic Modelling, Vol. 19 (4), pp. 611-639, **TD No. 358 (October 1999**).

D. FOCARELLI, F. PANETTA and C. SALLEO, *Why do banks merge?,* Journal of Money, Credit and Banking, Vol. 34 (4), pp. 1047-1066, **TD No. 361 (December 1999)**.

D. J. MARCHETTI, *Markup and the business cycle: Evidence from Italian manufacturing branches,* Open Economies Review, Vol. 13 (1), pp. 87-103, **TD No. 362 (December 1999**).

F. BUSETTI, *Testing for stochastic trends in series with structural breaks*, Journal of Forecasting, Vol. 21 (2), pp. 81-105, **TD No. 385 (October 2000)**.

F. LIPPI, *Revisiting the Case for a Populist Central Banker,* European Economic Review, Vol. 46 (3), pp. 601-612, **TD No. 386 (October 2000)**.

F. PANETTA, *The stability of the relation between the stock market and macroeconomic forces*, Economic Notes, Vol. 31 (3), **TD No. 393 (February 2001)**.

G. GRANDE and L. VENTURA, *Labor income and risky assets under market incompleteness: Evidence from Italian data,* Journal of Banking and Finance, Vol. 26 (2-3), pp. 597-620, **TD No. 399 (March 2001)**.

A. BRANDOLINI, P. CIPOLLONE and P. SESTITO, *Earnings dispersion, low pay and household poverty in Italy, 1977-1998,* in D. Cohen, T. Piketty and G. Saint-Paul (eds.), The Economics of Rising Inequalities, pp. 225-264, Oxford, Oxford University Press, **TD No. 427 (November 2001)**.

L. CANNARI and G. D'ALESSIO, *La distribuzione del reddito e della ricchezza nelle regioni italiane*, Rivista Economica del Mezzogiorno (Trimestrale della SVIMEZ), Vol. XVI (4), pp. 809-847, Il Mulino, **TD No. 482 (June 2003)**.

2003

F. SCHIVARDI, *Reallocation and learning over the business cycle,* European Economic Review, , Vol. 47 (1), pp. 95-111, **TD No. 345 (December 1998)**.

P. CASELLI, P. PAGANO and F. SCHIVARDI, *Uncertainty and slowdown of capital accumulation in Europe,* Applied Economics, Vol. 35 (1), pp. 79-89, **TD No. 372 (March 2000).**

P. ANGELINI and N. CETORELLI, *The effect of regulatory reform on competition in the banking industry,* Federal Reserve Bank of Chicago, Journal of Money, Credit and Banking, Vol. 35, pp. 663-684, **TD No. 380 (October 2000)**.

P. PAGANO and G. FERRAGUTO, Endogenous growth with intertemporally dependent preferences, Contribution to Macroeconomics, Vol. 3 (1), pp. 1-38, **TD No. 382 (October 2000).**

P. PAGANO and F. SCHIVARDI, *Firm size distribution and growth,* Scandinavian Journal of Economics, Vol. 105 (2), pp. 255-274, **TD No. 394 (February 2001)**.

M. PERICOLI and M. SBRACIA, *A Primer on Financial Contagion*, Journal of Economic Surveys, Vol. 17 (4), pp. 571-608, **TD No. 407 (June 2001)**.

M. SBRACIA and A. ZAGHINI, *The role of the banking system in the international transmission of shocks,* World Economy, Vol. 26 (5), pp. 727-754, **TD No. 409 (June 2001)**.

E. GAIOTTI and A. GENERALE, *Does monetary policy have asymmetric effects? A look at the investment decisions of Italian firms,* Giornale degli Economisti e Annali di Economia, Vol. 61 (1), pp. 29-59, **TD No**. **429 (December 2001)**.

L. GAMBACORTA, *The Italian banking system and monetary policy transmission: evidence from bank level data*, in: I. Angeloni, A. Kashyap and B. Mojon (eds.), Monetary Policy Transmission in the Euro Area, Cambridge, Cambridge University Press, **TD No. 430 (December 2001).**

M. EHRMANN, L. GAMBACORTA, J. MARTÍNEZ PAGÉS, P. SEVESTRE and A. WORMS, *Financial systems and the role of banks in monetary policy transmission in the euro area*, in: I. Angeloni, A. Kashyap and B. Mojon (eds.), Monetary Policy Transmission in the Euro Area, Cambridge, Cambridge University Press, **TD No. 432 (December 2001)**.

F. SPADAFORA, *Financial crises, moral hazard and the speciality of the international market: further evidence from the pricing of syndicated bank loans to emerging markets,* Emerging Markets Review, Vol. 4 (2), pp. 167-198, **TD No. 438 (March 2002)**.

D. FOCARELLI and F. PANETTA, *Are mergers beneficial to consumers? Evidence from the market for bank deposits,* American Economic Review, Vol. 93 (4), pp. 1152-1172, **TD No. 448 (July 2002)**.

E.VIVIANO, *Un'analisi critica delle definizioni di disoccupazione e partecipazione in Italia,* Politica Economica, Vol. 19 (1), pp. 161-190, **TD No. 450 (July 2002)**.

M. PAGNINI, *Misura e Determinanti dell'Agglomerazione Spaziale nei Comparti Industriali in Italia,* Rivista di Politica Economica, Vol. 3 (4), pp. 149-196, **TD No. 452 (October 2002)**.

F. BUSETTI and A. M. ROBERT TAYLOR, *Testing against stochastic trend and seasonality in the presence of unattended breaks and unit roots,* Journal of Econometrics, Vol. 117 (1), pp. 21-53, **TD No. 470 (February 2003)**.

2004

F. LIPPI, *Strategic monetary policy with non-atomistic wage-setters*, Review of Economic Studies, Vol. 70 (4), pp. 909-919, **TD No. 374 (June 2000)**.

P. CHIADES and L. GAMBACORTA, *The Bernanke and Blinder model in an open economy: The Italian case,* German Economic Review, Vol. 5 (1), pp. 1-34, **TD No. 388 (December 2000)**.

M. BUGAMELLI and P. PAGANO, *Barriers to Investment in ICT*, Applied Economics, Vol. 36 (20), pp. 2275-2286, **TD No. 420 (October 2001)**.

A. BAFFIGI, R. GOLINELLI and G. PARIGI, *Bridge models to forecast the euro area GDP*, International Journal of Forecasting, Vol. 20 (3), pp. 447-460,**TD No. 456 (December 2002)**.

D. AMEL, C. BARNES, F. PANETTA and C. SALLEO, *Consolidation and Efficiency in the Financial Sector: A Review of the International Evidence,* Journal of Banking and Finance, Vol. 28 (10), pp. 2493-2519, **TD No. 464 (December 2002).**

M. PAIELLA, *Heterogeneity in financial market participation: appraising its implications for the C-CAPM,* Review of Finance, Vol. 8, pp. 1-36, **TD No. 473 (June 2003)**.

E. BARUCCI, C. IMPENNA and R. RENÒ, *Monetary integration, markets and regulation,* Research in Banking and Finance, (4), pp. 319-360, **TD NO. 475 (June 2003)**.

E. BONACCORSI DI PATTI and G. DELL'ARICCIA, *Bank competition and firm creation*, Journal of Money Credit and Banking, Vol. 36 (2), pp. 225-251, **TD No. 481 (June 2003)**.

R. GOLINELLI and G. PARIGI, *Consumer sentiment and economic activity: a cross country comparison*, Journal of Business Cycle Measurement and Analysis, Vol. 1 (2), pp. 147-172, **TD No. 484 (September 2003)**.

L. GAMBACORTA and P. E. MISTRULLI, *Does bank capital affect lending behavior?*, Journal of Financial Intermediation, Vol. 13 (4), pp. 436-457, **TD NO. 486 (September 2003)**.

F. SPADAFORA, *Il pilastro privato del sistema previdenziale: il caso del Regno Unito,* Rivista Economia Pubblica, (5), pp. 75-114, **TD No. 503** (**June 2004**).

G. GOBBI e F. LOTTI, *Entry decisions and adverse selection: an empirical analysis of local credit markets,* Journal of Financial services Research, Vol. 26 (3), pp. 225-244, **TD NO. 535 (December 2004)**.

F. CINGANO e F. SCHIVARDI, *Identifying the sources of local productivity growth,* Journal of the European Economic Association, Vol. 2 (4), pp. 720-742, **TD NO. 474 (June 2003)**.

C. BENTIVOGLI e F. QUINTILIANI, *Tecnologia e dinamica dei vantaggi comparati: un confronto fra quattro regioni italiane*, in C. Conigliani (a cura di), *Tra sviluppo e stagnazione: l'economia dell'Emilia-Romagna,* Bologna, Il Mulino, **TD NO. 522 (October 2004)**.

2005

A. DI CESARE, *Estimating Expectations of Shocks Using Option Prices,* The ICFAI Journal of Derivatives Markets, Vol. II (1), pp. 42-53, **TD No. 506 (July 2004).**

M. OMICCIOLI, *Il credito commerciale: problemi e teorie,* in L. Cannari, S. Chiri e M. Omiccioli (a cura di), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD NO. 494 (June 2004)**.

L. CANNARI, S. CHIRI e M. OMICCIOLI, *Condizioni del credito commerciale e differenzizione della clientela,* in L. Cannari, S. Chiri e M. Omiccioli (a cura di), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD NO. 495 (June 2004)**.

P. FINALDI RUSSO e L. LEVA, *Il debito commerciale in Italia: quanto contano le motivazioni finanziarie?,* in L. Cannari, S. Chiri e M. Omiccioli (a cura di), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD NO. 496 (June 2004)**.

A. CARMIGNANI, *Funzionamento della giustizia civile e struttura finanziaria delle imprese: il ruolo del credito commerciale,* in L. Cannari, S. Chiri e M. Omiccioli (a cura di), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD NO. 497 (June 2004)**.

G. DE BLASIO, *Does trade credit substitute for bank credit?,* in L. Cannari, S. Chiri e M. Omiccioli (a cura di), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD NO. 498 (June 2004)**.

M. BENVENUTI e M. GALLO, P*erché le imprese ricorrono al factoring? Il caso dell'Italia,* in L. Cannari, S. Chiri e M. Omiccioli (a cura di), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD NO. 518 (October 2004)**.