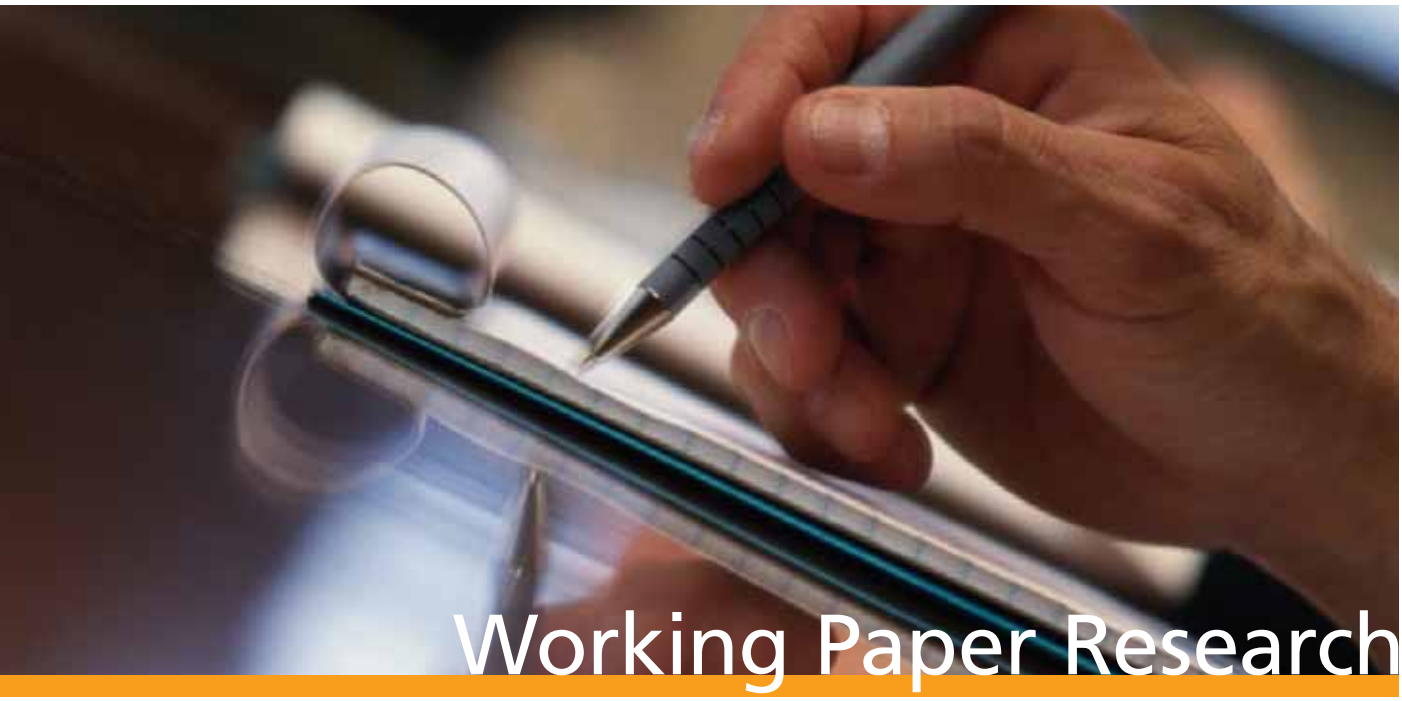
Searching for additional sources of inflation persistence : the micro-price panel data approach



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Abstract

It is often argued that the baseline New-Keynesian model, which relies solely on the notion of infrequent price adjustment, cannot account for the observed degree of inflation sluggishness. Therefore it is a common practice among macro modellers to introduce an ad hoc additional source of persistence to their models. Yet, the empirical validity of this practice has never been formally tested. This paper attempts to examine whether there is some additional persistence present in the data on micro-prices, beyond that implied by infrequent price adjustment. We consider two distinct sets of assumptions consistent with the existence of an intrinsic or extrinsic source of sluggishness and build and estimate two alternative models based on these assumptions. It is shown that in he case of certain product categories, particularly food, there is evidence of less sluggishness than what the standard assumptions underlying the New-Keynesian model would imply. We find certain support for the existence of an additional source of sluggishness for some industrial goods and services. Importantly however, the results are sensitive to the choice of the model. We conclude that some inconsistencies with the baseline New-Keynesian assumptions may be tracked in the price behavior. Yet, it is too early to assess their strength or the effect on macro aggregates. Therefore, at the current stage it would be premature to discard the baseline version of the New-Keynesian model based on evidence from micro-data. Similarly, the micro support for introducing an extra source of inflation sluggishness to macro-models is still relatively weak.

JEL-code : C51, C81, E13, E1, D21, L11.
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TABLE OF CONTENTS

1 Introduction.

The aim of this paper is to test whether there is some additional sluggishness present in the data on micro-prices, beyond that implied by the fact that prices adjust infrequently. This question lies at the core of an old and still unresolved economic puzzle: Why is there so much inertia and persistence in the dynamics of economic variables? The classic approach to explaining these phenomena points to the nominal frictions, in particular nominal price rigidity, as likely causes of the observed sluggishness. Yet, within the framework of the New-Keynesian model, which relies on the assumption of nominal frictions, it is not easy to reproduce the salient properties of the inflation and output dynamics. In the baseline version of this model, monopolistically competitive firms adjust their prices infrequently at random dates, as in Calvo (1983). The adjustment is complete: When it occurs, prices are always set to their desired level. Additionally, it is often postulated that the nominal marginal cost follows a random walk. Under these assumptions, the degree of price rigidity necessary to generate the observed output and inflation persistence is frequently found to be implausibly high. This observation gives rise to the two persistence puzzles discussed thoroughly by Taylor in his 1999 survey.

Although there is no consensus on this topic[1], many economists believe that in order to match dynamic properties of macroeconomic series, one has to seriously modify the structure of the New-Keynesian framework by allowing for additional sources of sluggishness, apart from the lumpiness of the process of price adjustment. Hence, many alternative models have been proposed, most of them making ad hoc assumptions about the way price-setters adjust their prices. For instance, Gali and Gertler (1999) supposed that a fraction of firms on the market is backward-looking, so that they set their prices based on the last period inflation rate. Christiano et al (2001) assumed instead that firms are allowed to review their prices only infrequently. In periods when price reviewing is not allowed, they set their prices according to some indexation rule. A yet different model was proposed by Mankiw and Reis (2002). These authors completely abandoned the standard price rigidity setting in favor of the assumption that it is the information about the marginal cost innovations that arrives to individuals at a Calvo rate. Their paper has given birth to an outburst of works on the so called sticky information models.

By construction, the departures from the standard setting listed above help in ratcheting up inflation and output persistence generated by macroeconomic models. Some of them seem to be helpful in enhancing properties of these models also in other dimensions[2]. None of these alternative assumptions however

[1] One of the controversies being the actual degree of inflation persistence, see the Cogley and Sargent (2001) vs. Stock (2001) debate and further papers on the issue, for instance Benati (2003), Clark (2003), Gadzinski and Orlandi (2004) and Levin and Piger (2004).

[2] For instance both, the sticky information model (Mankiw and Reis, 2002) and the model of Christiano, Eichenbaum and Evans (2001), were found to be helpful in generating realistic hump-shaped responses of output and inflation to monetary shocks. Unlike in the classic New-Keynesian model, announced and credible disinflations are contractionary in both models. They are also consistent with the so called accelerator hypothesis. See Trabant (2005) for an

can be argued to be directly implied by micro-level evidence[3]. Indeed, none of them looks plausible, when viewed as a literal description of firms' price setting behavior. We know that indexation schemes like in Christiano et al are rare or non-existent. Prices are changed infrequently, which contradicts the flexible price setting of Mankiw and Reis. It is also hard to believe that an important fraction of firms in the economy derives their pricing decisions based on information sets not having been updated for many periods, as it is demanded by the Calvo scheme assumption[4]. Also, the Gali and Gertler's assumption of a group of price-setters who constantly make an error when setting their prices does not seem overly realistic.

Nevertheless, some general micro-level evidence has already been collected suggesting that firms may indeed rely on limited information sets when setting their prices. Based on the results of an ad hoc survey on price-setting behavior conducted among Belgian firms, Aucremanne and Druant (2007) argue that an important fraction of firms uses a rule-of-thumb (indexation, fixed amount or percentage adaptation etc.) to choose a new price for their products. Similar findings were reported for several other European countries (see Fabiani et al, 2006, who provide a summary of conclusions from price-setting surveys conducted by nine Eurosystem central banks in years 2003-2004)[5]. These results suggest that, although there may be no point in searching for a micro-data confirmation of the validity of any of the specific modifications of the baseline New-Keynesian model proposed by the individual authors, it does make sense to investigate whether price data provide any support for a general 'another source of persistence' hypothesis. This is precisely the task we undertake in this paper. Our objective is to set up an econometric framework that would allow to test hypotheses about the potential sources of sluggishness present in the data. The data set used for the verification of these hypotheses is a large panel of consumer prices underlying the Belgian Consumer Price Index. The data span over most product categories used for constructing the Belgian CPI and over a relatively long period. The number of individual price setters within every product category is substantial.

Procedures developed in the paper permit to differentiate between the sluggishness implied by the existence of nominal price rigidity, as in the baseline New-Keynesian model, and that implied by a relaxation of one of this model's basic assumptions. First, we consider the consequences of relaxing the assump-

extended discussion and further references.

[3] It is in contrast to the nominal price rigidity, on which the standard approach is founded. The existence of this friction has been already thoroughly documented (see for instance Wolman, 2000 for a survey of research on the US micro-price data and Dhyne et al, 2005 and Alvarez et al, 2006 for two summaries of the IPN research on prices in the Euro Area).

[4] This criticism may be especially relevant, as Collard and Dellas (2003) have demonstrated that the sticky information model with information arrival a la Taylor loses its favorable properties vis a vis the standard New-Keynesian model.

[5] Fabiani et al emphasise however that the conclusions they draw mainly apply to producer prices. In contrast, the panel dataset used in this paper contains consumer prices. One has to bear in mind that a set of conclusions drawn based on the former does not necessarily transposes to the latter type of data.

tion of complete price adjustment. More precisely, we test the hypothesis that prices adjust to shocks only gradually. Then, we focus on alternative assumptions concerning the time series properties of these shocks, paying somewhat more attention to the distinction between their idiosyncratic and common innovations.

All works introducing modifications to the standard New-Keynesian model cited earlier relax the assumption of complete adjustment of prices, even though the way it is done differs from one model to another[6]. Therefore the first part of this paper is devoted to measuring the degree of price gradualism prevailing in the economy.

If there is some form of gradualism in the agents' behavior, the individual consecutive price changes will be serially correlated. It is not so in the simple New-Keynesian setting. In this setting the desired price level may be shown to be a constant mark-up over a discounted stream of expected future marginal costs. Hence, it contains all relevant information about the marginal cost shocks from today's and all past dates. Therefore, the size of a given price change is affected only by the information that has arrived since the previous price adjustment. Clearly then, under the assumption that the marginal cost innovations are white noise, the price changes are uncorrelated.

The above discussion indicates the route an investigator can take in order to test for the presence of gradualism in firms' pricing decisions. One should look closer at the autocorrelation of individual price changes. Crudely speaking, this is the approach we follow in this paper. Leaving aside the econometric difficulties related to the task of correctly measuring the autocorrelation of individual price changes, there is a pertinent question arising, what precise economic interpretation should be given to the obtained numbers. Not surprisingly, the answer to this question is model dependent. Delaying a more thorough discussion of this issue to the later sections, here we confine to saying that we impose a model formulation that allows for an easy interpretation of the obtained coefficients in terms of the speeds of adjustment. Our model is thus close in spirit to the celebrated partial adjustment model (see for instance Caballero and Engel, 2004). Its precise formulation is virtually identical to that used by Jorda (1999) to study the random-time aggregation bias.

The existence of autocorrelation between individual price changes does not have to be a result of incomplete price adjustment. We would observe a similar effect if the nominal marginal cost followed a process implying more persistence than a simple random walk. Hence, the sluggishness having an extrinsic origin[7] is a competing explanation for the positive autocorrelation potentially revealed in the data. Unfortunately, in order to decisively distinguish between the two

[6] For instance in Mankiw and Reis, managers set prices to a level that is optimal, given their current information set. They have to introduce price corrections however, as their information set is being updated. On the other hand, the behavior of the rule-of-thumb price setters of Gali and Gertler is completely erratic. They do not aim at hitting any precisely defined price level.

[7] This source of persistence will be called 'extrinsic' because it is an outcome of a mechanism that is treated as exogenous in the model.

possible (extrinsic or intrinsic) sources of sluggishness, one has to possess a proxy of the marginal cost (or, more precisely, of the desired price) on the level of particular firms. Needless to say, such a proxy is hard to be constructed. Therefore, in this paper we do not attempt to definitely discriminate one hypothesis against the other. Instead, we design an alternative model, which reformulates our research question to be consistent with a dual set of assumptions: In this dual approach, we return to the standard postulate that the price adjustment is complete. Instead, it is now assumed that the shocks to the (common component of the) desired individual price level follow an AR(1) process. We reestimate our model and reinterpret the results in the spirit of the alternative formulation.

Preferably, we would not make any specific assumptions about the timing of price adjustments in our models, so that they would be consistent with the time-dependent rules like in Taylor (1980) or Calvo (1983) as well as with the state-dependent rules studied in the menu-costs literature. Yet, totally ignoring factors that incite the firm to adjust its prices in a given period may have a detrimental effect on the unbiasedness of the estimates. Therefore, our framework contains a rule that permits to flexibly model the price adjustment mechanism. Its detailed description as well as the exact method used for estimation of the models are given further in the text.

The remainder of the paper is organized as follows. Next section discusses our data set and data-related issues. We develop the models and discuss their estimation strategy in Section 3. The results of estimation are presented in Section 4. This section also contains an attempt to assess the plausibility of both models. Section 5 concludes.

2 Data.

The data set used is a panel of micro-prices underlying the Belgian CPI. These data are collected on a monthly basis by the Federal Public Service Economy, Self-employed and Energy. We have at our disposal a sample of prices of products belonging to 368 different categories, with a monthly average number of observations per product category equal to 285. The weights of these products sum up to 70 % of the total basket of goods and services covered by the CPI. These are the products for which CPI compilers monitor prices nationwide in more than 10,000 outlets. The remaining 30 % covers items such as housing rents, telecommunications services, newspapers, insurance services, cars etc. for which prices are monitored centrally, either because a unique price applies or because these prices require a specific methodology of monitoring.

The data on 222 product categories are observed from July 1994 to February 2003 and hence for 104 months. The data on the other 146 product categories span over an 86 months period, that is from January 1996 to February 2003. A more detailed description of a sample of Belgian micro-prices covering an ovelapping although longer time span may be found in Aucremanne and Dhyne (2004).

The existence of promotions in our data introduces the effect of mean-

reversion to the individual price changes. In order to net out this effect we try to disregard price changes the most distinctly having to do with price discounts. Our definition of a promotion is the following:

$$p_{it} \text{ is a discount price} \Longleftrightarrow \begin{cases} p_{it} < p_{it-1} \\ p_{it-1} = p_{it+1} \end{cases}$$

where p_{it} is a price of the good and the index i refers to a particular price trajectory. This definition is identical to the one used by Aucremanne and Dhyne (2005) in their internal National Bank of Belgium paper on a similar issue.

3 Estimation Strategy.

Let us introduce some basic notation. T is the overall number of periods (months) covered by our sample. $t \in T$ is the t-th consecutive month considered. i refers to a particular outlet selling the product of interest. The price of this product in outlet i (we will say simply 'the price of item i') in period t will be denoted p_{it}. Prices are adjusted infrequently. We will denote the moment (month) of k-th price change of item i by τ_{ik}. When it is clear that we speak about the same item, we will often skip the subscript i and write τ_k. The k-th newly set price of item i will be denoted $x_{i\tau_k}$. Hence, we have:

$$p_{it} = \begin{cases} x_{it} \text{ for } t \in \{\tau_k\} \\ p_{it-1} \text{ otherwise} \end{cases}$$

Let p^*_{it} denote the price level in period t that maximizes the expected discounted value of profits of an individual i. We will refer to p^*_{it} as the desired or fundamental price of i. It may be decomposed as:

$$p^*_{it} \equiv f_t + g_{it} + g_i$$

where f_t is a component common to all outlets selling products from a given sector; g_{it} is an outlet's specific component; finally, g_i is an outlet's specific price level effect.

3.1 Model of Gradualism.

Our first goal is to construct a simple model of price setting with incomplete adjustment. This model will serve to examine whether firms, once they decide to make an adjustment, set their new price x_{it} equal to the fundamental price p^*_{it}. The alternative hypothesis is that firms adjust their prices to the desired price level only gradually.

As it was explained in the introduction, one implication of the incomplete price adjustment is that consecutive individual price changes are serially correlated. Therefore, in principle, measuring the magnitude of this correlation can reveal useful information about the degree of price setting gradualism in the

economy. There are two important issues related to this approach. First, the correct interpretation of the degree of autocorrelation of price changes in terms of the implied magnitude of incomplete adjustment will depend on a particular model of gradualism postulated. One may for instance assume that an amount of information about the magnitude of marginal cost innovation is revealed to the individuals every period at a constant rate. Then, it would be natural to relate the degree of autocorrelation found to the rate of information arrival. Alternatively, the newly set price of an individual might be a weighted average of the old price and the given period fundamental price. Then, the obtained autocorrelations should be interpreted in terms of the weights they are consistent with. We will stick to the latter approach since, as it will be shown, it mirrors the usual assumptions made in the partial adjustment literature.

Second, gradualism is not the only potential cause of autocorrelation in price changes. Autocorrelation may also have an extrinsic origin in the innovations of the fundamental price being themselves serially correlated. In order to distinguish between one form of sluggishness and the other, one needs a good proxy for the fundamental price innovations on the level of individual firms, a proxy that we do not have[8]. Note however that the assumption that the desired price follows a random walk is standard in the literature (see for instance Caballero and Engel, 2004). In particular, Bils and Klenow (2004) argue that the random walk is a good approximation of the nominal marginal cost process in the US. This, under fairly general conditions, implies that also the fundamental price process is a random walk. In the model of gradualism discussed here, we will follow the literature. We will take a closer look at the time series properties of the fundamental price in the model we are going to develop in the next subsection.

The above remarks imply the following two assumptions on which the model of gradualism will be founded:

A1. Random walk of the fundamental price. Consider the first order differences:

$$\Delta p^*_{it} \equiv p^*_{it} - p^*_{it-1} = \Delta f_t + \Delta g_{it}$$

Δf_t and Δg_{it} have the following properties:

1. $\Delta f_t \equiv v^A_t$ and v^A_t is iid with mean $Ev^A_t = \mu$ and variance $Var(v^A_t) = \sigma^2_A$.
2. $\Delta g_{it} \equiv v^I_{it}$ and v^I_{it}'s are independent across individuals and time, identically distributed with mean $Ev^I_{it} = 0$ and variance $Var(v^I_{it}) = \sigma^2_I$.
3. $cov(v^A_t, v^I_{it+s}) = 0$ for any i, t and s.

A2. Gradualism. The price adjustment, when it occurs, is incomplete:

$$x_{i\tau_k} \equiv (1-\alpha)p_{it-1} + \alpha p^*_{i\tau_k} \equiv (1-\alpha)x_{i\tau_{k-1}} + \alpha p^*_{i\tau_k}$$

where $2 > \alpha > 0$.

[8] Under the assumption of a constant desired mark-up, a proxy for the marginal cost process would suffice. Unfortunately, on the micro-level, this proxy is no less difficult to be constructed.

The notation adopted for the random walk assumption A1 is that of Caballero and Engel (2004). Following that work[9], it postulates that both, the common and the idiosyncratic components of the firm's fundamental price follow a random walk. Assuming that the firm's specific component is a random walk may be argued not to be entirely realistic, as it implies that prices of the same product sold in different outlets may be divergent in the long-run. We need this assumption however in order to set up a partial adjustment model that is feasible to estimate. We will consider a more realistic fundamental price process in the setting with extrinsic persistence, where it is easier to incorporate a less sluggish indiosyncratic component.

The gradualism assumption A2 relaxes the postulate of complete price adjustment made in the baseline New-Keynesian model. The particular form of gradualism considered is like in the Partial Adjustment Model of Jorda (1999)[10]: The newly set price $x_{i\tau_k}$ is a weighted average of the old price $p_{it-1} \equiv x_{i\tau_{k-1}}$ and the fundamental price $p^*_{i\tau_k}$. Then, the weight parameter α has the interpretation of the individual speed of adjustment, see Jorda (1999). For the adjustment to be gradual we must have $0 < \alpha < 1$. $\alpha \equiv 1$ is equivalent to complete adjustment. We also allow for having $\alpha > 1$, which would imply price overshooting.

In order to complete the set of assumptions underlying the model, we have to look closer at the rule according to which managers decide on the exact moment of a price adjustment. There is little doubt that the timing of adjustments is non-random. This fact must be taken into account in the specification of our model, or otherwise our estimates will not be reliable, an effect of the so called non-random selection bias[11]. Unfortunately, estimation of dynamic panel data models with incorporated a non-random selection mechanism is not an easy task. The most popular perhaps are the 2-steps estimation methods based on the path-breaking idea of James Heckman (e.g. Wooldridge, 1995; Semykina, 2007). These methods tend to trade off efficiency for computational tractability and hence cannot be our first choice[12]. The drawback of the more efficient methods is that they are usually highly computationally demanding and hence turn out infeasible for models with too many variables. This imposes the requirement of compactness on the model to be estimated. As an acceptable compromise between flexibility and estimability of our model, we have chosen to estimate it by the computationally expensive maximum likelihood method, but we consider only a reduced form of the timing-of-adjustment rule. This rule will be governed by what is usually called in the literature a selection equation, as it is defined in the assumption A3:

[9] Bils and Klenow in their 2004 paper appear to have implicitly made the same assumption.

[10] The only difference being that Jorda set up his model in the continuous time.

[11] For two text-book treatments of the non-random selection-related estimation issues see Maddala (1986) and more recent Hsiao (2003).

[12] The frequency of price adjustments is very low for many of the goods in the sample. In consequence, in the case of these goods, only a few data points per period may be used for the estimation of the dynamic properties of the price series. Hence, any loss of efficiency may heavily weigh on the reliability of the obtained results.

A3. Selection equation. Let define an indicator variable d_{it}, taking a value of 1 in the periods of price adjustments, and 0 otherwise. d_{it} is governed by a latent variable y_{it}^* according to the rule:

$$d_{it} = I\left[y_{it}^* \geq 0\right]$$

where $I\left[\cdot\right]$ is the usual indicator function. The latent variable y_{it}^* itself is a linear function of L state- and time-dependent variables w_{it}^l, and of an unobserved firm specific component c_i, so that:

$$y_{it}^* = \sum_{l=1}^{L} \beta_l w_{it}^l + c_i + \delta_{it}$$

where δ_{it} is an additional error term. Further, the individual unobserved effect c_i may be decomposed as a linear combination of M exogenous variables z_i^m and an error term u_i:

$$c_i = \sum_{m=1}^{M} \eta_m z_i^m + u_i$$

such that $cov(u_i, \delta_{jt}) = 0$ for all i, j and t.

Intuitively, the assumption A3 postulates that firms' decisions to adjust their prices depend on a set of variables w_{it}^l, such as the season of the year, the number of months since the last price adjustment, the cumulated inflation since the last price adjustment etc. The firms' individual probabilities of adjustment may differ one from another through the unobserved firm specific effect c_i. c_i is allowed to be correlated with the error terms δ_{it}. However, it is postulated that the part of c_i that is correlated with δ_{it} may be explained by a linear combination of variables z_i^m, being the observable characteristics of a given firm. The residual u_i is independent of δ_{it}. The assumption A3 implies that the selection equation may be written as:

$$y_{it}^* = \sum_{l=1}^{L} \beta_l w_{it}^l + \sum_{m=1}^{M} \eta_m z_i^m + u_i + \delta_{it}$$

For $t = \{\tau_{k-1} + 1, ..., \tau_k\}$, let define $\Delta_t^k \equiv t - \tau_{k-1}$, the number of periods since the last (kth-1) price adjustment. We will use the symbol $\Delta^k x_i$ to denote the size of the price change in period τ_k:

$$\Delta^k x_i \equiv x_{i\tau_k} - p_{i\tau_k - 1} \equiv x_{i\tau_k} - x_{i\tau_{k-1}}$$

In order to see what the assumptions A1-A2 imply for our model, construct an artificial latent variable $\Delta^k x_{it}^*$ that is explained by the following dynamic equation:

$$\Delta^k x_{it}^* \equiv \gamma \Delta^{k-1} x_{i\tau_{k-1}} + (1 - \gamma)\left(p_{it}^* - p_{i\tau_{k-1}}^*\right) \quad \text{for } t = \{\tau_{k-1} + 1, ..., \tau_k\}$$

where $\gamma \equiv 1-\alpha$ will be interpreted as a measure of the degree of gradualism and $|\gamma| < 1$. From the assumption A1, the above may be written as:

$$\Delta^k x_{it}^* = \gamma \Delta^{k-1} x_{i\tau_{k-1}} + (1-\gamma)\mu\Delta_t^k + \epsilon_{it} \quad \text{for } t = \{\tau_{k-1}+1, ..., \tau_k\}$$

where $\epsilon_{it} \equiv (1-\gamma) \sum_{s=\tau_{k-1}+1}^{t} \left(v_s^A + v_{is}^I - \mu\right)$ with mean $E(\epsilon_{it}) = 0$ and variance

$$var(\epsilon_{it}) \equiv \Delta_t^k (1-\gamma)^2 (\sigma_A^2 + \sigma_I^2) \equiv \Delta_t^k \sigma^2$$

Observe that in periods $t = \tau_k$ so defined latent variable $\Delta^k x_{it}^*$ is identical to the size of the price change in the same period:

$$\Delta^k x_i \equiv \Delta^k x_{i\tau_k}^*$$

Indeed, according to the partial adjustment equation in the assumption A2, we have:

$$\Delta^k x_i = \gamma \Delta^{k-1} x_i + (1-\gamma)(p_{i\tau_k}^* - p_{i\tau_{k-1}}^*)$$

and hence:

$$\Delta^k x_i = \gamma \Delta^{k-1} x_i + (1-\gamma)\mu\Delta_t^k + \epsilon_{i\tau_k}$$

This allows us to write all the equations of our partial adjustment model as[13]:

$$\begin{array}{r} y_{it}^* = \sum_{l=1}^{L} \beta_l w_{it}^l + \sum_{m=1}^{M} \eta_m z_i^m + u_i + \delta_{it} \\ \Delta^k x_{it}^* = \gamma \Delta^{k-1} x_{i\tau_{k-1}} + (1-\gamma)\mu\Delta_t^k + \epsilon_{it} \\ \Delta^k x_i = \Delta^k x_{it}^* * d_{it} \text{ with } d_{it} = I(y_{it}^* \geq 0) \end{array} \qquad \text{for } t = \{\tau_{k-1}+1, ..., \tau_k\} \quad \text{(M)}$$

All observable variables except $\Delta^k x_i$ are assumed strictly exogenous. The error terms are serially independent, and both, the error terms and the unobserved effects are independent across individuals. They are assumed to follow a joint normal distribution with the variance covariance structure:

$$\begin{pmatrix} \epsilon_{it} \\ \delta_{it} \\ u_i \end{pmatrix} \tilde{} N \left[\begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \Delta_t^k \sigma^2 & r\sqrt{\Delta_t^k}\sigma & 0 \\ r\sqrt{\Delta_t^k}\sigma & 1 & 0 \\ 0 & 0 & q^2 \end{pmatrix} \right]$$

for all i and t. The parameter r is the coefficient of correlation between ϵ_{it} and δ_{it} and q^2 is the variance of the individual effect u_i. Note also the normalization $var(\delta_{it}) = 1$. We assumed a priori independence between the individual effects and the main equation error term ϵ_{it}.

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[13] Note that the error term $\epsilon_{i\tau_k} \equiv (1-\gamma) \sum_{s=\tau_{k-1}+1}^{\tau_k} \left(v_s^A + v_{is}^I - \mu\right)$ in the transformed partial adjustment equation contains common shocks v_t^A's that are estimable in the panel model. However, an attempt to estimate them would require an additional hundred dummy variables to be added to the main equation (recall that for most products the sample extends over 104 periods). Hence, the fact that we do not have to estimate the common shocks under the chosen timing-of-adjustment rule is precisely what makes the maximum likelihood estimation of our model feasible.

3.2 Dual Approach: Relaxing the Random Walk Assumption.

The assumptions A1-A2 require that we interpret any potential evidence of serial correlation of price changes as related solely to the intrinsic source of persistence. Yet, as we argued before, the additional persistence may also be generated by a fundamental price process being more sluggish than the random walk. Since distinguishing between one source of sluggishness and the other is difficult, we would like at least to be able to reinterpret our findings in the other possible manner. Shortly speaking, we would like to be able to say what (extrinsic) persistence of the fundamental price would be necessary in order to produce the observed autocorrelation of the individual price changes. The most straightforward way to answer this question is to estimate directly a dual model that would be correctly specified under the extrinsic source of persistence hypothesis. To construct such a model, let us first propose a set of assumptions dual to those used in the previous setting:

D1. Dynamic properties of the fundamental price. Consider the fundamental price p^*_{it}. Its common component f_t and its idiosyncratic component g_{it} have the following properties:

1. $\Delta f_t \equiv \rho \Delta f_{t-1} + v^A_t$ and v^A_t is iid. The mean of v^A_t is $(1-\rho)\mu$ and its variance σ^2_A. The autoregressive parameter ρ is such that $|\rho| < 1$.

2. $g_{it} \equiv v^I_{it}$ and v^I_{it}'s are independent across individuals and time, identically distributed with mean $Ev^I_{it} = 0$ and variance $Var(v^I_{it}) = \sigma^2_I$.

3. $cov(v^A_t, v^I_{it+s}) = 0$ for any i, t and s.

D2. Completeness of the adjustment. The price adjustment, when it occurs, is complete:

$$x_{i\tau_k} = p^*_{i\tau_k}$$

The dual assumption D2, by postulating complete adjustment of prices, shuts down the intrinsic persistence channel. The assumption D1 changes the earlier assumption A1 in a double manner. First, by postulating that the common factor f_t follows an ARIMA(1,1,0) process, it explicitly opens up an extrinsic persistence channel. Second, it also assumes that the idiosyncratic factors do not have a unit root. This is introduced to add some reality to the model and guarantees that the differences between prices of different outlets selling the same good will not diverge to infinity in the long run, as it could be the case in our partial adjustment setting. Nonetheless, the particular white noise assumption we made here is somewhat extreme, in that it rules out any additional persistence introduced by the idiosyncratic shocks. While this assumption is consistent at least with a part of the literature (e.g. Dhyne et al, 2006), to the best of our knowledge it is not founded on any empirical evidence. As the individual shocks are likely to be moderately positively autocorrelated, our estimates of the magnitude of the persistence of the common factor may be biased

upward. Taking this into account, one should interpret the estimated values of parameter ρ as the upper bound of the true correlation.

We do not modify the form of the adjustment timing rule. Hence, the assumption A3 remains valid. However, we have to add one more assumption concerning the levels of prices of a particular product sold in different outlets. This step was not necessary when developing the model of gradualism, since the firm's specific price level effect g_i cancelled out in the model's main equation. Here we will have to be explicit about its properties. This is the motivation for the assumption D4:

D4. Decomposition of the individual price level effect g_i. The individual price level effect g_i may be decomposed as a sum of a linear combination of $\tilde{M}$ exogenous variables z_i^m and a zero-mean error term v_i:

$$g_i = \sum_{m=1}^{\tilde{M}} \tilde{\eta}_m z_i^m + v_i$$

The unobserved effect v_i of outlet i is independent of the idiosyncratic shocks v_{jt}^I, $cov(v_i, v_{jt}^I) = 0$ for all i, j and t as well as of the unobserved effects of the other outlets, $cov(v_i, v_j) = 0$ for all i, j.

In principle, the dual model could be estimated by the maximum likelihood method. However, as this time we cannot avoid an explicit estimation of the common factor for all periods, the maximum likelihood method becomes computationally too cumbersome. Therefore, for estimating the model, a 2-step procedure has been devised. The purpose of its first step is to consistently estimate the common factor f_t, ignoring its dynamics. To this end, we estimate the following model for every period separately:

$$\begin{aligned} y_{it}^* &= \sum_{l=1}^{L} \beta_l w_{it}^l + \sum_{m=1}^{M} \eta_m z_i^m + \xi_{it} \\ p_{it}^* &= f_t + \sum_{m=1}^{\tilde{M}} \tilde{\eta}_m z_i^m + \epsilon_{it} \\ x_{it} &= p_{it}^* \text{ for } y_{it}^* \geq 0 \end{aligned} \qquad \text{(MD step1)}$$

where we defined: $\xi_{it} \equiv \delta_{it} + u_i$ and $\epsilon_{it} \equiv g_{it} + v_i$. Since the model is estimated cross-section by cross-section, one cannot estimate the unobserved effects u_i and g_i and they simply enter the error terms ξ_{it} and ϵ_{it} respectively. The error terms are jointly normal:

$$\begin{pmatrix} \epsilon_{it} \\ \xi_{it} \end{pmatrix} \tilde{} N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2 & r\sigma \\ r\sigma & 1 \end{pmatrix} \right]$$

where r is the coefficient of correlation between ϵ_{it} and ξ_{it} and $var(\epsilon_{it}) = \sigma^2$. As usually, the variance of the error term in the selection equation is normalized to 1.

The first step may be estimated by the maximum likelihood and renders an estimate of the common factor $\hat{f}_t$. This estimate may be then decomposed as:

$$\hat{f}_t = f_t + e_t$$

where e_t is an error whose variance is inversely proportional to the number N_t of price changes in period t. Hence, in the second step one considers the model:

$$\begin{aligned} \hat{f}_t &= f_t + e_t,\ var\left(e_t\right) = N_t^{-1} q^2 \qquad \text{(MD step2)} \\ \Delta f_t &\equiv \rho \Delta f_{t-1} + v_t^A \end{aligned}$$

with the constant $q > 0$ to be estimated. Assuming that the error term e_t is serially independent, $cov\left(e_t, e_{t-s}\right) = 0$ for all t and s, the persistence parameter ρ, which is the parameter of interest, may be easily estimated by the Kalman filter. In the reality, the errors e_t are likely to be slightly positively autocorrelated due to the suppression of the unobserved effects v_i as well as because the idiosyncratic components may be positively autocorrelated on their own[14]. Hence, we stress again that, properly speaking, we only estimate the upper bound of the persistence parameter ρ.

The clear disadvantage of the proposed procedure is that in its first step it requires an estimation of a number of coefficients for every period separately. Therefore it cannot be applied to products characterized by an infrequent price adjustment, as there is not enough price changes for the estimation of L coefficients β_l and M coefficients η_m in the selection equation. Risking an improper specification of the model, one could try to decrease the number of variables in the selection equation. In the limit, one could resign at all of the selection equation, thus gaining the degrees of freedom at the risk of the emergence of the selection bias. For the sake of comparison, we will later present the results of this ad hoc estimation strategy. In this case we replace the first step model (MD step1) with a simple random effects panel regression:

$$x_{it} = f_t + \sum_{m=1}^{\tilde{M}} \tilde{\eta}_m z_i^m + v_i + g_{it} \quad \text{for all } t \in \{\tau_k\} \qquad \text{(MD step1b)}$$

3.3 The choice of explanatory variables.

The set of explanatory variables entering the selection and the main equations is the following. As the time-dependent factors influencing the adjustment decision, we considered:

- The cumulative number of months passed from the last price adjustment;
- The 1, 3, 6 and 12 months dummies.

As the state-dependent variable correlated with the adjustment timing decision we chose the absolute value of the cumulated inflation since the last price

[14] This will be the case if the assumption D1.2 does not exactly hold.

adjustment. For the parametrization of the individual unobserved effect u_i two variables were used. The first one is the average position of the particular firm on the low price - high price scale during the whole available period, normalized to lie between 0 and 1. This allows for different adjustment thresholds for outlets placed on the high/low end of the market. The second variable used was the proportion of price decreases in the overall number of price adjustments for a particular outlet. This variable is aimed to capture the possible distinct adjustment thresholds for firms that are more akin to decrease their prices. No variable was used for the parametrization of the individual effect v_i. Admittedly, the parametrizations of the effects may not be sufficiently rich. Other characteristics of the individual firms could be however added.

4 Results.

4.1 Model of gradualism: Do price-setters overshoot?

Table 1 contains a summary of the estimation results for the gradual adjustment model (M).

Table 1.

Product Group	##	$\gamma < 0$	γ insignif.	$\gamma > 0$	Average γ	Median γ
UNPF	66	66 (100%)	0 (0%)	0 (0%)	-0.441	–0.456
PF	73	69 (94.5%)	1 (1.4%)	3 (4.1%)	-0.413	-0.467
-Tobacco	4	1 (25%)	1 (25%)	2 (50%)	0.136	0.113
-Others	69	68 (98.6%)	0 (0%)	1 (1.4%)	-0.444	-0.472
Energy	11	7 (63.6%)	2 (18.2%)	2 (18.2%)	-0.042	-0.022
NEIG	170	46 (27.1%)	16 (9.4%)	108 (63.5%)	0.112	0.187
-Clothes etc.	50	2 (4.0%)	4 (8.0%)	44 (88.0%)	0.354	0.398
-Furniture etc.	17	0 (0%)	1 (5.9%)	16 (94.1%)	0.372	0.373
-House-care	7	7 (100%)	0 (0%)	0 (0%)	-0.299	-0.309
-Flowers	10	8 (80%)	0 (0%)	2 (20%)	-0.258	-0.338
-Cosmetics	11	9 (81.8%)	1 (9.1%)	1 (9.1%)	-0.262	-0.329
-Others	75	20 (26.6%)	10 (13.3%)	45 (60%)	0.036	0.126
Services	48	9 (18.8%)	16 (33.3%)	23 (47.9%)	0.107	0.078
-Food Services	16	3 (18.8%)	6 (37.5%)	7 (43.8%)	0.108	0.035
-Others	32	6 (18.8%)	10 (33.3%)	16 (47.9%)	0.107	0.117

The estimation results for the coefficient of gradualism γ.

##-number of products in a group.

Table 1 breaks the products into 5 main categories, a division that is frequently applied in the economic literature. These five categories are: Unprocessed Food (UNPF), Processed Food (PF), Energy, Non-Energy Industrial Goods (NEIG) and Services. This division was argued by some authors to well characterize the differences in the frequency of price adjustment between goods in Belgium (see Aucremanne and Dhyne, 2004), with prices of energy and unprocessed food changing the most frequently and prices of NEIGs and especially

services being adjusted very rarely. It does not however appear equally useful in grouping the products according to the degree of gradualism prevailing in the sector. Especially the NEIGs occur to be non-homogeneous with respect to this criterion. For this reason within some main categories we isolated a few groups of products that seem to display a bit more homogeneity across the estimated coefficients[15].

Table 1 reports three general categories of products, for which the gradualism parameter γ is typically estimated to be negative. This is unprocessed and processed foodstuff (γ virtually always negative) and energy goods. Under the assumption A2, negative γ's imply price overshooting (the speed of adjustment parameter $\alpha > 1$). Hence, the interpretation of the results is that the outlets offering these goods are likely to overreact to shocks to the fundamental price. What follows, for these product categories, there is no evidence of any additional source of intrinsic persistence above that implied by the infrequent adjustment. To the contrary, our findings suggest that in the case of the food and energy goods the sector inflation indices should be less sluggish than the baseline New-Keynesian model predicts. The only exception among these goods are tobacco products, whose price changes seem to be positively autocorrelated. This difference might be related to the fact that the prices of tobacco, in contrast to the prices of an overwhelming majority of products in the sample, are regulated.

For the remaining two product categories (non-energy industrial goods and services), we did find some evidence of gradualism. In the case of services, the gradualism coefficient γ was found to be significantly positive in almost 50% of cases. For comparison, only about 20% of services seems to be characterized by negative γ. The NEIG group is the largest and the least homogeneous among the main five product categories. This is reflected in the estimates of the gradualism parameter γ. For two sub-groups of NEIGs these estimates are virtually uniformly positive ('Furniture and Non-Mechanical House Equipment' and 'Clothes and Shoes'). However there are many goods in this category whose prices appear to be characterized by overshooting (house-care products, flowers, cosmetics). The dichotomy of NEIG category is well visible on Figure 1, which depicts the distribution of the estimates of γ for this group of products. One can see that the distribution is bimodal, with a higher mode on the positive side and a lower mode on the negative one.

In terms of the magnitudes of the estimates, Table 1 indicates that the departures from the complete adjustment in the case of the energy commodities (median $\gamma = -0.022$) and in the case of services (median $\gamma = 0.078$) are tiny and hence are likely to have a limited influence on the dynamics of aggregate inflation indices. On the other hand, the absolute values of the persistence coefficient appear substantial for the remaining product categories. The median parameter of gradualism for the whole industrial goods group is 0.187. Be-

[15] These groups are: Within the PF products: Tobacco (the COICOP codes 2.2.x.x). Within the NEIG products: Clothes and Shoes (3.1.1.x-3.1.3.x, 3.2.1.x), Furniture and Non-Mechanic House Equipment (5.1.x.x-5.2.x.x), House-care Goods (5.6.1.x), Flowers (9.1.7.x), Cosmetics (12.1.2.x). Within Services: Food Services (11.1.x.x).

hind this aggregate measure a considerable heterogeneity is concealed. Prices in certain NEIG sub-categories display a much higher degree of incomplete adjustment. The median estimate of γ within the 'Furniture etc.' subcategory equal to 0.37 implies the median speed of adjustment $\alpha = 0.63$. The median speed of adjustment of prices of 'Clothes and Shoes' is even lower and equals about 0.60. This implies that a median outlet offering clothes or shoes, when resetting its price, catches up with the shocks that have accumulated from the last price adjustment only in 60%. This is striking, as a large part of prices in this sector changes as infrequently as once every 2 years (see Aucremanne and Dhyne, 2004).

Other NEIG subcategories (in particular house-care goods, flowers and cosmetics) appear to be characterized by significantly negative values of the persistence parameter γ (the medians all below -0.3). The implied degree of overshooting in the case of these products makes them similar to food. For both food categories the median estimate of γ is found to be around -0.46. This gives the median degree of overshooting α equal to 1.46. If the model of gradualism is correct, this number would imply that the agents in the food industry overshoot the innovations to the fundamental price by almost as much as 50%.

Our findings so far suggest that there may be sectors (certain NEIG industries and potentially services) characterized by additional intrinsic price sluggishness due to gradualism. This feature is likely to add to the persistence of the sector inflation indices. However, other sectors (especially food industry) display price overshooting. This will decrease the inflation persistence. No theoretical model exists that allows for assessing which of these effects should prevail on the aggregate level. Hence, we do not find clear-cut evidence that would provide us with a rationale for introducing an additional source of intrinsic sluggishness into the one sector New-Keynesian model, as it is often done by economists (Christiano and Fisher, 2004; Christiano et al, 2001; Gali and Gertler, 1999; Smets and Wouters, 2003 and many more).

The above observation is strengthened by the fact that we transformed our initial panel in order to eliminate promotions. Clearly, the presence of promotions in the data introduces another source of negative correlation between individual price series. This obviously implies that the raw data would provide even less support for the existence of an additional source of persistence than the transformed data we are looking at in this paper did. The effect of not removing promotions is clearly visible on Figure 2, where the estimates based on the transformed data, against the estimates obtained on the basis of the raw data are depicted. The difference is especially pronounced for the industries where price overshooting appears to be common, while eliminating promotions does not have a very strong impact in the case of incomplete adjustment goods.

Note that our procedure of excluding discounts is unlikely to be fully effective. In particular, it does not exclude promotions longer than a month. Similarly, it omits discounts that are followed by a price higher than the price charged before the special was established. One could therefore wonder whether the non-eliminated promotions do not lie at the origin of the negative serial correlation of price changes found for many products. To test for this possibility

we experimented with a number of wider definitions of promotions. Although transforming the data according to these alternative definitions results in a quantitative change of estimates in the predicted direction, it does not alter the general pattern of our findings described in this section[16].

It should be stressed that, although not clearly supportive for the rule-of-thumb behavior of consumers, the obtained results challenge the deep fundaments of the baseline New-Keynesian model. In particular, if our model of gradualism is interpreted literally, the finding of price overshooting in certain wide product categories casts into question the consumers' full rationality presumption made in the standard versions of the New-Keynesian framework. Alternatively, it could be seen as a case against the frequently made in the literature (e.g. Caballero and Engel, 2004) 'random walk assumption'. This last interpretation is consistent with our dual framework, to which we now turn.

4.2 Dual model: Is the fundamental price a random walk?

The finding from the previous subsection that there is a considerable diversity of price setters' behavior in distinct product categories suggests an alternative explanation of our results. They may simply reflect differences of persistence of the fundamental price processes prevailing in different sectors. Under this hypothesis the gradualism assumption A2 is not correct. Instead, one should rather consider the model MD.

The results of the estimation are summarized in Table 2. The columns under the 'Model D' heading refer to the version of the dual model that incorporates the timing-of-adjustment rule. We could not estimate it for services. Also, for a vast majority of industrial goods there are not enough price changes in the sample for this version of the model to be estimated. For those product categories for which the estimation was feasible (these are mainly products characterized by a relatively high frequency of price adjustments), the estimates of the median persistence parameter ρ are usually very close to zero. This is so in particular for the foodstuff and the energy goods. There seems to be a little more diversity in the case of the NEIGs. Some evidence of mean-reversion (negative parameter ρ) was found for flowers. This result is consistent with our previous finding of negative gradualism parameter γ for the products belonging to this subcategory (see Table 1). Surprisingly, the persistence parameter appears slightly positive for the cosmetics, and more so for the house-care goods (median estimate of ρ equal to 0.086 and 0.328 respectively). None of these departures from the random walk assumption seems to be a robust finding however: Most of the

[16] The widest alternative definition of promotions considered was the following:

$$p_{it} \text{ is a discount price} \iff \begin{cases} p_{it} < p_{it-1} \\ p_{it-1} \leq p_{it+1} \end{cases} \quad or \quad \begin{cases} p_{it} = p_{it-1} < p_{it-2} \\ p_{it-2} \leq p_{it+1} \end{cases}$$

The application of the above definition of promotions to the house-care products, cosmetics and flowers results in an increase of the median of the estimated coefficient γ for each of these categories by about 0.1. Still, all these medians remain at the level lying very significantly below 0. The detailed results of the estimation are available from the author upon request.

estimates were found not to be significantly different from zero[17].

Table 2.

	Model D			Model Db		
Product Group	##	Average ρ	Median ρ	##	Average ρ	Median ρ
UNPF	65	-0.024	-0.050	66	-0.104	-0.202
PF	59	0.037	0.050	73	0.060	0.049
-Tobacco	0	—	—	4	0.390	0.368
-Others	59	0.037	0.050	69	0.041	0.032
Energy	8	0.098	0.085	11	0.110	0.072
NEIG	29	0.086	0.046	170	0.001	-0.019
-Clothes & Shoes	0	—	—	50	0.042	0.015
-Furniture etc.	0	—	—	17	-0.082	-0.024
-House-care	6	0.291	0.328	7	-0.035	-0.010
-Flowers	6	-0.052	-0.170	10	-0.074	-0.259
-Cosmetics	9	0.204	0.086	11	0.071	0.005
-Others	8	-0.175	-0.205	75	-0.005	-0.031
Services	0	—	—	48	-0.037	-0.013
-Food Services	0	—	—	16	-0.080	-0.030
-Others	0	—	—	32	-0.016	-0.008

The estimation results for the coefficient of persistence ρ.

##-number of products in a group.

It may be instructive to look at the relation between the frequency of price adjustments and the magnitude of the estimated ρ (Figure 3a). As expected, the average estimated value of the AR parameter is about zero regardless of the frequency of price adjustments. However the dispersion of the estimates is very high for the low frequencies, while it becomes visibly lower for the higher ones. This effect is likely to be caused by large standard errors of the estimates in the case of goods characterized by a low frequency of price adjustment, for which the number of price changes present in our sample is small. This is confirmed by inspection of Figure 3b, which depicts the estimates of ρ as a function of the number of price changes in the sample. Visibly, the dispersion of the estimates declines fast as the size of the sample increases.

For the sake of comparison, it is worth to also analyze the results of estimating the model MDb, which ignores the mechanism according to which outlets choose the timing of price adjustment. They are summarized in Table 2 under the heading 'Model Db'. As before, the estimates cluster around zero, with only a few product categories displaying any degree of persistence or mean-reversion. The most interesting for us are those groups of products, for which the fully-developed model could not be estimated. The median persistence parameter ρ

[17] This is the reason why we did not display the number of the significantly positive and negative estimates of ρ for every product type in Table 2, as we did for the partial adjustment model. One has to take into account however that, due to the inefficiency of our estimation method, the confidence intervals are wide.

is very close to zero for services as well as for the majority of non-industrial goods. In particular, unlike in the case of the gradual adjustment model (see Table 1), we do not find any evidence of additional persistence for the 'Clothes', 'Furniture' and 'other NEIG' subcategories. The only group characterized by a considerably positive degree of persistence of individual price changes are tobacco products, this being consistent with the findings from the previous model.

For the categories for which the results of the estimation of the model MD are also available, it might at first appear that the method that ignores the selection mechanism renders often relatively lower estimates of ρ. It is so in particular for the unprocessed food category and the house-care goods and flowers subcategories. However the converse is true for the 'other NEIG' subcategory[18]. As it is shown on Figure 4, the estimates obtained by the fully specified method can be very different from the ones obtained by the other, and the differences may be positive as well as negative. Hence, if there is any selection bias, its sign is unclear. Note also that the dispersion of the estimates is much higher in the case of the fully specified model[19]. While enriching the model with the selection equation helps to deal with the potential selection bias, it considerably inflates the standard errors of the estimates.

Regardless of the difficulties encountered during the estimation of the dual model, it should be clear that its results do not support the hypothesis that the innovations to the common factor of the fundamental price are mean-reverting or that they are highly persistent. This may be read as evidence in favor of the widespread in the econometric literature assumption (Caballero and Engel, 2004, Bils and Klenow, 2004) that the common factor of the fundamental price follows a random walk. This view is relatively strongly supported in the case of foodstuff and energy goods. There is less certitude as far as the services and non-energy industrial goods are concerned, since in their case the fully specified selection model could not be estimated. Nonetheless, the findings from the simplified model MDb suggest that there is not much persistence in the behavior of the fundamental price for these product categories either. One should therefore conclude that, according to the results of the dual model, there is only a little support for an 'extrinsic source of sluggishness' hypothesis. Hence, the results of estimation of this model do not provide evidence against the assumptions underlying the baseline New Keynesian framework.

4.3 Dual model vs. partial adjustment model: A comparison.

The messages implied by the results of estimation of the two competing models are clearly different. The findings obtained within the gradualism setting seem to be inconsistent with the assumptions underlying the standard New Keyne-

[18] Note however that the fully specified dual model could have been applied only to 8 products belonging to the 'other NEIG' subcategory, while the model MDb could be estimated for the whole set of 75 products.

[19] The standard deviation of the model MD estimates is 0.402, against the standard deviation of 0.317 for the model MDb.

sian framework. The contrary could be argued for the results from the dual model. This discrepancy is surprising: The models in this paper were constructed to offer two alternative interpretations of departures from the baseline New-Keynesian framework potentially revealed in the data. One would expect however that their results will not differ as to whether such departures exist or not. In this section we look for potential origins of this discrepancy. In particular, we examine whether a possibly misspecified structure of one of the models could not lead to a seeming inconsistency when the results of both models are compared.

Recall that one of the model M findings was that in the food sectors price setters overshoot, a behavior reflected in the negative correlation of consecutive price changes for products from these sectors[20]. This finding could be read as evidence that there is less persistence in the behavior of prices of foodstuff than what the New-Keynesian baseline would imply. Yet, this possibility was not corroborated by the estimation results of the dual model: We did not find a negative autocorrelation of innovations to the common factor for the food categories. One possible explanation of this discrepancy is that the assumption A1.2 of the gradual adjustment model, which postulates that the idiosyncratic component of the fundamental price is a random walk, is incorrect. We noted before that this postulate may be challenged on the theoretical grounds as implying an implausible behavior of prices. Here we will argue that its erroneous imposition on a model may in principle lead to an incorrect interpretation of the estimates of autocorrelation of price changes found in the data.

Consider an artificial economy similar to the one from the dual setting. In this economy the common component of the fundamental price follows a simple random walk, while its idiosyncratic component is a white noise. For the simplicity of exposition, assume that firms adjust their prices regularly every s periods. Let assume now that an econometrician erroneously believes that both components of the fundamental price contain a unit root and inquire the consequences of this belief. It is straightforward to show that the correlation between two consecutive price changes is:

$$corr\left(x_\tau - x_{\tau-s}, x_{\tau-s} - x_{\tau-2s}\right) = \frac{-\sigma_I^2}{s\sigma_A^2 + 2\sigma_I^2} = \frac{-1}{2 + s\frac{\sigma_A^2}{\sigma_I^2}}$$

where σ_A^2 is the variance of the common innovation to the fundamental price and σ_I^2 - the variance of its idiosyncratic component. The serial correlation is unambiguously negative, which will incite our econometrician to believe that the price setters overshoot. Note also that it converges to $-\frac{1}{2}$ as the ratio $\frac{s\sigma_A^2}{\sigma_I^2}$ tends to zero. Since the variance of the common shock is likely to be much smaller than the variance of the idiosyncratic component (see for instance Dhyne et al, 2006), we should expect the serial correlation to be strongly negative for those

[20] We found some evidence of negative autocorrelation of price changes also for the energy goods. As the estimated degree of overshooting in the case of this sector was tiny, we will not emphasize these results in the current section.

goods that are characterized by frequent price adjustments (and hence low s). This could potentially explain why we found (possibly spurious) evidence of price overshooting for the food sectors, in which prices change often, but not for most non-energy industrial goods and services. It could also explain why we do not find negative autocorrelations of innovations to the common component of the fundamental price when the dual model is estimated, as in this model we explicitly assumed the idiosyncratic components to be a white noise[21].

If the theory argued for above is to explain the discrepancies between the results of the two models, we should expect the goods characterized by more flexible prices to display price overshooting. Indeed, the inspection of Figure 5a, which depicts the estimates of the coefficient of gradualism γ as a function of the frequency of price adjustments suggests an inverse, hyperbolic-like relation between these two quantities. However, in order to find out whether there is a link between the frequency of price adjustment and the estimates of gradualism parameter, a more formal test is needed. Except for the above hypothesis, two other ones should be considered. First, since the frequency of price adjustments is tightly linked to the number of price changes within a given product category that are at our disposal, the apparent inverse relation between the former variable and the estimates of the coefficient γ could be simply a result of some type of small sample bias. Indeed, there seems to be an identical hyperbolic-like relation between the magnitude of the estimates of the gradualism parameter γ and the observed number of price changes of a given product, as it is clearly demonstrated on Figure 5b. Second, it could be also that the product categories characterized by a low (high) frequency of price adjustment are also characterized by partial adjustment (overshooting), but there is no link between these two dimensions within the categories. Then the pattern visible on Figure 5a would be a side effect of pooling together products having non-homogeneous characteristics.

In order to jointly test the three alternative hypotheses, the following regression was run for the whole set of products:

$$\hat{\gamma}_j = aFreq_j + bNoCh_j + \sum\nolimits_k c_k + error_j$$

where c_k's are dummy variables with k running over all 12 sub-categories[22]. The index j refers to a particular product. Except for the dummies, the estimates of γ are regressed on the variable $Freq_j$, which is a measure of the frequency

[21] The fact that a large non-persistent idiosyncratic component of the fundamental price can have a considerable downward effect on the autocorrelation of price changes could make a part of the explanation of the Bils and Klenow's (2004) puzzle that the sector inflation indices display significantly less sluggishness than it would be implied by the baseline New-Keynesian model calibrated to match the degree of nominal price rigidity prevailing in these sectors. Indeed, Bils and Klenow in their computations implicitly assume that it is the fundamental price, and not only its common component, that follows a random walk. This is a strong case for differentiating between the dynamics of the common and the idiosyncratic components in the empirical work.

[22] These categories are: UNPF products; Tobacco and other PF products; Energy; Clothes and Shoes, Furniture and Non-Mechanic House Equipment, House-care Goods, Flowers, Cosmetics and other NEIG products and Food Services and other Services.

of price adjustment for the product j and on the variable $NoCh_j$ counting the total number of price changes of a given product in the sample. If there is an inverse relationship between the estimates of the coefficient of gradualism and the frequency of price adjustments, the estimate of the parameter a should be negative and significant. Similarly, if in the result of a small number of data points available the estimates of γ tend to be biased upwards, the coefficient b should be significantly negative. Finally, if the observed differences in the estimates of the coefficient of gradualism are related to the category the products belong to, we should expect the dummies c_k to be jointly significant.

Table 3.

	$Freq$	$NoCh$	dummies c_k
Estimate	-.206	9E-6	—
t-Value (F-Value)	-1.11	1.97	56.95
p-value	0.266	0.049	<0.0001

For the dummy variables, the F-value reported is the III-type ANOVA test statistic. The estimates for the dummies c_k are not reported.

The results of the regression (reported in Table 3) seem to be most supportive for the last of the proposed hypotheses. The coefficient standing at the variable $Freq$ is negative, which could confirm a negative relation between this variable and the estimates of the degree of gradualism. However it is non-significant, which suggests that the relation is weak. The coefficient at the variable counting the number of price changes within a given product category appears marginally significant. Yet, contrary to the hypothesis we made, its sign is positive. Hence, the positive values of γ observed for categories characterized by a low frequency of price adjustments are unlikely to originate from a small sample bias. Finally, the analysis of variance reveals that the dummy variables indicating different product subcategories are jointly highly significant. Additionally, the overall fit of the regression is very good (adjusted R-squared equal to 0.69). We conclude that the division of our sample of products into more homogeneous product categories helps to explain the pattern of estimated measures of gradualism for different goods. At the same time, the relation between the frequency of price adjustment and the estimates of γ visible for the pooled sample of products is likely to be spurious. The link between the degree of price gradualism and the frequency of price adjustment stems from the fact that the product categories characterized by infrequent price changes are at the same time characterized by incomplete adjustment.

In the test proposed above, it has been assumed that the coefficients a and b are common across product categories. If this assumption is counterfactual, the results of the test may be misleading. An obvious solution to this problem is to estimate test regressions for every subcategory separately. Due to a small number of goods in some sub-categories, such a solution is not always feasible. The regressions performed for four biggest subcategories (UNPF, PF without tobacco, Clothes and Shoes and 'other NEIGs') fully support the results of the pooled test. In particular, the coefficient a at the variable measuring the

frequency of price adjustment is never found to be significant. This provides strong evidence against the hypothesis that the finding of overshooting for products characterized by frequent price adjustment in the model of gradualism was an effect of an incorrect assumption about the dynamics of the idiosyncratic component of the fundamental price.

The results of the models M and MD are hard to reconcile also in another dimension. If this is the dual model that is closer to reflect the behavior of price-setters in the true economy, the evidence of additional persistence found for some NEIGs and Services in the partial adjustment setting should indicate a positive autocorrelation of innovations to the common component of the fundamental price processes characterizing these products[23]. Unfortunately, no evidence of additional persistence was found for most industrial goods or services in the dual setting (see Table 2). However, this inconsistency could be a result of an insufficient number of price changes in the sample for these product categories, or of the non-random selection bias[24]. Therefore, we cannot definitely rule out the possibility that in the case of some non-energy industrial products and possibly also services there exists an extrinsic source of price sluggishness.

Given the observations above, it must be concluded that at this stage it is premature to decisively argue in favor or against the standard New-Keynesian framework. The findings from the partial adjustment setting may be interpreted as evidence against this framework. However, in order to make the estimation of the model of gradualism feasible, we had to assume a unit root in the specific component of the firms' fundamental price. We have shown that, if the true idiosyncratic component process of a firm does not contain a unit root, the imposition of this assumption might in principle lead to an incorrect interpretation of the estimated degree of autocorrelation of its price changes. This could suggest that we should rather rely on the findings of the dual model, whose assumptions about the dynamics of the idiosyncratic component are more realistic. As explained before, its results do not seem to contradict the assumptions of the standard New-Keynesian framework.

On the other hand, we have failed to decisively demonstrate that the estimates of autocorrelation obtained in the partial adjustment setting were driven by our particular specification of the dynamics of the idiosyncratic component. Moreover, the dual setting has problems in its own. The main difficulty with this model is that it is much harder to estimate. As we have seen, the feasible method of its estimation is very inefficient, which results in a very large variance of the estimates of the persistence parameter. Moreover, for a great number of goods, the estimation of the full version of the model cannot be performed. Due to the possibility of emergence of the non-random selection bias, the findings based on the simplified version of the model are less reliable. These factors might perhaps explain why the results from the dual model appear inconsistent

[23] We rule out the possibility that for these products the idiosyncratic innovations are more persistent than the random walk.

[24] Recall that for most non-energy industrial products and services the version of the dual model that takes into account the potentially non-random timing of price changes could not have been estimated.

with those obtained in the partial adjustment setting. Hence, until more efficient methods for estimating the Model MD are developed, the results of none of the two competing frameworks can be given our preference on the empirical grounds.

5 Conclusions.

In the previous sections serial correlation properties of the individual price changes in Belgium were documented. We estimated two models, each of them taking as the point of departure a different potential source of inflation sluggishness, beyond that implied by the existence of nominal price rigidity. In the partial adjustment model we assumed that the fundamental price of every good follows a random walk, as it is commonly done in the literature. We introduced an additional internal propagation mechanism however, by allowing price setters not to adjust completely to the innovations to the fundamental price, a departure from the standard framework devised to reflect the assumptions made in recent works estimating DSGE models. Surprisingly, so modelled autocorrelation of individual price changes appeared negative for many goods, in particular for the foodstuff. This, in line with the partial adjustment model logic, was interpreted as an outcome of price overshooting. We found also some evidence of positive autocorrelation in the case of several industries, especially within the non-energy industrial goods sector, implying a certain degree of gradualism in the price setters' behavior.

The other (dual) model retained the complete adjustment postulate. Instead, it allowed to differentiate between the dynamics of the idiosyncratic shocks to the fundamental price, which were supposed to be non-persistent, and the dynamics of its common component, which was permitted to be more persistent than the random walk. The outcome of the estimation of this model suggests that there is not much (positive or negative) autocorrelation of the innovations to the common factor of the fundamental price. This would be interpreted as evidence against the 'additional source of extrinsic persistence' hypothesis.

The results of the estimation of the dual model are not entirely consistent with these obtained with the partial adjustment one. We argued that the discrepancy, at least for goods characterized by frequent price adjustment, might be a result of an incorrect assumption about the dynamics of the idiosyncratic price component made in the model of gradualism. An attempt to back this hypothesis empirically did not turn up successful however. Hence, so far the causes of the inconsistencies between the models remain unclear.

It should be noted that, due to a small number of price changes for many product categories, the results of the estimation of the dual model are not fully reliable. This concerns in the first turn NEIGs and Services, as for these products the fully specified version of the dual model could not have been estimated. Therefore, for industrial goods and services, the results obtained within the gradualism framework should be given more weight. Taking into account that

evidence of partial adjustment was found in many NEIG sectors, this implies that the hypothesis that there are additional sources of (possibly intrinsic) persistence in certain industries cannot be eventually ruled out.

Even though the results from the estimation of the competing models are mixed, there are some broad implications that could be drawn from our analysis. First, while our findings from the gradualism framework seem to cast into question the main assumptions underlying the baseline New-Keynesian model, the results of estimation of the dual model are broadly in line with the random walk and perfect price adjustment hypotheses. Hence, at this stage we must conclude that the data on Belgian micro-prices do not provide decisive evidence against the basic New-Keynesian paradigm. Second, the results of estimation of the model of gradualism suggest that real inflation persistence may be higher or lower as compared to the baseline, depending on industry. Thus, there is no convincing evidence for an additional, economy-wide source of sluggishness, be it intrinsic or extrinsic. In this light, the attempts to enrich the standard one-sector New-Keynesian model with a source of inflation sluggishness different from the nominal rigidity (like in Christiano and Fisher, 2004; Christiano et al, 2001; Gali and Gertler, 1999; Mankiw and Reis, 2002; Smets and Wouters, 2003 and more) appear empirically unfounded. Finally, as said earlier, we could not rule out the possibility that prices of goods belonging to some product categories, in particular those of non-energy industrial products, do exhibit some additional persistence. This implies that the correct inflation modelling strategy might be to consider a number of economic sectors characterized by a different degree of intrinsic or extrinsic price sluggishness. It is entirely possible that a model containing a small sector of products whose fundamental price is more persistent than the random walk will produce enough inertia in the behavior of inflation to be able to explain its empirical dynamic properties. A piece of evidence supporting this view was provided by Altissimo et al (2004) who showed that a high persistence of a limited number of sector inflation indices may lead to a high degree of sluggishness of the aggregate index, even if the majority of sector indices is not very persistent.

There are several ways in which our results should be strengthened in the future. In particular, strengthening the econometric methods applied in the paper would be desirable. The task of a correct estimation of the speeds of price adjustment is challenging from the technical viewpoint. While standard panel data estimation methods can be used if there is no heterogeneity among agents within single product categories, nothing ensures that it is indeed the case. If the price setters differ as to the degree of gradualism, there are not many estimation methods available. Those that could be used, like the mean group estimator method of Pesaran and Smith (1995) demand relatively many observations per separate trajectory. Thus, they may produce misleading results when applied to sticky price products.

Similarly, we encountered many econometric problems with the dual model. Due to the computational constraints, we were unable to estimate it by the (efficient) maximum likelihood method. Even the two step method devised for the estimation of this model could not have been used for all product categories. Again, more work is needed to deal with these problems.

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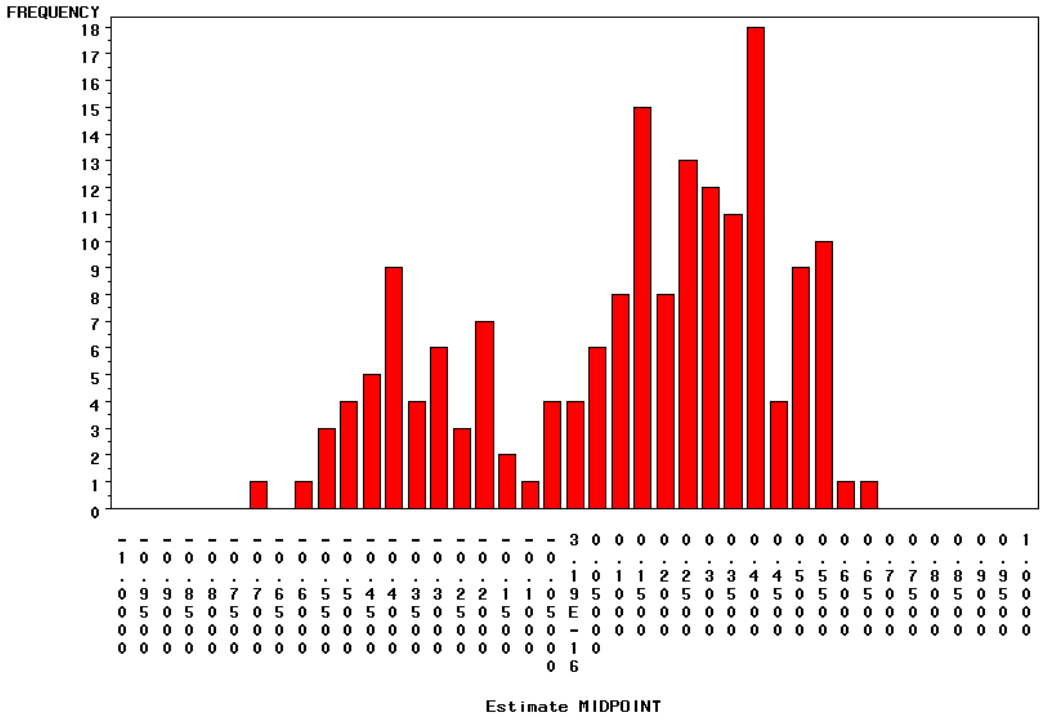
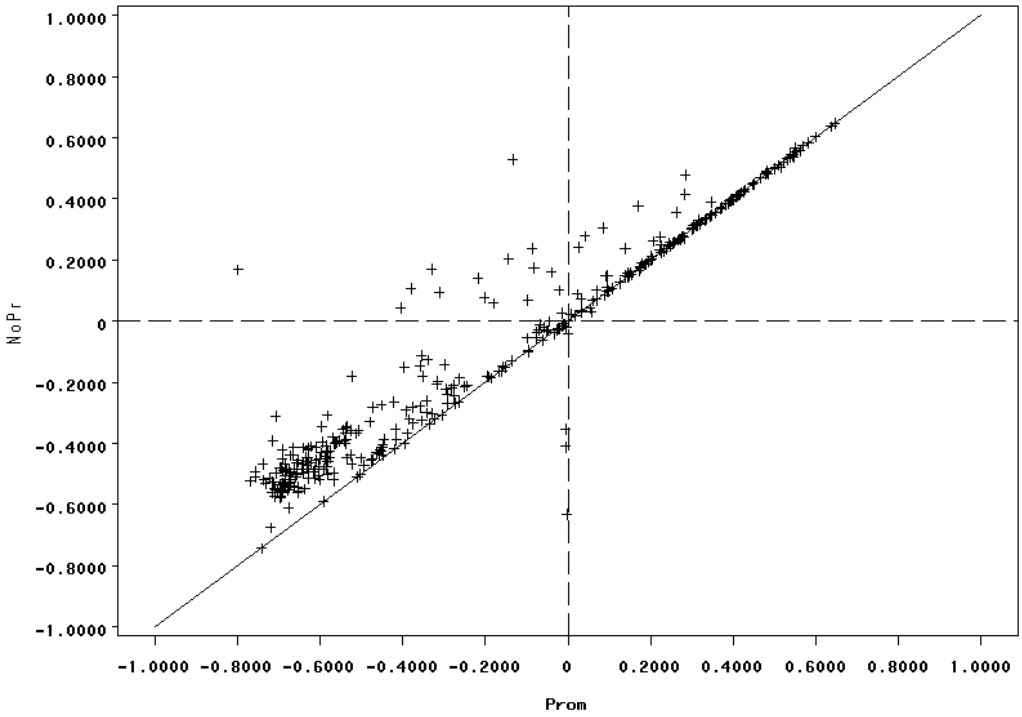
Figure 1.



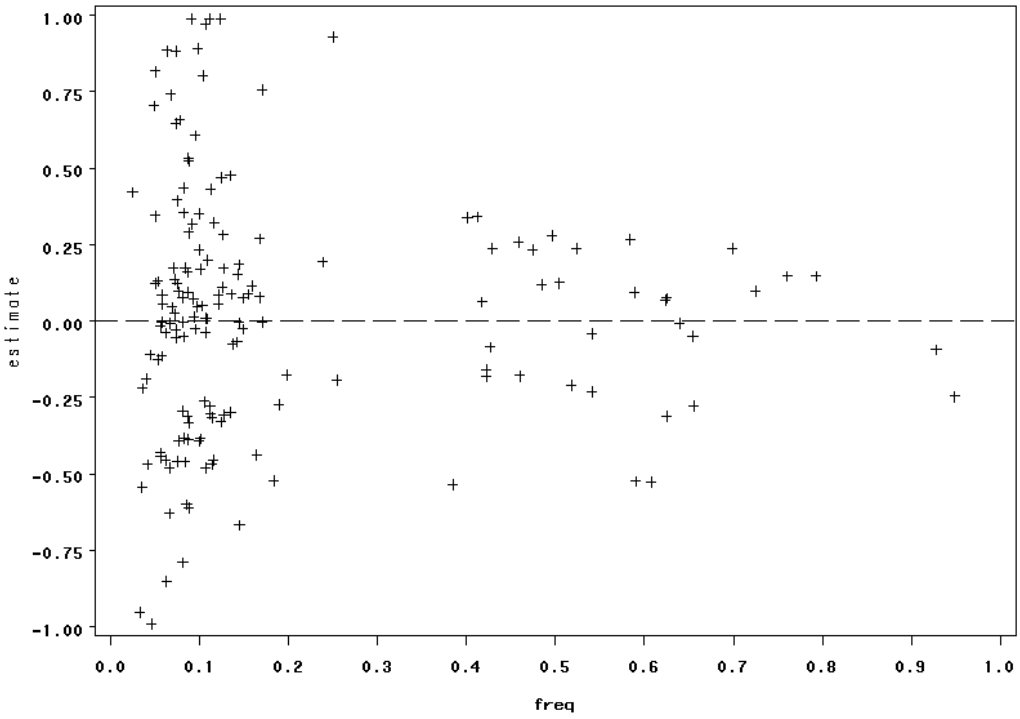
The histogram of the estimates of the gradualism parameter γ for the NEIG category.

Figure 2.



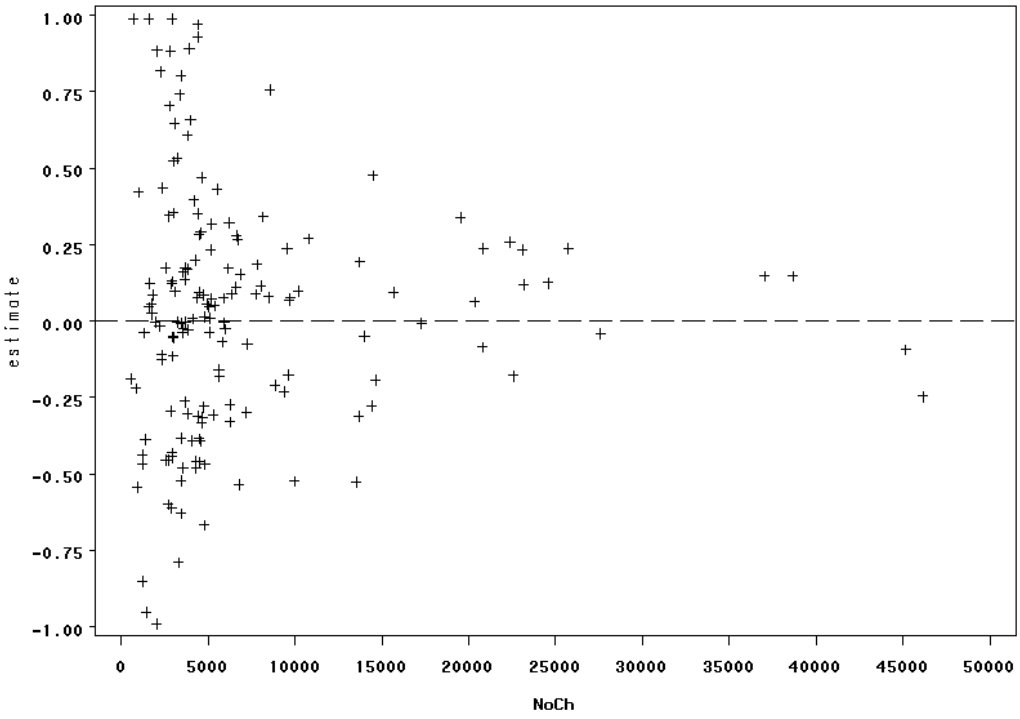
Scatter plot drawing the estimates of the coefficient of gradualism γ based on the data with the promotions removed against the estimates based on the full sample of prices.

Figure 3a.



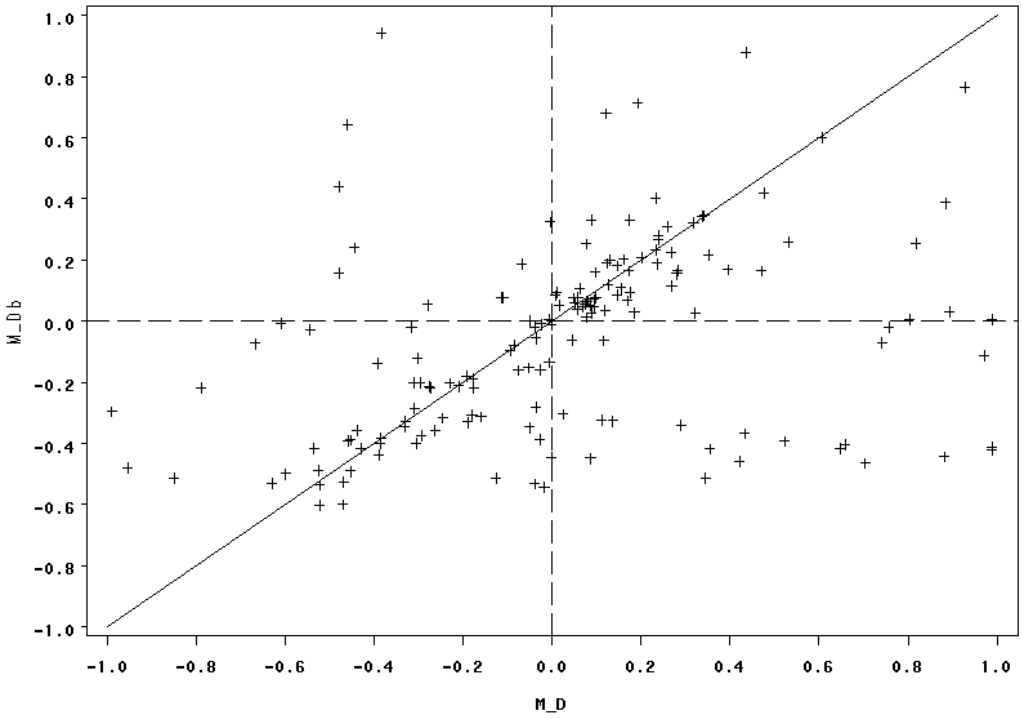
The dual model estimates of ρ as a function of the frequency of price adjustment.

Figure 3b.



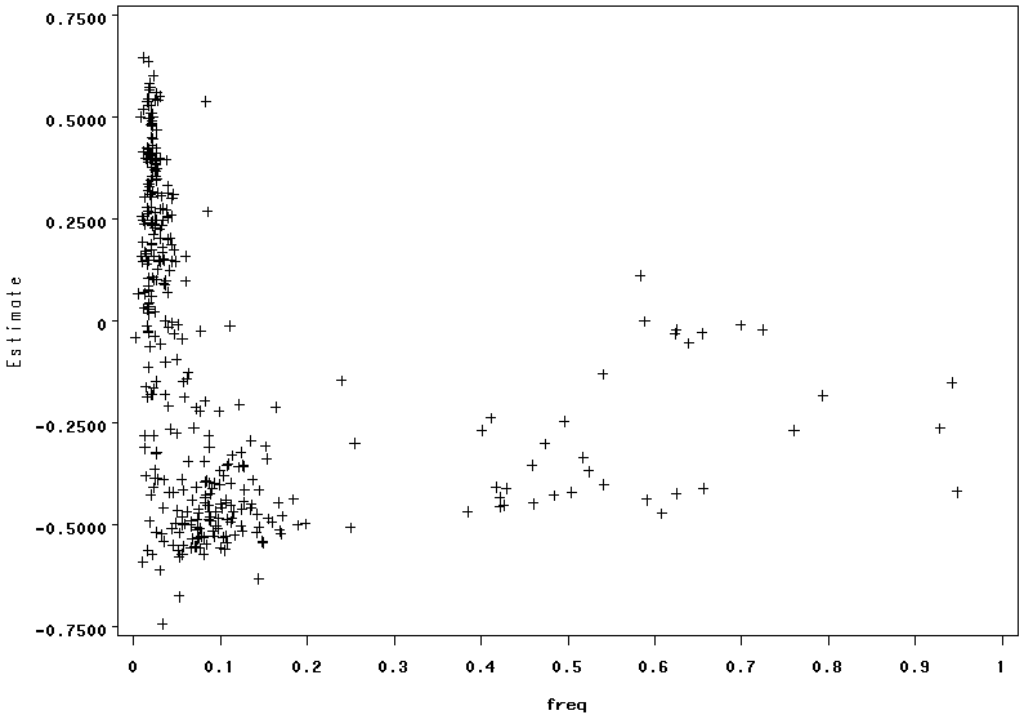
The dual model estimates of ρ as a function of the number of price adjustments of the given product in the sample.

Figure 4.



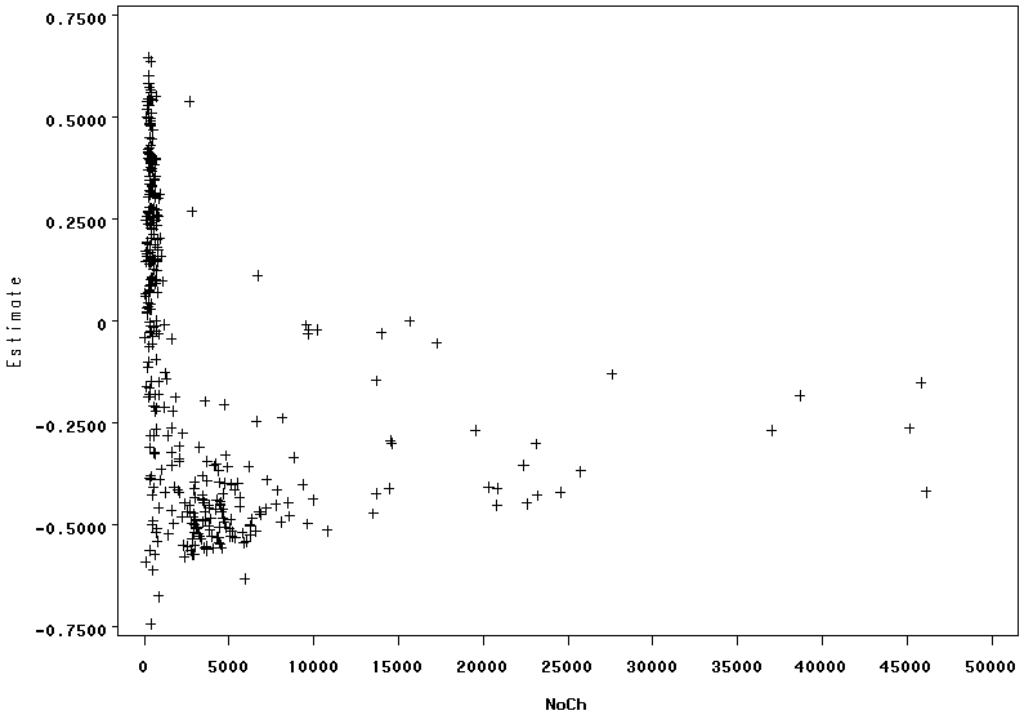
Scatter plot drawing the model MDb estimates of ρ against the estimates of ρ from the model MD.

Figure 5a.



The estimates of the coefficient of gradualism γ as a function of the frequency of price adjustment.

Figure 5b.



The estimates of the coefficient of gradualism γ as a function of the number of price adjustments of the given product in the sample.

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