MICROECONOMIC PRICE ADJUSTMENTS AND INFLATION: EVIDENCE FROM SPANISH SECTORAL DATA

Ángel Estrada and Ignacio Hernando

Banco de España

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Ángel Estrada and Ignacio Hernando

Banco de España, Research Department

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Abstract

The purpose of this paper is to illustrate the implications for aggregate price dynamics of alternative characterizations of microeconomic price adjustment policies. Within the hazard adjustment framework developed in Caballero and Engel (1993a), we present alternative models of individual price adjustment that are consistent with the predictions of the menu-cost models. We estimate these models using a data set of Spanish sectoral price indexes covering the period 1978-1998, and compare their ability to explain the behavior of the aggregate price level, taking the partial adjustment model as a benchmark. The main findings are the following: 1) models with non-constant hazard functions slightly outperform the partial adjustment model; 2) there is weak evidence regarding the existence of differences between the magnitudes of upward and downward price rigidities; and 3) the impact on the aggregate inflation rate stemming from the existence of nominal rigidities in several subsets of sectors is non-negligible.

Keywords: price adjustments, inflation, downward price rigidity.

JEL Classification: E31, D49.

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1- Introduction

The existence of nominal price rigidities is a standard assumption in macroeconomic modeling that allows for nominal shocks to influence real variables. In its turn, formal models of individual price adjustment are essential to understand aggregate price stickiness. An adequate formalization of the dynamic behavior of the aggregate price level would require both taking into account that firms are heterogeneous and adopting a realistic characterization of the price adjustment policies at the firm level. However, the standard framework to obtain a sluggish behavior of the aggregate price level has been to consider a representative agent model with quadratic adjustment costs of changing prices. In this setting, individual firms adjust their prices continuously and in a constant proportion. The lack of realism of this conduct has led to alternative characterizations of the aggregate price dynamics stemming from a more accepted microeconomic behavior. Time-dependent adjustment models, in which firms revise their prices at prespecified dates, represented a first step in this direction.

Models considering state-dependent pricing rules provide an alternative approach. Within this type of models, the adjustment hazard framework proposed by Caballero and Engel (1993a) tries to combine a more realistic characterization of the microeconomic behavior with more explanatory power at the aggregate level. In this framework, the adjustment policy at the firm level is characterized by means of the adjustment hazard function, which links the probability of a firm adjusting its price to the size of the deviation from its target price. Considering that firms are heterogeneous, the aggregate dynamics generated by non-constant hazard functions will depend on the cross-sectional distribution of price deviations.

The adjustment hazard framework is quite general and it encompasses the partial adjustment model or the menu-cost models as particular cases. Interestingly enough, the aggregate price dynamics arising from a wide range of hazard functions -i.e. those hazard functions that are increasing with respect to the absolute value of the price deviations- is consistent with the macroeconomic implications of the menu-cost models as those developed in Ball and Mankiw (1994,1995). Moreover, this approach is useful not only for testing reasonable conjectures about individual adjustment policies (for instance, the existence of downward price rigidity) but also for comparing alternative hazard functions regarding their ability to explain the evolution of the aggregate price level.

In this paper we try to illustrate the implications for aggregate price dynamics of alternative characterizations of microeconomic price adjustment policies, within the hazard adjustment framework developed by Caballero and Engel. For this purpose, we estimate alternative models of individual price adjustment, consistent with the predictions of the menu-cost models, using a data set of Spanish sectoral price indexes covering the period

1978-1998. A key element to estimate these adjustment hazard models is the sequence of cross-sectional distributions of price deviations. In this paper, we estimate them using two alternative procedures. First, we proxy the cross-sectional distribution of individual price deviations by the cross-sectional distribution of sectoral price deviations. Second, following Caballero and Engel (1993b), starting from an arbitrary cross-section density we simulate the whole sequence of price deviations using exclusively aggregate data. In both cases, we compare the ability of the alternative hazard adjustment models to explain the behavior of the aggregate price level, taking the partial adjustment model as a benchmark. Finally, we present simple simulation exercises in order to provide an assessment of the impact on the aggregate price level stemming from the existence of nominal rigidities in several subsets of sectors. The main results of the paper may be summarized as follows. Models with non-constant hazard functions slightly outperform the partial adjustment model, i.e. the aggregate behavior of prices depends not only on the past average deviation from target prices (as it is assumed when using a partial adjustment model) but also on higher moments of the cross-section distribution of price deviations. Besides, there is weak evidence regarding the existence of differences in the size of upward and downward price rigidities. Finally, the impact on the aggregate price level stemming from the existence of nominal rigidities in several subsets of sectors is non-negligible.

The rest of the paper is organized as follows. Section 2-a provides an informal review of the microfoundations of alternative theories explaining aggregate price dynamics. Section 2-b introduces the adjustment hazard framework and characterizes the implications of alternative hazard functions for the evolution of the aggregate price level. Section 3-a specifies the model for sectoral frictionless prices. Section 3-b describes the estimation methodology. Section 4 presents an application of the adjustment hazard framework for Spanish sectoral price data and section 5 provides a brief summary of the major findings.

2- Individual adjustment policies and aggregate price dynamics

2-a) Microfoundations for aggregate price stickiness

The fact that many individual prices are set in nominal terms and are not always readjusted when there is a change in the economic environment is far from controversial. These characteristics of the process of individual price setting are enough to justify the existence of aggregate nominal rigidities. However, the assumptions about price stickiness implicit in traditional Keynesian models have been criticized on the grounds of their inconsistency with any reasonable model of microeconomic behavior. Much recent theoretical research has been devoted to analyze the microeconomic foundations of the sluggish adjustment of prices and, as a consequence, the characterization of the aggregate price dynamics has been enriched[1]. This sub-section provides an informal review of alternative microeconomic adjustment policies and their implications for the behavior of the aggregate price level.

Standard models connecting the microeconomic behavior of individual prices with aggregate price dynamics include both partial adjustment models and time-dependent price-setting models. On the one hand, partial adjustment models are based on the introduction of quadratic adjustment costs of changing prices, this assumption leading to a continuous but partial adjustment at the firm level. Assuming identical structural parameters across firms, a partial adjustment equation is obtained for the aggregate price level. There is a clear lack of realism in the microeconomic behavior implicit in this framework given that firms do not adjust their prices continuously to respond to the shocks they suffer. An alternative interpretation of the partial adjustment model would suggest that a constant share of firms adjust fully (instead of all firms adjusting partially). As Rotemberg (1987) shows, both interpretations are observationally equivalent if the probability of adjusting is independent of the deviation between firms' prices and their corresponding targets. However, this assumption does not hold as long as the likelihood of a price change depends on the magnitude of the deviation between actual and target prices.

On the other hand, time-dependent adjustment models have relied on the assumption that each firm leaves its price unchanged for a fixed amount of time[2]. Considering time-dependent pricing rules provide a simple framework to derive the aggregate dynamics. Although these models allow for the realistic fact of nominal price changes occurring in discrete jumps, they share the unsatisfactory assumption that the time between price changes is preset, and firms are unable to adjust to any shock between price-adjustment dates.

An alternative approach to introduce price stickiness is provided by state-dependent adjustment models[3]. In these models a firm will change its price whenever it deviates a certain amount from its optimal value. The most obvious justification for this individual behavior is the existence of a fixed cost in the adjustment-cost function. This is the idea in which the menu-cost model of price adjustment is based. According to this model, a firm will change its actual price only if the desired adjustment is large enough to warrant paying the menu cost. The menu-cost model belongs to the class of non-convex adjustment cost

[1] See chapter 6 in Romer (1996) for a review of theoretical microfoundations of nominal rigidities and Sheshinski and Weiss (1993) for a collection of papers dealing with the interaction of individual price adjustment policies and aggregate price stickiness.

2 See, for instance, Taylor (1980).

[3] See, for instance, Sheshinski and Weiss (1977,1983) and Caplin and Leahy (1991).

models with a fixed (S,s) pricing rule[4]. Some degree of flexibility may be introduced in these models by allowing different thresholds for positive and negative deviations. This setting would be justified, as Ball and Mankiw (1994) suggest, by the presence of trend inflation. When there is trend inflation, prices will be more flexible upwards than downwards, i.e. positive shocks are more likely to trigger price adjustments than are negative shocks. The reason is that, in the face of a negative shock, the firm is more likely to avoid the menu cost expecting that some of the desired adjustment will come about as a result of trend inflation.

While menu-costs models provide realistic microfoundations for price dynamics at the firm level, they do not seem to be adequate to explain aggregate price dynamics since discontinuous adjustment is a feature of individual firms, not of the whole economy. The adjustment hazard models, developed by Caballero and Engel (1993a), try to address this problem. These authors introduce the adjustment hazard functions that link the probability of an individual adjustment to the size of the deviation between target and actual prices. This approach provides both a reasonable microeconomic behavior and, by allowing heterogeneity across individual firms, a realistic characterization of aggregate price dynamics. Furthermore, the partial adjustment model and the menu cost model may be considered as particular cases of the family of adjustment hazard models[5]. Thus, a direct comparison of these models to alternative hazard models is possible. For instance, Caballero and Engel (1993b) estimate an increasing hazard model for price adjustment that outperforms the partial adjustment model in explaining the dynamics of the aggregate manufacturing price level in the U.S.

2-b) The adjustment hazard model

In this sub-section, we briefly describe the basic framework underlying the adjustment hazard approach developed by Caballero and Engel (1992 and 1993a). This approach relies on two basic features: first, the probability of an individual price change depends on the departure of the actual price (P_{it}) from what it would be optimal ($\tilde{P}_{it}$), and, second, firms are heterogeneous. The main macroeconomic implication of this framework is that to explain aggregate price dynamics is necessary to keep track of the evolution of the cross-sectional distribution of individual price deviations.

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[4] A fixed (S,s) pricing rule indicates that the deviation of the actual price from its target floats freely between two predetermined bounds and the adjustment is triggered only if one of the bounds is reached. See Caballero and Engel (1991) for a study of the implications of individual (S,s) policies for aggregate dynamics.

[5] More precisely, in the partial adjustment model the hazard function is constant, i.e. the probability of adjustment is independent of the magnitude of the deviation. In the simplest menu-cost model the hazard function will take value one if the deviation (in absolute value) is larger than a specific threshold, and zero otherwise.

Let z_{it} denote the difference between actual and target price for firm i at time t and $F_t(z)$ the cross-section distribution of individual price deviations. The key element of this approach is the hazard function, $\Lambda(z)$, which links the probability of a firm adjusting its price to the size of the deviation from its target price. The transitional dynamics of individual price deviations will be driven by three elements: first, aggregate shocks shifting the whole distribution of price deviations; second, the hazard function that determines the price adjustment; and, third, idiosyncratic shocks affecting the determinants of individual price targets. Thus, the change in the aggregate price level during the time interval [t,t+1) would be equal to:

$$\Delta p_{t+1} = \int_{-\infty}^{\infty} (\Delta \tilde{p}_{t+1} - z)\Lambda(z - \Delta \tilde{p}_{t+1})\, f_t(z)dz \qquad [1]$$

where $\Delta \tilde{p}_{t+1}$ represents the aggregate shock (lower case letters meaning natural logarithms), and $f_t(z)$ is the cross-section density of price deviations at time t.

The intuition for this expression is the following. Let us assume that $f_t(z)$ is the cross-section density function of price deviations once firms have experienced their idiosyncratic shocks. After the realization of the aggregate shock ($\Delta \tilde{p}_{t+1}$), every individual price deviation z is shifted to ($z - \Delta \tilde{p}_{t+1}$). Then, a fraction $\Lambda(z - \Delta \tilde{p}_{t+1})$ of firms adjust their prices by ($z - \Delta \tilde{p}_{t+1}$). Expression [1] is obtained by adding over the whole distribution of z.

Looking at expression [1] it seems to be clear that, in this framework, the evolution of the aggregate price level depends on two elements: the cross-section distribution of price deviations and the microeconomic hazard function. Next, we try to illustrate how the shape of the adjustment hazard function conditions the aggregate price dynamics. For this purpose, first we present, as a benchmark case, the adjustment when the hazard function is constant, and then we derive the aggregate dynamics for alternative hazard functions that are consistent with a more realistic microeconomic behavior.

2-b)-a) Constant Hazard: The Partial-Adjustment Model

Let us start with the simple case where $\Lambda(z) = \lambda_0$. Replacing this expression for the hazard function in equation [1] yields:

$$\Delta p_{t+1} = l_0 (\Delta \tilde{p}_{t+1} - z_t^{(1)}) \qquad [2]$$

where $z^{(1)}$ represents the mean of the cross-sectional distribution of deviations. This equation shows that the change in the aggregate price level in every period is a constant share (λ_0) of the average price deviation[6]. This aggregate pattern is the result of a constant fraction of firms (λ_0) adjusting fully in every period. As Rotemberg (1987) showed, this aggregate dynamics is equivalent to that derived from a representative-agent model with quadratic adjustment costs (partial-adjustment model)[7]. It is worth noting that in this case the aggregate price dynamics depends only on the first moment of the cross-section distribution of firms.

2-b)-b) Alternative Hazards

In this section we introduce alternative hazard functions to relax the constant hazard assumption that the probability of an individual price adjustment is independent of the deviation between the actual and the target price. With these alternative hazards we look for a more realistic microeconomic behavior[8]. More precisely, we seek to allow for the probability of adjustment being: (i) increasing in the size of the price deviation and (ii) different for upwards and downwards price adjustments.

Let us consider the following non-constant hazard functions:

I. Simple Quadratic-Hazard model

First, as an example of an increasing hazard function, we consider the simple quadratic-hazard function:

$$\Lambda(z) = l_0 + l_2 z^2 \qquad [3]$$

Substituting the hazard function [3] in equation [1] yields the following expression for the aggregate price dynamics:

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[6] Note that $(\Delta\tilde{p}_{t+1} - z_t^{(1)})$ is the average price deviation after the realisation of the aggregate price shock $\Delta\tilde{p}_{t+1}$.

[7] Given that $z_t^{(1)} = p_t - \tilde{p}_t$, the level of observed prices would be:, $p_{t+1} = (1 - l_0)p_t + l_0\tilde{p}_{t+1}$, so equation [2] may be rewritten as: $\Delta p_{t+1} = (1 - l_0)\Delta p_t + l_0\Delta\tilde{p}_{t+1}$, that is the usual representation of the partial adjustment model.

[8] Nevertheless, we are still assuming that once a firm adjusts its price, it adjusts fully, i.e. the size of the adjustment is equal to the price deviation. See Ariga and Okhusa (1998) for a joint estimation of the probability of zero price change and the size of price changes, both conditional on the magnitude of the deviation of the price from the target level.

$$\Delta p_{t+1} = l_0 \Delta \tilde{p}^{pa}_{t+1} + l_2 \left\{ \left(\Delta \tilde{p}^{pa}_{t+1} \right)^3 + 3\Delta \tilde{p}^{pa}_{t+1} z^{(2)}_{c,t} - z^{(3)}_{c,t} \right\} \qquad [4]$$

where $\Delta \tilde{p}^{pa}_{t+1} = (\Delta \tilde{p}^{pa}_{t+1} - z^{(1)}_t)$ represents the partial adjustment term in equation [2], and $z^{(k)}_c$ denotes the kth centered moment. Unlike the constant hazard case, the mean is not the only moment of the cross-sectional distribution relevant for the aggregate dynamics. In this sense, two interesting features arise from expression [4]. First, the response of the aggregate price level to aggregate shocks is positively linked to the variance of the cross-sectional distribution. This prediction is consistent with Ball and Mankiw (1994). These authors show that more variability in the relative-price distribution increases the effects of asymmetric price adjustment -implied by the presence of trend inflation- on the aggregate price level. Second, for a given aggregate shock, increases in the aggregate price level are larger the more skewed to the left the distribution of deviations is. Again, this feature is consistent with Ball and Mankiw (1995) who show that inflation rises when the distribution of relative price changes is skewed to the side of relative price increases[9].

II. Asymmetric Hazard models

Next, we explicitly model the possible existence of asymmetries in the likelihood of price adjustments, that is, we allow the probability of adjustment being different for positive and negative price deviations as suggested by recent menu cost models. More precisely we try to capture the aforementioned Ball and Mankiw's (1994) prediction that, in the presence of trend inflation, the probability of adjusting prices is higher when the firm´s deviation is negative than when it is positive. In other terms, in the presence of trend inflation firms are more likely to increase their prices than to cut them[10]. To test this possibility we consider two alternative asymmetric-hazard functions, which allow for two different origins of the asymmetry.

a) Asymmetric-Constant-Hazard model

$$\Lambda(z) = \begin{cases} l_0^+, \text{ for } z > 0 \\ l_0^-, \text{ for } z < 0 \end{cases} \qquad [5]$$

[9] Note that in the comparisons between the implications of Caballero and Engel's hazard function approach and Ball and Mankiw's menu-cost models, we are identifying the relative price shocks of the later approach with the unadjusted portion of individual shocks (i.e., the price deviations) of the former.

[10] See Amano and Macklem (1997) for evidence, using Canadian data, supporting Ball and Mankiw's model of price-adjustment. See also, Yates (1998) for evidence contrary to the existence of downward nominal rigidity in price adjustment.

The aggregate price dynamics implied by the hazard function [5] is given by:

$$\Delta p_{t+1} = l_0^- F_t(\Delta \tilde{p}_{t+1})\left\{\Delta \tilde{p}_{t+1} - z_t^{(1-)}\right\} + l_0^+\left(1 - F_t(\Delta \tilde{p}_{t+1})\right)\left\{\Delta \tilde{p}_{t+1} - z_t^{(1+)}\right\} \qquad [6]$$

where $z^{(1-)}$ and $z^{(1+)}$ denote the means of the cross-sectional distribution conditional on z being smaller and larger than $\Delta \tilde{p}_{t+1}$.

b) Asymmetric-Quadratic-Hazard model

$$\Lambda(z) = l_0 + \begin{cases} l_2^+ z^2, & \text{for } z > 0 \\ l_2^- z^2, & \text{for } z < 0 \end{cases} \qquad [7]$$

Substituting the hazard function [7] in equation [1] yields the following expression for the aggregate price dynamics:

$$\begin{aligned} \Delta p_{t+1} = {} & l_0 \Delta \tilde{p}^{pa}_{t+1} + l_2^- F_t\left(\Delta \tilde{p}_{t+1}\right)\left\{\left(\Delta \tilde{p}^{pa-}_{t+1}\right)^3 + 3\Delta \tilde{p}^{pa-}_{t+1} z_{c,t}^{(2-)} - z_{c,t}^{(3-)}\right\} + \\ & l_2^-\left(1 - F_t\left(\Delta \tilde{p}_{t+1}\right)\right)\left\{\left(\Delta \tilde{p}^{pa+}_{t+1}\right)^3 + 3\Delta \tilde{p}^{pa+}_{t+1} z_{c,t}^{(2+)} - z_{c,t}^{(3+)}\right\} \end{aligned} \qquad [8]$$

where $\Delta \tilde{p}^{pa-}_{t+1} = (\Delta \tilde{p}_{t+1} - z_t^{(1-)})$ and $\Delta \tilde{p}^{pa+}_{t+1} = (\Delta \tilde{p}_{t+1} - z_t^{(1+)})$. $z_c^{(k-)}$ and $z_c^{(k+)}$ denote the kth centered moments of the cross-sectional distribution conditional on z being smaller and larger than $\Delta \tilde{p}_{t+1}$.

3- Empirical framework

Last section sets clearly the two pieces of information we need to implement the general approach to model the aggregate behavior of prices: first, the profile for aggregate target price (or aggregate shock, $\Delta \tilde{p}_{t+1}$); second, the individual deviations of observed prices from target prices (z_{it}), from which we can calculate the different moments of the distribution ($F_t(z)$)[11].

In this paper, we use two alternative procedures to estimate the alternative hazard models. Our first procedure exploits the information contained in a sectoral database. As a first step we estimate sectoral target prices. For this purpose, we first characterize how firms determine the frictionless prices (that is, the prices the firm would set if there were no rigidities) for the products they sell in the markets making use of the cointegration condition between frictionless and observed prices. Then we calculate the optimal price (having in mind that if the firms are not going to adjust every period their prices they need to incorporate their perspectives for the future evolution of the frictionless price determinants) and its deviations from the observed price. In a third stage, we define the aggregate shock as a weighted average of the changes in sectoral optimal prices. Finally, we take the path of the moments of the cross-sectional distribution of sectoral price deviations after the realization of the aggregate shock as a proxy for the different moments required to estimate the aggregate price dynamics equations corresponding to the models described in section 2.

As an alternative strategy, we follow Caballero and Engel (1993b). Their approach avoids the use of individual information about prices and its determinants. Using this second procedure, we use the same aggregate shock but we assume a specific distribution for the individual shocks. Thus, we only need an initial distribution of deviations from target prices to run the whole model[12]. We consider as initial distribution the average cross-sectional distribution of sectoral price deviations.

3-a) A model for frictionless prices and optimal price determination

The model we use to explain price formation at a sectoral level follows very closely a model proposed by Layard et al. (1991). Firms have market-power in the product market and they try to maximize their revenue. For this purpose, they decide optimally the selling price of their products, the employment and the intermediate inputs they use. Firms take wages, prices of the competing firms (including the price of imported goods), prices of the intermediate inputs, the sectoral level of demand and the capital stock as given.

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[11] It should be stressed that, if the hazard function is a constant, we only need to know the mean of this distribution to implement the model. As the mean is a linear combination of the individual z, we do not need individual information to calculate it.

[12] Caballero and Engel (1993b) considered as the initial density the ergodic density that would exist if aggregate shocks followed a random walk with drift equal to that of the aggregate shock process and (instantaneous) variance equal to the sum of the idiosyncratic variance and the variance of the series of aggregate shocks.

The representative firm (i) of a sector (s) has a technology represented by a Cobb-Douglass production function with constant returns to scale. This technology is the same for all the firms in the sector. The production is made with two factors, capital (K) -that we consider fixed-, and a variable factor (FV) that includes both labor and intermediate inputs. Thus, the output (Y) would be given by:

$$Y_{is} = A_{is} FV_{is}^{(1-b_s)} \qquad [9]$$

where

$$A_{is} = A_s^{'} K_{is}^{b_s} \qquad [10]$$

being A's the technological progress of the sector.

We assume that firms' demand functions depend positively on the prices of the other firms in the sector (P_j) -including the prices of the external goods that are substitutes for the domestic production- and negatively on the price charged by the own firm (P_i) in an isoelastic way:

$$Y_{is} = \left[\frac{P_{is}}{P_{js}}\right]^{-h_s} YD_s \qquad [11]$$

where YD_s is an indicator of the sectoral demand level.

In this setting, the problem faced by the representative firm can be stated as follows:

$$\begin{aligned} & Max \qquad \Pi_{is} = P_{is} Y_{is} - CV_s FV_{is} \\ & s.t. \\ & Y_{is} = A_{is} FV_{is}^{(1-b_s)} \\ & Y_{is} = \left[\frac{P_{is}}{P_{js}}\right]^{-h_s} YD_s \end{aligned} \qquad [12]$$

where CV represents the price of variable inputs, being equal for every firm of the sector. Assuming the existence of a symmetrical equilibrium in the sector (P_{is}=P_{js}), the expression for sectoral prices (in logs) can be written as follows:

$$p_s = cnst. + Ln\left(\frac{h_s}{h_s - 1}\right) + cvu_s \qquad [13]$$

with cvu representing the unit variable costs.

This equation implies that prices are a mark-up over marginal costs, which we proxy by the unit variable costs. In this model the mark-up is constant, but a growing theoretical and empirical body of research indicates that demand fluctuations can influence the mark-up. For example, Rotemberg and Woodford (1991) showed that the organization of the product market and the individual preferences had an impact on the size and sign of that coefficient. Besides, the openness degree of some sectors we are considering is very high, and it has been suggested that an increase in international competitiveness makes the domestic demand less elastic: when external prices are high, domestic firms can raise their prices and expand their mark-ups. We specify the following equation for the mark-up:

$$Ln\left(\frac{h_s}{h_s - 1}\right) = cnst. + g_{s1}(yd_s - y_s) + g_{s2}(p_s^x + e - p_s) \qquad [14]$$

where p^x_s is the external price in foreign currency and e the nominal exchange rate. Substituting this equation in [13], yields the following expression:

$$p_s = cnst. + m_s cvu_s + (1 - m_s)(p_s^x + e) + t_s(yd_s - y_s) \qquad [15]$$

Thus, sectoral prices are a weighted average of variable costs and external prices in domestic currency plus an additional term representing demand pressure in the sector. The relative weights of the first two components depend on the sectoral openness degree, and they have to add one. This constraint, called nominal homogeneity, is imposed in the empirical model. If the sector has no external competition we drop the external price component and we impose a coefficient of one for the unit variable costs. On the contrary, we have no a priori believe about the size and sign of τ, the parameter for demand pressure, so we will let the data to determine this question.

Now we can use the expression [15] to construct the frictionless price (p^*_{it}) for each sector. If firms are forward-looking, this price is not going to be the optimal price ($\tilde{p}_{it}$), because there is a non-null probability of not adjusting prices in the following periods. Thus, firms will also take into account the expectations of future developments in the frictionless price. Assuming that the firm knows the probability of adjusting its price and try to minimize the flow of deviations between the optimal and the frictionless prices Caballero and Engel (1993b) showed that the relation between this two prices is the following:

$$\tilde{p}_{it} = p^*_{it} + \sum_{m=t}^{\infty}\left(\frac{1-g}{1+b}\right)^{m+1-t} E_t(\Delta p^*_{im+1}) \qquad [16]$$

where γ is the probability of adjusting prices and β is the discount rate. This expression establishes that the frictionless price should be increased with the discounted flow (corrected by the probability of non-adjustment) of it expected increases to obtain the optimal price. Just aggregating through individual firms we get a similar result at the sectoral level.

3-b) Data and estimation procedures

Our database includes prices, unit variable costs, import prices and an indicator of the demand pressure for forty different categories of goods and services[13]. These prices correspond to wholesale prices of agricultural and industrial firms in domestic (IPPA and IPRI) and foreign markets (IVUX), retail prices for consumer goods (CPI-G) and consumer prices for services (CPI-S), all of them corrected from changes in indirect taxation. Thus, the aggregate price we are considering is slightly different from the price indexes we usually look at. Chart 1 compares the annual percentage changes of our aggregate price indicator, the CPI (that is, the aggregation of CPI-G and CPI-S plus indirect taxation) and the IPRI. Not surprisingly, our indicator lies, for most of the sample period, between the other two. Besides, the profiles are very similar and main differences appear in 1986, when the indirect taxation system was changed, and between 1989 and 1993, when the inflation differential in goods and services was the highest ever recorded.

Unit variable costs are an average of unit labor and unit intermediate costs, weights coming from the Input-Output tables for the Spanish economy. Unit intermediate costs include prices for all the intermediate goods and services each sector buys from other sectors and from the external markets. A special case is the intermediate costs for the different categories of consumer goods, because its main components are the domestic and imported wholesale prices for each category. The sectoral import prices are our proxy for external prices in domestic currency[14], and the indicator for demand pressure is obtained averaging intermediate and final demand for each particular sector, using, again, the Input-Output tables to calculate the weights.

Using this database, we implement a three-step procedure to estimate the different hazard models proposed in section 2. In the first stage, we estimate the frictionless price (p^*_{st}) for the forty goods and services we are considering. In order to do that, we exploit the cointegration condition that should be kept (till a constant) between observed and frictionless prices. As frictionless prices are given by the model described in last sub-

[13] See Appendix 1 for details on the sectoral classification.

[14] We assume external production and domestic production to be homogenous.

section, we run forty regressions in levels of observed prices on unit variable costs, external prices (if the sector is open to foreign markets) and demand pressure[15], and we define p_s^* as the fit of each particular regression. The detailed OLS estimates[16] of equation [15] could be found in table A.1 (Appendix 1). In general, most of the sectors satisfy the cointegration condition, but some of them require artificial variables to obtain stationary residuals. As expected, the influence of external prices is larger for export than for domestic prices. The impact of pressure demand is almost always positive for prices of goods (more than 50% of the cases), but for prices of services we have three sectors with pro-cyclical margins and one with counter-cyclical margins.

Defining the aggregate observed and frictionless prices (p, p*) as the weighted averages of the sectoral observed and frictionless prices (p_s, p_s*) respectively, we can compare the realized inflation with the inflation rate we would observe if firms did not face adjustment costs of changing prices. For that purpose, chart 2 compares the annual percentage changes of p and p*. As we would expect, the correlation coefficient between both series is very high (0.97) -in a long-term view, every shock should be incorporated in prices- and the series of frictionless price changes displays a slightly higher volatility[17] -in the short-term, nominal rigidities help to smooth the impact of shocks-. It is worth noting the existence of significant differences between both series in some periods. For instance, in the third quarter of 1984 the observed inflation rate was 2.2 % higher than the "frictionless" inflation rate, whereas in the second quarter of 1997 the observed rate was almost 2 % lower than the "frictionless" rate. Chart 2 suggests that nominal rigidities may have significant implications in terms of aggregate inflation, at least for some periods.

In the second step we construct the optimal sectoral price ($\tilde{p}_s$) using expression [16]. The sectoral expectations for the increases in the frictionless prices are derived from bivariate VARs of observed and frictionless prices for each sector. The discount parameter is equal across sectors (β=0.025) and the probability of adjustment (γ) is taken from a estimated partial adjustment model between observed and frictionless prices for each sector. Now we can calculate the sectoral deviations from target prices (z_s) as the difference between observed and optimal prices, so we can derive the cross-sectional distribution of z for every quarter and the aggregate shock ($\Delta \tilde{p}_{t+1}$) as the weighted average of sectoral optimal prices. Chart 3 displays the kernel density for the cross-sectional distributions of

[15] These regressions are run separately because the coefficients could depend on the sector; in the estimation we only impose that the parameters for unit variable cost and external price add to one.

[16] We tried more sophisticated methods as Fully-Modified Estimators, but the results were broadly the same.

[17] The standard deviations for the annual changes of the frictionless and realized prices are 4.3 and 4.2, respectively.

1980, 1989 and 1997[18]. As it could be seen, there are some differences not only in the mean of these distributions, but also in the variance and skewness; according to the model, these differences could help us to understand the aggregate behavior of prices. For example, although in 1989 the mean was higher and the variance was lower than in 1980, inducing a bigger downward adjustment, the distribution was less skew to the right, implying a movement to the opposite side. In any case, in interpreting the results, we must be specially cautious given the few observations (forty or even less in the case of asymmetrical hazard functions) available to calculate the moments.

In the third stage we consider two alternatives. Firstly, we use these quarterly cross-sectional distributions -to approximate the different moments required- and the aggregate observed and optimal prices ($p, \tilde{p}$), to estimate the models described in section 2. We estimate these models by the generalized method of moments (GMM) taking the third lag of the variables in the regression as instruments, to avoid problems of endogeneity with the aggregate shock. Secondly, we implement a procedure very similar to the one proposed by Caballero and Engell (1993b) assuming a normal distribution, with zero mean and standard deviation σ_I, for the individual shocks and using the same aggregate shock as before. Unlike these authors, we consider the average cross-sectional distribution of sectoral price deviations as the initial distribution. This alternative strategy is implemented as a grid-search procedure to estimate the parameters in the equations of section 2. This procedure selects those values of the parameters that minimize the squared sum of the residuals.

4- Results

4-a) Estimates of alternative hazard functions

Table 1 presents the estimates of the alternative adjustment hazard models for the period 1980:1-1998:4 obtained using the sequence of cross-sectional distributions estimated from sectoral data. The corresponding estimated hazard functions are displayed in Chart 4. Column 1 shows the estimates of equation [2] corresponding to the partial adjustment model ($\Lambda(z) = \lambda_0$). The estimated constant hazard parameter is 0.36 and it is significant. Column 2 presents the estimates of the simple quadratic-hazard model (equations [3] and [4]). The quadratic hazard parameter is positive and significant implying that the hazard function is increasing. However, the explanatory power for the aggregate price level does not improve. The sum of squared residuals is indeed larger than in the partial adjustment model. Column 3 presents the estimates of the asymmetric-constant-hazard model (equations [5] and [6]). The estimated non-constant parameter corresponding

[18] In the empirical work we use the quarterly cross-sectional distributions, but in order to reduce volatility, in this chart we consider the whole year.

to price increases (l_0^-) is 0.49 whereas the estimated non-constant parameter corresponding to price decreases (l_0^+) is 0.33. This result would suggest that the resulting hazard function is asymmetric: for a given price deviation in absolute value, firms are more likely to increase their prices than to decrease them. However, the hypothesis of equality of l_0^- and l_0^+ cannot be rejected. Again, the explanatory power with respect to the symmetric constant hazard model does not improve. Finally, column 4 presents the estimates of the asymmetric-quadratic-hazard model (equations [7] and [8]). In this case, although the parameter for price increases is larger than the corresponding for price decreases, it is not significant. Moreover, with this estimate, we also reject that the hazard function is asymmetric.

Table 2 displays the results for the adjustment hazard models estimated following the procedure proposed by Caballero and Engel (1993b). The corresponding estimated hazard functions are displayed in Chart 5. Column 1 shows the estimates corresponding to the partial adjustment model ($\Lambda(z) = \lambda_0$). The estimated long-run constant hazard parameter is 0.58, substantially higher than the estimated with the former approach. Column 2 presents the estimates of the simple quadratic-hazard model. The long-run quadratic hazard parameter is positive and significant implying that the hazard function is increasing. Moreover, the explanatory power for the aggregate price level slightly improves. The sum of squared residuals for the dynamic quadratic hazard model is around 3% smaller than for the partial adjustment model. Column 3 presents the estimates of the asymmetric-constant-hazard model. The estimated long-run non-constant parameters for price increases (l_0^-) is significantly larger than the estimated long-run non-constant parameters for price decreases (l_0^+). Thus, the asymmetry of the hazard function is not rejected in this case. Unlike what happened with the first estimation approach, the explanatory power with respect to the symmetric constant hazard model improves. The sum of squared residuals for the asymmetric-constant-hazard model is substantially smaller (12%) than for the partial adjustment model. Finally, column 4 presents the estimates of the asymmetric-quadratic-hazard model. In this case, the parameter for price increases (l_2^-) is larger than the corresponding for price decreases (l_2^+), but they are not significantly different. Thus, with this estimate, we cannot reject that the hazard function is increasing, but we reject the asymmetry property again. The dynamic asymmetric-quadratic-hazard model has a sum of square residuals larger than the partial adjustment model.

The difference between the two approaches to estimate the alternative hazard models relies on the way to compute the sequence of estimated cross-section distributions of individual price deviations. In the first approach, it is proxied by the sequence of cross-section distributions of sectoral price deviations. In this procedure, there is not needed an explicit assumption about the distribution of the idiosyncratic shocks. In fact, the aggregate shock, the hazard function and two consecutive cross-section distributions would implicitly allow estimating these idiosyncratic shocks. In the second approach, however, an explicit assumption about the distribution of the idiosyncratic shocks is used to derive the whole sequence of cross-sectional distributions. As the comparison of tables 1 and 2 reveals, both procedures generate, to a certain extent, different results. Overall, when we make use of the sectoral data we estimate models with higher explanatory power (lower sum of the squared residuals) than when we use only aggregate data, the only exception being the asymmetric-constant hazard model. It is also worth noting that the estimated probability of adjustment using the first approach tends to be smaller. This suggests that allowing for a time-varying distribution in the idiosyncratic shocks leads to a higher degree of stickiness at the individual level. Finally, the asymmetry property is not rejected in the asymmetric-constant hazard model estimated using the second approach.

4-b) Simulation: the impact of nominal rigidities on the aggregate price level

As it has been already mentioned, Chart 2 compares the observed inflation rate with the inflation rate that we would observe if firms did not face adjustment costs of changing prices. That chart shows that the differences between both rates may be fairly large, at least for some periods, suggesting that when price adjustment is costly, firms do not adjust to every shock in their price determinants. A straightforward implication is that the evolution of future inflation will be conditioned not only by the future shocks to the frictionless price determinants but also by the delayed price adjustments and the expectations of firms.

In order to assess the size of the delayed price adjustment, we present some simulation exercises. In these exercises we compare the evolution of the inflation rate in a model with adjustment costs of changing prices in every sector of the economy, with the inflation rate that we would observe in the absence of nominal rigidities in a subset of sectors. In other words, we allow for differences in the price-setting behavior across sectors. More precisely, in some sectors the price rule is given by the estimated hazard functions, while in others, prices immediately adjust to changes in their determinants.

Table 3 presents the results of these simulation exercises. Each of these exercises compares the sectoral contribution to the aggregate inflation rate that arise from the fitted quadratic hazard model[19] with the sectoral contribution to the aggregate inflation rate

[19] The results obtained using the fitted values of the constant hazard model are slightly different. This means that the choice of the specific hazard function matters. Although there is no improvement in the goodness-of-fit of the quadratic hazard model with respect to the constant hazard model, the results of the simulations using both models differ to some extent. Presumably, the explanatory power of the non-constant hazard models would be more relevant for periods in which shocks that have a large variance and are asymmetric across sectors predominate. Similarly, Caballero and Engel (1993b) found that the quadratic hazard model outperforms the constant hazard one during large shock episodes.

computed assuming that there are not nominal rigidities during the period 1996-1998 in that subset of sectors. Four subsets of sectors are considered: industrial and agricultural domestic prices, industrial and agricultural export prices, consumer prices of goods and consumer prices of services. In Table 3, for each of these subsets of sectors, the first row ('quadratic hazard') represents the contribution (of that subset of sectors) to the aggregate inflation rate that arises from the fitted quadratic hazard model. The second row ('full adjustment') represents that contribution computed assuming that there are not nominal rigidities during the period 1996-1998 in that subset of sectors. Finally, the following rows (third to fifth or third to fourth, depending on the sectors) represent the decomposition of this last contribution into the contributions of the different price determinants (unit variable costs, external prices and demand pressure). Therefore, we can read the difference between the first and the second row as the impact on the aggregate inflation rate stemming from the existence of nominal rigidities in a subset of sectors. A first look at the results suggests that the magnitude of the effect of the nominal rigidities on the aggregate inflation rate is non-negligible.

Let us now go into the details of the specific simulations. The first set of results corresponds to the simulation of absence of rigidities in the adjustment of agricultural and industrial domestic wholesale prices. We observe that if firms in these sectors did not face adjustment costs of changing prices during the period 1996-1998, the aggregate inflation rate would have been around 0.12% and 0.71% higher in 1996 and 1997, and almost 0.15% lower in 1998. The results of the simulation for this subset of sectors fit quite well with the existence of nominal rigidities. First, in 1997 firms in these sectors delayed the translation into their prices of the increment in demand pressure. Similarly, in 1998 these firms did not respond lowering their prices to the fall in the unit variable costs and demand.

The second set of results displays the impact on aggregate inflation of the existence of nominal rigidities in the adjustment of agricultural and industrial export wholesale prices. In the absence of nominal rigidities, the aggregate inflation rate would have been 0.23% lower in 1996 and 0.15% higher in 1997. This delayed adjustment implied that in 1998 firms in these sectors did not react fully to the fall in external prices and demand pressure by lowering their prices.

The third set of results, corresponding to the retail prices of consumer goods, suggests that in the absence of nominal rigidities the aggregate inflation rate would have been slightly lower (under 0.1 % on average). Finally, the last set of results corresponds to the simulation of absence of rigidities in the adjustment of retail prices of services. We observe that, in this case, the aggregate price level would have been higher in the first two years (0.22% and 0.31%, respectively) and lower in 1998(0.10%).

5- Conclusions

The partial adjustment model represents the standard approach to induce aggregate price stickiness. This model, based on the introduction of quadratic adjustment costs of changing prices, assumes a continuous but partial price adjustment at the firm level. This behavior is at odds with the uncontroversial fact that individual firms do not adjust continuously their prices to all the shocks they perceive in their price determinants. The search for alternative characterizations of the aggregate price dynamics stemming from a more realistic microeconomic behavior has led to models considering state-dependent pricing rules. The adjustment hazard framework proposed by Caballero and Engel (1993a) provides an example of this type of models. In this framework, the adjustment policy at the firm level is characterized by means of the adjustment hazard function, which links the probability of a firm adjusting its price to the size of the deviation from its target price. Considering that firms are heterogeneous, the aggregate dynamics generated by non-constant hazard functions will depend on the cross-sectional distribution of price deviations.

Using this hazard adjustment framework, this paper tries to illustrate the implications for aggregate price dynamics of alternative characterizations of microeconomic price adjustment policies. For this purpose, we follow two different approaches. Firstly, we estimate alternative models of individual price adjustments using a data set of Spanish sectoral price indexes covering the period 1978-1998. We use a three-step procedure to estimate the different models. In the first step, we estimate sectoral price equations and take the fit of each regression as the sectoral frictionless price, that is, the price that would be observed in the absence of nominal rigidities. In the second step we calculate the optimal prices, having in mind that if firms are forward-looking and they have a non-null probability of not adjusting prices the following periods, they need to take into account their expectations on price evolution. In the third stage, we use the cross-sectional distribution of the price deviations from optimal prices to proxy for the different moments required to estimate the aggregate price equation. In the second approach, we implement a similar procedure as in Caballero and Engel (1993b) that avoids the use of sectoral information. Under both approaches we compare the ability of the estimated hazard models to explain the behavior of the aggregate price level, taking the partial adjustment model as a benchmark. Finally, we present simple simulation exercises in order to provide an assessment of the impact on the aggregate inflation rate stemming from the existence of nominal rigidities in several subsets of sectors.

The main results of the paper may be summarized as follows. Models with non-constant hazard functions slightly outperform the partial adjustment model, i.e. the aggregate behavior of prices depends not only on the past average deviation from target prices (as it is assumed when using a partial adjustment model) but also on higher moments of the cross-section distribution of price deviations. Although the improvement in the goodness-of-fit with respect to the partial adjustment model is only modest, it may be important especially for those periods in which shocks that have a large variance and are asymmetric across sectors predominate. Besides, there is weak evidence of asymmetries in the price adjustment, that is, in the short-term the probability of adjusting prices upwards is higher than downwards. Finally, the magnitude of the effect of the nominal rigidities on the aggregate price level is non-negligible.

APPENDIX 1

SECTORAL DESEGREGATION

Our data base includes information for twenty one sectors of the economy (agriculture, energy, ten manufacturing sectors, building, retail trading and seven sectors from services) that covers the corporate non-financial side of the economy, but we manage to consider forty sectors. We do that splitting the selling prices for some of the sectors. In the case of manufacturing, agriculture and energy, we can separate prices of goods sold in the domestic and foreign markets, so we incorporate twelve more sectors. This has a drawback: the determinants of them in each sector (unit variable costs and external prices) are the same in the two markets. In retail trading we can distinguish the prices of eight different products, so we add other seven sectors. In this case, the unit variable costs in each category are different, because the most important part of them are the buying wholesale prices for the traded goods, but unit labor costs and intermediate input costs are the same.

Table A.1 displays the detailed results for the sectoral cointegration regressions that we use to obtain the sectoral frictionless prices. The regressions were run by OLS, and all of them included a constant and three seasonal dummies. The statistical tests we use to test cointegration are the most usual: Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP).

TABLE A.1. DETAILED RESULTS FOR COINTEGRATION REGRESSIONS. ESTIMATION METHOD: OLS							
	Cvu	ep^x	(yd-y)	deterministic component	Standard Deviation	ADF	PP
Agricultural and Industrial Domestic Wholesale Prices (IPPA and IPRI)							
Agriculture	1	-	0.47	-	5.87	-4.46	-4.54
Energy	0.92	0.08	-	S 86Q1	5.50	-4.55	-4.50
Basic Metal	1	-	0.50	-	3.52	-2.79	-2.94
Mineral	0.75	0.25	0.65	T	3.73	-3.64	-3.71
Chemical	0.85	0.15	-0.25	-	3.75	-3.09	-3.04
Equipment	0.91	0.09	0.11	-	2.42	-3.09	-2.96
Transp. Eq.	0.89	0.11	0.44	-	3.75	-3.40	-2.94
Food	1	-	1.27	-	3.25	-4.91	-4.83
Textile	0.85	0.15	0.53	-	3.31	-3.09	-2.80
Wood	1	-	0.53	T 91Q1	3.28	-5.70	-5.35
Paper	0.84	0.16	-	S 84Q1	3.77	-2.93	-3.07
Plastic	1	-	0.41	-	3.57	-4.38	-3.35
Agricultural and Industrial Export Wholesale Prices (IVUX)							
Agriculture	1	-	0.03	T	6.53	-4.60	-4.77
Energy	-	1	-	S 86Q1	13.74	-4.87	-4.54
Basic Metal	0.53	0.47	0.41	-	4.21	-3.23	-3.36
Mineral	0.58	0.42	-	S 82Q1	4.11	-5.22	-5.29
Chemical	0.49	0.51	0.84	-	6.47	-3.13	-2.96
Equipment	0.73	0.27	0.53	S82Q1	5.25	-2.94	-2.81
Transp. Eq.	0.61	0.39	0.45	T	4.67	-3.41	-3.22
Food	-	1	-	-	7.23	-4.50	-4.48
Textile	0.55	0.45	0.71	-	4.81	-2.63	-2.71
Wood	0.80	0.20	1.14	S 95Q1	4.66	-3.86	-3.69
Paper	0.59	0.41	-0.34	-	5.92	-2.92	-2.69
Plastic	0.53	0.47	-	T	6.62	-2.65	-2.16
Retail Prices for Consumer Goods and Services (CPI-B and CPI-S)							
Food	1	-	0.62	-	2.98	-2.94	-2.89
Energy	1	-	-	S 86Q1	4.54	-4.05	-4.06
Equipment	1	-	-	-	6.39	-3.48	-1.94
Chemical	1	-	-	S 86Q1	1.93	-4.00	-3.81
Transp. Eq.	1	-	1.17	S 86Q1	4.64	-3.87	-2.55
Textile	1	-	-	S 86Q1	2.58	-3.49	-3.60
Wood	1	-	0.23	S 80Q1	2.60	-4.72	-2.40
Paper	1	-	-	S 86Q1	2.15	-4.15	-4.08
Construction	1	-	0.18	T 94Q1	3.53	-3.03	-2.92
Repairs	1	-	0.71	S 89Q1	3.68	-3.15	-3.09
Hostess	1	-	-	T	2.15	-4.33	-4.33
Transp.Serv.	1	-	0.48	T86Q1	2.72	-7.93	-8.02
Communic.	1	-	2.36	-	10.85	-8.48	-8.61
Education	1	-	0.67	T	4.68	-7.88	-7.99
Health	1	-	-	T 84Q1	2.80	-7.89	-8.21
O. Services	1	-	-0.36	S 84Q1	2.57	-7.96	-8.21

Notes: ADF, augmented Dickey-Fuller test; PP, Phillips-Perron test; S, step dummy; T, trend dummy.

TABLE 1. GMM RESULTS FOR DIFFERENT HAZARD FUNCTIONS. DESAGREGATED APPROACH.				
	Constant-Hazard (Expression [2])	Quadratic-Hazard (Expression [4])	Asymmetric- Constant-Hazard (Expression [6])	Asymmetric- Quadratic-Hazard (Expression [8])
l_0	0.36 (17.36)	0.23 (4.32)	-	0.23 (3.47)
l_0^-	-	-	0.49 (1.54)	-
l_0^+	-	-	0.33 (3.77)	-
l_2	-	11.33 (2.52)	-	-
l_2^-	-	-	-	15.01 (1.23)
l_2^+	-	-	-	13.50 (1.84)
R^2	0.87	0.87	0.87	0.86
RSS*100	0.152	0.159	0.152	0.161
DW	1.44	1.43	1.46	1.41
Sargan Test [pv]	3.82 [0.28]	8.46 [0.21]	3.07 [0.22]	9.17 [0.42]
Test for $\lambda^-=\lambda^+$ [pv]	-	-	0.16 [0.69]	0.02 [0.89]

Notes: The instruments consist of lag 3 of the variables that appear in each equation. Estimation period: 1979:1 to 1998:4.

TABLE 2. CABALLERO & ENGEL PROCEDURE FOR THE ESTIMATION OF DIFFERENT HAZARD FUNCTIONS.				
	Constant-Hazard (Expression [2])	Quadratic-Hazard (Expression [4])	Asymmetric- Constant-Hazard (Expression [6])	Asymmetric- Quadratic-Hazard (Expression [8])
l_0	0.58 (21.71)	0.46 (4.82)	-	0.01 (0.08)
l_0^-	-	-	0.65 (12.07)	-
l_0^+	-	-	0.13 (1.33)	-
l_2	-	19.50 (1.71)	-	-
l_2^-	-	-	-	40.0 (5.38)
l_2^+	-	-	-	20.0 (0.69)
s_I	-	0.0325	0.0260	0.0325
R^2	-	-	-	-
RSS*100	0.170	0.166	0.150	0.177
Test for $\lambda^-=\lambda^+$ [pv]	-	-	3.60 [0.00]	0.75 [0.46]

t-ratios between brackets.

TABLE 3. SECTORAL CONTRIBUTIONS TO THE AGGREGATE INFLATION RATE

THE IMPACT OF NOMINAL RIGIDITIES

		1996	1997	1998	Average
Industrial Domestic Prices	Quadratic hazard	0.55	-0.06	-0.30	0.07
	Full adjustment	0.63	0.64	-0.45	0.28
	Unit variable cost	0.71	0.12	-0.40	0.15
	External price	0.04	-0.03	-0.04	-0.01
	Demand pressure	-0.12	0.55	-0.01	0.14
Industrial Export Prices	Quadratic hazard	0.19	0.19	0.04	0.11
	Full adjustment	-0.04	0.36	-0.03	0.10
	Unit variable cost	0.09	0.17	0.10	0.12
	External Price	-0.04	0.06	-0.11	-0.03
	Demand pressure	-0.09	0.13	-0.02	0.01
Consumer Prices of Goods	Quadratic hazard	0.78	0.27	0.12	0.39
	Full adjustment	0.50	0.46	0.01	0.32
	Unit variable cost	0.48	0.37	-0.03	0.27
	Demand pressure	0.02	0.09	0.04	0.05
Consumer Prices of Services	Quadratic hazard	1.10	0.95	0.84	0.96
	Full adjustment	1.22	1.24	0.74	1.07
	Unit variable cost	1.26	1.10	0.76	1.04
	Demand pressure	-0.04	0.15	-0.01	0.03

Notes: 'Quadratic hazard' represents the contribution (of the subset of sectors) to the aggregate inflation rate that arises from the GMM estimated quadratic hazard model. 'Full adjustment' represents that contribution computed assuming that there are not nominal rigidities during the period 1996-1998 in that subset of sectors. Finally, 'unit variable costs', 'external prices' and 'demand pressure' represent the decomposition of this last contribution into the contributions of the different frictionless price determinants.

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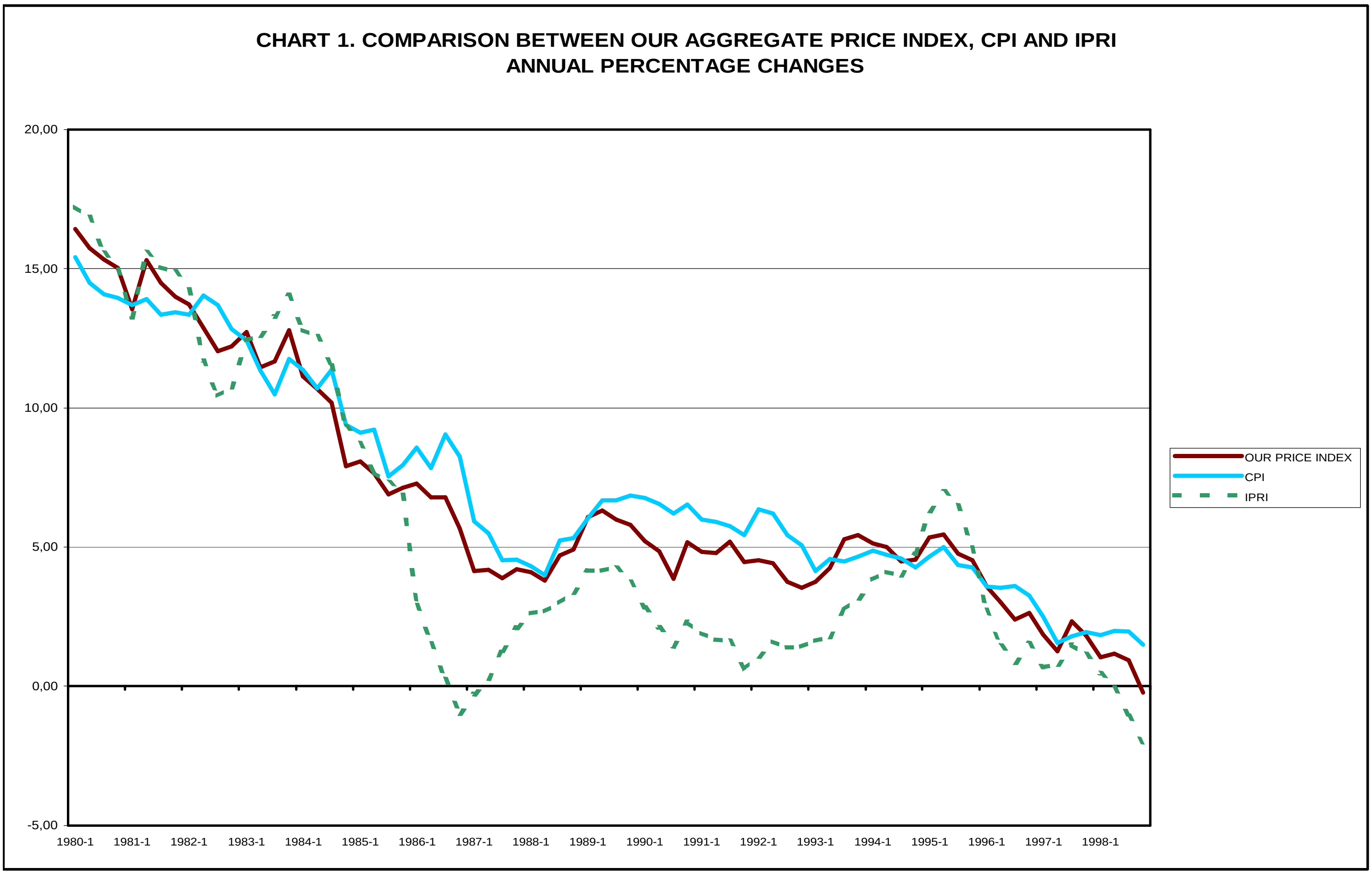
CHART 1. COMPARISON BETWEEN OUR AGGREGATE PRICE INDEX, CPI AND IPRI
ANNUAL PERCENTAGE CHANGES
20,00
15,00
10,00
5,00
0,00
-5,00
1980-1
1981-1
1982-1
1983-1
1984-1
1985-1
1986-1
1987-1
1988-1
1989-1
1990-1
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1998-1
OUR PRICE INDEX
CPI
IPRI

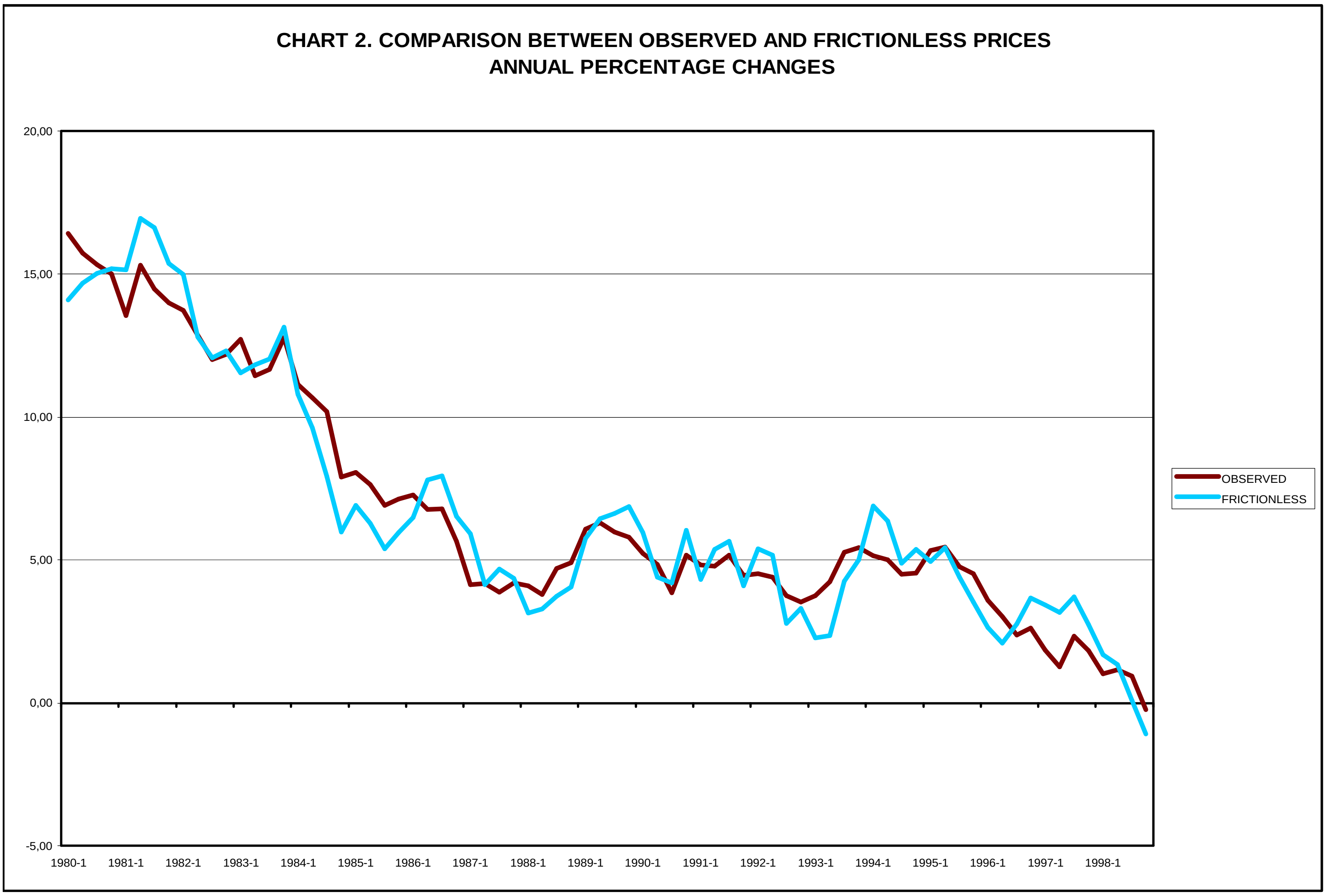
CHART 2. COMPARISON BETWEEN OBSERVED AND FRICTIONLESS PRICES
ANNUAL PERCENTAGE CHANGES
20,00
15,00
10,00
5,00
0,00
-5,00
1980-1
1981-1
1982-1
1983-1
1984-1
1985-1
1986-1
1987-1
1988-1
1989-1
1990-1
1991-1
1992-1
1993-1
1994-1
1995-1
1996-1
1997-1
1998-1
OBSERVED
FRICTIONLESS

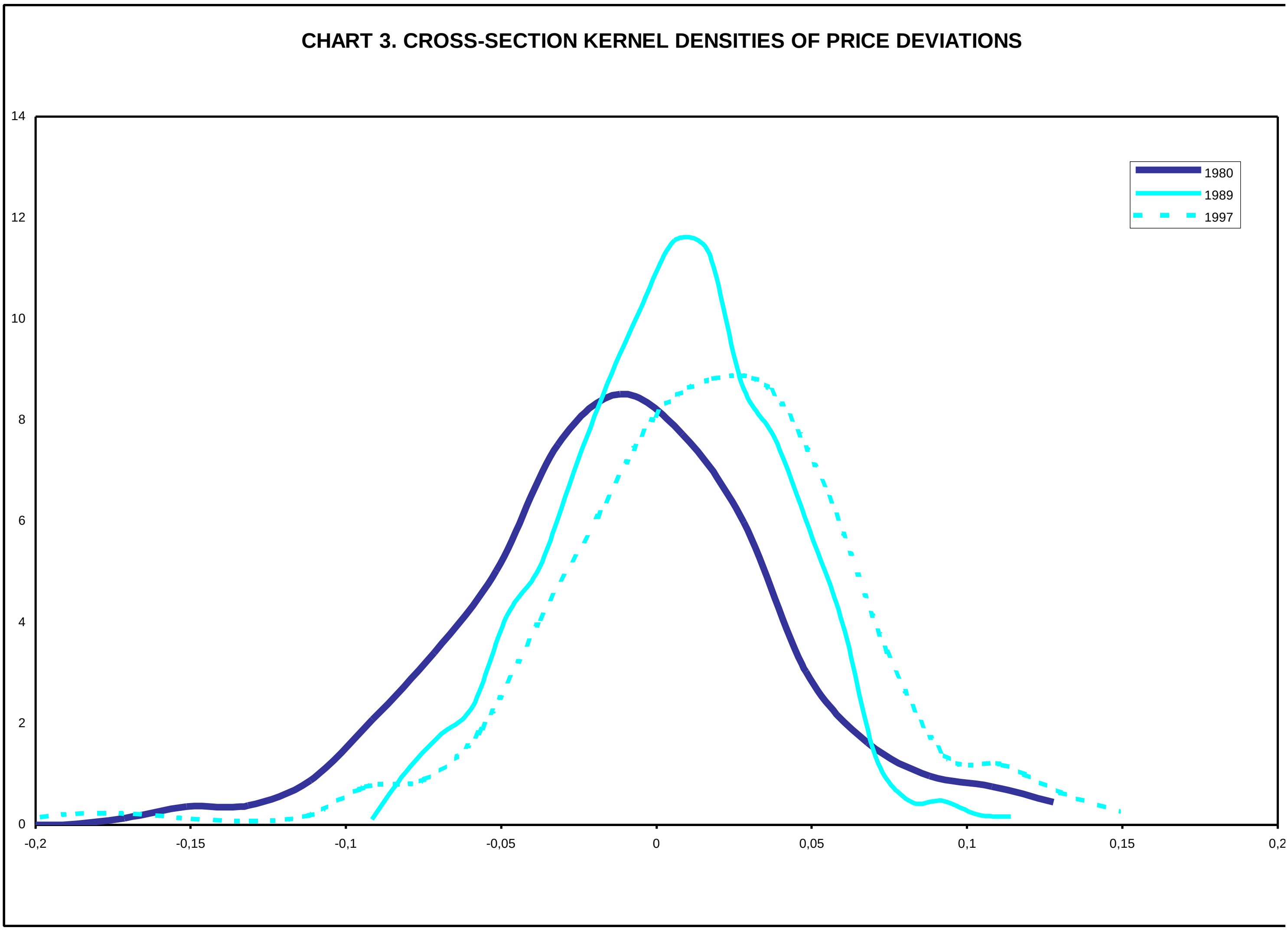
CHART 3. CROSS-SECTION KERNEL DENSITIES OF PRICE DEVIATIONS
14
12
10
8
6
4
2
0
-0,2
-0,15
-0,1
-0,05
0
0,05
0,1
0,15
0,2
1980
1989
1997

CHART 4. ESTIMATES OF HAZARD FUNCTIONS (GMM ESTIMATES)

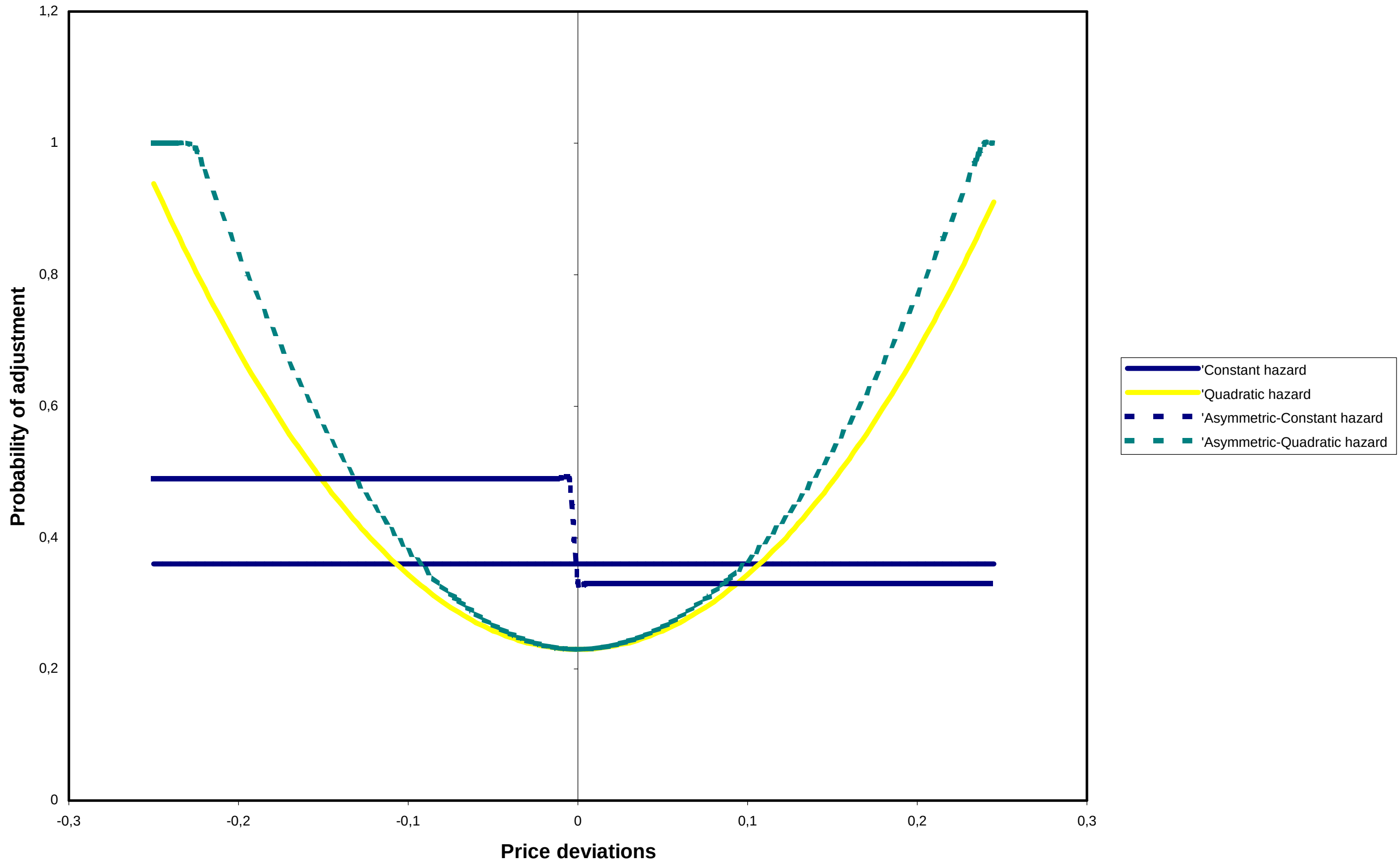

CHART 5. ESTIMATES OF HAZARD FUNCTIONS (CABALLERO & ENGEL APPROACH)