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IN THE EURO AREA

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Inflation and the Markup in the Euro Area*

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Abstract

The paper implements a consistent empirical strategy in order to investigate the behaviour of the markup over the cycle and its contribution to inflation movements. We model the price series as I(2) components and use polynomial cointegration in order to recover a long-run price schedule.

We do not reject statistically the reduction of the I(2) framework to an I(1) model as from the mid 1980s. We observe that the markup is fairly counter-cyclical and has a permanent effect on inflation through an error-correcting mechanism. Structural and forecasting equations exhibiting good performance are therefore estimated.

Key words *:* inflation, euro area, markup model, I(2) models, cointegration

JEL classification: C33, C53, E37

Résumé

L'article met en oeuvre une stratégie empirique cohérente afin d'étudier l'évolution du taux de marge au cours du cycle et sa contribution à l'inflation. Les variables de prix sont modélisées comme des processus I(2) et une courbe de fixation des prix à long terme est mise en évidence sous la forme d'une relation de cointégration de type polynomial.

La réduction du système I(2) sous la forme d'un modèle I(1) n'est pas rejetée par les données à partir du milieu des années 1980. Le taux de marge est assez largement contracyclique et a un effet permanent sur l'inflation par l'intermédiaire d'un mécanisme à correction d'erreur. Des équations structurelles possédant de bonnes performances en prévision sont ainsi estimées.

Mots clefs *:* inflation, zone euro, modèle de marge sur les coûts, modèles I(2), cointegration.

Classification JEL: C33, C53, E37.

Non technical summary

The paper studies the effects of the variations of the markup, defined as the difference between consumption prices and unit costs (unit labour cost and import prices). Such a variable is used to characterise the dynamics of inflation in the euro area. A structural model is estimated with two equations : a price schedule and a demand curve of the IS type. Such a structural model exhibits excellent out-of-sample forecasting properties. Special attention is given to the stability over time of the structural equations.

Two issues are more specifically investigated. On the one hand, the variations of the markup over the cycle -namely the pricing behaviour of producers and retailers when demand is high or when unit factor costs increase- has been extensively studied in the economic literature, often with conflicting results. Rotemberg and Woodford (1996) survey that literature and rather conclude that the markup is contracyclical, in the sense that it would decrease at the end of the upswing. On the other hand, the paper implements a consistent empirical strategy to take into account all possible sources of non - stationarity in the system (consumption prices, unit labour costs, import prices, interest rates, unemployment rate) on the period starting in the middle of the 1980s. Data are quarterly. On the basis of standard tests, price variables are seen as I(2), meaning that stationarity is only achieved after twice differentiation. Such a results is the consequence of the multiple shocks that affected euro area inflation over the period. Uncovering "polynomial" cointegration relations -where relationships among I(2) variables measured in level become stationary with the introduction of one of the variables in first difference- allows to estimate stationary relations on the levels of the variables, consistently with economic theory, in order to introduce error correction mechanisms. In our case, following Banerjee et al. (2001 and 2004), the main polynomial relationship is expressed as the link between, on the one hand, the markup, measured as the difference between the logarithms of consumption prices and a weighted sum of unit labour cost and import prices and, on the other hand, inflation. Statistical tests are available to check that such relationships can be expressed as an I(1) system without losing information. The paper shows that these " reduction " tests are more likely not to be rejected from the mid 1980s onwards, in particular when the price of oil is introduced as predetermined variable, or loosely speaking, when they are exogenous to the system under study. The paper uses the reduction tests in order to test possible breaks in the relationships.

Empirical results provide evidence in favour of the introduction of inflation in the cointegration relation, but one innovation of the paper is also to introduce the unemployment rate in the VAR system -and not to assume that it is exogenous-, in order to study the impact of the business cycle on the markup. It turns out that a reduction in the unemployment rate rather tends to decrease the markup in the short run, a conclusion which is consistent with a contracyclical behaviour. The last part of the paper studies the out-of-sample (one-year) forecasting performance of the structural model. It outperforms several alternative statistical models, indicating that the identification of cointegration

relationships enables to improve significantly forecasting on almost all the periods we consider.

Résumé non technique

L'article analyse l'impact des évolutions du taux de marge qui est défini comme l'écart entre les prix à la consommation et les coûts unitaires (coûts unitaires du travail et prix d'importations). Cette variable est utilisée pour caractériser la dynamique de l'inflation dans la zone euro. Un modèle structurel est estimé, comportant deux équations : une équation de fixation des prix et une courbe de demande, de type IS. Ce modèle structurel possède aussi d'excellentes performances en prévision hors échantillon. Une attention particulière est apportée à la stabilité dans le temps des relations structurelles ainsi dégagées.

Deux questions sont plus spécifiquement étudiées. D'une part, l'évolution du taux de marge au cours du cycle -c'est-à-dire le comportement de fixation des prix des producteurs et des distributeurs lorsque la demande s'accroît ou lorsque les coûts unitaires de production augmentent- a fait l'objet de nombreuses études dans la littérature économique, avec des conclusions souvent contradictoires. Rotemberg et Woodford (1996) en font la recension et concluent que le taux de marge est plutôt contra-cyclique, au sens où il tendrait à se réduire à la fin de la phase haute du cycle. D'autre part, il est mis en uvre une stratégie empirique cohérente des sources de non stationnarité présentes dans le système étudié (prix à la consommation, coûts salariaux unitaires, prix d'importation, taux d'intérêt et taux de chômage) sur la période débutant au milieu des années 1980. Les données utilisées sont trimestrielles. Sur la base des tests usuels, il apparaît en effet que les variables de prix ont tendance à ressortir comme I(2), c'est-à-dire qu'elles ne sont stationnaires que lorsqu'elles sont différenciées deux fois. Ce résultat découle de la multiplicité des chocs qui ont affecté l'inflation dans la zone euro sur la période. La mise en évidence des relations de cointégration de type " polynomial " -où des relations entre variables I(2) sont rendues stationnaires par l'introduction d'une des variables du système en différence première- permet de conserver des relations sur les niveaux des variables, comme suggérés par la théorie économique, tout en dégageant des équilibres véritablement stationnaires et s'interprétant comme des mécanismes à correction d'erreur. Dans notre cas, partant des travaux de Banerjee et al. (2001 et 2004), la principale relation polynomiale relie d'une part le taux de marge, mesuré par la différence entre le logarithme des prix de consommation, et une moyenne pondérée des coûts salariaux unitaires et des prix d'importation en euro, et d'autre part l'inflation. Des tests statistiques sont néanmoins disponibles pour vérifier si ces relations peuvent s'exprimer aussi sous la forme d'un système I(1) sans perdre d'information. L'article montrent que ces tests de " réduction " sont plus facilement acceptés à partir du milieu des années 1980, notamment si l'on introduit aussi le prix du pétrole comme variable pré-déterminée, c'est à dire à peu près exogène au système considéré. Ces tests de réduction sont utilisés dans le papier pour déterminer l'existence de ruptures éventuelles dans les relations statistiques.

Les résultats empiriques confirment la présence de l'inflation dans la relation de cointégration, mais l'innovation du papier est d'introduire aussi le taux de chômage dans le système VAR -et non pas de le considérer exogène-, de façon à étudier l'impact du cycle sur le taux de marge. Il ressort finalement que la baisse du taux de chômage conduit plutôt à réduire le taux de marge à court terme, ce qui est cohérent avec la thèse de son caractère contracyclique. La dernière partie de l'article étudie les performances en prévision hors échantillon (à un an) du modèle structurel. Ses performances sont supérieures à plusieurs types de modèle statistiques, montrant que l'identification de relations de cointégration permet d'améliorer la prévision de façon sensible sur presque toutes les périodes considérées.

1 Introduction

With the introduction of European Monetary Union, forecasting inflation and having a deep understanding of its main mechanisms have become key for the Eurosystem as well as economic agents in the euro area. Arguably, the issue of forecasting inflation in itself has already been addressed in many papers, including by some of the authors of the current paper (Bruneau, de Bandt, Flageollet, 2003). In addition, many theoretical models of inflation have been estimated or calibrated on the euro area (Clarida, Gali, Gertler, 1999, Jondeau and Le Bihan 2002, among others). However, there remains, in our view, a gap between these two approaches. Forecasting models, on the one hand, even if they exhibit excellent performance, often concentrate on the short run and do not convey a sufficiently "rich" explanation of the underlying developments. Theoretical models, on the other hand, do not address the issue of forecasting performance, even if they include forward-looking components. The aim of this paper is to bridge -at least partially- the gap between these two approaches.

We propose therefore a series of VECM models of euro area inflation, where the identification of the long-run relationships is made through explicit reference to economic theory. This is, in our view, a safeguard against selecting models that are unstable over time. We present structural models of euro area inflation and markup behaviour at the quarterly frequency on the basis of a consistent statistical approach. We derive variants of them -i.e. based on the same theoretical underpinnings- with good forecasting properties.

We also set out to investigate the relative "persistence" of inflation resulting from the combination of oil shocks, demand shocks and margin behaviour. In particular, our aim is to reassess margin behaviour over the business cycle. Rotemberg and Woodford (1999), in their survey of the literature on cyclical variations of prices and costs, note that "marginal costs rise more than prices during expansion, in particular late in expansion", implying a reduction in markups at the end of the cycle. In contrast, de Brouwer and Ericcson (1999) provide an empirical model of inflation in Australia and conclude that the markup is procyclical, when measured by the output gap.

The issue of margin behaviour has indeed a long history in economics, but we should distinguish between two strands of the literature, since we must take into account both the labour and the goods market. On the one hand, the literature on the profit share studies the distribution of value added between labour and capital, assuming the markup on the goods market is fixed. The latter results from the equilibrium on the labour market only (the "size of the pie" is given). Various explanation are given for the changes in the labour share: changes in factor costs, cyclical variations, etc. Represensative of the literature on the profit share, Prigent (1999) concludes that in the short run the profit share in France, for the 1964-1996 period, tended to decrease somewhat during periods of high growth[1]. Conversely, more recently, Baghli, Cette and Sylvain (2003) conclude, also in the case of France, that the profit share was, in the long run, positively correlated with the capacity utilisation rate over the 1970-2001 period. However, these papers do not investigate the implications for final prices, although a higher profit share may obviously be associated with higher prices if wages exhibit nominal rigidity. On the other hand, the literature on the markup considers that, depending on market power, companies might also wish to increase the total "size of the pie" by raising prices on the goods market, i.e. by varying the markup, which is not constant over time. Banerjee et al. (2001a and b and 2004), building on the conclusion of the literature on the markup, provide evidence that the markup, defined as the log-difference between consumption prices and a weighted sum of production costs, decreases as inflation increases.

However, while earlier papers, as mentioned above, introduced the business cycle among the determinants of the markup, it is difficult to do the same when the markup is itself non-stationary. In the work of Banerjee et al., the business cycle appears to be a predetermined variable, which only plays a short-run role, preventing any possible feedback from prices.

In what follows, we introduce directly unemployment in the long-run relationship, without detrending it, contrary to Banerjee et al. This enables us to relax the assumption that pricing is independent of demand conditions

More generally, we show that the choices made for statistical modelling are decisive in characterizing the dynamics of the prices and the behaviour of the markup over the cycle. In particular, modelling persistent and transitory comovements of the series studied is of great importance in deriving relevant empirical results in order to shed light on the questions under review.

From the empirical point of view, we propose a series of models based on the estimation of long-run relations between the markup and its determinants in the spirit of Banerjee (2001a and b, and 2004) and Juselius (2002), but we identify it as a pricing relation.

Given their high degree of persistence, the relevant variables, namely prices and costs, are taken to be I(2). Arguably, the alternative hypothesis, namely that prices are I(1), hence inflation is stationary or I(0), is a more standard assumption. However, euro area inflation appears to be statistically I(1), over a long period, due to its high degree of persistence. In addition, since the creation of EMU, over a short time span, the euro area was hit by many shocks (oil price shocks, changeover to the euro, long swings in the exchange rates) and adding a series of dummy variables representative of the different regimes would mean, from the statistical point of view,

[1] The author actually deals with the labor share, which is the complement of the profit share, and finds that it increases in periods of high growth.

that we would rapidly run out of degrees of freedom. The standard I(1) framework would not in principle be sufficient to accomodate relations derived from economic theory on variables in level. Nevertheless, using results from Konsted et al. (2002), it is possible to constrain the I(2) system into an I(1) system so as to benefit from the satistical tools developped in the I(1) framework.

The paper is organised as follows. Section 2 presents the theoretical underpinnings of the model. Section 3 describes our data set. Section 4 discusses the I(2) statistical framework and the I(2) to I(1) reduction tests. The empirical results are discussed in section 5.

2 A two-equation structural model of inflation

In order to study the dynamics of inflation in the euro area, we propose a very simple model where we include an IS curve and a wage-price schedule. The IS curve is fairly standard in the literature (see for example Gerlach and Smets, 1999)[2] :

$$y_t = A(L)y_{t-1} + \phi - \beta(C_t - E_{t-1}\Delta p_t) + e_t + \epsilon_t^y \tag{1}$$

which is an open-economy IS curve, where y_t is an indicator measuring real output (or the output gap). The other variables on the RHS are the real interest rate and the (log of the) price of the euro or the effective exchange rate.[3] At this stage, however, the exchange rate is omitted from the equation given the rather autarkic nature of the euro area and the fact that our model is only conditioned on the euro price of oil.

As regards the price schedule, we rely on literature on both the profit share and the markup. The literature on the profit share -or its complement, the labour share- studies the distribution of value added between labour and capital, assuming the markup on the goods market is fixed: to define the (optimal) labour share, the starting point is usually the (optimal) price equation, which is written as $P = \mu w / F_L'$, where P is the value added deflator, w is wage per head, L is employment, and F_L' is the marginal productivity of labour. This is equivalent, for a Cobb-Douglas production function, to $P = \mu w/(\alpha Y/L)$, where Y is value added, so that $\mu = \frac{\alpha PY}{wL}$ $= \alpha S_L^{-1}$. Changes in the markup μ are therefore directly explained by changes in the labour share S_L. On the other hand, the literature on the markup considers that, depending on market power, companies might also wish to increase the total "size

[2] See in particular Bank of England (1999), as well as as Jondeau and Lebihan (2000), who add forward-looking components to such a model. Here, for estimation purposes, only backward-looking components are used.

[3] C_t is the nominal interest rate.

of the pie" by raising prices on the goods market. Banerjee et al. (2001a and b, and 2004), building on the conclusion of the literature on the markup, provide evidence that the markup, defined as the log-difference between consumption prices and a weighted sum of production costs,[4] decreases as inflation increases, since companies lose market shares if they set prices inappropriately. However, while earlier papers, as mentioned above, introduced the business cycle among the determinants of the markup, Banerjee does not introduce activity in the equilibrium relationship defined on the level of the variables. It is only used as predetermined variable.

The wage and price setting equation may be expressed in levels, as in the Layard-Nickell Jackman (1992) tradition. The wage schedule is written as :

$$w_t = p_t + \alpha_1 \pi_t - \alpha_2 U_t + e_t^w \tag{2}$$

where w_t, p_t, π_t and U_t are wages, the price level, labour productivity and unemployment, respectively (lower-case variables are in logarithm). The price equation, or supply curve, is:

$$p_t = w_t - \gamma_1 \pi_t - \gamma_2 U_t - \gamma_3 \Delta p_t + \gamma_4 z_t^p + e_t^p \tag{3}$$

where γ_3 is the cost of inflation for firms (including, notably, menu costs, etc., see below), and z_t^p measures the effect of the business cycle on firms' competitive environment.[5] Indeed, the level of prices is directly affected by demand (or inversely by unemployment), while the business cycle and inflation also influence the markup. Regarding the effect of the business cycle on the markup, the effect is usually assumed to be negative if stronger demand requires firms to use less productive equipments leading to increasing marginal costs. Regarding inflation, one may argue that competition increases with inflation so that firms are less able to collude, yielding a negative effect of inflation on the markup. In addition, posting new prices is costly, so that the existence of menu costs leads firm to set suboptimal prices when inflation is high.

Equation (2) and (3) are obviously closely connected. Banerjee *et al.* (2001a) makes the assumption that pricing is independent of demand conditions, which is equivalent to $\gamma_2 = 0$. Empirically, we cannot adopt such an assumption as we find that the variables do not cointegrate according to a markup type relationship without introducing an activity variable (or unemployment). Moreover, we find, empirically, that $-\gamma_2 < 0$.

[4] The markup is defined as $p - \Sigma_{i=1}^{k} \varphi_i c_i$ where p is the logarithm of prices, the c_i's are the (logarithm of the) various costs of production and $\Sigma_i \varphi_i = 1$.

[5] See Appendix for the derivation of such an equation.

Actually, we use this result as an identifying constraint: indeed, $-\gamma_2 < 0$, implies that (3) is a pricing schedule. On the contrary, $-\gamma_2 > 0$ would imply that it is a wage schedule (or a Phillips curve, this would be consistent with $-\alpha_2 < 0$ in equation (2)).

In addition, we assume that $\gamma_1 = 1$.

$$mu_{1t} = p_t - w_t + \pi_t = \omega_0 - \omega_1 U_t - \omega_2 \Delta p_t \tag{4}$$

An extension of such a model is to define mu_{2t} such that

$$mu_{2t} = p_t - \frac{1}{1+\rho}(w_t - \pi_t) - \frac{\rho}{1+\rho} p_{mt} \tag{5}$$

which would also be explained by the same variables on the RHS of equation (4). The relative price between domestic prices and import prices in the domestic currency $p_t - p_{mt}$ is the real effective exchange rate. Note that such a formulation ensures a proper definition of the markup over costs, since linear homogeneity is satisfied.

According to Banerjee *et al.* (2001a), the third term on the RHS of equation (4) (i.e. $-\omega_2 \Delta p_t$) indicates that the markup decreases with inflation. Indeed, inflation, as well as unemployment has to be incorporated in the original equation for the relation to be cointegrated. This makes it possible to investigate the effect of activity on pricing behaviour.

These different long-run relationships are also used by Juselius (2002), with the additional distinction between producer prices and consumption prices (the latter being a weighted sum of producer prices and import prices). The main difference is that additional terms are used for the dynamic adjustement of the real exchange rate during the convergence (i.e. in the short-term component of the VECM).

Equation (4) is also a variant of the Ericsson and Brouwer(1998)'s markup model, where:

$$p_t = \ln(\mu_t) + \gamma_1'(w_t - \pi_t) + \gamma_2' p_{m,t} + \gamma_3' p_{oil,,t} + e_t^p \tag{6}$$

and homogeneity constraints are such that the $\gamma_i' s$ add up to one. μ_t still measures the markup which fluctuates with the business cycle (e.g. the output gap).

3 Data and stationarity tests

Variables are quarterly and defined for the euro area. $X_t = (p_t, ulc, p_{m,t,} U_t, C_t)'$ where p_t is (the logarithm of) euro area total HICP, backcasted from 1990 (see

Bruneau, de Bandt and Flageollet, 2003) and seasonally adjusted[6] ; ulc is (the logarithm of) unit labour cost, p_m is (the logarithm of) import deflator, U is the unemployment rate and C_t is a measure of nominal interest costs, defined as the equally weighted average of short- (3-month) and long-term (10-year) nominal interest rates. The price of Brent crude oil, expressed in euro, p_{oil} is used as an exogenous variable. Euro area data are computed on the basis of GDP PPP weights.[7] The sample period runs from 1976Q1 to 2002Q4, although we focus on the 1985-2002 period, which is posterior to the episode of strong disinflation that took place in many euro area countries. Data are displayed in Appendix B1.

The nominal variables p_t, ulc_t and $p_{m,t}$ clearly appear to be non-stationary. Implementing the usual unit root tests (Dickey-Fuller, as well as Elliott-Rothenberg and Stock) suggest that the price series p and ulc are well specified as $I(2)$ variables.[8] The import and oil price variables are less persistent, and might be seen as I(1). However we choose to specify the former as an $I(2)$ variable. Indeed, due to the succession of rather short-lived spikes, oil prices might be seen to be more stationary. This leaves open the possibility that the price levels share a second-order trend.

Relative prices $p_t - ulc_t$ and $p_t - p_{m,t}$ appear to be non-stationary but with a degree of persistence significantly lower than the price levels, that is $I(1)$. The unemployment rate and the interest rate are at most $I(1)$.

In general, as set out in the following section in fuller detail, there are two possibilities. Either relative price variables become stationary (i.e. the prices cointegrate from I(2) to I(0)), or they are only I(1) (the price variables cointegrate from I(2) to I(1)) but exhibit a stable relationship with other I(1) variables. Pricing according to a markup-type relationship as indicated above could be a long-run phenomenon implying cointegration between first-order non-stationary relative prices, with the possibility that the I(1) rate of inflation, the interest rate and the unemployment rate may have long-run effects in determining prices as emphasized by Banerjee et al. (2001), as well as Nielsen and Bowdler (2003).

The system of $n = 5$ variables can be decomposed into I(0), I(1) and I(2) directions, of dimension r, s and $n-r-s$, respectively. A formal test on the rank indices r and s and the associated number of unit roots is based on **a** joint rank statistic $S_{r,s}$, as explained in the next section. In the empirical part, we report the results with the 95% quantiles of the asymptotic distributions. A model is rejected only if all submodels are also rejected.

[6] In order to avoid an extra filter that might create artificial cycles, seasonal adjustment is computed by regressing the log-difference of the variable on a set of seasonal dummies (see, for example, the method described by Johnston, 1988).

[7] HICP data are monthly (as described in Bruneau, de Bandt and Flageollet, 2003), as well as oil prices, but averaged over the quarter. The other variables are quarterly and available since Q1:1970 from Fagan et al. (2001), updated using various issues of the ECB monthly bulletin.

[8] Detailed results of the tests are available from the authors upon request.

4 Econometric model

This section introduces the statistical framework that we adopt to explain and forecast inflation in the euro area.

4.1 The I(2) system

In order to introduce the I(2) system, we start with the more traditional I(1) system and motivate the differences.

If the n series studied are $I(1)$, if they obey a VAR model of order p and if they are cointegrated with a cointegration rank r, we obtain the well-known Error Correcting (EC) representation of the dynamics (Johansen (1988)):

$$\Delta X_t = \alpha\beta' X_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \varepsilon_t \tag{7}$$

where α and β are $n \times r$ denoting matrices of full rank. D_t represents the deterministic part: trend or dummies.

In the usual case, the cointegration relationships are candidates for steady-state relations. For example, if the three price variables p, ulc, p_m are integrated of order one, they could potentially cointegrate to stationarity such that the linearly homogeneous combination:

$$p - \gamma ulc - (1-\gamma) p_m$$

defines a stationary process.

The steady-state relation thus enables us to express to write the overall markup as the sum of the markup on unit labour costs, $p - ulc$, and the real exchange rate, $p - p_m$:

$$\gamma(p - ulc) + (1-\gamma)(p - p_m)$$

Accordingly, a partial adjustment model for the price level can be derived from the EC representation (1) as follows.

$$\Delta p_t = \delta_0 \Delta p_{t-1} - \delta_1[\gamma(p_{t-1} - ulc_{t-1}) + (1-\gamma)(p_{t-1} - p_{m,t-1}) - \delta_2] \tag{8}$$

where the conditions $\delta_0 > 0$ and $\delta_1 > 0$ express partial adjustement of the price level towards its steady-state value.

Now, if the variables p, ulc, p_m are recognized as $I(2)$, by substracting Δp_{t-1} from both sides of (8), we can write:

$$\Delta^2 p_t = (\delta_0 - 1)\Delta p_{t-1} - \delta_1[\gamma(p_{t-1} - ulc_{t-1}) + (1-\gamma)(p_{t-1} - p_{m,t-1}) - \delta_2] \tag{9}$$

This relation allows for an alternative intrepretation based on a dynamic steady-state relation. The price levels may still cointegrate from I(2) to I(1) such that the linearly homogeneous combination in (9) is an I(1) process which in turns becomes stationary by cointegrating with the I(1) inflation rate Δp. Thus, equation (9) -which is called a polynomially cointegrating relation- can be interpreted as describing a dynamic steady-state relation which influences the dynamics of $\Delta^2 p_t$ according to an error correcting mechanism.

More generally, the relevant representation of the dynamics of the system is given by:

$$\Delta^2 X_t = \Gamma \Delta X_{t-1} + \alpha\beta' X_{t-1} + \sum_{i=1}^{p-2} \widetilde{\Gamma}_i \Delta^2 X_{t-i} + \widetilde{\Phi} D_t + \varepsilon_t \tag{10}$$

where α and β in equation (10) are different from those in equation (7). In the I(2) representation, $\beta' X_t$ no longer defines r stationary processes but $I(1)$ processes which cointegrate with first differences of the processes.

More precisely, under the rank conditions (and denoting with a subindex $\perp$ the orthogonal complement of a given matrix):

$$\alpha \text{ and } \beta \text{ are } n \times r \text{ matrices}$$

$$\alpha'_\perp \Gamma \beta_\perp = \xi\eta' \text{with } \xi \text{ and } \eta \text{ denoting } (n-r) \times s \text{ matrices}$$

we can introduce the β_1 and β_2 matrices:

$$\begin{aligned} \beta_1 &= \overline{\beta_\perp}\eta = \beta_\perp(\beta'_\perp\beta_\perp)^{-1} \\ \beta_2 &= \beta_\perp \eta'_\perp \end{aligned}$$

with the columns of $\beta_\perp$ (resp. of $\eta_\perp$) denoting the $n-r$ (resp. $n-r-s$) directions which are orthogonal to those associated with the column vectors of β (resp. of η). β $(n \times r)$, β_1 $(n \times s)$ and β_2 $(n \times (n-r-s))$ are mutually orthogonal.

Thus, we can prove that the $r+s$ linear combinations $(\beta, \beta_1)' X_t$ both cointegrate from $I(2)$ to $I(1)$ although they differ in terms of further cointegration properties. The s-dimensional process $\beta_1{}' X_t$ remains $I(1)$ and enters the model only in first differences, while the r-dimensional process $\beta' X_t$ enters the model in levels and cointegrates to stationarity with the first differences in the so-called polynomial cointegrating relationships:

$$\beta' X_t - \overline{\alpha}' \Gamma \overline{\beta_2} \beta'_2 \Delta X_t$$

If $r > n-(r+s)$, there will exist linear combinations of variables that cointegrate directly to stationarity. Finally, there are three types of $I(0)$ processes:

$$\begin{array}{l} \Delta^2 X_t \text{ of dimension } n-r-s \\ \beta_1' \Delta X_t \text{ of dimension } s \\ \beta' X_t - \overline{\alpha}' \Gamma \overline{\beta_2} \beta_2' \Delta X_t \text{ of dimension } r \end{array}$$

4.2 Reduction from I(2) to I(1)

Now, under relevant conditions, the $I(2)$ representation can be transformed into an I(1) representation without losing information.The transformed process includes variables that reduce to $I(1)$ either by linear transformation or by first differencing.

Such a transformation is given by:

$$X_t \to \widetilde{X_t} = \begin{pmatrix} B' X_t \\ v' \Delta X_t \end{pmatrix} \tag{11}$$

with B and v denoting $n \times (r+s)$ and $n \times (n-(r+s))$ dimensional matrices provided the following conditions are satisfied:

$$\begin{aligned} \det(v' B_\perp) &\neq 0 \\ B_\perp'(\beta, \beta_1) &= 0 \end{aligned}$$

$B_\perp$ reflects the assumption about the way in which the components of X_t are affected by the $I(2)$ common trends : here, the three price components are supposed to share one $I(2)$ common trend while the other components (interest rates, and the unemployment rate) which are $I(1)$ are not supposed to be affected by the $I(2)$ trends. Thus, the loadings will be 1 for the former and 0 for the latter. Accordingly, $B_\perp = b$ is expected to be:

$$b = (1, 1, 1, 0, 0)' \tag{12}$$

and r and s are expected to satisfy:

$$n - (r+s) = 1 \tag{13}$$

Indeed, as indicated in the empirical part, we do not reject the hypothesis:

$$r = 2 \text{ and } s = 2$$

This implies the existence of one I(2) common trend ($n-(r+s) = 1$) and two possibly polynomial cointegration relations ($r = 2$). The $n \times (n-(r+s))$ matrix v is such that $v' B_\perp$ is invertible so that a full set of $n-(r+s)$ first differences of the initial variables is obtained from the transformed system $\widetilde{X_t}$, including those needed for polynomial cointegration: here $v'\Delta X_t$ reduces to Δp_t ,$v' = (1,0,0,0,0)$ and $v'B_\perp = 1$.

Given b, the matrix B has to be defined as:

$$B = \begin{bmatrix} 1 & 1 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & -1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

so that $B'X_t$ is the 4 dimensional vector:

$$B'X_t = \begin{bmatrix} p_t - ulc_t \\ p_t - p_{m,t} \\ U_t \\ C_t \end{bmatrix}$$

Finally $v = (1,0,0,0,0)'$ and $v'\Delta X_t = \Delta p_t$ and $\widetilde{X_t}$ is specified as:

$$\widetilde{X_t} = \begin{bmatrix} p_t - ulc_t \\ p_t - p_{m,t} \\ U_t \\ C_t \\ \Delta p_t \end{bmatrix}$$

Now, the transformed process $\widetilde{X_t}$ is $I(1)$ with cointegrating rank $\widetilde{r} = r$ if and only if the following constraints are satisfied (Kongsted (2002)):

$$b'\beta = b'\beta_1 = 0 \tag{14}$$

or, equivalently:

$$\begin{aligned} \beta &= B\phi \\ \beta_1 &= B\varphi \end{aligned}$$

The LR test statistic has to be compared to a $\chi^2((n-r-s)\times r)$ and $\chi^2((n-r-s)\times s)$ which are both, in our case distributed as $\chi^2(2)$.

It is worth noting that if we do not reject the hypothesis $b'\beta = 0$, the transformation includes the linear combinations of the initial variables needed for the polynomially cointegrating relations. Thus, there exists a VEC representation for $\widetilde{X_t}$, irrespective of the condition $b'\beta_1 = 0$.

$$\Delta\widetilde{X_t} = \widetilde{\Pi}\widetilde{X_{t-1}} + \sum_{i=1}^{p-1} \widetilde{\Gamma}_i \Delta\widetilde{X_{t-i}} + \widetilde{\varepsilon}_t \tag{15}$$

where the $\widetilde{\Pi}$ parameters are constrained according to :

$$\widetilde{\Pi} = (B, v)'(\alpha, \alpha_1) \left[\begin{array}{cc} \phi' & -\overline{\alpha}'\Gamma\overline{\beta_\perp}\beta'_\perp b(v'b)^{-1} \\ 0 & -\beta'_1 b(v'b)^{-1} \end{array} \right] \tag{16}$$

In any case, if the condition $b'(\beta, \beta_1) = 0$ is not satisfied, the transformed process $\widetilde{X_t}$ remains $I(2)$. Moreover, if $b'\beta = 0$ and $b'\beta_1 \neq 0$, the cointegration rank $\widetilde{r}$ of $\widetilde{X_t}$ is strictly greater than the cointegration rank r of the initial process X_t and the transformation does not provide a direct relationship between the two sets of cointegrating relations. In particular, if the transformation does not provide an $I(1)$ system, we may fail to estimate cointegrating combinations that enter the VEC model in terms of first differences (of the initial variables) only.

In what follows, we estimate directly the VEC representation of the transformed system $\widetilde{X_t}$ without referring to the estimation of equation (10). By doing so, we may run the risk of overestimating the cointegration rank, but the stability of the estimation results for the - possibly erroneous- $I(1)$ model as well as the good performances of this model in terms of out-of-sample forecasting *ex post* justify our choice.

5 Empirical results

We will now present the empirical results. The first subsection determines the rank of the I(2) system and runs the reduction tests described in the previous section. The second subsection estimates the reduced I(1) system, including the long-term relationship and its short-run dynamics.The final subsection focuses on the forecasting performance of these models.

5.1 I(2)analysis

5.1.1 Testing the rank[9]

To provide a benchmark, we first run the same kind of analysis as Banerjee (2001a and b) on the euro area, using a reduced system of 3 variables [p,$\widetilde{ulc}$, p_m], starting

[9]Critical values for the I2-rank test are given in Paruolo (1996).

in 1986, which corresponds to a period of lower inflation as compared to the 1970s and the early 1980s. In addition, as stated in the chart in Appendix B, oil prices expressed in euro returned to a lower level in nominal terms with the simultaneous decrease in oil prices in USD and the depreciation of the USD after the 1985 Plazza agreement.Unit labour costs are introduced as a moving average of t and $t-1$ values.[10] The model is estimated in levels with four lags (implying two lags in second difference). From table 1a, the conclusion is that we uncover one cointegrating vector ($r = 1$) and one I(1) common trend ($s = 1$). As a consequence, one I(2) common trend ($n - r - s = 1$) is suported by the data.

Table 1 a: *Trace test of cointegration ranks for I(2) model the case of a small system* $[p, \widetilde{ulc}, p_m]$

	Trace statistics*			
r = 0	98.14	58.88	44.6	43.47
1		54.06	**20.12**	16.54
2			10.01	5.69
	Critical Values at 1%			
r = 0	*70.87*	*51.35*	*38.82*	*29.68*
1		*36.12*	**22.6**	*15.41*
2			*12.93*	*3.84*
n-*r*-*s* =	*3*	*2*	**1**	*0*

* in bold we accept the integration indices at the 1% Level

When extending this model to our 5-variable system, namely $[p, \widetilde{ulc}, p_m, U, C]$, the model is also estimated with two lags in second difference. From table 1b, we conclude to the existence of two cointegration vectors ($r = 2$) and one, two or three I(2) common trends ($n-r-s = 1$, 2 or 3). However, since the latter two possibilities are borderline cases (they are just accepted by the tests), we prefer to conclude that there is only one I(2) common trend ($n-r-s = 1$). Using ulc instead of $\widetilde{ulc}$ would equally result in accepting one or two common trends ($n - r - s = 1$ or 2- table A.1.1.b in Appendix).

[10]The moving average on ULC is computed as the current and the lagged value of ULC, namely $\widetilde{ULC} = \frac{1}{2}(ULC_t + ULC_{t-1})$.

Table 1 b: *Trace test of cointegration ranks for I(2) model the case of the extended system* $[p, \widetilde{ulc}, p_m, U, C]$

	Trace statistic*					
$r = 0$	214.76	158.92	131.77	112.72	98.53	95.69
1		147.00	93.41	74.40	60.27	57.84
2			**70.40**	**50.45**	**34.38**	32.18
3				32.00	17.05	13.77
4					10.12	2.11
	Critical Values at 1%					
$r = 0$	*171.89*	*142.57*	*117.63*	*97.97*	*81.93*	*68.52*
1		*116.31*	*91.41*	*72.99*	*57.95*	*47.21*
2			**70.87**	**51.35**	***38.82***	*29.68*
3				*36.12*	*22.6*	*15.41*
4					*12.93*	*3.84*
n-r-s	*5*	*4*	**3**	**2**	**1**	*0*

* in bold we accept the integration indices at the 1% Level

5.1.2 Testing the reduction from I(2) to I(1)

We shall now consider whether our I(2) system can be reduced to an I(1) system, on the basis of the transformation introduced in Section 4, with the markup on unit labour cost and the real exchange rate. As indicated in Table 2, the constraints $b\prime\beta = 0$ and $b\prime\beta_1 = 0$ are globally accepted for the 1986-2002 sample period, without any predetermined variable: the p-value is above the 5% threshold.

We also extend the sample period backwards in order to test the robustness of our results. In both cases, we recursively run the joint test over the 1984Q4-2002Q4 period, by shifting progressively forward the beginning of the sample from 1984Q4 to 1987Q4, the end of the sample always being fixed at 2002Q4. In the right-hand side subpanel of Figure 1, the price of oil expressed in euro is used as predetermined variable, while no predetermined variable is used in the left-hand side panel. In comparison to the left-hand side subpanel, it appears that the introduction of the predetermined variable increases the p-value of the tests, which rises above the 5% threshold, especially when the sample starts in 1984-1985. For that period, it emerges therefore, from comparing the two subpanels of Figure 1, that the constraints are only accepted when we introduce oil prices as a predetermined variable.

Table 2: Reduction test $[p, ulc, p_m, U, C]$
(1986-2002, without predetermined variable)

Test	v	Stat*	P-value
$b'\beta = 0$	2	2.91	0.23
$b'\beta_1 = 0$	2	4.86	0.09

*Numbers in brackets are P-values according to $\chi^2(v)$

Figure 1: Reduction tests over (1984Q4-1987Q4) to 2002Q4
p-value of b'$\beta = 0$ (solid line), p-value of b'$\beta_1 = 0$ (dashed line)
without predetermined variables with Δoil as a predetermined variable

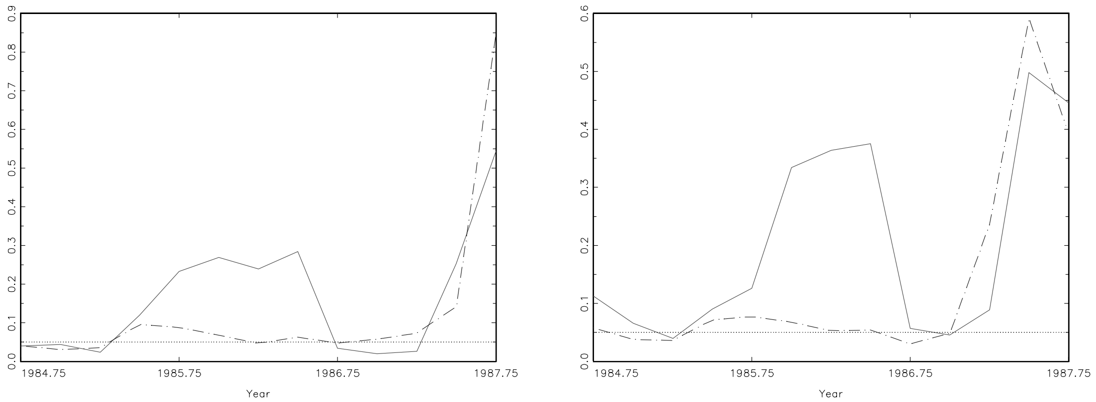

Figure 2: Loadings on the I(2) trends over (1984Q4-1987Q4) to 2002Q4
p (flat solid line equal to one), $\widetilde{ulc}$ (dashed line), p_m (black solid line)
U (short dashed line) and C (grey solid line).
without predetermined variables

At the same time, Figure 2 displays the loadings on the I(2) trends.From 1986 onwards, the loadings on the first three variables appear to be close to 1, while those for the other two variables are close to zero. Indeed for the whole 1986-2002 period, the loadings are $\beta_2 = (1, .0.84, 0.56, -0.11, -0.15)$ without any predetermined variables. The much more stable results after this date are also clear from the figure. Before that date, we need to introduce oil price as a predetermined variable.

From these various experiments, we conclude that our assumption of one I(2) common trend is generally corroborated by the data. The reduction to an I(1) system is also validated from 1986 onwards, but slightly more instability appears before, which requires the introduction of predetermined variables in order to take into account the various external shocks that hit the euro area economies.

5.2 Estimation of the I(1) real model

5.2.1 Estimating the long-run relationship in the I(1) system

On the basis of the reduction tests presented above, focusing on the sample period starting in 1986, we can now estimate the transformed systems $\widetilde{X}$ =[Δp, $p-$ $\widetilde{ulc}$, rer, U, C] which exhibit good statistical properties. As indicated in table A.1.2.b and A.1.2.c in the Appendix, two cointegration vectors are accepted at the 1% level, according to the trace test. In line with section 2, we assume that the two relationships are a markup and an IS curve. This structure is not rejected by the data since the restriction test, distributed as $\chi^2(1)$, has a value of 1.40, hence a p-value of 0.24.

For the system [Δp, $p-$ $\widetilde{ulc}$, rer, U, C], the markup type relation is estimated (with $rer = p - p_m$) as:

$$\begin{aligned}
0.072(p - \widetilde{ulc}) + 0.006(p - p_m) &= -0.133U - \Delta p + Z \\
&\Leftrightarrow \\
p - 0.93\widetilde{ulc} - 0.07p_m &= -1.68U - 12.89\Delta p + Z^{(1)}
\end{aligned}$$

with $Z^{(1)}$ denoting a stationary variable and Δp quarterly changes in HICP (Δp times 400 would be the equivalent of annual inflation in %). The negative impact of unemployment (akin to a positive impact of activity level) implies that we identify a pricing schedule, as opposed to a wage curve.

The IS-type equation is estimated as:

$$U = 0.69C - 7.09\Delta p + Z^{(2)}$$

The latter equation may be reformulated by introducing the real interest rate $C - 4\Delta p$ as:

$$U = 0.69(C - 4\Delta p) + (0.69 \times 4 - 7.09)\Delta p + Z^{(2)} = 0.69(C - 4\Delta p) - 4.33\Delta p + Z^{(2)}$$

As the dynamics are cointegrated, the estimates of the long-run parameters are superconsistent and, taking account of the whole system of 2 equations, we can evaluate the impact of the real interest rate (which expresses monetary policy) as well as inflation on the markup as follows:

- a one percentage point reduction in the real interest rate decreases the unemployment rate by 0.7 percentage point, leading to a 1.2 percentage point increase in the markup ($-1.68 \times -0.7 \approx 1.2$).
- a (permanent) 1 percentage point increase in annual inflation (Δp higher by 0.0025) implies, from the IS equation, a reduction in unemployment of around 1 p.p. ($0.0025 \times 4.33 = 1.08$), after correction for the interest rate (the nominal rate is assumed to be adjusted upaward with higher price so that the real interest rate would be constant), resulting in a decrease in the markup of 0.2 p.p. ($1.68 \times 0.18 - 12.89 \times 0.0025 = -0.0024$). This reduction is slightly smaller than the 0.5 p.p. reduction suggested by Banerjee et al. (2004).

A major difference with the latter analysis is that we provide evidence in favour of a permanent link between the markup and the level of activity, while confirming Banerjee's conclusions regarding the negative impact of Δp. Indeed, it should be noted that in our 5-variable system, the exclusion of unemployment from the long-run relationship would fail to provide a stationary cointegrating vector.

Both relations are fairly stable over time, as shown in the figures in Appendix B.2.1 (markup relation) and B.2.2. (IS curve).[11] See, in both cases, the charts "long-run coefficients β", which are displayed with their two standard deviation band. The full sample estimate presented above is very close to the central tendency provided by the confidence band. The coefficient for $p-\widetilde{ulc}$ is between 0.08 and 0.10. Although the confidence band associated with rer is sometimes close to zero, the coefficient is generally quite significant. This contrasts with the results of Banerjee *et al.* (2001b) who present non-significant results for rer in the largest euro area countries: Germany and France. The coefficient of U in the first equation fluctuates between 0.12 and 0.16. As regards the IS equation, the coefficients of Δp and C exhibit a spike in early 1998, but stabilise afterwards between 5 and 10 and around 0.5 and 1 respectively.

5.2.2 Assessing the overall fit

Using the long-run relationships we estimated on the reduced models we shall now discuss the complete model including the short-run dynamics. We focus on the first

[11] Our figures only provide the recursive coefficients for the final part of the sample, since our model is designed to be used for out-of-sample forecasting for the last observations.

two variables, namely price $\Delta^2 p$ and markup $\Delta(p - \widetilde{ulc})$. As mentioned above, the models are estimated in the form:

$$\begin{aligned}\Delta^2 p &= \alpha_{11}[\Delta p_{-1} + \beta_{12}(p_{-1} - \widetilde{ulc}_{-1}) + \beta_{13}(p_{-1} - p_{m,-1}) + \beta_{14}U_{-1}] \\ &\quad + \alpha_{12}[U_{-1} + \beta_{25}C_{-1} + \beta_{21}\Delta p_{-1}] + \sum_{i=1}^{2} \Gamma_{1i}\Delta\widetilde{X}_{t-i} + \Phi_1 D_t + \varepsilon_{1t} \quad (17)\end{aligned}$$

$$\begin{aligned}\Delta(p - \widetilde{ulc}) &= \alpha_{21}[\Delta p_{-1} + \beta_{12}(p_{-1} - \widetilde{ulc}_{-1}) + \beta_{13}(p_{-1} - p_{m,-1}) + \beta_{14}U_{-1}] \\ &\quad + \alpha_{22}[U_{-1} + \beta_{25}C_{-1} + \beta_{21}\Delta p_{-1}] + \sum_{i=1}^{2} \Gamma_{2i}\Delta\widetilde{X}_{t-i} + \Phi_2 D_t + \varepsilon_{2t} \quad (18)\end{aligned}$$

All the β_{ij}'s are positive, except β_{25}, as shown in the previous section. The error correction mechanism is different in the two equations.

When interpreting the results for the whole system -distinguishing between the effect of the error correction mechanism and the short-run dynamics- we find that unemployment (U) only has an indirect impact on the markup ($p - \widetilde{ulc}$), through the short run variations in inflation ($\Delta^2 p$). This is interpreted as a counter-cyclical effect of unemployment on the markup. In addition, as regards the indirect effect transmitted by the equilibrium error variables, it is worth stressing that there are no "short-run" indirect effects of unemployment on the variation of the markup that are transmitted by differences in the series of the system except for the inflation variable. In order to detect a direct effect, a longer sample period would have to be used, as shown in section 5.2.3.

Price equation $\Delta^2 p$ For the price equation, $\Delta^2 p$, the value of $\alpha_{11} < 0$ expresses an adjustment to the equilibrium level of inflation, which is significant as shown in Table A.2.1. and A2.2. Consequently, and as shown in Table 3 which focuses on the effects of the error correction mechanisms, an increase in unemployment has a negative and significant effect on $\Delta^2 p$ (measured by $[\alpha_{11}\beta_{14} + \alpha_{12}]$).

Table 3: Effect of inflation (Δp) and unemployment (U) through the error correction mechanisms ($\alpha\beta'$)

	$\Delta^2 p$	$\Delta(p - \widetilde{ulc})$
Δp	-0.80 (-3.98)	0.30 (0.72)
U	-0.10 (-3.66)	0.06 (0.97)

Markup equation $\Delta(p - \widetilde{ulc})$ For the markup on the labour cost equation, $\Delta(p-\widetilde{ulc})$, only the adjustment term to the first long-term relationship $(\alpha_{21}\beta_{12} < 0)$ can be viewed in terms of equilibrium correction. The sign of the α_{22} coefficient could be either positive or negative. We found it to be positive and significant..

When measuring the effect of unemployment on the markup through both correction mechanisms, as measured by $[\alpha_{21}\beta_{14} + \alpha_{22}]$, we find that it is non-significant (see bottom-right cell of Table 3).

However, the lagged variations in inflation, $\Delta^2 p_{-1}$ and $\Delta^2 p_{-2}$ have both a significant negative effect in the $\Delta(p - \widetilde{ulc})$-equation. There is therefore an indirect positive effect of unemployment on the markup, if we take into account the negative impact of unemployment on inflation, as previously described.

To summarize, the chain of events is as follows:

$$U \overset{(-)}{\rightarrow} \Delta^2 p \overset{(-)}{\rightarrow} \Delta(p - \widetilde{ulc})$$

This provides evidence showing a clear counter-cyclical effect of unemployment on the markup: a reduction in unemployment or an upward cyclical movement implies a reduction in the markup. [12] It is worth noting, as shown in next section, that the coefficient of ΔU_{-2} is significantly positive in the $\Delta(p - \widetilde{ulc})$-equation over the long period 1976-2002, giving direct empirical evidence of counter-cyclical behaviour of the markup.

As mentioned in the introduction, the modelling choice is decisive in deriving the empirical results showing counter-cyclical behaviour of the markup. Indeed, we find that, inside the long-run relationship, the markup is negatively related to the level of the unemployment rate. If we had chosen to model markup and unemployment as cyclical variables, by ignoring the persistence of these series, we would have found procyclical effects. In fact, they only share a common trend. From our model it appears that a fall in unemployment to below its long-run trend has a negative effect on the markup.

Our whole econometric modelling procedure is necessary to provide reliable empirical results showing counter-cyclical behaviour of the markup. First, the price variables are $I(2)$, while the other variables- in particular the activity variable,U- are less persistent, that is $I(1)$; second, there are dynamic (polynomially) cointegration relationships which express economic structural relationships between the different series; finally the $I(2)$ characterization of the joint dynamics can be replaced by an $I(1)$ one, which makes it possible to measure direct and indirect short-run -i.e. cyclical- links, between activity and markup.

The results are stable over time as shown in Appendix B.2.1 and B.2.2.

[12] The positive effect of unemployment on the markup appears clearly from an impulse response analysis.

5.2.3 Robustness check of the I(1) system

In order to assess the robustness of our findings, we run the analysis on an extended period, starting in 1976. It emerges that the overall structure of the model, in particular the long-run relationships, is similar, although the coefficients have a different magnitude. The main difference is that we need to include the contemporaneous value of the quarterly growth rate of the price of oil expressed in euro as a predetermined variable. Such a variable controls for oil shocks, which, as shown in Section 5.1, play a large role in particular in the first part of the sample period.

For the long-run relation, the markup-type relation is estimated as (the $\chi^2(1)$ restriction test is accepted with p-value of 0.81):

$$\begin{aligned} 0.058(p - ulc) + 0.018(p - p_m) &= -0.31U - \Delta p + Z \\ &\Leftrightarrow \\ p - 0.74ulc - 0.26p_m &= -4.1U - 13.2\Delta p + Z^{(1)} \end{aligned}$$

with $Z^{(1)}$ denoting a stationary variable and Δp expressed as monthly changes in the HICP, so that the negative coefficient on Δp is consistent with the first quarterly model. Given the higher frequency of import price shocks, the elasticity of the real exchange rate increases. In addition, the coefficient on unemployment is twice as large, due to the inclusion in our sample of the 1970s and the 1980s characterized by a higher level of inflation.

For the IS-type equation, the constraint on the real interest rate is almost verified:

$$\begin{aligned} U &= 0.39C - 2.27\Delta p + Z^{(2)} \\ U &= 0.39(C - 4\Delta p) - 0.71\Delta p + Z^{(2)} \end{aligned}$$

The elasticity of the markup with respect to inflation amounts to 0.9 p.p. which is higher than for the shorter period (0.2 p.p.) but it remains almost in line with Banerjee's estimates for the USA (0.75 p.p.). As shown in Appendix A.2.1, the short-run coefficients for the 1976-2002 period are fairly close to those obtained for the shorter period. In the price equation, only the correction term for the markup relationship has a significant impact (constraining the second error correction term to zero would leave the other coefficient almost unchanged). As regards the other short-term effects, the real exchange rate, lagged two periods, also has a negative impact. The markup equation exhibits a significant error correction term ($\alpha_{21} > 0$) while deviations from the IS curve have a positive effect ($\alpha_{22} > 0$). The overall effect of unemployment on the markup through both long-term relations is positive, but non-significant, as reported in Table 3 for the shorter period (bottom-right cell)

Nevertheless, as outlined above, the lagged variation of unemployment rate ΔU_{-2} has a significant positive coefficient, indicating counter cyclical behaviour of the markup on unit labour costs.

The figures reported in Appendix B.3.1. show that the models track inflation reasonably well in-sample, for both sample periods (1985-2002 and 1976-2002). The only caveat, from a statistical point of view, is that the I(2) reduction tests are not accepted for the 1976-2002 period, while they are satisfied for the shorter sample period.

5.3 Forecasting performance

The last criterion that we use to assess the validity of our approach is to run in-sample and recursive out-of-sample forecasts. Starting from 1985Q4, we progressively extend our estimation sample in order to provide rolling one-year ahead forecasts, starting projections in 1998Q1. It appears that our structural model exhibits excellent forecasting properties. From figures B.3.2, it appears that the model is able to track reasonably well the upturn in inflation in 1999-2000 following the rise in oil prices. From a purely forecasting point of view, the model starting in 1985, which includes oil prices as a predetermined variable, yields the best results and is able to capture more rapidly the upward movement in inflation[13].

In order to confirm the previous analysis, we also compare the forecasting properties of our structural VECM to three types of model. We consider[14]:

- an unconstrained VAR model in level,
- a VAR model in first difference,
- a simple univariate autoregressive model.

We also investigate the sensitiveness of our results to the sample period, by extending the estimation period backward. From figure B.3.3, it appears that in most cases, the structural VECM model (first block for each given year) exhibits the best forecasting performance in terms of smaller out-of-sample RMSE. The slightly worse performance in 1984 points to some instability and provides additional evidence in favour of a break in the regime of euro area inflation in the mid-1980[15].

[13] In all the experiments, future values of oil prices are forecast using a simple AR model.

[14] As indicated in Bruneau et al. (2003) the performance of the random walk for euro area inflation is inferior to the one of the Autoregressive model, on which we concentrated our analysis. To save space the random walk model does not appear in the tables.

[15] It should be noted that the predetermined variable may affect quite significantly the results, although not to the point where it would invalidate the results. For that purpose, and in order to run a more challenging comparison test, we select, for each class of models, the number of lags

6 Conclusion

In this paper we aim to investigate the dynamics of inflation in the euro area over the three last decades. We adopted a structural approach by estimating a cointegrated I(2) model, which can be further transformed into an I(1) cointegrated VAR model, in accordance with the results of a relevant transformation test. In the latter framework, we showed that there exist two long-run relationships: a price schedule and an IS-type long-run relation. Our findings are as follows. First, the markup appears to influence variations in inflation through the error correcting mechanism associated with the price schedule long-run relation. Second, we confirm recent results which tend to prove that the inflation rate persistently and negatively contributes to the markup inside the dynamic markup-type cointegration relation. However, contrary to earlier studies, the introduction of unemployment in the dynamic markup long-run relation is required in order to achieve stationarity of the equilibrium error variable.

Accordingly, we are able to capture the behaviour of the markup over the cycle only by focusing on the effects of unemployment on variations in the markup. We observe that activity does impact inflation through the error correcting mechanism, and that variations in inflation influence in turn the variations in the markup. All cyclical effects on the markup are channelled through the variations in inflation and only by this variable. By carefully measuring these indirect effects, we can conclude that the markup is counter-cyclical. It is worth stressing that treating markup, inflation and unemployment as stationary series would have been tantamount to concluding that the markup is procyclical. The empirical evidence highly depends on the econometric model specified to describe the dynamics of the system. We have recursively implemented a complete inference procedure equally using economic theory and statistical tools to obtain a reliable structural model which is quite robust and stable over time. Moreover our model provides very good forecasts of inflation one year ahead, as proved by recursive out-of-sample exercises.

7 References

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of the predetermined variable that yields the best forecasting performance and run a "horse race" between the best models. Obviously, this reinforces our conclusion regarding the superiority of the ECM model.

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A Appendix

We present here the results from tests on our different systems of variables. I(1) and I(2) models are estimated on a consistent sample period, but the reference period is the I(1) model. For example, an I(1) model in first difference estimated from 1985Q1 corresponds to an I(2) model in levels from 1984Q4. In that case, the sample period will appear in the tables from 1985Q1 onwards.

A.1 Rank tests

A.1.1 Some results with no filtered unit labour cost over 1986 to 2002

Table A.1.1 a: *Trace test of cointegration ranks for I(2) model the case of a small system* $[p, ulc, p_m]$

	Trace statistics*			
r = 0	97.30	58.25	41.43	40.23
1		48.07	**19.49**	15.78
2			9.92	5.97
	Critical Values at 1%			
r = 0	*70.87*	*51.35*	*38.82*	*29.68*
1		*36.12*	**22.6**	*15.41*
2			*12.93*	*3.84*
N-r-s =	*3*	*2*	**1**	*0*

* in bold we accept the integration indices at the 1% Level

Table A.1.1.b: *Trace test of cointegration ranks for I(2) model the case of an extended system* $[p, ulc, p_m, U, C]$

			Trace statistic*			
$r = 0$	214.05	169.23	142.29	120.02	104.80	101.98
1		142.06	100.00	77.43	62.24	59.81
2			83.71	**50.69**	**34.67**	32.84
3				34.73	18.43	14.86
4					10.15	4.14
			Critical Values at 1%			
$r = 0$	*171.89*	*142.57*	*117.63*	*97.97*	*81.93*	*68.52*
1		*116.31*	*91.41*	*72.99*	*57.95*	*47.21*
2			70.87	**51.35**	***38.82***	*29.68*
3				*36.12*	*22.6*	*15.41*
4					*12.93*	*3.84*
N-r-s	*5*	*4*	**3**	**2**	**1**	*0*

* in bold we accept the integration indices at the 1% Level

A.1.2 Rank tests for I(1) models for our three sub-periods

Table A.1.2: Trace test of cointegration ranks for I(1) model system: $[\Delta p,\ p - \widetilde{ulc}, p-\ p_m, U, C]$

a- Estimate carried out over 1976-2002

Hypothesized Rank*	Trace Statistic	5 Percents Critical Value
None	93.31	68.52
1	52.43	47.21
2	**27.04**	**29.68**
3	9.05	15.41
4	0.81	3.76

b- Estimate carried out over 1985-2002

Hypothesized Rank*	Trace Statistic	5 Percents Critical Value
None	104.87	68.52
1	54.03	47.21
2	**27.88**	**29.68**
3	9.36	15.41
4	0.10	3.76

c- Estimate carried out over 1986-2002

Hypothesized Rank*	Trace Statistic	5 Percents Critical Value
None	87.72	68.52
1	48.41	47.21
2	**28.86**	**29.68**
3	9.85	15.41
4	0.07	3.76

* in bold we accept the rank at the 5% level

A.2 Short-run coefficients $[\Delta p,\ p-\widetilde{ulc}, p-\ p_m, U, C]$

A.2.1 Model is estimated over 1985Q1-2002Q4 period

Dependent Variable	$\Delta^2 p$		$\Delta(\ p-\widetilde{ulc})$	
markup relation	-0.969	[-3.77]	-1.552	[-2.69]
Is equation	0.057	[1.03]	0.350	[2.81]
$\Delta^2 p_{-1}$	-0.658	[-4.75]	-1.080	[-3.47]
$\Delta^2 p_{-2}$	-0.497	[-5.68]	-0.252	[-1.28]
$\Delta(\ p-\widetilde{ulc})_{-1}$	-0.044	[-0.79]	0.546	[4.35]
$\Delta(\ p-\widetilde{ulc})_{-2}$	0.003	[0.06]	-0.401	[-3.20]
$\Delta(\ p-p_m)_{-1}$	-0.012	[-0.44]	-0.085	[-1.42]
$\Delta(\ p-p_m)_{-2}$	-0.073	[-3.38]	-0.078	[-1.61]
ΔU_{-1}	-0.122	[-0.71]	-0.232	[-0.60]
ΔU_{-2}	-0.337	[-1.83]	0.326	[0.79]
ΔC_{-1}	0.075	[1.20]	0.143	[1.03]
ΔC_{-2}	-0.089	[-1.36]	-0.124	[-0.84]
cst	0.000	[1.18]	0.001	[2.00]
Δoil_{-1}	0.004	[2.83]	-0.000	[-0.14]
Δoil_{-2}	-0.005	[-3.30]	-0.006	[-1.53]
Adj. R^2	0.76		0.67	

Test for normality (Lutkepohl):
$\chi^2(10)$=20.42, p-value =0.03.
Tests for serial correlation:
LM(1),$\chi^2(25)$=33.69, p-value=0.11,
LM(2),$\chi^2(25)$=27.46, p-value=0.3,
LM(4),$\chi^2(25)$=27.12, p-value=0.35.

A.2.2 Model is estimated over 1986Q1-2002Q4 period

Dependent Variable	$\Delta^2 p$		$\Delta(p - \widetilde{ulc})$	
markup relation	-0.884	[-2.18]	-1.626	[-2.09]
Is equation	0.018	[0.27]	0.267	[2.08]
$\Delta^2 p_{-1}$	-0.491	[-2.52]	-0.943	[-2.52]
$\Delta^2 p_{-2}$	-0.557	[-5.01]	-0.407	[-1.91]
$\Delta(p - \widetilde{ulc})_{-1}$	-0.032	[-0.48]	0.531	[4.20]
$\Delta(p - \widetilde{ulc})_{-2}$	-0.011	[-0.17]	-0.309	[-2.56]
$\Delta(p - p_m)_{-1}$	-0.037	[-1.27]	-0.128	[-2.29]
$\Delta(p - p_m)_{-2}$	-0.037	[-1.55]	-0.063	[-1.37]
ΔU_{-1}	-0.070	[-0.36]	-0.220	[-0.59]
ΔU_{-2}	-0.346	[-1.63]	0.355	[0.87]
ΔC_{-1}	0.078	[1.12]	0.108	[0.81]
ΔC_{-2}	-0.133	[-1.67]	-0.212	[-1.38]
cste	0.000	[1.01]	0.001	[2.50]
Δoil_{-1}	0.005	[2.50]	0.001	[0.36]
Adj. R^2	0.72		0.66	

Test for normality (Lutkepohl):
$\chi^2(10)$=16.40, p-value=0.09.
Tests for serial correlation:
LM(1),$\chi^2(25)$=28.33, p-value=0.29,
LM(2),$\chi^2(25)$=26.59, p-value=0.38,
LM(4),$\chi^2(25)$=24.66, p-value=0.48.

A.2.3 Model is estimated over 1976Q1-2002Q4 period

Dependent Variable	$\Delta^2 p$		$\Delta(p - \widetilde{ulc})$	
markup relation	-0.448	[-4.35]	-0.590	[-3.51]
Is equation	0.046	[1.44]	0.245	[4.74]
$\Delta^2 p_{-1}$	-0.444	[-3.90]	-0.430	[-2.30]
$\Delta^2 p_{-2}$	-0.388	[-4.33]	-0.199	[-1.36]
$\Delta(p - \widetilde{ulc})_{-1}$	-0.093	[-1.59]	0.357	[3.75]
$\Delta(p - \widetilde{ulc})_{-2}$	0.108	[1.79]	-0.263	[-2.66]
$\Delta(p - p_m)_{-1}$	-0.058	[-3.03]	-0.103	[-3.32]
$\Delta(p - p_m)_{-2}$	0.006	[0.35]	-0.076	[-2.53]
ΔU_{-1}	0.044	[0.23]	-0.128	[-0.41]
ΔU_{-2}	0.100	[0.53]	1.110	[3.58]
ΔC_{-1}	0.095	[1.59]	-0.037	[-0.38]
ΔC_{-2}	-0.009	[-0.17]	-0.057	[-0.63]
cst	0.000	[-1.05]	0.000	[-0.61]
Δoil	0.006	[4.47]	0.004	[1.89]
Adj. R^2	0.54		0.54	

Test for normality (Lutkepohl):
$\chi^2(10)$=15.21, p-value=0.12.
Tests for serial correlation:
LM(1),$\chi^2(25)$=24.81, p-value=0.47,
LM(2),$\chi^2(25)$=30.92, p-value=0.19,
LM(4),$\chi^2(25)$=16.86, p-value=0.89.

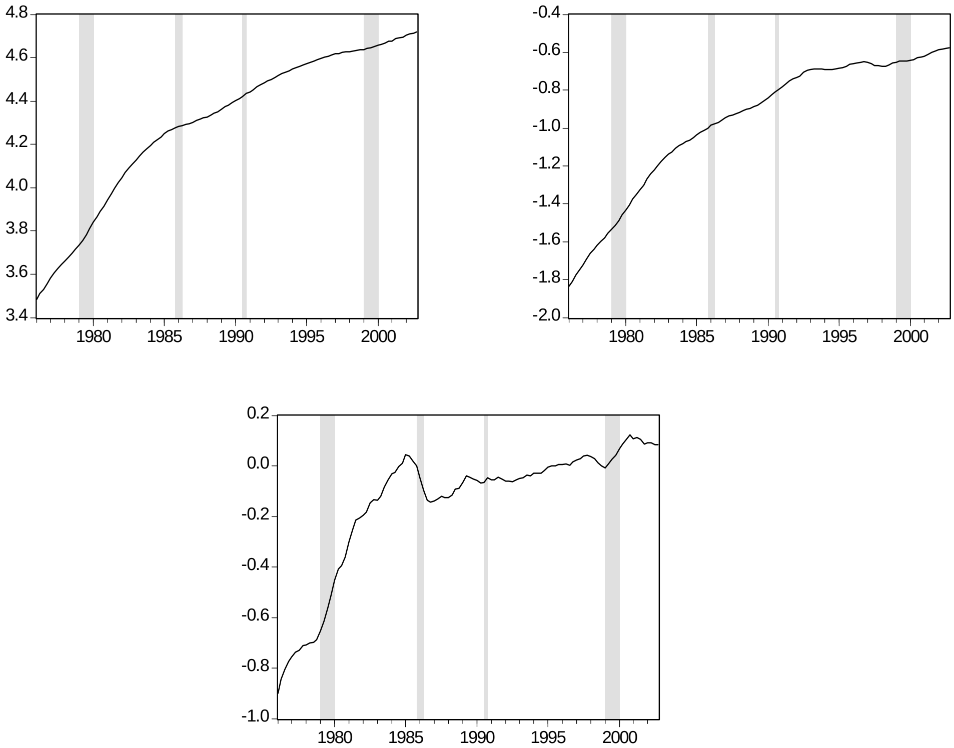
B Figures

B.1 Variables of the system[16]

Chart B.1.1: Endogenous and exogenous series since 1976

a - The I(2) variables expressed in log.

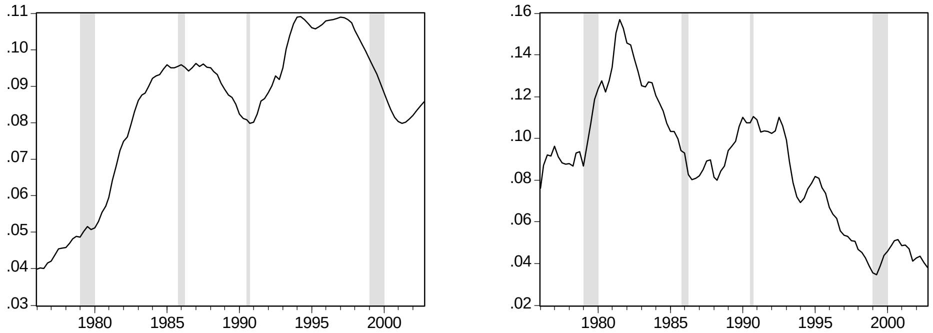
$$\begin{bmatrix} p & \widetilde{ulc} \\ \multicolumn{2}{c}{p_m} \end{bmatrix}$$



[16] Shaded areas match oil-shock periods.

b - The I(1) variables.

$$\begin{bmatrix} U & C \end{bmatrix}$$



c- The transformed variables and the exogenous oil price

$$\begin{bmatrix} \Delta p & oil\ in\ euro \\ p - \widetilde{ulc} & p - p_m \end{bmatrix}$$

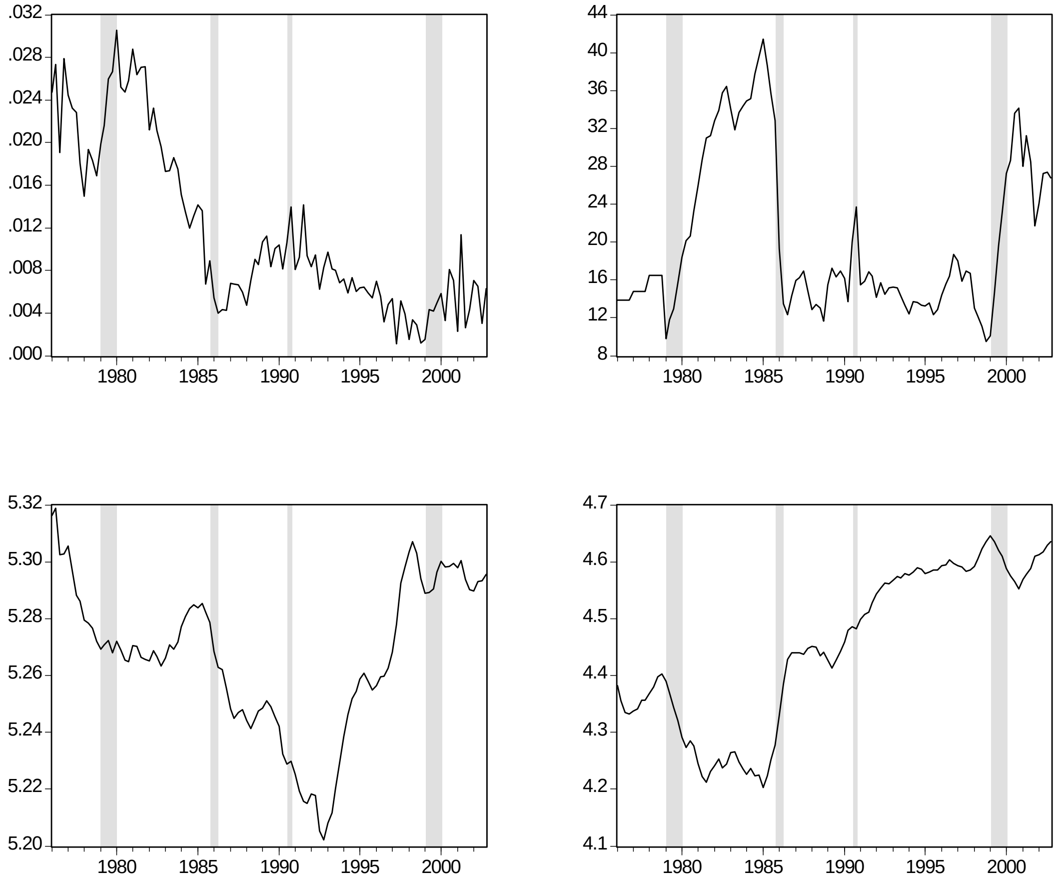
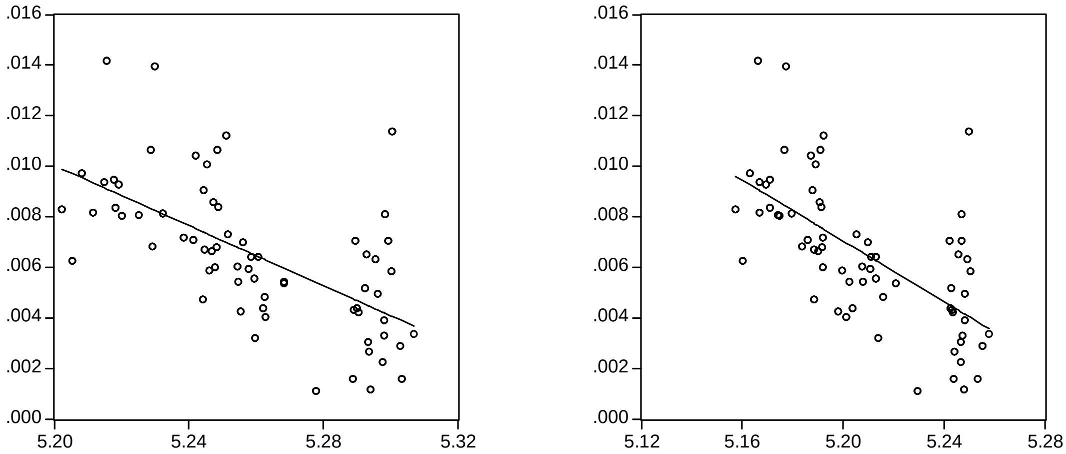


Chart B.1.2:Inflation and markup, scatter plots over 1986 to 2002

markup on unit labour cost $(p - \widetilde{ulc})$ *Estimated markup* $(p - \theta\widetilde{ulc} - (1-\theta)p_m)$
$\widehat{\theta}_{86} = 0.93$



B.2 Stability over time

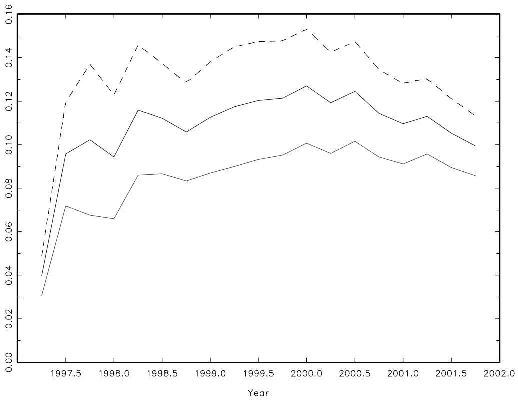
The following charts display the β and α coefficients for the price and markup equations when the systems are estimated recursively on a sample period progressively extended to include the years 1997 to 2001, in order to provide rolling one-year ahead forecasts. The charts are ordered by type of long-run relations (markup in B.2.1 and IS in B.2.2) and type of equation (price $\Delta^2 p$ or markup $\Delta(p - ulc)$). For example, Appendix B.2.1 describes the markup coefficients, starting with the model with unemployment, presenting the long-run coefficients β_1 to β_3, and the α_1 coefficients for the price equation $\Delta^2 p$ and the markup equation $p - ulc$.

B.2.1 The markup-type relation, $\Delta p = -\beta_1(p - \widetilde{ulc}) - \beta_2(p - p_m) - \beta_3 U$

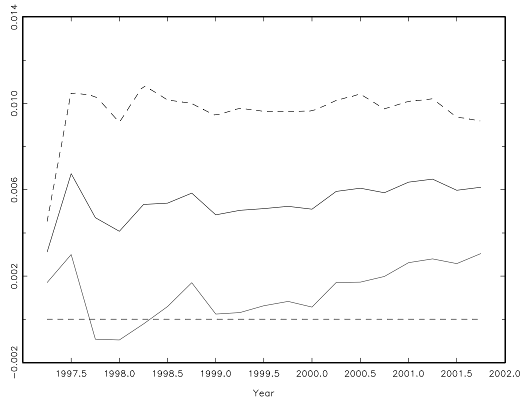
Chart B.2.1 : First long-run relation

a- Long-run coefficients

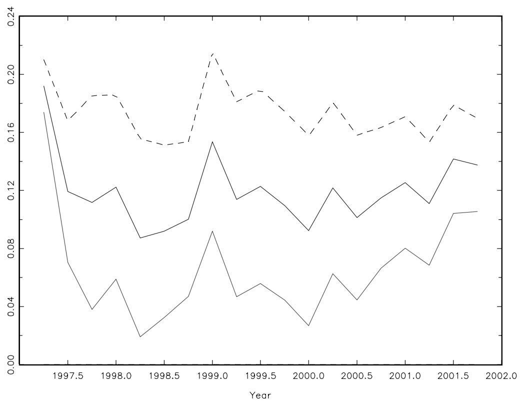
$p - \widetilde{ulc}$ (β_1)



$p - p_m$ (β_2)

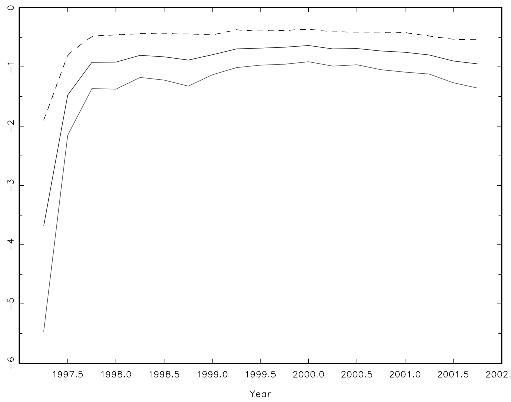


U (β_3)

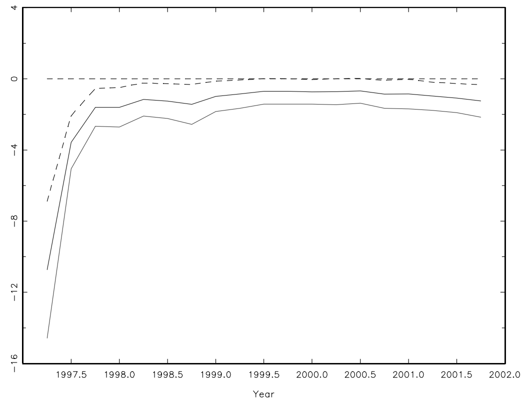


b- Short-run coefficients

$\Delta^2 p$



$\Delta(p - \widetilde{ulc})$

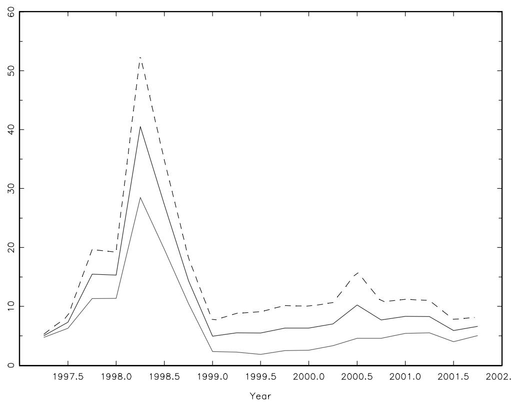


B.2.2 The IS-type equation $U = -\beta_4 \Delta p - \beta_5 C$

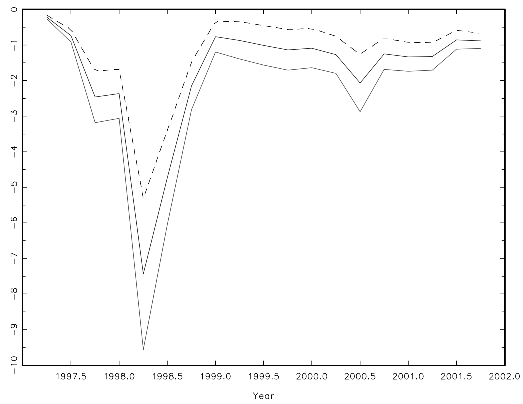
Chart B.2.2.: Second long-run relation

a- Long-run coefficients

$\Delta p\ (\beta_4)$

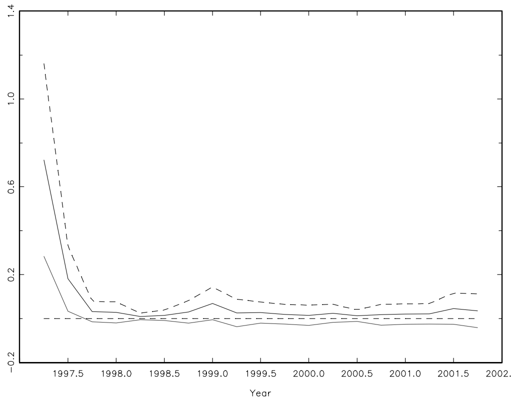


$C\ (\beta_5)$

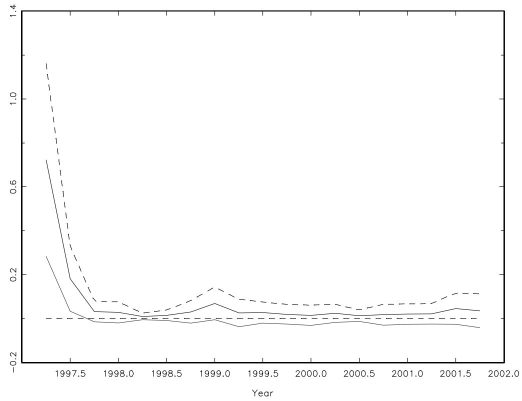


b- Short-run coefficients

$\Delta^2 p$



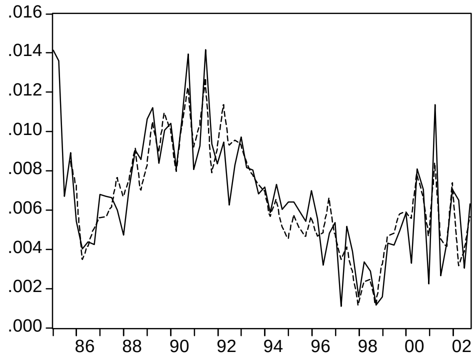
$\Delta(p - \widetilde{ulc})$



B.3 Forecasts as from 1998 onwards.

Chart B.3.1: In-sample simulated inflation

a- Model is estimated over 1985-2002



b- Model is estimated over 1976-2002

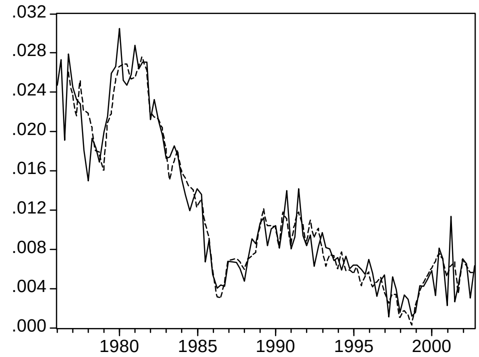
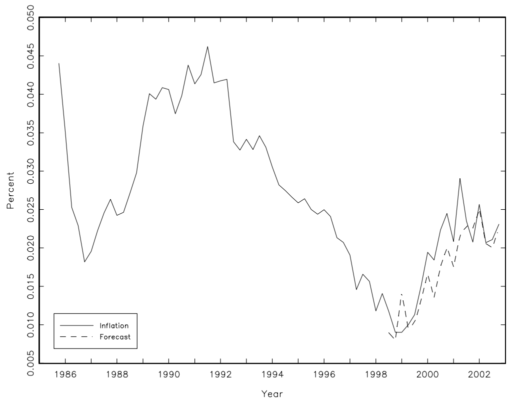
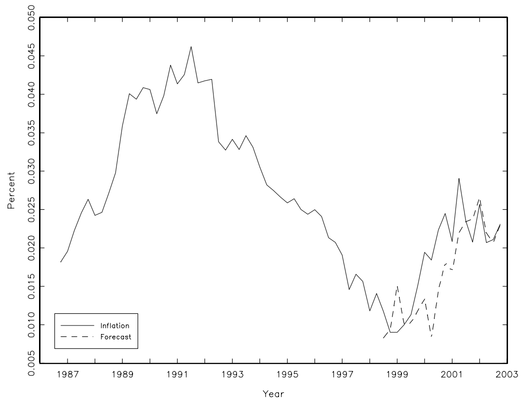


Chart B.3.2: One-year ahead forecasts*

a- Rolling estimation since 1985 [0.28]



b- Rolling estimation since 1986 [0.42]



* Numbers in brackets are RMSEs.

Chart B.3.3: Comparison of forecasting performances

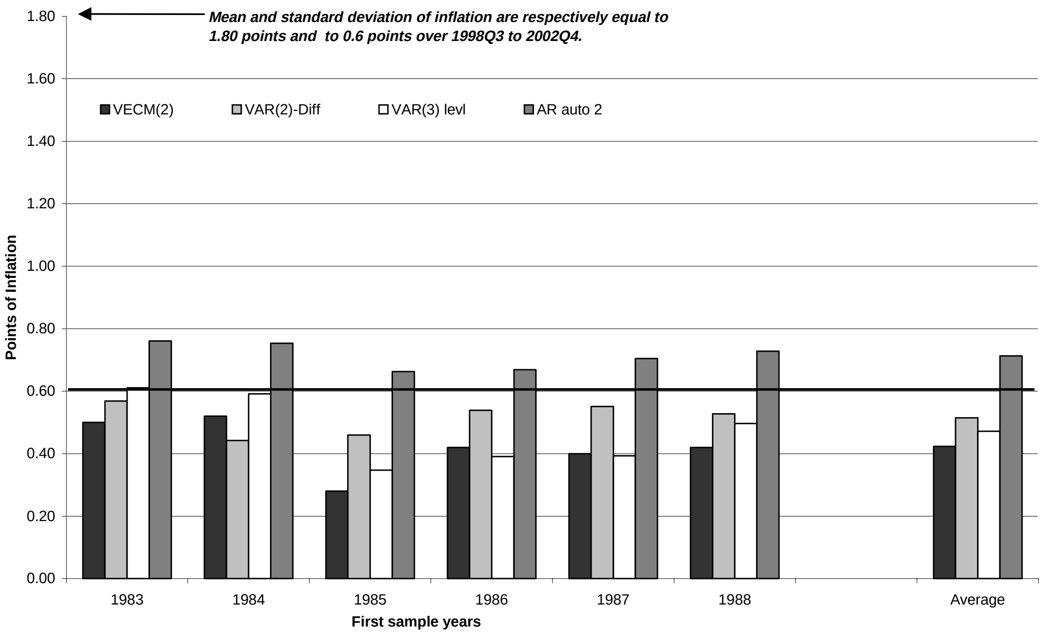


Table B.3.2: Root mean square error*

	VECM(2)	*VAR(2)-Diff*	*VAR(3) levl*	*AR (AIC)*
1983	0.50	0.57	0.61	0.76
1984	0.52	0.44	0.59	0.75
1985	0.28	0.46	0.35	0.66
1986	0.42	0.54	0.39	0.67
1987	0.40	0.55	0.39	0.70
1988	0.42	0.53	0.50	0.73
Average	0.42	0.51	0.47	0.71

*RMSEs are computed from out-sample forecast over the sample 1998Q3 to 2002Q4

C A simple price-setting schedule

The price setting equation (3) follows from Layard, Nickell and Jackman (1992). We suppose the economy is made of identical firms labelled i, with a technology using only labour ($\alpha > 0$). Firm i maximises profit defined as:

$$
\begin{aligned}
Max(L_i) &= P_i Y_i - W L_i && (19)\\
Y_i &= A L_i^{\alpha} && (20)\\
Y_i &= (\frac{P_i}{P})^{-\eta} && (21)
\end{aligned}
$$

with P_i is the price charged on the output of firm i, Y_i its supply of goods. $\eta > 1$. $\kappa = 1 - \frac{1}{\eta}$ is an index of product-market competitiveness: $\eta = 1$ or $\kappa = 0$ implies monopoly, while perfect competition occurs when $\eta = \infty$ or $\kappa = 1$.

The first order condition of programme (19)-(21) implies that:

$$[Y_i \frac{\partial P_i}{\partial Y_i} \frac{\partial Y_i}{\partial L_i} + P_i \frac{\partial Y_i}{\partial L_i}] - W = 0 \tag{22}$$

using $\frac{\partial LogP_i}{\partial LogY_i} = -\frac{1}{\eta}$, one gets:

$$L_i = \kappa \alpha \frac{P_i Y_i}{W} \tag{23}$$

Assuming the equilibrium is symmetric ($P = P_i, Y = Y_i, L = L_i$) and introducing the unemployment rate $U = 1 - \frac{L}{N}$, where N is total labour force:

$$1 - U = \alpha\kappa \frac{PY}{WN} \tag{24}$$

Taking logarithms on both sides, one gets (lower case variables are in logs):

$$p = w - Log(\frac{Y}{N}) - U - Log(\alpha\kappa) \tag{25}$$

where $Log(\frac{Y}{N})$ is labour productivity. Notice that, in such an equation, demand intervenes directly through U. In addition, the markup $Log(\kappa)$ depends on the business cycle as well as the level of inflation. An increase in competition implies that κ increases, hence the markup on unit labour costs is reduced.

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