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In 2004 all ECB publications will feature a motif taken from the €100 banknote.

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Abstract

Equilibrium correction models of the price level are often used to model inflation. Such models assume that the long-run markup of prices over costs is fixed, but this may not be true for the Euro area economy, which has undergone major structural reforms over the last 25 years. We allow for shifts in the markup factor through estimating an equation that includes a time-varying intercept. The model fits the data better than a linear alternative, and suggests that a reduction in the price-cost markup contributed to disinflation in the Euro area during the 1980s.

Keywords: *inflation, price-cost markup, cointegration, time-varying intercept, dynamic modelling.*

JEL classification: *C22,C32,E31*

Non-technical summary

This paper analyses inflation dynamics in the Euro area using a markup model of the price level. In such models the target price level is set as a markup on some combination of input prices, and fluctuations in inflation are then interpreted as partial adjustment of the price level towards that target, i.e. inflation is an equilibrium correcting process. This treatment of price adjustment is central to the modern literature on inflation, see, for instance, de Brouwer and Ericsson (1998). In the empirical analysis of the Euro area economy, markup models of inflation have been used in the Area Wide Model (AWM) developed by Fagan et al. (2001).

An important, but rarely challenged, assumption in this literature is that the target markup of prices over costs is fixed. We argue that in the case of the Euro area economy this is unlikely to be true, e.g. because greater product market competition may have decreased the percentage markup that firms can sustain in the long-run. As a result, a simple equilibrium correction model of the price level would fail to account for inflation adjustment that occurs following a reduction in the share of profits in total output.

One response to the problems posed by structural change is to generalise the equilibrium correction equation to include a time-varying intercept that is estimated jointly with the unknown coefficients of the model. If the estimation does not detect any time variation in the intercept then there is no evidence that breaks in the price-cost markup have affected inflation, and the linear equation is adequate. However, if the time-varying intercept turns out to be statistically significant then it follows that the equilibrium price level has changed independently of input costs, and that the target markup factor has shifted.

In the empirical analysis presented in this paper we use a cointegrated vector autoregression (VAR) to show that the Euro area price level can be modelled as a markup on unit labour costs, raw material prices and the tax wedge. The price level equation also includes a time trend in order to account for possible measurement errors in the data. Deviations from the long-run solution for the price level are used to explain the inflation rate in a model that also includes the output gap and dynamic adjustment terms. This model is then extended to include a time-varying intercept. It turns out that the time-varying intercept is significant at the 10% level, and helps to reduce the equation standard error by approximately one tenth. A graph of the estimated intercept term indicates that a reduction in the price-cost markup made an important contribution to the large disinflation that took place in the Euro area in the early 1980s. A shift in the price-cost markup is particularly likely to have occurred during the 1980s, as competition between suppliers in different Euro area countries increased following the removal of exchange rate uncertainty as part of the fixed exchange rate system introduced by the new European Monetary System (EMS) in 1979.

A comparison of the time-varying intercept model and the linear inflation equation is provided. The autoregressive coefficient in the time-varying intercept equation is smaller than that in the linear equation, suggesting that inflation persistence is less important after controlling for the effects of structural change. Secondly, we show that the right-handside of the equation that incorporates a time-varying intercept is more obviously stationary than the right-handside of the linear inflation equation. This is interpreted as a sign that changes in the price-cost markup induce additional persistence in the deviation of prices from costs, and that the time-varying intercept helps to control for such distortions.

1 Introduction

The introduction of the European single currency in January 1999 saw the responsibility for setting interest rates across the Euro area conferred upon the European Central Bank (ECB), which is charged with the task of achieving price stability across the region as a whole.[1] This restructuring of monetary policy institutions within Europe has obviously increased the need for empirical models of macroeconomic fluctuations in the Euro area. The study of inflation dynamics is a particularly important research topic given the central role of the inflation target in the constitution of the ECB, and there are now several papers in the literature that have applied well known techniques for the analysis of inflation to Euro area data, see Jansen (2004) for an overview. In this paper we contribute to that literature by estimating a model that clarifies the relative importance of the channels through which Euro area inflation adjustment has taken place.

The starting point for our analysis is a simple model in which the price level is set as a markup on input costs. As this is an equilibrium relationship, deviations from it can be used to explain variation in the inflation rate. In addition to this simple equilibrium correction relationship, the model conditions on a measure of the output gap, which controls for cyclical influences on firms' price-setting decisions. The novel feature of our analysis is the inclusion of a time-varying intercept in the inflation equation. This controls for influences on inflation that arise from economic events that are difficult to observe, e.g. shifts in the target markup of prices over costs caused by the process of economic and monetary integration within Europe.

The empirical methodology that we follow can be divided into three parts. In the first part we examine various steady-state representations of the price level using a Vector Equilibrium Correction framework, and an information set comprising unit labour costs, the import price deflator, a world market commodity price index and a measure of the tax wedge. We show that for the sample period 1980q4-2000q4 the price level, after correcting for the tax wedge and a time trend, can be expressed as a linearly homogeneous function of unit labour costs and commodity prices. In the second part of the analysis we use this representation in order to formulate a conditional model for inflation in which relative price measures are the main forcing variables. This specification includes very general dynamics and is tested down to a parsimonious form through deleting the least significant term, then re-estimating the model and then repeating the process until all terms remaining in the model are significant. This procedure yields a parsimonious equation, in which two relative price measures and the output gap exert powerful effects on the inflation rate. However, there is some evidence that the model is not congruent to the data generating process (DGP) for inflation, e.g. a test for the validity of the linear functional form results in a rejection. These test outcomes may be due to the fact that the model does not control for shifts in the price-cost markup that have occurred over the past quarter of a century as a result of structural change in the Euro area economy.

In the final part of the analysis we address this possibility by introducing a time-varying intercept, or local level, to the inflation equation. This is estimated as an unobserved component using the STAMP package of Koopman et al. (2000)

[1] See e.g. Coenen and Vega (2001) for a discussion of the monetary policy strategy of the ECB.

and is intended to capture the impact on inflation of shifts in the target markup of prices over costs. The extended model passes residual diagnostic tests, suggesting that the local level technique handles the effects of a time-varying markup factor, and effectively 'balances' inflation equations in which there are apparent, or real, non-stationarities due to the effects of structural change. The intercept drifts down over the first third of the sample, signalling a reduction in the target markup of prices over costs. This trend can be interpreted as the disinflationary effect of tougher price competition in the Euro area due to growth in trade following the creation of a new European Monetary System in 1979.

The remainder of the paper expands upon these points and is structured as follows. In Section 2 we set out the precise details of the empirical methods that we follow. Section 3 presents the data used for the analysis. Section 4 reports models for the Euro area price level and inflation rate, and, in light of the results, considers the importance of different channels for inflation adjustment. Section 5 rounds off the paper with a summary of the main arguments.

2 A Framework for Modelling Inflation

The approach to modelling inflation that we follow in this paper assumes that the equilibrium price level is set as a markup on some combination of input prices. Fluctuations in the inflation rate are then interpreted as partial adjustment of the actual price level towards a target value defined by the markup equation, i.e. inflation is an equilibrium correcting process. This treatment of price dynamics is at the heart of the modern approach to modelling inflation, both for forecasting and for policy analysis, see de Brouwer and Ericsson (1998), Aron and Muellbauer (2000), Banerjee and Russell (2001), Bank of England (1999), and Hendry (2001). In the empirical analysis of the Euro area economy, models of the inflation process based upon a markup equation for the price level have been used in the Area Wide Model (AWM) developed by Fagan et al. (2001).

A Model for the Price Level

In outlining the main features of a markup model of the price level we follow the exposition in de Brouwer and Ericsson (1998), though we amend their treatment slightly in order to accommodate the information set used in our empirical analysis.[2] A markup formula for the price level at time t can be written as follows:

$$P_t = \Psi ULC_t^{\kappa} PM_t^{\beta} PCOM_t^{\gamma} TAX_t^{\lambda} e^{\varphi trend} \tag{1}$$

where P represents the price level, ULC is the level of unit labour costs (defined as wages and salaries paid for each unit of production), PM is the level of import prices, $PCOM$ is the domestic currency price of oil and other raw materials, TAX is the percentage tax wedge and $trend$ is a time trend. Each of the time series referred to is in index form and the base period is common across series. Equation (1) states that the price level is obtained as a constant markup, Ψ, on a geometric weighting of

[2] Whilst the framework that we describe is most closely related to that in de Brouwer and Ericsson (1998), we should emphasise that its key features are central to many studies of the inflation process, see, for example, Banerjee and Russell (2001).

the three input cost indices and the tax wedge. This representation of the price level can be derived from a profit maximising model in which the production technology is Cobb-Douglas. We follow Hendry (2001) in adding a time trend to the model in order to control for measurement errors that may affect the basic markup relationship - these measurement errors may arise due to the exclusion of some component of wages and salaries from the unit labour cost index, e.g. the compensation of the self-employed.

The log-linear representation of (1) is as follows:

$$p_t = \log \Psi + \kappa ulc_t + \beta pm_t + \gamma pcom_t + \lambda tax_t + \varphi trend \tag{2}$$

where lower case variables represent logs of corresponding upper case variables. We allow the variables in (2) to be integrated of, at most, first order, I(1),[3] and therefore estimate the equation as the cointegrating relation within a Vector Autoregression (VAR) using the reduced rank regression technique due to Johansen (1988).[4] We then use the results to test two hypotheses. The first is that $\lambda = 1$, which implies that shifts in the tax wedge are passed on in full to the general level of prices. The second is that the price level is linearly homogeneous in unit labour costs, import prices and commodity prices, which corresponds to the restriction $\kappa + \beta + \gamma = 1$.

If both restrictions are found to be compatible with the data then deviations of the price level from steady-state can be written as the sum of a constant and a weighted average of three relative price terms, i.e.

$$\begin{aligned} \textit{deviation of price level from steady-state} \quad = \quad & \log \Psi + \kappa\left(ulc - p + tax + \varphi trend\right) \\ & +\beta\left(pm - p + tax + \varphi trend\right) \\ & +\gamma\left(pcom - p + tax + \varphi trend\right) \end{aligned} \tag{3}$$

These relative price terms measure the distance between the price level at time t and its steady-state value and hence define the scope for equilibrium correction effects to set the inflation rate.

A Single Equation Model for Inflation

In the second stage of the empirical analysis we move from the VAR to a single equation model for inflation. This reduction is shown to be compatible with the data and allows for a more detailed study of the dynamics of the inflation process. As there may be partial adjustment in the price level we consider an equation that allows for relative price effects on the inflation rate at lags ranging from one to four quarters. The model also permits transitory variations in the target markup of prices over costs through including a distributed lag in the output gap. This term, denoted

[3]The possibility of some variables being I(2) is further discussed in Section 4.1. The order of integration - either I(0) or I(1) - of each of the variables used in the analysis is indicated by the rank of the matrix of coefficients multiplying the explanatory variables in the Johansen framework, see Johansen (1988), and need not be tested for prior to the estimation.

[4]The model is not a pure VAR as we condition the analysis on the tax wedge. The critical values used for inference are adjusted to take account of this.

gap, measures the log ratio of actual GDP to its trend (permanent) value - see section 3 for details concerning the measurement of trend GDP.[5] The model excludes all t dated regressors in order to ensure that the set of explanatory variables is predetermined and that the coefficients are estimated free from endogeneity biases.[6] The unrestricted version of the Equilibrium Correction Model that we estimate in the second stage of our empirical analysis is therefore of the form:

$$\begin{aligned}\Delta p_t &= \log\Psi + \sum_{q=1}^{3}\varsigma_q gap_{t-q} + \sum_{i=1}^{3}\theta_i \Delta p_{t-i} + \sum_{j=1}^{3}\kappa_j \Delta ulc_{t-j} + \sum_{k=1}^{3}\beta_k \Delta pm_{t-k} \qquad (4)\\ &+ \sum_{l=1}^{3}\gamma_l \Delta pcom_{t-l} + \sum_{p=1}^{3}\lambda_p \Delta tax_{t-p} + \kappa^*(ulc - p + tax + \varphi trend)_{t-4}\\ &+\beta^*(pm - p + tax + \varphi trend)_{t-4} + \gamma^*(pcom - p + tax + \varphi trend)_{t-4}\end{aligned}$$

where $(\kappa^*, \beta^*, \gamma^*)$ equals (κ, β, γ) multiplied by α, the speed at which the price level converges on its equilibrium value (recall that $\kappa + \beta + \gamma = 1$). Note that although the levels terms in (4) enter at the fourth lag, any dating between $t-1$ and $t-4$ is possible because changing the dating only implies a reparametrisation of the short run coefficients, see Bårdsen (1989).

This version of the equilibrium correction model was tested down to a parsimonious form using a general-to-specific modelling strategy that entailed deleting from the initial regression estimate the least significant term, then re-estimating the model, and then repeating the procedure until each of the variables included in the regression were individually significant. For a given specification, further reductions in the parameter space could be achieved by changing the dating of the relative price effects. For instance, if one could not reject the hypothesis

$$\kappa_j = \lambda_p = -\theta_i = \kappa^* \qquad j, p, i = 1, 2, 3$$

then it was possible to use the fact that

[5] In the New Keynesian literature on inflation the output gap and real marginal unit labour costs are viewed as *alternative* control variables, see e.g. Galí and Gertler (1999). The idea is that real marginal unit labour costs behave cyclically, e.g. because overtime hours have to be paid at a premium rate during economic expansions, such that there is no need to control for the output gap. Equation (4) includes real average unit labour costs in the form of the first relative price term (with tax and trend corrections). As this series does not exhibit the same cyclicality as real marginal unit labour costs, we include in (4) an extra variable relative to the New Keynesian model, namely the output gap.

[6] We note that this parameter restriction rules out the logical possibility of perfect price flexibility. Whilst there is no prior justification for this restriction, the results reported later in this paper show that the coefficients on the dynamic adjustment terms dated $t-1$ through $t-3$ are relatively small (the tax term excluded). One would expect them to be large if contemporaneous adjustment terms had been erroneously excluded from the model, and we therefore conclude that the validity of our econometric specifications is not affected by this step.

$$
\begin{aligned}
(ulc - p + tax + \varphi trend)_{t-1} \quad &= \quad constant + \sum_{j=1}^{3} \Delta ulc_{t-j} + \sum_{p=1}^{3} \Delta tax_{t-p} \\
&\quad - \sum_{i=1}^{3} \Delta p_{t-i} + (ulc - p + tax + \varphi trend)_{t-4}
\end{aligned}
$$

in order to write the model as one in which disequilibrium in unit labour costs relative to the tax adjusted price level feeds into the inflation rate with a lag of one quarter rather than four quarters. This approach to dating the relative price terms allows for the possibility that disequilibrium in prices relative to unit labour costs induces a change in inflation with a different lag than does disequilibrium in prices relative to the cost of material inputs.

An Equilibrium Correction Model Augmented with an Unobserved Component

The use of (4) as a model for Euro area inflation may be problematic because the assumption of a fixed equilibrium price-cost markup factor, Ψ, is inappropriate for an economy that has experienced significant structural change over the past quarter of a century. Examples of these structural changes include the introduction of a fixed exchange rate regime in 1979 as part of the new European Monetary System (EMS), the marketisation of several southern European countries in preparation for accession to the European Union in 1986, and also the emergence of more rigorous competition policy as part of European Union law. These reforms are generally thought to have increased product market competition in the Euro area through stimulating intra-European trade, see Fontagne et al. (1998) for some empirical evidence on this issue. More competitive product markets are likely to have caused a reduction in the excess profits earned by suppliers, i.e. the Ψ factor may have drifted down over time (see Griffith (2001) for evidence of a negative relationship between product market competition and excess profits based on a panel of UK firms). If such a shift in the structure of price-setting has occurred in the Euro area during the post-1980 period, then we should expect to observe a downward trend in Euro area inflation beyond that which can be explained by the markup factor and the other explanatory variables in (4).

The fact that there are several prior reasons for expecting a time-varying price-cost markup in the Euro area suggests that (4) is likely to prove an inadequate model for inflation. At an economic level, the model will not provide any measure of the extent to which inflation adjustment occurs through changes in the share of profits in total output. At a statistical level, failure to model time-variation in the markup implies mis-measurement of the equilibrium factors in (4), which in turn means that the estimated equilibrium correction speeds may be biased.

These considerations point towards the need for some generalisation of the linear inflation equation. One option would be to construct a measure of the aggregate profit share and then condition on that variable in the empirical analysis. However, obtaining the relevant data is likely to prove a particularly demanding exercise. Alternatively, dummy variables that allow for a step shift in the relationship between inflation and the markup could be incorporated into the analysis. However, changes in competitive forces are likely to diffuse through the economy at a gradual pace, suggesting that non-stationarities in the relationship between inflation and the markup are not best modelled in this way. Instead, we adopt an approach closely related to that followed by Aron and Muellbauer (2000), which involves augmenting (2) with a time-varying intercept, or local level.[7] This is an unobserved component that is jointly estimated with the inflation equation. Its evolution is determined by the timing and magnitude of large shifts in the relationship between inflation and the markup, and it is therefore well suited to modelling relationships in which the precise dating of breaks is unknown. Further (as we discuss below) it is constructed as an I(1) process and therefore embodies the smoothness that one would associate with evolutionary changes in aggregate price-setting behaviour.

The Equilibrium Correction Mechanism extended to include a local level term, and also the error term that was not included in previous specifications, is as follows:

$$\begin{aligned}
\Delta p_t \;=\; & \sum_{q=1}^{3} \varsigma_q gap_{t-q} + \sum_{i=1}^{3} \theta_i \Delta p_{t-i} + \sum_{j=1}^{3} \kappa_j \Delta ulc_{t-j} + \sum_{k=1}^{3} \beta_k \Delta pm_{t-k} \\
& + \sum_{l=1}^{3} \gamma_l \Delta pcom_{t-l} + \sum_{p=1}^{3} \lambda_p \Delta tax_{t-p} \\
& + \alpha * \kappa (ulc - p + tax + \varphi trend)_{t-4} \\
& + \alpha * \beta (pm - p + tax + \varphi trend)_{t-4} \\
& + \alpha * \gamma (pcom - p + tax + \varphi trend)_{t-4} + \psi \mu_t + u_t \qquad (5)
\end{aligned}$$

$$(\kappa + \beta + \gamma = 1)$$
$$u_t \sim (0, \sigma_u^2)$$

$$\mu_t = \mu_{t-1} + \varepsilon_t, \varepsilon_t \sim (0, \sigma_\varepsilon^2) \qquad (6)$$

The local level component of the *measurement equation* in (5), μ_t, is constructed as an I(1) process using the STAMP package of Koopman et al. (2000). The evolution of μ_t is determined by the variance of the shocks that enter the *transition equation* in (6), σ_ε^2. The measurement and transition equations together constitute the *state space form* (SSF) of the model. The estimation of such a model proceeds in two steps. First, it is postulated that the two error terms, u_t and ε_t, are normally and independently distributed, so that maximum likelihood estimates of σ_u^2 and σ_ε^2 can be computed using numerical optimisation techniques. The coefficients of the measurement equation can then be retrieved using the Kalman filter, which performs the same function in the estimation of models in SSF as do least squares computations in the estimation of a standard regression model.

The local level estimation method is applied to the restricted version of the linear Equilibrium Correction Model that was obtained by testing down from the general specification in (4). We then check the validity of the model reductions

[7] In the remainder of the paper the terms 'time-varying intercept' and 'local level' are used interchangeably.

implemented in the linear case by adding to the specification a distributed lag in each of the variables, one variable at a time, and testing down once more. We opt for this strategy rather than repeating the entire general-to-specific modelling exercise because we find that highly parameterised model specifications tend to be associated with implausibly large variation in the local level term. The strategy that we follow, namely developing the best fitting linear model and then studying the properties of a non-linear version of that restricted specification, is consistent with the approach to non-linear modelling advocated by Teräsvirta (1998).

The fact that the local level term is constructed as an I(1) process raises some questions concerning the cointegrating properties of the model for inflation. In particular, it is unclear whether or not the equation is balanced in the sense that the variables on the left-handside and right-handside are integrated of the same order, see Granger (1990). The usual presumption is that the price level is I(1), or possibly I(1) with breaks in the mean, and cointegrates with measures of input costs such that any variation in the markup (which is equivalent to the sum of the three relative price terms in (3)) is stationary, or I(0). This I(0) term then explains the stationary variation in Δp_t, i.e. the model is balanced. The inclusion of the local level term complicates the picture, in that the right-handside of (5) must then be I(1), which is not balanced by the stationary inflation series on the left-handside of (5). We interpret the relationship between inflation and the markup as one in I(0) space that is subject to deterministic shifts that take effect gradually. In finite samples, these effects can be modelled as a low variance I(1) process (to see this point note that in the limiting case in which σ_ε^2 is set to zero, the local level term is simply a constant and is therefore obviously stationary). Thus, while (5) cannot represent the actual data generating process under the assumption that price inflation is stationary, it is a plausible model in finite samples, and, crucially, allows the relationship between inflation and the markup to be estimated free from the biases that would arise were we not to control for unobserved shifts in that relationship.[8]

3 The Data

The variables used in the empirical analysis are all measured at the quarterly frequency and expressed in seasonally adjusted form. Each variable is constructed as a weighted average of the corresponding national series, with the weight accorded to each country set equal to its share in constant 1995 Euro area GDP, measured at market prices using purchasing power parity (PPP) exchange rates, see Fagan et al. (2001) for further details.

The price level, P_t, is measured by the consumer expenditure deflator. ULC_t, is defined as whole economy wages and salaries paid per unit of GDP at factor cost

[8] A second logical possibility, which we find less plausible *a priori,* is that the price-cost markup is I(1), in which case the linear Equilibrium Correction Model in (4) is certainly unbalanced, in that the model would entail regressing an I(0) inflation rate on I(1) relative prices. If the local level term cointegrates with relative prices so that we obtain a stationary righthandside to equation (5), which then explains the stationary variation in the inflation rate, a balanced regression obtains (this point has also been made, albeit in a slightly different context, by Naug (2000)). In such instances the interpretation of the local level term is that it measures the I(1) contribution to inflation arising from permanent shifts in the price-cost markup.

(where wages and salaries include payroll taxes but exclude the compensation of the self-employed). PM_t is the implicit deflator for imported goods and services, obtained as the ratio of nominal import expenditures to constant price import expenditures. As this series is obtained as a weighted average of underlying national aggregates, it includes the average price of goods and services traded within the Euro area in addition to the price of goods and services imported from outside the area. Clearly, it is only the latter prices that should be included in the index, but it is not possible to remove the contribution from the average price of internally traded goods. As such, the import price index that enters our analysis contains an important measurement error component, and we shall take this point into account when analysing our empirical results in section 4.2 below.

The variable $PCOM_t$ is based upon the HWWA (Hamburgerisches Welt-Wirtschaftliches-Archiv) world commodity price index, which measures the US$ prices of 29 important raw materials at 36 market places around the world (the price of crude oil is the dominant component in the index and receives a weight of 55%). This index is then multiplied by the €-US$ exchange rate in order to give a measure of the cost of crude oil and raw material imports to the Euro area. The variable TAX is the ratio of GDP at market prices to GDP at factor cost. The output gap, gap_t, measures the deviation of the natural log of GDP from the natural log of trend GDP, where the latter is obtained using the Hodrick-Prescott filter with the smoothing parameter set to 1600.

In Figure 1 we plot the log transforms of P, ULC, PM, $PCOM$ and TAX and some relevant linear combinations of those variables. We also graph the first difference of the log consumer expenditure deflator, Δp_t, which is the measure of the quarterly inflation rate used in this paper, and the output gap variable, gap_t.

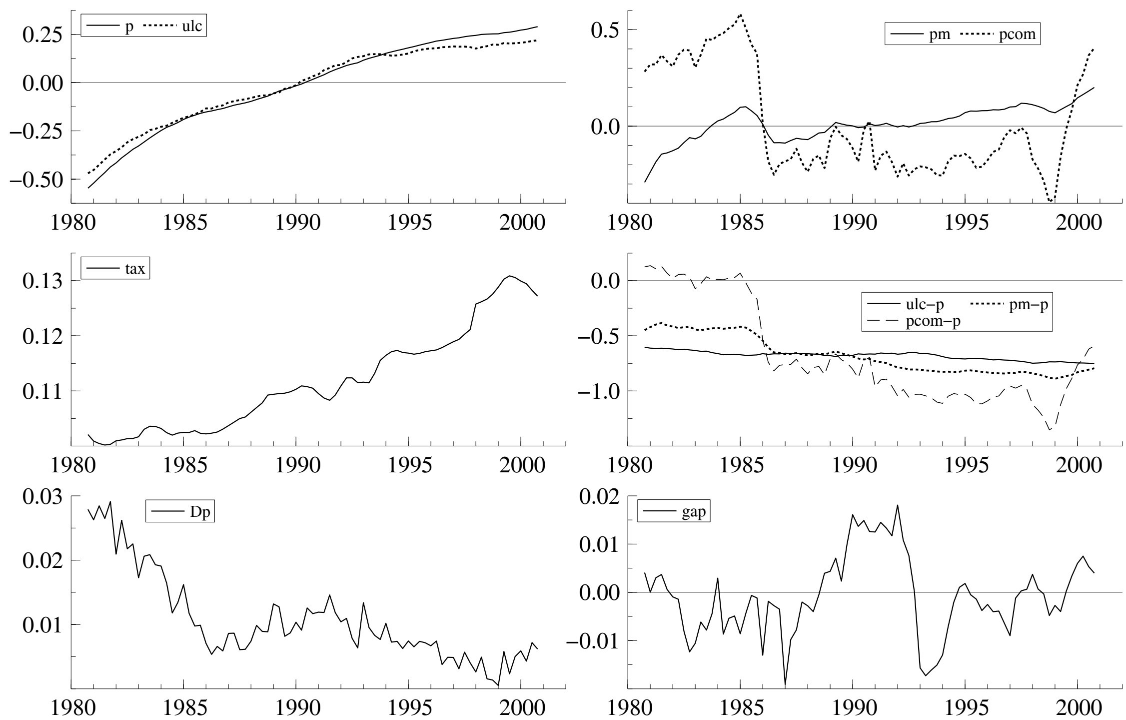

Figure 1: The data series used in the empirical analysis..

4 Empirical results for the Euro area

4.1 Modelling the long run

In the first part of the empirical analysis we estimate a VAR for the vector $[p\ ulc\ pm\ pcom]'$. Following the discussion in Section 2 we condition on the non-modelled variables tax and gap_{t-1}, which measure the tax wedge and the lagged output gap respectively. We also include a dummy variable set to unity in 1992(3) and 1992(4), and to zero otherwise. This dummy, denoted $i92q3q4$, is intended to control for any effects of the global recession that are not captured by the output gap variable, and also the events in foreign exchange markets that led to the break-up of the ERM.[9]

As alluded to in Section 2, we assume that the variables entering the price level function are I(1), or I(1) with breaks in means, rather than I(2). Hendry (2001) notes that if price levels are integrated once and are subject to major breaks then the measurements of those variables will be affected by I(1) deviations from their theoretical counterparts in Section 2. These measurement errors can lead to

[9] Strictly speaking, the use of such a dummy in the unrestricted part of the model, i.e. outside the cointegration space, implies that a step dummy should be included inside the cointegration space. However, we already include a time trend inside the cointegration space and this should be sufficient to control for any shift in the levels relationship.

the impression that price measures are I(2) when they are in fact I(1). Hendry (2001) therefore argues that the inflation rate be treated as I(0) with breaks, but (possibly) measured with an I(0) error, and that is the approach that we take in the analysis of Euro area inflation.

The VAR contains three lags in each of the endogenous variables and is estimated in the isomorphic Vector Equilibrium Correction Mechanism (VEqCM) form using data for 1980(4) to 2000(4). The estimation procedure is the Johansen (1988) cointegration technique, with a deterministic trend in the cointegration space.[10] The tax wedge is treated as I(1) and is therefore allowed to enter the cointegrating space that defines the long-run solution for the price level, while the output gap is treated as I(0) and does not enter the cointegration space. It is important to note that at this stage we maintain the assumption that the equilibrium price-cost markup is fixed, even though the discussion in section 2 indicated that it may vary over time (the time trend may control for some of the potential drift in the markup factor, but it can only be a very crude control). This caveat should be borne in mind throughout the remainder of this sub-section.

The results obtained for the unrestricted VEqCM are used to determine the cointegrating properties of the model. The *Trace*-statistic for the hypothesis that the number of cointegrating vectors, r, is less than or equal to one is 93.21, while that for r less than or equal to two is 53.51. The 5% critical values quoted in Harbo et al. (1998) are 71.7 and 49.6 respectively (see their Table 2 for the case in which one exogenous variable enters the cointegrating space). These results suggest that there is strong support for $r = 1$, but only weak support for $r = 2$. In view of the sample size being used in the present application, and also the absence of any economic interpretation for a second cointegrating vector, we maintain $r = 1$ in what follows.[11] If we normalise with respect to p_t, the cointegrating relation can be written as follows:

$$p_t = \underset{(0.05)}{0.73}ulc + \underset{(0.04)}{0.23}pm_t - \underset{(0.01)}{0.02}pcom_t \underset{(0.60)}{-1.00}\, tax_{t-1} \underset{(0.00015)}{+0.0025}\, trend_t$$

It can be seen that both the tax wedge variable, *tax*, and the commodity price index, *pcom*, are negatively signed. At the outset we would expect a positive unit coefficient for the tax wedge. However, we find that the hypothesis that the pass-through coefficient for the tax wedge is 50% yields a p-value of 0.01, suggesting that plausible settings for that parameter of the markup relation are rejected by the data.

A possible explanation for the negative sign of the commodity price index is that the import prices *pm* includes raw materials, and the negative sign corrects for double counting if what matters is in the inflation equation is the price of imported consumer goods.[12] Furthermore, recall that the aggregate deflator for import expenditures includes prices based on intra-European trade, which implies that the

[10] All of the empirical results in Section 4.1 were obtained using *PcFiml 9.3* — see Doornik and Hendry (1997)

[11] We note that the presence of the output gap in the unrestricted part of the model, i.e. outside the cointegration space, implies that the critical values used for inference here are approximate, see Rahbek and Mosconi (1999). This is a further reason for imposing a theoretical prior in setting the cointegrating rank.

[12] Assume that the price equation (1) were a weighted average of *ulc* and the import price of consumer goods *pmc*. The latter is not the information set, but it can be seen as the weighted

markup relation in (1) entails some double counting of all input costs. This may lead to biased estimates for both the import price coefficient and the coefficient on the commodity price index. In order to investigate these possibilities we analyse VARs that contain just three endogenous variables. The two cases that we consider are the following:

- Case 1: The vector of endogenous variables is $[p\ ulc\ pm]'$. The tax wedge, a time trend and *pcom* are not modelled but are allowed to enter the cointegrating space. All main variables enter the system with three lags, alongside the lagged level of the output gap.
- Case 2: The vector of endogenous variables is $[p\ ulc\ pcom]'$. The tax wedge, a time trend and *pm* are allowed to enter the cointegration space but are not modelled. Again, all main variables enter the system with three lags, alongside the lagged level of the output gap.

Using the same estimation procedure as that set out above, we find that in Case 1 there is strong support for the presence of one cointegrating vector and only marginal support for a second cointegrating vector.[13] Once again, we maintain the hypothesis that the VAR comprises just one cointegrating vector.

The economic identification of the cointegrating vector for case 1 is summarised in Table 1. Panel 1 shows that the results are almost the same as those obtained when there are four endogenous variables. Panel 2 shows that the highest setting for the tax wedge pass-through factor that is not rejected at the 5% significance level is 0.50. Conditional upon this restriction the coefficient on the world raw material price index, $pcom_{t-1}$, can be set to zero, and in panel 3 we implement this restriction. The hypothesis that p_t, ulc_t and pm_t form a linearly homogeneous relation cannot be rejected, a claim that is clear from an inspection of panel 3 and which is demonstrated formally by the second test outcome quoted beneath Panel 4, which tests the marginal restriction embodied in the model in panel 4 relative to that in panel 3. For the record, we note that linear homogeneity of the price level in unit labour costs alone is rejected by the data (results not reported here).

We now turn to case 2, i.e. a VAR model in $[p\ ulc\ pcom]$, with pm_{t-1} added to the set of conditioning variables. The mis-specification tests indicate that the new VAR is still congruent to the Data Generating Process (DGP). In contrast to the previous case, however, we now obtain formal support for the existence of just one cointegrating vector.[14] Table 2 summarises the results for this specification. In

---

average of aggregated import prices PM and the import price of raw materials $PCOM$, noting imports includes imports of raw materials. In logs we get $pcm = [pm - (1-a)pcom]/a$, where $(1-a)$ is the share of raw materials (ignoring other import components like capital goods etc). It is seen that *pcom* appears with a negative weight. We would like to thank the anonymous referee for pointing this out to us.

[13]The *Trace*-statistics are 78.15 and 37.47 (degrees of freedom corrected). Using Table 2 in Harbo et al. (1998) for the case with two exogeneous variables, the 5 percent critical values for rejecting the hypotheses that there are zero/at most one cointegrating vectors are 56.3 and 35.5, respectively.

[14]The *Trace*-statistics are 59.3 and 28.18 (degrees of freedom corrected). As in the previous case, the 5 percent critical values for rejecting the hypotheses that there are zero/at most one

Table 1: Testing steady-state hypotheses on the price level equation based on the VAR specification with unit labour cost and import prices as endogeous variables.

Panel 1: The just identified price level equation
$p_t = \underset{(0.05)}{0.73} ulc_t + \underset{(0.04)}{0.23} pm_t - \underset{(0.01)}{0.02} pcom_{t-1} \underset{(0.60)}{-1.00}\, tax_{t-1} \underset{(0.00015)}{+0.0026}\, trend_t$
Panel 2: 50 percent pass-through of indirect taxation
$p_t - 0.5\, tax_{t-1} = \underset{(0.04)}{0.84} ulc_t + \underset{(0.05)}{0.15} pm_t - \underset{(0.015)}{0.002} pcom_{t-1} \underset{(0.0002)}{+0.0014}\, trend_t$
$\chi^2(1) = 4.00[0.05]$
Panel 3: No effect from commodity prices and 50 % effect of indirect taxation
$p_t - 0.5\, tax_{t-1} = + \underset{(0.02)}{0.84} ulc_t + \underset{(0.02)}{0.16} pm_t \underset{(0.0002)}{+0.0015}\, trend_t$
$\chi^2(2) = 4.02[0.13],\ \chi^2(1) = 0.013[0.91]$
Panel 4: As Panel 3 and homogeneity
$p_t - 0.5\, tax_{t-1} = \underset{(-)}{0.84} ulc_t + \underset{(0.01)}{0.16} pm_t \underset{(0.0001)}{+0.0015}\, trend_t$
$\chi^2(3) = 4.02[0.25],\ \chi^2(1) = 0.001[0.97]$
The sample is 1980(4) to 2000(4), 81 observations.

System mis-specification tests for the underlying VAR:	
AR_v *1-5* $F(45, 119)$	1.14[0.28]
$Normality_v\ \chi^2(6)$	3.58[0.73]
$Heteroscedasticity_v\ F(144, 165)$	0.60[0.99]

References:
AR-test (Godfrey (1978) and Doornik (1996)),
Normality test (Doornik and Hansen (1994)), and
Heteroscedasticity test (White (1980) and Doornik (1996)).

the just identified case in which we do not impose theory-driven restrictions, the tax wedge effect is again negatively signed, though unlike in the previous case we cannot reject the hypothesis that there is full pass-through from the tax wedge to consumer prices - the imposition of this restriction (see panel 2) yields a test statistic of 1.20, which generates a p-value of 0.27 using a $\chi^2(1)$ distribution.

The import price term is the least significant in Panel 2 and we therefore delete it in moving to panel 3. This specification shows that after imposing the tax wedge and import price restrictions, the commodity price index enters the long-run relation with a positive and significant coefficient. In panel 4 the linear homogeneity of p_t in ulc_t and $pcom_t$ is imposed and it is clear from the test results quoted beneath panel 4 that such a restriction is compatible with the data. The relative weights in the long run equation are 0.96 and 0.04, the latter being small but clearly significant (as before, linear homogeneity of the price level in unit labour costs alone is rejected by the data). The time trend enters the long-run equation for the price level in panel 4 with a coefficient that implies autonomous annual growth in the price level equal to 0.36%. As suggested previously, this can be interpreted as the result of some form

cointegrating vectors are 56.3 and 35.5 respectively. Hence, there is no evidence of a second cointegrating vector in this case.

of measurement error such as the omission of the self-employed from the calculation for unit labour costs.

Table 2: Testing steady-state hypotheses on the price level equation based on the VAR specification with unit labour costs and commodity prices .

Panel 1: The just identified price level equation
$p_t = \underset{(0.07)}{0.82} ulc_t + \underset{(0.02)}{0.01} pcom_t + \underset{(0.06)}{0.09} pm_{t-1} - \underset{(0.97)}{0.27}\, tax_{t-1} + \underset{(0.0007)}{0.0021}\, trend_t$
Panel 2: Full effect of indirect taxation
$p_t - tax_{t-1} = \underset{(0.06)}{0.89} ulc_t + \underset{(0.02)}{0.02} pcom_t + \underset{(0.06)}{0.04} pm_{t-1} + \underset{(0.0003)}{0.0012}\, trend_t$
$\chi^2(1) = 1.20[0.27]$
Panel 3: No effect from import prices and full effect of indirect taxation
$p_t - tax_{t-1} = + \underset{(0.03)}{0.93} ulc_t + \underset{(0.01)}{0.04} pcom_t + \underset{(0.0003)}{0.0011}\, trend_t$
$\chi^2(2) = 1.46[0.48],\ \chi^2(1) = 0.26[0.61]$
Panel 4: As Panel 3 and homogeneity
$p_t - tax_{t-1} = \underset{(-)}{0.96} ulc_t + \underset{(0.01)}{0.04} pcom_t + \underset{(0.0002)}{0.0009}\, trend_t$
$\chi^2(3) = 1.86[0.60],\ \chi^2(1) = 0.39[0.53]$
The sample is 1980(4) to 2000(4), 81 observations.

System mis-specification tests for the underlying VAR:	
AR_v *1-5* $F(45, 122)$	1.24[0.18]
$Normality_v$ $\chi^2(6)$	8.04[0.24]
$Heteroscedasticity_v$ $F(144, 171)$	0.50[1.00]
References: See Table 1	

In the top two panels of Figure 2 we plot the residuals obtained from the final long-run pricing relations in Table 1 and Table 2, and the recursive estimates of the parameters corresponding to those relations in the remaining panels (see the notes to the figure for exact details). The latter plots indicate that coefficient stability is satisfactory over the later stages of the sample, though as this is a rather short period it is not an especially strong test of the congruency of the model. In particular, it does not cast any light on the possibility that there was time-variation in the price-cost markup during the 1980s (we do not report recursive coefficient estimates for the 1980s because they are difficult to interpret given that they are calculated using a very small number of observations). The residuals from the estimated cointegrating vectors indicates that some minor non-stationarities remain in both cases. Indeed, ADF(4) tests[15] for the presence of a unit root do not reject the null in either case (the test statistic is -2.27 for Case 1 (import prices) and -2.49 for Case 2 (world raw materials prices), while the 5% critical value is -2.90). While these test outcomes may reflect the low power of the ADF procedure (that the residuals from the cointegrating relations were identified as stationary using the Johansen analysis would support this view), they suggest some evidence of the local non-stationarities of the sort discussed in Section 2 above.

[15]The ADF specification chosen was that with the largest number of lags and a significant coefficient on the final lag (lag orders above 12 were not considered).

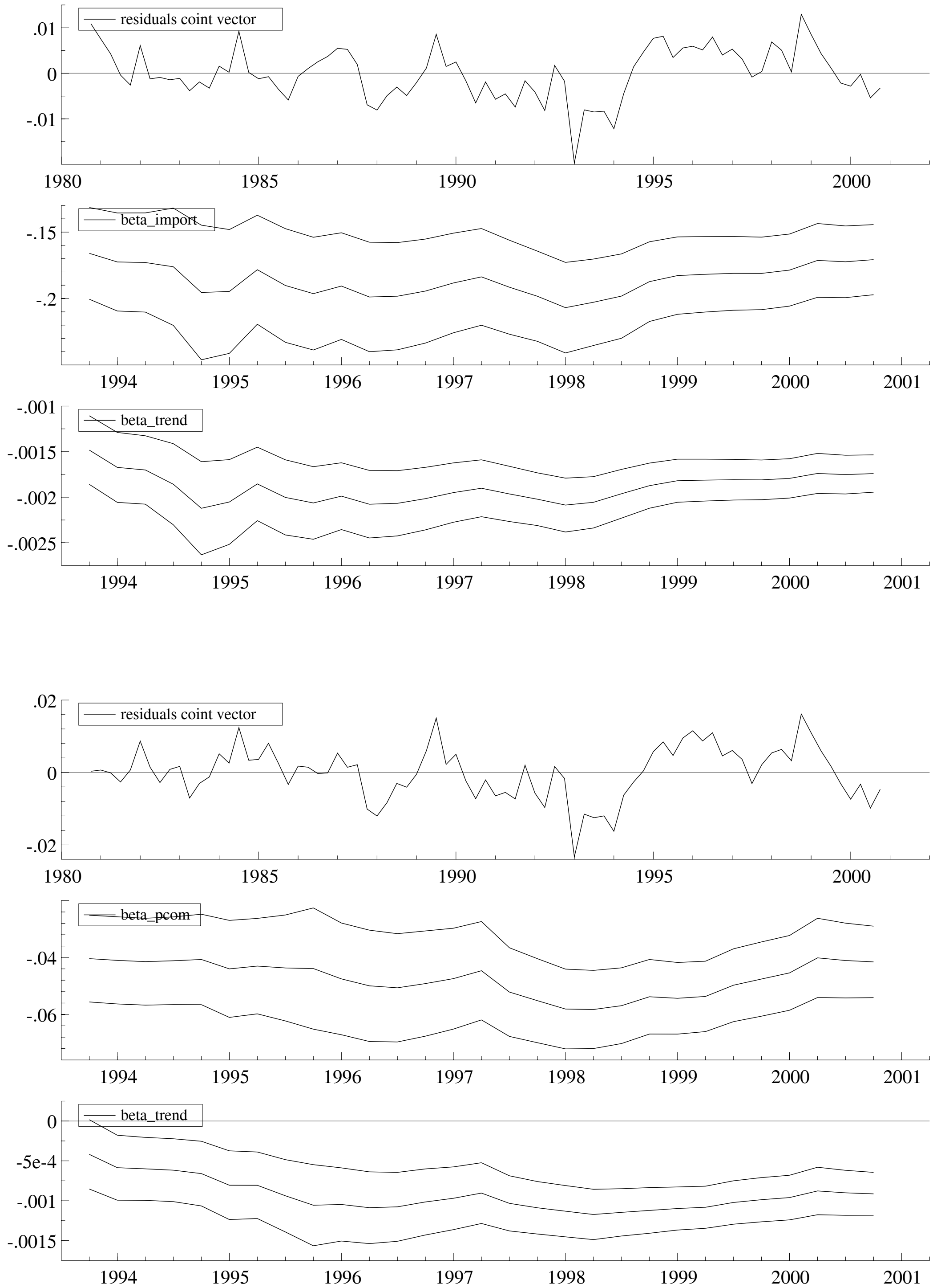

Figure 2: Residuals of the identified cointegration vectors. Recursively estimated parameters in the long run equation. The final specification with import prices top (cf. Panel 4, Table 1) and the specification with commodity prices bottom (cf. Panel 4, Table 2)

4.2 Modelling the short run (EqC) model of inflation

In order to analyse inflation fluctuations in detail we focus on the second of the trivariate VARs considered above, i.e. the model in which the steady state price level is a homogeneous function of unit labour costs and world commodity prices (after controlling for the tax wedge and a time trend). This model produced a more plausible steady-state pricing relation and also ensures that the import deflator, a variable that may be subject to important measurement biases, is moved to the conditioning set in the long-run analysis. A test for the weak exogeneity of unit labour costs and commodity prices within this VAR, i.e. a test for the loadings on the cointegrating vector in the unit labour cost and commodity price equations being zero, yields test outcomes of $\chi^2(1) = 1.62[0.20]$ and $\chi^2(1) = 1.10[0.30]$ respectively, and the joint hypothesis yields the outcome $\chi^2(2) = 2.44[0.30]$. Non-rejection of these hypotheses implies that switching from a system analysis to a single equation model for inflation can be done without any loss of efficiency, and we pursue that option in the next part of this analysis.

In order to model the short-run relationships we exploit the properties of the estimated steady-state in panel 4 of Table 2. Specifically, we formulate a dynamic equation for Δp_t conditional upon the relative prices $(ulc - p + tax + \varphi trend)$ and $(pcom - p + tax + \varphi trend)$ and the terms $\{\Delta p, \Delta ulc, \Delta pcom, \Delta tax, gap\}$. The exact specification from which the analysis begins corresponds to equation (4), which was discussed in section 2. The equation includes a time trend and the dummy $i92q3q4$ that was used in the cointegrated VAR analysis, and is estimated for the period 1981(1) to 2000(4). The parsimonious short-run model is derived using the general-to-specific modelling strategy described in section 2, and is reported below.[16] A series of residual diagnostic tests are reported along with the estimated regression. The references for these tests are given in Table 2, except for the *ARCH 1-4 F*-test, which is a test for autoregressive conditional heteroscedasticity, see Engle (1982), and the *RESET*-test, which is the specification test due to Ramsey (1969).

$$
\begin{aligned}
\Delta p_t = {} & \underset{(0.01)}{0.04} + \underset{(0.08)}{0.36}\, \Delta p_{t-2} + \underset{(0.002)}{0.010}\, \Delta pcom_{t-1} + \underset{(0.03)}{0.09}\, gap_{t-1} \\
& + \underset{(0.04)}{0.09}\, \Delta ulc_{t-1} + \underset{(0.29)}{0.62}\, \Delta tax_{t-1} - \underset{(0.001)}{0.006}\, i92q3q4 \\
& + \underset{(0.02)}{0.12}\, (ulc - p + tax + \varphi trend)_{t-2} \\
& + \underset{(0.001)}{0.006}\, (pcom - p + tax + \varphi trend)_{t-3}
\end{aligned} \tag{7}
$$

$$\sigma = 0.001901$$

AR 1-5 $F(5, 67)$	1.14[0.34]
ARCH 1-4 $F(4, 64)$	0.23[0.92]
Normality $\chi^2(2)$	0.73[0.70]
Heteroscedasticity $F(13, 58)$	0.54[0.91]
RESET $F(1, 71)$	4.51[0.04*]

[16] All of the empirical results in Section 4.2 were obtained using *PcGive10* - see Hendry and Doornik (2001).

The joint hypothesis comprising each of the zero restrictions imposed in deriving the model via general-to-specific modelling, which entails 11 exclusion restrictions, cannot be rejected (the p-value for the test is 0.97). This is the expected result given that the smallest p-value for any single reduction was 0.30. In Figure 3 we plot recursive estimates of the coefficients of the model. The behaviour of individual coefficients is fairly stable, though it should be noted that the intervention dummy $i92q3q4$ is unable to eliminate all parameter instability occurring in $1992(3) - 1992(4)$. The recursive residuals and recursive Chow tests plotted in Figure 4 confirm the stability of the estimated equation for the period running from the late 1980s, but (as was the case with the recursive output considered previously) we cannot infer whether or not the model is constant over the first half of the sample period. Indeed, the residual diagnostic tests suggest that the model is not congruent to the data generating process in some respects. For example, the RESET test signals mis-specified functional form, which may include omitted non-linearities, a possibility that we explore in further detail in Section 4.3. One explanation for this finding is that the model assumes a fixed markup of prices over costs in equilibrium. In order to investigate this possibility, we now estimate a model that controls for fluctuations in inflation that arise because of shifts in the target markup of prices over costs.

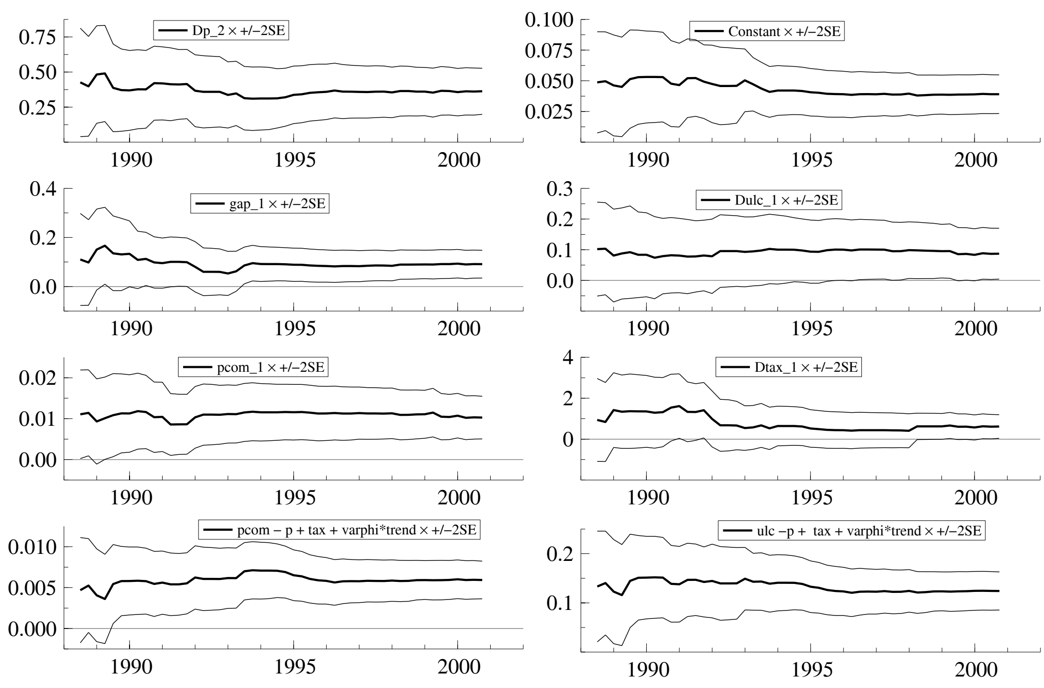

Figure 3: Recursive plots of the estimated coefficients of the equilibrium correction inflation model.

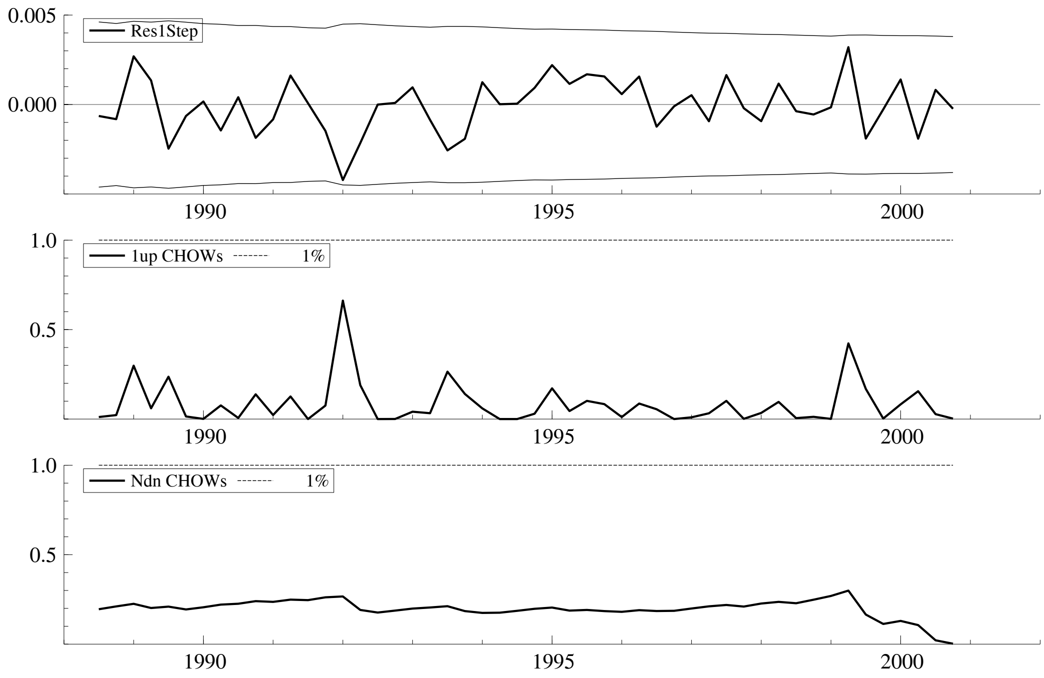

Figure 4: One step residuals and recursive Chow-tests for the linear inflation model.

4.3 A Local Level Equation for the Inflation Process

We apply the local level estimation technique to the tested down linear equilibrium correction model rather than the unrestricted specification in (5). The validity of the restrictions imposed in obtaining the parsimonious model are then checked by adding back in a distributed lag in each of the variables, one at a time, and testing down once more through repeated local level estimation. We opt for this strategy rather than repeating the entire general-to-specific modelling exercise because we find that highly parametrised models tend to be associated with implausibly large variation in the local level term. The strategy that we follow, namely developing the best fitting linear model and then studying the properties of a non-linear version of that restricted specification, is consistent with the approach to non-linear modelling advocated by Teräsvirta (1998).

The results are reported in panel 1 of Table 3.[17] The terms in unit labour cost inflation and the growth rate of the tax wedge turn out to be insignificant after controlling for the time-varying intercept (these two terms were only marginally significant in the previous case). In panel 2 we report a parsimonious model that excludes these two terms. Beneath each estimated equation we quote the Box-Ljung test for lack of residual serial correlation and the Doornik-Hansen test of residual normality, see Koopman et al. (2000). These tests refer to the *auxiliary residuals*, which are a smoothed version of the error processes u_t and ε_t in equations (5) and (6). Further details on the interpretation of auxiliary residuals can be found in Koopman

[17] All of the empirical results in section 4.3 are obtained using *STAMP 6* — see Koopman et al. (2000).

et al. (2000). The equation standard error (σ) is the square root of the variance of the model residuals (the u_t series in equation (5)), and is therefore comparable with the standard error reported for the linear inflation equation. In the lower half of Figure 5 we plot the time-varying intercept, which is the term $\psi\mu_t$ from equation (5) multiplied by the reciprocal of one minus the estimated autoregressive parameter in panel 2 of Table 3. This scaling factor is used in order to illustrate the steady-state contribution of the time-varying intercept to the inflation rate. In the upper graph the stretched time-varying intercept is plotted alongside inflation (a constant has been subtracted from the time-varying intercept in order to ensure that it has the same mean as the inflation series).

Table 3: Local Level estimates of the Equilibrium Correcting Model (Panel 1) and its simplification (Panel 2).

Panel 1: Local Level Estimate of the EqCM

$$\Delta p_t = \underset{(0.01)}{0.05} LL_t + \underset{(0.09)}{0.22}\Delta p_{t-2} + \underset{(0.042)}{0.080}\,\Delta ulc_{t-1} + \underset{(0.003)}{0.01}\,\Delta pcom_{t-1}$$

$$+\underset{(0.29)}{0.53}\Delta tax_{t-1} + \underset{(0.04)}{0.12}\Delta gap_{t-1} + \underset{(0.03)}{0.14}(ulc - p + tax + \varphi trend)_{t-2}$$

$$+\underset{(.002)}{0.005}(pcom - p + tax + \varphi trend)_{t-3} - \underset{(.002)}{0.005} i92q3q4$$

$$\sigma = 0.171\%$$

$$q\text{-}ratio^{*)} = 0.089$$

AR 1-5 statistic = 2.34, 95% *critical value is* 5.99 ($\chi^2(2)$)

Normality $\chi^2(2)$ 0.10[0.95]

Panel 2: Local Level Estimate of the Parsimonious EqCM

$$\Delta p_t = \underset{(0.014)}{0.05}\, LL_t + \underset{(0.09)}{0.23}\Delta p_{t-2} + \underset{(0.003)}{0.01}\,\Delta pcom_{t-1} + \underset{(0.04)}{0.11}\, gap_{t-1}$$

$$\underset{(0.03)}{0.13}(ulc - p + tax + \varphi trend)_{t-2}$$

$$+\underset{(0.002)}{0.005}(pcom - p + tax + \varphi trend)_{t-3} - \underset{(0.001)}{0.004} i92q3q4$$

$$\sigma = 0.172\%$$

$$q\text{-}ratio^{*)} = 0.108$$

AR 1-5 statistic = 2.34, 95% *critical value is* 5.99 ($\chi^2(2)$)

Normality $\chi^2(2)$ 0.12[0.94]

The sample is 1981(1) to 2000(4), 80 observations.

*) The *q-ratio* is defined as the ratio $\sigma^2_\varepsilon/\sigma^2_u$, where σ^2_ε is the residual variance of (6) and σ^2_u is the residual variance of (5)

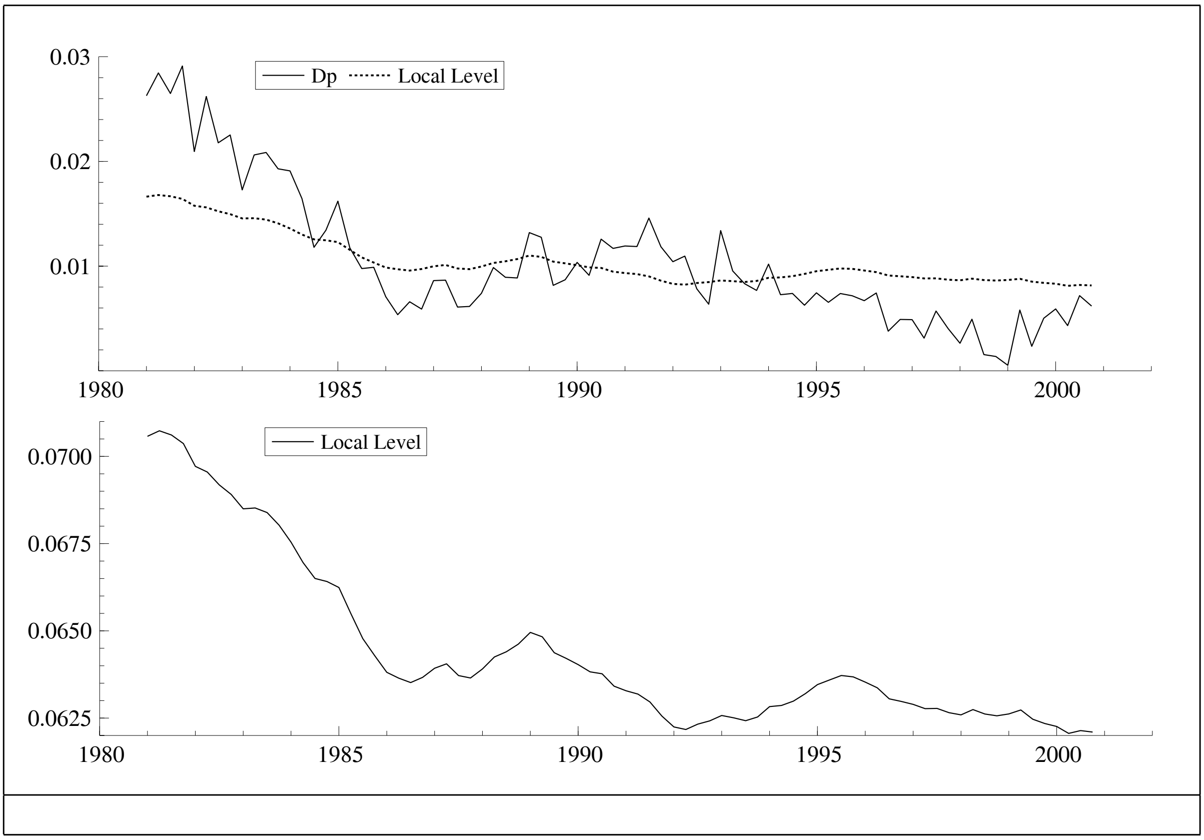

Figure 5: The upper graphs show the Local Level term plotted against inflation and the lower graph is a close-up of the Local Level term, scaled with the reciprocal of 1 minus the autoregressive parameter in Panel 2 of Table 3..

The addition of the local level term to the model reduces the regression standard error by approximately one tenth. The t-ratio for the coefficient multiplying the local level in the second model in Table 3 is 3.6. Given that the variables are not in zero mean form, this partly reflects the significance of the intercept. When the regression is estimated with all variables in zero mean form, the absolute t-ratio is 1.86 and the corresponding p-value is 0.067. The local level term therefore makes a contribution to inflation that is significant at the 10% level.

The autoregressive coefficient, which was estimated to be 0.36 in the linear model, is only 0.23 in the local level model. One interpretation of this parameter shift is that the autoregressive coefficient in the linear model is biased upwards because of the exclusion of the time-varying intercept. The additional explanatory power assigned to the autoregressive term in the linear model probably explains why that specification does not fail the residual diagnostic tests despite the exclusion of the time-varying intercept. Similarly, one explanation for the fact that the VAR estimated at the start of this section appeared to have plausible properties despite being based upon a fixed long-run markup assumption is that the autoregressive terms are able to proxy some of the effects of permanent changes in the percentage markup factor. Thus, one danger in applying a linear equilibrium correction model to data from an economy that has been subject to structural change appears to be that the autoregressive coefficients in the model are likely to be biased upwards, giving the impression that inflation is more persistent than is actually the case.

The parsimonious local level model in Table 3 yields a q-ratio of 0.108, which means that the variance of the shocks impacting the local level term is equal to 10.8 % of the variance of the residuals for the estimated model. An inspection of the time series plots for the local level term in Figure 5 casts some light on the role that it plays in explaining inflation. Its main feature is a downward trend over the first half of the 1980s, which accounts for between one quarter and one third of the eight percentage points reduction in Euro area inflation during the period 1980-85. Indeed, this may be an underestimate of the importance of the time-varying markup factor if some of its effects are captured by the weighted time trend derived from the VAR model. However, when we allow the time trend to enter the model unrestrictedly (rather than with a pre-determined coefficient, as part of the relative price terms), the results do not change in a meaningful way. In any case, the reduction in inflation (and hence the price level) that occurred independently of the costs of production is certainly important in an economic sense as well as a statistical sense.

This result is important because it suggests that the substantial reduction in Euro area inflation that occurred in the 1980s cannot be entirely explained in terms of well known disinflation strategies such as a deceleration of nominal wages, or an appreciation of the currency. Additionally, the price level fell independently of cost and demand conditions, and we attribute this part of the disinflation to firms reducing their claims on output through choosing a smaller percentage markup of prices over costs. Such a change in price-setting behaviour cannot be attributed, with certainty, to a single set of economic reforms. However, one possibility that we find plausible is that the inception of the new European Monetary System (EMS) in 1979, together with the marketisation of several southern European economies prior to their accession to the European Union, led to stronger product market competition and hence a reduction in the profit share.

If the local level term does account for shifts in the target markup of prices over costs, it should help to control for some of the local non-stationarities that occur in the markup when it is derived from the VAR model under the assumption that the long-run markup factor is fixed. Recall that this estimate of the markup is a weighted average of the relative price terms $(ulc - p + tax + \varphi trend)$ and $(pcom - p + tax + \varphi trend)$ and will therefore depend negatively on the percentage markup factor that firms target in equilibrium (since this is a component of p).[18] A regression of the markup on the (autoregressive coefficient adjusted) local level should therefore yield a residual series that is more obviously stationary than the markup implied by the linear model. An ADF(4) test based upon these residuals rejects the unit root hypothesis at the 5% level of significance. Recall that we were unable to reject the hypothesis that a weighted average of the equilibrium correction factors *alone* contains a unit root.[19] The reasons for the change in the test outcome can be seen in Figure 6, which plots autocorrelograms for the equilibrium correction factor before and after it has been regressed on the local level. There is clearly less persistence in the equilibrium correction factor after controlling for the local level

[18] The equilibrium correction term applies a weight of 0.96 to $(ulc - p + tax + \varphi trend)$ and .04 to $(pcom - p + tax + \varphi trend)$. These are the weights derived from equation (7) in the text.

[19] Of course, non-rejection of the unit root hypothesis most likely reflects the low power of the ADF procedure. Our point here is simply that a weighted average of the relative prices is more obviously stationary after controlling for the local level.

term. This illustrates an important second function performed by the local level: In addition to explaining negative drift in inflation arising from a reduction in the price-cost markup, it corrects for mis-measurement of relative prices arising when shifts in the equilibrium percentage price-cost markup are not taken into account.

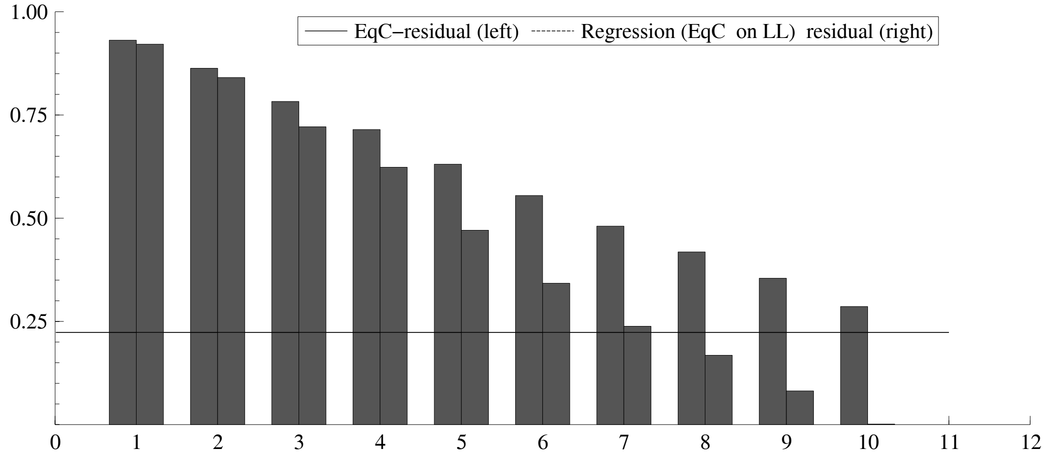

Figure 6: The autocorrelograms of the Equilibrium Correction factor of the model in Panel 2 of Table 3 (left) and of the residuals of that factor regressed on the Local Level term of that model (right). .

Features of price adjustment in the Euro area

We round off our discussion with some comments on the inflation process in the Euro area. Our results indicate that there is partial adjustment of the price level towards its long-run equilibrium. Deviations from steady-state arising due to fluctuations in the productivity adjusted real wage induce price adjustment six months later, whilst disequilibrium in commodity prices relative to the consumption deflator affects price inflation with a nine month lag (though note that commodity price inflation exerts a short-run effect after just one quarter).[20] Once these equilibrium correction effects have started, 12.2% of the disequilibrium in price levels is eliminated each quarter, with the autoregressive term inducing slightly faster adjustment after a further two quarters. Hence, the results are consistent with the view that homogeneity of the price level in input costs holds in the long-run but not the short-run.

The timing of price adjustment suggested by the model may reflect the fact that labour is directly purchased by most firms, whilst fuel and raw materials are indirect inputs for most firms, e.g. in the case of retailers that are at the end of a supply chain including manufacturers and wholesalers. This implies that movements in labour costs affect average profitability with a shorter lag than do shifts in commodity

[20]We note that the insignificance of the dynamic adjustment terms referred to in (4) indicates that it is not possible to reparameterise the local level equation such that there is an identical lag in the response of inflation to the two measures of relative prices.

prices, which in turn means that there is pressure for faster adjustment of consumer prices when cost increases originate in the labour market.

The pattern of price adjustment established for the Euro area can be compared with the results from other studies. Bowdler (2003) develops separate models of the inflation process for 20 OECD countries, including 9 of the 12 countries that constitute the Euro area economy (the three exceptions are Ireland, Luxembourg and Portugal).[21] In the largest Euro area economy, Germany, Bowdler finds that equilibrium corrections between prices and unit labour costs commence with a two quarter lag, matching the finding for the aggregate model documented in this paper. In contrast, unit labour costs relative to the consumption deflator start to increase inflation after one quarter in France and the Netherlands, but only after three quarters in Spain, and four quarters in Italy. In light of this evidence, the $t-2$ dating for unit labour costs that we identify here appears plausible. It is interesting to note that in the equations for the United States and the United Kingdom, Bowdler finds that the equilibrium correction terms in unit labour costs enter the equations at $t-1$. This suggests that price adjustment in the Euro area is quite slow by international standards. One interpretation of these results is that Euro area price dynamics may have been affected by state regulation of pricing decisions, which one would expect to have delayed pass-through from input costs to the price level.

Finally, we note that in the parsimonious model reported in Table 3, a 1% increase in output relative to trend raises the quarterly inflation rate by .11 percentage points in the next quarter. This is in the middle of the range of estimates obtained by Bowdler for individual countries, and therefore seems plausible.

5 Summary

This paper has drawn attention to the fact that standard equilibrium correction models of the price level assume that the markup of prices over costs is constant in the long-run. This assumption is particularly unattractive in the case of the Euro area economy because changes in trade and exchange rate policy are likely to have increased product market competition, and thereby decreased the markup factor. Consequently, linear equilibrium correction models may omit an important channel for Euro area inflation adjustment, and may therefore provide a poor fit to the data. One solution to this problem is to include a time-varying intercept, or local level, in the equilibrium correction model. In the case of the Euro area, such a term accounts for the negative drift in inflation that a linear model is unable to explain, and also corrects for measurement errors affecting relative price terms when shifts in the markup factor are overlooked. As a result, the model provides a better fit to the data, and is able to control for the local non-stationarities that affect relative prices. Although the structure of the inflation equation does not change dramatically compared to the linear case, the approach highlights the importance of controlling for the effects of structural change when modelling inflation.

[21] It should be noted that our approach departs from that in Bowdler (2003), in that the latter study tests down from general models that contain a time-varying intercept. This approach implies a non-nested testing procedure. However, the final equations reported by Bowdler appear to be well-specified, and the problem of a lack of comparability of models across studies is partly reduced by the fact that he reports OLS (constant intercept) estimates of the restricted models obtained at the end of the general-to-specific modelling exercise.

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