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EURO AREA INFLATION PERSISTENCE

BY NICOLETTA BATINI

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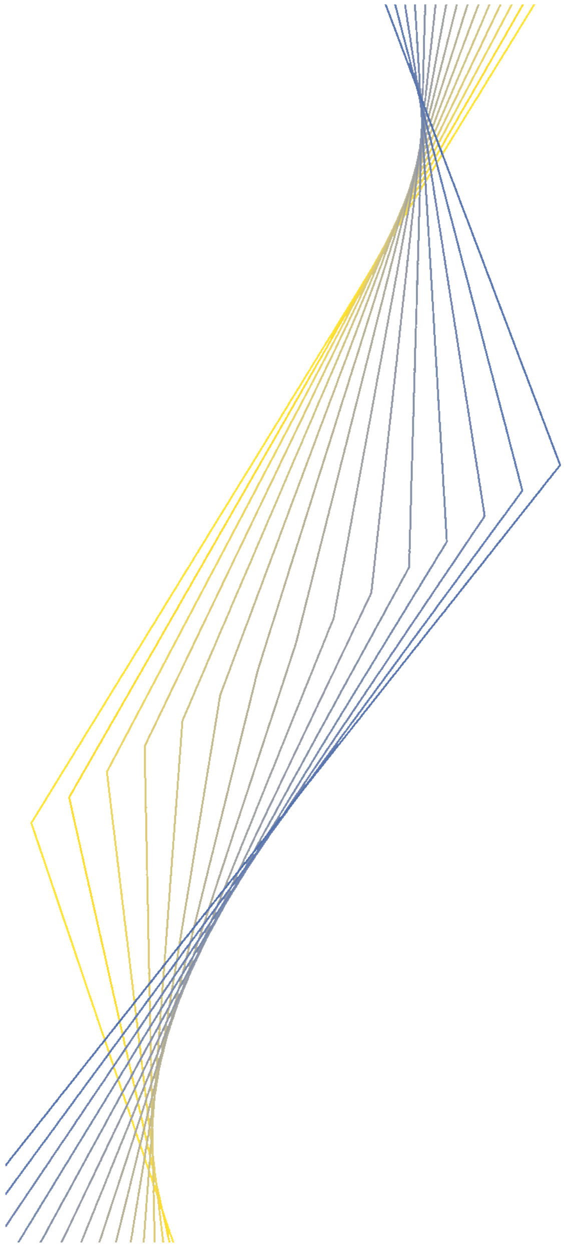

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EURO AREA INFLATION PERSISTENCE[1]

BY NICOLETTA BATINI[2]

December 2002

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Contents

Abstract

This paper presents evidence on the lag between monetary policy actions and the response of inflation in the euro area as a whole as well as in Germany, Italy and France. In line with previous findings for the US and the UK, results here show that this lag is longer than one year both in the euro area and in individual countries, and that a lag of this length has existed in Europe at least since the collapse of the Bretton Woods system, despite the numerous changes in European monetary policy regime thereafter. Results based on alternative definitions of inflation persistence support these findings, although, they suggest that at the country level, a drop in German inflation persistence and a sizeable shift in the mean of inflation—particularly in Italy and France—are beyond doubt. The paper shows that euro area inflation persistence could well be an intrinsic phenomenon rather than a 'statistical fluke' due to aggregation.

Key Words: Euro area, Europe, inflation persistence, HICP, monetary transmission, aggregation bias
JEL Classification Codes: E4-E5

Non-Technical Summary

The launch of the European Monetary Union has created an entirely new economic area. The exercise of monetary policy in such a new environment is a formidable task for European policymakers both because comprehensive and harmonised data for the area have not been collected in the past, and because the functioning of the area economy as a unified entity is yet largely unknown.

One matter on which uncertainty facing policymakers is particularly acute is that of the lag in effect of monetary policy at the area level. In general, the ability to quantify and, hence, model the sluggish response of inflation to changes in monetary conditions is important for monetary policymakers because it helps them understand how pre-emptive they should be in order to curb inflationary pressures at a minimum cost in terms of output gap variability. This is especially true for the recently-created European Central Bank (ECB), because it pursues the objective of price stability for the euro area as a whole, and it is not yet clear whether the observed delay in the response of euro area aggregate inflation to monetary stimuli is a mere product of aggregation of the individual countries' price indices, rather than an intrinsic phenomenon *per se*.

Recent empirical studies have emphasised that purely forward-looking NKPC specifications for inflation offer a good portrait of euro area inflation dynamics, suggesting that inflation persistence in the EMU is not a structural, policy-invariant feature of the data. On the other hand, European policymakers have adopted a definition of price stability which "is to be maintained over the medium term", in recognition of the existence of an intrinsic delay between monetary policy actions and their effect on inflation.

In line with previous work by Batini and Nelson (2001) on US and UK data, in this paper I looked at whether this conflict can be resolved by presenting model-free evidence on the delay between changes in monetary policy and their peak effect on inflation for the euro area as well as in a subset of its component countries. I also presented evidence of the autocorrelation properties of inflation and on the lag in the response of inflation to monetary policy shocks from simulations of a battery of identified parsimonious VARs.
I found that euro area inflation is persistent according to all definitions I use. For instance, results based on money, interest rate and inflation area-wide data show that it takes over a year before monetary policy actions have their maximum effect on inflation—a finding

which strongly validates the ECB medium-term policy orientation. I find relatively consistent results across different measures of monetary stance, including those based on monetary aggregates. This suggests that monetary aggregates in the euro area have a useful role as one set of measures of monetary conditions. Results based on alternative definitions of inflation persistence, like those on the degree of autocorrelation of inflation or on the lag between a monetary policy shock and the peak response in inflation from simulation of a small euro area estimated system in inflation deviations from target, the output gap and the short-term nominal interest rate, also indicate that European inflation is rather inertial.

Importantly, the persistence of European inflation seems to have varied only marginally over the past thirty years, despite the numerous monetary policy regime shifts occurred in Europe after the collapse of the Bretton Woods' exchange rate system.

A decomposition of the analysis at the country level reveals that some underlying inflation persistence is an inherent feature of way the single economies work—although I cannot exclude some changes over time in the countries' lead/lag relationship between inflation and prior money growth. However, important cross-country differences in the degree of inflation persistence emerge by looking at individual countries. Notably, inflation in Italy and France is highly serially correlated, a feature that has not changed greatly over time. By contrast, German inflation is much less autocorrelated today than in the past. And in Italy and France there has been a considerable drop in the mean and variance of inflation over the years, likely reflecting a shift in the implicit price stability objective in the countries following the French abandonment of the 'encadrement du credit' and the move to the 'Franc fort' policy, on one side, and the 'divorce' of the Banca d' Italia from the Italian Treasury, on the other side. I do not observe this for Germany, where inflation has been low historically.

In this sense and according to most definitions, the degree of inflation persistence in the euro area seems to be a halfway house between different degrees of inflation persistence at the country level. Put differently, area-wide persistence seems to result from a pure statistical averaging effect, rather than being a spurious phenomenon due to aggregation, as can be shown via alternative calculations based on the direct and indirect evidence presented in this paper.

1. Introduction

The launch of the European Monetary Union has created an entirely new economic area. The exercise of monetary policy in such a new environment is a formidable task for European policymakers both because comprehensive and harmonised data for the area have not been collected in the past, and because the functioning of the area economy as a unified entity is yet largely unknown.

One matter on which uncertainty facing policymakers is particularly acute is that of the lag in effect of monetary policy at the area level. In general, the ability to quantify and, hence, model the sluggish response of inflation to changes in monetary conditions is important for monetary policymakers because it helps them understand how pre-emptive they should be in order to curb inflationary pressures at a minimum cost in terms of output gap variability. This is especially true for the recently-created European Central Bank (ECB), because it pursues the objective of price stability for the euro area as a whole, and it is not yet clear whether the observed delay in the response of euro area aggregate inflation to monetary stimuli is a mere product of aggregation of the individual countries' price indices, rather than an intrinsic phenomenon *per se*.

Indeed, in accordance with conventional wisdom of the EMU's national central banks (NCBs), the ECB recognises the existence of lags in the monetary transmission mechanism. This emerges clearly from the wording of the mandate of the Governing Council of the European Central Bank, which states that "price stability is to be maintained over the medium term",[1] acknowledging the need for monetary policy to by-pass operational lags by means of a forward-looking, medium-term orientation.[2]

However, the actual extent of transmission lags from policy to inflation remains unclear, since, so far, available evidence on the characteristics of aggregate euro area inflation is still rather limited.[3]

When the EMU started, in 1999, the European Central Bank documented an approximate 6-quarter lag between money growth and inflation in the euro area.[4] This result was later

[1] See ECB *Monthly Bulletin*, January 1999, p. 46-7.
[2] Similarly, Issing et al. (2001) note that '[t]he medium-term orientation is partly a reflection of the time lag with which monetary policy affects prices'.

endorsed by systematic examinations of the leading indicator properties of money-based indicators, which found that money growth is particularly useful for forecasting inflation at horizons beyond one and a half years (see Nicoletti Altimari (2001)). Given the prominent role played by money growth in the stability-oriented strategy of the ECB, such evidence is presumably taken, among other facts, to lend formal support to the medium-term approach to price stabilisation embraced by the ECB's Governing Council (ECB (1999)).[5]

Parallel efforts aimed at improving the understanding of the properties of inflation in the EMU come from the empirical analysis on euro area inflation dynamics. A number of studies have modelled euro area aggregate and country-specific inflation behaviour starting from microprinciples, following the empirical work on fitting New Keynesian Phillips Curve (NKPC) to US data (see e.g. Sbordone (2002); Galí and Gertler (1999)). Results from this literature are mixed, although, at least for euro area, the majority of findings point to the fact that euro area inflation can be successfully modelled using the NKPC, whose structure does not imply inherent persistence in inflation.[6]

Looking at area-wide data, for instance, Galí, Gertler and López-Salido (2000) present evidence on the fit of the New Keynesian Phillips Curve (NKPC) over the period 1970-1998, and use it to compare the characteristics of European inflation dynamics with those observed in the US. They find that a purely forward-looking NKPC specification for inflation fits the euro area data remarkably well, possibly in a superior way than US data, and so infer that inflation in the euro area may be less inertial than in the US.[7] Re-estimates of the same curve by McAdam and Willman (2002), which attempt to treat the problem of non-stationarity in euro area data, also find weak support for the inclusion of backward-

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[3] Angeloni et al. (2001) summarise the available evidence on the monetary transmission mechanism in the euro area and in the area's individual countries.

[4] See ECB, *Monthly Bulletin*, February 1999, p. 38. Note that subsequent issues of the ECB *Monthly Bulletin* (notably, the July 2000 and May 2001 issues) revised this estimate downwards, suggesting a lag of 4 quarters between broad money growth and inflation for the euro area. This result is based on evidence obtained by matching 24-month moving average (MA) of annual percentage change in euro area *M3* and the 24-month MA of the annual percentage change in harmonised consumer prices (CPI until Jan 1996, the HICP thereafter).

[5] A medium-term orientation to policy is supported, indirectly, also by theoretical work. Using a two-equations model calibrated on the euro area, for instance, Smets (2000) finds that the optimal horizon for a forward-looking inflation constraint when policymakers place an equal weight on price stability and other growth objectives is between three and four years.

[6] The lack of inflation persistence in the NKPC was first pointed out by Ball (1994). Fuhrer (1997) and Ball et al (2002) again underlined this point, arguing that even the inclusion of a lagged inflation term in the NKPC fails to capture this stylised fact.

[7] Several works criticise the NKPC both on theoretical and empirical grounds. Among others, see for instance Ma (2002), Mavroeidis (2001), Bardsen et al (2002) and Guay, Luger and Zhu (2002).

looking elements. Smets and Wouters (2002) come to similar conclusions. They derive a model incorporating wage and price stickiness as well as wage and price indexation as in Christiano, Eichenbaum and Evans (2001), which implies a partially backward and forward-looking representation for inflation. They show that estimates of this model with Bayesian techniques can offer a good portrait of the behaviour of euro area inflation from 1970 to 1999, yet one in which inflation is only mildly backward-looking. Similarly, Coenen and Wieland (2000) estimate constrained and unconstrained VAR models for both the euro area and individual countries within the area. For the area as a whole, they find reasonably good fits both when they try to fit relative wage specifications à la Buiter and Jewitt (1981) and Fuhrer and Moore (1995), and when they use nominal wage specifications—although the data indicate a slight preference for the relative wage-contracting model, which implies inflation inertia.[8]

For individual countries the evidence is more conflicting. Amato and Gerlach (2000), for instance, fit NKPCs to a set of countries including a subset of euro area ones using non-linear least squares. Their results indicate that the purely forward-looking version of the Calvo price model of Sbordone (2002) and Galí-Gertler (1999) captures well inflation dynamics in most countries but there is evidence of parameter instability in France in the mid-90s. Similarly, Joendeau and Le Bihan (2001), present evidence on individual euro area countries—as well as the US and the UK—and conclude that the goodness of fit of the NKPC to inflation in these countries is not robust to changes in the specification and estimation procedures. Benigno and López-Salido (2002) estimate NKPCs for five individual countries in the euro area, namely Italy, Germany, France, Spain and the Netherlands, via GMM methods.[9] They find that the purely forward-looking model fits the data relatively nicely, but that there are important cross-country differences in the degree of inflation inertia, with German GDP deflator inflation ranking least inertial, and French and Italian inflation ranking most inertial. This result is in line with Coenen and Wieland (2000), who find that wage contracting structures that imply inertial components in inflation as in Fuhrer and Moore (1995) are better suited to capture the behaviour of

[8] In a recent contribution, Holden and Driscoll (2001) criticise the Fuhrer and Moore (1995) on the basis that this can only generate inflation persistence if workers are assumed to care about the past real wage of other workers. They demonstrate that making the more reasonable assumption that workers care about the current real wages of other workers, one obtains the standard (New Keynesian Phillips Curve) formulation with no inflation persistence. See also Rotemberg (1997).
[9] See Batini, Jackson and Nickell (2000) for evidence on the fit of NKPCs to UK data.

inflation high-inflation regime in countries such as Italy and France than in Germany, where a nominal wage contracting specification is to be preferred.[10]

Thus, although the ECB move towards a stability-oriented policy framework that takes inertia in inflation for granted is partly warranted by results on inflation persistence at the country level, it clashes with formal modelling for the aggregate euro area, which seems to prefer specifications where inflation persistence is not a structural, policy-invariant feature of the data.

In line with previous work by Batini and Nelson (2001) on US and UK data, this paper asks whether this conflict can be resolved by presenting model-free[11] evidence on the delay between changes in monetary policy and their peak effect on inflation for the euro area as well as in a subset of its component countries. It also presents evidence of the autocorrelation properties of inflation and on the lag in the response of inflation to monetary policy shocks from simulations of a battery of identified parsimonious VARs. Analysis of the characteristics of inflation both at the aggregate and at the country level controls for the impact of the aggregation bias on the measurement of persistence, and at the same time, to determine how intrinsic is persistence for the euro area as a whole.

The plan of the paper is as follows. Section 2 discusses various definitions of inflation persistence. It then reviews existing evidence for Europe as well as presents new evidence for the period 1970-2002, and debates what can be inferred from it about the bias in aggregation for the area-wide inflation data. Section 3 offers some concluding remarks and draws policy implications. A description of the data used and of their time-series properties is appended to the paper.

2. Measuring euro area inflation persistence

The economic literature is not unanimous about the definition of inflation persistence. To clarify matters, Batini and Nelson (2001) distinguished between three types of persistence: (1) positive serial correlation in inflation; (2) lags between *systematic* monetary policy

[10] An empirical analysis of NKPCs individual euro area countries at the sectoral level can be found in Sondergaard (2002). The empirical basis for NKPC at the sectoral level is investigated by Batini, Jackson and Nickell (2002) on UK quarterly data.
[11] I.e., purely statistical.

actions and their (peak) effect on inflation; and (3) lagged responses of inflation to *non-systematic* policy actions (i.e. policy shocks).[12]

The first type of persistence is a reduced-form property of inflation that manifests —simultaneously— the underlying pricing process, the conduct of monetary policy, and the expectations' formation process of price-setting agents. Changes in any of these three factors will influence the autocorrelation properties of inflation. This type of persistence has been documented in post-war data for various countries. In most cases, persistence measured according to this definition appeared to have been declining over time, especially following the onset of monetary regimes with strong anti-inflation credentials. In the US, for instance, Fuhrer and Moore (1995) detected high serial correlation of US inflation over the period 1965-1990, and used it to motivate the use for the US of the Phillips curves that impose inherent inertia in inflation. However, as shown using quarterly data by Sbordone (2002), Cogley and Sargent (2001a, b), Taylor (2000), Erceg and Levin (2001), and using monthly data by Batini and Nelson (2001), this correlation declined sharply after 1984, presumably as agents adjusted to the stabler Volcker-Greenspan monetary policy regime, weakening the empirical basis for assuming intrinsic inertia in inflation.[13] An alternative interpretation for the observed decline in the correlation of US inflation is that shocks hitting prices and quantities in the US have declined dramatically in amplitude in the past fifteen years.[14] Similarly, in the UK, inflation appears to have been strongly autocorrelated before 1992, the year of the adoption of an explicit inflation target. Yet, both Batini and Nelson (2001) and Kuttner and Posen (2001), find that there has been a dramatic decline in the autoregressive coefficient of monthly UK inflation since then.

[12] As Jeff Fuhrer pointed out to me, there is probably at least another important type of inflation persistence, namely the response of inflation to its 'own' shocks, as opposed to shocks to monetary policy innovations. I do not deal with this kind of persistence in this paper but I come back to it in my concluding remarks.

[13] These studies used various methods to examine the serial correlation properties of inflation, ranging from univariate to multivariate autoregressive representations.

[14] Evidence of the sharp decline in the volatility in US as well as world GDP and inflation over this period is discussed in McConnell and Perez-Quiros (2000), and Stock and Watson (2002), among others. The hypothesis that factors other than changes in policy regimes may be behind the breaks observed in inflation is endorsed, indirectly, by Sims (1999, 2001) who questions the evidence that policy rules have shifted in the post-Accord era. It is likely that with the more violent swings in inflation and real interest rates witnessed in Europe during the 1970s and 1980s, the absence of policy shifts is a less defensible position for the euro area. In more recent work (Sims and Zha (2002)), Sims is indeed able to reaffirm some aspects of conventional wisdom about changes in US monetary policy over time, but concludes that the changes in policy have been more subtle than dramatic.

Type 2 inflation persistence refers to the number of periods it takes for a change in monetary settings to have its maximum effect on inflation. Friedman (1972) pioneered the empirical analysis on this type of persistence for the US, looking at monthly data over the period 1953-1970. For that time span, he estimated that the money growth/CPI inflation relationship gave 'the highest correlation…[with] money leading twenty months for M1, and twenty-three months for M2' (p.15). In subsequent work, Friedman and Schwartz (1982) extended the analysis to the UK, and showed that monetary changes affect 'output after a brief lag (about six to nine months for the United States and the United Kingdom)…Later the impact shifts to prices (after another fifteen to twenty months for the United States and the United Kingdom' (p. 403). In their recent contribution, Batini and Nelson (2001) updated and extended Friedman's (1972) evidence on the lag between monetary policy actions and the response of inflation looking at UK and US data for the period 1953-2001 on monetary growth rates, inflation and interest rates, as well as annual data on money growth and inflation. They reaffirmed Friedman's result that it takes over a year before monetary policy actions have their peak effect on inflation both in the US and in the UK despite numerous changes in monetary policy arrangements in both countries.

Finally, type 3 of inflation persistence relates to the number of lags it takes for inflation to respond to a policy shock.[15] This type of persistence is often the only one consulted by economic modellers when validating models vis-à-vis the dynamics of real-world data generating processes. Rotemberg and Woodford (1997) and Christiano, Eichenbaum and Evans (2001), for instance, set their models' parameters so as to match US inflation (and output) responses to an estimated policy shock.

As discussed in Batini and Nelson (2001), a rigorous quantification of, combined with model accuracy regarding, type 2 inflation persistence is certainly the most relevant for monetary policymaking because this type of persistence determines the costs of disinflation. Knowing the length of the delay between policy action and their peak effect on inflation informs monetary policymakers on how pre-emptive they should be when responding to private sector shocks, and thereby helps to minimise the output gap variability costs of price stabilisation.

However, current practice in model evaluation does not attach much weight to type 2 persistence. Attention focuses mainly on type 1 or type 3 persistence. Yet a model that accounts for type 1 persistence does not necessarily account for type 2 persistence. Similarly, a model that accounts for type 3 persistence could fail to account for the delays in effect of the systematic component of policy, on the grounds that: (a) policymakers' decisions can hardly be characterised solely in terms of the non-systematic component of policy;[16] and (b) policy shocks are in any case not a major source of macrovariability (e.g. McCallum (1999) and Woodford (1998)) relative to technology and other real shocks, so their effect may not be a good guide to dynamic effects of policy.

In what follows, I briefly review existing evidence for the euro area and for three core countries within the area—notably Germany, France and Italy—on type 1 and type 3 persistence. I then provide fresh evidence for the period 1970-2002 for the area as a whole and new evidence for the period 1970-2000 for the individual countries. In addition, I present quantitative evidence on the extent of type 2 inflation persistence based on money growth rates, inflation and interest rates. The analysis of all types of persistence both for the region as a whole and for a subset of core countries enables me to infer indirectly whether the observed persistence in euro area inflation is a by-product of country data aggregation or an intrinsic euro area phenomenon.

2.1 Euro area and country-level type 1 inflation persistence

Evidence on type 1 inflation persistence for the euro area can be found in Galí, Gertler and López-Salido (2000, p. 1240. 'GGLS' hereafter). Regressing the log difference of the euro area GDP deflator on four lags of itself and detrended log real euro area GDP over the sample 1970 Q1-1998 Q2, they find that euro area quarterly inflation has been quite persistent in the past thirty years, with a weight on the value for the previous period of about 0.52. This is similar to what is found for US data (see Galí and Gertler (1999)). Related evidence is given by Coenen and Wieland (2000, 2002), who compute

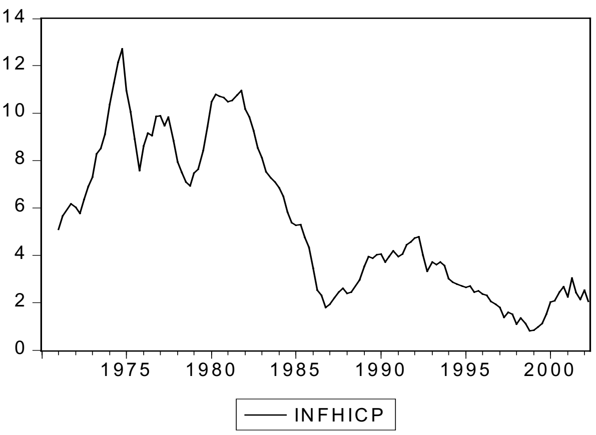
[15] See Christiano, Eichenbaum and Evans (1999) for an analysis of VAR evidence on the effects of monetary policy shocks.

[16] Indeed, major schools of economic thought, including monetarism and New Keynesianism assert that empirically, most real effects of monetary policy arise from the non-neutrality of policy responses to non-policy shocks (Woodford (1998)). In other words, the systematic component of policy has major effects too, so a model capable only of reproducing the data response to a policy shock is only partially validated against the data.

autocorrelation functions for GDP deflator inflation for the period 1974-1998 based on unconstrained estimated bivariate VARs in detrended inflation and the output gap.[17] Again, considering results obtained over the whole sample, they find that both detrended inflation and output are quite persistent, with significant positive autocorrelations out to lags of about 5 and 8 quarters.

Neither GGLS (2000) nor Coenen and Wieland (2000, 2002), however, discuss whether the serial correlation of euro area inflation has declined over time. Is this the case? Chart 1 below, plotting four-quarter changes in the HICP, seems to suggest that inflation may have behaved progressively less like a random walk as the sample progressed through the three major phases of European inflation, namely the post-Bretton Woods "Great Inflation" in the 1970s and 1980s, the disinflation of the early 1980s and the current period of low inflation.

Chart 1 HICP Inflation 1971 Q1-2002 Q2



To address this issue, in Table 1 and 2 below I provide more recent evidence on the first type of persistence for the whole euro area and for the three individual countries, respectively, in the form of univariate representations of annualised quarterly inflation in the Harmonised Index of Consumer Prices (π^{EA}_t) (euro area) or in each individual country's Consumer Price Index (π^{ITA}_t, π^{FRA}_t, π^{GER}_t for Italian, French and German CPI, respectively), since 1970.[18] In each case I present a regression of inflation on a constant, as well as split-

[17] Coenen and Wieland (2000, 2002) detrend inflation by taking deviations from a linear trend. They find that results obtained by taking deviations from an exponential rather than a linear trend are analogous.

[18] Datasets: For the euro area: ECB Area-Wide Model dataset, from Fagan et al (2001), series kindly updated for me to 2002 Q2 by Alastair Dieppe. For individual countries: macroeconomic database for the Monetary Transmission Network, kindly provided to me by Anna Maria Agresti. All individual country data are seasonally adjusted. However the HICP series in Fagan et al (2001) is not, so I adjust it using seasonal dummies. This may be preferable to other seasonal adjustment methods like the Census X-11 or X-12 which

sample AR(k) for $k = 5$ autoregressions for the inflation series. The regression on a constant provides useful summary statistics: its estimated parameter corresponds to the sample mean of inflation, while the residual standard error corresponds to inflation's standard deviation. The AR(5) specification of inflation summarises the degree of type 1 inflation persistence, with the estimated lagged-terms' autoregressive coefficient indicating the serial correlation of inflation in the data.[19]

For the euro area, the regressions are estimated over the sub-samples 1970 Q1-1984 Q2 and 1984 Q3-2002 Q2. There is substantial agreement in the literature in identifying a change in the nature of the ERM around 1984/1985 from a 'soft' to a 'hard' exchange rate parity arrangement (see Giavazzi and Spaventa (1990) and Angeloni and Dedola (1999)). Other events like the French abandonment of the 'encadrement du credit' and the move to the 'Franc fort' policy, as well as the 'divorce' of the Banca d' Italia from the Italian Treasury in the early 1980s are also believed to have contributed to the shift towards a more aggressive approach towards fighting inflation in Europe around that time, and hence to a break in the time series properties of inflation. The timing of the split is supported by results from a CUSUM stability test and by recursive Chow breakpoint tests.[20, 21] If I allow both the slope and the intercept of the equation to vary across these two regimes, the restriction of no structural change is rejected [F-statistics: $F(2, 124) = 12.94$, (p–value = 0.000)]. However, the restriction that the parameter non-constancy is confined to the equation's intercept is not rejected [F-statistics: $F(1, 124) = 0.539$, (p–value = 0.46)],[22]

utilise moving averages to remove seasonality and that may hence induce artificial persistence in the series that I analyse. Throughout the paper I look at evidence on HICP/CPI inflation, because the former (and hence, indirectly, the latter) are targets of monetary policy for the ECB. However results obtained using euro area and country-specific GDP deflators leads to similar conclusions to those put forth here.

[19] Given the quarterly frequency of the data and the existing evidence on autocorrelation in euro area inflation data, a $k = 5$ seems reasonable. See Coenen and Wieland (2000) for a discussion of the serial correlation properties of euro area inflation over the full 1970-1999 sample.

[20] The CUSUM test, by Brown, Durbin and Evans (1975), is based on the cumulative sum of recursive residuals. The test finds parameter instability if the cumulative sum goes outside the area defined by two 5% critical lines. The Chow breakpoint test instead, partitions the data in two (or more) subsamples. The test compares the sum of squared residuals obtained by fitting a single equation to the entire sample with the sum of squared residuals obtained when separate equations are fit to each subsample of data. Significant differences in the estimated equations indicate a structural change in the relationship. See Monticelli and Tristani (1999) for a discussion of the power of the Chow breakpoint test in the euro area context.

[21] This is not true of other frequently advocated timings like 1987 Q1—marking the end of exchange rate realignments in the ERM—for which the null of no structural change can be rejected at standard confidence levels.

[22] The intercept dummy takes the value 1.0 from 1984 Q3 onwards. The slope dummy is obtained by multiplying the equation's regressor by the intercept dummy.

indicating that the break might have been caused simply by a shift in the mean of inflation—likely reflecting a change in the implicit objective for price stability.

In contrast with findings for the US and the UK, Table 1 indicates that there has been very little change in the serial correlation of HICP inflation over time, with the AR(5) lagged-terms' coefficient sum for inflation actually rising from just below 0.7 before 1984, to just over 0.7. On the other hand, Table 1 shows that the variability of euro area inflation has dropped moving from the first to the second sample period (with inflation's standard deviation declining from 2.3% to 1.4%). Importantly, as illustrated by the regressions of inflation on constants alone, the mean of inflation has gone down significantly between periods, falling from a yearly average level of around 8.5% to 2.8% after 1984[23]—in accordance with the suggestion from the previous slope dummies' exclusion restrictions tests.

Table 1: Regression Evidence on Type 1 Inflation Persistence – Euro Area Aggregate (HICP)

Euro Area

Sample Period: 1970 Q1– 1984 Q2

$$\underset{}{\pi_t^{EA}} = \underset{(0.003)}{0.085},\ R^2 = 0,\ SEE = 0.023,\ DW = 0.62.$$

$$\pi_t^{EA} = \underset{(0.003)}{0.007} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{EA}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.149)}{0.679},\ R^2 = 0.478,\ SEE = 0.005,\ DW = 1.82.$$

Sample Period: 1984 Q3– 2002 Q2

$$\pi_t^{EA} = \underset{(0.002)}{0.028},\ R^2 = 0,\ SEE = 0.014,\ DW = 0.92.$$

$$\pi^{EA}{}_t = \underset{(0.001)}{0.002} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{EA}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.104)}{0.735},\ R^2 = 0.447,\ SEE = 0.003,\ DW = 1.98.$$

A number of reservations about empirical results with euro area data have been advanced. Prominent reservations include: (i) the data for the period preceding monetary union are synthetically derived by averaging data from the member countries; (ii) the member countries experienced different (and at times diverging) monetary policy regimes before

1999, and so an analysis of aggregate euro area developments is meaningless. So Table 2 below presents comparable evidence for Italy, France and Germany—the three largest EMU members in terms of area-wide GDP shares. Again, sample splits are chosen in line with times when it is plausible that each central bank changed its course of policy decisively towards fighting inflation.

More specifically, for Italy, I break the sample into 1970 Q1-1981 Q2 and 1981 Q3-2000 Q4, a split suggested by the 'divorce' of the Banca d'Italia and the Italian Treasury. This event put a halt to public debt monetisation in Italy, and so allowed the control of inflation to become a major focus of monetary policy (see Passacantando (1996) and Clarida, Galí and Gertler (1998, henceforth 'CGG')). For France, I chose 1984 Q3 as a breakpoint for the sample split. Around this time the policy of 'encadrement du credit' was *de facto* abandoned, and a gradual move away from the policy of a weak Franc started. So this date can be taken to mark, approximately, Banque de France's abjuration of its accommodating attitude towards inflation (see Artus, Avouyi-Dovi, Blenze and Lecointe (1991), and more recently, Mojon (1997, 1999) and CGG (1998)).[24] Finally, for Germany I partition the sample into 1970 Q1-1986 Q4 and 1987 Q1-2000 Q4. Although CGG (1998) take 1979 Q1, i.e. the time Germany when entered the ERM, as a watershed for German monetary policy pre- and post-oil shocks (with a renewal of its commitment to keep inflation low), I find less evidence of a break for this period. Indeed, German inflation reaches pre-1979 levels several times between 1979 and 1987, signalling no discontinuity, at least in data outcomes, in the properties of German inflation post-1979.[25] Rather, I find that, with Germany as a leader country, the change in the nature of the ERM arrangement from 'soft' to 'hard' plays more of an important role in the shift in the policy of the Bundesbank over the whole sample. One rationale for this is that, in 1987 the wedge between French and German money market interest rates started to shrink—a sign that the ERM set-up was increasingly more credible. This might have implied that since then, the Bundesbank could focus more

[23] This may explain also why the variance of inflation has dropped over time. There is, in fact, considerable evidence that inflation variability and the level of inflation are positively related across countries. Davis and Kanago (2000) review this evidence for the OECD countries.

[24] As for the euro area, in the case of Italy and France I find that I cannot reject the hypothesis of no structural change when I test for 1987 Q1—a date marking the end of realignments in the ERM—as a possible break (p–value = 0.46 and 0.48, respectively).

[25] Muscatelli, Tirelli and Trecroci (1998) discuss the dating of the institutional change in Germany. In practice, a Chow breakpoint test supports the presence of at least two breaks, one in 1979 Q4 (or in the previous 3 quarters) and one in 1987 Q1. But the test for the 1987 Q1 break is significant at the 1% level while the one for the 1979 break only at the 5% level, when these breaks are taken in isolation.

and more on price stability rather than moving rates in defence of the FF/DM parity.[26] Another—albeit far less plausible— possibility is that the switch by the Bundesbank in 1987 from announcing targets for the growth of Central Bank Money to *M3* implied a material change in policy stance at the time (see Neumann and von Hagen (1993), von Hagen (1995) and Issing (1997)).

As for the euro area data, in all cases the timing of the splits is backed by results from CUSUM stability tests and by recursive Chow breakpoint tests. As in the euro area case, if I allow both the slopes and the intercepts of the CPI inflation equations to vary across regimes, the restriction of no structural change is rejected.[27] However, for Italy and France, the restriction that the parameter non-constancy is limited to the equations' intercepts is not rejected,[28] suggesting that, in those countries, the breaks could be interpreted as mere shifts in the mean of CPI inflation. For Germany, by contrast, non-constancy post-1987 (but not post-1979) seems to extend to the equation's slope, since that restriction can be rejected at the 1% confidence level.[29]

Overall, results from Table 2 support those obtained using aggregate euro area data, but highlight important differences across countries in the extent of changes in persistence over time. More specifically, not only there seems to have been no reduction in the serial correlation of inflation over time for Italy or France, but this seems to have increased. The lagged-terms' coefficient sum for Italian inflation goes from around 0.7 to over 0.8 after the 1981 break. Similarly, there is a small rise in the serial correlation of French inflation, with the autoregressive coefficients' sum going from 0.69 to 0.7, before and after 1984. In contrast, for Germany, the serial correlation drops dramatically: from above 0.9 before 1987, to insignificantly different from zero. Furthermore, in contrast to the German case where inflation has approximately the same variance before and after the break, French and Italian inflation are way less volatile after the breaks (falling from 6.4% to 4.0% in Italy and from 3.0% to 1.6% in France).

[26] See Issing (1997) and Hetzel (2002) for a discussion of German monetary policy in the second half of the twentieth century.

[27] F-statistics: Italy: $F(2, 118) = 5.95$, (p–value = 0.00); France: $F(2, 118) = 10.15$ (p–value = 0.00); Germany (including dummies for both 1979 Q4 and 1987 Q1): $F(4, 116) = 4.86$, (p–value = 0.00).

[28] F-statistics: Italy: $F(1, 118) = 1.63$, (p–value = 0.20); France: $F(1, 118) = 0.723$ (p–value = 0.40).

[29] Joint restriction on 1979 Q4 and 1987 Q1 slope dummies: $F(2, 116) = 4.17$, (p–value =0.02). 1979 Q4 slope dummy only: $F(1, 116) = 3.19$, (p–value = 0.08). 1987 Q1 slope dummy only: $F(1, 116) = 8.31$, (p–value = 0.00).

For all countries, as in the case of the euro area, there is a substantial drop in the mean of inflation. The largest drop is for Italian inflation, going from a yearly average of 13.2 % to 5.9% in the most recent period. French inflation goes from 9.4% to 2.4% on a yearly basis; whereas German inflation almost halves on average falling from 4.1% to 2.2% (with the new mean level close to the 2.0% top range of the ECB's inflation target for yearly changes in HICP).

In summary, according to this first definition of inflation persistence, it appears that the degree of inertia in European inflation has not changed much over the past thirty years, a finding not always established at the country-level. In particular, there appears to have been a sizeable shift in inflation persistence in Germany, but only a little shift in the mean of German inflation; whereas Italy and France saw primarily large shifts in their average levels of inflation—which explains the the drop in the mean of euro area inflation—yet little if no variation in persistence over time.[30]

2.2 Euro area and country-level type 2 inflation persistence

In this section I offer some relatively model-free quantitative evidence on the extent of type 2 inflation persistence. The economic profession seems to have achieved a consensus on how to measure the non-systematic component of monetary policy,[31] but not on how to measure its systematic component, so neither the selection of the policy stance measure for this purpose, nor the appropriate statistic to calculate, is a straightforward issue. In particular, because the systematic component of policy is inherently endogenous, many of the conventionally desirable properties of measures of policy change, such as exogeneity, are inappropriate.

I follow Friedman (1972) and Batini and Nelson (2001) by using the correlation of inflation with money growth $k \geq 0$ periods earlier, a statistic denoted $\rho_{\pi\mu}(k)$, as one means of summarising evidence on type 2 inflation persistence. In using monetary aggregates for this purpose, I take no stand on whether money has a special role in the transmission mechanism. Rather, I view money as a quantity-side measure of the monetary conditions

[30] As discussed in Batini (2002), large shifts in the general level of inflation have implications for the reliability of NKPC, because these are usually derived assuming Calvo (1983) contracts, and this in turn presumes that firms face a fixed probability of resetting prices—which is unlikely in presence of such shifts.
[31] This consensus is not unanimous. Skeptical accounts of impulse response function analysis as a tool to evaluate policy can be found, for instance, in Ericsson et al (1998).

induced by central bank interest-rate policy. For example, open-market operations to alter short-term nominal interest rates tend also to change the growth rates of reserves and the money stock. On the other hand, one concern is that changes in the opportunity cost of holding money—such as an increase in the own-rate on broad money aggregates after financial liberalisation, or greater incentives for the private sector to hold purchasing power in the form of base money after a disinflation—potentially distort money growth.

Table 2: Regression Evidence on Type 1 Inflation Persistence – Individual Countries (CPIs)[1]

Italy

Sample Period: 1970 Q1– 1981 Q2

$$\underset{(0.009)}{\pi_t^{ITA} = 0.132}, \; R^2 = 0, \; SEE = 0.0637, \; DW = 0.64.$$

$$\pi_t^{ITA} = \underset{(0.025)}{0.051} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{ITA}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.177)}{0.682}, \; R^2 = 0.408, \; SEE = 0.048, \; DW = 2.01.$$

Sample Period: 1981 Q3– 2000 Q4

$$\underset{(0.004)}{\pi_t^{ITA} = 0.059}, \; R^2 = 0, \; SEE = 0.0397, \; DW = 0.28.$$

$$\pi_t^{ITA} = \underset{(0.003)}{0.005} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{ITA}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.046)}{0.845}, \; R^2 = 0.848, \; SEE = 0.016, \; DW = 1.80.$$

France

Sample Period: 1970 Q1– 1984 Q2

$$\underset{(0.004)}{\pi_t^{FRA} = 0.094}, \; R^2 = 0, \; SEE = 0.0305, \; DW = 0.56.$$

$$\pi_t^{FRA} = \underset{(0.014)}{0.031} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{FRA}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.142)}{0.687}, \; R^2 = 0.433, \; SEE = 0.023, \; DW = 2.00.$$

Sample Period: 1984 Q3– 2000 Q4

$$\underset{(0.002)}{\pi_t^{FRA} = 0.024}, \; R^2 = 0, \; SEE = 0.016, \; DW = 0.83.$$

$$\pi_t^{FRA} = \underset{(0.003)}{0.006} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{FRA}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.095)}{0.698}, \; R^2 = 0.480, \; SEE = 0.011, \; DW = 2.01.$$

(1) Standard errors in parenthesis.

Table 2: Regression Evidence on Type 1 Inflation Persistence – Individual Countries (CPIs) Continues[1]

Germany

Sample Period: 1970 Q1– 1986 Q4

$$\underset{(0.003)}{\pi_t^{GER} = 0.041}, R^2 = 0, SEE = 0.028, DW = 0.91.$$

$$\pi_t^{GER} = \underset{(0.006)}{-0.000} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{GER}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.135)}{0.948}, R^2 = 0.666, SEE = 0.017, DW = 1.96.$$

Sample Period: 1987 Q1– 2000 Q4

$$\underset{(0.003)}{\pi_t^{GER} = 0.022}, R^2 = 0, SEE = 0.023, DW = 1.77.$$

$$\pi_t^{GER} = \underset{(0.006)}{0.013} + \sum_{i=1}^{5} \hat{b}_i \pi_{t-i}^{GER}, \quad \sum_{i=1}^{5} \hat{b}_i = \underset{(0.236)}{0.431}, R^2 = 0.146, SEE = 0.023, DW = 1.96.$$

(1) Standard errors in parenthesis.

My calculation of $\rho_{\pi\mu}(k)$ across sub-samples allows for changes in steady-state velocity growth due to these factors. But in light of reservations about money growth, I also present correlations of inflation with r_t—the short-term real interest rate for the euro area—a variable chosen to capture the idea that monetary policy can influence the real rate over short periods.[32] In addition, I report results using a term-structure-based measure of monetary conditions.

Another concern in estimating dynamic relations between measures of systematic policy and inflation is that, if monetary policy adjusts completely and successfully to offset non-policy shocks, there should be no observed relation between policy measures and inflation. Several considerations, however, suggest that in practice such a relation will be present, that e.g. movements to monetary policy ease will be manifested in some upward movements in inflation in the future. Long-standing deviations of policymakers' specification of the economy from the true underlying economic process will tend to produce target misses that are attributable to policy actions.[33] Policy objectives other than

[32] Throughout, my r_t series is the quarterly average respective nominal rate minus $E_t\pi_{t+1}$, where π_t is quarter-on-quarter inflation. For all countries that I consider, the expectations $E_t(\bullet)$ are approximated by OLS projections of π_{t+i} on lags 1-4 of π_t and HP-filtered log real GDP (filter parameter 1,600), plus dummies for regime shifts.

[33] Prior to the 1970s, such specification errors might have included belief in a non-vertical Phillips curve and an overemphasis on 'special factors' theories of inflation. More recently, a candidate for specification error is

deviations of inflation from target (e.g. as occurs under flexible inflation targeting) tend to make it optimal to move policy in such a way that persistent but temporary deviations from target occur.[34] And the variability in the precise lag in effect of policy means that some target misses will be due to prior policy decisions. For all these reasons, in countries where the remit for monetary policy is price stability, some systematic deviations of inflation from target will be associated with systematic policy actions.

Table 3 lists the maximum values of *k* for 1970-2002 for selected sub-periods using four-quarter changes of euro area *M3*—the broad money aggregate in which is expressed the reference value for the 'first pillar' of the ECB's stability-oriented medium-term strategy[35]—and corresponding changes in euro area harmonised consumer prices—basis of the reference value for the 'second pillar'.[36] In most cases, sub-period correlations are very strong and highly significant both under standard and Newey-West-adjusted levels of confidence. The results with the interest-rate-based measure of policy largely support the timing evidence using money growth. This is true also when I use German *ex ante* short-term real interest rates, instead of the artificially-backed euro area rates, to account for the fact that these may be more representative for Europe—given the leading role of the Bundesbank within the ERM at least since the launch of the Exchange Rate Mechanism in 1979— although with German rates, peak correlations occur uniformly at slightly longer

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that the output gap series used in policymaking is conceptually very different from the output gap that is used in the theory underlying the NKPC.

[34] Rudebusch and Svensson (2002), for example, assume that the Eurosystem policymakers' objective function is similar to that of the Federal Reserve Board and penalise volatility both in inflation *and* in the output gap. Similarly, Coenen and Wieland (2000, 2002) assume that the Governing Council of the ECB has a flexible inflation target, i.e. acts to minimise deviations of inflation *and* output from desired values.

[35] See Issing et al. (2001) for a review of the monetary policy framework of the European Central Bank. In contrast to Batini and Nelson (2001), who compute correlations between narrow money growth and inflation for the UK and the US, here I do not look at evidence from such correlations because, at the time of writing, a narrow money series for the euro area going back to the 1970s is not available. Besides, the (short-run) stability and information content of euro area narrow money aggregates, like *M1*, have been questioned by a number of authors (see ECB (1999) and Coenen and Vega (1999)), so examining this extra evidence may not add any value to what is uncovered here.

[36] It is hard to assess times of possible past regime shifts for the euro area as a whole, since this only came into existence in 1999. So the idea here is to select sub-samples on the basis of developments in the ERM (the European monetary arrangement which pre-dated the European Monetary Union), as well as monetary events in major European countries which may have had implications for trend money velocity for the area as a whole. I chose the following sub-periods: 1970 Q1- 1978 Q4 (pre-ERM period); 1979 Q1- 1986 Q1 ('soft' ERM period); 1979 Q1- 1989 Q4 (ERM period, pre-German unification); 1991 Q1-2002 Q2 (ERM, post-German unification period); 1993 Q1-2002 Q2 (ERM/pre-EMU, post-German unification with lag). The logic behind the pre- and post-German unification split is that this has generated a large spike in German, and indirectly, in euro area money growth which may well have distorted average money velocity for the area.

lags.[37] Finally, in Table 3 I do not report correlations between inflation and measures of prior *ex ante* interest rates for the two last samples because these have a positive rather than the correct (negative) sign. This is most likely due to the simultaneous contraction in nominal interest and inflation rates differentials between Italy and France, on one side, and Germany, on the other side, over the second half of the 1990s, which imparted similar trends to inflation and real interest rates over that period for the corresponding euro area aggregates.[38]

Both for the period as a whole and for sub-samples, the euro area evidence suggests that money growth leads inflation by over a year. If anything, the lead of money growth over inflation is somewhat longer in recent decades. For 1970-2002, the lead is of the order of 7 quarters, slightly less than 2 years—a finding that reaffirms prior empirical results on the relationship between broad money growth and inflation in the euro area (see ECB (1999) and Nicoletti Altimari (2001)). The 1979-1986 and 1970-1989 period also suggest a long lead, with a peak of $\rho_{\pi\mu}(k)$ at k =10 and k = 9 respectively. For the period immediately following German reunification (which I take as starting in 1991 Q1) the lead of *M3* growth on HICP inflation is slightly shorter, at k = 6. This may reflect effects on measured average velocity in the euro area following the German reunification, which distorted the relationship between inflation and prior monetary change for data that overlap the pre- and post-regime change period. This explains why the sample starting in 1993, with data on both money and inflation generated well after German reunification, produces the maximum *M3*/inflation correlation at k =11, a lag analogous to those uncovered for samples pre-1989.

A remarkable message from Table 3 is the resilience of the relationship between inflation and prior *M3* growth to changes in sample period. Most notably, the correlation remains

[37] As a further robustness check, I have also derived dynamic correlations between *changes* in the annualised quarterly euro area and German *ex ante* short-term real interest rates and four-quarter changes of inflation. Using changes in r rather than its level may have the advantage that changes in r are not dominated by longer-term swings in the mean of r, which are likely to be determined by non-policy factors; and that cross correlations with inflation are less affected by the arithmetic link between the real rate and future inflation from the Fisher relation. However, I have found that, although all correlations exhibit the 'correct' (negative) sign, they are less strong (reaching, as a maximum, the value of – 0.523) and less significant than when I use *levels* of the corresponding variables.

[38] The fall in differentials was in turn driven by the gradual disappearance of the inflation premium incorporated in Italian and French short-term nominal interest rates relative to German rates ahead of the formation of a monetary union. See Gros and Thygesen (1992), De Grauwe (1996, 1997), Favero, Giavazzi and Spaventa (1997) and Angeloni and Dedola (1999), among others, for more details on the behaviour of interest and inflation rates differentials across Europe before the creation of European Monetary Union.

sizeable, positive and significant even for sub-samples that completely exclude the oil-shock-dominated pre-1980 data. These findings reaffirm the idea that money growth may serve adequately as a simple summary measure of monetary policy stance, and may have a strong relationship with future inflation, without implying that money contains any incremental information relative to real income growth and measures of excess demand. In this sense my results, based on the unconditional or bivariate relationship between inflation and prior money growth, are compatible with those in Trecroci and Vega (2000) and Gerlach and Svensson (2002), who evaluate the significance of area-wide broad money on the basis of the explanatory power it contributes to a regression for inflation after controlling for lagged inflation and lagged real GDP growth.[39]

Table 3: Correlations between HICP Inflation and Measures of Systematic Monetary Policy

Sample period	Monetary policy measure: Four-quarter money (*M3*) growth Maximum value of $\rho_{\pi\mu}(k)$	Monetary policy measure: short real rate Max. neg. value of $\rho_{\pi r}(k)$
	Euro Area	
1970 Q1 – 2002 Q2	0.868 [a,b] (k = 7)	– 0.863 [a,b] (k = 20)
1970 Q1 – 1978 Q4	0.336 (k = 6)	– 0.480 [a] (k = 7)
1979 Q1 – 1986 Q4	0.844 [a,b] (k = 10)	– 0.808 [a,b] (k = 8)
1979 Q1 – 1989 Q4	0.876 [a,b] (k = 9)	– 0.877 [a,b] (k = 10)
1991 Q1 – 2002 Q2	0.825 [a, b] (k = 6)	–
1993 Q1 – 2002 Q2	0.753 [a, b] (k = 11)	–

(a) Significantly different from zero using conventional *t*-test.
(b) Significantly different from zero using Newey-West *t*-test.

For individual countries, results for the 1970 Q1-2000 Q4 period suggest a lead of money growth of 11 quarters over inflation for France and Germany, and a corresponding lead of 3 quarters for Italy (see Table 4 below).[40] In other words, it seems as if the seven-quarter lead between aggregate *M3* and HICP inflation uncovered for the entire euro area is a halfway house between these disaggregated data extremes. Excluding the 1970s gives a longer lead for Italy and Germany (with k = 10 and k = 13, respectively), but a much shorter lead for France (k = 3). Finally, from the early 1990s onwards, the lead for money growth over

[39] Trecroci and Vega (2000) conduct Granger causality tests within a cointegrated VAR system comprising real money holdings, real GDP inflation and short- and long-term interest rates. They find little empirical support for rejecting Granger non-causality of money on inflation at standard confidence levels within that information set.

[40] Subsamples for individual countries are as follows. Italy: 1970 Q1-1981 Q2 (pre-BI 'divorce' period); 1981 Q3-1986 Q4 ('soft' ERM); 1987 Q1-1992 Q2 ('hard' ERM); 1992 Q3-2000 Q4 (pre-EMU period). France: 1970 Q1-1984 Q2 (pre-'largesse' period); 1984 Q3-1993 Q2 ('hard' ERM, pre-BF independence); 1995 Q1-2000 Q4 (post-BF independence with lag, pre-EMU). Germany: 1970 Q1-1978 Q4 (pre-ERM); 1979 Q1-

Italian CPI inflation is found to be over three years; less than a year for France; and about one and a half years for Germany, signalling that during the pre-EMU years, the lag between monetary policy actions and inflation lengthened substantially in Italy, whereas it shortened in both Germany and France. One possible explanation for these divergences may be found in the relative differences between consumption vis-à-vis output patterns across these three countries during that period, which may have caused in some cases relative growth in transaction demand for money to diverge from growth in nominal GDP or prices. In all circumstances, the results indicate that also at the individual country level, there still remains a clear delay in the reaction of inflation to changes in policy stance. This indicates that, in contrast to inflation persistence of the type 1 form which, notably in the case of Germany, did not appear to be invariant to the monetary policy rule, some underlying inflation persistence of the type 2 form (lasting around 1 year) is an inherent feature of way the single economies work, and is not an illusion generated by a particular policy rule.

Once more, the results with the interest-rate-based measure of policy largely support the timing evidence using money growth. Note however, that, like in Table 3, I do not report measures of correlation between short real rate and inflation for Italy and Germany for the 1990s. As discussed before for the aggregate case, these tend to exhibit a positive sign because of the (EMU convergence-driven) downward co-movement of nominal short-term interest rates and inflation over the second half of the 1990s, and so have no clear interpretation from the point of view of this analysis.

Table 5 and 6 check the robustness of the previous results to the use of different interest-rate-based measure of monetary policy, namely the spread of the nominal short-term interest rate over the nominal long-term interest rate. Galí, Gertler and López-Salido (2002) argue that 'the yield spread may be thought of as a rough measure of the stance of policy'. The expected relationship between this measure of the yield spread and inflation is negative. For example, suppose that the monetary authority attempts to peg the short-term interest rate in the face of expanding aggregate demand. The higher demand pressure will raise market-determined rates such as the long-term bond rate, and the spread of short rates over longs will fall, correctly indicating a loosening of monetary conditions. Nevertheless

1986 Q4 ('soft' ERM); 1987 Q1-2000 Q4 ('hard' ERM, pre-EMU); 1991 Q1-2000 Q4 (post-unification, pre-EMU).

one can imagine circumstances where the spread changes even though the degree of excess demand in the economy is unchanged so that the term-structure-based measures become imperfect indicators of monetary stance.

Table 4: Correlations between CPI Inflation and Measures of Systematic Monetary Policy

Sample period	Monetary policy measure: Four-quarter money (M3) growth Maximum value of $\rho_{\pi\mu}(k)$	Monetary policy measure: short real rate Max. neg. value of $\rho_{\pi r}(k)$
	Italy	
1970 Q1 – 2000 Q4	0.817 [a,b] (k = 3)	– 0.740 [a,b] (k = 16)
1970 Q1 – 1981 Q2	0.434 [a,b] (k = 4)	– 0.884 [a] (k = 4)
1981 Q3 – 1986 Q4	0.635 [a,b] (k = 10)	– 0.869 [a,b] (k = 12)
1987 Q1 – 1992 Q1	0.368 (k = 3)	– 0.550 (k = 10)
1992 Q3 – 2000 Q4	0.868 [a,b] (k = 14)	–
	France	
1970 Q1 – 2000 Q4	0.802 [a,b] (k = 11)	– 0.872 [a,b] (k = 18)
1970 Q1 – 1984 Q2	0.369 [a,b] (k = 16)	– 0.715 [a] (k =4)
1984 Q3 – 1993 Q2	0.708 [a,b] (k = 3)	*n.a.* [c]
1995 Q1 – 2000 Q4	0.279 (k = 2)	*n.a.* [c]
	Germany	
1970 Q1 – 2000 Q4	0.509 [a,b] (k = 11)	– 0.712 [a] (k = 20)
1970 Q1 – 1989 Q4	0.460 [a,b] (k = 7)	– 0.434 [a,b] (k = 4)
1979 Q1 – 1986 Q4	0.646 [a,b] (k = 13)	*n.a.* [c]
1987 Q1–2000 Q4	0.540 [a,b] (k = 6)	*n.a.* [c]
1991 Q1 – 2000 Q4	0.594 [a,b] (k = 6)	–

(a) Significantly different from zero using conventional *t*-test.
(b) Significantly different from zero using Newey-West *t*-test.
(c) Not Available. Indicates that the size of the sample does not provide enough observations to compute the correlation given the length of the lag.

Despite this caveat, results for the euro area in Table 5 and for individual countries in Table 6 do support—with some qualifications—my earlier findings. For the most recent period, movements in HICP inflation take (well) over a year on average to follow movements in systematic monetary policy. This lead-lag relationship appeared to be much shorter in the 1970s and 1980s, giving a short lead (k =1) between developments in the yield spread and inflation for the sample as a whole. At the individual country level, once more I find differences between the length of transmission in Italy, on one side, and Germany and France, on the other side, relative to the 1990s, with Italy displaying a longish lead of the spread over inflation (k =13), whereas France and Germany a considerably shorter one (k = 3).

Table 5: Correlations Between HICP Inflation and Yield Spread

Sample period	Monetary policy measure: Nominal short-term rate minus long-term government bond rate Maximum neg. value of $\rho_{\pi,sp}(k)$
	Euro Area
1970 Q1 – 2002 Q2	– 0.498 [a,b] (k = 1)
1970 Q1 – 1978 Q4	– 0.808 [a,b] (k = 2)
1979 Q1 – 1989 Q4	– 0.705 [a,b] (k = 2)
1991 Q1 – 2002 Q2	– 0.643 [a,b] (k = 9)
1993 Q1 – 2002 Q2	– 0.703 [a,b] (k = 9)
Memo: 1979 Q1 – 1986 Q4	– 0.791 [a,b] (k = 2)

Table 6: Correlations Between CPI Inflation and Yield Spread

Sample period	Monetary policy measure: Nominal short-term rate minus long-term government bond rate Maximum neg. value of $\rho_{\pi,sp}(k)$
	Italy
1970 Q1 – 2000 Q4	– 0.492 [a,b] (k = 2)
1970 Q1 – 1981 Q2	– 0.669 [a,b] (k = 2)
1981 Q3 – 1986 Q4	– 0.451 [a,b] (k = 7)
1987 Q1 – 1992 Q1	– 0.569 [a,b] (k = 12)
1992 Q3 – 2000 Q4	– 0.435 [a,b] (k = 13)
	France
	Maximum neg. value of $\rho_{\pi,sp}(k)$
1970 Q1 – 2000 Q4	– 0.234 [a] (k = 2)
1970 Q1 – 1984 Q2	– 0.568 [a,b] (k = 2)
1984 Q3 – 1993 Q2	– 0.635 [a,b] (k = 10)
1995 Q1 – 2000 Q4	– 0.368 (k = 3)
	Germany
	Maximum neg. value of $\rho_{\pi,sp}(k)$
1970 Q1 – 2000 Q4	– 0.453 [a,b] (k = 2)
1970 Q1 – 1978 Q4	– 0.434 [a,b] (k = 4)
1979 Q1 – 1986 Q4	– 0.635 [a,b] (k = 10)
1991 Q1 – 2000 Q4	– 0.368 (k = 3)

(a) Significantly different from zero using conventional t-test.
(b) Significantly different from zero using Newey-West t-test.

2.3 Euro area and country-level type 3 inflation persistence

As final evidence of inflation persistence in the euro area, I examine type 3 inflation persistence, that is, the lag that exists between the peak response of inflation and the time of the monetary policy shock which generated it.

As I discussed in the Introduction, this measure of inflation persistence is presumably not the most relevant for monetary policymaking because policy shocks are only a minor determinant of inflation and output fluctuations when compared to technology or other real shocks. Nevertheless, matching the dynamic properties of theoretically-derived models of inflation to the data is a popular way of validating theory-based models for policy analysis.[41] For this reason, evidence of type 3 persistence for the euro area can be found, alongside other validation methods, in a number of studies that either estimate or calibrate macromodels to capture aggregate euro-area data dynamics. Influential studies reporting on this type of persistence for the euro area include, for instance, Fagan, Henry and Mestre (2001), Smets and Wouters (2002), Coenen (2002) and Monticelli and Tristani (1999).[42] These models differ considerably both in the way they are derived and in the way they are estimated. More specifically, the models in Coenen (2002) and Monticelli and Tristani (1999) are identified vector autoregressions; the model in Fagan, Henry and Mestre (2001) is a macroeconometric model estimated via OLS, equation by equation. This model uses cointegration methods to estimate steady state relationships between variables and so has a long-run structure; whereas the model in Smets and Wouters (2002) is a dynamic stochastic general equilibrium model estimated via Bayesian techniques. Smets and Wouters (2002) provide the only model allowing for forward-looking behaviour.[43] In addition, although all these models are quarterly, the sample period over which they were estimated varies.[44]

One consequence of their heterogenous theoretical assumptions and estimation methods is that the impulse response functions of annualised inflation (expressed in terms of the GDP deflator) to a monetary policy shock of all these models differ substantially. The AWM

[41] See, among many others, Rotemberg and Woodford (1997) and Christiano, Eichenbaum and Evans (2001).
[42] Peersman and Smets (2001) also estimate an identified VAR on aggregate euro area data. However, their model specifies the behaviour of the price level and not inflation, so responses for inflation—which are needed to quantify type 3 persistence—are not shown, and so cannot be reported and/or compared here.
[43] The Fagan, Henry and Mestre (2001) Area-Wide Model (AWM) is currently being augmented with forward-looking components.

model of Fagan et al. (2001), together with the model of Smets and Wouters (2002), seems to suggest that inflation's maximum response to a monetary policy shock occurs within 4 or 5 quarters from the shock.[45] Depending on the wage contracting specification, the model in Coenen (2002) implies that the peak response in inflation occurs approximately 5 (Taylor contracts) to 10 quarters (Fuhrer and Moore contracts) after the shock.[46] By contrast, the structural VAR of Monticelli and Tristani (1999) indicates that following a monetary policy shock, inflation displays a small and short-lived movement on impact with the actual peak effect materialising after almost 3 years (12 quarters).

In what follows I estimate via Constrained Full Information Maximum Likelihood a battery of restricted parsimonious VARs for the euro area and the three countries examined in the previous sub-sections.[47] The way I specify these differs, yet again, from existing estimated models. The main points of departure from prior empirical studies are: (i) the minimalist, but policy-relevant constraints that I impose; (ii) the data that I use—which is filtered/detrended in an alternative way relative to available studies; (iii) the sample—which is quarterly in frequency and goes from 1970 Q1 to 2002 Q2 for the euro area and to 2000 Q4 for the single countries; and finally, (iv) the inclusion of regime shift dummies—which have not been incorporated in previous works.[48]

More specifically, I look at two-lag VARs, estimated over different sub-samples within the larger data-set sample. The choice of using two lags seems enough to capture most of the dynamics I am after (longer lag structures give similar persistence results across the board), minimises the problem of over-parametrisation over the shortest samples as well as

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[44] In particular, these works' respective sample periods are: Monticelli and Tristani (1999): 1978-1997; Fagan et al (2001): 1970-1999; Smets and Wouters (2002): 1970-1999; Coenen (2002): 1974-1998 (output and inflation equations), 1979-1998 (monetary policy rule).

[45] Strictly speaking, inflation responses to a nominal interest rate shock are not comparable for these two models because the experiment in Fagan et al. (2001) assumes that the shock is sustained for two years, whereas the shock discussed in Smets and Wouters (2002) lasts only one period. Indeed, simulation of a more updated version of the AWM in McAdam and Morgan (2001), where a Taylor rule is implemented after the temporary policy shock, gives much longer lags between the shock and the peak response in inflation (3 to 6 years depending on whether a fiscal rule is also in place and whether forward-looking elements in this new version of the model are allowed to operate).

[46] The shock in Coenen (2002) is a 50 basis point shock to the innovations of the estimated policy rule.

[47] These are estimated VARs where I test down in order to drop insignificant regressors. Note that Mojon and Peersman (2001) also estimate an identified VAR on aggregate euro area data. However, as in Peersman and Smets (2001), their model specifies the behaviour of the price level and not inflation, so responses for inflation—which are needed to quantify type 3 persistence—are not shown, and so cannot be reported and/or compared here.

[48] One exception being Fagan et al. (2001), who build in dummies for the money demand function, to keep with findings in Coenen and Vega (2001) on aggregate euro area broad money growth data.

warrants parsimonious representations of the data generating processes. However, on rare occasions, lengthier lag specifications cannot be formally rejected by means of conventional information criteria.[49] There are three endogenous variables in each model: the log of the deviation of GDP at factor costs from its Hodrick-Prescott trend (smoothing parameter 1,600) ($ybar_t$); the deviation of annual HICP (or CPI) inflation from an implicit/explicit inflation target ($\pi^{dev}{}_t$);[50] and the nominal interest rate, measured as an annualised fraction ($4*R_t$).[51]

As Rudebusch (1998) observes, the interest rate equation in a simultaneous equation model (or VAR) has a structural interpretation as a monetary policy reaction function. As he also notes, however, there is a danger that policy regime shifts may produce non-constant parameter estimates. As emphasised in subsections 2.1 and 2.2 above, over the entire sample, there have been several major breaks in the euro area's monetary policy regime, so my analysis is potentially vulnerable to this drawback. To minimise this danger I do two things. First, I carry out my estimation over specific sub-samples chosen to represent particular regime periods. Second, if some of the sub-samples over which I estimate the VARs do encompass periods of regime shifts, I augment each equation of the VAR with an appropriate number of regime-shift dummies, as detailed in the Data Appendix. For all models where I include them, these dummies are significant in the equations for the interest rate.[52]

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[49] Most criteria, including the sequential modified LR test statistic (LR), the final prediction error (FPE), the Schwarz information criterion (SCE), and the Hannan-Quinn information criterion (HQ), agree in suggesting a 1 or 2 lag specification for all samples for Italy, Germany, France and the euro area. The Akaike criterion (AIC) advocates longer lag specifications when I look at the model for France estimated over the sample from 1993 onwards. Finally all criteria suggest a longer lag structure (around 10 lags) for the euro area model when I estimate it over data from 1992 onwards.

[50] Unit root tests applied to four-quarter changes in the price levels used here for the various countries and the area as a whole cannot reject the null of non-stationarity in all cases at standard levels of significance. The opposite is true when I use deviations of inflation from its implicit/explicit target for each country.

[51] The Data Appendix provides a detailed description of these series and their time series properties. All my estimation output is available on request.

[52] The exception being a dummy for 1993Q3 marking France's central bank independence day when entered in the 1984-2000 model for France, and a dummy for 1992 Q3, marking the exit from the ERM, when I include it in the estimated models for Italy. In the latter case, however, this is significant only at around the 20% significance level. In addition, I generally find no evidence of parameter instability for individual country models, or for the estimated euro area model. Slope dummies appear highly insignificant in these models for all samples, at the exclusion of the case of the euro area model when I estimate it over the longest (1984-2002) sample. In that case I detect some borderline parametric instability, but it is hard to interpret the importance of its implications in the light of the fact that the data is recovered synthetically by aggregating individual country data.

On each VAR I impose two key restrictions. First, that the long-run response of the nominal interest rate to inflation is greater than one—the Taylor 'principle' (see Taylor (1999), Clarida, Galí and Gertler (2000), Woodford (2001) and Batini and Pearlman (2002)).[53] Given the story of successful disinflations in Europe after the oil shocks of the 1970s, I consider this restriction as relatively realistic. I find that this restriction is not rejected by the data at standard levels of significance for the euro area and individual countries for samples which exclude the 1980s.[54] So, in all cases apart from Germany (more on this below), I start estimation from the 1980s and impose this on the interest rate equation. Second, I impose that coefficients on the nominal interest rate and inflation in the output equation are equal in size and have opposite signs. This implies that it is the short-term real interest rate that matters for aggregate demand, as suggested by economic theory—a restriction again not rejected at standard levels of significance by the data I use for the various countries and the entire region. The 'VARs Results Appendix' at the end of the paper presents estimation results from this exercise.

To assess the degree of type 3 persistence I need to make a draw on each model's policy rule's innovations and quantify the number of periods between the timing of the shock and the maximum response in inflation. For that I need to identify the models' responses to shocks. I do this by means of a Cholesky decomposition, where the equation disturbances are assumed to follow the causal ordering (output innovation $\rightarrow$ inflation innovation $\rightarrow$ interest rate innovation). Under this identification scheme, no variable other than the interest rate responds contemporaneously to the monetary policy shock.

The responses of the estimated models broadly agree with conventional wisdom: following a temporary rise in the interest rate, output declines, but ultimately reverts to base, and annualised inflation also falls. However, the error bands around the estimated models' impulse responses are generally quite wide, indicating that these effects cannot be estimated with great precision. In particular, the magnitude of the responses should not be taken too literally, although the implications for the lag structure of the responses are not as dramatic, given that the paths of the error bands mimic quite closely those of the corresponding impulse responses.

[53] I do this by imposing that the long-run response is equal to 1.1.
[54] The inability to reject the 'Taylor principle' restriction after the 1970s but not before accords with finding for the US pre- and post-1979 in Clarida, Galí and Gertler (2000).

A summary of the results on type 3 persistence from this analysis is shown in Tables 7 and 8 below. Two things emerge by looking at the tables. First, both for individual countries and for the euro area, the delay of response in inflation to a policy shock is long-lasting—typically never shorter than three years. This result is in accordance with the findings of Monticelli and Tristani (1999) and is close to those in Coenen (2002), when using Fuhrer-Moore contracts, over a slightly shorter sample than mine. Second, this delay has changed little over time, as shown by results obtained by restricting the sample to cover the increasingly more recent past. For France, Italy and the euro area, the sign of the impulse response to a positive monetary policy shock is incorrect (positive) for samples including the 1970s; something that I do not find for Germany, where the sign of the response is negative, as expected, and suggests an 11-quarters lag between the shock and the peak response in inflation. Excluding the 1970s, none of the inflation responses derived here exhibits a 'price puzzle' (i.e. the perverse initial positive response of inflation to a contractionary nominal interest rate shock), a common finding for impulse responses of estimated unconstrained vector autoregressions. This is a valuable by-product of the two restrictions on the interest rate and output equations that I have imposed in estimation.

Table 7: Euro Area Type 3 Inflation Persistence

Sample	Lag between time policy shock occurs and maximum response in inflation (in quarters)
Euro Area	
1984 Q3-2002 Q2	$k = 15$
1987 Q1- 2002 Q2	$k = 21$
1992 Q4-2002 Q2	$k = 17$

Note: The policy shock is a unit start shock to the orthogonalised interest rate equation's innovation.

Table 8: Individual Countries Type 3 Inflation persistence

Italy	
1981 Q3-2000 Q4	$k = 12$
1987 Q1- 2000 Q4	$k = 12$
1992 Q3-2000 Q4	$k = 12$
France	
1984 Q3-2000 Q4	$k = 12$
1993 Q3- 2000 Q4	$k = 15$
1995 Q3-2000 Q4	$k = 15$
Germany	
1971 Q3-1989 Q4	$k = 11$
1979 Q1- 1989 Q4	$k = 14$
1991 Q1-2000 Q4	$k = 12$

Note: The policy shock is a unit start shock to the orthogonalised interest rate equation's innovation.

2.4 So does country data aggregation matter?

In the previous subsections I have analysed the properties of inflation persistence both at the euro area level and at the country level for three core EMU countries, namely Italy, Germany and France. We know from the work of Granger (1980) that, by construction, an aggregated series can display more persistence than the individual series from which it is derived.[55] So one concern with the results at the euro area level is that the lags observed empirically between policy actions and their effects on prices for EMU as a whole may be an artefact of country data aggregation as opposed to an inherent feature of the way the euro area economy responds to monetary impulses.[56]

The data that I used for the euro area comes from the Fagan et al. (2001) (updated to 2002 Q2) and has been compiled by transforming both real and nominal variables into logarithms. The aggregation was then performed using fixed weights, taken from GDP at market prices (PPP) for the EU11 for 1995. So the data are consistently aggregated, that is to say, data on GDP have been aggregated using the same method to aggregate data on prices etc. In addition, as noted by Fagan and Henry (1998), use of this 'index method' is advantageous because the measure of price inflation implied by this method—i.e. a weighted average of national inflation rates—corresponds to the standard definition employed by various international organisations and so the properties of the series can be straightforwardly compared with those of standard series.[57] Yet, within one class of aggregation (prices, say) there may be some bias coming from aggregating individual country price series, which either compounds or reduces actual area-wide persistence as measured here—a statistical 'fluke'.

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[55] In particular, Granger (1980) shows that the sum of AR(1) processes with coefficients drawn randomly from a suitable distribution approaches a long-memory process, as the number of terms in the sum increases.
[56] See the discussion in the February (1999) and July (2000) issues of the ECB *Monthly Bulletin*.
[57] The index method entails defining the log-level index for any series *X* (in this case individual countries' price series) as follows:

$$\ln X_Z = \Sigma_Z w_i * \ln X_i$$

where *w* is the weighting vector and *Z* is a subset of countries, say. The weights used to aggregate the individual country series in Fagan et al. (2001, and dataset updates), in line with corresponding values of constant GDP at market prices (PPP) for 1995, are: (1) Belgium = 3.9; (2) Germany = 30.529; (3) Spain = 10.233; (4) France = 21.003; (5) Ireland = 1.128; (6) Italy = 20.333; (7) Luxembourg = 0.233; (8) Netherlands = 5.585; (9) Austria = 3.023; (10) Portugal = 2.363; (11) Finland = 1.669.

Conducting a formal test for aggregation bias is beyond the scope of this paper.[58] Nevertheless, the analysis of the characteristics of inflation both at the aggregate and at the country level enables to make some indirect inferences on the impact of the aggregation bias on the measurement of persistence.

A back-of-the envelope calculation to check for this consists, for example, in looking at the degree of type 2 and type 3 persistence detected using euro area aggregate data and comparing it with the degree of persistence that would result by weight-averaging type 2 and type 3 persistence, respectively, across the three countries on which I have focused. To preserve comparability the weights to use must be the same used for aggregating the original area-wide data, namely the weights adopted by Fagan et al (2001). The sum of these for Italy, Germany and France amounts to 71.9% of the total EU11 region. So I rescale Italian, French and Germany weights so that they sum to 100% of the total EU11 region, assuming for simplicity that the remaining countires exhibit similar inflation persistence to these. Table 9 presents results from this simple experiment.

For brevity, relative to type 2 persistence, I only consider whole-sample persistence, but in principle, this exercise can be repeated to make the case for sub-samples. For type 3 persistence I also consider evidence only from one sample, the one covering the 1990s, because this offers the best overlap for individual countries given differences in the timing of regime shifts.

Table 9: Simple evidence on inflation persistence aggregation bias

	Type 2 Inflation Persistence[a]			Type 3 Inflation Persistence[b]
	Four-quarter M3 growth	*Short real rate*	*Yield spread*	*IRF's lags*
Euro area data: Direct evidence	7	20	1	17
Euro area weighted-average evidence	9	19	2	13

(a) Lags have been rounded up to the next unit.
(b) Lags from the impulse response functions from the 1990s samples.

[58] See Zellner (1997) and Shumway and Davis (2001) for a description and application of such tests to a variety of economic series.

The table reveals that, at least according to this basic comparison, there is no substantial exacerbation of the persistence deriving from aggregation.[59] If anything, for type 2 persistence, the lags suggested using ready-aggregated data suggest, less persistence in inflation than what can be recovered by aggregating *ex post* individual countries' persistence levels. This may be due to the fact that the assumption that I have made with regard to persistence across the residual 28.1% of the EU11 when rescaling the country weights to make them match the EU11 total is too conservative. Assuming a lower average degree of persistence than that of Italy, France and Germany for the remaining EU11 countries may deliver indirect evidence that is spot-on. In the case of type 3 persistence the reverse is true, but, as I already discussed, this measure of persistence is less meaningful. So this finding is not too worrisome even if confirmed by more formal tests.

An alternative, and perhaps more accurate, way to detect aggregation bias entails generating artificial country data via the three small country VARs of subsection 2.3, aggregate the data up at the area-wide level, and re-assess type 1, 2 and 3 persistence for the euro area on these synthethic data. It is possible, in fact, that the observed euro area inflation persistence is simply a by-product of a specific historical realisation of shocks at the country level. In practice, this may have implied specific patterns to price level' series in individual countries which, in turn, when aggregated, could have imparted a particularly persistent behaviour to area wide inflation. Deriving measures of persistence on data generated artificially via stochastic simulations (i.e. via 'bootstrapping') accounts for a large number of possible alternative realisations of the shocks—given the sampling distributions of each data series at the country level—and thus controls for such types of aggregation bias.

Table 10 presents results from this second experiment. To obtain these I generated inflation, output and nominal interest rate series for Italy, Germany and France by stochastically simulating 5,000 times the country VARs in subsection 2.3, with shocks drawn from the counties' respective VARs' variance-covariance matrices. I have then aggregated the series up using the same fixed-weights method employed to construct Table 9. Finally, I re-evaluated type 1, 2 and 3 persistence on these newly generated aggregate

[59] This result accords with findings in Granger (1980). He finds that an aggregate series can display more persistence than its individual component series when both: (i) the sum of component series approaches infinity, and (ii) many of the individual series display high first order autoregressive coefficients. This is not

data. More precisely, to evaluate type persistence 1, I run AR (5) on the synthetic euro area inflation series. To evaluate type 2 persistence I derive a measure of the short-run *ex ante* real rate as done in subsection 2.2 and then compute peak cross-correlations between the synthetic aggregated euro area inflation data and this synthetic short-run rate.[60] Finally, to evaluate type 3 persistence, I re-estimate an identified parsimonious VAR on the synthetic aggregate euro area data, applying the same constraints used when estimating the VAR on actual euro area data in subsection 2.3. All types of persistence are obtained over samples of around 130 observations, to replicate the length of the full 1970-2002 quarterly-frequency sample used in previous sections. However, to generate artificial country data by stochastically simulating the country VARs, I use coefficients from the VARs estimated over the 1990s, because this period provides the best overlap for individual countries given differences in the timing of regime shifts.

Table 10: Artificial data evidence on inflation persistence aggregation bias[a]

	Type 1 Inflation Persistence[b, c]	Type 2 Inflation Persistence[b]	Type 3 Inflation Persistence[d]
	$\sum_{i=1}^{5} \hat{b}_i \overline{\pi}_{t-i}^{EA}$	*Short-run ex ante real rate Max. neg. value of* $\rho_{\pi r}(k)$	π^{EA} *IRF's lags in response to monetary policy shock*
Euro area, actual data evidence	0.783 (0.083)	-0.863[e, f] ($k = 20$)	17
Euro area, artificial data evidence	0.748 (0.055)	-0.390[e, f] ($k = 1$)	16

(a) Standard errors in parentheses
(b) Calculated over sample 1970-2002
(c) Regressions on four-quarter changes in actual and artificially-derived HICP
(d) Lags derived over 1990-2002 samples
(e) Significantly different from zero using conventional *t*-test.
(f) Significantly different from zero using Newey-West *t*-test.

The more formal analysis presented in Table 10 reaffirms the finding that there appears to be no substantial exacerbation of the persistence deriving from aggregation. The serial correlation properties of the synthetic euro area inflation data are remarkably close to those from actual data. And the lag in the response of synthetic inflation to a shock to the policy innovation is almost as long as that implied by the VAR estimated on actual data. The only

the case with euro area data since the component series are just eleven, and the autoregressive coefficients are in general well below 1 (see subsection 2.1), warranting my finding of no substantial aggregation bias.

[60] I only look at short-run *ex ante* real rates as a measure of monetary stance because the VARs generating the data do not contain money or long-run rates among the set of endogenous variables, and so I cannot generate artifical series of these.

contradicting evidence comes from the analysis of type 2 persistence on synthetic versus actual data, with the lead between a change in the short-term *ex ante* real rate and its maximum (negative) effect on inflation being much shorter at $k = 1$ than the one implied by actual data ($k = 20$). However, this result may be easily explained by the fact that the dynamic specification of the VARs used to generate articficial country data where VAR(2), and so imparted a very short lead/lag relationship between inflation and interest rates. This may have 'distorted' the correlation results on synthetic data towards suggesting an extremely fast transmission between short real rates and changes in euro area prices—one not in line with correlations derived using raw real-world data.

In short, from these exercises comparing direct and indirect evidence on inflation persistence, it appears as if the aggregation bias does not pose major problem for inference from use of area-wide inflation data. Overall, the difference observed between individual countries' persistence and area-wide persistence seems to result from a pure statistical averaging effect. In addition, observed inflation persistence at the area-wide level appears to be independent to the particular realisation of shocks that occurred historically in Europe. These results concur, for example, with findings in other work investigating aggregation bias on other euro area data, like Fagan and Henry (1998).[61]

3. Conclusions

Recent empirical studies have emphasised that purely forward-looking NKPC specifications for inflation offer a good portrait of euro area inflation dynamics, suggesting that inflation persistence in the EMU is not a structural, policy-invariant feature of the data. On the other hand, European policymakers have adopted a definition of price stability which "is to be maintained over the medium term", in recognition of the existence of an intrinsic delay between monetary policy actions and their effect on inflation.

In line with previous work by Batini and Nelson (2001) on US and UK data, in this paper I looked at whether this conflict can be resolved by presenting model-free evidence on the delay between changes in monetary policy and their peak effect on inflation for the euro area as well as in a subset of its component countries. I also presented evidence of the

[61] Fagan and Henry (1998) find that aggregation bias is not a main driver of the diverse performance of area-wide vis-à-vis EMU member countries' equations. See also Weschke (1996).

autocorrelation properties of inflation and on the lag in the response of inflation to monetary policy shocks from simulations of a battery of identified parsimonious VARs. I found that euro area inflation is persistent according to all definitions I use. For instance, results based on money, interest rate and inflation area-wide data show that it takes over a year before monetary policy actions have their maximum effect on inflation—a finding which strongly validates the ECB medium-term policy orientation. I find relatively consistent results across different measures of monetary stance, including those based on monetary aggregates. This suggests that monetary aggregates in the euro area have a useful role as one set of measures of monetary conditions. Results based on alternative definitions of inflation persistence, like those on the degree of autocorrelation of inflation or on the lag between a monetary policy shock and the peak response in inflation from simulation of a small euro area estimated system in inflation deviations from target, the output gap and the short-term nominal interest rate, also indicate that European inflation is rather inertial.

Importantly, the persistence of European inflation seems to have varied only marginally over the past thirty years, despite the numerous monetary policy regime shifts occurred in Europe after the collapse of the Bretton Woods' exchange rate system.

On balance, these findings seem to clash with those in GGLS (2000) which imply that European inflation is not intrinsically persistent. One possible explanation is the claim by Rudd and Whelan (2001) that instrumental variable estimates (like the Generalised Method of Moments used by GGLS (2000)) could be biased towards forward-looking inflation formation even if the true model contains no such behaviour. McAdam and Willman (2002), for instance, show that alternative estimates of the GGLS (2000) 'hybrid' NKPC—i.e. one in which the instrumented variable is the weighted sum of leads of the mark-up over nominal marginal costs, as implied by Rudd and Whelan (2001), rather than the first lag of inflation—support the view that euro area inflation is strongly inertial. Such a finding does not deny the existence of forward-looking behaviour, but casts doubts on the ability of pure forward-looking equations to account for observed inflation dynamics.

A decomposition of the analysis at the country level reveals that some underlying inflation persistence of the type 2 form (lasting around 1 year) is an inherent feature of way the single economies work—although I cannot exclude some changes over time in the countries' lead/lag relationship between inflation and prior money growth like a longer transmission lag in France and Germany than in Italy when I include the 1970s and 1980s; and vice versa when I look at post-1990 samples. Type 3 persistence is also resilient in country data. However, important cross-country differences in the degree of type 1 inflation persistence emerge by looking at individual countries. Notably: (a) inflation in Italy and France is highly serially correlated, a feature that has not changed greatly over time. By contrast, German inflation is much less autocorrelated today than in the past; and (b) in Italy and France there has been a considerable drop in the mean and variance of inflation over the years, likely reflecting a shift in the implicit price stability objective in the countries following the French abandonment of the 'encadrement du credit' and the move to the 'Franc fort' policy, on one side, and the 'divorce' of the Banca d' Italia from the Italian Treasury, on the other side. I do not observe this for Germany, where inflation has been low historically.

In this sense and according to most definitions, the degree of inflation persistence in the euro area seems to be a halfway house between different degrees of inflation persistence at the country level. Put differently, area-wide persistence seems to result from a pure statistical averaging effect, rather than being a spurious phenomenon due to aggregation, as can be shown via alternative calculations based on the direct and indirect evidence presented in this paper.

The fact that I find persistence to be a pervasive feature of euro area inflation data has implications for the testing and development of macroeconomic models. The ability of models to match the various types of inflation persistence, and notably their ability to account for the lags in effect of systematic monetary policy actions appears to be a priority when validating structural macromodels for use in area-wide policy analysis.

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Data Appendix

A.1 Data description

Euro Area Data

- π^{EA}_t = annualised quarterly inflation in the s.a. Harmonised Index of Consumer Prices, where $\pi^{EA}_t = 4*\Delta p^{EA}_t = 4*(p^{EA}_t - p^{EA}_{t-1})$ and t indicates calendar quarters. (HICP source: Fagan et al. (2001)).

- Euro Area broad money aggregate = euro area harmonised M3 aggregate (Source: ECB).

- Euro Area short-run nominal interest rate = euro area harmonised short-term nominal interest rate. (Source: Fagan et al. (2001)).

- Euro Area long-run nominal interest rate = euro area harmonised long-term nominal interest rate. (Source: Fagan et al. (2001)).

- Euro Area real output = euro area harmonised GDP at factor costs. (Source: Fagan et al. (2001)).

- Euro Area π^{dev}_t = deviations of four-quarter changes in the HICP from an implicit/explicit inflation target. The target is set equal to the corresponding sample means during periods 1970 Q1-1984 Q2 and 1984 Q3-1998 Q4; and equal to 1% (representing the midpoint of the ECB inflation target) thereafter. Note that the dummies included in the euro area VAR comprise different periods from the ones chosen as shift periods in the construction of the implicit/explicit euro area inflation target. So using deviations of HICP inflation from target as an endogenous regressor in the VAR is not identical to using the undetrended four-quarter change in the HICP.

- Euro Area $ybar_t$ = deviations of the log of GDP at factor costs from its Hodrick-Prescott trend (smoothing parameter = 1,600).

Country Data

- π^{ITA}_t, π^{FRA}_t, π^{GER}_t = annualised quarterly inflation in the s.a. Italian, French and German Consumer Price Indices, respectively, where $\pi^i_t = 4*\Delta p^i_t = 4*(p^i_t - p^i_{t-1})$, t indicates calendar quarters, and i = '*ITA*', '*FRA*' and '*GER*', respectively. (CPIs source: BIS and ECB Monetary Transmission Network Macroeconomic Database).

- Broad money aggregates = Italian, French and German M3 aggregates (Source: BIS and NCB and ECB).

- Short-run nominal interest rates = Italian, French and German three-month money market rates. (Source: BIS and ECB Monetary Transmission Network Macroeconomic Database).

- Long-run nominal interest rates = Italian, French and German yields on long-term government bonds. (Source: BIS and ECB Monetary Transmission Network Macroeconomic Database).

- Real output = s.a. Italian, French and German GDP at factor costs. (Source: OECD and ECB Monetary Transmission Network Macroeconomic Database).

- Individual countries' π^{dev}_t = deviations of four-quarter changes inflation in the s.a. Italian, French and German CPI, respectively, from an implicit/explicit inflation target. For Italy, the target is set equal to the corresponding sample means during periods 1970 Q1-1981 Q2 and 1981 Q3-1998 Q4; and equal to 1% (representing the midpoint of the ECB inflation target) thereafter. For France, the target is set equal to the corresponding sample means during periods 1970 Q1-1984 Q2 and 1984 Q3-1998 Q4; and equal to 1% (representing the midpoint of the ECB inflation target) thereafter. For Germany, the target is set equal to the corresponding sample means during periods 1970 Q1-1986 Q4 and 1987 Q1-1998 Q4; and equal to 1% (representing the midpoint of the ECB inflation target) thereafter. Note that as the dummies included in the euro area VAR differ from the periods chosen as shifts in the implict/explicit target, it is not identical to use as an endogenous regressor the undetrended four-quarter change in the various CPIs.

- Individual countries' $ybar_t$ = deviations of the log of Italian, French and German GDP at factor costs, respectively, from their corresponding Hodrick-Prescott trend (smoothing parameter = 1,600).

A.2 Time series properties of the data used in subsection 2.3

For all countries' as well as the euro area model, I detrend output by taking deviations of log real GDP from its Hodrick-Prescott trend (smoothing parameter = 1,600) and together with π^{dev}_t I treat it as I(0) series. I also treat R_t as an I(0) series after controlling for key shifts in monetary policy regime. In the case of Germany, R_t is already I(0) even before controlling for shifts. As evidence, in Table 1A and 2A I present Augmented Dickey-Fuller (ADF) statistics that test the null of a unit root for the variables in my models. Two ADF statistics are calculated in each case for the nominal interest rate R_t: the first includes only a constant in the ADF regression, whereas the second includes a constant and various dummies, depending on the country under consideration. Excluding these dummy variables may bias the test toward suggesting a unit root in R_t.

Table 1A: ADF Tests for Estimated Euro Area Model	
$ybar_t$	– 6.036327**
π^{dev}_t	– 4.264342**
R_t (no shifts)	– 1.755989
R_t (shifts included: *D84Q2*,*D92Q3*)	– 3.360500*
Note: A lag length of four is used in the ADF regressions for each variable. A * denotes significance at 0.05 level according to the Dickey-Fuller distribution's critical values; a ** significance at the 0.01 level according to these values. To test stationarity when shifts are included in the R_t equation I used critical values for the appropriate limiting distribution in MacKinnon (1991).	

Table 2A: ADF Tests for Estimated Individual Country Models	
Italy	
$ybar_t$	– 3.873738**
π^{dev}_t	– 5.109643**
R_t (no shifts)	– 1.716699
R_t (shifts included: *D92Q3*)	– 2.676276*
France	
$ybar_t$	– 3.986229**
π^{dev}_t	– 3.937333**
R_t (no shifts)	– 2.000474
R_t (shifts included: *D84Q2*,*D93Q3*)	– 3.192164*
Germany	
$ybar_t$	– 4.702912**
π^{dev}_t	– 2.171782*
R_t (no shifts)	– 3.416699**
Note: *Ibidem* from Table 1A	

The tests generally reject the null of a unit root in favour of the alternative of stationarity (or in the case of R_t, an I(0) series with structural breaks). Thus I believe it is satisfactory to treat the elements of my estimated models as all I(0), and therefore not apply cointegration analysis.

A.3 Regime shift dummies included in the VARs estimated in subsection 2.3

Note: regime shift dummies take the value of 1 on the date indicate by the notation and are set equal to 1 thereafter. So, for instance, '*D87Q1*' is a variable equal to 0 from the beginning of the sample till 1986 Q4, and equal to 1 thereafter.

Euro Area VAR: 1984 Q3-2002 Q2: *D87Q1*, *D92Q3*; 1987 Q1-2002 Q2: *D92Q3*; 1992 Q3-2002 Q2: none

Italy VAR: 1981 Q3-2000 Q4: *D92Q3*; 1987 Q1-2000 Q4: *D92Q3*; 1992 Q3-2000 Q4: none

France VAR: 1984 Q3-2000 Q4: *D93Q3*; 1993 Q3-2000 Q4: none; 1992 Q3-2000 Q4: none

Germany VAR: 1970 Q1-1989 Q4: *D79Q1*; 1979 Q1-1989 Q4: none; 1991 Q1-2000 Q4: none

VARs' Results Appendix

Tables 3A-6A below show Constrained Full Information Maximum Likelihood estimates for the models of the euro area, Germany France and Italy discussed in Subsection 2.3 of the paper. In all cases the tables show only the models' estimates when the longer samples are used. Estimates over the shorter samples are available from the author on request.

Table 3A: Euro Area Data VAR Estimates
Sample period: 1984 Q3-2002 Q2[a]

	$ybar_t$	$\pi^{dev}{}_t$	R_t
$ybar_{t-1}$	0.8285 (0.0756)	—	0.0563 (0.0560)
$ybar_{t-2}$	—	0.1873 (0.0931)	—
$\pi^{dev}{}_{t-1}$	—	0.7189 (0.0710)	0.0867 (—)
$\pi^{dev}{}_{t-2}$	0.0130 (—)	—	—
R_{t-1}	—	—	0.9211 (0.0255)
R_{t-2}	−0.0130 (0.0341)	—	—
Constant	0.0013 (0.0037)	0.0037 (0.0022)	0.0034 (0.0027)
$D87Q1_t$	0.0002 (0.0021)	−0.0007 (0.0026)	0.0049 (0.0015)
$D92Q3_t$	−0.0018 (0.0018)	−0.0031 (0.0018)	−0.0058 (0.0014)
SD dep. Var.	0.0092	0.0120	0.0270
Equation's SD (σ)	0.0056	0.0067	0.0041
Log-Lik	1136.7354		

Note: (a) Standard errors in parentheses.

Table 4A: German Data VAR Estimates
Sample period: 1970 Q1-1989 Q4[a]

	$ybar_t$	$\pi^{dev}{}_t$	R_t
$ybar_{t-1}$	0.7937	0.1064	0.4698
	(0.0799)	(0.0469)	(0.0833)
$ybar_{t-2}$	—	—	—
$\pi^{dev}{}_{t-1}$	—	0.9519	0.3667
		(0.0401)	(—)
$\pi^{dev}{}_{t-2}$	0.1003	—	—
	(0.0606)		
R_{t-1}	—	—	0.6665
			(0.0633)
R_{t-2}	−0.1003	—	—
	(—)		
Constant	0.0061	−0.0004	0.0193
	(0.0018)	(0.0001)	(0.0045)
$D79Q1_t$	0.0009	0.0007	0.0076
	(0.0025)	(0.0001)	(0.0021)
SD dep. Var.	0.0146	0.0181	0.0277
Equation's SD (σ)	0.0096	0.0058	0.0091
Log-Lik	1082.2131		

Note: (a) Standard errors in parentheses.

Table 5A: French Data VAR Estimates
Sample period: 1984 Q3-2000 Q3[a]

	$ybar_t$	$\pi^{dev}{}_t$	R_t
$ybar_{t-1}$	0.9150	—	0.2744
	(0.0547)		(0.0972)
$ybar_{t-2}$	—	0.0614	—
		(0.1232)	
$\pi^{dev}{}_{t-1}$	—	0.6536	0.1605
		(0.0946)	(—)
$\pi^{dev}{}_{t-2}$	0.0631	—	—
	(0.0254)		
R_{t-1}	—	—	0.8540
			(0.0500)
R_{t-2}	−0.0631	—	—
	(—)		
Constant	0.0053	0.0033	0.0108
	(0.0006)	(0.0016)	(0.0048)
$D93Q3_t$	−0.0013	−0.0054	−0.0030
	(0.0010)	(0.0027)	(0.0029)
SD dep. Var.	0.0091	0.0136	0.0273
Equation's SD (σ)	0.0038	0.0085	0.0068
Log-Lik	1002.2455		

Note: (a) Standard errors in parentheses.

Table 6A: Italian Data VAR Estimates
Sample period: 1981 Q3-2000 Q3[a]

	$Ybar_t$	$\pi^{dev}{}_t$	R_t
$ybar_{t-1}$	0.6943 (0.0712)	0.2032 (0.1448)	0.2070 (0.1246)
$ybar_{t-2}$	—	—	—
$\pi^{dev}{}_{t-1}$	—	0.7388 (0.0706)	0.0688 (—)
$\pi^{dev}{}_{t-2}$	0.0617 (0.0192)	—	—
R_{t-1}	—	—	0.9373 (0.0323)
R_{t-2}	−0.0617 (—)	—	—
Constant	0.0080 (0.0007)	0.0014 (0.0018)	0.0067 (0.0047)
$D92Q3_t$	−0.0036 (0.0012)	−0.0018 (0.0027)	−0.0030 (0.0030)
SD dep. Var.	0.0090	0.0190	0.0448
Equation's SD (σ)	0.0051	0.0111	0.0092
Log-Lik	1120.4078		

Note: (a) Standard errors in parentheses.

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