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Working Paper no. 285

# The New Keynesian Phillips Curve in the United States and the euro area: aggregation bias, stability and robustness

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December 2005

Bank of England

# **The New Keynesian Phillips Curve in the United States and the euro area: aggregation bias, stability and robustness**

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The views expressed in this paper are those of the authors, and not necessarily those of the Bank of England. The authors would like to thank Ryan Banerjee, Nicoletta Batini, Martin Brooke, Karen Dury, Emilio Fernandez-Corugedo, Hashmat Khan, Simon Price and two anonymous referees for helpful comments. We thank David López-Salido for providing some of the programmes for conducting the econometric estimations. This paper was finalised on 6 December 2005.

The Bank of England's working paper series is externally refereed.

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## **Abstract**

In the recent past, the empirical literature on the New Keynesian Phillips Curve (NKPC) has grown rapidly. The NKPC has been shown to describe satisfactorily the relationship between inflation and marginal cost both for the United States and the euro area. However, little attention has been given so far to the stability and robustness of the parameters in the estimated NKPC. In this paper, we aim to help fill this gap. After estimating hybrid NKPCs on US and euro-area data using the generalised method of moments and having found that our results are broadly in line with previous findings, we subject our estimated NKPCs to a thorough stability analysis. We find that the estimated coefficients for the United States are stable, whereas those for the euro area are considerably less stable. We then investigate the possible reasons for this instability. One explanation, explored using the Andrews' test, is the presence of structural breaks. Another possibility is the presence of an aggregation bias, which we investigate by estimating NKPCs for the three largest euro-area economies: Germany, France and Italy. At this disaggregated level, the fit of the NKPC improves, but the coefficients are still unstable. Furthermore, the disaggregated analysis indicates the presence of structural breaks in the three largest euro-area economies.

JEL classification: C12, C13, E31.

## Summary

The traditional Phillips curve relates current inflation to lagged inflation and a cyclical indicator, such as the output gap or the unemployment rate. This specification has efficiently characterised the pattern of inflation over most of the post-war period in most industrialised economies. Two concerns have however been raised. First, the traditional Phillips curve is subject to the Lucas critique – its coefficients may not be invariant to changes in policy regimes. Second, the traditional Phillips curve explains recent data for the United States and the euro area less well, where inflation has been low despite positive output gaps. In an attempt to deal with the shortcomings of the traditional approach, the New Keynesian Phillips Curve (NKPC) literature uses microfoundations to derive a relationship between inflation, expectations of future inflation and the current value of the cyclical indicator. However, the pure NKPC lacks sufficient inertia to adequately explain the path of actual inflation. As a result, many authors attempt to improve the degree of fit of the NKPC by inserting a lagged inflation term. Such a curve is often referred to as a ‘hybrid’ NKPC. We estimate this using generalised method of moments techniques for the United States, the euro area and the three largest euro-area economies.

This relationship between output and inflation is of key importance to monetary policy authorities concerned with price stabilisation. In particular, understanding whether the relationship is stable and how it evolves over time is a primary concern. Therefore, it is somewhat surprising that while numerous papers have estimated the NKPCs for these countries there has been relatively little emphasis on testing the stability of their parameters over time. In other words, the robustness of the NKPC to the Lucas critique has not yet been subject to proper statistical testing. We aim to fill this gap by conducting comprehensive stability and structural break analysis, performing rolling and recursive estimation and applying standard tests for structural breaks.

Overall, our estimates of the structural and reduced-form coefficients on the lagged and expected future inflation terms are broadly in line with previous studies for both the United States and the euro area. One notable exception is the discount factor obtained for the euro area, which is lower than that found in most other studies. On the question of stability, rolling and recursive estimations produce stable and plausible estimates for the United States, but unstable parameters for the euro area. The breakpoint test analysis does not reveal any significant shift in any of the coefficients associated with past and expected future inflation and real marginal cost for the United

States. For the euro area, on the other hand, there is some tentative evidence of a structural break affecting the coefficients on past and expected future inflation in the late 1980s, possibly related to the German re-unification. In the disaggregated euro-area analysis, rolling estimation produces unstable estimates for Germany, and, albeit to a lower degree, for Italy. The estimates for France appear to be considerably more stable over the period considered. Consistent with these results, the breakpoint test analysis points to instability in the late 1970s for Italy and in the early 1980s for Germany. There is no evidence of structural breaks affecting inflation dynamics in France. These conflicting country-level results could indicate the presence of an aggregation bias in the results obtained with euro-area data, which could explain the implausibly low estimate of the discount factor obtained for the euro area.

There are several implications for monetary policy makers. Overall, our results suggest that policymakers should treat the forecasts generated by Phillips curves with some caution, as the structural parameters underlying the estimated relationships may be unstable over time. For the euro area, in particular, it may be useful to look at the results of individual countries, in addition to the aggregate results. Moreover, policymakers should examine the results of a broad range of estimation methodologies to assess whether the forecasts generated by a Phillips curve model agree with other evidence. This is consistent with the approach currently taken in most major central banks.

## 1 Introduction

The traditional Phillips curve relates inflation to lagged inflation and a cyclical indicator, such as the output gap or the unemployment rate.<sup>(1)</sup> This specification has efficiently characterised the pattern of inflation over most of the post-war period in a number of industrialised economies. Two concerns have however been raised. First, the traditional Phillips curve is subject to the Lucas critique, in that its coefficients may not be invariant across policy regimes. Specifically, coefficients on the lagged inflation terms may be considered as actually embedding expectations of future inflation. The Lucas critique is of particular concern when estimating a Phillips curve on euro-area data, due to the shift in policy regime coinciding with the transition to EMU. Second, the traditional Phillips curve explains recent data for the United States and the euro area less well, where inflation has been low despite positive output gaps.

In an attempt to deal with the shortcomings of the traditional approach, the New Keynesian Phillips Curve (NKPC) literature uses microfoundations to derive a relationship between inflation, expectations of future inflation and the current value of the cyclical indicator. Furthermore, many authors attempt to improve the degree of fit of the NKPC by inserting a lagged inflation term. Such a curve is often referred to as a ‘hybrid’ NKPC.

Conventional econometric techniques cannot be applied to estimate a Phillips curve (either traditional or of the new Keynesian generation) due to the problem of endogeneity of the regressors, namely the cyclical indicator and the forward-looking inflation term. A viable alternative is the use of instrumental variables techniques. In their seminal work, Galí and Gertler (1999) and Galí, Gertler and López-Salido (2001) have obtained plausible parameter estimates of the (hybrid) NKPC (based on US and the euro-area data) using the generalised method of moments (GMM). Recently, algorithms to solve linear rational expectation models embedded within a full information maximum likelihood (FIML) approach have been proposed as an alternative to GMM (see for example Lindé (2005) and Ireland (2001)). To date, the debate on the appropriate technique to use when estimating Phillips curves remains open.

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(1) This specification may stand some refinements. A restriction on the sum of the weights on lagged inflation coefficients to one leads to an assumption of no long-run trade-off between output and inflation. Additional lags of detrended output, as well as alternative cyclical indicators, such as unemployment rate or capacity utilisation may also be added to the model.

Despite the wealth of papers concentrating on producing estimates of NKPCs, there has been so far little emphasis on testing the stability of its parameters over time. In other words, the theoretical robustness of the NKPC to the Lucas critique has not yet been subject to proper econometric testing.<sup>(2)</sup> This paper aims at filling this gap by estimating NKPCs on US and euro-area data and then conducting a comprehensive stability analysis and an analysis of the presence of structural breaks. Understanding whether the relationship between output and inflation is stable and how it evolves over time is a primary concern for a monetary authority having the interest rate as the only instrument to achieve stabilisation. Therefore, the type of stability analysis on the NKPC conducted in this paper retains a high informative value for the policymaker.

To ensure that the starting point of our stability analysis is a meaningful one, we first estimate a hybrid NKPC using GMM on US and euro-area data and compare our results to those reported in the literature. We then refine the analysis of the euro-area inflation dynamics by performing similar estimations on data for the largest three euro-area countries: Germany, France and Italy. Next, we perform a comprehensive stability and structural break analysis, by applying the Andrews' test (Andrews (1993)) and performing rolling and recursive estimation.

The plan of the paper is as follows. In Section 2, we briefly recall the theoretical derivation of the New Keynesian Phillips Curve and discuss the difference with the traditional Phillips curve specification. We also illustrate the extension to an open-economy setting, which is used in the analysis on German, French and Italian data. Section 3 describes the empirical methodology used. In Sections 4 and 5, we discuss the data used and present the results of the GMM estimation and of the stability analysis. Section 6 concludes.

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(2) There are few exceptions to this. Jondeau and Bihan (2001), which tests the stability of the NKPC for the United States and euro area. We extend their stability analysis in several ways. First, we conduct rolling and recursive estimates of the NKPC. Second, we test the stability of each parameter, rather than just of the overall model. Finally, we examine the stability of the underlying structural model rather than just the reduced form. Other important differences are that, differently from Jondeau and Bihan, we do not impose super-neutrality, that is that the coefficients on the backward-looking and forward-looking inflation terms sum to one (this allows to avoid any possible instability problem arising when conducting the estimations); and that, for the disaggregated euro-area analysis, we control for the openness of economies in consideration. Also, Estrella and Fuhrer (2003) find the NKPC to be less stable than the traditional backward-looking Phillips curve in the United States.



## 2 Departing from the traditional Phillips curve

### 2.1 The New Keynesian Phillips Curve

The NKPC literature is part of the ever-growing New Keynesian macroeconomics literature. Among its many features, this literature emphasises the role of nominal rigidities as a way of generating real effects from nominal disturbances. A typical feature of the New Keynesian models is a monopolistically competitive corporate sector, that assigns price and/or wage-setting power to firms. Nominal rigidities are then introduced through either staggered price or wage setting. A commonly used assumption is that at any time there is a proportion of firms which do not reset their prices (wages).<sup>(3)(4)</sup> Consequently, after a shock, prices (respectively wages) are slow to adjust, allowing for substantial effects of nominal shocks upon real activity.

In this context, the NKPC is derived as the solution to the conventional profit maximisation problem of a representative monopolistically competitive firm subject to the assumed price (wage) resetting mechanism.<sup>(5)</sup> In its simplest form, the NKPC can be expressed as<sup>(6)</sup>

$$\pi_t = \beta E_t (\pi_{t+1}) + \lambda mc_t \quad (1)$$

where  $\pi_t$  denotes the inflation rate,  $mc_t$  the marginal cost and  $E_t (\cdot)$  is the expectation operator.<sup>(7)</sup> In other words, an increase in the current rate of inflation can be explained either in terms of an upward shift in inflation expectations or in terms of a rise in the marginal cost of production. The parameters of (1) bear a precise microeconomic interpretation. Specifically,  $\beta$  is the rate at which future profits are discounted. The parameter  $\lambda$  is defined as  $\lambda \equiv (1 - \theta)(1 - \beta\theta)\xi\theta^{-1}$ , where  $1 - \theta$  denotes the probability that a firm will reset its price in any period and  $\xi$  is a parameter depending on returns to scale. With respect to the traditional, reduced-form Phillips curve, the NKPC so

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(3) An alternative, equivalent assumption is that the time intervals over which the price of different goods (or alternatively, wages of the different agents) remain fixed overlap, rather than being perfectly synchronised.

(4) Although both price and wage stickiness have been considered in the literature, price rigidity has so far been the most widely adopted assumption. The derivation which follows is based on the Calvo specification of price stickiness (see Calvo (1983)). For alternative nominal rigidities specifications, see Rotemberg (1996), Taylor (1980), and Christiano, Eichenbaum and Evans (1997). For earlier insights, see Phelps (1978) and Taylor (1979).

(5) For a full derivation of the NKPC, see Galí *et al* (2001).

(6) In what follows, all variables are expressed as deviation from steady state.

(7) Under the assumption of a Cobb-Douglas or CES production technology, it can be proved that the marginal cost is proportional to the output gap. This leads to an alternative expression of the NKPC where the output gap is used as an indicator of business cycle movements. However, obtaining measures of the output gap which may satisfactorily capture the real activity pressures on inflation is a controversial task. This is one of the reasons for preferring the use of the marginal cost in econometric studies of the NKPC. In producing the results reported in this paper, we limit ourselves to the use of the marginal cost measure.

obtained should be less vulnerable to the Lucas critique, since its coefficients are derived from deep structural parameters and are thus unlikely to vary with shifts in monetary regimes. This makes the NKPC a more robust representation of inflation dynamics for policy analysis.

The ‘baseline’ NKPC **(1)** has proved to have two major limitations. First, by representing inflation as a purely forward-looking process, it implies that inflation responses anticipate those of the cyclical indicator and, in particular, that inflation responds quickly to monetary policy shocks. This is at odds with the empirical evidence.<sup>(8)</sup> Second, it cannot account for the type of inflation persistence witnessed both in the United States and in the euro area in recent years. This is particularly puzzling since empirical evidence seems to show that inflation dynamics tend to be characterised by a considerable degree of persistence.<sup>(9)</sup>

These limitations can be resolved by introducing in **(1)** one or more lagged inflation terms, typical of the traditional Phillips curve. Among the various ways proposed in the literature for deriving such a hybrid NKPC, one of the most commonly used is assuming that a fraction of the resetting firms follow a backward-looking rule of thumb. Specifically, it is assumed that, among the  $1 - \theta$  firms which at any time change their price, only a fraction  $(1 - \omega)$  of them reset their price optimally (that is, based on the expected future inflation rate), with the remaining fraction  $\omega$  choosing to set their price according to the past levels of inflation.

Under this assumption, the solution to the producer optimisation problem leads to a hybrid NKPC of the form:<sup>(10)</sup>

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t(\pi_{t+1}) + \lambda m c_t \quad (2)$$

With respect to the decomposition implied by **(1)**, **(2)** suggests a reduced inflation response ‘on impact’, as inflation developments are partly driven by past inflation movements.

As before, the reduced-form parameters  $\gamma_b, \gamma_f$  and  $\lambda$  can be expressed as functions of the structural parameters:

$$\lambda \equiv (1 - \theta)(1 - \beta\theta)(1 - \omega)\xi\phi^{-1} \quad (3)$$

$$\gamma_f \equiv \beta\theta\phi^{-1} \quad (4)$$

---

(8) This type of responses has been emphasised in the structural VAR literature, see for example Christiano *et al* (1997).

(9) See Fuhrer (1997).

(10) See Galí *et al* (2001) for a detailed derivation of the hybrid NKPC.

$$\gamma_b \equiv \omega\phi^{-1} \quad (5)$$

Specifically,  $\phi$  is a composite coefficient, such that:<sup>(11)</sup>

$$\phi = \theta + \omega[1 - \theta(1 - \beta)] \quad (6)$$

## 2.2 The open-economy hybrid NKPC

In the empirical analysis that follows, we base our estimate of the US and euro-area hybrid NKPCs on equation (2), where the real marginal cost is defined as the labour share. This is consistent with the consensual perception of the United States and euro area as closed economies. However, when estimating the hybrid NKPC for the largest three European economies, the model needs to be extended to account for the openness dimension associated with each of the individual countries considered. In the literature to date, two channels through which such a dimension may affect the evolution of inflation have been discussed. For example, Batini, Jackson and Nickell (2000) derive an open-economy NKPC in which both the marginal cost and the mark-up are affected by international prices. As shown by Bentolila and Saint-Paul (1999), in a strictly Cobb-Douglas framework, marginal cost is unaffected by the price of material inputs. However, if we assume more general technologies, an increase in the price of imported inputs raises the marginal cost of production. We model this channel using a linear specification:

$$mc_t = s_t^L + \delta p_t^m \quad (7)$$

where  $p_t^m$  denotes the relative price of imported goods (relative to the domestic price level) and  $s_t^L$  is the income share of labour. Moreover, the mark-up over marginal cost may also be a function of the degree of openness of the economy. This is because the demand elasticity and the extent of competitive pressures faced by the firm will be affected by the presence of foreign competition.

In the empirical analysis we conduct in the paper, we treat the United States and euro area as broadly closed economies, in line with assumption in most of the literature, but treat the individual German, French and Italian cases as open economies.<sup>(12)</sup> Specifically, we estimate

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t(\pi_{t+1}) + \lambda mc_t + \phi \Omega_t \quad (8)$$

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(11) Bakhshi, Burriel-Llombart, Khan and Rudolf (2003) and Ascari (2003) emphasise that equation (2) is derived under the assumption of a zero inflation rate in the steady state. This assumption is convenient, but not entirely realistic. Relaxing this assumption, however, adds to the complexity of equation (2). Although acknowledging this drawback in the derivation, we shall disregard it for the time being.

(12) Our results and conclusions are generally not sensitive to the inclusion or exclusion of the open-economy term.

where  $\Omega = p^m + \log\left(\frac{X+M}{GDP}\right)$ , with  $X$  and  $M$  denoting real export and import volumes. Hence  $\Omega$  is a measure of nominal trade share (under the assumption that export prices equal import prices).<sup>(13)(14)</sup>

### 2.3 The model

Estimating the coefficients in (2) requires dealing with the endogeneity problem of the explanatory variables, mainly the marginal cost and the expected future inflation term. In this paper, we adopt a partial equilibrium approach to estimating the NKPC. This involves estimating the coefficients in (2) as the parameters of the following data-generating processes for the inflation rate and the marginal cost:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t(\pi_{t+1}) + \lambda m c_t + \epsilon_t \quad (9)$$

$$m c_t = \rho m c_{t-1} + u_t \quad (10)$$

where  $\epsilon_t$  is a cost-push shock and (10) assumes that the marginal cost is generated by an autoregressive process which error term,  $u_t$ , is contemporaneously and serially uncorrelated with  $\epsilon_t$ .

Equation (2) cannot be directly estimated since  $E_t(\pi_{t+1})$  is not directly observable. One possibility is to assume that expectations are formed rationally, which means replacing the expectation of inflation with its future realisation plus a serially uncorrelated error term:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f \pi_{t+1} + \lambda m c_t + e_t \quad (11)$$

where it is

$$e_t = \epsilon_t - \gamma_f (\pi_{t+1} - E_t(\pi_{t+1})) = \epsilon_t - \gamma_f \eta_{t+1}$$

with  $\eta_{t+1}$  denoting the inflation forecast error. Under the rational expectation hypothesis, the forecast error  $\eta_t$  should be uncorrelated with the variables in equation (2).

A necessary condition for the existence of a closed-form solution in a model with forward-looking

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(13) In the estimations which follow, both the marginal cost and the nominal trade shares are measured in deviation from steady state (time-varying mean).

(14) An alternative approach to modelling the open-economy dimension has been proposed in Galí and Monacelli (1999), who argue that in an open economy the real marginal cost is a function of the ratio of foreign good prices to domestic good prices. Thus, shocks to foreign demand affect the terms of trade, and this in turn affects the domestic real marginal cost. But these foreign competition-led time-variations in the mark-up have proved not to be significant. We therefore disregard them in our analysis. Also, Adolfson, Laséen, Lindé and Villani (2004) show that if intermediate imported inputs are not directly used in the production function, then the standard NKPC still applies, but there is one Phillips curve for each sector of the economy.

expectations is that at most one of the roots is explosive, ie that at least one root is not explosive. When there is exactly one explosive root, the solution is unique. In the hybrid NKPC framework, these requirements translate into the assumptions:

$$\gamma_b, \gamma_f \geq 0 \quad \gamma_b + \gamma_f < 1 \quad (12)$$

This ensures that the roots are real, that a stationary solution always exists and that the solution is unique.<sup>(15)</sup>

Estimating (11) we obtain estimates of the reduced-form parameters  $\gamma_b$ ,  $\gamma_f$  and  $\lambda$ . But, using (3),(4), (5) we can rewrite equation (11) in terms of the underlying structural parameters:

$$\phi\pi_t = \omega\pi_{t-1} + \beta\theta\pi_{t+1} + (1 - \omega)(1 - \theta)(1 - \beta\theta)\xi mc_t + \phi e_t \quad (13)$$

Thus, estimating (13) produces estimates of the structural parameters. Notice that estimating (13) alongside (11) is essential for understanding how changes in the structural parameters affect the trade-off between inflation and the real economy as well as the degree of inflation persistence observed in the economy.

### 3 Empirical methodology

The presence of expected future inflation among the regressors of the NKPC invalidates the use of least squares estimation techniques, as this would yield inconsistent estimators. Two alternatives are commonly followed in the literature: estimating the model using the generalised method of moments (GMM), or estimating a stochastic linear rational expectations model by full information maximum likelihood methods (FIML). No consensus in the literature has been expressed so far on the comparative advantages of either of these methodologies.

With respect to other methodologies such as FIML, GMM does not require specific information about the distribution of the error terms, but only a specification of the moment conditions. However, the drawback of GMM is that it does not always make efficient use of the information in the sample. Furthermore, it has been demonstrated that GMM estimators are severely biased in small samples. For the benefit of comparison with other studies on inflation dynamics in the United States and the euro area, we perform only GMM estimations in this paper, leaving for

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(15) It can be shown that when the hybrid NKPC model is derived from the standard Calvo pricing model, restrictions (12) are always satisfied.

further research the application of the FIML methodology.<sup>(16)</sup>

### 3.1 *GMM pitfalls: some considerations*

The GMM methodology is subject to a number of pitfalls. Therefore, when applying this method, it is necessary to subject the results to a large number of robustness checks. In this section, we summarise the main pitfalls of the GMM methodology and discuss the way we have addressed them when performing estimations of the hybrid NKPC.

**Overidentifying restrictions and finite sample bias.** GMM is an instrumental variable (IV) methodology. When the number of orthogonality conditions (here the number of instruments) exceeds the number of parameters to be estimated, the model is overidentified, and the GMM estimates are biased. Monte Carlo simulations have demonstrated that the bias introduced by the use of an excessive number of instruments is sizable in finite samples. Furthermore, it can be shown that: the probability limit of the GMM estimator tends towards that of the least squares estimator when the number of instruments increases in proportion to the number of observations;<sup>(17)</sup> and the bias increases with the number of instruments used.<sup>(18)</sup>

To avoid the bias originated by the use of an excessive number of instruments, we have applied the Newey-West test for overidentifying restrictions in order to choose the proper number of instruments. This test is based on the minimised value of the objective function, the  $J$ -statistic. Under the null hypothesis that the overidentifying restrictions are satisfied, the  $J$ -statistic times the number of regression observations is asymptotically  $\chi^2$  distributed with number of degrees of freedom equal to the number of overidentifying restrictions. The results presented in this paper are based on a selection of instruments which does not introduce any overidentification bias according to the  $J$ -test.

**Misspecification of the orthogonality condition.** When specifying a hybrid NKPC for estimation purposes, two types of misspecification could occur. Misspecification could take the

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(16) When running GMM estimations, a kernel and a bandwidth have to be specified. The kernel is used to weight the covariances so that the weighting matrix ( $A_T$ ) is ensured to be positive semi-definite. The bandwidth determines how the weights given by the kernel change with the lags in the estimation of  $A_T$ . We follow Galí *et al* (2001) and choose a Bartlett kernel and a fixed bandwidth of 8.

(17) See Staiger and Stock (1997) and Woglom (2001).

(18) See Fuhrer, Moore and Schuh (1993), Rudebusch and Fuhrer (2002).

form of a measurement error or an omitted variables error. In the former case, some of the variables in the hybrid NKPC could be measured with an error. This is likely to occur when specifying the forcing variable. In the latter case, some dynamics, either of inflation or the forcing variable, could be mistakenly omitted. Overall, misspecification problems are known to produce asymptotically biased estimators.

Common types of measurement error involve using a forcing variable that is not the ‘true’ forcing variable; or using the ‘true’ forcing variable as an instrument only. For instance, while the output gap or the real marginal cost are often used as forcing variables in the (hybrid) NKPC, it is possible that the ‘true’ forcing variable is indeed a combination of the two. Lindé (2005), using Monte Carlo simulations, finds that even very small measurement errors lead to biased estimates of the NKPC parameters. There is unfortunately no simple way to avoid measurement errors. However, there is some evidence that output gap estimations can vary considerably depending on the methodology used. For this reason, and to use an approach consistent with the theoretical derivation of the (hybrid) NKPC, in the estimates we have conducted, we have used the real marginal cost as the forcing variable.

Omitted dynamics refers to the misspecification introduced in the model by omitting lags of, for instance, the inflation rate, or mistakenly assuming that the forcing variable is exogenous. Rudd and Whelan (2001) show that small specification errors in the assumed data-generating process can lead to highly misleading results. In the estimations we have conducted, in order to avoid both types of misspecification, we have instrumented the forcing variable and tested whether more than one lag of inflation should be included in the NKPC. Overall, we have found that only one lag of inflation is significant for the countries considered.

**The stationarity condition.** Application of the GMM method requires pre-whitening of non-stationary data (or possibly the use of cointegration techniques). In the empirical analysis that follows, we test for the presence of unit roots in the data. We find that, for inflation rates, although the presence of a unit root cannot be rejected, it can however be explained in terms of structural breaks. Furthermore, tests conducted on the real marginal cost and the variables used as instruments suggest mixed results. We therefore follow the convention adopted in the existing literature and assume stationarity for all the variables used in the estimations.

**Weak identification.** In the case of the hybrid NKPC model, the loss function at the base of the GMM minimisation problem, which solution produces the estimates of the structural parameters, is not necessarily quadratic. Consequently, the structural parameters in (13) are only weakly identified.<sup>(19)</sup> In the following, we take into account the identification issue by avoiding a strict mapping of the structural results into the reduced-form ones. Therefore, we conduct independent stability analysis on the structural coefficient estimates obtained from (13) and on the reduced-form coefficients obtained from (11).

**Normalisation dependence.** As Galí *et al* (2001) and Lindé (2003) show, GMM estimates of the structural parameters are dependent on the assumed specification for the hybrid NKPC. Specifically, renormalising (13) as

$$\pi_t = \omega\phi^{-1}\pi_{t-1} + \beta\theta\phi^{-1}\pi_{t+1} + \phi^{-1}(1 - \omega)(1 - \theta)(1 - \beta\theta)\xi mc_t + e_t \quad (14)$$

produces different structural parameter estimates. To control for the effect of the normalisation, we have estimated the structural parameters using both specifications (13) and (14). We have also conducted stability analysis on both sets of estimations. Consistent with previous findings in the literature, we find that the estimation results are affected by the normalisation. Overall, the results obtained using (13) appear to be more robust to the choice of instruments.<sup>(20)</sup> Consequently, we concentrate the discussion in the paper on these results only. Results based on estimations of (14) are available from the authors on request.

### 3.2 Testing for parameter instability in the GMM framework

The microfounded theoretical model underlying the (hybrid) NKPC makes it less vulnerable to the Lucas critique than traditional Phillips curves, since the coefficients are a function of structural parameters that are expected to be unaffected by policy regime shifts. However, in the empirical literature on the (hybrid) NKPC to date, there has been little emphasis on testing the stability of the estimated parameters over time. Such a test would provide support to the hypothesis of robustness of the (hybrid) NKPC to the Lucas critique.

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(19) For a discussion of this shortcoming of the GMM methodology, see Ma (2002). Mavroeidis (2002) demonstrates the empirical relevance of the weak identification problem of structural parameters and proposes a metric for the degree of identification of the parameters in forward-looking rational expectation models.

(20) Specifically, we found that with (14) the GMM estimator failed to converge with certain choices of instruments or estimation periods.



In this paper we aim to fill this gap by performing a comprehensive stability analysis on our NKPC estimated parameters. To the best of our knowledge, no similar analysis has been conducted to date.<sup>(21)</sup>

First, we apply the stability test developed by Andrews (1993).<sup>(22)</sup> Then, we perform rolling and recursive estimations. These estimations are based on successive subsample estimations. For rolling estimations, we use a subsample of a fixed number of observations  $T$ , and we perform GMM estimations for each subsample. The first subsample starts with the first observation in the full sample. We then move forward the starting observation of the subsample gradually until the last observation minus  $T$ . For recursive estimations, we use both an increasing and a decreasing number of observations.<sup>(23)</sup> With an increasing number of observations, we use the first observation and the smallest possible subsample, then gradually increase the sample, while keeping the first observation fixed. With a decreasing number of observations, we keep the last observation fixed, but vary the first observation of the subsample. For each type of estimation, we record the estimated coefficients for the GMM estimation performed on each subsample. Plotting them against time allows us to assess whether the estimated coefficients are changing over the sample.

#### 4 Data

For the empirical analysis which follows, we use the data set from the National Institute of Economic and Social Research (NIESR) in their world economy model, NiGEM.<sup>(24)</sup> In the following we refer to them as ‘NiGEM data’.<sup>(25)</sup> For the United States, the data set covers the period from 1965 Q1 to 2002 Q4, whereas the euro-area data sample extends from 1970 Q1 to 2002 Q4.<sup>(26)</sup> We conduct a disaggregate analysis using data from the three largest euro-area countries: Germany, France and Italy. For Germany and France, the data sample is 1965 Q1 to 2002 Q4, whereas the Italian data cover the period 1970 Q1 to 2002 Q4. In the estimation we report below, real marginal cost is used as the forcing variable. When estimating the hybrid NKPC

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(21) As remarked earlier, Jondeau and Bihan (2001) also conduct stability tests of the NKPC. See the earlier footnote 4 for a summary of the differences between that approach and the one followed in this paper.

(22) Details of the test are provided in Appendix A.

(23) This allows for testing the presence of both initial and terminal points effects.

(24) The database includes seasonally adjusted data and is updated quarterly. Here we use the version released in 2003 Q2.

(25) The original sources of the NiGEM database are the OECD, Eurostat and national statistics.

(26) For the euro area, the NiGEM database consists of data from the member countries aggregated using PPP weights.

for Germany, France and Italy, we introduce an open-economy dimension as discussed in Section 2.2 above.<sup>(27)</sup>

#### 4.1 Inflation rates

The inflation rates used in the estimations are year-on-year changes in the log of the GDP deflator.<sup>(28)</sup> This measure is more consistent with the theoretical derivation of the hybrid NKPC (from the representative firm optimisation problem) than consumer price inflation. Also, the use of this measure makes our estimates comparable to those of other studies (for example, Galí *et al* (2001), Galí and Gertler (1999), Fuhrer and Moore (1995)), and allows us to draw cross-country comparisons.

Over the period 1970 to 2002, inflation has followed similar patterns in the United States and the euro area (see Chart 1 below and, for quarterly data, Chart C.1 in Appendix C). However, over this period, the inflation rate has remained higher in the euro area than in the United States. And, over the 1970s, the variance of inflation was also higher in the euro area than in the United States. From the beginning of the 1980s, the level and the variance of inflation have fallen in both the United States and the euro area.

Over the period 1970-2002, the United States and the euro area appear to be characterised by highly persistent inflation rates. Performing the Augmented Dickey-Fuller (ADF) tests over the full sample indicate the presence of unit roots in the inflation rates (see Table A in Appendix B). However, visual inspection suggests the presence of a structural break in the early 1980s, which would indicate that the ADF test could suffer from misspecification. Choosing 1983 as the break point and hence splitting the sample in 1983, we have performed the ADF test separately over the two subsamples. In this case, the null hypothesis of non-stationarity in the subsamples can be rejected. Hence, we infer that the shifts in the level of inflation in both the United States and the euro area can account for the non-stationarity of the inflation rates and that the inflation rates, although persistent, do not necessarily have a unit root.

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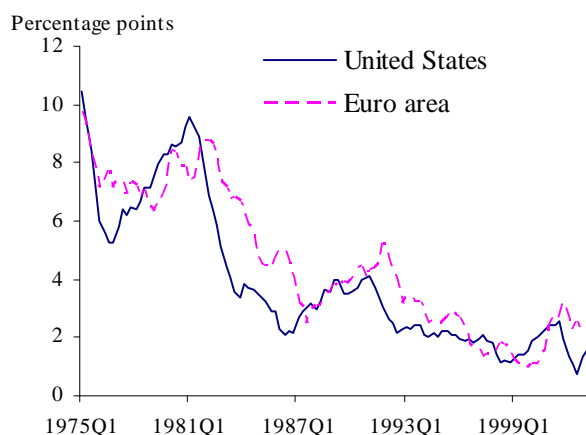
(27) In order to compare our euro-area results to other studies (specifically Galí *et al* (2001)), we also perform estimations using the area-wide model (AWM) data set, documented in Fagan, Henry and Mestre (2001), the latest update of which covers the period 1970 Q1 to 2002 Q4.

(28) For the United States, the euro area, Germany and France we have used annual inflation rates. For Italy, we have used quarterly changes in the log of the GDP deflator and added seasonal dummies to the instruments set.

As discussed earlier, applying the GMM methodology to non-stationary series might impair the estimation. On the other hand, estimating the hybrid NKPC on the subsamples over which the inflation rates are stationary might lead to important small-sample problems. Hence, we have chosen to estimate the hybrid NKPC on the whole available sample, with the inclusion of deterministic mean-shift dummies. Specifically, when estimating the NKPC on US data, we have included a dummy variable assuming value zero between 1970 Q1 and 1983 Q1 and value one thereafter in the instrument set. This would account for the shift in the conduct of the US monetary policy started with Volker in 1983 and continued by Greenspan from 1992 onwards. When estimating the NKPC on euro-area data, we have instead included in the instrument set a dummy for the German unification, assuming value zero until 1990 Q4 and value one thereafter.<sup>(29)</sup>

Notice that according to the theoretical definition of the NKPC, inflation should be measured in terms of deviation from its steady state. Bakhshi *et al* (2003) show that if one deviates from the conventional assumption in the literature of zero steady-state inflation, the form of the NKPC is considerably more complicated than that derived in the literature. In order to derive results comparable to those in the literature and, at the same time, take into account the arguments in Bakhshi *et al* (2003), we have chosen to use a demeaned measure of inflation, thus adopting a measure of steady-state inflation equal to the mean value of inflation. The use of deterministic dummies discussed earlier accounts for the shifts in the mean.

**Chart 1: Annual GDP deflator inflation rates**



(29) This is done to accommodate the fact that, in the NiGEM database, data for Germany are referred to West Germany only before the unification and to the whole of Germany after the unification.

## 4.2 Real marginal cost

The real marginal cost is usually measured by the labour share of income. However, following Galí *et al* (2001), we have corrected the labour share by a constant to account for decreasing returns to scale at the firm level.<sup>(30)</sup> Furthermore, in order to be consistent with the theoretical model underlying the (hybrid) NKPC, which uses deviations of the real marginal cost from its steady-state value, we use the demeaned labour share, where the mean is computed over the entire sample (this is analogous to the definition adopted in Galí *et al* (2001)).

Galí *et al* (2001) define compensation per employee as the ratio of compensation to the number of employees. In this paper, we use the NiGEM definition, as the ratio of compensation to the number of employee hours. In the NiGEM database, the aggregate euro-area series for compensation per employee hours exists only from 1983 Q1 to 1999 Q4 due to missing observations in the series for hours worked for some of the euro-area countries. To reduce the GMM estimation bias due to short samples, we have chosen to artificially extend the series for those countries with missing observations assuming they grow at the average growth rate of the countries for which the series are available.

## 4.3 German re-unification effect

Several time series in the NiGEM data set are not corrected for the German re-unification effect in 1991 Q1. In particular, the compensation and employment series exhibit a clear shift in 1991 Q1. This affects the aggregated euro-area measures for employment, compensation and compensation per employee hour quite strongly given the large weight of Germany in the euro area (32% by GDP at PPP weights in NiGEM data). To deal with this problem, we have introduced in the instrument list a deterministic intervention dummy that takes the value 0 until 1990 Q4 and 1 from 1991 Q1 onwards.

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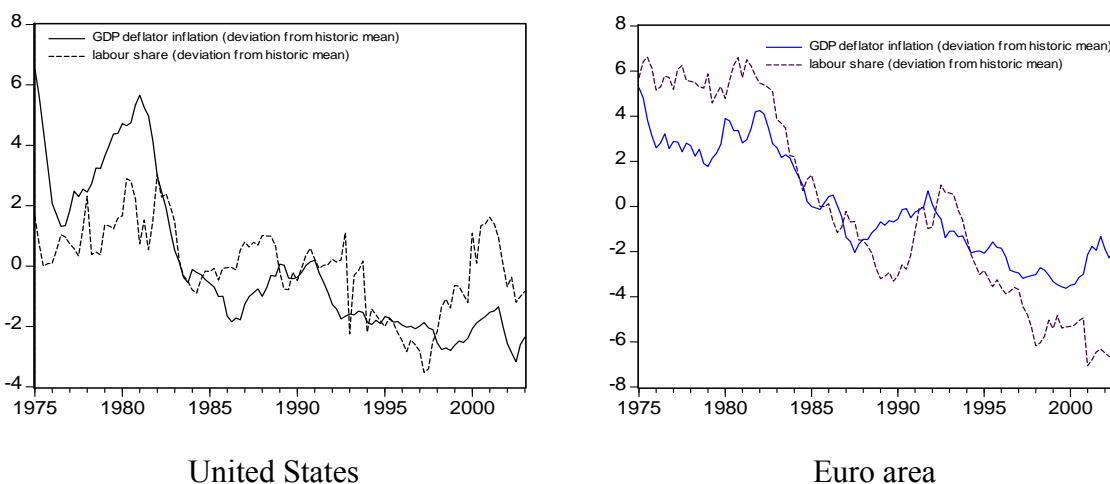
(30) The motivation proposed by Galí *et al* (2001) for this correction can be summarised as follows: cost minimisation implies that the firms' real marginal cost equals the real wage divided by the marginal product of labour. In the case of constant return to scale (Cobb-Douglas technology), the real marginal cost at time  $t + k$  for a firm that optimally sets prices at time  $t$ , is given by  $mc_{t,t+k} = \frac{1}{1-\alpha} \frac{W_{t,t+k}}{P_{t,t+k}} \frac{Y_{t,t+k}}{N_{t,t+k}}$ , where  $\alpha$  is the technology curvature parameter. In the absence of firm-level data, we define the observable average real marginal cost as  $mc_t = \frac{1}{1-\alpha} \frac{W_t}{P_t} \frac{Y_t}{N_t}$ , which depends only on aggregate variables. With decreasing returns to scale, the marginal cost is defined as  $mc_t = \frac{1}{1+\alpha(\varepsilon-1)} \frac{W_t}{P_t} \frac{Y_t}{N_t}$  where  $\varepsilon$  is the elasticity of demand. As it is assumed that  $\frac{1-\alpha}{1+\alpha(\varepsilon-1)} \neq 1$ , this implies that a measure of marginal cost can be obtained by multiplying the labour share by a constant (of value between 0 and 1). Hence, the correction required by the presence of non-constant returns to scale affects only the estimate of the coefficient of the real marginal cost.

## 5 Estimation

### 5.1 A comparison between the United States and the euro area

In the United States, the inflation rate traces the real marginal cost (as measured by labour share) closely over the period 1975-2003, perhaps with the exception of the period 1995-2003, when productivity growth was very high and the US economy experienced a prolonged boom. The same is true for the euro area (see Chart 2 below). Given this significant correlation,<sup>(31)</sup> it is not surprising that the fit of the hybrid NKPC is high, both for the United States and the euro area. Interestingly, it appears that for both the United States and the euro area, the gap between the real marginal cost and the inflation rate have been relatively closed from approximately the beginning of the 1980s until the middle 1990s. Since then, however, the gaps appear to have reopened again.<sup>(32)</sup> The analysis below is intended to explain these developments in the inflation/real marginal cost trade-off.

**Chart 2: Quarterly GDP deflator inflation rates and labour share, United States and euro area, 1975 - 2003**



(31) The correlation coefficients, computed over the available samples, are 0.63 for the United States and 0.93 for the euro area.

(32) On average, there seem to be a two years' lag between the euro area and the US inflation developments.

## 5.2 Estimation results

Tables B and C in Appendix B report the estimation results for equation (13) for the United States and the euro area obtained using data for the period 1975 Q1-2002 Q4.<sup>(33)</sup> The instruments used in the GMM estimations are a constant, four lags of GDP inflation, wage inflation, the labour share and the output gap.<sup>(34)</sup> As discussed earlier, deterministic dummies are also added to this set to account for the mean shift in inflation.

### 5.2.1 Reduced-form parameters

The reduced-form estimates are summarised in Table B in Appendix B. These are obtained by estimating (13) and then using (4), (5) and (3). Overall, our results are in line with those from previous studies,<sup>(35)</sup> notwithstanding the variability of the estimates reported in the literature, given the different periods, methodologies and specifications used in the different studies.

For the United States, the coefficient on the forward-looking inflation term  $\gamma_f$  (0.52) is large and highly significant. The coefficient on the backward-looking component  $\gamma_b$  (0.47) is also similar in size to the estimates reported in the literature, and equally significant. However, the estimate of  $\lambda$  (0.3) is slightly higher than other studies have previously found.

For the euro area, the coefficient on the forward-looking and backward-looking inflation terms are smaller than for the United States. This has also been found in several other studies. The estimate of the coefficient on the real marginal cost is also in line with other studies findings.<sup>(36)</sup>

These estimations are derived without imposing the restriction that  $\gamma_f + \gamma_b = 1$ . However, our results suggest that this restriction cannot be rejected at the 5% level for both the United States

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(33) We find that estimations of hybrid NKPCs using the output gap instead of the marginal cost tend to give less significant coefficients. This result is consistent with the findings of previous studies, including Galí *et al* (2001) and Jondeau and Le Bihan (2001). These results are not reported but available from the authors on request.

(34) The use of the output gap as an instrument is common in the empirical literature on (hybrid) NKPC. This is not inconsistent with the earlier discussion explaining our preference for a (hybrid) NKPC including the marginal cost as the forcing variable. Indeed, in the GMM procedure, the set of instruments is seen as representing the information available at the time agents form their expectation of the inflation rate. Thus, including the output gap in the instrument set is to be thought as making the model less prone to misspecification.

(35) See, for example, Galí *et al* (2001), Lindé (2003), Altig, Christiano, Eichenbaum and Lindé (2004) and Jondeau and Bihan (2003).

(36) In both estimations for the United States and the euro area, the Newey-West test indicates that our estimates are not overidentified.

and the euro area.

### 5.2.2 Structural estimates

As argued in Section 2 above, the main advantage of (hybrid) NKPCs on traditional Phillips curves is the structural interpretation of its reduced-form parameters. This allows to relate specific movements in inflation dynamics to the structure of the economy in consideration. For this reason, we complement the above estimates of the reduced-form parameters for the United States and the euro area with estimates of the structural parameters. These are reported in Table C in Appendix B.<sup>(37)</sup> The estimates of the structural parameters are highly significant, both for the euro area and the United States. For the United States, the value of the discount factor  $\beta$ , 0.98, is not significantly different (at the 5% level) from the conventional value (for quarterly data) of 0.99. For the euro area, the discount factor is implausibly small, suggesting that these results could be affected by an aggregation bias. The disaggregated analysis we conduct later aims at detecting any such possible bias.<sup>(38)</sup> In line with previous findings, the fraction of backward-looking agents ( $\omega$ ) is estimated to be quite small in the United States, although marginally larger than in the euro area. The small value of  $\omega$  for the euro area helps explain the small value of the backward-looking reduced-form parameter  $\gamma_b$ . The estimates of price stickiness ( $\theta$ ) are in line with previous findings, suggesting higher stickiness in the euro area. These estimates imply a contract duration ( $\frac{1}{1-\theta}$ ) of 4.2 quarters for the euro area and 2.9 quarters for the United States. In essence, these structural estimates suggest that although in the euro area the proportion of firms which, at any time, have a non zero probability of changing their price is smaller compared to the United States, the majority of those who can change their price choose to do so optimally. These results support the importance of computing structural estimates of hybrid NKPCs, as they are able to provide precise information about firms' pricing behaviour and their effect on inflation dynamics.

### 5.2.3 A cross-check using the Area Wide Model (AWM) data set

So far, the euro area has been proxied by the aggregation of the three largest euro-area economies.

As a cross-check, we have re-estimated the hybrid NKPC for the euro area using the ECB Area

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(37) Following Galí *et al* (2001), we have chosen not to estimate the parameter  $\zeta$ , which depends on the returns to scale. This choice is required to be able to identify the  $\beta$  and  $\theta$  parameters. We have set  $\zeta$  to the values commonly used in the literature: 0.2 for the euro area (as well as for Germany, France and Italy, in the disaggregated estimates) and 0.1 for the United States. We have also conducted sensitivity tests on these assumed values for  $\zeta$ . Overall, they suggest that our structural parameters estimations are robust to the choice of the value for  $\zeta$ .

(38) These results are however in line with Galí *et al* (2001).

Wide Model (AWM) data set, which covers the whole euro area and therefore is more complete than the NiGEM database. The results of the estimations are reported in Table D in Appendix B.<sup>(39)</sup> Overall, the results appear to be similar across the two databases. As for the reduced-form coefficients, the estimate of the coefficient on the backward-looking inflation term and that on the marginal cost term are lower than that obtained with the NiGEM data set, whereas the estimate of the coefficient on the forward-looking inflation term is higher. As for the structural coefficients, the major difference lies in the estimates of  $\omega$ , which is much lower than the estimate obtained with NiGEM data (which impacts on the estimates of  $\gamma_b$ ). Given that the proportion of the euro area covered by the NiGEM data is quite large, it is possible that there are other explanation than the different coverage for these discrepancies in the results. Among them, some differences in the methodology adopted to construct the data. Interestingly, the two databases bear identical (low) estimates of the discount factor  $\beta$ . This further reinforces the possibility that this economically implausible result might be indicative of an aggregation bias.<sup>(40)</sup>

### 5.3 *Stability analysis*

As recalled earlier, the reduced-form parameters in the (hybrid) NKPC are functions of structural parameters and as such expected to be time invariant, unless there are structural breaks. A failure of such time invariance would mean that policy decisions cannot be formulated on the type of inflation/real activity trade-off represented by the (hybrid) NKPC as they would indeed result in time-inconsistent estimates. In particular, the forecasting power of the (hybrid) NKPC would be severely undermined.

Notwithstanding its importance, a stability analysis of (hybrid) NKPC estimates has so far attracted little attention. By contrast, we make this analysis the focal contribution of this paper. Specifically, in the present section we first conduct a stability analysis using GMM rolling and recursive regressions on NiGEM data. That is, we conduct estimations over subsamples and analyse the variability across samples. Then, we test for the presence of structural breaks via recursive GMM estimations and a test for structural break with unknown breakpoints (Andrews’

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(39) Galí *et al* (2001) also perform an NKPC estimation on AWM data, although of a previous vintage than the one used here.

(40) Notice that only the NiGEM database enables us to investigate consistently the presence of this bias via a disaggregated, country-level analysis given that the AWM database does not explicitly include data for the individual member countries. It is for this reason that, in the paper, we have decided to concentrate on the euro-area results obtained with the NiGEM database rather than those obtained with the AWM database.



test). Rolling and recursive estimations, as well as Andrews' test, are performed for both reduced-form and structural parameters. The analysis on the structural parameters is indeed important for understanding how any observed instability or shift in the reduced-form parameters may be related to instabilities or shifts in the structural parameters.

### 5.3.1 *Rolling and recursive estimations*

Chart C.3 in Appendix C plots the output of the rolling estimations of the reduced-form coefficients  $\gamma_f$ ,  $\gamma_b$  and  $\lambda$  over time along with 95% confidence intervals.<sup>(41)</sup> The estimations are conducted using a window size of 60 quarters. Volatile lines indicate instability. In the rolling estimations conducted on US data (see Chart C.3a), the forward and backward-looking coefficients are remarkably stable at reasonable values. For the euro area (see Chart C.3b), both the implied forward and backward-looking coefficients are unstable from the late 1980s onwards. For both the United States and the euro area, the estimates of the coefficient for the real marginal cost are unstable. This is particularly true for the euro area, although in the United States the estimated coefficient is considerably more unstable than the coefficients on the forward and backward-looking inflation terms. In particular, in the United States the instability seems to start in the late 1980s. As for the euro area, the estimates of the three coefficients appear to be all remarkably characterised by one major 'episode' of large instability coinciding with the German unification.

Chart C.4 in Appendix C shows the rolling estimations of the structural parameters plotted against time. Overall the structural parameters also appear to be more stable for the United States than the euro area, but have relatively large standard errors for both countries. For the United States the estimated structural coefficients appear to be remarkably stable at plausible levels. For the euro area, the estimated coefficients become very unstable as soon as the date for the German re-unification is approached. The large confidence bands around the estimated discount factor ( $\beta$ ) cast some doubt over the implausibly small estimate we obtained and further points towards a technical (data-related) rather than an economic explanation of this result.

Recursive estimations, with either a decreasing or an increasing number of observations, confirm

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(41) Contrary to the results presented earlier, the rolling estimations were produced estimating **(11)** directly. This is because of the weak identification problem affecting the structural parameters, which implies that it is not possible to map exactly the stability results obtained for the structural parameters into those obtained for the reduced form.

the results of the above stability analysis.<sup>(42)</sup> Hence, the main conclusion remains that, whereas the estimates of the hybrid NKPC coefficients for the United States are stable and plausible (in both sign and size) from an economic point of view, the estimates for the euro area are unstable, with all the coefficients in the reduced-form hybrid NKPC more unstable and with larger standard errors than for the United States. Although a direct mapping of the structural coefficient results into the reduced-form ones is not possible, based on the results discussed here, it is however plausible to argue that the instability of the reduced-form relationship between inflation and real marginal cost for the euro area is due to structural causes rather than purely cyclical phenomena.

### 5.3.2 Breakpoint test

We perform the univariate Andrews' test for structural stability for each of the reduced-form parameters in the NKPC. The null hypothesis is that the value of a parameter that characterises the first  $t_s$  observations in the sample is significantly different from the value that characterises the last  $T - t_s$  observations, where  $t_s$  is an unknown time point and  $T$  denotes the time of the last observation. Test statistics are computed for each point in time in the sample (ie  $t_s$  varies between the first and the last time period).<sup>(43)</sup> Chart C.5 plots the sequences of test statistics over time, for the three reduced-form estimators, together with the critical values at the 1% and 5% significance level for the United States and the euro area. Given data availability and the requirements of the test, we perform the breakpoint test over the period 1976 Q2-1991 Q3 for the United States and 1981 Q2-1991 Q2 for the euro area.

The Andrews' test identifies a structural break as the period at which the highest value among the sequence of test statistics is obtained (see Andrews (1993)). However, the presence of a maximum (and therefore, a structural break) need not to exclude the presence of other structural breaks. Furthermore, structural breaks in the reduced-form estimates could possibly appear in the form of a slow transition from one parameter value to another. This would translate in intervals over which the statistics remain high. Hence, we will consider periods where the test statistics are higher than the critical values as an indication of a structural break.

As Chart C.5a shows, for the United States, a maximum in the series of test statistics for the

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(42) As the plots of the recursive estimations is very similar to those of the rolling estimations, we omit to include them. They are available from the authors on request.

(43) For details about the test statistic, see Appendix A.

coefficients on the lagged and forward-looking inflation terms is obtained in 1986, even though it is not significant at the 5% level. Consequently, we conclude that the test reveals no structural break in any of the coefficients of the Phillips curve. This result is consistent with the finding of Rudebusch and Svensson (1999).

As for the euro area (Chart C.5b), the sequences of test statistics for the coefficients on the lagged and forward-looking inflation terms both reach a maximum approximately at the time of the German unification. This is significant at the 5% level (even though not at the 1% level), for the coefficient on the forward-looking inflation term, but it is insignificant for the coefficient on the lagged inflation term. There is no evidence of a structural break for the coefficient on the marginal cost.

In summary, the results of the Andrews' test appear to support those of the rolling and recursive estimation analysis of a greater stability of the inflation/real marginal cost trade-off in the United States than the euro area, with the euro-area instability (partly) explained in terms of a German unification effect.

#### **5.4 Euro-area hybrid NKPC: a disaggregated analysis**

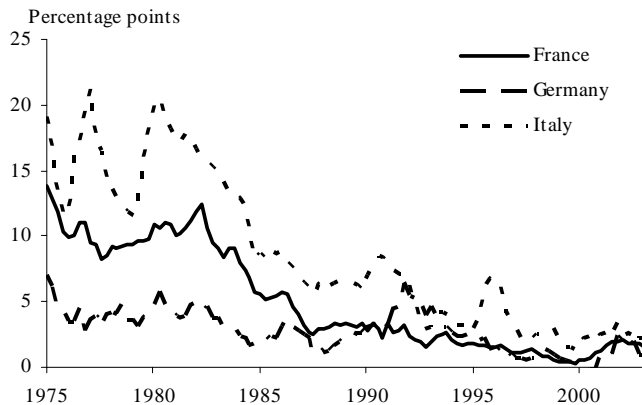
It is possible that the poor performance of the hybrid NKPC when estimated on euro-area data (at least compared with the hybrid NKPC estimated on US data) could be due to an aggregation bias.<sup>(44)</sup> The euro-area countries have experienced different inflation dynamics over the past 30 years (see Chart 3 below for annual data and Chart C.2 in Appendix C for quarterly data). Even if the (hybrid) NKPC provides a satisfactory description of the inflation dynamics in one member country, the aggregation over countries with different inflation dynamics could produce biased estimates of the hybrid NKPC coefficients at the euro-area level. To test the hypothesis of the existence of an aggregation bias, we estimate hybrid NKPCs for the three largest euro-area countries: Germany, France and Italy. Given the type of weights used to construct NiGEM euro-area data, these three countries account for 68% of the euro area.<sup>(45)</sup>

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(44) Another possible reason is the higher underlying steady-state inflation in the euro area with respect to the United States (see Bakhshi *et al* (2003)).

(45) The NiGEM weights are calculated from 1995 GDP under purchasing power parity for each of the countries.

**Chart 3: Annual GDP inflation rates**



#### 5.4.1 Estimation results

We estimated the reduced-form hybrid NKPC for Germany, France and Italy, for the period 1975 Q1-2002 Q4. We take into account the open-economy dimension of these countries by including the relative price of imported goods, adjusted by the ratio of the sum of exports and imports to GDP. That is, we estimate equation (8) on the individual-country data.

The instruments used in the estimations are: for Germany, a constant, five lags of GDP deflator inflation, four lags of wage inflation, oil price inflation, openness, relative import price and a deterministic dummy for the German unification, taking value zero until 1990 Q4 and value 1 thereafter; for France, five lags of GDP deflator inflation, four lags of wage inflation, oil price inflation, openness, marginal cost and a deterministic dummy for 1979 Q1 (to account for the second oil shock of the 1970s); for Italy, five lags of GDP deflator inflation, four lags of wage inflation, openness, marginal cost, output gap, seasonal dummies and a deterministic dummy for 1979 Q1.<sup>(46)(47)</sup> The results of the disaggregate analysis are reported in Table E.

The coefficients on the forward-looking inflation terms ( $\gamma_f$ ) are large and highly significant in all three countries. They average at approximately 0.46, which is lower than the value found for the

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(46) Contrary to Germany and France, in the estimation conducted on Italian data inflation is defined as quarterly change in the log of the GDP deflator. Hence the use of seasonal dummies. This definition is adopted because the measure of annual inflation for Italy is very volatile, which produces an implausible negative coefficient on the marginal cost.

(47) As before, we use a Bartlett kernel and a fixed bandwidth of 8.

euro area with either NiGEM data or the AWM database.

The coefficients on the backward-looking component are also large and significant for all three countries. However, they are considerably larger than the coefficients on the forward-looking inflation term for Germany and France, whereas it is well below it for Italy. One tentative explanation for the apparent smaller degree of backward-lookingness in Italy than in either Germany or France could be that the high inflation volatility experienced in Italy in the 1970s (relative to the other two countries) has made Italian firms less willing to anchor their prices to past levels of inflation. On the other hand, this same volatility makes the econometric results more uncertain.<sup>(48)</sup>

Finally, for Germany and France the estimated coefficients on the forward-looking and backward-looking inflation terms sum to a plausible value (0.97). For Italy, the two coefficients sum to a very low value, 0.87. The coefficient on the labour income share is only significant for France, whereas that on the open-economy dimension is only significant for Italy.

In conclusion, the hybrid NKPC parameters estimates appear considerably heterogeneous across the largest euro-area members. This may well explain the low or insignificant estimates obtained at the area-wide level.

#### 5.4.2 *Stability*

In this section, we repeat the stability analysis previously conducted on the aggregated euro area at the level of its largest members. The analysis is performed on the whole sample, 1975 Q1 - 2002 Q4.

Charts C.6, C.7 and C.8 in Appendix C show the results from rolling estimations of hybrid NKPC on French, Italian and German data respectively.<sup>(49)</sup> As for France, the coefficients on the forward and backward-looking inflation terms as well as on the marginal cost are relatively unstable, while the coefficient on the openness variable appear insignificant in all the subsamples. The result is

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(48) Another possible reason for the difference in these country-level results could be the different degrees of stickiness of information updating as discussed in Mankiw and Reis (2002).

(49) The window used in the chart is a 60-observation window. The dates refer to the end-point of the estimation window.

similar for Italy. In particular, the coefficients on the forward-looking inflation term appears considerably more unstable than the other coefficients: this result could be further interpreted as suggesting an adverse effect of the high inflation volatility in the 1970s on inflation expectations.<sup>(50)</sup> Finally, for Germany, all the estimated coefficients in the hybrid NKPC appear very unstable. Furthermore, even for Germany the open-economy dimension appear to be insignificant over all the subsamples considered.

Andrews' tests broadly confirm the instability of the reduced-form parameters; see Chart C.9 in Appendix C. For Germany, there is some tentative evidence of a break in the early part of the 1980s, but this is revealed only by the coefficient on the marginal cost, whereas the coefficients on the forward and backward-looking inflation terms appear stable. There are no specific signs of a structural break in the estimates for France. The estimated coefficients are quite unstable for Italy, with some evidence of a structural break in correspondence of the second oil shock in the 1970s. It could be argued that the size of the oil shock, which saw the oil price increasing from around US\$15 per barrel at the end of 1978 to around US\$40 per barrel by the end of 1979, combined with the almost concomitant introduction of the European Monetary System of pegged but adjustable exchange rates, might have had a significant impact on the Italian economy. We therefore sought further evidence of a structural break at the time of the second oil shock by re-estimating the hybrid NKPC for Italy with deterministic dummies that take the value 0 until 1978 Q4 and value 1 thereafter and that interact with the coefficients on the lead and lag of the inflation rate. However, the results do not sufficiently support the hypothesis of the existence of such a break.<sup>(51)</sup> A similar result is obtained also when re-estimating the hybrid NKPC for Germany and France with the same dummy.<sup>(52)</sup>

In summary, the combined results of the rolling estimations and the stability tests indicate that, over the period considered, there has not been any major structural shock which has contemporaneously affected the inflation dynamics in the three largest euro-area economies.

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(50) This result is therefore fully consistent with the GMM estimation results discussed above.

(51) The re-estimation produces insignificant coefficients on the lead and lag inflation terms over the period 1979 Q1-2002 Q1, whereas prior to 1979 Q1 the coefficients sum to a value below 0.8. Detailed estimation results are not reported in the paper but available from the authors on request.

(52) It could be argued that at the time of the second oil shock of the 1970s, Germany and France benefited of a more stable monetary regime than Italy, which soothed the effect of the shock.

## 6 Summary and conclusions

In this paper, we have estimated hybrid NKPCs using US and euro-area data. For the euro area, the analysis has been conducted both at an aggregated and disaggregated level. Importantly, we have subjected these estimates to a number of stability tests. The main results of our analysis are the following.

1. Our estimates of the coefficients on the lagged and expected future inflation terms are overall in line with previous studies for both the United States and the euro area. Specifically, the estimated coefficient on the forward-looking inflation term is larger than the one on the backward-looking term, this result being more true for the United States than the euro area, implying larger inflation persistence in the euro area with respect to the United States.
2. Estimates of the structural parameters underlying the reduced-form coefficients are also in line with previous findings. The fraction of backward-looking agents is estimated to be marginally smaller in the euro area than in the United States; price stickiness appears to be larger in the euro area. Concerning estimates of the discount factor, we found that while that for the United States is not significantly different from the value reported in other studies, our estimate is implausibly small for the euro area. This result appears to be robust across data sets (and in line with that of other studies).
3. Rolling (as well as recursive) estimation produces stable and plausible parameter estimates for the United States, but unstable parameters for the euro area. The breakpoint test analysis does not reveal any significant shift in any of the coefficients associated with either past and expected future inflation nor the real marginal cost for the United States. For the euro area, on the contrary, there is some tentative evidence of a structural break affecting the coefficients on past and expected future inflation in the late 1980s, possibly related to the German re-unification.
4. At the disaggregated level, the estimates are very similar for France and Germany, with the coefficient on the expected future inflation term smaller than that on the past inflation term. By contrast, the estimates based on Italian data reveal a considerably lower degree of inflation persistence than in the other two countries, a result possibly due to the larger inflation volatility experienced in Italy in the 1970s. Only in Italy does inflation appear to be significantly affected by the degree of openness of the economy.
5. Rolling estimations produce highly unstable estimates for Germany, and, albeit to a lower

degree, for Italy. The estimates for France appear to be considerably more stable over the period considered. The breakpoint test analysis evidences the possibility of instability of the coefficient estimates in the late 1970s for Italy and in the early 1980s for Germany. There is no evidence of structural breaks affecting inflation dynamics in France over the period considered. These conflicting results overall could indicate the presence of an aggregation bias in the results obtained with euro-area data. Specifically, we reckon this bias to be responsible for the implausibly low discount factor estimate for the euro area.

6. There are several implications for monetary policy makers from the results presented in this paper. Overall, the results discussed here suggest that policymakers should treat the forecasts generated by Phillips curves with some caution, as the structural parameters underlying the estimated relationships may be unstable over time. For the euro area, it also suggests that it may be useful to look at individual countries, in addition to the aggregate results. Moreover, policymakers should examine the results of a broad range of estimation methodologies to assess whether the forecasts generated by a Phillips curve model agree with other evidence. This is consistent with the approach currently taken in most major central banks.



## Appendix A: Andrews' test

Denote the full-sample length by  $T$ , and let  $t_s \in (1, T)$  denote the breakpoint date. The null hypothesis of the Andrews' test is that the parameters are stable.

$$H_0 : \gamma_t = \gamma_0 \quad \forall t \geq 1, \text{ for some } \gamma_0$$

If the point  $t_s$  is unknown, the alternative hypothesis is that the parameter is unstable at a point  $t_s$ , and takes the form

$$H_1 : \gamma_s \neq \gamma_T \quad \text{for some } s, T \geq 1$$

Given an alternative hypothesis of this form, we can use Andrews (1993) test and asymptotic critical values for the test statistics. We use the Wald test statistic, which is constructed from subsample and full-sample GMM estimations.

Assume that we want to test the stability of the parameter  $\gamma$  over the sample  $T$ . Let  $\widehat{S}_t$  denote the estimated weighting matrix in the GMM estimation over the sample  $T$ . Let  $\widehat{\gamma}_1(t_s)$ ,  $\widehat{S}_1(t_s)$  and  $\widehat{\gamma}_2(t_s)$ ,  $\widehat{S}_2(t_s)$  denote the estimated parameters and covariance matrices in the sample respectively  $(1, t_s)$  and  $(t_s, T)$ . The Wald-statistic is then defined as:

$$W_T(t_s) = T \times (\widehat{\gamma}_1(t_s) - \widehat{\gamma}_2(t_s))' \times \left( \frac{\widehat{S}_1(t_s)}{t_s} + \frac{\widehat{S}_2(t_s)}{T - t_s} \right)^{-1} \times (\widehat{\gamma}_1(t_s) - \widehat{\gamma}_2(t_s)) \quad \text{(A-1)}$$

We can then use the test statistics for  $t_s \in (1, T)$  at each point in the sample.

The test statistic at any point in time  $t_s$  can be interpreted as a test of whether the parameter is identical in the two samples  $(1, t_s)$  and  $(t_s, T)$ . When we reject the null hypothesis, we can consider the parameter estimates to be significantly different in the two samples  $(1, t_s)$  and  $(t_s, T)$ . Andrews (1993) suggested to look at the sequence of tests, and choose the highest value of the test statistic to identify the structural break.

## Appendix B: Tables

**Table A: ADF unit root tests. United States and euro area**

		Test statistic		
<b>United States</b>	$\pi_t$	1965-2003 -2.15	1965-1983 -2.95 (*)	1983-2003 -2.92 (*)
	$\Delta\pi_t$	-10.30 (**)	-10.16 (**)	-8.73 (**)
<b>Euro area</b>	$\pi_t$	-1.06	-3.14 (*)	-2.63
	$\Delta\pi_t$	-11.58 (**)	-3.99 (**)	-10.03 (*)

(\*) and (\*\*) means that the null hypothesis is rejected at the 5% and 1% significance level, respectively. The null hypothesis is non-stationarity.

**Table B: Reduced-form parameters. United States and euro area**

1975 Q1-2002 Q4	Coefficients with standard errors			Test for overidentifying restrictions
	$\gamma_f$	$\gamma_b$	$\lambda$	$J - statistic$ <i>p-value</i>
<b>United States</b>	0.52 [0.01]	0.47 [0.02]	0.04 [0.01]	11.88 [0.79]
<b>Euro area</b>	0.49 [0.03]	0.46 [0.03]	0.05 [0.02]	9.55 [0.77]

**Table C: Structural parameters. United States and euro area**

1975 Q1-2002 Q4	Coefficients with standard errors			Test for overidentifying restrictions
	$\beta$	$\theta$	$\omega$	$J - statistic$ <i>p-value</i>
<b>United States</b>	0.98 [0.02]	0.66 [0.03]	0.59 [0.05]	12.23 [0.79]
<b>Euro area</b>	0.82 [0.07]	0.76 [0.02]	0.58 [0.05]	9.55 [0.77]

**Table D: Comparative results across data sets, euro area**

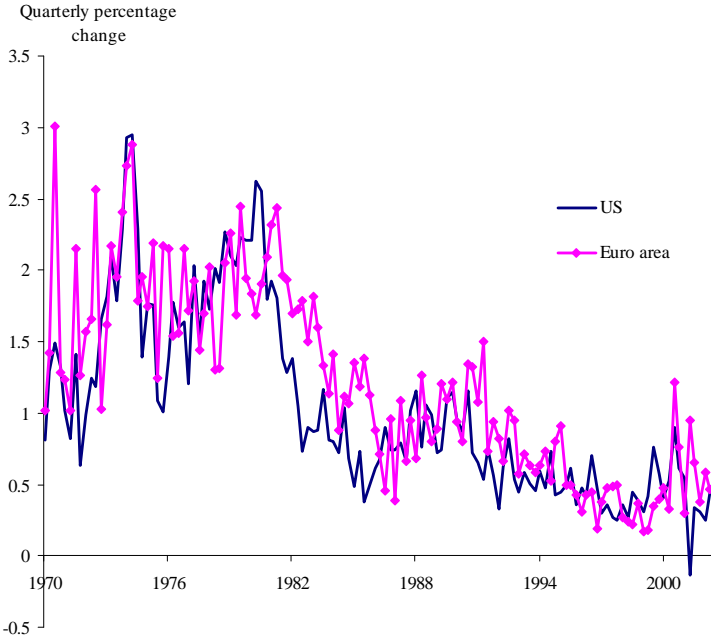
Sample: 1971 Q2-2002 Q3	<i>Parameters</i> <i>[standard errors]</i>	
	AWM data	NiGEM data
$\gamma_b$	0.35 [0.04]	0.45 [0.03]
$\gamma_f$	0.57 [0.02]	0.51 [0.02]
$\lambda$	0.05 [0.01]	0.06 [0.02]
$\beta$	0.82 [0.02]	0.82 [0.07]
$\theta$	0.75 [0.01]	0.76 [0.02]
$\omega$	0.38 [0.06]	0.58 [0.05]
J-statistic	18.37 [0.10]	9.55 [0.77]

**Table E: Reduced-form parameters. Germany, France, Italy**

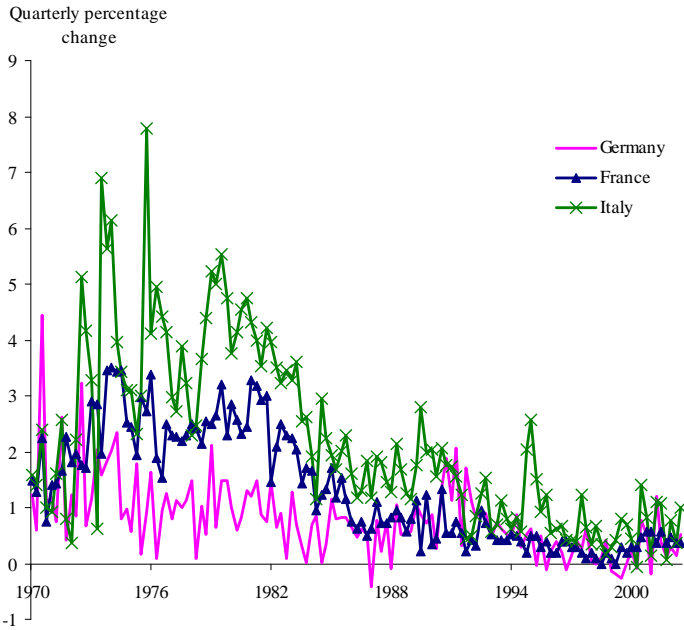
1975 Q1-2002 Q4	Coefficients with standard errors				Test for overidentifying restrictions
	$\gamma_f$	$\gamma_b$	$\lambda$	$\varphi$	<i>J - statistic</i> <i>p-value</i>
<b>Germany</b>	0.41 [0.04]	0.56 [0.03]	0.02 [0.01]	-0.00 [0.00]	10.54 [0.88]
<b>France</b>	0.41 [0.03]	0.56 [0.02]	0.05 [0.01]	0.00 [0.00]	9.62 [0.74]
<b>Italy</b>	0.57 [0.06]	0.30 [0.03]	0.02 [0.01]	0.01 [0.00]	10.52 [0.94]

Appendix C: Charts

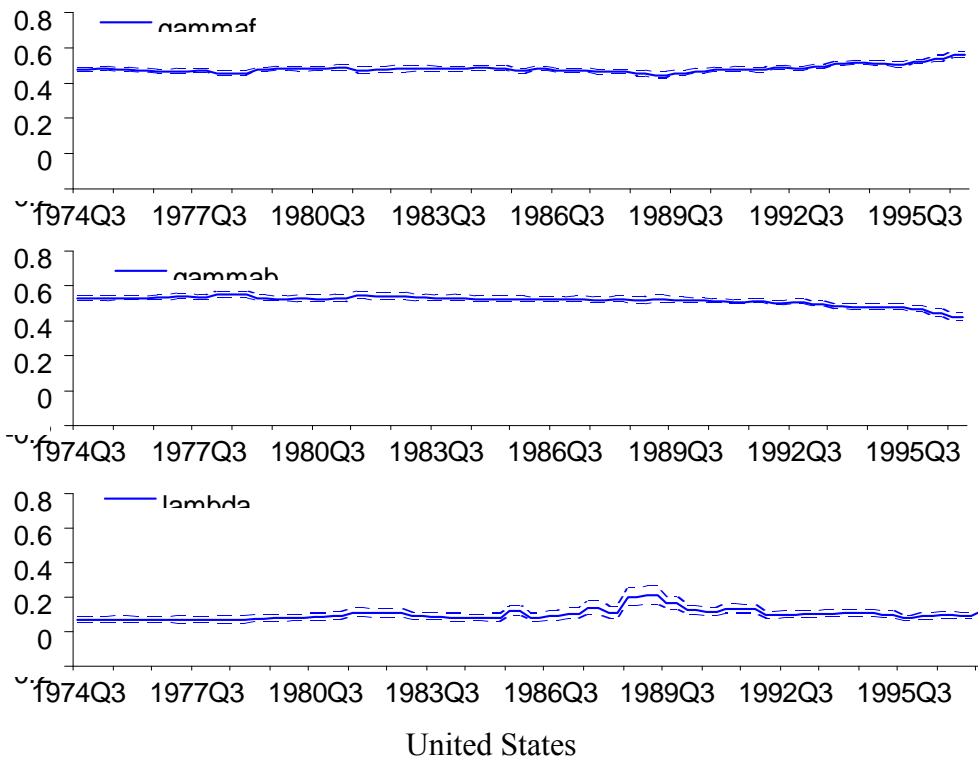
Chart C.1: Quarterly GDP deflator inflation in the euro area and the United States



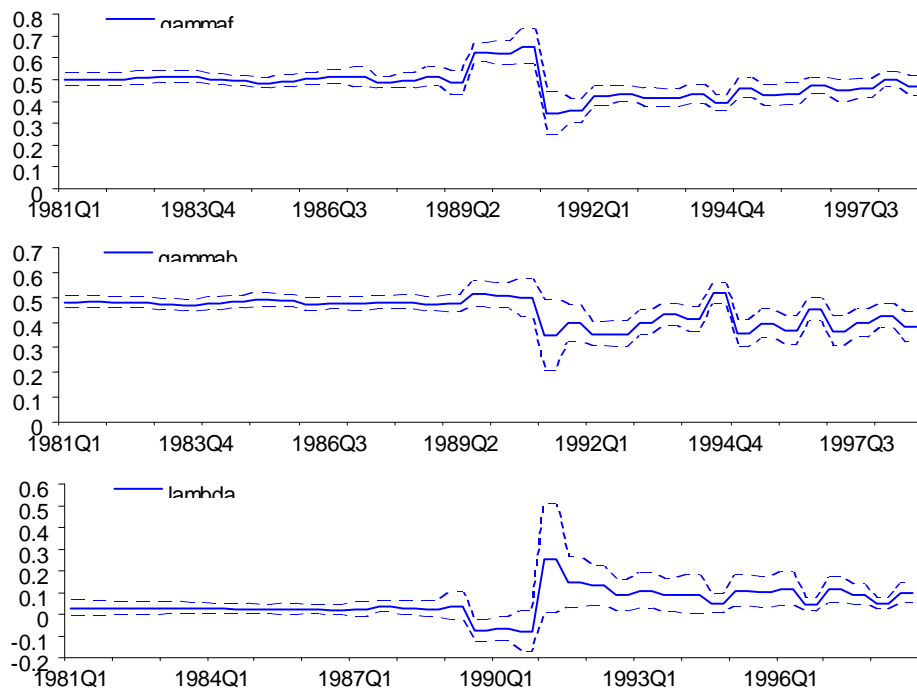
**Chart C.2: Quarterly inflation rates. Germany, France and Italy**



**Chart C.3: Rolling estimation. Reduced-form parameters, with 95% confidence intervals. United States and euro area**



United States



Euro area

**Chart C.4: Rolling estimation. Structural parameters, with 95% confidence intervals, United States and euro area**

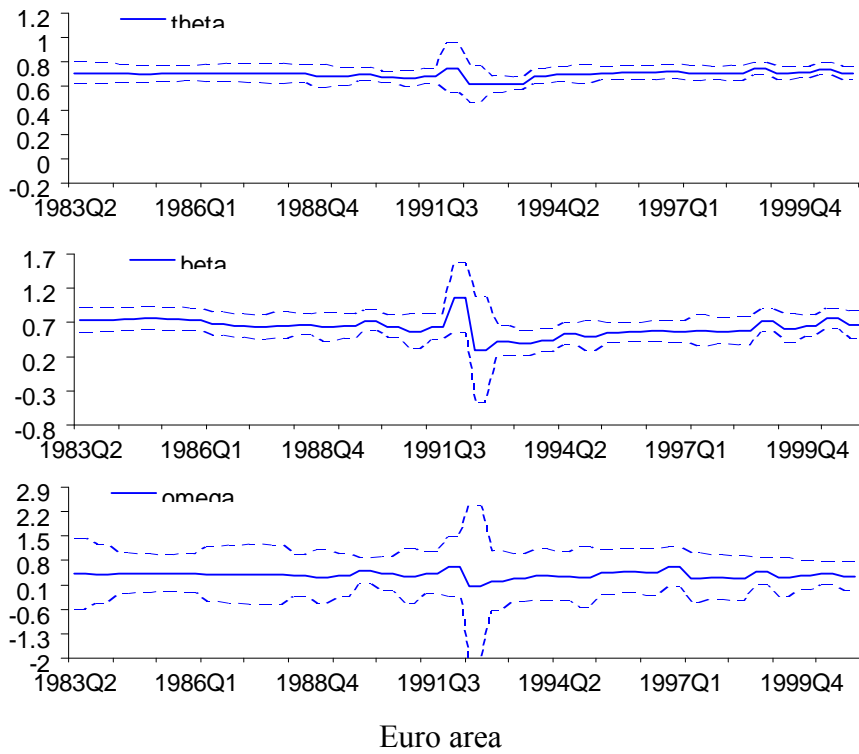
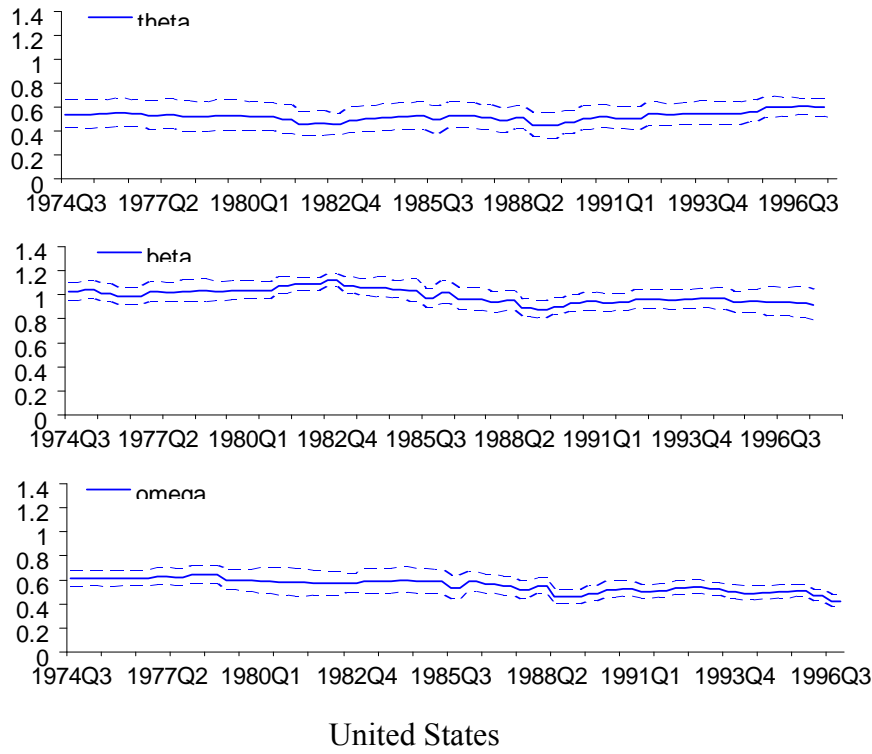
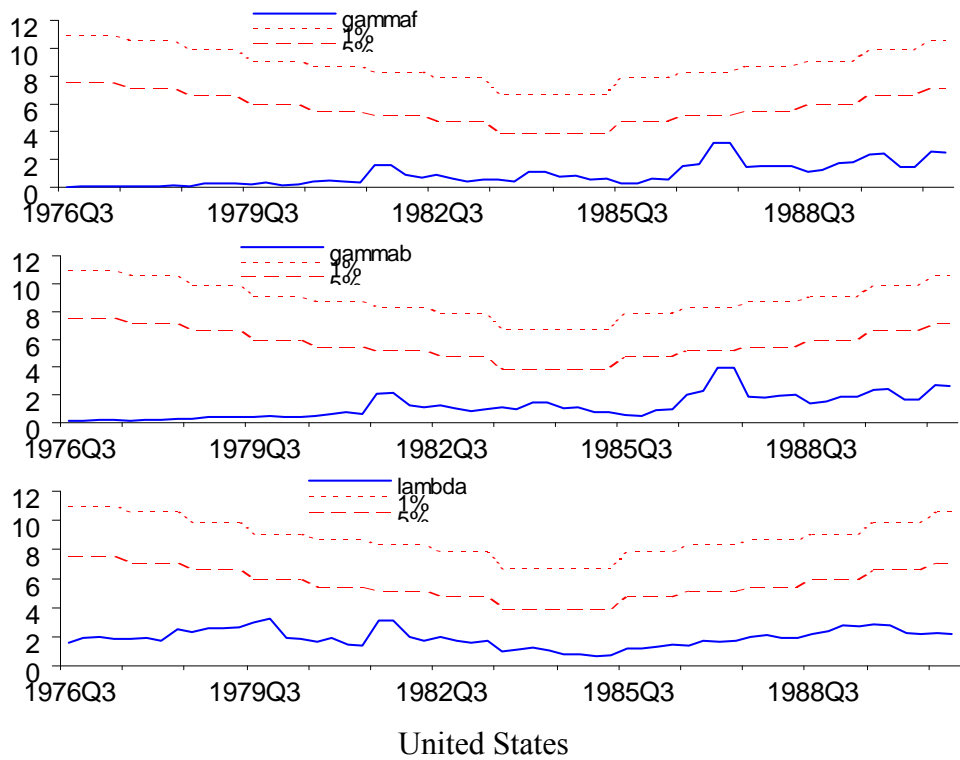
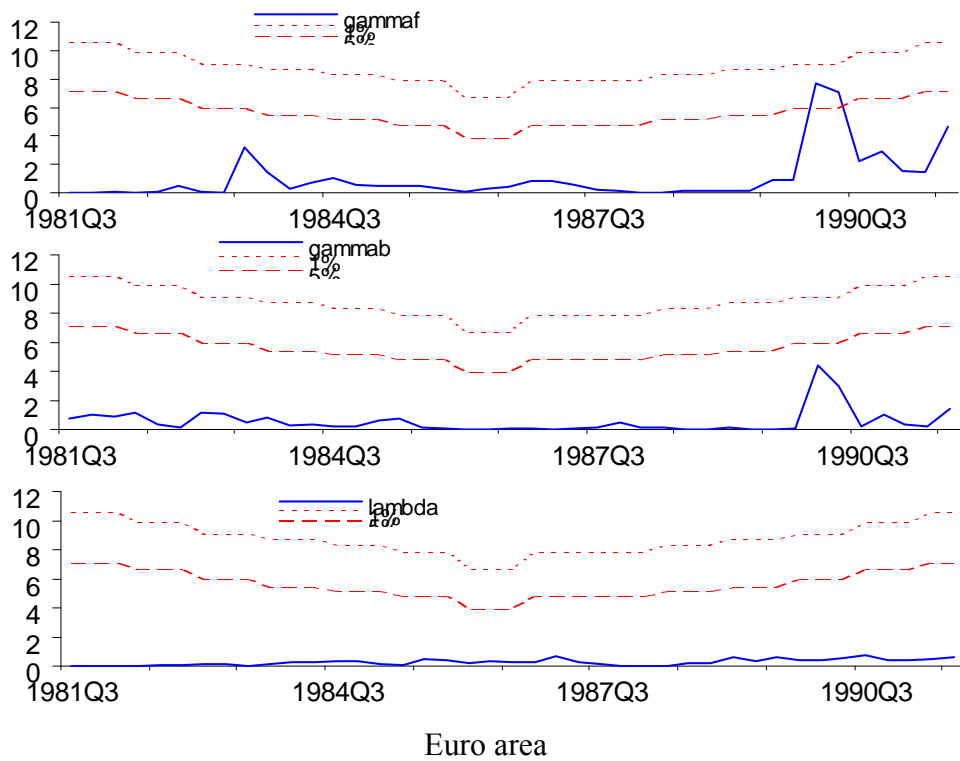


Chart C.5: Stability tests. United States and euro area



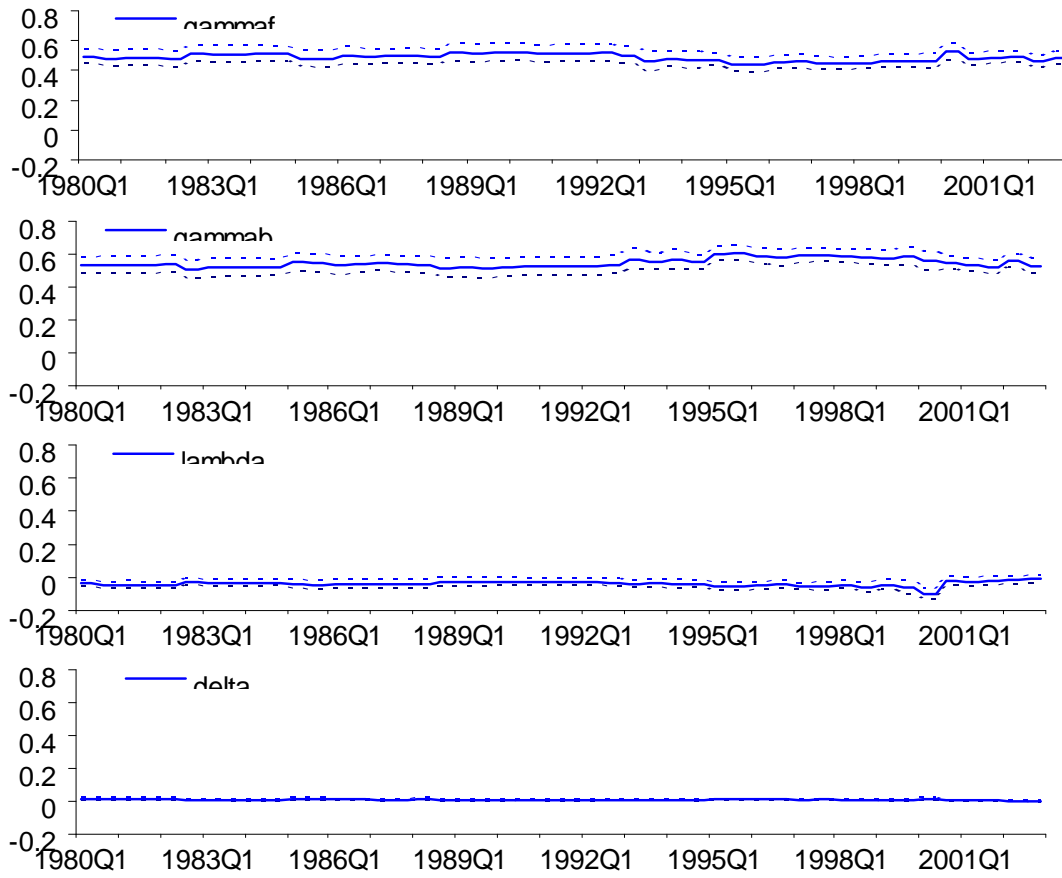
United States



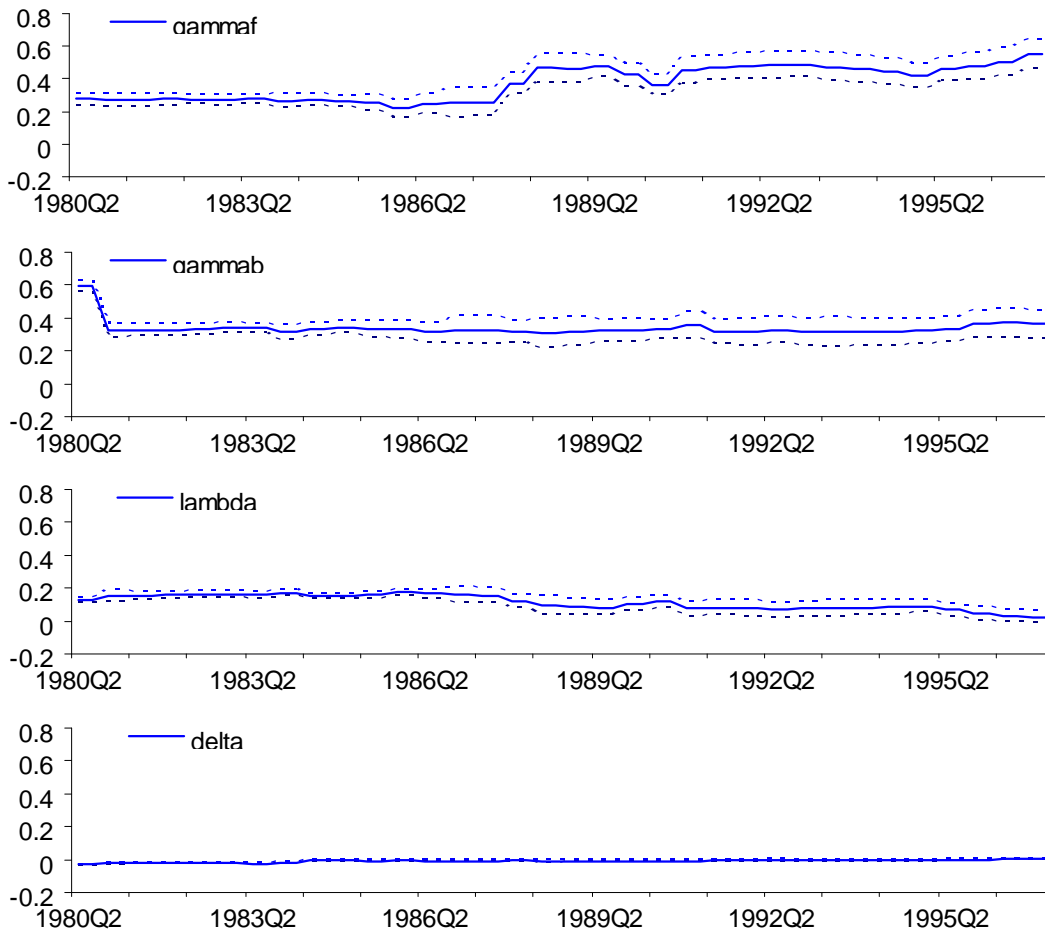
Euro area



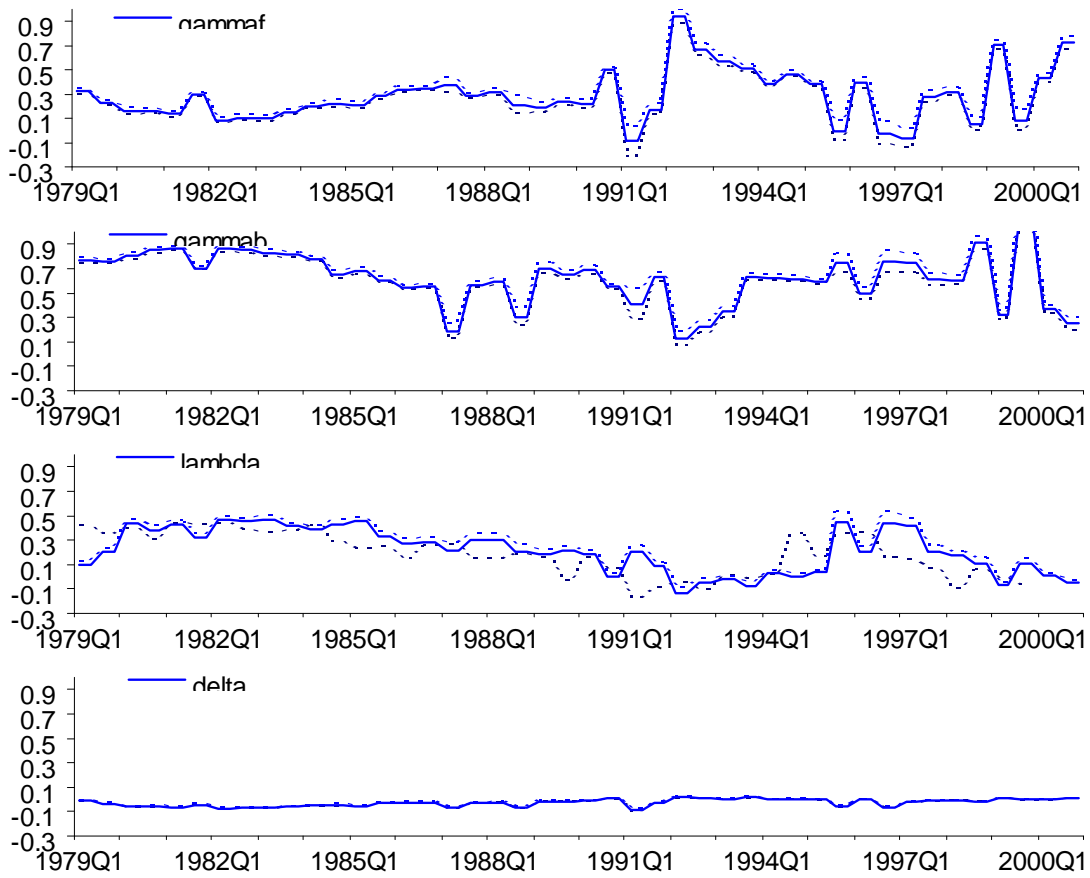
**Chart C.6: Rolling estimation. Reduced-form parameters, with 95% confidence intervals. France**



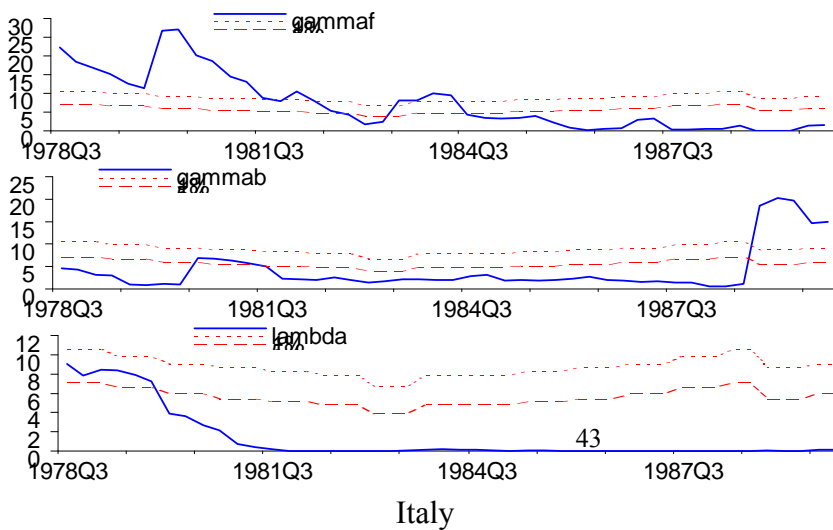
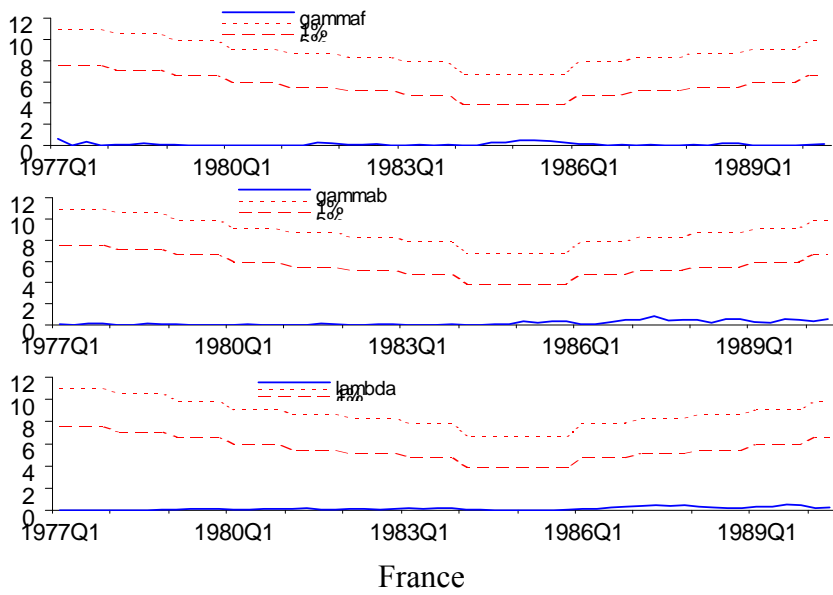
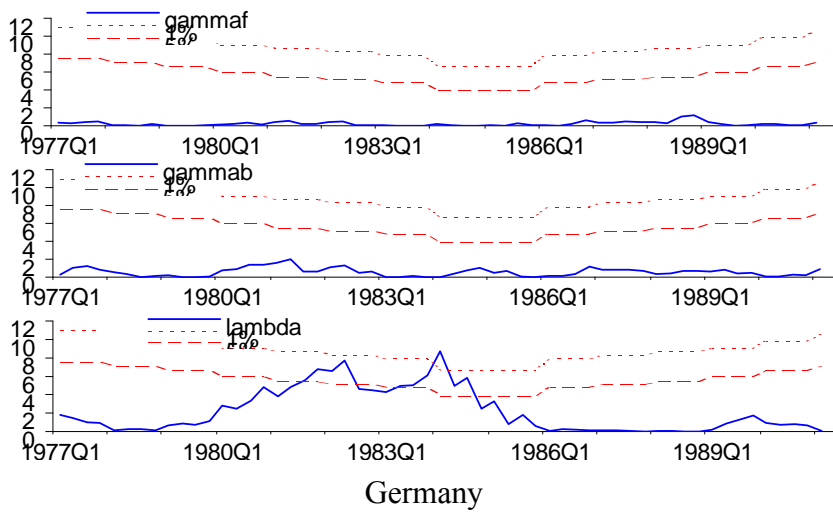
**Chart C.7: Rolling estimation. Reduced-form parameters, with 95% confidence intervals. Italy**



**Chart C.8: Rolling estimation. Reduced-form parameters, with 95% confidence intervals. Germany**



**Chart C.9: Stability tests. Germany, France and Italy**



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