NOTES D'ÉTUDES

ET DE RECHERCHE

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NER - E # 140

BANQUE DE FRANCE

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Is the Inflation-Output Nexus Asymmetric in the Euro Area?

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*We are grateful to Cécile Simon for her remarkable assistance and Jean-Guillaume Sahuc for his helpful comments. We also thank J.-M. Charpin who discussed in October 2002 a preliminary version of this paper at the Banque de France workshop chaired by M.-O. Strauss-Kahn. All remaining errors are ours. The views expressed herein are those of the authors and do not necessarily reflect those of the Banque de France. Corresponding author: GREQAM, Centre de la Vieille Charité, 2 rue de la Charité 13002 Marseille, FRANCE. Email: mustapha.baghli@univmed.fr. Tel: 33-4-91-14-07-23 and Fax: 33-4-91-90-02-27.

Résumé.

Cet article questionne l'hypothèse que l'inflation au sein de la zone euro est correctement appréhendée à l'aide d'une courbe de Phillips linéaire. Nous examinons dans un cadre non-paramétrique dans quelle mesure l'inflation est affectée par la croissance de l'économie. Un arbitrage inflation-croissance asymétrique est mis en évidence au sein de la zone euro aussi bien au niveau agrégé que des pays pris séparément.

Mots-Clés : Courbe de Phillips non linéaire ; stabilité des prix; méthode des noyaux.

Classification JEL : C14, C32, E31, E52

Abstract.

This paper challenges the assumption that the inflation process within the euro area is well-described by a linear Phillips curve and investigates in a nonparametric framework how inflation is sensitive to output growth. An asymmetric output-inflation trade-off is pointed out in the euro area at both aggregated and individual country levels.

Keywords: Nonlinear Phillips curve; price stability; kernel smoothing.

JEL classification: C14, C32, E31, E52

Résumé non technique.

Cet article s'intéresse à l'impact non linéaire du cycle d'activité sur l'inflation dans la zone euro et dans les principaux pays qui la composent (Allemagne, France et Italie). Il s'agit plus exactement d'identifier la forme fonctionnelle de ce lien potentiellement asymétrique et ainsi mettre en évidence un comportement particulier des autorités monétaires dans la zone Euro. En cas d'asymétrie, le banquier central confronté à une problématique de stabilité de l'inflation, pourrait adopter des décisions de politique monétaire différentes non seulement selon que l'économie se trouvera en récession ou en expansion mais aussi suivant l'ampleur de l'accélération ou du ralentissement d'activité. Dans ce but, nous avons étudié le processus dynamique d'inflation en Europe en vue de caractériser au sein d'une courbe de Phillips standard la forme de l'impact de l'écart de PIB sur l'inflation et cela sans aucun *a priori* sur cette forme.

Le principal enseignement de nos estimations est qu'il existe bien, tant au niveau agrégé de la zone euro qu'au niveau national, une telle asymétrie de sorte que les autorités monétaires pourraient répondre aux tensions inflationnistes de manière différenciée suivant le niveau et/ou le signe du décalage entre l'activité courante et potentielle. Ce lien asymétrique illustre le fait qu'en valeur absolue la sur-utilisation des capacités de production a un effet sur l'inflation supérieur à celui rencontré lorsqu'il y a une sous-utilisation des capacités productives. En outre, les tensions inflationnistes seront d'autant plus importantes que la production courante sera éloignée de son niveau d'équilibre. Enfin, il existe une région d'inaction (*watchful waiting*) autour de l'équilibre où l'inflation est sensiblement constante.

Non technical summary.

This paper focuses on the nonlinear effects of business cycle on inflation in the euro area and its major countries (France, Germany, and Italy). More precisely, we identify the functional form of this relationship, potentially asymmetric, and highlight a specific behaviour of monetary authorities in the euro area. As the central

banker aims at stabilising inflation, he may come to different decisions depending on the economy's position in the business cycle or the magnitude of the acceleration or slow down in activity. For this purpose, this paper investigates the inflationnary dynamics in Europe to identify the relationship between output gap and inflation in a Phillips curve framework without having a prior idea about its form.

The main findings of our estimates lead to establish such an asymmetric relationship in the euro area at both aggregated and individual country levels. This implies that monetary authorities may interact differently depending on the level and/or the sign of the gap between current and potential activity. This nonlinear nexus illustrates the fact that the effect of an over-utilization of productive capacities is higher in absolute value than in the case of an under-utilization. Moreover, the further the current production from its potential, the larger inflationnary pressures would be. Finally, our results reveal the existence of a *watchful waiting* zone around the equilibrium in which inflation is approximately constant.

1 Introduction

Although the original work of Phillips pointed out a nonlinear specification for inflation dynamics, the short-run trade-off between unemployment and inflation is traditionally assumed unchanged over time. However, many theoretical models of price-setting behaviour suggest that economic activity has a nonlinear effect on inflation. The theoretical arguments in support of a nonlinear specification are, for example, capacity constraints, agents confused by price shocks, menu costs, downward nominal wage rigidity or oligopolistic market.[1] As each of these arguments implies a particular nonlinearity in the Phillips curve (convex, “kinked”or concave function), from a policy perspective the choice of a source of nonlinearity is not neutral. Indeed, on the one hand, the output cost of fighting inflation will vary with the shape of the Phillips curve, and on the other hand, given the monetary policy transmission lags, stronger or weaker incentives for preemptive policy tightening will exist to counter expected inflationary pressure.

Within a nonparametric framework, this paper examines such possibilities for the Phillips equation in France, Germany, Italy and the euro area. Unlike the existing literature on this topic, we do not assume an *ad hoc* parametric form for the nonlinear Phillips curve. On the contrary, we resort to the kernel smoothing in order to uncover the correct functional form of policymakers’ inflation preferences.

The organization of the paper is as follows. In Section 2, the nonlinear Phillips curve is outlined, along with the kernel-based estimator. In Section 3, the compared results are discussed.

2 Nonlinear Phillips Curve and Kernel Smoothing

The nexus between inflation and economic activity has been increasingly investigated (Turner, 1995; Laxton, Meredith and Rose, 1995; Clark, Laxton and Rose, 1995 and 1996; Debelle and Laxton, 1997; Eisner, 1997; Laxton, Rose and Tambakis, 1998; Filardo, 1998). In this body of literature, the type of nonlinearity has been

[1] See Stiglitz (1997) and Dupasquier and Rickett (1998) for a comprehensive review.

usually assumed on the basis of policy implications without prior econometric testing. Following the lines of these different works, we study in this paper the following nonlinear specification:

$$\pi_t = \pi_t^e + F\left(y_{t-k} - y_{t-k}^*\right) + \xi_t = \sum_{i=1}^{4} \alpha_i \pi_{t-i} + F\left(y_{t-k} - y_{t-k}^*\right) + \xi_t \tag{1}$$

$$\Delta\pi_t = \sum_{i=1}^{3} \alpha_i^* \Delta\pi_{t-i} + F\left(y_{t-k} - y_{t-k}^*\right) + \xi_t \tag{2}$$

with π_t and π_t^e the actual and expected inflation rates, y_t and y_t^* the actual and potential GDPs (in logarithm), ξ_t a zero-mean supply shock and $k \geq 0$. $k > 0$ allows a delay in the apparition of inflationary pressures.[2] Unlike these studies, the inflation response to output gap, $F(.)$, has an unknown parametric form and therefore needs to be estimated from the data.

A nonparametric kernel type estimator allows to let the data themselves settle about their classification instead of stipulating in advance any functional form or smoothness constraints on $F(.)$. Nevertheless, in nonparametric regression the small number of observations with respect to the number of elements to estimate may distort the accuracy of the results. A useful and interesting approach enabling to overcome this so-called problem of "dimensionality curse" is the focus on:

$$E\left(\Delta\pi_t \middle| z_{t-k} = z, \Delta\boldsymbol{\pi}_t^{lag} = \mathbf{0}\right) \equiv F(z) \tag{3}$$

with $z_t = y_t - y_t^*$ the output gap, $\Delta\boldsymbol{\pi}_t^{lag} = (\Delta\pi_{t-1}, \Delta\pi_{t-2}, \Delta\pi_{t-3})$ and z a given point of the sample space of the z_{t-k}'s.

The idea behind the analysis of such a conditional expectation function is to reduce the dimension of the problem related to the estimation of function $F(z_{t-k})$. Thus, this amounts to examining the response of $\Delta\pi_t$ only in terms of the past output gap z_{t-k} without caring about the lagged changes in the inflation rate. By leaving some lagged inflation terms ($\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$ and $\Delta\pi_{t-3}$) out of the regression,

[2] Given the formulation of the Phillips curve we consider that the inflation process is integrated of order 1. For each of the measures of inflation, we tested this hypothesis and confirmed that the inflation rate in question is $I(1)$.

we consider that the inflation dynamics in the studied European countries are sufficiently well captured with the first-order lag of the inflation rate π_{t-1}. We impose that inflation has a purely backward-looking dynamics and that this specification fits the data for inflation quite well. As in the traditional linear framework, eliminating from the regression some lagged changes in the inflation rate could affect the estimate of the nonparametric part of the model because of a possible correlation between the output gap z_{t-k} and past inflation developments. Nevertheless, this potential bias on the estimate of $F(z_{t-k})$ is, in our context, less problematical than in the linear case. Indeed, as an alternative to proxy inflation expectations, we filtered out the linear component from the data. This method consisted in regressing $\Delta\pi_t$ on $\Delta\boldsymbol{\pi}_t^{lag}$ and in using the residuals resulting from this regression rather than $\Delta\pi_t$ to estimate the condition expectation (A.4). As it led broadly to the same conclusions obtained when we omit the lagged inflation terms, we can consider that $\Delta\boldsymbol{\pi}_t^{lag}$ yields no new significant information with regard to $\Delta\pi_{t-1}$ and can therefore be neglected.[3]

By assuming in (2) that

$$E\left(\xi_t \middle| z_{t-k} = z, \Delta\boldsymbol{\pi}_t^{lag} = \mathbf{0}\right) = 0$$

and

$$Var\left(\xi_t \middle| z_{t-k} = z, \Delta\boldsymbol{\pi}_t^{lag} = \mathbf{0}\right) = \sigma^2\left(z, \mathbf{0}\right),$$

the kernel-based estimator of (A.4) is given by the well-known Nadaraya-Watson estimator

$$\widehat{F}(z) = \frac{(hT_1)^{-1}\sum_{t=k+1}^{T}\Delta\pi_t\mathcal{K}(\frac{z-z_{t-k}}{h})}{(hT_1)^{-1}\sum_{t=k+1}^{T}\mathcal{K}(\frac{z-z_{t-k}}{h})} \tag{4}$$

where $T_1 = T - k$, $\mathcal{K}(.)$ is any univariate kernel and $h > 0$ is the bandwidth parameter.

To circumvent a potential endogeneity problem, we imposed the standard orthogonality condition. This restriction seems to be quite weak given the results

[3]Research on semiparametric estimation methods could be an important extension to the present study. Indeed, for example, the smoothing splines technique could constitue a relevant and informative methodology for estimating both the unknown nonlinear function $F(.)$ and the parametric part.

obtained with the pre-filtering stage. We adopted the gaussian density as $\mathcal{K}(.)$. As a comparison, alternative kernel functions were tested but this led to similar estimates of the function. We used mean-integrated-square-error driven bandwidth estimate and the validity of the estimated regression response was ensured by calculating the bias-corrected and accelerated (BC_a) bootstrap confidence intervals devised by Efron (1987).

3 Estimated Nonlinear Output Inflation Dynamics

At country level, quarterly seasonally adjusted series for real GDP and prices are drawn from the appropriate national accounts. The quarterly euro area data used in this paper are extracted from the AWM database compiled by Fagan, Henry and Mestre (2005). The sample period runs from 1973Q2 to 2003Q4 for the euro area, 1972Q1 to 2003Q4 for France and 1970Q1 to 2003Q4 for Germany and Italy.

For the three countries in question, the output gap z_t corresponds to the difference between the actual (log) gross domestic product (GDP) (y_t) and the (log) potential GDP (y_t^*) measured by means of an Hodrick-Prescott filter. We implemented alternative indicators of trend GDP (y_t^*), such as the Kalman filter based measure, a structural measure of potential GDP or the production function related measure. However, the results remained broadly identical. By contrast, for the euro area we retained a macroeconomic-model-based estimate of potential GDP, namely the one proposed by Fagan, Henry and Mestre (2005). As outlined by Galí and Gertler (1999), Orphanides (2001 and 2003) and Orphanides and Williams (2002), the use of an empirical proxy for the output gap leads inevitably to a kind of measurement error. Nonetheless, we obtained an asymmetric output-inflation trade-off in Europe irrespective of the empirical proxy for the gap used. The only difference was the degree of nonlinearity pointed out. Although this error-in-variable problem may affect the accuracy of the estimates, we can therefore consider that this particular issue is not of great significance in our context.

Figure 1 depicts the consecutive estimated response for the three major euro area

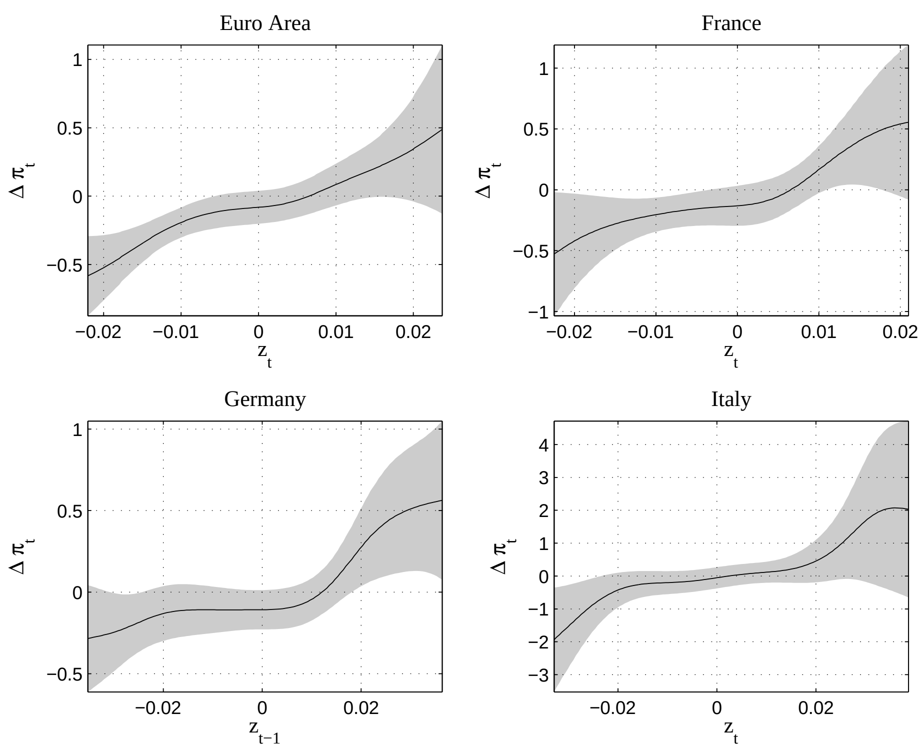

Figure 1: *Kernel-based estimation of the function response,* $\widehat{F}(z)$*, with a gaussian kernel,* $\mathcal{K}(u) = \frac{1}{\sqrt{2\Pi}} \exp(\frac{1}{2}u^2)$*, MISE-driven bandwidth estimates and the lower and upper limits of the 95% confidence interval.*

countries and for the aggregated zone. We selected the appropriate delay parameter $k \in \{0, 1\}$ for the output gap so that the presence of nonlinearity was the most evident and the most "economically" credible.[4] As a whole, the results confirm an expected asymmetry in aggregated data but also at country level. The inflation-output trade-off exhibits an outstanding S-shaped type nonlinearity reconciling the supporters of a convex Phillips curve (Turner, 1995; Laxton, Meredith and Rose, 1995; Clark, Laxton and Rose, 1995 and 1996; Debelle and Laxton, 1997) with those in favor of a concave curve (Eisner, 1997; Stiglitz, 1997). For the three countries, as well as for the euro area, this asymmetry illustrates the fact that (large) excess demand has a stronger effect in increasing inflation than (large) excess supply has in decreasing it. The theoretical arguments supporting that inflation responds strongly to positive excess demand may be the existence of capacity constraints and/or downward rigid nominal wages. Furthermore, a range of values is noticed for output gap where the German inflation rate is roughly constant. This may generate, in certain theoretical models, multiple equilibria. It is interesting to notice in Figure 1 that even under conditions of full utilization of capacity, *i.e.* $z_t = 0$, the change in the inflation rate may be significantly different from 0. This could illustrate the fact that the output gap has a more or less persistent lagged effect on inflation.

From a policy perspective, these empirical findings imply that European monetary authorities have to act quickly and significantly in the presence of an overheating of the economy, comparatively to a "linear" world. Indeed, if this extra inflation still persists, an increase in inflation stemming from an "excessive" contemporaneous economic activity will necessitate a larger recession in the future. Our estimated functions do not correspond to an exact rule for inflation-output trade-off, *i.e.* a reaction function indicating for example the thresholds where the policy adjustment need to be implemented.[5] Nevertheless, consistently with the characteristics of busi-

[4]From a theoretical standpoint, $k > 0$ can be linked to the existence of significant Calvo-type nominal rigidities so that a significant fraction of firms cannot optimally revise their prices in each period.

[5]To achieve this, in addition to the estimation of the Phillips curve, it would be convenient to estimate a small monetary model formed by an IS relationship and a loss function for the policymaker preferences. But again, the question arises of the parametric form of this function.

ness cycles in European economies, these kernel-based estimates give an interesting and immediate insight into the implications for demand management policies and bring a strong justification for preemptive policies, particularly in the presence of signs of overheating.

A Appendix

A.1 Data

For all countries and the euro area, the data are quarterly. The inflation rate π_t is defined by $\Delta_4 p_t = p_t - p_{t-4}$ with p_t the logarithmic price level. For the three countries, the output gap z_t corresponds to the difference between the actual (log) gross domestic product (GDP) (y_t) and the (log) potential GDP (y_t^*) measured by means of an Hodrick-Prescott (HP) filter.[6] By contrast, for the euro area we retained a structural measure of potential GDP, namely the one proposed by Fagan, Henry and Mestre (2005). For France, the sample period runs from 1972Q1 to 2003Q4. We used different indicators of prices such as the harmonized index of consumer prices (HICP), the consumption deflator (excluding or not energy) and the GDP deflator. As we obtained better results with the harmonized index, we finally retained this prices measure in our paper for France. For Germany and Italy, the sample period runs from 1970Q1 to 2003Q4. We also used different indicator of prices such as the HICP, the consumption deflator and the GDP deflator. Finally, the better results were obtained with the CPI based inflation for Germany and the HICP for Italy. For the euro area, the sample period runs from 1973Q2 to 2003Q4. The alternative indicators of prices used were again the HICP, the HICP excluding energy, the consumption deflator and the GDP deflator. As a result, we retained as indicator of inflationary pressures the HICP.

Traditional empirical work on the Phillips curve considers the output gap as the relevant indicator of real economic activity. Nevertheless, consistent with the theory it would have been better to instead use real marginal cost measures. Yet, as

[6] We used alternative indicators of trend GDP (y_t^*), such as the Kalman filter based measure or the production function related measure. But, the results globally remained identical.

showed by Galí and Gertler (1999), there is an approximate log-linear relationship between the real marginal cost and the output gap such that we used the output gap as the relevant determinant of inflation. Note also that the real marginal cost is usually approximated by the real unit labor cost which does not allow to take account of reductions in social security contributions implemented in Europe and France notably.

A.2 Stationarity Tests

Table 1: Stationarity Tests

	$KPSS$	R/S	S_1	S_2	S_3	S_4	S_5	Br	ADF
π_t^{FR}	2.05(4)	2.93(2)	47.73	115.17	59.87	7.28(5)	21.63	0.035	-1.18(3)
$\Delta\pi_t^{FR}$	0.11(3)	1.19(1)	-3.69	-3.75	-3.71	-3.71(3)	-1.36	0.003	-4.81(1)
z_t^{FR}	0.051(2)	0.91(2)	-2.63	-2.71	-2.74	-0.29(2)	-0.27	0.002	-6.51(0)
π_t^{GER}	2.24(2)	2.42(2)	50.52	50.31	45.09	13.45(3)	-14.22	0.051	-1.11(2)
$\Delta\pi_t^{GER}$	0.06(1)	1.15(1)	-2.22	-2.29	-2.23	-2.24(2)	-0.27	0.001	-6.79(1)
z_t^{GER}	0.056(2)	0.84(3)	-2.14	-2.62	-2.56	-0.29(2)	-0.72	0.001	-3.56(1)
π_t^{IT}	2.79(2)	2.92(2)	16.90	44.15	32.39	11.33(3)	-19.76	0.062	-1.13(3)
$\Delta\pi_t^{IT}$	0.14(1)	1.12(1)	5.32	3.24	4.81	25.41(2)	-1.13	0.002	-6.22(1)
z_t^{IT}	0.038(2)	0.77(2)	-1.14	-1.16	-1.17	-1.16(2)	0.36	0.002	-3.52(1)
π_t^{EA}	3.28(2)	2.46(3)	53.86	70.42	36.94	11.85(2)	-14.58	0.080	-1.49(2)
$\Delta\pi_t^{EA}$	0.12(1)	1.09(2)	10.27	10.12	9.43	8.69(1)	-0.57	0.016	-7.57(1)
z_t^{EA}	0.108(2)	1.39(2)	-1.06	-1.08	-1.21	-1.07(2)	-0.58	0.003	-3.26(3)

All tests were conducted at a 5% level. In parenthesis is provided the optimal truncation lag used for computing the corresponding statistic.

The statistics $KPSS$, R/S and $\{S_i\}_{i=1}^5$ correspond respectively to the Kwiatkowski-Phillips-Schmidt-Shin statistic (Kwiatkowski, Phillips, Schmidt and Shin, 1992), the modified rescaled range statistic (Lo, 1991) and the Bierens-Guo S_i statistics (Bierens and Guo, 1993). The R/S, S_1, S_2, S_3 and S_4 statistics test for the null of level stationarity. The S_5 statistic for testing for the trend stationarity null hypothesis is also presented. The $KPSS$ statistic has both null hypothesis. The $KPSS$ and S_2 tests are one-sided, the other tests are two-sided. Br is Breitung's (2002) test for the null hypothesis of non stationarity. It is a left-tailed test that applies to data demeaned/detrended by OLS or local-to-unity GLS. For the $KPSS$, R/S

and S_4 statistics, the optimal truncation lag were selected by means of information criteria and Andrews' data-dependent formulas. All tests were conducted by doing a distinction according to the type of null hypothesis, level or trend stationarity, and by evaluating the statistical significance of different forms of the deterministic trend in the testing regression.

In Table 1 are reported the results of testing for stationarity. The stationarity tests lead us to consider that the inflation rate π_t is $I(1)$ whereas the output gap z_t is $I(0)$. These testing procedures are robust to the existence of nonlinearity, especially Breitung's nonparametric test. Unlike the other testing devices, Breitung's (2002) testing procedure is a nonparametric methodology since it does not require the practitioner to specify the short-run dynamics of the process and to estimate the long-run variance (pivotal limiting distribution under the null of integration).

A.3 Kernel Smoothing

The kernel-based estimator is given by the well-known Nadaraya-Watson estimator:

$$\widehat{F}(z) = \frac{(hT_1)^{-1}\sum_{t=k+1}^{T}\Delta\pi_t\mathcal{K}(\frac{z-z_{t-k}}{h})}{(hT_1)^{-1}\sum_{t=k+1}^{T}\mathcal{K}(\frac{z-z_{t-k}}{h})} \tag{5}$$

where $T_1 = T - k$, $\mathcal{K}(.)$ is the gaussian kernel, $\mathcal{K}(u) = \frac{1}{\sqrt{2\Pi}}\exp(\frac{1}{2}u^2)$ and $h > 0$ is the bandwidth parameter that controls the degree of smoothness of the response $\widehat{F}(z)$.[7] Since (5) is the kernel-based estimator of the conditional mean $E(\Delta\pi_t \,\big|\, z_{t-k} = z, \boldsymbol{\Delta\pi}_t^{lag} = \mathbf{0}) = \frac{\int \Delta\pi f(\Delta\pi,z)d\Delta\pi}{\int f(\Delta\pi,z)d\Delta\pi}$, it can be interpreted as a weighted sum of the changes in the inflation rate $\Delta\pi$. The denominator, the probability density function of $\widehat{z}_{t-k}$ at z, ensures that the weights sum up to 1. The weight assigned to the t^{th} observation is high if the distance of $\widehat{z}_{t-k}$ from z is large. The consistency is ensured as both $\Delta\pi_t$ and $\widehat{z}_{t-k}$ are stationary processes. For further details concerning the asymptotic behaviour and the properties of such an estimator, we recommend consulting, from the very extensive literature available on this subject, Singh and Ullah (1985), Härdle (1990), Härdle and Linton (1994) and Wand

[7] As outlined by Silverman (1986), in kernel density estimation, it is generally well accepted that the choice of kernel is relatively unimportant in comparison to the choice of bandwidth.

and Jones (1995).

A critical step in the procedure is to determine in an optimal way the bandwidth h. As the bias (the variance) of kernel smoother declines (increases) with the smoothing parameter, we solve this classical bias/variance trade-off by selecting as optimal value for the bandwidth the value of parameter that minimizes the mean integrated squared error (MISE) criterion. Indeed, a growing body of literature dealing with this concern highly recommends using an automatic or hi-tech bandwidth selector (Härdle (1990); Härdle and Linton (1994)).

For the three countries, and the euro area as well, we selected the appropriate delay parameter $k \in \{0, 1\}$ for the output gap such that the presence of nonlinearity was the most evident and the most "economically" credible. Actually, the choice of a relatively "small" lag aims at having an underlying theoretical model in line with Calvo-type sticky price model. In particular, if $k > 0$, we expect that there exists some significant nominal rigidity such that only a fraction of firms can optimally revise their prices in each period. The construction of a reference model is beyond the scope of the present paper, but it is the topic of current research.

A.4 Bootstrap Confidence Intervals

The validity of the estimated regression response is ensured by calculating the bias-corrected and accelerated (BC_a) bootstrap confidence intervals devised by Efron (1987). The BC_a method adjusts the percentile method in such a way that it possesses higher-order improvements, notably in terms of coverage probability. Briefly sketched, the construction of the confidence intervals with a number B of bootstrap samples is, for $b = 1, ..., B$, as follows:

- Generate a random sample of length $T - k$ by resampling the $(\Delta\pi_t, \widehat{z}_{t-k})$ pairs, *i.e.* draw with replacement one of $\Delta\pi_t$'s coupled with its corresponding $\widehat{z}_{t-k}$ to form one observation of the bootstrap sample b. This operation yields a complete bootstrap sample, say $\{\Delta\pi^*(b), \widehat{z}^*(b)\}$.

- From the bootstrapped sample, estimate

$$E(\Delta\pi_t \Big| z_{t-k} = z, \Delta\boldsymbol{\pi}_t^{lag} = \mathbf{0}) \equiv F(z)$$

  by kernel smoothing as previously described and store the estimated regression response $\widehat{F}^{(b)}(z)$.

- Elaborate Efron's BC_a confidence interval for a given significance level.

The number of bootstrap replications for constructing the BC_a confidence interval is efficiently determined by means of the three-step method precisely conceived for this purpose by Andrews and Buchinsky (2002). This methodology is designed to choose the optimal B that minimizes the distance between the ideal or asymptotic bootstrap confidence band and the B repetitions-based BC_a band. As we can see, this extremely computer-intensive method led to an estimated confidence interval suffering a kind of "boundary effect". This edge effect is current in nonparametric regression and causes the estimated confidence band to sometimes broaden near the edges of the regression input space(Wand and Jones, 1995).[8]

A.5 Preliminary results

A.5.1 Pre-filtering stage

In a first stage we tested the unit root hypothesis for the inflation process π_t and we concluded that the inflation rate is indeed integrated of order 1. As a result, we opted for the following formulation of the Phillips curve:

$$\Delta\pi_t = \sum_{i=1}^{3} \alpha_i^* \Delta\pi_{t-i} + F\left(y_{t-k} - y_{t-k}^*\right) + \xi_t$$

The reason for using a nonparametric kernel type estimator was to let the data themselves settle about their classification instead of indicating in advance any functional form or smoothness constraints. Nevertheless, in nonparametric regression estimation the small number of observations with respect to the number of elements to

[8]The BC_a bootstrap confidence intervals were programmed in Gauss 3.2.32 and run on a 1.7Mhz Pentium IV.

estimate might distort the accuracy of the nonparametric estimates. In order to elude this potential problem of "dimensionality curse", we focused on:

$$E\left(\Delta\pi_t \middle| z_{t-k} = z, \Delta\boldsymbol{\pi}_t^{lag} = \mathbf{0}\right) \equiv F(z)$$

with $z_t = y_t - y_t^*$ the output gap and $\Delta\boldsymbol{\pi}_t^{lag} = (\Delta\pi_{t-1}, \Delta\pi_{t-2}, \Delta\pi_{t-3})$. We lessened the dimension of the problem of estimating the function $F(z)$ by considering that the inflation dynamics in the European countries in question were captured sufficiently well with the first-order lag of the inflation rate. Obviously, this needed to restrict further the model by imposing that the coefficient associated with π_{t-1} was equal to one. We assume that inflation has a purely backward-looking dynamics and that this specification fits the data for inflation quite well

As in the traditional linear framework, eliminating from the regression some lagged changes in the inflation rate ($\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$ and $\Delta\pi_{t-3}$) may affect the estimate of the nonparametric part of the model because of a potential correlation between the proxy for demand pressure (the output gap z_t for example) and past inflation developments[9]. Nevertheless, this potential bias on the estimate of the unknown function $F(z_{t-k})$ appeared less severe than in the linear case for two reasons.

First, instead of leaving lagged inflation ($\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$ and $\Delta\pi_{t-3}$) out of the regression, we used an alternative method that consisted in filtering out the linear component present in the model. More precisely, we regressed $\Delta\pi_t$ on $\Delta\boldsymbol{\pi}_t^{lag}$ and we computed and stored the associated residuals noted $\widehat{\epsilon}_t$. Following this linear prefiltering step, we then estimated the unknown function $F(z_{t-k})$ using the residuals $\widehat{\epsilon}_t$ rather than $\Delta\pi_t$. To achieve this purpose, we concentrated on the conditional expectation:

$$E\left(\widehat{\epsilon}_t \,|z_{t-k} = z\,\right) \equiv F(z)$$

The results obtained with this preliminary filtering step were globally similar to those obtained when we omit the lagged inflation terms $\Delta\boldsymbol{\pi}_t^{lag}$. Indeed, for each country we derived a S-shaped relationship between output and inflation largely similar to

[9]For a discussion concerning the empirical importance of the backward-looking component in inflation dynamics, see Jondeau and Le Bihan (2005) and Ruud and Whelan (2006), among others.

the function estimated without the pre-filtering stage. This confirms that even if our results will be conditional on this "omission", this is of no great significance since results appeared quite close with a linear pre-filtering step. The lagged inflation terms $\Delta\boldsymbol{\pi}_t^{lag}$ yielded no new significant information.

Research on semiparametric estimation methods in our context could be an important extension to the present study. But such an analysis is beyond the scope of our article since, given the number of available semiparametric estimation methods, it should constitute a separated study. Indeed, for example, the smoothing splines technique could constitue a relevant and informative methodology for estimating both the unknown nonlinear function $F(.)$ and the parametric part. Nevertheless, some issues can be immediately raised. It is more likely that this estimation method suffers from a problem of "dimensionality curse" and therefore that it necessitates a large sample size to avoid biased estimates. Besides, we can expect that the smoothing splines technique provides no significant results with regard to the kernel-smoothing-based results since Silverman (1984) proposed an asymptotic approximation to the spline estimator in terms of a kernel smoother.

Second, the kernel smoothing approach allowed us to estimate the regression response, and the function F in particular, without reference to a specific form. This flexible tool in analyzing unknown regression relationships is based on a local approximation of the condition mean $E\left(\Delta\pi \left| z_{-k}, \Delta\boldsymbol{\pi}^{lag}\right.\right)$. In order to identify the estimate of F, $\widehat{F}(z)$, at a given point z (*i.e.* to eliminate the linear part) we estimated this conditional mean around a particular point in the sample space of $\left(z_{t-k}, \Delta\boldsymbol{\pi}_t^{lag}\right)$, namely $(z, \mathbf{0})$. It consisted in locally weighted averaging the $\Delta\pi_t$'s around this particular point. The conditional mean restriction,

$$E\left(\xi_t \left| z_{t-k} = z, \Delta\boldsymbol{\pi}_t^{lag} = \mathbf{0}\right.\right) = 0,$$

constitues a sufficient condition that eliminates any endogeneity problem stemming from the omission of some lagged inflation terms. The study of performances of instrumental-variable-type estimators in a semiparametric context, particularly for the local averaging techniques, were subject to few studies[10]. As such an analysis is beyond the scope of this article, we circumvent this potential endogeneity problem by imposing this orthogonality condition. In addition, one should note that this restriction seems to be not too strong given the results stemming from the pre-filtering stage. Indeed, because the lagged inflation terms $\Delta\boldsymbol{\pi}_t^{lag}$ provided in our context no new significant information (with regard to $\Delta\pi_{t-1}$), we can expect that these lagged inflation terms are not significantly correlated with the lagged output gap z_{t-k} such that the error term ξ_t is mean independent of these lags.

Finally, we can also notice that, in addition to the preceding shortcoming related to lagged inflation terms, our kernel-based estimate of the function F suffers also from the well-known drawbacks inherent to the nonparametric estimators. Indeed, it gives the key features of the the function F but does not allow to explicitly derive the parametric expression of F. On the other hand, the bias (the variance) of kernel smoothers declines (increases) with the smoothing parameter (see below for details about the manner we solve this classical bias/variance trade-off).

The classic Phillips curve has been over the last decade disputed by the so-called New-Keynesian Phillips curve, especially by its "hybrid" version. Empirical estimates of the hybrid specification provided conflicting results and policy oriented implications (Jondeau and Le Bihan, 2005; Ruud and Whelan, 2006). A growing body of literature has been devoted to explain such a conclusion by determining the relevant forcing variable to introduce in the Phillips curve. Two alternative forcing variables are traditionally considered to proxy the real marginal cost, the output gap and the real unit labor cost. We opted for the output gap as the relevant indicator of real economic activity. Indeed, as showed by Galí and Gertler (1999), there is an approximate log-linear relationship between the real marginal cost and the output gap. And by selecting the real unit labor cost, we would have neglected in the

[10]Semiparametric IV estimators are still an extensive area for ongoing research. For example, recently Park (2006) proposed an IV-type estimator based on a semiparametric regression model in which the error term is correlated with the nonparametric part. This estimation method relies on a two-step procedure involving filtering steps on the basis of instruments. Note that this estimation technique was not used since it required a prior estimation of the nonparametric part and that it needed to select quite "arbitrarily" some instruments.

analysis the influence of reductions in social security contributions implemented in Europe and France notably. Besides, the average and marginal cost are likely to have different cyclical properties over the sample.

A.5.2 Output gap

It is well-documented (Galí and Gertler, 1999; Orphanides, 2001 and 2003; Orphanides and van Norden, 2002; Orphanides and Williams, 2002) that the use of detrended actual GDP (derived from the Hodrick-Prescott filter for example) or structural-macroeconomic-model-based estimate of GDP (as performed for the euro area by Fagan, Henry and Mestre, 2005) can lead to a kind of measurement error. Indeed, we do not observe a time series for the potential output. This error-in-variable problem may affect the accuracy of the estimates, but also impede the detection of nonlinearities present in the relationship. As a result, in order to get results unconditional on the output gap estimates, we used, for all countries and the euro area, different empirical measures for the output gap. We employed a wide array of methods that go from univariate approaches such as Hodrick-Prescott and band pass filters to multivariate or structural methods (such as the Kalman filter technique or the production function approach) for France and the euro area. Although the error-in-variable problem may remain more or less present according to the proxy of output gap retained, we obtained an asymmetric output-inflation trade-off in Europe irrespective of the empirical measure used. The only difference was the extent of nonlinearity in the relationship. Finally, we selected the measure of potential GDP based on a Hodrick-Prescott filter for individual countries and the Fagan *et al.* macroeconomic model for the Euro area, apparently because both resulting output gaps presented dynamics more in line with inflation developments.

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