Questioni di Economia e Finanza

(Occasional Papers)

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SURPRISE! EURO AREA INFLATION HAS FALLEN

by Marianna Riggi* and Fabrizio Venditti*

Abstract

Between 2013 and 2014, following the recession triggered by the sovereign debt crisis, euro-area inflation decreased sharply. Although a fall in the inflation rate was to be expected, given the severity of the recession, professional forecasters failed to anticipate it. A possible explanation for this forecast failure lies in a break in the cyclicality of inflation, which was unaccounted for in forecasting models. We probe this explanation in the context of a simple backward-looking Phillips curve and find that the sensitivity of inflation to the output gap has recently increased. We rationalize this result through a structural model, in which a steepening of the Phillips curve arises either from lower nominal rigidities (a decrease in the average duration of prices) or from fewer strategic complementarities in price-setting due to a reduction in the number of firms in the economy.

Contents

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The ECB never expects inflation to deviate from the target of just under 2 per cent. Yet each month inflation undershoots, and the ECB is apparently taken by surprise.

Münchau W., 2014 [1]

1 Introduction

Debate over the Phillips curve has gained momentum since the 2008 financial crisis. In the course of the recession that followed that crisis, a puzzle had emerged, in that inflation in advanced countries had not fallen as much as a traditional Phillips curve and past experiences would predict, given the severity and the length of the recession (Williams, 2010) and (Ball and Mazumder, 2011). The decline of euro area inflation between 2013 and 2014 is pointing in the opposite direction. Following the sovereign debt crisis, the euro area fell into a severe recession, which generated sizeable output losses in the countries more directly involved, in particular Greece, Spain, Portugal, Italy and Ireland. The recession was followed by a sharp fall in consumer price inflation, with core (net of food and energy) inflation dropping in the euro area to historically low levels in mid-2014. Two features stand out in this rapid inflation decline. First, it is broad based across countries, although relatively more intense in countries that have been hit the hardest by the sovereign debt crisis. Second, it was not anticipated by professional forecasters, neither for the euro area as a whole, nor for the larger member countries. This is particularly surprising if we consider that the fall in economic activity that most of the euro area countries have experienced after 2011 has generated significant gaps between actual and potential output in these economies. According to estimates from the European Commission the output gap in the euro area stood at 3% in 2012 and ranged from a minimum of 1% in Germany to a maximum of 13% in Greece. Estimates from the OECD point to an even larger slack. Given the large and persistent output gaps in these economies, it would have been possible to expect a prolonged period of inflation weakness. Why, then, were forecasters surprised? The question is a clear concern for policymakers since, in an environment in which expectations are not rational and monetary policy is constrained by the zero lower bound, sufficiently large negative shocks may start a deflationary spiral, in which downward revision of expectations and downward pressure on output and inflation follow each other (Evans and Honkapohja, 2009).

Drawing from the econometric literature, which has long identified structural breaks as the main cause of forecast failure, this paper asks whether the recent deep and long lasting fall of economic activity has been accompanied by an increased sensitivity of inflation to cyclical

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[1]Munchau, W (2014), Draghi is running out of legal ways to fix the euro, Financial Times, 17 August.

conditions (measured by the coefficient of the output gap in a backward looking Phillips curve) that, in turn, could account for the over-prediction of inflation.

Our empirical analysis focuses on the euro area and the four largest economies. Econometric tests of parameter instability and the estimation of time varying coefficient models yield three main conclusions. First, the relationship between inflation and the output gap displays significant instability in recent years in the euro area as a whole as well as in Italy, France and Spain (but not in Germany). Second, in Italy, France and the euro area as a whole the sensitivity of inflation to business cycle conditions has increased substantially in 2013. This is consistent both with the fact that the global recession of 2008-2009 had a muted effect on consumer price inflation in these countries, and with the fact that the sovereign debt crisis, albeit with some delay, significantly raised the prociclicality of inflation.[2] In Spain, on the other hand, such a break occurred much earlier, right after the first global recession of 2008-2009.[3] Third, looking at sub-aggregates of the consumer price index shows that the results are mainly driven by developments in the goods sector.[4]

Our findings are in line with the evidence put forward in a number of papers that investigate the inflation-unemployment relationship in the context of the U.S. economy. Stock and Watson (2010), for instance, find that unemployment is more useful for predicting inflation in recessions than in booms, a feature also highlighted in Olivei and Barnes (2004). Stella and Stock (2012), using a multivariate unobserved component model that implies a time varying Phillips curve, find that since 2008 the slope of the curve has become steeper.

Having documented in some detail the changing patterns of inflation cyclicality, we provide a tentative explanation on the basis of the New Keynesian Phillips curve that relates inflation to expected future inflation and to the output gap. This model suggests two possible mechanisms behind the rise in inflation cyclicality highlighted by our empirical analysis. First, that rise might reflect lower nominal rigidities, i.e., a higher frequency of price adjustment. Some support for this explanation comes from Fabiani and Porqueddu (2013) who show that in Italy, between 2006 and 2012, the average duration of consumer prices has indeed nearly halved to five months, from eight months between 1996 and 2001. Ongoing research at the Eurosystem level through the Wage Dynamics Network should provide better data and more evidence on this issue.

The other channel is a smaller degree of strategic complementarities in price setting, due to a significant drop in the number of firms in the economy. In the model an exogenous decrease

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[2]Estimates based on full fledged DSGE model indeed find that the Phillips curve was relatively flat up to 2012 in Italy, see Riggi and Santoro (2013)

[3]For Spain very similar results are obtained with a different methodology by the Bank of Spain, see the article *Variation in the cyclical sensitivity of Spanish inflation: an initial approximation*, in Banco de Espana, Economic Bulletin, July-August 2013.

[4]For simplicity of exposition in the paper we will simply call *goods* the *non-energy industrial goods* subcomponent of the consumer price index

in the number of firms implies a higher desired markup and a higher elasticity of prices with respect to the output gap. Whereas more work and better data is needed to shed light on the plausibility of this second line of explanation, preliminary analysis on firms demography suggests that the collapse of domestic demand induced by the sovereign debt crisis resulted indeed in the closure of many businesses.

A third factor at play, which cannot be taken into account by the model given the hypothesis of rational expectations, is a downward adjustment of inflation expectations, which could be feeding back to actual inflation. Although a robustness check (in which we control for inflation *forecasts* elicited from professional forecasters) leaves unaltered our baseline results, an adjustment of expectations to the prolonged inflation weakness can not be ruled out and needs to be carefully monitored.

The paper is structured as follows. Section 2 motivates the paper by discussing how forecasters overestimated inflation in 2013 and 2014. Section 3 presents the empirical analysis. Section 4 presents the theoretical model through which we interpret the empirical results. Section 5 concludes.

2   The inflation surprise

The slowdown of inflation in the course of 2013 and 2014 was not correctly predicted by forecasters. To illustrate this point, Figure 1 shows 4 steps ahead inflation forecast errors computed (as the difference between actual and expected inflation) on the basis of Consensus Economics surveys for the euro area and its four largest member countries (Germany, France, Italy and Spain), together with the price of oil (in euros). Data go from 2001 to the second quarter of 2014. Three interesting features emerge:

- Between 2001 and 2008 professional forecasters systematically under-predicted inflation in the euro area. Two plausible explanations for this outcome stand out. First, in order to comply with the Maastricht criteria at the end of the Nineties many euro area countries pursued disinflationary policies, mainly by restraining wage growth. However, after joining the Monetary Union, this policies were relaxed and wages outpaced productivity growth (Busetti et al., 2007), thus fostering inflation rates. Second, between 2003 and 2008, oil (and other commodity) prices increased continuously, pushing euro area inflation to a maximum of 4.0% in July 2008.
- Given the unexpected collapse of oil prices that followed the financial crisis, professional forecasters over-predicted inflation developments in 2009 by large amounts. As oil prices returned to pre-crisis levels starting in 2010, forecast errors once again turned positive.

- In 2013 and 2014, following the recession induced by the 2011 sovereign debt crisis, inflation slowed down much more sharply than anticipated by Consensus. The crucial difference between the recent over-prediction of inflation and the one in 2009 is that the former has occurred in a period of stable oil prices, indicating that a significant share of the recent inflation surprise is due to the weakening of core inflation.

As forecast failure in the econometric literature is frequently associated with structural breaks, a plausible explanation for this inflation surprise is a change in the slope of the Phillips curve. The next section turns to an extensive empirical investigation of this hypothesis.

3 Empirics

We start by positing a general specification of the Phillips curve:

$$\pi_t^j = \alpha + \beta \pi_{t-4}^j + \gamma y_{t-h}^i + \eta_t \tag{1}$$

where we interact different (year on year) inflation measures π_t, indexed by j, with different measures of economic slack y_t, indexed by i, and allow for the measure of slack to enter the equation with different lags h.

We estimate equation (1) for (i) different inflation measures (ii) different gaps between actual and trend economic activity (iii) different dynamic specifications on data from 1999Q1 to 2014Q2.[5] The idea is to test whether there is a change in the inflation/slack relationship and whether this feature is robust across different measures of both and also to uncertainties on the dynamic relationship between inflation and the business cycle.[6]

We consider nine different measures of inflation and fourteen measures of economic slack, as described in Table 1. The first two measures of inflation are based on the headline Harmonized Index of Consumer prices (HICP) and its *core* component (computed as the headline index net of food and energy). We then include the corresponding indicators net of the impact of indirect taxation, which are computed by Eurostat under the assumption that indirect tax increases are passed through fully and immediately to final consumer prices.[7] Next we consider

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[5]In our analysis we discard data prior to the inception of the euro, as member countries had very different monetary/fiscal/exchange rate regimes, so that any findings obtained using data before 1999 are unlikely to shed any light on current developments

[6]In a survey of the empirical New Keynesian Phillips curve estimation Mavroeidis et al. (2013) document around 1000 possible different specification for the Phillips curve arising from combinations of various inflation measure and forcing variables.

[7]The relevance of such an indicator has risen in recent years, owing to the sequence of indirect taxation hikes with which stressed countries have tried to reduce fiscal deficits to restore market confidence. Notice that in cases in which VAT increases are not passed through to final prices (like in Italy in October 2013 or in France in January 2014), these measures of inflation *understate* actual inflation, so that they must be interpreted with

the two core inflation sub-components (Goods and Services) that roughly proxy tradables and nontradables and therefore convey important information on which sector is bearing most of the adjustment at the current juncture. Finally, we consider compensations per employee, to check wether inflation weakness is feeding through wages. As for economic activity, we use measures of slack based on GDP, hours worked and the unemployment rate. To deal with the uncertainty on how to extract a trend from these series, we experiment with two different filters (Hp and Band-pass) and apply them both ex post or in pseudo real time. Finally we include the output gaps estimated from the European Commission and the OECD, which are based on more sophisticated approaches where a production function is used to compute potential output.

We show inflation rates and output gaps in Figure 2 and Figure 2.[8] Notice that a clear difference emerges between the gaps based on statistical filters and those computed by the EC and the OECD. Statistical filters attribute most of the fall in GDP induced by the sovereign debt crisis to its trend component, so that, at the end of the sample, output gaps are small. According to the measures provided by the EC and the OECD, on the other hand, output gaps are still largely negative.

3.1 End of sample instability tests

The first analysis we conduct is based on structural break tests. Since we are interested in parameters instability at the end of the sample,conventional break tests are of little use, as they have very low power when change-points occur towards the end of the sample. Busetti (2012) introduces a number of new tests designed to have high power in such circumstances. Among the many statistics proposed in the paper we consider Locally Most Powerful (LMP) tests and its extensions (Exp-L and Sup-L tests). The LMP statistics has the following form:

$$L_\pi = \hat{\sigma}^2 (T - \pi T)^{-2} \sum_{t=\pi T+1}^{T} S_t'^{-1} S_t$$

where $\hat{\sigma}^2 = \hat{u}_t' \hat{u}_t / (T - k)$, $S_t = \sum_{j=t}^{T} \hat{u}_j x_j$ and $V = T^{-1} \sum_{t=1}^{T} x_t x_t'$, and π is the last fraction of the sample where the break is supposed to have occurred. Busetti (2012) also derives two modifications of this L test, denoted Sup-L and Exp-L, computed as follows:

$$Sup - L = \underset{\pi \epsilon \Pi}{Sup}(L_\pi)$$

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due caution as a complement, rather than a substitute, of the actual index.

[8]Only ex post gaps are reported in the figure.

$$Exp - L = \log \int_{\pi \epsilon \Pi} \exp(L(\pi))\pi d\pi$$

The null hypothesis is that the parameters in (1) are time invariant, while the alternative hypothesis is that they evolve as driftless random walks. We test for a break in the last 25 and 10% of the sample. In Table 2 we report the results of these tests when applied to the euro area and to the four largest euro area countries. The table is organized in country specific panels where the columns identify the measure of inflation under investigation and the row the type of test conducted. In each cell we report the fraction of rejections of the stability hypothesis (at the 5% confidence level) in a regression of that specific inflation measure on a given measure of output gap and a given lag specification. To be clearer, let us consider the top left cell of the Table referring to (i) the euro area (ii) overall inflation (Overall)(iii) the L test (iv) the last 25% of the sample. For this specification 56 tests are conducted, corresponding to the 56 possible specification of the forecasting equation (2) that one can obtain given the 14 measures of output gap and the lags from 1 to 4 that we consider. The 0.14 figure reported in the top left cell is then telling us that in 14% of these 52 cases the L-test rejects the null of stability in the last 25% of the sample. When we move to headline inflation at constant fiscal impact (Overall-X) we obtain 16% of rejections and so on. A number of patterns may be identified:

1. Evidence of instability is quite widespread for the euro area, Italy, France and Spain, while for Germany a model with constant parameters cannot be rejected in almost all the cases.

2. Exp-L and Sup-L tests tend to reject much more than the simple L tests. For the euro area, for example, Exp-L and Sup-L tests reject in almost 100% of the cases for all inflation measures, excluding wage inflation. Similar evidence emerges for Italy, France and Spain. Since Monte Carlo results in Busetti (2012) suggest that the Exp-L test performs better than the other ones when the time of the break is unknown, we read these results as strongly supportive of a break having occurred.

3. Focussing on the core components the incidence of rejections is generally weaker for services than for goods inflation.

3.2 Time varying parameter models

Given the evidence of instability highlighted by the tests in the previous section, we now relax the assumption of constant parameters and specify a time varying coefficient model:

$$\pi_t^j = \alpha_t + \beta_t \pi_{t-4}^j + \gamma_t y_{t-h}^i + \eta_t \tag{2}$$

10

Parameter estimates will produce a path for the coefficients, therefore allowing us to gauge the direction of the change signalled by the break tests. Given the large number of models to be estimated (504 for each economic area) it is clear that the use of computationally intensive (bayesian) methods, customarily used to estimated time varying models, is not a viable empirical strategy. We therefore turn to a non-parametric estimator that is computationally much less cumbersome.

The nonparametric approach has long been used in econometrics in the context of deterministic structural change. It has been recently extended to the case of stochastic time variation by Giraitis et al. (2013) and Giraitis et al. (2014). The idea of this estimator is that, in the presence of structural change, older data should be discounted in favour of more recent information. Collecting the right hand variables of equation 2 in the vector x_t, the dependent variable in y_t and the time varying parameters in the vector γ_t, the estimator has the form:

$$\gamma_t = \left[\sum_{j=1}^{T} \omega_{j,t} x_j x_j'\right]^{-1} \left[\sum_{j=1}^{T} \omega_{j,t} x_j y_j\right]$$

where the sample moments are discounted by the function $\omega_{j,t}$:

$$\omega_{j,t} = cK\left(\frac{t-j}{H}\right) \tag{3}$$

where

- c is an integration constant

- $K\left(\frac{t-j}{H}\right)$ is the kernel function determining the weight of each observation j in the estimation at time t. This weight will depend on the distance to t normalized by the bandwidth H.

We use an Exponentially weighted moving average (EWMA) kernel:

$$\omega_{j,t} = \rho^{t-j}/\sum_{k=1}^{t-1} \rho^k, j = 1, 2, ..., t-1 \tag{4}$$

With this kernel only past observations are used to compute the sample moments and their importance is progressively discounted based on their distance from the current period t and on the tuning parameter ρ. To determine the value of ρ and the lag h for the output gap in equation (2) a criterion needs to be chosen. We chose ρ and h so as to minimize the one step head forecast error variance in a pseudo out of sample exercise that runs from 2008Q1 to 2014Q2.

The estimated evolution of the output gap coefficient for the euro area is shown in Figure 4. The blue line shows the median estimate obtained across the various estimated models where the value of ρ and the lag h is optimized based on the criterion described above (i.e., this is the median estimate across the 126 models that can be obtained combining the 14 gap measures with the 9 inflation measures). We also show, for completeness, (red line) the median estimate obtained across the whole model space (i.e., letting also h vary between 1 and 4 and ρ between 0.5 and 0.9, instead of fixing them at their optimal value). In both cases, the median output gap coefficient is quite flat at around 0.25/0.3 until the end of 2012, and then increases sharply from the last quarter of 2012 onwards, up to 0.45/0.5 at the end of the available sample. This result provides a plausible explanation for why the fall in inflation was understated by models estimated on data up to 2012, as the recent fast deceleration of consumer prices is at variance with the historical correlation between the output gap and price dynamics. One might suspect that this result is driven by the fact that in our sample the gaps obtained on the basis of statistical filters (i) behave quite similarly and imply small negative gaps at the end of the sample (ii) are over-represented compared to the estimates provided by Institutions (EC and OECD) and according to which the output gap was still largely negative in 2013/2014. This can be checked by limiting the analysis to the gaps obtained by the EC and the OECD. It turns out that the results are quite robust since, as shown in Figure 5, the output gap coefficient rises at the end of the sample also in the regressions that consider only gaps provided by the EC and the OECD.[9]

One caveat that must be stressed regarding these findings is that the uncertainty surrounding our estimates is large: Figure 6 shows, together with the median, also the first and the third quartile of the output gap coefficients. While there is an ample dispersion in the estimated slopes, a clear tendency to increase is apparent for all quartiles.

We further check the robustness of these findings across three dimensions.

First, we augment equation (2) with a forward looking inflation measure, i.e., expected inflation 6 quarters ahead, as surveyed by Consensus Economics. This modification makes the simple forecasting equation closer in spirit to structural inflation models, where a forward looking component contributes to driving current inflation dynamics. The results, shown in Figure 7, are in line with the baseline case, as the gap coefficient rises markedly at the end of the sample.

As a second exercise we control for possible nonlinear effects that could arise from the fact that the sample includes a large negative shock to the output gap. To take this into account we augment the baseline regression with the output gap to the power 3, that is we estimate the

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[9]In this picture we report only the median estimates across those obtained setting ρ and h at their optimal values.

modified equation:

$$\pi_t^j = \alpha_t + \beta_t \pi_{t-4}^j + \gamma_t y_{t-h}^i + \delta_t (y_{t-h}^i)^3 + \eta_t \qquad (5)$$

In this setup the response of inflation to the output gap depends on the level of the gap itself:

$$\frac{\partial \pi_t^j}{\partial y_t} = \gamma + \delta y_t^2 \qquad (6)$$

The inclusion of this term also leaves the main result unaffected as shown in Figure 8.

As a third check we add to the baseline regression the year on year rate of growth of the nominal effective exchange rate. Since the recent disinflation has also coincided with a strong currency appreciation, omitting this variable from the model could result in biased estimates. Although the estimates are more volatile, the tendency to increase towards the end of the sample is still visible, see Figure 9.

To assess which component of inflation is driving these results, we inspect the estimated gap coefficients for all the nine inflation measures separately, as shown in Figure 10. Together with the time varying coefficients we also plot (green flat line) the estimates that one would obtain from a fixed coefficient model using data up to 2012. Two comments are in order. First, focusing on core inflation, it can be noticed that the responsiveness to the cycle of this component has indeed increased since 2013 above the value estimated up to 2012 (see Core and Core-X panels of Figure 10). Second, this shift in the cyclical sensitivity of core inflation is entirely driven by goods prices (goods and goods-X panels of Figure 10) while the behaviour of Services prices has not changed significantly in recent quarters (Services and Services-X panels). This finding is consistent with the relatively higher stickiness of Services prices as stressed by a number of studies based on micro data (Alvarez et al., 2006), also due to the fact that Services prices include a number of tariffs set by local or central authorities and that therefore do not respond to the cycle.

Finally, we repeat the analysis at the country level and estimate the same model on data for Italy, Spain, Germany and France. The results, shown in Figure 11, indicate that:

- The increase in the cyclicality of inflation is observed in all the large countries, excluding Germany.
- The break is strong and sudden in Italy, more gradual in France.
- In Spain the break occurs much earlier, consistently with the fact that in this country the fall in consumer price inflation started with the global financial crisis and the burst of the housing bubble.

4 Interpretation of the evidence

We look for an interpretation of the evidence in the previous section within the theoretical model developed by Sbordone (2007). Model's details are provided in the Appendix.

The Phillips curve implied by this model is the following:

$$\pi_t = \beta \mathbb{E}_t \pi_{t+1} + \zeta \widehat{s}_t \tag{7}$$

where π_t denotes inflation, β is the discount factor, $\widehat{s}_t$ denotes real unit labor cost (where a hat indicates the log- deviation from the steady state) and ζ is a convolution of deep parameters capturing the sensitivity of price changes to variations in real unit labor cost, that is the determinant of price changes. In this class of models, the choice of using real marginal costs instead of the output gap is innocuous, since there is an approximate log-linear relationship between the two variables. Notice that the above equation is purely forward looking, while the model used in the empirical analysis has a backward looking nature. This is not a major issue, since our aim is not taking equation 7 to the data, but using it to organize a discussion on the possible sources of increased inflation cyclicality.

As shown in the Appendix, the slope coefficient can be defined as:

$$\zeta \equiv \frac{(1 - \alpha\beta)(1 - \alpha)}{\alpha} \frac{1}{1 + \overline{\theta}(N) \left[\overline{\epsilon}_\mu(N) + \overline{s}_y(N) \right]} \tag{8}$$

where β is the discount factor, α is the degree of price stickiness ($\frac{1}{1-\alpha}$ is the average price duration), N is the number of firms, $\overline{s}_y(N)$ denotes the elasticity of the marginal cost to the firm's own output, $\overline{\theta}(N)$ is the steady state elasticity of the firm's own output demand to its relative price and $\overline{\epsilon}_\mu(N)$ is the elasticity of the markup function to output evaluated at steady state.

We can thus disentangle the different theoretical channels that compose the inflation-marginal cost relationship as follows.

1. *Nominal rigidities.* More frequent prices changes (i.e., lower α) induce a steeper Phillips curve.

2. *The elasticity of marginal cost to the firm's own output.* The lower the sensitivity of the marginal costs to the level of output $\overline{s}_y(N)$ is, the steeper the Phillips curve. To understand this mechanism suppose there is a positive shock to real marginal costs $\widehat{s}_t$. This induces an increase in prices and a fall in demand. The latter, in turn, produces a fall in marginal costs (due to decreasing returns to scale) that will partially offset the initial shock and, therefore, reduce the need to adjust prices. It follows that a lower elasticity

14

of marginal costs to output requires a relatively larger price adjustment.

3. *The steady state elasticity of the firm's own output demand to its relative price.* A lower steady state elasticity of demand $\bar{\theta}(N)$ implies a steeper Phillips curve. The mechanism is akin to the one described in the previous point. For a lower steady state elasticity of demand the fall in demand induced by the initial adjustment to a shock to $\hat{s}_t$ is milder, hence requiring a relatively larger price adjustment.

4. *The elasticity of the markup function evaluated at steady state.* When the elasticity of substitution between differentiated goods is decreasing in the relative quantity consumed of the variety, firms face a price elasticity of demand that is increasing in their good's relative price. This makes the desired markup increasing on firm's relative market share (decreasing in firms' relative price). If the elasticity of the markup function evaluated at steady state $\bar{\epsilon}_\mu(N)$ decreases, the Phillips curve steepens. Indeed, when the elasticity of demand is increasing in the relative price, firms are reluctant to change their price as they would face a more elastic demand curve than firms whose relative price declines as a result of price fixity.

This model therefore suggests two possible explanations for an increase in ζ. One explanation is lower nominal rigidities, i.e., a higher frequency of price adjustment (smaller α). Empirical evidence, available only for Italy, shows that in the period between 2006 and 2012 the average duration of consumer prices has indeed declined to five months, from eight months between 1996 and 2001, indicating that increased sensitivity of prices to cyclical conditions might be partly be accounted for by lower nominal rigidities, see Fabiani and Porqueddu (2013). Structural reforms in stressed countries could indeed have played a role in this direction, by increasing flexibility in wage and price settings. Ongoing research at the Eurosystem level through the Wage Dynamics Network should provide better data and more evidence on this issue.

The second explanation rests on the three remaining channels, known in the literature as *strategic complementarities*. As shown in the Appendix in all these channels the number of firms N plays a crucial role. When the number of firms decreases, the steady state elasticity of demand $\bar{\theta}$ goes down (in line with the general intuition that the larger the number of goods that are traded in the market, the more likely it is that demand declines in response to a small increase in prices); this tends to increase inflation cyclicality. By contrast, the elasticity of the mark-up function $\bar{\epsilon}_\mu$ and the elasticity of the marginal cost to firm's own output $\bar{s}_y$ go up and

this tends to result in lower inflation cyclicality. To sum up:

$$\zeta \equiv \frac{(1 - \alpha\beta)(1 - \alpha)}{\alpha} \frac{1}{1 + \underset{+}{\overline{\theta}(N)} \left[\underset{-}{\overline{\epsilon}_\mu(N)} + \underset{-}{\overline{s}_y(N)} \right]} \tag{9}$$

The combination of these effects shapes the relationship between the slope of the Phillips curve and the number of firms, as shown in Figure 12. If the first effect dominates the other two, inflation cyclicality will increase as N falls.

A formal test of the hypothesis linking consumer prices and the number of firms in the economy is difficult because of poor data quality regarding business demography in the euro area. Keeping these caveats in mind, some preliminary analysis on available data indicates that, in the case of Italy and Spain, the sovereign debt crisis induced a significant reduction in the number of firms. This suggests that these two countries have been recently moving left-wise along the right portion of the curve shown in Figure 12, where a decrease in the number of firms increases inflation cyclicality. While more investigation is needed in this direction, the fact that strategic complementarities played a role in the steepening of the Phillips curve cannot be ruled out.

5 Conclusions

The bout of disinflation between 2013 and 2014 has been broad based across the euro area and more intense in those countries that have been hit the hardest by the sovereign debt crisis. Despite the persistent economic weakness, professional forecasters largely failed to predict the decline in inflation: those surveyed by Consensus Forecast systematically over predicted average inflation for 2013. This suggests that a structural break in the pattern of inflation cyclicality might have occurred. In this paper we explore, from an empirical point of view, whether this is the case and then offer a plausible interpretation through a theoretical model. Our empirical analysis uncovers a significant increase in the sensitivity of inflation to the business cycle since the second half of 2013, mainly due to developments in the tradable sector. The result applies to the euro area as a whole, as well as to Italy, France and Spain.

A steepening of the Phillips curve might have resulted from lower nominal rigidities (a decrease in the average duration of prices), a hypothesis supported for Italy by the results in Fabiani and Porqueddu (2013). Structural reforms in stressed countries could indeed have contributed to increasing flexibility in wage and price settings. Alternatively, we relate the increase in the short run response of inflation to real activity to the fact that the two consecutive recessions experienced by the euro area since 2008 resulted in the closure of many businesses,

thus inducing a decrease of strategic complementarities in price setting and a higher elasticity of consumer prices to the output gap.

Appendix: The theoretical model

We consider the theoretical framework developed by Sbordone (2007), that extends the Kimball's model in an environment where the number of firms is variable. Households' utility is defined over an aggregate C_t of differentiated goods $c_t(i)$, implicitly defined as

$$\int_\Omega \psi\left(\frac{c_t(i)}{C_t}\right) di = 1 \tag{10}$$

where $\psi\left(\cdot\right)$ is an increasing strictly concave function and Ω is the set of all potential goods produced. Note that the standard CES preferences are nested within this specification and the Kimball aggregator reduces to the Dixit-Stiglitz when $\psi\left(\frac{c_t(i)}{C_t}\right) = \left(\frac{c_t(i)}{C_t}\right)^{\frac{\theta-1}{\theta}}$ for some $\theta > 1$.

Each firm produces a differentiated good. We assume that the set of firms is $[0, N]$ and, thus, $c_t(i) = 0 \; \forall i > N$. The household must decide how to allocate its consumption expenditures among the different goods: $\min_{\{c_t(i)\}} \int_0^N p_t(i) c_t(i) \, di$ s.t. $\int_0^N \psi\left(\frac{c_t(i)}{C_t}\right) di = 1$. From the FOC to this problem one gets the demand for each good i:

$$c_t(i) = C_t \psi'^{-1}\left(p_t(i)\Lambda_t C_t\right) \; \forall i \in [0, N]$$

where Λ_t is the Lagrangian multiplier for constraint (10), that is implicitly defined by $\int_0^N \psi\left(\psi'^{-1}\left(p_t(i)\Lambda_t C_t\right)\right) di = 1$

The aggregate price index is the cost of a unit of the composite good: $P_t \equiv \frac{1}{C_t} \int_0^N p_t(i) c_t(i) di$.

We assume that firm i produces with the following technology:

$$y_t(i) = h_t(i)^{1-a} - \Phi \tag{11}$$

where Φ is a fixed cost. Accordingly firm's i real marginal cost $s_t(i)$ is:

$$s_t\left(y_t(i); \Gamma_t)\right) = \frac{1}{1-a}\frac{W_t}{P_t}\left(y_t(i) + \Phi\right)^{\frac{a}{1-a}} \tag{12}$$

where Γ_t indicates aggregate variables that enter into the determination of firms' marginal costs, W_t, is nominal wage and P_t is the aggregate price.

Following the formalism proposed in Calvo (1983), each fim may reset its price only with probability $(1-\alpha)$ in any given period, independently of the time elapsed since the last adjustment ($\frac{1}{1-\alpha}$ is the expected average duration of prices). A firm reoptimizing in period t will choose the price $p_t(i)$ by maximizing the expected string of profits over the life of the set price.

$$\mathbb{E}_t\left\{\sum_{j=0}^\infty \alpha^j Q_{t,t+j}\left[p_t(i)Y_{t+j}\psi'^{-1}\left(\frac{p_t(i)}{\widetilde{P}_{t+j}}\right) - C\left(Y_{t+j}\psi'^{-1}\left(\frac{p_t(i)}{\widetilde{P}_{t+j}}\right); \Gamma_{t+j}\right)\right]\right\} \tag{13}$$

where $C(\cdot)$ is the firm's cost function. Combining the the first order condition associated with the problem above with the the aggregate price dynamics yields the following Philips curve:

$$\pi_t = \beta\mathbb{E}_t\pi_{t+1} + \zeta\widehat{s}_t \tag{14}$$

where π_t denotes inflation, β is the discount factor, $\widehat{s}_t$ denotes real unit labor cost (where a hat indicates the log- deviation from the steady state) and ζ is a convolution of deep parameters. Our goal is to evaluate how the number of producing firms N affects the slope coefficient ζ. Let define $x = \psi'^{-1}\left(\frac{1}{N}\right)$ the relative share in the symmetric steady state, i.e. a steady state with symmetric prices $(p_t(i) = p_t \forall i)$. Then we can define the steady state elasticity of demand:

$$\overline{\theta} = -\frac{\psi'(x)}{x\psi'(x)} \tag{15}$$

and the elasticity of the mark-up function evaluated at steady state:

$$\overline{\epsilon}_\mu = \frac{x\mu'(x)}{\mu(x)}$$

The slope coefficient can be defined as:

$$\zeta \equiv \frac{(1-\alpha\beta)(1-\alpha)}{\alpha}\frac{1}{1+\overline{\theta}(N)\left[\overline{\epsilon}_\mu(N)+\overline{s}_y(N)\right]} \tag{16}$$

where $\overline{s}_y(N) = \frac{a}{1-a}\left[\frac{xY}{xY+\Phi}\right]$ denotes the elasticity of the marginal cost to firm's own output (Y is steady state aggregate output)[10]. We now turn to examine how the number of firms N affects these channels. To this aim we need to choose a functional form for $\psi(x)$. As in Sbordone (2007) we assume the one proposed by Dotsey and King (2005).

$$\psi(x) = \frac{1}{(1+\eta)\gamma}[(1+\eta)x-\eta]^\gamma - \frac{1}{(1+\eta)\gamma}(-\eta)^\gamma \tag{17}$$

In this case the steady state relative share x is:

$$x \equiv \psi^{-1}(\frac{1}{N}) = \frac{1}{1+\eta}\left\{\left(\frac{(1+\eta)\gamma}{N}+(-\eta)^\gamma\right)^{\frac{1}{\gamma}}+\eta\right\} \tag{18}$$

The latter is clearly decreasing in N. The steady state mark-up $\overline{\mu}$ is:

$$\overline{\mu} = \frac{\eta-(1+\eta)\psi^{-1}(\frac{1}{N})}{\eta-\gamma(1+\eta)\psi^{-1}(\frac{1}{N})}$$

In order to see the dependence of the slope on N, we need to study how $\overline{\theta}(N)$, $\overline{\epsilon}_\mu(N)$ and

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[10]Under a fairly standard log-utility: $u(C,h) = \log C - \frac{1}{1+v}h^{1+v}$, one gets $xY + \Phi = \left[\frac{1-a}{\overline{\mu}xYN^{1+v}}\right]^{\frac{1-a}{v+a}}$

$\bar{s}_y(N)$ vary with N. The steady state elasticity $\bar{\theta}$ is:

$$\bar{\theta} = \frac{\eta - (1+\eta)\,\psi^{-1}(\frac{1}{N})}{(\gamma-1)\,(1+\eta)\,\psi^{-1}(\frac{1}{N})} \tag{19}$$

which is decreasing in the steady state relative share x and, thus, increasing in N. This is in line with the general intuition that more goods are traded in a market more likely it is for the demand to decrease more in response to a small increase in prices. The elasticity of mark-up $\bar{\epsilon}_\mu$, that determines how much the steady state mark-up varies for small variation in N, is

$$\bar{\epsilon}_\mu = \frac{\eta\,(\gamma-1)\,(1+\eta)\,\psi^{-1}(\frac{1}{N})}{\left[\eta - (1+\eta)\,\psi^{-1}(\frac{1}{N})\right]\left[\eta - \gamma\,(1+\eta)\,\psi^{-1}(\frac{1}{N})\right]} \tag{20}$$

It can be demonstrated that this elasticity is a decreasing function of N.[11] Finally the elasticity of the marginal cost to firm's own output is:

$$\bar{s}_y(N) = \frac{a}{1-a}\left[\frac{xY}{xY+\Phi}\right] \tag{21}$$

It can be demonstrated that, assuming a fairly standard log-utility $u(C,h) = \log C - \frac{1}{1+\nu}h^{1+\nu}$, the steady state aggregate output is the solution to $xY + \Phi = \left[\frac{1-a}{\bar{\mu}xYN^{1+v}}\right]^{\frac{1-a}{v+a}}$ and $\bar{s}_y(N)$ is decreasing in N. To sum up:

$$\zeta \equiv \frac{(1-\alpha\beta)\,(1-\alpha)}{\alpha}\,\frac{1}{1+\underset{+}{\bar{\theta}(N)}\left[\underset{-}{\bar{\epsilon}_\mu(N)} + \underset{-}{\bar{s}_y(N)}\right]} \tag{22}$$

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[11]Indeed $\frac{\partial log\mu}{\partial \log N} = -\bar{\epsilon}_\mu \frac{\partial \log x}{\partial \log N}$ and $log\mu$ is a convex function of $logx$. Indeed because $\mu(x)$ is an increasing function of x, it is not possible for $log\mu$ to be a concave function of $logx$ as this would require $log\mu$ to be negative for positive and small enough x. If $log\mu$ must be convex at least for small values of x, it is convenient to assume that it is globally convex function of $logx$.

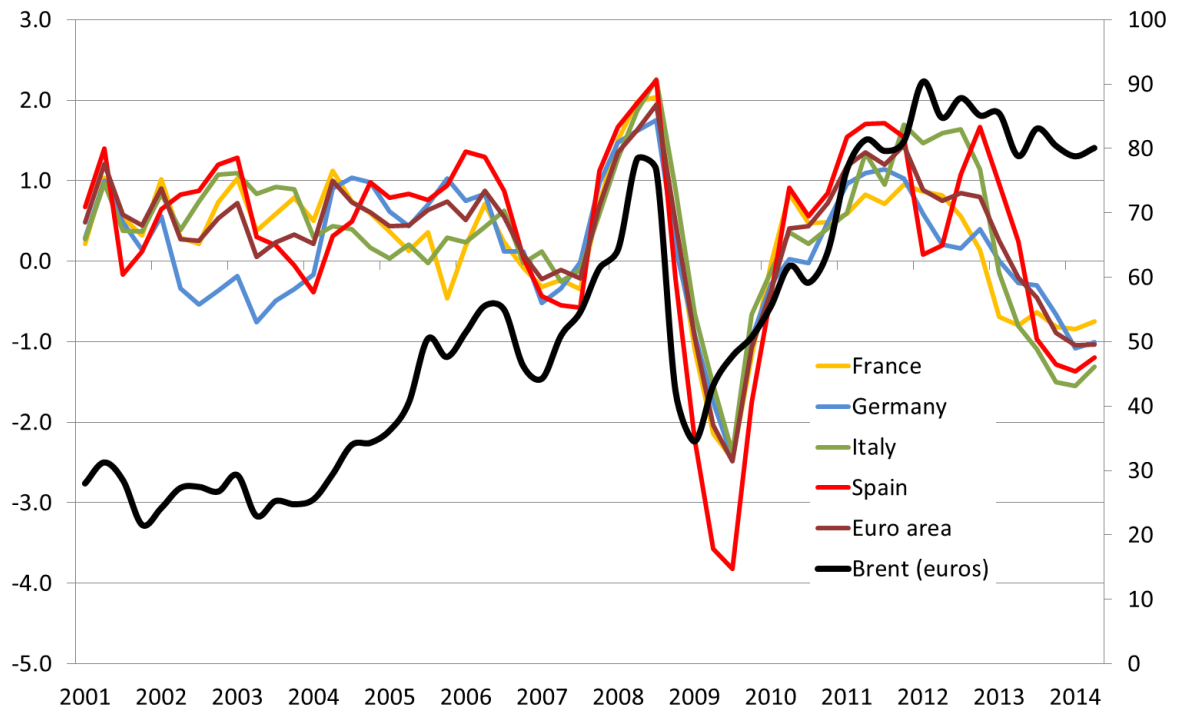Figure 1: Consensus Economics: 4 quarters ahead forecast errors

Table 1: Inflation measures and Activity gaps

	Overall
	Overall-X (net of fiscal impact)
	Core
	Core-X (net of fiscal impact)
Inflation measures	Compensations per employee
	Goods
	Goods-X (net of fiscal impact)
	Services
	Services-X (net of fiscal impact)
	GDP - Hp filter
	Hours - Hp filter
	Unemployment rate - Hp filter
	GDP - Hp filter (real time)
	Hours - Hp filter (real time)
	Unemployment rate - Hp filter (real time)
Gap measures	GDP - Band pass
	Hours - Band pass
	Unemployment rate - Band pass
	GDP - Band pass (real time)
	Hours - Band pass (real time)
	Unemployment rate - Band pass (real time)
	European Commission output gap
	OECD output gap

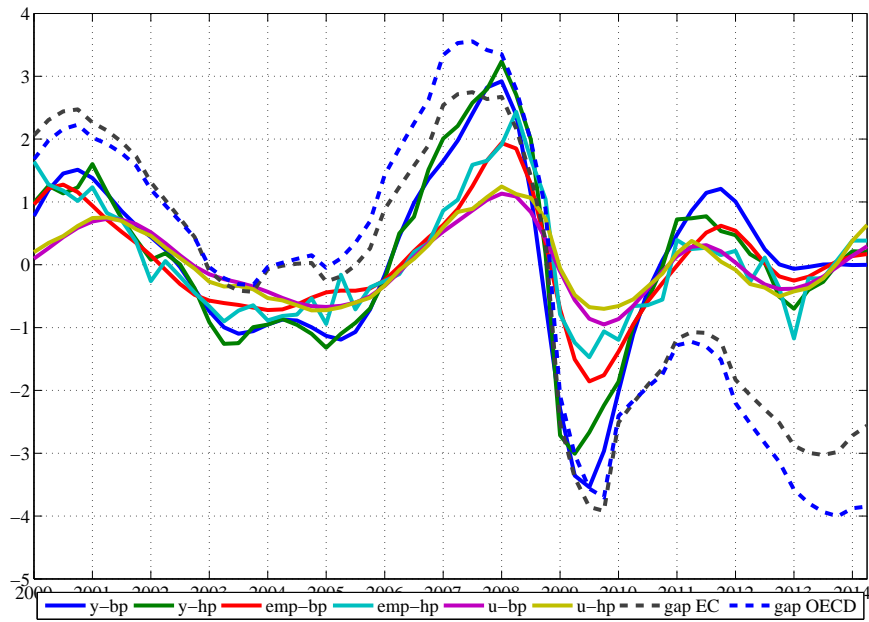Figure 2: Inflation rates - euro area
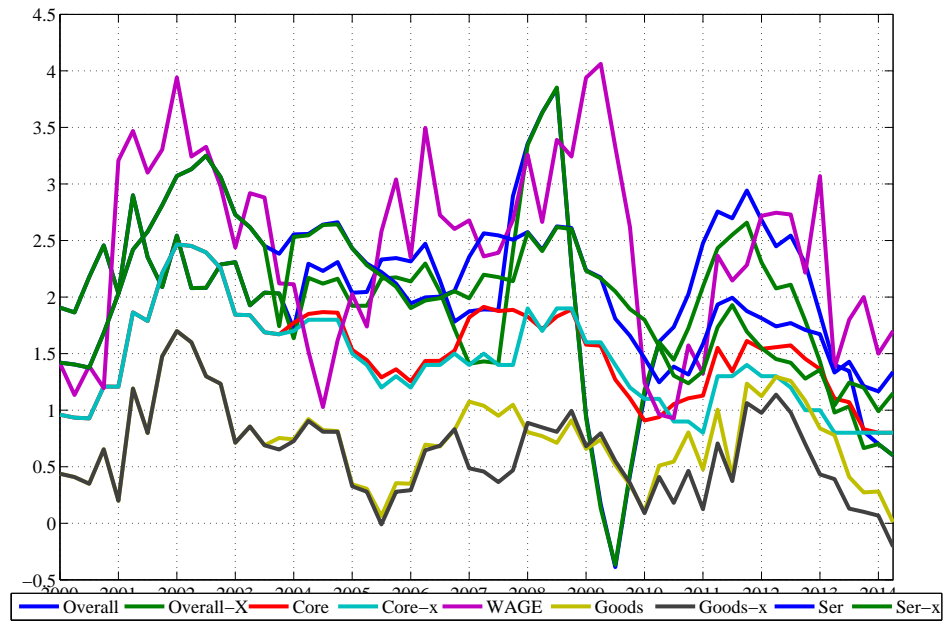
Figure 3: Activity gaps - euro area

23

Table 2: End of sample stability tests: percentage of rejections at the 5% confidence level

Test	Π	Overall	Overall-X	Core	Core-x	WAGE	Goods	Goods-x	Ser	Ser-x
					euro area					
L	25%	0.14	0.16	0.09	0.59	0.07	0.11	0.07	0.57	0.79
L	10%	0.18	0.70	0.64	0.79	0.07	0.00	0.57	0.77	0.77
exp-L	25%	1.00	1.00	1.00	0.82	0.11	1.00	1.00	0.89	0.75
exp-L	10%	1.00	1.00	1.00	0.82	0.09	1.00	1.00	0.89	0.68
sup-L	25%	1.00	1.00	1.00	0.80	0.11	1.00	1.00	0.91	0.71
sup-L	10%	1.00	1.00	1.00	0.80	0.09	1.00	1.00	0.91	0.64
					Germany					
L	25%	0.09	1.00	0.00	0.38	0.64	1.00	1.00	0.00	0.00
L	10%	0.00	0.95	0.00	0.02	0.14	0.00	0.91	0.00	0.00
exp-L	25%	0.00	1.00	0.00	0.21	0.25	0.61	1.00	0.00	0.00
exp-L	10%	0.00	0.09	0.00	0.00	0.05	0.00	0.20	0.00	0.00
sup-L	25%	0.04	1.00	0.00	0.05	0.21	0.23	1.00	0.00	0.00
sup-L	10%	0.04	0.09	0.00	0.00	0.07	0.00	0.27	0.00	0.00
					France					
L	25%	0.13	0.05	1.00	0.93	0.82	1.00	1.00	0.68	0.95
L	10%	0.05	1.00	1.00	1.00	0.09	0.86	0.86	0.79	0.79
exp-L	25%	1.00	1.00	1.00	1.00	0.21	1.00	1.00	0.88	1.00
exp-L	10%	1.00	1.00	1.00	1.00	0.04	1.00	1.00	0.91	0.70
sup-L	25%	1.00	1.00	1.00	1.00	0.04	1.00	1.00	0.84	1.00
sup-L	10%	1.00	1.00	1.00	1.00	0.00	1.00	1.00	0.86	0.70
					Italy					
L	25%	0.70	0.82	1.00	0.86	1.00	1.00	0.98	0.68	0.77
L	10%	0.70	0.98	1.00	1.00	0.86	1.00	1.00	0.89	0.84
exp-L	25%	1.00	1.00	1.00	1.00	0.93	1.00	1.00	0.96	1.00
exp-L	10%	1.00	1.00	1.00	1.00	0.86	1.00	1.00	0.98	1.00
sup-L	25%	1.00	1.00	1.00	1.00	0.93	1.00	1.00	0.98	1.00
sup-L	10%	1.00	1.00	1.00	1.00	0.86	1.00	1.00	0.98	1.00
					Spain					
L	25%	0.63	0.98	0.13	0.34	1.00	0.20	0.02	0.13	0.55
L	10%	0.84	0.93	0.66	0.86	1.00	0.98	0.09	0.23	0.93
exp-L	25%	1.00	0.96	1.00	0.89	1.00	1.00	0.41	1.00	0.91
exp-L	10%	1.00	0.93	1.00	0.91	1.00	1.00	0.70	1.00	0.91
sup-L	25%	1.00	0.96	1.00	0.89	1.00	1.00	0.66	1.00	0.91
sup-L	10%	1.00	0.93	1.00	0.91	1.00	1.00	0.73	1.00	0.93

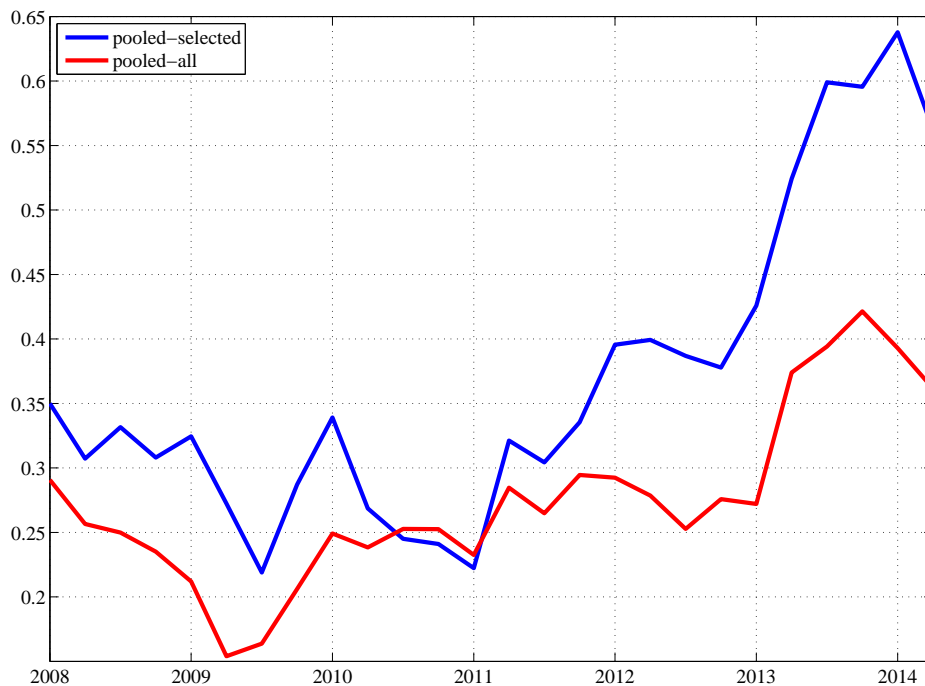Figure 4: Gap coefficients (median estimate)

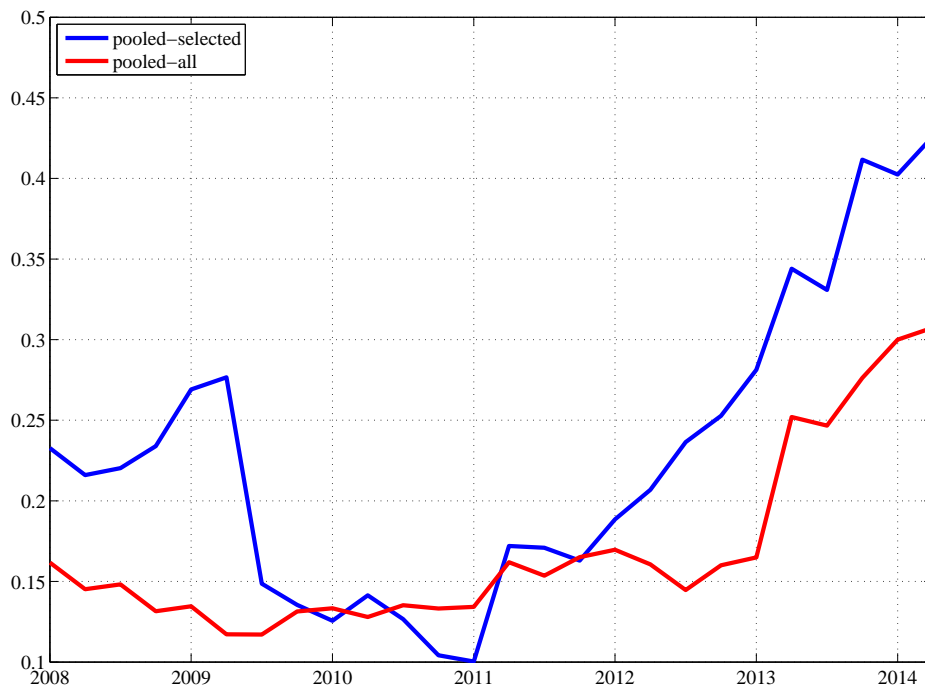Figure 5: Gap coefficients (median estimate): only EC and OECD gaps

Figure 6: Gap coefficients: 25th, 50th and $75^{th} percentiles$
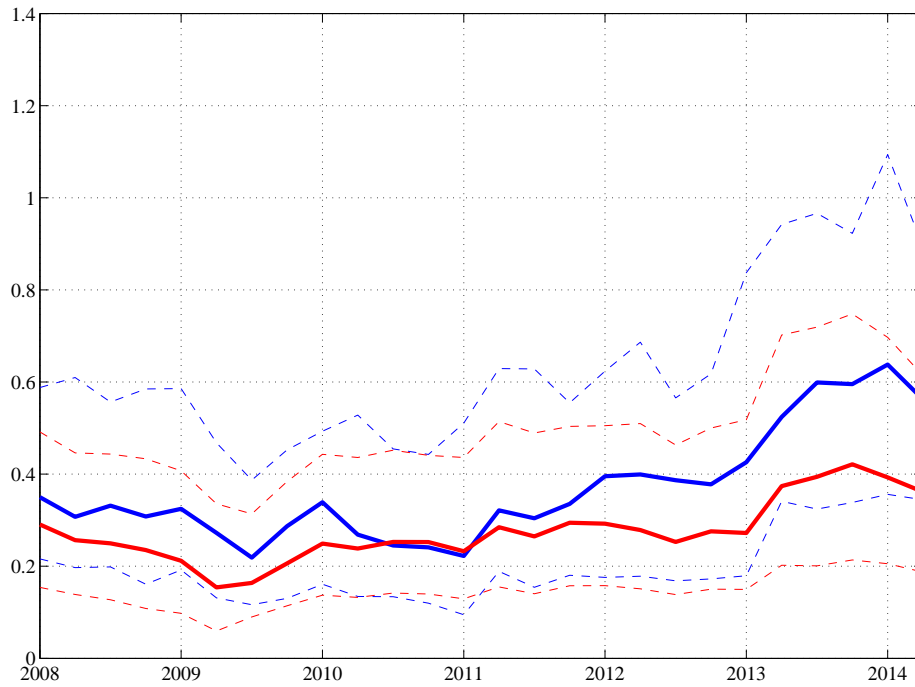

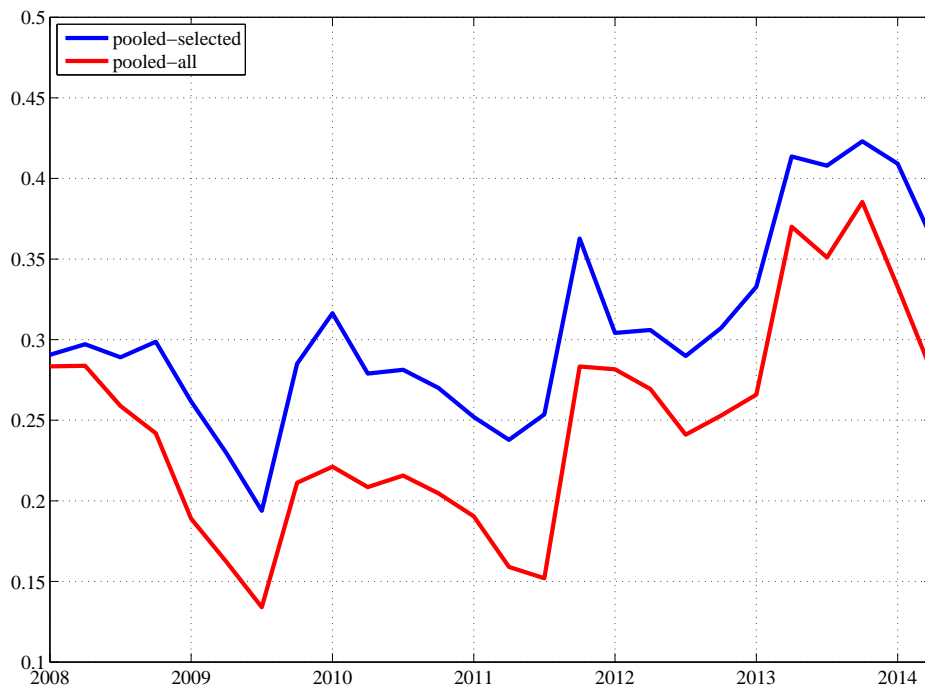Figure 7: Gap coefficients (median estimate): controlling for $E_t\pi_{t+6}$

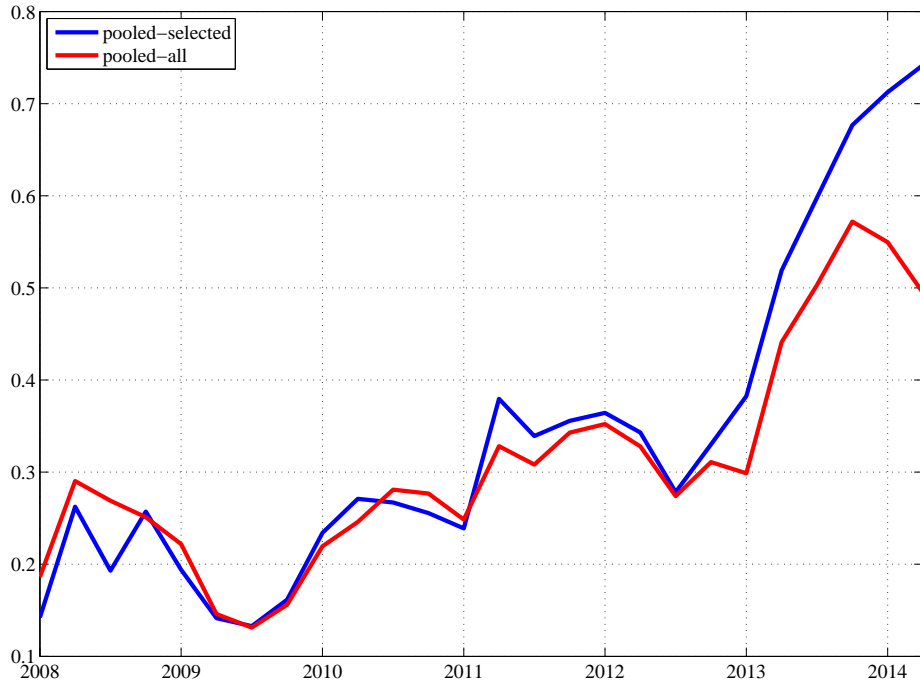Figure 8: Gap coefficients (median estimate): controlling for gap_t^3

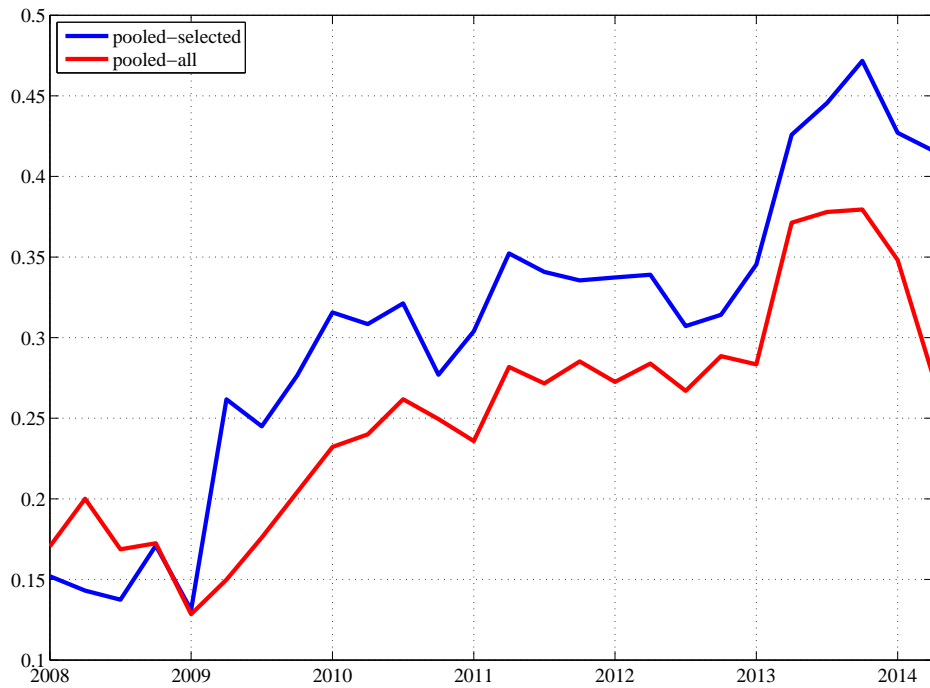Figure 9: Gap coefficients (median estimate): controlling for the exchange rate

Figure 10: Gap coefficients (median estimate): different inflation measures
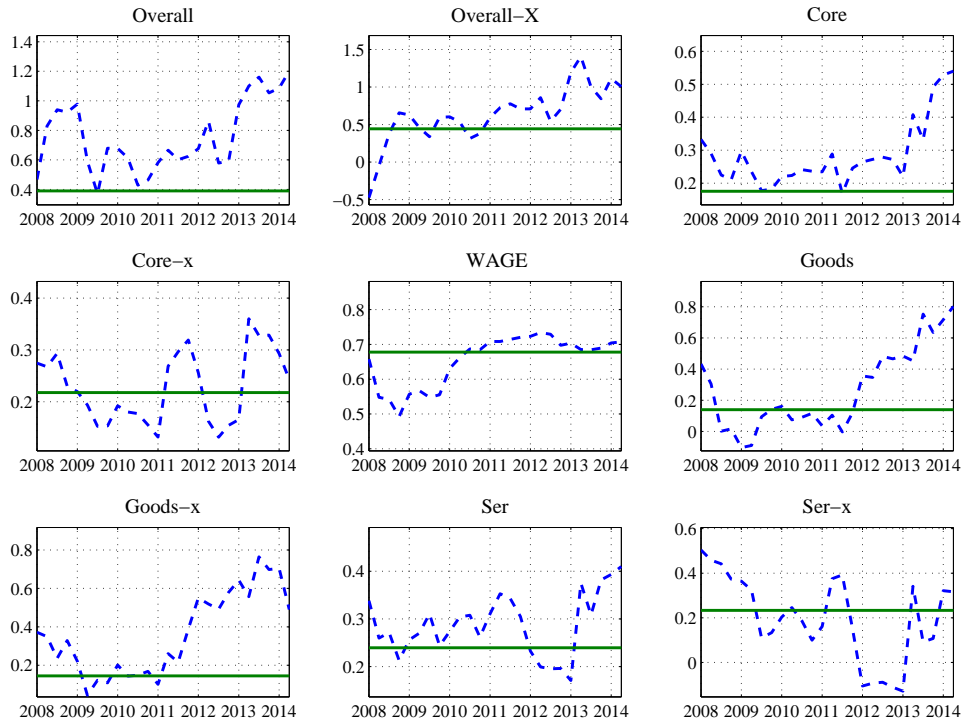

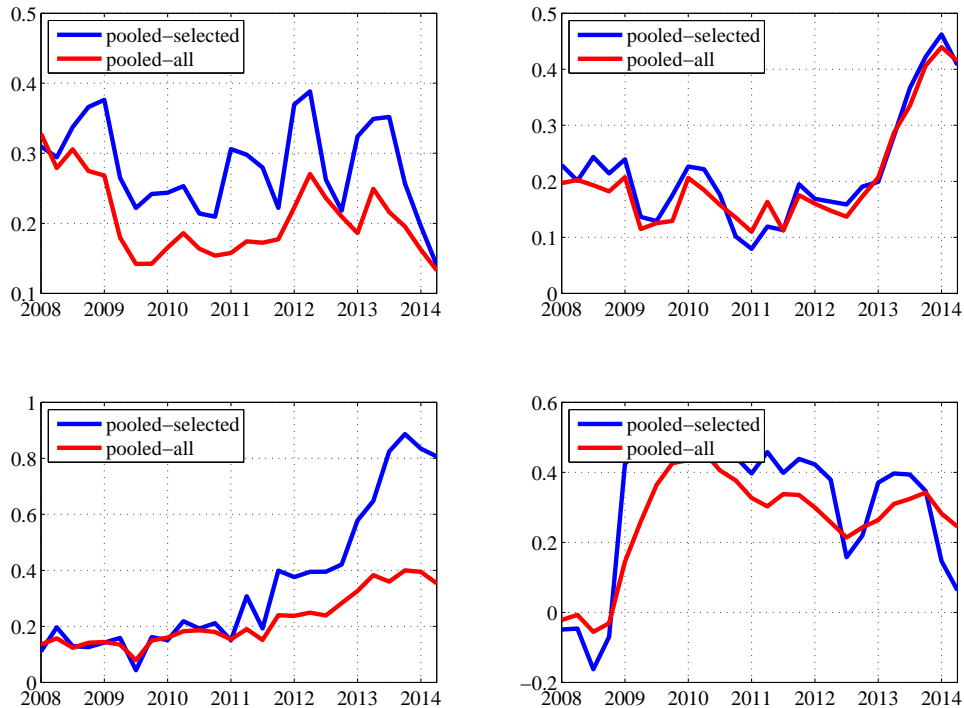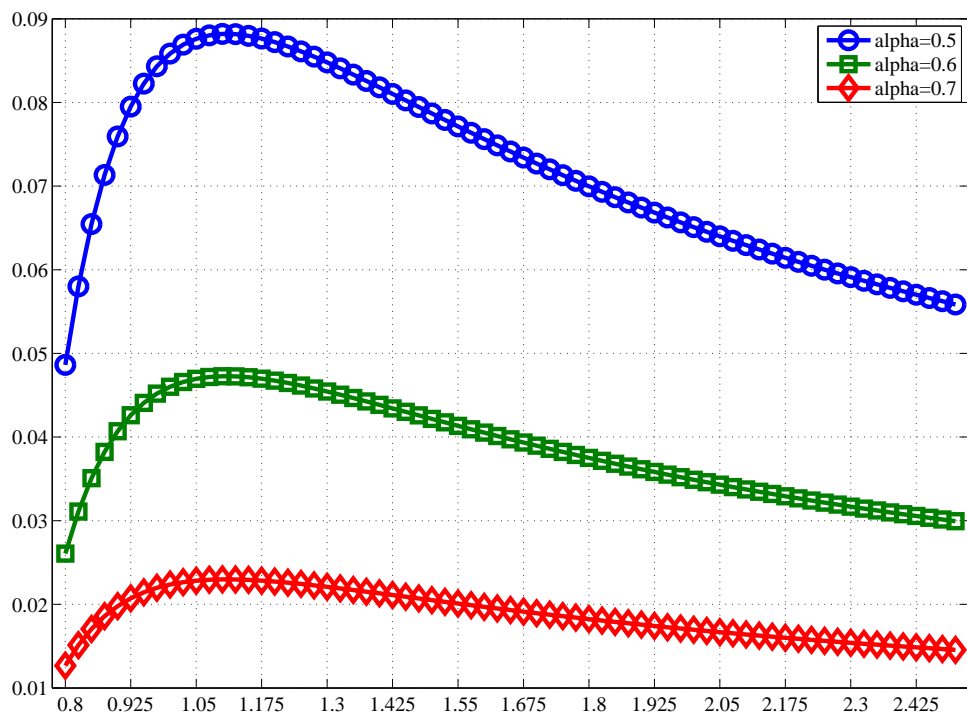Figure 11: Gap coefficients (median estimate): different countries

Figure 12: Slope of the Phillips curve and number of firms

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