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# The Great Recession and the Two Dimensions of European Central Bank Credibility*

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**Abstract**

A puzzle from the Great Recession is an apparent mismatch between a fall in the persistence of European inflation rates, and the increased variability of expert forecasts of inflation. We explain this puzzle and show how country specific beliefs about inflation are still quite close to the European Central Bank target of 2% (what we call official target credibility) but the degree of anchoring to this target has gone down, implying an erosion of what we call anchoring credibility. A decline in anchoring credibility can explain increased forecast variance independently of any changes in inflation persistence, contrary to standard time series models.

Keywords: central bank credibility, excess volatility, euro, inferential expectations, inflation.

JEL Classification: C51, D84, E31, E52.

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# 1 Introduction

Following the tumultuous events of 2008, inflation in the Euro zone has become less persistent. In standard models of inflation, a reduction in persistence should coincide with a reduction in the forecast error, but the reverse has been observed. This paper is about how changes in central bank credibility can alter the forecast error variance of inflation, independently of changes to its time series properties. We demonstrate that this has in fact occurred in Europe in recent years.

For decades economists have studied the time series properties of inflation (Fuhrer, 2011). Early work assumed inflation was inertial, and the primary objective was to accurately measure the degree of inertia by estimating inflation equations with several higher-order lags (Gordon, 1982). Only later, with the advent of rational expectations (RE), did researchers focus on the sources of this persistence. The first theoretical RE models, which predicted very low or no persistence, failed to find support in the inflation data. In order to add persistence to the RE models, it was necessary to introduce structural frictions, such as nominal price contracting (Fischer, 1977, Taylor, 1980, Rotemberg, 1982, and Calvo, 1983). Models with frictions have been extensively reworked (e.g. Dotsey et al., 1999; Mankiw and Reis, 2002) and still form the backbone of current monetary economics, including dynamic stochastic general equilibrium (DSGE) models.

Empirically, there have been numerous attempts to explain changes in reduced-form inflation persistence with reference to changes in the underlying determinants of inflation. The most common explanation is a systematic change in monetary policy, since it is well known that the coefficients of reduced form expressions are subject to instability from changing policies or beliefs (Friedman, 1968, and Lucas, 1972). Common methodological approaches include tests for structural breaks in single-equation models of inflation. As an alternative approach, Cogley et al. (2010) employ a time-varying VAR to estimate the trend component of inflation.

In an important contribution, Benati (2008) notes that international evidence points to a marked reduction in inflation persistence for countries that have adopted an official inflation targeting regime.[1] In the recent past, European inflation appears to have had a fall in its persistence, measured by the coefficient on the lagged dependent variable in a rolling 7-year annual regression of inflation on its lag.

*(Insert Figure 1 about here)*

Benati (2008) argues that the fall in inflation persistence for inflation targeting countries is the consequence of well anchored expectations. In the limit, the central bank's commitment to the official inflation target is so strong that any deviation of actual inflation from the target is assumed to be temporary and unplanned and therefore unforecastable. Williams (2006) concurs, arguing that inflation in recent years may be best described as the sum of a constant (the inflation target) and an i.i.d. error term.

[1] See also Darvas and Varga (2013) for central and Eastern European countries, Gamber et al. (2013) for the US, and Siklos (2013a) for Asia-Pacific countries.

Following this line of reasoning, one might conclude from Figure 1 that inflationary expectations in Europe have become more anchored in the recent past.

However, the results of Gerlach et al. (2011) pose a problem for this interpretation. They find that professional forecasters since the Great Recession (GR) are less reliant on past information, which is consistent with reduced aggregate autocorrelation implied by Figure 1. But they also find that experts' long term forecasts of inflation have become more dispersed.

The difficulty here is that in many sensible time series representations of inflation reduced persistence decreases the forecast error at any horizon, as well as the unconditional variance, the latter being an upper bound on the former.[2] So the coexistence of these two stylized facts (seemingly reduced variance based on declines in the persistence of inflation and higher variance based on expert forecasting) is deeply puzzling.[3] The puzzle—that there is a prediction of a correlation between these trends—would disappear if, in fact, the two are not related.

Of course, one possibility is that the expert forecasts do not affect the actual inflation process at all, in which case the dispersion of their forecasts is irrelevant. We eschew this explanation, which we find implausible, and instead propose that the increased variance of inflation forecasts could be explained by changes in beliefs about central bank credibility.[4] We model beliefs about central bank credibility by regarding the central bank inflation target as a null belief in a hypothesis test (Henckel et al., 2011), which can be modeled using the inferential expectations (IE) framework (Menzies and Zizzo, 2009).

Within this framework we define two types of credibility. A central bank has *official target credibility* if the null belief that agents hold corresponds to the official target value. In the case of the ECB we will treat this value as 2%.[5]

The null belief could, however, take some other value. *Anchoring credibility* refers to the capability of the central bank to anchor inflation to the value corresponding to the null belief, whether or not this is the official target. Table 1 shows the two dimensions of credibility.

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[2]For $y_t = \rho y_{t-1} + e_t$, the $r$-step-ahead forecast error variance is $\sigma^2\left(1+\rho^2+\rho^4+\ldots+\rho^{2(r-1)}\right)$ which is clearly increasing in $\rho$. As $r$ approaches infinity, the expression asymptotes to the unconditional variance $\sigma^2/\left(1-\rho^2\right)$.

[3]We are grateful to Petra Gerlach for encouraging us to demonstrate this with a more general structure (Appendix A).

[4]As Pierre Siklos pointed out to us in private correspondence, there may be other reasons for greater dispersion in inflation forecasts—e.g. forecasters' heightened attentiveness since the GR; less coordination among central bank forecasts; or greater central bank transparency (Siklos, 2013b).

For our empirical model the latter interpretation does not work, as transparency of the European Central Bank (ECB) has not changed since the GR (Siklos, 2013a). The first two interpretations do not necessarily contradict our proposition that a loss in central bank credibility leads to greater dispersion in inflation forecasts. In particular, heightened attentiveness and less coordination among forecasters most likely occur when the inflation rate moves away from its anchor, that is, precisely when central bank credibility is compromised.

[5]The European Central Bank (ECB) 'aims at inflation rates of below, but close to, 2% over the medium term' http://www.ecb.int/mopo/html/index.en.html. In the absence of a meaning for 'close' we regard 2% as the target in our discussion. However, we return to this issue in the conclusion, and show how one of our results is robust to changing 2% to what we will estimate the target to be for Germany (1.65%). Apart from this one result, the ECB's target is estimated econometrically in this paper, rather than assumed, and so the particular value chosen is irrelevant.

Borrowing the usual nomenclature for statistical hypothesis tests, in Table 1 a high $\beta$ means that an overwhelming majority of agents believe the null and a low $\beta$ means that few do.[6] Looking at the columns of Table 1, agents might believe that the ECB is *really* targeting, say, 1%, in spite of its official pronouncements. That would be the '$H_0$ not at official target' case which lacks official target credibility.

*(Insert Table 1 about here)*

We suspect that when most commentators refer to central bank credibility, or well anchored expectations, they mean the top left corner of Table 1 where a high proportion of agents believe the null, which, in turn, coincides with the official target (high $\beta$ and $H_0$ at target). Fuhrer (2011) furnishes a good example:

> "well-anchored inflation expectations require two things from the central bank. First, the central bank must have an inflation goal [target] that is known to the private agents in the economy, and second, the central bank must move its policy rate in a way that systematically pushes the inflation rate toward that goal." (p. 474)

Fuhrer's first requirement corresponds to our *official target credibility*, while his second corresponds to our *anchoring credibility*. Yet he only considers well anchored expectations to refer to a combination of the two when he refers to pushing inflation 'towards *that* goal' (our italics), rather than some other inflation target.[7]

We can agree with Fuhrer (2010) that when both requirements are met - the top left corner of Table 1 - a central bank is credible; we can also say that, when both requirements are not met - the bottom right corner of Table 1 -, a central bank is not credible. As soon as we leave the left diagonal of Table 1, however, it becomes harder to give a uni-dimensional designation 'credible' or 'not credible' to a central bank.

Suppose that the authorities accommodate a very high, even hyper-, rate of inflation which can be accurately predicted (top right). We could describe this as the central bank having anchoring credibility without official target credibility.[8]

For example, if experts during the Weimar Republic could predict inflation accurately, the central bank would be credible on the former definition. The reader might be uncomfortable using the word 'credibility' if the $H_0$ belief is too far from a socially desirable level, as it was in inter-war Germany, but this just strengthens our point that it is conceptually useful to have a multi-dimensional description of credibility rather than a binary choice along one dimension.

Naturally, it might be the case that *in practice* the two dimensions of credibility are highly correlated, so that either is a sufficient statistic for the other. If this proves to be the case, a single measure—a single dimension—will suffice for policymaking and

[6] In standard terminology, $\beta$ is the probabiliy of a type II error (falsely believing the null), and we will show later that the null is indeed false in our model. With respect to the size of 'high' and 'low' in Table 1, we are not trying to define any particular thresholds. Table 1 is for heuristic purposes.

[7] See also Demertzis et al. (2008).

[8] The analytic content of Table 1 can be made even richer by giving content to the alternative beliefs ($H_1$). We do so later.

other practical purposes. However, we will later present evidence that this is not the case in Europe, suggesting the distinction between the two dimensions is as valuable in practice as it is in theory.

The main result of this paper is that the European central bank still possesses a good deal of official targeting credibility, but that the GR has led to a loss of anchoring credibility.[9] Our estimated equations will show that anchoring credibility loss is most severe in Portugal, Ireland, Italy and Greece, which accords with our priors. In our theoretical model, outlined in section 2, we assume that price setters rely on a consensus (average) forecast by inflation experts (Carroll, 2003; Easaw and Golinelli, 2010) which we model with IE. In section 3 we explain how the proportion of agents who believe the target is related to the distribution of heterogeneous test sizes within the IE framework. In section 4 we estimate inflation equations and confirm that the surveys of forecasters in Gerlach et al. (2011) are consistent with an IE model of inflation. Section 5 justifies our approach by showing that the two measures of credibility are relatively uncorrelated, and concludes.

# 2 IE in the Country Specific Phillips Curve

The theoretical building block is an augmented Lucas supply function of the form,

$$\pi_t = \pi_t^e + \gamma_1 \pi_t^{EN} + \gamma_2 e_t, \qquad \gamma_1, \gamma_2 > 0, \tag{1}$$

where $\pi_t$ denotes actual inflation in period $t$, $\pi_t^e$ expected inflation, and $e_t$ real marginal cost. We augment the equation for the mechanical impact on inflation of energy prices $\pi_t^{EN}$, which is important empirically for Europe. We are not interested in the determinants of $e_t$ in general equilibrium. We therefore assume that $e_t$ has a zero mean but we leave unspecified whether it is autocorrelated. It is straightforward to use a mapping of real marginal cost to output or unemployment to rewrite (1) so that $e_t$ becomes an output gap. An example of a derivation of the supply function (1) is provided in Appendix B.

Expected inflation $\pi_t^e$ need not be the mathematical expectation. In particular, it may represent the average of a range of heterogeneous expectations, as in our case below. There is no forward-looking inflation expectation term, as in the New Keynesian Phillips curve, because prices are assumed to be perfectly flexible, providing firms with the opportunity to set prices anew in each period.

The coefficient $\gamma_2$ of real marginal cost may reflect institutional features or the relative weight attached to a noisy signal of $e$ as in a standard signal extraction problem.

In our model expected inflation is assumed, in the spirit of Carroll (2003), to be the consensus belief of experts about the true inflation target of the central bank, $\pi^{cb}$. In particular, we assume that in each Eurozone country a continuum of experts, after studying the economy and the central bank, separately guess $\pi^{cb}$. These within-country guesses are averaged and form a *consensus belief* which substitutes for $\pi^e$ in the first term on the RHS of (1) in each country's Phillips curve.

[9] Using different approaches, Galati et al. (2011) also conclude that inflation expectations in the Euro area have become less firmly anchored during the crisis while Albinowski et al. (2013) argue that trust in the ECB has suffered in this period.

We assume that experts have inferential expectations (IE), that is, they guess $\pi^{cb}$ using a statistical hypothesis test. We further assume that all expert groups in each Eurozone country use the same p-value $P$ of a test statistic as they guess $\pi^{e}$. We consider this a reasonable approximation as a common ECB monetary policy implies similar information sets. That said, we assume that each individual expert in a country group of experts has her own significance level $\alpha_i$. Thus, heterogeneity is located in the diverse extent of belief anchoring in each country group.

In many circumstances it might be natural to assume that the IE test statistic of a central bank is some function of the departure of inflation from the target (here, the 2% ECB target). For example, this was the assumption of Henckel et al. (2011). However, in the current crisis, it is both realistic and convenient to assume that the danger to credibility does not stem from the recent inflation history. It is realistic because inflationary concerns relate to the doubts about the fiscal sustainability of the Euro venture itself (Borgy et al., 2011),[10] and it is convenient because the independence of $\beta$ and lagged inflation makes our estimation below simpler.[11]

There would be no inference problem for the experts in the model described in (1) if $\pi^{cb}$ were common knowledge *and* credible. However, it is possible for the central bank to lie about—or anyway not follow through coherently with—monetary policy, in the sense of genuinely targeting a value of inflation different from the official value.[12] Using the notion of official target credibility, as previously defined, we denote $\pi^*$ the announced inflation target which may or may not equal $\pi^{cb}$. To the extent that agents believe $\pi^{cb}$ differs from $\pi^*$, official target credibility is eroded.

In the IE framework it is a standard discipline to adopt rational expectations as the alternative hypothesis (Menzies and Zizzo, 2009; Henckel et al., 2011).[13] Adopting that convention here suggests the following hypotheses:

$$\begin{aligned} H_0: &\quad \pi^{cb} = \pi^* \\ H_1: &\quad \pi^{cb} = E\left[\pi_t\right], \end{aligned}$$

where $E\left[\cdot\right]$ is the mathematical expectation operator. That is, the null hypothesis is that the central bank is really pursuing its announced target, but when experts do not believe that, they embrace RE.

We assume that all the experts combine their beliefs about inflation together, and their weighted average determines price setters' beliefs, with a proportion $\beta$ of experts believing $H_0$ and a proportion $(1-\beta)$ believing $H_1$. Substituting into (1) and allowing for an i.i.d. estimation error, $\eta_t$, we obtain:

$$\pi_t = \beta\pi^* + (1-\beta)\, E\left[\pi_t\right] + \gamma_1 \pi_t^{EN} + \gamma_2 e_t + \eta_t. \tag{2}$$

[10] We do not need to assume this however - inflationary concerns could be due to a range of economic factors and, as long they do include recent inflation history, this would be compatible with our framework.

[11] The dependence between the proportion of agents holding onto beliefs and lagged values of the dependent variable make the chartists/fundamentalists framework (Ahrens and Reitz, 2000, and Frankel and Froot, 1986) very challenging to estimate. Our assumption is a reasonable shortcut for this particular context which makes estimation straightforward.

[12] Alternatively, an inflation bias may be present, implying a wedge between the official inflation target and the actual equilibrium value, and this is imperfectly observed by experts. This latter possibility is in the spirit of Barro and Gordon (1983a, 1983b) and Henckel et al. (2011).

[13] This has the considerable advantages of model parsimony and of nesting RE as a special case of IE, for when $\alpha = 1$ the null must be rejected and RE obtains.

Agents who have rational expectations know the model and calculate the rational expectation of inflation based on (2). We can take mathematical expectations through (2), solve for $E[\pi_t]$, and then substitute back into (2) to obtain the IE Phillips curve.[14]

Solving for the expectation of (2) and substituting it back into (2) yields

$$\pi_t = \pi^* + \frac{\gamma_1}{\beta}\pi_t^{EN} + \frac{\gamma_2}{\beta}e_t + \eta_t. \tag{3}$$

Equation (2) shows how inflation is affected directly by $\pi^*$, $\pi^{EN}$ and $e$ and indirectly by beliefs about these variables embedded in $E[\pi]$. Let $Z_t = \gamma_1\pi_t^{EN} + \gamma_2 e_t$ and substitute the expectation of (2) back into (2):

$$\begin{aligned}
\pi_t &= \beta\pi^* + (1-\beta)E[\pi_t] + Z_t + \eta_t \\
&= \beta\pi^* + (1-\beta)(\beta E[\pi^*] + (1-\beta)E[E[\pi_t]] + E[Z_t]) + Z_t + \eta_t \\
&= \beta\{\pi^* + (1-\beta)E[\pi^*]\} + \{Z_t + (1-\beta)E[Z_t]\} + (1-\beta)^2 E[E[\pi_t]] + \eta_t.
\end{aligned}$$

Define the expectations operator such that $E[E[\cdot]] = E^2[\cdot]$. Repeated substitution of the expectation of (2) into $E^j[\pi_t]$, $j > 1$, demonstrates, upon rearranging, that inflation depends upon beliefs about beliefs, viz. higher-order beliefs:

$$\begin{aligned}
\pi_t &= \beta\left\{\pi^* + (1-\beta)E[\pi^*] + (1-\beta)^2 E^2[\pi^*]\right\} \\
&\quad + \left\{Z_t + (1-\beta)E[Z_t] + (1-\beta)^2 E^2[Z_t]\right\} + (1-\beta)^3 E^3[\pi_t] + \eta_t \\
&\quad \ldots \text{ making substitution } k \text{ times} \\
&= \beta\left\{\pi^* + (1-\beta)E[\pi^*] + \ldots + (1-\beta)^k E^k[\pi^*]\right\} \\
&\quad + \left\{Z_t + (1-\beta)E[Z_t] + \ldots + (1-\beta)^k E^k[Z_t]\right\} + (1-\beta)^{k+1} E^{k+1}[\pi_t] + \eta_t.
\end{aligned}$$

Appealing to the law of iterated expectations ($E^j = E$) and conditioning on $Z_t$ and $\pi^*$, we have that $E^j[Z] = Z$ and $E^j[\pi^*] = \pi^*$ and thus,

$$\begin{aligned}
\pi_t &= \beta\left\{\pi^* + (1-\beta)\pi^* + \ldots + (1-\beta)^k \pi^*\right\} \\
&\quad + \left\{Z_t + (1-\beta)Z_t + \ldots + (1-\beta)^k Z_t\right\} + (1-\beta)^{k+1} E^{k+1}[\pi_t] + \eta_t.
\end{aligned}$$

When $k$ goes to infinity we see that the impact on inflation of beliefs about beliefs concerning $Z$ compounds, since inflation is driven by the sum of a geometric progression $\left(1 + (1-\beta) + (1-\beta)^2 + \ldots\right) Z = Z/\beta$ when $\beta < 1$. In contrast, the impact on inflation of beliefs about beliefs concerning $\pi^*$ *cancels out*, since the geometric progression $\left(1 + (1-\beta) + (1-\beta)^2 + \ldots\right)\pi^* = \pi^*/\beta$ is pre-multiplied by $\beta$, and the $\beta$'s cancel. Writing out the relevant geometric progressions gives us an alternative expression for (3), which makes it clear why there is a unitary marginal impact of $\pi^*$ on $\pi_t$ and an

[14] The pronumeral $\pi^*$ could be anything that can be conditioned on at time $t$, including an autoregression in $\pi$. We assume $\pi^*$ is the official inflation target since that is the normal anchor in an inflation targeting regime.

impact of $Z$ which grows as $\beta$ decreases:

$$\begin{aligned}
\pi_t &= \lim_{k\to\infty}\Big(\beta\left\{\pi^* + (1-\beta)\,\pi^* + \ldots + (1-\beta)^k\,\pi^*\right\} \\
&\quad + \left\{Z_t + (1-\beta)\,Z_t + \ldots + (1-\beta)^k\,Z_t\right\} + (1-\beta)^{k+1}\,E^{k+1}\left[\pi_t\right]\Big) \\
&= \left(\beta\pi^* + Z_t\right)\left(1 + (1-\beta) + (1-\beta)^2 + \ldots\right) + 0 \\
&= \underbrace{\left(\beta\pi^* + Z_t\right)}_{\text{direct}} + \underbrace{\frac{1-\beta}{\beta}\left(\beta\pi^* + Z_t\right)}_{\text{beliefs about beliefs}} \qquad (4)
\end{aligned}$$

Intuitively, when $\beta$ is unity, all agents believe the inflation target, but inflation itself is buffeted about by mechanical forces which are represented by $Z$. This can be seen by direct substitution of $\beta = 1$ into (2), or by noting that there is no impact of beliefs about beliefs in the second term of the last line of (4).

When $\beta = 0$, direct substitution into (2) fails.[15] However, we can take the limit of (4) as $\beta$ goes to zero. A story emerges where agents who abandon $H_0$ recognize that *some other agents* still believe it, so the agents who have rejected $H_0$ still have to factor $\pi^*$ into their (rational) expectation.

If $\beta$ approaches zero there is a very strong effect of beliefs about beliefs in (4), because $(1-\beta)$ is close to unity, and so even if a vanishingly small group of agents anchor onto $H_0$, the effect of beliefs of agents who believe $H_1$ about the $H_0$ agents' beliefs offsets this, and $\pi^*$ appears in the Phillips curve with a unit coefficient. In contrast, the effect of beliefs about beliefs concerning $Z$ grows, because the increasing proportion of agents who believe $H_1$ have to take into account the mechanical effects on inflation of $Z$ in the beliefs about beliefs of other rational agents, and so the marginal impact of $Z$ grows with $(1-\beta)/\beta$ in the IE Phillips curve.

To flag the most germane insight of this section, if (3) is used to forecast inflation, a lower share of agents anchoring onto the null (i.e. a low $\beta$) leads to a higher forecast error variance in the future, for a given variance of energy prices and output gaps. This explains how forecast errors can be buffeted about by central bank anchoring credibility, irrespective of persistence, which is the main point of the paper.

A final comment on (3) is that inflation differs from its rational expectation ($E[\pi]$) by a white noise error $\eta_t$. That is, although the whole model cannot be called a rational expectations model, since some agents default to the null if $P$ is non-zero, inflation expectations themselves are model consistent and therefore rational. Two things follow directly from this.

First, the true state of the world (at least up to a white noise error) is that the null is false so $\beta$ is actually the probability of a type II error, in line with standard statistical terminology, as flagged earlier.

Second, standard evolutionary arguments for convergence to rational expectations could be made (Alchian, 1950). The validity of our model would then depend upon some implicit calculation costs, inattention, or belief inertia borne of psychological factors.[16]

[15] After making the substitution, a contradiction ensues if expectations of the resultant equation are taken.

[16] For a discussion of anchoring driven by belief conservatism see Lyons et al. (2012) or Edwards (1968). For a discussion of attention due to other factors expending the cognitive resources of agents

## 3 $\alpha$ and $\beta$

In this section we demonstrate how changes in beta relate to changes in anchoring credibility (that is, the tendency of agents to believe $H_0$). Readers who are prepared to accept that an increase in the proportion of agents who believe $H_0$, for a given $P$, corresponds to an increase in belief conservatism may wish to jump straight to section 4.

The conceptual challenge is that with heterogeneity across $\alpha$'s we cannot explain a change in $\beta$ by a change in a single representative agent's $\alpha$, which is the off-the-shelf measure of belief conservatism in IE. If $\beta$ falls by, say, 20%, does that mean that the whole distribution of $\alpha$'s is 'more conservative' in some sense? Answering questions like these require that we map changes in $\beta$ onto changes in the *shape* of an $\alpha$ distribution.

To do this, we need to find a way to describe whole distributions as being more, or less, belief conservative. To that end, we first outline the relationship between $\beta$ and the distribution of $\alpha$. In any hypothesis, $H_0$ is maintained if a p-value ($P$) fails to fall below the test size, that is if $P > \alpha$. This immediately gives us $\beta$:

$$\beta = \int_0^P f d\alpha = F(P) - F(0). \tag{5}$$

Can we now limit the set of densities $f$ such that an increase in $\beta$ implies more conservatism for any $P$? To be concrete, imagine several groups of experts, one for each country in the Eurozone. Thus, there is a group in the Netherlands, a group in Spain, and so on. We conceive of each individual expert in, say, Spain, drawing from a Spanish distribution of test sizes. In what sense can a Spanish distribution of $\alpha$'s become more belief conservative?

Intuitively, if a group of Spanish experts has a distribution of $\alpha$ with an increasing probability mass around zero, we could describe the group as becoming more 'conservative' in the sense that a randomly chosen $\alpha$ is likely to be closer to zero. But it is not hard to create cases where some probability mass shifts towards zero, and some shifts towards unity.

So, to pin down a clear taxonomy, we confine our attention to a change in the distribution which is *uniformly more conservative*. This is a switch to an $\alpha$ density $f^c$ that delivers a higher value of the integral in (5) compared to an original $\alpha$ density $f$ *for any value of* $P$:

$$\int_0^P f^c d\alpha > \int_0^P f d\alpha \quad \forall P \in (0,1). \tag{6}$$

This, then, is our way of mapping $\beta$ onto the shape of an $\alpha$ distribution. If Spanish $\beta$, say, increases and we assume $P$ is unchanged, then we can say that it is as if the Spanish distribution of $\alpha$ has become uniformly more conservative.

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see Andrade and Le Bihan (2010) and Luo and Young (2010), which relate to earlier papers (Reis, 2006; Sims, 2003; Mankiw and Reis, 2002). Another justification for IE based on the supposed logical consistency of economists is that it removes what we might call 'expectational asymmetry'. If we, as economists, use hypothesis tests in helping form our beliefs, then we should also expect agents in our models to do so. This is redolent of the Public Choice revolution which removed 'motivational asymmetry' from the analysis of government. It will be recalled that these scholars asked why the governed were assumed to be self interested while the government was benevolent.

Confining our attention to uniformly more conservative distributions means that we can make a claim about changed conservatism even if we do not know what $P$ is. That is, if we are examining a cross section of country expert groups under the same monetary regime, then all we need to do is assume they have the same $P$, even if we do not know what it is. If we observe estimated $\beta$'s change differentially across European economies, as we do in the next section, we can attribute the differential changes to changes in belief conservatism.[17]

We conclude this section by illustrating the simplest possible example of a uniformly more conservative distribution, and leaving more general pdfs to Appendix C. The simplest distribution is a straight line density, that is where $f'(\alpha)$ is a constant. The properties of pdf's allow us to derive an explicit formula:

$$\begin{aligned} f'(\alpha) &= \frac{df}{d\alpha} = \lambda \\ &\therefore \;\; f = \lambda\alpha + c \quad \text{and} \quad \int_0^1 f d\alpha = 1 \quad \Rightarrow \quad c = 1 - \frac{\lambda}{2} \\ &\therefore \;\; f = \lambda\alpha + 1 - \frac{\lambda}{2} \end{aligned}$$

If there are conditions under which $\beta$ will rise for any value of $P$, then according to (6) those conditions will define a change in $f$ that is uniformly more conservative:

$$\begin{aligned} \beta &= \int_0^P \left(\lambda\alpha + 1 - \frac{\lambda}{2}\right) d\alpha \qquad (7) \\ \frac{d\beta}{d\lambda} &= \int_0^P \frac{d}{d\lambda}\left(\lambda\alpha + 1 - \frac{\lambda}{2}\right) d\alpha = \int_0^P \left(\alpha - \frac{\lambda}{2}\right) d\alpha = \frac{-P(1-P)}{2} < 0. \end{aligned}$$

So, for any $P$, if $\lambda$ decreases, $\beta$ increases. Thus, a decrease in $\lambda$ makes $f$ more uniformly conservative. Graphically, a decrease in $\lambda$ shifts the probability mass towards zero everywhere on the distribution.

The analysis in this section allows us to connect increases in estimated $\beta$ to uniformly more conservative shifts in the distribution of $\alpha$. Since these explanations are not dependent on the particular value of $P$, we will be able to infer changes in the distribution of $\alpha$ in different countries when everyone has a common, but unknown, $P$. Naturally, this only applies as we look across countries at a point in time. Through time, $P$'s may change as the evidence against the null changes, but if that alone is the cause it will do so equally across all countries by assumption. Differential responses through time will indicate changes in belief conservatism across countries.

[17] This may be controversial since it is more natural to speak of the (singular) credibility of the single monetary authority, the ECB. We, however, speak of credibility in individual Eurozone countries meaning credibility as perceived by price setters within each country. As explained in the text this enables us to cleanly identify changes in cross-country belief conservatism while keeping monetary policy, and information about monetary policy, common across all countries. Another justification is that the Euro has been sufficiently imperiled by recent events that there is a 'shadow' central bank—with its own credibility—sitting within each country waiting to be (re)born, should the country concerned abandon the Euro. We take our later results, which identify a differential erosion of anchoring credibility in Portugal, Ireland, Italy and Greece, as providing support for our approach.

# 4 Estimates of Credibility

## 4.1 An Estimable Form of the Model

Equation (3) is the basis for our estimation. To obtain a measure of official target credibility we also allow the steady state inflation rate to change after the crisis (that is, from the September quarter of 2008). We estimate the following over March 2001 until December 2012:[18]

$$\pi_t = \theta_1 (1 - D) + \theta_2 D + \frac{\theta_3}{1 - \theta_5 D} \pi_t^{EN} + \frac{\theta_4}{1 - \theta_5 D} gap_t + \eta_t. \tag{8}$$

The parameters $\theta_3$ and $\theta_4$ in (8) equal $\gamma_1/\beta$ and $\gamma_2/\beta$ in (3) prior to the crisis, while $\theta_3/(1-\theta_5)$ and $\theta_4/(1-\theta_5)$ are the post crisis parameter ratios, after $\beta$ has fallen by a proportion $\theta_5$. They represent the impact of the energy price and the output gap on inflation, which operates both directly, and indirectly via expectations. The parameter $\theta_5$ should have an absolute value of no more than unity.[19] The proportional reduction applies equally to both coefficients, which is what is implied by (3). Inflation ($\pi$) and energy price growth relative to trend ($\pi^{EN}$) are both four-quarter ended.[20] The dummy $D$ takes on unity from September 2008, and the output gap is the deviation of the log of GDP from an OLS trend.[21] The parameters $\theta_1$ and $\theta_2$ are the pre- and post-recession inflation targets as perceived by experts.

The raw estimates are shown in Table 2. In 2A, we show both the Euro area model, where we assume a single group of experts over the whole Euro area and we use an aggregate price index and output gap, and countries with a high estimate of $\theta_5$. In Table 2B, we have the remaining countries with lower values of $\theta_5$. The models for Netherlands and Finland have a low adjusted $R^2$, whereas the others have adjusted $R^2$ values around or above 0.5.

*(Insert Table 2 about here)*

With the exception of $\theta_5$ for Ireland, which exceeds unity though not statistically significantly so (two-tailed $P = 0.117$), all other coefficients are within economically

[18] The estimation period is partly dictated by measurement issues in the output gap (described in a subsequent footnote) and partly by Greece's delayed entry into the Euro in 2001. We can interpret $\theta_1$ and $\theta_2$ as steady state inflation because we transformed the data for energy price inflation to be deviations from trend, which is zero in the steady state.

[19] Of course it could be negative, implying a rising $\beta$. The denominators are set up with minus signs only because we knew that $\beta$ was falling from exploratory regressions, but this is without loss of generality.

[20] Four-quarter-ended energy CPI is subtracted from the mean over March 2001-December 2012.

[21] Data was collected from 1998, and four-quarter ended price changes therefore began in 1999. Initial attempts to map annual output gap data from the OECD to quarterly OLS residuals of log output onto a linear trend from 1998 onwards ran into serious difficulties when the OLS residuals over 1998-2000 showed large negative gaps in Austria, France, Ireland, Italy, Luxembourg and the Netherlands, while the OECD ones showed positive or only marginally negative ones. In the end we used OLS residuals from 2001 based on a regression of log output on trend over the whole period, namely 1998-2012. These matched the OECD annual figures well over this truncated period.

reasonable ranges. The estimated inflation targets are all below 5% and many are close to the ECB's 2% target. All of the statistically significant coefficients on energy price inflation, $\theta_3$, are in the 0.05 to 0.10 range, which is not implausible considering energy has a weight of around 10 per cent in the CPI of European countries.[22] The only significant output gap term, for Germany, has an estimated coefficient of 0.2, which is also reasonable.

We set about simplifying the models in a number of ways. First, we deleted a number of insignificant coefficients. The Euro area equation was left untouched. Although the output gap was technically insignificant, the p-value was 0.1044, and the estimated effects before and after the recession (0.04 and 0.09) seemed economically reasonable. We also refrained from deleting $\theta_3$ and $\theta_4$ for Italy and Greece. Although the coefficients are insignificant, after the crisis they are divided by a very small number $(1-\theta_5)$ and these quotients have economically reasonable magnitudes. For Portugal and Ireland, the output gap terms are negative (though not significantly so) and we re-estimated both, constraining $\theta_5$ to be strictly between zero and unity and eliminating negative insignificant coefficients.[23]

Coming to the second group of countries, we eliminated a number of variables that were not significant at the 10% level. We left Finland and the Netherlands unchanged, as the models were already poorly fitted.

As part of our model reduction, we experimented with adding lags to the equations, because economically there is nothing stopping the null belief $\pi^*$ including lags of inflation. We also had econometric reasons for doing this, because the estimated equations had low Durbin-Watson statistics, indicating serial correlation in the errors. In the event, the lags did not change the results substantively (nor did they eliminate the autocorrelation), though that left us with the possibility of downward biased standard errors.

We included a moving average of the output gap in the equations.[24] This did not eliminate autocorrelation, although it provided us with some stronger output gap effects. The total effect of the output gap prior to the crisis, which is the sum of the MA(4) coefficients, was higher than in the un-lagged model for some countries (notably for Austria, Finland and the Netherlands). However, for the Euro area overall, the lags were jointly insignificant. Most importantly for the purposes of this paper, the proportional reduction in the number of agents believing the null, $\theta_5$, generally had very high standard errors, in contrast to the model without lags.[25]

We conducted feasible GLS to correct the standard errors for the model without lags.[26] Table 3 contains the simplified models, with feasible GLS standard errors be-

[22] The CPI weights for energy are 0.107, 0.09 and 0.095 for Germany, France and Italy (OECD, 2012).

[23] This was accomplished by estimating an unrestricted $r$ in $\theta_5 = r^2/\left(1+r^2\right)$.

[24] We are grateful to Adrian Pagan for suggesting this. A left hand side variable measured as a four quarter ended change is approximately equal to the sum of four quarter-on-quarter changes. Therefore, a single period impulse to an output gap might have a differential impact on each of these four quarterly changes. In theory, these different responses can be estimated by including up to four lags of the output gap on the right hand side. In practice, as in some of the regressions of Gruen et al. (1999), it did not improve the model overall. Autocorrelation can also be explained by unmodelled shifts in constant terms.

[25] These regressions are available in an electronic appendix.

[26] Feasible GLS, where the error structure is modelled rather than known *a priori*, is not necessarily

neath them.[27] There is also a column showing the ratio of feasible GLS standard errors to their OLS counterparts, averaged over all the countries with the correction. For $\theta_1$, $\theta_2$ and $\theta_4$ the feasible GLS standard errors are roughly twice as large as OLS. For the remaining coefficients they are roughly the same. The GLS correction does not change the significance of any coefficient at the 5% level, and the model simplifications do not change any coefficient by more than 0.2, with most altered coefficients changing by far less than that. The results are therefore robust to the simplifications.

*(Insert Table 3 about here)*

## 4.2 Official Target Credibility

We now ask how the estimated coefficients in Table 3 might address the central questions of the paper.[28] It is clear that official target credibility has not been a serious issue in the GR. A number of countries have seen increases in the null hypothesis inflation rates (Austria, Belgium, Finland and Germany) but the remaining countries have seen decreases, some substantially. For the Euro area overall (the first column of 3A), the point estimates indicate a mild improvement in official target credibility: prior to the GR the perceived target was 2.25% whereas afterwards it dropped to 2.07%. However, this difference is not significant ($t = 1.15$ in a reparameterized regression).[29]

## 4.3 Anchoring Credibility

Treating the Euro area as one block of experts with one $\alpha$ distribution, there has been a reduction in the proportion of agents believing the null of just over 60% (0.61 in first column of Table 3).[30] However, this loss of confidence has varied considerably across countries. At one extreme is Germany, a country where $\theta_5$ was not significantly different from zero (a proportional reduction of 0.06 in Table 2B). That is, the tumultuous events of 2008 have had no impact upon the proportion of agents holding onto the null belief

efficient so we adopt the common practice of using the OLS estimated coefficients. These are not corrupted by any errors in modelling the error structure, and are unbiased. We used a simple Cochrane-Orcutt procedure, and dropped the first observation. For each equation, we took the estimated residuals and ran the regression

$$\widehat{u}_t = \rho\widehat{u}_{t-1} + \eta_t.$$

The estimated AR(1) coefficient was then used to transform all the model variables by

$$z_t^* = z_t - \widehat{\rho} z_{t-1}.$$

[27] The Cochrane-Orcutt representations for Portugal and Ireland did not converge, so we used the original (OLS) standard errors.

[28] We will cast this discussion in terms of Table 3, though all the main points apply to the models of Table 2 as well, i.e. they are robust to our model simplification choices.

[29] The right hand side of the Euro area regression in Table 3 includes $\theta_1 D + \theta_2 (1 - D)$. The $t$-stat in the text (1.15) is for $\theta_1$ in a regression with $\theta_1 D + \theta_2$.

[30] As (8) makes clear, the 0.61 is a proportional reduction in $\beta$, which is itself a (latent) proportion. If $\beta$ were, say, 0.8 prior to the crisis, the new Euro area $\beta$ after the crisis would be $(0.8)(1 - 0.61)$ not $0.8 - 0.61$.

which, in turn, is close to the ECB's 2% target before and after the crisis (1.59 and 1.69 in Table 3). At the other end of the spectrum lie Portugal, Ireland, Italy and Greece, all of which have seen a reduction in $\beta$ of more than 80%.

Since we are not estimating $\beta$, but only its proportional reduction, it is hard to know how important these reductions are: if all the countries start with very low $\beta$'s then an 80% reduction might not amount to much economically. However, one thing we can say is that the proportional reduction in $\beta$ increases the coefficients in front of the output gap and energy prices in (3), and so increases the standard error associated with forecasts, by the same proportion.

Using the IE framework we can therefore provide a narrative for the way in which distressed economies in the Eurozone have processed information about inflation, which in turn feeds into higher forecast error variance. If we are prepared to maintain our assumption of common information across the economies at any point in time, we can explain the differential response of some countries at that point in time by differential changes in attitudes to evidence. It is interesting that the countries with the highest increase in $\beta$ are those who have fared badly in the crisis, and have even flagged a Euro exit.

Explaining the increase in forecast errors by a weaker attachment to the null decouples the variance of inflation from its persistence, allowing for a greater menu of possible explanations of either. The most likely sources for the fall in inflation persistence are to be found in the explanatory variables—output gaps and energy prices in our specification (8)—which both experienced reduced persistence in wake of the Great Recession (Figure 2). A definitive exploration of these sources of inflation persistence is not our current focus, though we could speculate that the reductions in the persistence of output gaps reflects the renaissance of activist counter-cyclical policy in the wake of the Great Recession.[31]

*(Insert Figure 2 about here)*

# 5 Discussion and Conclusion

A Phillips curve based on inferential expectations can explain how detachment from the null hypothesis can lead to greater inflation forecast variance using a mechanism flagged by Gerlach et al. (2011). Speaking of a greater spread in forecasts of inflation they say that "while this could reflect greater macroeconomic uncertainty in general, it could also suggest that expectations may be less firmly anchored than they used to be." (p. 49).

This mechanism operates independently of the extent of autocorrelation in the time series properties of inflation. Rather, forecasts based on equation (3) will exhibit greater volatility because of the changes in the coefficients as people lose their grip on the inflation target and become more attached to other information such as output gaps

[31] By definition, counter-cyclical fiscal policy seeks to close the output gap. In an analogous argument to Benati (2008) one might expect a renaissance of fiscal activism to result in less persistent output gaps.

and energy prices. As a result, we have a decoupling between experts' forecasts of inflation and a fall in persistence in European inflation rates.

A core part of the analysis has been to distinguish between official target credibility, which is whether agents treat the official target as the null hypothesis, and anchoring credibility, which is whether agents tend to anchor their belief to the null hypothesis, whatever that might be.

As we conclude, it is therefore worth considering whether that distinction is worthwhile in practice. That is, if the two dimensions of credibility are highly correlated, then either one could be a sufficient statistic for the other. If, for example, countries lost anchoring credibility the moment targeted inflation left the official target, we would be back in the world of uni-dimensional credibility, which could be described by either the proximity to the target, or the proportion of agents anchored to it.

Our econometric results give solid evidence that this is not the case in Europe. In Panel A of Figure 3, which is based on Table 3, we measure the percentage point loss in official target credibility on the horizontal axis.[32] This is measured by $|\theta_2 - 2| - |\theta_1 - 2|$, the proximity of the inflation target to the 2% official target after the crisis minus the proximity of the inflation target before it. If this is a positive number, it means the real inflation target is drifting further away from the official one. On the vertical axis, we plot the proportional reduction in $\beta$, which is estimated by $\theta_5$ in our models. A strong relationship (of either sign) would indicate that the distinction we have made in the paper is of little practical importance.

In Panel B we calculated the percentage point loss in official target credibility but do not use the 2% flagged by the ECB. Instead, we adopt the midpoint of Germany's 'before' and 'after' inflation target, namely 1.65%, giving the percentage point loss $|\theta_2 - 1.65| - |\theta_1 - 1.65|$. The reasoning here is that Germany is an important country in the Eurozone and that the ECB wants inflation 'below, but close to, 2 %' (http://www.ecb.int/mopo/html/index.en.html), so it is natural to pick the inflation target of a pivotal country which is, in fact, below, but close to, 2%.

*(Insert Figure 3 here)*

In neither panel is a strong correlation evident (for targets of 2 and 1.65 Pearson's $r$ is $-0.30$ and $-0.33$). To the extent that there is a weak negative association it is driven by Ireland and Portugal—both of whom experienced a very severe loss of anchoring credibility, in spite of the fact that their target inflation moved closer to the ECB target (measured either as 2% or 1.65%). Overall, the two measures are more or less independent, justifying a two-dimensional analysis of credibility.

The inferential expectations framework can therefore be used as a means of building a two-dimensional construct for credibility, and empirically as a means of interpreting a rise in estimated coefficients (in (3)) during the recent recession in the light of this framework.

Our IE analysis has shown that, as a result of the Great Recession, there is no evidence for an erosion of official target credibility in the Euro area: the ECB's official

[32] We use percentage points such that a fall of the inflation target from, say, 2% to 1% would be a 50% reduction but a 1% point reduction.

target of 2% is still, in its own way, credible in the sense that experts believe the monetary authorities are aiming for something close to this.

However, we have found evidence of an erosion of anchoring credibility, especially in Portugal, Ireland, Italy and Greece, as experts are more likely to consider new information on its merits and anchor their inflation forecasts less on the target of 2%.

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# 6 Appendixes

## 6.1 Appendix A

This appendix outlines a more general framework, and shows that lower persistence and a greater variance contradict each other in that framework too. Let

$$\pi_t = \pi_t^* + u_t$$

and

$$\pi_t^* = \rho \pi_{t-1}^* + e_t.$$

Using the lag operator these two equations may be rewritten as

$$(1 - \rho L)\,\pi_t = e_t + (1 - \rho L)\,u_t.$$

We then have the variance,

$$\gamma(0) = \sigma_u^2 + \frac{\sigma_e^2}{1-\rho^2} \qquad \text{and} \qquad \gamma(1) = E\left[\pi_t \pi_{t-1}\right] = \frac{\rho \sigma_e^2}{1-\rho^2}.$$

So,

$$\frac{\gamma(1)}{\gamma(0)} = \frac{\rho}{(1-\rho^2)\,\sigma_u^2/\sigma_e^2 + 1}.$$

If the variance of underlying inflation (say $\sigma_e^2$) increases, the variance of observed inflation $\gamma(0)$ becomes higher, but the autocorrelation $\gamma(1)/\gamma(0)$ increases, as in the simple model of the main text. Nor is the problem removed if $E\left[\pi_t^* \mid \pi_t\right] = \sigma_{\pi^*}^2 / \left(\sigma_{\pi^*}^2 + \sigma_u^2\right)\pi_t$ is somewhere in the model, since it inherits the properties of $\pi_t$. So, if $\sigma_e^2$ rises, so does $\sigma_{\pi^*}^2$, and with it the volatility of $E\left[\pi_t^* \mid \pi_t\right]$. However, the autocorrelation still increases, driven by the above properties of $\pi_t$.

## 6.2 Appendix B

The following appendix provides an example of how to derive the Lucas supply equation (6) used in the main text.

Nature draws a real marginal cost shock $e \sim N.i.i.d.\,(0, \sigma_e^2)$ and price setting agent $j$ observes a noisy signal: $x_j = e + v_j$ where $v_j \sim N.i.i.d.(0, \sigma_v^2)$. Considering an agent's belief about $e$ given $x_j$, we follow Lucas (1972) to obtain agent $j$'s optimal guess of $e$:

$$E\left[e \mid x_j\right] = \frac{\sigma_v^2}{\sigma_v^2 + \sigma_e^2} E\left[e\right] + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} x_j = \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} x_j. \qquad \text{(A1)}$$

A single producer in a centralized economy—the 'aggregator'—demands a continuum of substitutable intermediate goods $y_j$ indexed by $j$ over $[0,1]$ and combines them using CES production to produce a final good:

$$y = \left[\int_0^1 y_j^{\frac{\theta-1}{\theta}}\,dj\right]^{\frac{\theta}{\theta-1}}, \qquad \theta > 1. \qquad \text{(A2)}$$

Firm j maximizes profits, $py - \int_0^1 p_j y_j dj$, subject to (A2), which yields the demand for inputs as a function of relative prices,

$$y_j = y \left(\rho_j\right)^{-\theta}, \tag{A3}$$

where $\rho_j \equiv p_j/p$ is firm $j$'s relative price.

The aggregate price level $p$ is obtained by eliminating $y$ in (A2) and (A3):

$$p^{1-\theta} = \int_0^1 p_j^{1-\theta} dj. \tag{A4}$$

Intermediate firms all face the same production function with time-varying real marginal costs—equal across all firms—given by $rmc = rmc_{ss}(1+e)$ where $rmc_{ss}$ refers to steady state real marginal cost.

The economic interpretation of $e$ is that it is a demand shock arising from, say, fiscal or monetary policy. Thus, $e$ captures any movement *along* a Phillips curve. The zero mean of $e$ makes it consistent with the classic identification scheme where demand shocks have zero long run effects (for example, Blanchard and Quah, 1989) and the i.i.d. assumption is for simplicity.[33]

Price setters each receive a noisy signal of $rmc$, denoted $\widehat{rmc}_j$:

$$\widehat{rmc}_j = rmc_{ss}\left(1 + e + v_j\right) = rmc_{ss}\left(1 + x_j\right). \tag{A5}$$

Firms choose $\rho_j$ to maximize real current profits, $\Pi_j \left(\rho_j - rmc_j\right) y_j$. Using (A3) and assuming that firms treat aggregate output (A2) and the aggregate price level (A4) as exogenous and non-stochastic, one obtains the solution for firm $j$'s optimal relative price:

$$\rho_j = \frac{\theta}{\theta - 1} rmc_j,$$

or,

$$p_j = \frac{\theta}{\theta - 1} rmc_j p. \tag{A6}$$

Agents set prices (A6) after forming beliefs about $rmc_j$ . We assume they use the Lucas signal extraction solution (2) to infer $e$ from $x$. When it comes to calculating $p$, we do not require that the agents solve the whole system to arrive at a model-consistent $p$, since this would sit awkwardly with the assumption that they treat $p$ as exogenous in their maximization. Instead, we assume that firms use the price level implied by expected inflation, $\pi^e$, similar to Yun (1996). In the main text we assume that firms employ the publicly available consensus inflation forecast as an estimate of inflation but some other expectation scheme, such as adaptive expectations or rational expectations, is also possible. Introducing time subscripts, the relationship between the current price level and expected inflation is trivially given by $p_t = p_{t-1}\left(1 + \pi^e\right)$.

To derive the Phillips curve, use a Lebesgue integral to rewrite (A4) as an expectation over the noise $v_{jt}$ conditioning on $e_t$, which we will just denote $E_j[\cdot]$. That

[33] Nothing depends on this; it just makes the signal extraction solution (A1) simple. In the main text we allow it to exhibit autocorrelation.

is,

$$\begin{aligned} p_t^{1-\theta} &= \int_0^1 p_{jt}^{1-\theta} dj \\ &= E_j \left[ p_{jt}^{1-\theta} \right] \\ & E_j \left[ \left( \frac{\theta}{\theta - 1} rmc_j p_{t-1} \left(1 + \pi^e\right) \right)^{1-\theta} \right]. \end{aligned}$$

If $e$ and $v$ are small, $(1+y)^{1+\theta}$ approximates to $(1 + (1+\theta)\,y)$ and we have,

$$\begin{aligned} p_t^{1-\theta} &= E_j \left[ \left( \frac{\theta}{\theta-1} \left(1 + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} x_{jt}\right) p_{t-1} \left(1+\pi\right) \right)^{1-\theta} \right] \\ &= \left( \frac{\theta}{\theta-1} rmc_{ss} \left(1+\pi\right) p_{t-1} \right)^{1-\theta} E_j \left[ \left(1 + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} x_{jt}\right)^{1-\theta} \right] \\ &= \left( \frac{\theta}{\theta-1} rmc_{ss} \left(1+\pi\right) p_{t-1} \right)^{1-\theta} \left(1 + E_j \left[ \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} x_{jt} \right] \right)^{1-\theta} \\ &= \left( \frac{\theta}{\theta-1} rmc_{ss} \left(1+\pi\right) p_{t-1} \right)^{1-\theta} \left(1 + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} E_j \left[x_{jt}\right] \right)^{1-\theta} \\ &= \left( \frac{\theta}{\theta-1} rmc_{ss} \left(1+\pi\right) p_{t-1} \right)^{1-\theta} \left(1 + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} e_t \right)^{1-\theta} \end{aligned} \tag{A8}$$

In the first line of (A8) we use the fact that firms use the Lucas signal extraction solution (2) for $rmc$. In line 3, the small $e$ and $v$ approximation allows us to take a linearized expectation over $x$ and then reintroduce the exponential. In lines 4 and 5 we take expectations conditional on $e$, because we want the economy-wide effect on inflation after the realization of $e$. In a steady state, $e_t = 0$, $p_t = p_{t-1}\left(1+\pi^e\right)$, and $rmc_t = rmc_{ss}$. Substituting these terms into (A8) gives the steady-state solution to (A8):

$$\frac{\theta}{\theta-1} rmc_{ss} = 1. \tag{A9}$$

Substituting back into (A8) and noting $e\pi^e$ is second-order small we have:

$$p_t^{1-\theta} = \left(\left(1+\pi_t^e\right) p_{t-1}\right)^{1-\theta} \left(1 + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} e_t\right)^{1-\theta}.$$

Rearranging yields

$$\begin{aligned} \frac{p_t}{p_{t-1}} &\approx 1 + \pi_t^e + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} e_t \\ 1 + \pi_t &= 1 + \pi_t^e + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} e_t \\ \pi_t &= \pi_t^e + \frac{\sigma_e^2}{\sigma_v^2 + \sigma_e^2} e_t \end{aligned} \tag{A10}$$

Defining $\Phi = \sigma_e^2 / (\sigma_v^2 + \sigma_e^2)$ yields the Lucas supply function (2) in the main text. This supply function has the usual properties—there exists a positive relationship between inflation and an output measure and the curve shifts up or down with changes in expected inflation.

## 6.3 Appendix C

This appendix creates a broader family of uniformly more conservative distributions than the straight line with a higher gradient in the main body.

**Proposition 1** *If $f$ and $f^c$ are both continuous pdf's of the test size $\alpha$ over support $[0, 1]$ and the gap $f^c - f$ is decreasing as $\alpha$ increases, $f^c$ is uniformly more conservative.*

**Proof.** Since $f$ and $f^c$ are densities their difference must integrate to zero:

$$\int_0^1 (f^c - f)\, d\alpha = \int_0^1 f^c d\alpha - \int_0^1 f d\alpha = 0.$$

Figure 4 plots two functions for which the difference is decreasing and integrates to zero. It is obvious that the difference must start with a positive value and end with a negative value. Since the difference is monotonically decreasing, the curves cross only once.

Equation (6) in the text follows immediately for $P$-values no greater than $\alpha^*$. For any $P$ no greater than $\alpha^*$, say at $P_1$, a shift to the more belief conservative (dashed) distribution involves integrating over a more positive function.

For a $P$-value like $P_2$, which lies above $\alpha^*$, we note that the sum of the areas between the curves $f$ and $f^c$ over $[0, 1]$ is zero, and that it would be necessary to integrate from the intersection $\alpha^*$ to unity to fully offset the positive gap $f^c - f$ from 0 to $\alpha^*$. So for a range of integration less than that—such as 0 to $P_2$—the integral over the dashed function $f^c$ must be higher. ■

*(Insert Figure 4 about here)*

Our assumption about the decreasing gap $d(f^c - f)/d\alpha < 0$ implies $df^c/d\alpha < df/d\alpha$ for all $\alpha$ and if we let $df'$ refer to changing the function $f$ such that the slope $(f' = df/d\alpha)$ is everywhere more positive (going from the dashed to more solid line in Figure 4), then

$$\frac{d\beta}{df'} < 0 \quad \forall P \in (0, 1).$$

Geometrically, uniformly more conservative distributions have a greater probability mass around zero, and the above mathematics just rules out perverse cases where the probability mass shifts towards zero over some of the support $[0, 1]$ but away from it at other parts of the support. The point of this analysis is that if we empirically observe an increase in $\beta$ we can interpret it informally as a shifting of the probability mass towards zero, and formally from a solid to dashed Figure-4-style distribution.

**Table 1: Two Dimensions of Credibility**

| | | Official Target Credibility | |
|---|---|---|---|
| | | $H_0$ at official target | $H_0$ not at official target |
| Anchoring | High $\beta$ | credible | ? |
| Credibility | Low $\beta$ | ? | not credible |

**Table 2: Original Non-Linear Least Squares Estimation**
(estimation period: Mar-2001 to Dec-2012; quarterly data)

$$\pi_t = \theta_1 (1 - D) + \theta_2 D + \frac{\theta_3}{1 - \theta_5 D} \pi_t^{EN} + \frac{\theta_4}{1 - \theta_5 D} gap_t + \eta_t$$

**2A: Whole Euro Area and Large $H_0$ Reduction**

| | Euro Area | Ireland | Portugal | Italy | Greece |
|---|---|---|---|---|---|
| $\theta_1$ | 2.25** | 3.99** | 3.05** | 2.35** | 3.35** |
| | (0.08) | (0.28) | (0.16) | (0.08) | (0.11) |
| $\theta_2$ | 2.07** | 0.76 | 1.70** | 2.23** | 2.78** |
| | (0.13) | (0.54) | (0.24) | (0.16) | (0.14) |
| $\lvert\theta_2-2\rvert - \lvert\theta_1-2\rvert$ | −0.18 | −0.75 | −0.75 | −0.13 | −0.57 |
| $\theta_3$ | 0.05** | −0.07 | 0.00 | 0.01 | 0.02 |
| | (0.01) | (0.04) | (0.04) | (0.01) | (0.01) |
| $\theta_3/(1-\theta_5)$ | 0.12 | 0.35 | 0.17 | 0.11 | 0.10 |
| $\theta_4$ | 0.04 | −0.02 | 0.00 | 0.00 | 0.01 |
| | (0.02) | (0.02) | (0.02) | (0.01) | (0.01) |
| $\theta_4/(1-\theta_5)$ | 0.09 | 0.12 | −0.07 | 0.08 | 0.08 |
| $\theta_5$ | 0.61** | 1.20** | 0.98** | 0.94** | 0.82** |
| | (0.12) | (0.12) | (0.23) | (0.13) | (0.11) |
| $R^2$ | 0.73 | 0.75 | 0.59 | 0.66 | 0.73 |

**2B: Moderate $H_0$ Reduction**

| | Luxembourg | Austria | France | Spain | Belgium | Germany | Finland | Netherlands |
|---|---|---|---|---|---|---|---|---|
| $\theta_1$ | 2.41** | 1.98** | 1.90** | 3.34** | 2.27** | 1.60** | 1.25** | 2.18** |
| | (0.08) | (0.11) | (0.10) | (0.13) | (0.12) | (0.06) | (0.26) | (0.17) |
| $\theta_2$ | 2.40** | 2.39** | 1.54** | 1.80** | 2.22** | 1.69** | 2.80** | 2.17** |
| | (0.14) | (0.18) | (0.16) | (0.18) | (0.16) | (0.08) | (0.38) | (0.25) |
| $\lvert\theta_2-2\rvert - \lvert\theta_1-2\rvert$ | −0.01 | 0.36 | 0.36 | −1.14 | −0.06 | −0.09 | 0.05 | −0.02 |
| $\theta_3$ | 0.04** | 0.05** | 0.04** | 0.07** | 0.09** | 0.06** | 0.03 | 0.00 |
| | (0.01) | (0.02) | (0.01) | (0.02) | (0.01) | (0.01) | (0.02) | (0.04) |
| $\theta_3/(1-\theta_5)$ | 0.10 | 0.10 | 0.09 | 0.13 | 0.13 | 0.06 | 0.06 | 0.11 |
| $\theta_4$ | 0.00 | 0.08 | 0.01 | 0.01 | −0.02 | 0.20** | 0.12** | 0.00 |
| | (0.01) | (0.04) | (0.04) | (0.02) | (0.06) | (0.03) | (0.06) | (0.01) |
| $\theta_4/(1-\theta_5)$ | 0.01 | 0.17 | 0.01 | 0.02 | −0.03 | 0.21 | 0.25 | 0.03 |
| $\theta_5$ | 0.62** | 0.52** | 0.47** | 0.43** | 0.30** | 0.06 | 0.50 | 1.00** |
| | (0.09) | (0.16) | (0.17) | (0.16) | (0.12) | (0.17) | (0.25) | (0.37) |
| $R^2$ | 0.78 | 0.56 | 0.63 | 0.74 | 0.81 | 0.76 | 0.32 | 0.13 |

Notes: Significant at 5% and 1% denoted by * and **. Standard errors in parantheses.

## Table 3: Reduced Models

(estimation period: Mar-2001 to Dec-2012; quarterly data)

$$\pi_t = \theta_1 (1 - D) + \theta_2 D + \frac{\theta_3}{1 - \theta_5 D} \pi_t^{EN} + \frac{\theta_4}{1 - \theta_5 D} gap_t + \eta_t$$

### 3A: Whole Euro Area and Large $H_0$ Reduction

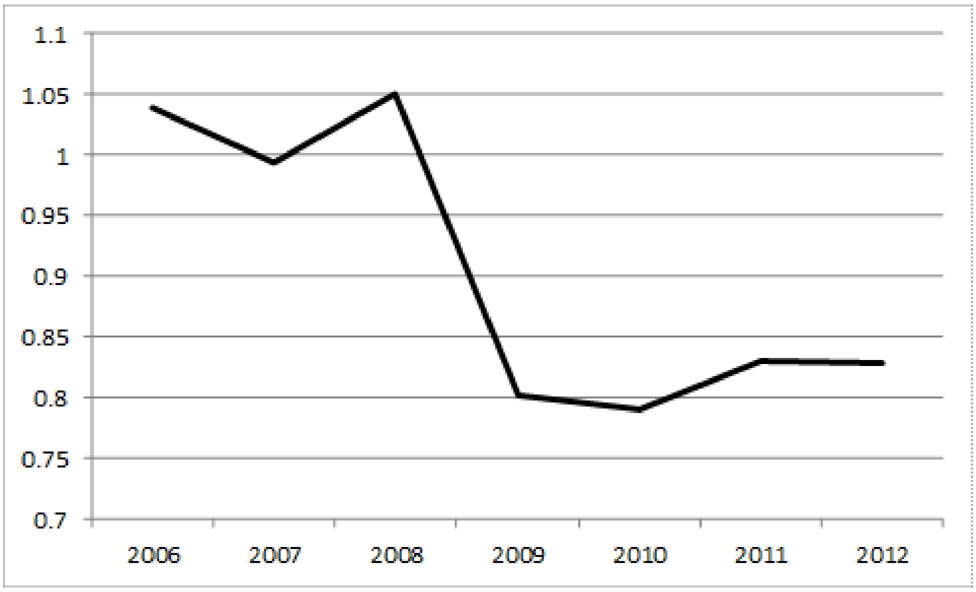
| | Euro Area | Ireland | Portugal | Italy | Greece | $\sigma_{GLS}/\sigma_{OLS}$ |
|---|---|---|---|---|---|---|
| $\theta_1$ | 2.25** | 3.84** | 3.05** | 2.35** | 3.35** | |
| | (0.28) | (0.28) | (0.16) | (0.17) | (0.19) | 2.4 |
| $\theta_2$ | 2.07** | 0.84 | 1.80** | 2.23** | 2.78** | |
| | (0.30) | (0.56) | (0.20) | (0.21) | (0.23) | 1.8 |
| $\|\theta_2-2\| - \|\theta_1-2\|$ | −0.18 | −0.68 | −0.85 | −0.13 | −0.57 | |
| $\theta_3$ | 0.05** | 0.00 | 0.00 | 0.01 | 0.01 | |
| | (0.02) | (0.04) | (0.04) | (0.01) | (0.01) | 0.9 |
| $\theta_3/(1-\theta_5)$ | 0.12 | 0.35 | 0.17 | 0.11 | 0.10 | |
| $\theta_4$ | 0.04 | 0.00 | | 0.00 | 0.01 | |
| | (0.07) | (0.02) | | (0.02) | (0.01) | 1.9 |
| $\theta_4/(1-\theta_5)$ | 0.09 | 0.14 | | 0.08 | 0.08 | |
| $\theta_5$ | 0.61** | 0.9999** | 0.9999** | 0.94** | 0.82** | |
| | (0.20) | | | (0.14) | (0.15) | 1.2 |
| $R^2$ | 0.73 | 0.75 | 0.59 | 0.66 | 0.73 | |

### 3B: Moderate $H_0$ Reduction

| | Luxembourg | Austria | France | Spain | Belgium | Germany | Finland | Netherlands |
|---|---|---|---|---|---|---|---|---|
| $\theta_1$ | 2.42** | 1.98** | 1.91** | 3.37** | 2.25** | 1.59** | 1.25** | 2.18** |
| | (0.16) | (0.28) | (0.19) | (0.25) | (0.27) | (0.11) | (0.67) | (0.44) |
| $\theta_2$ | 2.36** | 2.39** | 1.52** | 1.75** | 2.26** | 1.69** | 2.80** | 2.17** |
| | (0.17) | (0.31) | (0.20) | (0.29) | (0.29) | (0.13) | (0.56) | (0.41) |
| $\|\theta_2-2\| - \|\theta_1-2\|$ | −0.01 | 0.36 | 0.36 | −1.12 | 0.01 | −0.10 | 0.05 | −0.02 |
| $\theta_3$ | 0.04** | 0.05** | 0.05** | 0.08** | 0.09** | 0.06** | 0.03 | 0.00 |
| | (0.01) | (0.02) | (0.01) | (0.02) | (0.01) | (0.01) | (0.02) | (0.02) |
| $\theta_3/(1-\theta_5)$ | 0.10 | 0.10 | 0.09 | 0.13 | 0.12 | 0.06 | 0.06 | 0.11 |
| $\theta_4$ | | 0.08 | | | | 0.20** | 0.12** | 0.00 |
| | | (0.07) | | | | (0.04) | (0.09) | (0.03) |
| $\theta_4/(1-\theta_5)$ | | 0.17 | | | | 0.20 | 0.25 | 0.03 |
| $\theta_5$ | 0.62** | 0.52** | 0.46** | 0.39** | 0.30** | | 0.50 | 1.00** |
| | (0.09) | (0.20) | (0.15) | (0.16) | (0.11) | | (0.41) | (0.24) |
| $R^2$ | 0.78 | 0.56 | 0.63 | 0.77 | 0.82 | 0.76 | 0.32 | 0.13 |

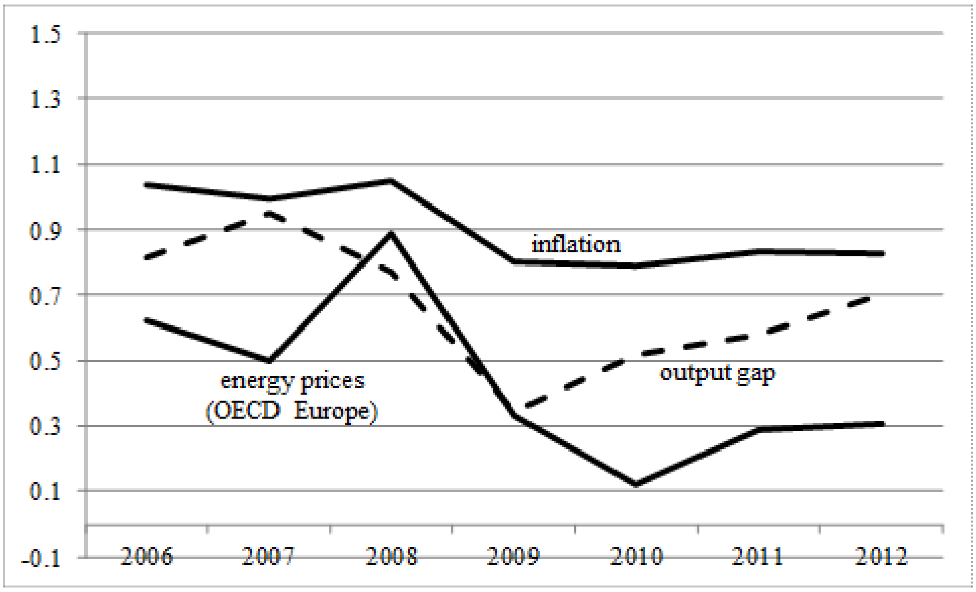
Notes: Significant at 5% and 1% denoted by * and **. Standard errors (in parantheses) are OLS for Ireland and Portugal and feasible GLS otherwise.

**Figure 1: Euro Area Inflation AR(1) Coefficients**



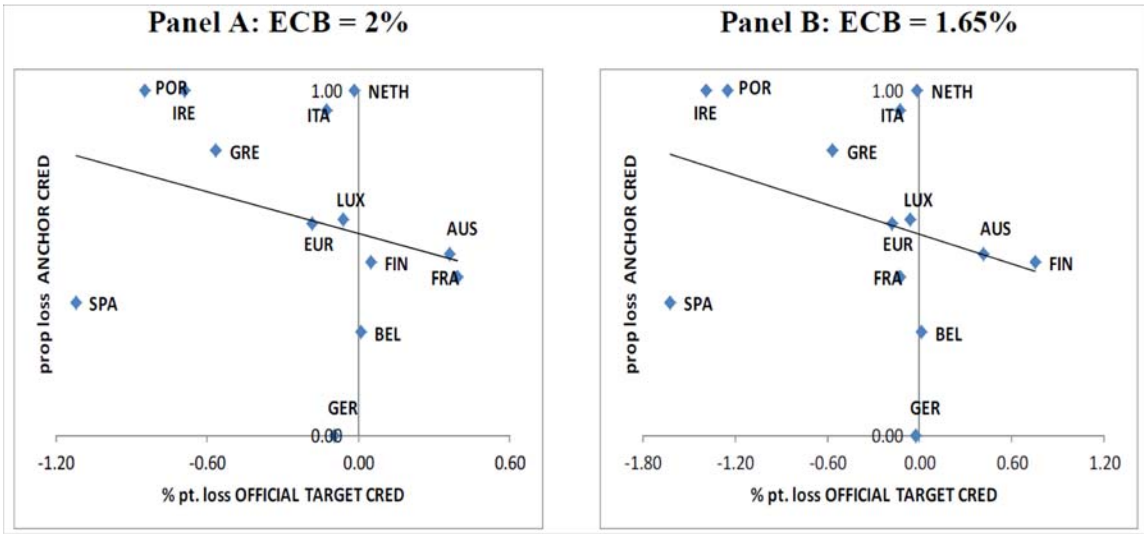

Notes: Year average inflation, 7-year rolling regression window, no constant. Datastream.

**Figure 2: Euro Area AR(1) Coefficients**




Notes: Year average variables, 7-year rolling regression window, no constant. Datastream.

**Figure 3: Independence of Anchoring Credibility and Official Target Credibility**




Notes: Regression output. % pt. loss OFFICIAL TARGET CRED is $|\theta_2 - 2| - |\theta_1 - 2|$ in the left panel and $|\theta_2 - 1.65| - |\theta_1 - 1.65|$ in the right panel, and positive values indicate a drift away from the ECB target. Prop loss ANCHOR CRED is the decline in $\beta$ estimated by $\theta_5$ in our econometric models.

**Figure 4: Two Examples of Probability Density Functions for $\alpha$**