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On the Macroeconomic Performance of the Euro Area ☆

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Abstract

This paper examines whether Euro Area countries would have faced a more favourable inflation output variability trade off without the Euro. We provide evidence that this claim is true for the periods of the Great Recession and the European Sovereign Debt Crisis. For the Euro Area as a whole, the deterioration of the trade off becomes insignificant with Draghi's *'whatever it takes'* announcement onwards. However, our more detailed analysis shows that the detrimental effect of the Euro is more severe for peripheral Euro Area countries and that ECB policies have been less effective for these countries. These findings point to structural differences among Euro Area countries as the explanation of the detrimental effect of the Euro as ECB policies can be denoted *'one size must fit all'* policies. We base our results on a novel empirical strategy that, consistent with monetary theory, models the joint determination of the variability of inflation and output conditional on structural supply shocks. Moreover, our findings are robust to potential endogeneity concerns related to adopting the Euro.

JEL Classification: C32, E50, F45.

Keywords: Euro Area, Monetary Policy, Difference-in-Differences.

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1. Introduction

The recent economic crises in Europe since 2007 have left member countries of the *Economic and Monetary Union* (EMU) in very heterogeneous economic conditions. While some members of the EMU by now experience modest growth and high employment, others are still in the process of recovery and suffer from unprecedented levels of unemployment. Clearly this experience is at odds with the goals of the EMU in general and the Euro Area in particular.[1]

A popular view on the economic events in the EMU during the last two decades is that, by construction of the EMU, structural heterogeneity, limited scope for fiscal policy and a union-wide monetary policy has amplified the effects of adverse shocks and lead to sub-optimal macroeconomic performance. For instance, an EMU country that adopted the Euro has chosen a monetary regime where monetary policy is delegated to the ECB (see, e.g., Ball, 2010). As a direct consequence, a Euro Area country can no longer offset country-specific shocks by a country-specific monetary policy. Moreover, the transmission of area-wide shocks may be heterogeneous due to structural differences among member countries. In consequence, ECB's monetary policy is believed to be sub-optimal: *'one size must fit all'* rather than *'one size fits all'* (Issing, 2001).

Consider the suggestive evidence in Figure 1 below.[2] Panels 1a and 1b compare the unconditional variances of inflation deviations from an estimated target and the output gap for Euro Area countries and non-Euro OECD countries over three periods: beginning of the Great Moderation until inception of the Euro, start of the Euro until the beginning of the Great Recession, and, the crisis period since then. The panels suggest that, according to these key indicators of macroeconomic performance, non-Euro OECD countries have been more successful in reducing output variability after the start of the Euro. Moreover, they have been more successful in stabilizing both inflation and output variability during the most recent period. However, instead of using suggestive evidence as provided in Figure 1, policy changes should be informed by thorough empirical analysis.

There is indeed a literature that tries to establish empirically if there exists a *'one size fits*

[1] According to Papademos (2009), proponents of the Euro Area have seen its adoption as a means to promoting trade and capital flows within the Euro Area with a subsequent increase in competition, the efficiency of resource allocation, and economic growth. A detailed description of the rationale behind the creation of the Euro Area is given by, for example, De Grauwe (2006).

[2] Note that these figures are based on calculations by the authors, which we detail in Section 3 below.

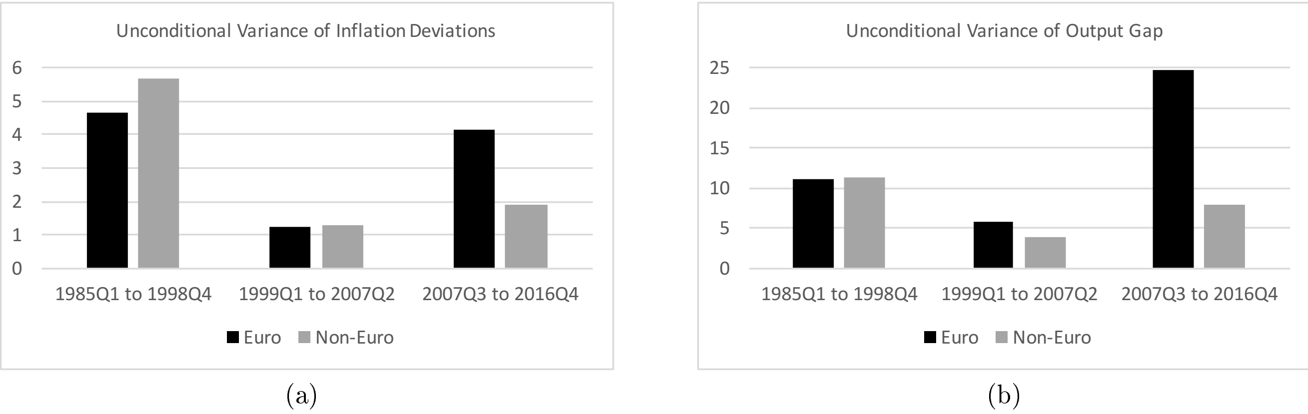


Figure 1: Comparison of the uncondional variances of inflation deviations from an estimated target and of the output gap between Euro Area and non-Euro OECD countries.

all' monetary policy for Euro Area countries. On the one hand, studies such as Peersman and Smets (2003), and Cecioni and Neri (2011) at the Euro Area level and Peersman (2004) in a cross-country set up do not find asymmetric effects due to monetary policy across Euro Area countries. On the other hand, Barigozzi et al. (2014) and Georgiadis (2015), in a cross-country empirical framework, show that the common monetary framework produces asymmetric effects driven by structural differences among Euro Area countries.[3]

Moreover, Ball (2010) finds that the Euro adoption had no significant effects on indicators of macroeconomic performance such as the level or variability of inflation or GDP. Nevertheless, the focus of Ball (2010) is on the effects of adopting inflation targeting (IT). In fact, the bulk of the empirical literature that quantifies the effect of a change in the monetary regime on macroeconomic performance focuses on IT. Two key themes in this literature stand out: first, this literature quantifies the effect of a change in the monetary regime on the moment of a single variable, e.g., the variability of inflation or GDP in isolation; second, a key challenge in this literature is endogeneity, as it is unanimously recognized that the choice of IT is affected by initial conditions.

However, through the lens of a theoretical IT framework the focus on a single variable does not seem fully appropriate. Measurement and comparison of macroeconomic performance in such a theoretical framework is routinely based on loss functions that involve inflation *and* output vari-

[3]The most recent data considered by this literature is from 2009. Therefore these papers do not take the European Sovereign Debt Crisis into account.

ability. Independently of whether one assumes optimal monetary policy or a simple Taylor (1993) rule, a central bank faces a long-run trade off between inflation and output variability. Moreover, the variability in these endogenous variables is jointly determined by structural supply shocks that move inflation and output in opposite directions.[4] Hence from the standpoint of a theoretical IT framework, evidence based on the variability of inflation *or* output in isolation appears problematic. For instance, if such research finds lower inflation variability for Euro Area countries compared to other countries, this might simply imply that the Euro Area countries are located on a different position of the inflation output variability trade off, but do not face an improved trade off due to the Euro adoption.[5]

Against this background, this paper seeks to examine the claim of whether the Euro Area countries would have faced a more favourable inflation output variability trade off without the Euro. To this end we propose a novel empirical research design that is coherent with the bulk of theoretical IT frameworks and tackles the endogeneity issue that comes along with a monetary regime change such as the adoption of the Euro.

Our research design involves several steps. First, we build a panel data set with observations on the unconditional variance of inflation deviations from target and of the output gap for twenty OECD countries over the sample period 1985 to 2016. We also estimate the variance of the structural supply shocks for each country by the help of a Structural Vector Autoregression (SVAR).

Second, as a clear novelty compared to the existing empirical literature on IT, we jointly condition the trade off between inflation and output gap variability on the variability of structural supply shocks using a set-up taken from the quantitative analysis of production processes (see, e.g., Kumbhakar, 2012). In brief, we interpret the variability of inflation and output gap of each country as jointly determined inputs and the variability of structural supply shocks of each country as an exogenous output, or, more generally as a shifter.

Third, in order to establish whether Euro adopters, on average, have been worse off by adopting this monetary regime, we first utilize a pure difference-in-differences (DiD) approach. However, as

[4] Taylor (1979) pioneered the empirical documentation of this long-run inflation output variability trade off based on the assumption of optimal monetary policy.

[5] In the discussion of Ball and Sheridan (2005), Stephen Cecchetti raised this issue in the context of the effects of IT on macroeconomic performance.

discussed in the IT literature (see, e.g., Ball, 2010), the choice of adopting the Euro might have been affected by initial conditions and therefore be subject to endogeneity. This can be interpreted as a violation of the parallel trends assumption between the treated (Euro Area countries after the adoption of the Euro) and the control group (the countries taken to construct the counterfactual), which is required by the DiD approach. In consequence, the estimates obtained via the DiD approach may be biased. Therefore, in order to corroborate our evidence, we also consider the lagged dependent variable (LDV) model (for a detailed discussion see Angrist and Pischke, 2009) that requires less stringent identification assumptions and controls for potential endogeneity.

We find that the Euro adoption worsened the macroeconomic performance of Euro countries on average. Moreover, when we account for the possibility that the effects of the Euro may vary over time, we find that the adoption of the Euro on average worsened macroeconomic performance only in the periods of the Financial Crisis and the European Sovereign Debt Crisis. Furthermore, the detrimental effect of the Euro ceases after 2012. This timing corresponds to Mario Draghi's announcement about *'whatever it takes to preserve the Euro'* and the ECB's announcements of more intense and additional unconventional policies such as the outright monetary transactions (OMTs) and the expanded asset purchase programme (EAPP) and its subsequent implementation (see, e.g., Fawley and Neely, 2013, for an overview on the events). Therefore we interpret our findings as evidence that these announcements have been effective in reducing inflation and output variability for the Euro Area as a whole. These measures may have credibly signalled that the ECB was going to act as *'buyer of last resort'* (Acharya et al., 2017b), i.e., what De Grauwe (2012) describes as a lender of last resort in the government bond markets.

A more detailed analysis shows that the detrimental effect of the Euro on macroeconomic performance in periods of crises is more severe in peripheral countries. In addition, while the detrimental effect of the Euro becomes insignificant for the core of the Euro Area after the above mentioned policy interactions, it remains significant for the peripheral countries. These findings suggest that structural differences among Euro Area countries explain the detrimental effect of the Euro and that monetary policy in the Euro Area is best characterized as a *'one size must fit all'* policy.

Our approach is related to the literature that uses a Taylor (1979) curve to evaluate macroeconomic performance. Cecchetti et al. (2006) evaluate macroeconomic performance for single countries, based on a comparison between two different subsamples of the radial distance of actual unconditional variances from the optimal variances implied by the Taylor (1979) curve. Mishkin and Schmidt-Hebbel (2007) extend the approach used by Cecchetti et al. (2006) to a multi-country level, utilizing a dynamic panel with fixed effects estimated through GMM in order to infer on the macroeconomic implications of IT. However, as illustrated by Angrist and Pischke (2009), identification in a panel with lagged variables and fixed effects is problematic when the policy is endogenous to initial conditions. Olson and Enders (2012) have also made use of a Taylor (1979) curve framework, but use a different metric to measure the distance between observed and optimal variances compared to Cecchetti et al. (2006). However their analysis is conducted exclusively for the US.

There are two main differences compared to the related literature. First, our research design does not require explicit assumptions on whether monetary policy in the examined countries is best described by optimal monetary policy or by a simple Taylor (1993) rule. Rather the opposite, our framework encompasses both the inflation output variability trade offs implied by optimal monetary policies *and* by simple Taylor (1993) rules. Second, in our empirical analysis we jointly model the variability of inflation and output gap determined by an exogenous supply shock.

The remainder of the paper is organized as follows. In Section 2 we outline the theoretical framework on which we base our empirical strategy. Section 3 describes the empirical implementation and the data in use. Section 4 presents the main results based on the DiD, while Section 5 contains our extensive robustness analyses. Section 6 concludes.

2. Theoretical Framework

We start out by briefly elaborating the theoretical inflation output variability trade off in the context of the New Keynesian model. We take the latter as our preferred benchmark for measuring and comparing macroeconomic performance. We argue that the inflation output variability trade off exists for optimal discretionary monetary policy as well as monetary policy described by a simple Taylor (1993) rule. Thereafter, we develop a theory-based empirical framework to estimate the inflation output variability trade offs in economies independent of any assumption about the type of monetary policy.

Frameworks for measuring and comparing macroeconomic performance in monetary theory are routinely based on loss functions. A popular approach is to consider *ad hoc* period loss functions such as

$$\mathcal{L} = \pi_t^2 + \omega_x x_t^2, \tag{1}$$

where π_t^2 denotes the deviation of inflation from an inflation target and x_t denotes the deviation of the output gap from steady state. Parameter ω_x captures the central bank's preference for output gap stabilization relative to inflation stabilization.

Moreover, assume that the aggregate economy is best approximated by a standard New Keynesian model under the rational expectations hypothesis, i.e.,

$$x_t = E_t x_{t+1} - \sigma^{-1} (i_t - E_t \pi_{t+1}) + g_t \tag{2}$$

$$\pi_t = \beta E_t \pi_{t+1} + \lambda x_t + e_t. \tag{3}$$

In this model, i_t denotes the nominal interest rate controlled by the central bank. β and σ are structural parameters, λ is a composite term comprising several structural parameters. g_t denotes an exogenous demand disturbance and e_t denotes an exogenous supply disturbance. In addition, the shocks are assumed to be $g_t \sim \text{iid}(0, \sigma_g^2)$, and $e_t \sim \text{iid}(0, \sigma_e^2)$.[6]

Now consider optimal monetary policy under discretion (as elaborated in Clarida et al., 1999). The central bank minimizes (1) subject to (3) in each period. The first-order necessary condition is

$$\pi_t = -(\omega_x / \lambda) x_t. \tag{4}$$

One can show that, under this policy, the model implies an inflation output variability trade off as first developed in Taylor (1979). In particular, solving the model for given parameters implies a

[6]See Galí (2015) or Woodford (2003) for more details on this model. In the theoretical literature e_t is usually denoted a cost-push shock. Notice that allowing for auto-correlation in the exogenous shocks would not alter any conclusion.

minimum state variable solution

$$\pi_t = a_\pi e_t \tag{5}$$

$$x_t = a_x e_t, \tag{6}$$

where $a_\pi \equiv \omega_x/(\omega_x + \lambda^2)$ and $a_x \equiv -\lambda/(\omega_x + \lambda^2)$. This solution implies the following long-run relationships in unconditional variances

$$\sigma^2_{\pi,*} = a^2_\pi \sigma^2_e \tag{7}$$

$$\sigma^2_{x,*} = a^2_x \sigma^2_e. \tag{8}$$

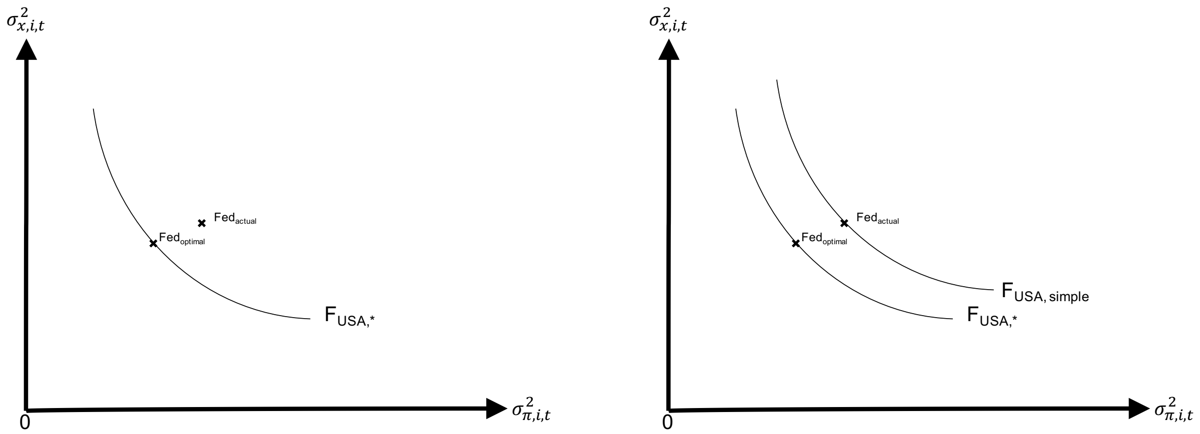
In short, (7) to (8) indicate that both the optimal variances of inflation and output gap depend on the variance of the supply shock. Moreover, the larger the central banks' preference for output gap stabilization, ω_x, the lower $\sigma^2_{x,*}$ and the larger $\sigma^2_{\pi,*}$.

Figure 2a depicts this concept for the case of the US. The variation of ω_x allows one to depict the Taylor (1979) curve, $\mathrm{F}_{\mathrm{USA}}$, which can be thought of as an efficient frontier. The idea is that country-specific supply shocks hit an economy and, given the structure of the economy, create a domestic trade off between inflation and output variability. *In theory*, a domestic central bank, e.g., the Federal Reserve (Fed), can conduct optimal policy and locate the economy on the trade off, see Figure 2a.

When it comes to measuring the macroeconomic performance, *in practice*, the efficient frontier can be estimated, for instance, via a parsimonious reduced form VAR with a supply and demand equation, including reduced form shocks and states an approximation of optimal monetary policy.[7] Actual observed variability in inflation and the output gap will routinely indicate that the economy is to the right of an estimated efficient inflation output variability trade off. Therefore a central bank's monetary policy can be classified as sub-optimal (see also Figure 2a).[8]

[7] *'Three or four estimated equations are crucial for the Taylor economic model but the economy as a whole is determined by millions of equations. At most, we could hope to get a rough picture of it.'*(Friedman, 2010, p.116).

[8] A frequently applied approach to measuring macroeconomic performance is based on the distance of actual variability in inflation and output gap, $\sigma^2_{\pi,\mathrm{USA},t}$ and $\sigma^2_{x,\mathrm{USA},t}$ from the model-implied optimal trade off, $\sigma^2_{\pi,\mathrm{USA},*}$ and $\sigma^2_{x,\mathrm{USA},*}$, at certain points in time (see, e.g., Cecchetti et al., 2006). One can then repeat such an exercise for a panel



(a) Stylized optimal trade off for the USA

(b) Stylized optimal and sub-optimal trade off for the USA

Figure 2: Illustration of the inflation output variability trade off

However, what if monetary policy in a country is not appropriately described by optimal monetary policy, but may be better approximated by a simple Taylor (1993) rule? The latter is a flexible way of describing monetary policies in theory. For instance, such simple interest rate rules can also involve terms for observed monetary policy inertia, feedback to real economic activity or exchange rates. For many countries, such rules may be a more suitable description of monetary policy.

As we discuss next, under such a Taylor (1993) rule, there is still an inflation output variability trade off. However, this trade off is neither optimal, nor is it captured by the approach pursued in Cecchetti et al. (2006) and related studies, which builds explicitly on the assumption of optimal monetary policy.

For instance, consider the simple interest rate rule[9]

$$i_t = \phi_\pi \pi_t, \qquad \phi_\pi > 0. \tag{9}$$

For simplicity, assume $g_t = 0$ for all t (i.e., we abstract from demand shocks).[10] Then, it can be

of countries and compare measures of macroeconomic performance for different countries at different points in time.

[9]Notice that the same arguments holds, if we would consider a rule that also involves feedback to the output gap, i.e., $i_t = \phi_\pi \pi_t + \phi_x x_t$, $\phi_x > 0$.

[10]This is solely for ease of exposition.

easily verified that the model (2) and (3) under policy rule (9) has the solution

$$\pi_t = b_\pi e_t \tag{10}$$

$$x_t = b_x e_t, \tag{11}$$

where $b_\pi \equiv (1 + \sigma^{-1}\phi_\pi\lambda)^{-1}$ and $b_x \equiv -\sigma^{-1}\phi_\pi/(1 + \sigma^{-1}\phi_\pi\lambda)$. This solution implies the following long-run relationships in unconditional variances

$$\sigma_\pi^2 = b_\pi^2 \sigma_e^2 \tag{12}$$

$$\sigma_x^2 = b_x^2 \sigma_e^2. \tag{13}$$

Thus, similar to (7) to (8), (12) to (13) show that both the variances of inflation and output gap depend on the variance of the supply shock. In addition, the smaller the central bank's coefficient on inflation, $\phi_\pi \in (1, \infty]$, the lower σ_x^2 and the larger σ_π^2.[11] Thus, there exists an inflation output variability trade off, although the latter is based on the simple interest rate rule (9). The challenge is then to develop an empirical specification that is flexible enough to encompass both the trade offs implied by optimal and simple monetary policy.

In this paper, we propose an empirical framework to tackle this challenge. We make the assumption that the inflation output variability trade off also exists independent of the particular monetary policy in a certain country. Coming back to the example of the Fed in Figure 2b, actual variances observed for the USA may be the result of optimal or sub-optimal monetary policy, but a trade off exists at any rate.

Our approach does not suppose the existence of an efficient inflation output variability trade off, but solely requires to assume that, consistent with the above theory, an exogenous supply shock determines both the variability of inflation and output. Moreover, a stronger central bank preference for inflation stabilization, i.e., lower ω_x, or, a higher coefficient on inflation in the interest rate rule, ϕ_π, implies a higher variability of output and a lower variability of inflation. As we detail below, our empirical specification models the relationship between the variance of the exogenous

[11] It is well known that this model does not have a determinate rational expectations equilibrium for $\phi_\pi \leq 1$.

structural supply shock and the endogenous variances of inflation and output gap consistently with the theoretical considerations above by positing a given functional form for the inflation output variability trade off. In sum, for observations for one or more countries at different points in time, our specification allows us to fit a curve as depicted in Figure 3.

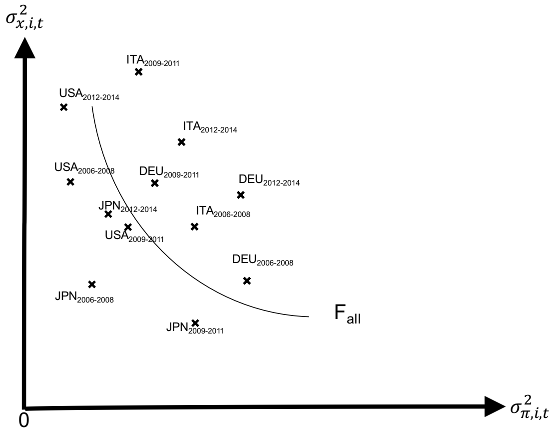


Figure 3: Estimation of the trade off in the sample by a translog transformation function

Our empirical strategy builds on tools developed in the quantitative analysis of production. In particular, we use a specification based on a translog transformation function (TTF). Kumbhakar (2012, 2013) shows that the input-oriented interpretation of a TTF can be used to model the determination of one or more endogenous production inputs, for exogenous production outputs, and technology.

We use the input-oriented TTF to model the joint determination of the endogenous variances of inflation and output gap (i.e., the two inputs in the TTF framework), for a given variance of an exogenous structural supply shock (i.e., a single output or shifter in the TTF framework) and a given monetary policy. Thus, the input-oriented TTF provides us with a functional form that captures the basic characteristics of an inflation output variability trade off.

Formally, one assumes that the relationship between the supply shock (as the single output, y) and the variances of inflation and output gap (as the $K = 2$ inputs, z) can be described by

$$Af(y, z) = 1 \qquad \Big| \ln(\cdot) \tag{14}$$

$$\Leftrightarrow \ln(A) + \ln(f(y, z)) = 0, \tag{15}$$

where we have one output y and K inputs z. Moreover, A captures factors that affect the TTF neutrally. We will be more specific about the assumptions further below.

Next, we assume a *translog* functional form, i.e.,

$$\begin{aligned}\ln(f(y,z)) = {} & \alpha_y \ln(y) + \frac{1}{2}\alpha_{yy}\ln(y)^2 \\ & + \sum_k \beta_k \ln(z_k) + \frac{1}{2}\sum_k \sum_l \beta_{k,l}\ln(z_k) \times \ln(z_l) \\ & + \sum_k \gamma_{k,y}\ln(y) \times \ln(z_k),\end{aligned} \tag{16}$$

where the following symmetry is imposed: $\beta_{k,l} = \beta_{l,k}$. Equation (16) requires $K+2$ additional identification, or, normalization restrictions. As discussed in Kumbhakar (2012), it is possible to impose the restrictions such that a single equation framework emerges that allows for simultaneous estimation of more than one endogenous input (e.g., input-oriented) or output (e.g., output-oriented).

Since, in the case of the inflation output variability trade off, we have simultaneous endogeneity of σ_π^2 and σ_x^2, while σ_e^2 is exogenous, we can consider the former two variances as inputs, while the latter variance is the output. Therefore, we adopt a normalization with respect to an input. This gives rise to an input-oriented TTF. Following Kumbhakar (2012), we rewrite (16) as

$$\begin{aligned}\ln(f(y,z)) = {} & \alpha_y \ln(y) + \frac{1}{2}\alpha_{yy}\ln(y)^2 \\ & + \sum_k \beta_k \ln(z_k/z_1) + \frac{1}{2}\sum_k \sum_l \beta_{k,l}\ln(z_k/z_1) \times \ln(z_l/z_1) \\ & + \sum_k \gamma_{k,y}\ln(y) \times \ln(z_k/z_1) + \Upsilon,\end{aligned} \tag{17}$$

where each input k has to be combined with the remaining inputs l as described in this equation. Υ is a composite term that follows from writing the second and third line in expression (16) in ratios (see, e.g., Kumbhakar, 2012, for the details).

Next we impose the normalization restrictions, $\sum_k \beta_k = -1$, $\sum_k \beta_{k,l} = 0 \; \forall k$, and, $\sum_k \gamma_{k,y} = 0$.[12] As a consequence, the composite term Υ is eliminated and we obtain the input-oriented TTF that

[12] The normalization restrictions imply homogeneity, symmetry and monotonicity properties of the TTF.

we use as our empirical specification

$$
\begin{aligned}
-\ln(z_1) = {} & \alpha_0 + \alpha_y \ln(y) + \frac{1}{2}\alpha_{yy}\ln(y)^2 \\
& + \sum_{k=2} \beta_k \ln(z_k/z_1) + \frac{1}{2}\sum_{k=2}\sum_{l=2} \beta_{k,l} \ln(z_k/z_1) \times \ln(z_l/z_1) \\
& + \sum_{k=2} \gamma_{k,y} \ln(y) \times \ln(z_k/z_1) + v,
\end{aligned} \tag{18}
$$

where $\ln(A) = \alpha_0 + v$. In this case we normalize our function on z_1. We would get exactly the same econometric results by normalizing the function on z_k. In the particular case of the inflation output variability trade off, we have $y = \sigma_e^2$, $z_1 = \sigma_x^2$ and $z_2 = \sigma_\pi^2$. Therefore our empirical specification is

$$
\begin{aligned}
-\ln(\sigma_x^2) = {} & \alpha_0 + \alpha_e \ln(\sigma_e^2) + (1/2)\alpha_{ee} \ln(\sigma_e^2)^2 \\
& + \beta_2 \ln(\sigma_\pi^2/\sigma_x^2) + (1/2)\beta_{2,1}\left[\ln(\sigma_\pi^2/\sigma_x^2)\right]^2 \\
& + \gamma_{2,e} \ln(\sigma_e^2) \times \ln(\sigma_\pi^2/\sigma_x^2) + v.
\end{aligned} \tag{19}
$$

The empirical specification and estimation of (19) allows one to empirically test whether a trade off between the variability of output and inflation actually exists in the data. Conditionally on the existence of this trade off, estimation of (19) uses the statistical information on macroeconomic performance more efficiently than the usual estimates based on either inflation *or* output gap variability alone. The reason is that in this regression set-up one can use the variability of the output gap (respectively inflation) to model the variability of inflation (respectively output gap). Besides, in (19) the macroeconomic performance of different countries is gauged while controlling for country-specific supply shocks, which cannot be offset by the monetary authorities.

3. Empirical Implementation

Our goal is to estimate the inflation output variability trade off for a number of countries $i = 1, ..., N$ over time $t = 1, ..., T$ based on (19). Moreover, we want to examine the effect of the Euro on macroeconomic performance of Euro Area countries after they have adopted the Euro relative to a comparable set of countries without the Euro. However, the empirical implementation of (19) is not that obvious. In principle, one can estimate the inflation output variability trade off by a two-way fixed-effect model, i.e.,

$$
\begin{aligned}
-\ln(\sigma^2_{x,i,t}) &= \alpha_e \ln(\sigma^2_{e,i,t}) + (1/2)\alpha_{ee} \ln(\sigma^2_{e,i,t})^2 \\
&+ \beta_2 \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) + (1/2)\beta_{2,1} \left[\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})\right]^2 \\
&+ \gamma_{2,e} \ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) + \alpha_i + \delta_t + \varepsilon_{i,t},
\end{aligned} \tag{20}
$$

where we assume $v_{i,t} = \alpha_i + \delta_t + \varepsilon_{i,t}$. $\varepsilon_{i,t}$ is a stochastic error term, α_i a fixed effect aimed at capturing unobserved time invariant country factors and δ_t can be thought of as a flexible (nonlinear) time trend, i.e., a common unobserved factor (shock) affecting all countries by the same amount (for further details see Smith and Fuertes, 2016).

Immediately two major challenges emerge with regard to our research question. First, consistently with theoretical inflation output variability trade off, we require observations of the variances of inflation deviation from target, of the output gap, and of the structural supply shock. This in turn implies some de-trending of the inflation and output data and the construction of structural supply shocks for each country. Second, we need to develop an identification strategy for the effect of the Euro on macroeconomic performance. We address these issues below.

3.1. Data

Our dataset includes quarterly observations of the consumer price index and real GDP for $N = 20$ member countries of the Organization for Economic Cooperation and Development (OECD) over the period 1984Q1-2016Q4. The source is the OECD database. The countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, UK and US. As in Ball (2010), we have chosen all countries with population above one million that were members of the OECD in 1985 (beginning of the Great Moderation). Inflation is calculated as the year-to-year percentage difference of the consumer price index (all items).

Moreover, further below in the estimation of the structural supply shocks, we also require a measure of monetary policy. For the nominal interest rate, we rely on the following indicators.

For the USA we use the Shadow Rate developed by Wu and Xia (2016) since the beginning of the sample. While in normal times it resembles, the effective federal funds rate, at the zero interest-rate lower bound it is aimed to capture unconventional policies adopted by the central bank. Next, for the UK, we use the treasury bill rate up to 2004Q3, then using the corresponding shadow rate developed by Wu and Xia (2016). Finally, for the Euro Area countries, we use the money market rate until 1998Q4. Up to 2004Q3, we use the common ECB refinancing rate. Thereafter we use the Euro Shadow Rate developed by Wu and Xia (2016). For the non-Euro OECD countries other than the UK or the US, we use the money market rate up to 2004Q3, and then, since afterwards no shadow rate is available, we use a quarterly measure of the overnight bank rate. Ciccarelli et al. (2013) make a similar choice and use the EONIA for Euro Area countries in recent years, since before the crisis it was indistinguishable from the main refinancing operations (MRO) rate, while after the adoption of unconventional policies it dropped below the MRO rate, being more sensitive to dynamics due to the unconventional policies.

3.2. Computing Variances of Output Gap and Inflation Deviation from Target

Our measure for the output gap is the difference in the log of real gross domestic product from its trend value computed through the Hamilton (2017) filter, while inflation is calculated as the year-to-year percentage difference of the consumer price index (all items) minus its trend value computed through the Hamilton (2017) filter. Here we assume that the filter-measured trend is able to capture explicit or implicit inflation target of the countries considered. This choice is motivated by the fact that we do not observe an explicit target in all countries of the sample. Moreover, in the short run, for instance, during the recent crisis, central banks may deliberately tolerate a deviation from the explicit target, which is a long-run concept by definition. Put differently, there may be an implicit short-run target different from the explicit long-run target and the implicit target represented by the filter-measured trend may provide a better representation of it. This approach is common in the literature (see, for instance, Olson and Enders (2012), where a Hodrick and Prescott (1997) (HP) filter is adopted). Our choice of the Hamilton (2017) filter with respect to the more traditional HP filter is motivated by the considerations made by Hamilton (2017), who shows that the persistence present in the cyclical part of the HP filter has nothing to do with the underlying data generating process.

Starting from the closed form solution of the HP minimization problem, with quarterly data and t more than 15 years from the start or end of a sample the cyclical component $c_t = \tilde{y}_t - \tilde{g}_t^*$ can be approximated by

$$c_t = \frac{\tilde{\lambda}(1-L)^4}{F(L)}\tilde{y}_{t+2}, \tag{21}$$

with $F(L) = 1+\tilde{\lambda}\left(1-L^{-1}\right)^2(1-L)^2$, which shows that the HP filter might be expected to produce a stationary series as long as the fourth differences of the original series is stationary, since it takes the fourth difference of $\tilde{y}_{t+2}$ and applies the operator $[F(L)]^{-1}$. However, De Jong and Sakarya (2016) have shown that there might still be some non stationarity coming from the beginning or end of the sample, while Phillips and Jin (2015) show that even with an I(1) series the HP filter might not be able to remove the trend. Cogley and Nason (1995) show that for a random walk $\tilde{y}_t = \tilde{y}_{t-1} + \varepsilon_t$ (where first differences are completely unpredictable) equation (21) near the middle of the sample can be approximated as

$$c_t = \frac{\tilde{\lambda}(1-L)^3}{F(L)}\varepsilon_{t+2}, \tag{22}$$

with $\tilde{\lambda} = 1600$ (as is usual for quarterly data) the HP filter will produce an ε_t, which is random, and a cycle which is predictable (as a function of past and future observations). The persistence of the cycle is due to the fact that the coefficients in $[F(L)]^{-1}$ depend on the value chosen for $\tilde{\lambda}$ and may not reflect any features of the data generating process (see Hamilton, 2017, for further details). Even when the properties of the data are such to make the HP filter optimal, Hamilton (2017) shows that the estimated value of λ is always much lower than the value of 1600 customarily used for quarterly data.

Hamilton (2017) proposes a new filter hinging on the forecast made two years in advance of $\tilde{y}_{t+h}$ on the basis of p current and past values. The suggested cycle (at time h), $\tilde{v}_{t+h}$, derived from the population linear projection, in case of quarterly data would be

$$\tilde{y}_{t+h} = \zeta_0 + \zeta_1 \tilde{y}_t + \zeta_2 \tilde{y}_{t-1} + \zeta_3 \tilde{y}_{t-2} + \zeta_4 \tilde{y}_{t-3} + \tilde{v}_{t+h}. \tag{23}$$

For quarterly data, we would have $h = 8$ and Hamilton (2017) suggests to take $p = 4$ in this case, interpreting the resulting forecast error $\tilde{v}_{t+h}$ as the cycle at time $t+h$. Den Haan (2000) has shown that such a forecast error would be stationary for a large class of nonstationary processes. Most importantly, Hamilton (2017) shows that the primary reason of wrongly predicting most of the macroeconomic and financial variables two years in advance would be due to cyclical factors.

As discussed by Hamilton (2017), it is not necessary to know the nature of the nonstationarity and to have the correct model for forecasting the time series. Even in the case of an I(2) variable, where the series should be differenced twice ($d = 2$) in order to have a stationary process, Hamilton (2017) shows that with $p > d$, equation (23) will use two coefficients to ensure that the residual are stationary and the remaining coefficients will be determined by the parameters that characterize the stationary variable $\tilde{v}_{t+h}$. This filter can deal with the same situation considered for the HP filter (in case we need to take the fourth difference of the time series to obtain stationarity), but without any of its previously discussed drawbacks, in particular without the introduction of spurious persistence in the cyclical component.

3.3. Estimation of the Structural Supply Shock

In order to derive our structural supply shock we consider a vector autoregression with exogenous variables (VARX), whose reduced form of order (p, q) can be represented by

$$\mathcal{Y}_t = C + \sum_{i=1}^{p} A_i \mathcal{Y}_{t-i} + \sum_{k=1}^{q} \Phi_k \mathcal{X}_{t-k} + u_t, \tag{24}$$

where C is a 3×1 vector of constant terms, $\mathcal{Y}_t$ is a 3×1 vector of endogenous variables including a measure of the output gap, inflation (as difference from target) and nominal interest rate, $\mathcal{X}_t$ is a 3×1 vector of exogenous variables including USA output gap, inflation (as difference from target) and a nominal interest rate, aimed at capturing world macroeconomic stance and u_t is a vector of reduced-form disturbances with $E[u_t] = 0$ and $E[u_t u_t'] = \Sigma_u$. We consider all countries with the exception of USA as open economies (for a more detailed discussion of this choice see Favero and Giavazzi (2008)). Identification of the VAR in equation (24), requires to impose enough restrictions to decompose u_t in order to obtain economically meaningful structural innovations. A matrix A is required such that $Ae_t = u_t$, where e_t represent the vector of structural shocks. At least $n \times (n-1)/2$ restrictions on A are required to obtain identification. Usually identification is obtained via Cholesky factorization of Σ_u

We adopt a sign restrictions identification strategy. Uhlig (2005) amongst others, shows how to obtain identification of the above VAR imposing sign restrictions on a (sub)set of the variables responses to shocks. The main advantage of this procedure is that only a minimum amount of economically meaningful sign restrictions are required in order to identify the structural shocks. In case of a single shock, Uhlig shows that any impulse vector a can be recovered if there is an n-dimensional vector q of unit length such that $a = \tilde{A}q$, where $\tilde{A}$ is the Cholesky factor of Σ_u.

More precisely, starting with estimation of the above reduced form model using ordinary least squares, identification of a single shock by sign restrictions (as in our case) can be obtained as follows:

1. derive the impulse-responses for the n variables corresponding to a given impulse vector a_j up to period f on which sign restrictions are intended to be imposed;
2. draw an n-dimensional q vector of independent $N(0,1)$ and divide it by its norm, obtaining a candidate draw q from which an impulse vector $a_j = Aq$ can be derived for then calculating the corresponding impulse responses;
3. if the resulting impulse responses meet the sign restrictions imposed accept the draw, otherwise discard it;
4. repeat 2 and 3 until a desired number of accepted draws is obtained.

Generally, the median of accepted draws is considered as the central estimate of interest for impulse responses. However, as shown by Fry and Pagan (2011), the median responses may combine information from several identification schemes (i.e., different q's). In order to overcome this problem, Fry and Pagan (2011) suggest the median target (MT) method, taking the responses to a shock, which are overall as close as possible to the median responses, while imposing that the responses are generated from a single identifying vector q. Adopting the MT method, in order to derive our structural supply shock of interest, we impose that for the first four quarters, there is a positive inflation and a negative output response based on 50,000 accepted draws.[13]

3.4. From Cross-Section to Panel Data

We create a panel dataset of dimension $N \times T = 20 \times 9 = 180$. Indeed, while our effective sample contains 128 quarterly observations from 1985Q1 to 2016Q4, we have divided the sample in $T = 9$ periods as can be seen from Table 1. We did this in order to compute variances for inflation deviations from target, output gap and supply shocks over a sufficiently long time window. Our chosen sub-periods seem to satisfy the need to identify different interesting economic episodes as shown in Table 1.

3.5. Identification of the Effect of the Euro on Macroeconomic Performance

As previously stated, our baseline specification is a pure DiD approach, where our aim is to infer whether the adoption of the Euro has on average improved the macroeconomic performance for the Euro Area countries. We augment the two-way fixed effect model (20) by a dummy $\mathcal{E}_{i,t}$, which is equal to zero for all countries and one for countries when the policy is implemented

$$\begin{aligned} -\ln(\sigma^2_{x,i,t}) = {} & \beta_{\mathcal{E}}\mathcal{E}_{i,t} + \alpha_e \ln(\sigma^2_{e,i,t}) + (1/2)\alpha_{ee}\ln(\sigma^2_{e,i,t})^2 \\ & + \beta_2 \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) + (1/2)\beta_{2,1}\left[\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})\right]^2 \\ & + \gamma_{2,e}\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) + \alpha_i + \delta_t + \varepsilon_{i,t}, \end{aligned} \tag{25}$$

where $\beta_{\mathcal{E}}$ represents our estimate of interest as it captures the impact of the Euro adoption on macroeconomic performance.

Equation (25) can be written more compactly as

$$Y_{i,t} = \alpha_i + \delta_t + \beta_{\mathcal{E}}\mathcal{E}_{i,t} + \Omega X_{i,t} + \varepsilon_{i,t}, \tag{26}$$

[13]For robustness purpose we have also derived structural supply shocks on a restricted sample ending in 2007Q4 for sign restrictions imposed for one and four quarters. The supply shocks derived for the two different samples show a correlation for the common period (1985Q1 to 2007Q4) that range between 0.83 (when we impose sign restrictions on four quarters) and 0.90 (when we impose sign restrictions on one quarter). From our point of view, this suggests that potential structural breaks following the Great Recession should not be particularly worrying.

where $X_{i,t}$ contains all the right hand side variables shown in equation (25) and Ω the corresponding vector of parameters to be estimated.

The key assumption here is that of parallel trends, i.e., the average outcome for treated and control group would have followed the same trend without treatment. With non random assignment, it must follow that

$$E(Y_{0,i,t} \mid \alpha_i, \delta_t,\ X_{i,t},\ \mathcal{E}_{i,t}) = E(Y_{0,i,t} \mid \alpha_i, \delta_t,\ X_{i,t}). \tag{27}$$

Equation (27) is the conditional independence assumption (CIA) in case of parallel trends. It states that unobserved confounders are fully captured with the two way fixed effect model. In other words, it says that assignment is determined only by the country plus time fixed effects and control variables $X_{i,t}$, where the latter cannot be influenced by the policy. $\beta_{\mathcal{E}}$ represents the average treatment on the treated (ATT).[14]

However, often, to suppose that the effects of unobserved confounders is fully controlled by the two way fixed effect model may be restrictive. Ashenfelter (1978) and Ashenfelter and Card (1985) have for example found, in a labour market context, that participants to a government-sponsored training programme have earning histories that have a pre-program-dip. Indeed, the literature aimed at evaluating the change in macroeconomic performance due to IT almost unanimously consider its adoption as endogenous. In particular, policy assignment is seen as dictated by previous economic conditions (see Ball, 2010, for an interesting survey of the literature). Since the Euro adoption might not be exempt from such considerations, we also consider an alternative specification in order to corroborate our results.

We therefore consider the LDV model, where it is possible to not make the parallel trends assumption required in the DiD setting and at the same time to control for past outcomes. It can generally be specified as follows

$$Y_{i,t} = \alpha_i + \delta_t + \beta_{\mathcal{E}}\mathcal{E}_{i,t} + \Omega X_{i,t} + \theta Y_{i,t-h} + \varepsilon_{i,t}, \tag{28}$$

[14] The non-random DiD identifies the ATT, see, e.g., Athey and Imbens (2006).

where $X_{i,t}$ contains all the right hand side variables shown in equation (25) and Ω the corresponding vector of parameters to be estimated. The sample in this case would start at the date of the Euro adoption denoted as T_0. Moreover, note that this is not a dynamic model, since we are conditioning on a fixed vector of pre-treatment responses $Y_{i,t-h}$, where $t-h$ spans the period from $t-1$ to the earliest available observation. In this case, the less stringent conditional independence assumption would be

$$E(Y_{0,i,t} \mid \delta_t, Y_{i,t-h}, X_{i,t}, \mathcal{E}_{i,t}) = E(Y_{0,i,t} \mid \delta_t, Y_{i,t-h}, X_{i,t}), \tag{29}$$

where we assume that conditional on past outcomes and time fixed effects, the potential outcomes are independent of treatment status.[15] Given that past outcomes are influenced by observed and unobserved components, with a long pre-treatment period, as in our case, the pre-treatment variables (i.e the fixed vector of pre-treatment responses $Y_{i,t-h}$) represents a proxy for controlling for unobserved time-varying heterogeneity.

The pure DiD and the LDV model are not nested. So, we cannot take one of the two as special case of the other if needed. But if they give broadly similar results, we might be more confident about evidence obtained on our estimate of interest.

An apparently ideal strategy, where for simplicity we do not consider time fixed effects, would be to condition on both LDV and unobserved time invariant effects (i.e., fixed effects), to obtain an even weaker CIA

$$E(Y_{0,i,t} \mid \alpha_i,\ Y_{i,t-h},\ X_{i,t},\ \mathcal{E}_{i,t}) = E(Y_{0,i,t} \mid \alpha_i,\ Y_{i,t-h},\ X_{i,t}). \tag{30}$$

However, as discussed in Angrist and Pischke (2009), this combined approach requires very stringent econometric conditions for identification. In this empirical study, we will therefore utilize the DiD and LDV approaches and compare the results obtained through each of them.

When there is a treatment regarding a multiplicity of periods, one way to assess the appro-

[15] There are also other empirical strategies that share this feature such as, for example, the synthetic control method (Abadie and Gardeazabal, 2003; Abadie et al., 2010).

priateness of the parallel trends assumption within the DiD is to allow for leads and lags of the treatment, which can be written as

$$Y_{i,t} = \alpha_i + \delta_t + \sum_{j=-m}^{q} \beta_{\mathcal{E},j} \mathcal{E}_{i,t=T_0+j} + \Omega X_{i,t} + \varepsilon_{i,t}, \tag{31}$$

where T_0 is the implementation date of the Euro, i.e., 1999Q1. Thus, instead of estimating a single post-treatment effect of the policy, we estimate m leads (pre-treatments) and q lags (post-treatments) of the policy effect. If k coincides with the date of the Euro adoption,, m to $k-1$ would coincide with the pre-treatment period (i.e., the leads). Proposed for the first time by Autor (2003), this is defined by the literature as a placebo experiment, where one pretends that the implementation of the policy took place earlier than in reality. The test proposed by Autor (2003) would then be $\beta_{\mathcal{E},j} = 0 \ \forall j < 0$.

Keeping in mind that this cannot be considered as a proper (over)identification test since it is based only on the pre-treatment period (i.e., there is no guarantee that trends continue to be parallel after the treatment), the null can be rejected because of two not mutually exclusive reasons:

1. the policy effect might have been anticipated by the economy, and thus cannot be safely ascribed to the policy itself;
2. the parallel trends assumption is not a satisfactory basis for the identification of policy effects.

Note on the other hand that if the ATT is not constant over time after policy implementation, the modelling of $\beta_{\mathcal{E},j}, j \geq 0$, which is compatible both with the DiD and the LDV models, allows us to have estimates of the time-varying impact of the policy. This specification is of great policy interest in our empirical application as it allows one to assess whether the impact of the Euro adoption changes after the inception of the Great Recession. Equation (32) provides a companion to (31) for the LDV model,

$$Y_{i,t} = \alpha_i + \delta_t + \sum_{j=-m}^{q} \beta_{\mathcal{E},j} \mathcal{E}_{i,t=T_0+j} + \Omega X_{i,t} + \theta Y_{i,t-h} + \varepsilon_{i,t} \tag{32}$$

4. Main Results

We present our main results for specification (25) in the first column in Table 2. This is a pure DiD approach and requires the parallel trends assumption. Thus, we assume that the evolution of the inflation output variability trade off in countries that did not adopt the Euro (control group) serves as a proxy (with a potentially different conditional mean) for the counterfactual outcome. In our case the counterfactual outcome is the evolution of the inflation output variability trade off for the countries that adopted the Euro (treatment group), which would have occurred in absence of the introduction of the Euro.

First, notice that the coefficient for the ratio of the variability of inflation to output gap, $\hat{\beta}_2$, is highly significant. This means that the inflation output variability trade off exists. One can see this by ignoring all other terms in (25) except for the one involving $\hat{\beta}_2$, thus

$$\begin{aligned} -\ln(\sigma^2_{x,i,t}) &= \hat{\beta}_2 \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) = \hat{\beta}_2 \left[\ln(\sigma^2_{\pi,i,t}) - \ln(\sigma^2_{x,i,t})\right] \\ \Leftrightarrow [-1+\hat{\beta}_2] \ln(\sigma^2_{x,i,t}) &= \hat{\beta}_2 \ln(\sigma^2_{\pi,i,t}) \\ \Leftrightarrow \ln(\sigma^2_{x,i,t}) &= [\hat{\beta}_2/(-1+\hat{\beta}_2)] \ln(\sigma^2_{\pi,i,t}), \end{aligned} \tag{33}$$

where $\hat{\beta}_2 \in [0,1)$ implies that $[\hat{\beta}_2/(-1+\hat{\beta}_2)] < 0$, i.e., a higher inflation variability implies a lower output variability. An estimate of $\hat{\beta}_2 \in [0,1)$ significantly different from zero already implies a non-linear inflation output variability trade off. Equation (33) is linear in the natural logarithms of variances. However, if we apply $\exp(\cdot)$ on both sides, one can see that the relationship between the variances in inflation and output is non-linear and convex as suggested by economic theory.

Second, the coefficient $\hat{\alpha}_e$ shows that the variance of the supply shock has a highly significant impact on the location of the trade off. It has a negative sign, $\hat{\alpha}_e < 0$, therefore, the larger the variance of the supply shock, the larger the variance of inflation and the output gap. Why? Ignore all other terms in (25) apart from the ones involving $\hat{\alpha}_e$ and $\hat{\beta}_2$, thus

$$-\ln(\sigma^2_{x,i,t}) = \hat{\alpha}_e \ln(\sigma^2_{e,i,t}) + \hat{\beta}_2 \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$$

$$\Leftrightarrow [-1+\hat{\beta}_2]\ln(\sigma^2_{x,i,t}) = \hat{\alpha}_e \ln(\sigma^2_{e,i,t}) + \hat{\beta}_2 \ln(\sigma^2_{\pi,i,t})$$

$$\Leftrightarrow \ln(\sigma^2_{x,i,t}) = [\hat{\alpha}_e/(-1+\hat{\beta}_2)]\ln(\sigma^2_{e,i,t}) + [\hat{\beta}_2/(-1+\hat{\beta}_2)]\ln(\sigma^2_{\pi,i,t}) \quad (34)$$

$$\Leftrightarrow \ln(\sigma^2_{\pi,i,t}) = -[\hat{\alpha}_e/\hat{\beta}_2]\ln(\sigma^2_{e,i,t}) + [(-1+\hat{\beta}_2)/\hat{\beta}_2]\ln(\sigma^2_{x,i,t}) \quad (35)$$

and, as $\hat{\alpha}_e < 0$ and $\hat{\beta}_2 \in [0,1)$, it follows that $[\hat{\alpha}_e/(-1+\hat{\beta}_2)], -[\hat{\alpha}_e/\hat{\beta}_2] > 0$. We conclude from (34) and (35) that the relationship between the variance of the supply shock and the variances of the output gap and inflation is positive, which is consistent with the economic theory discussed above.

Next, the coefficients for the non-linear terms $\ln(\sigma^2_{e,i,t})^2$ and $\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$ are not significantly different from zero. This would be evidence in favour of a Cobb-Douglas input transformation function, which is also convex in its inputs. Thus, we can conclude that our data definitely supports a convex inflation output variability trade off. Next, $\hat{\gamma}_{2,e} < 0$ is significant at the ten percent level, which is evidence against the separability between inputs and outputs in the transformation function.[16] In sum, the significant coefficient estimates provide evidence for the existence of an inflation output variability trade off broadly consistent with the theoretical trade off developed in Section 2.

Most importantly, the coefficient for the dummy on Euro adoption, $\hat{\beta}_{\mathcal{E}}$, is highly significant. A negative sign for $\hat{\beta}_{\mathcal{E}}$ means that, on average, adopters of the Euro have been worse off due to adopting this monetary regime. Why? Similar arguments as above, yield

$$\ln(\sigma^2_{x,i,t}) = [\hat{\alpha}_e/(-1+\hat{\beta}_2)]\ln(\sigma^2_{e,i,t}) + [\hat{\beta}_2/(-1+\hat{\beta}_2)]\ln(\sigma^2_{\pi,i,t}) + [\hat{\beta}_{\mathcal{E}}/(-1+\hat{\beta}_2)]\mathcal{E}_{i,t} \quad (36)$$

$$\ln(\sigma^2_{\pi,i,t}) = -[\hat{\alpha}_e/\hat{\beta}_2]\ln(\sigma^2_{e,i,t}) + [(-1+\hat{\beta}_2)/\hat{\beta}_2]\ln(\sigma^2_{x,i,t}) - [\hat{\beta}_{\mathcal{E}}/\hat{\beta}_2]\mathcal{E}_{i,t}, \quad (37)$$

and, as $\hat{\beta}_{\mathcal{E}} < 0$ and $\hat{\beta}_2 \in [0,1)$, it follows that $[\hat{\beta}_{\mathcal{E}}/(-1+\hat{\beta}_2)], -[\hat{\beta}_{\mathcal{E}}/\hat{\beta}_2] > 0$.

We can provide some further interpretation to $\hat{\beta}_{\mathcal{E}} = -0.827$ as it follows that $\exp(\hat{\beta}_{\mathcal{E}}) \approx 0.44$.

[16] The separability hypothesis, which implies that the marginal rates of substitution between inputs are independent of the outputs in the transformation function, is not crucial for the present analysis. Our main results always carry true regardless of its validity.

The latter can be interpreted as the ratio of the Euro Area (post-Euro introduction) and control group transformation functions. Thus, the inverse of this ratio is ≈ 2.29 and means that, conditionally on the supply shock variances, the Euro Area has a joint variance of inflation and output, which is around 129% larger than that of the control group.

Finally, notice that the high p-value for the Ramsey (1969) Reset test suggests that the null hypothesis of the test, i.e., omitted variables being orthogonal with respect to the included variables, cannot be rejected. Therefore, we cannot find evidence that regression equation (25) suffers from misspecification.

5. Robustness

The results from Section 4 suggest that the Euro had a detrimental effect on the macroeconomic performance of the Euro Area countries. The purpose of this section is threefold. First, we assess the robustness of this finding by contrasting the pure DiD approach with the LDV approach. Second, we conduct a more elaborate robustness analysis for both the DiD and LDV approach. To this end we present results for the pure DiD approach where we check for the existence of anticipation and/or divergent trends and lagged heterogeneous effects as explained in Subsection 3.5. Next, we compare these results with the ones estimated in the LDV model introducing the possibility of lagged heterogeneous effects (recall that in this model there cannot be ATT leads). These exercises also allow us to provide a plausible and more detailed economic interpretation of the basic finding of a detrimental effect of the Euro. As we discuss in detail below, our findings can be interpreted as evidence that monetary policy in the Euro Area is best characterized as a *'one size must fit all'* policy. Third, we present several additional robustness analyses to show that none of our results depends on the assumptions made in Section 3.

5.1. Lagged Dependent Variable Approach

Table 3 provides estimates obtained through the LDV approach. We adopt this approach, because it is one way to account for the possibility that policy choices such as the adoption of the Euro depend on past economic performance. This is indeed a plausible scenario and would create an endogeneity problem for the pure DiD approach. Further advantages of the LDV approach are that we do not require the parallel trends assumption and that it is a way of controlling for unobserved time-varying heterogeneity.

Inspection of the first column in Table 3, relating to the estimation of equation (28), reveals that all our findings are qualitatively unchanged vis-à-vis the previous ones. The estimates directly related to the inflation output variability trade off are consistent with the previous findings. Moreover, we find that the Euro dummy, $\hat{\beta}_{\mathcal{E}}$, is significantly different from zero. Clearly, this is further evidence that, relative to the control group, countries with the Euro faced a worse trade off. The usual calculation, $\exp(\hat{\beta}_{\mathcal{E}}) \approx 0.58$, implies a joint variance of inflation and output in the Euro Area, which is more than 1.73 times or 73% larger compared to the one in the control group. Thus, the quantitative implications of the treatment are in the same ballpark of those obtained under the parallel trends assumption. The high p-values for the Ramsey (1969) Reset test imply that our estimates are not misspecified. In sum, our previous findings are robust to relaxing the parallel trends assumption and following an alternative empirical strategy that also controls for potential endogeneity of policy choices and unobserved time-varying heterogeneity.

5.2. Difference-in-Differences with Leads and Lags of the Treatment

The second column in Table 2 presents the results for the specification (31), introducing leads and lags, as previously explained. Compared to the pure DiD results, three observations stand out. First, the coefficients characterizing the existence of the inflation output variability trade off validate the results obtained with specification (25). There is again significant and well-specified evidence in favour of an inflation output variability trade off consistent with the theoretical trade off developed in Section 2.

Second, the coefficients on the leads, $\hat{\beta}_{\mathcal{E},-3}$, $\hat{\beta}_{\mathcal{E},-2}$, and $\hat{\beta}_{\mathcal{E},-1}$, are insignificant. We also tested the null hypothesis that these coefficients are all zero and could not reject it. Thus, anticipation effects and divergent trends between treated and control group appear to play no role in explaining our basic finding from above.

Third, inspection of the coefficients in the second column of Table 2 also reveals that $\hat{\beta}_{\mathcal{E},0}$ and the coefficients on the lags, $\hat{\beta}_{\mathcal{E},2}$ and $\hat{\beta}_{\mathcal{E},3}$, are significant, while $\hat{\beta}_{\mathcal{E},4}$ is insignificant. The former two coefficients relate to periods including the Financial Crisis and the European Sovereign Debt Crisis, while the latter coefficient relates to the time *after* Mario Draghi's *'whatever it takes'*. We interpret these findings as evidence that the treatment effect has changed over time. Apparently, the trade off for the Euro Area countries has significantly deteriorated after the inception of the Euro (relative to the control group) until the Draghi and OMTs announcements as well as the EAPP announcement and implementation, but not thereafter. The coefficient estimates can be interpreted as follows: $\exp(\hat{\beta}_{\mathcal{E},0}) \approx 0.49$, $\exp(\hat{\beta}_{\mathcal{E},2}) \approx 0.33$, and $\exp(\hat{\beta}_{\mathcal{E},3}) \approx 0.24$ imply that the Euro Area had a joint variance of inflation and output, which is more than 105% (205%, 323%) larger during 1999Q1 to 2002Q2 (2006Q1 to 2009Q2, 2009Q3 to 2012Q4) compared to the control group.

5.3. Lagged Dependent Variable Approach with Lagged Effects

Consider now the second column of Table 3, which provides the results from estimation of equation (32). In this specification we cannot have ATT leads, hence $m = 0$.

As in Subsection 5.2, the interaction of the Euro dummy with both the crises periods as well as the periods of Draghi and OMTs announcement as well as EAPP matter. The adoption of the Euro worsens the trade off at the occurrence of the Financial Crisis and the European Sovereign Debt Crisis, but this detrimental effect becomes insignificant in correspondence of the period which starts with the Draghi announcement and includes the OMTs announcement and the EAPP announcement and implementation.

The Euro Area had a joint variance of inflation and output, which is more than 124% (170%) larger during 2006Q1 to 2009Q2 (2009Q3 to 2012Q4) compared to the control group. Once more, the Ramsey (1969) Reset test does not highlight any potential model misspecification. Summing up, Table 3 shows that all our main findings are qualitatively unchanged vis-à-vis the ones from Table 2, regardless of whether the ATT is captured by a single coefficient, $\hat{\beta}_{\mathcal{E}}$, or by a string of lagged variables.

5.4. Discussion

Summing up, we consistently find a detrimental effect of the Euro on macroeconomic performance in periods 7 (2006Q1 to 2009Q2) and period 8 (2009Q3 to 2012Q4). In keeping with most of the previous literature, we surmise an interpretation of these results in terms of a monetary policy for the Euro Area during these periods of crises that can be denoted a *'one size must fit all'* rather than a *'one size fits all'* monetary policy. Moreover, we think that our finding that the detrimental effect of the Euro ceases in period 9 has two implications.

First, it suggests that the detrimental effect during periods 7 and 8 is directly related to periods of crises and so is our *'one size must fit all'* judgement. This would be consistent with De Grauwe's (2012) *Eurozone fragility hypothesis*. The ECB did not immediately react to solvency concerns regarding some peripheral Euro Area countries by signalling its willingness to act as *'buyer of last resort'* on the market for bonds of European governments, although this would have been a natural policy for independent national central banks in these peripheral Euro Area countries. Thus, the ECB policy during this time was rather *'one size must fit all'* than *'one size fits all'*, which resulted in turmoil on European sovereign debt markets that transmitted into worse macroeconomic performance.

Second, we interpret our finding that the detrimental effect of the Euro ceases in period 9 meaning that the ECB acted in such a way to make clear that it was willing to act as *'buyer of last resort'*. In turn, European sovereign debt markets calmed down, leading to the disappearance of the detrimental effect of the Euro on macroeconomic performance for the Euro Area on average. This narrative is also supported by empirical work on the effects of these announcements on European sovereign debt markets (see, e.g., De Grauwe and Ji, 2013; Saka et al., 2015).

A more detailed analysis of the relative effectiveness of such announcements and unconventional monetary policy measures is beyond the scope of this paper. Nevertheless, one indicator to gauge the effectiveness of the monetary policies during periods 7, 8 and 9 is the evolution of the shadow rate developed in Wu and Xia (2016). We depict this measure for the USA and the UK as examples of the control group on the one side and the Euro Area on the other side in Figure 4 below. We observe that the shadow rates in the control group countries moved below zero much earlier than in the Euro Area. In particular, the shadow rate for the Euro Area is continuously negative only since mid 2013, which corresponds to our period 9 and the Draghi and OMTs announcements as well as the EAPP announcement and its subsequent implementation. This stylized fact is broadly consistent with our interpretation from above.

Finally, our discussion of results so far suggests that ECB monetary policy is *'one size must fit all'* only when it comes to its role of *'buyer of last resort'* in periods of crises. However, right

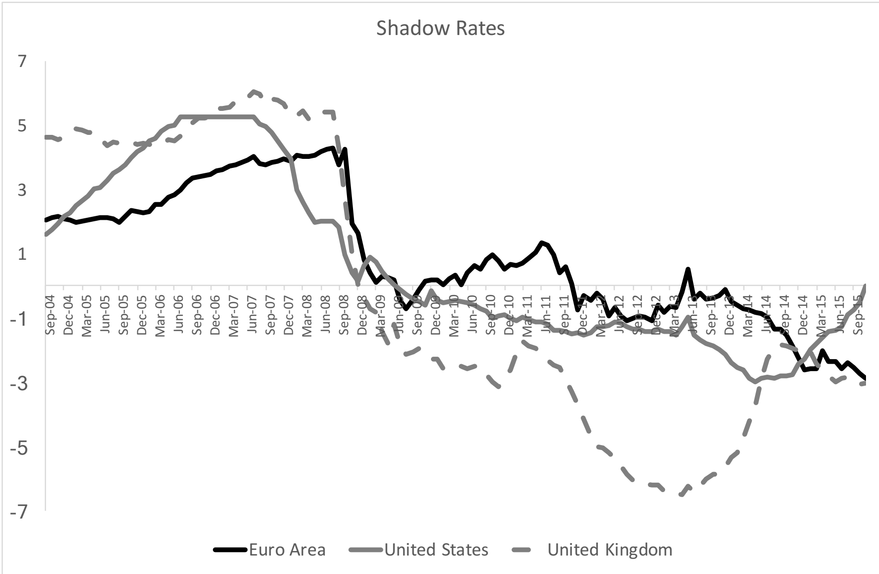


Figure 4: Shadow Rates

below in Subsection 5.5 we present a set of results for the effects of the Euro on the *'core'* and the *'periphery'* of the Euro Area. We find that the detrimental effect of the Euro on macroeconomic performance is more severe in peripheral countries and does disappear in period 9 only in the core, but not in the periphery. This suggests that our *'one size must fit all'* judgement is also plausible outside periods of crises.

5.5. Effects for the Core and the Periphery

It is often argued that there are rather homogeneous *'core'* and *'periphery'* whithin the Euro Area. The periphery is believed to be structurally different from the core in many aspects. This could be a potential explanation of why, monetary policies by the ECB face a problem of *'one size must fit all'* rather than being characterized by *'one size fits all'* (Issing, 2001). In order to examine whether the core or periphery of the Euro Area drives the results, we include interaction terms in our specification. In our analysis, the core consists of Austria, Belgium, Finland, France, Germany, and the Netherlands. The periphery is comprised by Italy, Ireland, Portugal, and Spain, which are the only countries in our sample that experienced a sovereign debt crisis. Subsequently, the regression equations are augmented with an additional explanatory variable that takes a value of one for all Euro countries belonging to the periphery starting in period 5 and zero otherwise. All other assumptions are as in the baseline analysis. This means that we again consider 9 periods as illustrated in Table 1. When computing the variances for the structural supply shocks, we again impose sign restrictions for the horizon $K = 4$, but no sign restriction on the nominal interest rate. Moreover, we again estimate the pure DiD model, the LDV approach, the DiD with lead and lagged effects, and the LDV approach with lagged effects.

Table 4 reports the results for this analysis, whose main features can be summed up by the following four remarks. First, we observe that throughout all specifications, the estimated coefficients regarding the inflation output variability trade off are similar to our previous results. Second, again the Ramsey (1969) Reset test does not give rise to concerns related to omitted variables problems. Third, the effect of the Euro on the core is in line with our findings above. There is again a detrimental effect of the Euro. Once we allow for heterogeneous effects of the treatment over time, the effect is observed in periods 7 and 8, but not in period 9.

However, and fourth, the coefficient estimates for the periphery reveal differences vis-à-vis the former results. During the periods of crises the detrimental effect of the Euro is more severe in the periphery as the respective coefficients are larger in absolute value. This is consistent with the asymmetric effects of shocks in the Euro Area as found in Barigozzi et al. (2014) and Georgiadis (2015). More crucially, the detrimental effect of the Euro does not cease in period 9. This finding suggests that while Draghi and OMTs announcements as well as the EAPP announcement and implementation had a beneficial effect on the macroeconomic performance on the core, this was not the case for the periphery. We also carried out an analysis with separate core and periphery subsamples, where the treatment group consists only of the core *or* the periphery of the Euro Area. These results are reported in Table 5 and are consistent with the results in this subsection.

One plausible explanation for this finding is the observation of zombie lending or forbearance, i.e., measures that keep non-performing loans performing. This is associated with a misallocation of bank credit and may prevent economic recovery. Acharya et al. (2017a) present evidence that such zombie lending in the Euro Area was particularly severe in the peripheral countries. However, there is also evidence for zombie lending in the control group, e.g., Japan (see, Caballero et al., 2008). Moreover, Schivardi et al. (2017) provide evidence for Italy that questions the importance of zombie lending in the Euro Area. Thus, while zombie lending appears to be a plausible candidate explanation, the empirical literature has not yet reached a consensus on the quantitative importance of this candidate explanation.

5.6. Alternative Identification of the Structural Supply Shock

Horizon. Ex ante it is not clear for what horizon K one should restrict the signs on the impulse response functions in Subsection 3.3. For instance, the theoretical model in Section 2 above suggests that, under discretionary optimal monetary policy, the effects of a structural supply shock are offset within one quarter if the shock is purely transitory. However, if the shock is serially correlated, the deviations from steady state can last for many quarters and that would justify to impose sign restrictions for a longer horizon. But, in the case of optimal monetary policy under commitment, independent of the persistence of the shock, one would expect inflation to revert the sign after one or some initial periods. This reversion of sign would suggest to impose the sign restriction for a rather short horizon. In order to address these issues, we examine robustness of our previous results obtained (with horizon $K = 4$) and set $K = 1$. Thereafter in this subsection we refer to the results obtained in Section 4 and Subsections 5.1 to 5.3 as our baseline set of results.

Table 6 displays results for imposing sign restrictions for one quarter instead of four quarters, when computing the variance of the structural supply shock. A comparison of Table 6 with Tables 2 and 3 enables us to conclude that all the previously obtained results carry through.

Once we account for lead and lagged effects of the treatment in the estimation with the DiD estimator, there emerges one major difference to our previous set of results. The detrimental effect of the Euro on macroeconomic performance is present starting in 1999Q1 to 2002Q2 and does not disappear in the period 2013Q1 to 2016Q4. This finding goes against our suggested interpretation that the Draghi and OMTs announcements as well as the EAPP announcement and implementation of the ECB have had a beneficial effect on macroeconomic performance in the Euro Area countries. However, once we account for potential endogeneity of the treatment and unobserved time-varying heterogeneity in the LDV approach, we find again that the detrimental effect vanishes in the period 2013Q1 to 2016Q4.

Thus, although we find one incident that can be held against our suggested interpretation of the baseline set of results, we conclude that our baseline set of results is, overall, robust to shortening the horizon for which we impose sign restrictions on inflation and output gap in the computation of the variances of the structural supply shock.

SIGN. Our baseline set of results uses the variance of a structural supply shock that is computed based on an identification scheme that does not restrict the response of the nominal interest rate. The underlying consideration is that we do not want to make any explicit assumption on the particular monetary policy in the countries on our sample. To make this clear, consider the theoretical model outlined above in Section 2. Independent of whether monetary policy is best described by a simple Taylor (1993) rule or optimal monetary policy under discretion, the model implies a positive response of the nominal interest rate in response to a structural supply shock. In contrast, under the assumption of optimal monetary policy under commitment, the model can imply a negative response of the nominal interest rate.

However, once we allow for serially correlated shocks, even optimal monetary policy under commitment typically exhibits an increase of the nominal interest rate on impact and initial periods followed by undershooting of the steady state. Moreover, there is empirical evidence that, since 1979, monetary policy in many developed countries is characterised by a strong positive response of the nominal interest rate to increases in inflation (see, e.g., Clarida et al., 1998). Thus, it seems natural to examine whether our baseline findings are robust to imposing an additional identification restriction on the nominal interest rate when computing the variances of the structural supply shock. In particular, we restrict the response of the nominal interest rate to be positive for $K = 4$ periods.

Table 7 displays results for imposing a sign restriction on the nominal interest rate for $K = 4$ quarters when computing the variance of the structural supply shock. Once we compare Table 7 to Tables 2 and 3, we can conclude that our baseline set of results and our suggested interpretation is entirely robust to imposing also a sign restriction on the nominal interest when computing the variances of the structural supply shock.

5.7. Different Number of Periods in the Panel Estimation

So far, we have split our quarterly data into 9 periods as described in Subsection 3.4. While we believe that this choice is plausible, one may argue that it is arbitrary and that a different choice may potentially yield different results. In order to address this concern, we provide an alternative choice of periods in Table 8, where we split the sample into 7 instead of 9 periods. As a consequence, both the financial crisis and the European sovereign debt crisis are in period 6 and the Draghi and OMTs announcements as well as the EAPP announcement and implementation are in period 7. We then estimate the pure DiD model, the LDV approach, the DiD with lead and lagged effects, and the LDV approach with lagged effects. All other assumptions are the same as in the baseline set-up.

Table 9 displays these estimation results. Across the different regression equations, three observations stand out. First, the coefficient estimates related to the inflation output variability trade off are in line with our previous estimates and confirm the existence of the this trade off consistent with economic theory. Second, for both the pure DiD and the LDV approach, there is a detrimental effect of the Euro. Third, when we control for lead and lagged effects, there is no effect for the DiD model, but the usual effect for the LDV approach. In periods of crises there is a detrimental effect of the Euro in period 6 that ceases with the Draghi and OMTs announcements as well as the EAPP announcement and implementation in period 7. Fourth, the Ramsey (1969) Reset test is in line with the baseline results. Thus, we conclude that our findings are robust to different choices of the periods in the analysis.

5.8. Further Discussion

Could bad luck be the explanation of the detrimental effect of the Euro on macroeconomic performance in periods 7 and period 8 and the disappearance of the detrimental effect in period 9? Put differently, did the Euro Area countries experienced a more severe sequence of shocks than the economies in the control group during periods 7 and 8, while this was not the case in period 9? In principle, we are confident that this is not a reasonable explanation, as we have controlled for any (supply) shock that generates an inflation and output variability trade off. In consequence, we are confident that our results are not driven by such shocks.

However, it may be argued that other shocks that create variability of inflation and output, but do not necessarily create an inflation and output variability trade off, can explain the detrimental effect of the Euro during periods 7 and 8. For instance, in the context of our theoretical New Keynesian model, technology shocks, discount factor shocks, or, shocks to the financial market conditions may create variability of inflation and output at the zero interest-rate lower bound as they can no longer be offset by conventional monetary policy. In consequence, the detrimental effect of the Euro in periods 7 and 8 could be due to an idiosyncratic sequence of shocks, for instance, to financial market conditions. Yet this is not a likely explanation for the following reasons.

First, the Ramsey (1969) Reset misspecification test is never significant. It has long been known (Thursby, 1981, 1982) that in this case we can safely expect that any omitted variable is orthogonal to the included regressors and hence does not bias results. Second, we have obtained similar results with the pure DiD and the LDV approaches, where the latter also controls for unobserved time-varying heterogeneity. Therefore, we conclude that bad luck does not seem to be a very plausible explanation for the detrimental effect of the Euro in periods 7 and 8. The very same considerations suggest that good luck is not a very plausible explanation for the disappearance of the detrimental effect of the Euro in period 9.

6. Concluding Remarks

This paper conducts a counterfactual analysis providing evidence that Euro Area countries would have experienced a more favourable inflation output variability trade off without the Euro. The deterioration of the trade off for Euro Area countries is related to the period of the Great Recession and the European Sovereign Debt Crisis and ceases with the Draghi and OMTs announcements as well as the EAPP announcement and implementation.

Our findings are based on a novel empirical strategy that is consistent with monetary theory that implies an inflation output variability trade off whose position is influenced by structural supply shocks. We develop a panel data set for twenty OECD countries. The computation of variances for output gap and inflation deviations from target is based on a novel detrending method, which has been shown to be more robust than previous filters used in the literature such as the HP filter. The variance for the structural supply shock is estimated via a structural model that uses sign-restrictions to identify the shock. In the estimation of the trade off, we model the joint determination of the variability of inflation and output by the structural supply shock. The counterfactual evidence is identified via the ATT effect and is robust to various assumptions on the empirical specification.

We interpret the higher inflation and output variability in the Euro Area during the periods of crisis as evidence that the ECB measures during these periods have not been effective to reduce inflation and output variability to levels comparable with other economies. Moreover, the disappearance of this detrimental effect cannot be found after the Draghi announcement onwards. This suggests that the policy moves subsequent to Draghi's *'whatever it takes'* announcement have been effective in reducing inflation and output variability in the Euro Area on average. We argue that this is the case, because these moves credibly signalled that the ECB was going to act as *'buyer of last resort'*.

Our more detailed analysis shows that the detrimental effect of the Euro is more severe for peripheral countries of the Euro Area. Moreover, while the Draghi and OMTs announcements as well as the EAPP announcement and implementation had a beneficial effect on the macroeconomic performance of the core, this was not the case for the periphery. Hence, structural differences among Euro Area countries may be the underlying reason for the detrimental effect of the Euro.

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Tables

Table 1: Overview on the periods considered in the panel regressions

Period	From	To	# of Obs.	Comments
1	1985Q1	1988Q2	14	Beginning of the Great Moderation
2	1988Q3	1991Q4	14	
3	1992Q1	1995Q2	14	
4	1995Q3	1998Q4	14	
5	1999Q1	2002Q2	14	Start of the Euro
6	2002Q3	2005Q4	14	
7	2006Q1	2009Q2	14	Financial Crisis
8	2009Q3	2012Q4	14	European Sovereign Debt Crisis Draghi announcement (July 26th, 2012) Outright Monetary Transactions (OMTs) announcement (September 6th, 2012)
9	2013Q1	2016Q4	16	Expanded Asset Purchase Programme (EAPP) (January 22nd, 2015)

Table 2: Estimated parameters for model with fixed effects and time dummy[a]

Variables	Coefficient	Estimates[b] (25)	Estimates[b] (31)
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.827** (0.310)	
$\mathcal{E}_{i,T_0-3}$	$\beta_{\mathcal{E},-3}$		0.030 (0.348)
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$		-0.054 (0.389)
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$		-0.419 (0.339)
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$		-0.718* (0.363)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$		-0.584 (0.413)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$		-1.115** (0.490)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$		-1.442*** (0.484)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$		-0.813 (0.494)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.508*** (0.089)	-0.540*** (0.095)
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	-0.191 (0.269)	-0.233 (0.228)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.530*** (0.048)	0.559*** (0.055)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	-0.021 (0.029)	-0.016 (0.029)
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.206* (0.114)	-0.221* (0.110)
Country fixed effect		yes	yes
Time fixed effect		yes	yes
N		20	20
Number of observations		180	180
R^2		0.808	0.820
Specification tests[c]:			
Ramsey (1969) Reset		0.606	0.730
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$			0.492

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.
[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).
[c] p-values are reported for all tests.

Table 3: Estimated parameters for model with fixed effects, time dummy, and lagged dependent variables[a]

Variables	Coefficient	Estimates[b]	
		(28)	(32)
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.550*	
		(0.302)	
$\mathcal{E}_{i,T_0-3}$	$\beta_{\mathcal{E},-3}$		
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$		
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$		
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$		-0.336
			(0.307)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$		-0.120
			(0.370)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$		-0.805**
			(0.328)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$		-0.995**
			(0.389)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$		-0.421
			(0.477)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.537**	-0.630**
		(0.233)	(0.237)
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	-0.195	-0.337
		(0.419)	(0.351)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.563***	0.604***
		(0.088)	(0.092)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	0.014	0.039
		(0.045)	(0.049)
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.020	-0.001
		(0.116)	(0.098)
Country fixed effect		yes	yes
Time fixed effect		yes	yes
N		20	20
Number of observations		100	100
R^2		0.818	0.833
Specification tests[c]:			
Ramsey (1969) Reset		0.9197	0.8595
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$			

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.

[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).

[c] p-values are reported for all tests.

Table 4: Estimated parameters with interaction terms for core and periphery for model with fixed effects and time dummy as well as the former augmented with lagged dependent variables[a]

Variables	Coefficient	Estimates[b]							
		(25)		(28)		(31)		(32)	
		Core	Periphery	Core	Periphery	Core	Periphery	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.759**	-0.924**	-0.095	-0.913***				
		(0.331)	(0.361)	(0.295)	(0.280)				
$\mathcal{E}_{i,T_0-3}$	$\beta_{\mathcal{E},-3}$					0.126	-0.075		
						(0.487)	(0.437)		
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$					-0.094	-0.065		
						(0.512)	(0.439)		
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$					-0.237	-0.616		
						(0.299)	(0.577)		
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$					-0.695*	-0.708	-0.068	-0.529
						(0.369)	(0.591)	(0.325)	(0.399)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$					-0.710	-0.349	-0.047	-0.118
						(0.471)	(0.433)	(0.378)	(0.387)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$					-1.019*	-1.185	-0.459	-1.035**
						(0.507)	(0.717)	(0.352)	(0.446)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$					-1.267**	-1.693***	-0.577	-1.393***
						(0.531)	(0.563)	(0.418)	(0.384)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$					-0.335	-1.653***	0.240	-1.430***
						(0.545)	(0.439)	(0.563)	(0.281)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.519***		-0.522**		-0.515***		-0.521**	
		(0.088)		(0.225)		(0.112)		(0.235)	
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	-0.192		0.161		-0.303		-0.056	
		(0.270)		(0.405)		(0.208)		(0.354)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.530***		0.536***		0.543***		0.538***	
		(0.048)		(0.083)		(0.063)		(0.087)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	-0.018		-0.027		0.015		0.028	
		(0.028)		(0.043)		(0.029)		(0.046)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.210*		-0.102		-0.237*		-0.100	
		(0.113)		(0.117)		(0.116)		(0.108)	
Country fixed effect		yes		yes		yes		yes	
Time fixed effect		yes		yes		yes		yes	
N		20		20		20		20	
Number of observations		180		100		180		100	
R^2		0.8086		0.8348		0.8355		0.8672	
Specification tests[c]:									
Ramsey (1969) Reset		0.5422		0.9528		0.7449		0.8859	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$									

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.
[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).
[c] p-values are reported for all tests.

Table 5: Estimated parameters for core or periphery of Euro Area for model with fixed effects and time dummy as well as the former augmented with lagged dependent variables[a]

Variables	Coefficient	Estimates[b]							
		(25)		(28)		(31)		(32)	
		Core	Periphery	Core	Periphery	Core	Periphery	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.753** (0.340)	-0.937** (0.377)	0.378 (0.368)	-1.286*** (0.276)				
$\mathcal{E}_{i,T_0-3}$	$\beta_{\mathcal{E},-3}$					0.186 (0.488)	-0.093 (0.449)		
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$					-0.044 (0.524)	-0.086 (0.452)		
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$					-0.192 (0.287)	-0.702 (0.610)		
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$					-0.664* (0.363)	-0.781 (0.588)	0.533 (0.317)	-0.930** (0.362)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$					-0.656 (0.473)	-0.382 (0.434)	0.618 (0.443)	-0.460 (0.348)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$					-0.968* (0.511)	-1.233 (0.722)	0.153 (0.424)	-1.415*** (0.456)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$					-1.234** (0.537)	-1.807*** (0.549)	0.044 (0.452)	-1.846*** (0.319)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$					-0.262 (0.560)	-1.597*** (0.432)	0.939 (0.562)	-1.744*** (0.370)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.495*** (0.115)	-0.559*** (0.095)	-0.457 (0.278)	-0.635** (0.249)	-0.517*** (0.125)	-0.581*** (0.105)	-0.520* (0.260)	-0.663** (0.250)
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	-0.351 (0.256)	-0.180 (0.272)	0.242 (0.526)	0.052 (0.444)	-0.351 (0.242)	-0.296 (0.189)	0.230 (0.412)	-0.187 (0.403)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.546*** (0.055)	0.563*** (0.059)	0.534*** (0.114)	0.567*** (0.087)	0.579*** (0.068)	0.581*** (0.057)	0.582*** (0.115)	0.577*** (0.070)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	0.027 (0.046)	0.004 (0.031)	0.068 (0.112)	0.027 (0.049)	0.034 (0.050)	0.030 (0.029)	0.095 (0.113)	0.074 (0.043)
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.198 (0.159)	-0.278** (0.116)	-0.138 (0.134)	-0.192 (0.126)	-0.221 (0.166)	-0.291** (0.115)	-0.158 (0.139)	-0.167 (0.100)
Country fixed effect		yes	yes	yes	yes	yes	yes	yes	yes
Time fixed effect		yes	yes	yes	yes	yes	yes	yes	yes
N		16	14	16	14	16	14	16	14
Number of observations		144	126	80	70	144	126	80	70
R^2		0.7931	0.8066	0.8515	0.8703	0.8051	0.8338	0.8682	0.9006
Specification tests[c]:									
Ramsey (1969) Reset		0.3089	0.8333	0.9686	0.7905	0.6218	0.9737	0.5576	0.7694
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$						0.7133	0.3675		

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.
[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).
[c] p-values are reported for all tests.

Table 6: Estimated parameters for model with fixed effects and time dummy as well as the former augmented with lagged dependent variables with sign restrictions imposed for one quarter[a]

Variables	Coefficient	Estimates[b]			
		(25)	(28)	(31)	(32)
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.870**	-0.598*		
		(0.324)	(0.298)		
$\mathcal{E}_{i,T_0-3}$	$\beta_{\mathcal{E},-3}$			-0.181	
				(0.384)	
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$			-0.204	
				(0.424)	
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$			-0.517	
				(0.330)	
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$			-0.846**	-0.357
				(0.371)	(0.301)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$			-0.782*	-0.206
				(0.442)	(0.373)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$			-1.245**	-0.810**
				(0.511)	(0.324)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$			-1.552***	-1.006**
				(0.503)	(0.385)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$			-1.071*	-0.543
				(0.544)	(0.453)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.438***	-0.595**	-0.453***	-0.633***
		(0.088)	(0.220)	(0.100)	(0.216)
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	-0.300	-0.370	-0.336	-0.446
		(0.254)	(0.483)	(0.240)	(0.442)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.518***	0.540***	0.536***	0.566***
		(0.050)	(0.092)	(0.057)	(0.095)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	-0.023	0.027	-0.022	0.048
		(0.028)	(0.059)	(0.029)	(0.063)
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.139*	-0.059	-0.155**	-0.043
		(0.067)	(0.127)	(0.057)	(0.106)
Country fixed effect		yes	yes	yes	yes
Time fixed effect		yes	yes	yes	yes
N		20	20	20	20
Number of observations		180	100	180	100
R^2		0.801	0.820	0.812	0.833
Specification tests[c]:					
Ramsey (1969) Reset		0.8218	0.7781	0.8662	0.5792
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$				0.4611	

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.
[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).
[c] p-values are reported for all tests.

Table 7: Estimated parameters for model with fixed effects and time dummy as well as the former augmented with lagged dependent variables with additional sign restriction on the nominal interest rate[a]

Variables	Coefficient	Estimates[b]			
		(25)	(28)	(31)	(32)
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.727** (0.300)	-0.482* (0.277)		
$\mathcal{E}_{i,T_0-3}$	$\beta_{\mathcal{E},-3}$			0.157 (0.386)	
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$			0.053 (0.496)	
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$			-0.340 (0.385)	
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$			-0.526 (0.395)	-0.272 (0.325)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$			-0.459 (0.420)	-0.126 (0.353)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$			-0.927* (0.484)	-0.712** (0.315)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$			-1.237** (0.471)	-0.894** (0.360)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$			-0.561 (0.507)	-0.333 (0.492)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.286* (0.159)	-0.450 (0.262)	-0.311* (0.163)	-0.513* (0.293)
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	0.121 (0.313)	-0.172 (0.342)	0.140 (0.296)	-0.240 (0.324)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.499*** (0.039)	0.544*** (0.086)	0.532*** (0.049)	0.576*** (0.088)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	-0.032 (0.033)	-0.002 (0.046)	-0.028 (0.035)	0.012 (0.048)
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.183 (0.116)	0.039 (0.103)	-0.205* (0.109)	0.036 (0.094)
Country fixed effect		yes	yes	yes	yes
Time fixed effect		yes	yes	yes	yes
N		20	20	20	20
Number of observations		180	100	180	100
R^2		0.791	0.810	0.802	0.822
Specification tests[c]:					
Ramsey (1969) Reset		0.5478	0.8502	0.7399	0.7694
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$				0.5123	

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.
[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).
[c] p-values are reported for all tests.

Table 8: Overview on the modified periods considered in the panel regressions

Period	From	To	# of Obs.	Comments
1	1985Q1	1989Q3	19	Beginning of the Great Moderation
2	1989Q4	1994Q2	19	
3	1994Q3	1998Q4	18	
4	1999Q1	2003Q2	18	Start of the Euro
5	2003Q3	2007Q4	18	
6	2008Q1	2012Q2	18	Financial Crisis, European Sovereign Debt Crisis
7	2012Q3	2016Q4	18	Draghi announcement (July 26th, 2012), Outright Monetary Transactions (OMTs) announcement (September 6th, 2012), Expanded Asset Purchase Programme (EAPP) (January 22nd, 2015)

Table 9: Estimated parameters for model with fixed effects and time dummy as well as the former augmented with lagged dependent variables with seven periods in the panel[a]

Variables	Coefficient	Estimates[b]			
		(25)	(28)	(31)	(32)
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.503* (0.0.281)	-0.586** (0.0.214)		
$\mathcal{E}_{i,T_0-2}$	$\beta_{\mathcal{E},-2}$			0.192 (0.415)	
$\mathcal{E}_{i,T_0-1}$	$\beta_{\mathcal{E},-1}$			-0.037 (0.338)	
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$			-0.294 (0.404)	-0.446 (0.290)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$			-0.345 (0.472)	-0.460 (0.336)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$			-0.786 (0.473)	-0.968*** (0.259)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$			-0.339 (0.465)	-0.395 (0.320)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.778*** (0.093)	-0.791*** (0.168)	-0.807*** (0.104)	-0.847*** (0.171)
$\ln(\sigma^2_{e,i,t})^2$	α_{ee}	-0.442 (0.538)	-0.355 (0.492)	-0.412 (0.544)	-0.335 (0.395)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.558*** (0.061)	0.604*** (0.113)	0.584*** (0.067)	0.647*** (0.117)
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,1}$	-0.006 (0.037)	-0.027 (0.083)	-0.004 (0.036)	-0.013 (0.084)
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{2,e}$	-0.084 (0.102)	-0.101 (0.120)	-0.087 (0.109)	-0.105 (0.107)
Country fixed effect		yes	yes	yes	yes
Time fixed effect		yes	yes	yes	yes
N		20	20	20	20
Number of observations		140	80	140	80
R^2		0.8240	0.8435	0.8289	0.8513
Specification tests[c]:					
Ramsey (1969) Reset		0.8328	0.7117	0.6340	0.7090
$\beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$				0.7887	

[a] The dependent variable is the variance of the output gap, i.e., $-\ln(\sigma^2_{x,i,t})$.
[b] ***p<0.01; **p<0.05; *p<0.10; Standard errors are in parentheses (cluster-robust standard errors, robust to serial correlation and heteroskedasticity).
[c] p-values are reported for all tests.