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Does One Size Fit All in the Euro Area? Some Counterfactual Evidence.[⋆]

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Abstract

This paper examines whether Euro Area countries would have faced a more favourable inflation output variability tradeoff without the Euro. We provide evidence supporting this claim for the periods of the Great Recession and the European Sovereign Debt Crisis. The deterioration of the tradeoff becomes insignificant only after Draghi's *'whatever it takes'* announcement. A more detailed analysis shows that the detrimental effect of the Euro is more severe and long-lasting for peripheral countries, pointing to structural differences among Euro Area countries as a key element of the Euro's impact during crises. We base our results on a novel empirical strategy that, consistently with monetary theory, models the joint determination of the variability of inflation and output conditional on structural supply and demand shocks. Moreover, our findings are robust to potential endogeneity concerns related to adoption of the Euro.

JEL Classification: C32, E50, F45.

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1. Introduction

The economic crises that occurred 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 experienced modest growth and high employment in the second half of the last decade, others remained in the process of recovery, suffering 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 have amplified the effects of adverse shocks and lead to sub-optimal macroeconomic performance. 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'* (Issing, 2001) rather than *'one size fits all'* (Issing, 2005).

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 period of Issing's (2005) judgment roughly before the beginning of the Great Recession, and, the crisis period since then. The panels suggest that, according to these key indicators

[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, 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]These figures are based on our own calculations, which we detail in Section 3 below.

of macroeconomic performance, non-Euro OECD countries have been more successful in reducing inflation and 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, this kind of evidence leaves many questions unanswered. Do these results depend on different shocks hitting the two country groups? Are they uniform across Euro countries, time or policy changes? In this paper, we address these questions through an empirical set-up drawing from both counterfactual analysis and the analysis of productive processes.

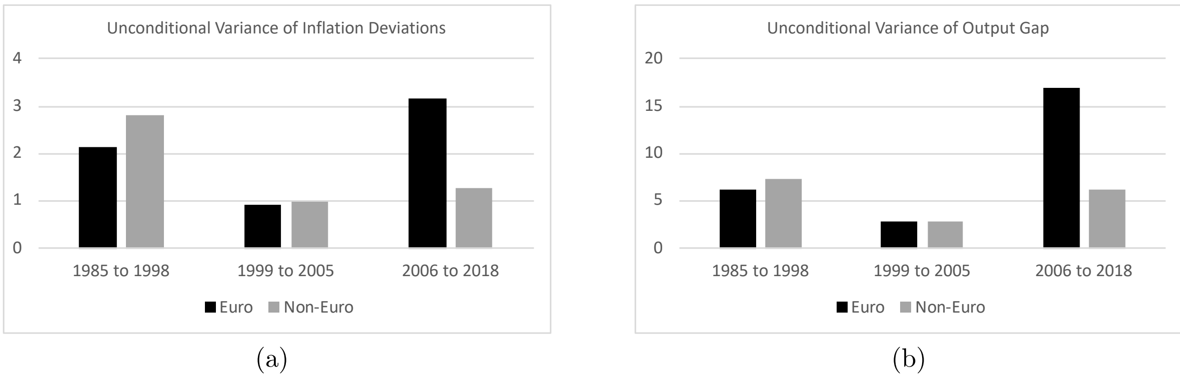


Figure 1: Comparison of the uncondional variances of inflation deviations from an estimated target and of the output gap between Euro Area as well as non-Euro OECD countries.

There is already a literature that tries to establish empirically if there exists a *'one size fits all'* monetary policy for Euro Area countries. Most of these papers focus on the transmission of monetary policy shocks before and after the introduction of the Euro in 1999, with somewhat ambiguous results. 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), Georgiadis (2015) and Burriel and Galesi (2018), 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; second, a key challenge in this literature is endogeneity, as it is unanimously recognized that the choice of IT is affected by initial conditions.

Conceptually, the focus on a single variable does not seem fully appropriate. Measurement and comparison of macroeconomic performance in a theoretical IT framework is routinely based on loss functions that involve inflation *and* output variability. Independently of whether one assumes optimal monetary policy or a simple Taylor (1993) rule, a central bank faces a long-run tradeoff 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 , and also by structural demand shocks.[4] Hence from the standpoint of such a framework, evidence based on the variability of inflation *or* output in isolation appears problematic. In case 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 tradeoff, but do not face an improved tradeoff due to the Euro.[5]

Against this background, this paper seeks to examine the claim of whether the Euro

[3]The most recent data considered by this literature, except for Burriel and Galesi (2018), is from 2009. Therefore most papers do not take the European Sovereign Debt Crisis into account.

[4]Taylor (1979) pioneered the empirical documentation of this long-run inflation output variability tradeoff 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.

Area countries would have faced a more favorable inflation output variability tradeoff 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. Furthermore, we take advantage of a more extensive dataset than that available to previous work to explore possible heterogeneities in the responses to monetary regimes across countries and time.

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 2018. We also estimate the variance of the structural supply and demand 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 tradeoff between inflation and output gap variability on the variability of structural 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 and demand shocks of each country as exogenous outputs, or, more generally, as shifters. Third, in order to establish whether Euro adopters, on average, have been worse off by adopting this monetary regime, we first utilize a difference-in-differences (DiD) approach. However, as 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, we test for the parallel trends assumption and also consider the lagged dependent variable (LDV) model (for a detailed discussion see

Angrist and Pischke, 2009). The latter requires less stringent identification assumptions and controls for potential endogeneity of the Euro adoption.

We find that adopting the Euro worsened the macroeconomic performance of Euro countries on average. More precisely, 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, for the Euro Area as a whole the detrimental effect of the Euro ceases after Mario Draghi's announcement about *'whatever it takes to preserve the Euro'* and the ECB's enactment of more intense and additional unconventional policies such as the outright monetary transactions (OMTs), the targeted longer-term refinancing operations (TLTROs) and the expanded asset purchase programme (EAPP). Therefore we interpret our findings as evidence that these measures 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., 2017), i.e., what De Grauwe (2012) describes as a lender of last resort in the government bond markets.

Disaggregating the analysis across country groups shows that the detrimental effect of the Euro is more severe in peripheral countries. In addition, while this effect of the Euro becomes insignificant for the core of the Euro Area in the period of the above mentioned policy interventions, it remains significant for the peripheral countries until 2015Q3. These findings suggest that structural differences among Euro Area countries are a key element of the detrimental effect of the Euro and that monetary policy in the Euro Area during the crises period was best characterized as a *'one size must fit all'* policy. The findings are consistent with Burriel and Galesi (2018) who find that the effects of ECB's unconventional monetary policy measures on Euro Area countries are heterogeneous and related to Barigozzi et al. (2014) and Georgiadis (2015) who provide evidence for asymmetric effects at the country level to common monetary policy shocks in the Euro Area.

Our study is also 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). Furthermore their analysis is conducted exclusively for the US.

Unlike this literature, 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 tradeoffs implied by optimal monetary policies *and* by simple Taylor (1993) rules. Besides, in our empirical analysis we explicitly link the jointly determined variability of inflation and output gap to exogenous supply and demand shocks.

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 7 concludes.

2. Theoretical Framework

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

Frameworks for measuring and comparing macroeconomic performance in 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 relative to inflation stabilization. Moreover, assume that the aggregate economy is best approximated by a standard New Keynesian model under the rational expectations hypothesis.[6] Under optimal discretionary monetary policy (as elaborated in Clarida et al., 1999), the central bank minimizes (1) subject to the aggregate economy in each period. One can show that the minimum state variable solution of this model then implies the following long-run relationships in unconditional variances

$$\sigma_{\pi,*}^2 = a_\pi^2 \sigma_e^2 \tag{2}$$

$$\sigma_{x,*}^2 = a_x^2 \sigma_e^2, \tag{3}$$

[6]See Galí (2015) or Woodford (2003) for more details on this model.

where e_t is an exogenous supply disturbance assumed to be $e_t \sim \text{iid}(0, \sigma_e^2)$.[7] In short, (2) to (3) show that both the optimal variances of inflation and output gap depend on the supply shock variance. Moreover, one can verify that 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, $\text{F}_{\text{USA},*}$, 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 tradeoff 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 tradeoff, $\text{Fed}_{\text{optimal}}$ in Figure 2a.

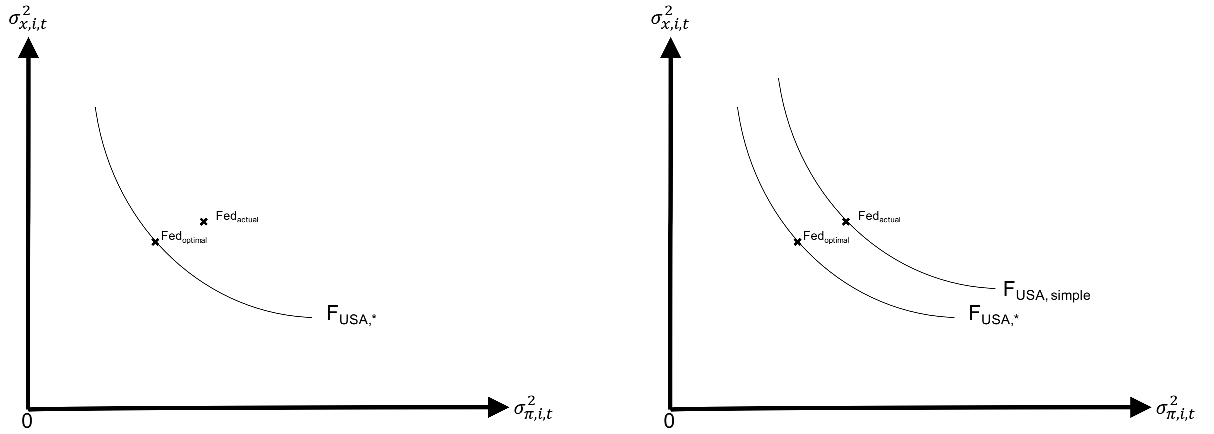
When it comes to measuring the macroeconomic performance, *in practice*, the efficient frontier can be estimated, for instance, via a parsimonious reduced form VAR with a supply and demand equation, including reduced form shocks. Such a frontier states an approximation of optimal monetary policy.[8] 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 tradeoff, $\text{Fed}_{\text{actual}}$ in Figure 2a. Therefore a central bank's monetary policy can be classified as sub-optimal.[9]

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, it can also involve terms for observed monetary policy inertia, feedback to real economic activity or

[7] See Appendix A for the details.

[8] *'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).

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



(a) Stylized optimal tradeoff for the USA

(b) Stylized optimal and sub-optimal tradeoff for the USA

Figure 2: Illustration of the inflation output variability tradeoff

exchange rates. Therefore, such rules may be a more suitable description of monetary policy for many countries.

As we discuss next, under such a Taylor (1993) rule, there is still an inflation output variability tradeoff. However, this tradeoff is neither optimal, nor is it captured by the approach pursued in Cecchetti et al. (2006) and related studies, which explicitly assume optimal monetary policy. For instance, consider the simple interest rate rule[10]

$$i_t = \phi_\pi \pi_t, \qquad \phi_\pi > 0. \tag{4}$$

The minimum state variable solution under policy (4) then implies the following long-run relationships in unconditional variances

$$\sigma^2_\pi = b^2_{\pi,e}\sigma^2_e + b^2_{\pi,g}\sigma^2_g \tag{5}$$

$$\sigma^2_x = b^2_{x,e}\sigma^2_e + b^2_{x,g}\sigma^2_g, \tag{6}$$

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

In contrast to (2) and (3), also demand shocks affect the unconditional variances of inflation and output gap in (5) to (6). The reason for this is, that unlike in the case of optimal policy, a simple rule does not necessarily offset the effects of demand shocks. 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 tradeoff, although the latter is based on the simple interest rate rule (4). The challenge is then to develop an empirical framework that is flexible enough to encompass both the tradeoffs implied by optimal and simple monetary policy.

In this paper, we propose an empirical framework to tackle this challenge. We assume that a inflation output variability tradeoff exists independently of the specific monetary policy in a certain country. Coming back to the example of the Fed in Figure 2, actual variances observed for the USA may be the result of optimal or sub-optimal monetary policy, but a tradeoff exists at any rate, see Figure 2b. We solely assume that, consistently with the above theory, an exogenous supply shock shifts output and inflation in opposite directions and 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. Using observations for more countries at different points in time, our framework allows us to fit a convex curve as depicted in Figure 3, which is shifted by changes in the variance of supply shocks.

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

[11] It is well known that this model lacks a determinate rational expectations equilibrium for $\phi_\pi \leq 1$.

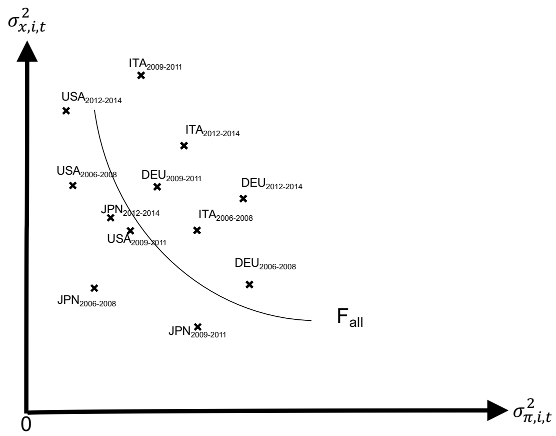


Figure 3: Illustration of the estimation of the tradeoff by a translog transformation function

nous 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. In this way, the macroeconomic performance of different countries can be gauged controlling for country-specific supply shocks. As we show in Appendix A, we can obtain the following empirical specification

$$\begin{aligned}
-\ln(\sigma_x^2) = {} & \alpha_0 + \alpha_e \ln(\sigma_e^2) + \alpha_g \ln(\sigma_g^2) \\
& + (1/2)\left[\alpha_{e,e}\ln(\sigma_e^2)^2 + 2\alpha_{e,g}\ln(\sigma_e^2)\times\ln(\sigma_g^2) + \alpha_{g,g}\ln(\sigma_g^2)^2\right] \\
& + \beta_2\ln(\sigma_\pi^2/\sigma_x^2) + (1/2)\beta_{2,2}\left[\ln(\sigma_\pi^2/\sigma_x^2)\right]^2 \\
& + \gamma_{e,2}\ln(\sigma_e^2)\times\ln(\sigma_\pi^2/\sigma_x^2) + \gamma_{g,2}\ln(\sigma_g^2)\times\ln(\sigma_\pi^2/\sigma_x^2) + v. \qquad (7)
\end{aligned}$$

The above function is normalized with respect to σ_x^2, but exactly the same econometric results would be obtained by normalizing on σ_π^2. Furthermore, provided that shifter σ_e^2 is exogenous, the presence of σ_x^2 (in the σ_π^2-σ_x^2-ratio) among the regressors does not make OLS estimates inconsistent (see Kumbhakar, 2012, in particular, for a formal treatment of this issue for a single shifter). Estimating (7) allows one to test whether a tradeoff between the variability of output and inflation actually exists in the data. Note that conditionally on the existence of this tradeoff, estimation of (7) 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 set-up one can use the variability of the output gap (respectively inflation) to model the variability of inflation (respectively output gap).

3. Empirical Implementation

Our goal is to estimate the inflation output variability tradeoff for a number of countries $i = 1, ..., N$ over time $t = 1, ..., T$ based on (7). However, the empirical implementation of (7) is not obvious. In principle, as said above, one can consistently estimate the inflation output variability tradeoff by a two-way fixed-effect model, i.e.,

$$\begin{aligned}
-\ln(\sigma^2_{x,i,t}) &= \alpha_e \ln(\sigma^2_{e,i,t}) + \alpha_g \ln(\sigma^2_{g,i,t}) \\
&\quad + (1/2)\left[\alpha_{e,e} \ln(\sigma^2_{e,i,t})^2 + 2\alpha_{e,g} \ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t}) + \alpha_{g,g} \ln(\sigma^2_{g,i,t})^2\right] \\
&\quad + \beta_2 \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) + (1/2)\beta_{2,2}\left[\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})\right]^2 + \gamma_{e,2} \ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) \\
&\quad + \gamma_{g,2} \ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t}) + \alpha_i + \delta_t + \varepsilon_{i,t}, \qquad (8)
\end{aligned}$$

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).

Yet, some major challenges emerge with regard to the estimation of (8). First, consistently with theoretical inflation output variability tradeoff, 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 derivation of structural supply and demand shocks for each country. We estimate the latter by SVARs. Moreover, we want to examine the effect of the Euro monetary policy on macroeconomic performance of Euro Area countries after they have adopted the Euro relative to a comparable set of countries without the Euro. Hence, we need to develop an identification strategy for the effect of the Euro on macroeconomic performance. We address these issues below.[12]

3.1. Data and Estimation of Structural Shocks

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 1984Q1-2018Q3. 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).

In the estimation of the structural supply shocks, we 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 shadows the effective federal funds rate, at the zero interest-rate lower bound it is aimed to represent 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

[12]We are aware that in principle this approach (see, e.g., Peersman, 2004; Olson and Enders, 2012) is subject to the Lucas critique. The empirical significance of this critique is however an unsettled issue. In a set-up comparable to ours, Olson and Enders (2012, p.1290) argue, relying on Favero and Hendry (1992), Ericsson and Irons (1995), Hendry (2002), Estrella and Fuhrer (2003), that the Lucas critique is likely to have little effect on the estimation of SVARs. We address the Lucas critique and related robustness concerns in Section 5.

use the money market rate until 1998Q4. Then, up to 2004Q3, we use the common ECB refinancing rate, and 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.[13]

In order to compute the variances of output gap and inflation deviation from target, we use a filter recently proposed by Hamilton (2018), which avoids the potentially spurious persistence in the cyclical component implied by the traditional Hodrick and Prescott (1997) (HP) filter. However, Hodrick (2020) shows that, while Hamilton's (2018) filter performs better in identifying the cyclical components in time series environments in which the first-differenced series is stationary, the HP performs better if the time series become more complex such as when it is generated from an unobserved components model. Hence, choosing among the two filters depends on the nature of the series under scrutiny, which may not always easy to ascertain. Given this uncertainty, we also assess robustness of our main results with respect to the choice of filter by employing a one-sided HP filter.[14]

Our main results are based on a measure for the output gap that is the difference in the log of real gross domestic product from its trend value computed through the Hamilton (2018) filter, while inflation deviations from target are calculated as the year-to-year percentage difference of the consumer price index (all items) minus its trend value computed through the same filter. Here we assume that the filter-measured trend is able to capture the 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

[13] 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 the unconventional policies.

[14] In contrast to the two-sided HP filter, the one-sided HP filter only uses current past states to compute the current observation.

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 HP filter is adopted).

Next, we base the derivation of the structural supply and demand shocks on a Bayesian VAR model with stochastic volatility (Cogley and Sargent, 2005). Taking stochastic volatility of the underlying VARs into account is especially important in our empirical setup since the variance of demand and supply shocks may change over time. The structural shocks are identified via sign restrictions. To identify the supply shock, we impose that the impulse response is positive for inflation, while it is negative for output throughout the first four quarters. In contrast, to identify the demand shock, we impose that the impulse response of both inflation and output are positive for the first four quarters. Our identification strategy may be viewed as another source of uncertainty surrounding our main results. For this reason, we also assess robustness of our main results with regard to the identification strategy. In this case, as an alternative we employ timing restrictions to derive the structural shocks. Notably, we identify the two shocks via a Cholesky decomposition, where we order inflation as first, output gap as second and the nominal interest rate as last variable. Further details on the specification of the VAR and the shock identification assumptions can be found in Appendix B.

Finally, it must be pointed out that the structural shocks and their variance are by construction orthogonal to the information set available to policy decision makers. It follows that the shock variances are exogenous shifters (regressors) in the sense assumed by Kumbhakar (2012). Accordingly equation (8) and its variants to be considered below can be consistently estimated through OLS.

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

It will be recalled that our basic observations are the variances of inflation deviations from target, output gaps, and structural supply and demand shocks. In order to compute these variances, we have divided the sample in the ten periods highlighted in Table 1. The intention is to compute variances over a sufficiently long time window. Our chosen sub-periods also seem to satisfy the need to single out interesting economic episodes. We end up with a panel dataset of dimension $N \times T = 20 \times 10 = 200$.

Our baseline specification is a DiD approach, where our aim is to infer whether the adoption of the Euro and its new monetary framework has on average improved the macroeconomic performance for the Euro Area countries. We augment the two-way fixed effect model (8) by a dummy $\mathcal{E}_{i,t}$, which is equal to zero for all countries and one for Euro countries when the policy is implemented. Equation (8) can be written more compactly as

$$Y_{i,t} = \alpha_i + \delta_t + \beta_{\mathcal{E}}\mathcal{E}_{i,t} + \Omega' X_{i,t} + \varepsilon_{i,t}, \tag{9}$$

where $\beta_{\mathcal{E}}$ represents our estimate of interest as it captures the impact of the Euro adoption on macroeconomic performance. $Y_{i,t}$ is the dependent variable, $X_{i,t}$ contains all the right hand side variables shown in equation (8) and Ω is 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. Under this assumption and with non-random policy assignment, the following conditional independence assumption (CIA) can be written:

$$E(Y_{0,i,t} \mid \alpha_i,\ \delta_t,\ X_{i,t},\ \mathcal{E}_{i,t}) = E(Y_{0,i,t} \mid \alpha_i,\ \delta_t,\ X_{i,t}). \tag{10}$$

Equation (10) states that unobserved confounders are fully captured by the two way fixed effect model. In other words, it means that the 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).[15]

When there is a treatment regarding a multiplicity of periods, one way to assess the appropriateness of the parallel trends assumption within the DiD is to allow for leads and lags of the treatment, which can be written as

$$Y_{i,t} = \alpha_i + \delta_t + \sum_{j=-m}^{q} \beta_{\mathcal{E},j} \mathcal{E}_{i,t=T_0+j} + \Omega' X_{i,t} + \varepsilon_{i,t}, \tag{11}$$

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 $\beta_{\mathcal{E},j} = 0 \ \forall j < 0$. This cannot be considered 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: first, the policy effect might have been anticipated by the economy, and thus cannot be safely ascribed to the policy itself; second, the parallel trends assumption is not a satisfactory basis for the identification of policy effects. In both cases, however, the estimation of (11) yields useful information on the appropriateness of the DiD approach.

Moreover, if the ATT is not constant over time after policy implementation, the modelling of $\beta_{\mathcal{E},j}, j \geq 0$, allows us to have estimates of the time-varying impact of the policy regime change. This specification is of great policy interest in our empirical application as

[15] The treatment effect identified by a non-random DiD is the ATT (see, e.g., Athey and Imbens, 2006).

it allows one to assess whether the impact of the monetary policy regime changes after the inception of the Great Recession.

At any rate supposing that the effects of unobserved confounders is fully controlled by the two way fixed effect model is often restrictive. Ashenfelter (1978) and Ashenfelter and Card (1985) have for example found, in a labor market context, that participants to a government-sponsored training program 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 adoption of the Euro may be subject to such considerations, we also resort to an alternative specification to corroborate our results.

In order to check the sensitivity of the ATT estimates, we therefore consider the LDV model, where it is possible to avoid reliance on the parallel trends assumption, and, at the same time, to control for past outcomes (O'Neill et al., 2016). The LDV model can generally be specified as follows

$$Y_{i,t} = \alpha_i + \delta_t + \beta_{\mathcal{E}}\mathcal{E}_{i,t} + \Omega' X_{i,t} + \theta' Y_{i,t-h} + \varepsilon_{i,t}. \tag{12}$$

In this case the sample starts 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 is

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

where we assume that conditional on past outcomes and time fixed effects, the potential outcomes are independent of the treatment status. Moreover, 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 $Y_{i,t-h}$) represent a proxy for controlling for unobserved time-varying heterogeneity.

Also in the LDV case, adopting the Autor (2003) multiple-effect framework makes it possible to analyze whether the Euro monetary policy has had heterogeneous effects through time. Equation (13) provides a companion to (11) for the LDV model,

$$Y_{i,t} = \delta_t + \sum_{j=0}^{q} \beta_{\mathcal{E},j} \mathcal{E}_{i,t=T_0+j} + \Omega' X_{i,t} + \theta' Y_{i,t-h} + \varepsilon_{i,t}. \tag{13}$$

The DiD and the LDV model are not nested. So, we cannot take one of the two as a special case of the other. But if they give broadly similar results, we might be more confident about evidence obtained on our estimate of interest.[16]

Finally, it is often argued that there are two rather distinct country groups within the Euro Area. The periphery is believed to be structurally different from the core in many aspects (see, e.g., De Grauwe and Ji, 2013). This could be a potential explanation of why monetary policies by the ECB face a problem of *'one size must fit all'* (Issing, 2001) rather than being characterized by *'one size fits all'* (Issing, 2005). In order to shed light upon this important matter, we conduct a subsample analysis for the core and periphery countries, and present results for these groups as well as for the Euro Area as a a whole. 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 countries in our sample that experienced a sovereign debt crisis.

[16] 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: $E(Y_{0,i,t} \mid \alpha_i,\ Y_{i,t-h},\ X_{i,t},\ \mathcal{E}_{i,t}) = E(Y_{0,i,t} \mid \alpha_i,\ Y_{i,t-h},\ X_{i,t})$. However, as discussed in Angrist and Pischke (2009), this combined approach requires very stringent econometric conditions for identification. In this empirical study, we therefore utilize the DiD and LDV approaches and compare the results obtained through each of them.

4. Main Results

The first three columns of Table 2 present the baseline results for the DiD specification (9). First, notice that the coefficient for the ratio of the variability of inflation to output gap, $\hat{\beta}_2$, is highly significant in all columns, in line with the considerations developed at the end of Section 2. Next, the coefficient for the non-linear term, $\hat{\beta}_{2,2}$, is not significantly different from zero. Yet, this is not necessarily evidence against a convex inflation output variability tradeoff for the Euro Area as a whole, as well as for the core and the periphery.[17] Second, the coefficient $\hat{\alpha}_e$ shows that the variance of the supply shock has a highly significant impact on the location of the tradeoff in all the Euro Area, the core and the periphery. 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.[18] Finally, demand shocks appear to have some impact on the location of the tradeoff in all the Euro Area, the core and the periphery. In sum, the significant coefficient estimates are evidence for the existence of an inflation output variability tradeoff for the Euro Area consistent with the theoretical tradeoff discussed in Section 2.

Most importantly, the coefficient for the dummy on Euro adoption, $\hat{\beta}_{\mathcal{E}}$, is highly significant for all the Euro Area, the core and the periphery. A negative sign for $\hat{\beta}_{\mathcal{E}}$ means that, on average, Euro adopters have been worse off due to adopting this monetary regime.

We can provide some further interpretation for the coefficient estimates. For instance, for $\hat{\beta}_{\mathcal{E}} = -0.586$ follows that $\exp(\hat{\beta}_{\mathcal{E}}) \approx 0.56$. 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 ≈ 1.80 and means that, conditionally on the supply and demand

[17] Based on Diewert and Wales (1987), we have assessed whether the estimated coefficients imply a violation of the convexity assumption for our sample. To that end, we have checked whether the observation-specific Hessian matrix with respect to the variances of inflation and output is negative semidefinite for each observation. This is the case for the vast majority of observations and we conclude that our estimates are consistent with a convex tradeoff.

[18] Appendix C contains details on the interpretation of the coefficient estimates.

shock variances, the Euro Area has a joint variance of inflation and output, which is around 80% larger than that of the control group. Likewise for the core and the periphery this variance was respectively 70% and 94% larger.

Finally, notice the p-values for the Ramsey (1969) Reset test. Under this test, rejection of the null provides clear evidence of misspecification, especially in the form of omitted variables (Godfrey and Orme, 1994). However, the null, i.e., omitted variables being orthogonal with respect to the included variables, is never rejected. Therefore, we cannot find evidence that estimates of equation (9) suffer from misspecification.

In the next three columns of Table 2, we present results for the specification (11), where we include lags of the treatment, along the lines of Autor (2003). This exercise allows us to provide a more articulated economic interpretation of the basic finding of a detrimental effect of the Euro. Consider the fourth column in Table 2, which relates to the Euro Area as a whole. Compared to the previous results, three observations stand out. First, there is again significant and well-specified evidence in favor of an inflation output variability tradeoff consistent with the theoretical tradeoff developed in Section 2. Second, coefficients $\hat{\beta}_{\mathcal{E},0}$ and, $\hat{\beta}_{\mathcal{E},2}$, $\hat{\beta}_{\mathcal{E},3}$ and $\hat{\beta}_{\mathcal{E},4}$ relating to periods including the Financial Crisis, the European Sovereign Debt Crisis, and the period of Mario Draghi's *'whatever it takes'* and OMTs announcements and the EAPP announcement and implementation are significant, while $\hat{\beta}_{\mathcal{E},5}$, relating to the time *after* this policy turnaround, is insignificant. We interpret these findings as evidence that the impact of the Euro adoption has changed over time. The inflation output variability tradeoff for the Euro Area countries is worse during the crisis period (relative to the control group) until the ECB's 2012 policy turnaround, but not thereafter. The coefficient estimates can be interpreted as follows: $\exp(\hat{\beta}_{\mathcal{E},0}) \approx 0.57$, $\exp(\hat{\beta}_{\mathcal{E},2}) \approx 0.47$, $\exp(\hat{\beta}_{\mathcal{E},3}) \approx 0.33$, and $\exp(\hat{\beta}_{\mathcal{E},4}) \approx 0.51$ imply that the Euro Area had a joint variance of inflation and output, which is more than 75% (111%, 200%, 95%) larger during 1999Q1 to 2002Q2 (2006Q1 to 2009Q2, 2009Q3 to 2012Q2, 2012Q3 to 2015Q3) compared to the control group.[19]

The fifth and sixth column in Table 2 relate to the analysis on the core and periphery. Their main features can be summed up by the following four remarks. First, throughout all specifications the estimated coefficients regarding the inflation output variability tradeoff are similar to the previous results. Second, again the Reset test does not give rise to concerns about omitted variables or specification problems. Third, the effect of the Euro on the core is less detrimental than for the Euro Area as a whole. There is again a detrimental effect of the Euro, but once we allow for heterogeneous effects of the treatment over time, the effect is observed in periods 8, but not in periods 9 and 10.

However, and fourth, the coefficient estimates for the periphery reveal differences vis-à-vis the former results. From 2009Q3 onwards 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), Georgiadis (2015) and Burriel and Galesi (2018). More crucially, the detrimental effect of the Euro ceases in the periphery only in period 10. Our findings can be interpreted as evidence that the core has benefited significantly earlier from the monetary policy turnaround occurring in 2012.

5. Robustness

The purpose of this section is threefold. First, we consider the LDV approach, which is an alternative identification strategy for the ATT. Second, we assess the robustness of our findings obtained via the DiD and LDV approach with regard to choice of the de-trending method. Third, we present an additional robustness analysis relating to the identification

[19] As anticipated in Subsection 3.2, we also tested the significance of the coefficients on the leads, $\hat{\beta}_{\mathcal{E},-3}$, $\hat{\beta}_{\mathcal{E},-2}$, and $\hat{\beta}_{\mathcal{E},-1}$. The null hypothesis that these coefficients are all zero is not rejected either for the baseline model, or for the robustness checks. Thus, anticipation effects and/or divergent trends between treated and control group appear to play no role in driving our basic findings.

scheme in the estimation of the structural shocks.[20]

5.1. Lagged Dependent Variable Approach

As already explained in Section 3, the LDV approach is one way to account for the possibility that policy choices such as the Euro adoption depend on past economic performance. This is indeed a plausible scenario and would imply an endogeneity bias for the DiD approach. Further advantages of the LDV approach are that it does not require the parallel trends assumption and that it controls for unobserved time-varying heterogeneity.

Inspection of the first, second and third column in Table 3, relating to the estimation of equation (12), reveals that our key findings are qualitatively unchanged vis-à-vis the previous ones. The estimates directly related to the inflation output variability tradeoff are consistent with the previous findings, but demand shocks have a larger effect on the tradeoff (in absolute value). Moreover, we find that the Euro dummy, $\hat{\beta}_{\mathcal{E}}$, is significantly different from zero for the Euro Area as a whole, the core, and the periphery. Clearly, this is further evidence that, relative to the control group, countries with the Euro faced a worse tradeoff. The usual calculation, $\exp(\hat{\beta}_{\mathcal{E}}) \approx 0.55$, implies a joint variance of inflation and output for all the Euro Area, which is more than 81% larger compared to the one in the control group. For the core this variance is 46% larger and for the periphery it is 130% larger. Thus, the quantitative implications of the treatment are in the same ballpark of those obtained under the parallel trends assumption. 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.

[20]Furthermore, in order to address the Lucas critique, we ran several Chow tests on (9) and the other specifications and found no evidence of parameter instability. The potential breakpoints were taken at the inception of the Euro, at the beginning of the Financial Crisis, and also at a previous stage when inflation targeting had gained preeminence among central bankers. These tests are available upon request.

Consider now the fourth, fifth and sixth column of Table 3, which provide the results from estimation of equation (13). Notice also that in this specification we cannot have ATT leads. As found above, the adoption of the Euro worsens the tradeoff at the occurrence of the European Sovereign Debt Crisis for the Euro Area as a whole, its core and its periphery. Next, this detrimental effect remains significant for the Euro Area as a whole and the periphery during the period, which starts with the Draghi announcement and includes the OMTs announcement and the EAPP announcement and implementation. Notice also, that the detrimental effect ceases already during this period for the core and that it starts already during the Financial Crisis for the periphery. These findings suggest that, relative to the control group, countries with the Euro faced a worse tradeoff during the crises, and that this result is to a large extent driven by the detrimental effect of the Euro on the periphery.

In sum, 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 modelled through a single variable or through a string of lagged variables.

5.2. Using the HP filter

Another potential concern to the above presented results is that they depend on a particular filter, i.e., the Hamilton (2018) filter. We address this concern by repeating our analysis with the HP filter. In Tables 4 and 5 we provide estimates for this robustness exercise. While the DiD coefficients of the first three columns in Table 4 directly related to the inflation output variability tradeoff are broadly consistent with the previous evidence, we now remarkably find that the Euro dummy, $\hat{\beta}_{\mathcal{E}}$, only approaches significance at the 10% level for the Euro Area as a whole and the periphery.

The LDV estimates from the first three columns in Table 5 are more in line with the previous evidence. The Euro dummy, $\hat{\beta}_{\mathcal{E}}$, is significantly different from zero for the Euro Area as a whole and the periphery. In these estimates, coefficients $\hat{\gamma}_{g,2}$ and $\hat{\gamma}_{e,2}$

are significant in some cases. However, evidence against the separability hypothesis is very weak.[21] The time patterns of the Euro coefficients from the last three columns in Tables 4 and 5 are not perfectly aligned with the corresponding coefficients in Tables 2 and 3. Also, more than in the previous tables, this kind of evidence differs across DiD and LDV estimates. In no case, however, we obtain evidence going against the main gist of the previous story, namely that the Euro Area's tradeoff worsens across the periods corresponding to the Financial Crisis and the European Sovereign Dect Crisis, and then improves again.

5.3. Shock Identification with Timing Restrictions

Finally, our results may depend on the identification strategy for the structural shocks. Recall that all results reported so far are based on imposing sign restrictions on the impulse response functions for inflation and output for four quarters. In this subsection, we assess, whether our results critically hinge on this particular identification approach. To this end, we repeat the analysis, but impose timing restrictions by applying the Cholesky decomposition when identifying the structural supply and demand shocks. Therefore, the sign and duration of the impulse response functions are unrestricted.

Tables 6 and 7 display the results for the DiD and LDV estimates. The coefficient estimates related to the inflation output variability tradeoff are in line with our previous estimates. Focusing on the treatment effects in the first three columns of both tables, it appears that our main result, the detrimental effect of the Euro, is fully confirmed by both the DiD and LDV estimates.

Likewise, columns four to six in both tables fully confirm the previously reported time patterns from Tables 2 and 3. The Euro Area as a whole experienced a worse tradeoff

[21] The separability hypothesis implies that the marginal rates of substitution between inputs are independent of the outputs in the transformation function. While this not crucial if there is just one output, the estimates' properties could on the other hand be affected in the case of multiple outputs (see on this Kumbhakar, 2013). Non-separability implies the joint significance of coefficients $\hat{\gamma}_{g,2}$ and $\hat{\gamma}_{e,2}$, which only occurs once. Thus, we can safely conclude that the separability hypothesis is supported by our estimates.

between inflation and output variability during the Financial and European Sovereign Debt Crisis relative to the control group. With Mario Draghi's *'whatever it takes'* and OMTs announcements and the EAPP announcement and implementation, this worsening of the tradeoff becomes insignificant for the core, but not for the periphery (and, as said above, for the Euro Area as a whole). Only after one more period the tradeoff significantly improves again in those cases. All in all, we conclude that our baseline findings are left untouched by the change in the identification strategy.

6. Discussion

Summing up, we consistently find a detrimental effect of the Euro on macroeconomic performance in period 8 (2009Q3 to 2012Q2), and, depending on the specification and filter also in periods 5 (1999Q1 to 2002Q2), 7 (2006Q1 to 2009Q2) and 9 (2012Q3 to 2015Q3). Moreover, we find that the detrimental effect of the Euro on macroeconomic performance during crises is more severe in peripheral countries and does already disappear in period 9 in the core, but not in the periphery.

It is worth emphasizing that these findings are conditional on controlling for country-specific supply and demand shocks in Euro Area and control group countries. Therefore, bad luck does not strike us as a reasonable explanation of our findings.[22]

We surmise an interpretation of these results in terms of a monetary policy for the Euro Area that can be denoted a *'one size must fit all'* (Issing, 2001) monetary policy during the start of the Euro Area (period 5) and a *'one size fits all'* policy in period 6 (2002Q3 to 2005Q4) as the stabilizing channels of a currency union become more effective (Issing, 2005). However, during the subsequent periods of crises (periods 7, 8 and 9) the

[22] It may be argued that in the context of our theoretical New Keynesian model above, 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 might no longer be offset by monetary policy. Yet, to the extent these other shocks move output and inflation in the same direction, we explicitly control for them, as our demand shocks are identified by imposing exactly this sign restriction.

ECB's policy can again be qualified as *'one size must fit all'*. Next, our finding that the detrimental effect of the Euro for the whole Euro Area definitely ceases in period 10 has two implications. First, it suggests that the detrimental effect during periods 7, 8 and 9 is directly related to periods of crises and so is our *'one size must fit all'* judgment. This is 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 signaling 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. Second, we interpret our finding that the detrimental effect of the Euro ceases for the whole Euro Area in period 10 meaning that the ECB finally 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 not only for the core, but also for the periphery. 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) and broadly consistent with the interpretation of events, monetary policies and evidence surveyed in Dell'Ariccia et al. (2018).

To gain further understanding on these issues, it is also interesting to consider, throughout periods 7 to 10, the evolution of the shadow rate developed in Wu and Xia (2016). We depict this rate for the US 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. On the other hand, the shadow rate for the Euro Area becomes consistently negative only since mid 2013 and lines up with the other rates only toward the end of period 9 (early 2015; of course since that period, the shadow rate of the US - but not of the UK - began to rise again). This stylized fact is broadly consistent with the interpretation we gave above

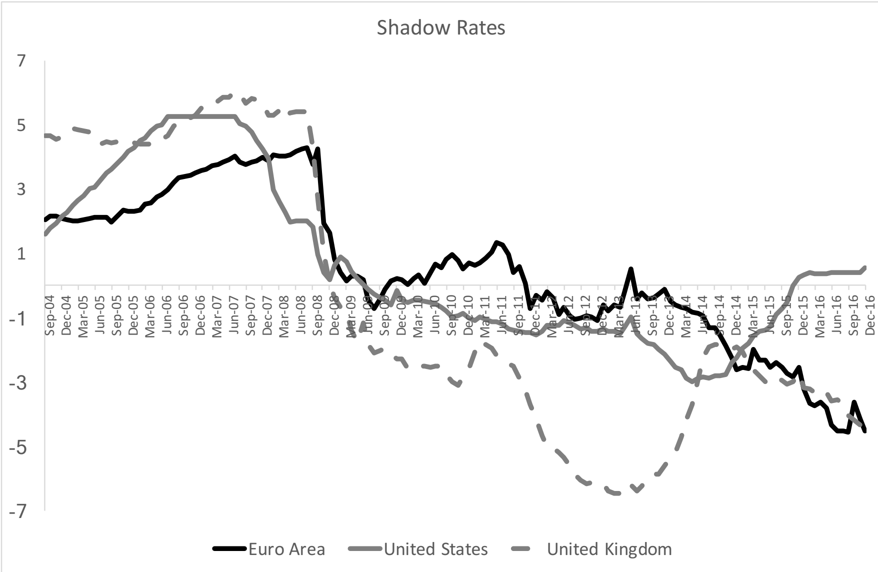


Figure 4: Shadow Rates

for our empirical evidence.

7. Concluding Remarks

This paper conducts a counterfactual analysis providing evidence that the inflation output variability tradeoff of Euro Area countries deteriorated during the periods of the Great Recession and the European Sovereign Debt Crisis. This deterioration was more severe and long-lasting for the peripheral countries of the Euro Area, while basically ceasing with the Draghi and OMTs announcements as well as with the EAPP announcement and implementation.

Our findings are based on a novel empirical strategy that is consistent with a theory-based inflation output variability tradeoff whose position is influenced by structural supply and demand shocks. We develop a panel data set for twenty OECD countries which consists of variances for output gap and inflation deviations from target as well as variances for the structural supply and demand shocks. The shock variances are estimated via a structural model that uses sign restrictions to identify shocks. In the estimation of the tradeoff, we model the joint determination of the variability of inflation and output by structural shocks through a transformation function taken from the quantitative analysis of production processes. The counterfactual evidence relating to ATT's is robust throughout various empirical specifications.

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. The detrimental effect attenuates and finally vanishes after the Draghi announcement. 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'*.

Most importantly, our more detailed analysis shows that the detrimental effect of the Euro was more severe for peripheral countries of the Euro Area. Moreover, while the Draghi and OMTs announcements as well as the EAPP announcement and implementation immediately had a beneficial effect on the macroeconomic performance of the core, this effect occurred much later for the periphery. Hence, structural differences between core and periphery countries are likely to be the driving force behind the impact of the Euro. Investigation about the nature and role of these structural differences must however be left to future research.

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Tables

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

Baseline choice of periods				
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	2012Q2	12	European Sovereign Debt Crisis
9	2012Q3	2015Q3	13	European Sovereign Debt Crisis (continued) Draghi announcement (July 26th, 2012) Outright Monetary Transactions (OMTs) announcement (September 6th, 2012) Expanded Asset Purchase Programme (EAPP) (January 22nd, 2015)
10	2015Q4	2018Q3	12	

Table 2: Estimated parameters for all countries, core and periphery (with interaction terms), for DiD model.[a] Observations are based on the Hamilton (2018) filter. Shocks are identified with sign restrictions.

Variables	Coefficient	Estimates[b]					
		(9)			(11)		
		All	Core	Periphery	All	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.586**	-0.532*	-0.663*			
		(0.266)	(0.297)	(0.343)			
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$				-0.561*	-0.654*	-0.400
					(0.290)	(0.341)	(0.390)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$				-0.404	-0.660	-0.132
					(0.390)	(0.484)	(0.417)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$				-0.748*	-0.685	-0.847
					(0.427)	(0.461)	(0.540)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$				-1.099***	-1.097**	-1.128**
					(0.377)	(0.397)	(0.441)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$				-0.668*	-0.223	-1.515***
					(0.340)	(0.369)	(0.265)
$\mathcal{E}_{i,T_0+5}$	$\beta_{\mathcal{E},5}$				-0.104	-0.196	-0.267
					(0.300)	(0.333)	(0.581)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.405***	-0.407***		-0.425***	-0.393***	
		(0.085)	(0.085)		(0.081)	(0.086)	
$\ln(\sigma^2_{g,i,t})$	α_g	-0.185*	-0.192*		-0.135	-0.152	
		(0.096)	(0.101)		(0.101)	(0.105)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t})$	$\alpha_{e,g}$	0.377*	0.373*		0.391*	0.256	
		(0.181)	(0.183)		(0.197)	(0.191)	
$\ln(\sigma^2_{e,i,t})^2$	$\alpha_{e,e}$	-0.483***	-0.482***		-0.463***	-0.453***	
		(0.130)	(0.133)		(0.142)	(0.140)	
$\ln(\sigma^2_{g,i,t})^2$	$\alpha_{g,g}$	-0.227*	-0.229*		-0.204	-0.124	
		(0.130)	(0.131)		(0.156)	(0.157)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.535***	0.533***		0.548***	0.525***	
		(0.046)	(0.046)		(0.045)	(0.042)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,2}$	0.007	0.009		-0.004	0.017	
		(0.042)	(0.041)		(0.037)	(0.041)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{e,2}$	-0.009	-0.012		-0.025	-0.058	
		(0.059)	(0.060)		(0.061)	(0.067)	
$\ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{g,2}$	-0.093	-0.090		-0.104	-0.055	
		(0.070)	(0.070)		(0.070)	(0.081)	
Country fixed effect		yes	yes		yes	yes	
Time fixed effect		yes	yes		yes	yes	
N		20	20		20	20	
Number of observations		200	200		200	200	
R^2		0.843	0.843		0.850	0.863	
Specification tests[c]:							
Ramsey (1969) Reset		0.127	0.118		0.141	0.213	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$					0.737	0.547	
$\beta_{\mathcal{E},0} = \beta_{\mathcal{E},1} = \cdots = \beta_{\mathcal{E},5} = 0$					0.003	0.001	

[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 all countries, core and periphery, for LDV model.[a] Observations are based on the Hamilton (2018) filter. Shocks are identified with sign restrictions.

Variables	Coefficient	Estimates[b]					
		(12)			(13)		
		All	Core	Periphery	All	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.593***	-0.379**	-0.834***			
		(0.195)	(0.167)	(0.247)			
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$				-0.657**	-0.599**	-0.556
					(0.279)	(0.262)	(0.330)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$				-0.323	-0.416	-0.196
					(0.322)	(0.298)	(0.424)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$				-0.459	-0.171	-0.805**
					(0.317)	(0.329)	(0.344)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$				-0.922**	-0.756*	-1.191***
					(0.358)	(0.392)	(0.365)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$				-0.770**	-0.152	-1.858***
					(0.315)	(0.247)	(0.190)
$\mathcal{E}_{i,T_0+5}$	$\beta_{\mathcal{E},5}$				-0.386	-0.461	-0.683
					(0.278)	(0.297)	(0.430)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.425***	-0.460***		-0.409***	-0.375**	
		(0.118)	(0.134)		(0.138)	(0.144)	
$\ln(\sigma^2_{g,i,t})$	α_g	-0.469***	-0.441***		-0.435**	-0.445**	
		(0.150)	(0.145)		(0.174)	(0.165)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t})$	$\alpha_{e,g}$	0.449*	0.431*		0.529**	0.276	
		(0.231)	(0.227)		(0.224)	(0.220)	
$\ln(\sigma^2_{e,i,t})^2$	$\alpha_{e,e}$	-0.546***	-0.552***		-0.513***	-0.439**	
		(0.153)	(0.171)		(0.174)	(0.183)	
$\ln(\sigma^2_{g,i,t})^2$	$\alpha_{g,g}$	-0.386**	-0.420***		-0.420***	-0.323**	
		(0.138)	(0.143)		(0.141)	(0.148)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.638***	0.611***		0.650***	0.608***	
		(0.057)	(0.063)		(0.052)	(0.043)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,2}$	0.109***	0.090**		0.110***	0.121***	
		(0.038)	(0.043)		(0.036)	(0.039)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{e,2}$	0.093	0.079		0.092	0.018	
		(0.094)	(0.099)		(0.092)	(0.102)	
$\ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{g,2}$	-0.020	-0.030		-0.022	0.062	
		(0.073)	(0.071)		(0.072)	(0.074)	
Country fixed effect		no	no		no	no	
Time fixed effect		yes	yes		yes	yes	
N		20	20		20	20	
Number of observations		120	120		120	120	
R^2		0.890	0.894		0.895	0.919	
Specification tests[c]:							
Ramsey (1969) Reset		0.059	0.033		0.068	0.139	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$							
$\beta_{\mathcal{E},0} = \beta_{\mathcal{E},1} = \cdots = \beta_{\mathcal{E},5} = 0$							

[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 for all countries, core and periphery, for DiD model.[a] Observations are based on the HP filter. Shocks are identified with sign restrictions.

Variables	Coefficient	Estimates[b]					
		(9)			(11)		
		All	Core	Periphery	All	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.377	-0.306	-0.493			
		(0.229)	(0.256)	(0.297)			
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$				-0.478*	-0.512	-0.459
					(0.269)	(0.327)	(0.326)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$				-0.288	-0.482	-0.001
					(0.295)	(0.361)	(0.326)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$				-0.630*	-0.649*	-0.570
					(0.341)	(0.348)	(0.431)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$				-0.578*	-0.588*	-0.509
					(0.322)	(0.314)	(0.460)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$				-0.240	0.266	-1.335***
					(0.472)	(0.486)	(0.463)
$\mathcal{E}_{i,T_0+5}$	$\beta_{\mathcal{E},5}$				0.138	0.205	-0.149
					(0.356)	(0.335)	(0.567)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.428***	-0.429***		-0.428***	-0.388***	
		(0.065)	(0.064)		(0.062)	(0.064)	
$\ln(\sigma^2_{g,i,t})$	α_g	-0.194*	-0.203*		-0.189*	-0.189*	
		(0.108)	(0.105)		(0.107)	(0.092)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t})$	$\alpha_{e,g}$	0.362	0.343		0.352	0.355	
		(0.220)	(0.212)		(0.245)	(0.219)	
$\ln(\sigma^2_{e,i,t})^2$	$\alpha_{e,e}$	-0.227*	-0.220*		-0.231*	-0.251**	
		(0.111)	(0.107)		(0.125)	(0.118)	
$\ln(\sigma^2_{g,i,t})^2$	$\alpha_{g,g}$	-0.224	-0.216		-0.135	-0.152	
		(0.153)	(0.152)		(0.172)	(0.165)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.530***	0.529***		0.527***	0.470***	
		(0.048)	(0.049)		(0.054)	(0.053)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,2}$	0.066**	0.066**		0.059**	0.074***	
		(0.024)	(0.024)		(0.024)	(0.025)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{e,2}$	0.018	0.013		0.027	-0.007	
		(0.061)	(0.058)		(0.070)	(0.059)	
$\ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{g,2}$	-0.097	-0.088		-0.118**	-0.101**	
		(0.059)	(0.059)		(0.055)	(0.044)	
Country fixed effect		yes	yes		yes	yes	
Time fixed effect		yes	yes		yes	yes	
N		20	20		20	20	
Number of observations		200	200		200	200	
R^2		0.859	0.860		0.864	0.881	
Specification tests[c]:							
Ramsey (1969) Reset		0.172	0.187		0.381	0.997	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$					0.680	0.791	
$\beta_{\mathcal{E},0} = \beta_{\mathcal{E},1} = \cdots = \beta_{\mathcal{E},5} = 0$					0.155	0.018	

[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 all countries, core and periphery, for LDV model.[a] Observations are based on the HP filter. Shocks are identified with sign restrictions.

Variables	Coefficient	Estimates[b]					
		(12)			(13)		
		All	Core	Periphery	All	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.476*	-0.179	-0.630***			
		(0.274)	(0.222)	(0.194)			
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$				-0.625	-0.480	-0.749**
					(0.388)	(0.391)	(0.319)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$				-0.479	-0.493	-0.261
					(0.343)	(0.332)	(0.435)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$				-0.508	-0.253	-0.644**
					(0.344)	(0.325)	(0.301)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$				-0.620**	-0.446	-0.508
					(0.288)	(0.338)	(0.341)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$				-0.292	0.262	-1.515***
					(0.516)	(0.458)	(0.413)
$\mathcal{E}_{i,T_0+5}$	$\beta_{\mathcal{E},5}$				-0.059	0.217	-0.477
					(0.350)	(0.348)	(0.333)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.665***	-0.681***		-0.674***	-0.590***	
		(0.096)	(0.097)		(0.094)	(0.100)	
$\ln(\sigma^2_{g,i,t})$	α_g	-0.364**	-0.391***		-0.379**	-0.420***	
		(0.142)	(0.133)		(0.145)	(0.121)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t})$	$\alpha_{e,g}$	0.783***	0.701**		0.769**	0.706**	
		(0.252)	(0.250)		(0.289)	(0.275)	
$\ln(\sigma^2_{e,i,t})^2$	$\alpha_{e,e}$	-0.599***	-0.603***		-0.618**	-0.610**	
		(0.204)	(0.198)		(0.231)	(0.256)	
$\ln(\sigma^2_{g,i,t})^2$	$\alpha_{g,g}$	-0.633***	-0.606***		-0.553***	-0.605***	
		(0.153)	(0.148)		(0.149)	(0.136)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.589***	0.575***		0.587***	0.455***	
		(0.071)	(0.073)		(0.079)	(0.064)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,2}$	0.104*	0.101*		0.094	0.075*	
		(0.050)	(0.049)		(0.055)	(0.039)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{e,2}$	0.131*	0.124*		0.134*	0.073	
		(0.070)	(0.070)		(0.075)	(0.064)	
$\ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{g,2}$	-0.130*	-0.112		-0.142*	-0.135**	
		(0.067)	(0.070)		(0.072)	(0.049)	
Country fixed effect		no	no		no	no	
Time fixed effect		yes	yes		yes	yes	
N		20	20		20	20	
Number of observations		120	120		120	120	
R^2		0.891	0.899		0.895	0.921	
Specification tests[c]:							
Ramsey (1969) Reset		0.739	0.877		0.932	0.735	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$							
$\beta_{\mathcal{E},0} = \beta_{\mathcal{E},1} = \cdots = \beta_{\mathcal{E},5} = 0$							

[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 6: Estimated parameters for all countries, core and periphery, for DiD model.[a] Observations are based on the Hamilton (2018) filter. Shocks are identified with timing restrictions.

Variables	Coefficient	Estimates[b]					
		(9)			(11)		
		All	Core	Periphery	All	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.701**	-0.668**	-0.750**			
		(0.248)	(0.274)	(0.321)			
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$				-0.694**	-0.752**	-0.548
					(0.315)	(0.349)	(0.403)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$				-0.511	-0.717	-0.183
					(0.376)	(0.465)	(0.367)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$				-0.739**	-0.697**	-0.795*
					(0.286)	(0.294)	(0.403)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$				-1.111***	-1.021***	-1.158***
					(0.314)	(0.333)	(0.386)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$				-0.745**	-0.333	-1.481***
					(0.350)	(0.372)	(0.351)
$\mathcal{E}_{i,T_0+5}$	$\beta_{\mathcal{E},5}$				-0.323	-0.426	-0.352
					(0.385)	(0.406)	(0.721)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.457***	-0.459***		-0.469***	-0.424***	
		(0.063)	(0.063)		(0.063)	(0.068)	
$\ln(\sigma^2_{g,i,t})$	α_g	-0.254***	-0.258***		-0.214***	-0.242***	
		(0.064)	(0.067)		(0.066)	(0.084)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t})$	$\alpha_{e,g}$	-0.023	-0.026		0.003	0.005	
		(0.200)	(0.202)		(0.225)	(0.223)	
$\ln(\sigma^2_{e,i,t})^2$	$\alpha_{e,e}$	-0.182	-0.177		-0.140	-0.160	
		(0.173)	(0.184)		(0.168)	(0.214)	
$\ln(\sigma^2_{g,i,t})^2$	$\alpha_{g,g}$	-0.072	-0.078		-0.086	-0.106	
		(0.086)	(0.084)		(0.099)	(0.100)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.592***	0.591***		0.605***	0.570***	
		(0.043)	(0.042)		(0.048)	(0.058)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,2}$	0.007	0.007		0.012	0.026	
		(0.035)	(0.034)		(0.033)	(0.037)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{e,2}$	-0.017	-0.017		-0.042	-0.037	
		(0.051)	(0.050)		(0.050)	(0.053)	
$\ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{g,2}$	-0.065	-0.066		-0.052	-0.052	
		(0.056)	(0.056)		(0.057)	(0.058)	
Country fixed effect		yes	yes		yes	yes	
Time fixed effect		yes	yes		yes	yes	
N		20	20		20	20	
Number of observations		200	200		200	200	
R^2		0.861	0.861		0.865	0.876	
Specification tests[c]:							
Ramsey (1969) Reset		0.354	0.349		0.253	0.202	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$					0.872	0.706	
$\beta_{\mathcal{E},0} = \beta_{\mathcal{E},1} = \cdots = \beta_{\mathcal{E},5} = 0$					0.020	0.040	

[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 all countries, core and periphery, for LDV model.[a] Observations are based on the Hamilton (2018) filter. Shocks are identified with timing restrictions.

Variables	Coefficient	Estimates[b]					
		(12)			(13)		
		All	Core	Periphery	All	Core	Periphery
$\mathcal{E}_{i,t}$	$\beta_{\mathcal{E}}$	-0.572***	-0.398***	-0.750***			
		(0.143)	(0.119)	(0.183)			
$\mathcal{E}_{i,T_0}$	$\beta_{\mathcal{E},0}$				-0.493	-0.522*	-0.337
					(0.294)	(0.256)	(0.432)
$\mathcal{E}_{i,T_0+1}$	$\beta_{\mathcal{E},1}$				-0.290	-0.428	-0.069
					(0.268)	(0.268)	(0.376)
$\mathcal{E}_{i,T_0+2}$	$\beta_{\mathcal{E},2}$				-0.538***	-0.385	-0.763***
					(0.185)	(0.227)	(0.192)
$\mathcal{E}_{i,T_0+3}$	$\beta_{\mathcal{E},3}$				-0.875***	-0.611*	-1.073***
					(0.243)	(0.346)	(0.211)
$\mathcal{E}_{i,T_0+4}$	$\beta_{\mathcal{E},4}$				-0.696**	-0.135	-1.765***
					(0.292)	(0.253)	(0.162)
$\mathcal{E}_{i,T_0+5}$	$\beta_{\mathcal{E},5}$				-0.407	-0.484	-0.548
					(0.287)	(0.287)	(0.436)
$\ln(\sigma^2_{e,i,t})$	α_e	-0.408***	-0.429***		-0.402***	-0.288***	
		(0.083)	(0.089)		(0.079)	(0.081)	
$\ln(\sigma^2_{g,i,t})$	α_g	-0.430**	-0.426**		-0.345**	-0.419**	
		(0.153)	(0.152)		(0.160)	(0.174)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{g,i,t})$	$\alpha_{e,g}$	0.217	0.126		0.264	0.237	
		(0.244)	(0.250)		(0.260)	(0.216)	
$\ln(\sigma^2_{e,i,t})^2$	$\alpha_{e,e}$	-0.491***	-0.454***		-0.476***	-0.470***	
		(0.132)	(0.133)		(0.146)	(0.145)	
$\ln(\sigma^2_{g,i,t})^2$	$\alpha_{g,g}$	-0.209	-0.198		-0.191	-0.252	
		(0.184)	(0.180)		(0.206)	(0.174)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	β_2	0.607***	0.598***		0.635***	0.546***	
		(0.072)	(0.076)		(0.085)	(0.074)	
$\ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})^2$	$\beta_{2,2}$	0.110*	0.094		0.132**	0.147**	
		(0.058)	(0.060)		(0.062)	(0.066)	
$\ln(\sigma^2_{e,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{e,2}$	0.041	0.034		0.029	0.024	
		(0.062)	(0.064)		(0.060)	(0.062)	
$\ln(\sigma^2_{g,i,t}) \times \ln(\sigma^2_{\pi,i,t}/\sigma^2_{x,i,t})$	$\gamma_{g,2}$	-0.002	0.001		0.032	0.028	
		(0.101)	(0.104)		(0.111)	(0.100)	
Country fixed effect		no	no		no	no	
Time fixed effect		yes	yes		yes	yes	
N		20	20		20	20	
Number of observations		120	120		120	120	
R^2		0.898	0.901		0.901	0.925	
Specification tests[c]:							
Ramsey (1969) Reset		0.423	0.291		0.460	0.409	
$\beta_{\mathcal{E},-3} = \beta_{\mathcal{E},-2} = \beta_{\mathcal{E},-1} = 0$							
$\beta_{\mathcal{E},0} = \beta_{\mathcal{E},1} = \cdots = \beta_{\mathcal{E},5} = 0$							

[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.

Online Appendix

Does One Size Fit All in the Euro Area? Some Counterfactual Evidence.

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A. More on the Theoretical Framework

Optimal Monetary Policy. Assume that the aggregate economy is best approximated by a standard New Keynesian model under the rational expectations hypothesis., i.e.,

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

$$\pi_t = \beta E_t \pi_{t+1} + \lambda x_t + e_t. \tag{A.2}$$

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. We assume $g_t \sim \text{iid}(0, \sigma_g^2)$, $e_t \sim \text{iid}(0, \sigma_e^2)$ and $\sigma_{eg} = 0$.[1]

Consider optimal monetary policy under discretion (as elaborated in Clarida et al., 1999). The central bank minimizes (1) subject to (A.2) in each period. The first-order necessary condition is $\pi_t = -(\omega_x/\lambda)x_t$ and one can show that, under this policy, the model implies an inflation output variability tradeoff as first developed in Taylor (1979). In particular, solving the model for given parameters implies a minimum state variable solution $\pi_t = a_\pi e_t$ and $x_t = a_x e_t$, where $a_\pi \equiv \omega_x/(\omega_x + \lambda^2)$ and $a_x \equiv -\lambda/(\omega_x + \lambda^2)$. This implies the long-run relationships in unconditional variances (2) and (3).

[1] In the theoretical literature e_t is usually denoted a cost-push shock. Notice that allowing for autocorrelation in the exogenous shocks would not alter any conclusion.

TAYLOR (1993) RULE. It is easy to verify that the model (A.1), (A.2) and (4) has the solution $\pi_t = b_{\pi,e}e_t + b_{\pi,g}g_t$ and $x_t = b_{x,e}e_t + b_{x,g}g_t$, where $b_{\pi,e} \equiv (1 + \sigma^{-1}\phi_\pi\lambda)^{-1}$, $b_{\pi,g} \equiv (1 + \sigma^{-1}\phi_\pi\lambda)^{-1}\lambda$, $b_{x,e} \equiv -\sigma^{-1}\phi_\pi/(1 + \sigma^{-1}\phi_\pi\lambda)$ and $b_{x,g} \equiv (1 + \sigma^{-1}\phi_\pi\lambda)^{-1}$. This solution implies the long-run relationships in unconditional variances given by (5) and (6).

INPUT-ORIENTED TTF. Next we show that the input-oriented TTF provides us with a functional form that captures the basic characteristics of an inflation output variability tradeoff. Formally, one assumes that the relationship between the supply and demand shock (as the $M = 2$ outputs, y) and the variances of inflation and output gap (as the $K = 2$ inputs, z) can be described by $Af(y, z) = 1$, where we have M outputs 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.,

$$\begin{aligned}
\ln(f(y,z)) = &\sum_m \alpha_m \ln(y_m) + \frac{1}{2}\sum_m\sum_n \alpha_{m,n}\ln(y_m)\times\ln(y_n) \\
&+ \sum_k \beta_k \ln(z_k) + \frac{1}{2}\sum_k\sum_l \beta_{k,l}\ln(z_k)\times\ln(z_l) \\
&+ \sum_m\sum_k \gamma_{m,k}\ln(y_m)\times\ln(z_k),
\end{aligned} \tag{A.3}$$

where the following symmetry is imposed: $\beta_{k,l} = \beta_{l,k}$ and $\gamma_{m,n} = \gamma_{n,m}$. Equation (A.3) requires $M + K + 2$ 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 tradeoff, we have simultaneous endogeneity of σ_π^2 and σ_x^2, while σ_e^2 and σ_g^2 are exogenous, we can consider the former two variances as inputs, while the latter two variances are the outputs. Therefore, we adopt a normalization with respect to an input. This gives rise to an input-oriented TTF. Following Kumbhakar (2012), we rewrite (A.3) as

$$
\begin{aligned}
\ln(f(y,z)) = & \sum_m \alpha_m \ln(y_m) + \frac{1}{2}\sum_m \sum_n \alpha_{m,n} \ln(y_m) \times \ln(y_n) \\
& + \sum_k \beta_k \ln(z_k/z_1) + \frac{1}{2}\sum_k \sum_l \beta_{k,l} \ln(z_k/z_1) \times \ln(z_l/z_1) \\
& + \sum_m \sum_k \gamma_{m,k} \ln(y_m) \times \ln(z_k/z_1) + \Upsilon,
\end{aligned}
$$

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 (A.3) in ratios (see, e.g., Kumbhakar, 2012, for the details).

Next we impose the normalization restrictions, $\sum_k \beta_k = 1$, $\sum_l \beta_{k,l} = 0$ $\forall k$, and, $\sum_k \gamma_{m,k} = 0$ $\forall m$.[2] As a consequence, the composite term Υ is eliminated and we obtain the input-oriented TTF that we use as our empirical specification

$$
\begin{aligned}
-\ln(z_1) = & \alpha_0 + \sum_m \alpha_m \ln(y_m) + \frac{1}{2}\sum_m \sum_n \alpha_{m,n} \ln(y_m) \times \ln(y_n) \\
& + \sum_{k=2} \beta_k \ln(z_k/z_1) + \frac{1}{2}\sum_{k=2} \sum_{l=2} \beta_{k,l} \ln(z_k/z_1) \times \ln(z_l/z_1) \\
& + \sum_m \sum_k \gamma_{m,k} \ln(y_m) \times \ln(z_k/z_1) + v,
\end{aligned}
$$

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 tradeoff, we have $y_1 = \sigma_e^2$, $y_2 = \sigma_g^2$, $z_1 = \sigma_x^2$ and $z_2 = \sigma_\pi^2$.

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

B. Estimation of Structural Shocks

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

$$\mathcal{Y}_t = \sum_{i=1}^{p} A_i \mathcal{Y}_{t-i} + C\mathcal{X}_t + u_t, \tag{B.1}$$

where $\mathcal{Y}_t$ is a 3×1 vector of endogenous variables including a measure of the output gap, inflation (as difference from target) and nominal interest rate. 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). Therefore, the models for the other countries in the sample include $\mathcal{X}_t$, which is a 3×1 vector of exogenous variables including US output gap, inflation (as difference from target) and a nominal interest rate, aimed at capturing the world macroeconomic stance. $u_t \sim N(0, \Sigma_{u,t})$ is a vector of reduced-form disturbances with $E[u_t] = 0$ and $E[u_t u_t^{'}] = \Sigma_{u,t}$. u_t is independently but not identically distributed across time. $\Sigma_{u,t}$ is time-varying, rendering volatility stochastic and introducing heteroskedasticity. We model stochastic volatility as in Cogley and Sargent (2005). It is assumed that $\Sigma_{u,t}$ can be decomposed as

$$\Sigma_{u,t} = F \Lambda_t F^{'},$$

where F is a lower triangular matrix with ones on the main diagonal. Λ_t is a time-varying diagonal matrix equal to $(\overline{s_1}\exp(\lambda_{1,t}), \overline{s_2}\exp(\lambda_{2,t}), ...\overline{s_n}\exp(\lambda_{n,t}))$, where n is the number of endogenous variables. $\overline{s_1}, \overline{s_2}, ...\overline{s_n}$ are known scaling terms and $\lambda_{1,t}, \lambda_{2,t}, ...\lambda_{n,t}$ are dynamic processes generating the heteroskedasticity that are characterized by the autoregressive process

$$\lambda_{i,t} = \gamma\lambda_{i,t-1} + \nu_{i,t} \qquad \nu_{i,t} \sim N(0, \phi_i).$$

The parameters to be estimated are: the parameters of the reduced form VAR, the elements of the F matrix, the dynamic coefficients $\lambda_{i,t}$ and the heteroskedasticity parameters ϕ_i. With regard to the priors, for A_i, C and covariance matrix Ω_0 we adopt classical Minnesota priors. Considering the inverse of F, that is, F^{-1} we adopt a multinormal diffuse prior with mean f_{i0}^{-1}(which is set as a vector of zeros) and covariance diagonal matrix Υ_{i0} with large diagonal entries. The prior for $\pi(\lambda_i \mid \phi_i)$ are not simple to formulate since each term $\lambda_{i,t}$ depends on its previous value. The solution to this problem is given by separating $\pi(\lambda_i \mid \phi_i)$ into T different priors, with prior in each period t conditional on period $t-1$. To obtain the conditional posterior for ϕ_i, the previous prior will be combined with a joint prior for $\lambda_{i,1}, ..., \lambda_{i,T}$, since the joint formulation is faster for ϕ_i. We make 5000 draws of which the first 1000 are discarded as burn in draws.

The identification structural shocks requires to impose restrictions. Our preferred choice in this paper are sign restrictions. Uhlig (2005) and others, show how to obtain identification of the above VAR (B.1) by imposing sign restrictions on a (sub)set of the variables responses to shocks as discussed in the main text. An 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 (2005) shows that any impulse vector a can be recovered if there is an $n-$dimensional vector q of unit length such that $a = \widetilde{A}q$, where $\widetilde{A}$ is the Cholesky factor of $\Sigma_{u,t}$. More precisely, starting with estimation of the above reduced form model, identification of a single shock by sign restrictions 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.

For robustness purposes we also consider zero short run restrictions via Cholesky decomposition of Σ. Exact identification requires that $(n^2 - n)/2$ restrictions must be placed between the regression residuals and structural innovations. Given that the Cholesky decomposition is triangular, it forces exactly $(n^2 - n)/2$ elements of the matrix of contemporaneous relationships to be zero. The resulting recursive structure impose a causal ordering on the variables in the VAR: shocks to one equation contemporaneously affect variables below that equation but only affect variables above that equation with a lag. With this interpretation in mind, the causal ordering one chooses reflects his beliefs about the relationships among variables in the VAR.

C. Details on the Interpretation of Coefficient Estimates

$\hat{\boldsymbol{\beta}}_{\mathbf{2}}$: Ignore all other terms in (9) except for the one involving $\hat{\beta}_2$, thus

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

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 tradeoff. The equation 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.

$\hat{\boldsymbol{\alpha}}_{\boldsymbol{e}}$: Ignore all other terms in (9) apart from the ones involving $\hat{\alpha}_e$ and $\hat{\beta}_2$, thus

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

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 the equations above 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. The same arguments and conclusions apply to $\hat{\boldsymbol{\alpha}}_{\boldsymbol{g}} < 0$.

$\hat{\boldsymbol{\beta}}_{\boldsymbol{\mathcal{E}}}$: Similar arguments as above, yield

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

and, as $\hat{\beta}_{\mathcal{E}} < 0$ and $\hat{\beta}_2 \in [0,1)$, it follows that $[\hat{\beta}_{\mathcal{E}}/(-1+\hat{\beta}_2)], -[\hat{\beta}_{\mathcal{E}}/\hat{\beta}_2] > 0$. Therefore, a significant coefficient estimate $\hat{\beta}_{\mathcal{E}} < 0$ is interpreted as a deterioration of the tradeoff.