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INFLATION DYNAMICS AND THE OUTPUT-INFLATION TRADE-OFF: INTERNATIONAL PANEL DATA EVIDENCE

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Abstract

We explore the impact of inflation and its variability on the output-inflation trade-off using a unified single-step approach in a panel data context. A limitation of earlier empirical approaches is that they focus on either cross-country or country-by-country time-series analyses. This paper employs a dynamic heterogeneous panel data specification and uses an all-encompassing estimation framework that accounts for parameter heterogeneity, cross-sectional dependence, dynamics, and non-stationarity. Our sample covers 60 countries from 1970 to 2010. While inflation variability reduces the trade-off for specific periods and country groups, an unambiguous and more pronounced negative relation emerges between the inflation rate and the responsiveness of real output to nominal shocks. The findings are in line with the New Keynesian view of a negative association between the rate of inflation and the output-inflation trade-off, as well as with the observed flattening of the Phillips curve over the past decades.

Keywords: *Output-Inflation Trade-off, Inflation Dynamics, Dynamic Panel Data, Parameter Heterogeneity, Cross-Sectional Dependence.*

JEL Classification: *E31, E12, C23.*

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1 Introduction

The slope of the Phillips curve (or output-inflation trade-off) constitutes one of the most intensively explored empirical relationships in modern macroeconomics (Mishkin, 2007; Ball and Mazumder, 2011; Gordon, 2011; among others).[1] This is not surprising, since the related empirical evidence can (in)validate contrasting views on the real effects of nominal aggregate demand shocks. In addition, the recent economic crisis and the puzzling behavior of inflation that followed have further fueled the interest in the output-inflation trade-off (e.g., Arias *et al.*, 2016; Blanchard, 2016; Laseen and Sanjani, 2016).[2] For example, some analysts proclaim the accelerationist hypothesis dead (Krugman, 2015), others find that the puzzle can disappear depending on the modeling of inflation expectations (Gordon, 2013; Coibion and Gorodnichenko, 2015), while a spate of recent contributions document a flattening of the Phillips curve since the period of Great Moderation (Roberts, 2006; Mishkin, 2007; De Veirman, 2009; Kuttner and Robinson, 2010).

Recent theoretical (Benigno and Ricci, 2011; Eggertsson and Giannoni, 2013; Daly and Hobijn, 2014; Alex Ho and Yetman, 2008) and empirical contributions (Abbott and Martinez, 2008; De Veirman, 2009; Fendel and Rulke, 2012; Sun, 2014) consider the output-inflation trade-off and its implications for the real effects of nominal aggregate demand shocks. A number of papers show that significant changes occurred in the slope of the short-run Phillips curve (e.g., Mishkin, 2007; Kuttner and Robinson, 2010; Ball and Mazumder, 2011; Davig, 2016), suggesting the flattening of the Phillips curve (Stock and Watson, 2010; Matheson and Stavrev, 2013; Simon *et al.*, 2013), as well as changes in the dynamics of inflation persistence (Carlstrom *et al.*, 2009) and inflation volatility (Summers, 2005). Moreover, the changing shape of the Phillips curve may reflect time variation in the slope, which motivates further scrutiny on the determinants of the trade-off (e.g., Kuttner and Robinson, 2010; Ball and Mazumder, 2011; Murphy, 2014).

This paper considers the effect of inflation dynamics (the rate of inflation and its variability) on the slope of the trade-off, over a period that encompasses a variety of episodes in the global economy from the 1970s to the aftermath of the Great Recession. Lucas (1972; 1973) shows that the effect of inflation surprises on the aggregate level of output weakens when inflation variability is high.

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[1] The output-inflation trade-off is one of the principal cornerstones in modern macroeconomics and a key component in models for applied monetary policy analysis (see Rudd and Whelan, 2007; King, 2008; Gordon, 2011; Meade and Thornton, 2011 for a review of the literature and a historical overview).

[2] A growing discussion in the literature of inflation dynamics and the Phillips curve examines the absence of a persistent decline in the level of inflation in the wake of the recent economic crisis (i.e. the 'missing disinflation' puzzle) as well as the absence of any signs of recovery of inflation after the crisis, where inflation rate has been persistently observed below its target (i.e. the 'missing inflation' puzzle). For an extensive examination of the puzzling behavior of inflation since the crisis, see Ball and Mazumder (2011), Matheson and Stavrev (2013), Coibion and Gorodnichenko (2015) and Bobeica and Jarociński (2017) among others.

In contrast, a new Keynesian rebuttal (Ball *et al.*, 1988; hereafter BMR) suggests that it is the higher average inflation rates, and not increased inflation variability, that diminishes the response of output to aggregate demand shocks.[3] The empirical literature has produced inconclusive evidence, typically utilizing a two-step procedure in the context of cross-country and time-series analyses. The first step of these analyses consists in characterizing the trade-off on a country-by-country basis while the second step considers the determinants of the Phillips curve slope across countries. DeFina (1991) introduces a one-step procedure, implemented in a time-series context, where the output-inflation trade-off coefficient is a function of both the level and the volatility of the inflation rate. His results show a significant effect of the average inflation rate on the output-inflation trade-off coefficient for a considerable number of countries corroborating the cross-country evidence of BMR.[4] No empirical work exists, however, to the best of our knowledge, that exploits both time-series and cross-section dimensions in a single step panel framework. This paper, thus, tries to address the shortcomings of the previous studies by using an all-encompassing panel econometric framework and providing a unified one-step approach to the understanding of the determinants of the output-inflation trade-off.

Extensive empirical evidence exists on inflation dynamics and the output inflation trade-off, following the works of Lucas and BMR. The findings, however, are mixed and inconclusive (e.g., DeFina, 1991; De Veirman, 2009; Sun, 2014). While the results of early tests support the basic implications of the neo-Classical model (Alberro, 1981; Koskela and Viren, 1980; Jung, 1985), subsequent research challenges them (Froyen and Waud, 1985; Katsimbris, 1990a; b). In general, the support for the Neo-Classical model diminishes as the analysis moves from cross-country to time series methods (Katsimbris and Miller, 1996). More recently, a renewed interest in the empirical research has emerged. Abbott and Martinez (2008) examine both the BMR and Lucas hypotheses by employing a two-step approach, producing evidence which is consistent with the Neo-Classical view, but is inconclusive concerning the impact of mean inflation. De Veirman (2009) produces evidence from Japan for the competing BMR and Lucas models. His results are consistent with the endogenous pricing models, which imply that declining trend inflation causes the Phillips curve to flatten. Fendel

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[3] Lucas (1972) suggests that a trade-off emerges between inflation and output because of the producers' inability to distinguish between changes in relative prices and changes in the aggregate price level. Thus, producers' expectations depend on the variability of relative prices to the general price level, implying a negative relationship between the Phillips curve slope and the variability in nominal aggregate demand. Lucas (1973) tests this hypothesis for a cross-section of OECD countries and finds supporting evidence. Ball *et al.*, (1988) shift the focus to average inflation as a potential determinant of the output-inflation relationship, developing a model where the real effects of nominal aggregate demand shocks depend on the frequency of price changes. A higher inflation rate increases the frequency of firm's price adjustments and decreases the real effects of a nominal shock. As a result, the output inflation trade-off depends also on the mean rate of inflation. They consider an extended sample of 43 countries to test the hypothesis of a negative relationship between the average rate of inflation and the output-inflation trade-off, while controlling for the variability of aggregate demand. Their results point to a diminishing effect of the average inflation rate on the short-run output-inflation trade-off, that is supportive to the New Keynesian hypothesis.

[4] The approach of DeFina (1991) overcomes limitations of the two-stage cross-country methodology by allowing for time-series variation of the independent variables as well as for a time-varying trade-off parameter.

and Rulke (2012) focus on inflation surprises to consider the empirical validity of the Lucas supply function and point to a negative relation between the slope parameter and inflation variability. Sun (2014) evaluates four alternative estimation approaches that have been used in the empirical literature to examine the New Keynesian hypothesis. Her results suggest that nominal rigidity is the main determinant of the trade-off and that the one-stage country-by-country time-series procedure emerges as superior. Benati (2007) explores the BMR hypothesis for 12 OECD countries and the Eurozone using complex demodulation techniques and shows that changes in trend inflation cause analogous changes in the frequency of price adjustment, and thus, affect the slope of the Phillips curve. His results confirm the documented flattening of the Phillips curve over the recent period. Ball and Mazumder (2011) examine inflation dynamics in the United States from 1960 through 2010, focusing on the Great Recession period. They find that two simple modifications of the Phillips curve fit the entire period very well. First, approximating core inflation with the weighted median of price changes, and second allowing the slope of the Phillips curve to change with the level and variance of inflation. Simon *et al.* (2013) reiterate that inflation in advanced economies has become less responsive to changes in economic slack and that the longer-term inflation expectations have become more anchored. Lopez-Villavicencio and Mignon (2015) study the time instability of the Phillips curve by focusing on price stickiness and the inflation environment, finding that the last affects the slope of the Phillips curve. In addition, they find that the Phillips curve has become flatter around a lower mean of inflation. Finally, Davig (2016) develops a model with a time-varying slope coefficient Phillips curve, providing another theoretical explanation for the observed flattening of the Phillips curve.

The related empirical literature has been slow in catching up with recent advances in panel data econometrics. In particular, the literature typically assumes that the impact of the trade-off determinants across countries is characterized by homogeneous dynamics. The evidence, however, shows that the estimated trade-off coefficient varies across countries (e.g., Akerlof *et al.*, 1988; Benati, 2007) and that significant variation exists in the estimated size of the effects of inflation and its variability on the trade-off (e.g., DeFina, 1991; Altissimo *et al.*, 2006). Likewise, Imbs *et al.* (2011) and Byrne *et al.* (2013) highlight the heterogeneity in inflation dynamics and the Phillips curve. Ignoring heterogeneity, thus, can lead to inconsistent estimates (Pesaran and Smith, 1995). Furthermore, macroeconomic variables are interconnected across countries (Bailey *et al.*, 2016), and therefore the errors of panel data regressions can be cross-sectionally correlated. Such interdependence can emerge because of a limited number of strong factors, associated with global and/or aggregate common shocks that have heterogeneous impact across countries, and an infinite number of weak factors, such as local spillover effects between countries or regions (Chudik *et al.*, 2011). Therefore, conventional panel estimators that ignore the cross-sectional dependence of errors can also lead to inconsistent estimates of the slope parameters and misleading inference.

We consider the determinants of the output-inflation trade-off for a panel of 60 countries over the period 1970-2010, along the lines of Lucas' (1973) and Ball *et al.*'s (1988) hypotheses. Unlike previous analyses on the factors that determine the output-inflation trade-off, we exploit the panel nature of the data and use estimation methodologies, which produce consistent estimates by allowing for cross-country heterogeneity and cross-sectional dependence. Our approach also allows for a time-varying impact of the trade-off determinants on the slope of the Phillips curve. We use four decades of data covering the Great Moderation, the high inflation period preceding it, and the Great Recession period that followed it.

We contribute to the literature in several ways. First, we examine the determinants of the output-inflation trade-off in a unified panel framework, which incorporates a one-step approach, extending the procedure of DeFina (1991) to a panel data context. Second, we introduce an all-encompassing panel specification, which allows to account for heterogeneity in the slope coefficients of the trade-off determinants and for cross-sectional dependence in the panel to provide more robust estimates. Third, we consider a wide sample, which covers the samples of Lucas (1973) and of Ball *et al.* (1988) as special cases, but extends to recent years to consider the Great Moderation and the Great Recession periods. Finally, we control for the nonstationarity of real and nominal output (Abbott and Martinez, 2008) and we model the time-varying variance of inflation using a GARCH approach (Asai, 1999). To the best of our knowledge, this is the first attempt in the literature that produces evidence utilizing the single-step estimation approach into a heterogeneous dynamic panel framework.

Our main findings can be summarized as follows. First, there is strong support for the New Keynesian view of a negative association between the rate of inflation and the output-inflation trade-off. Moreover, this finding is consistent with the recently observed flattening of the Phillips curve under low inflation conditions. Second, there is fairly weak evidence of a significant effect of inflation variability on the trade-off for the OECD countries over the full sample, consistent with the Neo-Classical hypothesis. Third, the non-stationarity of real/nominal GDP does not affect the significance of the estimates. Fourth, the findings remain robust to alternative samples, when we employ the variability of nominal GDP instead of, or in addition to, the variability of inflation and when we consider alternative specification forms, including those of Ball *et al.* (1988), Akerlof *et al.* (1988), and Khan (2004). Finally, our analysis highlights the importance of considering typical features of macroeconomic panels, such as parameter heterogeneity and cross-sectional dependence, when exploring the output-inflation trade-off in a panel setting.

The remainder of the paper is organized as follows. Section 2 discusses the econometric model and outlines the estimation methodology. Section 3 describes the data and their properties. Section 4 provides the estimation results and discusses the empirical findings. Section 5 presents several robustness checks. Finally, Section 6 concludes.

2 Econometric Model

The existing literature typically relies on a two-step empirical framework introduced by Lucas (1973) and extended by Ball *et al.* (1988) to explore the determinants of the output-inflation trade off. The first stage of this approach estimates a conventional time-series aggregate supply curve by regressing real output (y_t) onto its own lagged value, the rate of change in nominal output (Δx_t), and a time trend for each country i separately. The second stage employs a cross-section regression of the estimated output-inflation trade-off coefficients across countries onto the rate of inflation (π_i) and the variability of the inflation rate and/or the variability of aggregate demand (σ_i).

Our estimation framework starts with a well-established reduced-form specification used in the empirical literature (Lucas, 1973; Ball *et al.*, 1988; DeFina, 1991; Abbott and Martinez, 2008; Sun, 2014; among others). In order to explore both the time-series and cross-section dimension of the data, we write the first stage equation as a dynamic heterogeneous panel data model:

$$y_{i,t} = \alpha_i + \eta_i y_{i,t-1} + \tau_i \Delta x_{i,t} + \gamma_i t + u_{i,t}, \tag{1}$$

where $y_{i,t}$ is the logarithm of real GDP and $x_{i,t}$ is the logarithm of nominal GDP for country i at time t, α_i is a set of country-specific fixed effects that captures the influence of unobserved country-specific heterogeneity and $u_{i,t}$ is the error term.

Following DeFina (1991) and assuming a time-varying trade-off (τ) that depends on the inflation rate ($\pi_{i,t}$) and its variability ($\sigma_{i,t}$) in each time period, we can write the output-inflation trade-off coefficient as:

$$\tau_{i,t} = f(\pi_{i,t}, \sigma_{i,t}). \tag{2}$$

Assuming, also, a linear relationship between the trade-off parameter and the inflation rate and/or its variability we combine the two steps into a unified single-stage panel regression as follows:[5]

$$y_{i,t} = \alpha_i + \eta_i y_{i,t-1} + \beta_{0,i}\Delta x_{i,t} + \beta_{1,i}\pi_{i,t} \times \Delta x_{i,t} + \beta_{2,i}\sigma_{i,t} \times \Delta x_{i,t} + \gamma_i t + u_{i,t}. \tag{3}$$

Finally, in order to take into account the possibility of real and nominal output ($y_{i,t}$ and $x_{i,t}$ respectively) non-stationarity, we adopt a stationary specification form (see Abbott and Martinez, 2008):

$$\Delta y_{i,t} = c_i + \rho_i \Delta y_{i,t-1} + \beta'_{0,i}\Delta x_{i,t} + \beta'_{1,i}\pi_{i,t} \times \Delta x_{i,t} + \beta'_{2,i}\sigma_{i,t} \times \Delta x_{i,t} + u_{i,t}. \tag{4}$$

As Abbott and Martinez (2008) show, the interpretation of the trade-off coefficients ($\beta' s$) in the stationary specification (Equation 4) is consistent with the interpretation of the non-stationary specification (Equation 3).[6] Thus, the stationary specification (Equation 4) allows to explore the determinants of real output's responsiveness to nominal changes.

To allow for cross-sectional dependence in the disturbances, $u_{i,t}$, we assume that they follow a multi-factor error structure:

$$u_{i,t} = \gamma'_{\mathbf{i}}\mathbf{f_t} + \varepsilon_{i,t}, \tag{5}$$

where $\mathbf{f_t}$ is a $m \times 1$ vector of unobserved common factors that capture cross-sectional dependencies across countries, and $\gamma'_{\mathbf{i}}$ are the country specific associated factor loadings. The idiosyncratic errors, $\varepsilon_{i,t}$, are assumed to be independently distributed across i and t with zero mean and constant variance.

The interaction terms measure the degree to which the coefficients of nominal aggregate demand vary as in response to changes in inflation and inflation variability. Both, the New-Keynesian and the Neo-Classical, models suggest that an increase in inflation rate variability causes a decline in the output-inflation trade-off. That is, both models suggest that $\beta'_{2,i} < 0$. According to the New-

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[5] To our knowledge, only Yates and Chapple (1996) hint to a panel extension of a one-step procedure to explore the output-inflation trade-off, but no explicit evidence or discussion is provided due to methodological constraints of the time. Loungani *et al.* (2001) also consider a panel data extension of their open economy analysis on how capital controls may impact the trade-off.

[6] For example, in the model $y_{i,t} = \alpha_i + \eta_i y_{i,t-1} + \beta_i z_{i,t} + v_{i,t}$, the trade-off coefficient is $\frac{\partial y_{i,t}}{\partial z_{i,t}} = \beta_i$, and in the model $\Delta y_{i,t} = c_i + \rho_i \Delta y_{i,t-1} + \beta_i{}' \Delta z_{i,t} + v_{i,t} \Rightarrow y_{i,t} = c_i + (1+\rho_i) y_{i,t-1} - \rho_i y_{i,t-2} + \beta_i{}' z_{i,t} - \beta_i{}' z_{i,t-1} + v_{i,t}$, the trade-off coefficient is $\frac{\partial y_{i,t}}{\partial z_{i,t}} = \beta_i{}'$.

Keynesian model, however, the coefficient of the interaction term between nominal aggregate demand and the inflation rate is negative, that is $\beta'_{1,i} < 0$. In other words, the sign and significance of the estimated coefficient of this interaction term indicates whether or not the coefficient varies systematically as the New-Keynesian model suggests.

Finally, to obtain a time-varying measure of inflation variability, $\sigma_{i,t}$, we estimate a generalized autoregressive conditional heteroscedasticity GARCH (1,1) model for $\pi_{i,t}$. The conditional variance of $\pi_{i,t}$ is a parametric proxy for inflation variability.[7],[8]

2.1 Estimation Methodology

A first concern that emerges in estimating macroeconomic panel data models, such as the specification of Equation 4, is the heterogeneity of the estimated parameters. The fixed effects (FE) estimator allows for individual heterogeneity through different intercepts across countries and can be estimated using ordinary least squares, but restricts slopes to be homogeneous. In addition, the slope homogeneity assumption of the traditional FE estimator is quite restrictive, since the application of pooled estimation methods can lead to substantial 'heterogeneity' bias in the estimated parameters (Pesaran and Smith, 1995). Pesaran and Smith (1995) propose the Mean Group (MG) estimator that allows for slope heterogeneity in the panel. The MG estimator consists in estimating separate OLS regressions for each country and then calculating averages of the specific coefficients over groups.[9]

A second concern, typical in the analysis of macroeconomic panel data models, is the cross-sectional correlation among countries. Estimators that fail to account for cross-sectional dependence turn out to be inefficient and inconsistent (Sarafidis and Wansbeek, 2012). To remedy for the presence of dependence across countries, we use the Common Correlated Effects (CCE) procedure (Pesaran, 2006) that accounts for unobserved common factors and produces consistent estimates. The CCE estimator uses cross-section averages of the dependent and independent variables as proxies for the unobserved factors and includes them as additional regressors. Specifically, we use the Mean Group version of the CCE estimator (CCEMG) which is found to perform better in small samples (Coakley *et al.*, 2006).

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[7] We also employ a five-year rolling window standard deviation of $\pi_{i,t}$ as a benchmark case.

[8] Asai (1999) similarly, follows a GARCH approach to measure inflation variability on a time-series context.

[9] In the main analysis of the paper, we present the results based on the benchmark FE estimator while in the robustness section, we present additional evidence using the system Generalized Method of Moments (GMM) estimator of Blundell and Bond (1998) that accounts for dynamics and controls for endogeneity, and the MG estimator of Pesaran and Smith (1995) that accounts for slope heterogeneity.

Finally, to control for the presence of the lagged dependent variable in the specification of Equation 4, and taking into account all preceding issues, we use the recent extension of the CCEMG estimator, proposed by Chudik and Pesaran (2015), which allows to include lagged values of the dependent variable and/or weakly exogenous regressors in heterogeneous panel data models with cross-sectional dependence. In particular, Chudik and Pesaran (2015) show that the dynamic CCEMG (dynCCEMG) extension of the CCE Mean Group estimator performs well and remains valid asymptotically when (i) the individual equations of the panel in addition to the cross-section averages include a sufficient number of lags of cross-section averages, (ii) the number of cross-section averages is at least as large as the number of unobserved common factors, and (iii) the time dimension is large enough to allow the regression to be estimated for each cross-section. In order to apply the dynCCEMG estimator we augment Equation 4 with the cross-section averages of the dependent and independent variables as well as their lags as additional regressors:

$$\Delta y_{i,t} = c_i + \rho_i \Delta y_{i,t-1} + \beta_{\mathbf{i}}' \mathbf{z}_{\mathbf{i},\mathbf{t}} + \sum_{l=0}^{p_T} \delta_{il}' \mathrm{v}_{\mathbf{t}-\mathbf{l}} + u_{i,t}, \tag{6}$$

where $\mathbf{z}_{\mathbf{i},\mathbf{t}}$ is a vector that contains the set of regressors, $\mathbf{z}_{\mathbf{i},\mathbf{t}} = (\Delta x_{i,t}, \pi_{i,t} \times \Delta x_{i,t}, \sigma_{i,t} \times \Delta x_{i,t})'$, and $\mathrm{v}_{\mathbf{t}} = N^{-1} \sum_{i=1}^{N} \mathrm{v}_{\mathbf{i},\mathbf{t}} = (\Delta y_t, \mathbf{z}_{\mathbf{t}})'$. We estimate the model using different numbers of lags of cross-section averages up to the p_T, where p_T is equal to the integer part of $T^{1/3}$, as suggested by Chudik and Pesaran (2015).

The impact of inflation and its variability on the trade-off may vary across countries, depending on country specific structural factors as well as on spillover effects across them.[10] Therefore, dynamic panel estimations examining the determinants of output inflation trade-off should control for cross-country heterogeneity and cross-sectional dependence.

3 Data

We use annual data over the period 1970-2010 for a sample of 60 countries. The variables we consider are the logarithm of the real GDP ($y_{i,t}$) and the logarithm of the nominal GDP ($x_{i,t}$), both measured in U.S. dollars, and the GDP price deflator ($p_{i,t}$). Table 1 presents the countries considered in our analysis and their abbreviations. In order to ensure a comparable and balanced data set we obtain the data from the United Nations *National Accounts Main Aggregates Database*. We calculate

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[10] For example, Daly and Hobijn (2014) consider the role of downward nominal rigidities.

the inflation rate using the GDP price deflator ($\pi_{i,t} = \Delta log(p_{i,t})$). We construct two time-varying measures for the volatility of inflation ($\sigma_{i,t}$) using, first, a five-year rolling window standard deviation for inflation and, second, a generalized autoregressive conditional heteroscedasticity (GARCH (1,1)) model, where the estimated 'conditional variance' is a proxy of inflation variability.

We also consider a sub-sample of the 42 countries originally analyzed in Ball *et al.* (1988) and Akerlof *et al.* (1988) and a sub-sample of OECD countries (as an approximation to the Lucas (1973) sample) as benchmarks to which we compare our full sample empirical results.[11]

3.1 Data Properties

Prior to estimating Equation 4 we conduct various tests to explore the properties of our data.[12] We consider the assumption of cross-sectional independence among countries (cross-sectional dependence tests) in Table 2, the assumption of slope homogeneity across countries (poolability tests) in Table 3, and the order of integration of the series (panel unit root tests) in Table 4. The CD_p test (Pesaran, 2004) for cross-sectional dependence (Table 2) shows that the null hypothesis of cross-sectional independence among the countries in our panel is strongly rejected at the 5% level of significance, indicating that our data are subject to considerable cross-section dependence. The Delta test of Pesaran and Yamagata (2008) is used to test the assumption of homogeneous slope parameters in the panel. As the results in Table 3 show, the hypothesis of common slopes is strongly rejected, suggesting that notable heterogeneity among countries exists. Finally, the results from the Pesaran's (2007) CIPS panel unit root test, reported in Table 4, indicate that real and nominal GDP series are integrated of order one, $I(1)$, while the inflation and inflation variability series are stationary, $I(0)$.

4 Empirical Results

4.1 Lucas' Variance Hypothesis

We use the Lucas' (1973) 'variance hypothesis' as the initial benchmark specification for our analysis, and consider whether a significant negative relationship exists between the variance of nominal shocks and the output-inflation trade-off. Table 5 presents the results for the Lucas hypothesis based

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[11] The only exception is Zaire, which we exclude from the original BMR dataset due to data unavailability.

[12] A detailed description of these tests is provided in the Appendix.

on alternative estimation approaches. To capture nominal shocks' variability we consider both nominal demand and the inflation rate. We obtain two volatility measures for each, namely a five year rolling standard deviation (Panel A) and a GARCH time varying measure (Panel B). Columns 1 and 4 report the estimates from the (homogeneous) Fixed Effects estimator, while Columns 2-3 and 5-6 report the results from the Mean Group CCE estimators that account for heterogeneity and cross-sectional dependence.[13] In particular, Columns 2 and 5 present the standard (CCEMG) version and Columns 3 and 6 present the dynCCEMG extension as suggested by Chudik and Pesaran (2015).

The estimates of the impact of nominal shocks' variability (inflation or nominal demand) on the output-inflation trade-off, are not statistically significant, with the only exception the case of Column 1 in Panel A, where the variability of nominal shocks emerges as statistically significant.[14] Therefore the evidence does not provide any support for the Lucas' (1973) 'variance hypothesis' and cannot confirm the previous findings of Apergis and Miller (2004). This motivates the broadening of our focus to cover not only the variability but also the average rate of inflation as a potential determinant of the output-inflation trade-off (Ball *et al.*, 1988).

4.2 Determinants of the Output-Inflation Trade-off

The heart of our analysis considers the role of inflation dynamics (both the average rate and its variability) in determining the output-inflation trade-off, by estimating the one step dynamic regression of Equation 4 in a panel context. Table 6 presents our main results based on alternative estimation approaches and using the two measures for the variability of inflation, namely, the 5-year rolling standard deviation, reported in Panel A and the GARCH time varying measure, reported in Panel B. Column 1 shows the estimates from the (homogeneous) Fixed Effects estimator, which consists our benchmark case for the analysis. Columns 2-3 report the results from the heterogeneous estimators that account for cross-sectional dependence, using both the standard (CCEMG) and the dynamic extension (dynCCEMG) as suggested by Chudik and Pesaran (2015).[15] All estimates indicate that the coefficient of nominal demand is statistically significant and positive, confirming the conventional understanding of the short-run Phillips curve. A diagnostic check of the alternative

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[13] Following the tests presented in Tables 2 and 3, the preferred estimation technique is the usage of Mean Group CCE estimators, which account for the cross-country heterogeneity and cross-sectional dependence that are present in our data. We complement the analysis with the benchmark FE estimator for comparison purposes, and in the next section we also employ additional estimation techniques for robustness purposes. We would like to thank an anonymous referee for drawing our attention to this point.

[14] The estimates, however, in this case may be subject to bias because the estimation methods used (FE) does not take into account cross-sectional dependence and heterogeneity.

[15] The dynamic CCEMG estimator of Chudik and Pesaran (2015), performed here, is augmented with one lag of the cross-section averages. In the robustness section we employ, additionally, the version of the estimator that is augmented with three lags of the cross-section averages.

estimators (FE, CCEMG and dynCCEMG) based on the root mean squared error (RMSE) statistic shows that the dynamic CCEMG models provide the lower values for the measure of goodness of fit. This indicates that estimators accounting for dynamics, cross-sectional dependence and heterogeneity in the specification form, e.g., the dynCCEMG model, fits better the data.

When we move to the heterogeneous estimation approaches, however, the value of the trade-off coefficient, is considerably larger as compared to that of the (pooled) FE estimates. Turning our attention to the interaction terms of the trade-off determinants, all estimates point to the inflation rate as the main source of nominal aggregate demand changes. The interaction term coefficients between the rate of inflation and nominal aggregate demand are negative and statistically significant. While the coefficient of the interaction term between inflation variability and nominal demand is also negative, as the Lucas (1973) hypothesis suggests, it is not statistically significant.[16,17] The results presented in Panel B of Table 6 show that the value of the interaction term coefficient for inflation is substantially lower under the homogeneous approach as compared to those from heterogeneous approaches. Specifically, the estimate from the Fixed Effects method is -0.078 and becomes -0.623 when allowing for slope heterogeneity and controlling for cross-sectional correlation across countries using the CCE Mean Group estimator. The preferred estimation technique is the Chudik and Pesaran's dynCCEMG estimator (Column 3), which accounts for dynamics, cross-sectional dependence and heterogeneity. This method reports a negative and significant coefficient for the interaction term of inflation with an estimated value of -0.559, suggesting that an increase in the rate of inflation will cause a significant decline in the output-inflation trade-off. Therefore, our findings show that the average level of inflation affects negatively the slope of the Phillips curve, as measured by the trade-off between output and inflation. These results render the inflation rate the main determinant of real output's responsiveness to nominal demand changes, as the New Keynesian hypothesis suggests.[18,19] This evidence corroborates previous studies such as Ball *et al.* (1988), DeFina (1991) and Khan (2004), while calls into question some recent evidence (Abbott and Martinez, 2008; Fendel and Rulke, 2012). In addition, our results reinforce the recent findings of Ball and Mazumder (2011), who show that a backward-looking Phillips Curve with a time-varying slope (where the slope is affected by the level and the variability of inflation) is a better approximation of

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[16] The only exception where inflation variability is significant refers to Column 1 in Panel A that uses the 5-years rolling measure of inflation variability.

[17] The joint significance of the two interaction terms is high ($p < 0.05$), especially on the heterogeneous panel data estimations, indicating that the Phillips curve slope is determined by inflation dynamics ($\pi_{i,t}$ and $\sigma_{i,t}$). This is in line with the findings of Ball and Mazumder (2011) and Murphy (2014).

[18] Akerlof *et al.* (1988) argue that one cannot discriminate the effect on the steepness of the Phillips curve that is due to the volatility from the effect that results from the level of inflation, due to the high correlation among the two variables. The time varying measure for the variability of inflation we use shows that this is not true for our data, since the correlation among $\pi_{i,t}$ and $\sigma_{i,t}$ is very low and equals -0.0058.

[19] The performance of the system GMM estimation in the robustness section, additionally, indicate that our results are not subject to endogeneity between the explanatory variables ($\pi_{i,t}$ and $\sigma_{i,t}$), and thus, confirm that the main source of the effect results from the level of inflation.

inflation dynamics during the recent period. Our findings, however, overcome the limitations faced by previous time-series and 'pooled' cross-sectional analyses and highlight the role of heterogeneity and cross-sectional dependence in the estimation of alternative specifications of the Phillips curves. Similar concerns regarding heterogeneity in estimating Phillips curves have beed also raised by Imbs *et al.* (2011) and Byrne *et al.* (2013).

To ensure that the stationary specification of Equation 4 does not affect our results we perform the panel analysis using $y_{i,t}$ instead of $\Delta y_{i,t}$ (Equation 3). Columns 4-6 in Table 6 report results based on the three estimation approaches (FE and CCEMG in static and dynamic form). As the estimates reveal, our results are not affected by the non-stationary treatment of the real GDP. This is consistent with the findings of Abbott and Martinez (2008) regarding non-stationarity.[20] The impact of nominal growth on real GDP is weaker but displays the correct sign. The sign of the interaction terms coefficients is consistent with the stationary specification. The estimates based on the preferred dynCCEMG method (Column 6), yield again a negative and significant coefficient for the interaction term of inflation ranging from -0.58 to -0.63, that is very close to the estimates from the stationary specification (Column 3). Hence, our main conclusions of a significant negative relation between the rate of inflation and the output-inflation trade-off is upheld.

5 Robustness Checks

We consider a number of robustness checks regarding the estimation techniques, the period considered, the grouping of the cross-sectional sample, and the functional form of the specification. Table 7 extends Table 6, by reporting the results on the determinants of the output-inflation trade-off based on three additional estimation methods (GMM, MG and an alternative version of the dynCCEMG estimator). The results from Table 7 strongly reaffirm our main findings (Table 6), which point to the inflation rate as the main determinant of the output-inflation trade-off. Table 8 documents the robustness of our previous findings, by reporting the dynamic CCEMG estimates under alternative time periods and country samples. We examine the robustness of the results over the entire sample period (Columns '70 - '10), the period from the mid-1980s until the recent crisis of 2007 (Columns '82 - '07), and finally, the full sample excluding the recent crisis and the Great Recession (Columns '70 - '07).[21],[22] The second sample period practically corresponds to the Great Moderation. We consider

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[20] Abbott and Martinez (2008) present evidence that the use of non-stationary specifications in earlier studies does not invalidate the results.

[21] A substantial reduction in both the level and the volatility of inflation for most of the OECD countries occurs during the Great Moderation (see among others, Summers, 2005; Davis and Kahn, 2008).

[22] As Gilchrist *et al.* (2017) observe the recent economic crisis have caused changes in inflation dynamics.

three different cross-sectional groups. The full sample of 60 countries (Columns FULL), the sub-sample of the 42 countries, corresponding to the sample of Ball *et al.* (1988) (Columns BMR), and the sub-sample of 24 OECD countries (Columns OECD), which approximates the Lucas (1973) sample. Table 8 builds on Table 6 and consists of nine sub-samples, combining the partitions of the full sample over alternative time periods and country groupings. The estimated parameters of primary interest, i.e., the coefficients of the interaction terms between nominal aggregate demand and the level of inflation and its volatility, display the correct sign and are consistent with those reported in Table 6. There is substantial variation, however, in terms of significance, depending on the sub-sample we consider. The results from the entire sample period (Columns 1, 4 and 7) reveal a significant negative impact of inflation on the trade-off, which endures when we exclude the recent years of the financial crisis (Columns 3, 6 and 9). These results are not affected by consideration of alternative country groupings. The impact of inflation on the slope of the Phillips curve, however, vanishes when we focus on the Great Moderation period for the OECD countries (Column 8). This result reaffirms previous evidence (Summers, 2005; Davis and Kahn, 2008), showing that inflation has been less sensitive to demand shocks during the Great Moderation. This reduced sensitivity leads to the non-significance of the interaction term (Column 8).

When we consider the implications of alternative country groupings in our sample, we find that the FULL sample, (Columns 1, 2 and 3), as well as the BMR sample (Columns 4, 5 and 6) produce similar results, with those reported in Table 6. In contrast, when we examine the OECD sub-sample over the extended periods of data (Columns 7 and 9), quite different results emerge. While inflation has a consistently negative impact on the trade-off, a significant and stronger negative impact of inflation volatility on the trade-off now emerges. The inclusion of the 1970s period, when the OECD countries experienced high and volatile inflation rates (e.g., Summers, 2005), drives this result. A significant effect of inflation dynamics, both in level and volatility, on the output-inflation trade-off, becomes evident for the OECD economies when considering the full sample period, while when we concentrate on the Great Moderation, the results reflect the reduced sensitivity of the trade-off to inflation. Thus, while our results provide strong support for the New-Keynesian hypothesis as laid out by Ball *et al.* (1988), are also reconcilable with the Lucas' (1973) when focusing on the OECD countries and relatively high inflation periods.

As a further robustness check of the results, we consider alternative specifications of our baseline regression. Table 9 presents the results of the dynamic CCEMG estimator on various extensions and alternatives of a generalized form of Equation 4 as follows:

$$\Delta y_{i,t} = c_i + \rho_i \Delta y_{i,t-1} + \beta'_{0,i} \Delta x_{i,t} + f(\pi_{i,t}, \sigma_{i,t}) \Delta x_{i,t} + \beta'_{\mathbf{i}} \mathbf{X_{i,t}} + u_{i,t}. \qquad (7)$$

This functional form allows inflation and its variability to affect the trade-off through a non-linear specification, $\tau = f(\pi_{i,t}, \sigma_{i,t})$.[23] In particular, we consider both a quadratic representation and the inverse of the trade-off parameter, following Ball *et al.* (1988) and Akerlof *et al.* (1988), respectively. Furthermore, we augment the specification with a set of additional explanatory variables (vector $\mathbf{X_{i,t}}$) including supply shocks and inflation persistence.

Estimating the general form of Equation 7 through the dynamic CCEMG estimation method we can explore a number of research questions raised in the literature. Column 1 in Table 9 shows the results of the specification including only the rate of inflation as a potential trade-off determinant, while the model in Column 2 contains only the variability of inflation as the trade-off determinant. Column 3 repeats the specification with both determinants, $\pi_{i,t}$ and $\sigma_{i,t}$, as provided in Column 3 of Panel B at Table 6. It emerges that the main determinant of the output-inflation trade-off is the rate of inflation. Moreover, our results are not affected by the multicollinearity problem among the two factors, an issue raised by Ball and Mazumder (2011). Column 4 present the results of the quadratic specification of BMR, while Columns 5 and 6 reports results based on the ARY specification in linear and quadratic form, respectively. These estimations reiterate the significance of inflation rate's coefficient versus the inflation variability coefficient. To explore further alternatives to the specification form we consider as a possible determinant the variability of nominal demand, $\sigma_{i,t}^{x}$. Columns 7 to 10 present the results from this extension. It emerges that replacing the variability of inflation, $\sigma_{i,t}$, with that of nominal demand, $\sigma_{i,t}^{x}$, does not affect the validity of our main findings. The only exception is in Column 9 where we include both inflation and demand variability and we also find a significant effect of inflation's volatility on the trade-off. Column 11 shows the results of an extended specification that includes a supply-side effect. We use the growth rate of oil price as a proxy for the supply side effects, as suggested by DeFina (1991). While the estimation results show that the supply side effect is statistically significant, the inflation rate still emerges as the key determinant of the trade-off, a finding that contradicts the evidence of Apergis and Miller (2004). Finally, the last two Columns, 12 and 13, examine the impact of inflation on output persistence, as suggested by Khan (2004). Indeed, using Khan's (2004) specification, the results show a significant negative effect of inflation on the persistence of output, while the role of inflation as the main determinant of the output-inflation trade-off remains robust.

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[23] Several papers document the non-linearity of the Phillips curve using a wide range of non-linear specification forms, e.g., Dolado *et al.* (2005); Lopez-Villavicencio and Mignon (2015).

Some key findings emerge from our analysis. First and foremost, there is evidence of a significant negative impact of the rate of inflation on the output-inflation trade-off. To our knowledge, this is the first set of evidence from a unified one-step panel estimation procedure. This result is consistent with the New Keynesian hypothesis (Ball *et al.*, 1988; DeFina, 1991; Khan, 2004; Lopez-Villavicencio and Mignon, 2015), as well as with the flattening of the Phillips curve since the Great Moderation (Mishkin, 2007; De Veirman, 2009; Simon *et al.*, 2013). Moreover, our findings remain robust to alternative time series and cross-sectional samples and specifications. Second, there is evidence of a significant impact of inflation volatility on the trade-off, shyly resuscitating the New Classical hypothesis, but this is confined to the high inflation experience of the OECD countries. A major implication is that inflation dynamics (both the inflation rate and its variability) are key determinant factors of the Phillips curve slope in the OECD countries (Kuttner and Robinson, 2010; Ball and Mazumder, 2011; Lopez-Villavicencio and Mignon, 2015). Our analysis reinforces the evidence of time variation in the slope of the Phillips Curve (Ball and Mazumder, 2011; Murphy, 2014; Coibion and Gorodnichenko, 2015) and validates the use of a one-step approach that follows the procedure of DeFina (1991) in a panel data context. Finally, the convincing empirical evidence of considerable heterogeneity in the dynamics of inflation (e.g., Imbs *et al.*, 2011; Byrne *et al.*, 2013), the extended cross-country variation of the trade-off parameter (e.g., Akerlof *et al.*, 1988; DeFina, 1991; Benati, 2007), and the documented cross-sectional correlation across countries (e.g., Byrne *et al.*, 2013; Bailey *et al.*, 2016), motivate the use of an all-encompassing panel data approach that controls for such typical features of macroeconomic panel datasets when exploring the trade-off determinants.

6 Conclusions

This paper considers the effects of inflation and inflation variability on the output-inflation trade-off. Two contrasting hypotheses have been developed on the relationship between inflation dynamics and the slope of the Phillips curve. One, emanating from Lucas' (1972; 1973) research program, identifies inflation variability as the key determinant of the trade-off. A New Keynesian rebuttal, advanced by Ball *et al.* (1988), however, shifts focus to the rate of inflation. In revisiting empirically the Phillips curve determinants, we develop a unified single step approach in a panel data context. In addition, we use recently developed tools to explore a number of econometric issues that have not been addresed by the earlier literature. In particular, we employ an all-encompassing dynamic heterogeneous panel data specification and consider estimation techniques that account for parameter heterogeneity, cross-sectional dependence, dynamics, and non-stationarity. We use a sample of 60 countries from 1970 to 2010, which covers the sub-samples of earlier influential studies. Although we detect evidence of a negative association between inflation volatility and the trade-off, these findings are sensitive to the period and the set of countries considered. Our results, however, suggest an unambiguous and more pronounced role for inflation in determining the trade-off. Specifically, strong evidence of a negative association between the rate of inflation and the slope of the Phillips curve emerges. The results are robust to alternative periods, country groups, and model specifications. These findings are in line with the New Keynesian view that high inflation rates result in a declining responsiveness of real output to nominal shocks. Moreover, the evidence regarding the effects of inflation dynamics on the trade-off is consistent with the observed flattening of the Phillip curve over the past decades.

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Appendix

A1 Cross-Sectional Dependence Tests

We apply the Cross-Sectional Dependence (CD) test, suggested by Pesaran (2004), which is based on the average of pair-wise correlation coefficients ($\hat{\rho}_{ij}$) of the OLS residuals, obtained from ADF regressions for each country. The CD_P test is given by:

$$CD_P = \sqrt{\frac{2T}{N(N-1)}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right)}. \tag{1A}$$

The CD_P statistic, under the null hypothesis of independence across errors, follows a two-tailed standard normal distribution, i.e. $CD_P \sim N(0,1)$ for $T_{ij} > 3$ and sufficient large N.

A2 Poolability Tests

The assumption of homogeneity of the slope coefficients across countries is crucial for the consistency of the estimates in panel data models (Pesaran and Smith, 1995). Pesaran and Yamagata (2008) propose a standardized version of Swamy's (1970) statistic to test for slope homogeneity in large panels. The standardized Delta test statistic ($\widetilde{\Delta}$) can be defined as

$$\widetilde{\Delta} = \sqrt{N}\left(\frac{N^{-1}\tilde{S}-k}{\sqrt{2k}}\right), \tag{2A}$$

and the bias adjusted version of the Delta test statistic is

$$\widetilde{\Delta}_{adj} = \sqrt{N}\left(\frac{N^{-1}\tilde{S}-E(\tilde{z}_{iT})}{\sqrt{Var(\tilde{z}_{iT})}}\right), \tag{3A}$$

where $E(\tilde{z}_{iT}) = k$, $Var(\tilde{z}_{iT}) = \frac{2k(T-k-1)}{T+1}$, and $\tilde{S}$ is the modified version of Swamy's (1970) statistic that is based on the dispersion of individual slope estimates from a weighted Fixed Effects pooled estimator. The Delta test and its bias adjusted version have an asymptotic standard normal distribution under homogeneity null and as $(N,T) \to \infty$ with $\sqrt{N}/T^2 \to 0$.

A3 Panel Unit Root Tests

To check for the order of integration, we use the CIPS panel unit root test of Pesaran (2007), which accounts for cross-sectional dependence across countries.

Pesaran (2007) uses a single-factor model with heterogeneous factor loadings for residuals and proposes the augmentation of the standard ADF regression with cross-section averages of lagged levels and first-differences of the individual series. The regression used for the i^{th} cross-section unit is defined as:

$$\Delta y_{i,t} = \alpha_i + \phi_i y_{i,t-1} + c_i y_{t-1} + \sum_{j=0}^{p_i} \theta_{i,j} \Delta y_{t-j} + \sum_{j=1}^{p_i} \theta_{i,j} \Delta y_{i,t-j} + e_{i,t}, \tag{4A}$$

where $y_{t-1} = N^{-1} \sum_{i=1}^{N} y_{i,t-1}$ and $\Delta y_t = N^{-1} \sum_{i=1}^{N} y_{i,t} = y_t - y_{t-1}$. The CIPS test statistic is based on the average of the individual cross-sectionally augmented ADF statistic, given by the t-ratio of the ϕ_i coefficient, as follows:

$$CIPS = \frac{1}{N} \sum_{i=1}^{N} t_i(N,T). \tag{5A}$$

The inverse normal version of the CIPS test statistic (CIPS$_Z$) can be defined as:

$$CIPS_Z = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1} p_{iT}. \tag{6A}$$

where p_{iT} is the p-value of the individual cross-sectionally augmented ADF statistic. The CIPS$_Z$ statistic follows a standard normal distribution under the null hypothesis of nonstationarity.

Tables

Table 1: Country Details and Abbreviations

Country	Abbrev.	BMR	OECD	Country	Abbrev.	BMR	OECD
Argentina	ARG	*		Japan	JPN	*	*
Australia	AUS	*	*	Korea	KOR		
Austria	AUT	*	*	Kuwait	KWT		
Belgium	BEL	*	*	Luxembourg	LUX		*
Bolivia	BOL	*		Malaysia	MYS		
Brazil	BRA	*		Mexico	MEX	*	
Canada	CAN	*	*	Morocco	MAR		
Chile	CHL			Netherlands	NLD	*	*
China	CHN			New Zealand	NZL		*
Colombia	COL	*		Nicaragua	NIC	*	
Costa Rica	CRI	*		Nigeria	NGA		
Denmark	DNK	*	*	Norway	NOR	*	*
Dominican Rep	DOM	*		Pakistan	PAK		
Ecuador	ECU	*		Panama	PAN	*	
Egypt	EGY			Peru	PER	*	
El Salvador	SLV	*		Philippines	PHL	*	
Finland	FIN	*	*	Portugal	PRT	*	*
France	FRA	*	*	Puerto Rico	PRI		
Germany	DEU	*	*	Singapore	SIN	*	
Greece	GRC	*	*	South Africa	ZAF	*	
Guatemala	GTM	*		Spain	ESP	*	*
Hong Kong	HKG			Sweden	SWE	*	*
Iceland	ISL	*	*	Switzerland	CHE	*	*
India	IND			Syria	SYR		
Indonesia	IDN			Thailand	THA		
Iran	IRN	*		Tunisia	TUN	*	
Ireland	IRL	*	*	Turkey	TUR		*
Israel	ISR	*		United Kingdom	GBR	*	*
Italy	ITA	*	*	United States	USA	*	*
Jamaica	JAM	*		Venezuela	VEN	*	

N = 60
T = 41 (1970-2010)
Obs = 2460

Notes: BMR refers to the sample of countries used in the original Ball, Mankiw, and Romer (1988) paper (excluding Zaire due to data unavailability). OECD refers to the sample of OECD countries used as an approximation for Lucas (1973) sample.

Table 2: Cross-Sectional Dependence Tests

	$y_{i,t}$	$\Delta y_{i,t}$	$x_{i,t}$	$\Delta x_{i,t}$	$\pi_{i,t}$	$\sigma_{i,t}$
CD_P test	244.81*	45.24*	249.82*	81.28*	82.44*	17.47*
p-value	0.000	0.000	0.000	0.000	0.000	0.000
Abs$(Corr)$	0.920	0.229	0.939	0.320	0.329	0.204

Notes: The CD_P test refers to the cross-sectional dependence test of Pesaran (2004). Abs$(Corr)$ refers to the average absolute pair-wise correlation coefficients. * indicates rejection of the null hypothesis at *5%* significance level.

Table 3: Poolability Tests

	$y_{i,t}$		$\Delta y_{i,t}$	
	BMR	ARY	BMR	ARY
$Delta$ test	19.068*	15.038*	18.064*	14.335*
p-value	0.000	0.000	0.000	0.000
$Delta_{adj}$ test	20.638*	16.276*	19.593*	15.549*
p-value	0.000	0.000	0.000	0.000

Notes: $Delta$ and $Delta_{adj}$ refer to the standardized Delta test and the bias adjusted version of the Delta test for slope homogeneity in large panels, proposed by Pesaran and Yamagata (2008). BMR refers to the Ball, Mankiw, and Romer (1988) specification, while ARY refers to the Akerlof, Rose, and Yellen (1988) specification. * indicates rejection of the null hypothesis at *5%* significance level.

Table 4: Panel Unit Root Tests

	$y_{i,t}$		$\Delta y_{i,t}$		$x_{i,t}$		$\Delta x_{i,t}$		$\pi_{i,t}$		$\sigma_{i,t}$	
	Statistic	p-value	Statistic	p-value	Statistic	p-value	Statistic	p-value	Statistic	p-value	Statistic	p-value
$CIPS$	3.062	0.999	−20.260*	0.000	1.516	0.935	−25.620*	0.000	−26.736*	0.000	−18.293*	0.000

Notes: CIPS refers to the $CIPS_Z$ panel unit root test of Pesaran (2007). * indicates rejection of the null hypothesis at 5% significance level.

Table 5: The Output-Inflation Trade-off: Lucas Variance Hypothesis

	PANEL (A): 5-years St. Dev. Variability					
	(1)	(2)	(3)	(4)	(5)	(6)
	FE	CCEMG	dynCCEMG	FE	CCEMG	dynCCEMG
$\Delta y_{i,t-1}$	0.146**	0.219**	0.213**	0.162**	0.231**	0.229**
	(3.58)	(7.28)	(6.40)	(4.73)	(7.74)	(6.90)
$\Delta x_{i,t}$	0.182**	0.204**	0.198**	0.141**	0.169**	0.164**
	(5.51)	(7.37)	(6.83)	(9.16)	(7.35)	(6.37)
$\sigma^{\pi}_{i,t} \times \Delta x_{i,t}$	−0.593*	−11.637	−11.054			
	(−1.99)	(−1.33)	(−1.29)			
$\sigma^{x}_{i,t} \times \Delta x_{i,t}$				0.292	−1.038	−2.983
				(0.78)	(−0.35)	(−0.87)
Obs	2160	2160	2100	2160	2160	2100
N	60	60	60	60	60	60
T	36	36	35	36	36	35
RMSE	0.0344	0.0250	0.0232	0.0345	0.0245	0.0230
	PANEL (B): GARCH Variability					
	(1)	(2)	(3)	(4)	(5)	(6)
	FE	CCEMG	dynCCEMG	FE	CCEMG	dynCCEMG
$\Delta y_{i,t-1}$	0.154**	0.208**	0.204**	0.156**	0.204**	0.201**
	(4.35)	(7.04)	(6.53)	(4.24)	(6.86)	(6.49)
$\Delta x_{i,t}$	0.154**	0.289**	0.327**	0.137**	0.222**	0.220**
	(6.29)	(2.51)	(2.53)	(5.80)	(4.47)	(3.52)
$\sigma^{\pi}_{i,t} \times \Delta x_{i,t}$	−0.025	−0.860	−1.211			
	(−0.15)	(−1.10)	(−1.17)			
$\sigma^{x}_{i,t} \times \Delta x_{i,t}$				0.074	−0.523	−0.465
				(0.86)	(−1.50)	(−0.98)
Obs	2340	2340	2340	2340	2340	2340
N	60	60	60	60	60	60
T	39	39	39	39	39	39
RMSE	0.0354	0.0263	0.0250	0.0354	0.0265	0.0252

Notes: FE – Fixed Effects estimator. CCEMG – Pesaran's (2006) Common Correlated Effects Mean Group estimator. dynCCEMG – Chudik and Pesaran's (2015) Dynamic Common Correlated Effects Mean Group estimator (augmented with one lag of the cross-section averages). *t*-statistics in parentheses. *RMSE* refers to the root mean squared error. * and ** denotes significance at the 10% and 5% significance levels, respectively.

Table 6: Determinants of the Output-Inflation Trade-off

	PANEL (A): 5-years St. Dev. Variability					
	$\Delta y_{i,t}$			$y_{i,t}$		
	(1)	(2)	(3)	(4)	(5)	(6)
	FE	CCEMG	dynCCEMG	FE	CCEMG	dynCCEMG
$\Delta y_{i,t-1}$	0.138**	0.191**	0.189**			
	(3.51)	(6.13)	(5.64)			
$y_{i,t-1}$				0.963**	0.852**	0.838**
				(90.59)	(47.40)	(39.31)
$\Delta x_{i,t}$	0.192**	0.255**	0.249**	0.204**	0.276**	0.273**
	(5.89)	(7.12)	(6.75)	(6.96)	(7.29)	(7.25)
$\pi_{i,t} \times \Delta x_{i,t}$	−0.096**	−0.584**	−0.537**	−0.113**	−0.647**	−0.583**
	(−2.40)	(−2.58)	(−2.32)	(−3.05)	(−3.05)	(−2.64)
$\sigma_{i,t} \times \Delta x_{i,t}$	−0.957**	−7.932	−7.859	−1.148**	−4.721	−3.616**
	(−2.98)	(−1.45)	(−1.57)	(−4.03)	(−1.25)	(−2.26)
Obs	2160	2160	2100	2160	2160	2100
N	60	60	60	60	60	60
T	36	36	35	36	36	35
Trend	No	No	No	Yes	Yes	Yes
RMSE	0.0342	0.0231	0.0210	0.0340	0.0212	0.0185

	PANEL (B): GARCH Variability					
	$\Delta y_{i,t}$			$y_{i,t}$		
	(1)	(2)	(3)	(4)	(5)	(6)
	FE	CCEMG	dynCCEMG	FE	CCEMG	dynCCEMG
$\Delta y_{i,t-1}$	0.150**	0.175**	0.183**			
	(4.03)	(5.89)	(5.87)			
$y_{i,t-1}$				0.970**	0.897**	0.892**
				(96.39)	(70.17)	(61.93)
$\Delta x_{i,t}$	0.168**	0.306**	0.340**	0.185**	0.249**	0.299**
	(6.49)	(2.84)	(3.09)	(7.23)	(2.59)	(2.03)
$\pi_{i,t} \times \Delta x_{i,t}$	−0.078**	−0.623**	−0.559**	−0.082**	−0.648**	−0.626**
	(−2.01)	(−2.86)	(−2.69)	(−2.18)	(−2.94)	(−2.80)
$\sigma_{i,t} \times \Delta x_{i,t}$	−0.125	−0.421	−0.874	−0.173	0.103	−0.537
	(−0.72)	(−0.58)	(−1.07)	(−0.93)	(0.15)	(−0.44)
Obs	2340	2340	2340	2400	2400	2340
N	60	60	60	60	60	60
T	39	39	39	40	40	39
Trend	No	No	No	Yes	Yes	Yes
RMSE	0.0350	0.0242	0.0225	0.0350	0.0229	0.0212

Notes: FE – Fixed Effects estimator. CCEMG – Pesaran's (2006) Common Correlated Effects Mean Group estimator. dynCCEMG – Chudik and Pesaran's (2015) Dynamic Common Correlated Effects Mean Group estimator (augmented with one lag of the cross-section averages). *t*-statistics in parentheses. *RMSE* refers to the root mean squared error. * and ** denotes significance at the 10% and 5% significance levels, respectively.

Table 7: Determinants of the Output-Inflation Trade-off:

Robustness I – Alternative Estimation Methods

PANEL (A): 5-years St. Dev. Variability

	$\Delta y_{i,t}$			$y_{i,t}$		
	(1)	(2)	(3)	(4)	(5)	(6)
	GMM	MG	dynCCEMG$_3$	GMM	MG	dynCCEMG$_3$
$\Delta y_{i,t-1}$	0.146**	0.171**	0.153**			
	(3.23)	(6.29)	(3.62)			
$y_{i,t-1}$				0.952**	0.854**	0.713**
				(124.24)	(46.77)	(19.40)
$\Delta x_{i,t}$	0.192**	0.225**	0.249**	0.217**	0.258**	0.259**
	(4.28)	(6.39)	(5.68)	(5.89)	(7.14)	(6.31)
$\pi_{i,t} \times \Delta x_{i,t}$	−0.156**	−0.711**	−0.689**	−0.193**	−0.743**	−0.630**
	(−2.29)	(−2.97)	(−2.48)	(−2.91)	(−3.22)	(−2.92)
$\sigma_{i,t} \times \Delta x_{i,t}$	−0.549	−3.661	−0.347	−0.837**	−1.908	−5.416
	(−1.31)	(−0.73)	(−0.10)	(−2.36)	(−0.60)	(−1.09)
Obs	2160	2160	1980	2160	2160	1980
N	60	60	60	60	60	60
T	36	36	33	36	36	33
Trend	No	No	No	Yes	Yes	Yes
RMSE	0.0315	0.0276	0.0157	0.0338	0.0254	0.0126

PANEL (B): GARCH Variability

	$\Delta y_{i,t}$			$y_{i,t}$		
	(1)	(2)	(3)	(4)	(5)	(6)
	GMM	MG	dynCCEMG$_3$	GMM	MG	dynCCEMG$_3$
$\Delta y_{i,t-1}$	0.167**	0.182**	0.158**			
	(4.11)	(6.92)	(4.23)			
$y_{i,t-1}$				0.961**	0.873**	0.876**
				(94.92)	(71.76)	(36.15)
$\Delta x_{i,t}$	0.157**	0.329**	0.344**	0.198**	0.274**	0.148
	(5.18)	(3.36)	(2.11)	(6.64)	(3.04)	(1.01)
$\pi_{i,t} \times \Delta x_{i,t}$	−0.116*	−0.808**	−0.610**	−0.138**	−0.801**	−0.550**
	(−1.74)	(−3.28)	(−2.65)	(−2.17)	(−3.38)	(−2.49)
$\sigma_{i,t} \times \Delta x_{i,t}$	−0.004	−0.504	−1.222	−0.136	0.134	0.597
	(−0.03)	(−0.72)	(−0.92)	(−0.66)	(0.22)	(0.51)
Obs	2340	2340	2220	2400	2400	2220
N	60	60	60	60	60	60
T	39	39	37	40	40	37
Trend	No	No	No	Yes	Yes	Yes
RMSE	0.0324	0.0285	0.0180	0.0350	0.0269	0.0166

Notes: GMM – Blundell and Bond's (1998) system GMM estimator. MG – Pesaran and Smith's (1995) Mean Group estimator. dynCCEMG$_3$ – Chudik and Pesaran's (2015) Dynamic Common Correlated Effects Mean Group estimator (augmented with three lags of the cross-section averages). t-statistics in parentheses. $RMSE$ refers to the root mean squared error. * and ** denotes significance at the 10% and 5% significance levels, respectively.

Table 8: Determinants of the Output-Inflation Trade-off:

Robustness II – Alternative Samples

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	’70 - ’10	’82 - ’07	’70 - ’07	’70 - ’10	’82 - ’07	’70 - ’07	’70 - ’10	’82 - ’07	’70 - ’07
	(FULL)	(FULL)	(FULL)	(BMR)	(BMR)	(BMR)	(OECD)	(OECD)	(OECD)
$\Delta y_{i,t-1}$	0.183**	0.191**	0.157**	0.211**	0.211**	0.177**	0.243**	0.311**	0.226**
	(5.87)	(4.49)	(4.83)	(5.48)	(3.79)	(4.29)	(4.90)	(4.42)	(4.38)
$\Delta x_{i,t}$	0.340**	−0.219	0.354**	0.238*	0.162	0.268**	0.374**	0.165*	0.417**
	(3.09)	(−0.49)	(3.29)	(1.85)	(1.40)	(2.04)	(3.09)	(1.91)	(3.11)
$\pi_{i,t} \times \Delta x_{i,t}$	−0.559**	−0.856**	−0.547**	−0.476**	−0.776*	−0.463**	−0.450*	−0.761	−0.438*
	(−2.69)	(−2.28)	(−2.60)	(−2.66)	(−1.74)	(−2.57)	(−1.71)	(−1.21)	(−1.65)
$\sigma_{i,t} \times \Delta x_{i,t}$	−0.874	3.593	−0.913	−0.283	0.250	−0.512	−2.234**	−0.425	−2.470**
	(−1.07)	(0.96)	(−1.13)	(−0.27)	(0.26)	(−0.48)	(−2.10)	(−0.67)	(−2.06)
Obs	2340	1560	2160	1638	1092	1512	936	624	864
N	60	60	60	42	42	42	24	24	24
T	39	26	36	39	26	36	39	26	36
RMSE	0.0225	0.0178	0.0218	0.0187	0.0148	0.0185	0.0138	0.0105	0.0133

Notes: Estimations are based on the dynCCEMG – Chudik and Pesaran’s (2015) Dynamic Common Correlated Effects Mean Group estimator (augmented with one lag of the cross-section averages). FULL refers to the full sample of 60 countries. BMR refers to the sample of countries used in the Ball, Mankiw, and Romer (1988). OECD refers to the sample of OECD countries used as an approximation for Lucas (1973) sample. *t*-statistics in parentheses. *RMSE* refers to the root mean squared error. * and ** denotes significance at the 10% and 5% significance levels, respectively.

Table 9: Determinants of the Output-Inflation Trade-off:

Robustness III – Alternative Specification Forms

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
$\Delta y_{i,t-1}$	0.189**	0.204**	0.183**	0.161**	0.191**	0.175**	0.181**	0.187**	0.187**	0.145**	0.194**	0.254**	0.300**
	(6.49)	(6.53)	(5.87)	(5.19)	(5.95)	(5.15)	(5.97)	(5.35)	(5.75)	(4.41)	(5.86)	(7.06)	(8.10)
$\Delta x_{i,t}$	0.230**	0.327**	0.340**	−0.003	0.105	1.119	0.216**	−1.024	1.222	4.401	0.415**	0.406**	0.454**
	(6.17)	(2.53)	(3.09)	(−0.01)	(1.62)	(1.07)	(3.44)	(−0.98)	(1.56)	(0.87)	(2.97)	(3.62)	(3.61)
$\pi_{i,t} \times \Delta x_{i,t}$	−0.648**		−0.559**	−0.859*			−0.592**	−0.839**	−0.564**	−0.826*	−0.559**	−0.392*	−0.480*
	(−2.55)		(−2.69)	(−1.91)			(−2.50)	(−2.11)	(−2.48)	(−1.82)	(−2.73)	(−1.69)	(−1.81)
$\sigma_{i,t} \times \Delta x_{i,t}$		−1.211	−0.874	1.311					−10.520*	−78.850	−1.440	−0.864	−0.931
		(−1.17)	(−1.07)	(0.17)					(−1.90)	(−0.70)	(−1.33)	(−1.08)	(−1.09)
$(\pi_{i,t} \times \Delta x_{i,t})^2$				0.500				0.960		−0.954			
				(0.28)				(0.52)		(−0.51)			
$(\sigma_{i,t} \times \Delta x_{i,t})^2$				20.532						291.611			
				(0.44)						(0.63)			
$\pi_{i,t}/\Delta x_{i,t}$					0.001*	0.001*							
					(1.68)	(1.69)							
$\sigma_{i,t}/\Delta x_{i,t}$					0.006	−0.258							
					(0.79)	(−1.00)							
$(\pi_{i,t}/\Delta x_{i,t})^2$						0.000							
						(0.98)							
$(\sigma_{i,t}/\Delta x_{i,t})^2$						0.018							
						(1.09)							
$\sigma^x_{i,t} \times \Delta x_{i,t}$							0.058	23.935	1.438	21.172			
							(0.13)	(1.29)	(0.64)	(0.41)			
$(\sigma^x_{i,t} \times \Delta x_{i,t})^2$								−115.166		−98.795			
								(−1.36)		(−0.43)			
Δz_t											−0.013**		
											(−2.00)		
$\pi_{i,t} \times \Delta y_{i,t-1}$												−2.293**	−2.193**
												(−6.67)	(−6.31)
$\Delta x_{i,t-1}$													−0.031**
													(−2.03)
Obs	2340	2340	2340	2340	2340	2340	2340	2340	2340	2340	2340	2280	2280
N	60	60	60	60	60	60	60	60	60	60	60	60	60
T	39	39	39	39	39	39	39	39	39	39	39	38	38
RMSE	0.0246	0.0250	0.0225	0.0184	0.0234	0.0204	0.0219	0.0180	0.0200	0.0134	0.0201	0.0194	0.0187

Notes: Estimations are based on the dynCCEMG – Chudik and Pesaran's (2015) Dynamic Common Correlated Effects Mean Group estimator (augmented with one lag of the cross-section averages). *t*-statistics in parentheses. * and ** denotes significance at the 10% and 5% significance levels, respectively.

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