The Small Open Economy New Keynesian Phillips Curve: Specification, Structural Breaks and Robustness*

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This version: December 23, 2019

Abstract

This paper empirically assesses the concern on whether the slope of the Phillips curve with respect to the output gap has decreased (i.e. the Phillips curve has "flattened"). We derive a generalized lag-augmented version of the New-Keynesian Phillips Curve for a small open economy (Galí and Monacelli, 2005) in order to specify a semi-structural estimation equation. For the Peruvian economy, such equation is estimated *via* the Generalized Method of Moments for the Inflation-Targeting regime (January 2002 - March 2019) and the post-crisis (January 2008 - March 2019) periods. We found that the slope parameter has remained stable for both estimation periods. Moreover, the expectation channel has gained more relevance for the post-crisis period, a result that is consistent with a lower persistence of inflation dynamics. Our results are also consistent with the presence of long run nominal homogeneity across estimation samples.

JEL Classification: C22, C51, E31.

Keywords: New-Keynesian Phillips curve, small open economy, Generalized Method of Moments.

*I thank Marco Vega, Paul Castillo, Youel Rojas, and Raimundo Chirinos for comments and suggestions, and María Gracia Ramos who additionally provided support with data. The usual disclaimer applies.

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1 Introduction

One of the most important ways in which monetary policy affects inflation is through its effects on economic activity. Such channel is usually represented by a (positive) relation between inflation and a measure of "inflationary pressures" known as the Phillips curve and inspired by the work of Phillips (1958). For the case of Peru, Figure 1 depicts the quarterly evolution of the four-quarter core inflation and the cyclical component of GDP (a.k.a. output gap) since 1999. It can be noticed that, since the adoption of the Inflation Targeting regime in 2002, such relation has apparently prevailed until 2013 (gray shaded area). However, it can also be noticed that the same relation has apparently eroded from 2014 onwards (dark gray shaded area) which naturally raises concerns about the effectiveness of monetary policy.

From a technical point of view, the previous discussion is often organized in terms of the New-Keynesian Phillips Curve (henceforth, NKPC)

$$\pi_t = \beta E_t \pi_{t+1} + \kappa(y_t - g_t) + \varepsilon_t$$

(see Clarida et al., 1999) where π_t denotes the inflation rate, $E_t \pi_{t+1}$ denotes the expected inflation rate, y_t denotes the output level, g_t denotes the potential output level, β and κ are positive constants, and ε_t is a random disturbance term. In this regard, the recent episodes of economic contraction and lack of deflationary pressures led to a concern on whether the Phillips curve has "flattened" or, equivalently, the slope parameter κ has decreased.

In this paper, we perform a semi-structural estimation of a NKPC for Peru in order to answer whether or not the Phillips curve has flattened. Our approach incorporates some novel features. First, our reduced-form specification (and model-based sign restrictions) arises from our derivation of a hybrid (lag-augmented) version of the small open economy NKPC by Galí and Monacelli (2005) in order to account for inflation persistence. Second, our extension is compatible with monthly data available for the entire Inflation Targeting regime adopted by the Central Reserve of Peru in 2002. Third, our estimates are obtained *via* the Generalized Method of Moments (GMM) developed by Hansen (1982) and the moment-selection criteria proposed by Andrews (1999). Finally, we report our results for two estimation samples, 2002-2019 (entire Inflation Targeting regime) and 2008-2019 (post-crisis period), in order to check for parameter stability.

Our findings are summarized as follows. First, our estimates support the theory-based sign restrictions. Second, for both estimation samples, the slope parameter has remained stable and thus the corresponding channel of monetary policy is unaltered. Third, compared to the full Inflation-Targeting regime, for the post-crisis period the expectation channel has gained more relevance and this finding is consistent with a lower persistence of inflation. Finally, our results are consistent with the presence of long run nominal homogeneity for both estimation samples.

The rest of this paper is organized as follows. Section 2 provides a (non-exhaustive) review of related literature, with a special emphasis on alternative derivations of the NKPC. Section 3 presents the theoretical framework that leads to the (semi-structural) specification to be employed in the estimation process. Section 4 briefly describes both the GMM estimator and our testable hypotheses of interest. Section 5 reports our estimation results. Section 6 concludes.

2 Related Literature

According to the relationship known as the Phillips curve, strengthening of the economy is commonly associated with increasing inflation. With inflation having only modestly picked up in the aftermath of the Great Recession around the world, many believe the Phillips curve relationship has weakened, with the curve becoming flatter. The implications of such a change include that a positive output gap would be less inflationary, but the cost of reducing inflation, once established, would increase. Some argue that the flattening of the Phillips curve (observed in the industrial countries) has been attributed to globalization, in contrast with the traditional explanation centered on monetary policy credibility. The empirical literature is in general not conclusive. Other argue that micro data is needed to identify whether changes in the slope of the Phillips curve are structural.

For the case of Spain, Bentolila et al. (2008) argue that over the 1995-2006 period the New Keynesian Phillips curve was shifted by inmigration as natives' and immigrants' labor supply elasticities and bargaining power differed.

Kuttner and Robinson (2010) review the evidence and possible explanations for the flattening of the Phillips curve within the New Keynesian framework. Using data for the United States and Australia, they find that the flattening is evident in the baseline (structural) New Keynesian Phillips curve. They also consider several reasons for the structural flattening such as: data problems, globalization and alternative definitions of marginal cost, and none of them is entirely satisfactory. On the contrary, their estimates suggest that there has been a change in the price-setting behavior of firms, potentially because of the lower-trend inflation resulting from the improved conduct of monetary policy. Alternatively, they argue that the expectations formation process may have changed, a point that is also stressed out by Dennery (forthcoming).

Iakova (2007) estimates a small macroeconomic model for the UK economy in order to analyze the implication of a reduction in the responsiveness of inflation to domestic demand pressures due to globalization. The author concludes that the monetary policy implications of a flattening Phillips curve will be different than those when the flattening is related to increased monetary policy credibility (which is likely to be a factor in the initial years after the introduction of an inflation targeting regime). She also stresses out the importance of empirically differentiating between the possible causes of a structural change at any point in time when setting policy, and that the implications of the structural change for the volatility and speed of adjustment of macroeconomic variables have to be communicated clearly to the public to ensure that inflation expectations remain anchored around the target.

Gaiotti (2010) takes advantage of a unique data set including about 2,000 Italian firms, and tests i) whether a change in the link between capacity utilization and prices is confirmed at the company level and ii) whether such change is concentrated among those firms that are more exposed to foreign competition. The answer is either inconclusive or negative in all cases. The results do not lend support to the view that the flattening of the Phillips curve is due to globalization.

Coibion and Gorodnichenko (2015) argue that if firms' inflation expectations track those of households, then the missing disinflation during the Great Recession can be explained by the rise in their inflation expectations between 2009 and 2011. These authors present new survey evidence (consistent with firms having similar expectations as households) where the rise in household inflation expectations from 2009 to 2011 can be explained by the increase in oil prices over the same time period.

Recently, Bullard (2018) used a standard New Keynesian model (dynamic IS equation, a structural New Keynesian Phillips curve and a Taylor-type monetary policy rule) to show that, under the (constrained optimal) promise to react aggressively to deviations of inflation from target in conducting monetary policy, the Phillips curve becomes flat. He concludes that although the model economy considered still has a positively-sloped structural Phillips curve, it is only the (bivariate) empirical Phillips curve that is "disappearing."

Karlsson and Österholm (2018) investigate the stability of the US Phillips curve by assessing the importance of time-varying parameters and stochastic volatility. The authors employ bivariate Bayesian VARs of Personal Consumption Expenditures (PCE) inflation and the unemployment rate (under a number of different assumptions concerning the dynamics and covariance matrix) for quarterly data from 1990Q1 to 2017Q3 and find support for both time-varying parameters and stochastic volatility. After interpreting the Phillips curve as the inflation equation of a Bayesian VAR, they conclude that the US Phillips curve has been unstable and may have been somewhat flatter between 2005 and 2013 than in the decade preceding that period. In a similar exercise for the Swedish unemployment rate and inflation (Karlsson and Österholm, forthcoming) with quarterly data from 1995Q1 to 2018Q3, the same authors find that the evidence in favor of a stable dynamic relationship between the unemployment rate and inflation is rather mixed.

Alternatively, Gagnon and Collins (2019) argue that the Phillips curve may be nonlinear when inflation is low, with the US economy having operated in the flat region of the curve for most of the past 20 years. In this regard, a flat Phillips curve implies little change in inflation going forward, but a nonlinear curve implies moderate increases in inflation over the next few years.

Jacob and van Florenstein Mulder (2019) investigate the potential causes of the flattening of the New Zealand Phillips curve by relying on a simple structural model in which inflation and economic activity move in the same direction conditional on demand shocks (reflecting random changes in the economy's rate of time preference, in the financial sector or monetary policy, or in demand components such as government spending, investment or export demand), and in opposite directions conditional on supply shocks (which may capture random shifts in firms' market power, labor market frictions, in import prices, or in price or wage inflation expectations). The overall correlation between inflation and activity in the model is influenced by the relative strength of the two types of shocks, which in turn is determined by the respective volatilities of the shocks and by structural features of the economy that amplify or weaken shock transmission. They show that the Phillips curve can flatten in an economy in which supply shocks are more dominant.

Occhino (2019) shows that the flattening of the Phillips curve can be due to either changes in the structure of the economy unrelated to policy or changes in the behavior of monetary policy itself. In this regard, knowing the type of change that has occurred is crucial for choosing the appropriate monetary policy (that is, simply knowing that the Phillips curve has flattened is not enough). The study also shows how the adoption of a new monetary policy rule, unresponsive to output and slightly more aggressive toward inflation, can have opposite effects on household welfare, depending on the cause of the flattening.

For McLeay and Tenreyro (2019), a targeting monetary policy rule that aims to minimize welfare losses (subject to a Phillips curve) will transmit a negative correlation between inflation and the output gap, which in turn blurs the identification of a (positively sloped) Phillips curve. The authors discuss several strategies to overcome the former identification problem and present evidence of a robust Phillips curve for the US. Moreover, Murphy (2018) suggests that the slopes of the price and wage Phillips Curves in the US are low and have gotten a little flatter. The dynamic forecasts obtained from the wage and price Phillips curves suggest that the low current inflation is not that surprising and that factors such as increased globalization, increased e-commerce activity, changes in concentration, the aging of the US population and mismeasurement of the NAIRU are not that relevant for explaining this phenomenon.

Pickering and Valle (2008) derive a Phillips curve with imported commodities as an additional input in the production process. The Phillips curve becomes flatter relative to the New Keynesian framework. Moreover, the empirical evidence supports the hypothesis that greater imported commodity intensity in production increases the slope of the Phillips curve. Watson (2016) evaluates the impact of trade openness on the Phillips curve by accounting for the effects of product market competition on price flexibility, and develops a New-Keynesian open-economy DSGE model with non-constant price elasticity of demand and Calvo price-setting in which the frequency of price adjustment is endogenously determined. Within such framework, trade openness has two opposing effects on the sensitivity of inflation to output fluctuations because it raises strategic complementarity in firms' pricing decisions and the degree of real price rigidities (which makes inflation less responsive to changes in real marginal cost) and it also strengthens firms' incentives to adjust their prices, thereby reducing the degree of nominal price rigidities and increasing the sensitivity of inflation to changes in marginal cost.

Recently, Laseen and Sanjani (2016) used multivariate, possibly time-varying, time-series models and show that changes in shocks are a more salient feature of the data than changes in coefficients (i.e. the global financial crisis did not break the Phillips curve). They also show that financial and external variables have the highest forecasting power for inflation and unemployment after the global financial crisis. In this latter regard, Lieberknecht (2018) proposes an explanation for the missing disinflation after the Global Financial Crisis: the interplay between financial frictions, the Phillips curve and the optimal response by central banks. The theoretical framework is a tractable financial accelerator New Keynesian DSGE model that allows for closed-form solutions. Therefore, the presence of financial frictions decreases the slope of the structural Phillips curve via a counter-cyclical credit spread that reduces the pro-cyclicality of marginal costs. Such feature worsens the central bank's trade-off between output gap and inflation stabilization, rendering the former costlier. Within such environment, optimal monetary policy is strongly geared towards inflation stabilization, regardless of the policy regime. Following large contractionary shocks, the optimal response by central banks is thus to mitigate disinflation to a large extent.

3 Theoretical Framework

The starting point towards our econometric model specification is the New-Keynesian framework for a small open economy by Galí and Monacelli (2005) and the reader is referred for further details to Galí (2015, Chapter 8) which is the exposition we borrowed the notation from. Specifically, $\beta \in (0,1)$ is the domestic households' discount factor, $\upsilon \in [0,1]$ represents the share of foreign goods in domestic composite consumption and therefore can be interpreted as a measure of openness, $\eta > 0$ measures the substitutability between domestic and foreign goods, $\epsilon > 1$ denotes the elasticity of substitution between varieties produced domestically, $\sigma > 0$ is the inverse of the intertemporal elasticity of substitution, $\varphi > 0$ is the inverse of the (real) wage elasticity of domestic households' labor supply, $1-\alpha \in (0,1)$ represents the elasticity of domestic output with respect to labor and $\theta \in (0,1)$ measures the fraction of domestic firms that cannot set new prices each period.

Our extension to the previous framework is described as follows: producers who are not allowed to reset their prices rather index them to the last q realizations of the domestic inflation rate $\pi_{H,t-1}$, $\pi_{H,t-2}$, ... and $\pi_{H,t-q}$ with non-negative coefficients ρ_1, ρ_2, ... and ρ_q, respectively. Following Sbordone (2005) and Magnusson and Mavroeidis (2014), it is easy to show that our extension leads to the following hybrid New-Keynesian Phillips Curve for the domestic inflation rate $\pi_{H,t}$

$$\pi_{H,t} = \frac{\rho(L) - \beta\rho_\Delta(L)}{1+\beta\rho_1}\pi_{t-1} + \frac{\beta}{1+\beta\rho_1}E_t\pi_{H,t+1} + \kappa'_\upsilon \tilde{y}_t \tag{1}$$

where the polynomials $\rho(L) = \rho_1 + \rho_2 L + \ldots + \rho_q L^{q-1}$ and $\rho_\Delta(L) = \rho_2 + \rho_3 L + \ldots + \rho_q L^{q-2}$ are expressed in terms of the lag operator L. Also, the slope of (1) with respect to the output gap

$\tilde{y}_t$ is given by $\kappa'_\upsilon = \lambda'(\sigma_\upsilon + \frac{\varphi+\alpha}{1-\alpha}) > 0$ where the terms $\lambda' = \frac{(1-\theta)(1-\beta\theta)}{\theta(1+\beta\rho_1)}\Theta$, $\Theta = \frac{1-\alpha}{1-\alpha+\alpha\epsilon}$, $\sigma_\upsilon = \sigma\Phi$, $\Phi = \frac{1}{1+\upsilon(\varpi-1)}$ and $\varpi = \sigma\eta + (1-\upsilon)(\sigma\eta - 1)$ are all positive in the parameter space.

For $\beta \approx 1$ and $q = 3$ we obtain

$$\pi_{H,t} = \frac{\rho_1 - \rho_2}{1+\rho_1}\pi_{H,t-1} + \frac{\rho_2 - \rho_3}{1+\rho_1}\pi_{H,t-2} + \frac{\rho_3}{1+\rho_1}\pi_{H,t-3} + \frac{1}{1+\rho_1}E_t\pi_{H,t+1} + \kappa'_\upsilon\tilde{y}_t \tag{2}$$

which is expressed in terms of the deep parameters in $(\rho_1, \rho_2, \rho_3, \upsilon, \eta, \epsilon, \sigma, \varphi, \alpha, \theta)$. Some comments are in order. First, (2) imposes no sign restriction on the coefficients associated to either $\pi_{H,t-1}$ or $\pi_{H,t-2}$. Second, the coefficient associated to $\pi_{H,t-3}$ is allowed to be greater than or equal to zero. Third, the coefficients corresponding to the expected domestic inflation $E_t\pi_{H,t+1}$ and the output gap $\tilde{y}_t$ are both strictly positive and provide testable hypotheses. Fourth, the coefficients associated to the lagged and expected (domestic) inflation add up to 1 (i.e. there is long run nominal homogeneity) and this feature also provides a testable hypothesis. Finally, for the case of no indexation ($\rho_1 = \rho_2 = \ldots = \rho_q = 0$), equation (1) leads to the canonical representation of the New-Keynesian Phillips Curve in Galí (2015, Chapter 8, equation 37).

4 Empirical Strategy

4.1 Generalized Method of Moments (GMM) Estimator

The equilibrium New-Keynesian Phillips Curve (2) underlies the following reduced-form equation for estimation purposes:

$$\pi_{H,t} = c_0 + c_1\pi_{H,t-1} + c_2\pi_{H,t-2} + c_3\pi_{H,t-3} + c_{exp}\pi_{H,t+1} + c_{gap}\tilde{y}_t + u_t, \tag{3}$$

where c_0 is a constant term, c_i is the coefficient of the i-th lag of the domestic inflation $\pi_{H,t-i}$ ($i = 1,2,3$) and is intended to capture inflation inertia, c_{exp} is the coefficient of the future domestic inflation $\pi_{H,t+1}$ and is intended to capture the expectation channel,[1] c_{gap} is the coefficient of the output gap $\tilde{y}_t$ (i.e. the "slope" of the New-Keynesian Phillips Curve) and u_t contains the forecasting error $\pi_{H,t+1} - E_t\pi_{H,t+1}$.

Let $x_t \equiv (\pi_{H,t}, \pi_{H,t-1}, \pi_{H,t-2}, \pi_{H,t-3}, \pi_{H,t+1}, \tilde{y}_t)$ contain the variables involved in (3) and let

[1] Although there exists an available series on agents' expectations since the beginning of the Inflation Targeting regime, such information is not employed for two reasons. First, it provides agents' expected headline inflation, whereas our model is posed in terms of domestic inflation. And second, it consists of a 12-month ahead expectation, whereas our model is posed in terms of a one-month ahead expectation.

$c \equiv (c_0, c_1, c_2, c_3, c_{exp}, c_{gap})$ contain the reduced-form coefficients in (3). Also, let

$$m(x_t; c) \equiv \pi_{H,t} - (c_0 + c_1\pi_{H,t-1} + c_2\pi_{H,t-2} + c_3\pi_{H,t-3} + c_{exp}\pi_{H,t+1} + c_{gap}\tilde{y}_t) \qquad (4)$$

denote the forecasting error u_t and c^0 denote the coefficient vector of the data generating process. Under rational expectations, the equation (3) evaluated at $c = c^0$ implies that the unconditional expectation of the forecasting error u_t equals zero (i.e. $E\left[m(x_t; c^0)\right] = 0$). Also, under rational expectations such forecasting error is uncorrelated to any variable in the agents' information set. Let $z_{j,t}$ $(j = 1, \ldots, p)$ represent such variable. Then, the previous description leads to p moment conditions for p (instrumental) variables $\{z_{1,t}, \ldots, z_{p,t}\}$ in the information set the form $E\left[z_{j,t}m(x_t; c^0)\right] = 0$ $(j = 1, \ldots, p)$ or, compactly,

$$E\left[Z_t m(x_t; c^0)\right] = 0 \qquad (5)$$

where $Z_t = [z_{1,t} \ \ldots \ z_{p,t}]'$ is the vector of instrumental variables. Since the vector c contains six coefficients, we restrict to the case of over-identification by assuming $p > 6$. The Generalized Method of Moment (GMM) estimator by Hansen (1982) estimates c^0 by finding the c that makes the sample analogue of (5) as close to zero as possible through the use of a weighing matrix. Specifically, for a sample of size T the GMM estimator $\hat{c}_{GMM}$ minimizes

$$\mathcal{L}_{GMM}(c) \equiv g'_T(c)\,\hat{N}_u^{-1} g_T(c)\,, \text{ with } g_T(c) = T^{-1}\sum_{t=1}^{T} g_t(c), \qquad (6)$$

where $g_t(c) = Z_t m(x_t; c)$ and $\hat{N}_u \xrightarrow{p} N_u = \texttt{lim}_T\left[\texttt{Var}[\sqrt{T}g_T(c^0)]\right] \equiv \texttt{AVar}[\sqrt{T}g_T(c^0)]$.

The reader should notice that the estimator is based on the assumption that the vector of instruments Z_t satisfies the over-identifying conditions (5). For large T and under the null hypothesis that such over-identifying restrictions are all valid, the Sargan's J-statistic $J_T(\hat{c}_{GMM}) \equiv Tg'_T(\hat{c}_{GMM})\,\hat{N}_u^{-1} g_T(\hat{c}_{GMM})$ is chi-squared distributed with $p - 6$ degrees of freedom and cumulative distribution denoted by F. Let $\tilde{\alpha}$ denote the chosen significance level. Therefore, we reject the null hypothesis of over-identification if the calculated p-value $1 - F(J_T(\hat{c}_{GMM}))$ is greater than $\tilde{\alpha}$ and cannot reject it otherwise.

Our consistent moment selection follows Andrews (1999) as it involved a search along vectors Z_t that contain a constant and instruments within the set $\{\pi_{H,t-k}, \tilde{y}_{t-k}\}_{k=1}^{k_{max}}$ and GMM analogues of the Bayesian, Akaike and Hannan-Quinn Information Criteria for moment selection were employed. We refer to them as GMM-BIC, GMM-AIC and GMM-HQIC and define them by

$$
\begin{aligned}
\text{GMM-BIC} &: \quad MSC_{BIC,T}\left(Z_t\right) = J_T\left(Z_t\right) - (p-6)\log T; \\
\text{GMM-AIC} &: \quad MSC_{AIC,T}\left(Z_t\right) = J_T\left(Z_t\right) - 2\times(p-6)\log T; \\
\text{GMM-HQIC} &: \quad MSC_{HQIC,T}\left(Z_t\right) = J_T\left(Z_t\right) - 2.01\times(p-6)\log\log T;
\end{aligned}
$$

where Z_t is a vector containing p instruments, and log denotes natural logarithm.

4.2 Data

Our theoretical framework implies an econometric specification involving only two variables: the period-to-period domestic inflation rate and the output gap. We consider monthly data for the period from January of 2002 to March of 2019 for two reasons. First, a (New-Keynesian) Phillips Curve is one of the key ingredients of the Inflation Targeting (IT) regime that the Central Reserve Bank of Peru (BCRP) adopted in 2002, in which monetary policy decisions are made on a monthly basis and there is available data for the same period and frequency. Second, the use of monthly data implies 207 observations from January of 2002 to March of 2019 which, unlike the 69 quarterly observations for the same time span, provides more data variability in the estimation process. The latter feature will allow our statistical inference to rely on large-sample distributions as we assume that they fairly approximate finite-sample distributions of tests statistics.

Our theoretical framework is also explicit regarding the variables to include and the transformations to perform. Also, all raw variables were obtained from the Central Reserve Bank of Peru's database. The variable representing domestic inflation $\pi_{H,t}$ is given by $100\times\Delta\log(\text{IPC\_h})$, the (natural) logarithm of the domestic component of the monthly Consumer Price Index in first differences. On the other hand, a proxy for the output gap $\tilde{y}_t$ is given by $100\times\text{output\_gap}$, the difference between the (natural) logarithm of the seasonally-adjusted monthly Economic Activity Index and its Hodrick-Prescott (HP) filter trend.[2] It is worth to mention that our seasonal adjustment made use of the automatic mode of the programs TRAMO and SEATS

[2] For this purpose, the standard value of the smoothing parameter ($\lambda = 14,400$) was employed. Also, in order to mitigate the end-point bias, the calculations also included the ARIMA forecasts from April of 2019 to December of 2019. Finally, it is worth to emphasize that, unlike $\tilde{y}_t$, the use of filtered data implies that the error term u now also contains the irregular components of the flexible-price output level such as preference and technology shocks.

which implement the methodology proposed by Gomez and Maravall (1994) and available on the Bank of Spain's website.

4.3 Specification and testable hypotheses

Therefore, a semi-structural specification based on (2) and suitable for estimation is

$$\begin{aligned}\Delta\log(\text{IPC\_h}) &= c_0 + c_1 \times \Delta\log(\text{IPC\_h(-1)}) + c_2 \times \Delta\log(\text{IPC\_h(-2)}) \\ &+ c_3 \times \Delta\log(\text{IPC\_h(-3)}) + c_{exp} \times \Delta\log(\text{IPC\_h(+1)}) + c_{gap} \times \text{output\_gap} + \text{u} \quad (7)\end{aligned}$$

where u is an error term that contains preference and technology shocks and domestic inflation forecasting errors as we include the actual future domestic inflation rate $\Delta\log(\text{IPC\_h(+1)})$ instead of its conditional expectation. In addition to the usual tests of significance, there are three hypotheses we are interested in:

1. $H_0 : c_{exp} \leq 0$ against $H_1 : c_{exp} > 0$ (expectations matter in the NKPC),
2. $H_0 : c_{gap} \leq 0$ against $H_1 : c_{gap} > 0$ (positive slope of the NKPC), and
3. $H_0 : c_1 + c_2 + c_3 + c_{exp} = 1$ (long run homegeneity) against $H_1 : c_1 + c_2 + c_3 + c_{exp} \neq 1$.

It is in this regard that rejecting the null hypothesis in 1 would support the alternative hypothesis that the expectations are relevant for domestic inflation dynamics. A similar description applies to the hypotheses in 2 regarding the slope of the New-Keynesian Phillips Curve and thus the effect of the output gap. Finally, the null hypothesis in 3 is consistent with long run nominal homogeneity as specified by our theoretical model.

5 Results

5.1 Unit Root Testing

As it is customary, the detection of unit roots becomes relevant for the specification of our empirical model. This so happens because all the variables included in (7) are assumed to be stationary. For this reason, the unit root tests by Dickey and Fuller (1979), Said and Dickey (1984) and Phillips and Perron (1988) are reported in Table 1. We reject the null hypothesis that output_gap contains a unit root and cannot reject the null hypothesis that log(IPC_h) contains a unit root. Both results hold for all conventional significance levels (1%, 5% and 10%) and regardless of the specification of the deterministic component. The same results are obtained for the efficient test developed by Elliott et al. (1996) in Tables 2 and 3, and for the class of M-test by Ng and Perron (2001) in Table 4 that overcome a series of well known limitations involving the power loss of unit root tests against local alternatives.

Nevertheless, it can be noticed in Figure 2 that log(IPC_h) seems to exhibit a trend shift. A similar pattern is observed for output_gap in Figure 6. According to Perron (1989), such abrupt shifts distort conventional unit root tests and lead to an over acceptance of the unit root hypothesis. For this reason, Table 5 reports the unit root tests proposed by Perron and Rodríguez (2003) which allow for the presence of a structural change. That is, a trend shift is allowed and "controlled" in a robust fashion while testing for unit roots. Once again, we reject the null hypothesis that output_gap contains a unit root and cannot reject the null hypothesis that log(IPC_h) contains a unit root at all conventional significance levels.

However, and by construction, the tests by Perron (1989) pre-assume the existence of a break with non-trivial effects on its power. Moreover, a detected break date can turn out to be spurious. For this reason. the tests by Cavaliere et al. (2011) pre-test for the existence of a break in the trend function. At the 5% level of significance, we reject the null that output_gap as has a unit root with a structural break. Also, at the same level of significance we cannot reject the null that log(IPC_h) has a unit root with a structural break.

5.2 GMM Estimation and Hypothesis Testing

Table 7 summarizes our estimates of the coefficients in equation (7) for two estimation periods and several instrument sets. Columns I, II and III contain estimates for the period from January of 2002 to March of 2019 (i.e. from the beginning of the Inflation Targeting regime) whereas columns IV, V and VI contain estimates for the period from January of 2008 onwards (i.e. consistent with the last financial crisis) since the univariate unit root test by Perron and Rodríguez (2003) estimates a structural break for the (log of) domestic prices as occurring during January of 2008. The Generalized Method of Moments (GMM) estimator by Hansen (1982) was employed for all of the equations. Also, for all cases, the effective numbers of observations are lower that those implied by the original time span because of the lagged variables being employed as regressors and/or instruments. For each estimation period, we set $k_{max} = 7$ (i.e. the maximum lag used as an instrument) and an exhaustive search for instruments was performed.

Results for instrument sets exhibiting the three lowest moment selection criteria (GMM-BIC, GMM-AIC and GMM-HQIC) are also reported in Table 8. In our search, we filtered out any instrument vector such that the null hypothesis of over-identification is rejected. It is worth to notice that for both estimation periods, each moment selection criterion monotonically decreases (from III to I and from VI to IV). This reflects that the different bonus terms (that reward selection vectors that utilize more moment conditions) have no impact on the corresponding moment-selection criterion and therefore the problem of moment selection reduces to minimize the Sargan's J-statistic with respect to the instrument vector.

On the one hand, from column I it can be asserted that, regarding the sign-unrestricted coefficients, the lagged domestic inflation Δlog(IPC_h(-1)) is significant at any of the conventional significance levels (either 1%, 5% or 10%) and has a positive marginal effect that equals 0.32. On the contrary, neither Δlog(IPC_h(-2)) nor Δlog(IPC_h(-3)) are individually significant at any of the conventional significance levels. The point estimate of the marginal effect of Δlog(IPC_h(-2)) is negative (a possibility captured by the theoretical model). Also, we cannot reject the null hypothesis that Δlog(IPC_h(-3)) is not significant. On the other hand, regarding the sign-restricted coefficients, we reject that the domestic inflation expectation (output gap) has a lower-than-or-equal-to-zero effect at the 10% significance level and conclude that there exists a positive and significative effect. Such conclusion is reflected by a one-sided p-value lower than 0.10. Finally, we cannot reject the null hypothesis that there is long run nominal homogeneity at the 10% significance level, which is reflected by a two-sided p-value that equals 0.13 (greater than a conservative 0.10). A similar analysis applies to both columns II and III.

For the post-crisis period, from column IV it is found that, regarding the sign-unrestricted coefficients, the lagged domestic inflation Δlog(IPC_h(-1)) is again significant at any conventional significance level although the point estimate of the marginal effect now equals 0.17 which is lower than 0.32 for the full-sample estimation. From a standpoint based on a theoretical framework, this sheds light on what structural feature might be driving the change in the inflation dynamics. Namely, after the financial crisis this is consistent with a lower fraction of firms indexing their prices to the previous domestic inflation. Again, neither Δlog(IPC_h(-2)) nor Δlog(IPC_h(-3)) are significant at any of the conventional significance levels. The point estimate of the marginal effect of Δlog(IPC_h(-2)) is again negative but the point estimate for the coefficient of Δlog(IPC_h(-3)) is negative as well, which is at odds with our theoretical formulation. Regarding the sign-restricted coefficients, we once again reject that the domestic inflation expectations (output gap) have (has) a lower-than-or-equal-to-zero effect at the 10% significance level and conclude that there exists a positive and significative effect. Compared with the full sample estimation, the point estimate of the coefficient of the expectations is higher which in turn suggests that the expectations channel has gained more relevance after the financial crisis, even when the marginal effect of the output gaps has remained unaltered for both estimation samples. We cannot reject the null hypothesis that there is long run nominal homogeneity at the 10% significance level, which is reflected by a two-sided p-value equal to 0.11 for columns IV and V. However, such type of result is not reflected in column VI and this partly reflects that the adjusted number of observations (135) is considerably lower than the one originally employed. Under such situation, the large-sample distributions might not constitute an acceptable approximation to their finite-sample counterparts. That being said, the results for the post-crisis period should be interpreted with caution.

6 Conclusions

In this paper, we estimated a reduced-form version of the NKPC for the Peruvian economy and the 2002-2019 period. Our empirical evidence supports the argument that the slope of the Phillips curve for Peru has remained stable. At the same time, the expectation channel has gained more relevance in the aftermath of the last financial crisis and this fact is consistent with a lower fraction of producers indexing their prices. Of course, a model-consistent explanation requires an estimation of the model parameters. In this sense, the GMM estimator under structural change by Antoine and Boldea (2018) is particularly promising for both semi-structural and structural estimation.

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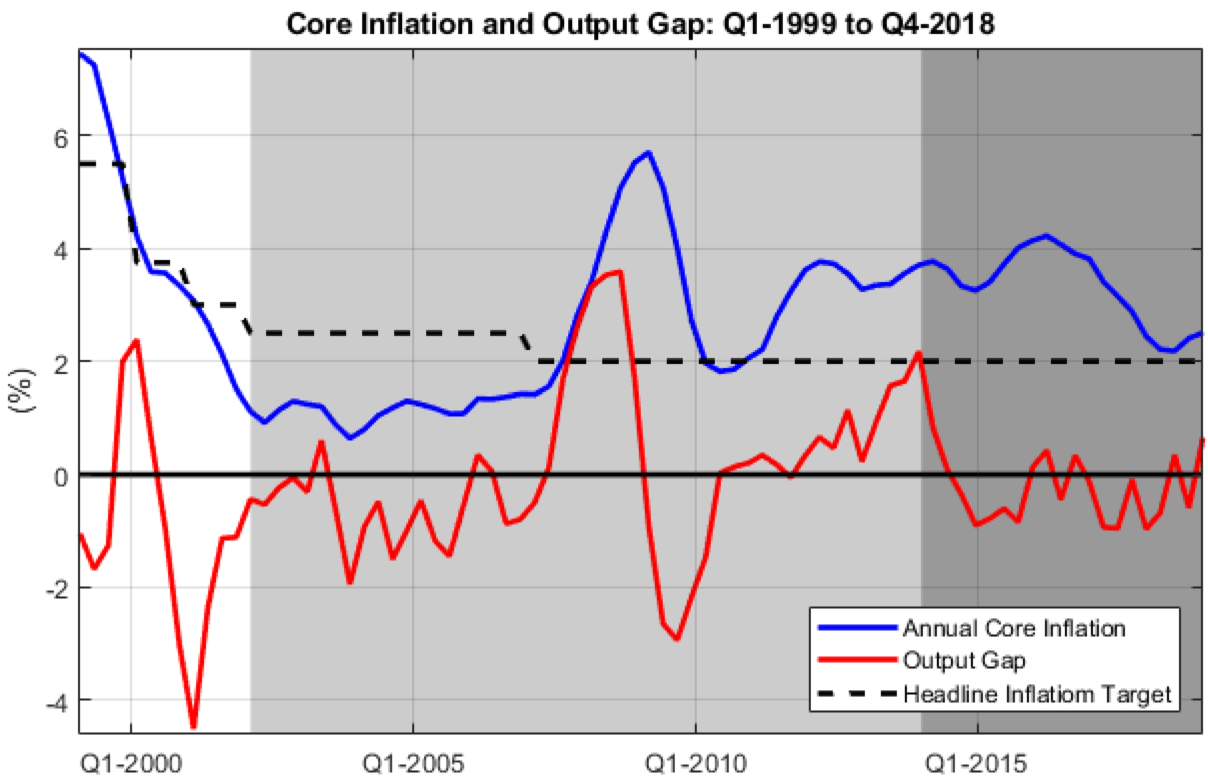

Figure 1: Quarterly Core Inflation and Output Gap

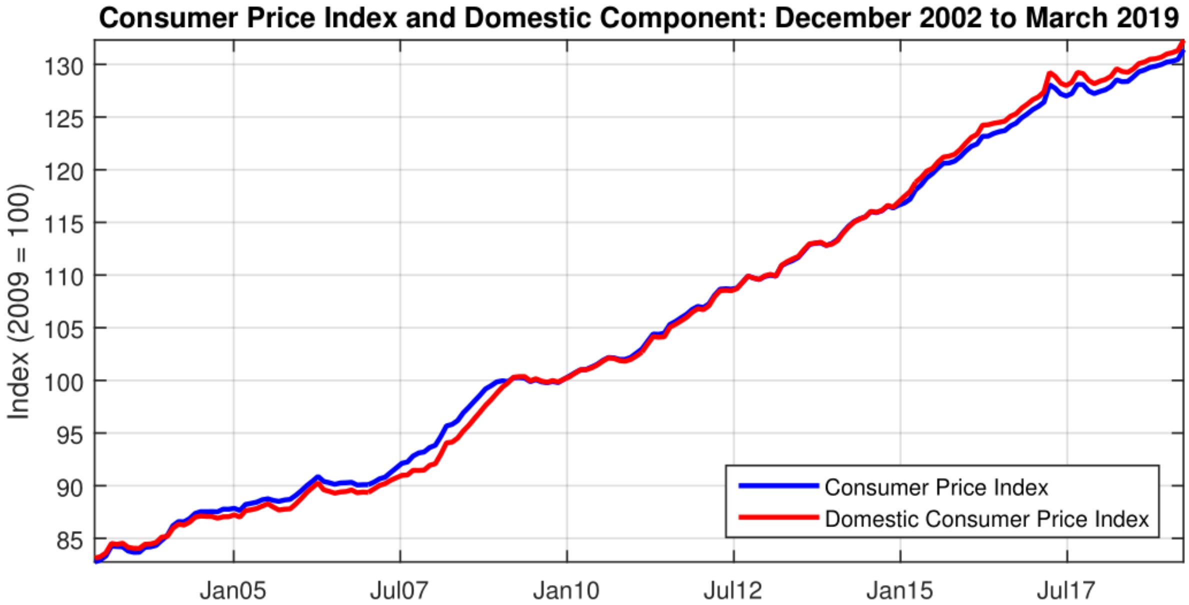

Figure 2: Consumer Price Index and Domestic Component

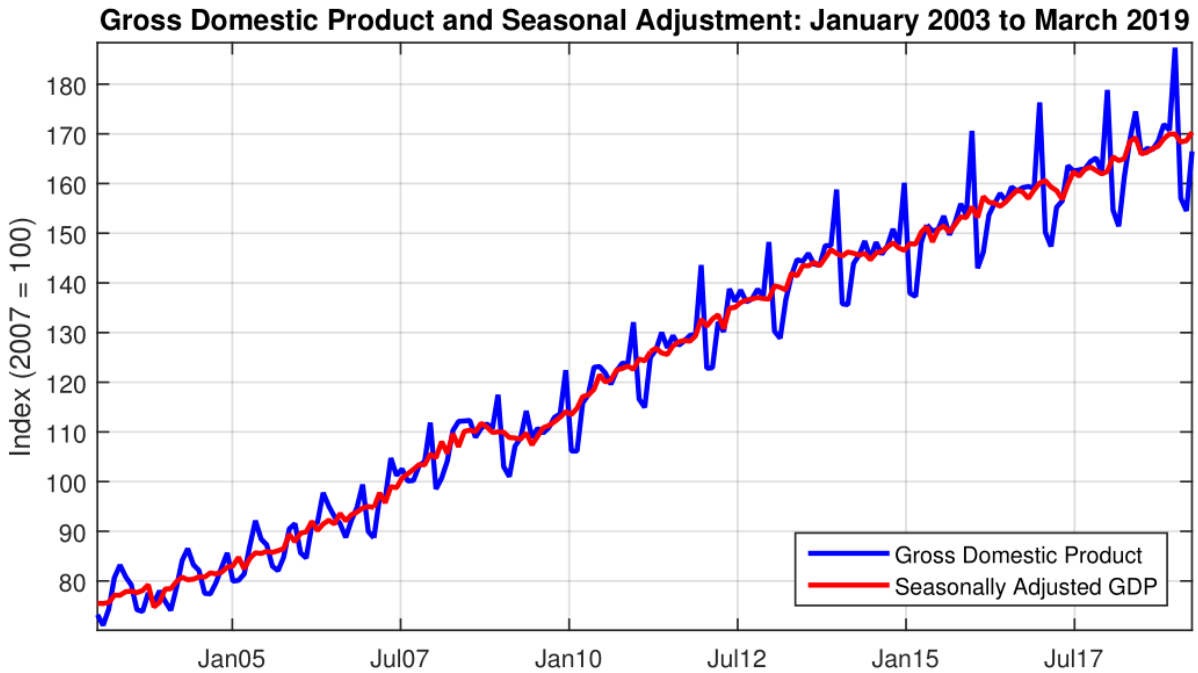

Figure 3: Gross Domestic Product and Seasonal Adjustment

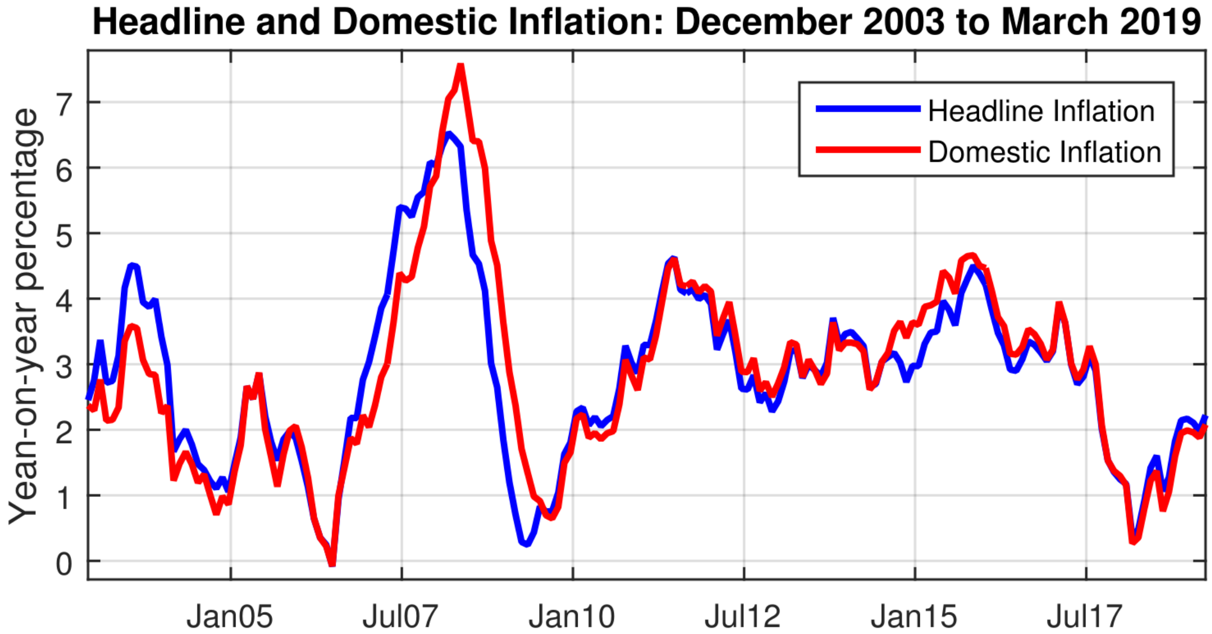

Figure 4: Headline and Domestic Inflation

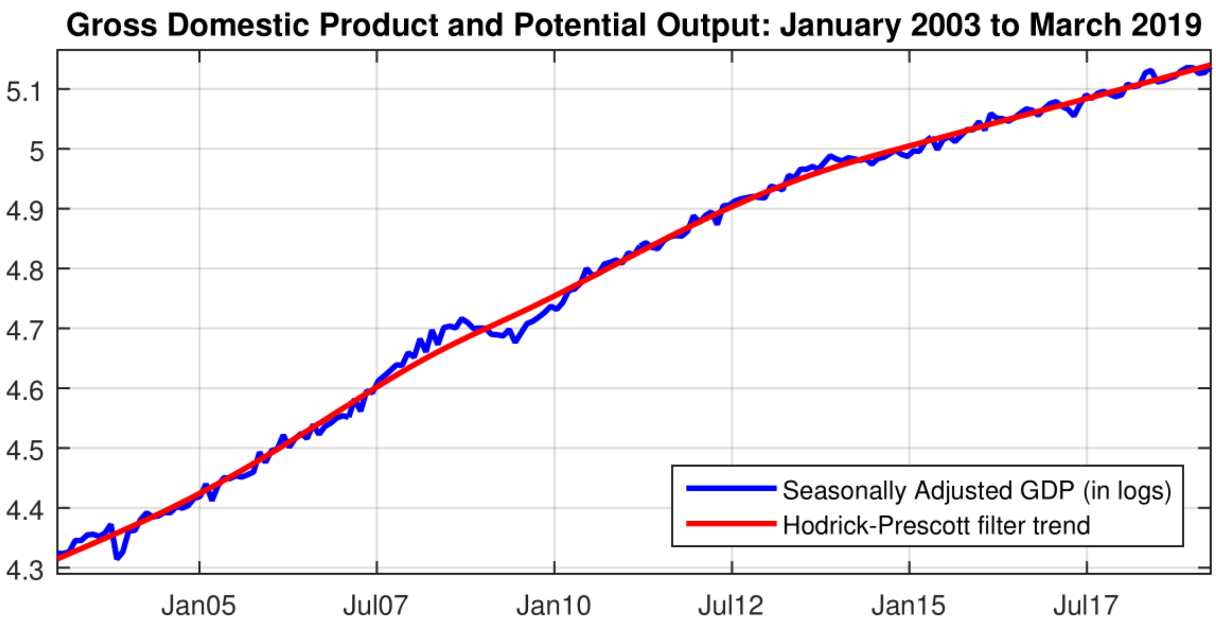

Figure 5: Gross Domestic Product and Potential Output

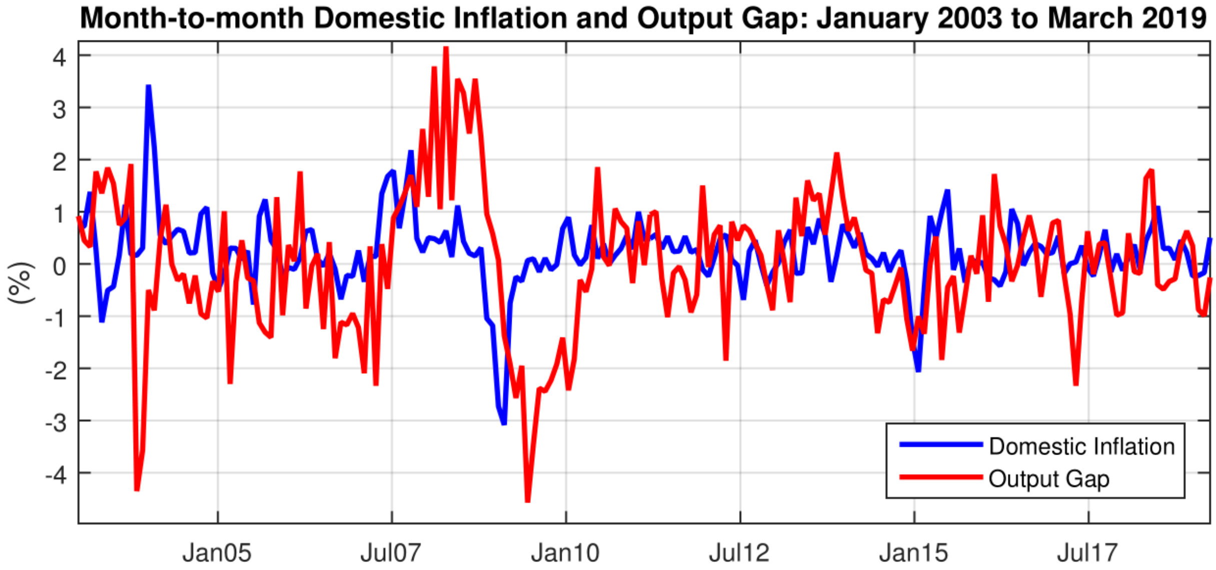

Figure 6: Month to month Domestic Inflation and Output

Table 1: Augmented Dickey-Fuller and Phillips-Perron Unit Root Tests[a]

Augmented Dickey-Fuller Tests				
		No drift nor trend	Drift, no trend	Drift and trend
output_gap		-4.6286***	-4.6167***	-4.6047***
log(IPC_h)		6.3658	0.2604	-2.5142
Critical values[b]	1%	-2.5770	-3.4643	-4.0063
	5%	-1.9425	-2.8764	-3.4333
	10%	-1.6156	-2.5747	-3.1405
Phillips-Perron Tests				
		No drift nor trend	Drift, no trend	Drift and trend
output_gap		-7.9415***	-7.9244***	-7.9066***
log(IPC_h)		8.5337	0.1812	-2.2680
Critical values[b]	1%	-2.5769	-3.4641	-4.0061
	5%	-1.9425	-2.8763	-3.4332
	10%	-1.6156	-2.5747	-3.1404

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.
[b] MacKinnon (1996) one-sided p-values.

Table 2: Elliott-Rothenberg-Stock Unit Root Tests[a]

		Intercept
output_gap		1.0243***
log(IPC_h)		1218.5560
Asymptotic critical values[b]	1%	1.9120
	5%	3.1670
	10%	4.3320
		Trend and Intercept
output_gap		2.8870***
log(IPC_h)		15.0534
Asymptotic critical values[b]	1%	4.0605
	5%	5.6590
	10%	6.8565

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.
[b] Elliott et al. (1996, Table 1).

Table 3: Elliott-Rothenberg-Stock DF-GLS Unit Root Tests[a]

		Intercept
output_gap		-3.8804***
log(IPC_h)		4.2886
Asymptotic critical values[b]	1%	-2.5770
	5%	-1.9425
	10%	-1.6156
		Trend and Intercept
output_gap		-4.3942***
log(IPC_h)		-1.8133
Asymptotic critical values[b]	1%	-3.4684
	5%	-2.9370
	10%	-2.6470

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.
[b] Elliott et al. (1996, Table 1).

Table 4: Ng-Perron Unit Root Tests[a]

		Intercept			
		MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MPT^{GLS}
output_gap		-25.9777***	-3.5904***	0.1382***	0.9886***
log(IPC_h)		1.6456	4.8807	2.9660	639.6520
Asymptotic critical values[b]	1%	-13.8000	-2.5800	0.1740	1.7800
	5%	-8.1000	-1.9800	0.2330	3.1700
	10%	-5.7000	-1.6200	0.2750	4.4500
		Trend and Intercept			
		MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MPT^{GLS}
output_gap		-31.4988***	-3.9684***	0.1260***	2.8938***
log(IPC_h)		-6.4635	-1.7976	0.2781	14.0984
Asymptotic critical values[b]	1%	-23.8000	-3.4200	0.1430	4.0300
	5%	-17.3000	-2.9100	0.1680	5.4800
	10%	-14.2000	-2.6200	0.1850	6.6700

[a] *, ** and *** indicate rejection of the $I(1)$ null hypothesis at the 10%, 5% and 1% level of significance, respectively. Modified or M-tests are described in Ng and Perron (2001). For the the case of the MZ_{α}^{GLS}, MZ_{t}^{GLS} and MSB^{GLS} tests, a statistic lower than the critical value leads to a rejection of the $I(1)$ null hypothesis.

[b] Ng and Perron (2001, Table 1).

Table 5: Perron-Rodríguez Unit Root Tests[a]

		$\sup MZ_{\alpha}^{GLS}$	$\sup MZ_{t}^{GLS}$	$\sup MSB^{GLS}$
output_gap		-27.0414***	-3.6770***	0.1360**
log(IPC_h)		-15.4456	-2.6837	0.1738
Critical values[b]	1%	-27.0000	-3.6600	0.1340
	5%	-22.9000	-3.3500	0.1450
	10%	-20.7000	-3.1900	0.1540

[a] *, ** and *** indicate rejection of the $I(1)$ null hypothesis at the 10%, 5% and 1% level of significance, respectively. Modified or M-tests under structural change are described in Perron and Rodríguez (2003). In the case of $\sup MZ_{\alpha}^{GLS}$, $\sup MZ_{t}^{GLS}$ and $\sup MSB^{GLS}$ tests, a statistic lower than the critical value leads to a rejection of the $I(1)$ null hypothesis.

[b] Perron and Rodríguez (2003), Table 2.

Table 6: Cavaliere-Harvey-Leybourne-Taylor Unit Root Tests[a]

		MZ_{α}	MZ_{t}	MSB	$t(\bar{\tau})$
output_gap		-25.4480**	-3.5670**	0.1400**	-3.7580**
Critical values[b]	(5%)	-16.2190	-2.8150	0.1730	-2.9110
log(IPC_h)		-15.7110	-2.7100	0.1720	-2.7330
Critical values[b]	(5%)	-23.1760	-3.3930	0.1460	-3.6380

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.

[b] Critical values are computed via the bootstrap algorithm by Cavaliere et al. (2011, Section 4).

Table 7: Estimation results[a]

Equation	(I)	(II)	(III)	(IV)	(V)	(VI)
Estimation method	GMM	GMM	GMM	GMM	GMM	GMM
Period	Jan2002-Mar2019	Jan2002-Mar2019	Jan2002-Mar2019	Jan2008-Mar2019	Jan2008-Mar2019	Jan2008-Mar2019
Dependent variable	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)
Constant	0.1613	0.1509	0.1654	0.1921	0.1921	0.1994*
	(0.1016)	(0.1046)	(0.1039)	(0.1306)	(0.1321)	(0.1024)
Δlog(IPC_h(-1))	0.3237***	0.3205***	0.3261***	0.1692***	0.1680***	0.1977***
	(0.1048)	(0.1050)	(0.1062)	(0.0954)	(0.0963)	(0.0864)
Δlog(IPC_h(-2))	-0.5265	-0.5183	-0.5587	-0.0508	-0.0142	-0.0700
	(0.3304)	(0.3312)	(0.3525)	(0.1112)	(0.1113)	(0.0869)
Δlog(IPC_h(-3))	0.1027	0.1008	0.1119	-0.5864	-0.6320	-0.4099
	(0.0916)	(0.0919)	(0.1000)	(0.3356)	(0.3279)	(0.2799)
Δlog(IPC_h(+1))	0.4386+	0.4710+	0.4369+	0.6887+	0.6931+	0.5084+
	(0.3213)	(0.3310)	(0.3274)	(0.4270)	(0.4285)	(0.3463)
One-sided p-value[b]	0.0869	0.0783	0.0919	0.0546	0.0541	0.0722
output_gap	0.0738+	0.0720+	0.0757+	0.0834+	0.0846+	0.0628+
	(0.0473)	(0.0476)	(0.0483)	(0.0586)	(0.0592)	(0.0444)
One-sided p-value[c]	0.0604	0.0661	0.0595	0.0786	0.0777	0.0799
P-value (homogeneity)[d]	0.1326	0.1623	0.1311	0.1088	0.1094	0.0379
No. of observations	189	189	183	135	135	135
No. of instruments	12	11	11	13	12	12
Sargan's J-statistic	0.5903	0.4059	0.5204	4.7417	3.0685	3.9533
Prob(J-statistic)	0.9966	0.9952	0.9914	0.6914	0.8002	0.6830
GMM-BIC	-30.8602	-25.8029	-25.6883	-29.5952	-26.3631	-25.4784
GMM-AIC	-11.4097	-9.5941	-9.4796	-9.2583	-8.9315	-8.0467
GMM-HQIC	-19.3889	-16.2435	-16.1290	-17.6339	-16.1106	-15.2259

[a] *, ** and *** indicate rejection of the null hypothesis of a zero coefficient at the 10%, 5% and 1% level of significance, respectively. Also, +, ++ and +++ indicate rejection of the null hypothesis of a lower-than-or-equal-to-zero coefficient at the 10%, 5% and 1% level of significance, respectively. Reported moment-selection criteria (GMM-BIC, GMM-AIC and GMM-HQIC) are computed as described by Andrews (1999, Section 3).

[b] One-sided p-value for testing $H_0 : c_5 \leq 0$ against $H_1 : c_5 > 0$ in (7).

[c] One-sided p-value for testing $H_0 : c_6 \leq 0$ against $H_1 : c_6 > 0$ in (7).

[d] Two-sided p-value for testing $H_0 : c_2 + c_3 + c_4 + c_5 = 1$ against $H_1 : c_2 + c_3 + c_4 + c_5 \neq 1$ in (7).

Table 8: Instruments[a]

Equation	(I)	(II)	(III)	(IV)	(V)	(VI)
Constant	Yes	Yes	Yes	Yes	Yes	Yes
Δlog(IPC_h(-1))	Yes	Yes	Yes	Yes	Yes	Yes
Δlog(IPC_h(-2))	No	No	No	Yes	Yes	Yes
Δlog(IPC_h(-3))	Yes	Yes	Yes	No	No	No
Δlog(IPC_h(-4))	No	No	No	Yes	Yes	Yes
Δlog(IPC_h(-5))	Yes	Yes	No	Yes	Yes	Yes
Δlog(IPC_h(-6))	Yes	No	Yes	Yes	Yes	Yes
Δlog(IPC_h(-7))	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-1)	Yes	Yes	Yes	No	No	Yes
output_gap(-2)	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-3)	Yes	Yes	Yes	Yes	Yes	No
output_gap(-4)	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-5)	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-6)	Yes	Yes	Yes	Yes	No	No
output_gap(-7)	No	No	No	Yes	Yes	Yes
GMM BIC	-30.8602	-25.8029	-25.6883	-29.5952	-26.3631	-25.4784
GMM AIC	-11.4097	-9.5941	-9.4796	-9.2583	-8.9315	-8.0467
GMM HQIC	-19.3889	-16.2435	-16.1290	-17.6339	-16.1106	-15.2259

[a] Reported moment-selection criteria (GMM-BIC, GMM-AIC and GMM-HQIC) are computed as described by Andrews (1999, Section 3).