On the Slope and the Persistence of the Italian Phillips Curve*

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We investigate the determinants of inertia in Italian inflation, estimating a Phillips curve derived from a general equilibrium business-cycle model that allows for intrinsic and extrinsic sources of inflation persistence, along with trend inflation, and that encompasses both nominal and real rigidities as key factors of the output-inflation trade-off. We perform the estimation over two different sub-samples, 1981:Q1–1998:Q4 and 1999:Q1–2012:Q3, to take into account the structural break represented by the starting of the Economic and Monetary Union. We find that in the period between 1999:Q1 and 2012:Q3, the dependence of Italian inflation on its own past diminished and the slope of the Phillips curve dropped relative to the years before 1999. The latter is a consequence of increased strategic complementarity in price setting, due in turn to higher sensitivity of demand elasticity to firms' relative prices, on top of lower trend inflation and an increase in the average duration of prices.

JEL Codes: E31, E32.

1. Introduction

During the Great Recession inflation in advanced countries hardly responded to conditions in product and labor markets. According to many economists,[1] based on past experiences and given the depth and the duration of the worldwide recession, advanced economies should have experienced severe disinflation—perhaps even deflation.

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[1] See section 2 for a literature review.

Hall (2011) called for a fundamental reconsideration of theories in which inflation depends on a measure of slack, reading the recent experience as a contradiction to the axiomatic relationship between slack and declining inflation and concluding that inflation behaves in a nearly exogenous fashion. A debate has emerged to find possible tweaks to the traditional Phillips curve and explain why extreme economic slump has done little to reduce inflation in most advanced economies.

In this paper we focus on Italy. After the outburst of the sovereign debt crisis in the euro area, GDP in Italy contracted bluntly, by almost 5 percent from mid-2011 to the end of 2013, more than the euro-area average. However, at the beginning of this prolonged recession, inflation remained stubbornly high, above 3 percent year on year until the last months of 2012. Although the rise in indirect taxes contributed by around 1 percentage point, even the price index computed net of the increase in indirect taxation grew at a faster pace than in the euro area. In 2013 inflation started moderating, and it dropped below 1 percent in the second half of the year. However, core inflation (i.e., net of food and energy prices) remained hovering around 1 percent, despite an unprecedented contraction of domestic demand. This outturn suggests that in the short run the relationship between economic activity and price dynamics (i.e., the Phillips curve) might be relatively weak in Italy, hence justifying such a slow and, so far, moderate response of inflation to the collapse of output.

Nevertheless, this reduced-form evidence should be considered with caution, since it is plagued by the Lucas critique, as it might be driven by monetary policy. Moreover, various factors may lie behind a slow adjustment of inflation. Some of them are associated with what Altissimo, Ehrmann, and Smets (2006) define as "extrinsic persistence," namely that inherited from persistent fluctuations in the determinants of inflation such as marginal costs or the output gap. Additionally, inflation can show "intrinsic persistence," if it depends on its own past. In a New Keynesian framework the latter might be generated by different channels, like the assumption of indexation or rule-of-thumb behavior on the part of the price or wage setters. Finally Altissimo, Ehrmann, and Smets (2006) highlight that if expectations are not rational, there might be persistence due to the formation of inflation expectations ("expectations-based persistence").[2]

We want to document the slope of the Italian Phillips curve and identify the source of persistence of Italian producer price inflation over two periods: 1981–98 and 1999–2012.[3] To do so, we estimate a New Keynesian structural model of the Italian inflation that allows for both extrinsic and intrinsic persistence. To model the former, we draw on the more recent literature on New Keynesian monetary economics that has emphasized the role of the following factors: the frequency of price revisions, the economy's returns to scale, and the sensitivity of the elasticity of demand to the firm's relative price. The latter gives rise to strategic complementarities in price setting. In addition, we allow for positive trend inflation, needed given our strong empirical focus. Intrinsic persistence is instead due to partial indexation of prices, which introduces a backward-looking element in inflation determination.

There is an ample literature on the estimation of the Phillips curve. In recent years, following the seminal contribution of Galí and Gertler (1999), many papers have adopted the New Keynesian framework and, estimating a single-equation model via the GMM, have investigated the importance of intrinsic inertia to account for the observed inflation dynamics. Benigno and López-Salido (2006) and Rumler (2007) apply this technique to several euro-area countries, including Italy, finding that in a sample which covers mainly the period before 1999 in Italy there is a significant role for "backward-lookingness" in inflation, and thus for intrinsic inflation persistence. In these papers the only source of extrinsic persistence is the existence of nominal rigidities; instead, as mentioned above, we allow also for the presence of strategic complementarities in price setting, following the formulation introduced in Sbordone (2010). A different approach is implemented by Gaiotti (2010), who uses micro data to assess the role of globalization for the flattening of the Phillips curve in Italy.

[2] See Molnár and Santoro (2014) for the implications of non-rational beliefs on monetary policy.

[3] We follow most of the theoretical New Keynesian literature in concentrating on the link between real unit labor cost and producer price inflation rather than consumer price inflation.

Rather than relying on a single-equation estimation, we estimate the deep parameters that determine the slope and the persistence of the Phillips curve generated by our model using full-information Bayesian techniques. This allows us to disentangle the role played by the different factors in the relationship that links the dynamics of domestic inflation to its determinants. We use quarterly data on industrial producer price inflation,[4] GDP, short-term interest rate, and real wages over two distinct sub-samples: 1981:Q1–1998:Q4 and 1999:Q1–2012:Q3. The rationale for splitting the sample in connection with the start of the Economic and Monetary Union is the need to account for the above-mentioned importance of monetary policy conduct in shaping the Phillips curve.

In the empirical implementation, we adopt a measure of unit labor cost as the relevant determinant of producer price changes, instead of an ad hoc output gap. In the class of model we consider, this choice is innocuous, since there is an approximate log-linear relationship between the two variables; moreover, we avoid the pitfall of using measures of the output gap constructed with some type of filters that are not consistent with the theoretical structure of the model.[5]

We find significant changes in the Phillips curve in the more recent years. More in detail, our model implies a Phillips curve of the following form: $\pi_t^p = \lambda_1 \pi_{t-1}^p + \lambda_2 \mathbb{E}_t \left[\pi_{t+1}^p\right] + \lambda_3 mc_t + \lambda_4 \mathbb{E}_t \phi_{t+1} + \lambda_5 \mathbb{E}_t \varphi_{t+1}$, where π_t^p denotes price inflation, mc_t is real unit labor cost, and ϕ_{t+1} and φ_{t+1} are auxiliary variables. We get that the slope of the Phillips curve (namely, the short-term sensitivity of inflation to marginal costs, λ_3) has dropped sharply, from 0.74 to 0.05.

According to our estimates, the increased strategic complementarity in price setting, due to its turn to higher sensitivity of demand elasticity to firms' relative prices, has played an important role in determining the vanishing trade-off between producer price inflation and unit labor cost. In other words, the story suggested by our estimates is the following.

[4]Hence we abstract from the service sector.

[5]See Galí and Gertler (1999). They also show that estimating the New Keynesian Phillips curve using filtered GDP data as a proxy for the output gap delivers misleading results.

Whereas in the pre-1999 period the price elasticity of demand was almost constant, after 1999 the elasticity of substitution between a given variety and others increased in the firm's good's relative price, leading to strategic complementarities in price setting and to a flatter Phillips curve. Indeed, when a repricing firm faces a higher unit labor cost, say due to an economic boom, the firm will temper its price increase relative to the pre-1999 period because this would result in a more elastic demand curve than for those firms whose relative price declines as a result of price fixity. Symmetrically, during recessions firms are more reluctant to reduce their price relative to the pre-1999 period, as they would face a less elastic demand curve than their competitors whose relative price increases as a result of price rigidity. In other words, as Basu (2005), Dotsey and King (2005), and Klenow and Willis (2006) show, a price elasticity of demand that increases with the firm's relative price generates a smoothed version of a "kink" in the demand curve facing a given firm. Indeed, it implies that consumers flee from individual items with high relative prices, but do not flock to inexpensive ones, creating "rigidity" in the relative price a firm wants. Hence, the higher sensitivity of the demand elasticity to the firm's relative price (and the consequent increase in strategic complementarities in price setting) in the post-1999 period has weakened the short-term relationship between economic activity and price dynamics in Italy and explains the slow adjustment of inflation during a severe and long-lasting recession.[6]

Our model is not designed to capture the primitive sources driving this increase in strategic complementarities; however, in a similar framework Sbordone (2010) points out that it can be associated with an increase in market competition. This seems to be consistent with some fundamental changes in the context surrounding Italy: Brandolini and Bugamelli (2009) argue that the shift in the technological paradigm, ushered in by the new information and communications technologies (ICT); "globalization," that is, the global integration of the real and financial markets; and the process of European integration had as a common consequence a strong and sharp increase in competitive pressure.

[6]Let us recall that this is a business-cycle model; hence it is silent on the long-run relationship between producer prices and labor cost.

On top of increased strategic complementarities, we estimate an increase in the degree of price stickiness that flattens the Phillips curve. According to our results, the average duration of a price is nearly 1.7 quarters in the first sub-sample and 2.9 in the second one, consistent with the common wisdom that as inflation has decreased over time, one would expect price setters to change prices less often.

The vanishing trade-off between producer price inflation and unit labor cost also stems from the decrease in trend inflation, which we calibrate to match the decline in the average sample inflation. Importantly, whereas in a Dixit-Stiglitz world the slope of the New Keynesian Phillips curve becomes steeper under lower trend inflation (Ascari 2004), the presence of strategic complementarities inverts the sign of the derivative: the Phillips curve flattens as the trend inflation rate declines.

The degree of intrinsic persistence λ_1 has decreased in the second sub-sample (from 0.45 to 0.18), suggesting that the observed reduced-form persistence of inflation in recent years cannot be attributed to intrinsic inflation persistence but rather to an increase in extrinsic persistence due to stronger real rigidities.

A flatter Phillips curve involves a higher sacrifice ratio than otherwise, i.e., a longer spell of GDP below its natural level for every desired reduction in inflation. However, it also implies that an overheated economy will tend to induce a milder rise in inflation and, vice versa, reduces the risk of deflation in the face of severe downturns.

The paper is organized as follows. Section 2 indicates how our study relates to the literature. Section 3 presents the theoretical model. Section 4 provides a discussion of the role of nominal rigidities, strategic complementarities, backward indexation, and trend inflation for the Phillips curve implied by our theoretical framework. Section 5 briefly summarizes the Bayesian technique we employ. Section 6 describes data and prior distributions. Section 7 discusses our results and section 8 concludes.

2. Related Literature

The absence of a severe disinflation during the recent deep worldwide recession was not peculiar to Italy. Since the financial crisis of 2008–9, extreme economic slack has done little to reduce inflation in most advanced economies, insomuch that Hall (2013) reads the recent experience as a contradiction to the fundamental macroeconomic axiom that inflation declines during periods of economic weakness. In September 2010 John Williams stated, "The surprise [about inflation] is that it's fallen so little, given the depth and duration of the recent downturn. Based on the experience of past severe recessions, I would have expected inflation to fall by twice as much as it has." The missing disinflation between 2009 and 2011 is particularly striking when compared with the strong deflation that hit the United States during the Great Depression (cumulative deflation between 1930 and 1932 was nearly 25 percent). Ball and Mazumder (2011) document that when Phillips curves estimated over 1960–2007 are used to predict inflation over 2008–10, a puzzle emerges, as inflation should have fallen by much more than it did. Fostered by this evidence, a large literature has emerged to find possible tweaks to the traditional Phillips curve and explain why inflation has fallen so little in spite of the severity and duration of the recent crisis.

Ball and Mazumder (2011) show that the puzzle can be solved with two modifications of the Phillips curve, both suggested by theories of costly price adjustment: first by measuring core inflation with the median CPI inflation rate, and second by allowing the slope of the Phillips curve to change with the level and variance of inflation.[7]

According to Bernanke (2010), the credibility of modern central banks has succeeded in relegating high inflation and high deflation as very unlikely outcomes in people's conviction and, thus, in stabilizing actual inflation through expectational channels. However, while there is evidence that inflation expectations in the United States and in the euro area have indeed become more anchored over time (Williams 2006), Coibion and Gorodnichenko (2013) convincingly show that this "anchored expectations" hypothesis is not sufficient to explain the full extent of the missing disinflation in the United States between 2009 and 2011, as the latter is still present in the data even after conditioning on the "anchored" expectations of professional forecasters. Coibion and Gorodnichenko (2013) propose a novel explanation based on the idea that household forecasts are likely to be a better proxy for firm inflation expectations (i.e., the ones relevant for pricing decisions and thus

[7]The second tweak is consistent with our results.

for the Phillips curve) than either professional or backward-looking forecasts. When household inflation expectations—measured by the Michigan Survey of Consumers—are considered in an expectations-augmented Phillips curve, then the puzzle of the missing disinflation in the United States since 2009 is solved. Indeed, while households' inflation expectations rose sharply since 2009, pushed by oil price increases, other measures of inflation expectations, such as those from financial markets or professional forecasters, have stayed stable in the neighborhood of 2 percent over the same period.[8]

Studying the Swedish case, Svensson (2013) points out that if inflation expectations are irrationally anchored at the inflation target even when average inflation is systematically lower than the target, the long-run Phillips curve becomes downward sloping and the undershooting of the inflation target brings higher average unemployment than if average inflation had been kept on target.

Other authors focus on the wage formation process as the driving force behind the lack of disinflation. Among them, Daly, Hobijn, and Lucking (2012) point to downward nominal wage rigidity as an important factor that has shaped the dynamics of inflation during and after the last three U.S. recessions. In fact, the fraction of workers receiving zero wage changes in the U.S. economy increases around business-cycle downturns and has risen to historical highs since the 2007 recession. The resistance to reducing wages, especially at low levels of inflation, might have bent the short-run Phillips curve.

Finally, the interpretation favored by the International Monetary Fund's (2013) analysis is that, on the top of more firmly anchored long-term inflation expectations, the relation linking the rate of change in prices to the level of economic activity in advanced countries is considerably flatter today than in the past. By estimating

[8]Coibion and Gorodnichenko (2013) show that more than half of the historical differences in inflation forecasts between households and professionals can be accounted for by oil price dynamics, and that the run-up in oil prices since 2009 can fully explain the rise in household inflation expectations since then. The authors conjecture that households adjust their inflation forecasts more in response to oil price changes than professional forecasters because gasoline prices are among the most visible prices to consumers. Consistently, they show that according to data from the Michigan Survey of Consumers, inflation forecasts of those who spend more money on gasoline (in dollar terms) react more in response to oil price variations than expectations of those who spend less money on gasoline.

an unemployment-based Phillips curve for twenty-one advanced economies, the IMF's (2013) study concludes that the average slope of the Phillips curve has flattened as inflation rates declined and suggests, as an explanation, that the prices' adjustment costs may induce firms to vary prices less frequently when inflation is lower (Ball, Mankiw, and Romer 1988 and Klenow and Malin 2010). On the other hand, the evidence on the role of globalization in affecting the slope of the Phillips curve is either inconclusive or negative (Ball 2006; Gaiotti 2010).

3. The Model

We consider an environment of monopolistically competitive firms each producing a differentiated good. The price adjustment rule is time dependent, following the formalism proposed in Calvo (1983). With respect to a basic dynamic New Keynesian setup, the distinctive features of our theoretical framework can be summarized as follows.

We assume a trend component in technology, so that data do not need to be detrended before estimation and the dynamics of the model is evaluated with respect to a balanced growth path. To improve the empirical performance of the model, we consider positive trend inflation (on the empirical relevance of trend inflation, see Cogley and Sbordone 2008).[9] In the spirit of Galí and Gertler (1999), we augment the Calvo model in price setting by the assumption that

[9]The canonical New Keynesian sticky-price model, which has emerged as the workhorse for monetary policy analysis, hinges on inflation being a stationary variable. In recent years, there has been a great deal of research focusing on inflation persistence. Some empirical works question the hypothesis of inflation being I(0) and find that log prices are I(2) and that inflation therefore is I(1). A few of these are Johansen (1992), Bardsen, Jansen, and Nymoen (2004), O'Reilly and Whelan (2005), Fanelli (2008), and Mavroeidis, Plagborg-Møller, and Stock (2013). The empirical evidence on the stationarity of inflation is quite mixed. For instance, Zivot and Andrews (1992) and Lumsdaine and Papell (1997), among others, conclude that inflation rates are stationary. Concerning European countries, Culver and Papell (1997) apply time-series unit-root tests to thirteen OECD countries and find overwhelming evidence in favor of inflation being I(0). Here we stick to the hypothesis of inflation being a stationary variable consistent with the standard New Keynesian model and with a successful implementation of inflation targeting. However, we appropriately take into account the empirical and theoretical relevance of a positive steady inflation rate.

prices that are not reoptimized are partially indexed to past inflation rates. As stressed by Galí and Gertler (1999), the estimation of the degree of indexation allows to measure the residual inertia that the forward-looking Phillips curve leaves unexplained.[10] We allow for variable demand elasticity (Kimball 1995) with the aim to take into account the role of strategic complementarity in shaping the relationship that links the dynamics of inflation to unit labor cost. Four structural shocks are considered in order to avoid stochastic singularity and achieve an exact identification of the model: a monetary policy shock, a technology shock, and two preference shocks, one intertemporal and one intratemporal.[11]

3.1 Preferences

We consider a continuum of households uniformly distributed on the unit interval. The household is the relevant decision unit and has an objective function given by

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \varsigma_t \beta^t \left\{ U\left(C_t\left(j\right)\right) - \chi_t V\left(N_t\left(j\right)\right)\right\}, \tag{1}$$

where $C_t(j)$ denotes household's j consumption; $N_t\left(j\right) \in [0,1]$ indicates the fraction of household members who are employed; and ς_t and χ_t are two preference shocks which are assumed to follow stationary first-order autoregressive (AR(1)) processes $\varsigma_t = \varsigma_{t-1}^{\rho_\varsigma} e^{\varepsilon_t^{\varsigma_t}}$ and $\chi_t = \chi_{t-1}^{\rho_\chi} e^{\varepsilon_t^{\chi}}$, respectively.

We specify the household's period utility to be given by

$$U\left(C_t\left(j\right)\right) \equiv \log C_t\left(j\right) \tag{2}$$

$$V\left(N_t\left(j\right)\right) \equiv \frac{N_t^{1+\phi}(j)}{1+\phi}. \tag{3}$$

The period budget constraint takes the form

$$P_t C_t\left(j\right) + Q_t B_t\left(j\right) = B_{t-1}\left(j\right) + W_t N_t(j) + \Pi_t,$$

[10] Galí and Gertler (1999) suggest that adding rule-of-thumb price setters allows to measure the departure from the baseline forward-looking model in the spirit of the way that Campbell and Mankiw (1989) use rule-of-thumb consumers to test the life-cycle/permanent-income hypothesis.

[11] See below for more details.

where P_t is the price index for goods; W_t is the nominal wage; Π_t are profits; Q_t is the price of a one-period riskless bond, paying one unit of currency; and B_t denotes the quantity of that bond purchased in period t.

We assume a flexible variety aggregator à la Kimball (1995), which allows for the possibility that firms face price elasticity of demand which is increasing in firms' relative prices. Whereas in the Dixit-Stiglitz preferences the elasticity of substitution between a given variety and others is constant, in Kimball's world the elasticity of substitution between differentiated goods is decreasing in the relative quantity consumed of the variety, implying that firms' desired markup is decreasing in their relative price. The consumption aggregate is defined by

$$\int_{\Omega} \psi \left(\frac{C_t\,(i)}{C_t} \right) di = 1, \tag{4}$$

where $\psi(\cdot)$ is an increasing, strictly concave function, and Ω is the set of all potential goods produced. The standard constant elasticity of substitution (CES) preferences are nested within this specification, and the Kimball aggregator reduces to the Dixit-Stiglitz when $\psi\left(C_t\,(i)\,/C_t\right) = \left(C_t\,(i)\,/C_t\right)^{\frac{\epsilon-1}{\epsilon}}$. In order to allocate its consumption expenditures among the different goods, the household maximizes the consumption index C_t for any given level of expenditures $\int_0^1 P_t\,(i)\,C_t\,(i)\,di$. The solution to that problem yields the set of demand equations $P_t\,(i) = \frac{1}{\Lambda_t C_t}\psi'\left(\frac{C_t(i)}{C_t}\right)$ for each $i \in [0,1]$, where Λ_t is the Lagrangian multiplier for constraint (4). Accordingly, we can write the demand curve for good i as

$$C_t\,(i) = C_t \psi^{'-1}\left(\frac{P_t\,(i)}{P_t'}\right), \tag{5}$$

where $P_t' \equiv \frac{1}{\Lambda_t C_t}$ for any set of prices $\{P_t\,(i)\}$. We define the price index as the cost of a unit of the composite good:

$$P_t = \frac{1}{C_t}\int_0^1 P_t\,(i)\,C_t\,(i)\,di = \int_0^1 P_t\,(i)\,\psi'^{-1}\left(\frac{P_t\,(i)}{P_t'}\right). \tag{6}$$

Following Dotsey and King (2005) and Sbordone (2010), we adopt a functional form that implies $P_t' = \left[\int_0^1 P_t\,(i)^{1-(1+\eta)\epsilon}\,di\right]^{\frac{1}{1-(1+\eta)\epsilon}}$,

$P_t = \frac{1}{1+\eta} P_t' + \frac{\eta}{1+\eta} \int_0^1 P_t(i)\, di$. Hence, the relative demand for good i (5) is

$$\frac{C_t(i)}{C_t} = \frac{1}{1+\eta} \left[\left(\frac{P_t(i)}{P_t'} \right)^{-(1+\eta)\epsilon} + \eta \right]. \tag{7}$$

Clearly the Dixit-Stiglitz preferences are nested within this specification for $\eta = 0$.

The optimal consumption/savings and labor supply decisions are described by the following conditions:

$$Q_t = \beta \mathbb{E}_t \frac{\varsigma_{t+1}}{\varsigma_t} \frac{C_t}{C_{t+1}} \frac{P_t}{P_{t+1}} \tag{8}$$

$$\frac{W_t}{P_t} = \chi_t N_t^{\phi} C_t. \tag{9}$$

3.2 Technology

There is a continuum of firms distributed uniformly on the unit interval. Each firm is indexed by $i \in [0, 1]$ and produces a differentiated good with a technology

$$Y_t(i) = Z_t N_t^{1-\alpha}(i). \tag{10}$$

Z_t is an aggregate technology index which follows the trend-stationary process

$$Z_t = \gamma^t A_t, \tag{11}$$

where γ is the deterministic growth rate of the economy and A_t is a stationary AR(1) process, i.e., $A_t = A_{t-1}^{\rho_A} e^{\varepsilon_t^A}$.

3.3 Price Setting

Each firm may reset its price only with probability $(1-\theta)$, independently of the time elapsed since the last adjustment (Calvo 1983). With probability θ the firm adjusts its price mechanically according to $P_t(i) = (\gamma_{p,t-1})^{\tau_p} P_{t-1}(i)$, where $\gamma_{p,t-1} \equiv \left(\frac{P_{t-1}}{P_{t-2}}\right)$

denotes last-period inflation. We denote with $\overline{\Pi}$ steady-state inflation.[12]

A firm reoptimizing in period t will choose the price P_t^* by solving the following maximization problem:

$$\max_{P_t} \sum_{k=0}^{\infty} \theta^k \mathbb{E}_t \left\{ \Lambda_{t,t+k} \left(P_t^* \left(\frac{P_{t+k-1}}{P_{t-1}} \right)^{\tau_p} Y_{t+k/t} - \Psi_{t+k} \left(Y_{t+k/t} \right) \right) \right\} \tag{12}$$

subject to the sequence of demand constraints (7),

$$Y_{t+k/t} = \frac{1}{1+\eta} \left[\left(\frac{P_t^* \left(\frac{P_{t+k-1}}{P_{t-1}} \right)^{\tau_p}}{P'_{t+k}} \right)^{-(1+\eta)\epsilon} + \eta \right] Y_{t+k}. \tag{13}$$

Combining the first-order conditions for this problem with the aggregate price index and log-linearizing, we get the following system of equations that characterizes the generalized Phillips curve under trend inflation, strategic complementarities, and backward indexation:

$$\begin{aligned}
\pi_t &= \frac{\tau_p (1 + kDD)}{1 + kDD + \widetilde{\beta}\tau_p} \pi_{t-1} + \frac{\widetilde{\beta}}{1 + kDD + \widetilde{\beta}\tau_p} \mathbb{E}_t (\pi_{t+1}) \\
&\quad + \frac{k}{1 + kDD + \widetilde{\beta}\tau_p} \left(\widetilde{w_t^r} - \widetilde{mpn}_t \right) + \frac{\omega_1}{1 + kDD + \widetilde{\beta}\tau_p} \mathbb{E}_t \widehat{\phi}_{t+1} \\
&\quad + \frac{\omega_2}{1 + kDD + \widetilde{\beta}\tau_p} \mathbb{E}_t \widehat{\varphi}_{t+1} \\
\widehat{\phi}_t &= \left[1 - \beta\theta \Pi^{(1+\eta)\epsilon(1-\tau_p)} \right] \left\{ \left(\widetilde{w_t^r} - \widetilde{mpn}_t \right) + \left[(1+\eta)\epsilon BB - DD \right] \Delta_t \right\} \\
&\quad + \beta\theta \overline{\Pi}^{(1+\eta)\epsilon(1-\tau_p)} \mathbb{E}_t \left[\widehat{\phi}_{t+1} + (1+\eta)\epsilon \Delta_{t+1} \right] \\
\widehat{\varphi}_t &= \beta\theta \overline{\Pi}^{-(1-\tau_p)} \mathbb{E}_t \left(\widehat{\varphi}_{t+1} - \Delta_{t+1} \right),
\end{aligned} \tag{14}$$

[12] If one imposed full price indexation to past inflation, the theoretical relevance of a positive average inflation rate would be muted, as in this extreme case there would not be price dispersion in the steady state (see Ascari and Sbordone 2013).

where π_t denotes inflation, $\widetilde{w_t^r}$ the detrended real wage, and $\widetilde{mpn}_t$ the marginal product of labor, log-linearized around the stationarized steady state. Also we have denoted $\Delta_t \equiv \pi_t - \tau_p \pi_{t-1}$. $\widehat{\phi}_t$ and $\widehat{\varphi}_t$ are auxiliary variables. The coefficients are convolutions of parameters, inter alia, trend inflation $\overline{\Pi}$, price indexation τ_p, and the degree of strategic complementarities η.[13] See the appendix for details.

3.4 Monetary Policy

The model is closed by assuming the following forward-looking rule linking the interest rate to economic activity and inflation:

$$\frac{R_t}{\overline{R}} = \left(\frac{R_{t-1}}{\overline{R}}\right)^{\rho_r} \left[\left(\frac{\mathbb{E}_t\left(\Pi_{t+1}\right)}{\overline{\Pi}}\right)^{\phi_\pi} \left(\widetilde{Y}_t\right)^{\phi_y}\right]^{1-\rho_r} \mu_t, \qquad (15)$$

where $\overline{R}$ is the steady-state nominal gross rate, $\widetilde{Y}_t$ is the cyclical component of GDP obtained by detrending Y_t with the level of technology γ^t, and μ_t is a monetary policy shock, which is assumed to follow the stationary $AR(1)$ process $\mu_t = \mu_{t-1}^{\rho_\mu} e^{\varepsilon_t^\mu}$.

A caveat is required. Starting from 1999, monetary policy has been set uniformly by the European Central Bank (ECB) for the euro area as a whole. Equation (15) should be read as a positive, rather than normative, relationship capturing the correlation of the interest rate to the country-specific economic conditions.

4. The Phillips Curve

The Phillips curve implied by our theoretical model can be written as

$$\pi_t = \lambda_1 \pi_{t-1} + \lambda_2 \mathbb{E}_t\left[\pi_{t+1}\right] + \lambda_3\left(\widetilde{w_t^r} - \widetilde{mpn}_t\right) + \lambda_4 \mathbb{E}_t \phi_{t+1} + \lambda_5 \mathbb{E}_t \varphi_{t+1}.$$

The dynamics of inflation depends on three major driving forces: past realized values of inflation, the expectations of future inflation, and the real unit labor cost. The latter affects price dynamics through several channels.

[13] Our Phillips curve is an extension of the one derived in Ascari and Ropele (2007, 2009) to the case in which the Dixit-Stiglitz constant elasticity does not necessarily hold.

- *Nominal rigidities.* The longer prices are kept unchanged, the flatter the Phillips curve. Indeed, the smaller the frequency of price changes, the more nominal disturbances translate into real effects rather than aggregate inflation.
- *The degree of indexation to past inflation.* The larger the degree of indexation to past inflation, the flatter the Phillips curve. The reason is that a smaller fraction of inflation dynamics is due to movements in real unit labor cost, since past inflation exerts a mechanical pressure on contemporaneous inflation.
- *The degree of trend inflation.* One established result in the macroeconomics of trend inflation (see Ascari and Sbordone 2013, among others) is that the Phillips curve becomes steeper under lower steady inflation. Such a result ceases to hold when one leaves the Dixit-Stiglitz world (see figure 1): if the elasticity of substitution between a given variety and others is increasing in the firm's good's relative price, the slope of the curve flattens when trend inflation falls, consistently with most economists' priors (see Shirota 2007).
- *The sensitivity of the marginal costs to the firm's level of production.* The higher the sensitivity of the marginal costs to the level of production, the flatter the Phillips curve. The intuition is as follows. When a firm faces decreasing returns to scale, the marginal cost is increasing in its own output. Hence, in the face of an increase in $\left(\widetilde{w_t^r} - mpn_t\right)$, the desired price increase is smaller because the firm takes into account the decline in marginal cost due to the loss in demand associated with the price increase.
- *The steady-state sensitivity of the firm's own output demand to its relative price.* The higher the steady-state elasticity of demand, the larger the loss in demand incurred for a price increase, the smaller the desired price increase for any given rise in $\left(\widetilde{w_t^r} - mpn_t\right)$, and the flatter the Phillips curve.
- *The sensitivity of the elasticity of substitution between differentiated goods to the relative quantity consumed of the variety.* When the elasticity of substitution between differentiated goods is decreasing in the relative quantity consumed of the variety, firms face a price elasticity of demand that is

> increasing in their good's relative price. This makes the desired markup decreasing in firms' relative price. The larger the sensitivity of the elasticity of substitution to the relative price (the more negative is η), the flatter the Phillips curve. Indeed, when the elasticity of demand is increasing in the relative price, firms are reluctant to change their price, as they would face a more elastic demand curve than firms whose relative price declines as a result of price fixity.

Following the literature (Woodford 2003; Sbordone 2010), we call the last three components "strategic complementarity" channels.

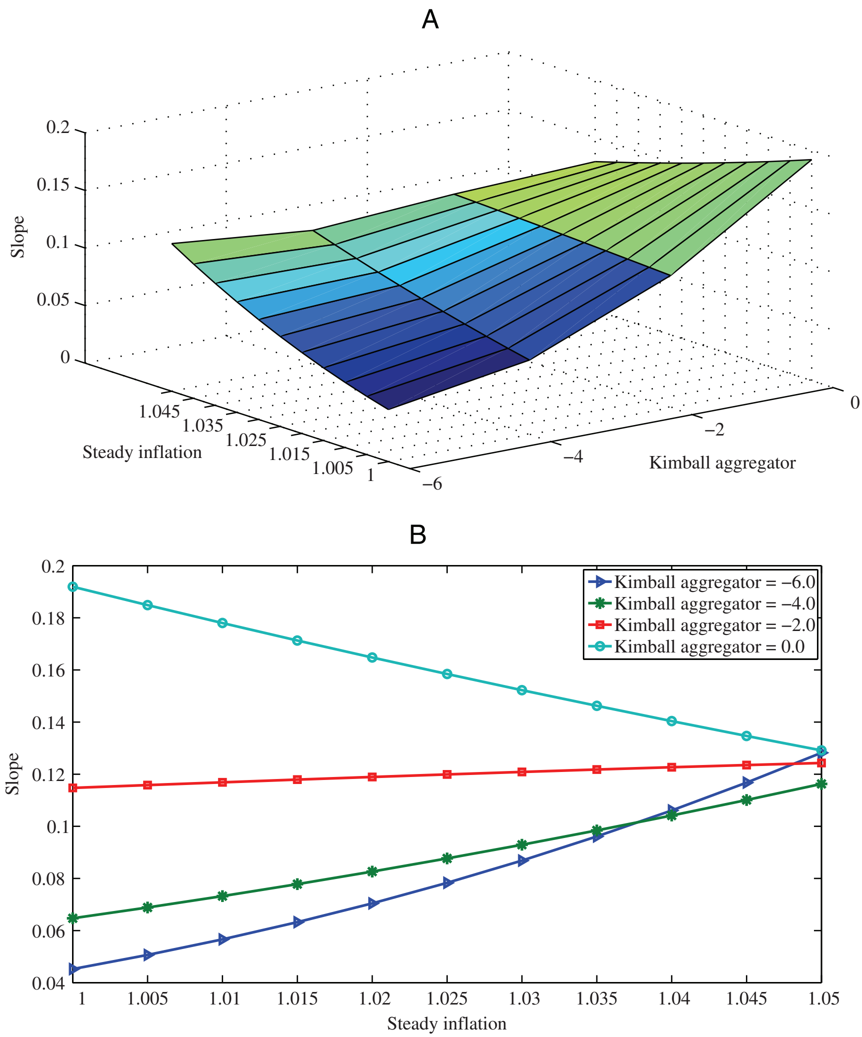
Figures 1, 2, 3, and 4 show the slope of the Phillips curve under different combinations of η and $\overline{\Pi}$, η and θ, τ_p and θ, and $(1-\alpha)$ and θ.

In order to evaluate the role of nominal rigidities and strategic complementarity in determining the slope of the Phillips curve, we implement a Bayesian Markov chain Monte Carlo (MCMC) estimation procedure in the spirit of Schorfheide (2000), Smets and Wouters (2003, 2007), and Fernandez-Villaverde and Rubio Ramirez (2004).

5. Bayesian Estimation

When dealing with the estimation of dynamic general equilibrium models, the performances of the full-information maximum-likelihood (FIML) estimator may be significantly compromised. Indeed, the mapping of structural parameters to the coefficients of the reduced form of the model is highly non-linear and non-identification is frequent, as different sets of parameters may yield nearly the same value for the likelihood function (Canova and Sala 2009). A viable solution would be to use a constrained FIML estimator that restricts the estimates within a reasonable range. However, recent developments rely on Bayesian estimation, where restrictions (priors) are defined in terms of probability distributions and the posterior distribution for the model parameters is obtained by nesting the formalized prior distribution for the vector parameters $\xi \in \Xi$ and the likelihood of the data. The posterior distribution can be read as a weighted average of prior non-sample information and the

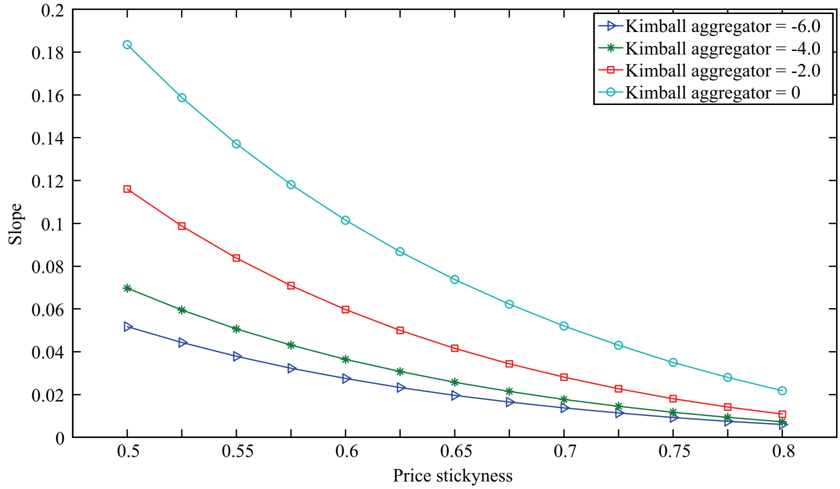
Figure 1. Trend Inflation, Kimball Aggregator, and the Slope of the Phillips Curve



Notes: Panel A: On the x-axis η. On the y-axis $\overline{\Pi}$. Panel B: On the x-axis $\overline{\Pi}$. Each curve is drawn under a different calibration for η.

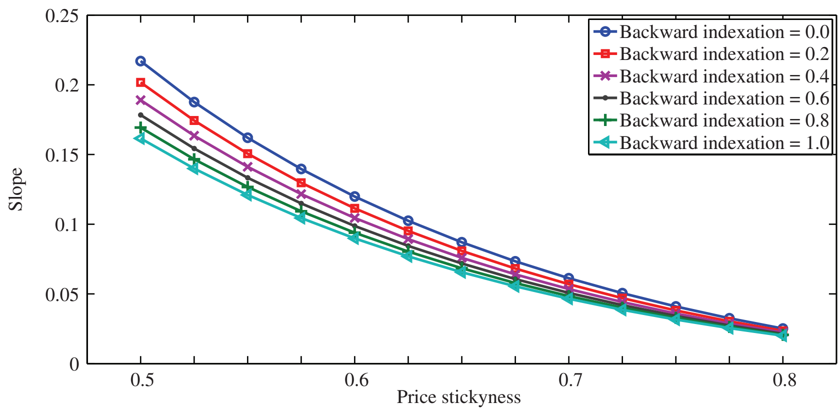
conditional distribution, where the weights are inversely related to the variance of the prior distributions and the variance of the sample information, respectively.

Figure 2. Price Stickiness, Kimball Aggregator, and the Slope of the Phillips Curve



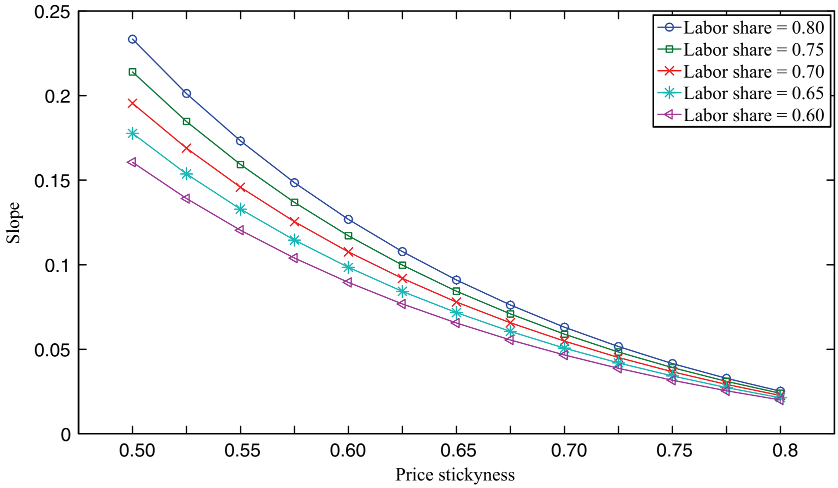
Notes: On the x-axis θ. Each curve is drawn under a different calibration for η.

Figure 3. Price Stickiness, Backward Indexation, and the Slope of the Phillips Curve



Notes: On the x-axis θ. Each curve is drawn under a different calibration for τ_p.

Figure 4. Price Stickiness, Returns to Scale, and the Slope of the Phillips Curve



Notes: On the x-axis θ. Each curve is drawn under a different calibration for $(1-\alpha)$.

More in detail, let $P(\xi, M)$ denote the prior beliefs on parameters ξ given model M and let $P(X_T/\xi, M)$, $X_T = \{x_t\}_{t=1}^{T}$ denote the conditional distribution (likelihood). By using the Bayes rule, the posterior density $P(\xi/X_T, M)$ can be written as

$$P(\xi/X_T, M) = \frac{P(X_T/\xi, M)P(\xi, M)}{P(X_T, M)}. \tag{16}$$

Following the literature, we get the Bayesian posterior estimates by using the Kalman filter to form the likelihood function and the Metropolis-Hastings algorithm for Monte Carlo integration (three chains of 1,000,000 draws) to optimize the posterior density function. We formally check whether estimation is possible or whether there are serious identification issues by verifying the necessary and sufficient conditions for local identification discussed by Iskrev (2010b) for the parameters declared in the prior definition at the prior mean; see the appendix for details.

6. Data and Prior Distribution

Our model is estimated by using Italian quarterly data over two sample periods: 1981:Q2–1998:Q4 and 1999:Q1–2012:Q2. We consider four time series: the log-differences of real GDP and of the real wage,[14] the short-term interest rate, and the percent change in the producer price index as a measure of inflation. As argued before, the rationale for splitting the sample in connection with the start of the Economic and Monetary Union is the need to account for the importance of monetary policy conduct in shaping the Phillips curve. In order to verify the presence of a structural break in the correlation among our observed series, we apply a Chow test to the VAR representation of inflation with different lags, and the test always rejects the null hypothesis that there is no structural break in 1999:Q1 at the significance level of 1 percent.

In order to enhance estimation of the key structural parameters, we impose the following dogmatic priors. The discount factor β is fixed at 0.99 and the Frisch elasticity on labor supply $1/\phi$ is set at 1, a standard calibration for macroeconomic models. The "share" parameter on labor in the production function $(1-\alpha)$ is fixed at 2/3 (consistent with computation in Giordano and Zollino 2013). We set $\epsilon = 6$, consistent with Dotsey and King (2005).[15]

Our sample data feature an average quarterly GDP growth rate of almost 0.5 percent in the pre-1999 period and 0.2 percent in the post-1999 period. We use this information to calibrate the steady-state value of productivity growth γ. We calibrate steady inflation to match the observed decline in the average annual sample producer price inflation, which is nearly 4.9 percent in the first sample and 2.4 percent in the second sample. All remaining parameters are estimated.

[14]The real wage is obtained by deflating earned income in manufacturing with the consumer price index.

[15]As discussed in the appendix, in order to check formally whether estimation is possible or whether there are serious identification issues, we verify the necessary and sufficient conditions for local identification discussed by Iskrev (2010b) for the parameters declared in the prior definition at the prior mean. When we include ϵ in the set of the parameters to be estimated, we get that ϵ and η are collinear. We thus calibrate ϵ consistently with the literature and estimate the extent to which the model departs from the Dixit-Stiglitz constant elasticity by allowing data to choose the value of η.

Considering the shape of the prior distributions, following standard practice, a beta distribution is adopted for parameters theoretically defined in a $[0-1]$ range, whereas a normal distribution is assumed for priors on parameters theoretically defined over the $\mathbb{R}$ range. As for the structural shocks, the reference distribution is the inverted gamma which is defined over the $\mathbb{R}^+$ range.

The prior mean for the Calvo parameter θ is set at 0.5 (implying that the average duration of a price is two quarters), with a standard deviation of 0.1. This relatively weak prior is based on Fabiani et al. 2010. For the sensitivity of firm demand elasticity to market share (the Kimball aggregator) η, we assume a very uniformative prior centered on zero with a standard error of 5. In other words, our prior is consistent with the hypothesis of a standard Dixit-Stiglitz constant demand elasticity and we let the estimates determine whether the data favor a Kimball aggregator characterized by a variable demand elasticity. The prior for the degree of backward indexation is assumed to be centered on 0.5 with a standard error of 0.1.

The parameters describing monetary policy are based on a standard Taylor rule: the prior mean of the Taylor coefficient on inflation ϕ_π and real activity ϕ_y is set at 1.5 and 0.5, with a standard deviation of 0.2 and 0.1, respectively, so as to guarantee a unique solution path when solving the model. The autoregressive coefficient ρ_r, capturing interest rate smoothness, has a diffuse prior mean of 0.5 with a prior standard error of 0.2.

The priors on the stochastic processes are harmonized and weakly informative, reflecting the very imprecise opinion about the dimensionality and the persistence of shocks (see table 2). The standard errors of the innovations have a prior mean of 0.01 with two degrees of freedom. All shocks are assumed to be serially correlated with autoregressive coefficient having a prior mean of 0.5 and a prior standard deviation of 0.2.

The measurement equations that link data to the corresponding model variables are described in the appendix.

7. Results

Tables 1 and 2 show, together with the prior distribution, the posterior mean and the 5th and 95th percentile of the posterior distribution of the parameters obtained through the Metropolis-Hastings

Table 1. Prior and Posterior Distribution of Structural Parameters

	Prior Distribution		Posterior Distribution 1981:Q2–1998:Q4	Posterior Distribution 1999:Q1–2012:Q2
	Distr.	**Mean (St. Dev.)**	**Mean (5%; 95%)**	**Mean (5%; 95%)**
θ Price Stickiness	$\mathcal{B}$	0.5 (0.1)	0.42 (0.26; 0.55)	0.66 (0.59; 0.72)
τ_p Degree of Backward Indexation	$\mathcal{B}$	0.5 (0.1)	0.45 (0.29; 0.61)	0.22 (0.12; 0.32)
η Kimball Aggregator	$\mathcal{N}$	0.0 (5.0)	0.85 (0.66; 0.95)	−2.00 (−2.64; −1.34)
ρ_r Monetary Policy Inertia	$\mathcal{B}$	0.5 (0.2)	0.18 (0.06; 0.30)	0.83 (0.79; 0.88)
ϕ_π Taylor Coefficient on Inflation	$\mathcal{N}$	1.5 (0.2)	1.82 (1.55; 2.13)	1.74 (1.46; 2.02)
ϕ_Y Taylor Coefficient on Real Activity	$\mathcal{N}$	0.5 (0.1)	0.03 (−0.03; 0.11)	0.60 (0.44; 0.75)

Notes: The posterior distribution is obtained using the Metropolis-Hastings algorithm.

Table 2. Prior and Posterior Distribution of Shock Processes

	Prior Distribution		**Posterior Distribution 1981:Q2–1998:Q4**	**Posterior Distribution 1999:Q1–2012:Q2**
	Distr.	**Mean (St. Dev.)**	**Mean (5%; 95%)**	**Mean (5%; 95%)**
Autoregressive Parameters				
ρ_μ	$\mathcal{B}$	0.50	0.06	0.75
Monetary Policy Shock		(0.20)	(0.01; 0.10)	(0.70; 0.81)
ρ_ς	$\mathcal{B}$	0.50	0.90	0.53
Intertemporal Shock		(0.20)	(0.84; 0.98)	(0.39; 0.66)
ρ_χ	$\mathcal{B}$	050	0.97	0.53
Intratemporal Shock		(0.20)	(0.95; 0.99)	(0.42; 0.65)
ρ_A	$\mathcal{B}$	0.50	0.99	0.95
Technology Shock		(0.20)	(0.98; 0.99)	(0.91; 0.99)
Standard Errors of Innovations				
σ_μ	$\mathcal{IG}$	0.01	0.003	0.004
Monetary Policy Shock		(2)*	(0.002; 0.004)	(0.002; 0.005)
σ_ς	$\mathcal{IG}$	0.01	0.016	0.044
Intertemporal Shock		(2)*	(0.005; 0.038)	(0.035; 0.053)
σ_χ	$\mathcal{IG}$	0.01	0.028	0.029
Intratemporal Shock		(2)*	(0.024; 0.032)	(0.024; 0.034)
σ_A	$\mathcal{IG}$	0.01	0.015	0.003
Technology Shock		(2)*	(0.013; 0.017)	(0.002; 0.004)

Notes: The posterior distribution is obtained using the Metropolis-Hastings algorithm. *For the inverted gamma distributions, degrees of freedom are indicated.

sampling algorithm, based on three chains of 1,000,000 draws.[16] Acceptance rates over the three chains are [0.236, 0.234, 0.228] for the first sample and [0.255, 0.254, 0.254] for the second one, close to the optimal acceptance rate of 0.234 (see Roberts, Gelman, and Gilks 1997). Markov chain Monte Carlo convergence diagnostics is discussed in the appendix.

The following conclusions can be drawn. Our estimates show that in the more recent years, the degree of intrinsic persistence (λ_1) has decreased, from 0.45 to 0.18, suggesting that the observed reduced-form inertia of inflation in recent years cannot be attributed to intrinsic inflation persistence but rather to an increase in extrinsic persistence. In particular, we document a vanishing trade-off between inflation and economic activity. Indeed, the slope of the Phillips curve (λ_3) was 0.74 in the pre-1999 period and has reduced to 0.05 in the post-1999 period, which is in line with the estimates reported in Massidda and Mattana (2010). According to our estimates, such a flattening has been driven by increased strategic complementarity in price setting, on top of lower steady inflation and a longer average duration of prices, from 1.7 to 2.9 quarters.

Whereas data in the pre-1999 period favor the hypothesis of a constant demand elasticity, our estimates for the post-1999 period signal a much less convex demand curve compared with the Dixit-Stiglitz case.[17] When $\eta < 0$, the elasticity of demand faced by the firm depends inversely on its relative market share, hence the desired markup over marginal cost is decreasing in the firm's relative price. This flattens the Phillips curve, as for any given rise in unit labor cost, the firm will temper its price increase because of the endogenous drop in its desired markup. Indeed, for any given price increase, the firm will face a more elastic demand curve than for firms whose relative price declines as a result of prices remaining

[16]Note that in both of the sub-periods the estimated posterior means of the parameters are consistent with the condition $\beta\theta\overline{\Pi}^{(1+\eta)\epsilon\left(1-\tau_p\right)} < 1$. This condition is needed to ensure that in the deterministic steady state the infinite sums in the optimal pricing equation of the firms converge; see Ascari and Ropele (2009) for a similar condition in a setting with CES demand.

[17]The larger the η in absolute value, the more concave will be the demand curve.

fixed. Symmetrically, during recessions the firm will temper its price decrease because of the endogenous increase in its desired markup. Indeed, firms are more reluctant to reduce their price, as they would face a less elastic demand curve than those competitors whose relative price increases as a result of price fixity.

In other words, Basu (2005), Dotsey and King (2005), and Klenow and Willis (2006) show that a price elasticity of demand that is increasing in the firm's relative price leads to a smoothed version of a "kink" in the demand curve facing a given firm because it implies that consumers flee from individual items with high relative prices, but do not flock to inexpensive ones, thus creating "rigidity" in the relative price a firm wants. Hence, the higher sensitivity of the demand elasticity to the firm's relative price in the post-1999 period has weakened the relationship between economic activity and price dynamics in Italy and can account for the inertial adjustment of inflation. This change in real rigidities might be due to the increase in competitive pressure documented in Brandolini and Bugamelli (2009), as a consequence of the ICT revolution, globalization, and the process of European integration.[18] However, our model does not allow to capture the primitive sources driving this structural change, as it is not designed to disentangle the role of a change in preferences from that of a change in competition.

We have calibrated steady-state inflation to match the observed decline in the average annual sample producer price inflation, from 4.9 percent in the first sample to 2.4 percent in the second one. Whereas in a Dixit-Stiglitz world such a reduction would result in a steeper Phillips curve, when the elasticity of substitution between a given variety and others is increasing in the firm's good's relative price, the slope of the curve flattens when trend inflation falls. Our finding that the average duration of prices has increased by almost one quarter is consistent with the evidence documented using micro data that inflation correlates positively with the frequency of price adjustments: e.g., see Nakamura and Steinsson (2008) for results obtained using U.S. data, and Wulfsberg (2009) for an investigation carried out with Norwegian data. Also, our result is consistent

[18]See Sbordone (2010) for an extensive theoretical treatise of the effect of increased market competition on the degree of strategic complementarities in price setting and the slope of the Phillips curve in a framework similar to ours.

with the idea suggested by Ball, Mankiw, and Romer (1988) that the flattening of the Phillips curve at low levels of inflation might reflect the fact that there are costs associated with adjusting nominal prices that lead firms to change prices less frequently when inflation is lower. Cross-country evidence documented by Klenow and Malin (2010) confirms that firms do change prices less frequently when inflation is lower.

It is worth noting that another stark difference between the two sub-samples is the first-order autocorrelation of the monetary policy shock: in the pre-1999 period it was very small (0.06), so that the shock was close to a white noise; in the post-1999 period, after the start of the EMU, it increased sharply to 0.75. This evidence should be interpreted with caution: as mentioned above, the Taylor rule after 1999 cannot be interpreted, as usual, as the policy reaction function of the central bank, since monetary policy was conducted by the ECB at the euro-area level. Hence, the high level of the shock's persistence can capture a systematic difference in economic conditions between Italy and the average of the euro area.

8. Conclusions and Future Research

Debate over the Phillips curve has gained momentum since the financial crisis began in 2007. Indeed a puzzle has emerged, as inflation in advanced countries has not fallen as much as a traditional Phillips curve and past experiences would predict, given the severity and the duration of the recession. In this paper we focus on Italy and investigate the changes that occurred in the Italian Phillips curve.

According to our estimates, after 1999 the degree of "intrinsic persistence" in Italian inflation—i.e., the dependence of inflation on its own past—has diminished. In other words, in the more recent years, the degree of backwardness in price setting needed to account for the observed inflation persistence, while statistically significant, is not quantitatively important.

The slow response of inflation to economic activity observed in Italy in recent years may then be accounted for by the flattening of the Phillips curve. The latter is mainly due to the increase in strategic complementarities in price setting, on top of lower trend inflation and higher average duration of prices. The story that emerges from our estimates is as follows.

Whereas in the pre-1999 period the price elasticity of demand was almost constant, after 1999 the elasticity of substitution between a given variety and others is increasing in the firm's good's relative price, leading to strategic complementarities in price setting and to a flatter Phillips curve. Indeed, in a boom the firm will temper its price increase relative to the pre-1999 period because this will result in a more elastic demand curve than for those firms whose relative price declines as a result of price fixity. During recessions firms are more reluctant to cut their price relative to the past, as they would face a less elastic demand curve than their competitors whose relative price increases as a result of price rigidity.

In other words, as Basu (2005), Dotsey and King (2005), and Klenow and Willis (2006) show, a price elasticity of demand that is increasing with the firm's relative price generates a smoothed version of a "kink" in the demand curve facing a given firm. Indeed, it implies that consumers flee from individual items with high relative prices, but do not flock to inexpensive ones, creating "rigidity" in the relative price a firm wants. Hence, the higher sensitivity of the demand elasticity to the firm's relative price in the post-1999 period has weakened the relationship between economic activity and price dynamics in Italy and explains the slow adjustment of inflation during a severe and long-lasting recession.

A flatter Phillips curve involves a higher sacrifice ratio than otherwise, i.e., a longer spell of GDP below its natural level for every desired reduction in inflation. However, it also implies that an overheated economy will tend to induce a milder rise in inflation and, vice versa, reduces the risk of deflation in the face of severe downturns.

We see as fruitful directions for future research an investigation of the deep forces behind the increase in strategic complementarities we found in the estimation. Moreover, a cross-country analysis could shed light on the determinants underlying the differences between Italian and euro-area inflation dynamics.

Appendix

Phillips-Curve Coefficients

The generalized Phillips curve under trend inflation, strategic complementarities, and backward indexation can be described by the following system of equations:

$$
\begin{aligned}
\pi_t = {} & \frac{\tau_p \left(1 + kDD\right)}{1 + kDD + \widetilde{\beta}\tau_p}\pi_{t-1} + \frac{\widetilde{\beta}}{1 + kDD + \widetilde{\beta}\tau_p}\mathbb{E}_t\left(\pi_{t+1}\right) \\
& + \frac{k}{1 + kDD + \widetilde{\beta}\tau_p}\left(\widetilde{w_t^r} - \widetilde{mpn}_t\right) + \frac{\omega_1}{1 + kDD + \widetilde{\beta}\tau_p}\mathbb{E}_t\widehat{\phi}_{t+1} \\
& + \frac{\omega_2}{1 + kDD + \widetilde{\beta}\tau_p}\mathbb{E}_t\widehat{\varphi}_{t+1} \\
\widehat{\phi}_t = {} & \left[1 - \beta\theta\Pi^{(1+\eta)\epsilon(1-\tau_p)}\right]\left\{\left(\widetilde{w_t^r} - \widetilde{mpn}_t\right) + \left[\left(1+\eta\right)\epsilon BB - DD\right]\Delta_t\right\} \\
& + \beta\theta\overline{\Pi}^{(1+\eta)\epsilon(1-\tau_p)}\mathbb{E}_t\left[\widehat{\phi}_{t+1} + \left(1+\eta\right)\epsilon\Delta_{t+1}\right] \\
\widehat{\varphi}_t = {} & \beta\theta\overline{\Pi}^{-(1-\tau_p)}\mathbb{E}_t\left(\widehat{\varphi}_{t+1} - \Delta_{t+1}\right).
\end{aligned}
\tag{17}
$$

The coefficients $\widetilde{\beta}$, k, ω_1, and ω_2 are the following convolutions of deep parameters:

$$
\widetilde{\beta} \equiv \frac{\beta\theta}{CC}\left\{\begin{array}{c} \overline{\Pi}^{(1+\eta)\epsilon(1-\tau_p)}\left(1+\eta\right)\epsilon + \overline{\Pi}^{[(1+\eta)\epsilon-1](1-\tau_p)} \\ \times\left[\eta_{\varphi'}\left(\left(1+\eta\right)\epsilon - 1\right) + AA\left(1-\eta_P\right)\right] - \eta_{\varphi}\overline{\Pi}^{-(1-\tau_p)} \end{array}\right\}
\tag{18}
$$

$$
k \equiv \frac{\overline{\widetilde{P}}'^{(1+\eta)\epsilon}\,\overline{\psi}^{R}}{\overline{\phi}}\frac{1}{CC}
\tag{19}
$$

$$
\omega_1 = \frac{\beta\theta}{CC}\overline{\Pi}^{[(1+\eta)\epsilon-1](1-\tau_p)}\left\{\overline{\Pi}^{(1-\tau_p)} - 1\right\}
\tag{20}
$$

$$
\omega_2 = \frac{\beta\theta}{CC}\eta_{\varphi}\overline{\Pi}^{[(1+\eta)\epsilon-1](1-\tau_p)}\left\{\overline{\Pi}^{-(1+\eta)\epsilon(1-\tau_p)} - 1\right\},
\tag{21}
$$

where we have defined

$$AA \equiv \frac{\theta\left\{\left[(1-\theta)\,\widetilde{\overline{P}}^{*1-(1+\eta)\epsilon}+\theta\left(\Pi^{\tau_p-1}\right)^{1-(1+\eta)\epsilon}\right]^{\frac{1}{1-(1+\eta)\epsilon}-1}\times\left(\Pi^{\tau_p-1}\right)^{1-(1+\eta)\epsilon}+\eta\Pi^{\tau_p-1}\right\}}{(1-\theta)\left\{\left[(1-\theta)\,\widetilde{\overline{P}}^{*1-(1+\eta)\epsilon}+\theta\left(\Pi^{\tau_p-1}\right)^{1-(1+\eta)\epsilon}\right]^{\frac{1}{1-(1+\eta)\epsilon}-1}\times\widetilde{\overline{P}}^{*1-(1+\eta)\epsilon}+\eta\widetilde{\overline{P}}^{*}\right\}} \tag{22}$$

$$\begin{aligned} BB \equiv\ & (1-\theta)\left[(1-\theta)\,\widetilde{\overline{P}}^{*1-(1+\eta)\epsilon}+\theta\left(\Pi^{[(\tau_p-1)(1-(1+\eta)\epsilon)]}\right)\right]^{-1} \\ & \times\widetilde{\overline{P}}^{*[1-(1+\eta)\epsilon]}AA-\theta\left[(1-\theta)\,\widetilde{\overline{P}}^{*[1-(1+\eta)\epsilon]}\right. \\ & \left.+\theta\left(\Pi^{[(\tau_p-1)(1-(1+\eta)\epsilon)]}\right)\right]^{-1}\theta\left(\Pi^{[(\tau_p-1)(1-(1+\eta)\epsilon)]}\right) \end{aligned} \tag{23}$$

$$DD \equiv \frac{\epsilon\alpha}{1-\alpha}\left[\frac{\left(\frac{\widetilde{\overline{P}}}{\widetilde{\overline{P}}'}\right)^{-(1+\eta)\epsilon}}{\frac{1}{1+\eta}\left(\frac{\widetilde{\overline{P}}}{\widetilde{\overline{P}}'}\right)^{-(1+\eta)\epsilon}+\eta}\right](AA-BB)\,. \tag{24}$$

Also, we have denoted with $\widetilde{\overline{P}}^{*} \equiv \frac{P}{P}$ the relative price of the optimizing firm (which is different from 1 as long as the steady-state trend inflation is positive and $\tau_p < 1$) and with $\widetilde{\overline{P}}' = \frac{P'}{P}$ the steady-state ratio between $P_t' = \left[\int_0^1 P_t(i)^{1-(1+\eta)\epsilon}\,di\right]^{\frac{1}{1-(1+\eta)\epsilon}}$ and P_t.

The Measurement Equation

We consider four time series: the log-differences of real GDP ($\Delta \log Y_t$) and of the real wage ($\Delta \log W_t^r$) obtained by deflating the earned income in manufacturing with the consumer price index, and the log-levels of producer price inflation $\log(P_t/P_{t-1})$ and of the short-term rate evaluated on a quarterly basis $\log(1+\frac{R_t}{400})$. Because the model is expressed in log-deviations around the deterministic growth path γ, the following measurement equation relates the set of observables (on the left side) to the corresponding model variables (on the right side):

$$\begin{bmatrix} \Delta \log Y_t \\ \Delta \log W_t^r \\ \log(P_t/P_{t-1}) \\ \log(1+\frac{R_t}{400}) \end{bmatrix} = \begin{bmatrix} \widetilde{y}_t - \widetilde{y}_{t-1} + \log\gamma \\ \widetilde{w}_t^r - \widetilde{w}_{t-1}^r + \log\gamma \\ \pi_t^p + \log\overline{\Pi} \\ r_t - \log\beta + \log\gamma + \log\overline{\Pi} \end{bmatrix},$$

where $\log\overline{\Pi}$ is the average sample quarterly inflation.

An Analysis of Local Identification Base on Iskrev (2010a, 2010b)

The last decade has seen a notable development in the specification and estimation of DSGE models with Bayesian full-information methods. Pure maximum likelihood is rarely used because the mapping of DSGE models' structural parameters to the coefficients of the reduced form is highly non-linear, weak identification is frequent, and Bayesian priors help to deal with likelihood functions presenting flat surfaces in the economically reasonable portion of the parameter space. However, as Canova and Sala (2009) stress, Bayesian methods, when improperly used, may conceal identification problems when they exist (on identification of Bayesian DSGE models, see also Kleibergen and Mavroeidis 2013 and Koop, Pesaran, and Smith 2013). The latter may arise for different reasons: a structural parameter might disappear from a log-linearized solution; two structural parameters might enter the objective function only proportionally, thus being separately unrecoverable; or the curvature of the objective function might be insufficient.

These problems are hard to tackle because, in most cases, the models at stake can only be solved numerically. It is often unfeasible to derive explicitly the relationship between the deep parameters and the statistical model used to estimate them and, thus, parameters' identification can only be assessed indirectly and with the use of numerical methods. In other words, it is generally impossible to establish what is called "global identification."

The latter can be defined as in Iskrev (2010b). Let $\xi \in \Xi$ be a k-dimensional vector of deep parameters, where ξ is a point in the parameter space $\Xi \subset R^k$. Let σ_T be the vector collecting the second moments of the data on which the estimation of ξ is based. Let X_T be the observed data. Then Iskrev (2010b) gives the following definition: *Suppose that the data X_T is generated by the model with parameter vector ξ_0. Then ξ_0 is globally identified by the second moments of X_T if and only if $\sigma_T(\xi) = \sigma_T(\xi_0) \Leftrightarrow \xi = \xi_0$ for any $\xi \in \Xi$. If the previous condition is true only for values in an open neighborhood of ξ_0, the identification of ξ_0 is only local.*

While it is often not possible to establish global identification, Iskrev (2010b) develops a rank condition for local identification that applies to identification with limited- as well as full-information methods. Local identification by itself does not guarantee that a model is globally identified. Nevertheless, as Iskrev (2010b) stresses, it is relevant to establish whether a model is locally identified for two main reasons: first, local identification is sufficient for the asymptotic properties of classical estimators to hold, and, second, parameters that are globally unidentifiable everywhere in the parameter space, either because they do not appear in the likelihood function at all or because they are indistinguishable from other parameters, are also locally unidentifiable. The rank condition developed by Iskrev (2010b) reads as follows: *Suppose that σ_T is a continuously differentiable function of ξ, and let ξ_0 be a regular point of the Jacobian matrix $J(T) \equiv \frac{\partial \sigma_T}{\partial \xi'}$. Then ξ_0 is locally identifiable if and only if $J(T)$ has a full column rank at ξ_0.*

In order to check formally whether estimation is possible or whether there are serious identification issues, we verify the necessary and sufficient conditions discussed by Iskrev (2010b) for the parameters declared in the prior definition at the prior mean. To do that, we make use of the routines developed by Ratto (2011).[19] In particular, we consider the chain rule: $J(T) = \underbrace{\frac{\partial \sigma_T}{\partial \tau'}}_{J_1(T)} \underbrace{\frac{\partial \tau}{\partial \xi'}}_{J_2}$, where J_2

[19] We refer to a computational tool developed by M. Ratto (Joint Research Centre, European Commission), with the contribution of Nikolai Iskrev (see Ratto 2011). The routines are a set of algorithms for identification analysis. The Dynare package, version 4, includes this tool for identification analysis of the DSGE models (Juillard 1996).

measures the effect of perturbations in ξ on the parameters characterizing the equilibrium of the model. Finding that matrix J_2 is rank deficient at ξ means that this particular point is unidentifiable in the model. Finding that J_2 has full rank but $J(T)$ does not means that ξ cannot be identified given the set of observed variables and the number of observations.

We get that both J_2 and $J_1(T)$ have full rank: all estimated parameters are locally identified at the prior mean.

We now turn to analyze how well identified are the identifiable parameters. Indeed, even if the likelihood is not completely flat, it can exhibit very low curvature with respect to some parameters, which are said to be, in this case, weakly identified. We evaluate the strength of identification by using the measures proposed by Iskrev (2010a) and the set of routines developed by Ratto (2011) for the Dynare package.

In particular, given the asymptotic information matrix $\mathcal{I}_T(\xi) = J_2'\Sigma(m_T)J_2$, where $\Sigma(m_T)$ is the covariance matrix of simulated moments, we compute two measures of the strength of identification. The first one is

$$s_i = \sqrt{\xi_i^2/(\mathcal{I}_T(\xi)^{-1})_{(i,i)}}. \tag{25}$$

This measure can be thought of as composed of two components: "sensitivity" and "correlation": weak identification can be attributable to the fact that changing ξ_i does not change significantly the likelihood, or that the effect on the likelihood of changing ξ_i can be offset by changing other parameters (multicollinearity). The sensitivity component is defined as

$$\triangle_i = \sqrt{\xi_i^2 \mathcal{I}_T(\xi)_{(i,i)}}. \tag{26}$$

The second one is obtained by using the prior standard deviation $\sigma(\xi_i)$ to normalize the identification strength:

$$s_i^{prior} = \sigma(\xi_i) / \sqrt{(\mathcal{I}_T(\xi)^{-1})_{(i,i)}}. \tag{27}$$

In this case the "sensitivity" component reads as follows:

$$\triangle_i^{prior} = \sigma(\xi_i) \sqrt{\mathcal{I}_T(\xi)_{(i,i)}}. \tag{28}$$

Figure 5. Identification Strength with Asymptotic Information Matrix (log-scale)

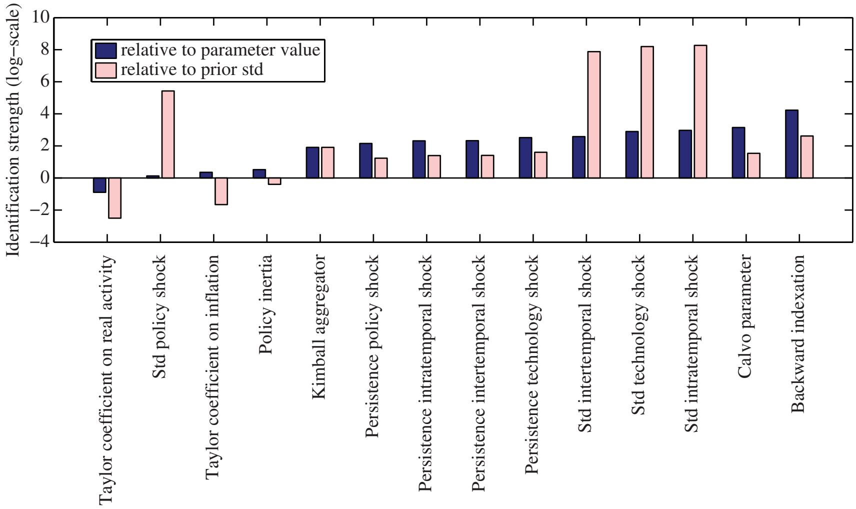
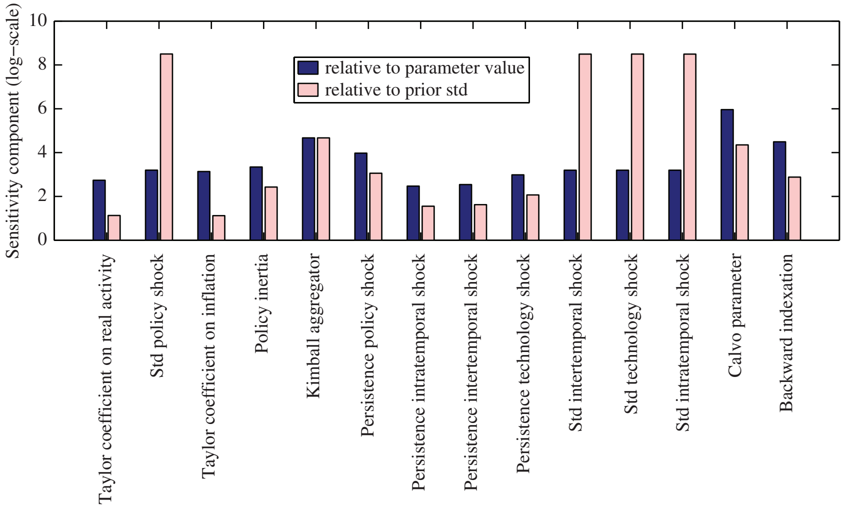


Figure 5 plots the strength of identification for all estimated parameters relative to both the parameter's value and its prior standard deviation (the model parameters on the x-axis are ranked in increasing order of strength of identification). Figure 6 shows the sensitivity component. We can conclude that all parameters are identified and have a non-negligible effect on the moments. However, identification is quite weak for the four parameters in the Taylor rule: ϕ_Y, σ_μ, ϕ_π, and ρ_r. This result is in line with an ample literature that has recently pointed out the weak identifiability of Taylor coefficients (see, for example, Cochrane 2011 and Qu and Tkachenko 2012). As expected, the comparison between figures 5 and 6 suggests that such weak identifiability depends on multicollinearity.

Markov Chain Monte Carlo Diagnostics

As specified in the main text, for each sample we run three chains of 1,000,000 Metropolis-Hastings simulations. If the results are sensible, they should be similar within any of the 1,000,000 iterations of Metropolis-Hastings simulations and close across chains. We refer to Brooks and Gelman (1998) to check for convergence.

Figure 6. Sensitivity Component with Asymptotic Information Matrix (log-scale)



More in particular, let ξ_{ij} be the i^{th} draw out of I, in the j^{th} sequence out of J. Let $\overline{\xi}_{\bullet j}$ be the mean of the j^{th} sequence and let $\overline{\xi}_{\bullet\bullet}$ be the mean across all available data. We define with $\widehat{B} = \frac{1}{J-1}\sum_{j=1}^{J}\left(\overline{\xi}_{\bullet j} - \overline{\xi}_{\bullet\bullet}\right)^2$ the estimate of the "between" variance of the mean σ^2/I, and with $B = \widehat{B}I$ an estimate of the variance. We denote with $\widehat{W} = \frac{1}{J}\sum_{j=1}^{J}\frac{1}{I}\sum_{i=1}^{I}\left(\overline{\xi}_{ij} - \overline{\xi}_{\bullet j}\right)^2$ and with $W = \frac{1}{J}\sum_{j=1}^{J}\frac{1}{I-1}\sum_{i=1}^{I}\left(\overline{\xi}_{ij} - \overline{\xi}_{\bullet j}\right)^2$ two estimates of "within" variance.

To have sensible results, one should have $lim_{I\to\infty}\widehat{B} \to 0$ and $lim_{I\to\infty}\widehat{W} \to$ constant. These can obviously be done for any moments, not just the variance.

Figures 7 and 8 report W (line with circles) and $(\widehat{W} + \widehat{B})$ (line with stars) of three measures of parameters moments: "m2," being a measure of the variance; "m3" based on third moments; and "interval," being constructed from the 80 percent confidence interval around the parameter mean. As was said before, to have reliable results, these should be relatively constant and should converge. Figures 9 and 10 show an aggregate measure based on the eingenvalues of the variance-covariance matrix. The horizontal axis represents

Figure 7. Convergence Diagnostic. Univariate Analysis. Pre-1999

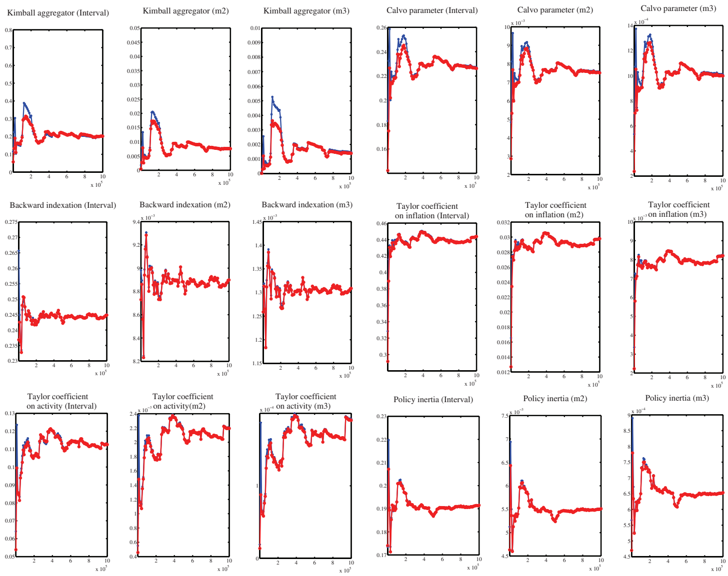


Figure 8. Convergence Diagnostic. Univariate Analysis. Post-1999

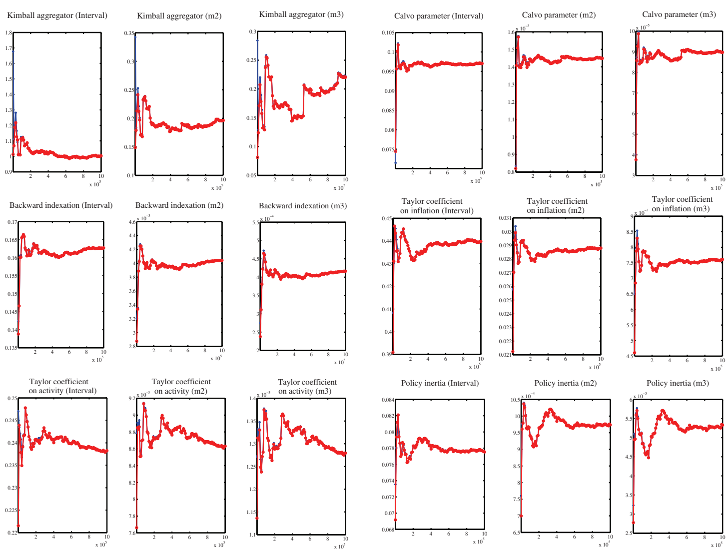


Figure 9. Convergence Diagnostic. Multivariate Analysis. Pre-1999

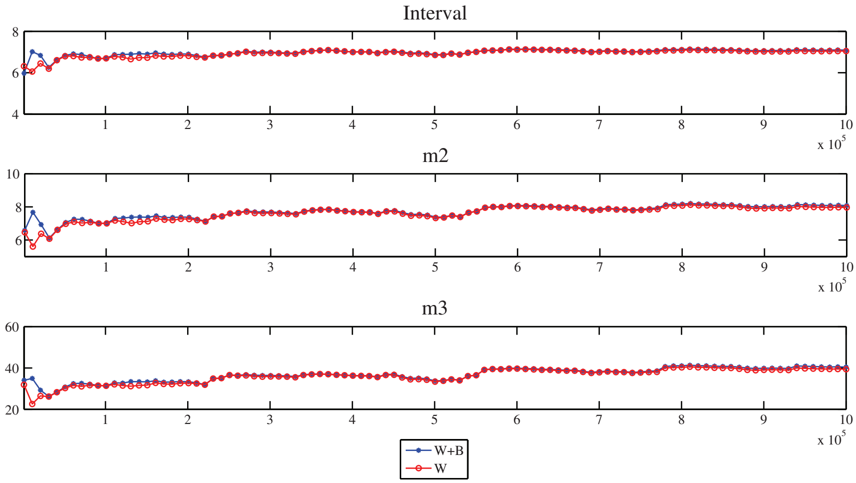
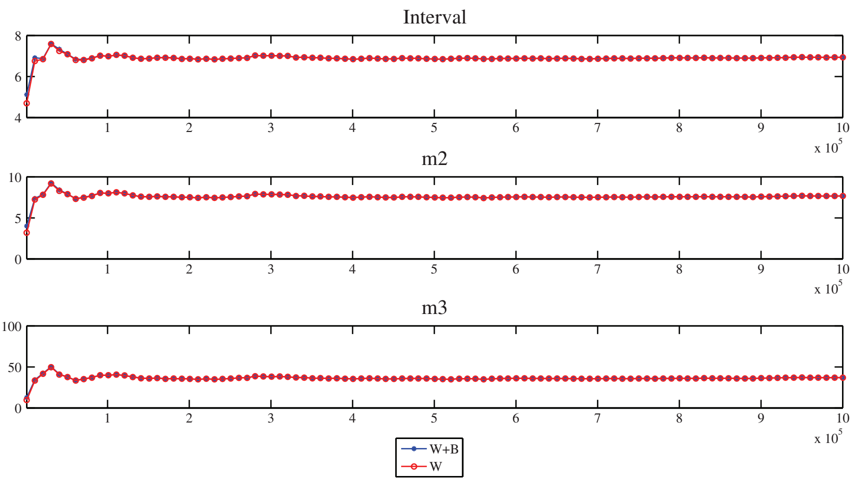


Figure 10. Convergence Diagnostic. Multivariate Analysis. Post-1999



the number of Metropolis-Hastings iterations, whereas the vertical axis represents the measure of the parameter moments.

Both univariate and multivariate analysis confirm that we obtain convergence and stability in all measures of the parameter moments.

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