



Calvo vs. Rotemberg in a trend inflation world: An empirical investigation

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ARTICLE INFO

Article history:

Received 5 August 2010

Accepted 12 May 2011

Available online 12 June 2011

JEL classification:

C32

E3

E52

Keywords:

Calvo

Rotemberg

Trend inflation

Bayesian estimations

ABSTRACT

This paper estimates and compares New-Keynesian DSGE monetary models of the business cycle derived under two different pricing schemes—Calvo (1983) and Rotemberg (1982)—under a positive trend inflation rate. Our empirical findings (i) support trend inflation as an empirically relevant feature of the U.S. great moderation; (ii) provide evidence in favor of the statistical superiority of the Calvo setting; (iii) point to a substantially lower degree of price indexation under Calvo. We show that the superiority of the Calvo model is due to the restrictions imposed by such a pricing scheme on the aggregate demand equation.

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1. Introduction

The Calvo (1983) and Rotemberg (1982) models are the two most popular pricing schemes in the New-Keynesian business cycle literature.¹ Under the typically employed linear approximation around “zero inflation in steady state”, these two pricing mechanisms lead to the very same reduced-form macroeconomic dynamics (Rotemberg, 1987; Roberts, 1995) and to equivalent welfare indications (Nisticò, 2007).² Given such a model equivalence, the choice of the Calvo vs. Rotemberg pricing scheme has typically been no more than a matter of macroeconomists’ taste.

In a recent contribution, Ascari and Rossi (2009) show that, contrary to conventional wisdom, the Calvo and Rotemberg models may imply substantially different macroeconomic dynamics if log-linearized around a positive steady state inflation rate, that is assuming trend inflation.³ The two models, then, have very different policy implications regarding the inflation–output relationships, the determinacy conditions, and the disinflation dynamics. Given that (i) the Calvo and the Rotemberg models are the two most popular way of modeling nominal price rigidities, (ii) they result in different log-linearized dynamic macroeconomic models under positive trend inflation, and that (iii) positive mean inflation is an undeniable empirical fact in

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¹ For in-depth analyses of the New-Keynesian model of the business cycle, see King (2000) and Woodford (2003).

² Lombardo and Vestin (2008) discuss the conditions under which welfare costs might be different under these two pricing schemes.

³ As in the literature, trend inflation indicates a positive steady state level of inflation.

OECD countries in the post-WWII sample, it seems natural to proceed to a comparative quantitative investigation of these two pricing schemes.⁴

This paper fits the Calvo and Rotemberg frameworks derived under positive trend inflation to 1984:I–2008:II U.S. macroeconomic data. Several findings arise. First, models acknowledging a positive trend inflation rate display a better (or, at least, no worse) fit than a baseline “zero inflation in steady state” framework.⁵ Given the different policy implications stemming from a framework incorporating trend inflation (as opposed to the baseline model) in terms of optimal policy and determinacy of simple monetary policy rules (Ascari and Ropele, 2007, 2009), our results push towards the employment and development of macroeconomic frameworks consistently accounting for a positive steady state inflation rate. Second, the U.S. data support Calvo as empirically superior with respect to the standard Rotemberg framework, where adjustment costs are measured in terms of the final-good. We stress that this result emerges out of a set of empirical exercises conducted with macro data. Of course, the Rotemberg model can easily be rejected by a quick look at micro-data, which suggest the presence of price dispersion. The Rotemberg model is just ill-suited to replicate such dispersion. However, it is potentially well-equipped to replicate the dynamics of inflation at a macro level. The philosophy we adopt here is that a useful macro model of price stickiness should not primarily aim to match any feature of micro-data, but it should be thought instead as a (*de facto* reduced form) model of some form of cost of changing prices, and should be primarily judged on its ability to fit the data at a macro level. What we show in this paper is that macroeconomic, aggregate data favor the Calvo framework when compared with the standard Rotemberg model. To our knowledge, this result is novel in the literature.

We also verify that the Calvo model calls for a substantially lower degree of price indexation with respect to the standard Rotemberg framework. As in Smets and Wouters (2007), this result arises in *absence* of any stochastic model for the low frequency of the inflation rate, i.e., without appealing to any exogenous process modeling the possibly time-varying trend inflation as in Ireland (2007) and Cogley and Sbordone (2008). Further exercises reveal that this result is due to the different wedge characterizing the two pricing framework, which affects differently the dynamics and particularly the demand side of the two models. More in detail, while in the Calvo model price dispersion creates a wedge between output and labor, in the standard Rotemberg model price adjustment costs introduce a wedge between consumption and output, because these costs are measured in terms of output. As shown in Ascari and Rossi (2009), it turns out that with trend inflation the two wedges are responsible for a different IS schedule. Indeed, the Calvo model is consistent with a standard IS schedule, whereas the standard Rotemberg framework implies the presence of current inflation together with expected inflation in the IS equation. As a matter of fact, this alternative IS schedule leads to a worse empirical performance and it induces a higher degree of indexation.

The standard formulation of the Rotemberg pricing scheme is just a shortcut for modeling some sort of “customer anger costs”, as originally suggested by Rotemberg (1982). Considerable uncertainty thus remains about the presence and exact nature of this wedge.⁶ We then consider an alternative version of the Rotemberg model where the adjustment costs are measured in labor units. This model implies a standard IS schedule and it fits the data as good as Calvo, calling for an equally low degree of price indexation. Two novel conclusions follow from this result. First, from a macroeconomic perspective, the standard Rotemberg specification should be substituted by a far less common formulation where the adjustment costs are measured in terms of units of labor inputs rather than units of output. Second, more generally, price indexation and the performance of the Phillips curve should be assessed with full-system strategies, exploiting informations that would not be taken into account in a single-equation approach.

We conduct further investigations along two dimensions. First, we introduce a deterministic trend in technology as in Mattesini and Nisticò (2010), which enables us to estimate the model with the growth rate of output as an observable. Then, we consider a policy rule featuring two lags of the policy rate among the regressors and a systematic reaction to the output growth as in Coibion and Gorodnichenko (2011a,b). Our main results are not overturned under these perturbations of our baseline analysis.

Note that our findings are also relevant from a policy standpoint. Ascari and Rossi (2009) show that the regions of the space of the Taylor rule parameters that induce a determinate rational expectation equilibrium are quite different under Calvo and under Rotemberg when trend inflation is taken into account. Hence, establishing which of the two models is more reliable from an empirical standpoint is important for a correct assessment of the set of implementable monetary policy rules.

Other papers stress the importance of considering trend inflation in empirical work. Benati (2008) estimates a NKPC for a variety of countries, and shows that price indexation à la Christiano et al. (2005) is not stable across different samples in countries that explicitly adopted an inflation targeting scheme. He relates this instability to different policy regimes, so demonstrating that indexation is “not structural in the sense of Lucas”. Elaborating on this paper, Benati (2009) estimates different NKPCs derived under alternative pricing schemes. His results corroborate and extend his previous findings,

⁴ The zero steady state assumption, which is empirically problematic, has already been questioned from a theoretical standpoint. Ascari (2004) and Yun (2005) show that first-order effects arise on the Calvo price setting setup under trend inflation. Elaborating further with the Calvo setup, Ascari and Ropele (2007, 2009) study the implications of different trend inflation levels for the optimal monetary policy and for the Taylor principle.

⁵ What we mean here is that over and above its use in accounting for the sample mean of inflation, a positive rate of trend inflation influences the local dynamics around the steady state. Therefore, keeping track of this influence improves the performance of the model.

⁶ See Rotemberg (forthcoming) for an attempt to provide some more microfoundations on the source of the anger cost.

i.e., the degree of price indexation is not invariant across different policy regimes, and it tends to zero under the more recent, stable regimes. Notably, Benati (2009) supports, among others, Ascari and Ropele's (2007) derivation of the Calvo model under trend inflation for a variety of countries.⁷ He considers a step-function to model possible drifts in the inflation target. Differently from Benati (2009), who works with a fully fledged New-Keynesian DSGE framework, Cogley and Sbordone (2008) estimate a NKPC embedding a drifting trend inflation coupled with a TVC-VAR model. They find that, once drifts in trend inflation are accounted for, price indexation in the U.S. is zero, i.e., a purely forward looking NKPC fits the data well without the need for *ad hoc* backward-looking components.⁸ Coibion and Gorodnichenko (2011a,b) estimate a Taylor rule with trend inflation for the post-WWII, and couple it with a calibrated New-Keynesian model derived under positive trend inflation. They show that the U.S. economy switched to determinacy during the Volcker disinflation because of changes in the Federal Reserve's response to macro variables and the decline in trend inflation.⁹ Paciello (2009) estimates a Calvo-based NKPC with constant trend inflation for the post-WWII via indirect inference, and shows that such a model is able to match the dynamic responses of inflation to monetary policy and technology shocks even in absence of indexation, an ability not enjoyed by the standard, zero steady state inflation framework.

Our investigation departs from the ones above along different dimensions. First and foremost, our paper focuses on the estimation of, and the empirical comparison between, two different frameworks, i.e., the standard Rotemberg and Calvo pricing models. To our knowledge, this is the only contribution to date assessing the relative empirical relevance of these two very widely employed pricing schemes under trend inflation. Second, we focus on two models displaying a *constant* trend inflation rate, i.e., displaying no exogenous random-walk type of process for the Fed's inflation target. Still, the version of the Calvo model preferred by the data is that with no-price indexation. With respect to Benati (2009), we provide evidence for the U.S. case, therefore complementing his battery of estimates. With respect to Cogley and Sbordone (2008) and Paciello (2009), we consider a structural representation of the demand side of the economy, rather than a reduced-form TVC-VAR. This is obviously important from an econometric standpoint, because the identification of forward and backward looking terms in the NKPC also depends on how the remaining structural equations are modeled. When such equations are not specified, as in the NKPC-VAR approach, the meaning of the economic restrictions imposed to the estimation is unclear, as pointed out by Cogley and Sbordone (2008) themselves. Also from a theoretical point of view, our analysis shows the importance of estimating the full model equations, because the assumed pricing scheme may affect not only the supply side of the model, but also the other model equations, as in the case of the Rotemberg model. Moreover, differently from Paciello (2009), we conduct our empirical analysis with Bayesian techniques. Our choice is driven by the possibly superior performance against indirect inference (impulse response matching) as far as this class of DSGE models is concerned (Canova and Sala, 2009). Finally, we concentrate on a stable subsample (great moderation), which is likely to feature a unique equilibrium even under historically plausible values for trend inflation (Coibion and Gorodnichenko, 2011a), and a more stable low-frequency component of inflation. This sample choice makes our assumption of a constant trend inflation more palatable. We see our contribution as complementary to those presented above.

The paper is structured as follows. Section 2 describes the two frameworks we deal with and highlights the relevant differences. Section 3 presents and discusses our empirical findings, with particular emphasis on the estimated degree of price indexation. Section 4 scrutinizes further the two pricing schemes, and discusses the reasons of Calvo's superiority. This section introduces and assesses an alternative specification of Rotemberg adjustment costs measured in labor units. Section 5 scrutinizes the two models further by either allowing for a deterministic trend in technology or by modeling a more general policy rule. Section 6 concludes, and draws some directions for further research.

2. The theoretical models

In this section we sketch a small-scale New-Keynesian model in the two versions of the Rotemberg (1982) and the Calvo (1983) price setting scheme. The model economy is composed of a continuum of infinitely lived consumers, producers of final and intermediate goods. Households have the following instantaneous and separable utility function:

$$U(C_t, N_t) = \frac{C_t^{1-\sigma}}{1-\sigma} - d_n \frac{N_t^{1+\varphi}}{1+\varphi},$$

where C_t is a consumption basket (with elasticity of substitution among goods ε) and N_t are labor hours.

Final good market is competitive and the production function is given by $Y_t = [\int_0^1 Y_{i,t}^{(\varepsilon-1)/\varepsilon} di]^{1/(\varepsilon-1)}$. Final good producers demand for intermediate inputs is therefore equal to $Y_{i,t+j} = (P_{i,t}/P_{t+j})^{-\varepsilon} Y_{t+j}$. The intermediate inputs $Y_{i,t}$ are produced by a continuum of firms indexed by $i \in [0, 1]$ with the following simple constant return to scale technology $Y_{i,t} = A_t N_{i,t}$, where

⁷ The list considered by Benati (2009) includes the Euro area, West Germany, Germany, France, Italy, UK, Canada, Sweden, Australia, New Zealand, and Switzerland.

⁸ Barnes et al. (2009) use a different estimation methodology and a more flexible indexation scheme with respect to Cogley and Sbordone (2008), and show that indexation to past inflation may be substantial in the post-WWII sample. See also the reaction to the former paper by Cogley and Sbordone (2010).

⁹ Schorfheide (2005) and Ireland (2007) also embed a time-varying inflation target in their models, but without consequences for the specification of the NKPC due to the assumption of full indexation.

labor is the only input and $\ln A_t = a_t$ is an exogenous productivity shock, which follows an AR(1) process. The intermediate good sector is monopolistically competitive.

2.1. Firms and price setting: Rotemberg (1982) vs. Calvo (1983)

The Calvo model: The Calvo price setting scheme assumes that in each period there is a fixed probability $1-\theta$ that a firm can re-optimize its nominal price, i.e., $P_{i,t}^*$. With probability θ , instead, the firm automatically and costlessly adjust its price according to an indexation rule. The price setting problem is

$$\max_{P_{i,t}^*} E_t \sum_{j=0}^{\infty} \mathcal{D}_{t,t+j} \theta^j \left[\frac{P_{i,t}^* (\bar{\pi}^{\chi j})^{1-\mu} (\Pi_{t,t+j-1}^{\chi})^{\mu}}{P_{t+j}} Y_{i,t+j} - MC_{i,t+j} Y_{i,t+j} \right], \tag{1}$$

$$\text{s.t. } Y_{i,t+j} = \left[\frac{P_{i,t}^* (\bar{\pi}^{\chi j})^{1-\mu} (\Pi_{t,t+j-1}^{\chi})^{\mu}}{P_{t+j}} \right]^{-\varepsilon} Y_{t+j} \tag{2}$$

and

$$\Pi_{t,t+j-1} = \begin{cases} \left(\frac{P_t}{P_{t-1}} \right) \left(\frac{P_{t+1}}{P_t} \right) \times \dots \times \left(\frac{P_{t+j-1}}{P_{t+j-2}} \right) & \text{for } j = 1, 2, \dots, \\ 1 & \text{for } j = 0, \end{cases} \tag{3}$$

where $\mathcal{D}_{t,t+j} \equiv \beta^j Y_t / Y_{t+j}$ represents firms' stochastic discount factor, $MC_{i,t+j} = W_{i,t+j} / P_{t+j} A_{t+j}$ is the real marginal cost function, and $\bar{\pi}$ denotes the central bank's inflation target and it is equal to the level of trend inflation. The indexation scheme in (1) is very general. In particular: (i) $\chi \in [0, 1]$ allows for any degree of price indexation; (ii) $\mu \in [0, 1]$ allows for any degree of (geometric) combination of the two types of indexation usually employed in the literature: to steady state inflation (e.g., Yun, 1996) and to past inflation rates (e.g., Christiano et al., 2005).

In the Calvo price setting framework, prices are staggered because firms charging prices at different periods will set different prices. Then, in each given period t , there will be a distribution of different prices. Price dispersion results in an inefficiency loss in aggregate production. Formally:

$$N_t^d = \frac{Y_t}{A_t} \underbrace{\int_0^1 \left(\frac{P_{i,t}}{P_t} \right)^{-\varepsilon} di}_{s_t} = s_t \frac{Y_t}{A_t}. \tag{4}$$

Schmitt-Grohé and Uribe (2007) show that s_t is bounded below at one, so that s_t represents the resource costs due to relative price dispersion under the Calvo mechanism. Indeed, the higher s_t , the more labor N_t^d is needed to produce a given level of output. Note that price dispersion creates a wedge between aggregate output and aggregate employment. To close the model, the aggregate resource constraint is simply given by

$$Y_t = C_t. \tag{5}$$

The standard Rotemberg model: The standard Rotemberg model assumes that a monopolistic firm faces a quadratic cost of adjusting nominal prices, that can be measured in terms of the final-good and given by

$$\frac{\varphi_p^Y}{2} \left(\frac{P_{i,t}}{(\pi_{t-1}^{\chi})^{\mu} (\bar{\pi}^{\chi})^{1-\mu} P_{i,t-1}} - 1 \right)^2 Y_t, \tag{6}$$

where $\varphi_p^Y > 0$ determines the degree of nominal price rigidity. As stressed in Rotemberg (1982), the adjustment cost seeks to account for the negative effects of price changes on the customer–firm relationship. These negative effects increase in magnitude with the size of the price change and with the overall scale of economic activity, Y_t . As for the Calvo model, (6) includes a general specification for the adjustment cost used by, e.g., Ireland (2007), among others. In particular, the adjustment cost will depend on the ratio between the new reset price and the one set during the previous period, adjusted by a (geometric) combination of steady state inflation and of past inflation. The parameters μ and χ play a parallel role as in the indexation scheme in the Calvo model.

The problem for the firm i is then:

$$\max_{\{P_{i,t+j}\}_{j=0}^{\infty}} E_t \sum_{j=0}^{\infty} \mathcal{D}_{t,t+j} \left\{ \frac{P_{i,t+j}}{P_{t+j}} Y_{i,t+j} - MC_{i,t+j} Y_{i,t+j} - \frac{\varphi_p^Y}{2} \left(\frac{P_{i,t+j}}{(\pi_{t+j-1}^{\chi})^{\mu} (\bar{\pi}^{\chi})^{1-\mu} P_{i,t+j-1}} - 1 \right)^2 Y_{t+j} \right\}, \tag{7}$$

$$\text{s.t. } Y_{i,t+j} = \left[\frac{P_{i,t+j}}{P_{t+j}} \right]^{-\varepsilon} Y_{t+j}, \tag{8}$$

where the notation is as above. Firms can change their price in each period, subject to the payment of the adjustment cost. Therefore, all the firms face the same problem, and thus will choose the same price and output. In other words the equilibrium is

symmetric: $P_{i,t} = P_t, Y_{i,t} = Y_t, W_{i,t} = W_t$ and $MC_{i,t} = MC_t \forall i$. Given this symmetry, and differently with respect to the Calvo model, in the Rotemberg model the aggregate production function features no inefficiency due to price dispersion, therefore:

$$Y_t = A_t N_t. \quad (9)$$

In this version of the Rotemberg model, which is the one typically adopted in the literature, the adjustment cost enters the aggregate resource constraint, which creates an inefficiency wedge between output and consumption:

$$Y_t = \left[1 - \frac{\varphi_p^Y}{2} \left(\frac{P_t}{(\pi_{t-1}^\chi)^\mu (\bar{\pi}^\chi)^{1-\mu} P_{t-1}} - 1 \right)^2 \right]^{-1} C_t = \Psi_t C_t. \quad (10)$$

Some key-differences between the Calvo and the Rotemberg models arise. In the Calvo model, the cost of nominal rigidities, i.e., price dispersion, creates a wedge between aggregate employment and aggregate output, making aggregate production less efficient. In the Rotemberg model, instead, the cost of nominal rigidities, i.e., the adjustment cost, creates a wedge between aggregate consumption and aggregate output, because part of the output goes in the price adjustment cost. As shown in Ascari and Rossi (2009), and evident from (4) and (10), both these wedges are *non-linear functions of inflation* and they increase with trend inflation. However, both wedges take the same unitary value in steady state under two particular cases: (i) a net steady state inflation equals zero, and/or (ii) full indexation to past or trend inflation.

2.2. The log-linearized frameworks

We now present the log-linearized versions of the two pricing frameworks we deal with (for a full derivation, see Ascari and Ropele, 2007, 2009; Ascari and Rossi, 2009). Again, we stress that the derivation allows for a non-zero value for the inflation rate in steady state, which may be interpreted as the target pursued by the Federal Reserve in conducting the U.S. monetary policy.

The Calvo model: The Calvo model is described by the following first-order difference equations:

$$\Delta_t = [\beta \bar{\pi}^{1-\chi} + \eta(\varepsilon - 1)] \Delta_{t+1|t} + \kappa \hat{y}_t - \lambda \varphi a_t + \lambda \varphi \hat{s}_t + \eta \hat{\phi}_{t+1|t}, \quad (11)$$

$$\hat{\phi}_t = (1 - \sigma)(1 - \theta \beta \bar{\pi}^{(\varepsilon-1)(1-\chi)}) \hat{y}_t + \theta \beta \bar{\pi}^{(\varepsilon-1)(1-\chi)} [(\varepsilon - 1) \Delta_{t+1|t} + \hat{\phi}_{t+1|t}], \quad (12)$$

$$\hat{s}_t = \xi \Delta_t + \theta \bar{\pi}^{\varepsilon(1-\chi)} \hat{s}_{t-1}, \quad (13)$$

$$\hat{y}_t = \hat{y}_{t+1|t} - \sigma^{-1} (\hat{i}_t - \hat{\pi}_{t+1|t}) + g_t, \quad (14)$$

where $\Delta_t \equiv \hat{\pi}_t - \chi \mu \hat{\pi}_{t-1}$, $\hat{\pi}$ stands for the inflation rate, \hat{y} for output, a is the technological shock, g is the demand shock.¹⁰ Hatted variables indicate percentage deviations with respect to steady state values. The notation $x_{t+1|t}$ indicates the expectation in t of x_{t+1} . σ is the relative risk aversion parameter, φ the labor supply elasticity, β the discount factor, ε the Dixit–Stiglitz elasticity of substitution among goods, θ the Calvo parameter, χ the degree of price indexation, μ the relative weight of indexation to past inflation vs. trend inflation, and $\bar{\pi}$ the steady state, trend inflation rate. Finally, λ, η, κ , and ξ in Eqs. (11)–(14) are the following convolutions of parameters:

$$\lambda \equiv \frac{(1 - \theta \bar{\pi}^{(\varepsilon-1)(1-\chi)}) (1 - \theta \beta \bar{\pi}^{\varepsilon(1-\chi)})}{\theta \bar{\pi}^{(\varepsilon-1)(1-\chi)}},$$

$$\eta \equiv \beta (\bar{\pi}^{1-\chi} - 1) [1 - \theta \bar{\pi}^{(\varepsilon-1)(1-\chi)}],$$

$$\kappa \equiv (\lambda_{(\bar{\pi}, \varepsilon)} (\sigma + \varphi) - \eta_{(\bar{\pi}, \varepsilon)} (1 - \sigma)),$$

$$\xi \equiv \frac{\varepsilon \theta \bar{\pi}^{(\varepsilon-1)(1-\chi)} (\bar{\pi}^{1-\chi} - 1)}{1 - \theta \bar{\pi}^{(\varepsilon-1)(1-\chi)}}.$$

Notably, all the convolutions of the log-linearized model are a function of the trend inflation rate $\bar{\pi}$, that generally tends to increase the coefficients on the forward looking variables (see Ascari, 2004; Yun, 2005; Hornstein and Wolman, 2005; Kiley, 2007). Moreover, the log-linearized NKPC is influenced by the price dispersion process s_t . Under Calvo, just a fraction $(1 - \theta)$ of firms is allowed to re-optimize in each period, then price dispersion arises. Conditional on a strictly positive trend inflation rate, price dispersion assumes a first-order relevance and influences the evolution of the log-linearized inflation rate. Moreover, price dispersion has a backward-looking dynamics. The forward looking auxiliary process ϕ_t also participates to the determination of inflation.

¹⁰ The demand shock captures the discrepancies between the predicted output level according to the deterministic part of the Euler equation (14) and output's realizations. Possible interpretations of this shock involve stochastic variations in households' preferences and/or perturbations to the national account equation.

The aggregate demand equation (14) is standard in this framework. Importantly, this does not hold in the Rotemberg model, in which adjustment costs enter the log-linearized expression for output, as explained below. Therefore, also the formulation of the IS schedule carries relevant information on the pricing mechanism prevailing in the economic system.¹¹

The Rotemberg model: The Rotemberg model is characterized by the following difference equations:

$$\hat{\pi}_t = \gamma_p \hat{\pi}_{t-1} + \gamma_f \beta \hat{\pi}_{t+1|t} + \gamma_{dy} \beta (1-\sigma) \Delta \hat{y}_{t+1|t} + \gamma_{mc} \widehat{mc}_t, \tag{15}$$

$$\widehat{mc}_t = (\sigma + \varphi) \hat{y}_{t-\zeta_c} \sigma \hat{\pi}_t + \zeta_c \sigma \mu \chi \hat{\pi}_{t-1} - (1 + \varphi) a_t, \tag{16}$$

$$\hat{y}_t = \hat{y}_{t+1|t} - \zeta_c \Delta \hat{\pi}_{t+1|t} + \zeta_c \mu \chi \Delta \hat{\pi}_t - \sigma^{-1} (i_t - \hat{\pi}_{t+1|t}) + g_t, \tag{17}$$

where \widehat{mc} stands for real marginal costs, and the notation has the same interpretation as in the previous subsection. The coefficients γ_p , γ_f , γ_{dy} , γ_{mc} , and ζ_c are convolutions of the structural parameters of the model:

$$\zeta_c = \varphi_p^Y (\bar{\pi}^{1-\chi} - 1) \frac{\bar{\pi}^{1-\chi}}{\left[1 - \frac{\varphi_p^Y}{2} (\bar{\pi}^{1-\chi} - 1)^2\right]},$$

$$\frac{C}{\bar{Y}} = \left(1 - \frac{\varphi_p^Y}{2} (\bar{\pi} - 1)^2\right),$$

$$\varrho \equiv (2\bar{\pi}^{2(1-\chi)} - \bar{\pi}^{(1-\chi)}) (1 + \beta \chi \mu) \frac{C}{\bar{Y}} + \beta [(\bar{\pi}^{1-\chi} - 1) \bar{\pi}^{1-\chi}]^2 \sigma \varphi_p^Y (1 + \mu \chi),$$

$$\gamma_p = \frac{(2\bar{\pi}^{2(1-\chi)} - \bar{\pi}^{(1-\chi)}) \chi \mu \frac{C}{\bar{Y}} + \beta [(\bar{\pi}^{1-\chi} - 1) \bar{\pi}^{1-\chi}]^2 \sigma \varphi_p^Y \mu \chi}{\varrho},$$

$$\gamma_f = \frac{(2\bar{\pi}^{2(1-\chi)} - \bar{\pi}^{(1-\chi)}) \frac{C}{\bar{Y}} + [(\bar{\pi}^{1-\chi} - 1) \bar{\pi}^{1-\chi}]^2 \sigma \varphi_p^Y}{\varrho},$$

$$\gamma_{dy} = \frac{(\bar{\pi}^{2(1-\chi)} - \bar{\pi}^{1-\chi}) \frac{C}{\bar{Y}}}{\varrho},$$

$$\gamma_{mc} = \frac{(\varepsilon - 1 + \varphi_p^Y (\bar{\pi}^{2(1-\chi)} - \bar{\pi}^{1-\chi})) (1 - \beta) \frac{C}{\bar{Y}}}{\varphi_p^Y \varrho}.$$

As often assumed in the literature,¹² it is possible to draw a relationship between the Rotemberg adjustment cost φ_p^Y and the Calvo parameter θ imposing the condition $\varphi_p^Y = (\varepsilon - 1)\theta / (1 - \theta)(1 - \beta\theta)$, that implies the same first-order dynamics of the two models in the case of zero steady state inflation. Such relationship will enable us to indirectly estimate the Rotemberg adjustment cost by focusing on the Calvo parameter θ and to use the very same prior densities for the structural parameters of the two models we ultimately aim at comparing, i.e., Calvo and Rotemberg.¹³

A few comments are in order. First, the impact of trend inflation is evident when looking at Eqs. (15)–(17) and their convolutions of parameters. As in the Calvo model, trend inflation alters the inflation dynamics by directly affecting the NKPC coefficients. Higher trend inflation increases the coefficient relative to expected and past inflation as well as the coefficient of real marginal costs (Ascari and Rossi, 2009). Notice that the presence of past inflation in (15) is due to indexation to past inflation. With no indexation to past inflation, i.e., with $\mu = 0$, however, the coefficient γ_p equals zero and the NKPC becomes completely forward looking. Recall that this is not the case in the Calvo model, because even if past inflation disappears when $\mu = 0$, price dispersion dynamic introduces a backward-looking component. Furthermore, the expected difference of detrended output appears in the NKPC, Eq. (15), because of the influence that trend inflation exerts over firms' discount factor.

Second, because of the presence of the price adjustment cost, in the Rotemberg model the log-linearized resource constraint can be written as

$$\hat{c}_t = \hat{y}_t - \varphi_p^Y (\bar{\pi}^{1-\chi} - 1) \frac{\bar{\pi}^{1-\chi}}{\left[1 - \frac{\varphi_p^Y}{2} (\bar{\pi}^{1-\chi} - 1)^2\right]} \hat{\pi}_t + \varphi_p^Y (\bar{\pi}^{1-\chi} - 1) \frac{\bar{\pi}^{1-\chi} \mu \chi}{\left[1 - \frac{\varphi_p^Y}{2} (\bar{\pi}^{1-\chi} - 1)^2\right]} \hat{\pi}_{t-1}. \tag{18}$$

¹¹ The formulation proposed in this paper assumes no habit formation. We experimented with alternative models admitting for lagged realizations of output to enter the IS schedules. Our results turned out to be unaltered. These additional exercises are available upon request.

¹² See Rotemberg (1987) or Lombardo and Vestin (2008) for details.

¹³ We also estimated a version of the Rotemberg model in which the adjustment cost is a free parameter, and verified no appreciable variations in the marginal likelihood.

This equation shows that to a first-order approximation the Rotemberg model: (i) implies a wedge between output and consumption; (ii) this wedge depends positively on current and past inflation level; (iii) the coefficients of $\hat{\pi}_t$ and $\hat{\pi}_{t-1}$ in (18) increase with trend inflation; (iv) the wedge disappears with zero steady state inflation rate or with full indexation, i.e., with $\chi = 1$. Such a wedge affects also the amount of resources produced in the economy. Consequently, the IS, Eq. (17), features the first difference in inflation rates. For the same reasons, the real marginal cost to depend also on actual inflation and past inflation (see the additional terms $\zeta_c \sigma \hat{\pi}_t$ and $\zeta_c \sigma \mu \chi \hat{\pi}_{t-1}$ in (16)).

Notably, under the peculiar case of zero trend inflation, i.e., $\bar{\pi} = 1$, both the Rotemberg and the Calvo frameworks lines up with the standard hybrid New-Keynesian formulation allowing for price indexation to past/steady state inflation. The same holds true in a full indexation scenario, i.e., when $\chi = 1$.

To sum up, the different wedges which characterize the Calvo and the standard version of the Rotemberg model induce three main differences in the two log-linearized representations. First, the Calvo model displays price dispersion, which enters the NKPC as an endogenous predetermined variable. By contrast, given the symmetry in the Rotemberg economy, price dispersion is absent in the Rotemberg model. Second, the presence of the price adjustment cost in the Rotemberg model causes the real marginal cost to depend also on actual and past inflation. Finally, the price adjustment cost generates a wedge between output and consumption in the resource constraint (10), which is reflected in the IS curve (17). As shown by Ascari and Rossi (2009), these differences are relevant from a policy standpoint, because of their impact on the definition of the determinacy territory associated to simple, implementable Taylor-type rules.

2.3. Closing the models

The two models are closed by a common set of equations:

$$\hat{i}_t = \alpha_i \hat{i}_{t-1} + (1 - \alpha_i)(\alpha_\pi \hat{\pi}_t + \alpha_y \hat{y}_t) + m_t, \quad (19)$$

$$z_t = \rho_z z_{t-1} + \varepsilon_{zt}, \quad \varepsilon_{zt} \sim N(0, \sigma_z^2), \quad z \in \{a, g, m\}. \quad (20)$$

Eq. (19) is a standard policy rule postulating a smoothed reaction of the policy rate \hat{i}_t to fluctuations in inflation and output, with stochastic deviations driven by the monetary policy shock m_t . Eq. (20) defines the stochastic properties of the mutually uncorrelated shocks hitting the system.

3. Econometric exercise

Our investigation focuses on U.S. data. We employ three “observables”, i.e., the log of the quarterly gross growth rate of the GDP deflator π_t^{obs} , the log-deviation of real GDP with respect to its long-run trend y_t^{obs} ,¹⁴ and the Federal Funds Rate i_t^{obs} .¹⁵ Our measure of detrended output, being mainly statistical, is robust to model misspecification, and it is also justified by the absence in this model of physical capital, which would probably return a severely misspecified model-consistent measure of natural output. Proxies of detrended output robust to model misspecification have recently been employed by, among others, Lubik and Schorfheide (2004); Boivin and Giannoni (2006); Benati (2008, 2009) Benati and Surico (2008, 2009).¹⁶

Several authors (Clarida et al., 2000; Lubik and Schorfheide, 2004; Boivin and Giannoni, 2006, Benati and Surico, 2009, and Mavroeidis, 2010) have documented a break in the U.S. monetary policy conduct in correspondence to the advent of Paul Volcker as chairman of the Federal Reserve. Changes in the U.S. macro-dynamics possibly consequential to such a monetary policy shift have also been investigated by D’Agostino et al. (forthcoming), Benati and Surico (2008) and Cogley et al. (2010), who document a variation in inflation predictability when entering the 1980s, and by Castelnovo and Surico (2010), who show how VAR impulse response functions may be affected by a drift towards a more hawkish monetary policy. Importantly, Coibion and Gorodnichenko (2011a) show that the switch from multiple equilibria to uniqueness is supported also in a context in which trend inflation is allowed to be positive. In particular, they show that the U.S. economy has entered the unique equilibrium territory at the end of the “Volcker experiment” because (i) the Fed has engaged in a stronger systematic reaction against inflation fluctuations and (ii) trend inflation has fallen. Therefore, we condition our analysis on the “great moderation” period 1984:I–2008:II. Our end-of-sample choice enables us to avoid dealing with the acceleration of the financial crises began with the bankruptcy of Lehman Brothers in September 2008, which triggered non-standard policy moves by the Fed (Brunnermeier, 2009).

¹⁴ We employ the Hodrick–Prescott detrended output (relative weight of the smoothing component: 1600). Further investigations with a model featuring a deterministic trend in technology and estimated with growth rates are proposed in Section 5.

¹⁵ The source of the data is the Federal Reserve Bank of St. Louis’ website, i.e., <http://research.stlouisfed.org/fred2/>. Quarterly observations of the federal funds rate were constructed as averages of monthly observations. The detrended output and the policy rate were demeaned prior to estimation.

¹⁶ For an alternative approach, based on a model-consistent treatment of the real GDP trend, see e.g. Smets and Wouters (2007), Justiniano and Primiceri (2008), Castelnovo and Nisticò (2010), and Riggi and Tancioni (2010).

3.1. Bayesian inference

We estimate the Calvo (11)–(14), (19)–(20) and the Rotemberg (15)–(17), (19)–(20) models with Bayesian techniques (see e.g., An and Schorfheide, 2007). Canova and Sala (2009) show that this technique is less prone to identification issues with respect to alternatives in the context of DSGE models.¹⁷

The following measurement equations link our observables to the latent factors of our models:

$$\begin{bmatrix} \pi_t^{obs} \\ y_t^{obs} \\ i_t^{obs} \end{bmatrix} = \begin{bmatrix} \log(\bar{\pi}) \\ \bar{y} \\ \bar{i} \end{bmatrix} + \begin{bmatrix} \hat{\pi}_t \\ \hat{y}_t \\ \hat{i}_t \end{bmatrix}, \quad (21)$$

where \bar{y} and \bar{i} are the sample means of, respectively, detrended output and the federal funds rate, and π_t^{obs} is computed as $\log(P_t^{obs}/P_{t-1}^{obs})$, where P_t^{obs} is the GDP deflator.

Eq. (21) identifies the quintessence of a trend inflation model, i.e., its ability to shape the steady state inflation rate. Clearly, different trend inflation values will lead to different empirical performances of the different models we will investigate. Our empirical investigation exactly aims at discriminating such models on the basis of their ability to replicate inflation's long-run value on top of its dynamics. Our microfounded models, which are log-linearized around a general trend inflation level, can treat trend inflation in a model-consistent way. This consideration is important when searching for the encompassed “baseline New-Keynesian model”. Indeed, an obvious way to collapse to such model would be that of setting the gross trend inflation rate $\bar{\pi} = 1$, so reconstructing the “zero steady state” assumption typically employed in the literature when deriving such model. However, Eq. (21) makes it clear that, while being logically grounded, this choice would force us to leave the mean of observed inflation unmodeled, so condemning the standard New-Keynesian model to a poor empirical performance. To circumvent this issue, one could demean observed inflation prior to estimation. However, this would probably penalize, in relative terms, the trend inflation models, one of their edges being their ability to model the first moment of observed inflation. To estimate the encompassed baseline, “zero trend inflation” framework, we then set the indexation parameter $\chi = 1$, as in Christiano et al. (2005). In doing so, we switch off the trend inflation-related “extra terms”, and we mute the impact of trend inflation on the relative weights of inflation expectations and marginal costs in the NKPC and IS schedules. We can then assign a positive trend inflation rate (with which we model the inflation mean) to the baseline New-Keynesian model in a theoretically consistent manner.

3.2. Calibrated parameters

We calibrate a subset of models' parameters. We set the discount factor β to 0.99, the elasticity of substitution among goods $\varepsilon = 6$, and the inverse of the labor elasticity φ to 1.

The indexation parameters μ and χ deserve particular attention. Intuitively, these two parameters should have a different impact on the dynamics of the system. Indeed, when looking at the model equations, one may easily notice that they do not always appear in the same joint combination, a necessary condition for these two parameters to be separately identified. However, preliminary attempts to jointly estimate these two parameters invariably missed to achieve convergence to the ergodic posterior density according to the standard convergence diagnostics developed by Brooks and Gelman (1998).¹⁸ For this reasons, we calibrate the relative degree of indexation μ to 1, and we concentrate on the estimated degree of indexation to past inflation in line with Benati (2009) (further discussions on indexation to past vs. steady state inflation are offered in Section 3.3).¹⁹ To verify the solidity of our findings, we also engage in exercises conditional on $\mu = 0$.

We set $\bar{y} = 0.0012$ and $\bar{i} = 0.0131$ (sample means of the corresponding observables employed in the estimation).²⁰ We also calibrate the trend inflation rate to its sample mean, i.e., we set $\bar{\pi} = 1.0065$. This choice is not crucial, because estimating (rather than calibrating) the rate of trend inflation leaves the results unchanged.²¹ Note that, according to the theory, inflation, the nominal interest rate and the real interest rate are linked via the Fisher equation, and the real interest rate is related to the discount factor and the growth rate of output. However, these theoretical relationships hardly hold empirically. Therefore, we did not impose them when conducting our empirical exercises.

¹⁷ An Appendix available from authors' webpages offers details on our estimation strategy.

¹⁸ We experienced convergence problems even with long chains of 1,000,000 accepted draws, a very large number for models of the size we focus on in this paper. Differently, when calibrating the relative degree of indexation μ , 200,000 accepted draws turned out to be largely sufficient to achieve convergence.

¹⁹ A second reason relates to our intention of comparing the estimated degree of price indexation across different models. Such comparison is much more correct if conditional on a single indexation parameter. In other words, the joint estimation of μ and χ would likely lead us to compare a pair $(\mu^{Calvo}, \chi^{Calvo})$ to a pair $(\mu^{Rotemberg}, \chi^{Rotemberg})$ with $\mu^{Calvo} \neq \mu^{Rotemberg}$ and $\chi^{Calvo} \neq \chi^{Rotemberg}$, so rendering the comparison of the different degrees of indexation difficult.

²⁰ Following Lubik and Schorfheide (2004), we employed a longer sample (1954:III–2008:II) to compute the Hodrick–Prescott detrended measure of output than the one we employ in our empirical exercise (1984:I–2008:II). We did so to increase the number of observations exploited to compute the cyclical component of output. As a consequence, the mean of our detrended output measure as for the sample 1984:I–2008:II is different from zero.

²¹ See the working paper version of the paper: Ascari et al. (2010).

Table 1

Priors and posteriors for structural parameters. Models estimated assuming indexation to past inflation. The Table reports the posterior means and the [5th,95th] percentiles. The log-Marginal Likelihoods are computed with the modified harmonic mean estimator by Geweke (1998). Details on the model estimation are reported in the text.

Param.	Description	Prior dens.	Prior mean (Std. dev.)	Posterior mean [5th pct, 95th pct]				
				NK, $\chi = 1$	Calvo, $\mu = 1$	Rotemb., $\mu = 1$	Calvo, $\mu = 0$	Rotemberg, $\mu = 0$
χ	Indexation	Beta	0.50 (0.285)	1.00 [–]	0.13 [0.00,0.28]	0.33 [0.01,0.62]	0.48 [0.00,0.88]	0.75 [0.49,1.00]
θ	Calvo	Beta	0.50 (0.15)	0.43 [0.34,0.53]	0.66 [0.57,0.76]	0.58 [0.48,0.70]	0.70 [0.63,0.78]	0.70 [0.62,0.78]
σ	Risk aversion	Normal	2.50 (0.25)	2.18 [1.74,2.59]	2.35 [1.92,2.76]	2.25 [1.85,2.66]	2.42 [2.01,2.83]	2.38 [1.97,2.82]
α_π	T. rule inflation	Normal	2.00 (0.50)	3.35 [2.79,3.91]	3.31 [2.74,3.86]	3.38 [2.81,3.94]	3.30 [2.68,3.90]	3.31 [2.74,3.90]
α_y	T. rule output	Gamma	0.125 (0.05)	0.14 [0.06,0.22]	0.15 [0.06,0.24]	0.14 [0.06,0.22]	0.15 [0.06,0.24]	0.15 [0.06,0.23]
α_i	T. rule smooth.	Beta	0.50 (0.285)	0.74 [0.66,0.82]	0.77 [0.70,0.84]	0.75 [0.68,0.82]	0.78 [0.72,0.84]	0.77 [0.71,0.84]
ρ_a	Tech. shock pers.	Gamma	0.90 (0.05)	0.97 [0.95,0.99]	0.97 [0.95,0.99]	0.97 [0.95,0.99]	0.97 [0.95,0.99]	0.97 [0.95,0.99]
ρ_m	M. pol. shock pers.	Beta	0.50 (0.15)	0.43 [0.32,0.54]	0.43 [0.32,0.54]	0.42 [0.31,0.53]	0.43 [0.33,0.54]	0.42 [0.31,0.53]
ρ_g	IS shock pers.	Gamma	0.90 (0.05)	0.92 [0.87,0.96]	0.90 [0.85,0.95]	0.91 [0.87,0.96]	0.89 [0.85,0.94]	0.90 [0.85,0.94]
σ_a	Tech. shock std.	IGamma	0.005 (2.00)	0.0080 [0.0064,0.0094]	0.0088 [0.0071,0.0106]	0.0084 [0.0067,0.0100]	0.0089 [0.0073,0.0106]	0.0089 [0.0072,0.0108]
σ_m	M. pol. shock std.	IGamma	0.005 (2.00)	0.0021 [0.0015,0.0026]	0.0019 [0.0014,0.0023]	0.0020 [0.0015,0.0025]	0.0018 [0.0014,0.0022]	0.0019 [0.0014,0.0023]
σ_g	IS shock std.	IGamma	0.005 (2.00)	0.0009 [0.0007,0.0011]	0.0009 [0.0007,0.0011]	0.0008 [0.0007,0.0010]	0.0009 [0.0007,0.0011]	0.0009 [0.0007,0.0011]
$\log(ML)$				–35.37	–30.45	–33.00	–28.90	–29.97

3.3. Prior densities

Table 1 reports the prior densities for the estimated parameters. We place a weakly informative beta distribution centered at 0.5 on the indexation parameter χ . Our choice is motivated by the large variety of estimates/calibrations in the literature, which range from the extreme calibration by Christiano et al. (2005), who adopt a unitary value, to the point estimate by Cogley and Sbordone (2008), who find support for a zero value. Moving to the remaining parameters, we adopt weakly informative priors as for the Calvo parameter θ , the interest rate smoothing parameter α_i , and the monetary policy shock autoregressive parameter ρ_m . Calibrations for the degree of risk aversion σ typically range between 1 and 5: we center our prior at 2.5, and we allow the 95 percent prior interval to range between 2 and 4. Consistent with the literature, the coefficients of the policy rule imply a substantial interest rate reaction to inflation and a positive reaction to output. Finally, our priors are centered to a high degree of persistence of the technology and demand shocks, in line with the evidence in Smets and Wouters (2007).

3.4. Posterior densities and model comparison

Table 1 collects our posterior estimates.²² The posterior means of the Calvo parameter and the degree of relative risk aversion are very standard. The estimated Taylor rule parameters suggest a strong long-run systematic reaction to the inflation gap—in line with recent estimates provided by Blanchard and Riggi (2009)—and a more moderate reactivity to output, both tempered in the short run by a fairly large amount of policy gradualism. As in previous studies, (e.g., Smets and Wouters, 2007), the estimated persistence of the technological shock is large.²³

In terms of model comparison, the marginal likelihood (computed with the modified harmonic mean estimator developed by Geweke, 1998) points towards the superiority of models taking trend inflation into account. The Bayes factor involving the baseline and the Calvo models (unrestricted) reads $\exp(4.92) \approx 137$, which suggests a “strong” support for

²² We estimated our models with Dynare, a set of algorithms developed by Michel Juillard and collaborators (Adjemian et al., 2010). Dynare is freely available at <http://www.dynare.org>.

²³ The data appear to be informative as for the estimated parameters. A figure contrasting our priors and posteriors, not reported here for the sake of brevity, may be found in the working paper version of this contribution, i.e., Ascari et al. (2010).

the trend inflation model.²⁴ Interestingly, the Rotemberg model is also supported by the marginal likelihood comparison, even if the wedge with the baseline New-Keynesian model is smaller.

Conditional on a positive trend inflation rate, one may also detect two important differences when comparing Calvo and Rotemberg. First, a comparison based on their power of fit speaks in favor of the Calvo model, with a log-difference that translates into a Bayes factor of about 12.81. Second, while both models point towards a degree of indexation clearly lower than 1 (the calibration suggested by e.g., [Christiano et al., 2005](#)), there is a clear difference in the estimated degree of indexation χ . The estimated posterior mean associated to the Calvo model clearly points towards a negligible value for price indexation (0.13). Moreover, the zero value belongs to the 90% credible set. By contrast, the Rotemberg model calls for a more than double posterior mean, 0.33, and it calls for a very high 95th percentile reading 0.62.

The theoretical justification for the introduction of indexation in a macroeconomic model is somewhat questionable. Moreover, as shown by [Benati \(2008, 2009\)](#) and [Cogley and Sbordone \(2008\)](#), such a parameter is hardly structural in the sense of Lucas, so that policy exercises conducted with models appealing to indexation may very well be misleading. Then, our posterior estimates point to the Calvo model as the more appealing from a “structural” standpoint.²⁵

Our findings (i) offer empirical support to models featuring trend inflation; (ii) point to the empirical superiority of the Calvo pricing scheme as opposed to the standard Rotemberg framework; (iii) suggest a lower degree of indexation to past inflation under Calvo. Interestingly, [Table 1](#) shows that this set of findings is confirmed by our exercises conditional on indexation to the steady state inflation rate, i.e., under $\mu = 0$. In particular, the estimated indexation parameters feature higher posterior means with respect to the previous scenario. Nevertheless, the value associated to the Calvo model is substantially lower than the one called for by the standard Rotemberg framework. Moreover, the 5th percentile of the posterior density of this parameter under Calvo hits the zero lower bound. Quite differently, the 95th percentile in the standard Rotemberg case hits the upper bound.

To be clear, our results are in line with those proposed in previous contributions as far as the estimated degree of price indexation is concerned. When referring to the Calvo framework, the estimated degree of price indexation we find is not different with respect to the point estimate obtained in [Smets and Wouters \(2007\)](#). [Cogley and Sbordone \(2008\)](#) obtain an even lower value of inflation indexation (basically zero) when allowing for drifting trend inflation, an ingredient which we do not consider in our analysis. We stress, however, that our analysis is not concerned with the estimation of the degree of price indexation per se. Differently, we are interested in assessing the relevance of trend inflation for different pricing schemes, with a particular attention in gauging the ability of the Calvo model to replicate the data as opposed to Rotemberg’s proposals.

4. Understanding the superior empirical performance of the Calvo model

The differences between Calvo and Rotemberg are fundamentally two: (i) the different order of the dynamics because of the presence of price dispersion \hat{s}_t in Calvo but not in Rotemberg; (ii) the different structure (“regressors”) in the NKPC and IS schedules of the two models. We will study the relevance of these ingredients in turn by contrasting some perturbations of the Calvo and standard Rotemberg models.²⁶

4.1. Exploring the mechanism

The presence of the price dispersion process \hat{s}_t in the Calvo model affects the persistence of inflation, in that price dispersion depends by its own past realizations as well as current and past realizations of inflation. Therefore, in principle, price dispersion should be of help to capture inflation dynamics and the autocovariance structure of the observables we aim to model in general. To isolate the role played by price dispersion, we estimate a “reduced-form” version of the Calvo model in which we set $\hat{s}_t = 0$ at all times.²⁷

Our results are reported in [Table 2](#). The first two columns report the estimates related to the baseline Calvo model, whose figures are replicated here to ease comparisons with alternative models, and those regarding the model with no-price dispersion, respectively. One can hardly notice a difference between the two set of estimates. Hence, we conclude that the difference between Calvo and the standard Rotemberg framework is not due to the presence of price dispersion in the former model.

²⁴ According to [Kass and Raftery \(1995\)](#), a Bayes factor between 1 and 3 is “not worth more than a bare mention”, between 3 and 20 suggests a “positive” evidence in favor of one of the two models, between 20 and 150 suggests a “strong” evidence against it, and larger than 150 “very strong” evidence.

²⁵ We also conducted pairwise model comparisons between unrestricted models and models estimated under the “no indexation” constraint $\chi = 0$. We obtain evidence “against” indexation as for the Calvo framework. Differently, the marginal likelihood of the standard Rotemberg framework typically deteriorates under such constraint. More information on these comparisons may be found in [Ascari et al. \(2010\)](#).

²⁶ A third difference between the two models regards the different non-linear impact of trend inflation on the convolutions of parameters characterizing the two systems. However, for low levels of trend inflation as the one realized during the Great Moderation, such difference is very unlikely to be an important driver of our results. For a quantitative assessment of the impact of trend inflation restrictions in the Calvo framework, which confirms that such impact is moderate in the sample at hand, see [Cogley and Sbordone \(2008\)](#).

²⁷ While being intuitive, this exercise represents a violation of the structural model, in that price dispersion under positive trend inflation is not a zero-process at a first-order level.

Table 2

Priors and posteriors for structural parameters. Models estimated assuming indexation to past inflation. The table reports the posterior means and the [5th,95th] percentiles. The log-marginal likelihoods are computed with the modified harmonic mean estimator by Geweke (1998). Details on the model estimation are reported in the text.

Param.	Description	Prior dens.	Prior mean (Std. dev.)	Posterior mean [5th pct, 95th pct]			
				Calvo	Calvo, no-price dispersion	Calvo, Std. Rotemberg IS	Rotemberg, costs in labor units
χ	Indexation	Beta	0.50 (0.285)	0.13 [0.00,0.28]	0.12 [0.00,0.26]	0.33 [0.01,0.62]	0.14 [0.00,0.31]
θ	Calvo	Beta	0.50 (0.15)	0.66 [0.57,0.76]	0.66 [0.57,0.75]	0.58 [0.47,0.70]	0.66 [0.56,0.76]
σ	Risk aversion	Normal	2.50 (0.25)	2.35 [1.92,2.76]	2.38 [1.98,2.79]	2.24 [1.84,2.63]	2.36 [1.91,2.75]
α_π	T. rule inflation	Normal	2.00 (0.50)	3.31 [2.74,3.86]	3.31 [2.74,3.88]	3.38 [2.86,3.97]	3.30 [2.76,3.89]
α_y	T. rule output	Gamma	0.125 (0.05)	0.15 [0.06,0.24]	0.15 [0.06,0.23]	0.14 [0.06,0.22]	0.15 [0.06,0.23]
α_i	T. rule smooth.	Beta	0.50 (0.285)	0.77 [0.70,0.84]	0.77 [0.71,0.84]	0.75 [0.68,0.82]	0.77 [0.70,0.84]
ρ_a	Tech. shock pers.	Gamma	0.90 (0.05)	0.97 [0.95,0.99]	0.97 [0.95,0.99]	0.97 [0.95,0.99]	0.97 [0.95,0.99]
ρ_m	M. pol. shock pers.	Beta	0.50 (0.15)	0.43 [0.32,0.54]	0.44 [0.33,0.55]	0.42 [0.31,0.53]	0.43 [0.33,0.55]
ρ_g	IS shock pers.	Gamma	0.90 (0.05)	0.90 [0.85,0.95]	0.90 [0.85,0.94]	0.91 [0.86,0.94]	0.90 [0.86,0.95]
σ_a	Tech. shock std.	IGamma	0.005 (2.00)	0.0088 [0.0071,0.0106]	0.0088 [0.0072,0.0104]	0.0083 [0.0068,0.0098]	0.0088 [0.0070,0.0104]
σ_m	M. pol. shock std.	IGamma	0.005 (2.00)	0.0019 [0.0014,0.0023]	0.0019 [0.0014,0.0023]	0.0020 [0.0016,0.0025]	0.0019 [0.0015,0.0023]
σ_g	IS shock std.	IGamma	0.005 (2.00)	0.0009 [0.0007,0.0011]	0.0009 [0.0007,0.0011]	0.0009 [0.0007,0.0010]	0.0009 [0.0007,0.0011]
$\log(ML)$				-30.45	-30.72	-32.97	-30.79

We then turn to the distinct structures of the two models. Recall that the two pricing schemes under scrutiny have different implications also as for the IS curves. This is due to the different implications on the relationship between consumption and output. We then implement an exercise to investigate if the difference between the two IS curves is responsible for the fit of the overall frameworks. In particular, we estimate a “Calvo NKPC—Std. Rotemberg IS” model, which is composed by Eqs. (11)–(13), (17), (19) and (20). If the structure of the IS curve is the reason behind the worse fit of the standard Rotemberg model and the higher degree of indexation associated to it, we should observe a deterioration of the marginal likelihood and a higher estimated degree of indexation.

We report our results in Table 2 under the label “Calvo, Std. Rotemberg IS”. Interestingly, the estimated indexation parameter for the “Calvo NKPC—Std. Rotemberg IS” turns out to be $\chi = 0.33$, i.e., the posterior mean of the indexation parameter is almost three times as large as that associated to the Calvo model. Moreover, the empirical fit deteriorates, with the marginal likelihood reading -32.97 . These findings suggest that the empirical superiority of the Calvo proposal is not necessarily due to its implications for the NKPC *per se*, but to the implications of the Rotemberg pricing for the demand side of the economy. Therefore, empirical studies dealing with different pricing mechanisms should adopt a full-system approach able to account for all the model equations, i.e., an investigation on the NKPCs *per se* is not necessarily exhaustive.

4.2. An alternative Rotemberg specification

We investigate the relevance of the demand side-specification further by deriving an alternative version of the Rotemberg model where the adjustment costs are measured in labor units. In other words, the Rotemberg adjustment costs now introduces a wedge between production and labor inputs, as in Calvo. In particular, the intermediate good-producing firms i faces the following production function:

$$Y_t(i) = A_t N_t(i) - \frac{\varphi_p^N}{2} \left(\frac{P_{i,t}}{(\pi_t^\chi)^\mu (\bar{\pi}^\chi)^{1-\mu} P_{i,t-1}} - 1 \right)^2 Y_t. \quad (22)$$

In this formulation, Rotemberg adjustment costs result in an inefficiency loss in aggregate production. It is possible to show that, in equilibrium, the log-linearized representation of the economy reads²⁸:

$$\hat{\pi}_t = \gamma_p \hat{\pi}_{t-1} + \gamma_f \beta \hat{\pi}_{t+1|t} + \gamma_{dy} \beta (1-\sigma) \Delta \hat{y}_{t+1|t} + \gamma_{mc} \widehat{mc}_t - \beta \gamma'_{mc} \widehat{mc}_{t+1}, \quad (23)$$

$$\widehat{mc}_t = (\sigma + \varphi) \hat{y}_t - (1 + \varphi) a_t + \frac{\varphi \Psi}{1 + \Psi} \varphi_p^N (\pi^{1-\chi} - 1) \pi^{1-\chi} [\hat{\pi}_t - \chi \mu \hat{\pi}_{t-1}], \quad (24)$$

$$\hat{y}_t = \hat{y}_{t+1|t} - \sigma^{-1} (\hat{i}_t - \hat{\pi}_{t+1|t}) + g_t, \quad (25)$$

where

$$\varphi_p^N = \frac{\varepsilon}{\varepsilon - 1} \varphi_p^Y,$$

$$\gamma_p = \frac{\chi \mu}{(1 + \beta \chi \mu)},$$

$$\gamma_f = \frac{1}{(1 + \beta \chi \mu)},$$

$$\gamma_{dy} = \frac{(\pi^{1-\chi} - 1)}{(1 + \beta \mu \chi)(2\pi^{1-\chi} - 1)},$$

$$\gamma_{mc} = \frac{(\varepsilon - \varphi_p^N (\pi^{1-\chi} - 1) \pi^{1-\chi})}{\varphi_p^N \pi^{1-\chi} (2\pi^{1-\chi} - 1)(1 + \beta \mu \chi)},$$

$$\gamma'_{mc} = \frac{(\pi^{1-\chi} - 1)}{(1 + \beta \mu \chi)(2\pi^{1-\chi} - 1)},$$

$$\Psi = \frac{\varphi_p^N}{2} (\pi^{1-\chi} - 1)^2.$$

Table 2 reports our estimates of this alternative Rotemberg framework, labeled as “Rotemberg, Costs in labor units”. Three results arise. First, this alternative Rotemberg pricing scheme fits the aggregate U.S. data better than the standard Rotemberg proposal. To our knowledge, this is the first empirical comparison involving these two alternative Rotemberg schemes in the literature. Second, the alternative Rotemberg model fits the data as good as the Calvo model. Third, the estimated degree of price indexation is comparable to the one associated to the Calvo model. This set of results confirm the conjecture we put forward in the previous subsection, i.e., a full-system approach is needed for a correct assessment of different pricing schemes.

5. Further investigations

In contrasting Calvo with the standard Rotemberg proposal, our empirical exercises support (i) models featuring trend inflation, (ii) the empirical superiority of the Calvo model, and (iii) the lower degree of indexation to past inflation called for by the Calvo model. We conduct further investigations along two relevant dimensions: (i) a different transformation of output as an observable; (ii) a policy rule featuring a reaction to output growth and two lags of the interest rate à la Coibion and Gorodnichenko (2011a,b). We consider each of these two scenarios as a departure with respect to our baseline analysis developed in Section 3.²⁹

5.1. Output growth

In our benchmark exercise, we apply the Hodrick-Prescott filter to isolate the cyclical component of output.³⁰ Data-prefiltering has the advantage of identifying a cyclical component of output which is robust to model misspecification. However, as argued by Canova (1998), Burnside (1998), and Gorodnichenko and Ng (2010) among others, filtering may also induce spurious results. An alternative to filtering is to use the growth rate of output as an observable. Then, we modify our model by introducing a deterministic trend in technology as for example in Mattesini and Nisticò (2010). Therefore, we assume:

$$Y_t = \mathcal{A}_t N_t = \gamma^t A_t N_t, \quad (26)$$

²⁸ For the sake of brevity, we confine the full derivation of this set of difference equations to an Appendix available upon request.

²⁹ We thank both referees for suggesting us the analysis in this section.

³⁰ Our result are also robust to an alternative detrending strategy involving the piecewise linear trend put forward by Perron and Wada (2009) (see footnote 34).

where

$$\ln A_t = (1 - \rho_a) \ln A + \rho_a \ln A_{t-1} + \varepsilon_{a,t} \quad (27)$$

generates stationary fluctuations of $A_t \equiv \gamma^t A_t$ around the trend process whose deterministic growth rate in steady state is γ . Consequently, the following relationship holds:

$$\gamma_{yt}^{obs} = \gamma + \Delta \hat{y}_t, \quad (28)$$

where γ_{yt}^{obs} is the growth rate of output, γ is the growth rate of output in steady state, and $\Delta \hat{y}_t$ is the first difference of the (log of the) stationarized output level. Notice that, given that our utility function is separable in consumption and leisure, we set $\sigma = 1$ to ensure the existence of a balance growth path (see King and Rebelo, 1999).³¹

Table 3 reports our estimates (columns identified with “Output gr. rate”). The marginal likelihoods of these two models are lower than in our baseline case. They are, however, higher than the one obtained with the version of the model featuring no role for trend inflation (model estimated under $\chi = 1$), which reads -43.61 .³² The estimated degree of price indexation in our Calvo model is higher than our baseline estimate, while the degree of price stickiness turns out to be lower. This set of estimates reveal that the empirical fit of the Calvo and the standard Rotemberg proposal is comparable. However, the Rotemberg model calls for a higher and more uncertain degree of price indexation, as in our baseline analysis. Overall, we can state that our output growth-based exercises do not overturn our main conclusions.

5.2. Taylor rule à la Coibion and Gorodnichenko (2011a,b)

Coibion and Gorodnichenko (2011a) estimate a Taylor rule with time-varying coefficients for the U.S. economy. They detect a tangible shift from a strong response to the output gap to a response to the output growth rate when entering the great moderation sample, which is the one we consider in our analysis. Moreover, they show that two lags of interest rate smoothing are needed to remove the serial correlation in the Taylor rule residual. Coibion and Gorodnichenko (2011b) support these results in the context of a horseshoe involving different distributed-lags policy rules. We then re-estimate our baseline models by replacing our policy rule (19) with the following formulation:

$$\hat{r}_t = \alpha_{i1} \hat{r}_{t-1} + \alpha_{i2} \hat{r}_{t-2} + (1 - \alpha_{i1} - \alpha_{i2})(\alpha_\pi \hat{\pi}_t + \alpha_{\Delta y} \Delta \hat{y}_t) + m_t, \quad (29)$$

which features a systematic reaction to the growth rate of output and an AR(2) interest rate smoothing structure.

Some results stand out (see Table 3). First, the marginal likelihood of both Calvo and Rotemberg under trend inflation is again higher than the one associated to the model with no effects deriving from trend inflation, which reads -18.04 .³³ As in the “output growth” case, the difference between the two models is not as large as the one recorded in our baseline analysis. However, the fit of the Calvo model is marginally superior than that of the Rotemberg framework, and the estimated degree of indexation lower. Again, this set of results mitigates, but it does not overturn, our main findings. Last but not least, our full-system estimates offer support to the single-equation estimates put forth by Coibion and Gorodnichenko (2011a,b), because this alternative formulation of the policy rule leads to a remarkably superior empirical performance for both the pricing schemes under investigation. This can be appreciated by contrasting the marginal likelihoods associated to the Calvo and Rotemberg models (under $\mu = 1$) in Table 1 with those reported in Table 3, which are substantially higher. Hence, our empirical exercises support Coibion and Gorodnichenko’s proposal to model the policy rate with an AR(2) process enriched with a systematic reaction to inflation and the growth rate of output (19).³⁴

6. Conclusions

This paper compares two New-Keynesian DSGE monetary models of the business cycle derived under different pricing schemes—Calvo (1983) and Rotemberg (1982)—and a positive trend inflation rate. We exploit the different reduced-form dynamics of the two models, derived in Ascari and Rossi (2009), to assess their relative empirical fit for the 1984:I–2008:II U.S. data.

³¹ Mattesini and Nisticò (2010) show that, in presence of non-separable, non-logarithmic preferences, a positive trend growth affects the dynamics of inflation and influences the design of the optimal policy reaction to cost-push shocks. In light of our empirical results and their theoretical derivations, it would be interesting to work out a model with non-separable-non-logarithmic preferences and contrast the empirical performance of the Calvo vs. Rotemberg settings. We leave this endeavor to future research.

³² The empirical estimates of the models estimated under $\chi = 1$ in this section are not presented for the sake of brevity, but are available upon request.

³³ The empirical estimates of the models estimated under zero trend inflation in this section are not presented for the sake of brevity, but are available upon request.

³⁴ We performed further robustness checks to verify the solidity of our results. With respect to the exercises documented in this section, we considered: (a) a more informative prior for the indexation parameter, i.e., $\chi \sim \text{Beta}(0.25, 0.10)$; (b) a different measure of inflation, i.e., the “inflation gap” computed by taking raw inflation in deviations with respect to its Hodrick–Prescott “trend”. This exercise is conducted to account for possible drifts in trend inflation in the sample at hand; (c) a different proxy of the business cycle, computed by modeling output’s low-frequency component with a piecewise quadratic trend with a break in 1973:I as suggested by Perron and Wada (2009); (d) alternative calibrations of the Frisch labor supply elasticity φ drawn from the interval [0.5, 1.5]. Our findings (i)–(iii) turn out to be robust to these perturbations. These robustness checks are reported in the working paper version (Ascari et al., 2010) and in an Appendix posted on our websites.

Table 3

Priors and posteriors for structural parameters: further investigations. Models estimated assuming indexation to past inflation. The table reports the posterior means and the [5th,95th] percentiles. The log-marginal likelihoods are computed with the modified harmonic mean estimator by Geweke (1998). Details on the model estimation are reported in the text.

Param.	Description	Prior dens.	Prior mean (Std. dev.)	Posterior mean [5th pct, 95th pct]			
				Calvo, output gr. rate	Rotemberg, output gr. rate	Calvo <i>TRgr</i> , <i>AR</i> (2) <i>irs</i>	Rotemberg, <i>TRgr</i> , <i>AR</i> (2) <i>irs</i>
χ	Indexation	Beta	0.50 (0.285)	0.27 [0.00,0.61]	0.35 [0.00,0.70]	0.20 [0.00,0.45]	0.25 [0.00,0.52]
θ	Calvo	Beta	0.50 (0.15)	0.47 [0.34,0.59]	0.45 [0.33,0.57]	0.50 [0.39,0.61]	0.48 [0.37,0.59]
σ	Risk aversion	Normal	2.50 (0.25)	1.00 [–]	1.00 [–]	2.19 [1.79,2.60]	2.17 [1.78,2.58]
α_π	T. rule inflation	Normal	2.00 (0.50)	3.48 [2.98,4.02]	3.53 [2.98,4.08]	3.18 [2.68,3.69]	3.19 [2.68,3.73]
α_y	T. rule output	Gamma	0.125 (0.05)	0.07 [0.03,0.12]	0.08 [0.03,0.12]	–	–
$\alpha_{\Delta y}$	T. rule output growth	Gamma	0.125 (0.05)	–	–	0.19 [0.08,0.28]	0.18 [0.07,0.27]
α_i	T. rule smooth.	Beta	0.50 (0.285)	0.64 [0.52,0.76]	0.64 [0.52,0.76]	–	–
α_{i1}	T. rule smooth.	Normal	1.10 (0.50)	–	–	1.45 [1.21,1.69]	1.46 [1.21,1.70]
α_{i2}	T. rule smooth	Normal	–0.20 (0.50)	–	–	–0.84 [–1.17,–0.53]	–0.86 [–1.17,–0.53]
ρ_a	Tech. shock pers.	Gamma	0.90 (0.05)	0.98 [0.96,0.99]	0.98 [0.96,0.99]	0.96 [0.94,0.99]	0.96 [0.94,0.99]
ρ_m	M. pol. shock pers.	Beta	0.50 (0.15)	0.35 [0.25,0.45]	0.34 [0.24,0.43]	0.37 [0.25,0.50]	0.37 [0.25,0.49]
ρ_g	IS shock pers.	Gamma	0.90 (0.05)	0.95 [0.91,0.98]	0.95 [0.91,0.98]	0.90 [0.85,0.96]	0.91 [0.86,0.96]
σ_a	Tech. shock std.	IGamma	0.005 (2.00)	0.0057 [0.0050,0.0064]	0.0057 [0.0050,0.0064]	0.0078 [0.0064,0.0091]	0.0077 [0.0063,0.0090]
σ_m	M. pol. shock std.	IGamma	0.005 (2.00)	0.0028 [0.0020,0.0037]	0.0029 [0.0020,0.0038]	0.0028 [0.0016,0.0038]	0.0029 [0.0018,0.0040]
σ_g	IS shock std.	IGamma	0.005 (2.00)	0.0013 [0.0011,0.0015]	0.0013 [0.0011,0.0016]	0.0010 [0.0008,0.0012]	0.0010 [0.0008,0.0012]
$\log(ML)$				–41.40	–41.98	–14.23	–15.12

Several findings arise. First, we find empirical support in favor of models that incorporates trend inflation as opposed to models derived under the commonly used “zero inflation in steady state” assumption. Second, the data support the Calvo model as empirically superior with respect to the standard Rotemberg pricing scheme. This superiority is suggested both by comparisons based on marginal likelihoods and by the economic plausibility of the estimates we obtain. In particular, we find the estimated degree of indexation in the Calvo model to be close to zero, in line with the results in Smets and Wouters (2007) and Cogley and Sbordone (2008). This is a *plus* of the Calvo model with trend inflation, in that price indexation is theoretically questionable and empirically at odds with micro-data evidence (Bils and Klenow, 2004; Nakamura and Steinsson, 2008). Interestingly, Calvo’s superior empirical performance is related to the different cross-equation restrictions affecting the demand side of the economy with respect to the ones implied by the standard Rotemberg price setting framework. This result stresses the importance of conducting empirical investigations with a fully structural system, rather than only with a NKPC generated from a particular price setting mechanism.

Overall, this paper offers support to the Calvo pricing scheme for the modeling of inflation dynamics from a macroeconomic perspective. Admittedly, the Calvo-parameter is hardly structural, and the policy implications stemming from the Calvo-world should be carefully assessed. While offering some empirical support to the Calvo-mechanism conditional on their micro-data analysis, Costain and Nakov (2008) call for further explorations of state-dependent pricing models, which can potentially provide policymakers with more reliable policy suggestions. We are sympathetic with this call, and welcome contributions engaging in the design of more realistic pricing schemes.

Acknowledgments

We thank Christopher Otrok (Editor) and two anonymous referees for their very insightful comments and suggestions. We also thank Gianni Amisano, Paul Beaudry, Martin Ellison, Jeff Fuhrer, Francesco Furlanetto, Fabià Gumbau-Brisa,

Martin Eichenbaum, Giovanni Lombardo, Tommaso Monacelli, Giovanni Olivei, Luigi Paciello, Giorgio Primiceri, Antti Ripatti, Tiziano Ropele, Tommy Sveen, Jouko Vilmunen, and participants at the IV PIERO MONCASCASCA Money-Macro Workshop (Bocconi University), the II Economic Policy and the Business Cycle Workshop (Milano Bicocca), the Monetary and Fiscal Policy for Macroeconomic Stability (Pavia), and at seminars held at Oxford University, EIEF, Goethe University, and Norges Bank for valuable feedbacks. Part of this research was carried out while the second author was visiting the Boston University Department of Economics, whose kind hospitality is gratefully acknowledged. The opinions expressed in this paper do not necessarily reflect those of the Bank of Finland. Financial support from MIUR, Grant PRIN 2007P8MJ7P is gratefully acknowledged.

Appendix A. Supplementary data

Supplementary data associated with this article can be found in the online version at doi:10.1016/j.jedc.2011.06.002.

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