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# The New Keynesian Phillips Curve in Europe: does it fit or does it fail?


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**Abstract:** The canonical New Keynesian model specifies inflation as the present-value of future real marginal cost. This paper tests this New Keynesian Phillips Curve and exploits projections of future real marginal cost generated by VAR models to assess the model's ability to match the behavior of actual inflation. In accordance to the literature, the model fits Euro data well at first sight. However, analyses of this kind disregard the considerable degree of uncertainty surrounding VAR forecasts. A set of bias-corrected bootstrapped confidence bands reveals that this result is consistent with both a well fitting and a completely failing model. Allowing for inflation inertia through backward-looking indexation narrows confidence bands around measures of the model's fit but, still, cannot generate sufficiently precise estimates. Hence, we cannot say whether the model fits or fails.

**Keywords:** New Keynesian Phillips Curve, present-value model, marginal cost, VAR, bootstrap

**JEL classification:** E31, E32

## Non technical summary

Sticky-price models with monopolistic competition have become the canonical framework to study inflation and monetary policy. Under this New Keynesian paradigm, inflation dynamics are forward-looking. The New Keynesian Phillips Curve (NKPC) relates current inflation to expected future inflation and a measure of current real activity. Moreover, it can be shown that the inflation rate is given as the present-value of the entire expected path of future real marginal cost.

This present-value relation implicit in models of staggered price setting is the central topic of this paper as it lends itself to a well-established empirical approach. The approach to assess the model's empirical fit is similar to empirical studies of the intertemporal model of the current account or the relation between stock prices and future dividends. Specifically, we can employ VAR based forecasts to generate a series of model-consistent or "fundamental" inflation that is supposed to match the behavior of actual inflation if the model is correct. It is frequently argued that fundamental inflation explains actual inflation quite well. In this paper we shed light on this finding using data for the Euro area. In particular, we use bootstrapped confidence bands to quantify the degree of estimation uncertainty around these estimates. Huge confidence bands preclude any meaningful interpretation of conventionally employed measures of fit. We show that the result of the forward-looking model cannot be interpreted as it is done in the literature due to immensely wide confidence intervals. The baseline specification is consistent with both a completely failing model where the correlation coefficient between actual and fundamental inflation is negative and, at the same time, with a remarkably well fitting model where actual and fundamental inflation exhibit an almost perfect positive correlation.

Supplementing the model with backward-looking inflation and, thus, allowing for inflation inertia substantially improves the model's fit and narrows confidence bands around the correlation coefficient. Nevertheless, while the point estimates can replicate actual inflation quite well, we are left with wide confidence bands around the relative volatility of fundamental and actual inflation. Hence, a large degree of uncertainty remains that impedes a reasonable interpretation of the model's empirical performance.

# Nicht technische Zusammenfassung

Makroökonomische Modelle mit nominalen Rigiditäten und monopolistischer Konkurrenz sind mittlerweile zum Referenzrahmen für die Analyse der Inflationsentwicklung und geldpolitischer Fragestellungen geworden. Derartige neu-keynesianische Modelle implizieren, dass die Inflationsdynamik vorausschauend ist. Die Inflationsrate ist also von gegenwärtigen realen Variablen und von der erwarteten Inflationsrate der nächsten Periode abhängig. Weiterhin kann gezeigt werden, dass die Inflationsrate dem Gegenwartswert der zukünftigen erwarteten realen marginalen Kosten entspricht.

Diese Gegenwartswertbeziehung, die von Modellen mit gestaffeltem Preissetzungsverhalten impliziert wird, soll in dieser Arbeit empirisch analysiert werden. Der Zusammenhang zwischen gegenwärtiger Inflation und zukünftigen marginalen Kosten ist analog zu anderen makroökonomischen Modellen, bspw. dem intertemporalen Modell der Leistungsbilanz oder der Gegenwartswertbeziehung zwischen Aktienkursen und zukünftigen Dividendenzahlungen und kann in einem ähnlichen empirischen Ansatz untersucht werden. Dieser empirische Ansatz leitet aus einem vektorautoregressiven Modell eine theoriekonsistente oder "fundamentale" Inflationsrate ab, die der tatsächlich beobachteten Inflationsrate entspricht, sofern das Modell korrekt ist. Jüngste Beiträge zu dieser Forschungsrichtung interpretieren die Ergebnisse dieses Ansatzes im Sinne einer Bestätigung des zugrunde liegenden theoretischen Modells.

In dieser Arbeit sollen diese Ergebnisse im Hinblick auf die Erklärung der Inflationsentwicklung im Euro-Raum hinterfragt werden. Zu diesem Zweck soll vor allem der Schätzunsicherheit Rechnung getragen werden. Die Konfidenzbänder, die die Streuung der errechneten Kriterien zur Messung der Erklärungskraft des Modells beschreiben, zeigen, dass die empirischen Ergebnisse nicht eindeutig interpretiert werden können. Die Ergebnisse des vorausschauenden Modells können aufgrund der weiten Konfidenzbänder nicht so interpretiert werden, wie es in der Literatur geschieht. Die zentrale Spezifikation des Modells ist vielmehr konsistent mit einem vollständig versagenden theoretischen Modell, das eine negative Korrelation zwischen tatsächlicher und fundamentaler Inflationsrate aufweist und einem bemerkenswert guten Modell, in dem die Inflationsraten perfekt miteinander korreliert sind.

Eine Ergänzung des Modells um die verzögerte Inflationsrate, also die Berücksichtigung der empirisch zu beobachtenden Trägheit der Inflationsentwicklung, führt zu einer deutlichen Verengung der Konfidenzbänder um die errechnete Korrelation und erhöht somit die Erklärungskraft des Modells. Die relative Volatilität der beiden Inflationsraten hingegen kann weiterhin nicht hinreichend genau abgebildet werden.

# Contents

## List of Tables

# List of Figures

The New Keynesian Phillips Curve in Europe:
does it fit or does it fail?[1]

# 1 Introduction

Sticky-price models with monopolistic competition have become the canonical framework to study inflation and monetary policy. Despite the diversity of assumptions about the specific sources of nominal rigidity within this field, most approaches share a common building block. Under the New Keynesian paradigm, this common element claims that inflation dynamics are to a certain extent forward-looking. Hence, the workhorse New Keynesian Phillips Curve (NKPC) relates current inflation to expected future inflation and a measure of current real activity. Moreover, it can be shown that the inflation rate is given as the present-value of the entire expected path of future real marginal cost.

This present-value relation implicit in any off-the-shelf New Keynesian model is the central topic of this paper as it lends itself to a well-established empirical approach. Like other present-value relations, e.g. the intertemporal model of the current account or the relation between stock prices and future dividends, this model can straightforwardly be assessed using the seminal framework laid out by Campbell and Shiller (1987). The advantage of this empirical approach is that it circumvents controversial issues involved in standard GMM estimates of the NKPC, i.e. small-sample problems and the choice of appropriate instruments.[2]

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[2]See, among others, Mavroeidis (2004, p. 632): "The existing empirical analyses of such [forward looking] models should be treated with caution."

Specifically, we can employ VAR based forecasts to generate a series of model-consistent or "fundamental" (Galí, Gertler, and López-Salido 2001, henceforth GGL) inflation that is supposed to match the behavior of actual inflation if the model is correct. Prominent contributions that exploit the present-value structure for U.S. data are Galí and Gertler (1999), Sbordone (2002, 2004), and Kurmann (2003).[3] In their influential paper, GGL (2001) derive a series of fundamental inflation of the Euro area and argue that "fundamental inflation tracks the behavior of actual inflation quite well" (p. 1260). And in their abstract they argue that "the NKPC fits Euro data very well, possibly better than U.S. data".

In this paper we assess the empirical fit of the present-value relation implied by the Calvo price setting scheme, critically assess the validity of the aforementioned results of the literature and contribute to the literature in five respects:

First, we follow Kurmann (2003) and take account of estimation uncertainty. Since forecasts derived from VAR estimates are mere point estimates, plotting the implied inflation rate disguises the uncertainty involved in the estimation process. Hence, we assess whether the model indeed fits or whether it poorly fails.

Second, while bootstrapping confidence bands for major measures of the model's fit, we correct the bias due to the nonlinear nature of conventionally used measures of fit by employing Kilian's (1998) bias correction.

Third, we do not only provide evidence on the pure forward-looking version of the model, but also estimate a backward-looking model in which lagged inflation enters the Phillips Curve through indexation following, among others, Christiano, Eichenbaum, and Evans (2005).

Fourth, we use the latest set of data available for the aggregate Euro area. The benchmark study by GGL (2001) uses only pre-EMU data.

Fifth, we estimate parameters and provide evidence on the duration of Calvo contracts.

In contrast to GGL (2001) we find that the forward-looking NKPC fits Euro data even worse than U.S. data. Huge confidence bands preclude any meaningful interpretation of conventionally employed measures of fit. Supplementing

[3]GGL (2001) and Jondeau and Le Bihan (2001, 2003) are the main contributions for evidence on the New Keynesian Phillips Curve with European data.

the model with backward-looking elements substantially improves the model's fit and narrows confidence bands around the correlation coefficient. While the point estimates can replicate actual inflation quite well, we are left with wide confidence bands around the relative volatility of actual and fundamental inflation. Nevertheless, a large degree of uncertainty remains that impedes a reasonable interpretation of the model's adequacy.

The present paper is organized as follows. The next section derives the New Keynesian Phillips Curve and the present-value relation for inflation from a standard model of staggered price setting. Section three presents the estimation strategy, discusses estimation uncertainty and elaborates the bootstrap approach to calculate confidence intervals around standard measures of fit. Section 4 presents the results and, finally, section 5 concludes.

# 2 The New Keynesian model of inflation

Models with staggered price setting and monopolistic competition are frequently referred to as New Keynesian models.[4] In this section we use a stylized log-linear model to derive the basic present-value relation for inflation that is central to most specifications of the New Keynesian Phillips Curve.

## 2.1 A stylized model of staggered prices

Under imperfect competition, firms' price setting behavior is driven by the behavior of their marginal cost of production. This implies that the aggregate price level and, hence, the overall inflation rate are determined by individual firms' marginal cost.

Consider the case of staggered price setting following the seminal work of Calvo (1983).[5] All variables are in logs. Each firm adjusts its price during the current period with a fixed probability $1 - \mu$, where $0 < \mu < 1$. With a probability $\mu$

[4] See Woodford (2003) for a systematic and profound overview.

[5] We concentrate here on Calvo-style price setting behavior. Roberts (1995) shows that fixed length contracts proposed by Taylor (1980) result in similar inflation dynamics and Sbordone (2002) shows in her appendix that both models of price setting imply a similar common trend retrictions.

the price is kept fixed. Firms minimize the discounted future deviations of their price from the price they would set if prices were fully flexible. It can be shown that this optimizing problem results in an optimal reset price $p_t^*$ given by

$$p_t^* = (1-\phi\mu)\sum_{k=0}^{\infty}(\phi\mu)^k E_t\{nmc_{t+k}\} \tag{1}$$

with a subjective discount factor $\phi$. The optimal reset price is set equal to a weighted average of the prices that it would have expected to set in the future if there weren't any price rigidities. In a frictionless market this price would equal a fixed markup over marginal cost. For simplicity, the markup is set to zero in the theoretical considerations. In setting prices at time $t$, each firm takes the expected path of future nominal marginal cost, $nmc_t$, into account.
The price level $p_t$ is then given as a convex combination of the lagged price level $p_{t-1}$ and the optimal reset price $p_t^*$

$$p_t = \mu p_{t-1} + (1-\mu)p_t^* \tag{2}$$

Combining these two equations gives the aggregate price level as the present-value of expected future nominal marginal cost

$$p_t = \mu p_{t-1} + (1-\mu)(1-\phi\mu)\sum_{k=0}^{\infty}(\phi\mu)^k E_t\{nmc_{t+k}\} \tag{3}$$

The higher the probability $\mu$, the more persistent is the price level. In the limiting case of perfectly flexible prices (i.e. $\mu \to 0$), the optimal reset price and, thus, the price level are determined only by the current level of marginal cost, $p_t = nmc_t$. The higher the probability is that prices remain unchanged, the more important the forward looking element.

## 2.2 The forward-looking Phillips Curve

The New Keynesian Phillips Curve (NKPC) can be derived from the model presented in the previous section (see, e.g. Galí and Gertler, 1999, and GGL, 2001). Inflation is determined by expected future inflation and current real activity proxied by real marginal cost, where $\pi_t = P_t - P_{t-1}$ is the inflation rate, $rmc_t$ denotes a measure of real marginal cost and $E_t$ is the expectations operator

$$\pi_t = \phi E_t \pi_{t+1} + \gamma rmc_t \tag{4}$$

The composite parameter $\gamma$ is given by $\frac{(1-\mu)(1-\phi\mu)}{\mu}$. Repeated substitution then yields

$$\pi_t = \gamma \sum_{k=0}^{\infty} \phi^k E_t rmc_{t+k} \tag{5}$$

Equation (5) says that the inflation rate at time $t$ is a fraction of the present-value of the expected path of future real marginal cost. We will later follow the literature and proxy real marginal cost with the labor share of income.

## 2.3 The role of backward-looking indexation

The Phillips Curve equation derived in the previous section has been frequently criticized for a lack of inflation inertia that is present in U.S. and other countries' data (see, e.g. Fuhrer and Moore, 1995). To capture inflation persistence, we modify the model following Christiano, Eichenbaum, and Evans (2005), Sbordone (2004), Smets and Wouters (2003), and Walsh (2004) by assuming that firms that are not selected to reset prices according to the Calvo price setting scheme are allowed to index their price to past inflation.[6] All beforementioned authors argue that partial or full indexation of the Calvo model improves the empirical fit of their models. While equally ad-hoc, partial price indexation appears to be a more reasonable assumption to motivate the relevance of lagged inflation than to resort to rule-of-thumb consumers as in the hybrid specification of GGL (2001).

Let the degree of indexation be denoted by $\kappa$. The aggregate price level is then given by

$$p_t = \mu\left(p_{t-1} + \kappa\pi_{t-1}\right) + (1-\mu)p_t^* \tag{6}$$

Christiano, Eichenbaum, and Evans (2005) and Eichenbaum and Fisher (2004) assume an indexation parameter $\kappa = 1$, while Woodford (2003) only requires $0 \leq \kappa \leq 1$. Even when setting $\kappa = 1$ we have one quarter lag-dynamics in the inflation rate. The Phillips Curve in the presence of price indexation, see

[6] Woodford (2003, p. 214) argues that "it is far more plausible, then, to imagine a policy of automatic indexation of one's price (between the occasions on which a full review of the optimality of the price is undertaken) to the change in an overall price index over some *past* time interval."

Woodford (2003, p. 215), becomes

$$\pi_t = \kappa\pi_{t-1} + \phi\left(E_t\pi_{t+1} - \kappa\pi_t\right) + \gamma rmc_t \tag{7}$$

or

$$\pi_t = \frac{\kappa}{1+\phi\kappa}\pi_{t-1} + \frac{\phi}{1+\phi\kappa}E_t\pi_{t+1} + \frac{\gamma}{1+\phi\kappa}rmc_t \tag{8}$$

which has the same form as the hybrid expression proposed by GGL (2001). Solving this equation forward yields a NKPC with backward-looking elements as the discounted stream of expected real marginal cost plus a proportion of lagged inflation

$$\pi_t = \kappa\pi_{t-1} + \frac{(1-\mu)(1-\phi\mu)}{\mu}\sum_{k=0}^{\infty}\phi^k E_t rmc_{t+k} \tag{9}$$

that nests the purely forward-looking model when $\kappa$ is set to zero. We will later discuss appropriate values of $\kappa$.[7]

# 3 The present-value relation under estimation uncertainty

Campbell and Shiller (1987, 1988) propose a well-known framework to assess the fit of forward-looking present-value models.[8] These considerations, that are originally developed for term structure applications, are easily transferred to the present-value relation specified by the New Keynesian model of inflation. As a clear advantage, this approach does not involve making assumption about the structure of the whole economy in the application of maximum likelihood methods or the choice of appropriate instruments in an instrumental variables estimation. In a first step, we derive the cointegration restriction implied by the forward-looking model. In a second step, we present the estimation strategy based upon VAR projections as a proxy for market expectations.

[7]Similar NKPC specifications with backward-looking indexation can be found in Dotsey (2002) and Rudd and Whelan (2003) for the case of U.S. inflation.

[8]See Engsted (2002) for an extensive survey of various techniques and applications.

## 3.1 The common trends restriction

To derive a clear-cut testable restriction, we undertake some algebraic steps that are well known from the literature on present-value models. In particular, subtract $nmc_t$ from (3) and recognize that

$$(1-\phi\mu)\left(\sum_{k=0}^{\infty}(\phi\mu)^k E_t\{nmc_{t+k}\}\right)-nmc_t=\sum_{i=1}^{\infty}(\phi\mu)^i E_t\{\Delta nmc_{t+i}\} \qquad (10)$$

with the difference operator given by $\Delta$. After some steps we obtain the following expression for a given indexation parameter $\kappa$

$$rmc_t = nmc_t - p_t = \left(\frac{\mu}{1-\mu}\right)(\Delta p_t - \kappa\Delta p_{t-1}) - \sum_{i=1}^{\infty}(\phi\mu)^i E_t\{\Delta nmc_{t+i}\} \quad (11)$$

Equation (11) specifies real marginal cost, $rmc_t$, as the present-value of the future path of changes in nominal marginal cost. This expression imposes a crucial restriction on the joint dynamics of the price level and the level of nominal marginal cost. If $nmc_t$ and $p_t$ are nonstationary, $I(1)$, their first differences must by definition be stationary, $I(0)$. Thus, equation (11) says that the linear combination $nmc_t - p_t$ must be stationary. Assume a linear combination $\boldsymbol{\beta}'\mathbf{x}_t$ with a $(1 \times 2)$ vector $\boldsymbol{\beta}$ and the data vector $\mathbf{x}_t' = (nmc_t, p_t)$. The testable implication of the Calvo model of inflation is that $\boldsymbol{\beta}' = (1, -1)$. In other words, nominal marginal cost and the price level are cointegrated or share a common trend, respectively. Thus, real marginal cost must be stationary.

## 3.2 Inflation forecasts from VAR projections

To assess the explanatory power of the Calvo model of staggered price setting, we construct an implied series for the forward-looking terms and contrast model-consistent inflation rates with actually observed inflation rates. As mentioned before, this approach is identical to GGL (2001), Sbordone (2002, 2004), Kurmann (2003) and others. We assume that the information contained in a small atheoretical bivariate VAR is a subset of the market's full information set.[9]
Let the information set of agents be described by past realizations of inflation and real marginal cost. The vector $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \pi_t, ..., \pi_{t-q+1}]'$ follows a

[9] An early contribution to this empirical strategy is Sargent (1979).

VAR($q$) in companion form

$$\mathbf{Z}_{t+1} = \mathbf{A}\mathbf{Z}_t + \mathbf{\Gamma}_{Z_{t+1}} \tag{12}$$

where $\mathbf{\Gamma}_{Z_{t+1}} = [u_{1t,}0, ..., 0, u_{2t}, 0, ...0]'$ represent innovations to agents' information sets and $\mathbf{A}$ is the $2q \times 2q$ matrix

$$\mathbf{A} = \begin{bmatrix} a_{111} & a_{112} & \cdots & a_{11q-1} & a_{11q} & a_{121} & a_{122} & \cdots & a_{12q-1} & a_{12q} \\ 1 & 0 & \cdots & 0 & 0 & 0 & 0 & \cdots & 0 & 0 \\ \vdots & \vdots & & \vdots & \vdots & \vdots & \vdots & & \vdots & \vdots \\ 0 & 0 & \cdots & 1 & 0 & 0 & 0 & \cdots & 0 & 0 \\ a_{211} & a_{212} & \cdots & a_{21q-1} & a_{21q} & a_{221} & a_{222} & \cdots & a_{22q-1} & a_{22q} \\ 0 & 0 & \cdots & 0 & 0 & 1 & 0 & \cdots & 0 & 0 \\ \vdots & \vdots & & \vdots & \vdots & \vdots & \vdots & & \vdots & \vdots \\ 0 & 0 & \cdots & 0 & 0 & 0 & 0 & \cdots & 1 & 0 \end{bmatrix} \tag{13}$$

We will later check for the robustness of the results and will use the alternative forecasting VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \Delta nmc_t, ..., \Delta nmc_{t-q+1}]'$, i.e. a VAR with current and lagged realizations of the level of real marginal cost and changes of nominal marginal cost. Forecasts based on the econometrician's information set $\mathbf{H}_t$, which includes only current and lagged values of the variables in $\mathbf{Z}_t$, are given by the multi-period forecasting formula

$$E_t \left[\mathbf{Z}_{t+k} | \mathbf{H}_t\right] = \mathbf{A}^k \mathbf{Z}_t \tag{14}$$

The vector of the discounted future paths of the variables can be calculated using the summation formula for infinite geometric series

$$\begin{aligned} \sum_{k=0}^{\infty} \phi^k E_t \mathbf{Z}_{t+k} &= \left(\mathbf{I} + \phi\mathbf{A} + \phi^2\mathbf{A}^2 + ...\right) \mathbf{Z}_t \\ &= \left(\mathbf{I} - \phi\mathbf{A}\right)^{-1} \mathbf{Z}_t \end{aligned} \tag{15}$$

We map these forecasts into the present-value representation of the Calvo pricing model to obtain an expression for the model-consistent inflation rate. This theoretical or "fundamental" (Galí and Gertler 1999, p. 217) inflation rate is given by

$$\begin{aligned}\pi_t^{fund} &= \gamma \sum_{i=1}^{\infty} \phi^i E_t \left\{ rmc_{t+i} \right\} \\ &= \gamma \sum_{i=1}^{\infty} \mathbf{h}'_{rmc} \phi^i \mathbf{A}^i \mathbf{Z}_t \\ &= \gamma \mathbf{h}'_{rmc} \left( \mathbf{I} - \phi \mathbf{A} \right)^{-1} \mathbf{Z}_t \end{aligned} \tag{16}$$

where $\mathbf{h}'_{rmc}$ denotes a selection vector that singles out the forecast of real marginal cost, i.e. the first element of $\left( \mathbf{I} - \phi \mathbf{A} \right)^{-1} \mathbf{Z}_t$. The NKPC thus predicts that inflation at time $t$ should be a scalar multiple of the first entry in the vector $\left( \mathbf{I} - \phi \mathbf{A} \right)^{-1} \mathbf{Z}_t$, which is currently observable. This equation is central to the empirical approach pursued in this paper. We will assess the fit of the Calvo model by comparing actual inflation $\pi_t$ with fundamental inflation $\pi_t^{fund}$. If the model provides an accurate description of European inflation dynamics, these two series must closely coincide.
Rudd and Whelan (2003) propose to infer the slope coefficient $\gamma$ from an OLS regression of actual inflation on the present-value of future real marginal cost $\mathbf{h}'_{rmc} \left( \mathbf{I} - \phi \mathbf{A} \right)^{-1} \mathbf{Z}_t$ and a constant.
We plot actual inflation against fundamental inflation and compute standard measures of fit. Following the literature on present-value models, Kurmann (2003) proposes two measures that indicate the extent to which the model is able to replicate actual inflation rates.[10] The first measure is the ratio of standard deviations

$$\frac{std.dev.\left(\pi_t^{fund}\right)}{std.dev.\left(\pi_t\right)} = \gamma \sqrt{\frac{\mathbf{h}_{rmc} \mathbf{M} \boldsymbol{\Sigma} \mathbf{M}' \mathbf{h}'_{rmc}}{\mathbf{h}_{\pi} \boldsymbol{\Sigma} \mathbf{h}'_{\pi}}} \tag{17}$$

where $\mathbf{M} = \left[ \mathbf{I} - \phi \mathbf{A} \right]^{-1}$, $\boldsymbol{\Sigma} = E\left[ \mathbf{Z}_t \mathbf{Z}'_t \right]$, and $\pi_t = \mathbf{h}'_{\pi} \mathbf{Z}_t$ where $\mathbf{h}_{\pi}$ is an appropriate selection vector. A perfect fit would result in a standard deviation ratio of unity. In that case the New Keynesian model would explain all the variation in the

[10] Campbell and Shiller (1987) argue in favor of a graphical comparison of actual and fundamental inflation since "tests of predictability ... are highly sensitive to deviations from the ... theory - so sensitive, in fact, that they may obscure some of the merits of the theory" (p. 1080).

actual inflation rate. Due to that fact that $\hat{\gamma}$ is on OLS regressor, the ratio of standard deviations is (in the absence of small sample problems) equal to the correlation coefficient $corr(\pi_t^{fund}, \pi_t)$ and therefore bounded by unity. Under the bootstrap exercise presented below, however, the ratio can exceed unity since we keep $\hat{\gamma}$ fixed when computing this ratio for artificially created series of the present value of real marginal cost. The second measure is the correlation coefficient between fundamental and actual inflation

$$\begin{aligned} corr(\pi_t^{fund}, \pi_t) &= \frac{E\left[\pi_t \pi_t^{fund}\right]}{\sqrt{E\left[\pi_t^2\right] E\left[\left(\pi_t^{fund}\right)^2\right]}} \\ &= \frac{\gamma \mathbf{h}_{rmc} \mathbf{M} \mathbf{\Sigma} \mathbf{h}_\pi'}{\gamma \sqrt{\mathbf{h}_{rmc} \mathbf{M} \mathbf{\Sigma} \mathbf{M}' \mathbf{h}_{rmc}' \mathbf{h}_\pi \mathbf{\Sigma} \mathbf{h}_\pi'}} \\ &= \frac{\mathbf{h}_{rmc} \mathbf{M} \mathbf{\Sigma} \mathbf{h}_\pi'}{\sqrt{\mathbf{h}_{rmc} \mathbf{M} \mathbf{\Sigma} \mathbf{M}' \mathbf{h}_{rmc}' \mathbf{h}_\pi \mathbf{\Sigma} \mathbf{h}_\pi'}} \end{aligned} \tag{18}$$

It is important to note that $\gamma$ cancels out. Hence, this measure of fit is independent of the estimated composite parameter $\gamma$ that to some extent reflects the degree of price rigidity specified by Calvo contracts.[11]

Note that these measures of fit do not reflect the degree of uncertainty about the model's fit. Previous applications of the empirical approach to the New Keynesian model of inflation dynamics sketched above , i.e. GGL (2001) and many others, neglect this issue at all. In fact, hardly any of the present-value applications surveyed in Engsted (2002) take account of estimation uncertainty apart from performing standard Wald tests. On the contrary, Kurmann (2003) provides evidence on the uncertain fit of the NKPC for U.S. data. In this paper we follow his approach and compute confidence bands around the measures of fit. We will return to that issue after we discussed the estimation of the backward-looking model in the following section.

[11] Both measure are widely used in the literature on present-value relations. See, e.g., Ghosh (1995) for an application to assess the fit of the intertemporal model of the current account.

### 3.3 The case of partial indexation

For the case of backward-looking price indexation lagged inflation enters the model-consistent inflation rate. Hence, fundamental inflation is now given by

$$\pi_t^{fund} = \kappa\pi_{t-1} + \gamma \mathbf{h}'_{rmc} \left(\mathbf{I} - \phi\mathbf{A}\right)^{-1} \mathbf{Z}_t \tag{19}$$

with a prespecified $\kappa$. To calibrate $\kappa$, we refer to recent general equilibrium models for the Euro area economy. We follow the results of the benchmark general equilibrium model developed and estimated by Smets and Wouters (2003) and set $\kappa = 0.46$. Onatski and Williams (2004) re-estimate the Smets-Wouters model with their own set of Bayesian priors. They obtain a degree of indexation of $\kappa = 0.32$, which we also include as an alternative specifications. Adolfson et al. (2004) estimate an open-economy variant of the Smets-Wouters model and find $\kappa = 0.23$. The equation for inflation dynamics collapses to the purely forward-looking NKPC once we set the degree of indexation to zero.[12]

### 3.4 Bootstrapping confidence bands for measures of fit

The crucial motivation of the empirical analysis in this paper is the fact that the series of fundamental inflation is merely a point estimate that disguises the degree of estimation uncertainty. Nevertheless, several contributions to the literature, e.g. GGL (2001), argue that fundamental inflation matches actual inflation quite well.

To assess the accuracy of the model's fit to the actual data, we employ a bootstrap approach that infers the distribution of our measures of fit, i.e. the ratio of standard deviations and the correlation coefficient, from estimating the model with artificially created data.

We obtain confidence intervals by drawing from the residuals of the estimated VAR model and generating new observations for the $\mathbf{Z}_t$ vector using the estimated companion matrix $\hat{\mathbf{A}}$. Using the artificially created observations the VAR

[12] A closely related empirical exercise can be found in Gruber (2004). He uses a similar empirical approach to assess the explanatory power of the intertemporal model of the current account. In his model, the lagged current account enters the present-value relation due to habit persistence. He also uses the ratio of standard deviations between model-consistent and actual data series as a measure of the model's fit.

model is estimated again and a new coefficient matrix is computed. From this we compute the series of expected real marginal cost and regress actual inflation on the present value of future real marginal cost to infer the slope coefficient. Finally, the ratio of standard deviation and the correlation coefficient is computed. Repeating this procedure 10000 times provides us with an empirical distribution for the ratio of standard deviations and the correlation coefficient from which an interval that includes 90 per cent of the estimates can be calculated.[13]

However, Kilian (1998) shows that this standard bootstrap algorithm performs poorly when it is used to compute distributions of statistics that are nonlinear functions of VAR parameters. Note that both the ratio of standard deviations and the correlation coefficient are indeed highly nonlinear functions of the estimated VAR coefficients. Therefore, we cannot rely on the conventional bootstrap approach here since the small sample distributions of the measures of fit are likely to be biased.

Therefore, we follow Kurmann (2003) and Adler (2003) and apply Kilian's bias-corrected bootstrap algorithm. Basically, he proposes to replace the estimated VAR coefficients $\hat{\mathbf{A}}$ by bias-corrected estimates $\bar{\mathbf{A}}$ before running the bootstrap to compute the measures of fit. Details about this bias-correction can be found in Kilian (1998) and Kurmann (2003) and are briefly sketched in the appendix. The algorithm also includes a procedure for shrinking the bias estimates in case the bias-corrected VAR estimates imply that the resulting VAR becomes nonstationary

Moreover, Kilian (1998) proposes a second bias-correction because the OLS estimates are themselves biased away from their population values. We therefore should replace $\hat{\mathbf{A}}$ prior to generating artificial data series. The approach amounts to a bootstrap-after-bootstrap technique. In a first step we do a bootstrap to approximate the OLS small-sample bias. In a second step we replace the coefficients with bias-corrected coefficients, use the first stage bias-correction again and use a second bootstrap-round to generate the distribution of our estimates and measures of fit.

Since the correlation coefficient is independent of the estimate of $\gamma$, we hold $\gamma$ fixed across bootstrap replications. Otherwise, we would have two unidentifiable

[13]By definition, this bootstrap approach respects the boundedness of the correlation coefficient. Furthermore, this approach allows for skewness and does not impose symmetry.

influences on changes in the correlation across bootstrap-rounds, namely the forecast of future real marginal cost obtained from VAR estimates and the slope coefficient $\gamma$.

# 4 Results

We use quarterly data for the Euro area from 1970:1 to 2003:4. The data construction is explained in the appendix. While unit root tests have some problems to reject the null of a unit root, we nevertheless assume inflation and real marginal cost to be stationary and attribute the difficulty to reject the null to well-known power problems of unit root tests in the (very likely) presence of structural breaks. Moreover, many other papers use the same data set for GMM applications and also assume stationarity given that stationary inflation rates and, hence, the existence of a steady state are a main prerequisite to model staggered price setting.

## 4.1 Testing the present-value relation

The Campbell-Shiller approach pursued in this paper requires the price level and the level of nominal marginal cost to be cointegrated with a cointegrating vector $\beta' = (1, -1)$. A standard Johansen trace or eigenvalue test within a vector error-correction model finds a cointegrating relation but rejects the unit cointegrating coefficients. The results of the more powerful test for prespecified cointegration proposed by Horvath and Watson (1995), however, supports this cointegrating relation, see table (1). In other words, the powerful Horvath-Watson test confirms that real marginal cost, i.e. the cointegrating relation, is stationary. Note that this test amounts to a standard Likelihood-Ratio test for the presence of the candidate error-correction terms in a first difference VAR and is more powerful than conventional unit-root tests applied to the real marginal cost term.

## 4.2 Forward-looking inflation dynamics

To the extend that inflation is forward-looking, equation (5) suggests that current inflation provides information about future real marginal cost. This forecasting property implies that Granger causality should run from inflation to real activity. Hence, the information incorporated in inflation should help forecasting real marginal cost. Table (2) reports the results from testing these Granger causality propositions. We find that inflation consistently Granger causes *rmc* for alternative VAR orders.

The accuracy of the bootstrap approach relies on VAR errors to be serially uncorrelated and homoscedastic. We estimate the two alternative auxiliary VAR models with five lags in order to minimize serial correlation and heteroscedasticity in the estimated residuals. Hence, we balance the suggestions of standard information criteria with the results of specification tests, see table (3).

The resulting VAR parameters are reported in table (4). We proceed by calculating fundamental inflation according to the model laid out before. Following Rudd and Whelan (2003), actual inflation is regressed on the present-value of future real marginal cost in order to infer the parameter $\gamma$. We then compare the series of fundamental inflation with actual inflation by means of the ratio of their standard deviations and their correlation coefficient.

The results for the baseline model, i.e. the pure forward-looking model, are presented in table (5) for alternative values of the discount factor $\phi$. Throughout the alternative specifications we report estimates under different values of $\phi$ to check the robustness of the results. We set the discount factor to 0.99, 0.98, 0.95, and 0.91. While the first three values are fairly standard assumptions, the low value of 0.91 corresponds to the specification of GGL (2001) for European data and is used here for reasons of comparability.

The forward-looking specification for $\phi = 0.99$ yields a series of fundamental inflation, see figure (1), that, at first sight, tracks the actual European inflation rate quite well. The ratio of standard deviations is 0.89 and the correlation coefficient between actual and fundamental inflation is 0.88, see table (5). This is exactly the result put forward in the literature. However, this impressive fit is merely a point estimate. The confidence bands obtained from the bootstrap approach reveal that both measures are associated with an extremely large degree of uncertainty. In fact, the confidence band shows that a correlation of $-0.31$ is as likely (within a 90% band) as a correlation of 0.96. Moreover, the confidence band includes ratios of standard deviations between 0.14 and 1.651. Hence we cannot say whether the model fits or fails. It could equally likely explain 15% of the variation of the inflation rate and more than 150% of the variation in inflation. Moreover, it turns out that the performance of the Calvo model for European data is even worse than for U.S. data. Kurmann (2003) finds a standard deviation ratio for the U.S. within the interval $[0.01, 1.57]$ and a correlation coefficient within the range $[0.40, 0.99]$. While the width of the former interval is more or less equal to the uncertainty surrounding the standard deviation ratio for European data, the latter interval is substantially smaller than the corresponding interval for European data. Given that the baseline forecasting VAR contains some insignificant coefficients, see table (4), we check for the robustness of these results by restricting the lag order to $q = 3$. As shown in table (6), the results are remarkably similar. These findings are also robust to the choice of the forecasting VAR model. Table (7) reports results of an alternative VAR model that includes real marginal cost and changes in nominal marginal cost. This model yields equally wide confidence bands.

The fit slightly improves under lower discount factors. However, even the smallest confidence band (under $\phi$ equal to 0.91) covers an explanatory power of the variation in inflation between 45% and more than 100% and, hence, is too wide to be interpretable.

### 4.3 Hybrid inflation dynamics

It is frequently argued that allowing for inflation inertia within the NKPC generates a well fitting description of actual inflation. When estimating the backward-looking model we need to set the degree of indexation. As a first guess we specify $\kappa = 0.46$ as suggested by Smets and Wouters (2003) for European data. As mentioned above, we also use the estimate Onatski and Williams (2004). They estimate the Euro area model following Smets and Wouters under an own set of priors and get a degree of indexation of $\kappa = 0.32$.

Allowing for inflation inertia through backward-looking indexation does indeed substantially improve the model's fit. Setting $\kappa = 0.46$ gives a correlation coefficient of 0.91 (for a discount factor of 0.99), see table (8). Interestingly, these estimates are much more reliable since the confidence bands narrow. The interval now covers a correlation of 0.82 up to 0.94. For all values of the discount factor, the confidence bands narrow considerably. Under a slightly lower degree

of price indexation of $\kappa = 0.32$ we obtain intermediate results (see table 9) with wider confidence bands than under $\kappa = 0.46$ but wider bands than under a pure forward-looking specification. The same holds for the estimates under $\kappa = 0.23$ (see table 10). However, while the correlation is now more precisely computed, we are still left with wide confidence bands around the ratio of standard deviations that impede a reasonable interpretation of the model's adequacy.
Figures (2) to (5) show the density of the ratio of standard deviations of fundamental and actual inflation across 10000 bootstrap replications. Clearly, the densities become much more narrowly centered around a standard deviation ratio and a correlation coefficient of unity if the degree of indexation is increased.

## 4.4 The duration of sticky-price contracts

The estimated slope coefficient $\gamma$ in conjunction with a fixed discount factor $\phi$ allows us to infer the average duration of sticky-price contracts under the Calvo price setting scheme. In the baseline specification with no indexation we obtain an estimate of $\gamma$ of 0.012, which implies a duration of fixed prices of 10 quarters. For lower values of the discount factor we obtain durations between 5.82 and 9.14 quarters.[14] These numbers are perfectly in line with those obtained from GGL (2001). Their estimated slope coefficient is 0.014 with a duration between 10 and 12 quarters. Gagnon and Khan (2001) also obtain similar results. Recently provided micro evidence indicates a slightly higher frequency of price adjustment. Interestingly, however, Hoffmann and Kurz-Kim (2004) find for the case of retail consumer prices in Germany that prices are fixed for two years on average, which is broadly consistent with the results of this paper.[15] Under partial price indexation with $\kappa = 0.46$ the duration of Calvo contracts changes to values between 8 and 14 quarters. Smets and Wouters (2003) argue for this case

[14] This is roughly consistent with the results of Benigno and López-Salido (2002). These authors estimate country-specific hybrid NKPC models and aggregate their results across major European economies. They find an area-wide average duration between 7 and 8.3 quarters.

[15] Recently, extensive research on price stickiness in EMU countries based on micro data was carried out under the auspices of the Eurosystem's Inflation Persistence Network. See Aucremanne and Dhyne (2004) for Belgium, Dias, Dias, and Neves (2004) for Portugal, Fabiani, Gattulli, and Sabbatini (2004) for Italy, Baudry, Le Bihan, Sevestre, and Tarrieu (2004) for France, and Álvarez and Hernando (2004) for Spain.

that "the greater stickiness of prices is somewhat counterintuitive, but turns out to be a very robust outcome of the estimated model" (p. 1144). Nevertheless, these spells of price stickiness are implausibly long. While the model's fit in term of the ratio of standard deviations of actual and fundamental inflation worsens when the degree of indexation increases from 0.32 to 0.46, the average duration of sticky prices becomes more realistic.

To summarize, we find that the pure forward looking model can hardly be interpreted as tracking "... the behavior of actual inflation quite well" (GGL 2001, p. 1260). The confidence bands preclude any meaningful assessment of the model's empirical performance. However, allowing for inflation inertia through backward-looking indexation narrows confidence bands around the point estimates and shows that we can match actual inflation quite accurately in terms of the correlation coefficient, but that a large portion of the variance of actual inflation cannot be explained.[16]

# 5 Conclusions

The standard New Keynesian Phillips Curve specifies current inflation as the present-value of the future stream of real marginal cost. Previous contributions to the literature exploited VAR projections of future real marginal cost to proxy market expectations and to derive a series of model-consistent or fundamental inflation rates. It is frequently argued that this series of fundamental inflation explains actual inflation quite well.

In this paper we shed light on this finding using data for the Euro area. In particular, we used bootstrapped confidence bands to quantify the degree of estimation uncertainty around these estimates. We show that the result for the purely forward-looking model cannot be interpreted as it is done in the literature due to the immensely wide confidence intervals. The baseline specification is consistent with both a complete failure of the model where the correlation coefficient between actual and fundamental inflation is -0.31 and, at the same time, with a remarkably well fitting model where actual and fundamental inflation exhibit an

[16]Hence, partial indexation improves the fit, but is far from generating a well fitting NKPC as e.g. Sahuc (2004) suggests for European data.

almost perfect positive correlation.

Once we allow for inflation inertia in the sense that past inflation enters the Phillips curve, e.g. through price indexation, the model's fit improves. Hence, we cannot interpret the evidence unless we supplement the model with backward looking inflation. Even then, i.e. once we include lagged inflation, the explanatory power of the staggered price setting scheme for European inflation is limited due to a large degree of estimation uncertainty.

Hence, the results presented in this paper further question the appropriateness of standard New Keynesian models based on staggered price setting to adequately describe inflation. Given that even the specification that allows for inflation inertia does a poor job in replicating inflation dynamics, the lesson from the evidence presented in this paper is to focus on other schemes of staggered price setting to model nominal rigidities.

# 6 Appendix A: The data set

We use quarterly data for the Euro area obtained from the ECB's Area Wide Model database covering 1970:1 -2003:4. Inflation is measured as the first difference (multiplied by 400 to obtain annualized inflation rates in percentage points) of the logarithms of the implicit GDP deflator. Real marginal costs can be shown to be proportional to labor's share of income. Suppose a conventional Cobb-Douglas technology $Y_t = AK_t^{\alpha_k} N_t^{\alpha_n}$ with constant returns-to-scale. Real marginal costs in logs, $rmc_t$, are then given by

$$\log(RMC) = rmc = \log\left(\frac{W_t N_t}{P_t Y_t}\right) - \log(\alpha_n)$$

where $W_t N_t$ denotes compensation to employees, $P$ is the deflator and $Y$ is real GDP. Hence, we use the log ratio of compensation to employees to nominal GDP in deviations from the mean (multiplied by 100 to obtain percentage points).

# 7 Appendix B: The bootstrap algorithm

This algorithm implements Kilian's (1998) bias-corrected bootstrap approach. The documentation roughly follows Kurmann (2003).

1. Estimate the bivariate VAR system

   $$\mathbf{Z}_{t+1} = \mathbf{A}\mathbf{Z}_t + \mathbf{\Gamma}_{Z_{t+1}}$$

   by OLS and consider the estimated companion matrix $\hat{\mathbf{A}}$ and the vector of residuals $\left\{\hat{\mathbf{\Gamma}}_{Z_{t+1}}\right\}$, which is rescaled as follows

   $$\hat{\mathbf{\Gamma}}'_t = \frac{1}{T}\sum_{t=1}^{T} \hat{\mathbf{\Gamma}}_t \sqrt{\frac{T}{T-2q}}$$

   where $q$ is the lag length of the VAR system (here we set $q = 5$).

2. For each artificial series, fit a VAR and estimate the coefficients $\hat{\mathbf{A}}_i^*$.

3. Approximate the OLS small-sample bias term $\boldsymbol{\psi} = E\left(\hat{\mathbf{A}} - \mathbf{A}\right)$ by

$$\hat{\boldsymbol{\psi}} = \frac{1}{N}\sum_{i=1}^{N}\left(\hat{\mathbf{A}}_i^* - \hat{\mathbf{A}}\right)$$

   where $N$ is the number of bootstrap replications, which we set to $N = 10000$.

4. Construct the bias-corrected coefficient estimate $\bar{\mathbf{A}} = \hat{\mathbf{A}} - \hat{\boldsymbol{\psi}}$. Compute the roots of $\bar{\mathbf{A}}$ and, if necessary, adjust the bias-correction following the procedure laid out in Kilian (1998) to avoid a non-stationary VAR.

5. Replace $\hat{\mathbf{A}}$ with $\bar{\mathbf{A}}$ in the auxiliary VAR and generate $N$ new artificial series $\left\{\boldsymbol{\Gamma}^*_{Z_{t+1}}\right\}$ from this bias-corrected data-generating process.

6. Fit a VAR to each artificial data series and estimate a companion matrix $\hat{\mathbf{A}}^*$. To reduce computational requirements, use the first-stage bias approximation again and, if necessary, adjust the bias-correction again to avoid a non-stationary VAR

$$\bar{\mathbf{A}}^* = \hat{\mathbf{A}}^* - \hat{\boldsymbol{\psi}}$$

7. For each artificial series, compute the series of fundamental inflation. Calculate the ratio of standard deviations of fundamental and actual inflation as well as the correlation coefficient.

8. Calculate the 5% and the 95% fractiles of the distributions of the ratio of standard deviations and correlation coefficients.

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Table 1: Results of cointegration tests

| lag order | $\ln L$ | test statistic $2(\ln L_{VECM} - \ln L_{VAR})$ | critical value (5%) |
|---|---|---|---|
| $q = 1$ | | | |
| VECM | -184.12 | | |
| VAR | -201.23 | 34.22 | 10.18 |
| $q = 2$ | | | |
| VECM | -172.93 | | |
| VAR | -183.87 | 21.88 | 10.18 |
| $q = 3$ | | | |
| VECM | -166.05 | | |
| VAR | -179.27 | 26.44 | 10.18 |
| $q = 4$ | | | |
| VECM | -152.41 | | |
| VAR | -163.14 | 21.46 | 10.18 |
| $q = 5$ | | | |
| VECM | -143.69 | | |
| VAR | -154.08 | 20.78 | 10.18 |

*Notes:* The Horvath-Watson test of the null hypothesis of no cointegration against the known alternative of rank $r = 1$ with $\beta' = (1, -1)$ corresponds to a Wald test for the inclusion of error-correction terms, i.e. real marginal cost, in a bivariate VAR in first differences of order $q$ with $\Delta nmc_t$ and $\pi_t$ and a constant. The critical value for the case of an unrestricted constant is from Horvath and Watson (1995), table 1.

Table 2: Granger causality tests

| | $H_0$ | $F-Statistic$ | $Prob$ |
|---|---|---|---|
| $q=1$ | $\pi \nRightarrow rmc$ | 25.19 | 0.00 |
| | $rmc \nRightarrow \pi$ | 17.10 | 0.00 |
| $q=2$ | $\pi \nRightarrow rmc$ | 11.14 | 0.00 |
| | $rmc \nRightarrow \pi$ | 2.48 | 0.09 |
| $q=3$ | $\pi \nRightarrow rmc$ | 9.48 | 0.00 |
| | $rmc \nRightarrow \pi$ | 1.57 | 0.20 |
| $q=4$ | $\pi \nRightarrow rmc$ | 7.21 | 0.00 |
| | $rmc \nRightarrow \pi$ | 0.67 | 0.61 |
| $q=5$ | $\pi \nRightarrow rmc$ | 5.80 | 0.00 |
| | $rmc \nRightarrow \pi$ | 1.25 | 0.29 |
| $q=6$ | $\pi \nRightarrow rmc$ | 5.12 | 0.00 |
| | $rmc \nRightarrow \pi$ | 0.98 | 0.44 |
| $q=7$ | $\pi \nRightarrow rmc$ | 4.42 | 0.00 |
| | $rmc \nRightarrow \pi$ | 0.85 | 0.54 |
| $q=8$ | $\pi \nRightarrow rmc$ | 4.21 | 0.00 |
| | $rmc \nRightarrow \pi$ | 0.76 | 0.64 |

*Notes:* Pairwise Granger causality tests for alternative lag orders. The notation $\nRightarrow$ means "does not Granger cause".

Table 3: Choosing the lag order of the auxiliary VAR

| | AIC($q$) | SC($q$) | HQ($q$) | LM(1) | LM(4) | White |
|---|---|---|---|---|---|---|
| | | | VAR system: $[rmc_t, \pi_t]'$ | | | |
| $q=1$ | -3.77 | -3.64 | -3.72 | 15.08*** | 20.07*** | 18.87 |
| $q=2$ | -3.79 | -3.57 | -3.70 | 15.19*** | 12.38** | 61.84** |
| $q=3$ | -3.83 | -3.51 | -3.70 | 21.53*** | 11.39** | 88.70 |
| $q=4$ | -3.85 | -3.45 | -3.68 | 15.11*** | 4.13 | 152.30 |
| $q=5$ | -3.83 | -3.43 | -3.63 | 3.87 | 7.96* | 181.92 |
| | | | VAR system: $[rmc_t, \Delta nmc_t]'$ | | | |
| $q=1$ | 3.03 | 3.16 | 3.08 | 47.69*** | 14.26*** | 18.73 |
| $q=2$ | 2.62 | 2.85 | 2.71 | 9.07*** | 13.83*** | 51.63 |
| $q=3$ | 2.65 | 2.96 | 2.78 | 26.22*** | 11.15** | 101.34* |
| $q=4$ | 2.61 | 3.01 | 2.77 | 14.78*** | 8.59* | 137.98 |
| $q=5$ | 2.62 | 3.11 | 2.82 | 2.40 | 3.64 | 212.56 |

*Notes:* AIC($q$), SC($q$), and HQ($q$) denote the Akaike information criterion, the Schwartz criterion, and the Hannan-Quinn information criterion, respectively, for a bivariate VAR of order $q$. These criteria compare the goodness of the fit of maximum likelihood estimations and correct for the loss of degrees of freedom when additional lags are added. LM($h$) is a multivariate Lagrange-Multiplier test for residual correlation up to order $h$. Under the null hypothesis of no serial correlation of order $h$, the LM statistic is asymptotically $\chi^2$ distributed with 4 degrees of freedom. White denotes the $\chi^2$ test statistic of a White test that includes cross terms. The null is the absence of heteroscedasticity. A significance level of 1%, 5%, and 10% is indicated by ***, **, and *.

Table 4: Estimated VAR parameters

| | VAR system: $[rmc_t, \pi_t]'$ dependent variable | | | VAR system: $[rmc_t, \Delta nmc_t]'$ dependent variable | |
|---|---|---|---|---|---|
| | $rmc_t$ | $\pi_t$ | | $rmc_t$ | $\Delta nmc_t$ |
| $rmc_{t-1}$ | 1.11 (0.08) | -0.32 (0.24) | $rmc_{t-1}$ | 0.93 (0.16) | -0.71 (0.19) |
| $rmc_{t-2}$ | 0.08 (0.14) | 0.59 (0.33) | $rmc_{t-2}$ | 0.41 (0.28) | 0.93 (0.34) |
| $rmc_{t-3}$ | -0.19 (0.15) | -0.29 (0.29) | $rmc_{t-3}$ | -0.54 (0.26) | -0.32 (0.30) |
| $rmc_{t-4}$ | 0.18 (0.14) | 0.21 (0.33) | $rmc_{t-4}$ | 0.53 (0.31) | 0.14 (0.36) |
| $rmc_{t-5}$ | -0.19 (0.08) | -0.18 (0.18) | $rmc_{t-5}$ | -0.34 (0.20) | -0.04 (0.23) |
| $\pi_{t-1}$ | 0.06 (0.03) | 0.62 (0.09) | $\Delta nmc_{t-1}$ | 0.20 (0.13) | 0.78 (0.15) |
| $\pi_{t-2}$ | -0.04 (0.03) | 0.17 (0.08) | $\Delta nmc_{t-2}$ | -0.16 (0.13) | 0.02 (0.16) |
| $\pi_{t-3}$ | 0.06 (0.03) | -0.07 (0.07) | $\Delta nmc_{t-3}$ | 0.22 (0.11) | 0.12 (0.13) |
| $\pi_{t-4}$ | -0.01 (0.04) | 0.40 (0.10) | $\Delta nmc_{t-4}$ | -0.14 (0.16) | 0.20 (0.20) |
| $\pi_{t-5}$ | -0.06 (0.03) | -0.13 (0.09) | $\Delta nmc_{t-5}$ | -0.12 (0.06) | -0.13 (0.08) |
| $R^2$ | 0.99 | 0.86 | | 0.99 | 0.70 |

*Notes:* OLS Estimates of the auxiliary VAR system. Standard errors in parenthesis.

Table 5: The uncertain fit of the forward-looking NKPC using the baseline forecasting VAR with five lags

| | $\phi = 0.99$ | | $\phi = 0.98$ | |
|---|---|---|---|---|
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.(\pi^{fund})}{std.dev.(\pi^{actual})}$ | 0.89 | $\begin{bmatrix} 0.14 & 1.51 \end{bmatrix}$ | 0.88 | $\begin{bmatrix} 0.19 & 1.29 \end{bmatrix}$ |
| $corr(\pi^{fund}, \pi^{actual})$ | 0.88 | $\begin{bmatrix} -0.31 & 0.96 \end{bmatrix}$ | 0.88 | $\begin{bmatrix} 0.27 & 0.95 \end{bmatrix}$ |
| $D$ | 10.63 | | 9.14 | |
| | $\phi = 0.95$ | | $\phi = 0.91$ | |
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.(\pi^{fund})}{std.dev.(\pi^{actual})}$ | 0.88 | $\begin{bmatrix} 0.31 & 1.12 \end{bmatrix}$ | 0.89 | $\begin{bmatrix} 0.45 & 1.02 \end{bmatrix}$ |
| $corr(\pi^{fund}, \pi^{actual})$ | 0.88 | $\begin{bmatrix} 0.70 & 0.93 \end{bmatrix}$ | 0.87 | $\begin{bmatrix} 0.79 & 0.91 \end{bmatrix}$ |
| $D$ | 7.03 | | 5.82 | |

*Notes:* Results using the baseline VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \pi_t, ..., \pi_{t-q+1}]'$ and $q = 5$ to generate the present value of future real marginal cost. $D$ denotes the average duration (in quarters) of fixed-price Calvo contracts. The parameter $\gamma$ is estimated by regressing actual inflation on the present-value of the future path of real marginal cost and a constant and is kept fixed across bootstrap replications for reasons explained in the text. Note that the standard errors of this regression are not valid due to a generated-regressor problem. The confidence bands denote the 5% and the 95% fractiles of the distribution of the respective measure of fit across 10000 bias-corrected bootstrap replications as explained in the text.

Table 6: The uncertain fit of the forward-looking NKPC using the baseline forecasting VAR with three lags

| | $\phi = 0.99$ | | $\phi = 0.98$ | |
|---|---|---|---|---|
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.85 | $\left[\begin{array}{cc} 0.13 & 1.12 \end{array}\right]$ | 0.85 | $\left[\begin{array}{cc} 0.20 & 0.99 \end{array}\right]$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.85 | $\left[\begin{array}{cc} 0.81 & 0.87 \end{array}\right]$ | 0.85 | $\left[\begin{array}{cc} 0.81 & 0.87 \end{array}\right]$ |
| $D$ | 12.37 | | 10.03 | |
| | $\phi = 0.95$ | | $\phi = 0.91$ | |
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.85 | $\left[\begin{array}{cc} 0.36 & 0.92 \end{array}\right]$ | 0.85 | $\left[\begin{array}{cc} 0.49 & 0.90 \end{array}\right]$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.85 | $\left[\begin{array}{cc} 0.81 & 0.87 \end{array}\right]$ | 0.84 | $\left[\begin{array}{cc} 0.81 & 0.87 \end{array}\right]$ |
| $D$ | 7.31 | | 5.94 | |

*Notes:* Results using the baseline VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \pi_t, ..., \pi_{t-q+1}]'$ and $q = 3$ to generate the present value of future real marginal cost. $D$ denotes the average duration (in quarters) of fixed-price Calvo contracts. The parameter $\gamma$ is estimated by regressing actual inflation on the present-value of the future path of real marginal cost and a constant and is kept fixed across bootstrap replications for reasons explained in the text. Note that the standard errors of this regression are not valid due to a generated-regressor problem. The confidence bands denote the 5% and the 95% fractiles of the distribution of the respective measure of fit across 10000 bias-corrected bootstrap replications as explained in the text.

Table 7: The uncertain fit of the forward-looking NKPC using the alternative forecasting VAR

| | $\phi = 0.99$ | | $\phi = 0.98$ | |
|---|---|---|---|---|
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.(\pi^{fund})}{std.dev.(\pi^{actual})}$ | 0.88 | $\begin{bmatrix} 0.12 & 1.37 \end{bmatrix}$ | 0.88 | $\begin{bmatrix} 0.17 & 1.25 \end{bmatrix}$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.88 | $\begin{bmatrix} -0.35 & 0.96 \end{bmatrix}$ | 0.87 | $\begin{bmatrix} 0.25 & 0.95 \end{bmatrix}$ |
| $D$ | 11.73 | | 9.73 | |
| | $\phi = 0.95$ | | $\phi = 0.91$ | |
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.(\pi^{fund})}{std.dev.(\pi^{actual})}$ | 0.87 | $\begin{bmatrix} 0.29 & 1.11 \end{bmatrix}$ | 0.87 | $\begin{bmatrix} 0.43 & 1.02 \end{bmatrix}$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.87 | $\begin{bmatrix} 0.67 & 0.92 \end{bmatrix}$ | 0.86 | $\begin{bmatrix} 0.77 & 0.90 \end{bmatrix}$ |
| $D$ | 7.23 | | 5.90 | |

*Notes:* Results using the alternative VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \Delta nmc_t, ..., \Delta nmc_{t-q+1}]'$ and $q = 5$ to generate the present value of future real marginal cost. $D$ denotes the average duration (in quarters) of fixed-price Calvo contracts. The parameter $\gamma$ is estimated by regressing actual inflation on the present-value of the future path of real marginal cost and a constant and is kept fixed across bootstrap replications for reasons explained in the text. Note that the standard errors of this regression are not valid due to a generated-regressor problem. The confidence bands denote the 5% and the 95% fractiles of the distribution of the respective measure of fit across 10000 bias-corrected bootstrap replications as explained in the text.

Table 8: The uncertain fit of the backward-looking indexation NKPC using the baseline forecasting VAR

| | $\phi = 0.99, \kappa = 0.46$ | | $\phi = 0.98, \kappa = 0.46$ | |
|---|---|---|---|---|
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.91 | $\left[\begin{array}{cc} 0.46 & 1.24 \end{array}\right]$ | 0.91 | $\left[\begin{array}{cc} 0.50 & 1.137 \end{array}\right]$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.91 | $\left[\begin{array}{cc} 0.82 & 0.94 \end{array}\right]$ | 0.91 | $\left[\begin{array}{cc} 0.85 & 0.93 \end{array}\right]$ |
| $D$ | 14.51 | | 12.59 | |
| | $\phi = 0.95, \kappa = 0.46$ | | $\phi = 0.91, \kappa = 0.46$ | |
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.91 | $\left[\begin{array}{cc} 0.60 & 1.04 \end{array}\right]$ | 0.91 | $\left[\begin{array}{cc} 0.68 & 0.98 \end{array}\right]$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.91 | $\left[\begin{array}{cc} 0.88 & 0.93 \end{array}\right]$ | 0.91 | $\left[\begin{array}{cc} 0.89 & 0.92 \end{array}\right]$ |
| $D$ | 9.85 | | 8.28 | |

*Notes:* Results using the baseline VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \pi_t, ..., \pi_{t-q+1}]'$ and $q = 5$ to generate the present value of future real marginal cost. $D$ denotes the average duration (in quarters) of fixed-price Calvo contracts. The parameter $\gamma$ is estimated by regressing actual inflation on the present-value of the future path of real marginal cost and a constant and is kept fixed across bootstrap replications for reasons explained in the text. Note that the standard errors of this regression are not valid due to a generated-regressor problem. The confidence bands denote the 5% and the 95% fractiles of the distribution of the respective measure of fit across 10000 bias-corrected bootstrap replications as explained in the text.

Table 9: The uncertain fit of the backward-looking indexation NKPC using the baseline forecasting VAR

| | $\phi = 0.99, \kappa = 0.32$ | | $\phi = 0.98, \kappa = 0.32$ | |
|---|---|---|---|---|
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.90 | $\begin{bmatrix} 0.33 & 1.32 \end{bmatrix}$ | 0.90 | $\begin{bmatrix} 0.38 & 1.18 \end{bmatrix}$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.91 | $\begin{bmatrix} 0.76 & 0.94 \end{bmatrix}$ | 0.91 | $\begin{bmatrix} 0.81 & 0.94 \end{bmatrix}$ |
| $D$ | 12.940 | | 11.15 | |
| | $\phi = 0.95, \kappa = 0.32$ | | $\phi = 0.91, \kappa = 0.32$ | |
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.89 | $\begin{bmatrix} 0.50 & 1.06 \end{bmatrix}$ | 0.89 | $\begin{bmatrix} 0.60 & 0.99 \end{bmatrix}$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.90 | $\begin{bmatrix} 0.86 & 0.93 \end{bmatrix}$ | 0.90 | $\begin{bmatrix} 0.88 & 0.92 \end{bmatrix}$ |
| $D$ | 8.66 | | 7.23 | |

*Notes:* Results using the baseline VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \pi_t, ..., \pi_{t-q+1}]'$ and $q = 5$ to generate the present value of future real marginal cost. $D$ denotes the average duration (in quarters) of fixed-price Calvo contracts. The parameter $\gamma$ is estimated by regressing actual inflation on the present-value of the future path of real marginal cost and a constant and is kept fixed across bootstrap replications for reasons explained in the text. Note that the standard errors of this regression are not valid due to a generated-regressor problem. The confidence bands denote the 5% and the 95% fractiles of the distribution of the respective measure of fit across 10000 bias-corrected bootstrap replications as explained in the text.

Table 10: The uncertain fit of the backward-looking indexation NKPC using the baseline forecasting VAR

| | $\phi = 0.99, \kappa = 0.23$ | | $\phi = 0.98, \kappa = 0.23$ | |
|---|---|---|---|---|
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.89 | $\left[\ 0.26 \quad 1.37\ \right]$ | 0.89 | $\left[\ 0.31 \quad 1.21\ \right]$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.90 | $\left[\ 0.66 \quad 0.95\ \right]$ | 0.90 | $\left[\ 0.76 \quad 0.94\ \right]$ |
| $D$ | 12.11 | | 10.45 | |
| | $\phi = 0.95, \kappa = 0.23$ | | $\phi = 0.91, \kappa = 0.23$ | |
| | estimate | 90% conf. band | estimate | 90% conf. band |
| $\frac{std.dev.\left(\pi^{fund}\right)}{std.dev.\left(\pi^{actual}\right)}$ | 0.89 | $\left[\ 0.44 \quad 1.07\ \right]$ | 0.88 | $\left[\ 0.55 \quad 0.99\ \right]$ |
| $corr\left(\pi^{fund}, \pi^{actual}\right)$ | 0.90 | $\left[\ 0.84 \quad 0.93\ \right]$ | 0.90 | $\left[\ 0.86 \quad 0.92\ \right]$ |
| $D$ | 8.09 | | 6.73 | |

*Notes:* Results using the baseline VAR with $\mathbf{Z}_t = [rmc_t, ..., rmc_{t-q+1}, \pi_t, ..., \pi_{t-q+1}]'$ and $q = 5$ to generate the present value of future real marginal cost. $D$ denotes the average duration (in quarters) of fixed-price Calvo contracts. The parameter $\gamma$ is estimated by regressing actual inflation on the present-value of the future path of real marginal cost and a constant and is kept fixed across bootstrap replications for reasons explained in the text. Note that the standard errors of this regression are not valid due to a generated-regressor problem. The confidence bands denote the 5% and the 95% fractiles of the distribution of the respective measure of fit across 10000 bias-corrected bootstrap replications as explained in the text.

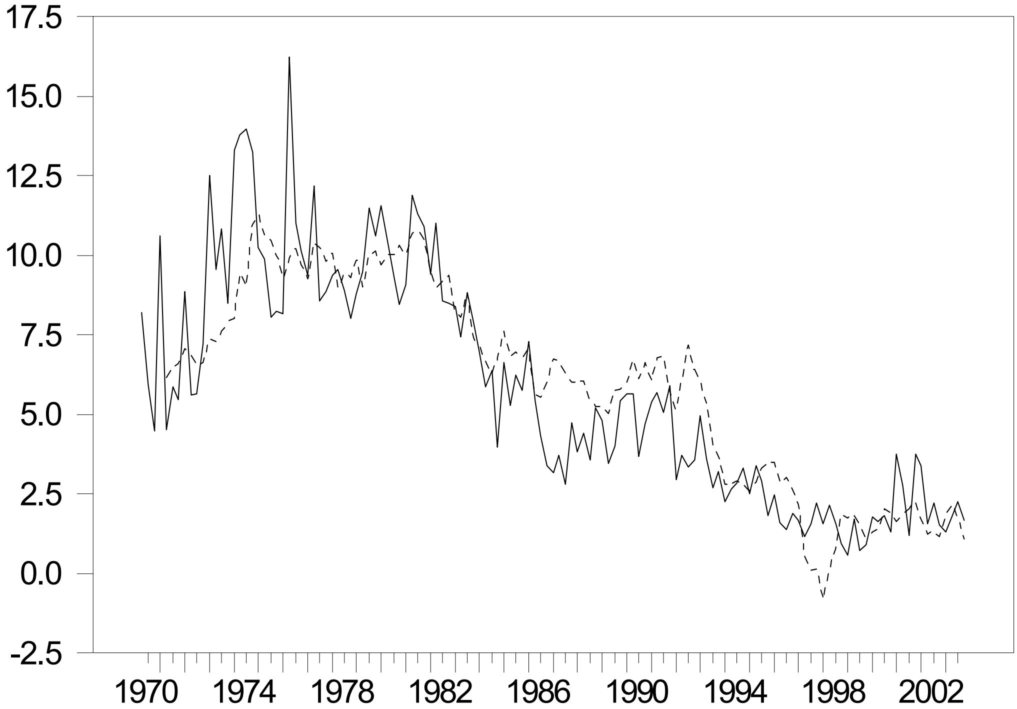


Figure 1: Actual (bold line) and fundamental (dotted line) inflation in the Euro area (in % p.a.) for $\phi = 0.99$ and $\kappa = 0$

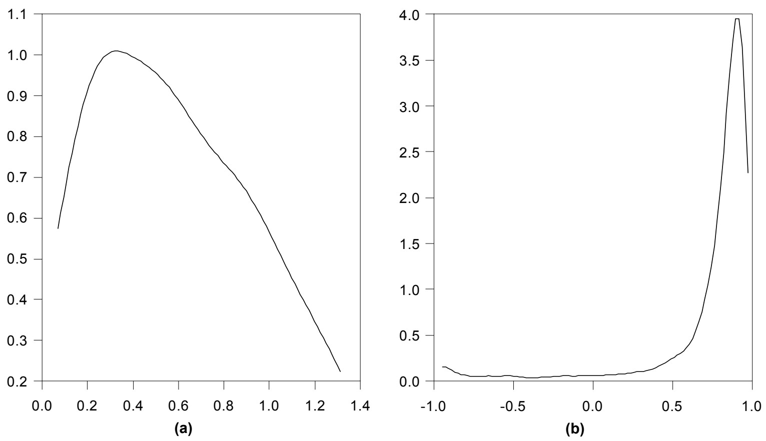


Figure 2: Distribution of (a) ratio of standard deviations and (b) correlation coefficient across bias-corrected bootstrap replications for $\phi = 0.99$ and $\kappa = 0$

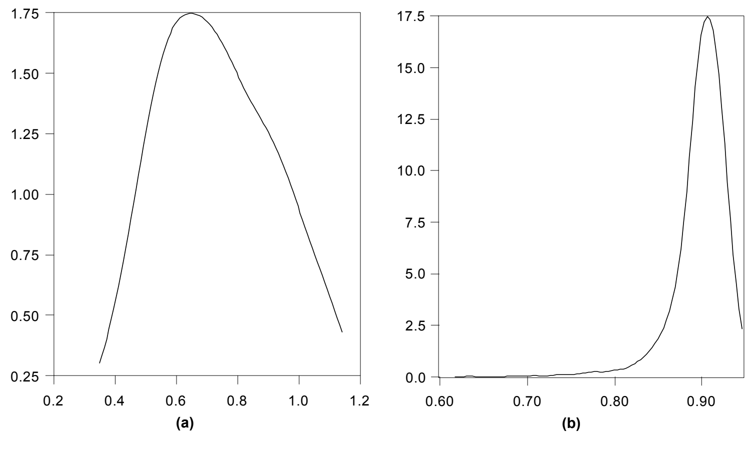


Figure 3: Distribution of (a) ratio of standard deviations and (b) correlation coefficient across bias-corrected bootstrap replications for $\phi = 0.99$ and $\kappa = 0.46$

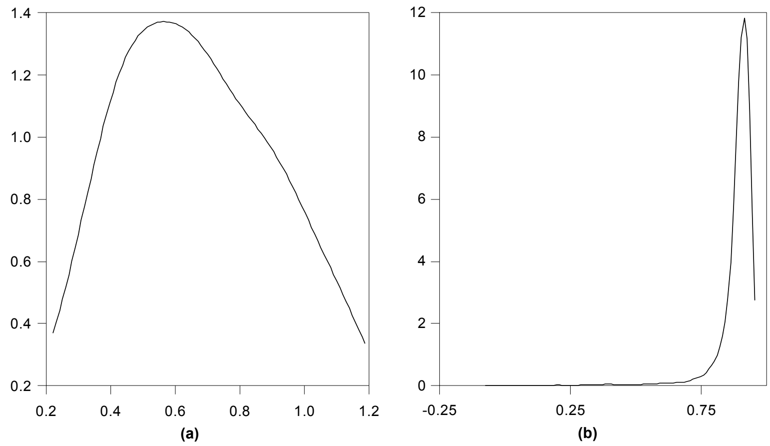


Figure 4: Distribution of (a) ratio of standard deviations and (b) correlation coefficient across bias-corrected bootstrap replications for $\phi = 0.99$ and $\kappa = 0.32$

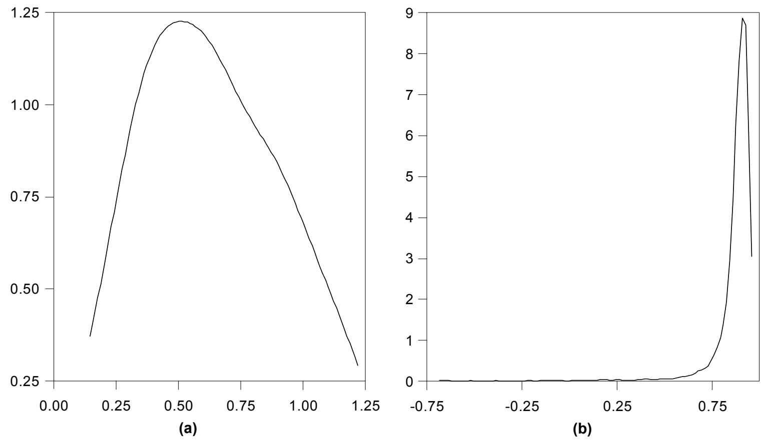


Figure 5: Distribution of (a) ratio of standard deviations and (b) correlation coefficient across bias-corrected bootstrap replications for $\phi = 0.99$ and $\kappa = 0.23$

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