Investigating Inflation Persistence Across Monetary Regimes*

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Abstract

Under inflation targeting inflation exhibits *negative* serial correlation in the U.K., and little or no persistence in Canada, Sweden and New Zealand, and estimates of the indexation parameter in hybrid New Keynesian Phillips curves are equal to 0.3 in Canada, and *zero* in the other three countries. Analogous results hold for the U.S., the U.K. and Sweden under the Gold Standard: under stable monetary regimes with clearly defined nominal anchors, inflation appears to be (nearly) purely forward-looking, so that no mechanism introducing backward-looking components is necessary to fit the data.

These results question the notion that the intrinsic inflation persistence found in post-WWII U.S. data—captured, in hybrid New Keynesian Phillips curves, by a significant extent of backward-looking indexation—is *structural* in the sense of Lucas (1976). Further, they suggest that 'hardwiring' inflation persistence in macroeconomic models is potentially misleading. In particular, both assessing alternative monetary regimes and computing optimal monetary policies based on models featuring structural persistence might deliver incorrect results.

Keywords: Inflation; inflation targeting; Gold Standard; Lucas critique; median-unbiased estimation; Markov Chain Monte Carlo.

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One might well expect that some years of experience with a regime in which inflation is more stable than it was in the period 1965-85 would reduce the extent to which inflation expectations vary with temporary variations in the inflation rate [...].

But under this view, it would be a mistake, in choosing a policy rule, to treat the 'intrinsic inflation inertia' that may have existed between 1970 and 1990 as inevitable; for it should dissipate before too long under precisely the kind of policy rule that it would be best to adopt. (emphasis added)

—Woodford (2006)

Evidence of persistence in the inflation rate suggests that market agents expect that monetary authorities will continue to pursue an inflationary policy; its absence would be consistent with market agents' beliefs that the authorities will pursue a stable monetary rule such as the gold standard's convertibility rule.

—Bordo and Schwartz (1999)

1 Introduction

Inflation persistence has been, over the last decade, one of the most intensely investigated topics in macroeconomics. The inability of New Keynesian Phillips curve models to replicate the high inflation persistence found in post-WWII U.S. data—first documented by Fuhrer and Moore (1995)—spawned a vast effort aimed at 'hardwiring' inflation persistence into macroeconomic models. From Fuhrer and Moore (1995) to Sheedy (2006), several authors have proposed different mechanisms, some *ad hoc*, some highly ingenious, to make inflation persistence *structural*, i.e. intrinsic to the deep structure of the economy and impossible to eradicate without incurring a recession.

1.1 'Hardwiring' intrinsic persistence in macro models

Building on the work of Buiter and Jewitt (1981), Fuhrer and Moore (1995) proposed a contracting model in which workers care about the level of their real wage relative to those of previous and successive cohorts of workers, while Gali and Gertler (1999a) postulated that a fraction of firms set their prices based on a backward-looking rule of thumb, thus automatically introducing a backward-looking component in aggregate inflation dynamics. More recently, increasing attention has been paid to the implications of relaxing a key (and highly counterfactual) assumption of the Calvo (1983) model, that of a constant probability that firms receive the 'Calvo signal' allowing them to re-optimise their prices. As shown by Dotsey (2002),[1] in particular, if the hazard function is increasing in the length of time expired since the price was last re-optimised—as it is found in the data—rather than being constant as in the original

[1] See also the theoretical work of Wolman (1999) and Sheedy (2006).

Calvo model, estimated hybrid New Keynesian Phillips curves will point towards a significant, and entirely spurious, extent of backward-looking indexation even if the data generation process is, by assumption, purely forward-looking. So far, the most popular mechanism for hardwiring inflation persistence into the structure of macroeconomic models is however the one originally devised by Christiano, Eichenbaum, and Evans (2005) (henceforth, CEE), and extensively applied, *inter alia*, by Smets and Wouters and their co-authors,[2] based on the notion that firms which are not allowed to re-optimise their price will change it nonetheless reflecting, either fully or partly, past average inflation.[3]

With the notable exception of the recent literature on non-flat (i.e., increasing) hazard functions for the probability of a price change, several of these 'solutions' are however well known for being sometimes less than fully satisfactory. The Buiter and Jewitt (1981)-Fuhrer and Moore (1995) (henceforth, BJFM) relative contracting model, for example, by implicitly assuming that labor supply is a function of the rate of change of the real wage[4]—rather than of its level—is in fact incompatible with basic microeconomics, while Gali and Gertler's (1999) assumption that a fraction of firms sets prices according to a rule of thumb introduces a backward-looking component by 'brute force'. As for full or partial indexation mechanisms along the lines of CEE or Smets and Wouters, again, as stressed by Cogley and Sbordone (2005), '[...] from a theoretical point of view [they are] not too satisfactory, since dependence on past inflation is introduced as an *ad hoc* feature.' Further, as stressed by Woodford (2006), there are several reasons to be skeptical of models featuring indexation to past inflation.

> 'One is the lack of direct microeconomic evidence for the indexation of prices [...]. Another is the lack of a plausible argument for why such a practice should be adopted universally—not only in environments with large and persistent swings in the inflation rate, but also when inflation is stable which is what one must assume if the model is treated as structural for purposes of policy analysis.'

The present paper is conceptually related to Woodford's second point. Abstracting from the specific way in which intrinsic inflation persistence is hardwired into macro models, the crucial question that ought to be asked of this entire literature is—in our view—a simple one:

[2]See in particular Smets and Wouters (2003) and Smets and Wouters (2007).

[3]See also the 'sticky-information' model of Mankiw and Reis (2002), and Sims' conceptually related 'rational inattention' approach—see Sims (2003) and Sims (2006).

[4]This is extensively discussed in Roberts (1997). As stressed by Roberts, a straighforward way of seeing this is to simply notice that the BJFM model is nothing but the 'slipped derivative' version of the Taylor (1980) model. Given that in the Taylor model labor supply is a function of the real wage *level*, it automatically follows that in the BJFM it is a function of its *rate of change*.

Are we really sure that the intrinsic persistence found in U.S. post-WWII data—captured, in hybrid New Keynesian Phillips curves, by a significant extent of backward-looking indexation—is structural in the sense of Lucas (1976)?

1.2 Is intrinsic persistence structural in the sense of Lucas (1976)?

As we previously stressed, a common theme among the vast majority of the papers in this literature is that the high inflation persistence found in post-WWII U.S. data is regarded—either explicitly or implicitly—as a *structural* feature, to be hardwired into the deep structure of the economy. But are we really sure that high inflation persistence truly is structural in the sense of Lucas (1976), i.e. it is invariant with respect to changes in the monetary regime? *What is, in fact, the empirical evidence in favor of such a position?* As a simple matter of logic, the *only* way to provide evidence in favor of the notion that intrinsic persistence is structural in the sense of Lucas (1976) is to show that its extent remains virtually unchanged across different monetary regimes. The reason, quite simply, is that only a significant extent of variation in the monetary policy rule allows the researcher to disentangle what is structural in the sense of Lucas (1976) from what it is not.[5]

Drawing on the experience of stable monetary regimes with clearly defined nominal anchors—historically, the Gold Standard and inflation targeting[6]—and conceptually in line with the recent work of Coenen and Levin (2004) based on German post-WWII data, the present work

- documents the (near) white-noiseness of inflation in the U.K., Canada, Sweden and New Zealand under inflation targeting, and reasserts the well-known lack of serial correlation of inflation in the U.S., the U.K. and Sweden under the Gold standard; and
- shows how estimates of the indexation parameter in hybrid New Keynesian Phillips curves are around 0.3 in Canada under inflation targeting, and virtually *zero* in both the other three inflation-targeting countries, and in the U.S., the U.K. and Sweden under the Gold standard.

In line with Coenen and Levin's (2004) conclusion that '[...] backward-looking price-setting behavior (such as indexation to lagged inflation) is not needed in explaining the aggregate data, at least in an environment with a stable monetary policy

[5] To put it differently, only a significant extent of variation in the (monetary) policy rule provides a proper *test* of the Lucas-invariance of a structural parameter.

[6] We ignore the price level targeting regime adopted by Sweden in 1931, following its abandonement of the Gold Standard, because of its short duration. After only two years Sweden switched to an exchange rate targeting regime, pegging the *krona* first to the British pound, and then, after 1939, to the U.S. dollar. On the Swedish interwar experience, see Jonung (1979).

regime and a transparent and credible inflation objective', our results question the notion that the intrinsic inflation persistence found in the post-WWII U.S. is structural in the sense of Lucas (1976). Further, they suggest that the widespread notion of inflation as a highly persistent process stems from the macroeconomic profession's dominant focus on the post-WWII U.S. experience. In particular, the fact that both in the U.S. itself under the Gold Standard, and in the U.K., Canada, Sweden, and New Zealand under either the Gold Standard or inflation targeting, (i) inflation was, or is, (nearly) white noise, and (ii) estimates of the indexation parameter in hybrid New Keynesian Phillips curves are (close to) zero, suggests that, as far as inflation persistence is concerned, *the post-WWII U.S. experience should not be regarded as the 'benchmark'.* To put it differently, there appears to be something peculiar concerning such experience, so that the properties exhibited by U.S. inflation over this period should not be regarded as representative or indicative of *general structural* properties of inflation that DSGE models should replicate. Although speculating on *what* exactly this peculiarity might be is beyond the scope of this paper, in section 4 we briefly discuss the recent work of Aoki and Kimura (2005), suggesting that the lack of a clearly defined nominal anchor, which has been the hallmark of the post-Bretton Woods U.S., might have played a crucial role.

The paper is organised as follows. The next section presents preliminary reduced-form evidence on inflation's statistical persistence from fixed-coefficients AR(p) models based on the Hansen (1999) 'grid bootstrap' MUB estimator of the sum of the autoregressive coefficients. In section 3 we present structural evidence based on Bayesian estimates of New Keynesian sticky-price DSGE models featuring Phillips curves with indexation. Section 4 discusses the implications of our results for macroeconomics, both in terms of modelling, and in terms of policy. Section 5 concludes, and outlines several future directions for research.

2 Reduced-Form Evidence: A Look at Statistical Persistence Across Monetary Regimes

In order to motivate the next section's structural investigation, we start by presenting evidence on inflation's *statistical* persistence—i.e., its extent of serial correlation—for each country and monetary regime[7] based on fixed-coefficients AR(p) models estimated *via* the Hansen (1999) 'grid bootstrap' MUB estimator of the sum of the

[7]Appendix B discusses and motivates our choices of how to break down the overall sample periods. In several cases such choices are obvious, being dictated by clear changes in the monetary framework—this is the case, for example, of the Gold Standard resumption in the United States, in January 1879, or the introduction of inflation targeting in the U.K., Sweden, New Zealand, and Canada, at the beginning of the 1990s. In other cases the choice is less clear-cut, as is the case of the post-Bretton Woods era for the U.S., which we break into the Great Inflation episode and the period following the Volcker stabilisation.

autoregressive coefficients. Specifically, for each country, regime/period, and inflation series[8] we estimate via OLS the AR(p) model[9]

$$\pi_t = \mu + \phi_1 \pi_{t-1} + \phi_2 \pi_{t-2} + ... + \phi_p \pi_{t-p} + u_t \tag{1}$$

selecting the lag order based on the Akaike information criterion (AIC) for a maximum possible number of lags P=6.[10] We then compute both median-unbiased estimates of our preferred measure of persistence—which, following Andrews and Chen (1994), we take it to be the sum of the autoregressive coefficients, ρ [11]—and 90%-coverage confidence intervals, *via* the Hansen (1999) 'grid bootstrap' procedure.[12]

2.1 United Kingdom

Starting from the U.K., whose experience is especially interesting for the present purposes, Table 1 shows a small subset of the results reported in Table B of Benati (2006), updating the results for the inflation-targeting regime up to 2005:4. As the table clearly shows, historically in the U.K. high inflation persistence appears to have been the *exception*, rather than the rule, with inflation estimated to have been very highly persistent only during the period between the floating of the pound, in June 1972, and the introduction of inflation targeting, in October 1992.[13] In particular, persistence appears to have been entirely absent under the Gold Standard,[14] either *de*

[8] We compute inflation as the non-annualised quarter-on-quarter rate of change of the relevant price index. Appendix A describes in detail our dataset.

[9] In the case of seasonally unadjusted series, we augment (1) with three seasonal dummies.

[10] Specifically, for each inflation series we select the lag order based on the model estimated over the full sample. The only exception is U.S. GNP deflator inflation, for which AIC selects, for the period 1879:1-2005:4, p=2. In order to make results for GNP deflator inflation fully comparable to those based on the GDP deflator for the post-WWII era, we set p=4, the lag order selected by AIC for GDP deflator inflation.

[11] As shown by Andrews and Chen (1994), the sum of the autoregressive coefficients maps one-to-one into two alternative measures of persistence, the cumulative impulse-response function to a one-time innovation and the spectrum at the frequency zero. Andrews and Chen (1994) also contain an extensive discussion of why an alternative measure favored, e.g., by Stock (1991), the largest autoregressive root, may provide a misleading indication of the true extent of persistence of the series depending on the specific values taken by the other autoregressive roots.

[12] Specifically, following Hansen (1999, section III.A) we recast (1) into the augmented Dickey-Fuller form $\pi_t = \mu + \rho\pi_{t-1} + \gamma_1 \Delta\pi_{t-1} + ... + \Delta\gamma_{p-1}\pi_{t-(p-1)} + u_t$ and we simulate the sampling distribution of the t-statistic t=($\hat{\rho}$-ρ)/$\hat{S}(\hat{\rho})$, where $\hat{\rho}$ is the OLS estimate of ρ, and $\hat{S}(\hat{\rho})$ is its estimated standard error, over a grid of possible values [$\hat{\rho}$-3$\hat{S}(\hat{\rho})$; $\hat{\rho}$+3$\hat{S}(\hat{\rho})$], with step increments equal to 0.01. For each of the possible values in the grid, we consider 1,999 replications. Both the median-unbiased estimates of ρ, and the 90% confidence intervals, are based on the bootstrapped distribution of the t-statistic.

[13] The overall picture is confirmed by Benati's (2006) more extensive analysis.

[14] One *caveat* is that the likely presence of measurement error in old (log) price series automatically introduces negative serial correlation in their first differences, thus biasing downwards persistence estimates. Unfortunately, it is not clear at all how to even gauge an idea of the likely extent of such an effect.

facto or *de jure*, while under inflation targeting inflation is estimated to have been so far *negatively* serially correlated,[15] although statistically indistinguishable from white noise, based on all the three price indices we consider.[16] In stark contrast with both the Gold Standard and the current regime, the period between 1972 and 1992 exhibits very high persistence based on each single series, with point estimates of ρ ranging from 0.88 to 0.91, and upper limits of 90% confidence intervals ranging between 0.98 and 1.04. Bretton Woods displays some evidence of serial correlation, with point estimates of ρ equal to 0.44 to 0.56, and upper limits of the 90% confidence intervals equal to 0.83, but the evidence is clearly nowhere nearly as strong and consistent as that for the 1972-1992 period. Finally, and intriguingly, the turbulent interwar period only displays some mild evidence of serial correlation.

2.2 Canada, Sweden, and New Zealand

Moving to the other three inflation-targeting countries (see Table 2) evidence clearly points, once again, against the notion of inflation as a *uniformly* highly persistent process. While the period between the collapse of Bretton Woods and the introduction of inflation targeting was indeed characterised by very high persistence for each single country and each inflation series—with high point estimates of ρ (with the exception of Sweden), and upper limits of the 90% confidence intervals ranging between 1.01 and 1.04—under inflation targeting serial correlation appears to have essentially disappeared. Focussing on the GDP deflator, in particular (which, since we are dealing with small open economies, should be regarded as the more appropriate measure of inflation) inflation is estimated to have been almost exactly white noise in Sweden and New Zealand, and to have exhibited very little persistence in Canada. Based on the CPI the picture changes very little, with negative serial correlation in Canada, and slightly positive correlation in the two other countries (to be noticed, though, that for Sweden estimates are so imprecise that we still can't reject the null of a unit root). As for the other regimes/periods, Bretton Woods exhibits very little persistence for both Sweden and New Zealand, while for Canada serial correlation appears to have been somewhat higher. Finally, while the interwar era does not exhibit any clear-cut pattern, evidence for Sweden under the Gold Standard points, once again, towards very little persistence.

2.3 United States

Let's finally turn to the U.S.. Table 3 presents what is, to the very best of our knowledge, the most extensive investigation to date on changes in U.S. inflation persistence

[15] This is consistent with the notion that the current monetary regime contains, *de facto*, a slight component of mean-reversion in the price level.

[16] It is interesting to notice that, compared to the results reported in Benati's (2006) Table B, for the inflation-targeting regime the extent of negative serial correlation has increased for all the three series. This suggests that the evidence is becoming stronger as the data accumulate.

over time and across monetary regimes since the Colonial era. Several findings clearly emerge from the table. In particular, persistence appears to have been virtually absent under metallic standards[17]—either *de facto* or *de jure*, and based on either gold or silver—with Hansen MUB estimates of ρ ranging from -0.06 to 0.24, and the lower and upper limits of bootstrapped 90% confidence intervals ranging from -0.27 to 0.37 respectively. The Colonial era, too, exhibited no persistence whatsoever, but given the sheer peculiarity of Colonial monetary arrangements[18], the implications of this finding are not clear-cut. Finally, the 'greenback period', during which the gold standard was temporarily suspended, and the U.S. was operating under a fiat money regime, exhibits very little persistence too. Such a finding appears as especially intriguing given that this period comprises the single most catastrophic event in U.S. history, the Civil War.[19]

As for the interwar period point estimates of ρ, ranging from 0.49 to 0.77, decisively point towards higher persistence than under metallic standards, but the extent of serial correlation is nowhere nearly as comparable to that prevailing over the post-WWII era.

Turning to the post-WWII era, results for the full Bretton Woods regime (December 1946-August 1971) point towards a comparatively low extent of persistence, while those for the the more proper sub-sample March 1951-August 1971 (following the FED-Treasury Accord) uniformly point towards higher persistence, with larger point estimates, and three series for which it is not possible to reject the null of a unit root. The Great Inflation episode clearly appears to have been characterised, to date, by the highest persistence in the entire U.S. history, with only one series out of five for which it is possible to marginally reject the null of a unit root at the 90% level.[20]

2.3.1 The controversy between Cogley-Sargent and Stock-Pivetta-Reis

Finally, as for the period following the end of the Volcker stabilisation, for which the controversy over a possible decrease in persistence is more heated,[21] our results suggest one possible explanation for the difference between the results of Cogley and Sargent and those of Stock, and Pivetta and Reis. As the table shows, over the most recent period there has been a quite significant difference between persistence estimates

[17] The white-noiseness of U.S. inflation under the Classical Gold Standard had already been documented by, e.g., Shiller and Siegel (1977) and Barsky (1987).

[18] On this, see e.g. Smith (1985a) and Smith (1985b).

[19] One possible explanation for the lack of serial correlation during the greenback period is that, under rational expectations, the expectation that after the Civil War the Gold Standard would be restored at the prewar parity—as it actually happened in January 1879—would have been sufficient to anchor inflation expectations, thus preventing inflation from 'taking off'.

[20] This is in line with the finding of Cogley and Sargent (2002) and Cogley and Sargent (2005) of a peak in U.S. inflation persistence around the end of the 1970s.

[21] See in particular Cogley and Sargent (2002), Cogley and Sargent (2005) and Cogley and Sargent (2006) on the one hand, and Stock (2002) and Pivetta and Reis (2006) on the other.

based on the CPI—the price index used by Cogley and Sargent (2002, 2005)—and those based on either the GDP (GNP), or the PCE deflators. While the CPI points towards a fall in persistence compared to the Great Inflation episode, both the GDP deflator (the price index used by Stock (2002) and Pivetta and Reis (2006)) and the PCE deflator consistently suggest an *increase* in terms of both point estimates, and the upper limits of the 90% confidence intervals. Further, based on neither the GDP nor the PCE deflator it is possible to reject the null of a unit root in inflation.[22] Our results therefore suggest that, different from what has been conjectured by Stock (2002), the difference between the results of Cogley and Sargent and those of Stock and Pivetta and Reis is not the product of the respective econometric methodologies—Bayesian as opposed to Classical—but rather of the quite significant divergence, over the most recent years, between the serial correlation properties of the CPI and those of other price indices.

2.4 High statistical persistence?

The results reported in Tables 1-3 provide a sobering assessment of the widespread notion of inflation as a very highly persistent process, suggesting such a notion to find its origin in the macroeconomic profession's dominant focus on one country, the U.S., and one specific sample period, the post-WWII era. As our results show, indeed, first, even focusing exclusively on the U.S., before WWII inflation's serial correlation has been either non-existent (under metallic standards) or comparatively lower (during the interwar era), while under the full Bretton Woods regime (1946-1971) it clearly appears to have been lower than post-1971.

Second—and crucially—the current experience of inflation targeting countries clearly shows that the widespread notion that inflation 'inertia [...] is central to inflation dynamics *in modern economies*'[23] is mistaken: Canada, Sweden, the U.K., and New Zealand *are* modern economies, but in these countries inflation is today (close to) white noise. To put it differently, historically lack of persistence has *not* been uniquely associated with the distant experience of metallic standards, and it is here with us today ...

[22] An important point to stress, however, is that all of our estimates ignore the possibility of a break in the mean of inflation within individual sub-samples (while they allow each sub-sample to have its own mean). As shown by Levin and Piger (2003), allowing for a break in the mean around the beginning of the 1990s significantly decreases persistence estimates.

[23] See Ball, Mankiw, and Reis (2005) (emphasis added).

3 Structural Evidence from Estimated Sticky-Price DSGE Models

Although intriguing, the reduced-form results reported in the previous section are subject to the obvious criticism that, as stressed by, e.g., Sheedy (2006), 'statistical inflation persistence is neither necessary nor sufficient for structural inflation persistence'[24] which, at the end of the day, is what truly matters, especially for monetary policy purposes. In the spirit of the recent work of Linde (2005) and Rabanal and Rubio-Ramirez (2005), in this section we therefore proceed to a structural investigation based on estimated sticky-price DSGE models featuring hybrid Phillips curves. Our main goal is to ascertain whether the intrinsic, structural component of inflation persistence—captured, in hybrid Phillips curves, by the extent of backward-looking indexation—does or does not change systematically across monetary regimes. Evidence that the extent of backward-looking indexation is not structurally stable across different regime would indeed provide clear evidence that—in line with Woodford's (2006) quotation in Section 1.1—intrinsic persistence is not structural in the sense of Lucas (1976).

3.1 The model

The model we use in what follows is given by

$$y_t = \gamma y_{t+1|t} + (1-\gamma)y_{t-1} - \sigma^{-1}(R_t - \pi_{t+1|t}) + \epsilon_{y,t}, \qquad \epsilon_{y,t} = \rho_y \epsilon_{y,t-1} + \tilde{\epsilon}_{y,t} \tag{2}$$

$$\pi_t = \frac{\beta}{1+\alpha\beta}\pi_{t+1|t} + \frac{\alpha}{1+\alpha\beta}\pi_{t-1} + \kappa y_t + \epsilon_{\pi,t}, \qquad \epsilon_{\pi,t} \sim WN(0, \sigma_\pi^2) \tag{3}$$

where π_t and y_t, are inflation and the output gap,[25] γ is the forward-looking component in the intertemporal IS curve, α is price setters' extent of indexation to past inflation, and $\epsilon_{\pi,t}$ and $\epsilon_{y,t}$ are reduced-form disturbances to the two variables. (All of the variables in (2)-(3) are expressed as log-deviations from a non-stochastic steady-state.) By imposing the white noiseness of $\epsilon_{\pi,t}$ in (3) we are essentially 'stacking the cards against ourselves', thus forcing all existing inflation persistence to be absorbed by the structural component. Our key result of (near) absence of structural persistence under inflation-targeting and the Gold Standard will therefore be, in a sense, all the more striking ...

For the post-WWII era we close the model with a Taylor rule with smoothing,

$$R_t = \rho R_{t-1} + (1-\rho)[\phi_\pi \pi_t + \phi_y y_t] + \epsilon_{R,t}, \qquad \epsilon_{R,t} = \rho_R \epsilon_{R,t-1} + \tilde{\epsilon}_{R,t} \tag{4}$$

[24] For an effective discussion of why this is the case, see Sheedy (2006, Section 4.2).

[25] The robustness of our results to replacing the output gap with a measure of the marginal cost is currently being investigated

where R_t is the nominal rate, and $\epsilon_{R,t}$ is a disturbance to the monetary rule, while for the Gold Standard we close it with an AR(1) specification for the rate of growth of base money, and a money demand equation coming from log-linearisation of the first-order conditions on the money-bonds market,

$$\mu_t = \rho_\mu \mu_{t-1} + \epsilon_{\mu,t}, \qquad \epsilon_{\mu,t} = \rho_\mu \epsilon_{\mu,t-1} + \tilde{\epsilon}_{\mu,t} \tag{5}$$

$$m_t - p_t = \psi y_t + \delta R_t + \epsilon_{MD,t}, \qquad \epsilon_{MD,t} = \rho_{MD} \epsilon_{MD,t-1} + \tilde{\epsilon}_{MD,t} \tag{6}$$

with $\epsilon_{\mu,t}$ and $\epsilon_{MD,t}$ being a shock to base money growh and a money demand disturbance respectively.

Our specification for the monetary rule under the Gold Standard—postulating that the money stock evolves according to an *exogenous* process—deserves some discussion. As it is well known,[26] under the Gold standard the evolution of the stock of base metal was indeed partly exogenous and partly endogenous. The former component reflected exogenous influences on gold production, e.g., the invention of the cyanide process in the second half of the XIX century, or the discovery of California's gold fields. The latter originated from the self-correcting mechanism intrinsic to metallic standards—see, e.g., Fisher (1922) and Barro (1979)—with a negative shock to the price level giving rise to both an increase in extraction activity, and a switch of base metal from non-monetary to monetary uses. Finally, it is important to stress that, *in the short-run*, the central banks of the countries that adhered to the Gold Standard tended to smooth the impact of changes in the stock of gold on corresponding changes in the stock of base money, so that the short-run link between the two was not one-to-one. (This argument, however, does not apply to the United States, as the Federal Reserve System was founded in 1913, just a few months before the collapse of the Gold Standard.) So our assumption that the stock of base money evolved according to an entirely exogenous process must necessarily be regarded as an approximation. The main justification for adopting this approach (essentially, a short-cut) is that a more satisfactory approach would have required to set up a sticky-price DSGE model of a commodity standard, which is beyond the scope of this paper.

Finally, given that the model is a closed-economy one, while the inflation-targeting countries are small open economies, for all countries and sample periods we estimate it based on GDP deflator inflation, which is the proper measure of domestically-generated inflation. The sample periods are the following. As for the full post-WWII period, 1954:3-2005:4 for the U.S., 1963:1-2005:4 for the U.K., and 1957:1-2005:4 for Canada. As for the Gold Standard, 1875:1-1914:2 for the U.S., 1834-1913 for the U.K., and 1861-1913 for Sweden.

3.2 Bayesian estimation

We estimate (2)-(3), augmented either by (4) or by (5)-(6), *via* Bayesian methods. Our preference for the use of a Bayesian, as opposed to a Classical approach is motivated

[26] See e.g. Barro (1979) and Barsky and Summers (1988).

by the fact that estimation attempts based on pure maximum likelihood tended to produce, in general, fragile results, with some parameters' estimates 'drifting away' towards completely implausible values. So, as we discuss in the next sub-section, we only use MLE results to calibrate the priors for the standard deviations of the structural disturbances—for which we obtained, in general, sensible estimates—but we make sure to make these prior sufficiently uninformative, calibrating the prior's standard deviation to a comparatively large value.

The following four sub-sections describe our choices for the priors, the numerical optimisation algorithm we use to numerically maximise the log-posterior, the Random-Walk Metropolis algorithm we use to get draws from the posterior, and the methodology we use to calibrate the covariance matrix's scale factor in the Random-Walk Metropolis algorithm.

3.2.1 Priors

Following, e.g., Lubik and Schorfheide (2004) and An and Schorfheide (2006), all structural parameters are assumed, for the sake of simplicity, to be *a priori* independent from one another. Table 4 reports the parameters' prior densities, together with two key objects characterising them, the mode and the standard deviation. While most of our choices are standard in the literature, a few details of our parameterisation deserve some discussion.

First, different from the vast majority of the literature, we calibrate both the Gamma and the inverse Gamma prior densities in terms of the *mode* of the distribution, rather than in terms of the *mean*—specifically, we calibrate the densities so that our 'preferred values' for the parameters of interest are equal to the mode. The key reason for doing so is in order to give the maximal amount of prior weight to our 'preferred values', which, on the other hand, would not be the case if calibration were performed in such a way as to make the densities' means equal to such values.[27]

Second, as we mentioned, we calibrate the inverse Gamma densities for the variances of the structural shocks—σ_R^2, σ_π^2, σ_y^2, and σ_{MD}^2—based on the results from maximum likelihood estimation, setting the densities' modes equal to the MLE estimates, and the standard deviations to 10.

Finally, different from several papers in the literature,[28] we adopt a perfectly flat (i.e., uniform) prior for the indexation parameter, α. The key reason for doing so is

[27] A simple example may clarify the practical implications of doing this. Consider the parameter κ, for which, based on results from previous studies—see e.g. Table 5 in Linde (2005)—we hold a prior that the 'right' value should be around 0.05. If we calibrated the prior for κ by setting the mean of the Gamma prior density to 0.05 (and the same standard deviation reported in table 4), we would end up with a prior with a mode at 0.0375, equal to 75% of the value we regard as the 'right one'. This would automatically imply that the Random Walk Metropolis algorithm we use in order to generate draws from the posterior distribution of the model's structural parameters would inevitably 'drift away' from our preferred value.

[28] See e.g. Smets and Wouters (2007).

in order to 'let the data speak' as freely as possible on the crucial feature of interest.

3.2.2 Getting draws from the posterior *via* Random-Walk Metropolis

We numerically maximise the log posterior—defined as $\ln L(\theta|Y) + \ln P(\theta)$, where θ is the vector collecting the model's structural parameters, $L(\theta|Y)$ is the likelihood of θ conditional on the data, and $P(\theta)$ is the prior—*via* a two-step numerical optimisation procedure combining simulated annealing, in the first step, and the Nelder and Mead (1965) simplex algorithm, in the second. (For a full description of the methodology, see Appendix C.1.) We then generate draws from the posterior distribution of the model's structural parameters *via* the Random Walk Metropolis (henceforth, RWM) algorithm as described in, e.g., An and Schorfheide (2006). In implementing the RWM algorithm we exactly follow An and Schorfheide (2006, Section 4.1), with the single exception of the method we use to calibrate the covariance matrix's scale factor—the parameter c below—for which we follow the methodology we describe in Appendix C.2.

Let then $\hat{\theta}$ and $\hat{\Sigma}$ be the mode of the maximised log posterior and its estimated Hessian, respectively.[29] We start the Markov chain of the RWM algorithm by drawing $\theta^{(0)}$ from $N(\hat{\theta}, c^2\hat{\Sigma})$. For $s = 1, 2, ..., N$ we then draw $\tilde{\theta}$ from the proposal distribution $N(\theta^{(s-1)}, c^2\hat{\Sigma})$, accepting the jump (i.e., $\theta^{(s)} = \tilde{\theta}$) with probability $\min\{1, r(\theta^{(s-1)}, \theta|Y)\}$, and rejecting it (i.e., $\theta^{(s)} = \theta^{(s-1)}$) otherwise, where

$$r(\theta^{(s-1)}, \theta|Y) = \frac{L(\theta|Y)\ P(\theta)}{L(\theta^{(s-1)}|Y)\ P(\theta^{(s-1)})}$$

We run a burn-in sample of 200,000 draws which we then discard. After that, we run a sample of 500,000 draws, keeping every draw out of 10 in order to decrease the draws' autocorrelation.

3.3 Empirical evidence

Figure 2 plots the posterior distributions for the indexation parameters for all the countries and sample periods, while Tables 4 and 5 report the modes of the posterior distributions for all the model's structural parameters, together with the 90%-coverage confidence intervals, for the post-WWII era and the Gold Standard, respectively.

Starting from the United States, consistent with, e.g., Linde (2005) structural persistence is estimated to have been very high over the full post-WWII sample period[30], with the mode at 0.827 and the 90% confidence interval equal to [0.749;

[29] We compute $\hat{\Sigma}$ numerically as in An and Schorfheide (2006).

[30] Based on full-information maximum likelihood and GDP deflator inflation, Linde (2005, Table 5) estimates the forward-looking component of the New Keynesian Phillips curve at 0.282, with a standard deviation on 0.057.

0.899], while in line with Gali and Gertler (1999b) it is estimated to have decreased—although not dramatically—after the Volcker stabilisation, with the mode falling to a still respectable 0.511. The contrast between the post-WWII U.S. experience and the alternative experiences of the Gold Standard and of inflation-targeting countries is simply striking. First, with the partial exception of Canada, which still exhibits a mild backward-looking component, under inflation targeting intrinsic persistence *appears to have all but disappeared*. Second, under the Gold Standard the degree of indexation is estimated to have been uniformly *close to zero* for all the three countries, with the modes of the distributions ranging between 0.023 and 0.087, and the upper limits of the 90%-coverage confidence intervals between 0.053 and 0.111. Finally, it is important to stress how, in line with the United States, both the United Kingdom and Canada exhibit very high structural persistence based on the full post-WWII sample period.

The results reported in Figure 2 and in Tables 4-5 suggest to us several reflections. In particular

- they question the notion that intrinsic persistence is *structural in the sense of Lucas (1976)*. This has important implications for both macroeconomic modelling and policy that we will discuss in Section 5.
- In line with Coenen and Levin (2004) these results show that under stable monetary regimes with clearly defined nominal anchors intrinsic persistence is *not necessary to fit the data*, so that—to a first approximation, and from a strictly macroeconomic point of view[31]—purely forward-looking New Keynesian Phillips curves fare (near) perfectly well.
- The contrast between the post-Bretton Woods U.S. experience and the experience of inflation-targeting countries suggests to us that—as far as inflation persistence is concerned—there has been something peculiar about the post-Bretton Woods U.S.,[32] so that such experience should not be regarded as the benchmark for understanding inflation dynamics. Rather, the benchmark should be taken to be the experience of inflation-targeting countries.

3.3.1 Robustness to alternative priors for the coefficient on inflation in the monetary rule

One important dimension along which these results ought to be checked is robustness to alternative priors for the coefficient on inflation in the monetary rule. Although our prior for ϕ_π, centered at 1.5, is standard in the literature, in principle one could

[31] By this we mean that, as it is well known, there are several microeconomic stylised facts concerning price-setting behavior that simple New Keynesian models fail to replicate.

[32] In line with Aoki and Kimura (2005), in Section 4 we will conjecture that this 'something' may have been the lack of a clearly defined nominal anchor, which has been, so far, the hallmark of the post-Bretton Woods experience.

envisage a situation in which in reality the central bank were significantly more aggressive, so that having a less aggressive prior than in reality might bias downward the estimates of the extent of indexation. In order to check for this possibility we have re-estimated the models for the four inflation-targeting countries with two alternative modes for the prior distribution of ϕ_π, 3 and 4.5. Results, reported in Table 7, clearly show our previously discussed findings to be robust along this important dimension.

3.3.2 Robustness to using marginal cost measures

A second important dimension along which the robustness of these results should be checked is the use of marginal cost measures in the Phillips curve equation—after all, the authentic driving variable in the New Keynesian Phillips curve is the marginal cost, and only under special circumstances it can legitimately be replaced by the output gap. While we have not been able to obtain reasonable proxies for the marginal cost for all countries and all sample periods, for two countries, the United States and the United Kingdom, we have been able to construct measures of the labor share,[33] the most extensively used marginal cost proxy, for the post-WWII era. Table 8 reports the modes of the posterior distributions of the indexation parameters,[34] together with the 90%-coverage confidence intervals, both for the full post-WWII sample periods, and for the most recent sub-period, based on model (2)-(4), where (3) has been replaced by

$$\pi_t = \frac{\beta}{1+\alpha\beta}\pi_{t+1|t} + \frac{\alpha}{1+\alpha\beta}\pi_{t-1} + \kappa mc_t + \epsilon_{\pi,t}, \qquad \epsilon_{\pi,t} \sim WN(0, \sigma_\pi^2) \tag{7}$$

where mc_t is the marginal cost. As the table clearly shows, replacing the output gap with the labor share does not produce any significant change in the previously discussed results. Focussing on the modes of the distributions, in particular, as for the United States indexation is still estimated to have been remarkably high, in excess 80%, over the full sample period, and to have instead been lower, but still comparatively high, after the Volcker disinflation. As for the United Kingdom, the dramatic contrast between the full-sample results and those for the inflation-targeting regime is still there, thus clearly showing that—at least for these two countries—our results are completely independent of the specific driving variable that is used in the Phillips curve.

[33] For a detailed discussion of how we constructed the labor share measures, see Appendix A.

[34] For reasons of space we do not report results for all parameters. The full set of results is however available from the author upon request.

4 Why Is Inflation Still Persistent in the United States?

As we pointed out, when compared with the experience of inflation-targeting countries the post-Bretton Woods U.S. experience appears, in a sense, peculiar. *Why*, exactly, is inflation still persistent in the United States, even under a regime that, for the last two decades and a half has been clearly oriented at fighting inflation? Although providing an answer to this question is beyond the scope of this paper, we conjecture, in line with Aoki and Kimura (2005), that the lack of a clearly defined nominal anchor, which has been so far the hallmark of post-Bretton Woods U.S., might have played a crucial role.

As Aoki and Kimura (2005) have shown, indeed, in the absence of a clearly defined anchor (in their specific case, of a publicly announced and credible numerical target for inflation) the public's inevitable learning process about the central bank's inflation objective causes *higher order expectations* to become relevant, thus dramatically increasing inflation's persistence and volatility. The key mechanism at work here is a sort of 'hall of mirrors' effect. On the one hand the public has to learn about the central bank's inflation objective. On the other hand the central bank, in order to filter out, for stabilisation purposes, movements in the Wicksellian rate of interest, ought to estimate the public's estimate of its own inflation objective. As Aoki and Kimura (2005) show, this mechanism is capable of generating high inflation persistence even within a model without intrinsic persistence, and conditional on white noise structural shocks.

5 Implications for Macroeconomic Modelling and Policy

Suppose that this paper's position is correct, so that (*i*) inflation is not intrinsically persistent (i.e., it is purely forward-looking), and (*ii*) the backward-looking price setting behavior found in post-WWII U.S. data reflects the lack, over the sample period, of a stable and clearly defined nominal anchor. This has several important implications, which we discuss in turn.

5.1 Evaluating alternative monetary regimes

In line with Woodford's opening quotation, the first danger of treating as structural something which is not has to do with assessing the desirability of adopting alternative monetary regimes. To make things concrete, suppose that Ben Bernanke wanted to assess the desirability of introducing an inflation (or a price level) target in the United States,[35] and therefore asked FED staff to investigate this issue *via* an estimated sticky-price DSGE model of the U.S. economy. If the model treated the persistence found in post-WWII U.S. as entirely structural, chances are that such an exercise would deliver an incorrect answer, precisely because, as conjectured by Woodford (2006), the very introduction of a clearly defined nominal anchor would cause persistence to disappear.

5.2 The width and shape of model-generated 'fan charts' for inflation

Suppose now that an inflation (or price level) target has been introduced in the United States, but the FED regards post-WWII inflation persistence as entirely structural, so that it hardwires it into a model used for generating fan charts for inflation. Given that, as it is well known, the extent of future uncertainty concerning a specific stochastic process is positively related to its persistence,[36] this would automatically imply that model-generated fan charts would exhibit the incorrect width and shape. Specifically, under the new regime the model would predict inflation uncertainty to be greater than it would in fact be, because the very introduction of the new regime, by 'killing off' persistence, would actually shrink the width of the fan charts, and would make their shape flatter.

6 Conclusions

Eleven years after the publication of Fuhrer and Moore (1995), a significant portion of the macroeconomic profession is still hard at work exploring new mechanisms to make the inflation persistence found in post-WWII U.S. data structural, i.e. intrinsic to the deep structure of the economy and impossible to eradicate without incurring a recession. Building on the experience of stable monetary regimes with clearly defined nominal anchors—historically, the Gold Standard and inflation targeting—in this paper

- we have documented the (near) white-noiseness of inflation in the U.K., Canada, Sweden and New Zealand under inflation targeting, and we have reasserted the well-known lack of serial correlation of inflation in the U.S., the U.K. and Sweden under the Gold standard; and
- we have shown how estimates of the indexation parameter in hybrid New Keynesian Phillips curves are below 0.3 in Canada under inflation targeting, and

[35] Let's forget that decisions over the details of the U.S. monetary constitution are in the hands of the Congress ...

[36] For example, for a white noise process, uncertainty is identical at all future horizons, while for a random walk it increases linearly in the length of the horizon.

virtually *zero* in both the other three inflation-targeting countries, and in the U.S., the U.K. and Sweden under the Gold standard.

These results question the notion that the intrinsic inflation persistence found in post-WWII U.S. data—captured, in hybrid New Keynesian Phillips curves, by a significant extent of backward-looking indexation—is structural in the sense of Lucas (1976). Further, they suggest that 'hardwiring' intrinsic persistence in macroeconomic models is potentially misleading. In particular, assessing alternative monetary regimes based on models featuring structural persistence might deliver incorrect results.

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A The Data

A.1 United States

A monthly seasonally unadjusted series for the wholesale price index, all commodities, from Warren and Pearson (1933), is available from July 1748 to December 1932 (the periods March 1782-December 1784, January 1788-December 1788, and January 1792-March 1793 are missing). A monthly seasonally unadjusted 'index of the general price level', available from January 1860 to November 1939, is from the *NBER Historical Database* on the web (the series' NBER code is 04051; the original data are from the Federal Reserve Bank of New York for the period 1860-1933, and from the *Monthly Review of Credit and Business Conditions* after that). A quarterly seasonally adjusted series for the GNP deflator has been constructed by linking the GNP deflator series from Balke and Gordon (1986), appendix B, Table 2, available from 1875:1 to 1983:4, to the series for the GNP implicit price deflator from the U.S. Department of Commerce's *Bureau of Economic Analysis* (henceforth, *BEA*), which starts in 1947:1. The overall sample period is 1875:1-2005:4 (specifically, the linked series is made up of the Balke-Gordon index up to 1946:4, and of the latter index after that).[37] A monthly seasonally unadjusted series for the CPI for all urban consumers, all items (acronym is CPIAUCNS), is from the U.S. Department of Labor, *Bureau of Labor Statistics* (henceforth, *BLS*), and is available since January 1946. We linked this series to a seasonally unadjusted series for the CPI, all items, from the *NBER Historical Database* (NBER code is 04128)—specifically, the overall linked series is made up of the series from the NBER database up to December 1945, and of the series from the *BLS* after that. The overall sample period is January 1913-December 2005. The corresponding seasonally adjusted series has been constructed by linking the seasonally adjusted series for the CPI for all urban consumers from the *BLS* (CPIAUCSL), available since January 1947, to series 04128 from the *NBER Historical Database* after adjusting it *via* ARIMA X-12. The seasonally adjusted series for the GDP and the PCE deflators, both available for the period 1947:1-2005:4, are from the *BEA*.

We construct the labor share measure as in Groen and Mumtaz (2007), as

$$\begin{aligned} s_t &= \ln\left[\frac{W_t N_t}{P_t Y_t}\right] = \\ &= \ln\left[\frac{\text{Unambiguous Labor Income + Private Share Supplements}}{\text{GDP - Ambiguous Labor Income - Government Labor Income}}\right] \end{aligned}$$

Following Groen and Mumtaz (2007), 'Unambiguous Labour Income' is given by 'wages and salary accruals, other' (row B203RC1, Table 1.12 in the NIPA); 'Private Share Supplements' is equal to 'supplements to wages and salaries' (row A038RC1,

[37] One possible problem with this series is that, over the period 1875-1946, it has been interpolated *via* the Chow-Lin method, based on an annual GNP deflator series and a quarterly index for the wholesale price level (see Balke and Gordon, 1986, 'Section 2 Source Notes', p. 809).

Table 1.12 in NIPA) times the ratio of 'wages and salary accruals, other' (row B203RC1, Table 1.12 in NIPA) to 'wages and salary accruals' (row A034RC1, Table 1.12 in NIPA); GDP is nominal GDP (row A191RC1, Table 1.1.5 in NIPA); 'Ambiguous Labour Income' equals the sum of 'Proprietors' income with IVA and CCAdj' (row A041RC1, Table 1.12 in NIPA).and the difference between nominal GDP (row A191RC1, Table 1.1.5 in NIPA) and national income (row A032RC1, Table 1.12 in NIPA); 'Government Labor Income' equals the sum of 'Wages and salary accruals, government' (row A553RC1, Table 1.12 in NIPA).and 'Supplements to wages and salaries' (row A038RC1, Table 1.12 in NIPA) times the ratio of 'wages and salary accruals, government' (row A553RC1, Table 1.12 in NIPA) to 'wages and salary accruals' (row A034RC1, Table 1.12 in NIPA).

As for the Gold Standard period, we approximate the output gap by the HP-filtered logarithm of the real GNP series from Balke and Gordon (1986), appendix B, Table 2. As for the post-WWII period, we compute it as the difference between the logarithms of GDPC1 ('Real Gross Domestic Product, 1 Decimal'), from the *BEA*, and GDPPOT ('Real Potential Gross Domestic Product') from the Congressional Budget Office.

The interest rate is the Federal Funds rate from the Federal Reserve Board for the post-WWII period, and the commercial paper rate from Balke and Gordon (1986), appendix B, Table 2 for the Gold Standard. Finally, the base money series for the Gold Gold Standard is again from Balke and Gordon (1986), appendix B, Table 2.

A.2 United Kingdom

The price series are a small subset of the series analysed in Benati (2006), which we updated for the most recent period. The annual composite price index, available for the period 1750-2003, is from the *Office for National Statistics* (henceforth, ONS)—see O'Donoghue, Goulding, and Allen (2004). A quarterly series for the GDP deflator from the ONS is available since 1955:1. A GNP deflator series for the period 1830-1913 has been computed as the ratio between nominal and real GNP based on National Accounts' Tables 5 and 6 of Mitchell (1988). Monthly seasonally unadjusted series for the retail price index are available from Table III.(11) of Capie and Webber (1985) for the period July 1914-December 1982, and from the ONS for the period since June 1947. A monthly series for the CPI from the ONS is available since January 1975.

An annual series for real GNP, available for the period 1830-1913, is from Table 6 of Mitchell (1988). An annual series for the three-month rate for the period 1830-1913 has been computed by linking Gurney's rate for first-class three-months bills found in Mitchell (1988), available until 1844, and a three-months banks bills series found again in Mitchell (1988) after that. An annual series for base money forthe period 1833-1913 has been computed by linking the high-powered money series from Huffman and Lothian (1980) and the base money series from Capie and Webber (1985), available since 1870. A quarterly series for real GDP, available since 1955:1, and a monthly series for the average discount rate on Treasury bills, available since January 1963, are both from the ONS.

We construct the labor share measure as in Batini, Jackson, and Nickell (2005), as the ratio of the total compensation of employees to nominal GDP at factor costs, correcting both for self-employees' jobs, and for the presence of the government sector. (For a full description of the methodology and of the data sources, see the Appendix in Batini, Jackson, and Nickell (2000).) Specifically, the Batini et al.'s (2005) labor share measure is defined as

$$s_t = \ln\left\{\frac{[(HAEA_t \text{ - } NMXS_t) \cdot A_t]}{(ABML_t \text{ - } GGGVA_t)}\right\}$$

where $HAEA_t$ is the compensation of employees (including the social contributions payable by the employer); $ABML_t$ is is gross value added at current prices, measured at basic prices, excluding taxes less subsidies on products.; $NMXS_t$ is the compensation of employees paid by the general government; and $GGGVA_t$ is a measure of the part of gross value added attributable to the general government. When needed, both monthly interest rate and price series have been converted to the quarterly frequency by taking averages within the quarter and, respectively, by keeping the last observation from each quarter.

A.3 Canada

A monthly seasonally unadjusted series for the consumer price index, available from January 1914 to December 2005, is from *Statistics Canada*. Quarterly seasonally adjusted series for the GDP deflator and real GDP, and a series for the Treasury Bill Rate are from the IMF's *International Financial Statistics* (henceforth, *IFS*).

A.4 Sweden

A quarterly seasonally unadjusted series for the consumer price index, available from 1917:3 to 2005:4, is from *Statistics Sweden*. Quarterly seasonally adjusted series for the GDP deflator and real GDP, and a series for the interest rate on 3 months Treasury notes are from the IMF's *IFS*.

A.5 New Zealand

A quarterly seasonally unadjusted series for the consumer price index, available from 1925:3 to 2005:2, is from *Statistics New Zealand*. Quarterly seasonally adjusted series for the GDP deflator and real GDP, and a series for the interest rate on 3 months Treasury notes are from the IMF's *IFS*.

The analysis in the paper is entirely based on quarterly inflation rates. In the case of series originally available at the monthly frequencies, the relevant price indices have been converted to the quarterly frequency by keeping the last observation from each quarter.

B Identifying Monetary Regimes

This appendix discusses and motivates our choices of how to break the overall sample periods by monetary regimes.

B.1 United States

We consider the following breakdown of the overall sample period.

- *De facto* silver standard: from the U.S. Congress' 'Mint and Coinage Act' of April 2, 1792, up to the 'Act Regulating the Gold Content of U.S. Coins' of June 28, 1834. The former act established a bimetallic standard based on gold and silver, with a legal mint ratio of gold to silver of 15:1. At this ratio gold was undervalued at the mint compared with its market price, and the system soon degenerated into a *de facto* silver standard. The act of 1834 changed the mint ratio to 16:1, with the result that gold became now overvalued at the mint, and the U.S. switched to a *de facto* gold standard.

- *De facto* gold standard: from June 28, 1834 up to December 30, 1861, when, following the outbreak of the Civil War, the Treasury suspended convertibility of its notes into gold.[38]

- The greenback period: from December 30, 1861 until the gold standard resumption, on January 2, 1879.

- Classical Gold Standard: from January 2, 1879 up to the beginning of WWI, on August 6, 1914.

- The interwar period, from November 11, 1918, when Germany signed the armistice agreement with the Allies, up to the Declaration of War on Japan, on December 8, 1941. An obvious objection to such a choice is that a more appropriate breakdown of the interwar era would distinguish between the two sub-periods preceding and respectively following April 19, 1933, when President Roosevelt took the dollar off gold.[39] In what follows we therefore also report results for

[38] Although the Civil War had already started on April 12, 1861—the date of the attack on Fort Sumter on the part of Confederate troops—as pointed out by, e.g., Officer (1981), it was only at the end of December 1861 that, in the northern states, 'almost all banks ceased to convert their notes and deposits into gold coin', while on December 30 the Treasury 'suspended the rights of holders of its demand notes to redeem the notes in gold'.

[39] See Eichengreen (1992).

the sub-period November 1918-April 1933, which, following Meltzer (1986), we label as 'gold exchange standard with a central bank'.[40]

- The Bretton Woods regime, from December 18, 1946, the date in which the 32 member countries declared their par values *vis-a-vis* the U.S. dollar, up to August 15, 1971, when President Nixon finally closed the 'gold window'. An objection to adopting December 1946 as its official starting date of the Bretton Woods regime is that, until the Treasury-Federal Reserve accord of March 4, 1951, the FED was legally obliged to support the market for U.S. government's bonds, and could not therefore follow an independent monetary policy. In what follows we therefore also report results for the 1951-1971 sub-period.

- The Great Inflation episode, from the the collapse of Bretton Woods up to the end of the Volcker stabilisation, which, following, e.g., Clarida, Gali, and Gertler (2000), we date in the fourth quarter of 1982.

- The period following the Volcker stabilisation.

B.2 United Kingdom

Following Benati (2006), we divide the overall sample period into the following monetary regimes/historical periods.

- *De facto* gold standard, from 1718[41] up until the beginning of the suspension period associated with the wars with France, on February 26, 1797.

- *De jure* gold standard: from the gold standard official resumption, on May 1, 1821, up to the beginning of the second suspension period associated with the outbreak of WWI, on August 6, 1914.

- Interwar period: from the constitution of the Irish Free State as a British dominion,[42] on December 6, 1921, to the U.K.'s declaration of war on Germany, on September 3, 1939.

- Bretton Woods regime: from December 18, 1946 up to the floating of the pound *vis-a-vis* the U.S. dollar, on June 23, 1972.

[40] See Meltzer (1986, Table 4.1), although he adopts slightly different beginning and end dates. On the other hand, we ignore the period between the floating of the dollar and the declaration of war on Japan because of its comparatively short length.

[41] In 1717 the United Kingdom accidentally switched from a *de facto* silver standard to a *de facto* gold standard due to a mistake of the then Master of the Mint, Sir Isaac Newton, in fixing the official parity between gold and silver (the switch was therefore due to the operation of Gresham's law).

[42] For a discussion of why taking this specific date as the begininng of the interwar period, see Benati (2006).

- From June 23, 1972 to the introduction of inflation targeting, on 8 October 1992. An obvious objection to treating the 1972-1992 period as a single 'regime' is that it was characterised by a succession of several different frameworks and arrangements: the period of monetary targets; a brief period without any clearly defined monetary policy strategy; then the period of shadowing the Deutsche Mark; and finally, membership of the European Monetary System (EMS). The key reason behind our choice is simply a practical one: given the frequency of the changes in the monetary framework, breaking the sample period every few years would have made it impossible to perform any econometric analysis.

- Inflation targeting regime: from 8 October 1992 to the present.

B.3 Canada, Sweden, and New Zealand

As for Canada, Sweden, and New Zealand, we consider the following regimes/periods.

- Interwar period, from the armistice of November 11, 1918, up to WWII. In particular, for New Zealand and Canada we consider the date of their declaration of war on Germany, September 3, 1939 and respectively September 10, 1939. As for Sweden, which was neutral during the war, we preferred to exclude the WWII years for reasons of consistency with the other countries, and we therefore end the interwar period in September 1939.

- The Bretton Woods regime, from December 18, 1946 to August 15, 1971, the exception being Sweden, which joined Bretton Woods on November 5, 1951.[43] The case of Canada is unfortunately problematic, as it suspended the par value, thus allowing the Canadian dollar to float, on September 30, 1950, re-entering the system at a new parity on May 2, 1962. For practical reasons, in what follows we are going to treat the period between 1946 and 1971, for Canada, as a unique 'regime/period', but it is important to keep in mind that it comprises several alternative arrangements.

- The period between the collapse of Bretton Woods and the introduction of inflation targeting: on January 15, 1993 in Sweden; on February 1, 1990 in New Zealand; and on February 26, 1991 in Canada.

- The inflation targeting regime.

[43] See Bordo (1993).

C Two Technical Aspects of the Bayesian Estimation Procedure

This appendix discusses in detail two technical aspects the Bayesian estimation procedure we used in section 3.

C.1 Numerical maximisation of the log posterior

We numerically maximise the log posterior—defined as $\ln\, L(\theta|Y) + \ln\, P(\theta)$, where θ is the vector collecting the model's structural parameters, $L(\theta|Y)$ is the likelihood of θ conditional on the data, and $P(\theta)$ is the prior—*via* a two-step numerical optimisation procedure combining simulated annealing, in the first step, and the Nelder and Mead (1965) simplex algorithm, in the second.[44] Specifically, following Goffe, Ferrier, and Rogers (1994) we implement simulated annealing *via* the algorithm proposed by Corana, Marchesi, Martini, and Ridella (1987), setting the key parameters to T_0=100,000, r_T=0.9, N_t=5, N_s=20, ϵ=10^{-6}, N_ϵ=4, where T_0 is the initial temperature, r_T is the temperature reduction factor, N_t is the number of times the algorithm goes through the N_s loops before the temperature starts being reduced, N_s is the number of times the algorithm goes through the function before adjusting the stepsize, ϵ is the convergence (tolerance) criterion, and N_ϵ is number of times convergence is achieved before the algorithm stops. Finally, initial conditions were chosen stochastically by the algorithm itself, while the maximum number of functions evaluations, set to 1,000,000, was never achieved. For robustness reasons—i.e, in order to be reasonably sure that we had identified the global maximum of the log posterior—for each country, each sample, and each single model we implemented the two-step numerical optimisation routine 15 times,[45] choosing the solution associated with the maximum of the log posterior. With a few exceptions, results were, overall, remarkably stable across the 15 runs, thus clearly suggesting that our two-step procedure has indeed identified a global maximum.

[44] The reason for sequentially combining the two optimisation algorithms has to do with the fact that, according to our own experience, they should be regarded as largely complementary. Simulated annealing is excellent at producing a broad stochastic mapping of the function of interest in order to identify the most promising region, but is less good at 'closing in' on the minimum. The Nelder-Mead simplex algorithm, on the other hand, is extremely powerful at converging to the right solution once you are broadly in that neighborhood, but is less good at identifying a function's most promising region (and indeed, in large-scale optimisation problems, it suffers from a well-known tendency to 'get stuck' around the initial conditions). By identifying the most promising neighborhood *via* simulated annealing, and then 'closing in' *via* Nelder-Mead, we therefore use each algorithm to perform what it does best.

[45] Overall optimisation was performed by eight Pentium IV machines linked in a parallel computing network *via* MATLAB's Distributed Computing Toolbox^{TM}.

C.2 Calibrating the covariance matrix scale factor

A key problem in implementing Metropolis algorithms is how to calibrate the covariance matrix's scale factor—the parameter c in subsection 3.2.2—in order to achieve an acceptance rate of the draws close to the ideal one (in high dimensions) of 0.23.[46] Typically the problem is tackled by starting with some 'reasonable' value for c, and adjusting it after a certain number of iterations during the initial burn-in period. Specifically, given that the draws' acceptance rate is decreasing in c, c gets increased (decreased) if the initial acceptance rate was too high (low). A problem with this approach is that it does not guarantee that after the adjustment the acceptance rate will be reasonably close to the ideal one. The approach for calibrating c used in this paper, on the other hand, is based on the idea of *estimating* a reasonably good approximation to the inverse relationship between c and the acceptance rate by running a pre-burn-in sample. Specifically, let C be a grid of possible values for c—in what follows, we consider a grid over the interval [0.1, 1.3] with increments equal to 0.05. For each single value of c in the grid—call it c_j—we run n draws of the RWM algorithm as described in section 3.2.2, storing, for each c_j, the corresponding fraction of accepted draws, f_j. Based on n=2,000, Figure 1 shows the f_j's for the U.K., Canada, New Zealand and Sweden under inflation targeting, for the U.K., Sweden, and the U.S. under the Gold Standard, and for the U.S. for the period following the Volcker stabilisation. As the Figure makes clear, even a relatively low number of draws as 2,000 allows to trace out the inverse relationship between c and the acceptance rate with remarkable accuracy. We then fit a third-order polynomial to the f_j's *via* least squares, and letting $\hat{a}_0$, $\hat{a}_1$, $\hat{a}_2$, and $\hat{a}_3$ be the estimated coefficients, we choose c by solving numerically the equation $\hat{a}_0+\hat{a}_1c+\hat{a}_2c^2+\hat{a}_3c^3$=0.23.

[46] See Gelman, Carlin, Stern, and Rubin (1995).

Table 1 Inflation persistence in the United Kingdom: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

	De facto gold standard	*De jure* gold standard	Interwar period	Bretton Woods	Bretton Woods to inflation targeting	Inflation targeting
	Annual series:					
GNP/GDP deflator	–	0.05 [-0.13; 0.22]				
Composite price index	-0.17 [-0.62; 0.29]	-0.21 [-0.45; 0.02]				
	Quarterly series:					
Retail price index[a]	–	–	0.37 [-0.05; 0.80]	0.56 [0.33; 0.83]	0.91 [0.72; 1.03]	-0.10 [-0.79; 0.68]
Consumer price index					0.93 [0.89; 0.98]	-0.19 [-0.54; 0.15]
GDP deflator	–	–	–	0.44 [0.07; 0.83]	0.88 [0.70; 1.04]	-0.31 [-0.69; 0.10]

[a] For the interwar period, retail price index from Capie and Webber (1985). After that, from ONS.
Source: Benati (2006), Table B, updated for the inflation targeting period.

Table 2 Inflation persistence in Canada, New Zealand, and Sweden: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

Monetary regime/ historical period	*Canada*		*Sweden*	
	CPI[a]	GDP deflator[b]	CPI[a]	GDP deflator[b]
Gold standard				0.30 [0.03; 0.53]
Interwar period	0.58 [0.38; 0.81]		0.69 [0.42; 1.02]	
Bretton Woods	0.71 [0.54; 0.89]	0.77 [0.46; 1.05]	0.29 [-0.01; 0.59]	
1971 to inflation targeting	0.91 [0.72; 1.04]	1.00 [0.78; 1.04]	0.53 [0.12; 1.02]	
Inflation targeting	-0.25 [-0.73; 0.22]	0.34 [0.00; 0.73]	0.37 [-0.24; 1.05]	0.06 [-0.47; 0.63]
	New Zealand			
	CPI[a]	GDP deflator[b]		
Interwar period	0.98 [0.69; 1.06]			
Bretton Woods	0.29 [0.02; 0.56]			
1971 to inflation targeting	0.82 [0.67; 1.01]			
Inflation targeting	0.41 [0.14; 0.72]	0.01 [-0.63; 0.70]		

[a] Seasonally unadjusted. [b] Seasonally adjusted. *: RPI;#: CPI.

Table 3 Inflation persistence in the United States: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

	Seasonally unadjusted series:			
Monetary regime/ historical period	Warren-Pearson wholesale price index	Index of general price level	Consumer price index	
Colonial era	-0.04 [-0.20; 0.13]			
De facto silver standard	0.24 [0.11; 0.37]			
De facto gold standard	0.22 [0.06; 0.37]			
Greenback period	0.28 [0.08; 0.48]	0.43 [0.24; 0.62]		
Classical Gold Standard	0.11 [-0.04; 0.25]	0.05 [-0.09; 0.19]		
Interwar period		0.49 [0.32; 0.66]	0.70 [0.53; 0.88]	
'Gold exchange standard with a central bank'[a]	0.67 [0.49; 0.87]	0.52 [0.31; 0.75]	0.76 [0.56; 1.01]	
Bretton Woods			0.50 [0.32; 0.65]	
March 1951-August 1971			0.69 [0.49; 0.92]	
Great Inflation			0.72 [0.44; 1.02]	
Post-Volcker stabilisation			0.35 [0.12; 0.61]	
	Seasonally adjusted series:			
	Consumer price index	GNP deflator	GDP deflator	PCE deflator
Classical Gold Standard		-0.01 [-0.29; 0.26]		
Interwar period	0.72 [0.57; 0.89]	0.62 [0.44; 0.78]		
'Gold exchange standard with a central bank'[a]	0.77 [0.59; 1.01]	0.71 [0.32; 1.04]		
Bretton Woods	0.46 [0.30; 0.64]	0.63 [0.45; 0.81]	0.55 [0.36; 0.75]	0.54 [0.36; 0.72]
March 1951-August 1971	0.69 [0.49; 0.94]	1.02 [0.86; 1.10]	1.02.[0.86; 1.09]	0.91 [0.72; 1.04]
Great Inflation	0.78 [0.51; 1.03]	0.74 [0.48; 1.02]	0.74 [0.49; 1.02]	0.74 [0.53; 0.99]
Post-Volcker stabilisation	0.41 [0.17; 0.63]	0.91 [0.71; 1.05]	0.91 [0.70; 1.05]	0.84 [0.65; 1.03]

Results are based on the Hansen (1999) 'grid bootstrap' procedure. [a] See Meltzer (1986).

Table 4 Prior distributions for the structural parameters

Parameter	Domain	Density	Mode	Standard deviation
σ_R^2, σ_π^2, σ_y^2, σ_{MD}^2	$\mathbb{R}^+$	Inv. Gamma	$\sigma_{x,MLE}^2$	10
κ	$\mathbb{R}^+$	Gamma	0.05	0.025
σ	$\mathbb{R}^+$	Gamma	10	5
α	[0, 1]	Uniform	–	0.577
γ	[0, 1]	Uniform	–	0.289
ρ, ρ_R, ρ_π, ρ_y, ρ_μ, ρ_{MD}	(-1, 1)	Uniform	–	0.289
ϕ_π	$\mathbb{R}^+$	Gamma	1.5	0.25
ϕ_y	$\mathbb{R}^+$	Gamma	0.5	0.25
ψ	$\mathbb{R}^+$	Gamma	1	0.1
δ	$\mathbb{R}^+$	Gamma	0.5	0.05

Table 5 Post-WWII era, Bayesian estimates of the New Keynesian model's structural parameters, posterior mode and 90% confidence interval

	United States	United Kingdom	Sweden	Canada	New Zealand
	Full sample				
σ_R	1.120 [1.032; 1.157]	1.010 [0.945; 1.135]		0.932 [0.860; 1.036]	
σ_π	0.805 [0.790; 0.930]	2.784 [2.736; 2.892]		1.609 [1.502; 1.818]	
σ_y	0.205 [0.166; 0.275]	0.192 [0.190; 0.198]		0.179 [0.155; 0.237]	
κ	0.005 [0.003; 0.006]	0.027 [0.013; 0.064]		0.030 [0.017; 0.059]	
σ	28.705 [20.630; 32.311]	33.293 [25.746; 46.575]		32.919 [25.698; 34.054]	
α	0.827 [0.749; 0.899]	0.653 [0.589; 0.700]		0.716 [0.608; 0.781]	
γ	0.624 [0.577; 0.704]	1.000 [0.934; 0.998]		0.901 [0.791; 0.979]	
ρ	0.827 [0.801; 0.870]	0.905 [0.878; 0.919]		0.882 [0.861; 0.905]	
ϕ_π	1.312 [1.165; 1.678]	0.919 [0.798; 0.990]		1.077 [0.942; 1.407]	
ϕ_y	0.710 [0.598; 1.157]	1.992 [1.465; 2.769]		1.621 [1.170; 2.280]	
ρ_R	0.122 [0.027; 0.257]	0.087 [0.064; 0.153]		0.049 [0.030; 0.066]	
ρ_y	0.715 [0.584; 0.806]	0.844 [0.823; 0.860]		0.796 [0.733; 0.845]	
	Post-1982	*Inflation targeting*			
σ_R	0.746 [0.689; 0.903]	0.478 [0.414; 0.587]	1.001 [0.887; 1.237]	1.064 [1.020; 1.125]	1.171 [1.037; 1.415]
σ_π	0.612 [0.555; 0.766]	1.572 [1.344; 1.898]	1.254 [1.159; 1.378]	1.901 [1.699; 2.156]	3.186 [2.852; 3.790]
σ_y	0.187 [0.160; 0.251]	0.089 [0.071; 0.130]	0.295 [0.219; 0.453]	0.100 [0.071; 0.114]	0.339 [0.285; 0.612]
κ	0.013 [0.007; 0.027]	0.043 [0.025; 0.098]	0.051 [0.024; 0.095]	0.063 [0.037; 0.126]	0.056 [0.029; 0.113]
σ	9.109 [6.767; 23.111]	17.276 [12.619; 37.214]	17.421 [11.674; 30.756]	28.677 [21.081; 37.429]	26.324 [16.422; 40.496]
α	0.511 [0.333; 0.756]	0.016 [0.010; 0.250]	0.004 [0.003; 0.125]	0.286 [0.137; 0.432]	0.023 [0.005; 0.155]
γ	0.726 [0.650; 0.866]	0.660 [0.511; 0.870]	0.993 [0.830; 0.995]	0.503 [0.463; 0.628]	0.462 [0.460; 0.963]
ρ	0.839 [0.753; 0.879]	0.930 [0.883; 0.952]	0.924 [0.869; 0.945]	0.867 [0.819; 0.903]	0.895 [0.832; 0.921]
ϕ_π	1.530 [1.194; 2.047]	1.190 [0.956; 1.643]	1.312 [1.010; 1.706]	1.372 [1.105; 1.860]	1.186 [0.923; 1.592]
ϕ_y	0.734 [0.468; 1.503]	0.943 [0.556; 1.790]	0.738 [0.407; 1.459]	0.782 [0.665; 0.890]	1.204 [0.626; 1.920]
ρ_R	0.467 [0.275; 0.653]	0.126 [0.026; 0.326]	0.054 [0.011; 0.241]	0.032 [0.024; 0.034]	0.072 [0.018; 0.282]
ρ_y	0.815 [0.689; 0.868]	0.769 [0.605; 0.839]	0.692 [0.546; 0.810]	0.664 [0.594; 0.820]	0.669 [0.060; 0.741]

Table 6 Gold Standard era: Bayesian estimates of the New Keynesian model's structural parameters, mode and 90% confidence interval

	United States	United Kingdom	Sweden
σ_π	2.817 [2.484; 3.587]	3.132 [2.701; 3.827]	2.812 [2.575; 3.656]
σ_y	0.993 [0.815; 1.691]	1.376 [1.105; 1.949]	0.806 [0.675; 1.182]
σ_μ	6.251 [5.531; 7.212]	6.615 [6.040; 7.831]	10.616 [9.271; 11.662]
σ_{MD}	4.616 [4.028; 5.704]	2.149 [2.007; 2.388]	9.315 [8.449; 11.054]
κ	0.028 [0.017; 0.058]	0.052 [0.034; 0.091]	0.031 [0.020; 0.044]
σ	0.404 [0.335; 0.526]	0.425 [0.340; 0.513]	0.435 [0.366; 0.552]
α	0.087 [0.025; 0.111]	0.023 [0.003; 0.053]	0.046 [0.018; 0.064]
γ	1.000 [0.943; 1.000]	1.000 [0.921; 0.998]	1.000 [0.949; 1.000]
ρ_μ	0.480 [0.255; 0.600]	-0.046 [-0.148; 0.348]	0.423 [0.318; 0.505]
ψ	0.794 [0.688; 0.974]	0.826 [0.708; 0.974]	0.772 [0.680; 0.954]
δ	0.698 [0.619; 0.805]	0.714 [0.623; 0.811]	0.731 [0.658; 0.848]
ρ_y	0.306 [-0.061; 0.614]	0.605 [0.354; 0.807]	0.749 [0.556; 0.906]
ρ_{MD}	0.349 [0.181; 0.520]	0.265 [0.101; 0.479]	0.472 [0.325; 0.543]

Table 7 Robustness checks for the inflation-targeting countries: Bayesian estimates of α for alternative priors for ϕ_π, mode and 90% percentiles

Mode of the prior for ϕ_π:	United Kingdom	Sweden	Canada	New Zealand
3	0.001 [0.009; 0.320]	0.000 [0.002; 0.136]	0.279 [0.118; 0.511]	0.000 [0.003; 0.141]
4.5	0.040 [0.012; 0.355]	0.000 [0.003; 0.149]	0.283 [0.109; 0.489]	0.000 [0.003; 0.146]

Table 8 Robustness to using marginal cost proxies: posterior modes for α and 90% percentiles

	United States		United Kingdom	
	Full sample	*Post-1982*	*Full sample*	*Inflation targeting*
α	0.805 [0.744; 0.866]	0.457 [0.295; 0.612]	0.617 [0.535; 0.696]	0.024 [0.003; 0.056]

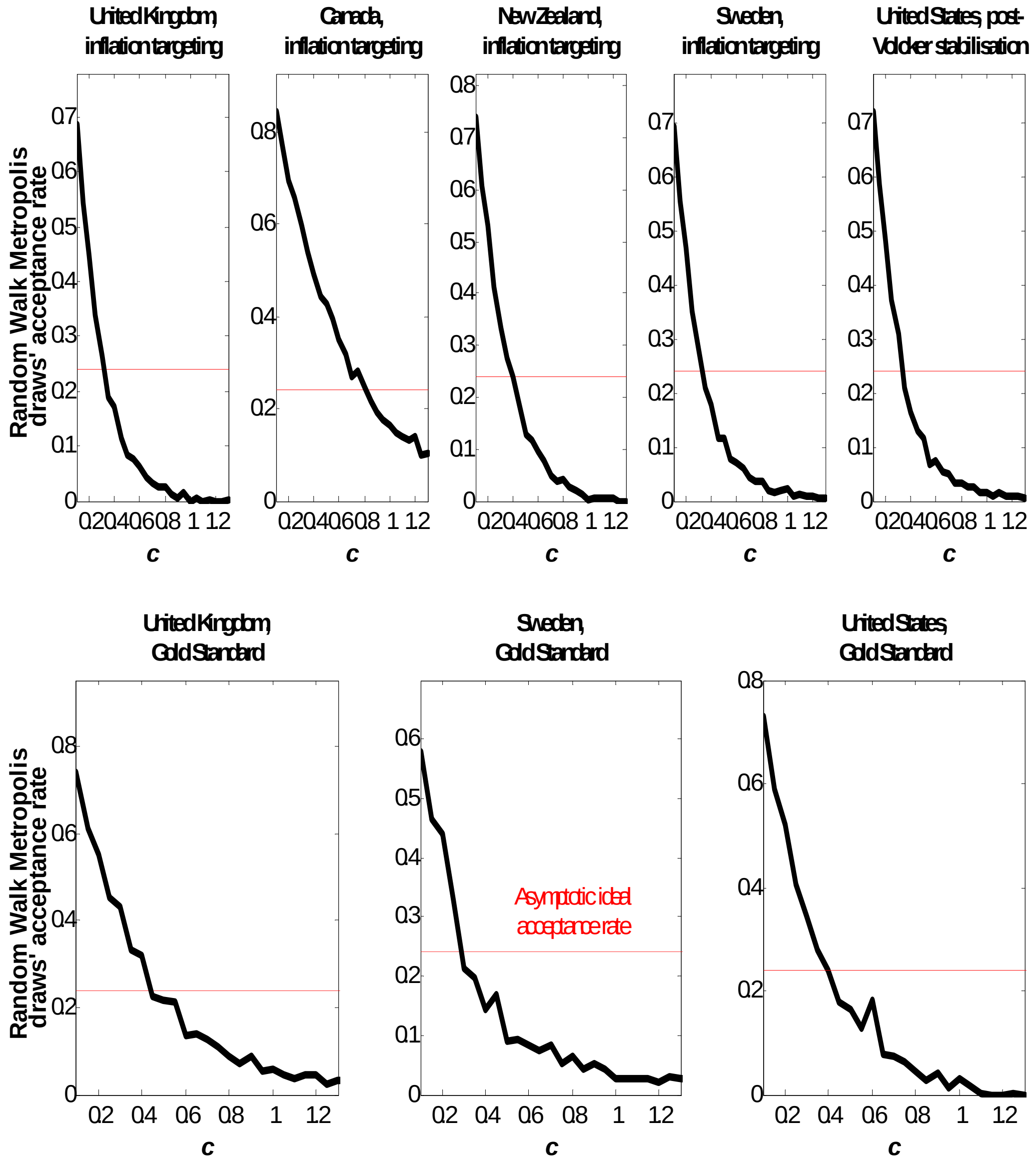

Figure 1: Calibrating the covariance matrix scale factor in the Random Walk Metropolis algorithm

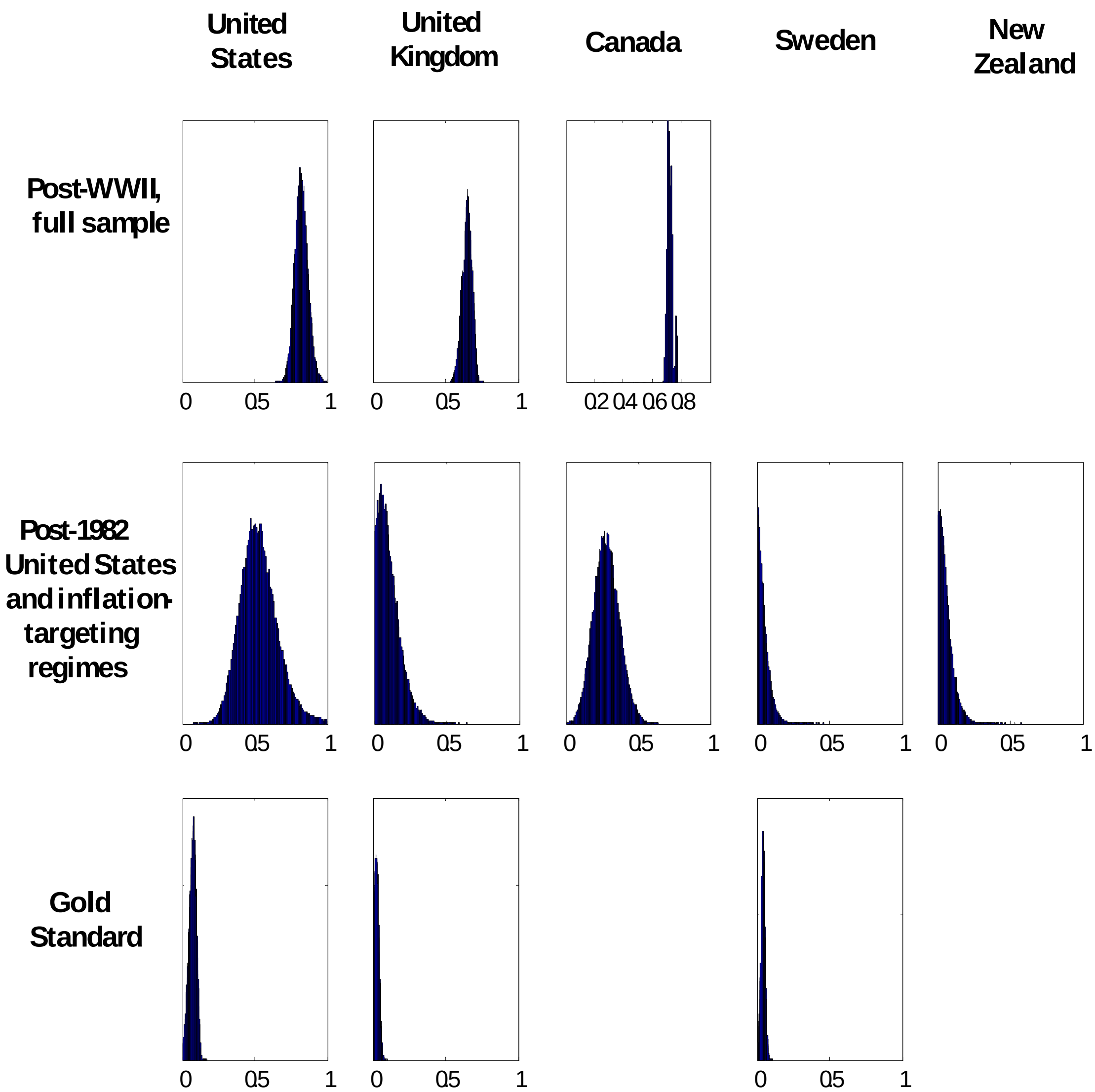

Figure 2: Posterior distributions for the indexation parameter, α