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In 2008 all ECB publications feature a motif taken from the €10 banknote.

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Abstract

Under inflation targeting inflation exhibits *negative* serial correlation in the United Kingdom, and little or no persistence in Canada, Sweden and New Zealand, and estimates of the indexation parameter in hybrid New Keynesian Phillips curves are either equal to *zero*, or very low, in *all* countries. Analogous results hold for the Euro area—and for France, Germany, and Italy—under European Monetary Union; for Switzerland under the new monetary regime; and for the United States, the United Kingdom 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), and suggest that building inflation persistence into macroeconomic models as a structural feature is potentially misleading.

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Non Technical Summary

Over the last twelve years inflation persistence has been 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 building inflation persistence into macroeconomic models. From Fuhrer and Moore (1995) to Blanchard and Gali (2007), several authors have proposed different mechanisms to build inflation persistence into the the deep structure of the economy, thus making it invariant to changes in the monetary regime.

Drawing on the experience of the European Monetary Union, of inflation-targeting regimes, of the new Swiss 'monetary policy concept', and of the Gold Standard, the present work documents how estimates of the indexation parameter in hybrid New Keynesian Phillips curves are either equal to *zero*, or very low, in *all* inflation targeting regimes; in the Euro area, Germany, Italy, and France under European Monetary Union; in Switzerland under the new monetary regime; and in the United States, the United Kingdom, and Sweden under the Gold standard.

These results question the notion that the intrinsic component of inflation persistence many researchers have found in the past—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 building inflation persistence into macroeconomic models as a structural feature is potentially misleading. In particular, both the computation of optimal monetary policies, and the evaluation of the pros and cons of alternative monetary regimes, cannot be performed based on models featuring intrinsic persistence. The reason, quite obviously, is that in both cases the researcher needs a model which can be reasonably be regarded as structural in the sense of Lucas (1976), and as this paper's results show, intrinsic persistence models clearly fail under this crucial dimension.

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.

—Michael Woodford[1]

[T]he launch of European Monetary Union and the establishment of a clearly defined nominal anchor [was] the defining event that changed the very nature of the inflationary process in the euro area. This institutional break has eradicated the "intrinsic" component of the inflation formation mechanism, namely the extent to which economic agents—in re-setting prices or negotiating wages—look at the past history of inflation, rather than into the eyes of the central bank.

—Jean-Claude Trichet[2]

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 building inflation persistence into macroeconomic models. From Fuhrer and Moore (1995) to Blanchard and Gali (2007), several authors have proposed different mechanisms to build inflation persistence into the the deep structure of the economy, thus making it invariant to changes in the monetary regime.

Building on the work of Buiter and Jewitt (1981), Fuhrer and Moore (1995) originally 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 (1999) 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. So far, the most popular mechanism for building 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,[3]

[1] See Woodford (2006).
[2] See Trichet (2007).
[3] See in particular Smets and Wouters (2003) and Smets and Wouters (2007).

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.[4]

As stressed by Cogley and Sbordone (2005), however,

> '[...] from a theoretical point of view [indexation mechanisms 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 my view, a simple one:

> *Is 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—structural in the sense of Lucas (1976)?*

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

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. Notable exceptions to this position are represented by the work of Andrew Levin and his co-authors,[5] and especially by that of Tim Cogley and Argia Sbordone,[6] who show that, once controlling for shifts in the low-frequency component of inflation, it is not possible to reject the null that U.S. post-WWII inflation is purely forward-looking.

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

[5] See Erceg and Levin (2003), Levin and Piger (2003) and Coenen, Levin, and Christoffel (2007).

[6] See Cogley and Sbordone (2005) and Sbordone (2007).

But are we really sure that high inflation persistence truly is structural in the sense of Lucas (1976)? What is, in fact, the empirical evidence in favor of such position? As a simple matter of logic, the *only* way to provide evidence in its favor would be to show that its extent remains virtually unchanged across different monetary regimes. The reason, quite obviously, 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.[7]

1.2 This paper: main results, and implications for macroeconomic modelling and policy

Drawing on the experience of the European Monetary Union, of inflation-targeting regimes, and of the Gold Standard, the present work

- documents, within a reduced-form context, the *negative* serial correlation of U.K. inflation, and the (near) white noiseness of Canadian, Swedish, and New Zealand's inflation under inflation targeting; it documents the low-to-non-existent extent of inflation persistence in the Euro area—and in Germany, France, and Italy—under European Monetary Union, and in Switzerland under the new monetary regime;[8] and it reasserts the well-known absence of inflation persistence in the United States, the United Kingdom and Sweden under the Gold standard.

- Working within a structural context, it documents how estimates of the indexation parameter in hybrid New Keynesian Phillips curves are either equal to *zero*, or very low, in *all* inflation targeting regimes; in the Euro area, Germany, Italy, and France under European Monetary Union; in Switzerland under the new monetary regime; and in the United States, the United Kingdom, and Sweden under the Gold standard.

These results question the notion that the intrinsic component of inflation persistence many researchers have found in the past—captured, in hybrid New Keynesian Phillips curves, by a significant extent of backward-looking indexation—is structural

[7] This is conceptually in line with Friedman and Schwartz (1963)'s position that the lead of money growth over inflation, having remained largely unchanged since the Gold Standard era, should be regarded, for all practical purposes, as a structural feature of the economy.

[8] After implementing monetary targeting almost uninterruptedly since 1974, at the end of 1999 the *Swiss National Bank* (*SNB*) announced a new 'monetary policy concept'. While, before January 2000, the *SNB* only had a strong but generic committment to price stability, without ever providing a specific numerical objective, a key element of the new concept is an *explicit* definition of price stability. Under the new concept, the *SNB* "[...] defines price stability in the same manner as the *European Central Bank*, [as] a CPI inflation rate of less than 2 per cent per year." (Jordan, Peytrignet, and Rich (2000)).

in the sense of Lucas (1976). Further, they suggest that, for several important applications, models featuring 'hardwired' inflation persistence may deliver incorrect answers.[9] In particular, while—in the spirit of Leeper and Zha (2003)'s work on 'modest' policy interventions—it might still be possible to use such models to analyse the consequences of 'small' policy changes *within* a specific regime, the analysis of drastic *regime changes*, like the introduction of a price-level targeting regime, would most likely produce unreliable results. The reason, quite obviously, is that for these applications the researcher needs a model which can be reasonably be regarded as structural in the sense of Lucas (1976), and as my results show, models featuring 'hardwired' persistence clearly fail under this crucial dimension. By the same token, it is not clear to which extent these models can be used for the computation of optimal monetary policies. To the extent that the optimal policy conditional on the estimated model (and therefore, conditional on a specific estimated extent of intrinsic persistence) turns out to be significantly different from the policy which had been conducted over the sample period used for estimation, the very implementation of such policy would render it—quite paradoxically—not optimal any longer, for the simple reason that it would change the structure of the economy. So, quite bizarrely, an optimally computed policy could be regarded as reliable if and only if it were sufficiently close to the policy which had already been followed over the sample period, with the consequence that the reliable computation of optimal policies could only be used to validate policies which are already being followed!

The paper is organised as follows. The next section presents, for the post-WWII period, 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 AR coefficients; and structural evidence based on Bayesian estimates of New Keynesian sticky-price DSGE models featuring Phillips curves with indexation. Section 3 presents analogous evidence for the metallic standards era, while Section 4 concludes.

2 The Post-WWII Period

2.1 Reduced-form evidence

In order to motivate the next sub-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[10] based on fixed-coefficients AR(p) mod-

[9] On this, see in particular Sbordone (2007). A symmetric risk, as it was pointed out by a referee, 'is the risk of conducting monetary policy under the assumption that inflation is not structurally persistent, when in fact it is.' The impact of the degree of structural persistence on optimal monetary policy is analysed by Coenen and Smets (2003).

[10] 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.

els estimated *via* the Hansen (1999) 'grid bootstrap' median-unbiased (MUB) estimator of the sum of the autoregressive coefficients. Specifically, for each country, regime/period, and inflation series[11] we estimate via OLS the AR(p) model[12]

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

selecting the lag order based on the Schwartz information criterion (SIC) for a maximum possible number of lags P=6.[13] 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 [14]—and 90%-coverage confidence intervals, *via* the Hansen (1999) 'grid bootstrap' procedure.[15] Finally, for each sub-period we deconvolute the probability density function (henceforth, PDF) of the median-unbiased estimate of ρ *via* the procedure described in Appendix C, and based on the deconvoluted PDFs we compute p-values for the test of no change in ρ between successive sub-periods.

2.1.1 Inflation-targeting countries

Starting from the United Kingdom, whose experience is especially interesting for the present purposes, Table 1 shows MUB estimates of ρ by regime/period, while Table 2 reports the bootstrapped p-values for testing the null of no change in ρ across sub-periods. As Table 1 clearly shows, under inflation targeting inflation is estimated to have been, so far, *negatively* serially correlated,[16] although statistically indistin-

This is the case, for example, of the introduction of inflation targeting in the U.K., Sweden, New Zealand, and Canada, and of the launch of European Monetary Union. 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.

[11]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.

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

[13]Specifically, for each inflation series we select the lag order based on the model estimated over the full sample.

[14]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.

[15]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 [$\hat{\rho}$-4$\hat{S}(\hat{\rho})$; $\hat{\rho}$+4$\hat{S}(\hat{\rho})$], 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.

[16]This is consistent with the notion that the current monetary regime contains, *de facto*, a slight component of mean-reversion in the (log) price level. It is interesting to notice that, compared to the

guishable from white noise, based on all the three price indices we consider.[17] In stark contrast with the current regime, the period between 1972 and 1992—a significant portion of which had been characterised by the complete lack of any nominal anchor—exhibited very high persistence based on each single series, with point estimates of ρ ranging from 0.89 to 0.95, and upper limits of 90% confidence intervals ranging between 0.98 and 1.04. Bretton Woods displayed low serial correlation, with point estimates of ρ around 0.40, and upper limits of the 90% confidence intervals equal to 0.69 and 0.79. Finally, and intriguingly, the turbulent interwar period only displayed some mild evidence of serial correlation. The p-values reported in Table 2 uniformly point towards changes in persistence following both the floating of the pound *vis-a-vis* the U.S. dollar, in June 1972, and the introduction of inflation targeting, in October 1992, while there is no evidence that persistence under Bretton Woods was any different from what it was during the interwar era.

Moving to the other three inflation-targeting countries (see Tables 3 and 4) 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 in both Canada and New Zealand—with very high estimates of ρ, and upper limits of the 90% confidence intervals beyond one—under inflation targeting, in line with the U.K. experience, inflation persistence has been, so far, low-to-non-existent in *all* three countries.[18] In particular, for all countries, and based on either the GDP deflator or the CPI,[19] it is not possible to reject the null that inflation is currently white noise, while the null of a unit root can uniformly be rejected. 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, the interwar era does not exhibit any clear-cut pattern.

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 time passes, and the data accumulate.

[17] From the perspective of New Keynesian models, the finding that inflation is statistically indistinguishable from white noise may appear at first sight puzzling. Given that the driving variable in the Phillips curve is either the marginal cost or the output gap, and given that a change in the monetary regime cannot logically be expected to render any of the two variables white noise, it is not clear, at first sight, how inflation can become completely serially uncorrelated. The key issue here, however, is that, as it is well known, estimates of the elasticity with respect to the marginal cost or the output gap—the parameter κ in equation (3) below—are very low, being usually around 0.05. As a result, the dominant influence on the serial correlation properties of inflation comes from the indexation component, which, as this paper shows, changes indeed systematically across monetary regimes.

[18] It is important to stress that a median-unbiased estimate of ρ equal to 0.5—which, based on the estimates reported in Tables 1 and 3, is the highest value among inflation-targeting regimes—implies a half-life of reduced-form inflation shocks of *just three months*.

[19] With the only exception of New Zealand based on the CPI.

2.1.2 The Euro area

Tables 5 and 6 report median-unbiased estimates of ρ, and bootstrapped p-values for testing the null of no change in ρ across sub-periods, for both the Euro area considered as a whole and its three largest economies. In spite of the significant extent of uncertainty associated with the estimates for the EMU period—which, given the short sample, is simply unavoidable—the picture emerging from the two tables points towards a fall in inflation persistence following the introduction of European Monetary Union. The fall is especially apparent for the Euro area considered as a whole and for France, based on both point estimates and bootstrapped p-values, while evidence for Germany and Italy is mixed. For Italy, in particular, there is a clear decrease in the point estimates based on either price index, but the p-value is significant only based on the GDP deflator. For Germany the fall in $\hat{\rho}$ is apparent based on the CPI, but the p-values are never significant.

2.1.3 The United States

Tables 7 and 8 present what is—to the very best of my knowledge—the most extensive investigation to date on changes in U.S. inflation persistence over time and across monetary regimes since the Colonial era. Focussing, for the time being, on the post-WWII period, several findings emerge from the tables. Results for the full Bretton Woods regime (December 1946-August 1971) point towards a comparatively low extent of persistence, while those for the more proper sub-sample March 1951-August 1971 (following the FED-Treasury Accord) point towards very high persistence, with point estimates of ρ in excess on one for four series out of five. As for the Great Inflation episode, although for four series out of five the upper limit of the 90% confidence interval is beyond one, point estimates, ranging between 0.72 and 0.77, are quite surprisingly comparatively low, thus highlighting the importance of controlling for the dramatic shift in equilibrium inflation which took place in the 1970s.[20] 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, our results point towards very high persistence based on either the GDP(GNP) or the PCE deflator—with point estimates equal to either 0.80 or 0.91, and upper limits of the 90%-coverage confidence intervals equal to either 1.03 or 1.04—and much lower persistence based on the CPI. Based on the seasonally unadjusted CPI series, in particular, U.S. inflation appears, today, indistinguishable from white noise, with a point estimate of 0.08, and a 90% confidence interval stretching from -0.17 to 0.33. Although intriguing, we regard the CPI-based results as anomalous and puzzling. Anomalous because, different from, e.g., the United Kingdom, for which results based on any of the three price indices we consider uniformly point in the same direction, results for the CPI are here in contrast

[20] This is in line with Cecchetti and Debelle (2004)'s conclusion that 'the conventional wisdom that inflation has a high level of persistence is not robust. Once one controls for a break in the mean of inflation, measured persistence is considerably lower.'

with those based on the GDP(GNP) and PCE deflators; and puzzling because the PCE deflator and the CPI share many individual sub-indices,[21] so that we should not expect to find such a marked difference between their serial correlation properties.

How do my results for the post-Volcker stabilisation period relate to the previous literature? Based on a Bayesian time-varying parameters VAR, Cogley and Sargent (2002) first produced evidence suggesting a fall in U.S. inflation persistence around the time of the Volcker stabilisation, subsequently showing[22] such result to be robust to Stock (2002)'s criticism that it was the figment of not controlling for changes in the VAR's covariance matrix. Pivetta and Reis (2006), however, questioned Cogley and Sargent's findings on the grounds of statistical significance, arguing that once one takes into account the uncertainty surrounding median estimates, the null of no change in U.S. inflation persistence over the post-WWII era cannot be rejected. Cogley and Sargent (2006), however, show that once one focusses on the deviation of inflation from a time-varying equilibrium (what they label as the 'inflation gap') the evidence of a fall in persistence is very strong, especially based on multivariate methods.[23] So, since I am here focussing not on Cogley and Sargent's 'inflation gap', bur rather, as in Pivetta and Reis (2006), on inflation itself, my finding of high persistence for the period following the Volcker stabilisation is in principle compatible with the findings reported in both papers.

2.1.4 Japan and Switzerland

Finally, Tables 9 and 10 report results for Switzerland and Japan. For our purposes, the Swiss experience is especially interesting because exactly like Germany, it pursued, over the entire post-WWII period, a consistently 'hard-money, low-inflation' policy, but only in January 2000 it introduced an *explicit* numerical target for inflation. A comparison between the period before 2000 and the current monetary regime may therefore provide evidence on whether an explicit numerical target, added on the top of an already strongly counter-inflationary stance, may make a difference. Empirical evidence—although inevitably tentative, given the short sample period for the current regime—overall suggests that it does, with point estimates clearly suggesting a fall in persistence under the current regime. It is important to stress, however, how uncertainty is currently still very high, so that more data are needed before it is possible to be sufficiently confident of such fall in persistence.

On the other hand, evidence for Japan—a country which, since the collapse of Bretton Woods, has not had any nominal anchor whatsoever—does not point towards any change in persistence over the most recent years. While persistence is estimated

[21] I wish to thank an anonymous referee for bringing this issue to my attention.

[22] See Cogley and Sargent (2005).

[23] In a conceptually related paper, Stock and Watson (2007) model U.S. post-WWII inflation as the sum of a stochastic trend and a serially uncorrelated transitory component, identifying large changes in the variance of the innovation to the permanent component, which is currently' estimated to be at a record low since 1954'.

to have significantly increased following the collapse of Bretton Woods, the period following the end of the Great Inflation still exhibits, in line with the U.S. experience, comparatively high persistence, and the null of a unit root cannot be rejected for either of the two series.

2.2 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 statistical inflation persistence is neither necessary nor sufficient for *structural* inflation persistence, which, at the end of the day, is what truly matters, especially for monetary policy purposes. In this section I therefore proceed to a structural investigation based on estimated sticky-price DSGE models featuring hybrid Phillips curves. The 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 the Introduction, intrinsic persistence is not structural in the sense of Lucas (1976).

2.2.1 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,[24] 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 persistence to be absorbed by the structural component. Our key result of (near) absence of structural persistence under EMU, inflation-targeting, the new Swiss monetary regime and, in Section 3, the Gold Standard will therefore be all the more striking. 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] The robustness of our results to replacing the output gap with a measure of the marginal cost is investigated in section 2.2.3.

where R_t is the nominal rate, and $\epsilon_{R,t}$ is a disturbance to the monetary rule.

2.2.2 Bayesian estimation

We estimate (2)-(4) *via* Bayesian methods. Our preference for the use of a Bayesian, as opposed to a Classical approach is motivated by the fact that estimation attempts based on pure maximum likelihood tended to produce, in general, fragile results. We eschew GMM, on the other hand, because within the present context the quality of the instruments, and therefore the reliability of the estimates, is in principle not independent of monetary policy, and on the contrary is crucially affected by it. Intuitively, under monetary regimes which are very successful at stabilising inflation—like European Monetary Union, inflation targeting regimes, and the new Swiss 'monetary policy concept'—the quality of the instruments for inflation should be *logically* expected to be low, for the simple reason that, in principle, any information such variables may contain on future inflation fluctuations will be used by the monetary authority to move interest rates in order to counter deviations of inflation from equilibrium. As a consequence, these variables will exhibit, *ex post*, little informational content, *precisely* because the monetary authority has already exploited part or all of such information to keep inflation under control. This point—which is conceptually in line with Woodford (1994)—is extensively analysed by Mavroedis in two recent papers. As stressed by Mavroeidis (2004),

> 'the main sources of weak identification in forward-looking models [...] are that: (i) the forcing variables have limited dynamics, and/or (ii) the unpredictable variation in future endogenous variables is large relative to what is predictable based on available instruments.'

As he stresses in his discussion of identification of forward-looking Taylor rules (the problem for New Keynesian Phillips curves is, conceptually, exactly the same),

> 'the more successful the policy, the more inflation forecasts converge to the actual inflation target, and the less they depend on current and past data, which is a necessary condition for a forward-looking Taylor rule to be empirically identified.'

So, given the lack of reliability, in principle, of GMM estimates within the present context, and the previously mentioned fragility or results based on FIML, we regard a full-information Bayesian approach as the only valid alternative left.

The following two sub-sections describe our choices for the priors and the Random-Walk Metropolis algorithm we use to get draws from the posterior.

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 11 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 the Gamma, inverse Gamma, and Beta 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.

Second, different from several papers in the literature, we adopt a perfectly flat (i.e., uniform) prior for the indexation parameter, α. The key reason for doing so is in order to 'let the data speak' as freely as possible on the crucial feature of interest.

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* simulated annealing (for a full description of the methodology, see Appendix D.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 described in Appendix D.2. As the fractions of accepted draws reported in Tables 12 and 15 show, our methodology is quite successful at producing fractions close to the ideal one (in high dimensions) of 0.23.[25]

Let then $\hat{\theta}$ and $\hat{\Sigma}$ be the mode of the maximised log posterior and its estimated Hessian, respectively.[26] 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

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

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

a sample of 500,000 draws, keeping every draw out of 100 in order to decrease the draws' autocorrelation.

2.2.3 Empirical evidence[27]

Table 12 reports, for each country and sample period, the modes and the 5th and 95th percentiles of the posterior distributions for all the model's structural parameters.

Starting from the Euro area considered as a whole, the contrast between the full-sample results and those for the EMU sub-sample could not be starker. While, over the entire sample, indexation is estimated to have been remarkably high—with the mode of the posterior distribution at 0.864—since January 1999 structural persistence has all but disappeared, with a modal estimate equal to just 0.025.[28] Qualitatively the same results hold for the three largest EMU countries, with the (near) complete disappearance of structural persistence under the current regime. Under this respect the case of Germany is especially interesting, because, exactly like Switzerland, it has pursued a consistently counter-inflationary policy over the entire post-WWII period. And indeed its modal estimate of the indexation parameter over the full sample, at 0.427, is clearly lower than those for France and Italy, equal to 0.826 and 0.679, respectively. At the same time, however, a comparison between estimates of indexation for Germany for the full sample and for the current regime clearly point towards a decrease in indexation, with the modal estimate dropping to 0.019. Although inevitably tentative, given the short sample length for the current regime, these results therefore clearly suggest that the launch of EMU *did* indeed make a difference even for Germany.

The same qualitative picture also holds for inflation-targeting countries, with comparatively high estimated indexation for both the United Kingdom and Canada based on the full sample, and a near-complete lack of structural persistence under inflation targeting for all countries, with the single possible exception of Canada, which, with indexation estimated at 0.19, displays some very weak evidence of persistence.

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[27] All of the results reported in Tables 12-15 are based on the GDP deflator, but an alternative set of qualitatively similar results based on the CPI (in the case of the United States, the PCE deflator) are available upon request.

[28] As pointed out by an anonymous referee, an important point to stress is that changes in the monetary policy rule across regimes are incapable—by themselves—to replicate the changes in inflation's statistical persistence we see in the data, so that changes in the indexation parameter are necessary in order replicate such feature of the data. This can be easily illustrated based on the models estimated for the Eurozone. Simulating the model estimated for full sample conditional on ϕ_π=5, a remarkably aggressive monetary rule, and estimating AR(p) models for inflation (selecting the lag order based on AIC) we obtain a median of the distribution of the sum of the AR coefficients equal to 0.662. (This is based on 1,000 simulations of length T=1,000.) By the same token, simulating the model estimated for the EMU period conditional on ϕ_π=0.85 (close to the boundary between determinacy and indeterminacy) we get a median of the distribution equal to 0.331. This clearly shows that the impact of shifts in monetary policy is by no means sufficient to replicate the changes in the serial correlation properties of inflation across regimes.

Finally, results for Switzerland are exactly in line with those for Germany, with indexation estimated at 0.377 over the whole sample, and virtually at zero under the current monetary regime. Although once again inevitably tentative, given the short length of the most recent sample period, results for Switzerland therefore suggest that the introduction of an explicit numerical objective for inflation, on the top of an already strongly counter-inflationary policy, *does* indeed make a difference. Intriguingly, this finding is conceptually in line with the theoretical analysis of Aoki and Kimura (2007). Based on a simple New Keynesian model, they show that uncertainty on the part of the public on the central bank's inflation objective creates a complicated 'hall of mirrors' effect involving higher order expectations. On the one hand the public is compelled to learn about the central bank's objective. On the other hand, however, the central bank, for stabilisation purposes, has to form an estimate of the public's estimate of its own objective. As Aoki and Kimura show, this mechanism, by causing higher order expectations to become relevant, gives rise to high persistence and high volatility of macroeconomic time series even within an environment of white noise structural shocks.

The contrast between the previous results and those for the United States and Japan is simply striking. For both countries results based on the full samples produce, unsuprisingly, strong evidence of high structural persistence, which is especially clear for the United States. Following the end of the Great Inflation, however, the estimated extent of indexation, although lower than that based on the full sample, is still clearly there for both countries, with modal estimates of α equal to 0.619 and 0.457 respectively.

The results reported in Table 12—and in particular, the contrast between those for inflation-targeting countries, EMU, and the Swiss new monetary regime, on the one hand, and those for the United States and Japan—which, following the end of the Great Inflation, are only characterised by a *generic* commitment to price stability, but without any clearly defined anchor—suggest to us several reflections. In particular

- they question the notion that the intrinsic persistence many researchers have previously found in post-WWII U.S. data is *structural in the sense of Lucas (1976).*

- They 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[29]—purely forward-looking New Keynesian Phillips curves fare (near) perfectly well.

- Finally, conceptually in line with the analysis of Aoki and Kimura (2007), they suggest that the introduction of an explicit numerical target on the top of an already strongly counterinflationary monetary policy does make a difference.

[29] 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.

Robustness to alternative priors for the coefficient on inflation in the monetary rule An 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 envisage a situation in which in reality the central bank were significantly more aggressive, so that having a prior less aggressive than the true value might bias downward the estimates of the indexation parameter. In order to check for this possibility we have re-estimated the models for the Euro area and its three largest countries under EMU, for Switzerland under the new monetary regime, and for the four inflation-targeting countries with two alternative modes for the prior distribution for ϕ_π, 3 and 4.5. Results, reported in Table 13, clearly show our previously discussed findings to be robust along this dimension.

Robustness to using marginal cost measures A second important dimension along which the robustness of these results should be checked is the use of the output gap 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,[30] the most extensively used marginal cost proxy, for the post-WWII era.

Table 14 reports the modes of the posterior distributions of the indexation parameters,[31] 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 the output gap in (3) has been replaced by the labor share. 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, for the United States indexation is still estimated to have been remarkably high, in excess of 0.80, 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.

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

[31] 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.

3 The Metallic Standards Era

Turning to the metallic standards era, several issues ought to be discussed before turning to the empirical evidence.[32]

First, to the extent that old prices series are most likely contaminated by a non-negligible random measurement error, this automatically introduces negative serial correlation in their first differences, thus biasing downwards persistence estimates. Although the problem is obviously potentially there, unfortunately it is not clear at all how to even gauge an idea of the likely extent of its impact, and in what follows I will therefore ignore it. In principle, it is however important to keep this *caveat* in mind when assessing the empirical evidence that follows.

A second important issue is to what extent price stickiness under metallic standards was different from the modern era. Given that, *ceteris parisbus*, an increase in price stickiness automatically causes an increase in inflation persistence within sticky-price models, one possible explanation for the lack of persistence under metallic standards is that, under those regimes, prices used to be significantly more flexible than today. Although intuitively sensible, this conjecture is however at odds with available empirical evidence. Kackmeister (2008), in particular, compares price micro-data for the years 1889-1891 with corresponding micro-data for the years 1997-1999 for matching sets of goods. Compared with the more recent period, the former one exhibits a *lower*—rather than higher—frequency of price changes. As Kackmeister points out, '[t]he number of price changes in the 1889-1891 data is one-fifth of that in the 1997-1999 data despite a similar number of first-differenced observations.' Although Kackmeister's evidence is limited to a specific set of goods, taken at face value it suggests that, if price stickiness were the *only* determinant of inflation persistence, persistence under the Gold Standard should have been *higher* than it is today, rather than lower.

A third issue is the nature of metallic standards compared with contemporary regimes, and its likely impact on the extent of indexation we should logically expect to find in the data. As it is well known—see, e.g., Barro (1982), or Barro (1979)—a fundamental difference between metallic and contemporary standards is that, to a first approximation, the former used to render stationary the price level, rather than the inflation rate. As a consequence, we might logically expect to detect, under these regimes, a *negative* extent of indexation, as rational economic agents expect that any positive shock to the price level will be reversed in the future. Because of this, in this section the uniform prior for α is defined over [-1; 1], rather than over [0; 1] as in the previous section.

Finally, as we will discuss in Section 3.2, given the obvious implausibility of a Taylor rule as reasonable description of the workings of a metallic standards regime, in the structural part we will use a short-cut, by simply modelling the evolution of the monetary base, which under such standards maintained, to a first approximation,

[32] I wish to thank the Editor and three anonymous refereees for bringing these issues out.

a one-to-one link with the stock of base metal.

3.1 Evidence on statistical persistence

Starting from the reduced-form results, in line with previous evidence for the United States,[33] Tables 1, 3, and 7 clearly show statistical persistence to have been entirely absent from the pre-1914 world, and to only have appeared after the collapse of the Classical Gold Standard.

In the United States, in particular, persistence appears to have been virtually absent under metallic standards—either *de facto* or *de jure*, and based on either gold or silver—with Hansen MUB estimates of ρ ranging from -0.02 to 0.24, and the lower and upper limits of bootstrapped 90% confidence intervals ranging from -0.27 to 0.36. The Colonial era, too, exhibited no persistence whatsoever, but given the sheer peculiarity of Colonial monetary arrangements[34], 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.[35]

Results for the United Kingdom clearly show that, historically in the U.K. high inflation persistence appears to have been the *exception*, rather than the rule: consistent with the results for the U.S., persistence appears to have been entirely absent under the Gold Standard, either *de facto* or *de jure*. Finally, results for Sweden are in line with those we just mentioned, with a median-unbiased estimate of ρ equal to just 0.28.

3.2 Structural estimation

Given, as we mentioned, the implausibility of a Taylor rule as a reasonable description of a metallic standard, we (2)-(3) 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,

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

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

[33] 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).

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

[35] 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' and acquiring persistence.

with $\epsilon_{\mu,t}$ and $\epsilon_{v,t}$ being a shock to base money growth and a disturbance to the rate of growth of velocity, respectively. Such specification for the monetary rule—postulating that the money stock evolves according to an *exogenous* process—deserves some discussion. As it is well known,[36] 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 exactly 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.[37]

Table 15 reports the results for the United States, the United Kingdom, and Sweden. In line with the previously discussed evidence for the European Monetary Union, inflation-targeting regimes, and the new Swiss monetary regime, the evidence reported in the table points towards very little persistence, with inflation very close to being a purely forward-looking process.

4 Conclusions, and Directions for Future Research

Twelve 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 invariant to changes in the monetary

[36] See e.g. Barro (1979).

[37] An anonymous referee touched upon the issue of how a monetary rule specified in terms of the rate of growth of a monetary aggregate guarantees equilibrium determinacy. Although I do not have a formal proof of this, I suspect that a monetary rule of this kind—and even more, a money level rule—is such to guarantee global equilibrium determinacy. At any rate, in estimation I programmed the Random Walk Metropolis algorithm to automatically reject all draws which did not guarantee equilibrium determinacy (i.e., uniqueness of the solution), so that, for the present purposes, this is not an issue.

regime.[38] Building on the experience of the European Monetary Union, inflation-targeting regimes, the new Swiss monetary regime, and the Gold Standard, in this paper, first, we have documented the (near) white-noiseness of inflation in the U.K., Canada, Sweden and New Zealand under inflation targeting; the near absence of serial correlation under European Monetary Union and the new Swiss monetary regime; and we have reasserted the well-known lack of serial correlation of inflation under the Gold standard. Second, we have shown how estimates of the indexation parameter in hybrid New Keynesian Phillips curves are extremely low or close to zero under EMU, inflation-targeting regimes, the new Swiss regime, and 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 building inflation persistence in macroeconomic models as a structural feature 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.

In terms of directions for future research, the results reported in this paper are, in my view, indicative of a much deeper problem with DSGE models, which is starting to get more and more recognized within the profession. The key rationale behind the developement of such models was that, by capturing supposedly deep features of the economy (preferences, technology, etc. ...) they should be, in principle, invulnerable to the Lucas critique, and could therefore be used to compute optimal policies, assess the desirability of alternative monetary regimes, etc. As Cogley (2007) put it, however, 'Lucas (1976) said that Cowles Commission models are not structural, but he did not say that DSGE models are.' To put it differently, up until very recently the fact that DSGE models' parameters are structural in the sense of Lucas (1976) has uniformly been *assumed*, rather than *tested*. This paper has however conclusively demonstrated such assumption to be mistaken for at least one crucial parameter, while the recent work of Fernández-Villaverde and Rubio-Ramírez (2007) suggests that such problems may affect, in principle, the entire structure of New Keynesian models. Finally, by the very same token, the fact that New Keynesian models may suffer from this problem naturally suggests that the sticky-information general equilibrium models popularised by Mankiw and Reis[39] may suffer, too, from structural instability of the parameters. Given that macroeconomic models whose parameters are not structural in the sense of Lucas (1976) are not much use, this naturally suggests that exploring the extent to which general equilibrium models' parameters are—or are not—structurally stable across monetary regimes should feature high on the research agenda.

[38] See e.g. Sheedy (2007)

[39] See e.g. Mankiw and Reis (2002) and Mankiw and Reis (2007).

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A The Data

A.1 Canada

A monthly seasonally unadjusted series for the consumer price index 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 *IFS*. The overall sample period is 1957:1-2006:4.

A.2 Euro area

Quarterly seasonally adjusted series for real GDP, the consumption and GDP deflators, and the short-term rate are from the European Central Bank's database. The sample period is 1970:1-2006:4.

A.3 France

Quarterly seasonally adjusted series for real GDP and the GDP deflator, and quarterly seasonally unadjusted series for the CPI and the 3-month Treasury bill rate are from the *IMF*'s *IFS*. The sample period is 1970:1-2006:4.

A.4 Germany

Quarterly seasonally adjusted series for real GDP and the GDP deflator, and quarterly seasonally unadjusted series for the CPI and the government bond yield are from the *International Monetary Fund*'s *International Financial Statistics*. The sample period is 1960:1-2006:4. As for the discontinuity introduced by the German reunification, we treated it in the following way. As for the output gap proxy, we HP-filtered the log of real GDP by sub-sample (i.e., before and after the reunification) and we link the two HP-filtered series. As for inflation rates, given that reunification creates a jump in price indices, and therefore 'spikes' in inflation rates, we follow Stock and Watson (2002) and we replace the observation corresponding to the spike with the median of the 6 adjacent values. Finally, as for the government bond yield we do not adjust it in any way.

A.5 Italy

A quarterly series for the government bond yield isfrom the *IMF*'s *IFS* (series' code is 13661...ZF...). A quarterly seasonally adjusted series for the GDP deflator is from the BIS dataset. Finally, a quarterly seasonally adjusted series for real GDP is from the OECD Economic Outloook (series acronym is: OEO.Q.ITA.GDPV). The sample period is 1970:1-2007:2.

A.6 Japan

A quarterly seasonally adjusted series for real GDP, and quarterly seasonally unadjusted series for the CPI and the discount rate (end of period) are from the *IMF*'s *IFS*. The sample period is 1957:1-2006:4.

A.7 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 dataset ends in 2007:1.

A.8 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*. The dataset ends in 2006:4.

A.9 Switzerland

Quarterly seasonally adjusted series for real GDP and the GDP deflator, and quarterly seasonally unadjusted series for the CPI and the 3-month Treasury bill rate are from the *IMF*'s *IFS*. The sample period is 1970:1-2006:4.

A.10 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.11 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).[40] A monthly

[40] 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).

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, 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.

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.

When neeeded, CPI series have been seasonally adjusted *via* the ARIMA X-12 procedure as implemented in *Eviews*.

B Identifying Monetary Regimes

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

B.1 Euro area, France, Germany, and Italy

For the Euro area, France, Germany, and Italy we divide the post-WWII era into the following regimes/periods.

- The Bretton Woods regime, from December 18, 1946 up to August 15, 1971.
- The period betwen the collapse of Bretton Woods and the start of European Monetary Union, on January 1, 1999. For Germany, we consider the period between August 15, 1971 and on October 3, 1990, the date of the reunification.
- European Monetary Union.

B.2 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.[41]

[41] Although the Civil War had already started on April 12, 1861—the date of the attack on Fort

- 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.[42] In what follows we therefore also report results for the sub-period November 1918-April 1933, which, following Meltzer (1986), we label as 'gold exchange standard with a central bank'.[43]
- 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.3 United Kingdom

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

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'.

[42] See Eichengreen (1992).

[43] 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.

- *De facto* gold standard, from 1718[44] 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,[45] 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.

- 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.4 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 and September 10, 1939, respectively. 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.[46]

[44] 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).

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

[46] See Bordo (1993).

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.

B.5 Switzerland

We consider the following regimes/periods.

- Bretton Woods, defined as for the Euro area.
- The period between the collapse of Bretton Woods and the introduction of the new 'monetary policy concept', on January 1, 2000.
- The post-January 2000 regime.

B.6 Japan

As for Japan, we consider the Bretton Woods and the post-Bretton Woods periods, and the period following the end of the Great Inflation episode, which we set to January 1983-present.

C Deconvoluting the Probability Density Function of the Hansen (1999) 'Grid Bootstrap' Median-Unbiased Estimate of ρ

This appendix describes the procedure we use to deconvolute the probability density function of the Hansen (1999) 'grid bootstrap' median-unbiased estimate of ρ.[47] Following Hansen (1999, section III.A), based on the augmented Dickey-Fuller form of (1) we simulate the sampling distribution of the t-statistic t=($\hat{\rho}$-ρ)/$\hat{S}(\hat{\rho})$—where

[47] The procedure described herein is conceptually very similar to the one I used in Benati (2007) to deconvolute the probability density functions of the Stock and Watson (1996) and Stock and Watson (1998) time-varying parameters median-unbiased estimates of λ.

$\hat{\rho}$ is the simple (i.e., not median-unbiased) OLS estimate of ρ, and $S(\hat{\rho})$ is its estimated standard error—over a grid of possible values [$\hat{\rho}$-4$\hat{S}(\hat{\rho})$; $\hat{\rho}$+4$\hat{S}(\hat{\rho})$], with step increments equal to 0.01.[48] For future reference, let's define the grid as L. Based on the bootstrapped distribution of the t-statistic, and following the procedure for the construction of a (1-α)% confidence interval for the median-unbiased estimate of ρ (henceforth, $\hat{\rho}_{MUB}$), detailed in Hansen (1999, section III.A), we then identify the cutoff points of the percentiles of the cumulative density function (CDF) of $\hat{\rho}_{MUB}$ over the domain L. Finally, we fit a logistic function to the CDF of $\hat{\rho}_{MUB}$ *via* non-linear least squares, and we compute the implied estimate of the probability density function of $\hat{\rho}_{MUB}$, scaling its elements so that they sum up to one.

D Two Technical Aspects of the Bayesian Estimation Procedure

This appendix discusses in detail two technical aspects the Bayesian estimation procedure.

D.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* simulated annealing. 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.

D.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 2.2.2—in order to achieve an acceptance rate of the draws close to the ideal one (in high dimensions) of 0.23. 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.

[48] For each of the possible values in the grid, we consider 1,999 replications.

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] 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 2.2.2, storing, for each c_j, the corresponding fraction of accepted draws, f_j. 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 $\hat{a}_0+\hat{a}_1 c+\hat{a}_2 c^2+\hat{a}_3 c^3=0.23$. As the fractions of accepted draws reported in Tables 12 and 15 shows, the procedure works quite remarkably well in the vast majority of cases.

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]	–	–	0.31 [-0.16; 0.99]	0.40 [0.12; 0.69]	0.95 [0.71; 1.04]	-0.07 [-0.71; 0.60]
Consumer price index	–	–	–	–	0.78 [0.62; 0.98]	-0.13 [-0.45; 0.18]
GDP deflator	–	–	–	0.39 [0.03; 0.79]	0.89 [0.68; 1.04]	-0.37 [-0.83; 0.06]

[a] For the interwar period, retail price index from Capie and Webber (1985). After that, from ONS.

Table 2 Inflation persistence in the United Kingdom: bootstrapped *p*-values for testing the null of no change in ρ across sub-periods

	De facto gold standard ⇒ *de jure gold* standard	Interwar period ⇒ Bretton Woods	Bretton Woods ⇒ Bretton Woods to inflation targeting	Bretton Woods to inflation targeting ⇒ inflation targeting
	Annual series:			
Composite price index	0.815	–	–	–
	Quarterly series:			
Retail price index[a]	–	0.588	0.002	4.5E-6
Consumer price index	–	–	–	2.0E-5
GDP deflator	–	–	0.004	2.9E-7

[a] For the interwar period, retail price index from Capie and Webber (1985). After that, from ONS.

Table 3 Inflation persistence in Canada, Sweden, and New Zealand: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

Monetary regime/ historical period	*Canada* CPI	*Canada* GDP deflator	*Sweden* CPI	*Sweden* GDP deflator
Gold standard	–	–	–	0.28 [0.04; 0.55]
Interwar period	0.59 [0.38; 0.80]	–	0.70 [0.42; 1.02]	–
Bretton Woods	0.71 [0.54; 0.88]	0.78 [0.43; 1.05]	0.29 [-0.01; 0.58]	–
1971 to inflation targeting	0.90 [0.72; 1.04]	1.00 [0.79; 1.04]	0.52 [0.12; 1.02]	–
Inflation targeting	-0.33 [-0.76; 0.14]	0.28 [-0.05; 0.63]	0.36 [-0.11; 0.92]	-0.05 [-0.20; 0.13]
	New Zealand CPI	*New Zealand* GDP deflator		
Interwar period	1.00 [0.68; 1.05]	–		
Bretton Woods	0.29 [0.03; 0.55]	–		
1971 to inflation targeting	0.82 [0.68; 1.01]	–		
Inflation targeting	0.41 [0.00 0.86]	-0.04 [-0.60; 0.61]		

Table 4 Inflation persistence in Canada, Sweden, and New Zealand: bootstrapped p-values for testing the null of no change in ρ across sub-periods

	Canada	*Sweden*		*New Zealand*
	CPI[a]	GDP deflator[b]	CPI[a]	CPI[a]
Interwar period ⇒ Bretton Woods	0.329	–	0.012	0.000
Bretton Woods ⇒ 1971 to inflation targeting	0.152	0.135	0.107	0.003
1971 to inflation targeting ⇒ inflation targeting	5.4E-7	0.000	0.238	–

[a] Seasonally unadjusted. [b] Seasonally adjusted.

Table 5 Inflation persistence in the Euro area, Germany, France, and Italy: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

Monetary regime/ historical period	*Euro area*		*Germany*	
	GDP deflator[a]	Consumption deflator[a]	GDP deflator[a]	CPI[a]
Bretton Woods	–	–	1.01 [0.50; 1.08]	0.74 [0.16; 1.07]
Bretton Woods to EMU[c]	1.01 [0.94; 1.04]	1.01 [0.93; 1.03]	0.81 [0.55; 1.04]	0.95 [0.81; 1.03]
After EMU	0.35 [-0.23; 1.04]	0.10 [-0.36; 0.66]	0.70 [0.12; 1.07]	0.26 [-0.70; 1.09]
	France		*Italy*	
	GDP deflator[a]	CPI[a]	GDP deflator[a]	CPI[b]
Bretton Woods	–	0.45 [0.18; 0.72]	0.83 [0.39; 1.06]	0.75 [0.43; 1.04]
Bretton Woods to EMU	1.00 [0.88; 1.03]	1.01 [0.96; 1.04]	0.94 [0.80; 1.03]	0.95 [0.84; 1.02]
After EMU	0.22 [-0.02; 0.48]	0.14 [-1.37; 1.13]	-0.25 [-0.96; 0.47]	0.52 [-0.03; 1.03]

[a] Seasonally adjusted. [b] Seasonally unadjusted. [c] For Germany, Bretton Woods to German unification.

Table 6 Inflation persistence in the Euro area, Germany, France, and Italy: bootstrapped *p*-values for testing the null of no change in ρ across sub-periods

	Euro area		*Germany*	
	GDP deflator[a]	Consumption deflator[a]	GDP deflator[a]	CPI[a]
Bretton Woods $\Rightarrow$ Bretton Woods to EMU[b]	–	–	0.110	0.191
Bretton Woods to EMU[b] $\Rightarrow$ after EMU	0.000	1.2E-5	–	0.238
	France		*Italy*	
	GDP deflator[a]	CPI[a]	GDP deflator[a]	CPI[b]
Bretton Woods $\Rightarrow$ Bretton Woods to EMU[b]	–	0.001	0.357	0.183
Bretton Woods to EMU[b] $\Rightarrow$after EMU	5.3E-5	3.1E-5	0.001	0.074

[a] Seasonally adjusted. [b] Seasonally unadjusted. [c] For Germany, Bretton Woods to German unification.

Table 7 Inflation persistence in the United States: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

Monetary regime/ historical period	Warren-Pearson wholesale price index	Index of general price level	Consumer price index	
	Seasonally unadjusted series:			
Colonial era	-0.05 [-0.20; 0.13]	–	–	
De facto silver standard	0.24 [0.12; 0.36]	–	–	
De facto gold standard	0.21 [0.05; 0.36]	–	–	
Greenback period	0.28 [0.08; 0.49]	0.43 [0.24; 0.62]	–	
Classical Gold Standard	0.11 [-0.03; 0.25]	0.05 [-0.09; 0.19]	–	
Interwar period	–	0.50 [0.33; 0.67]	0.69 [0.52; 0.87]	
'Gold exchange standard with a central bank'[a]	0.67 [0.48; 0.86]	–	0.71 [0.51; 0.95]	
Bretton Woods	–	–	0.49 [0.32; 0.65]	
March 1951-August 1971	–	–	0.69 [0.50; 0.92]	
Great Inflation	–	–	0.72 [0.44; 1.02]	
Post-Volcker stabilisation	–	–	0.08 [-0.17; 0.33]	
	Seasonally adjusted series:			
	Consumer price index	GNP deflator	GDP deflator	PCE deflator
Classical Gold Standard		-0.02 [-0.27; 0.25]	–	–
Interwar period	0.45 [0.27; 0.60]	0.61 [0.44; 0.80]	–	–
'Gold exchange standard with a central bank'[a]	0.48 [0.28; 0.65]	0.63 [0.42; 0.87]	–	–
Bretton Woods	0.61 [0.42; 0.81]	0.63 [0.44; 0.83]	0.55 [0.35; 0.76]	0.54 [0.37; 0.71]
March 1951-August 1971	1.01 [0.75; 1.06]	1.02 [0.86; 1.09]	1.02 [0.87; 1.09]	1.01 [0.76; 1.05]
Great Inflation	0.77 [0.50; 1.03]	0.74 [0.50; 1.02]	0.75 [0.50; 1.02]	0.74 [0.53; 0.98]
Post-Volcker stabilisation	0.49 [0.19; 0.79]	0.90 [0.70; 1.04]	0.90 [0.69; 1.04]	0.81 [0.62; 1.03]

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

Table 8 Inflation persistence in the United States: bootstrapped p-values for testing the null of no change in ρ across sub-periods

	Seasonally unadjusted series:			
	Warren-Pearson wholesale price index	Index of general price level	Consumer price index	
Colonial era ⇒ *de facto* silver standard	0.547	–	–	
De facto silver standard ⇒ *de facto* gold standard	0.815	–	–	
De facto gold standard ⇒ Greenback period	0.538	–	–	
Greenback period ⇒ Classical Gold Standard	0.191	0.017	–	
Classical Gold Standard ⇒ Interwar period	0.002	0.007	–	
Interwar period ⇒ Bretton Woods	–	–	0.170	
Bretton Woods ⇒ Great Inflation	–	–	0.107	
March 1951-August 1971 ⇒ Great Inflation	–	–	0.876	
Great Inflation ⇒ Post-Volcker stabilisation	–	–	0.001	
	Seasonally adjusted series:			
	Consumer price index	GNP deflator	GDP deflator	PCE deflator
Classical Gold Standard ⇒ Interwar period	–	0.001	–	–
Interwar period ⇒ Bretton Woods	0.213	0.876	–	–
Bretton Woods ⇒ Great Inflation	0.213	0.404	0.170	0.152
March 1951-August 1971 ⇒ Great Inflation	0.107	0.044	0.044	0.066
Great Inflation ⇒ Post-Volcker stabilisation	0.046	0.238	0.238	0.588

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

Table 9 Inflation persistence in Switzerland and Japan: Hansen (1999) 'grid-bootstrap' median-unbiased estimates of ρ, and 90% confidence intervals

Monetary regime/ historical period	*Switzerland*		Monetary regime/ historical period	*Japan*	
	CPI	GDP deflator		GDP deflator	CPI[a]
Bretton Woods	0.76 [0.54; 1.02]	–	Bretton Woods	-0.13 [-0.60; 0.40]	0.38 [0.01; 0.85]
Bretton Woods to 1999	0.89 [0.76; 1.02]	0.67 [0.47; 0.91]	Post Bretton Woods	0.91 [0.82; 1.01]	0.93 [0.80; 1.03]
New monetary regime	0.13 [-0.55; 1.02]	-0.21 [-0.99; 0.72]	*Post-1982*	0.88 [0.70; 1.04]	0.77 [0.48; 1.05]

[a] Seasonally adjusted.

Table 10 Inflation persistence in Switzerland and Japan: bootstrapped *p*-values for testing the null of no change in ρ across sub-periods

Monetary regime/ historical period	*Switzerland*		Monetary regime/ historical period	*Japan*	
	CPI	GDP deflator		GDP deflator	CPI
Bretton Woods $\Rightarrow$ Bretton Woods to 1999	0.329	–	Bretton Woods $\Rightarrow$ post Bretton Woods	1.9E-06	0.002
Bretton Woods to 1999 $\Rightarrow$ New monetary regime	0.000	0.012	Bretton Woods $\Rightarrow$ post 1982	5.1E-06	0.019

[a] Seasonally adjusted.

Table 11 Prior distributions for the structural parameters

Parameter	Domain Gold Standard	Domain Post-WWII	Density	Mode Gold Standard	Mode Post-WWII	Standard deviation Gold Standard	Standard deviation Post-WWII
σ_R^2	$\mathbb{R}^+$		Inverse Gamma	–	1	–	2
σ_π^2	$\mathbb{R}^+$		Inverse Gamma	10	1	10	2
σ_y^2	$\mathbb{R}^+$		Inverse Gamma	10	1	10	2
σ_μ^2, σ_v^2	$\mathbb{R}^+$		Inverse Gamma	10	–	10	–
κ	$\mathbb{R}^+$		Gamma	0.05		0.01	
σ	$\mathbb{R}^+$		Gamma	2		2	
α	[-1, 1]	[0, 1]	Uniform	–		0.2887	
γ	[0, 1]		Uniform	–		0.2887	
ρ, ρ_x	[0, 1]		Beta	0.5		0.25	
ϕ_π	$\mathbb{R}^+$		Gamma	1.5		0.25	
ϕ_y	$\mathbb{R}^+$		Gamma	0.5		0.25	
ψ	$\mathbb{R}^+$		Gamma	1		0.025	
δ	$\mathbb{R}^+$		Gamma	0.1		0.025	

$x = R, \pi, y, \mu, v$

Table 12 Post-WWII era, Bayesian estimates of the New Keynesian model's structural parameters, posterior mode and 90%-coverage percentiles

	Euro area	Germany	France	Italy
	Full sample			
σ_R	0.621 [0.556; 0.695]	0.932 [0.853; 1.087]	0.881 [0.783; 1.019]	0.653 [2.792; 0.087]
σ_π	1.064 [0.976; 1.224]	2.702 [2.442; 3.092]	1.336 [1.184; 1.543]	0.596 [2.464; 0.056]
σ_y	0.096 [0.072; 0.142]	0.298 [0.229; 0.435]	0.106 [0.080; 0.180]	0.764 [3.229; 0.163]
κ	0.044 [0.031; 0.061]	0.056 [0.039; 0.073]	0.042 [0.029; 0.066]	0.043 [0.028; 0.064]
σ	27.979 [21.788; 38.514]	10.203 [8.076; 16.962]	30.550 [23.062; 45.014]	29.585 [24.911; 43.973]
α	0.864 [0.771; 0.933]	0.427 [0.406; 0.542]	0.826 [0.731; 0.940]	0.723 [0.605; 0.833]
γ	0.850 [0.775; 0.988]	1.000 [0.851; 0.997]	0.980 [0.743; 0.992]	0.794 [0.689; 0.912]
ρ	0.852 [0.816; 0.884]	0.853 [0.810; 0.885]	0.815 [0.756; 0.859]	0.923 [0.903; 0.941]
ϕ_π	1.000 [0.971; 1.226]	0.888 [0.771; 1.220]	1.031 [0.956; 1.209]	1.018 [0.917; 1.228]
ϕ_y	1.602 [1.201; 2.522]	1.743 [1.092; 2.227]	1.858 [1.125; 2.958]	1.836 [0.931; 2.587]
ρ_R	0.336 [0.220; 0.479]	0.134 [0.044; 0.286]	0.369 [0.217; 0.518]	0.172 [0.042; 0.309]
ρ_y	0.854 [0.775; 0.898]	0.773 [0.691; 0.882]	0.812 [0.705; 0.883]	0.887 [0.786; 0.940]
Fraction of accepted draws	0.263	0.225	0.203	0.284
	European Monetary Union			
σ_R	0.362 [0.304; 0.519]	0.702 [0.615; 0.744]	0.327 [0.264; 0.466]	0.340 [2.259; 0.269]
σ_π	0.797 [0.643; 1.128]	0.429 [0.360; 0.660]	1.192 [0.892; 1.600]	0.288 [1.887; 0.207]
σ_y	0.143 [0.103; 0.221]	0.272 [0.223; 0.426]	0.252 [0.198; 0.396]	0.489 [2.822; 0.431]
κ	0.050 [0.036; 0.068]	0.053 [0.039; 0.071]	0.051 [0.036; 0.070]	0.041 [0.032; 0.063]
σ	10.379 [6.465; 17.511]	11.742 [8.096; 17.710]	10.031 [5.881; 17.573]	11.710 [6.813; 18.592]
α	0.026 [0.009; 0.279]	0.019 [0.003; 0.107]	0.033 [0.018; 0.387]	0.006 [0.002; 0.076]
γ	0.382 [0.346; 0.453]	0.376 [0.342; 0.427]	0.369 [0.345; 0.469]	0.378 [0.354; 0.448]
ρ	0.856 [0.760; 0.915]	0.793 [0.695; 0.856]	0.901 [0.822; 0.931]	0.944 [0.904; 0.963]
ϕ_π	1.249 [0.982; 1.691]	1.383 [1.157; 1.922]	1.246 [0.968; 1.755]	1.285 [0.993; 1.765]
ϕ_y	0.890 [0.467; 1.461]	0.704 [0.435; 1.267]	0.845 [0.478; 1.613]	0.850 [0.515; 1.558]
ρ_R	0.309 [0.118; 0.683]	0.921 [0.867; 0.950]	0.431 [0.141; 0.695]	0.232 [0.058; 0.478]
ρ_y	0.653 [0.323; 0.768]	0.421 [0.191; 0.616]	0.360 [0.086; 0.661]	0.340 [0.117; 0.614]
Fraction of accepted draws	0.255	0.218	0.231	0.253

Table 12 (continued) Post-WWII era, Bayesian estimates of the New Keynesian model's structural parameters, posterior mode and 90%-coverage percentiles

	United Kingdom	Sweden	Canada	Australia	New Zealand
	Full sample				
σ_R	1.087 [0.995; 1.225]	—	0.905 [0.850; 1.016]	1.051 [0.968; 1.210]	—
σ_π	3.064 [2.737; 3.441]		1.733 [1.584; 1.938]	2.751 [2.493; 3.179]	
σ_y	0.277 [0.215; 0.388]		0.171 [0.138; 0.236]	0.285 [0.224; 0.409]	
κ	0.048 [0.030; 0.062]		0.046 [0.032; 0.060]	0.047 [0.035; 0.067]	
σ	25.893 [20.022; 33.748]		28.451 [21.634; 37.085]	20.306 [16.480; 31.053]	
α	0.606 [0.511; 0.722]		0.665 [0.566; 0.757]	0.548 [0.438; 0.672]	
γ	1.000 [0.918; 0.999]		0.881 [0.805; 0.988]	0.993 [0.891; 0.997]	
ρ	0.890 [0.858; 0.910]		0.876 [0.850; 0.903]	0.889 [0.858; 0.913]	
ϕ_π	0.941 [0.842; 1.082]		1.074 [0.962; 1.412]	0.955 [0.875; 1.265]	
ϕ_y	2.083 [1.440; 2.845]		1.451 [1.003; 2.126]	2.156 [1.359; 2.863]	
ρ_R	0.105 [0.029; 0.246]		0.159 [0.061; 0.274]	0.132 [0.051; 0.284]	
ρ_y	0.733 [0.664; 0.825]		0.815 [0.738; 0.866]	0.743 [0.673; 0.838]	
Fraction of accepted draws	0.245		0.298	0.245	
	Inflation targeting				
σ_R	0.463 [1.693; 0.113]	1.041 [0.874; 1.290]	1.040 [0.903; 1.264]	0.463 [0.385; 0.571]	1.170 [3.029; 0.415]
σ_π	0.400 [1.437; 0.087]	2.750 [2.511; 3.467]	2.107 [1.812; 2.585]	2.198 [1.839; 2.744]	1.032 [2.692; 0.321]
σ_y	0.600 [2.068; 0.191]	0.347 [0.261; 0.592]	0.104 [0.085; 0.171]	0.274 [0.195; 0.443]	1.429 [3.644; 0.657]
κ	0.053 [0.035; 0.069]	0.050 [0.036; 0.069]	0.055 [0.037; 0.071]	0.048 [0.037; 0.070]	0.051 [0.037; 0.072]
σ	11.211 [6.691; 17.733]	9.268 [5.496; 16.216]	15.779 [12.555; 23.609]	10.822 [5.619; 16.720]	12.004 [7.717; 22.708]
α	0.021 [0.006; 0.179]	0.006 [0.003; 0.117]	0.190 [0.044; 0.275]	0.056 [0.011; 0.278]	0.015 [0.005; 0.148]
γ	0.484 [0.463; 0.755]	0.990 [0.846; 0.997]	0.493 [0.466; 0.591]	1.000 [0.524; 0.993]	0.988 [0.505; 0.981]
ρ	0.908 [0.845; 0.936]	0.887 [0.834; 0.925]	0.844 [0.780; 0.889]	0.908 [0.862; 0.938]	0.855 [0.799; 0.900]
ϕ_π	1.120 [0.866; 1.566]	1.146 [0.929; 1.562]	1.237 [0.999; 1.661]	1.043 [0.873; 1.510]	1.092 [0.840; 1.445]
ϕ_y	1.245 [0.783; 2.102]	1.069 [0.554; 1.811]	0.990 [0.566; 1.445]	1.216 [0.758; 2.195]	1.510 [0.865; 2.325]
ρ_R	0.189 [0.059; 0.421]	0.071 [0.021; 0.230]	0.079 [0.025; 0.258]	0.058 [0.018; 0.238]	0.153 [0.033; 0.286]
ρ_y	0.746 [0.512; 0.842]	0.660 [0.535; 0.824]	0.695 [0.529; 0.810]	0.584 [0.248; 0.751]	0.632 [0.178; 0.756]
Fraction of accepted draws	0.222	0.226	0.241	0.234	0.247

Table 12 (continued) Post-WWII era, Bayesian estimates of the New Keynesian model's structural parameters, posterior mode and 90%-coverage percentiles

	United States	Japan	Switzerland
		Full sample	
σ_R	0.927 [0.882; 1.040]	0.509 [2.447; 0.339]	0.338 [0.297; 0.397]
σ_π	0.777 [0.721; 0.854]	0.465 [2.213; 0.233]	2.959 [2.581; 3.771]
σ_y	0.186 [0.149; 0.238]	0.566 [2.768; 0.461]	0.237 [0.149; 0.370]
κ	0.024 [0.02; 0.032]	0.037 [0.026; 0.055]	0.046 [0.027; 0.070]
σ	25.748 [21.587; 37.114]	25.255 [18.589; 34.257]	14.873 [10.517; 25.855]
α	0.908 [0.836; 0.972]	0.666 [0.552; 0.758]	0.377 [0.156; 0.510]
γ	0.690 [0.637; 0.787]	1.000 [0.917; 0.999]	1.000 [0.754; 0.994]
ρ	0.827 [0.774; 0.857]	0.935 [0.918; 0.950]	0.958 [0.941; 0.973]
ϕ_π	1.456 [1.308; 1.667]	0.990 [0.905; 1.191]	0.927 [0.742; 1.183]
ϕ_y	0.708 [0.531; 1.118]	1.417 [0.929; 2.155]	1.698 [1.130; 2.769]
ρ_R	0.182 [0.080; 0.309]	0.219 [0.099; 0.353]	0.180 [0.068; 0.392]
ρ_y	0.794 [0.667; 0.833]	0.779 [0.666; 0.845]	0.797 [0.677; 0.904]
Fraction of accepted draws	0.243	0.266	0.323
	After the Volcker stabilisation	*After the Great Inflation*	*New Monetary Regime*
σ_R	0.667 [0.607; 0.799]	0.501 [1.122; 0.1258]	0.401 [1.453; 0.295]
σ_π	0.626 [0.567; 0.797]	0.451 [1.025; 0.0966]	0.340 [1.219; 0.227]
σ_y	0.172 [0.142; 0.241]	0.594 [1.349; 0.2112]	0.620 [2.073; 0.498]
κ	0.034 [0.025; 0.047]	0.052 [0.037; 0.0669]	0.050 [0.035; 0.068]
σ	7.248 [5.329; 11.384]	9.945 [5.875; 18.0949]	6.893 [3.914; 12.711]
α	0.619 [0.363; 0.962]	0.457 [0.301; 0.6750]	0.013 [0.006; 0.220]
γ	0.774 [0.677; 0.877]	1.000 [0.814; 0.9957]	0.366 [0.340; 0.561]
ρ	0.808 [0.760; 0.861]	0.874 [0.825; 0.9019]	0.879 [0.784; 0.920]
ϕ_π	1.814 [1.436; 2.359]	1.243 [1.064; 1.6039]	1.379 [1.039; 1.807]
ϕ_y	0.688 [0.446; 1.217]	0.794 [0.531; 1.4003]	0.792 [0.418; 1.347]
ρ_R	0.374 [0.251; 0.536]	0.091 [0.032; 0.2855]	0.096 [0.029; 0.420]
ρ_y	0.832 [0.762; 0.891]	0.849 [0.774; 0.9435]	0.333 [0.128; 0.656]
Fraction of accepted draws	0.231	0.230	0.248

Table 13 Robustness of estimates of α to alternative priors for ϕ_π for inflation-targeting countries, and the Euro area, Germany, and France under European Monetary Union: mode and 90%-coverage percentiles

Mode of the prior for ϕ_π:	United Kingdom	Sweden	Canada	New Zealand	Switzerland
3	0.014 [0.006; 0.170]	0.015 [0.004; 0.149]	0.266 [0.166; 0.355]	0.013 [0.005; 0.158]	0.000 [0.007; 0.211]
4.5	0.022 [0.006; 0.179]	0.031 [0.006; 0.183]	0.341 [0.219; 0.459]	0.013 [0.005; 0.144]	0.000 [0.006; 0.253]
	Euro area	Germany	France	Italy	
3	0.012 [0.010; 0.274]	0.015 [0.004; 0.150]	0.010 [0.005; 0.187]	0.010 [0.004; 0.145]	
4.5	0.118 [0.010; 0.338]	0.008 [0.003; 0.095]	0.038 [0.021; 0.408]	0.009 [0.005; 0.166]	

Table 14 Robustness to using marginal cost proxies: posterior modes for α and 90%-coverage percentiles

United States	
Full sample	*Post-1982*
0.820 [0.746; 0.867]	0.473 [0.309; 0.607]
United Kingdom	
Full sample	*Inflation targeting*
0.576 [0.479; 0.649]	0.014 [0.006; 0.171]

Table 15 Gold Standard era: Bayesian estimates of the New Keynesian model's structural parameters, mode and 90%-coverage percentiles

	United States	United Kingdom	Sweden
σ_π	2.844 [2.422; 3.545]	3.34 [2.864; 4.145]	2.648 2.[169; 3.263]
σ_y	0.356 [0.275; 0.517]	0.325 [0.273; 0.519]	0.276 [0.228; 0.413]
σ_μ	5.389 [4.718; 6.716]	5.807 [5.085; 6.957]	9.861 [8.965; 11.434]
σ_V	4.225 [3.729; 5.135]	2.817 [2.425; 3.587]	9.162 [8.465; 10.746]
κ	0.044 [0.034; 0.063]	0.041 [0.031; 0.060]	0.047 [0.033; 0.063]
σ	2.014 [1.617; 2.624]	2.140 [1.671; 2.731]	1.881 [1.569; 2.473]
α	0.137 [0.090; 0.217]	0.040 [0.023; 0.096]	0.190 [0.128 0.233]
γ	0.976 [0.873; 0.996]	1.000 [0.926; 0.998]	0.846 [0.808; 0.916]
ρ_μ	0.246 [0.153; 0.476]	0.043 [0.009; 0.134]	0.484 [0.304; 0.631]
ψ	0.200 [0.159; 0.253]	0.2122 [0.162; 0.261]	0.227 [0.185; 0.288]
δ	1.001 [0.965; 1.046]	1.009 [0.968; 1.050]	1.015 [0.969; 1.052]
ρ_y	0.585 [0.408; 0.904]	0.726 [0.518; 0.919]	0.762 [0.496; 0.921]
ρ_{MD}	0.219 [0.134; 0.404]	0.084 [0.051; 0.170]	0.442 [0.305; 0.634]
Fraction of accepted draws	0.218	0.218	0.216

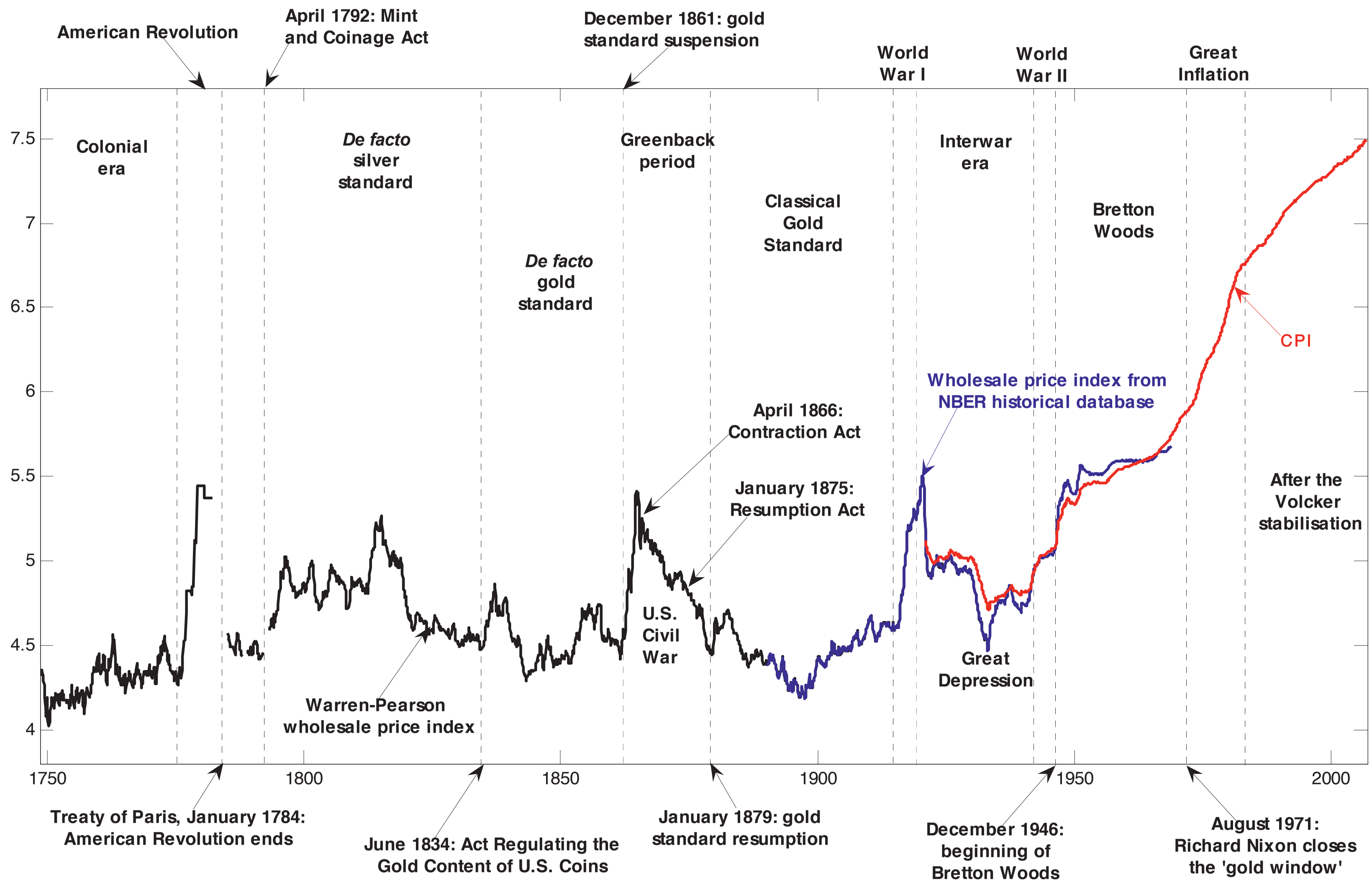

Figure 1 Logarithm of the monthly U.S. price level, Colonial times to the present (July 1748-December 2006)

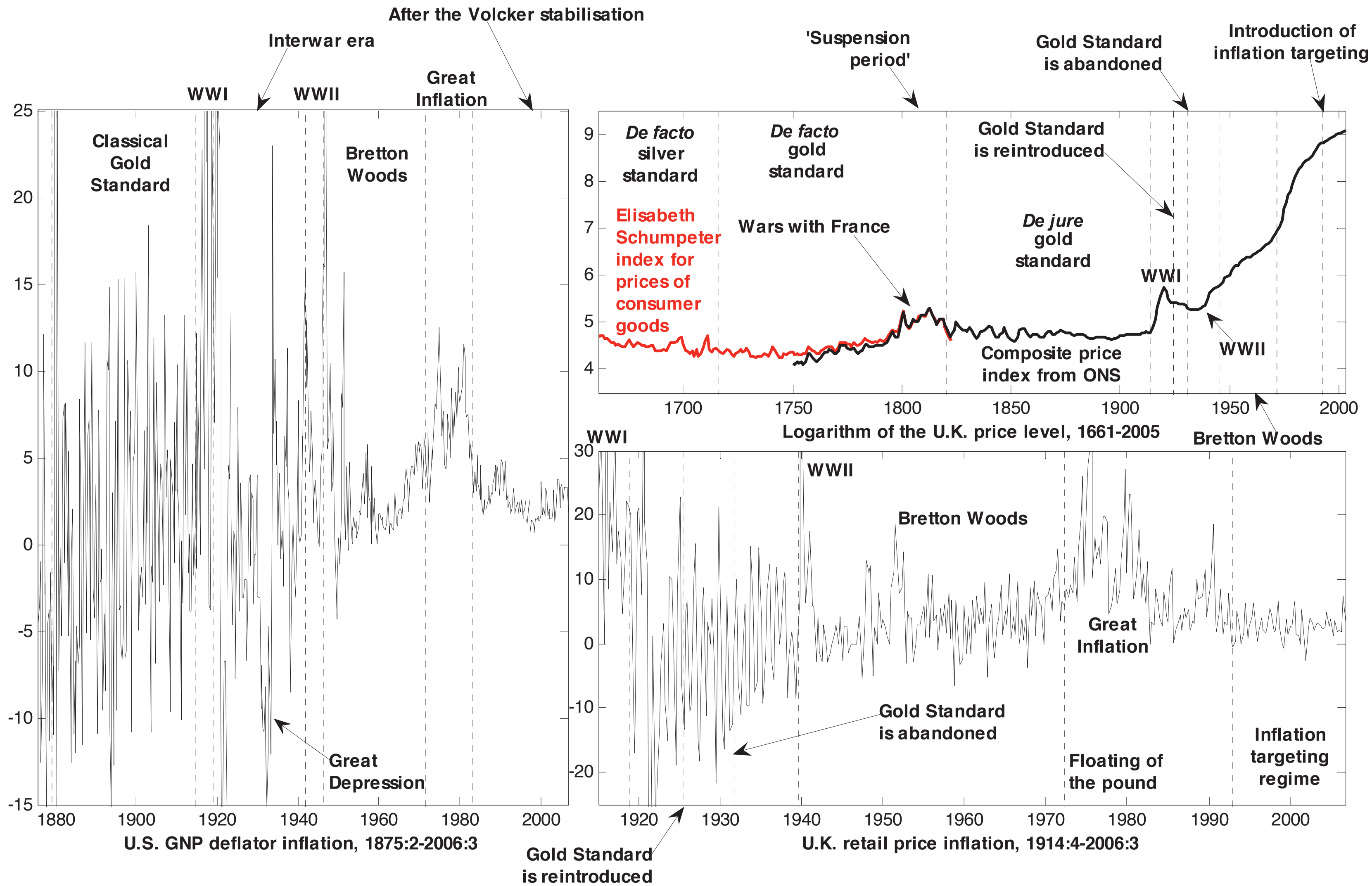

Figure 2 U.S. GNP deflator inflation (1875:2-2006:4), logarithm of the U.K. price level (1661-2005), and U.K. RPI inflation (1914:4-2006:4)

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