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LONG RUN EVIDENCE ON MONEY GROWTH AND INFLATION [1]

by Luca Benati [2]

In 2009 all ECB publications feature a motif taken from the €200 banknote.

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Abstract

Over the last two centuries, the cross-spectral coherence between either narrow or broad money growth and inflation at the frequency ω=0 has exhibited little variation–being, most of the time, close to one–in the U.S., the U.K., and several other countries, thus implying that the fraction of inflation's long-run variation explained by long-run money growth has been very high and relatively stable. The cross-spectral gain at ω=0, on the other hand, has exhibited significant changes, being for long periods of time smaller than one. The unitary gain associated with the quantity theory of money appeared in correspondence with the inflationary outbursts associated with World War I and the Great Inflation–but not World War II–whereas following the disinflation of the early 1980s the gain dropped below one for all the countries and all the monetary aggregates I consider, with one single exception. I propose an interpretation for this pattern of variation based on the combination of systematic velocity shocks and infrequent inflationary outbursts. Based on estimated DSGE models, I show that velocity shocks cause, ceteris paribus, comparatively much larger decreases in the gain between money growth and inflation at ω=0 than in the coherence, thus implying that monetary regimes characterised by low and stable inflation exhibit a low gain, but a still comparatively high coherence. Infrequent inflationary outbursts, on the other hand, boost both the gain and coherence towards one, thus temporarily revealing the one-for-one correlation between money growth and inflation associated with the quantity theory of money, which would otherwise remain hidden in the data.

Keywords: Quantity theory of money, inflation, frequency domain, cross-spectral analysis, band-pass filtering, DSGE models, Bayesian estimation, trend inflation.

JEL Classification: E30, E32

Non-technical summary

Over the last two centuries, the fraction of inflation's long-run variation explained by long-run money growth has been very high, and relatively stable, in the United States, the United Kingdom and several other countries. The proportionality between the long-run components of money growth and inflation, on the other hand, has exhibited significant changes, being for long periods of time lower than one-for-one. In particular, the one-for-one correlation between the long-run components of money growth and inflation associated with the quantity theory of money appeared in correspondence with the inflationary outbursts associated with World War I and the Great Inflation, whereas following the disinflation of the early 1980s the correlation dropped below one for all the countries and all the monetary aggregates considered herein, with one single exception.

The paper proposes an interpretation for the identified stylised facts based on the combination of systematic velocity shocks and infrequent inflationary outbursts. Based on estimated structural macroeconomic models, it is shown that velocity shocks cause, ceteris paribus, comparatively much larger decreases in the correlation between the long-run components of money growth and inflation than in the fraction of inflation's variance explained by money growth. This implies that monetary regimes characterised by low and stable inflation exhibit a low correlation between the long-run components of the two series, whereas the fraction of inflation's variance explained by money growth is still comparatively high. Infrequent inflationary outbursts, on the other hand, boost the correlation towards one, thus temporarily revealing the one-for-one association between the long-run components of money growth and inflation, which would otherwise remain hidden in the data.

> The central predictions of the quantity theory are that, in the long run, money growth should be neutral in its effects on the growth rate of production, and should affect the inflation rate on a one-for-one basis.
>
> —R.E. Lucas, Jr.[1]

> [...] think of velocity shocks as the noise that obscures the signal from monetary aggregates. In a regime in which changes in [...] inflation and the money supply are subdued, the signal-to-noise ratio is likely to be low [...]. However, in other economies or in other time periods in which we experience more pronounced changes in money and inflation, the velocity shocks might become small relative to the swings in money growth, thus producing a higher signal-to-noise ratio.
>
> —A. Estrella and F.S. Mishkin[2]

1 Introduction

This paper is an investigation of the relationship between money growth and inflation at the very low frequencies over the last two centuries. Based on data for inflation and the rates of growth of both narrow and broad monetary aggregates for the United States, the United Kingdom, and several other countries since the Gold Standard era, I use frequency-domain techniques to address the following questions.

(1) 'In the very long run—which I identify with the frequency ω=0—do money growth and inflation move one-for-one, as predicted by the quantity theory of money?'
(2) 'What is the fraction of long-run—i.e., frequency-zero—inflation variance which is explained by long-run money growth?'
(3) 'Are the relationships identified in (1) and (2) stable over time? And if the answer is 'No', have they historically exhibited any systematic pattern of time-variation?'

1.1 Main results

I document several facts. In particular,

- over the last two centuries the cross-spectral gain between either narrow or broad money growth and inflation at the frequency ω=0 has exhibited important changes in most countries, and it has been, for long periods of time, significantly smaller than one at conventional levels. Taken at face value, this result implies that, for long periods of time, the long-run component of inflation has moved *less than one-for-one* with the long-run component of money growth.

[1] See Lucas (1995).
[2] See Estrella and Mishkin (1997).

- The cross-spectral coherence between the two series, on the other hand, has exhibited much less variation, and it has been, most of the times, close to one, thus implying that the fraction of inflation's long-run variation explained by long-run money growth has consistently been close to 100 per cent.

- Evidence for the United States and the United Kingdom (the only two countries for which high-frequency data for both money and prices are available since before World War I) suggests that the unitary gain at ω=0 conceptually associated with the quantity theory of money appeared in correspondence with the inflationary upsurges associated with World War I and the Great Inflation—but *not* World War II.[3] Following the disinflation of the early 1980s, the gain at zero dropped below one for *all* the countries and *all* the monetary aggregates I consider, with the single exception of M3 for Sweden.

- Finally, a comparison between the pre-1914 metallic standard era and the post-WWII period suggests that Rolnick and Weber's (1997) key finding—based on the raw data—of a weaker correlation between money growth and inflation under metallic than under *fiat* standards also holds, most of the times, for the low-frequency components of the two series.

1.2 Interpreting the results

Conceptually in line with Estrella and Mishkin (1997), I propose an interpretation for the identified pattern of variation in the gain and the coherence at ω=0 based on the combination of systematic velocity shocks, and infrequent inflationary upsurges due to either policy mistakes or major geo-political upheavals. Based on estimated DSGE models, I show that velocity shocks cause, *ceteris paribus*, comparatively much larger decreases in the gain between money growth and inflation at ω=0 than in the coherence, so that monetary regimes characterised by low and stable inflation exhibit, in general, a low gain, but a still comparatively high coherence. On the other hand, infrequent inflationary upsurges—such as those associated with World War I and the Great Inflation—by 'swamping' the velocity growth noise away, *temporarily* reveal the one-for-one long-run relationship between the two series, which would otherwise remain hidden in the data.

I also show that three alternative mechanisms which—at least in principle—might account for the identified pattern of variation, do not offer satisfactory explanations. Specifically, *first*, I show that changes in monetary policy have a hard time in reproducing the identified pattern of variation. Indeed, on the one hand, under *determinacy* both the gain and the coherence at zero are largely invariant to changes in the systematic component of monetary policy. On the other hand, indeterminacy is typically

[3] As I discuss in Section 3.2.3, the most logical explanation for the different pattern between WWI and WWII is the presence of extensive price controls during the latter conflict, but not during the former.

associated with lower values of both the gain and the coherence, thus implying that (i) this explanation would still fail to account for the different pattern of variation in the two cross-spectral statistics, and (ii) it would crucially hinge on the notion that comparatively more stable monetary regimes—such as the Gold Standard, or the period following the disinflation of the first half of the 1980s—are characterised by indeterminacy, whereas during periods such as the Great Inflation the economy was under determinacy, which most macroeconomists would most likely find unappealing.

Second, in line with Lucas (1988) and Reynard (2006), I explore the possibility that a gain at zero lower than one may result from systematic, *endogenous* shifts in velocity growth due to Fisherian movements in interest rates—and therefore, in the opportunity cost of money—caused by shifts in the low-frequency (and therefore, highly predictable) component of inflation. This mechanism, too, appears as incapable of explaining the pattern of variation seen in the data.

Third, I show that changes in the elasticity of money demand with respect to either output or the interest rate has essentially *no* impact on either the gain or the coherence at ω=0.

The paper is organised as follows. The next section presents, in the spirit of Lucas' (1980) classic analysis of the low-frequency association between money growth and inflation, evidence from band-pass filtering. Section 3 presents evidence from cross-spectral analysis, whereas Section 4 discusses several possible explanations for our findings. Section 5 concludes.

2 Lucas (1980) *Redux*: Evidence from Band-Pass Filtering

In 'Two Illustrations of the Quantity Theory of Money', Robert Lucas used linear filtering techniques[4] to extract low-frequency components from U.S. M1 growth and CPI inflation over the period 1955-1975, uncovering a near one-for-one correlation between the two series at the very low frequencies.[5] He interpreted his evidence as

> '[...] additional confirmation of the quantity theory, as an example of one way in which the quantity-theoretic relationships can be recovered via atheoretical methods from time-series which are subject to a variety of other forces [...].'

Figure 1 shows some of the series in our dataset, which is described in detail in Appendix A. Figure 2 presents evidence in the spirit of Lucas (1980), by plotting, for

[4] Specifically, a 'precursor' of the Hodrick-Prescott filter.

[5] As I will discuss more extensively in Section 3.2.1, however, McCallum (1984), and especially Whiteman (1984), pointed out how Lucas' results, being based on reduced-form methods, were in principle vulnerable to the Lucas (1976) critique, and as such they could not be interpreted as evidence in favor of the quantity theory of money.

the series shown in Figure 1, the components of broad money growth and inflation with a frequency of oscillation beyond 30 years.[6] The approximated band-pass filter we use in order to extract the low-frequency components from the raw data is the one proposed by Christiano and Fitzgerald (2003).[7] Consistent with the evidence reported in Lucas (1980), Figure 2 points towards a very close correlation between the low-frequency components of broad money growth and inflation since the metallic standards era. Evidence appears especially strong for the United States, Norway, the United Kingdom, Australia, and Portugal—for which the 'eyeball metric' suggests the correlation to be very close to one-for-one—less so for the remaining countries, and in particular for Switzerland, Italy, and the Netherlands, for which the low-frequency components of the two series appear to sometimes diverge to a non-negligible extent.

How should we interpret these results? At first sight, the fact that the correlation between money growth and inflation at the very low frequencies appears to have remained stable across such a marked variation in monetary arrangements over the last two centuries—from a *de jure* or *de facto* Gold Standard, to, in most cases, *de jure* or *de facto* inflation targeting—seems to suggest that such correlation is indeed structural in the sense of Lucas (1976), and it is therefore 'hardwired' into the deep structure of the economy.[8] A crucial problem with this kind of evidence, however, is that strictly speaking band-pass filtering (or, more generally, linear filtering) is not a proper econometric method, in the sense that—different from, e.g., cross-spectral analysis—it does not provide numerical estimates, and measures of uncertainty around such estimates, for key objects of interest capturing the relationship between the two series. As a result, even evidence at first sight strong—such as that for the United States or Norway—ought necessarily to be regarded as purely *suggestive* of stable relationship between money growth and inflation at the very low frequencies since the metallic standard era.

In the next section we therefore turn to cross-spectral methods, which will allow us to compute precise numerical estimates of key objects of interest, and—crucially—to

[6] A legitimate question is: 'Why choosing 30 years as the lower bound of the frequency band of interest?' To be fair, there is no compelling reason why 30 years should be preferred to, say, 29 or 37. Christiano and Fitzgerald (2003, Section 5), for example, consider three frequency bands, with the lowest frequencies being associated with fluctuations between 20 and 40 years.

[7] To be precise, for a sample of size T, Christiano and Fitzgerald (2003) provide formulas for filtered observations from t=3 to t=T-2, thus losing 2 observations at the beginning and 2 at the end of the sample. I worked out the formulas for the first two and the last two observations, so that in performinng band-pass filtering I do not lose any observation. (Needless to say, because of end-of-sample problems, these additional observations are filtered quite imprecisely.)

[8] To the very best of my knowledge, this argument was first made by Batini and Nelson (2002) based on the analysis of raw U.K. and U.S. data since the Gold Standard. As discussed in Batini and Nelson (2002), Friedman (1961) originally made the conceptually related argument that *'[f]or the United States for nearly a century [...] cyclical movements in money have apparently had much the same relation in both timing and amplitude to cyclical movements in business under very different monetary arrangements, though of course the movements in money or in business alone have been very different'.*

characterise the extent of econometric uncertainty associated with such estimates.

3 Evidence from Cross-Spectral Methods

3.1 Methodology

3.1.1 Computing cross-spectral objects

Let x_t and y_t be two jointly covariance-stationary series, with x_t being the 'input' series (in the language of transfer function models), and y_t being the 'output' series (in our case, x_t and y_t are money growth and, respectively, inflation); let $F_x(\omega_j)$ and $F_y(\omega_j)$ be the smoothed spectra of the two series at the Fourier frequency ω_j; and let $C_{x,y}(\omega_j)$ and $Q_{x,y}(\omega_j)$ be the smoothed co-spectrum and, respectively, quadrature spectrum between x_t and y_t at the Fourier frequency ω_j. We estimate both the spectral densities of x_t and y_t, the co-spectrum, and the quadrature spectrum, by smoothing the periodograms and, respectively, the cross-periodogram in the frequency domain by means of a Bartlett spectral window. We select the spectral bandwidth automatically *via* the procedure proposed by Beltrao and Bloomfield (1987). For a specific Fourier frequency ω_j, the estimated smoothed gain and coherence[9] are then defined as[10]

$$\Gamma(\omega_j) = \frac{[C_{x,y}(\omega_j)^2 + Q_{x,y}(\omega_j)^2]^{\frac{1}{2}}}{F_x(\omega_j)} \tag{1}$$

$$K(\omega_j) = \left\{ \frac{C_{x,y}(\omega_j)^2 + Q_{x,y}(\omega_j)^2}{F_x(\omega_j) \cdot F_y(\omega_j)} \right\}^{\frac{1}{2}} \tag{2}$$

[9] On the other hand, we disregard the phase angle. There are several reasons for doing so. *First*, the interpretation of the phase angle statistic is intrinsically *conceptually* tricky, as we are dealing with sine and cosine waves. To illustrate the problem in the simplest possible way, does the sine lead the cosine, or vice-versa? Given that cos(ω)=sin(ω+π/2), it is conceptually impossible to establish which of the two series leads the other one. *Second*, given that the phase angle is defined as $P(\omega_j)$=-arctan[$Q_{x,y}(\omega_j)/C_{x,y}(\omega_j)$], and given that the tangent function is periodic, it is technically impossible to compute confidence intervals for the phase statistics *via* spectral bootstrapping, as the arctangent function converts a sufficiently *large* bootstrapped realisation of the ratio $Q_{x,y}(\omega_j)/C_{x,y}(\omega_j)$ into a comparatively *small* realisation of the phase, precisely because after a certain threshold the periodicity kicks in. (An alternative would be to use asymptotic confidence bands, but based on my own experience—based on extensive Monte Carlo—they have, unsurprisingly, a very poor.coverage.)

[10] It is to be noticed that the literature presents alternative, slightly different definitions of the gain and the coherence—on this, see (e.g.) Hamilton (1994), page 275. The gain, for example, is sometimes defined as the numerator of (1), whereas the coherence is defined as the square of (2).

3.1.2 Monte Carlo evidence on the performance of the cross-spectral estimators

Since, ultimately, our results are only as reliable as our estimators, in this section we present some Monte Carlo evidence on the their performance conditional on a widely representative data generation process (henceforth, DGP). For three sample lenghts, T = 200, 500, 1000, we simulate N times (with N = 5,000) the following DGP:

$$\begin{aligned} x_t &= \rho_x x_{t-1} + u_t + \theta_{ux} u_{t-1} + v_t + \theta_{vx} v_{t-1} & (3) \\ y_t &= \rho_y y_{t-1} + u_t + \theta_{uy} u_{t-1} + v_t + \theta_{vy} v_{t-1} & (4) \end{aligned}$$

with u_t and v_t unit-variance white noise.[11] For each of the N simulations, we draw the key parameters[12]—ρ_x, ρ_y, θ_{ux}, θ_{vx}, θ_{uy}, and θ_{vy}—as follows. θ_{ux}, θ_{vx}, θ_{uy}, and θ_{vy} are drawn from uniform distributions defined over [-0.5; 0.5].[13] The AR parameters (ρ_x and ρ_y) on the other hand, are drawn from a uniform distribution defined over [0; 0.9]. For each single stochastic simulation we estimate the gain and the coherence, and we then compute the difference between the *estimated* objects and the *theoretical* ones, which we compute by Fourier-transforming the DGP conditional on the very same random configuration of parameters, ρ_x, ρ_y, θ_{ux}, θ_{vx}, θ_{uy}, and θ_{vy} which we used to perform the simulation.

Figure 3 shows the medians of the Monte Carlo distributions of the differences between estimated and theoretical gains and coherences, together with the 90 per cent lower and upper percentiles. As the figure clearly shows, the performance of the gain and coherence estimators is uniformly excellent, with the medians of the distributions virtually flat at zero for all the three sample lengths, thus pointing towards no systematic bias in either object at any frequency.

3.1.3 Computing confidence bands

We compute confidence bands *via* the non-parametric multivariate spectral bootstrap procedure introduced by Berkowitz and Diebold (1998)—more precisely, *via* the first of the two procedures they propose. As they show via Monte Carlo, such a procedure generates confidence intervals with superior coverage properties compared to those based on the approximated asymptotic formulas found for example in Koopmans (1974), ch. 8. The Berkowitz-Diebold spectral bootstrap—a multivariate generalisation of the Franke and Hardle (1992) univariate bootstrap—can be briefly described as follows. Let Z_t=[x_t, y_t]$'$, and let $S(\omega_j)$, $I(\omega_j)$, and $\hat{S}(\omega_j)$ be the population spectral density matrix; the unsmoothed sample spectral density matrix; and the smoothed

[11] Specifically, for each single simulation we start x_t and y_t at 0, we generate T+50 realisations for each series, and we then discard the first 50 realisations in order to make the results as independent on initial conditions as possible.

[12] In randomly drawing the coefficients ρ_x, ρ_y, θ_{ux}, θ_{vx}, θ_{uy}, and θ_{vy}, we follow Forni, Hallin, Lippi, and Reichlin (2000), Section V, pages 457-458.

[13] The rationale behind this is in order to rule out 'too large' MA roots.

sample spectral density matrix (i.e., the consistent estimator of $S(\omega_j)$), for the random vector Z_t, all corresponding to the Fourier frequency ω_j. As it is well known[14], $I(\omega_j)$ converges in distribution to a N-dimensional complex Wishart distribution with one degree of freedom and scale matrix equal to $S(\omega_j)$, namely

$$I(\omega_j) \xrightarrow{d} W_{N,C}\left(1, S\left(\omega_j\right)\right) \tag{5}$$

where $W_{s,C}\left(h, H\right)$ is a s-dimensional complex Wishart distribution with h degrees of freedom and scale matrix H. Berkowitz and Diebold (1998) propose to draw from

$$I^k(\omega_j) = S\left(\omega_j\right)^{\frac{1}{2}} W^k_{2,C}\left(1, I_N\right) S\left(\omega_j\right)^{\frac{1}{2}} \tag{6}$$

for all the Fourier frequencies ω_j=$2\pi j/T$, j=1,2, ..., $[T/2]$, with T being the sample length, and $[\cdot]$ meaning 'the largest integer of'. Confidence bands are computed by first getting a smoothed estimate of the spectral density matrix, $S\left(\omega_j\right)$. Then, for each ω_j=$2\pi j/T$, j=1,2, ..., $[T/2]$, we generate 10,000 random draws from (6), thus getting bootstrapped, artificial (unsmoothed) periodograms, we smooth them exactly as we previously did with $I(\omega_j)$, and based on the bootstrapped, smoothed spectral density matrices that we thus obtain we compute gains and coherences according to the traditional formulas, thus building up their empirical distributions. Finally, we compute the confidence bands based on the percentiles of the distribution.[15]

3.2 Empirical evidence

3.2.1 Full-sample estimates

Tables 1 and 2 show estimates of the cross-spectral gain and coherence of money growth onto inflation based on long-run data, whereas Tables 3 and 4 show the same objects based on post-WWII data. Specifically, Tables 1 and 3 both report, for the frequency zero and for the frequency band beyond 30 years,[16] the simple estimate of

[14] See for example Brillinger (1981).

[15] A subtle issue here is the following. Since, in general, the medians of the bootstrapped distributions of the gain and the coherence at each frequency ω_j are not numerically identical to the simple estimates of the two objects based on (1) and (2), we rescale the two distributions so that their medians are indeed equal to such estimates. Given that the gain is, by construction, greater than or equal to zero, whereas the coherence is between 0 and 1, we perform such rescaling based on the log and, respectively, the logit transformations. To be clear, this implies that (e.g.) for the gain, for each frequency ω_j we subtract from the log of the bootstrapped distribution of the gain at ω_j its median, we add to it the log of the simple estimate of the gain at ω_j, and we then take the exponential of the resulting distribution, thus obtaining a botstrapped distribution which, by construction, is *exactly* centered around the simple estimate. For the coherence we follow the same procedure, with the only difference that we use the logit, instead of the log, transformation.

[16] We report results for the frequency band associated with fluctuations with periodicities beyond 30 years as a robustness check. Given that the frequency zero pertains to the infinite long run, and given the inevitable uncertainty associated with estimating objects pertaining to this frequency from a finite data set, results for a set of 'low' frequencies provide a useful check of the reliability of the results for the frequency zero.

the gain based on (1) and the average gain over such band, respectively; the 90%-coverage percentiles of the bootstrapped distributions of the gain and average gain; and the p-values for rejecting the null hypothesis that the gain, and the average gain, respectively, be equal to one. Tables 2 and 4, on the other hand, both report, for the frequency zero and for the frequency band beyond 30 years, the simple estimate of the coherence based on (2) and the average coherence over such band, respectively; the 90%-coverage percentiles of the bootstrapped distributions of the coherence and average coherence; and the mass of the bootstrapped distribution which is beyond 0.99. The reason for reporting this object is that the coherence is bounded, by construction, between zero and one, so that, strictly speaking, it is not technically possible to perform a test that the coherence is equal to one. As a result, we have decided to report the fraction of the mass of the bootstrapped distribution which is greater than 0.9 as a simple indicator of how much the distribution is clustered towards one.

Several facts readily emerge from Tables 1-4.

Gain Starting from the results based on long-run data, for 30 series out of a total of 40 the simple estimate of the gain at ω=0 is lower than one, and in 15 cases this is statistically significant at the 10 per cent level. For the frequency band beyond 30 years, 34 series out of 40 have an estimated average gain lower than one, and in 21 cases—that is, for *more than half* of the series—this is significant at the 10 per cent level. By contrast, for only 10 and 6 series the estimated gain at zero and, respectively, the average gain at the frequencies beyond 30 years are greater than one.

For the post-WWII period results are more in line with our *ex ante* expectation of a unitary gain at the very low frequencies, with 14 series out of 25 having a gain at zero lower than one, and with 7 cases being significant at the 10 per cent level, whereas the corresponding figures for the frequency band beyond 30 years are 18 and 6. Finally, 11 series have a gain at zero greater than one, whereas the corresponding figure for the frequency band beyond 30 years is 7.

Coherence Based on long-run data, 26, 22, and 16 series out of 40 have a simple estimate of the coherence at zero greater than 0.8, 0.9, and 0.95, respectively, whereas for 24 series the fraction of the mass of the bootstrapped distribution of the gain at zero which is beyond 0.99 is greater that 0.01. For the frequency band beyond 30 years, 22, 9, and 3 series, respectively, have an estimated average coherence greater than 0.8, 0.9, and 0.95, respectively, whereas for only 3 series the fraction of the mass of the bootstrapped distribution of the average gain which is beyond 0.99 is greater that 0.01. Once again, results for the post-WWII era are more in line with our *ex ante* expectations, with 23, 20, and 18 (16, 14, and 7) series out of 25 having a simple estimate of the coherence at zero (average estimate of the coherence for the frequencies beyond 30 years) greater than 0.8, 0.9, and 0.95, respectively, and 20 (10) series having a fraction of the mass of the bootstrapped distribution of the gain at zero which is beyond 0.99 being greater than 0.01.

Interpreting the evidence How should we interpret these results? In particular, how should we interpret the fact that, for a comparatively large fraction of series, both the cross-spectral gain at zero, and the average gain at the frequencies beyond 30 years, are estimated to be significantly smaller than one? As originally stressed by Whiteman (1984) and McCallum (1984) in their criticism of Lucas (1976), the interpretation of results produced by frequency-domain methods is, in general, *intrinsically* difficult, as these techniques are reduced-form, and therefore in principle vulnerable to the Lucas (1976) critique. As a consequence, in the same way as Lucas' (1980) results could not—as a matter of principle—be taken as a confirmation of the quantity theory of money, the results reported herein cannot, by the same token, be taken as a refutation of quantity-theoretic arguments. This conceptual problem is—at last potentially—especially severe for the results based on long-run data reported in Tables 1-2, as for many of these series the sample period encompasses radically different monetary regimes, from the XIX century's Gold Standard to contemporary regimes, which are, in most cases, a form of either *de jure* or *de facto* inflation targeting. In order to be able to meaningfully interpret the evidence contained in Tables 1-4 we will therefore need to use structural macroeconomic models, which we will do in Section 4.

Before turning to that, however, let's consider additional empirical evidence, starting from a comparison between results for the Gold Standard and those for post-WWII regimes.

3.2.2 Rolnick and Weber (1997) reconsidered: the Gold Standard *versus* the post-WWII era

In a well-known paper, Rolnick and Weber documented how

> '[...] under *fiat* standards, the growth rates of various monetary aggregates are more highly correlated with inflation [..] than under commodity standards'[17]

Rolnick and Weber 's results were based on an analysis of the *raw* data. Given this paper's focus on the low-frequency components of the data, a question that naturally arises is then: *'What if we focus on the low frequencies? Do Rolnick and Weber's results still hold?'*

Figures 5 and 6 report the bootstrapped distributions of the gain and the coherence, respectively, for both the Gold Standard and the post-WWII period, for all of the series in the dataset with more than 30 years of observations for either regime. We report results both for the simple estimates of the gain and the coherence at ω=0, and—as a robustness check—for the average gains and coherences within the frequency band beyond 30 years. Several results clearly emerge from the Figures.

[17] See Rolnick and Weber (1997, p. 1308).

For the post-WWII era, the bootstrapped distribution of the coherence at ω=0 is very tightly clustered towards one for all countries and monetary aggregates—with the single exception of Italy's M2—thus implying that, over this period, the long-run component of money growth has explained virtually 100 per cent of the long-run variance of inflation. For the Gold Standard, with the single exception of Australia's M3, the probability mass is less clustered towards one than for the post-WWII period. However, with the exception of Italy's and Portugal's M2, and M0 for the United States, the distribution's modes are still very close to one, and the fractions of the masses of the distributions which are clustered towards one are still substantial. Results for the frequency band beyond 30 years paint an overall similar picture, with the fraction of the long-run variance of inflation which is explained by long-run money growth being very close to one for the post-WWII era, and lower—although, in many cases, still quite close to one—for the Gold Standard. As we will discuss in Section 4.2., whereas results for the post-WWII period can be explained very easily based on standard DSGE models, results for the Gold Standard are intriguing, given that one of its key features was its tendency to stabilise the *price level*, as opposed to the inflation rate. As we will show, indeed, under regimes making the (logarithm of the) price level $I(0)$, both the gain and the coherence between money growth and inflation should be expected to be essentially zero.

Turning to the gain, for both the frequency ω=0 and the frequency band beyond 30 years evidence points, overall, towards a smaller gain under the Gold Standard than during the post-WWII period. Two things, however, ought to be noticed. First, estimates for the Gold Standard are, in several cases, clearly greater than zero (this is especially clear, e.g., for Sweden and Norway), and only in a few cases the mass of the bootstrapped distribution is clustered towards zero. Second, as for the post-WWII era, in 7 and 8 cases out of 11 the modal estimate of the gain at zero, and of the average gain for the frequency band beyond 30 years, respectively, are lower than one. As it readily appears from the figure, however, only in a few cases the null of a unitary gain can be rejected at conventional levels.

Finally, let's turn to estimates for rolling samples.

3.2.3 Estimates for rolling samples

A methodological issue: *'Why cross-spectral estimates for rolling samples?'* Before examining the empirical evidence, it is worth spending a few words discussing the methodological reasons behind our choice of exploring time-variation in the cross-spectral gain and coherence between money growth and inflation at the very low frequencies based on estimates for rolling samples. The fundamental reason for doing so is that all other alternatives are—in our view—either unfeasible or still unproven.

Complex demodulation[18]—which I used in Benati (2007) to explore time-variation

[18]Invented by John Tukey—see Tukey (1961)—complex demodulation was introduced in eco-

in the unemployment-inflation trade-off at the business-cycle frequencies—is not a feasible option for strictly *conceptual* reasons. Given that the entire notion behind complex demodulation is to demodulate a specific frequency (or set of frequencies) to ω=0, and then to low-pass filter the resulting complex exponential, quite obviously this cannot be done for the frequency ω=0 itself, and can only reliably work for frequencies which are sufficiently far away from zero.

The so-called 'evolutionary' spectral and cross-spectral analysis developed in a series of papers by Maurice Priestley and his co-workers[19] produces, based on my own experience, results very much in line with those obtained by simply performing (cross-)spectral analysis for rolling samples, but suffers from the fundamental drawback that confidence interval can only be computed based on asymptotic theory, as no bootstrapping method has been developed. Given that for traditional (i.e., time-invariant) spectral analysis asymptotic intervals have an extremely poor coverage,[20] we should logically expect this problem to carry over to the 'evolutionary' spectral analysis, which automatically implies that this methodology should be expected to be, overall, inferior to the one adopted herein.

Finally, time-varying cross-spectral objects could, in principle, be recovered from a time-varying parameters VAR along the lines of, e.g., Cogley and Sargent (2005). Under this respect, however, the key problem is that—as discussed by Christiano, Eichenbaum, and Vigfusson (2006), and as we previously mention—given VARs' focus on fitting short-run dynamics, they should not, in general, be expected to be able to precisely estimate the spectral density matrix of the data at the frequency zero, which is the key object of interest here.

Evidence Figures 8-15 show, for rolling windows of length equal to 25 years,[21] estimates of the gain and coherence between money growth and inflation at the frequency zero, together with 90%-coverage bootstrapped confidence intervals (in the first row); and the p-value for rejecting the null hypothesis that the gain is equal to one, together with the mass of the bootstrapped distribution of the coherence which is behind 0.99

nomics by Granger and Hatanaka (1964), but since then it has been all but forgotten by the profession, while it is used, e.g., in the natural sciences and in medicine—see, e.g., Kim and Euler (1997), Nearing and Verrier (1993), Wilhelm, Grossman, and Roth (1997), Hayano, Taylor, Yamada, Mukai, Hori, Asakawa, Yokoyama, Watanabe, Takata, and Fujinami (1993), and Kessler, McPhaden, and Weickmann (1995). Expositions of complex demodulation techniques can be found in Granger and Hatanaka (1964, chapters 10-12), Bloomfield (1976), Priestley (1981), Hasan (1983), and Priestley (1998).

[19] See in particular Priestley (1965) and Priestley and Tong (1973). For an extended exposition, see Priestley (1998).

[20] On this, see e.g. the discussion in Berkowitz and Diebold (1998).

[21] The reliability of frequency-zero estimates based on 25 years of data is, quite obviously, lower than that of estimates based on samples of 50 or 100 years, or even longer. Unfortunately, we face an unavoidable trade-off between being able to identify time-variation in the data (which requires comparatively a short rolling window), and being able to reliably estimate objects pertaining to the very long run (which requires a comparatively long window).

(in the second row). Each cross-spectral object is plotted in correspondence with the middle of the rolling window based on which it has been computed.[22]

United States Starting from the United States, two key features clearly emerge from Figures 7-9.

First, for either of the two M0 series, the 'money stock'[23] aggregate, M1 or M2 the coherence has been uniformly very high, with its simple estimate being almost always greater than 0.9, and very large (although time-varying) fractions of the mass of the bootstrapped distribution being beyond 0.99. For all series such mass exhibits a clear hump-shaped pattern around the time of the Great Inflation—or slightly after that, as in the case of M2—with large fractions of the mass being clustered towards one.

Second, for all series the gain has been significantly smaller than one for large portions of the sample period, as illustrated by the evolution of the p-values for rejecting the null hypothesis that the gain be equal to one. Since the second half of the XIX century, the gain has exhibited a clear hump-shaped pattern around the time of both World War I (based on the Balke-Gordon M0 and the M2 series) and the Great Inflation (based either the St. Louis adjusted monetary base, M1, or M2).

United Kingdom As for the United Kingdom, the coherence is estimated to have been uniformly very high based on the monetary base—with the only exception of the very first years of the XX century, and of the two decades between the beginning of WWII and the early 1960s—with a large mass of the bootstrapped distribution beyond 0.99 around the time of the Great Inflation. Results based on M3 are broadly in line with those based on M0, with two periods (between mid-1940s and mid-1950s, and between mid-1980s and mid-1990s) in which estimates of the coherence are significantly below one, and the mass of the bootstrapped distribution which is beyond 0.99 collapses to essentially zero, and for the rest of the sample estimates very close to one. Results for M4 (in Figure 15) are in line with those for M3.

Based on M0, the gain is not significantly different from one until mid-1930s, if becomes significantly lower than one between mid-1930s and the end of the 1950s, and it increases beyond one after that. Based on M3, the time-profile of the gain is qualitatively the same as that for M2 in the United States, with about two decades, between the end of the 1930s and the beginning of the 1960s, in which the gain was significantly lower than one, an increase around the time of the Great Inflation episode, and a statistically significant decrease below one over the most recent period. Once again, results for M4 closely mirror those for M3, with the simple estimate of the gain falling from about 2 in mid-1970s, to around 0.5 over the most recent years,

[22] I.e., it has been plotted against t-12.5.

[23] This aggregate is defined as 'money stock, commercial banks plus currency held by the public' (see Appendix A).

with a p-value for rejecting the null that the gain be equal to one virtually equal to zero.

Canada Based on either M0, M1, or M2, results for Canada (see Figures 12 13) are in line with those for the United States. First, the coherence is uniformly very high, and close to one, with a hump-shape in the mass of the bootstrapped distribution which is beyond 0.99 around the time of the Great Inflation. Second, especially for M1 and M2 the time-profile for the gain is very close to the one for M2 for the United States, with a long period starting at the beginning of the 1930s during which the gain has been significantly lower than one, and a clear hump-shaped pattern around the time of the Great Inflation episode. Over the most recent years the gain has fallen significantly lower than one based on M2, whereas based on M1 results are less clear-cut.

Norway For Norway (see Figure 14) results for the coherence are less strong than those seen up until now, with a relatively long period of time—between teh beginning of the 1940s and the end of the 1960s—during which it was quite significantly below one. During the most recent period, on the other hand, the coherence has been uniformly very high and close to one. Results for the gain are in line with those seen up until now, with a clear hump-shape around the time of the Great Inflation, and estimates significantly lower than one over the most recent period.

Euro area, Japan, Sweden Finally, for both the Euro area and Japan the coherence has been remarkably high and, based on the simple estimates, very close to one during the entire period, whereas following the Great Inflation episode the gain has decreased below one—this is especially clear for Japan. Sweden, on the other hand, exhibits essentially no time-variation in either the gain (which, since mid-1970s, has almost never been significantly lower than one) or the coherence (which, based on the simple estimate, has consistently been very close to one, with a large mass of the bootstrapped distribution beyond 0.99).

Summing up Two main features emerge from Figures 7-16. *First*, for all series estimates of the coherence have been consistently very high, and very close to one, for most of the sample periods.[24] *Second*, estimates of the gain, on the other hand, have been significantly lower than one for long periods of time, and they have exhibited a hump-shaped pattern around the time of both WWI and the Great Inflation episode.

[24] As suggested by a referee, an important issue is how to reconcile the comparatively high and stable coherence with the fact that recent work—see e.g. Fischer, Lenza, Pill, and Reichlin (2006) for the Euro area—has documented that monetary aggregates do not provide a satisfactory out-of- sample forecasting performance for inflation. One possible explanation is that all the results reported herein are, by definition, in-sample, so that, in principle, they are in no way incompatible with the evidence of weak out-of-sample forecasting performance of monetary aggregates.

Intriguingly, the significent increases in the rates of growth of monetary aggregates around the time of WWII were not accompanied by corresponding increases in inflation, thus resulting in comparatively low estimates of the gain, which, for all countries and monetary aggregates, were significantly lower than one. The most logical explanation for this is the presence of extensive price controls during the latter conflict, which were instead largely absent during the former.

We now turn to an interpretation of the evidence produced in this section.

4 Interpreting the Evidence

As previosuly mentioned, the interpretation of evidence based on frequency-domain methods is, in general, difficult, as these methods are reduced-form, and therefore in principle vulnerable to the Lucas critique. As a consequence, in order to be able to meaningfully interpret this kind of evidence we will need structural macroeconomic (i.e., DSGE) models.

4.1 The impact of monetary policy

The model I use in order to assess the ability of changes in the systematic component of monetary policy to reproduce the pattern of variation in the gain and the coherence at zero seen in the data is the one I estimated in Benati (2008), which is given by

$$y_t = \gamma y_{t+1|t} + (1-\gamma)y_{t-1} - \sigma^{-1}(R_t - \pi_{t+1|t}) + \epsilon_{y,t}, \qquad \epsilon_{y,t} = \rho_y \epsilon_{y,t-1} + \tilde{\epsilon}_{y,t} \tag{7}$$

$$\pi_t = \frac{\beta}{1+\alpha\beta}\pi_{t+1|t} + \frac{\alpha}{1+\alpha\beta}\pi_{t-1} + \kappa y_t + \epsilon_{\pi,t}, \qquad \epsilon_{\pi,t} \sim WN(0, \sigma^2_\pi) \tag{8}$$

where R_t is the nominal rate, π_t and y_t, are inflation and the output gap, 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 (7)-(8) are expressed as log-deviations from a non-stochastic steady-state. In what follows I calibrate the structural parameters based on the Bayesian estimates for the Euro area and the United States for the full sample periods found in Table XII of Benati (2008).[25] In order to investigate the

[25] An important point to stress here is the following. In Benati (2008) I argued that estimates based on the full sample periods are much less reliable than those based on the most recent (and more stable) monetary regimes, because failure to control for shifts in the low-frequency component of inflation over the full samples tend to artificially 'blow up' the estimated extent of indexation, thus giving the illusion that the 'intrinsic' component of inflation persistence is sizeable. Within the present context I calibrate the models I use based on the *full-sample* estimates found in Benati (2008) because (i) the results I obtain based on these estimates are qualitatively the same as those obtained based on estimates for the most recent periods; and (ii) given that my position on inflation's 'intrinsic' persistence may be regarded as contentious by part of the profession, I want the results in this paper to be completely independent of the issue of whether intrinsic persistence is high, low, or non-existent.

relationship between money growth and inflation, we append to the model (7)-(8) the log-linearised money demand equation

$$m_t - p_t = \eta_y y_t + \eta_R R_t - v_t, \qquad \Delta v_t = \rho_v \Delta v_{t-1} + \tilde{\epsilon}_{v,t} \tag{9}$$

where m_t, p_t, and v_t are the log-deviations of the nominal money stock, the price level, and velocity from the steady-state. We calibrate η_y and η_R at the standard values of 1 and -0.1.

Figure 16-19 explore the impact of monetary policy on the gain and the coherence at the frequency ω=0, conditional on Benati's (2008) modal estimates for the Euro area and the United States, and setting ρ_v=0 and σ_v^2=1. I consider four monetary rules: a standard Taylor rule with smoothing,

$$R_t = \rho R_{t-1} + (1-\rho)[\phi_\pi \pi_t + \phi_y y_t] + \epsilon_{R,t}, \qquad \epsilon_{R,t} = \rho_R \epsilon_{R,t-1} + \tilde{\epsilon}_{R,t} \tag{10}$$

where $\epsilon_{R,t}$ is an AR(1) disturbance; a price-level targeting rule,

$$R_t = \rho R_{t-1} + (1-\rho)[\phi_\pi p_t + \phi_y y_t] + \epsilon_{R,t}, \qquad \epsilon_{R,t} = \rho_R \epsilon_{R,t-1} + \tilde{\epsilon}_{R,t} \tag{11}$$

where p_t is the log-deviation of the price level from a non-stochastic steady-state; a money-growth rule along the lines of Andres, López-Salido, and Nelson (2008), in which the interest rate responds to μ_t rather than to π_t,

$$R_t = \rho R_{t-1} + (1-\rho)[\phi_\pi \mu_t + \phi_y y_t] + \epsilon_{R,t}, \qquad \epsilon_{R,t} = \rho_R \epsilon_{R,t-1} + \tilde{\epsilon}_{R,t} \tag{12}$$

where $\mu_t = m_t$ - m_{t-1} is the growth rate of money, and an alternative specification for the money growth rule

$$\mu_t = \rho_\mu \mu_{t-1} + (1-\rho)[\phi_\pi \pi_t + \phi_y y_t] + \epsilon_{R,t}, \qquad \epsilon_{R,t} = \rho_R \epsilon_{R,t-1} + \tilde{\epsilon}_{R,t} \tag{13}$$

In Figure 16 I consider grids of values for the long-run coefficients on inflation and the output gap in the Taylor rule (10). For each combination of values of ϕ_π and ϕ_y we Fourier-transform the VAR(MA) representation of the DSGE model,[26] thus getting the model's theoretical spectral density matrix, and based on this object we compute the theoretical cross-spectral gain and coherence between money growth and inflation at the frequency zero. By the same token, Figure 19 shows the theoretical cross-spectral gain and coherence as a function of the parameters of the price level targeting rule (11), whereas Figures 17-18 show the same objects as functions of the parameters of two alternative specifications for the money growth rule, (13) and (12).

Several facts readility emerge from the figures. In particular,

[26] Under indeterminacy we solve the model via the Lubik and Schorfheide (2003) extension of the Sims (2002) method. In particular, the solution we use is the one Lubik and Schorfheide (2003) label as 'continuity', which is based on the notion if minimising the difference between the model's impulse-responses on impact when crossing the boundary between determinacy and indeterminacy. The model has a pure VAR representation under determinavy, and a VARMA one under indeterminacy.

- a price level targeting rule causes both the gain and the coherence at zero to essentially *vanish*. It can be easily shown[27] that this result is not unique to such rule, but hold for all monetary rules—such as, e.g., a *money* level targeting rule—which cause the price *level*, as opposed to the inflation rate, to be $I(0)$.

- Under a Taylor rule, both the gain and the coherence are very close to one, and virtually unaffected by changes in the systematic component of monetary policy, under *determinacy*.[28] Under indeterminacy, results are crucially dependent on the specific parameterisation, with a very limited impact based on the estimates for the Euro area, and a comparatively greater one based on those for the United States. An important point to stress is that, irrespective of the parameterisation, a less aggressively counterinflationary policy leads to a *decrease* in both the gain and the coherence at zero.

- As for money growth rules, results for rule (12) are qualitatively the same as those for the Taylor rule, whereas those for the alternative specification (Figure 18) exhibit a different pattern, but still point towards little impact of policy on the cross-spectral statistics at zero.

A crucial point to stress is that for two monetary rules out of three—and, in particular, for the Taylor rule, which is usually regarded as a reasonable characterisation of the way monetary policy has been conducted over the most recent period—a low value of the gain at zero *requires the economy to be under indeterminacy*. This automatically implies that, for monetary policy to be able to explain the fact that, following the disinflation of the first half of the 1980s, the gain at zero has systematically decreased for al countries and monetary aggregates (with the single exception of M3 for Sweden) we ought to believe that the economy was under determinacy around the time of the Great Inflation, and it has instead been under indeterminacy following the disinflation of the first half of the 1980s, a notion that the vast majority of macroeconomists would probably find hard to accept.

4.2 Velocity shocks and infrequent inflationary outbursts

If changes in the systematic component of monetary policy cannot explain the pattern of time-variation in the gain at zero we see in the data, what else can explain it? In this section we discuss one mechanism which can rationalise the broad features of the historical changes in the gain at zero documented in Section 3.2.3—in particular, the fact that (*i*) the coherence appears to have been, most of the times, quite close to one, whereas (*ii*) the gain has often been significantly smaller than one, and it has increased around the time of the inflationary upsurges associated with WWI

[27] These results are available upon request.

[28] So my results are very different from those of Sargent and Surico (2008), who identify a significant impact of policy on the Lucas (1980) estimator at the frequency ω=0.

nd the Great Inflation. Conceptually in line with Estrella and Mishkin (1997), the mechanism is based on the combination of velocity shifts and infrequent inflationary upsurges due to either policy mistakes or major geo-political upheavals. Under 'normal' circumstances—that is, under (reasonably) stable monetary regimes—such mechanism can indeed generate a comparatively low gain at zero, whereas it still produces a coherence quite close to one. On the other hand, infrequent inflationary upsurges—such as those associated with World War I and the Great Inflation—by 'swamping' the velocity growth noise away, would *temporarily* reveal the long-run relationship between the two series, which would otherwise remain hidden in the data. In order to do that, we need to work with a New Keynesian model which has been log-linearised around a non-zero steady-state inflation rate.

4.2.1 A New Keynesian model with non-zero trend inflation

The model I use in this section is the one proposed by Ascari and Ropele (2007), which generalises the standard New Keynesian model analysed by Clarida, Gali, and Gertler (2000) and Woodford (2003) to the case of non-zero trend inflation, nesting it as a particular case.

The Phillips curve block of the model is given by

$$\Delta_t = \psi \Delta_{t+1|t} + \eta \phi_{t+1|t} + \kappa \frac{\sigma_N}{1+\sigma_N} s_t + \kappa y_t + \epsilon_{\pi,t} \tag{14}$$

$$\phi_t = \chi \phi_{t+1|t} + \chi(\theta\text{-}1)\Delta_{t+1|t} \tag{15}$$

$$s_t = \xi \Delta_t + \alpha \bar{\pi}^{\theta(1\text{-}\epsilon)} s_{t-1} \tag{16}$$

where $\Delta_t \equiv \pi_t\text{-}\tau\epsilon\pi_{t-1}$; π_t, y_t, and s_t are the log-deviations of inflation, the output gap, and the dispersion of relative prices, respectively, from the non-stochastic steady-state; θ>1 is the elasticity parameter in the aggregator function turning intermediate inputs into the final good; α is the Calvo parameter; $\epsilon \in$[0,1] is the degree of indexation; $\tau \in$[0,1] parameterises the extent to which indexation is to past inflation as opposed to trend inflation (with τ=1 indexation is to past inflation, whereas with τ=0 indexation is to trend inflation); Δ_t and ϕ_t are auxiliary variables; σ_N is the inverse of the elasticity of intertemporal substitution of labor, which, following Ascari and Ropele (2007), I calibrate to 1; and $\psi \equiv \beta\bar{\pi}^{1-\epsilon}+\eta(\theta\text{-}1)$, $\chi \equiv \alpha\beta\bar{\pi}^{(\theta\text{-}1)(1\text{-}\epsilon)}$, $\xi \equiv (\bar{\pi}^{1-\epsilon}\text{-}1)\theta\alpha\bar{\pi}^{(\theta\text{-}1)(1\text{-}\epsilon)}[1\text{-}\alpha\bar{\pi}^{(\theta\text{-}1)(1\text{-}\epsilon)}]^{-1}$, $\eta \equiv \beta(\bar{\pi}^{1-\epsilon}\text{-}1)[1\text{-}\alpha\bar{\pi}^{(\theta\text{-}1)(1\text{-}\epsilon)}]$, and $\kappa \equiv (1+\sigma_N)[\alpha\bar{\pi}^{(\theta\text{-}1)(1\text{-}\epsilon)}]^{-1}[1\text{-}\alpha\beta\bar{\pi}^{\theta(1\text{-}\epsilon)}][1\text{-}\alpha\bar{\pi}^{(\theta\text{-}1)(1\text{-}\epsilon)}]$, where $\bar{\pi}$ is gross trend inflation measured on a quarter-on-quarter basis.[29] In what follows we uniquely consider the case of indexation to *past* inflation, and we therefore set τ=1.

[29] To be clear, this implies that (e.g.) a steady-state inflation rate of 4 per cent per year maps into a value of $\bar{\pi}$ equal to $1.04^{1/4}$=1.00985.

Following Andres, López-Salido, and Nelson (2008), we add a forward- and backward-looking specification for the demand for log real balances, which expressed in first-difference form, becomes

$$\mu_t = \mu_y(y_t - y_{t-1}) + \mu_R(R_t - R_{t-1}) + \mu_1\mu_{t-1} + \beta\mu_1\mu_{t+1|t} + \beta\mu_1(\mu_t - \mu_{t|t-1}) + \epsilon_{v,t} \quad (17)$$

where $\mu_t \equiv \Delta\tilde{m}_t$—with $\tilde{m}_t \equiv m_t$-p_t being the log-deviation of real balances from the steady-state—and $\epsilon_{v,t} = \rho_v\epsilon_{v,t-1} + \tilde{\epsilon}_{v,t}$ is a stochastic disturbance.

We close the model with the intertemporal IS curve (7) and the monetary policy rule[30] (10).

4.2.2 Estimation issues

We estimate the model *via* Bayesian methods as in Benati (2008). We now briefly discuss two important estimation issues arising from the fact that the model has been log-linearised around a non-zero steady-state inflation rate.

The issue of indeterminacy In a string of papers,[31] Guido Ascari has shown that, when standard New Keynesian models are log-linearised around a non-zero steady-state inflation rate, the size of the determinacy region is, for a given parameterisation, 'shrinking' (i.e., decreasing) in the level of trend inflation.[32] Ascari and Ropele (2007) in particular show that, conditional on their calibration, it is very difficult to obtain a determinate equilibrium for values of trend inflation beyond 4 to 6 per cent. Given that, for both the Euro area and the United States, inflation has been beyond this threshold for a significant portion of the sample period (first and foremost, during the Great Inflation episode), the imposition of determinacy in estimation (*i*) is, *ex ante*, hard to justify, and (*ii*) might end up distorting the estimates of the parameters encoding the intrinsic component of inflation persistence.[33] In what follows we

[30] The imposition of a *single* monetary policy rule over the entire sample period represents an obvious shortcoming of the present version of the paper. For the United States, for example, Clarida, Gali, and Gertler (2000) have convincingly argued for a significant shift in the monetary stance around October 1979, whereas (e.g.) for the United Kingdom it is extremely hard to believe that monetary policy after the introduction of inflation targeting, in October 1992, has not been significantly different from what it had been during the 1960s and 1970s. (The dramatic changes in the intellectual climate surrounding U.K. monetary policy-making have been extensively documented by Edward Nelson in a string of papers—see in particular Nelson and Nikolov (2004) and Batini and Nelson (2005).) In future versions of this work I plan to relax this restriction, by allowing for different monetary policy rules across regimes/periods.

[31] See in particular Ascari (2004) and Ascari and Ropele (2007).

[32] On this, see also Kiley (2007).

[33] The reason is that, as extensively discussed by Lubik and Schorfheide (2004), under indeterminacy the economy exhibits greater volatility and greater persistence across the board, so that part of the high inflation persistence characterising a significant portion of the post-WWII era may simply originate from the fact that, during those years, the economy was operating under indeterminacy. If this is true, but the econometrician imposes, in estimation, determinacy over the entire sample

therefore estimate the model given by (7), (10), and (8)-(16) by allowing for the possibility of one-dimensional indeterminacy,[34] and further imposing the constraint that, when trend inflation is lower than 3 per cent, the economy is within the determinacy region.[35]

Modelling time-variation in trend inflation Within the present context, an important modelling choice is how to specify time-variation in trend inflation. My first choice of modelling it as a random walk—conceptually in line with the work of, e.g., Stock and Watson (2007) and Cogley, Primiceri, and Sargent (2006)—entails, unfortunately, a staggering computational burden, as it implies that trend inflation takes a different value in each single quarter. Since the model's solution crucially depends on the specific value taken by trend inflation—through its impact on the parameters ψ, χ, ξ, η, and κ in (14)-(16)—this means that the model has to be solved *for each single quarter*, which (e.g.) in the case of the United States implies that it takes about 30 seconds to compute the log-likelihood under determinacy (under indeterminacy it takes even more.). Although unwillingly, in what follows I have therefore adopted the shortcut of modelling trend inflation as a step function, allowing it to change every five years, both in the first quarter of each decade, and in the first quarter of the middle year of each decade (so, to be clear, e.g., in 1950Q1, 1955Q1, 1960Q1, etc..).[36] Finally,

period, the immediate consequence will be, quite obviously, to artificially 'blow up' the estimated extent of intrinsic persistence.

[34] This is in line with Justiniano and Primiceri (2008). As they stress (see Section 8.2.1), '[t]his means that we effectively truncate our prior at the boundary of a multi-dimensional indeterminacy region'.

[35] The constraint that, below 3 per cent trend inflation, the economy is under determinacy was imposed in order to rule out a few highly implausible estimates we obtained when no such constraint was imposed. In particular, without imposing any constraint, in a few cases estimates would point towards the economy being under indeterminacy even within the current low-inflation environment, which we find *a priori* hard to believe. These results originate from the fact that, as stressed e.g. by Lubik and Schorfheide (2004), (in)determinacy is a *system* property, crucially depending on the interaction between *all* of the (policy or non-policy) structural parameters, so that parameters' configurations which, within the comparatively simple New Keynesian model used herein, produce the best fit to the data may produce such undesirable 'side effects'.

[36] At first sight, a better alternative might have seemed to run tests for structural breaks at unknown points in the sample in the mean of inflation—based, e.g., on the Bai and Perron (1998) and Bai and Perron (2003) method—and then to impose these breaks in estimation of the New Keynesian model. This, however, would violate the rules of the Bayesian game, as the sample would be used twice, first to get the breaks in the mean of inflation, and then to estimate the model. The solution I devised, on the other hand, does not suffer from this shortcoming because the rule for choosing the break dates in the mean of inflation is independent of the data, and it uniquely depends on calendar time. Further, it produces very reasonable estimates of trend inflation. For the United States, for example, Cogley and Sargent (2002) estimate trend inflation to have reached about 8 per cent in the second half of the 1970s (see their Figure 3.1.), whereas Cogley and Sargent (2005) estimate it between 7 and 8 per cent. By comparison, the methodology adopted herein estimates it slightly above 7 per cent, which provides *prima facie* evidence—admittedly, however, only *prima facie* evidence—that the time-profile of trend inflation produced by the specification adopted herein

I assume that each 'jump' in the step function which represents trend inflation is (i) unanticipated by economic agents, (ii) immediately and perfectly understood when it takes place,[37] and (iii) expected to last forever. Although such assumptions are quite obviously extreme, two things ought to be stressed. First, assumptions (i) and (iii) are compatible with the trend inflation specification upon which the macroeconomic profession has converged upon, i.e. a random-walk. Second, although relaxing (ii) is in principle possible, it would introduce severe complications into the analysis, as (a) it would introduce a distinction between actual trend inflation and the inflation trend which is perceived by economic agents, which would most likely be constantly learning about the time-varying trend; and (b) since such learning would in general imply that the perceived trend changes from one quarter to the next, it would imply the same staggering computational burden of a random-walk specification.

Table 5 reports, for each of the model's structural parameters, the prior density together with two key objects characterising it, the mode and the standard deviation; and the mode and the 90%-coverage percentiles of the posterior distribution generated by the Random-Walk Metropolis algorithm.

4.2.3 The impact of the persistence and innovation variance of velocity growth on the gain and coherence at ω=0

Figure 22 shows, based on the estimated model of Ascari and Ropele for the Euro area,[38] the impact of the persistence and innovation variance of velocity growth on the gain and the coherence at zero between money growth and inflation. I consider grids of values for ρ_v, from 0 to 0.99, and for σ_v^2, from 0 to 10, and for each combination of values in the grid I stochastically simulate the model 200 times for a sample length equal to 1,000, conditional on the value of trend inflation being set to zero. For each stochastic simulation I compute the gain and the coherence between money growth and inflation at zero. The figure shows, for each combination of values for ρ_v and σ_v^2, the medians of the distributions of the gain and the coherence at zero across the 200 replications. As it is apparent, both ρ_v and σ_v^2 have a significantly greater impact on the gain than on the coherence. This implies that, under monetary regimes characterised by low and stable inflation, we should expect to find a comparatively low gain, whereas the coherence should still be expected to be comparatively high. On the other hand, as I now show, infrequent inflationary upsurges, by causing large fluctuations in the low-frequency—i.e., trend—components of both inflation and money growth, boost both the gain and the coherence towards one, thus temporarily 'revealing' the one-for-one correlation between money growth and inflation associated with the quantity theory of money.

is broadly in line with the one that would result from a more appropriate random-walk specification.

[37] So I rule out, by assumption, the need, on the part of economic agents, to learn about shifts in trend inflation, which, on the other hand, might have played a non-trivial role in reality.

[38] I do not report results based on estimates for the United States because they are qualitatively the same as those shown in Figure 22, but these results are available upon request.

4.2.4 Combining velocity shifts and infrequent inflationary outbursts

We now explore the impact of infrequent inflationary upsurges—such as those associated with WWI and the Great Inflation—on the gain and coherence at zero between money growth and inflation. Conditional on the estimates for the Euro area reported in Table 5,[39] and setting, again, ρ_v=0 and σ_v^2=1, I stochastically simulate the Ascari-Ropele model for 130 years at the quarterly frequency (i.e. for 520 periods). Exactly as I did in Section 3.2.3 based on actual data, I then estimate the gain and the coherence between money growth and inflation at zero for rolling sample of 25 years. I model trend inflation by means of the 'bell curve' used for the normal distribution, calibrating the mean and the standard deviation in such a way that (*i*) the peak of the simulated 'Great Inflation episode', equal to about 1.5 per cent, takes place in the 310th period of the simulation, and (*ii*) the length of the inflationary upsurge,[40] from the beginning to the end, is 25 years. The first panel of Figure 23 shows money growth and inflation based on a single stochastic simulation, whereas the second and the third panels show the medians and the 90 per cent lower and upper percentiles of the distributions of the rolling estimates of the gain and the coherence at zero, respectively. The two panels are clearly reminiscent of the pattern we have seen in the data in Section 3.2.3, with the gain at zero being compararatively lower than the coherence under zerom trend inflation, and both statistics being boosted towards one by the 'Great Inflation' episode. The intuition behind this is straightforward. In line with Estrella and Mishkin (1997), if inflation is low and relatively stable—so that trend inflation does not exhibit large fluctuations—velocity shocks weaken the relationship between money growth and inflation, and in particular they weaken the gain much more than the coherence. Large and infrequenct fluctuations in trend inflation, on the other hand, 'swamp' the velocity noise away, thus revealing the one-for-one relationship between money growth and inflation which is 'hardwired' into the deep structure of the model.

4.3 Two further 'non-explanations'

4.3.1 Endogenous shifts in velocity growth

Lucas (1988) and Reynard (2006) suggest that *endogenous* shift in money velocity due to Fisherian movements in interest rates—and therefore in the opportunity cost of money—caused, in turn, by fluctuations in the low-frequency component of inflation, may account for departures from the one-for-one relationship between money growth

[39] Without loss of generality, I rescale all the innovation variances by dividing them by 100. This simplifies the stochastic simulation, because it allows me to consider a much smaller peak—about 1.5 per cent—for the simulated 'Great Inflation episode', which eliminates technical nuisances associated with the possibility that, for large values of the inflation trend, the economy may jump from determinacy to indeterminacy rom one period to the next.

[40] I define the beginning and the end of the upsurge as the quarters in which trend inflation exceeds and falls below, respectively, 0.01 per cent.

and inflation associated with the quantity theory of money. In This section we explore the ability of the Lucas-Reynard hypothesis to account for the pattern of variation in the gain at zero documented in Section 3.2.3. In the standard New Keynesian model of Section 4.1.1 we replace the AR(1) specification for velocity growth of equation (9) with

$$\Delta v_t = \theta R_t + \tilde{\epsilon}_{v,t} \tag{18}$$

where θ>0, so that velocity growth increases with increases in the nominal interest rate. Figure 24 shows the gain and the coherence between money growth and inflation at zero as functions of θ, conditional on Benati's (2008) modal estimates for the Euro area and the United States. As the figure showm, the Lucas-Reynard hypothesis appears as unable to explain the differential pattern between the gain and the coherence at zero. Specifically,

- up to a certain threshold value of θ—depending on the calibration, between 0.55 and 0.7—this mechanism causes the coherence to fall below one, and the gain to *increase above one*.
- Beyond that threshold, the gain and the coherence decrease in tandem, but the gain stays *systematically above the coherence*.

Overall, this mechanism does not appear, therefore, a promising one. In particular, the fact that, for all values of θ, the gain is above the coherence, is very difficult to square with the fact that, historically, for long periods of time the opposite appears to have been true.

4.3.2 Changes in the elasticity of the demand for real balances with respect to output or the interest rate

A final possibility we explored is that the pattern seen in the data may have been the result of changes in the elasticity of the demand for real balances with respect to the interest rate and/or real output. Evidence clearly rejects this possibility too, with essentially *no* impact of changes in either elasticity on the gain and coherence at zero.

5 Conclusions

Over the last two centuries, the cross-spectral coherence between either narrow or broad money growth and inflation at the frequency ω=0 has exhibited little variation—being, most of the time, close to one—in the U.S., the U.K., and several other countries, thus implying that the fraction of inflation's long-run variation explained by long-run money growth has been very high and relatively stable. The cross-spectral gain at ω=0, on the other hand, has exhibited significant changes, being for long periods of time smaller than one. The unitary gain associated with the quantity theory of money appeared in correspondence with the inflationary outbursts associated with World War I and the Great Inflation—but not World War II—whereas following the disinflation of the early 1980s the gain dropped below one for all the countries and all the monetary aggregates I have considered, with one single exception.

I have proposed an interpretation for this pattern of variation based on the combination of systematic velocity shocks and infrequent inflationary outbursts. Based on estimated DSGE models, I show that velocity shocks cause, ceteris paribus, comparatively much larger decreases in the gain between money growth and inflation at ω=0 than in the coherence, thus implying that monetary regimes characterised by low and stable inflation exhibit a low gain, but a still comparatively high coherence. Infrequent inflationary outbursts, on the other hand, boost both the gain and coherence towards one, thus temporarily revealing the one-for-one correlation between money growth and inflation associated with the quantity theory of money, which would otherwise remain hidden in the data.

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A The Data

A.1 Euro area

Money Monthly seasonally unadjusted series for Euro area M1, M2, and M3 are from the *European Central Bank*'s (henceforth, *ECB*) database. The sample periods are January 1970-August 2008 for M1 and M3, and January 1980-August 2008 for M2. Quarterly seasonally adjusted series for M2 and M3, available for the period 1970Q1-2008Q2, are from the same source.

Prices A monthly seasonally unadjusted series for the HICP, available for the period January 1980-August 2008, and a quarterly seasonally adjusted series available for the period 1970Q1-2008Q2, are from the *ECB*'s database.

Output A quarterly seasonally adjusted series for real GDP, available for the period 1970Q1-2008Q2, is from the *ECB*'s database.

A.2 United States

Money A monthly seasonally unadjusted series for 'Total currency outside the Treasury', available from June 1878 to December 1914, is from the *NBER Historical Database* (the series' code is 14135). An annual M0 series for the period 1869-1983 is from Balke and Gordon (1986). An annual M2 series for the period 1820-1993 is from the Rolnick and Weber (1997) dataset, which I downloaded from the Minneapolis FED website. The years 1867 to 1877 included, though, are missing, so I only use this series for the period 1820-1866 (as for the period after 1877, I use the quarterly series which is detailed below). A quarterly seasonally adjusted[41] M0 series for the period 1875Q1-1983Q4 is from Balke and Gordon (1986, henceforth, BG). A quarterly seasonally adjusted M2 series for the period 1875Q1-2008Q2 has been constructed by linking the M2 series from BG with M2SL,[42] which I dowloaded from FRED (at the St. Louis FED website) and I converted to the quarterly frequency by taking averages within the quarter. Specifically, the linked series consists of the BG series up to 1958Q4, and of M2SL after that. A monthly seasonally adjusted series for the St. Louis adjusted monetary base (acronym is AMBSL) for the period January 1918-August 2008 is from FRED. A monthly seasonally adjusted series for the 'money stock, commercial banks plus currency held by the public' for the period May 1907-June 1969 is from Friedman and Schwartz (1970) until December 1946,

[41] For both this series and the next two—M1 and M2—Balke and Gordon (1986) do not mention whether they are seasonally adjusted or unadjusted. An analysis of the spectral densities of their log-differences, however, clearly points towards the absence of sizeable seasonal components in either series.

[42] 'M2SL, M2 Money Stock, Board of Governors of the Federal Reserve System, H.6 Money Stock Measures, Seasonally Adjusted, Monthly, Billions of Dollars'.

and from the Federal Reserve Board after that.[43].Monthly seasonally adjusted series for M1, M2, and M3—acronyms are M1SL, M2SL, and M3SL, respectively—are from FRED. The sample periods are January 1959-August 2008 for M1 and M2, and January 1959-February 2006 for M3, whose publication has been discontinued.

Prices 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). An annual CPI series for the period 1790-1991 is from the Rolnick and Weber (1997, henceforth RW) dataset, whereas an annual GNP deflator series for the period 1869-1983 is from BG. A quarterly seasonally adjusted series for GNP deflator the period 1875Q1-2008Q2 has been constructed by linking the GNP deflator series from BG with GNPDEF,[44] which I dowloaded from FRED. The linked series consists of the BG series up to 1946Q4, and of GNPDEF after that. A monthly seasonally adjusted CPI series, available for the period January 1947-August 2008, is from FRED (acronym is CPIAUCSL). A monthly seasonally unadjusted CPI series for the period January 1913-August 2008 has been constructed by linking CPIAUCNS ('Consumer Price Index for All Urban Consumers: All Items') from FRED, which is available starting from January 1921, to the monthly seasonally unadjusted CPI series from the *NBER Historical database* (the series' code is 04128).

Output A quarterly seasonally adjusted series for real GNP, available for the period 1875Q1-2008Q2, has been constructed by linking the real GNP series found in BG to the real GNP series produced by the U.S. Department of Commerce's *Bureau of Economic Analysis* (the series' acronym is GNPC96). Specifically, the overall series consists of the BG series up to 1946Q4, and of the one from the *BEA* after that.

A.3 United Kingdom

Money An annual M0 series for the period 1833-1879 is from Huffman and Lothian (1980). Annual series for M0 and M3, available for the period 1871-2007, are from Capie and Webber (1985) until 1969, and from the *Bank of England* database after that. A quarterly seasonally unadjusted series for M3, available for the period 1922Q1-2008Q2, is from Capie and Webber (1985) until 1968Q4, and from the Bank of England after that. Quarterly seasonally adjusted series for M0 and M4, available for the periods 1969Q2-2006Q1 and 1963Q1-2008Q2, respectively, are from the *Bank of England* database. A monthly seasonally unadjusted M0 series, available for the period January 1870-April 2006, has been constructed by linking the series from Capie and Webber (1985), which is available until December 1969, to an M0 series from the Bank of England. A monthly seasonally unadjusted M1 series for the period

[43]Both series are available from the *NBER Historical Database* on the web. Their NBER series' codes are 14144a and 14144c respectively.

[44]'Gross National Product: Implicit Price Deflator, U.S. Department of Commerce: Bureau of Economic Analysis, Seasonally Adjusted'.

December 1921-December 1969 is from Capie and Webber (1985).[45]

Prices An annual series for the retail price index, available for the period 1750-2007, is from the *Office for National Statistics* (henceforth, *ONS*). Specifically, O'Donoghue, Goulding, and Allen (2004) contains a series for the period 1750-2003, and data for the most recent years have been dowloaded from the *ONS* website. A quarterly seasonally adjusted series for the GDP deflator, available for the period 1955Q1-2006Q3, is from the *ONS*. A monthly seasonally unadjusted series for the wholesale price index, available from January 1885 to May 1951, is from the *NBER Historical Database* (NBER series' code is 04053). A monthly seasonally unadjusted series for the retail price index, available from August 1914 to August 2008 is from Capie and Webber (1985) until June 1947, and from the *ONS* after that. In those cases I needed it at the quarterly frequency, I converted the original monthly figures by sampling the last month of each quarter.

Output An annual series for real GDP for the period 1870-2007 has been constructed by linking the GDP series found in RW (available for the period 1870-1994) and that produced by the *ONS* (available since 1948). The linked series consists of the former until 1947, and of the latter after that. quarterly series for real GDP, available for the period 1955Q1-2008Q2, is from the *ONS* (acronym is ABMI). GDP (from Rolnick-Weber until 1947, from *ONS* after that)

A.4 France

An annual M2 series for the period 1909-1994 is from the RW dataset.

Prices An annual prices series for the period 1820-1994 is from the RW dataset.

A.5 Switzerland

Money All series for Swiss monetary aggregates are from the *Swiss National Bank* (henceforth, *SNB*), and they have been kindly provided by Samuel Reynard. Annual series for M0, M1, and M3 are available for the period 1907-2005. Monthly seasonally unadjusted series for M1, M2, and M3 are available for the periods January 1950-December 2006 (M1), and June 1975-December 20076 (M2 and M3).

Prices An annual CPI series is from *Global Financial Database* from 1880 until 1920, and from the *Swiss National Bank* (henceforth, *SNB*) after that. A monthly seasonally unadjusted CPI series for the period January 1921-December 2006 is from the SNB.

Output An annual real GDP series for the period 1929-2005 is from the *SNB*.

[45] I could not update this series, as, different from M0 and M3, the Bank of England does not publish any M1 series starting in (or before) December 1969.

A.6 Italy

Money An annual M2 series for the period 1862-1998 is from the RW dataset. A monthly seasonally unadjusted M2 series for the period January 1948-September 1997 has been reconstructed by Eugenio Gaiotti, and it has been kindly provided by Alberto Baffigi (both of Banca d'Italia).

Prices An annual CPI series for the period 1861-2007 is from the RW dataset until 1947 and from the IMF's IFS after that. A monthly seasonally unadjusted CPI series produced by the Italian statistical agency, ISTAT, has been kindly provided by Alberto Baffigi of Banca d 'Italia. The sample period is January 1948-February 2005.

A.7 Japan

Money Monthly seasonally adjusted series for M1 and M2, available for the period January 1957-March 2008, are both from the *IMF*'s *IFS* (the two series' codes are 15859MACZF... and 15859MBFZF...).

Prices A monthly seasonally unadjusted series for the CPI ('General index excluding imputed rent'), from Japan's national statistical agency's website, is available from August 1946 to December 2006. The series suffers from a discountinuity in April 1997, and in our analysis—in line with Stock and Watson (2002)—we have therefore replaced the log-difference of the CPI for that month with the median value of the 6 adjacent values.

A.8 Canada

Money An annual M2 series for the period 1871-2006 is from RW until 1967, and from the *IMF*'s *IFS* (the series' code is 15659MB.ZF...) after that. Monthly seasonally unadjusted series for M0, M1, and M2, available for the period January 1871-December 1967, are from Metcalf, Redish, and Shearer (1996), and have been extended based on data from *Statistics Canada*.

Prices An annual CPI series for the period 1870-2007 is from RW until 1947, and from the *IMF*'s *IFS* (the series' code is 15664...ZF...) after that. A monthly seasonally unadjusted series for the CPI, available from January 1914 to May 2008, is from *Statistics Canada*.

Output An annual real GDP series for the period 1870-1994 is from RW.

A.9 Netherlands

Money Annual M0 and M2 series for the period 1864-192 are from the RW dataset. A monthly seasonally unadjusted M2 series for the period January 1957-December 1997 is from the *IMF*'s *IFS* (the series' acronym is 13859MB.ZF...). The series has a break in December 1982, and—as we do for the Japanese CPI (see above), and in line with Stock and Watson (2002)—we replace the log-differences of the series for that month with the median value of the 6 adjacent values.

Prices An annual CPI series for the period 1880-1992 is from *GlobalFinancialData* until 1899, and from the RW dataset after that. A monthly seasonally unadjusted CPI series for the period January 1957-August 2008 is from the *IMF*'s *IFS* (the acronym is 13864...ZF...).

Output An annual real GDP series for the period 1900-1992 is from the RW dataset.

A.10 Australia

All historical annual series for monetary aggregates and the CPI were kindly provided by James Holloway and Cathie Close of the *Reserve Bank of Australia* (henceforth, *RBA*).

Money An annual M2 series for the period 1841-1983 is originally from Pope (1986). The annual series for M1 is available for the period 1900-1973 (the *RBA* code is R7701.14A). The historical M3 series (*RBA* code: R7701.14B) is available for the period 1900-1973, and it has been extended back in time by linking it to the series from Pope (1986), and forward to 2006 based on data from the *RBA*. Quarterly series for M1 and M3, available for the periods 1975Q1-2008Q2 and 1959Q3-2008Q2 respectively, are from the *IMF*'s *IFS* database (the series' codes are 19359MA.ZF... and 19359MC.ZF... respectively).

Prices The CPI historical annual series is available for the period 1850-2006. Original sources are: 1850-1900: Sydney Retail Prices, The Labour Report; 1901-1949: Retail Price Index Year Book Australia 1986 ABS 1301.0; 1950-present: Consumer Price Index ABS 6401.0. A quarterly seasonally adjusted series for the GDP deflator for the period 1959Q3-2008Q2 is from the *IMF*'s *IFS* (the series' code is 19399BIRZF...).

Output A real GDP annual series for the period 1885-2007 is from *GlobalFinancialData*. A quarterly seasonally adjusted real GDP series for the period 1959Q3-2008Q2 is from the *IMF*'s *IFS* (the series' code is 19399BVRZF...).

A.11 West Germany

All data for West Germany are from the *IMF*'s *IFS*.

Money Monthly seasonally adjusted series for M1, M2, and M3 are available for the periods January 1957-October 1989 (M1), and January 1969-October 1989 (M2 and M3). The series' IFS codes are 13459MACZF..., 13459MBCZF..., and 13459MCCZF... respectively.

Prices A monthly seasonally unadjusted CPI series is available for the period January 1957-October 1989 (the series' code is 13464.D.ZF...).

A.12 Norway

Money Annual series for the M0 and M2, available for the period 1819-2004 and 1819-2007, respectively, are from the *Bank of Norway*'s website. Monthly seasonally unadjusted series for M0, M1, and M2, available for the periods January 1850-June 2008, January 1913-December 1980, and January 1913-June 2008, respectively, are from the *Bank of Norway*'s website.

Prices An annual series for the CPI, available for the period 1516-2006, is from the *Bank of Norway*'s website. A monthly seasonally unadjusted CPI series, available for the period January 1940-August 2008, is from the *Bank of Norway*'s website. Monthly seasonally unadjusted series for the CPI and the WPI, available for the periods January 1920-August 2008 and January 1914-August 2008 respectively, are from *GlobalFinancialData*.

Output An annual real GDP series for the period 1830-2006, is from the *Bank of Norway*'s website.

A.13 Sweden

Money Annual series for M0 and M3, available for the period 1871-2006, are from the *Riksbank*'s website. A monthly seasonally unadjusted series for M3, available for the period January 1961-May 2008, is from the *IMF*'s *IFS* (the code is 14459MC.ZF...).

Prices An annual series for the CPI, available for the period 1871-2006, is from the *Riksbank*'s website. A monthly seasonally unadjusted CPI series, available for the period January 1961-May 2008, is from the *IMF*'s *IFS* (the code is 14464...ZF...).

Output An annual real GDP series, available for the period 1861-1994, is from the RW dataset.

Table 1 Cross-spectral gain between money growth and inflation at the low frequencies based on long-run data

	Estimate of the gain and 90%-coverage percentiles of the bootstrapped distribution, and p-value for rejecting H_0: gain=1			
	$\omega = 0$		Frequencies beyond 30 years	
	Gain	p-value	Average gain	p-value
United States				
Currency outside the Treasury, WPI (Jun. 1878-Jul. 1914)	1.94 [0.96; 3.41]	0.06	1.51 [0.82; 2.51]	0.14
M0, GNP deflator (1869-1983)	0.94 [0.40; 1.80]	0.43	0.63 [0.29; 1.18]	0.11
M0, GNP deflator (1875Q2-1983Q4)	0.82 [0.29; 1.64]	0.33	0.45 [0.17; 0.97]	0.04
St. Louis adj. monetary base, CPI (Jan. 1921-Aug. 2008)	0.48 [0.13; 1.58]	0.16	0.38 [0.16; 0.97]	0.04
'Money stock',# CPI (Jan. 1913-Jun. 1969)	0.27 [0.13; 0.56]	0.00	0.58 [0.31; 1.06]	0.07
M1, GNP deflator (1915Q2-2008Q2)	0.82 [0.31; 2.00]	0.36	0.50 [0.21; 1.00]	0.05
M2, CPI (1821-1866)	0.76 [0.22; 1.37]	0.23	0.81 [0.24; 1.44]	0.29
M2, GNP deflator (1875Q2-2008Q2)	1.58 [0.55; 3.30]	0.23	0.95 [0.59; 1.50]	0.43
United Kingdom				
M0, RPI (1834-1879)	0.35 [0.06; 1.37]	0.10	0.22 [0.06; 0.70]	0.01
M0, RPI (1872-2005)	1.42 [0.75; 2.02]	0.14	1.19 [0.74; 1.64]	0.23
M0, WPI (Jan. 1885-Dec. 1969)	0.30 [0.10; 0.82]	0.02	0.74 [0.20; 1.49]	0.26
M0, RPI (Jan. 1914-May 1951)	0.94 [0.49; 1.72]	0.44	0.62 [0.25; 1.06]	0.07
M1, RPI (Dec. 1921-Dec. 1969)	0.45 [0.32; 0.79]	0.01	0.41 [0.28; 0.71]	0.00
M3, RPI (1872-2007)	0.83 [0.64; 1.04]	0.08	0.79 [0.59; 0.98]	0.03
M3, RPI (1922Q2-2008Q2)	0.90 [0.62; 1.21]	0.27	0.81 [0.53; 1.07]	0.11
Norway				
M0, CPI (1831-2007)	1.08 [0.58; 1.56]	0.38	0.91 [0.58; 1.25]	0.33
M1, CPI (Jan. 1920–Dec. 1980)	0.30 [0.14; 0.75]	0.02	0.22 [0.10; 0.47]	0.00
M2, CPI (Jan. 1920–Jun. 2008)	0.64 [0.45; 0.80]	0.00	0.61 [0.42; 0.75]	0.00
M2, CPI (1831-2007)	1.32 [0.71; 1.90]	0.20	0.82 [0.60; 1.12]	0.15

Commercial banks plus currency held by the public.

Table 1 (continued) Cross-spectral gain between money growth and inflation at the low frequencies based on long-run data

	Estimate of the gain and 90%-coverage percentiles of the bootstrapped distribution, and p-value for rejecting H_0: gain=1			
	$\omega = 0$		Frequencies beyond 30 years	
	Gain	p-value	Average gain	p-value
Canada				
M0, CPI (Jan. 1914-Dec. 2003)	0.33 [0.08; 0.77]	0.01	0.20 [0.05; 0.50]	0.00
M1, CPI (Jan. 1914-May. 2008)	0.28 [0.09; 0.88]	0.03	0.25 [0.12; 0.53]	0.00
M2, CPI (Jan. 1914-May. 2008)	0.50 [0.33; 0.99]	0.05	0.35 [0.21; 0.58]	0.00
M2, CPI (1872-2006)	1.06 [0.48; 1.71]	0.44	0.70 [0.51; 0.97]	0.03
Australia				
M1, CPI (1900-1973)	0.54 [0.42; 0.83]	0.00	0.48 [0.39; 0.68]	0.00
M2, CPI (1850-1983)	0.86 [0.45; 1.35]	0.27	0.76 [0.40; 1.11]	0.12
M3, CPI (1860-2006)	0.90 [0.63; 1.22]	0.28	0.87 [0.59; 1.16]	0.22
France				
M2, CPI (1910-1994)	1.01 [0.47; 1.62]	0.48	1.07 [0.50; 1.64]	0.41
Italy				
M2, CPI (1863-1998)	0.42 [0.15; 0.76]	0.01	0.44 [0.16; 0.70]	0.00
Netherlands				
M0, CPI (1881-1992)	0.55 [0.35; 0.85]	0.01	0.67 [0.43; 0.95]	0.03
M2, CPI (1881-1992)	0.96 [0.36; 1.63]	0.46	0.82 [0.31; 1.33]	0.27
Portugal				
M2, CPI (1855-1912)	0.21 [0.08; 0.60]	0.01	0.19 [0.09; 0.48]	0.00
M2, CPI (1932-1998)	1.04 [0.69; 1.31]	0.39	1.08 [0.74; 1.35]	0.32
Chile				
M0, CPI (1862-1913)	1.96 [0.36; 4.69]	0.22	1.60 [0.29; 3.93]	0.30
M2, CPI (1861-1913)	0.80 [0.29; 1.60]	0.32	0.78 [0.32; 1.40]	0.27

Table 1 (continued) Cross-spectral gain between money growth and inflation at the low frequencies based on long-run data

	Estimate of the gain and 90%-coverage percentiles of the bootstrapped distribution, and p-value for rejecting H_0: gain=1			
	$\omega = 0$		Frequencies beyond 30 years	
	Gain	p-value	Average gain	p-value
Switzerland				
M0, CPI (1908-2005)	0.72 [0.11; 1.68]	0.29	0.43 [0.10; 0.83]	0.01
M1, CPI (1908-2005)	0.12 [0.02; 0.36]	0.00	0.60 [0.18; 1.12]	0.10
M3, CPI (1908-2005)	0.63 [0.19; 1.37]	0.16	0.73 [0.23; 1.51]	0.25
Sweden				
M0, CPI (1872-2006)	0.79 [0.56; 1.20]	0.17	0.73 [0.50; 1.07]	0.09
M2, CPI (1872-1988)	0.97 [0.59; 1.34]	0.44	0.84 [0.56; 1.12]	0.16
M3, CPI (1872-2006)	1.85 [0.65; 2.90]	0.12	1.13 [0.54; 1.77]	0.36

Table 2 Cross-spectral coherence between money growth and inflation at the low frequencies based on long-run data

	Estimate of the coherence and 90%-coverage percentiles of the bootstrapped distribution, and fraction of the mass of the distribution beyond 0.99			
	$\omega = 0$		Frequencies beyond 30 years	
	Coherence	Mass beyond 0.99	Average coherence	Mass beyond 0.99
United States				
Currency outside the Treasury, WPI (Jun. 1878-Jul. 1914)	0.88 [0.66; 0.97]	0.00	0.85 [0.65; 0.96]	0.00
M0, GNP deflator (1869-1983)	0.97 [0.87; 1.00]	0.19	0.61 [0.27; 0.90]	0.00
M0, GNP deflator (1875Q2-1983Q4)	0.96 [0.83; 0.99]	0.11	0.50 [0.20; 0.87]	0.00
St. Louis adj. monetary base, CPI (Jan. 1921-Aug. 2008)	0.22 [0.05; 0.74]	0.00	0.46 [0.18; 0.84]	0.00
'Money stock',# CPI (Jan. 1913-Jun. 1969)	0.27 [0.07; 0.78]	0.00	0.75 [0.39; 0.95]	0.00
M1, GNP deflator (1915Q2-2008Q2)	0.63 [0.33; 0.93]	0.00	0.68 [0.33; 0.94]	0.00
M2, CPI (1821-1866)	0.92 [0.47; 0.99]	0.04	0.84 [0.31; 0.97]	0.00
M2, GNP deflator (1875Q2-2008Q2)	0.68 [0.27; 0.94]	0.00	0.80 [0.56; 0.93]	0.00
United Kingdom				
M0, RPI (1834-1879)	0.50 [0.08; 0.94]	0.01	0.47 [0.16; 0.85]	0.00
M0, RPI (1872-2005)	0.98 [0.78; 1.00]	0.25	0.89 [0.64; 0.97]	0.00
M0, WPI (Jan. 1885-Dec. 1969)	0.52 [0.17; 0.91]	0.00	0.75 [0.26; 0.95]	0.00
M0, RPI (Jan. 1914-May 1951)	0.82 [0.54; 0.96]	0.00	0.70 [0.30; 0.89]	0.00
M1, RPI (Dec. 1921-Dec. 1969)	0.93 [0.75; 0.98]	0.02	0.88 [0.69; 0.97]	0.00
M3, RPI (1872-2007)	0.98 [0.89; 1.00]	0.29	0.93 [0.79; 0.98]	0.00
M3, RPI (1922Q2-2008Q2)	0.97 [0.85; 1.00]	0.17	0.95 [0.80; 0.99]	0.02
Norway				
M0, CPI (1831-2007)	0.98 [0.81; 1.00]	0.26	0.87 [0.63; 0.96]	0.00
M1, CPI (Jan. 1920–Dec. 1980)	0.47 [0.14; 0.89]	0.00	0.71 [0.33; 0.95]	0.00
M2, CPI (Jan. 1920–Jun. 2008)	0.98 [0.87; 1.00]	0.31	0.97 [0.86; 0.99]	0.13
M2, CPI (1831-2007)	0.96 [0.74; 0.99]	0.08	0.90 [0.75; 0.97]	0.00

Commercial banks plus currency held by the public.

Table 2 (continued) Cross-spectral coherence between money growth and inflation at the low frequencies based on long-run data

	Estimate of the coherence and 90%-coverage percentiles of the bootstrapped distribution, and fraction of the mass of the distribution beyond 0.99			
	$\omega = 0$		Frequencies beyond 30 years	
	Coherence	Mass beyond 0.99	Average coherence	Mass beyond 0.99
Canada				
M0, CPI (Jan. 1914-Dec. 2003)	0.69 [0.18; 0.95]	0.00	0.43 [0.10; 0.84]	0.00
M1, CPI (Jan. 1914-May. 2008)	0.34 [0.10; 0.80]	0.00	0.52 [0.24; 0.87]	0.00
M2, CPI (1872-2006)	0.95 [0.67; 0.99]	0.08	0.91 [0.73; 0.97]	0.00
M2, CPI (Jan. 1914-May. 2008)	0.84 [0.65; 0.96]	0.00	0.75 [0.52; 0.93]	0.00
Australia				
M1, CPI (1900-1973)	0.92 [0.80; 0.98]	0.02	0.94 [0.86; 0.98]	0.01
M2, CPI (1850-1983)	0.99 [0.90; 1.00]	0.39	0.80 [0.44; 0.93]	0.00
M3, CPI (1860-2006)	0.98 [0.89; 1.00]	0.30	0.89 [0.66; 0.96]	0.00
France				
M2, CPI (1910-1994)	0.95 [0.68; 0.99]	0.06	0.85 [0.43; 0.96]	0.00
Italy				
M2, CPI (1863-1998)	0.97 [0.67; 1.00]	0.25	0.77 [0.27; 0.95]	0.00
Netherlands				
M0, CPI (1881-1992)	0.93 [0.73; 0.99]	0.03	0.90 [0.69; 0.97]	0.00
M2, CPI (1881-1992)	0.92 [0.57; 0.99]	0.03	0.80 [0.32; 0.95]	0.00
Portugal				
M2, CPI (1855-1912)	0.37 [0.18; 0.75]	0.00	0.46 [0.28; 0.75]	0.00
M2, CPI (1932-1998)	0.98 [0.84; 1.00]	0.35	0.97 [0.81; 0.99]	0.14
Chile				
M0, CPI (1862-1913)	0.77 [0.29; 0.96]	0.00	0.63 [0.18; 0.91]	0.00
M2, CPI (1861-1913)	0.80 [0.38; 0.96]	0.00	0.80 [0.42; 0.95]	0.00

Table 2 (continued) Cross-spectral coherence between money growth and inflation at the low frequencies based on long-run data

	Estimate of the coherence and 90%-coverage percentiles of the bootstrapped distribution, and fraction of the mass of the distribution beyond 0.99			
	$\omega = 0$		Frequencies beyond 30 years	
	Coherence	Mass beyond 0.99	Average coherence	Mass beyond 0.99
Switzerland				
M0, CPI (1908-2005)	0.68 [0.11; 0.95]	0.00	0.72 [0.17; 0.94]	0.00
M1, CPI (1908-2005)	0.12 [0.01; 0.59]	0.00	0.61 [0.19; 0.90]	0.00
M3, CPI (1908-2005)	0.78 [0.28; 0.97]	0.01	0.67 [0.22; 0.90]	0.00
Sweden				
M0, CPI (1872-2006)	0.98 [0.91; 1.00]	0.18	0.85 [0.66; 0.95]	0.00
M2, CPI (1872-1988)	0.99 [0.92; 1.00]	0.42	0.93 [0.73; 0.98]	0.01
M3, CPI (1872-2006)	0.94 [0.53; 0.99]	0.04	0.82 [0.42; 0.95]	0.00

Table 3 Cross-spectral gain between money growth and inflation at the low frequencies, post-WWII period

	Estimate of the gain and 90%-coverage percentiles of the bootstrapped distribution, and p-value for rejecting H_0: gain=1			
	$\omega = 0$		Frequencies beyond 30 years	
	Gain	p-value	Average gain	p-value
Euro area				
M1, HICP (Jan. 1980-Aug. 2008)	1.28 [0.93; 1.65]	0.11	–	–
M2, HICP (Jan. 1980-Aug. 2008)	1.04 [0.48; 1.63]	0.44	–	–
M2, HICP (1970Q3-2008Q2)	0.88 [0.71; 1.09]	0.14	0.88 [0.71; 1.09]	0.14
M3, HICP (Jan. 1980-Aug. 2008)	0.82 [0.27; 1.41]	0.23	–	–
M3, HICP (1970Q2-2008Q1)	0.98 [0.72; 1.29]	0.44	0.97 [0.70; 1.27]	0.41
United States				
M1, CPI (Feb. 1959-Jul. 2008)	0.82 [0.34; 1.37]	0.24	0.79 [0.34; 1.35]	0.23
M2, CPI (Feb. 1959-Jul. 2008)	1.06 [0.61; 1.61]	0.41	1.04 [0.60; 1.57]	0.45
M3, CPI (Feb. 1959-Feb. 2006)	1.06 [0.57; 1.70]	0.43	0.96 [0.50; 1.56]	0.45
Japan				
M1, CPI (Feb. 1957-Dec. 2007)	0.86 [0.62; 1.16]	0.21	0.80 [0.58; 1.09]	0.12
M2, CPI (Feb. 1957-Dec. 2007)	0.53 [0.36; 0.79]	0.01	0.51 [0.33; 0.75]	0.01
West Germany				
M1, CPI (Feb. 1957-Oct. 1989)	1.68 [0.63; 4.46]	0.23	0.58 [0.26; 1.58]	0.20
M2, CPI (Feb. 1969-Oct. 1989)	0.65 [0.47; 0.92]	0.02	–	–
M3, CPI (Feb. 1969-Oct. 1989)	0.75 [0.48; 1.23]	0.16	–	–
United Kingdom				
M0, RPI (Jun. 1969-Apr. 2006)	2.34 [0.69; 4.40]	0.11	2.11 [0.62; 3.81]	0.12
M4, GDP deflator (1963Q2-2008Q2)	1.23 [0.42; 1.97]	0.28	1.20 [0.43; 1.90]	0.31

Table 3 (continued) Cross-spectral gain between money growth and inflation at the low frequencies, post-WWII period

	Estimate of the gain and 90%-coverage percentiles of the bootstrapped distribution, and p-value for rejecting H_0: gain=1			
	$\omega = 0$		Frequencies beyond 30 years	
	Gain	p-value	Average gain	p-value
Australia				
M1, GDP deflator (1975Q2-2008Q1)	1.69 [0.65; 2.39]	0.12	1.63 [0.75; 2.27]	0.10
M3, GDP deflator (1959Q4-2008Q1)	1.81 [0.73; 3.34]	0.14	1.79 [0.77; 3.37]	0.13
Italy				
M2, CPI (Jan. 1948-Sep. 1997)	1.31 [0.88; 1.77]	0.15	1.20 [0.83; 1.65]	0.24
Netherlands				
M2, CPI (Jan. 1957-Dec. 1997)	0.73 [0.37; 1.00]	0.05	0.73 [0.43; 0.98]	0.04
Switzerland				
M1, CPI (Feb. 1950-Dec. 2006)	0.79 [0.33; 1.81]	0.33	0.67 [0.28; 1.51]	0.23
M2, CPI (Jul. 1975-Dec. 2006)	0.36 [0.28; 0.55]	0.00	0.33 [0.26; 0.51]	0.00
M3, CPI (Jul. 1975-Dec. 2006)	0.36 [0.27; 0.55]	0.00	0.33 [0.26; 0.51]	0.00
Canada				
M1, CPI (Feb. 1946-Dec. 2006)	0.33 [0.08; 0.77]	0.01	0.20 [0.05; 0.50]	0.00
M3, CPI (Jan. 1970-Feb. 1970-May 2008)	0.50 [0.33; 0.99]	0.05	0.35 [0.21; 0.58]	0.00
Sweden				
M3, CPI (Feb. 1961-Jun. 2008)	1.20 [0.46; 1.73]	0.29	1.12 [0.47; 1.62]	0.36

Table 4 Cross-spectral coherence between money growth and inflation at the low frequencies, post-WWII period

	Estimate of the coherence and 90%-coverage percentiles of the bootstrapped distribution, and fraction of the mass of the distribution beyond 0.99			
	$\omega = 0$		Frequencies beyond 30 years	
	Coherence	Mass beyond 0.99	Average coherence	Mass beyond 0.99
Euro area				
M1, HICP (Jan. 1980-Aug. 2008)	0.97 [0.91; 0.99]	0.12	–	–
M2, HICP (Jan. 1980-Aug. 2008)	0.98 [0.76; 1.00]	0.31	–	–
M2, HICP (1970Q3-2008Q2)	1.00 [0.97; 1.00]	0.75	0.99 [0.94; 1.00]	0.43
M3, HICP (Jan. 1980-Aug. 2008)	0.99 [0.71; 1.00]	0.39	–	–
M3, HICP (1970Q2-2008Q1)	1.00 [0.97; 1.00]	0.82	0.98 [0.88; 1.00]	0.32
United States				
M1, CPI (Feb. 1959-Jul. 2008)	0.97 [0.73; 1.00]	0.18	0.91 [0.57; 0.99]	0.03
M2, CPI (Feb. 1959-Jul. 2008)	0.96 [0.90; 0.99]	0.12	0.94 [0.76; 0.99]	0.04
M3, CPI (Feb. 1959-Feb. 2006)	0.96 [0.81; 0.99]	0.10	0.91 [0.68; 0.98]	0.01
Japan				
M1, CPI (Feb. 1957-Dec. 2007)	0.98 [0.92; 1.00]	0.24	0.97 [0.89; 0.99]	0.0879
M2, CPI (Feb. 1957-Dec. 2007)	0.99 [0.93; 1.00]	0.51	0.96 [0.80; 0.99]	0.1395
West Germany				
M1, CPI (Feb. 1957-Oct. 1989)	0.80 [0.58; 0.96]	0.00	0.52 [0.31; 0.85]	0.00
M2, CPI (Feb. 1969-Oct. 1989)	0.93 [0.76; 0.99]	0.03	–	–
M3, CPI (Feb. 1969-Oct. 1989)	0.97 [0.86; 1.00]	0.15	–	–
United Kingdom				
M0, RPI (Jun. 1969-Apr. 2006)	0.91 [0.47; 0.99]	0.05	0.86 [0.37; 0.98]	0.01
M4, GDP deflator (1963Q2-2008Q2)	0.97 [0.69; 1.00]	0.24	0.93 [0.49; 0.99]	0.06

Table 4 (continued) Cross-spectral coherence between money growth and inflation at the low frequencies, post-WWII period

	Estimate of the coherence and 90%-coverage percentiles of the bootstrapped distribution, and fraction of the mass of the distribution beyond 0.99			
	$\omega = 0$		Frequencies beyond 30 years	
	Coherence	Mass beyond 0.99	Average coherence	Mass beyond 0.99
Australia				
M1, GDP deflator (1975Q2-2008Q1)	0.83 [0.56; 0.96]	0.00	0.83 [0.61; 0.95]	0.00
M3, GDP deflator (1959Q4-2008Q1)	0.97 [0.65; 1.00]	0.16	0.96 [0.69; 0.99]	0.09
Italy				
M2, CPI (Jan. 1948-Sep. 1997)	0.99 [0.97; 1.00]	0.56	0.94 [0.86; 0.98]	0.01
Netherlands				
M2, CPI (Jan. 1957-Dec. 1997)	0.98 [0.75; 1.00]	0.23	0.97 [0.80; 0.99]	0.16
Switzerland				
M1, CPI (Feb. 1950-Dec. 2006)	0.73 [0.43; 0.93]	0.00	0.68 [0.41; 0.90]	0.00
M2, CPI (Jul. 1975-Dec. 2006)	0.97 [0.89; 0.99]	0.13	0.92 [0.80; 0.98]	0.02
M3, CPI (Jul. 1975-Dec. 2006)	0.9673 [0.89; 0.99]	0.14	0.92 [0.79; 0.98]	0.02
Canada				
M1, CPI (Feb. 1946-Dec. 2006)	0.6913 [0.18; 0.95]	0.00	0.43 [0.10; 0.84]	0.00
M3, CPI (Jan. 1970-Feb. 1970-May 2008)	0.8442 [0.65; 0.96]	0.00	0.75 [0.52; 0.93]	0.00
Sweden				
M3, CPI (Feb. 1961-Jun. 2008)	0.98 [0.72; 1.00]	0.26	0.95 [0.60; 0.99]	0.08

Table 5 Bayesian estimates of the structural parameters for the model of Ascari and Ropele (2007)

Parameter	Domain	Density	Prior distribution		Posterior distribution: mode and 90%-coverage percentiles	
			Mode	St. dev.	*Euro area*	*United States*
σ_R^2	$\mathbb{R}^+$	Inverse Gamma	0.5	5	1.10 [0.60; 1.84]	1.24 [1.00; 1.62]
σ_π^2	$\mathbb{R}^+$	Inverse Gamma	0.5	5	3.71 [2.23; 6.96]	1.49 [1.22; 1.91]
σ_y^2	$\mathbb{R}^+$	Inverse Gamma	0.5	5	0.14 [0.05; 0.46]	0.19 [0.14; 0.27]
σ_v^2	$\mathbb{R}^+$	Inverse Gamma	0.5	5	0.57 [0.21; 3.10]	0.62 [0.39; 1.34]
θ-1	$\mathbb{R}^+$	Gamma	10	5	5.71 [1.45; 12.41]	12.52 [6.34; 17.92]
α	(0; 1]	Beta	0.59	0.02	0.59 [0.52; 0.67]	0.65 [0.63; 0.68]
ϵ	[0; 1]	Uniform	–	0.29	0.13 [0.01; 0.34]	0.45 [0.36; 0.53]
σ	$\mathbb{R}^+$	Gamma	2	1	8.74 [2.97; 13.34]	7.77 [5.21;11.19]
δ	[0; 1]	Uniform	–	0.29	0.46 [0.03; 0.94]	0.58 [0.48; 0.69]
δ_0	[0; 100]	Uniform	–	29	4.75 [0.50; 8.92]	3.11 [1.48; 4.68]
γ_1	$\mathbb{R}^+$	Gamma	1	0.05	0.99 [0.80; 1.36]	1.01 [0.92; 1.09]
γ_2	$\mathbb{R}^+$	Gamma	0.1	0.01	0.10 [0.07; 0.30]	0.10 [0.08; 0.12]
ρ	[0; 1)	Beta	0.8	0.1	0.83 [0.73; 0.89]	0.75 [0.70; 0.79]
ϕ_π	$\mathbb{R}^+$	Gamma	1	0.5	2.36 [1.53; 3.50]	2.70 [2.28; 3.18]
ϕ_y	$\mathbb{R}^+$	Gamma	0.1	0.5	2.66 [0.55; 5.68]	1.75 [1.27; 2.39]
ρ_R	[0; 1)	Beta	0.25	0.1	0.33 [0.12; 0.62]	0.34 [0.22; 0.48]
ρ_y	[0; 1)	Beta	0.25	0.1	0.69 [0.51; 0.83]	0.56 [0.44; 0.66]
ρ_v	[0; 1)	Beta	0.25	0.1	0.20 [0.06; 0.56]	0.26 [0.12; 0.43]

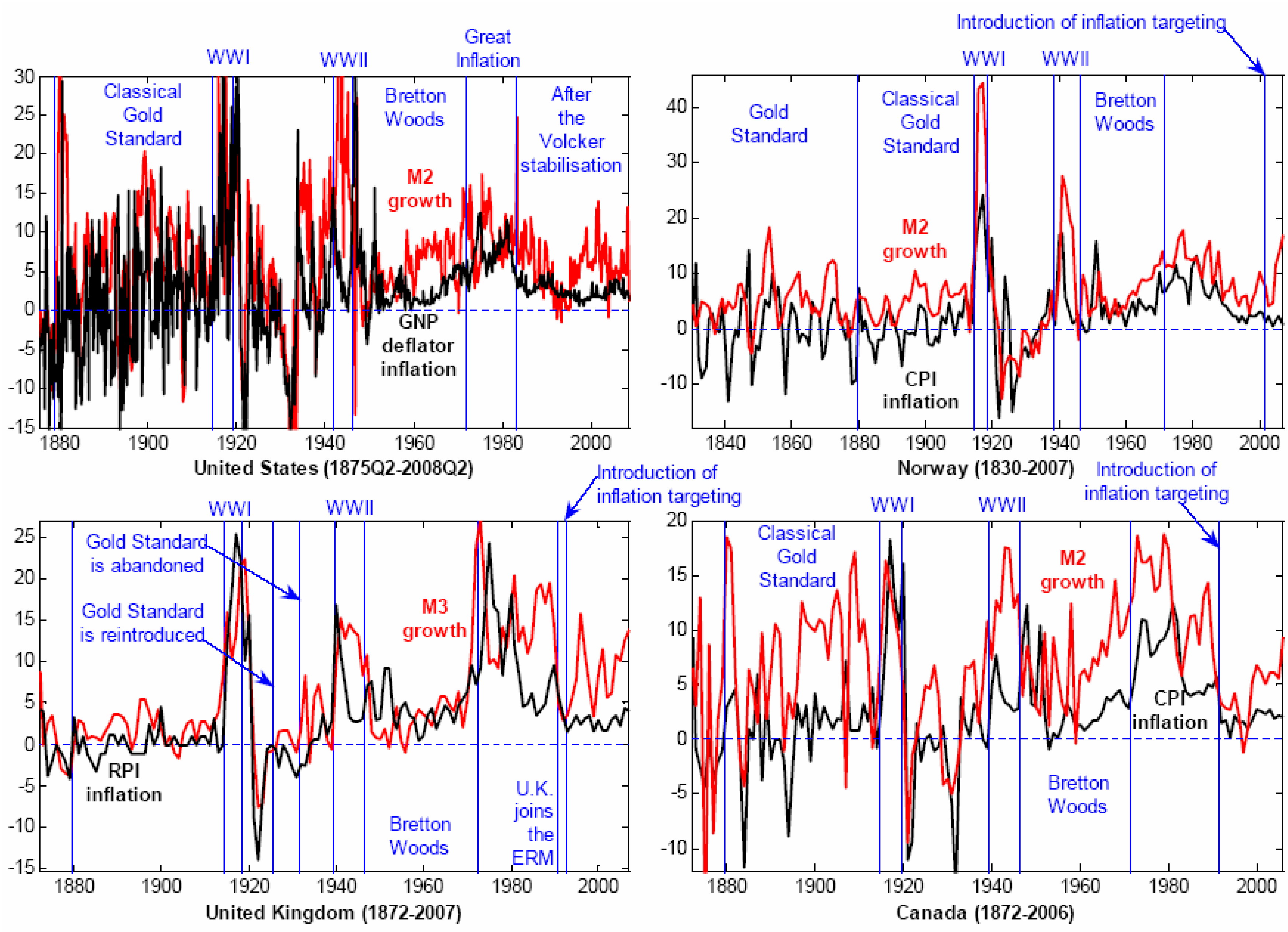

Figure 1 Inflation and broad money growth, long-run data

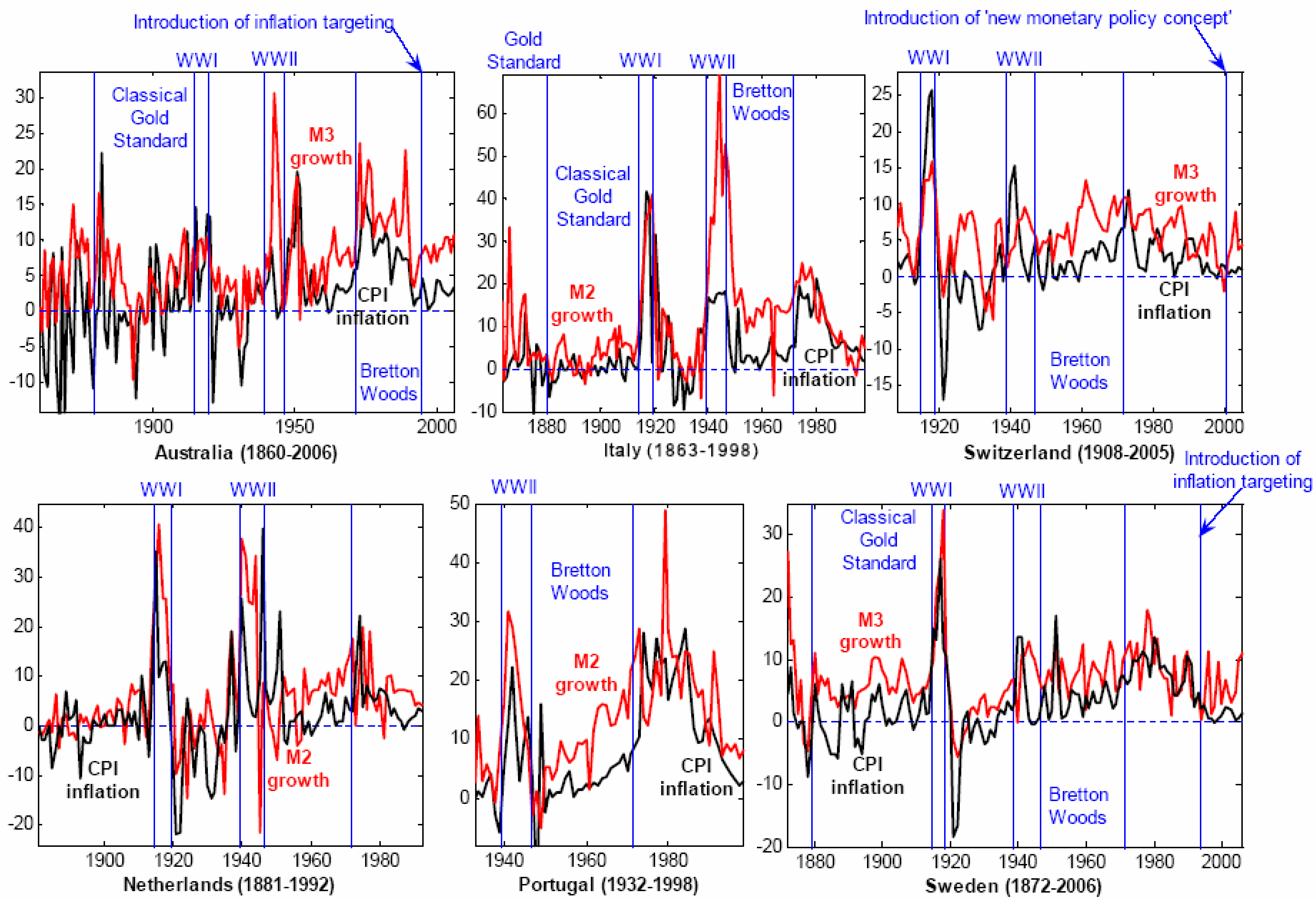

Figure 1 (continued) Inflation and broad money growth, long-run data

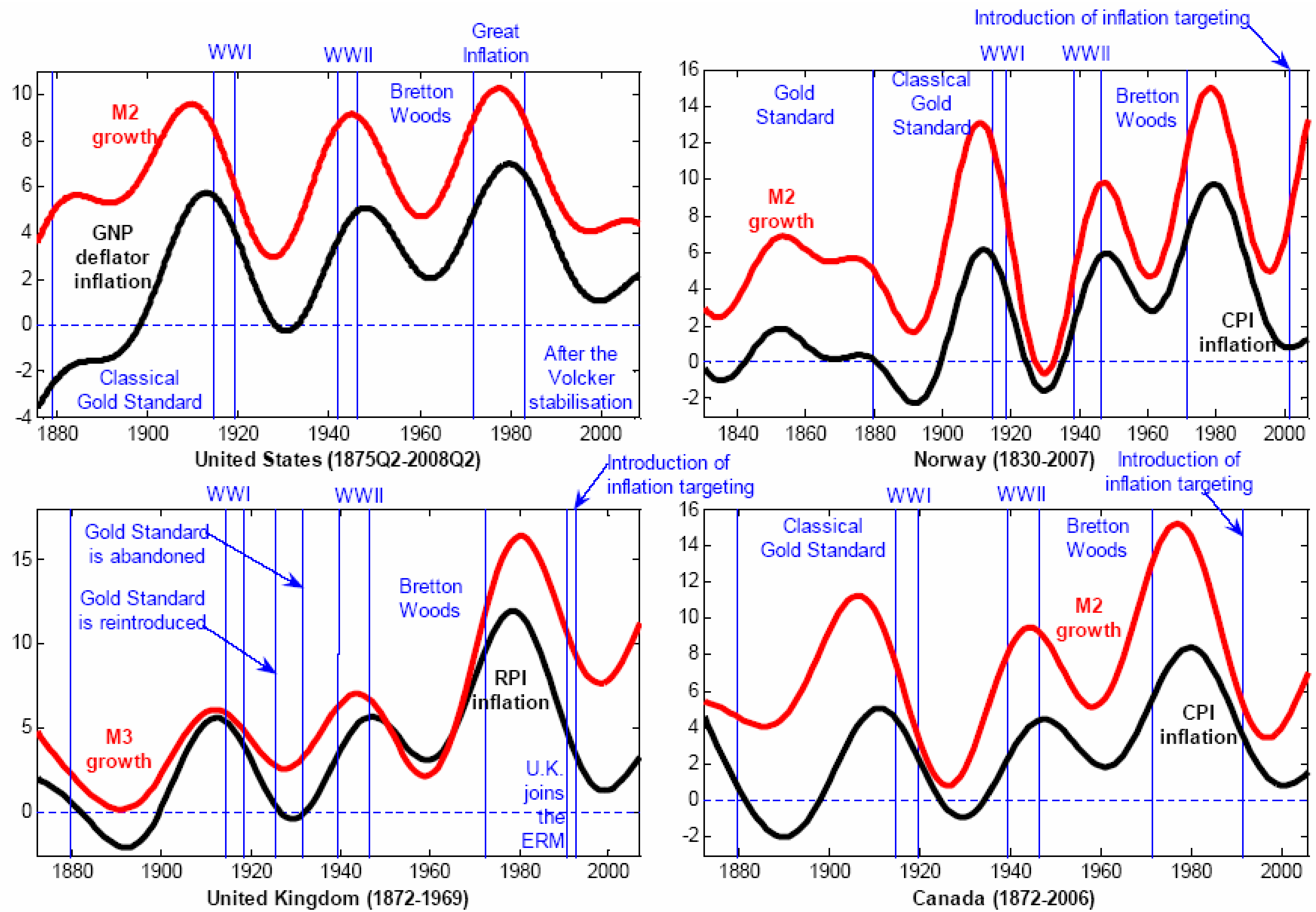

Figure 2 Low-frequency components of inflation and broad money growth, long-run data

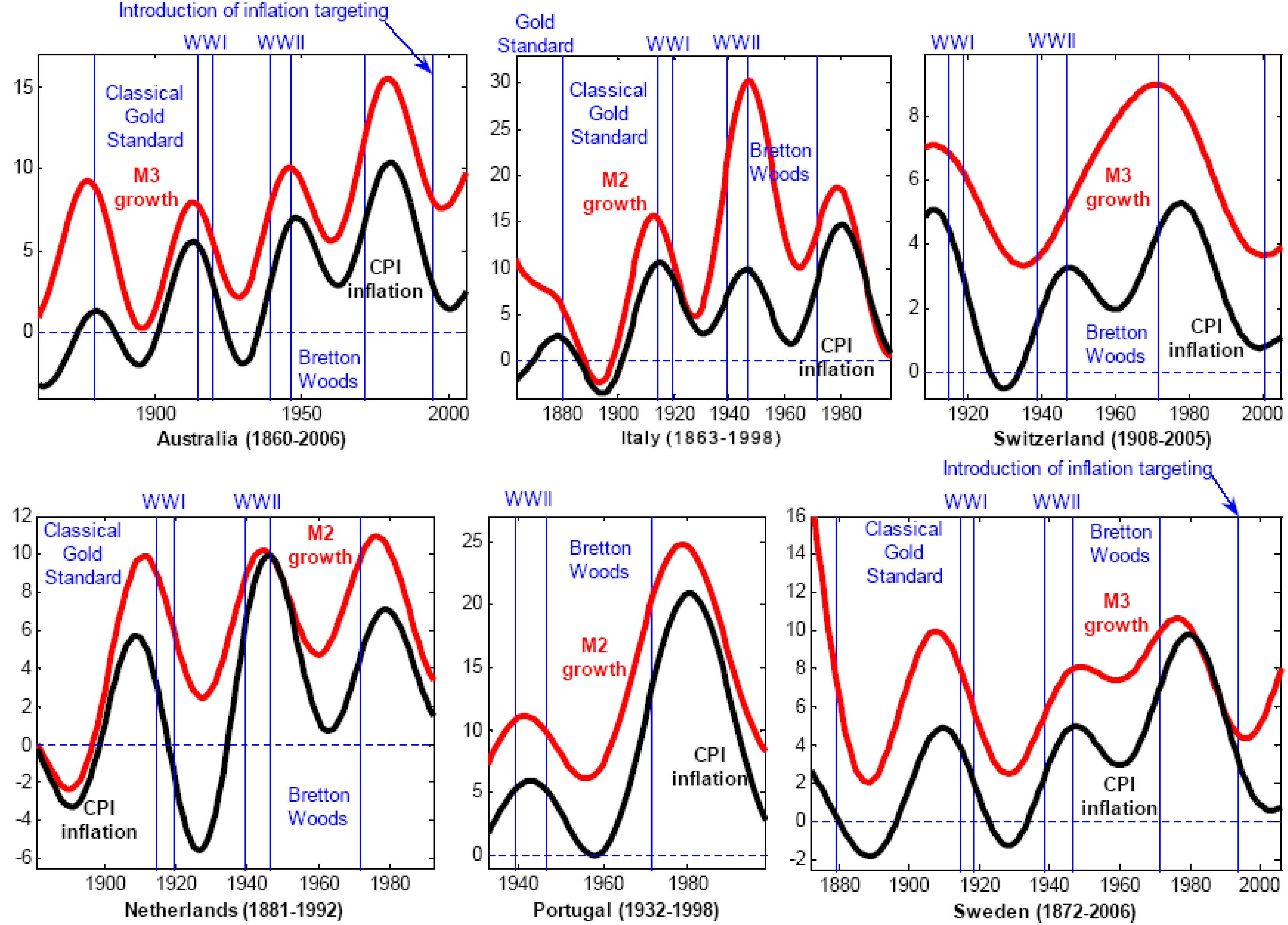

Figure 2 (continued) Low-frequency components of inflation and broad money growth, long-run data

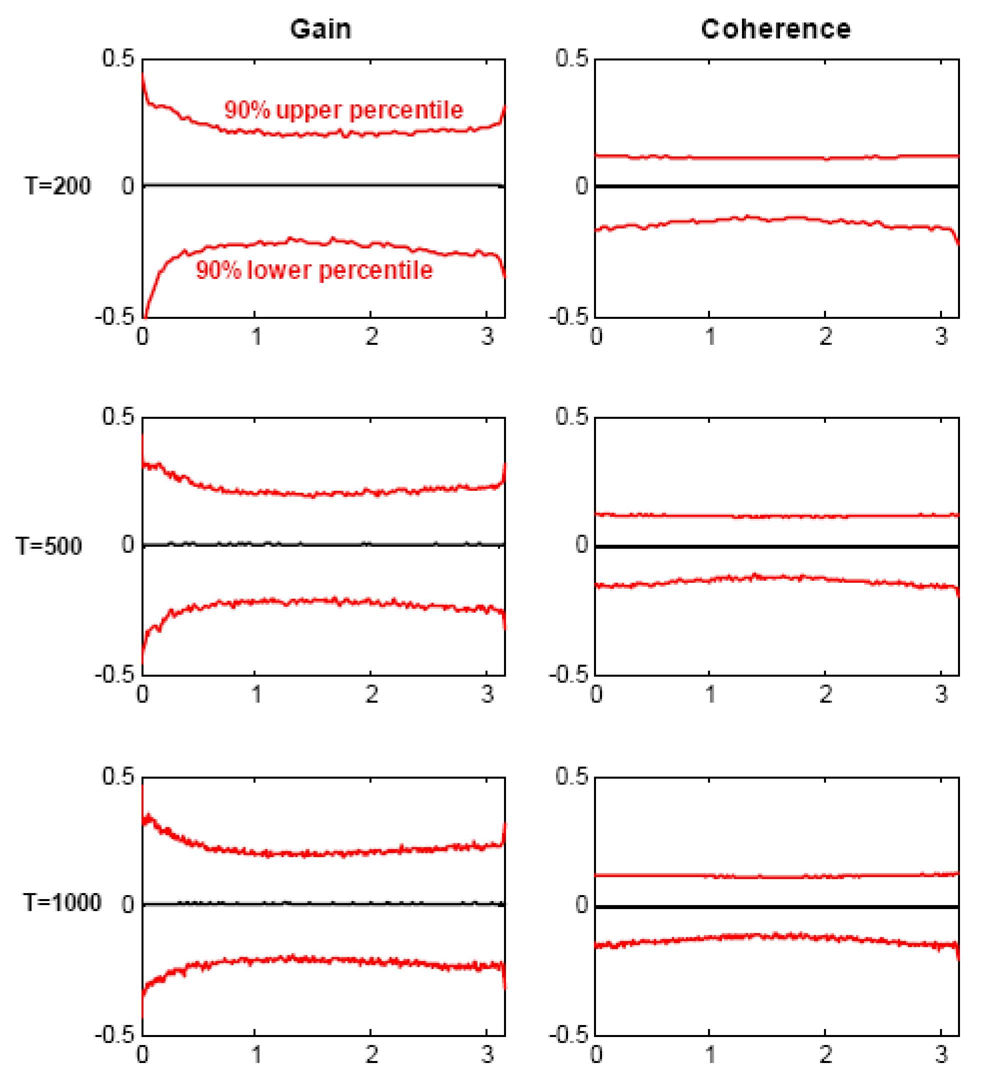

Figure 3 Monte Carlo evidence on the performance of the gain and coherence estimators: medians of the distributions of the difference between estimated and theoretical cross-spectral objects, and 90% lower and upper percentiles

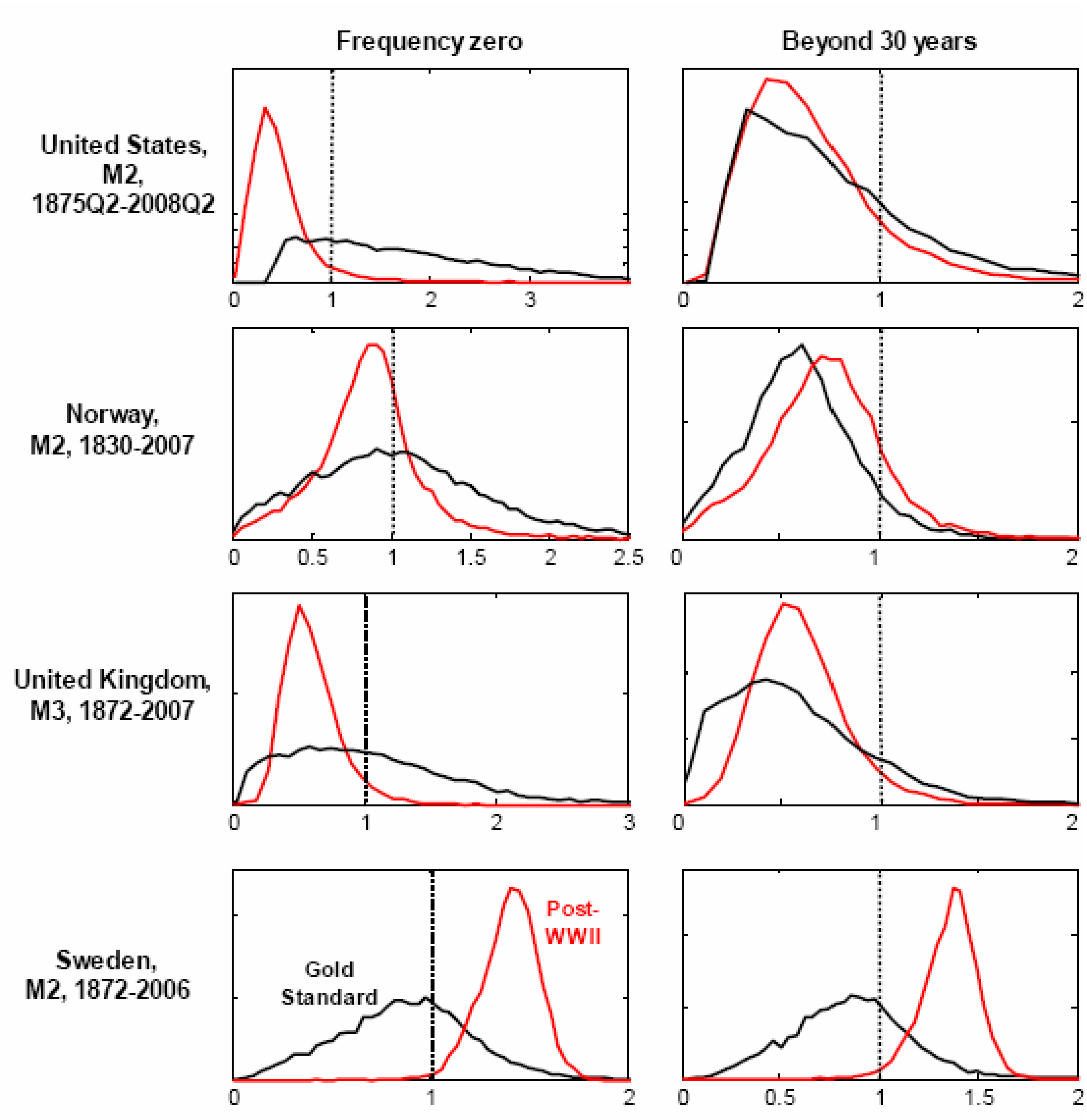

Figure 4 Comparing the Gold Standard and the post-WWII period: bootstrapped distributions of the cross-spectral gain between broad money growth and inflation, frequency zero, and average within the frequency bands beyond 15 and 30 years

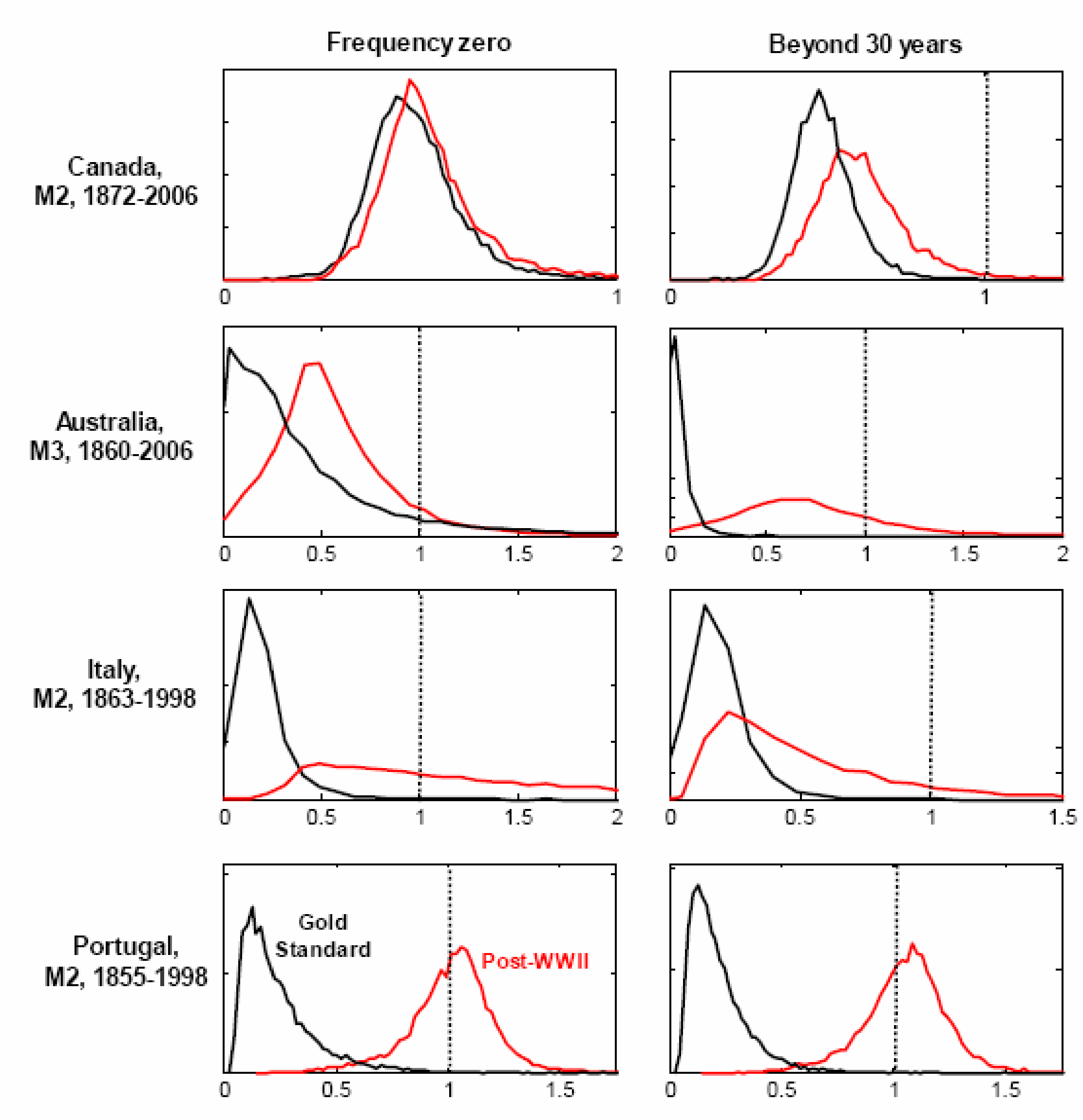

Figure 4 (continued) Comparing the Gold Standard and the post-WWII period: bootstrapped distributions of the cross-spectral gain between broad money growth and inflation, frequency zero, and average within the frequency bands beyond 15 and 30 years

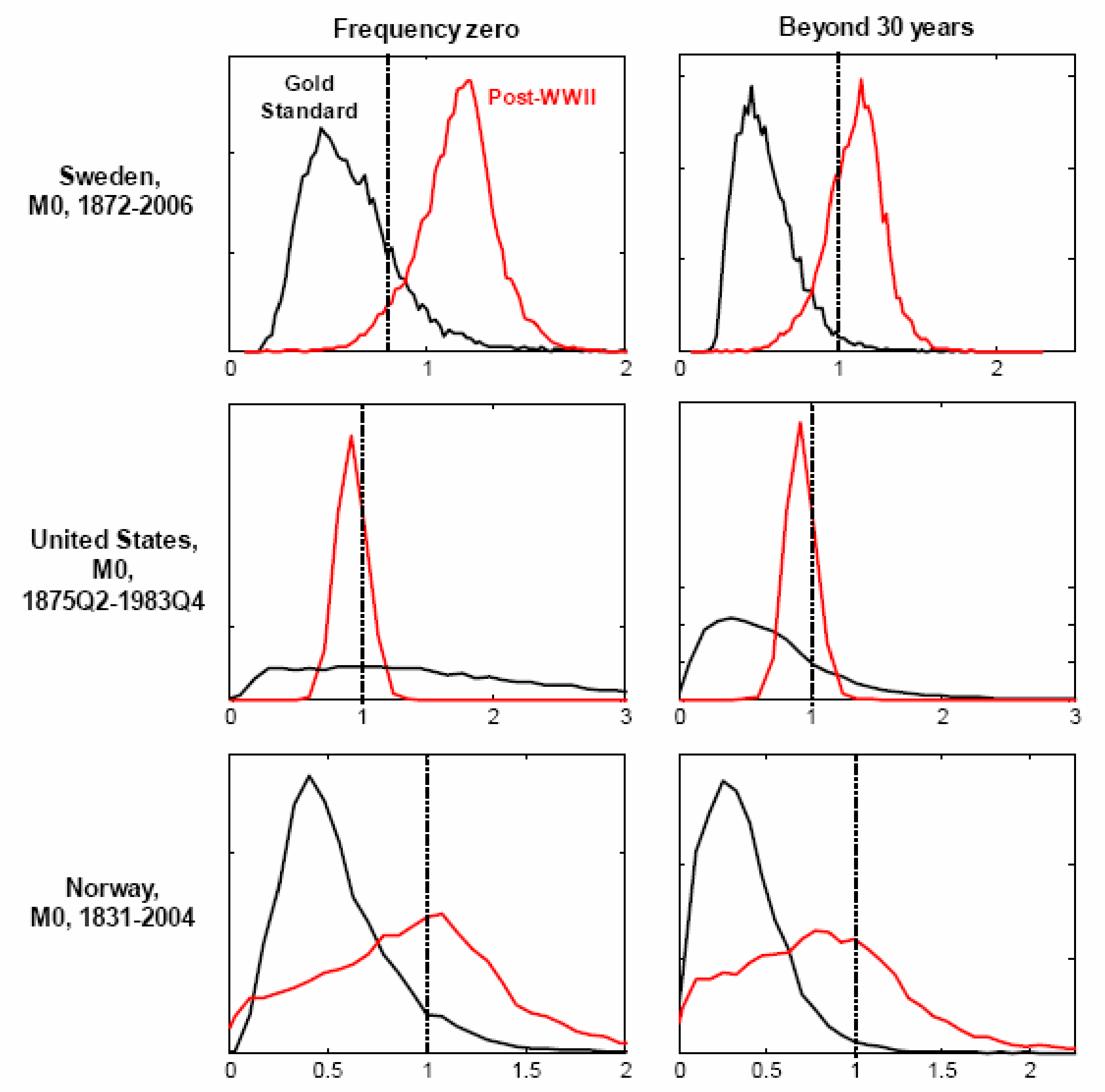

Figure 4 (continued) Comparing the Gold Standard and the post-WWII period: bootstrapped distributions of the cross-spectral gain between broad money growth and inflation, frequency zero, and average within the frequency band beyon 30 years

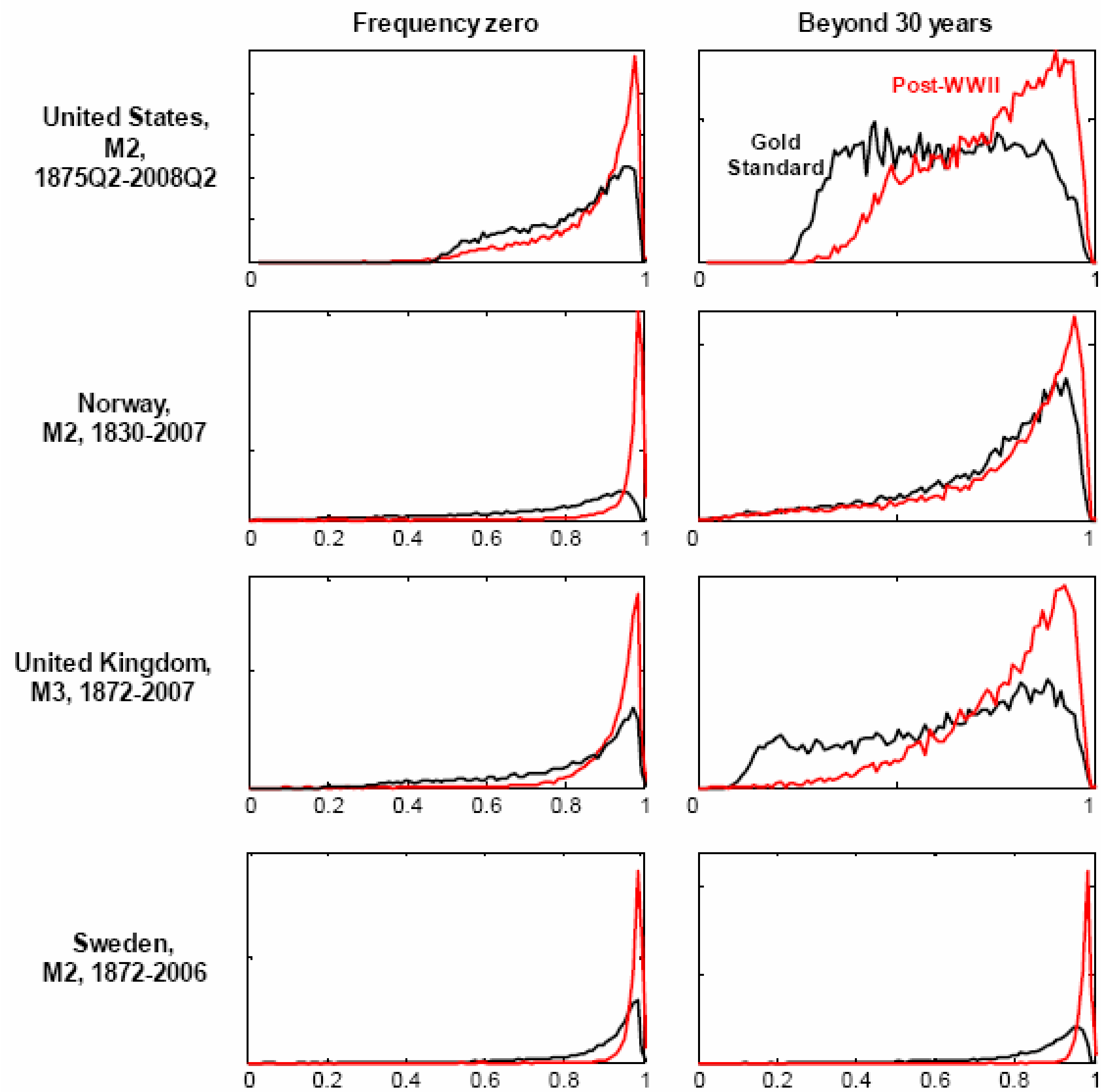

Figure 5 Comparing the Gold Standard and the post-WWII period: bootstrapped distributions of the cross-spectral coherence between broad money growth and inflation, frequency zero, and average within the frequency bands beyond 15 and 30 years

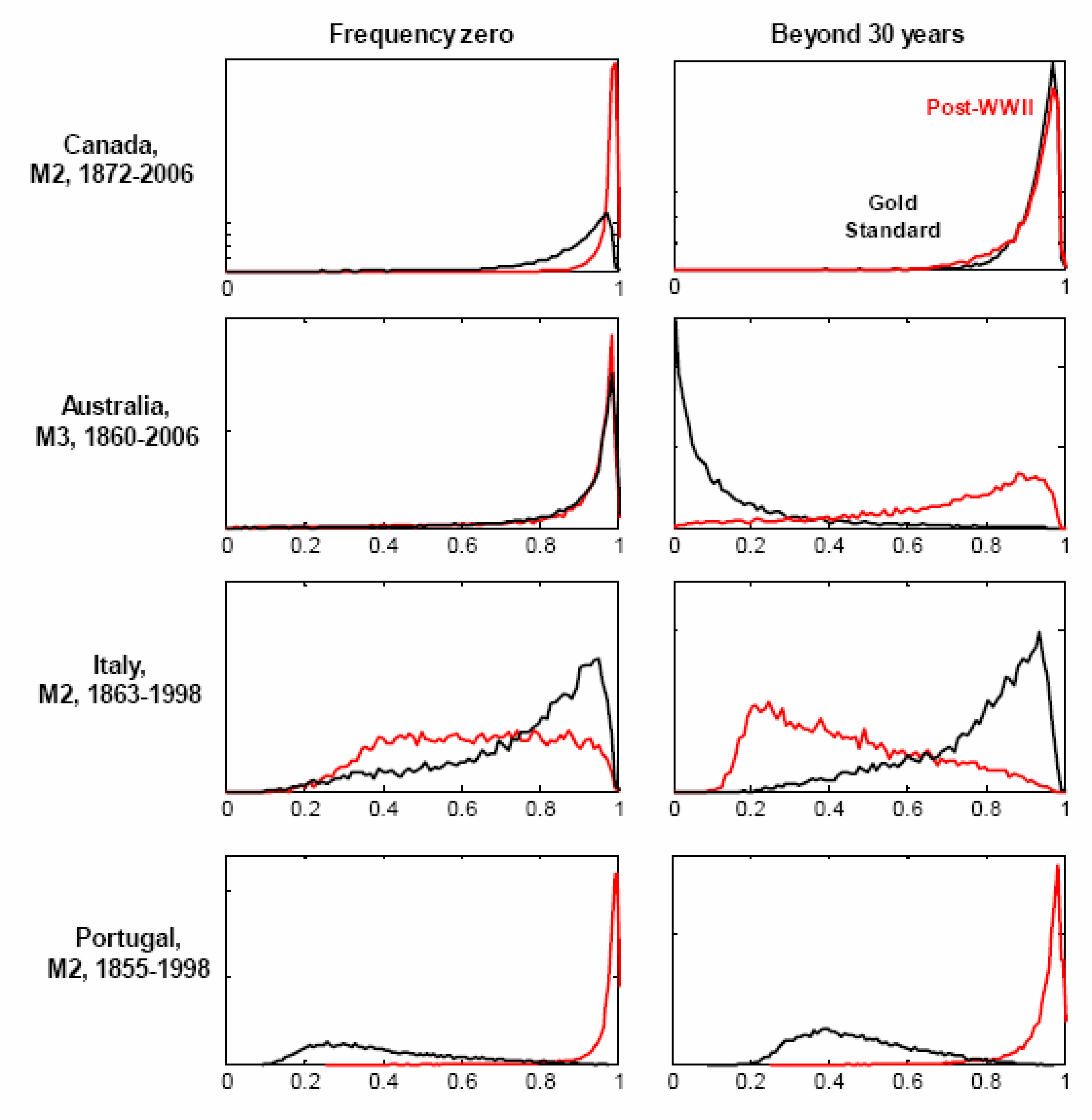

Figure 5 (continued) Comparing the Gold Standard and the post-WWII period: bootstrapped distributions of the cross-spectral coherence between broad money growth and inflation, frequency zero, and average within the frequency bands beyond 15 and 30 years

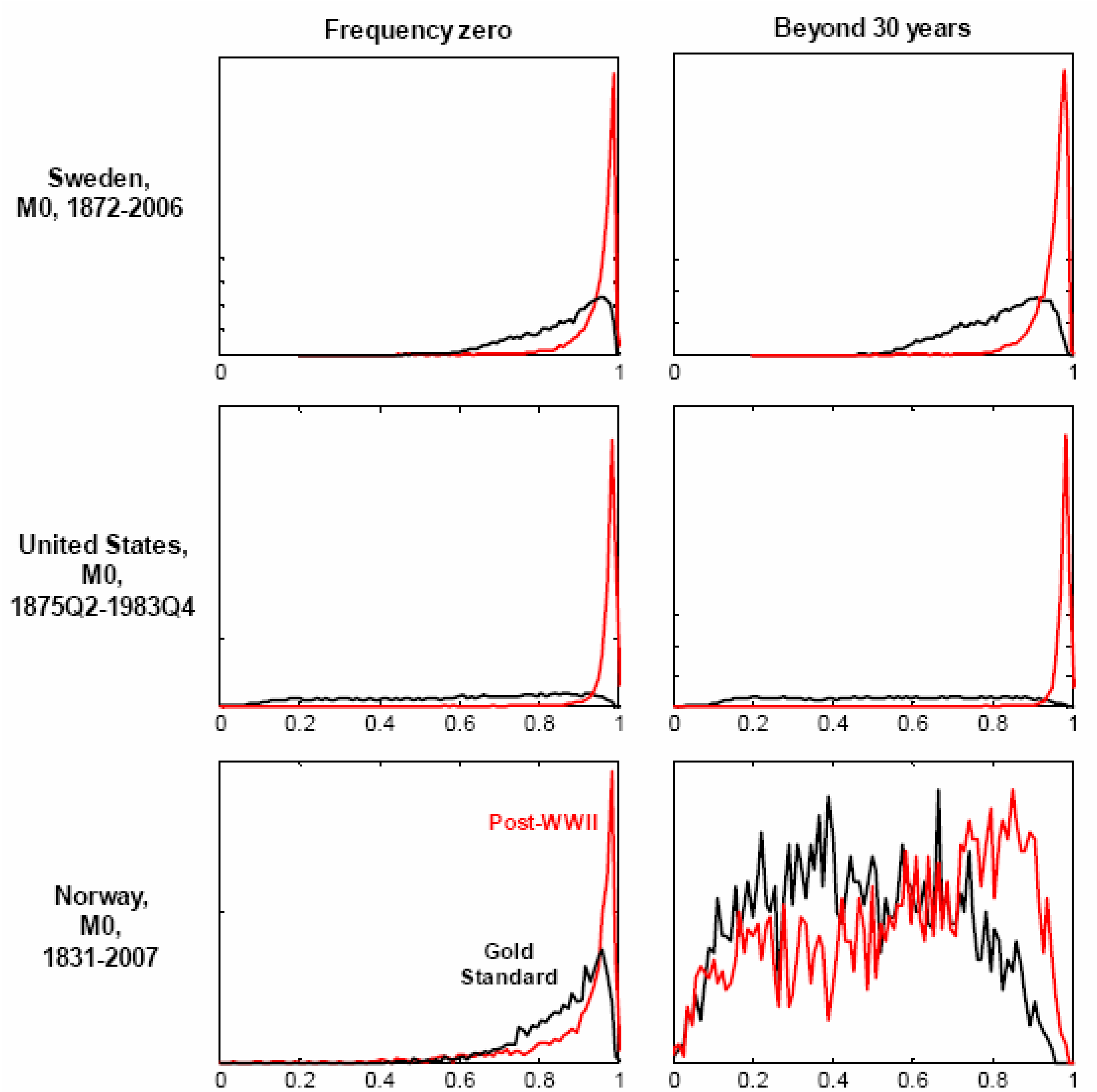

Figure 5 (continued) Comparing the Gold Standard and the post-WWII period: bootstrapped distributions of the cross-spectral coherence between broad money growth and inflation, frequency zero, and average within the frequency band beyond 30 years

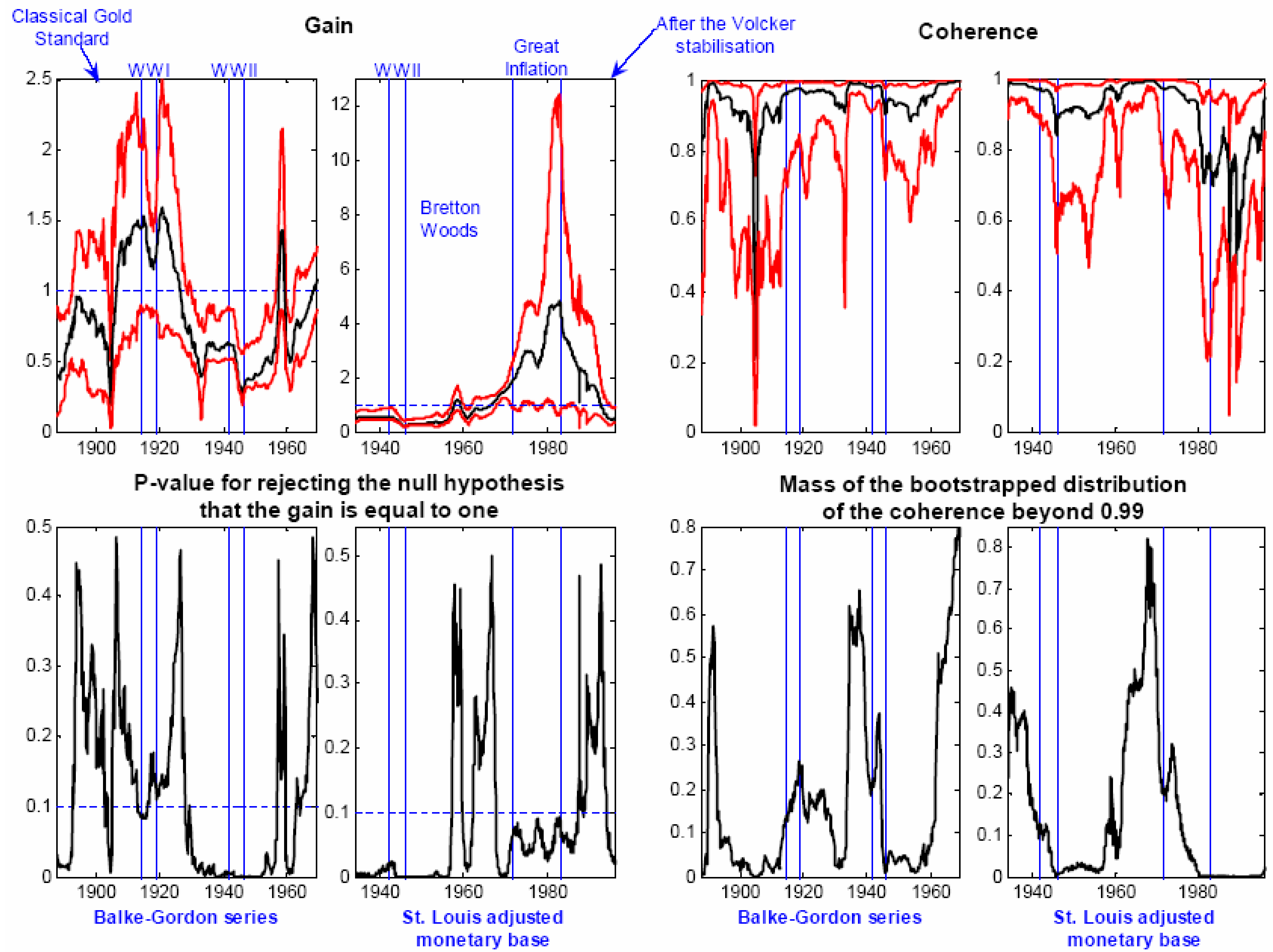

Figure 6 United States, rolling estimates of the cross-spectral gain and coherence between M0 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

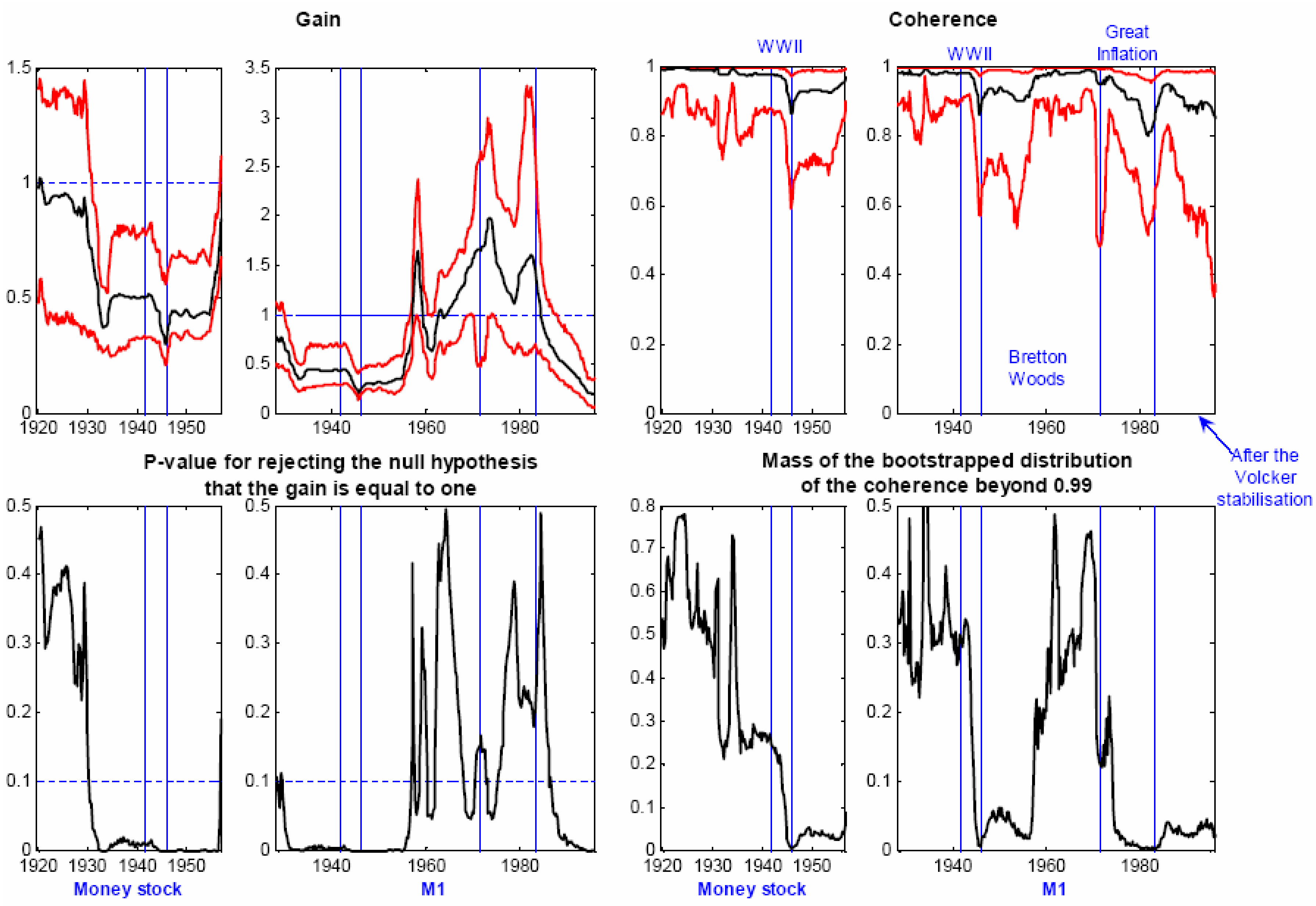

Figure 7 United States, rolling estimates of the cross-spectral gain and coherence between inflation and the rates of growth of M1, and the 'money stock' monetary aggregate at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

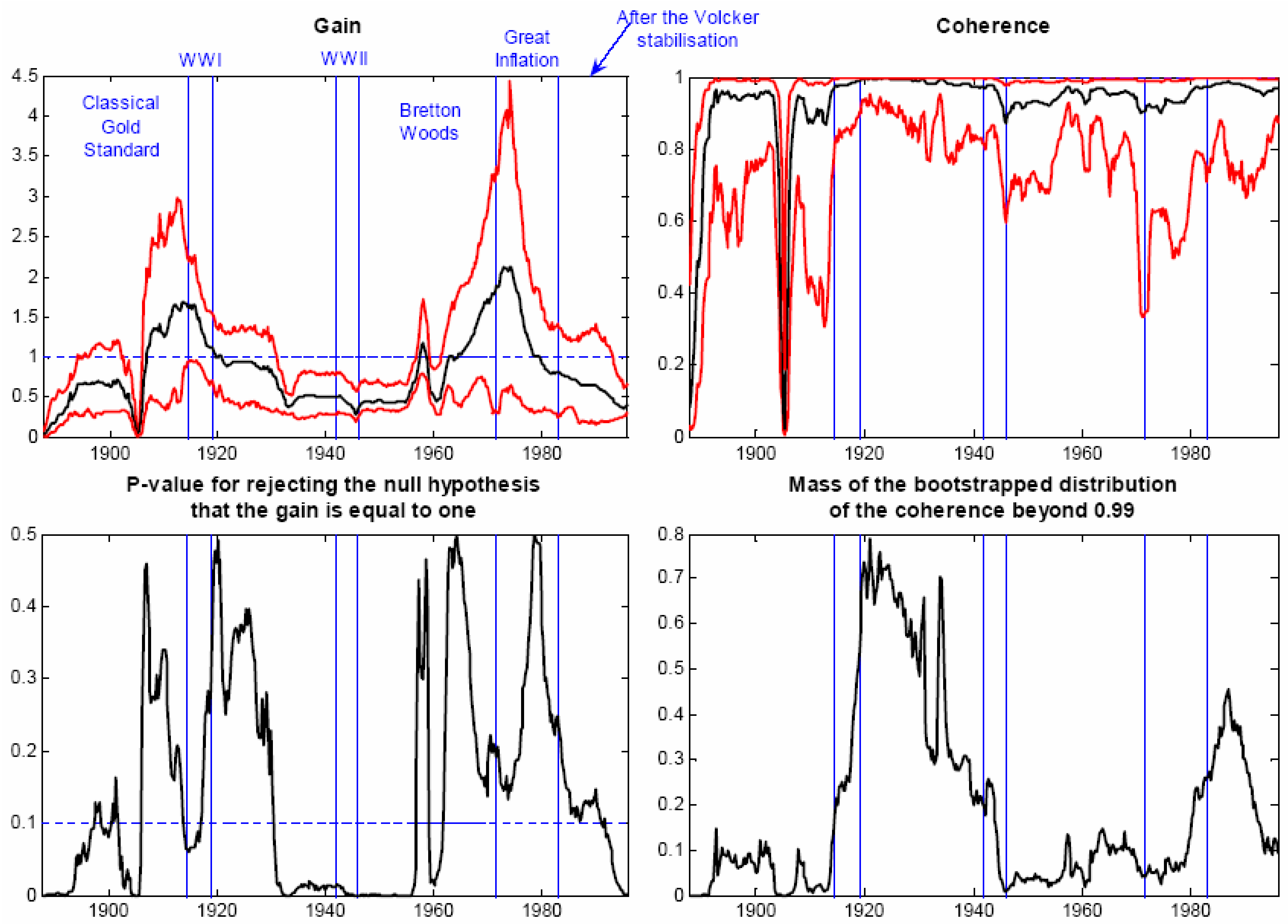

Figure 8 United States, rolling estimates of the cross-spectral gain and coherence between M2 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

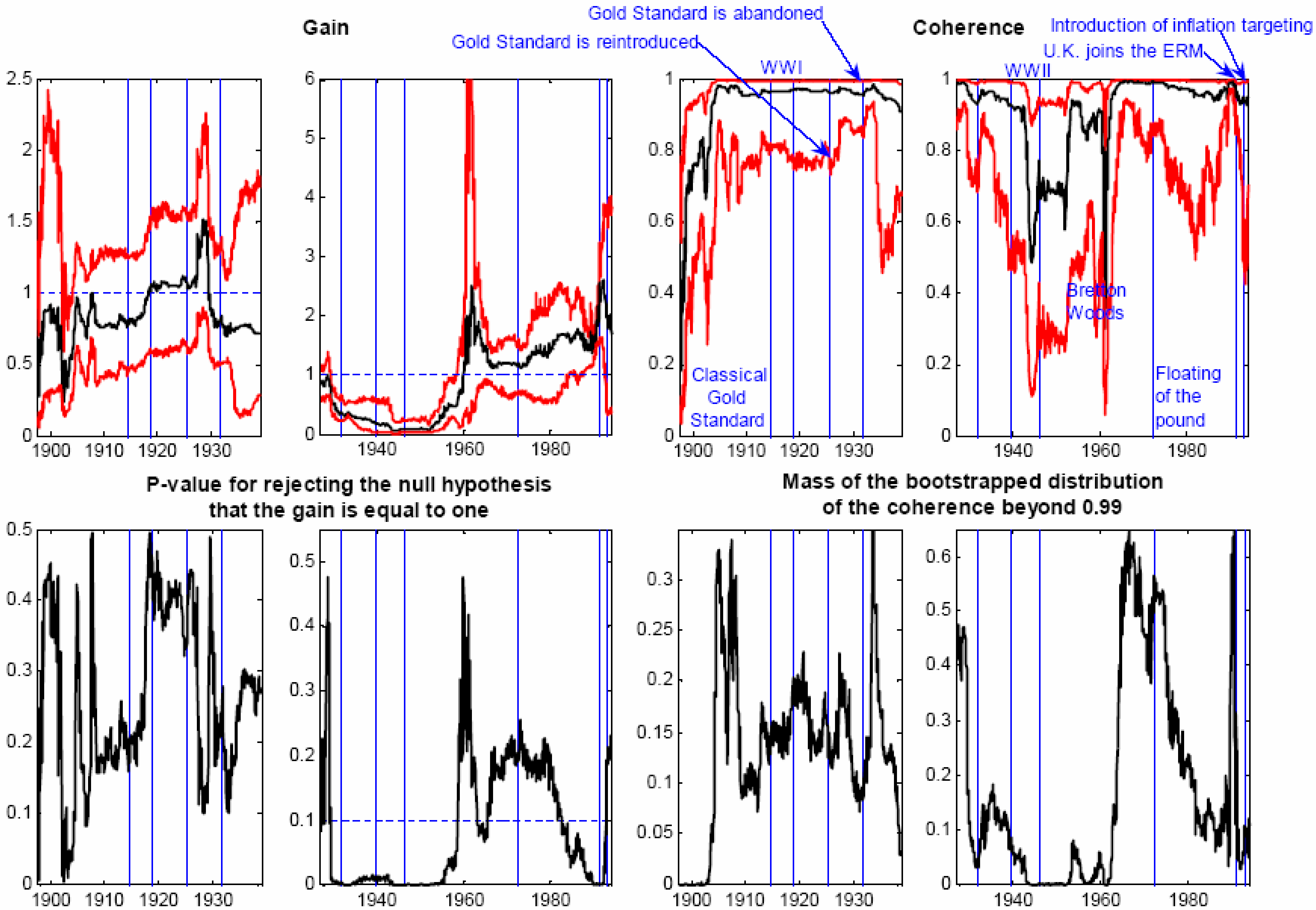

Figure 9 United Kingdom, rolling estimates of the cross-spectral gain and coherence between M0 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

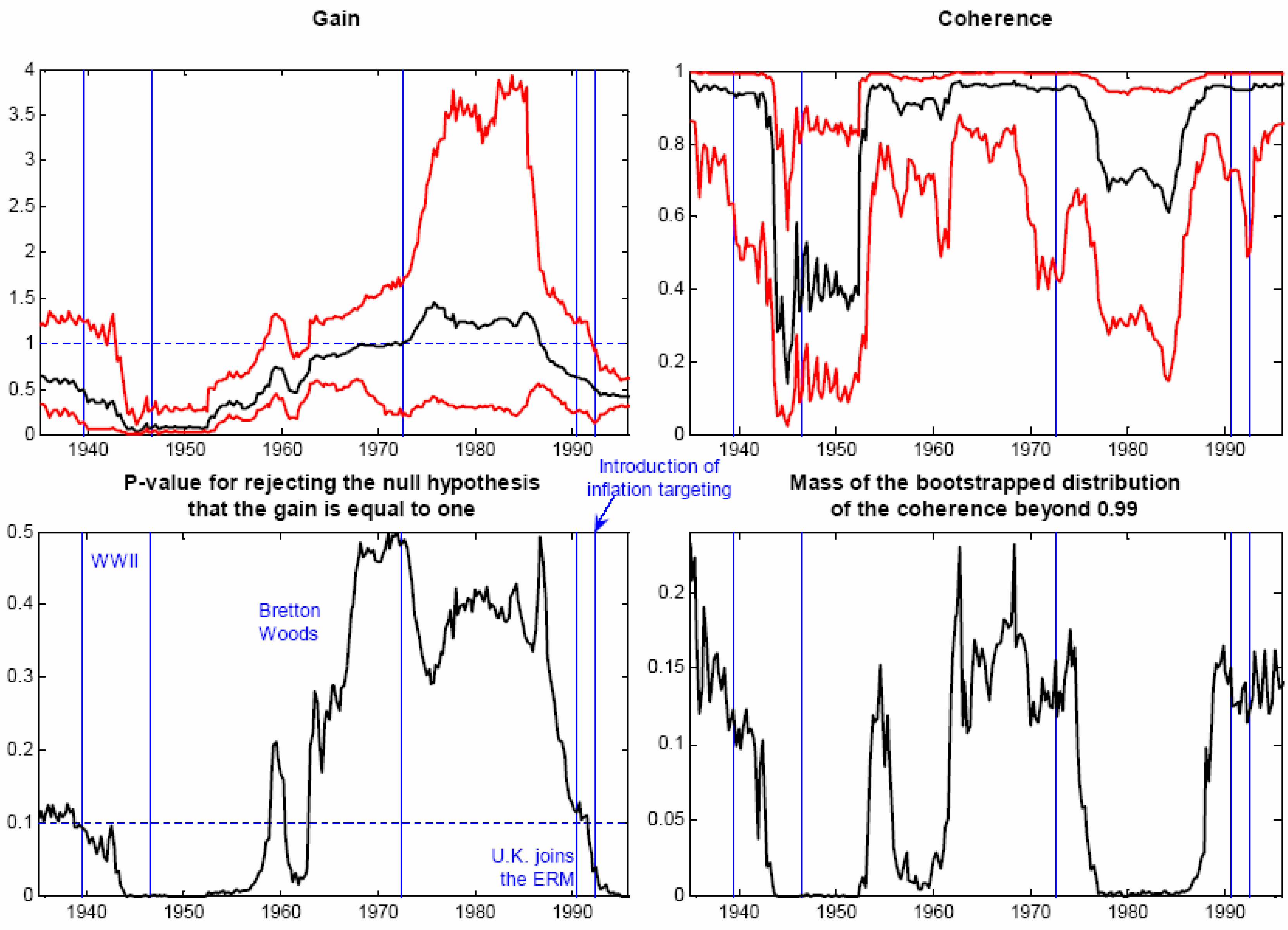

Figure 10 United Kingdom, rolling estimates of the cross-spectral gain and coherence between M3 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

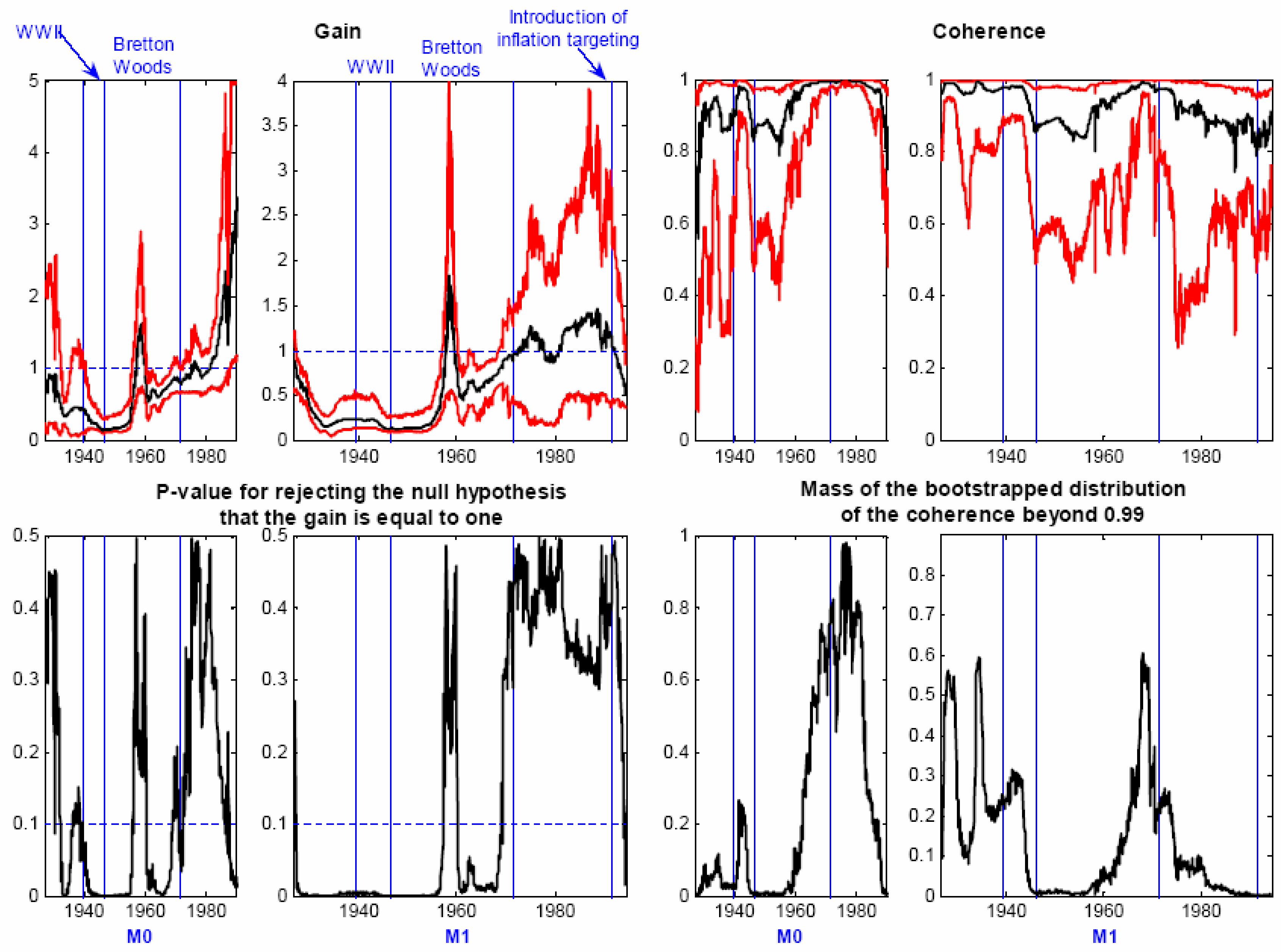

Figure 11 Canada, rolling estimates of the cross-spectral gain and coherence between M0 and M1 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

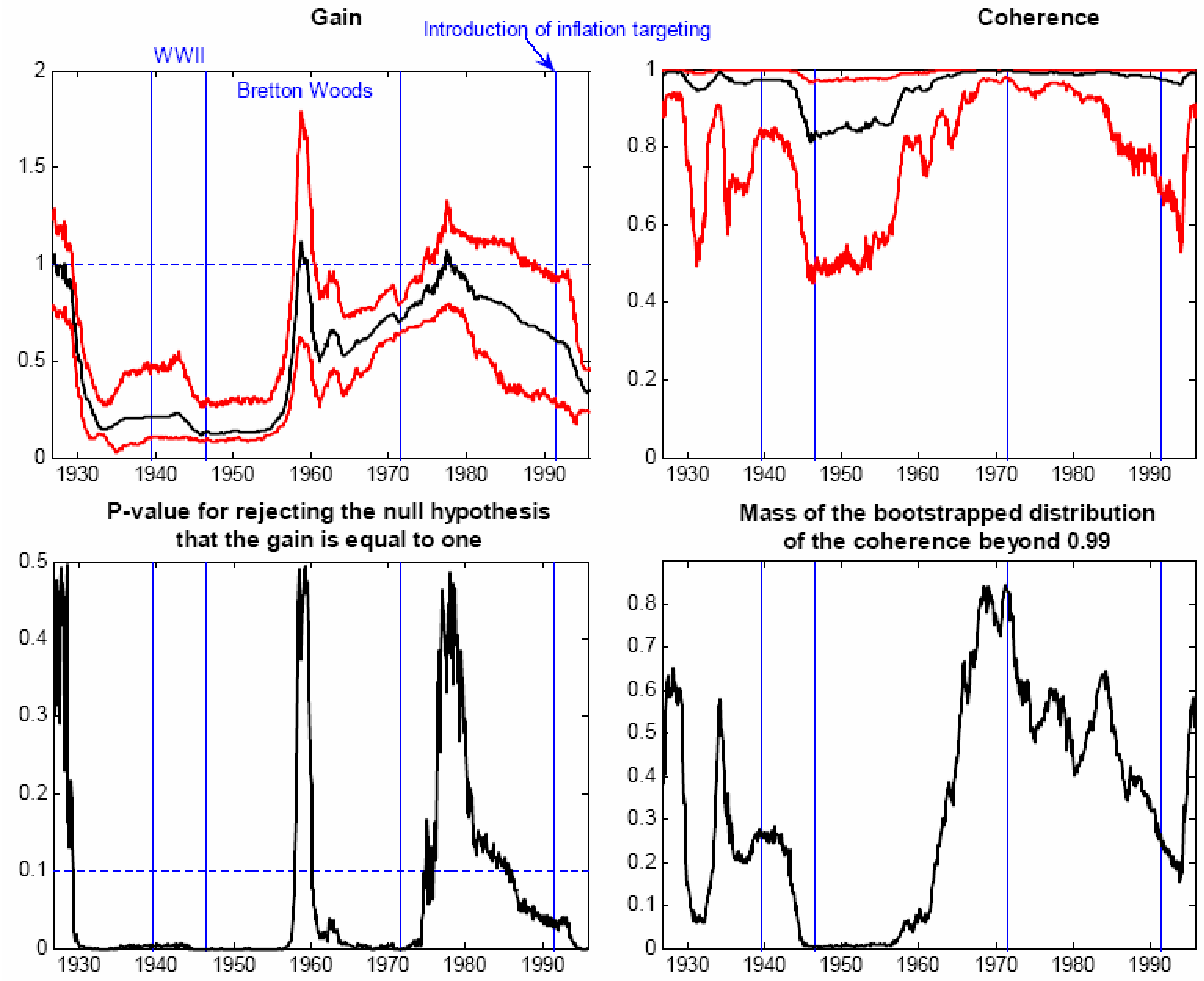

Figure 12 Canada, rolling estimates of the cross-spectral gain and coherence between M2 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

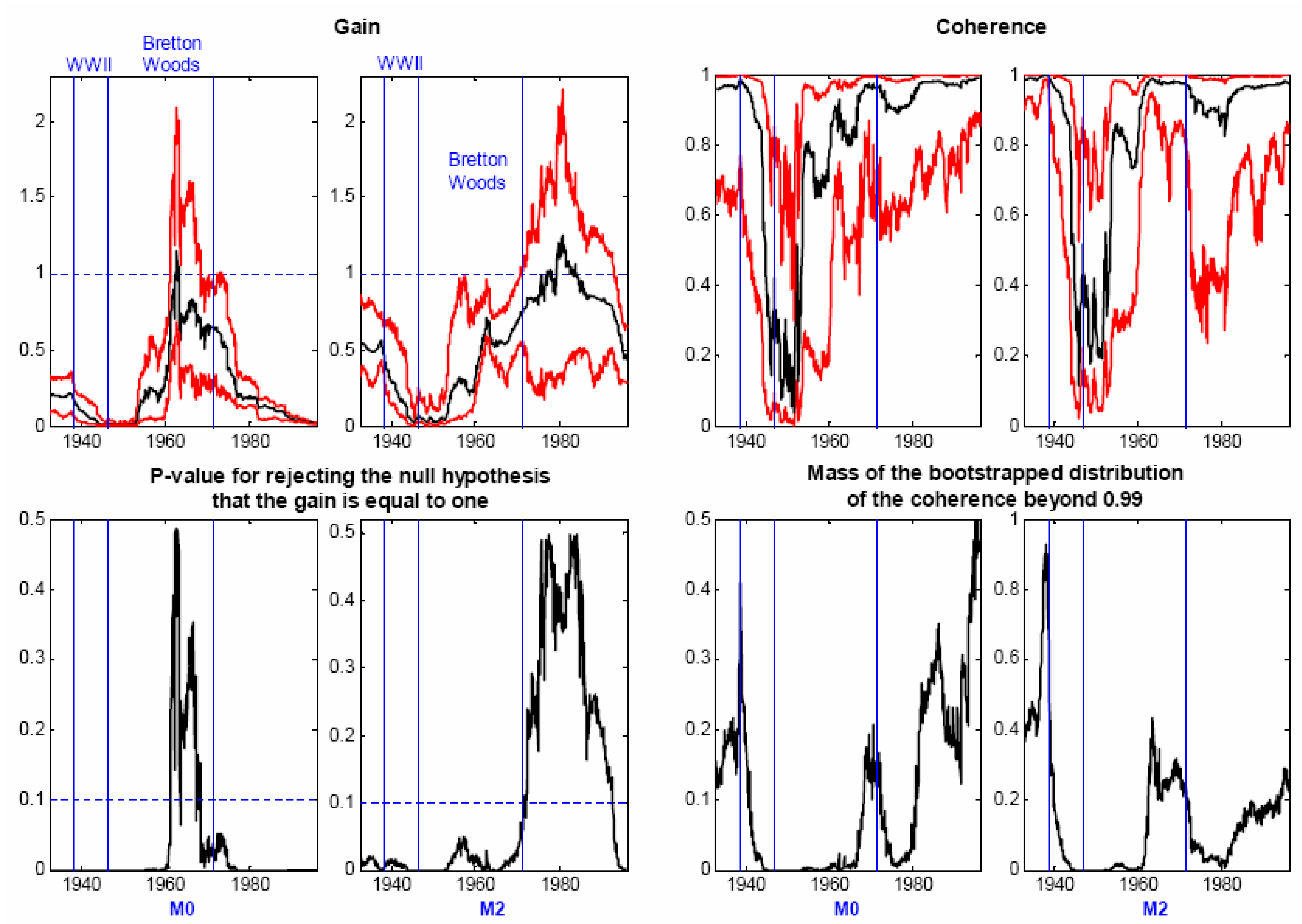

Figure 13 Norway, rolling estimates of the cross-spectral gain and coherence between M0 and M2 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

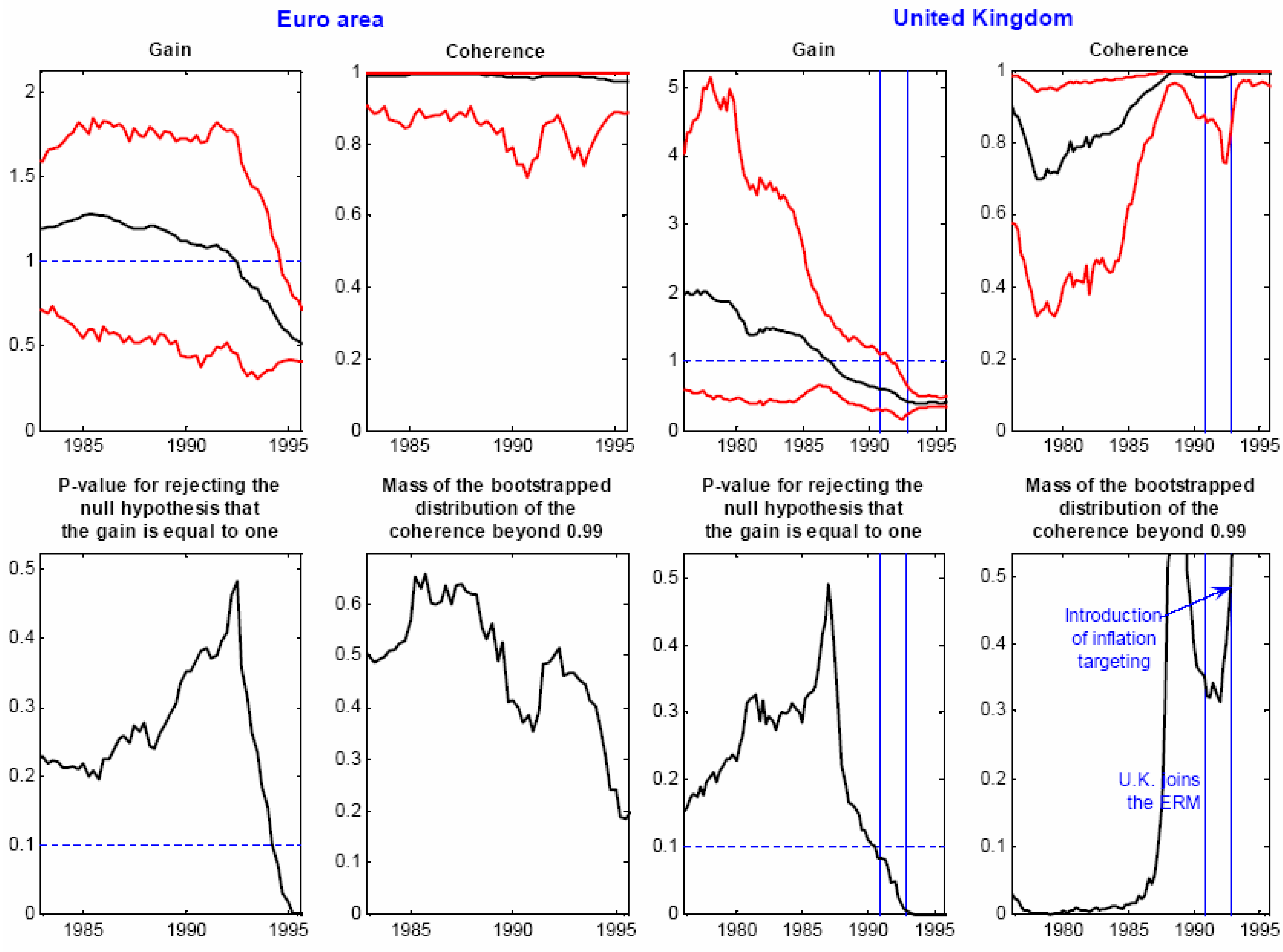

Figure 14 Euro area and United Kingdom, rolling estimates of the cross-spectral gain and coherence between M3 and M4 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

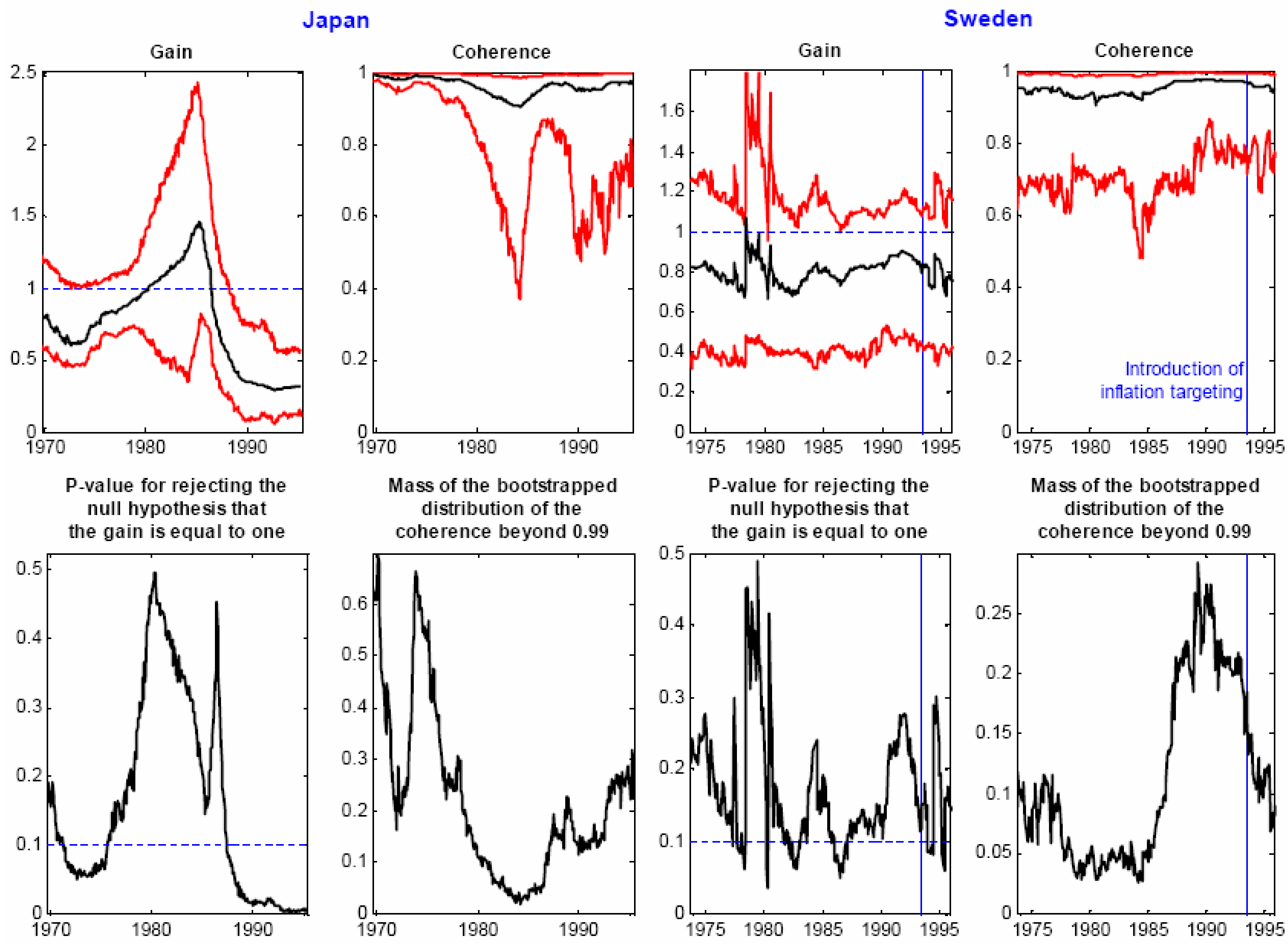

Figure 15 Japan and Sweden, rolling estimates of the cross-spectral gain and coherence between M2 and M3 growth and inflation at the frequency zero, median of the bootstrapped distribution and 90%-coverage percentiles

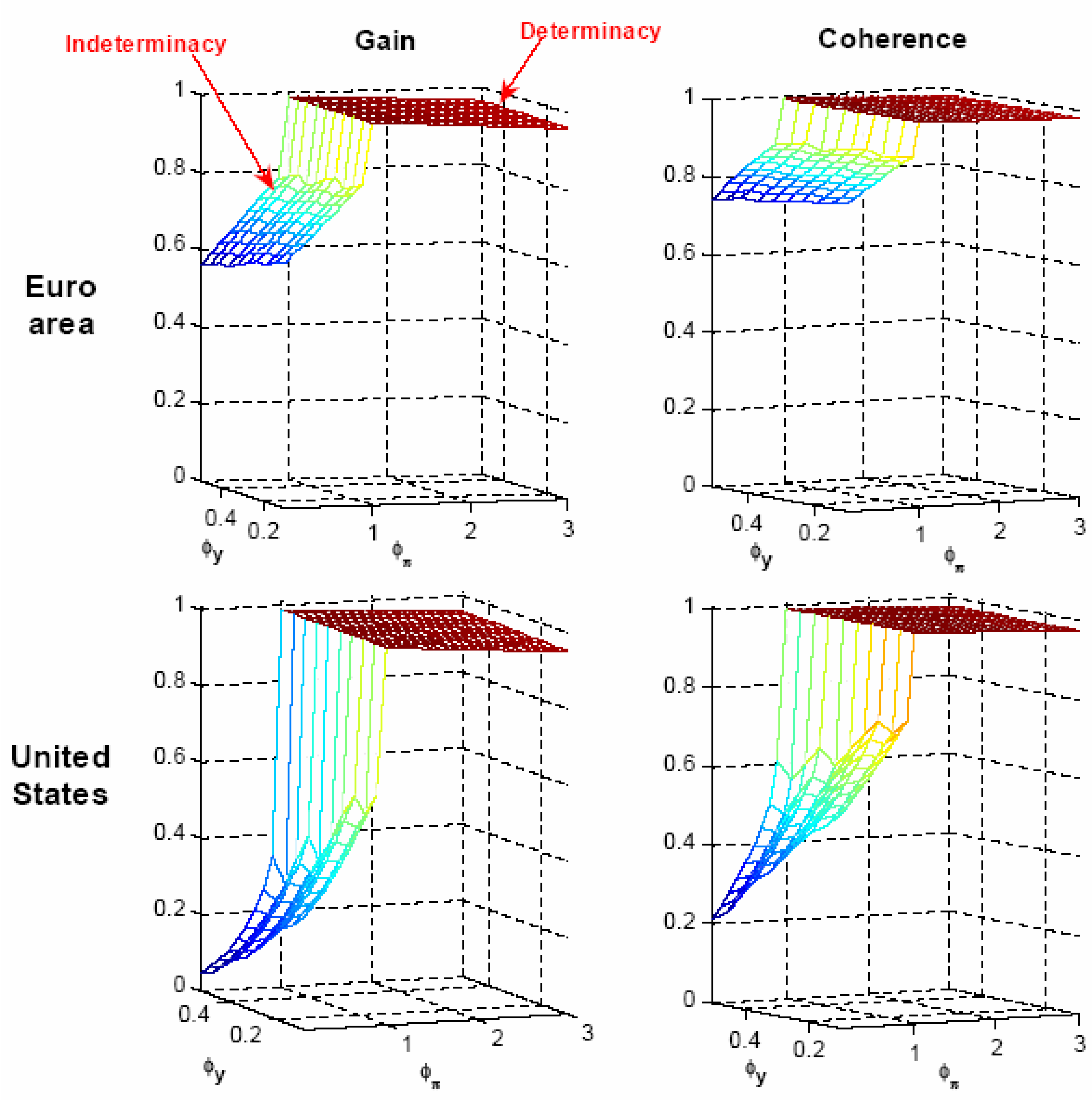

Figure 16 New Keynesian model, Taylor rule: cross-spectral gain and coherence of money growth onto inflation at the frequency zero with velocity growth white noise (based on Benati's (2008) estimates for the Euro area for the full sample)

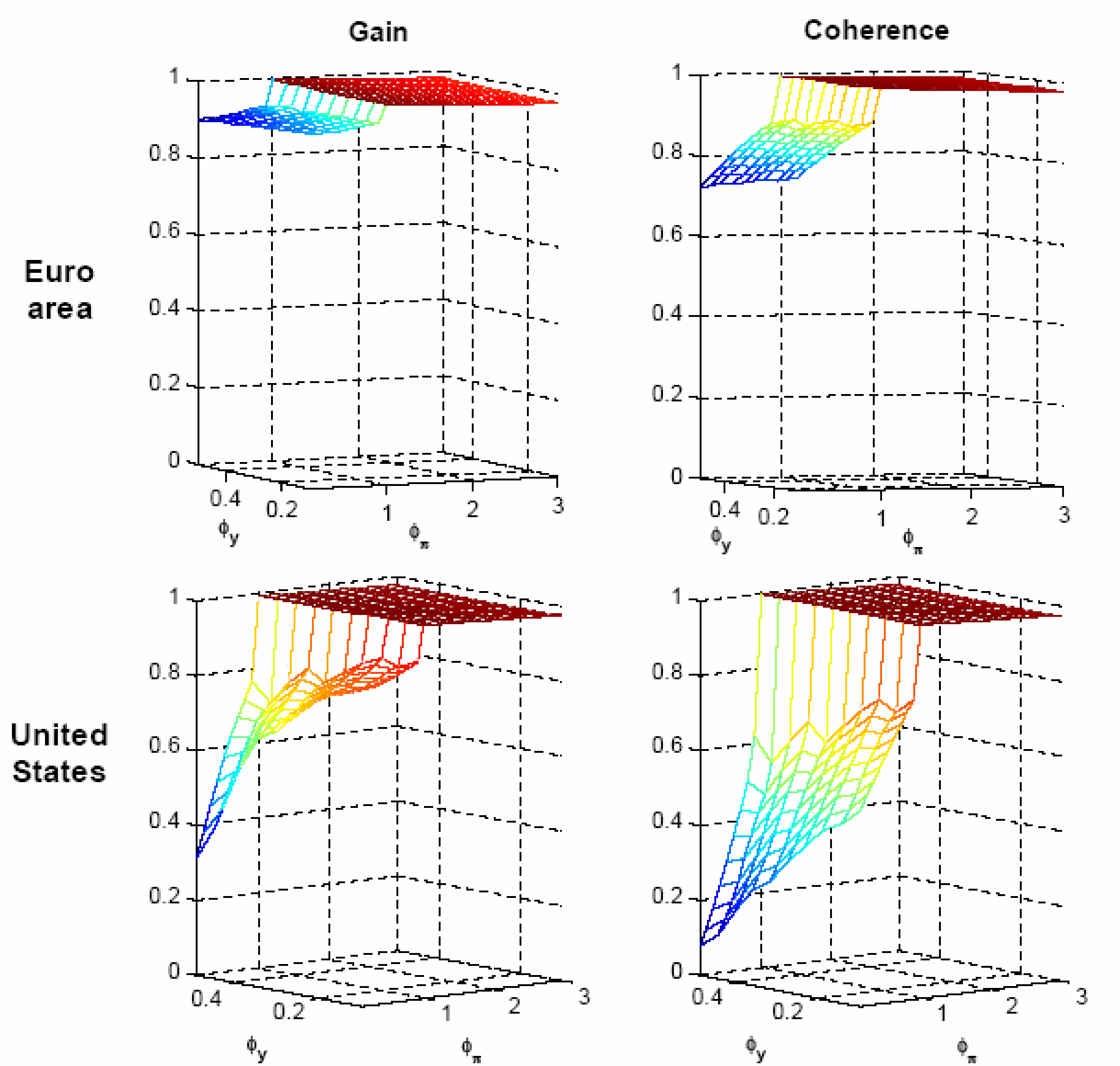

Figure 17 New Keynesian model, first money growth rule: cross-spectral gain and coherence of money growth onto inflation at the frequency zero with velocity growth white noise (based on Benati's (2008) estimates for the Euro area for the full sample)

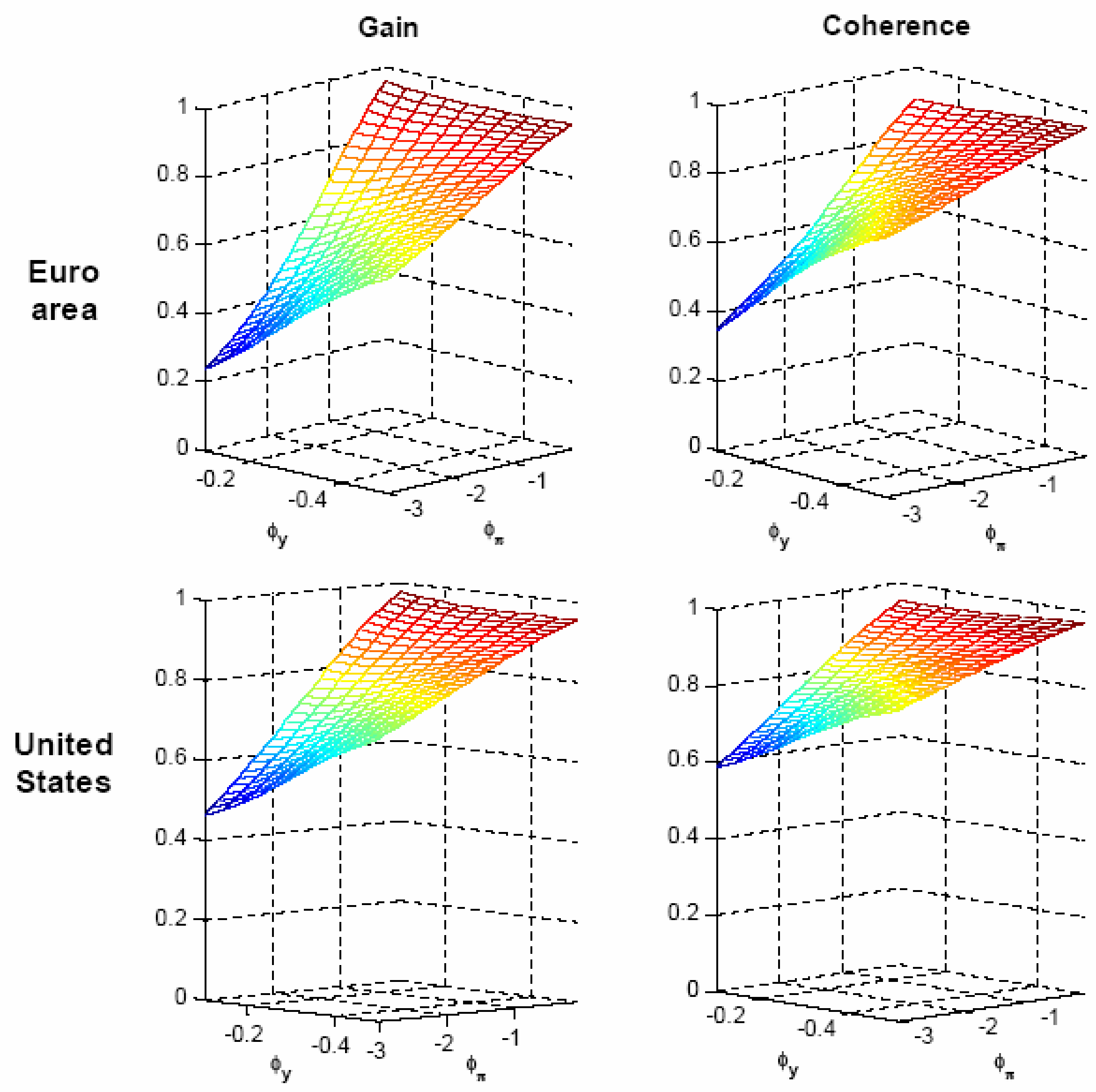

Figure 18 New Keynesian model, second money growth rule: cross-spectral gain and coherence of money growth onto inflation at the frequency zero with velocity growth white noise (based on Benati's (2008) estimates for the Euro area for the full sample)

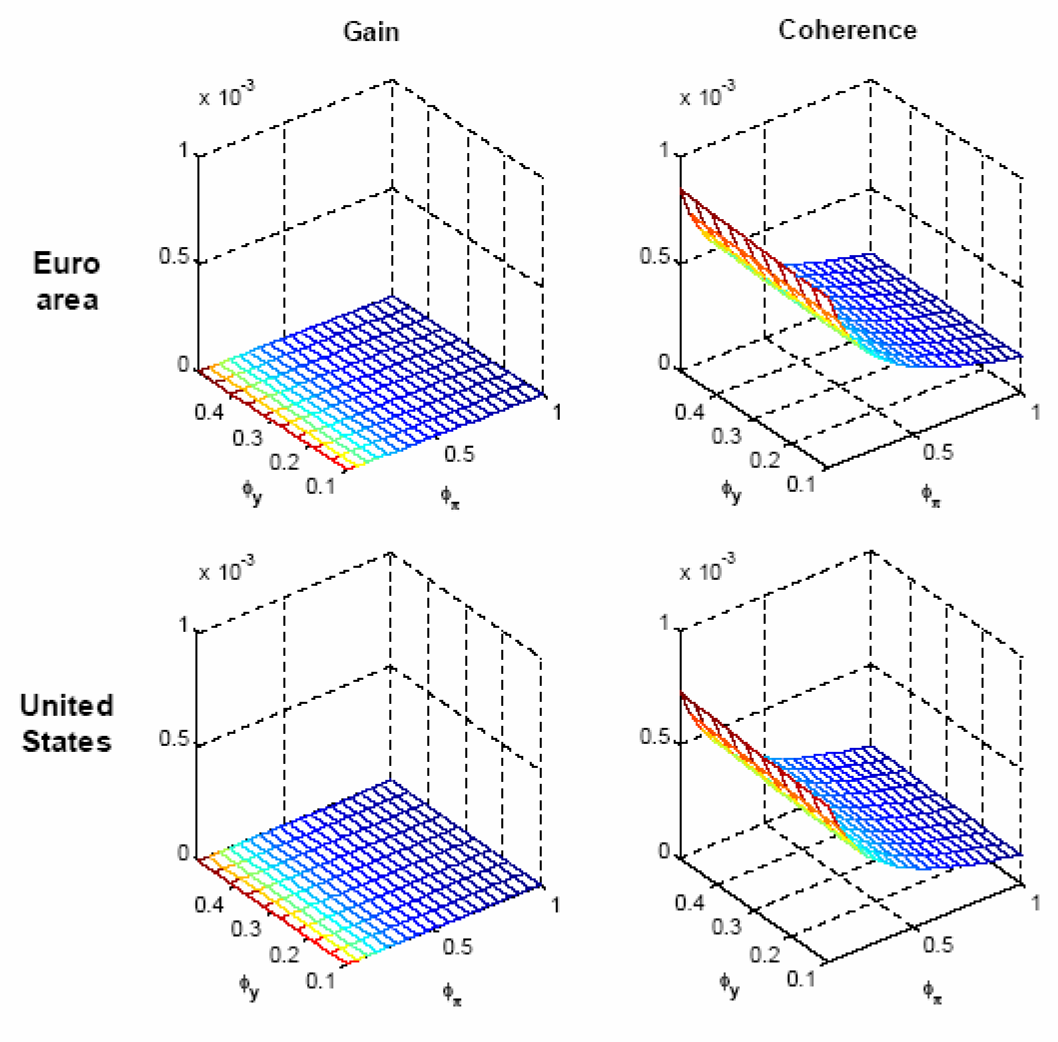

Figure 19 New Keynesian model, price level targeting rule: cross-spectral gain and coherence of money growth onto inflation at the frequency zero with velocity growth white noise (based on Benati's (2008) estimates for the Euro area for the full sample)

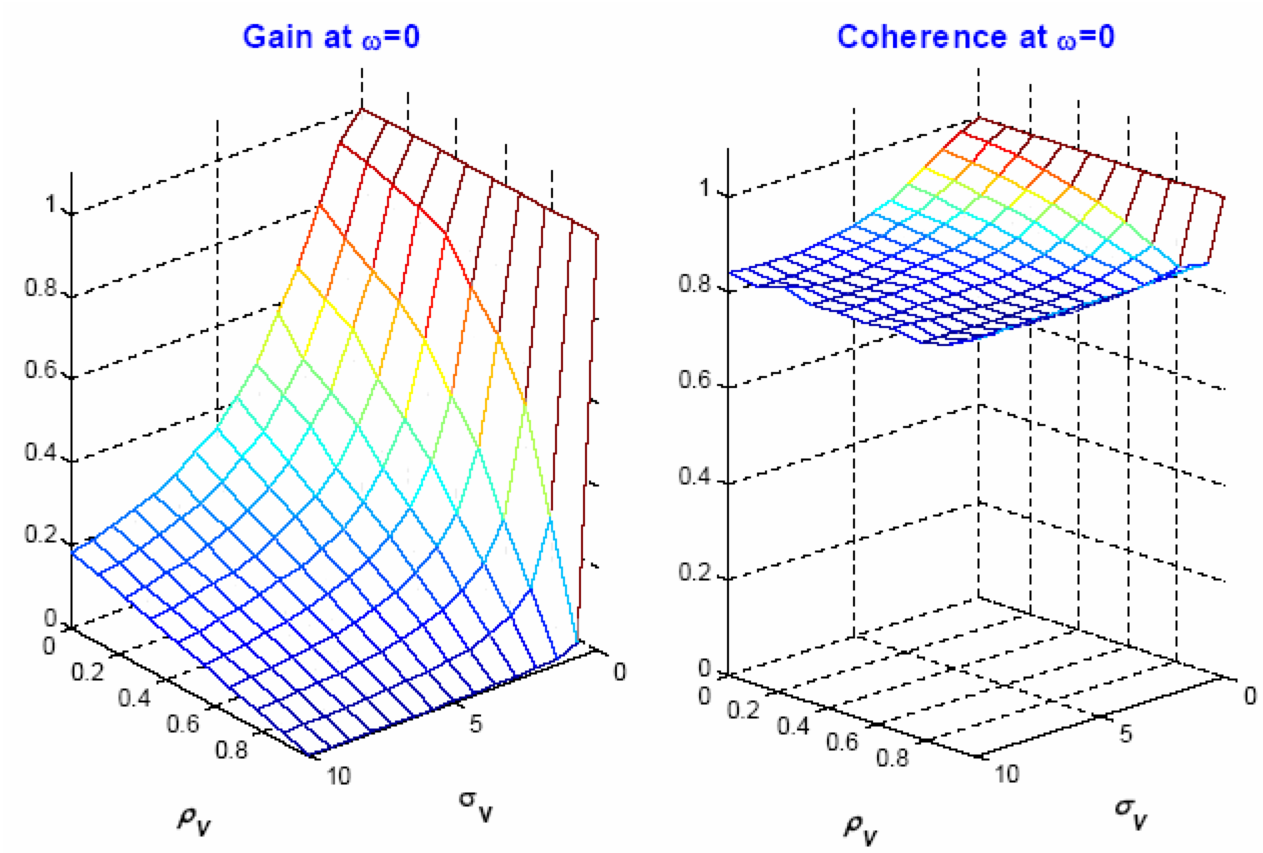

Figure 20 The impact of velocity shifts on the gain and the coherence between money growth and inflation at ω=0 (Taylor rule, based on the estimated model of Ascari and Ropele (2007) for the Euro area)

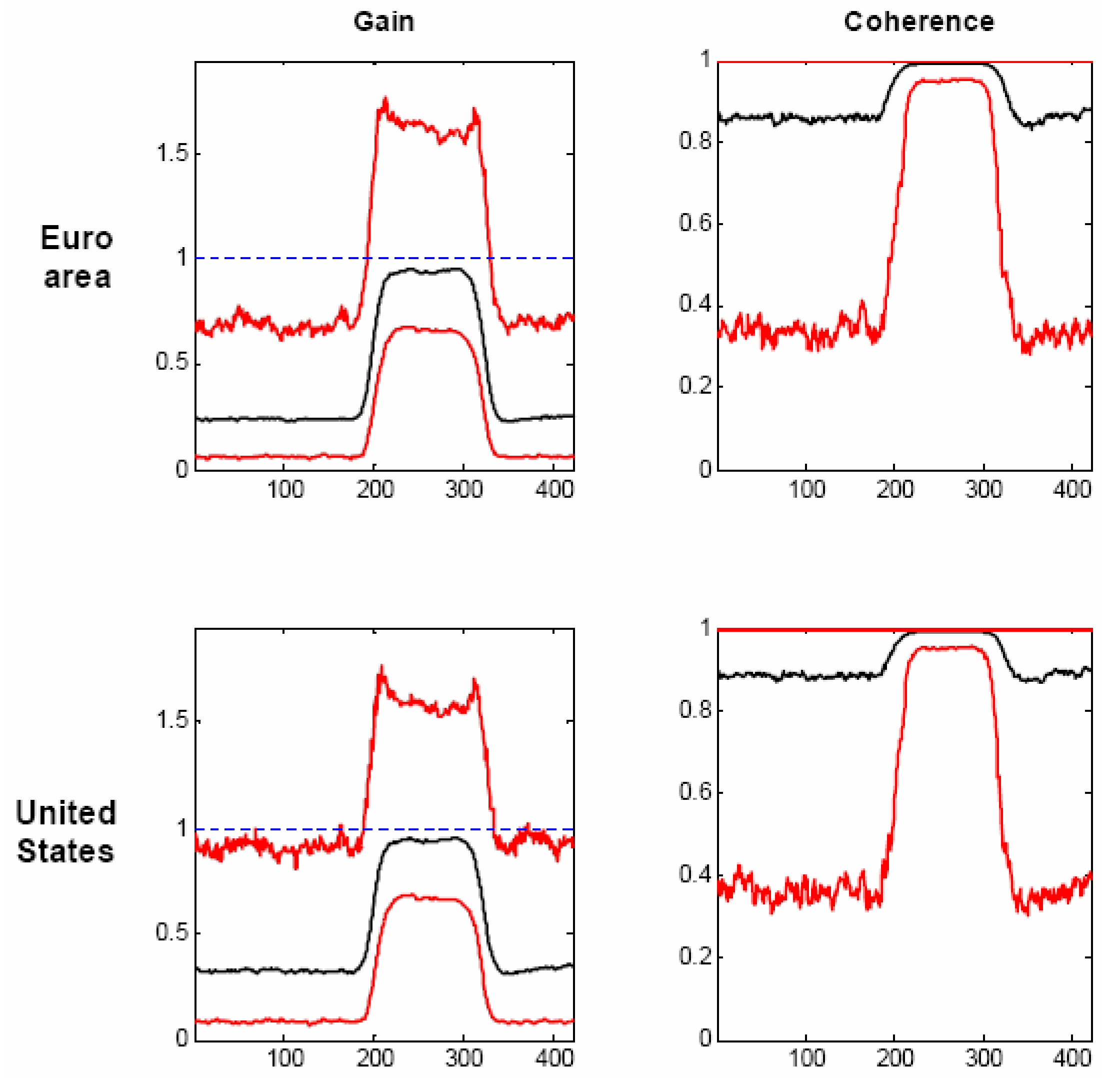

Figure 21 The impact of infrequent inflationary upsurges in the presence of velocity shifts: medians and 90% lower and upper percentiles of the distributions of the rolling estimates of the gain and coherence between money growth and inflation at ω=0 (Taylor rule, based on the estimated model of Ascari and Ropele (2007))

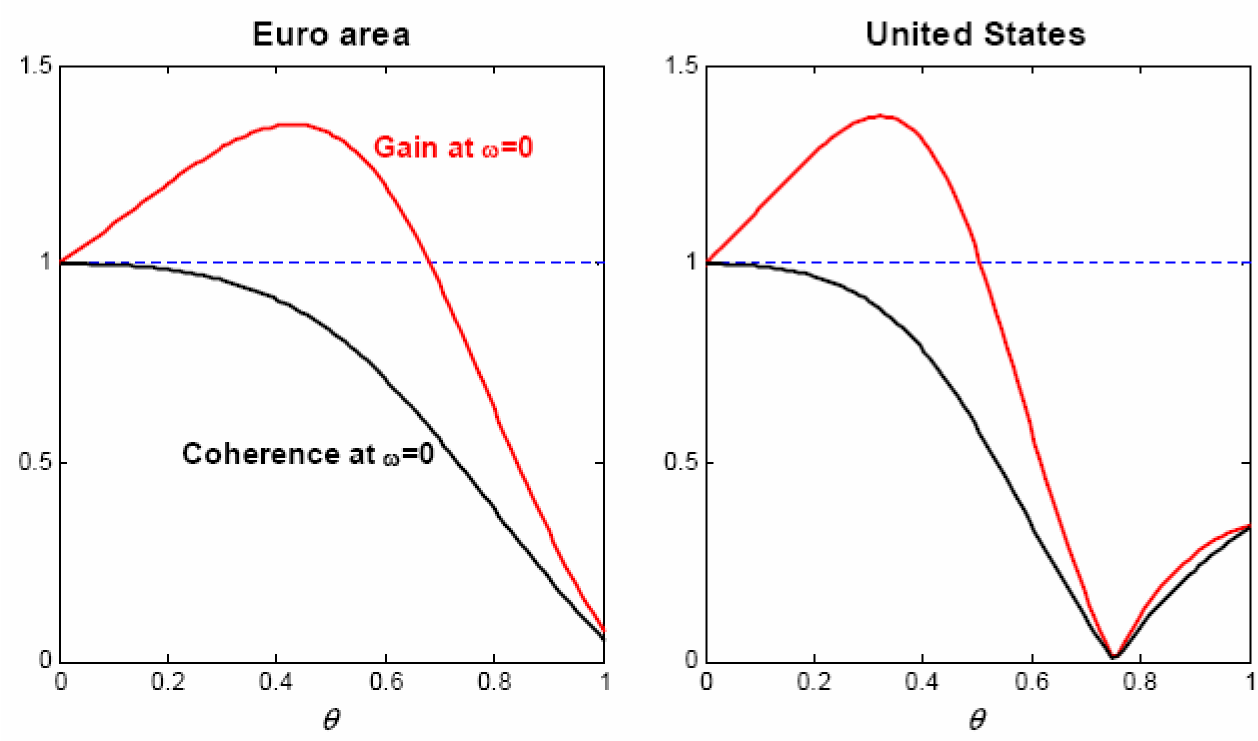

Figure 22 The impact of endogenous shifts in velocity growth on the gain and the coherence between money growth and inflation at the frequency zero (Taylor rule, calibration based on Benati's (2008) estimates for the Euro area and the United States for the full sample)

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