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1 The views expressed in this paper are those of the author and do not necessarily reflect those of the European Central Bank.

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Abstract

Policy counterfactuals based on estimated structural VARs routinely suggest that bringing Alan Greenspan back in the 1970s' United States would not have prevented the Great Inflation. We show that a standard policy counterfactual suggests that the Bundesbank–which is near-universally credited for sparing West Germany the Great Inflation–would also not have been able to prevent the Great Inflation in the United States. The sheer implausibility of this result sounds a cautionary note on taking the outcome of SVAR-based policy counterfactuals at face value, and raises questions on the very reliability of such exercises.

Keywords: Bayesian VARs; time-varying parameters; stochastic volatility; identified VARs; Great Inflation; policy counterfactuals.

JEL Classification: E32, E47, E52, E58

Non Technical Summary

A standard result produced by structural VAR-based studies of the U.S. Great Moderation is that imposing over the entire post-WWII sample the structural monetary policy rule associated with the more recent, and more stable, period—in the literature jargon, 'bringing Alan Greenspan back in time'—would not have prevented the Great Inflation, and, more generally, would only have exerted a limited impact on U.S. post-WWII macroeconomic dynamics. Because of the comparatively wide range of VAR specifications and identification schemes conditional on which it has been produced, the result that 'bringing Alan Greenspan back in time' would only have exerted a limited impact on U.S. post-WWII macroeconomic dynamics is regarded as a very robust outcome of the structural VAR methodology, and is routinely taken to imply that sheer 'luck' has played a dominant role in shaping U.S. post-WWII macroeconomic fluctuations.

The *Bundesbank* is near-universally credited, within both academia and central banking, for preventing the Great Inflation in West Germany. As a consequence, we would logically expect that policy counterfactuals based on estimated structural VARs for the United States and West Germany would suggest that—in the very same way as the *Bundesbank* was able to successfully counter the 1970s' inflationary impulses for West Germany—it would have been able to prevent the Great Inflation if it had been put in charge of U.S. monetary policy.

As this paper shows, this is not the case: imposing the structural monetary policy rule estimated for West Germany's *Bundesbank* in the post-WWII United States (*i*) would only have exerted a limited impact on overall macroeconomic dynamics, and (*ii*) crucially, it would *not* have prevented the Great Inflation. The results produced by this counterfactual are therefore qualitatively the same as those obtained by 'bringing Alan Greenspan back in time'. The key difference is that, whereas in the case of Alan Greenspan (or, more generally, of FED officials who have been in charge of U.S. monetary policy over the most recent years) we have no way of knowing how they would have performed had they been in charge of U.S. monetary policy in the 1970s, this is obviously not the case for the *Bundesbank*. West Germany's central bank *was* indeed there, and its monetary policy is widely credited for sparing West Germany the Great Inflation. The notion that, if it had been put in charge of post-WWII U.S. monetary policy, it would have been unable to successfully counter the 1970s' inflationary upsurge in the United States is therefore *extremely* hard to believe. As a logical corollary, this result raises serious questions on the very reliability of policy counterfactuals based on estimated structural VARs, a reliability which, it is important to stress, has always been assumed, rather than demonstrated.

The start of inflation occurred under the Bretton Woods system of fixed exchange rates. [...] Once the fixed exchange rate system ended, Japan, Germany, Switzerland, and Austria reduced their inflation rates. Others permitted inflation to continue or increase. [...] The start of the Great Inflation—the sustained increase in the price level—was a monetary event. Monetary policy could have mitigated or prevented the inflation but failed to do so.

—Allan Meltzer[1]

In the 1970s and 80s there were few central banks whose policy responses to inflation provided a sufficient tightening of policy in the face of inflation to anchor public beliefs around low and stable inflation. [...] [A]n exception to the general picture was the Bundesbank which kept stable and positive real interest rates over this period with the result that German inflation remained low and stable even though it was subject to the same international cost shocks as the other countries [...].

—Timothy Besley[2]

[D]ue to the vigorous action by the Bundesbank, Germany experienced much lower inflation rates than did the United States. In fact, after its peak in 1981, when the inflation rate stood at 6.3 percent, the German inflation rate swiftly declined, reaching values of around 2 percent at the end of 1985 [...].

—Otmar Issing[3]

1 Introduction

A standard result produced by structural VAR-based studies of the U.S. Great Moderation is that imposing over the entire post-WWII sample the structural monetary policy rule associated with the more recent, and more stable, period—in the literature jargon, 'bringing Alan Greenspan back in time'—would not have prevented the Great Inflation, and, more generally, would only have exerted a limited impact on U.S. post-WWII macroeconomic dynamics. This result has been obtained based on either Markov-switching[4] or time-varying parameters VARs,[5] and based on several alternative identification schemes—specifically, Cholesky, as in Primiceri (2005); sign restrictions, as in the work of Fabio Canova and his co-authors; and based on the alternative identification scheme of Sims and Zha (2006). Because of the comparatively wide range of VAR specifications and identification schemes conditional on which it has been produced, the result that 'bringing Alan Greenspan back in time' would only have exerted a limited impact on U.S. post-WWII macroeconomic dynamics is

[1] See Meltzer (2005).
[2] See Besley (2008).
[3] See Issing (2005).
[4] See e.g. Sims and Zha (2006).
[5] See e.g. Primiceri (2005), Gambetti, Pappa, and Canova (2006), and Canova and Gambetti (2008).

regarded as a very robust outcome of the structural VAR methodology, and is routinely taken to imply that sheer 'luck' (that is, shocks), has played a dominant role in shaping U.S. post-WWII macroeconomic fluctuations.

With a few exceptions,[6] such result is typically *not questioned*.[7] The reason for this is quite obvious, although never explicitly mentioned in the literature: during the 1970s, Alan Greenspan was not Chairman of the FED,[8] and as a result there is simply *no way* of knowing whether, facing those very same shocks, he would have been able to spare the U.S. economy the Great Inflation.[9] There is however at least one important exception to this logic:[10] the *Bundesbank* is near-universally credited, within both academia and central banking, for preventing the Great Inflation in West Germany. As a consequence, we would logically expect that policy counterfactuals based on estimated structural VARs for the United States and West Germany would suggest that—in the very same way as the *Bundesbank* was able to successfully counter the 1970s' inflationary impulses for West Germany—it would have been able to save the day if it had been put in charge of U.S. monetary policy. Quite stunningly, as this paper shows, *this is not the case*: imposing the structural monetary policy rule estimated for West Germany's *Bundesbank* in the post-WWII United States (*i*) would only have exerted a limited impact on overall macroeconomic dynamics, and (*ii*) crucially, it would *not* have prevented the Great Inflation. The results produced by this counterfactual are therefore qualitatively the same as those obtained by 'bringing Alan Greenspan back in time'. The key difference is that, whereas—as we previously discussed—in the case of Alan Greenspan (or, more generally, of FED officials who have been in charge of U.S. monetary policy over the most recent years) we have no way of knowing how they would have performed had they been in charge of U.S. monetary policy in the 1970s, this is obviously not the case for the *Bundesbank*. West Germany's central bank *was* indeed there, and its monetary policy is widely credited

[6] See in particular DeLong (2003) and Bernanke (2004). In a paper conceptually related to the present work, Benati and Surico (2009) produce a simple example in the spirit of Clarida, Gali, and Gertler (2000) and Lubik and Schorfheide (2004) in which (*i*) a shift in the systematic component of monetary policy such as to move the economy from the indeterminacy to the determinacy region is sufficient, by itself, to replicate the key qualitative features of the transition from the Great Inflation to the Great Moderation; and (*ii*) structural VAR methods, when applied to this data-generation process, fail to point towards a change in policy as the cause of the changes in the reduced-form properties of the economy, and—in line with the results produced by structural VAR-based studies of the Great Moderation—point instead towards a fall in the volatilities of the structural shocks.

[7] More generally, doubts about the reliability of SVAR-based policy counterfactuals are seldom expressed. An exception is represented by Christiano (1998)'s discussion of Sims (1998).

[8] Indeed, during those years he was either working in the private sector, as a macroeconomic forecaster, or working for the U.S. Government in a number of jobs (e.g., as Chairman of the Council of Economic Advisers) which, however, were all outside of the Federal Reserve System.

[9] After all, this is *precisely* the reason why the counterfactual is performed in the first place ...

[10] We say 'at least' because, as stressed (e.g.) by Meltzer's initial quotation, West Germany was not the only country to escape the 1970s largely unscathed: Japan, Austria, and especially Switzerland, too, were equally successful under this respect. Quite obviously, the very same logic underlying the present work could equally be applied to those three countries.

for sparing West Germany the Great Inflation. The notion that, if it had been put in charge of post-WWII U.S. monetary policy, it would have been unable to successfully counter the 1970s' inflationary upsurge in the United States is therefore extremely hard to believe. As a logical corollary, this result raises serious questions on the very reliability of policy counterfactuals based on estimated structural VARs, a reliability which, it is important to stress, has always been *assumed*, rather than demonstrated.

The paper is organised as follows. The next section discusses the reduced-form specification for the time-varying parameters VAR with stochastic volatility we will use throughout the paper, and the identification strategy, which is based on sign restrictions, whereas (standard) technical aspects of the Bayesian inference—in particular, our choices for the priors, and the Markov chain Monte Carlo algorithm we use to simulate the posterior distribution of the hyperparameters and the states conditional on the data—are relegated to an appendix. Section 3 presents results from the policy counterfactual in which we 'bring the Bundesbank to the post-WWII United States', by imposing over the entire U.S. post-WWII sample period the structural monetary rule estimated for West Germany's central bank. Section 4 draws some implications of these results for macroeconomics. In particular, we argue that the sheer implausibility of the outcome produced by this counterfactual—an outcome which, it is important to stress, has been obtained based on standard methodology—raises questions on the very reliability of SVAR-based policy counterfactuals. Section 5 concludes, and outlines directions for future research.

2 Methodology

2.1 A Bayesian time-varying parameters VAR with stochastic volatility

In what follows we work with the following time-varying parameters VAR(p) model:

$$Y_t = B_{0,t} + B_{1,t}Y_{t-1} + ... + B_{p,t}Y_{t-p} + \epsilon_t \equiv X_t^{'}\theta_t + \epsilon_t \tag{1}$$

where the notation is obvious, and Y_t (which is an $N \times 1$ vector) is defined as either $Y_t \equiv [r_t, \pi_t, y_t, m_t]'$, or $Y_t \equiv [r_t, \pi_t, y_t, m_t, neer_t]'$, with r_t, π_t, y_t, m_t, and $neer_t$ being a short-term interest rate (specifically, the Federal Funds rate for the United States, and a call money rate for West Germany), GDP deflator inflation, and the rates of change of real GDP, nominal M2, and the nominal effective exchange rate (henceforth, NEER), respectively (for a description of the data, see Appendix A).[11] The overall sample periods are 1959:2-2008:1 for the United States, and 1960:2-1990:25 for West

[11] GDP deflator inflation and the rates of growth of real GDP, nominal M2, and the NEER have been computed as the annualised quarter-on-quarter rates of growth of the relevant series.

Germany.[12] For reasons of comparability with other papers in the literature[13] we set the lag order to p=2. Following, e.g., Cogley and Sargent (2002), Cogley and Sargent (2005), Primiceri (2005), and Gambetti, Pappa, and Canova (2006) the VAR's time-varying parameters, collected in the vector θ_t, are postulated to evolve according to

$$p(\theta_t \mid \theta_{t-1}, Q) = I(\theta_t)\ f(\theta_t \mid \theta_{t-1}, Q) \tag{2}$$

with $I(\theta_t)$ being an indicator function rejecting unstable draws—thus enforcing a stationarity constraint on the VAR—and with $f(\theta_t \mid \theta_{t-1}, Q)$ given by

$$\theta_t = \theta_{t-1} + \eta_t \tag{3}$$

with $\eta_t \sim N(0,\ Q)$. The VAR's reduced-form innovations in (1) are postulated to be zero-mean normally distributed, with time-varying covariance matrix Ω_t which, following established practice, we factor as

$$\mathrm{Var}(\epsilon_t) \equiv \Omega_t = A_t^{-1} H_t (A_t^{-1})' \tag{4}$$

The time-varying matrices H_t and A_t are defined as:

$$H_t \equiv \begin{bmatrix} h_{1,t} & 0 & ... & 0 \\ 0 & h_{2,t} & ... & 0 \\ ... & ... & ... & ... \\ 0 & 0 & ... & h_{N,t} \end{bmatrix} \qquad A_t \equiv \begin{bmatrix} 1 & 0 & ... & 0 \\ \alpha_{2,1,t} & 1 & ... & 0 \\ ... & ... & ... & ... \\ \alpha_{N,1,t} & \alpha_{N,2,t} & ... & 1 \end{bmatrix} \tag{5}$$

with the $h_{i,t}$ evolving as geometric random walks,

$$\ln h_{i,t} = \ln h_{i,t-1} + \nu_{i,t} \tag{6}$$

For future reference, we define $h_t \equiv [h_{1,t},\ h_{2,t},\ ...\ ,h_{N,t}]'$. Following Primiceri (2005), we postulate the non-zero and non-one elements of the matrix A_t—which we collect in the vector $\alpha_t \equiv [\alpha_{2,1,t},\ \alpha_{3,1,t},\ ...,\ \alpha_{N,N-1,t}]'$—to evolve as driftless random walks,

$$\alpha_t = \alpha_{t-1} + \tau_t\ , \tag{7}$$

and we assume the vector $[u_t',\ \eta_t',\ \tau_t',\ \nu_t']'$ to be distributed as

$$\begin{bmatrix} u_t \\ \eta_t \\ \tau_t \\ \nu_t \end{bmatrix} \sim N\left(0, V\right), \text{ with } V = \begin{bmatrix} I_4 & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & Z \end{bmatrix} \text{ and } Z = \begin{bmatrix} \sigma_1^2 & 0 & ... & 0 \\ 0 & \sigma_2^2 & ... & 0 \\ ... & ... & ... & ... \\ 0 & 0 & ... & \sigma_N^2 \end{bmatrix} \tag{8}$$

[12]The first 8 years of data are however used to calibrate the Bayesian priors, based on a time-invariant version of the same VAR used in estimation.

[13]See e.g. Primiceri (2005), Gambetti, Pappa, and Canova (2006), and Canova and Gambetti (2008).

where u_t is such that $\epsilon_t \equiv A_t^{-1} H_t^{\frac{1}{2}} u_t$.[14] Finally, following, again, Primiceri (2005) we adopt the additional simplifying assumption of postulating a block-diagonal structure for S, too—namely

$$S \equiv \mathrm{Var}\left(\tau_t\right) = \begin{bmatrix} S_1 & 0_{1\times 2} & \ldots & 0_{1\times(N-1)} \\ 0_{2\times 1} & S_2 & \ldots & 0_{2\times(N-1)} \\ \ldots & \ldots & \ldots & \ldots \\ 0_{(N-1)\times 1} & 0_{(N-1)\times 2} & \ldots & S_{N-1} \end{bmatrix} \tag{9}$$

with $S_1 \equiv \mathrm{Var}(\tau_{21,t})$, $S_2 \equiv \mathrm{Var}([\tau_{31,t}, \tau_{32,t}]')$, ..., and $S_{N-1} \equiv \mathrm{Var}([\tau_{N,1,t},\ \tau_{N,2,t},\ \ldots,\ \tau_{N,N-1,t}]')$, thus implying that the non-zero and non-one elements of A_t belonging to different rows evolve independently. As discussed in Primiceri (2005, Appendix A.2), this assumption drastically simplifies inference, as it allows to do Gibbs sampling on the non-zero and non-one elements of A_t equation by equation.

We estimate (1)-(9) *via* Bayesian methods. Appendix B discusses our choices for the priors, and the Markov-Chain Monte Carlo algorithm (specifically, Gibbs-sampling) we use to simulate the posterior distribution of the hyperparameters and the states conditional on the data.

2.2 Identification

In either VAR we identify N shocks by imposing sign restrictions[15] on the estimated reduced-form VAR on a period-by-period basis. Specifically, in the smaller VAR we identify four shocks—a monetary policy shock (ϵ_t^M), a supply shock (ϵ_t^S), a demand non-policy shock (ϵ_t^D), and a money demand shock (ϵ_t^{MD})—wheres in the larger one we identify an additional shock (ϵ_t^{NEER}) which can be given several alternative interpretations. For example, it might reflect either a shock to the foreign exchange risk premium, or the impact of a foreign monetary policy shock. The following table summarises the sign restrictions we impose on the estimated VAR. A '+' and a '-' mean 'greater than or equal to zero' and 'smaller than or equal to zero', respectively, whereas a '?' means that the sign of this specific impact has been left unconstrained. Sign restrictions are imposed only on impact (that is, at zero). It can be trivially shown that these sign restrictions are sufficient to identify the shocks. We compute the time-varying structural impact matrix, $A_{0,t}$, *via* the procedure proposed by Rubio-Ramirez, Waggoner, and Zha (2005).[16]

[14] As discussed in Primiceri (2005, pp. 6-7), there are two justifications for assuming a block-diagonal structure for V_t. First, parsimony, as the model is already quite heavily parameterized. Second, 'allowing for a completely generic correlation structure among different sources of uncertainty would preclude any structural interpretation of the innovations'.

[15] Sign restrictions have been used in the studies of the U.S. Great Moderation of Gambetti, Pappa, and Canova (2006) and Canova and Gambetti (2008).

[16] Specifically, let $\Omega_t = P_t D_t P_t'$ be the eigenvalue-eigenvector decomposition of the VAR's time-varying covariance matrix Ω_t, and let $\tilde{A}_{0,t} \equiv P_t D_t^{\frac{1}{2}}$. We draw an $N \times N$ matrix, K, from the N(0,

Variable:	Shock: ϵ_t^M	ϵ_t^D	ϵ_t^S	ϵ_t^{MD}	ϵ_t^{NEER}
short rate	+	+	?	+	−
inflation	−	+	−	−	−
output growth	−	+	+	−	−
money growth	−	+	?	+	?
rate of change of NEER	+	+	+	+	+

We eschew Cholesky, on the other hand—although it has been used in one of the best-known studies of the U.S. Great Moderation[17]—because, as it is well known, DSGE models exhibit a recursive ordering of the variables only under very special circumstances (basically, for a DSGE model to exhibit a recursive structure it has to be specifically 'engineered' for that purpose). To put it differently, within a DSGE context the 'normal state of affairs' is for all structural shocks to have an impact at zero on all variables, so that the imposition of a Cholesky structure for the impact matrix at zero can safely be regarded, in general, as being pretty far away from the truth, *whatever the truth in fact is*. Further, as shown by Canova and Pina (2005), the imposition of a recursive ordering of the variables when such ordering is, in fact, false, can lead to a dramatic distortion of the inference.[18] On the other hand, precisely because, within a DSGE context, the impact at zero of all structural shocks is, in general, non-zero on all variables, sign restrictions appear as the most natural choice, since they do not impose 'incredible' zero restrictions on impact.

Finally, another possibility—which we however leave to future research—would be to consider the identification scheme used by Sims and Zha (2006). Although such scheme is almost entirely based on zero restrictions on impact,[19] their use of *monthly*, as opposed to quarterly, data, makes the imposition of such restrictions much more plausible, notwithstanding what we just said about the impacts of all structural shocks at zero on all variables as being in general non-zero within a DSGE context. For example, it appears as entirely reasonable to assume that variables capturing real economic activity—in the case of Sims and Zha (2006), the unemployment rate and monthly interpolated real GDP—do not react to monetary shocks within the month, whereas financial variables instead do.

1) distribution, we take the QR decomposition of K—that is, we compute matrices Q and R such that K=$Q \cdot R$—and we compute the time-varying structural impact matrix as $A_{0,t}$=$\tilde{A}_{0,t} \cdot Q'$. If the draw satisfies the restrictions we keep it, otherwise we discard it and we keep drawing until the restrictions are satisfied, as in the Rubio-Waggoner-Zha code `SRestrictRWZalg.m` which implements their algorithm. (See at http://home.earthlink.net/~tzha02/ProgramCode/SRestrictRWZalg.m.)

[17] See Primiceri (2005).

[18] Canova and Pina (2005) stochastically simulate standard DSGE models and apply Cholesky to the artificial data. They show that, *first*, in general, impulse-response functions are dramatically mis-estimated; and *second*, in several instances the use of Cholesky gives rise to an estimated 'price puzzle' which was not in the original data-generation process.

[19] See their Table I, and the discussion in Section IV.

3 Would the *Bundesbank* Have Prevented the Great Inflation in the United States?

Figures 1-4 show the results from the counterfactual simulations[20] in which we 'bring the Bundesbank to the United States', by imposing the structural monetary policy rule estimated for West Germany's central bank for the period 1973:2-1990:1 in the estimated structural VAR for the U.S. economy over the entire sample period.[21] Specifically, Figures 1 and 2 show results based on the four-variables VAR, whereas Figures 3 and 4 show those based on the larger VAR, which also includes the rate of change of the NEER. Figures 1 and 3 show, for all the series used in the VAR, the actual historical values together with the medians and the 16th and the 84th percentiles of the distributions of the simulated counterfactuals series, whereas Figures 2 and 4 show the medians and the 16th and the 84th percentiles of the distributions of the difference between the actual and the counterfactual series. Since a comparison between Figures 1 and 3, and 2 and 4, respectively, shows that results are qualitatively the same for the two alternative VAR specifications, in what follows we will exclusively focus on those produced by the four-variables VAR.

Focusing in particular on inflation, the upper right panel of Figure 2 shows how the difference between the actual and the counterfactual inflation series during the Great Inflation episode is, overall, quite modest, with median estimates oscillating between slightly below -2 per cent and slightly above 2 per cent. The overall negligible impact of putting the *Bundesbank* in charge of post-WWII U.S. monetary policy—when seen through the lenses of the SVAR-based counterfactual—is even more apparent from the upper right panel of Figure 1, in which the median of the counterfactual simulation closely tracks actual inflation most of the times. During the 1970s the Federal Funds rate would have been, on average, lower than the actual one—this is especially clear during the second half of the decade—whereas starting from the beginning of the

[20] For each simulation j=1, 2, ..., 1,000, at each quarter t=p+1, p+2, ..., T we draw three random numbers, τ, indexing the quarter between 1973:2 and 1990:1 (included) from which we draw the elements of the *Bundesbank*'s structural monetary rule; and κ_t and κ_τ, indexing the iterations of the Gibbs sampler at times t and, respectively, τ from which we draw the state of the economy. (All three numbers are defined over appropriate uniform distributions.) We then take all of the elements of the monetary rule from iteration κ_τ of the Gibbs sampler for quarter τ of the estimated SVAR for West Germany, while we take everything else from iteration κ_t for quarter t of the estimated SVAR for the United States. We start each counterfactual simulation conditional on the first p actual historical values of the vector Y_t for the U.S.. Finally, we convert the quarter-on-quarter rates of growth of the GDP deflator, real GDP, M2, and the NEER into annual rates of growth by simply computing the convolutions of the quarter-on-quarter rates of growth at time t and in the previous three quarters. Specifically, letting x_t^A and x_t^Q be the annual and quarterly rates of growth of variable X in quarter t, we have x_t^A=$(1+x_t^Q)\cdot(1+x_{t-1}^Q)\cdot(1+x_{t-2}^Q)\cdot(1+x_{t-3}^Q)$-1.

[21] The reason for focusing on the period following 1973:1 is that in March 1973 West Germany abandoned the parity with respect to the U.S. dollar, so that the Bundesbank was free to fully pursue a counter-inflationary policy without the impedements coming from such external constraint (on this, see Issing, 2005, p. 329).

following decade (with the exception of the first half of 1980) becomes smaller. Finally, for real GDP growth and M2 growth, too, the 1970s are not characterised by any systematic difference between actual and counterfactuals series.

4 Implications for Macroeconomics

Given that, within the present work, we are not considering *any* plausible identification scheme,[22] we obviously cannot claim that failure, on the part of the *Bundesbank*, to prevent the Great Inflation in the United States is a *robust* implication of policy counterfactuals based on estimated structural VARs. So it is entirely possible that, had we performed the counterfactual based on an alternative identification strategy—for example, Sims and Zha's (2006)—we might have obtained the alternative result that the *Bundesbank* could have prevented the U.S. Great Inflation.

The key point of this paper, however, is *not* to claim that policy counterfactuals based on structural VARs are *in general* unreliable. Rather, it is to show, by example, that *standard* structural VAR methodology can produce results which the vast majority of macroeconomists would regard as extremely hard to believe. The notion that the *Bundesbank*—which burnished its reputation as a hard-nosed, hard-money central bank by preventing the Great Inflation in West Germany—would have been unable to deliver an analogous performance had it been put in charge of U.S. monetary policy is indeed one such, highly implausible result.[23]

So, two things ought to be stressed here. *First*, the reliability of policy counterfactuals based on estimated structural VARs has never been demonstrated in any way, and it has rather always been *assumed*. As this paper has shown by means of a simple example, however, such assumption appears to be, in general, unwarranted, as such counterfactuals can indeed produce 'incredible' results.[24] *Second*, whereas, for a specific methodology to be regarded as reliable, it has to be shown to perform well conditional on a wide range of plausible circumstances, a *single* example of an unsatisfactory performance under relatively 'normal' circumstances is sufficient to raise doubts on its reliability. Under this respect, the counterfactual associated with 'bringing the *Bundesbank* to the United States' ought to be regarded a 'standard' one: in particular, the only difference with the traditional counterfactual of 'bringing Alan Greenspan back in time' is that, instead of being performed within a single country and across time it is performed across countries. So the fact that this counterfactual produces such an implausible outcome sounds a cautionary note on taking

[22] By 'plausible' we mean schemes which, different from Cholesky, can be justified/defended on conceptual grounds, in particular with reference to standard macroeconomic theory.

[23] Unfortunately, 'bringing the *Bundesbank* to the United States' is the only 'test' of the reliability of policy counterfactuals based on structural VARs we could come up with.

[24] As we previously mentioned in footnote 7, Benati and Surico (2009) produce a simple DSGE-based example in which policy counterfactuals based on the theoretical structural VAR(MA) representation of the model dramatically fail.

results from SVAR-based policy counterfactuals at face value, to the point of raising questions on their very reliability.

5 Conclusions, and Directions for Future Research

Since the structural VAR methodology came to essentially dominate applied macroeconomic research, around mid-1980s, policy counterfactuals have been one of its main applications. As we have discussed, the outcome of such counterfactuals is seldom questioned, and the results they produce are usually taken at face value. In this paper we have shown that standard structural VAR methodology, when applied to a specific policy counterfactual—'bringing the *Bundesbank* to the post-WWII United States'—produces a result which the vast majority of macroeconomists would likely find extremely hard to believe: the very same central bank which burnished its 'hard-money', anti-inflation reputation by successfully countering the 1970s' inflationary impulses in West Germany would not have been able to deliver a comparable performance had it been put in charge of U.S. monetary policy. The fact that (i) such counterfactual is a 'standard' one—in the specific sense that, instead of being performed within a single country and across time, it is performed across countries—and (ii) it has been produced based on 'off-the-shelf' methods (in terms of both estimation and identification), sounds a cautionary note on taking the outcome of SVAR-based policy counterfactuals at face value, and raises questions on their very reliability.

Where could the problems come from? More generally, what might go wrong when performing policy counterfactuals based on structural VARs? An issue which is sometimes mentioned—the possible relevance of the Lucas critique for such counterfactuals—is not, in our view, a relevant one. The reason for this is quite simple: the key theme of the Lucas critique is the impact of changes in policy on the *reduced-form* properties of the economy. To the extent that the *structural* VAR correctly captures the *true underlying structure* of the economy—as defined, for example, by a standard New Keynesian model—the Lucas critique should therefore not be, as a simple matter of logic, a problem at all.[25] So the fundamental problem, in our view, is rather the very ability of structural VAR methods to correctly capture the true underlying structure of the economy, or, to just rephrase the same concept, the mapping between the underlying true (DSGE) model and the theoretical structural VAR representation which is implied by that very model. The *implicit* presumption behind SVAR-based policy counterfactuals is that switching the estimated coefficients of the interest rate equations in the structural VAR provides a reasonable approximation to the *authentic* policy counterfactual, i.e. the one the researcher wold obtain if (s)he were to switch the parameters of the monetary policy rule in the underlying DSGE model. Benati and Surico (2009), however, produce a simple example in which this presumption is

[25] To put it differently, structural VARs have been 'sold' as *structural*, so they can't possibly suffer from a problem which plagues *reduced-form* models.

dramatically violated, so that changes in the interest rate equation of the structural VAR bear no clear-cut relationship with changes in the parameters of the monetary policy rule in the underlying DSGE model. So the key issue here, in our view, is the mapping between the underlying true model of the economy and its structural VAR representation, and in particular the ability of counterfactuals based on the latter to correctly capture the true counterfactuals based on the former. Both issues are currently being investigated in our work in progress.

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A The Data

A.1 United States

Quarterly seasonally adjusted series for the GDP deflator ('GDPCTPI: Gross Domestic Product: Chain-type Price Index, Seasonally Adjusted, Quarterly, Index 2000=100') and real GDP ('GDPC96: Real Gross Domestic Product, 3 Decimal, Seasonally Adjusted Annual Rate, Quarterly, Billions of Chained 2000 Dollars') are both from the U.S. *Department of Commerce*, *Bureau of Economic Analysis*, and are both available since 1947:1. A monthly series for the Federal Funds rate ('FEDFUNDS: Effective Federal Funds Rate, Monthly, Percent') from the *Board of Governors of the Federal Reserve System* is available since July 1954, and has been converted to the quarterly frequency by taking averages within the quarter. A monthly seasonally adjusted series for M2 ('M2SL: M2 Money Stock, H.6 Money Stock Measures, Seasonally Adjusted, Monthly, Billions of Dollars'), available since January 1959, is from the from the *Board of Governors of the Federal Reserve System* and has been converted to the quarterly frequency by taking averages within the quarter. Finally, a monthly seasonally adjusted series for the nominal effective exchange rate computed based on unit labor costs (series' code is 111..NEUZF...), available since January 1957, is from the *IMF*'s *IFS*, and it has been converted to the quarterly frequency by taking averages within the quarter. The overall sample period is from 1959:1 to 2008:1.

A.2 West Germany

Quarterly seasonally adjusted series for the GDP deflator and real GDP are both from the *IMF*'s *IFS*, and are both available since 1960:1 (series' codes are 13499BIRZF... and 13499BVRZF... respectively). A monthly seasonally unadjusted series for the call money rate from the *IMF*'s *IFS*, available since January 1957, has been converted to the quarterly frequency by taking averages within the quarter (series' code is 13460B..ZF...). A quarterly seasonally adjusted series for M2, available since 1948:4, is from the *Bundesbank*. Finally, a monthly seasonally adjusted series for the nominal effective exchange rate computed based on unit labor costs (acronym is 134..NEUZF...), available since January 1960, is from the *IMF*'s *IFS*, and it has been converted to the quarterly frequency by taking averages within the quarter. The overall sample period is from 1960:1 to 1990:1.

B Details of the Markov-Chain Monte Carlo Procedure

We estimate (1)-(9) *via* Bayesian methods. The next two subsections describe our choices for the priors, and the Markov-Chain Monte Carlo algorithm we use to simulate the posterior distribution of the hyperparameters and the states conditional on the data, while the third section discusses how we check for convergence of the Markov chain to the ergodic distribution.

B.1 Priors

For the sake of simplicity, the prior distributions for the initial values of the states—θ_0, α_0, and h_0—which we postulate all to be normal, are assumed to be independent both from one another, and from the distribution of the hyperparameters. In order to calibrate the prior distributions for θ_0, α_0 and h_0 we estimate a time-invariant version of (1) based on the first 8 years of data, from 1959:3 to 1966:4, and we set

$$\theta_0 \sim N\left[\hat{\theta}_{OLS}, 4 \cdot \hat{V}(\hat{\theta}_{OLS})\right] \tag{B1}$$

As for α_0 and h_0 we proceed as follows. Let $\hat{\Sigma}_{OLS}$ be the estimated covariance matrix of ϵ_t from the time-invariant VAR, and let C be the lower-triangular Choleski factor of $\hat{\Sigma}_{OLS}$—i.e., $CC' = \hat{\Sigma}_{OLS}$. We set

$$\ln h_0 \sim N(\ln \mu_0, 10 \times I_4) \tag{B2}$$

where μ_0 is a vector collecting the logarithms of the squared elements on the diagonal of C. We then divide each column of C by the corresponding element on the diagonal—let's call the matrix we thus obtain $\tilde{C}$—and we set

$$\alpha_0 \sim N[\tilde{\alpha}_0, \tilde{V}(\tilde{\alpha}_0)] \tag{B3}$$

where $\tilde{\alpha}_0$—which, for future reference, we define as $\tilde{\alpha}_0 \equiv [\tilde{\alpha}_{0,11},\ \tilde{\alpha}_{0,21},\ ...,\ \tilde{\alpha}_{0,61}]'$—is a vector collecting all the non-zero and non-one elements of $\tilde{C}^{-1}$ (i.e, the elements below the diagonal), and its covariance matrix, $\tilde{V}(\tilde{\alpha}_0)$, is postulated to be diagonal, with each individual (j,j) element equal to 10 times the absolute value of the corresponding j-th element of $\tilde{\alpha}_0$. Such a choice for the covariance matrix of α_0 is clearly arbitrary, but is motivated by our goal to scale the variance of each individual element of α_0 in such a way as to take into account of the element's magnitude.

Turning to the hyperparameters, we postulate independence between the parameters corresponding to the three matrices Q, S, and Z—an assumption we adopt uniquely for reasons of convenience—and we make the following, standard assumptions. The matrix Q is postulated to follow an inverted Wishart distribution,

$$Q \sim IW\left(\bar{Q}^{-1}, T_0\right) \tag{B4}$$

with prior degrees of freedom T_0 and scale matrix $T_0\bar{Q}$. In order to minimize the impact of the prior, thus maximizing the influence of sample information, we set T_0 equal to the minimum value allowed, the length of θ_t plus one. As for $\bar{Q}$, we calibrate it as $\bar{Q}= \gamma \times \hat{\Sigma}_{OLS}$, setting γ=3.5×10^{-4}, the same value used by Cogley and Sargent (2005).

The three blocks of S are assumed to follow inverted Wishart distributions, with prior degrees of freedom set, again, equal to the minimum allowed, respectively, 2, 3 and 4:

$$S_1 \sim IW\left(\bar{S}_1^{-1}, 2\right) \tag{B5}$$

$$S_2 \sim IW\left(\bar{S}_2^{-1}, 3\right) \tag{B6}$$

$$S_3 \sim IW\left(\bar{S}_3^{-1}, 4\right) \tag{B7}$$

As for $\bar{S}_1$, $\bar{S}_2$ and $\bar{S}_3$, we calibrate them based on $\tilde{\alpha}_0$ in (B3) as $\bar{S}_1$=$10^{-3} \times |\tilde{\alpha}_{0,11}|$, $\bar{S}_2$=10^{-3}×diag($[|\tilde{\alpha}_{0,21}|, |\tilde{\alpha}_{0,31}|]'$) and $\bar{S}_3$=10^{-3}×diag($[|\tilde{\alpha}_{0,41}|, |\tilde{\alpha}_{0,51}|, |\tilde{\alpha}_{0,61}|]'$). Such a calibration is consistent with the one we adopted for Q, as it is equivalent to setting $\bar{S}_1$, $\bar{S}_2$ and $\bar{S}_3$ equal to 10^{-4} times the relevant diagonal block of $\tilde{V}(\tilde{\alpha}_0)$ in (B3). Finally, as for the variances of the stochastic volatility innovations, we follow Cogley and Sargent (2002, 2005) and we postulate an inverse-Gamma distribution for the elements of Z,

$$\sigma_i^2 \sim IG\left(\frac{10^{-4}}{2}, \frac{1}{2}\right) \tag{B8}$$

B.2 Simulating the posterior distribution

We simulate the posterior distribution of the hyperparameters and the states conditional on the data *via* the following MCMC algorithm, combining elements of Primiceri (2005) and Cogley and Sargent (2002, 2005). In what follows, x^t denotes the entire history of the vector x up to time t—i.e. $x^t \equiv [x_1', x_2', \quad, x_t']'$—while T is the sample length.

(a) Drawing the elements of θ_t Conditional on Y^T, α^T, and H^T, the observation equation (1) is linear, with Gaussian innovations and a known covariance matrix. Following Carter and Kohn (2004), the density $p(\theta^T|Y^T, \alpha^T, H^T, V)$ can be factored as

$$p(\theta^T|Y^T, \alpha^T, H^T, V) = p(\theta_T|Y^T, \alpha^T, H^T, V)\prod_{t=1}^{T-1} p(\theta_t|\theta_{t+1}, Y^T, \alpha^T, H^T, V) \tag{B9}$$

Conditional on α^T, H^T, and V, the standard Kalman filter recursions nail down the first element on the right hand side of (B9), $p(\theta_T|Y^T, \alpha^T, H^T, V) = N(\theta_T, P_T)$, with P_T being the precision matrix of θ_T produced by the Kalman filter. The remaining elements in the factorization can then be computed via the backward recursion algorithm found, e.g., in Kim and Nelson (2000), or Cogley and Sargent (2005, appendix B.2.1). Given the conditional normality of θ_t, we have

$$\theta_{t|t+1} = \theta_{t|t} + P_{t|t}P_{t+1|t}^{-1}\left(\theta_{t+1} - \theta_t\right) \tag{B10}$$

$$P_{t|t+1} = P_{t|t} - P_{t|t}P_{t+1|t}^{-1}P_{t|t} \tag{B11}$$

which provides, for each t from T-1 to 1, the remaining elements in (1), $p(\theta_t|\theta_{t+1}, Y^T, \alpha^T, H^T, V) = N(\theta_{t|t+1}, P_{t|t+1})$. Specifically, the backward recursion starts with a draw from $N(\theta_T, P_T)$, call it $\tilde{\theta}_T$ Conditional on $\tilde{\theta}_T$, (B10)-(B11) give us $\theta_{T-1|T}$ and $P_{T-1|T}$, thus allowing us to draw $\tilde{\theta}_{T-1}$ from $N(\theta_{T-1|T}, P_{T-1|T})$, and so on until t=1.

(b) Drawing the elements of α_t Conditional on Y^T, θ^T, and H^T, following Primiceri (2005), we draw the elements of α_t as follows. Equation (1) can be rewritten as $A_t\tilde{Y}_t \equiv A_t(Y_t\text{-}X_t'\theta_t)$=$A_t\epsilon_t \equiv u_t$, with Var($u_t$)=$H_t$, namely

$$\tilde{Y}_{2,t} = -\alpha_{21,t}\tilde{Y}_{1,t} + u_{2,t} \tag{B12}$$

$$\tilde{Y}_{3,t} = -\alpha_{31,t}\tilde{Y}_{1,t} - \alpha_{32,t}\tilde{Y}_{2,t} + u_{3,t} \tag{B13}$$

$$\tilde{Y}_{4,t} = -\alpha_{41,t}\tilde{Y}_{1,t} - \alpha_{42,t}\tilde{Y}_{2,t} - \alpha_{43,t}\tilde{Y}_{3,t} + u_{4,t} \tag{B14}$$

—plus the identity $\tilde{Y}_{1,t} = u_{1,t}$—where $[\tilde{Y}_{1,t}, \tilde{Y}_{2,t}, \tilde{Y}_{3,t}, \tilde{Y}_{4,t}]' \equiv \tilde{Y}_t$. Based on the observation equations (B12)-(B14), and the transition equation (7), the elements of α_t can then be drawn by applying the same algorithm we described in the previous paragraph separately to (B12), (B13) and (B14). The assumption that S has the block-diagonal structure (9) is in this respect crucial, although, as stressed by Primiceri (2005, Appendix D), it could in principle be relaxed.

(c) Drawing the elements of H_t Conditional on Y^T, θ^T, and α^T, the orthogonalised innovations $u_t \equiv A_t(Y_t\text{-}X_t'\theta_t)$, with Var($u_t$)=$H_t$, are observable. Following Cogley and Sargent (2002), we then sample the $h_{i,t}$'s by applying the univariate algorithm of Jacquier, Polson, and Rossi (1994) element by element.[26]

(d) Drawing the hyperparameters Finally, conditional on Y^T, θ^T, H^T, and α^T, the innovations to θ_t, α_t, the $h_{i,t}$'s are observable, which allows us to draw the hyperparameters—the elements of Q, S_1, S_2 S_3, and the σ_i^2—from their respective distributions.

Summing up, the MCMC algorithm simulates the posterior distribution of the states and the hyperparameters, conditional on the data, by iterating on (a)-(d). In what follows we use a burn-in period of 50,000 iterations to converge to the ergodic distribution, and after that we run 10,000 more iterations sampling every 10th draw in order to reduce the autocorrelation across draws.[27]

B.3 Assessing the convergence of the Markov chain to the ergodic distribution

Following Primiceri (2005), we assess the convergence of the Markov chain by inspecting the autocorrelation properties of the ergodic distribution's draws. Specifically, in

[26] For details, see Cogley and Sargent (2005, Appendix B.2.5).

[27] In this we follow Cogley and Sargent (2005). As stressed by Cogley and Sargent (2005), however, this has the drawback of 'increasing the variance of ensemble averages from the simulation'.

what follows we consider the draws' inefficiency factors (henceforth, IFs), defined as the inverse of the relative numerical efficiency measure of Geweke (1992),

$$RNE = (2\pi)^{-1} \frac{1}{S(0)} \int_{-\pi}^{\pi} S(\omega) d\omega \tag{B15}$$

where $S(\omega)$ is the spectral density of the sequence of draws from the Gibbs sampler for the quantity of interest at the frequency ω. We estimate the spectral densities by smoothing the periodograms in the frequency domain by means of a Bartlett spectral window. Following Berkowitz and Diebold (1998), we select the bandwidth parameter automatically *via* the procedure introduced by Beltrao and Bloomfield (1987).

Figure 5 and 6 show, for the United States and West Germany, respectively, the draws' IFs for the models' hyperparameters—i.e., the free elements of the matrices Q, Z, and S—and for the states, i.e. the time-varying coefficients of the VAR (the θ_t), the volatilities (the $h_{i,t}$'s), and the non-zero elements of the matrix A_t. As the figure clearly shows, the autocorrelation of the draws is uniformly very low, being in the vast majority of cases around or below 3—as stressed by Primiceri (2005, Appendix B), values of the IFs below or around twenty are generally regarded as satisfactory.

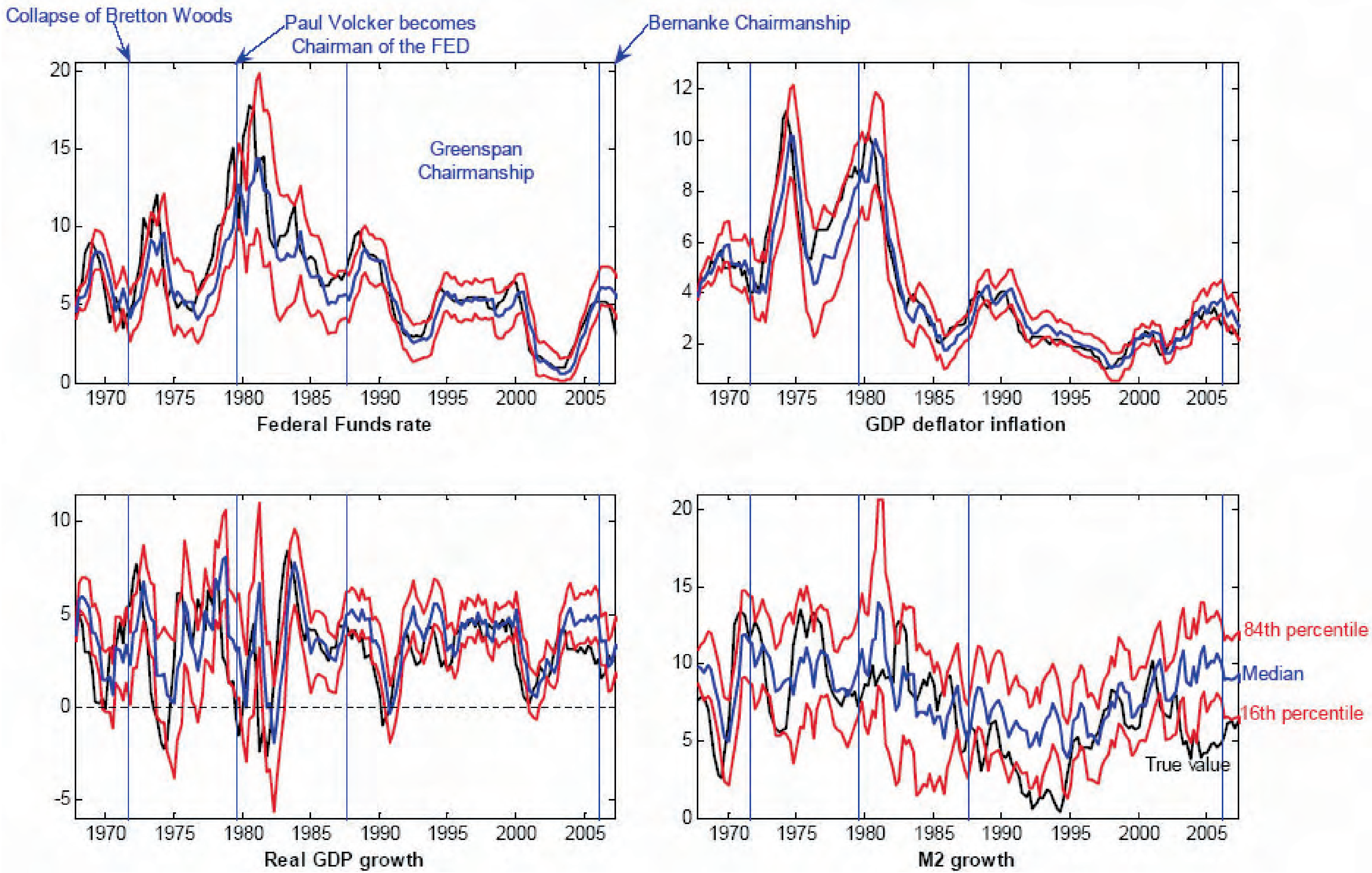

Figure 1 Bringing the *Bundesbank* to the United States: results from counterfactual simulations based on the four-variables VAR (medians and 16th and 84th percentiles), together with the actual series

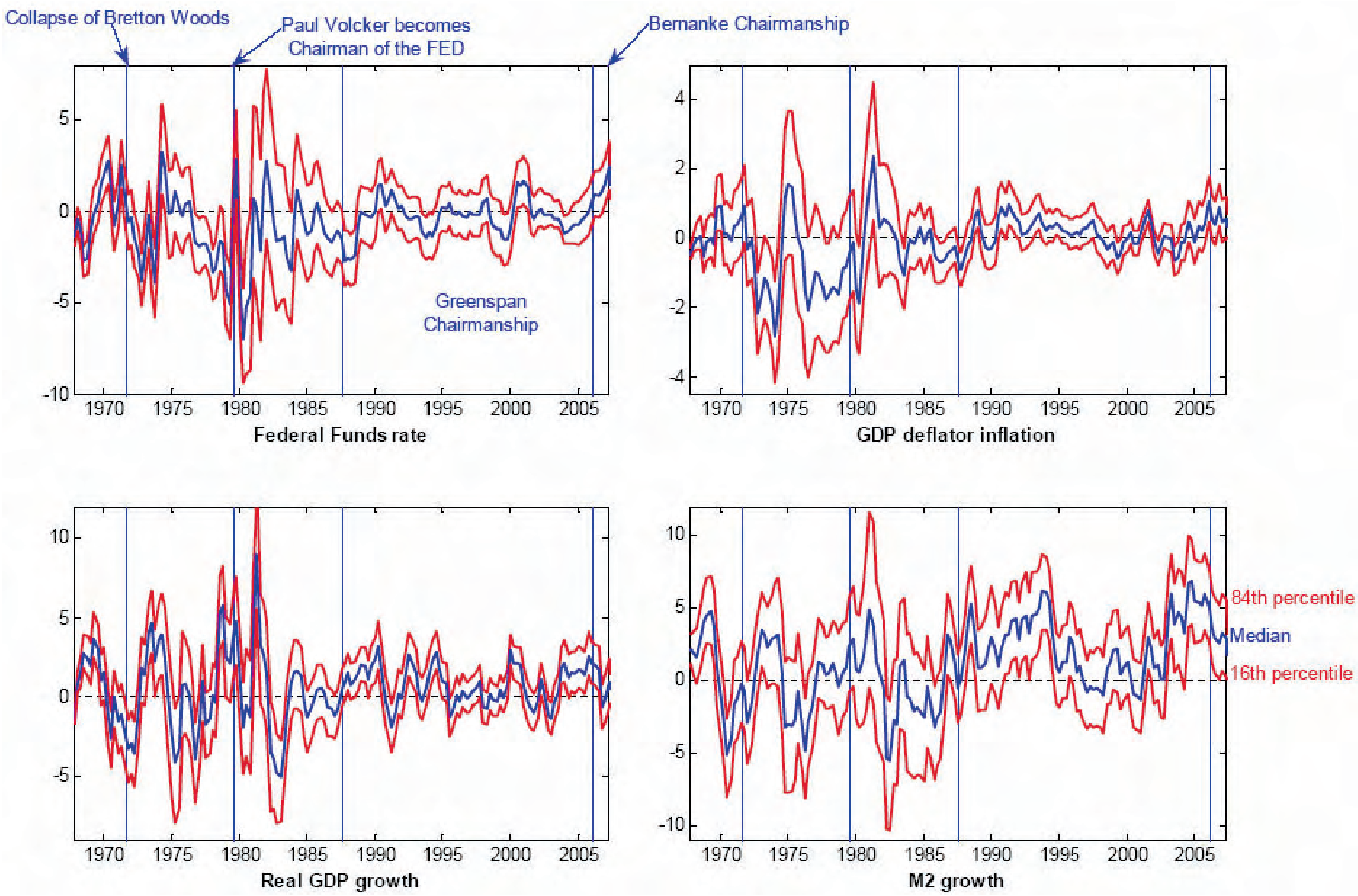

Figure 2 Bringing the *Bundesbank* to the United States: results from counterfactual simulations based on the four-variables VAR (medians and 16th and 84th percentiles of the difference between actual and counterfactual

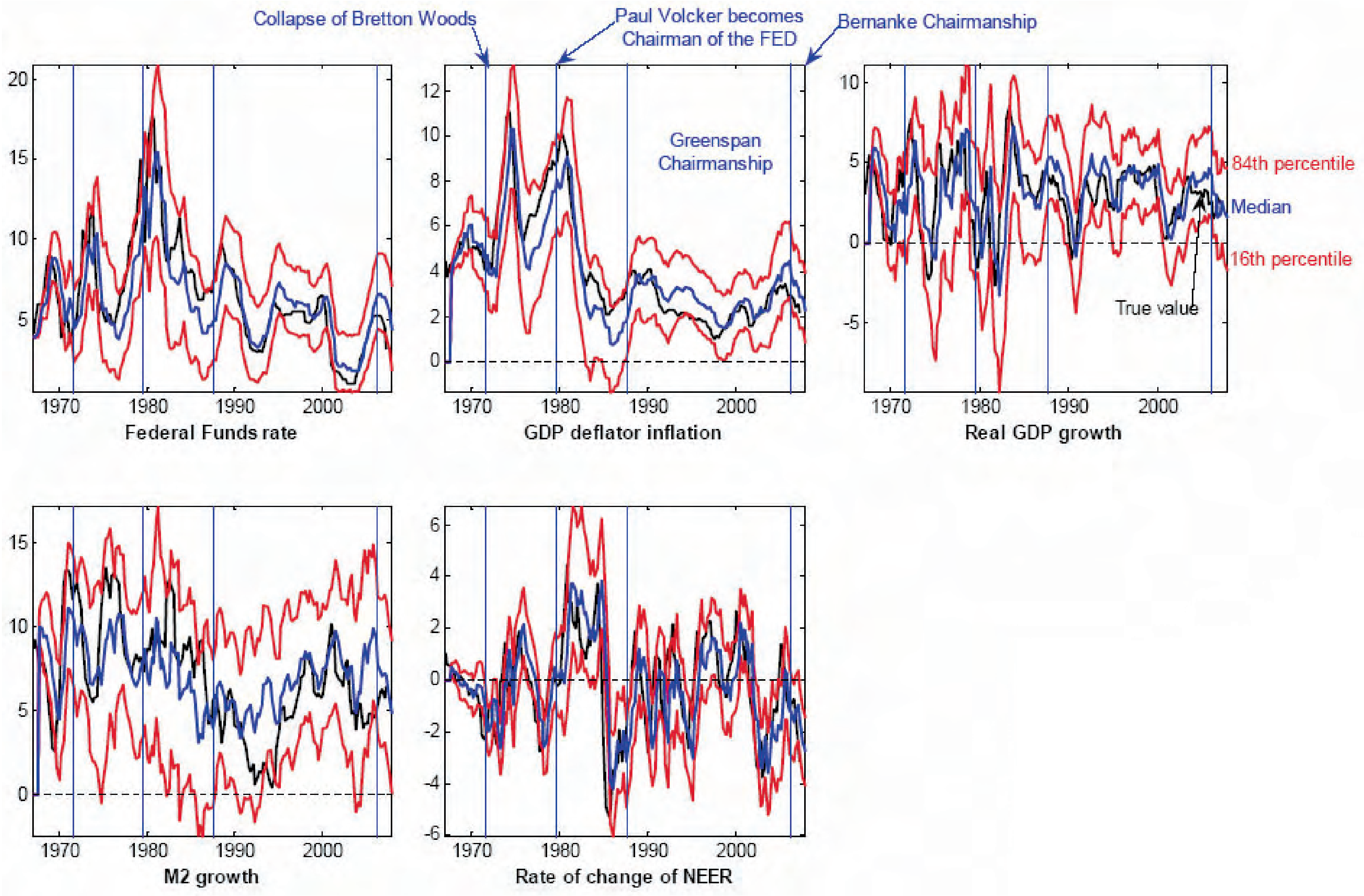

Figure 3 Bringing the *Bundesbank* to the United States: results from counterfactual simulations based on the five-variables VAR (medians and 16th and 84th percentiles), together with the actual series

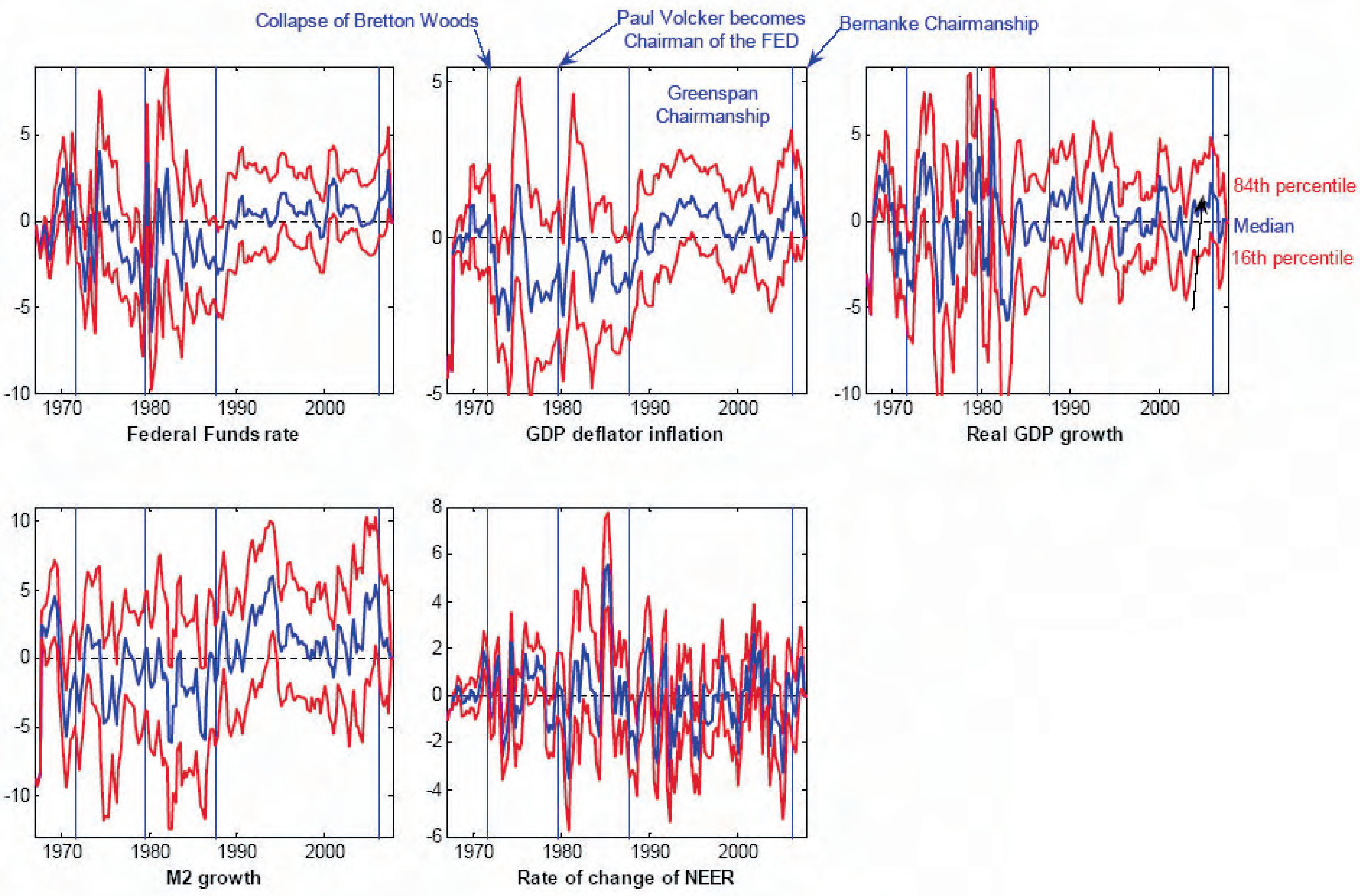

Figure 4 Bringing the *Bundesbank* to the United States: results from counterfactual simulations based on the five-variables VAR (medians and 16th and 84th percentiles of the difference between actual and counterfactual series)

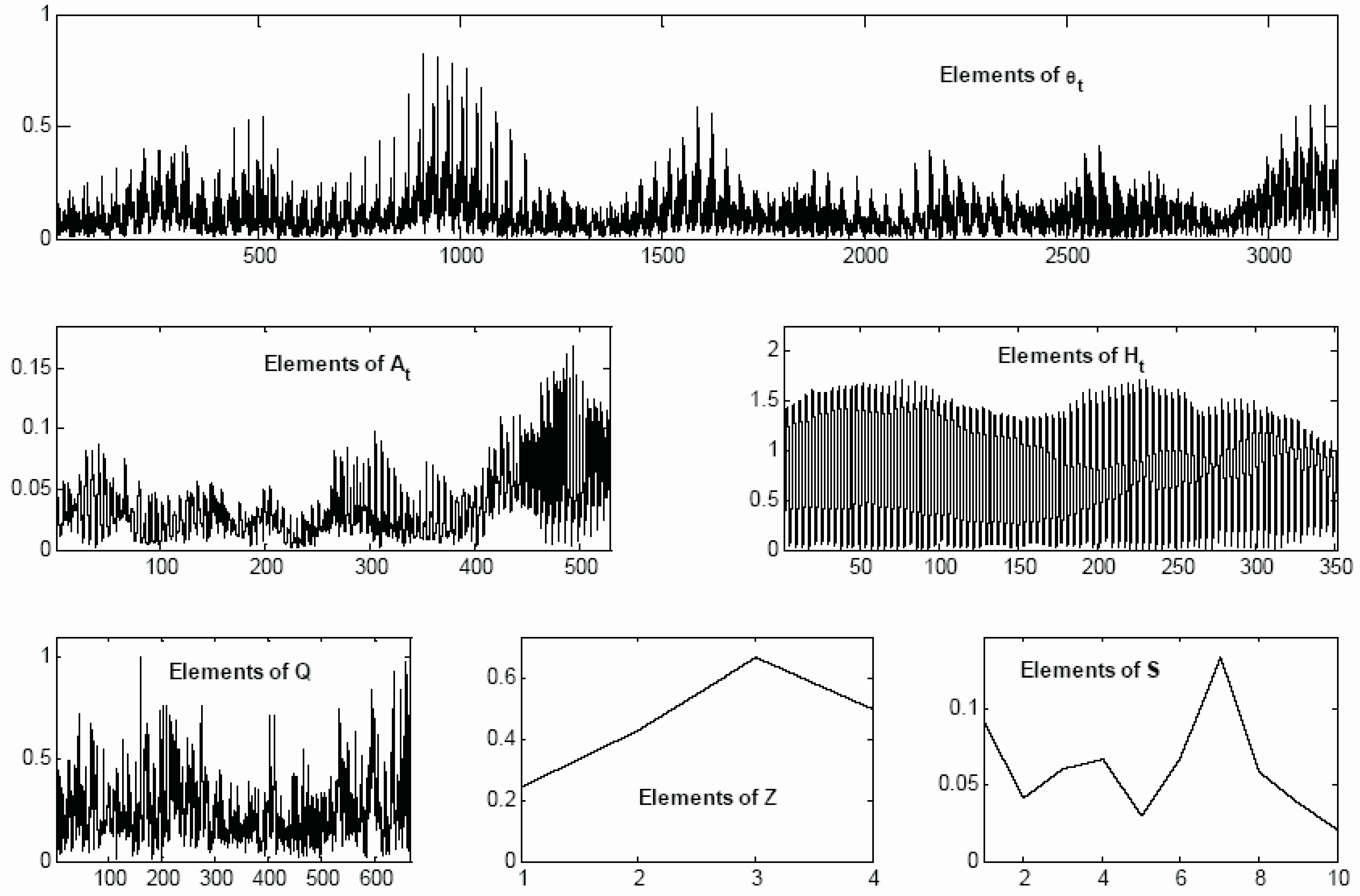

Figure 5 Checking for the convergence of the Markov chain: inefficiency factors for the draws from the ergodic distribution for the hyperparameters and the states (West Germany)

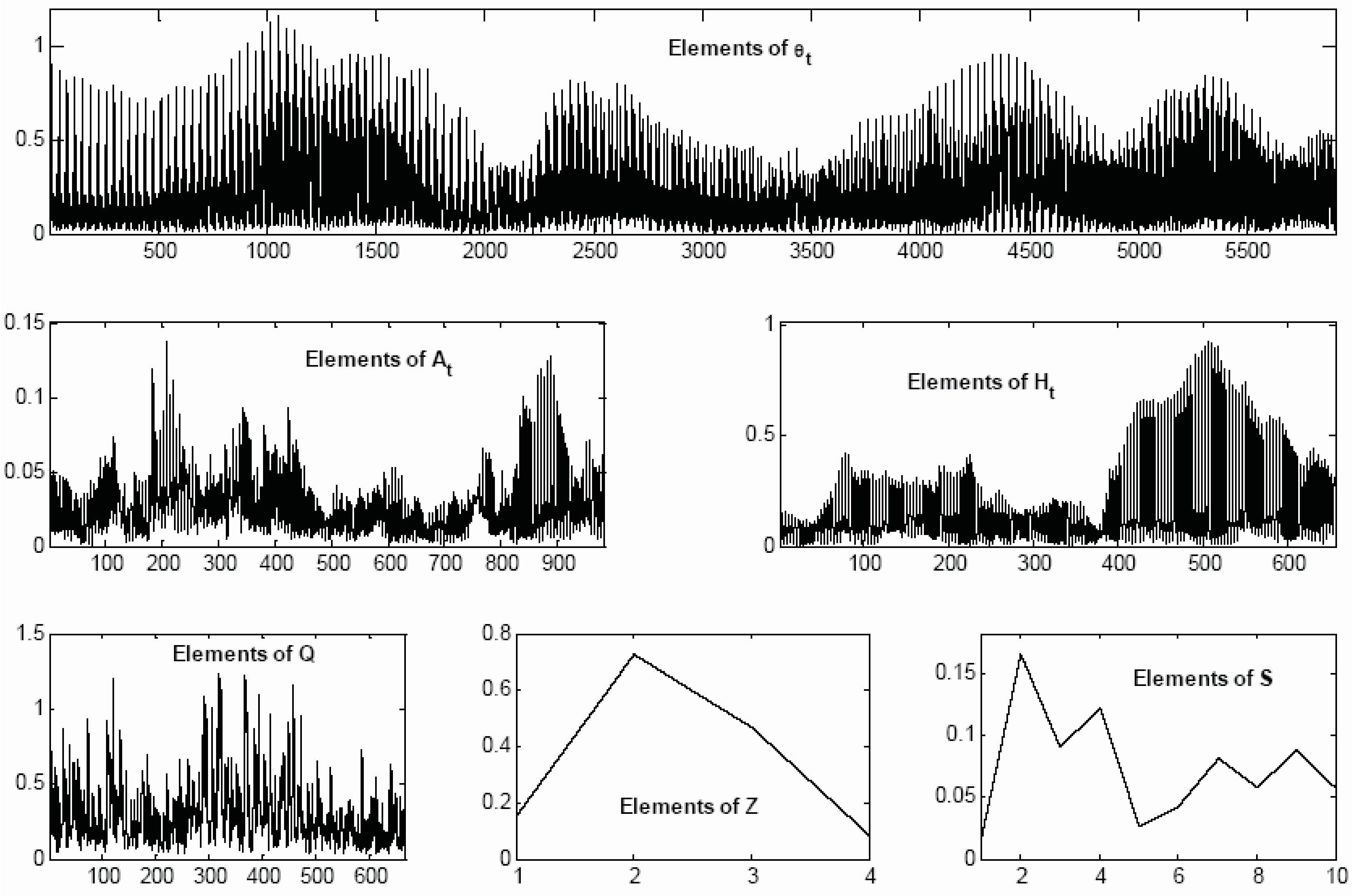

Figure 6 Checking for the convergence of the Markov chain: inefficiency factors for the draws from the ergodic distribution for the hyperparameters and the states (United States)

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