Policy Shifts, Heteroskedasticity and the Lucas Critique*

Thomas A. Lubik
Department of Economics
Johns Hopkins University

Paolo Surico
Bank of England and
University of Bari

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Abstract

This paper re-considers the empirical relevance of the Lucas critique using a DSGE sticky price model in which a *weak* central bank response to inflation generates equilibrium indeterminacy. The model is calibrated on the magnitude of the historical shift in the Federal Reserve's policy rule and is capable of generating the decline in the volatility of inflation and real activity observed in postwar U.S. data. Using Monte Carlo simulations and a backward-looking model of aggregate supply and demand, we show that shifts in the policy rule induce breaks in both the reduced-form coefficients and the reduced-form error variance. The statistics of popular parameter stability tests are severely downward biased if such heteroskedasticity is neglected. In contrast, when the instability of the reduced-form error variances is accounted for the Lucas critique is found to be empirically relevant on both artificial and actual data.

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1 Introduction

The Lucas critique comes as close to a natural law as seems possible in macroeconomics. While its theoretical validity is largely uncontested there is, however, a surprising lack of empirical support in the literature. A host of studies has argued that the relevance of the Lucas critique is limited in practice. In particular, shifts in policy seem not to have any significant effects on the stability of backward representations of macroeconomic models for various documented historical episodes. We argue in this paper that the evident inapplicability of the Lucas critique is due to problems with the size and power of the econometric tests used. We use a simple structural model of the U.S. economy as a data-generating process (DGP) to illustrate these issues both conceptually and by means of a Monte Carlo analysis. Our empirical findings confirm that the Lucas critique is valid for the shift in U.S. monetary policy behavior in the early 1980s.

It is difficult to underestimate the importance of Lucas (1976) in the development of modern macroeconomic thought. The introduction of rational expectations in macroeconomics at the middle of the 1970s represented an intellectual revolution for the profession and a serious challenge for large-scale, backward-looking econometric models that were used for policy analysis. Lucas argued that changes in policy have an immediate effect on agents' decision rules since they are inherently forward-looking and adapt to the effects of the new policy regime. An important corollary of this argument is that any policy evaluation based on backward-looking macroeconomic models is misleading whenever such policy shifts occur.

Empirical studies of the validity of the Lucas critique (e.g. Estrella and Fuhrer, 2003, Rudebusch, 2005) suggest that it is unimportant in practice. Tests for parameter stability in backward-looking specifications or reduced forms of macroeconomic relationship typically fail to reject the null of structural stability in the presence of well-documented policy shifts. This evidence would support the conclusion that policy changes are, in the terminology of Leeper and Zha (2003), 'modest' enough not to alter the behavior of private agents in a manner that is detectable for the econometrician. A further implication is that backward-looking monetary models of the type advocated by Rudebusch and Svensson (1999), that perform well empirically, are safe to use in policy experiments.

We argue in this paper that the failure of these studies to detect Lucas critique effects rests on the use of parameter stability tests such as the Chow break-point test and the superexogeneity test. These tests implicitly assume equality of the variances of the reduced-form innovations across sub-samples. It is, however, well known, at least since Toyoda (1974), that even moderate heteroskedasticity has a considerable impact on the significance levels of the null of parameter stability, which leads to an incorrect acceptance of parameter invariance. Moreover, and similar to Lindé (2001), we also find evidence of a small-sample bias which tends to hide the instability of backward-looking specifications.

At the heart of Lucas' critique lies an emphasis on the use of fully-specified, optimisation-based dynamic stochastic general equilibrium (DSGE) models for policy analysis. We follow this approach and work within the confines of a structural New Keynesian model that has been widely used for monetary policy analysis. We treat this model as our DGP which we use for generating simulated time series for a Monte Carlo analysis. This allows us to control the environment in which the policy change occurs. Otherwise, it might be difficult to distinguish between actual policy shifts and changes in the economy's driving processes.

Our main experiment is to model the monetary policy change in the U.S. that occurred at the turn of the 1980s. We follow Clarida et al. (2000) and Lubik and Schorfheide (2004) in assuming that policy during the 1970s responded only weakly to inflation, whereas with the start of Volcker's tenure policy became much more aggressively anti-inflationary. In the context of our model, this implies that in the first sub-sample the rational expectations (RE) equilibrium was indeterminate, and determinate later on. On the basis of the approach developed by Lubik and Schorfheide (2003) to describe indeterminate equilibria, we show, first, that a policy shift of the magnitude observed in the Federal Reserve's monetary policy rule is capable of reproducing the historical decline in the volatility of real activity, inflation and the interest rate. Furthermore, we demonstrate that such a decline invalidates the assumption of constant reduced-form innovation variances in a widely used backward-looking model of aggregate supply and aggregate demand estimated over the sub-samples associated with the regime shift.

Monte Carlo simulations show that the instability of the error variances severely affects the power of the Chow test and the superexogeneity test to the effect that it prevents the rejection of the *incorrect* null hypothesis of parameter stability. When the heteroskedasticity in the estimated backward-looking specifications is accounted for, we find robust evidence in favour of the empirical relevance of the Lucas critique. An application for U.S. data reveals that controlling for the decline in the variance of the estimated error terms matters also in practice. We show that this overturns the results from parameter stability tests which are based on the (incorrect) assumption of equal errors variance between sub-samples. Hence, we conclude that the Lucas critique is alive and well.

To our knowledge, this is the first paper that investigates the effect of equilibrium indeterminacy on the stability of reduced-form models. The message of the paper is, however, not predicated on policy changes that induce switches between determinacy and indeterminacy, but it helps to sharpen our argument. We also report results from simulations that preserve a specific regime.

The paper is organized as follows. Section 2 presents and discusses analytical results in the context of simple examples on how policy changes in structural models affect their reduced-form representations and how this relates to the debate about the empirical validity of the Lucas critique. Section 3 introduces the model we use for the Monte Carlo study and explains our simulation strategy. In the following Section we show that the policy-induced shift from indeterminacy to determinacy is capable of explaining the Great Moderation. In Section 5 we report the Monte Carlo evidence on the structural stability tests based on a backward-looking model of aggregate supply and aggregate demand. We also provide a sensitivity analysis for variations in the values of some model parameters. The sixth Section contains empirical results obtained on actual U.S. data, while the last section concludes.

2 A Primer on Indeterminacy

We assume throughout in our paper that the data are generated by a DSGE model. Analyzing the effects of parameter changes thus requires an understanding of the reduced-form properties of structural linear rational expectations models. In this sense, our paper is similar to Lindé (2001) and Rudebusch (2005). However, we go further than the earlier literature in analyzing the effects of policy break that changes the equilibrium properties of the economy, specifically a change from indeterminacy to determinacy. In this section, we therefore give a brief overview on indeterminacy in linear rational expectations models.

2.1 Determinate and Indeterminate Equilibria

Consider the simple expectational difference equation:

$$x_t = aE_t x_{t+1} + \varepsilon_t,$$

where a is a parameter, ε_t is a white noise process with mean zero and variance σ^2, and E_t is the rational expectations operator conditional on information at time t. It is well known that the type of solution depends on the value of the parameter a. If $|a| < 1$ there is a unique ('determinate') solution which is simply:

$$x_t = \varepsilon_t.$$

On the other hand, if $|a| > 1$, there are multiple solutions and the rational expectations equilibrium is indeterminate.

In order to derive the entire set of solutions we follow the approach developed by Lubik and Schorfheide (2003). For this purpose it is often convenient to rewrite the model by introducing endogenous forecast errors $\eta_t = x_t - E_{t-1}x_t$. Define $\xi_t = E_t x_{t+1}$ so that:

$$\xi_t = \frac{1}{a}\xi_{t-1} - \frac{1}{a}\varepsilon_t + \frac{1}{a}\eta_t. \tag{1}$$

Under indeterminacy this is a stable difference equation which imposes no further restriction on the evolution of the endogenous forecast error η_t.[1] Hence, any covariance-stationary stochastic process for η_t is *a* solution for this linear rational expectations model. The forecast error can then be expressed as a linear combination of the model's fundamental disturbances and extraneous sources of uncertainty, typically labeled 'sunspots'. The coefficients on the shocks in this decomposition generally depend on parameters of the model and parameters that index a specific sunspot equilibrium. To wit, the forecast error can be written as:

$$\eta_t = m\varepsilon_t + \zeta_t,$$

where the sunspot ζ_t is a martingale-difference sequence, and m is an unrestricted parameter.[2] Substituting this into Eq. (1) yields the solution under indeterminacy:

$$x_t = \frac{1}{a}x_{t-1} + m\varepsilon_t - \frac{1}{a}\varepsilon_{t-1} + \zeta_t.$$

The evolution of x_t now depends on an additional (structural) parameter m which indexes specific rational expectations equilibria.

Indeterminacy affects the behavior of the model in three main ways. First, indeterminate solutions exhibit a much richer lag structure and more persistence than the corresponding determinate solution. This feature has been exploited by Lubik and Schorfheide (2004) in distinguishing between the two types of rational expectations equilibria in U.S. data.

[1] In the case of determinacy, the restriction imposed is that $\xi_t = 0$, all t, which implies $\eta_t = \varepsilon_t$.

[2] There is a technical subtlety in that ζ_t is actually, in the terminology of Lubik and Schorfheide (2003), a reduced-form sunspot shock, with $\zeta_t = m_\zeta \zeta_t^*$. Furthermore, in less simple models, there would be additional restrictions on the coefficients which would depend on other structural parameters.

In the simple example, this is strikingly evident: under determinacy the solution for x_t is white noise, while under indeterminacy the solution is described by an ARMA(1,1) process. Second, under indeterminacy sunspot shocks can affect equilibrium dynamics. Other things being equal, variables generated by sunspot equilibria are inherently more volatile than their determinate counterparts. The third implication, especially emphasized by Lubik and Schorfheide (2003), is that indeterminacy affects the response of the model to *fundamental* shocks, whereas the response to sunspot shocks is uniquely determined. In the example, innovations to ε_t could either increase or decrease x_t depending on the sign of m.

2.2 Indeterminacy and the Lucas Critique

We now try to provide some insight into the analytics of the Lucas critique by means of a simple example. We consider the following simple two equation model which describe the evolution of an economic variable x_t and a policy variable y_t:

$$x_t = aE_t x_{t+1} + bE_t y_{t+1} + \varepsilon_{1t}, \tag{2}$$

$$y_t = cx_t + \varepsilon_{2t}. \tag{3}$$

ε_{1t}, ε_{2t} are white noise processes with variances σ_1^2, σ_2^2, respectively; a, b are structural parameters, and c is a policy parameter. We assume for simplicity that all parameters are positive. Eq. (3) is a feedback rule of the type often used in monetary policy models.

The model has a unique rational expectations equilibrium if $0 < c < \frac{1-a}{b}$, the solution of which is:

$$\begin{aligned} x_t &= \varepsilon_{1t}, \\ y_t &= c\varepsilon_{1t} + \varepsilon_{2t}. \end{aligned}$$

As in the simple example above, both variables are white noise processes. If $c > \frac{1-a}{b}$ the solution is indeterminate. The laws of motion for the two variables are as follows:

$$\begin{aligned} x_t &= \frac{1}{a+bc} x_{t-1} + m\varepsilon_{1t} - \frac{1}{a+bc}\varepsilon_{1t-1} + \zeta_t, \\ y_t &= \frac{1}{a+bc} y_{t-1} + mc\varepsilon_{1t} - \frac{c}{a+bc}\varepsilon_{1t-1} + \varepsilon_{2t} - \frac{1}{a+bc}\varepsilon_{2t-1} + c\zeta_t. \end{aligned}$$

Again, the change in the stochastic properties of the variables, when moving across the parameter space is quite evident. If (2) - (3) is the data generating process, then these equations are the reduced-form representations upon which tests for the empirical relevance of the Lucas critique are based.

The type of experiment we are interested in is an exogenous, unanticipated change in the policy parameter c. We can distinguish four different scenarios: a shift from determinacy to indeterminacy, from indeterminacy to determinacy, a change that preserves a previously determinate equilibrium, and an indeterminate equilibrium. If the break in c is such that the solution stays determinate, the behavior of x_t is unaffected while the variance of y_t changes. However, an econometrician could not distinguish between a change in σ_2^2 and the policy coefficient by observing y_t. A change in the variance of ε_{1t}, on the other hand, could be deduced from observations on both variables.

The more interesting case is a change in policy behavior that moves c across the boundary $\frac{1-a}{b}$ that separates the determinacy from the indeterminacy region. Suppose the economy is initially in an indeterminate equilibrium associated with a policy parameter c_0. An unexpected (and believed to be permanent) policy shift to $c_1 < \frac{1-a}{b} < c_0$ moves the economy to a determinate equilibrium. This implies a dramatic change in the nature of the stochastic processes for x_t and y_t, as they switch from persistence to white noise. Specifically, both their variance and the degree of auto-correlation decline. Furthermore, sunspot shocks, and the extra volatility introduced, no longer affect the dynamics. It is this type of scenario we have in mind in our empirical analysis. The standard example is the change in monetary policy that occurred with Volcker's tenure at the Federal Reserve. Naturally, similar reasoning applies for the opposite case when a policy change moves the economy from a determinate to an indeterminate equilibrium.

Yet even if equilibrium properties are unaffected by shifts in policy, the properties of the reduced-form model are not. Crucially, the variance of the error term in a regression of the endogenous variables on their first lag would (i) change with the shift in policy, and (ii) be correlated with the regressor. The latter issue is, of course, well known, and methods to deal with this, such as IV-estimation, are now widely used.[3] The former is not as well appreciated, however. Consequently, the main thrust of our paper is directed at this form of heteroskedasticity, which reduces the power of parameter stability tests based on reduced-form regressions.[4]

[3] See, however, Lubik and Schorfheide (2005) for a set of examples in the context of DSGE-models where IV-methods fail.

[4] Structural estimation methods are immune against this problem since the structure of the reduced-form error would be reflected in, say, the likelihood function. This is likely the main reason why the GMM-based

3 Testing for the Lucas Critique

In this section, we lay out a simple algorithm to investigate the empirical relevance of the Lucas critique using a small-scale DSGE monetary model. To focus on the importance of a change in monetary policy, we simulate a shift in the coefficients of the policy rule while keeping the parameters describing the structure of the economy fixed across simulations. For each simulation, we thus obtain two sets of artificial data. The first set is generated from an indeterminate equilibrium and is associated with pre-1979 estimates of the policy rule typically found in the literature. The second sample is generated under the assumption of determinacy and it corresponds to post-1979 description of policy behavior. Any difference in the econometric results from the simulations of the two sets of artificial data is thus only attributable to changes in policy.

3.1 The Model

Our simulation analysis is based on a log-linearized, microfounded New-Keynesian sticky price model of the business cycle of the kind popularized by Clarida, Galí and Gertler (1999), King (2000) and Woodford (2003) among others. The model consists of three aggregate relationships which describe the dynamic behavior of output y_t, inflation π_t, and the nominal interest rate R_t:

$$y_t = E_t y_{t+1} - \tau(R_t - E_t\pi_{t+1}) + g_t, \tag{4}$$

$$\pi_t = \beta E_t \pi_{t+1} + \kappa\,(y_t - z_t)\,, \tag{5}$$

$$R_t = \rho_R R_{t-1} + (1-\rho_R)\left(\psi_\pi \pi_t + \psi_y (y_t - z_t)\right) + \varepsilon_{R,t} \tag{6}$$

All variables are expressed in percentage deviations from a steady state.

Eq. (4) is a log-linearized IS curve derived from a household's intertemporal optimization problem in which consumption and nominal bond holdings are the control variables. Since there is no physical capital in this economy consumption is proportional to total resources up to an exogenous process g_t. The latter is typically interpreted as a government spending shock or a shock to preferences.[5] The parameter $\tau > 0$ represents the intertemporal elasticity of substitution.

results of Collard et al. (2002) are an outlier in the literature in that they find strong evidence for the empirical relevance of the Lucas critique.

[5]The IS curve can easily be reinterpreted as a schedule explaining the behavior of the 'output gap' defined as the difference between the stochastic components of output and the flexible price level of output (see Clarida, Galí, and Gertler, 1999). In this case, the shock g_t is also a source of potential output variations.

The Phillips-curve relationship (5) describes inflation dynamics as a function of output. It captures the staggered feature of an economy with Calvo-type price setting in which firms adjust their optimal price with a constant probability in any given period, independently of the time elapsed from the last adjustment. The discrete nature of price setting creates an incentive to adjust prices more the higher the future inflation expected at time t is. The contemporaneous rate of inflation is thus related to the difference between output and the stochastic marginal cost of production z_t via the parameter $\kappa > 0$, which can be interpreted as the inverse of the sacrifice ratio. $0 < \beta < 1$ is the agents' discount factor.

The policy rule (6) characterizes the behavior of the monetary authorities according to which the central bank adjusts the policy rate in response to current inflation and the output gap $(y_t - z_t)$. ψ_π, $\psi_y \geq 0$ are the policy coefficients. These adjustments are implemented smoothly, with $0 < \rho_R < 1$ measuring the degree of interest rate smoothing. The random variable $\varepsilon_{R,t}$ stands for the monetary policy shock, which can be interpreted either as unexpected deviations from the policy rule or as policy implementation error. We assume it to be a white noise process with mean zero and variance σ_ε^2.

The model description is completed by specifying the stochastic properties of the of the exogenous shocks g_t and z_t. We assume they are first-order autoregressive process with lag-coefficients $0 \leq \rho_g, \rho_z < 1$. Their innovations are assumed to be *i.i.d.* with variances σ_g^2 and σ_z^2, respectively. The equation system (4) - (6) together with the two processes for the exogenous shocks describes a linear rational expectations model that can be solved with the methods described in Sims (2002). We solve the model for both determinacy and indeterminacy. In the latter case, we follow the approach developed by Lubik and Schorfheide (2003).

3.2 Calibration

We choose parameter values based on the estimates in Lubik and Schorfheide (2004) who analyzed a model similar to ours. The values of the structural parameters are reported in Panel A of Table 1. We assign the same values in both subsamples to focus on changes in the policy coefficients. Unlike Lubik and Schorfheide (2004), however, we set the variance of sunspot shocks to zero in the case of indeterminacy. This is designed to facilitate the comparison between our simulations and the previous literature which typically does not take the presence of sunspot fluctuations into account.

A large body of empirical literature has documented the dramatic change in the conduct of U.S. monetary policy at the turn of the 1980s. The established consensus is that the nominal interest rate response to inflation in estimated policy rules became substantially more aggressive in the early 1980s than it was before. As the number of available estimates is quite large, we set the coefficients in the policy rules for Periods I and II to values that are broadly in line with the evidence pioneered by Clarida, Galì and Gertler (2000). Within the indeterminacy and determinacy regions, the results are fairly robust to deviations from the baseline values reported in Panel B of Table 1.

The interest rate response to inflation over the indeterminacy sample does not guarantee a unique rational expectations equilibrium because $\psi_\pi = 0.4$ violates the Taylor principle.[6] On the other hand, the parameter constellation associated with Period II guarantees a determinate equilibrium. As detailed above, under indeterminacy the solution of the model is affected along various dimensions. Most importantly, equilibrium dynamics can be driven by sunspot fluctuations. Secondly, the transmission of fundamental shocks can be altered relative to the unique rational expectations solution. Having set to zero the variance of sunspot disturbances, we will thus focus on the transmission mechanism effect.

3.3 Simulation Strategy

The procedure in the simulation analysis is as follows:

1. Solve the model under both indeterminacy and determinacy, and generate two samples of 82 and 61 observations[7] for output, inflation and the interest rate.[8]

2. For each artificial sample, estimate a reduced-form, backward-looking equation for output, inflation, and the interest rate.

3. Perform tests for error variance constancy and parameter stability across the two periods, computing the relevant statistics and probability values.

[6] The added requirement is that the other policy coefficients in the rule are not 'too large' to compensate for the weak inflation response. The analytics and intuition of indeterminate equilibria in the New Keynesian monetary model are discussed extensively by Lubik and Marzo (2006).

[7] The number of observations has been chosen to match the quarterly data points which are typically used in sub-sample analysis on US data (see Lubik and Schorfheide, 2004; and Clarida, Gali and Gertler, 2000). The first period ranges from 1960:1 to 1979:2 while the second corresponds to 1982:4 to 1997:4. In each simulated sample, 100 extra-observations are produced to provide us with a stochastic vector of initial conditions, which are then discarded.

[8] The solution of the model under indeterminacy is computed using the continuity assumption in Lubik and Schorfheide (2004). Similar results, not reported but available upon request, are obtained by imposing that the structural shocks are orthogonal to the sunspot shocks.

4. Repeat steps 1. to 3. $20,000$ times; for each parameterization select the median values of the statistics of interest.

5. Prior to the simulations, steps 1. - 4. are carried out under the (true) assumption of no policy regime shift and small-sample critical values are computed from the actual distributions of the relative statistics under such scenario. This ensures a correct size of the tests for parameter instability.

4 Replicating the Stylized Facts

Before evaluating the empirical relevance of the Lucas critique we assess the extent to which the magnitude of historical policy shifts can account for the decline in volatility and persistence of output, inflation and interest rate observed in U.S. data at the beginning of the 1980s. A large number of papers, including Kim and Nelson (1999), McConnell and Perez-Quiros (2000), and Blanchard and Simon (2001), document a sharp decline in the volatility of many U.S. macroeconomic variables. The variances of inflation and output declined by more than two thirds for inflation and by just less than one half for output. Following Stock and Watson (2002), this set of facts has come to be known as the Great Moderation.

There is, however, a simmering debate whether this was due to good luck (a decline in the volatility of exogenous shocks) or good policy (a shift to a more stabilizing policy regime). Studies by Boivin and Giannoni (2002), Stock and Watson (2002), and Kim et al. (2004) show that a sizable part of such decline was driven by a reduction in the volatility of the reduced-form innovations of estimated backward-looking specifications for inflation and output. This is invariably interpreted as *prima facie* evidence in favor of good luck.[9]

We thus ask the question whether the good luck and good policy scenario are capable of explaining this decline in volatilities in the context of our model. We present results for two sets of model calibrations. In the first experiment, we change the policy parameters such that the equilibrium switches from indeterminacy to determinacy. Differences in the dynamics are therefore driven only by a shift in the monetary policy rule as all non-policy parameters of the model are kept fixed across regimes. In the second calibration, we simulate a decline in the variance of the structural shocks such as to match the magnitude of the

[9] Benati and Surico (2006) show, however, that this interpretation may be unwarranted as also a move from indeterminacy to determinacy is consistent with a decline in the innovation variances of reduced-form VARs.

decline in the volatility of inflation and output observed in the data. All other parameters are fixed in the second simulation. In particular, the Taylor-principle applies and the equilibrium is therefore always determinate.

Simulation results for the good policy hypothesis are reported in Table 2. Panel A contains the standard deviation of output, inflation and the interest rate. All variables are far less volatile in the post-1982 simulated sample, with the most dramatic change occurring for the latter two. The volatility of output and the interest rate declined by 41%, while the standard deviation of inflation was reduced by 67% when moving from indeterminacy to determinacy. Panel B reports a measure of persistence obtained as the sum of autoregressive coefficients from the estimates of an univariate $AR(p)$ process. In line with the evidence in Cogley and Sargent (2005), the most dramatic change occurs for inflation whose inertia drops by almost 58%. The impact on the interest rate is considerably smaller. This can be explained by the higher smoothing parameter under determinacy which outweighs the impact of the 'extra' persistence induced by monetary policy under indeterminacy.[10]

We contrast this with the statistics obtained under the good luck hypothesis, reported in Table 3. While a decline in the variance of the structural shocks can replicate the drop in volatility, it cannot match the observed decrease in persistence, in particular the well-documented decline in inflation persistence. This strongly points toward a change in the transmission mechanism. A similar conclusion has been reached by McConnell and Perez-Quiros (2000), Boivin and Giannoni (2005) and Canova, Gambetti and Pappa (2005), pinpointing monetary policy changes. Our specific explanation is the shift from an indeterminate to a determinate equilibrium induced by a change in the Federal Reserve's policy behavior. As we demonstrated in Section 2 such a shift changes the dynamic properties of the reduced form of the model.

We find that both explanations are capable of explaining the Great Moderation. However, only the good policy hypothesis successfully reproduces the decline in inflation persistence. A shift from passive to active monetary policy can account *on its own* for the sharp decline in the volatility and persistence of inflation and output observed in the early 1980s. Furthermore, the decline in the volatility of the series is likely to contaminate the

[10] On the other hand, output persistence actually increases. This can be explained by the behavior of the ex-ante real interest rate. Across regimes the persistence of the nominal interest rate is almost unchanged and close to 0.8. In contrast, the persistence of expected inflation, whose behavior is similar to actual inflation, declines remarkably, from 0.7 to 0.3. This implies that the persistence of the ex-ante real interest rate is higher in the second sub-sample, which makes output more persistent.

properties of the error terms in estimated reduced-form models. As many widely used break point tests like the Chow test are based on the assumption of constancy of the innovations variance, we show below that the neglected heteroskedasticity leads to incorrect inference on the empirical relevance of the Lucas critique.

5 Results on Simulated Data

This section presents the results of the Monte Carlo simulations. We then present the relevant statistics of the Chow test and the superexogeneity test together with some robustness checks of our results.

5.1 Backward-Looking Models and Heteroskedasticity

We now look at the empirical importance of the Lucas critique through the lens of a widely used backward-looking model of aggregate supply and aggregate demand. The model is a quarterly version of the specification in Svensson (1997) and similar to the one in Rudebusch (2005):

$$\pi_t = \alpha_1 \pi_{t-1} + \alpha_2 \pi_{t-2} + \alpha_3 \pi_{t-3} + \alpha_4 \pi_{t-4} + \alpha_y y_{t-1} + u_t^{\pi}, \tag{7}$$

$$y_t = \beta_1 y_{t-1} + \beta_2 y_{t-2} + \beta_3 y_{t-3} + \beta_4 y_{t-4} + \beta_r (R - \pi)_{t-1} + u_t^{y}. \tag{8}$$

Inflation depends on its own past values within a year and on the lagged value of real activity. Aggregate demand is characterized by an autoregressive structure with four lags augmented by the lagged value of the real interest rate. The model is closed with a standard Taylor-type rule with interest rate smoothing:

$$R_t = \gamma_R R_{t-1} + \gamma_\pi \pi_t + \gamma_y y_t + u_t^R. \tag{9}$$

This model is estimated on simulated data under the assumption that the DGP is the structural model (4)-(6). We want to emphasize that we do not restrict the backward-looking model to the reduced-form representation of the DSGE model in Section 3. In particular, we do not impose the lag structure of the DSGE model. Instead, we specify a more general form of the kind used in earlier contributions and let the data decide whether some additional regressors, including the past values of the dependent variables, have explanatory power.

Table 4 reports residual standard deviations from the estimated model (7) - (9) together with the statistics of the Goldfeld-Quandt test of innovations variance constancy.[11] The variance of reduced-form innovations in the output and inflation equation exhibit a dramatic decline from the indeterminacy to the determinacy period of more than 50%. The decline for the interest rate equation is less dramatic. The null hypothesis of stability across samples is overwhelmingly rejected for the output and inflation equations.

This leads us to conclude that heteroskedasticity is present in the reduced-form equations over the two sub-samples. From the point of view of a structural DSGE model this is not surprising as we demonstrated in Section 2: it is precisely what would be expected in the presence of a monetary policy shift. What makes this observation potentially relevant for the empirical literature on the Lucas critique is that inference is based on parameter stability tests that neglect this feature of the data. More specifically, the Chow test and the superexogeneity test, used by Favero and Hendry (1992), Lindé (2001), and Rudebusch (2005) among many others, implicitly assume homoskedasticity of the error variances between sample regimes and are therefore subject to our criticism.

Toyoda (1974) has demonstrated that the size and power of the Chow test can be considerably affected by neglecting policy-induced differences in reduced-form error variances. The problem is most serious when samples of similar size are used. When overall sample size is small even moderate degrees of heteroskedasticity reduce the power of the test significantly. Both scenarios apply to the historical episode of the monetary policy shift in the U.S.. Perhaps not surprisingly, the conclusions drawn in the literature almost always go against the statistical importance of Lucas' critique.[12] We consequently ask the question whether the failure to detect any effects is due to a lack of power of these tests.

5.2 Parameter Stability Tests

To quantify the importance of heteroskedasticity for the results of conventional parameter stability tests, we follow the literature and use the Chow and superexogeneity tests on the two simulated samples. The superexogeneity test measures the probability of rejecting

[11] Estimates of the reduced-form coefficients are available from the authors upon request. We chose not to report these as the actual estimates are immaterial to our discusssion. What matters is the joint significance (summarized in the Table), since it is well known that two sets of parameters can be jointly different from each other even though the parameters are not statistically different individually.

[12] An exception is Lindé (2001) who argues in favor of its empirical relevance based on an empirical money demand equation. He shows, however, that the superexogeneity tests suffers from a serious small-sample problem.

the null of parameter stability in the output and inflation equations conditional on having rejected the null hypothesis of parameter stability in the interest rate equation.[13] We present results for three versions of the parameter stability and superexogeneity tests.

The first version is based on the residuals of the OLS estimates of equations (7) and (8), and therefore implicitly (and incorrectly) assumes homoskedasticity. A prominent example of this class of tests is the Chow statistics used by Lindé (2001) and Rudebusch (2005):

$$Chow = \frac{\left[RSS_{full} - (RSS_1 + RSS_2)\right]/m}{(RSS_1 + RSS_2)/(T-k)}, \tag{10}$$

where RSS represents the residual sum of squares of either the full sample or the first and second sub-periods, and m is the number of restrictions in an equation of k parameters.

In contrast, the second and third versions of the tests account for possible heteroskedasticity across sub-samples. In particular, we use a GLS version based on a two-step procedure. In the first step, Eqs. (7) and (8) are estimated by OLS over the two sub-samples. In the second step, the variables are normalized by the square root of the estimated innovation variance-covariance matrix. Statistics of interest are then computed using the residuals of the OLS estimates of (7) and (8) whereby the original variables are replaced with the transformed variables.

The third version is based on the difference of estimated parameters across sub-samples weighted by the variance-covariance matrix of the estimated parameters, rather than on the residual sum of squares. The test statistic is:

$$Wald = \left(\hat{\theta}_1 - \hat{\theta}_2\right)' \left(\hat{V}_1 + \hat{V}_2\right)^{-1} \left(\hat{\theta}_1 - \hat{\theta}_2\right), \tag{11}$$

where V_i are Newey-West corrected estimates of the variance-covariance matrix of the parameters $\hat{\theta}_i$ in regime $i = 1, 2$. Parameter stability tests based on the RSS such as (10) and on the differences of estimated parameters such as (11) give the same results when the restrictions on the parameters of the model are linear (see Hamilton, 1994, Ch. 8). The Newey-West correction for heteroskedasticity and autocorrelation, however, has no influence on the RSS, implying that only an expression such as (11) is suited for dealing with heteroskedasticity.

Table 5 reports the results of the three versions of the parameter stability and superexogeneity tests of 5% empirical size. The size-corrected power of the tests, i.e. the probability of rejecting the null hypothesis when it is false using a empirical 5% significance level, are

[13] See Favero and Hendry (1992) for the strategy behind this test.

computed using the simulation strategy described in Section 3.3. According to the standard Chow test (labeled 'OLS-based') there is little evidence of parameter instability in the output and inflation equations with probability that are never larger than 0.10. Yet, the ability of the Chow test to detect instability in the interest rate equation suggests that this test does not suffer any particular small-sample problem, at least as far as the policy rule is concerned.

The power of the test to detect parameter instability increases noticeably for our second and third versions which take the instability of the reduced-form innovations variance into account. When a GLS-based correction is used, the probability of (correctly) rejecting the null hypothesis increases by a factor of five in the case of the output equation, and four times for inflation. While the power of the test for the inflation specification is good, it is still less than satisfactory in the case of output. This need not to be surprising per se for two reasons. First, there may still be a small-sample issue. We investigate this possibility further below. Second, it may very well be the case that the effect of the policy parameter change on the behavior of the output specification is statistically small. In other words, the policy intervention might be modest in the Leeper-Zha sense. Additionally, the estimation of a very similar model in Lubik and Schorfheide (2004) reveals that output over the sample period (and conditional on the structural model) is almost exclusively driven by technology shocks, and that the feedback from the policy equation is minor.

The simulation evidence on the performance of the superexogeneity tests corroborate the notion of a bias induced by neglecting heteroskedasticity. Incidentally, the probabilities of rejecting the null of superexogeneity reported in Panel B of Table 5 are virtually identical.[14] We conclude that the test for parameter stability using either a GLS-based or a Newey-West correction for the presence of heteroskedasticity always have higher power than the tests based on OLS.

5.3 Sensitivity Analysis

We assess the robustness of our conclusion along three dimensions. We first modify our baseline calibration with respect to a few key parameters, while maintaining the regime shift from indeterminacy to determinacy. Secondly, we keep the baseline specification for

[14] These results are consistent with the findings in Toyoda and Ohtani (1986), who show that neglecting the change in disturbance variances when testing for parameter stability makes the actual probability of a type I error smaller than the chosen significance level.

the structural parameters, but simulate a policy shift between determinate regimes. Finally, we address the small-sample problems emphasized by Lindé (2001).

The results for variations of our baseline calibration are reported in Table 6. We consider three cases. The simulation using a smaller Phillips-curve slope coefficient $\kappa = 0.2$ reinforces the neglected heteroskedasticity bias as the power of the tests on the stability of the output and inflation equations become 0.18 and 0.60, respectively. Similar conclusions are reached on the basis of the superexogeneity tests in the last two columns or using even smaller values of κ. When aggregate demand is less sensitive to interest rate movements, $\tau^{-1} = 3$, the GLS-based probabilities of rejecting the null for the output and inflation equations are three times as large as those neglecting the instability of the reduced-form innovations variance in the column OLS-based. Lastly, to appreciate fully the effect of a shift from indeterminacy to determinacy, the bottom panel reports the results based on the value of 0.2 for σ_ζ, the standard deviation of sunspot shocks estimated by Lubik and Schorfheide (2004). This modification does not vary the baseline results much and confirms the conclusion reached by Castelnuovo and Surico (2006) that the passive monetary policy regime influenced U.S. aggregate fluctuations through a change in the transmission mechanism rather than through sunspot shocks.

The second robustness check analyzes to what extent our results are driven by the change in the equilibrium properties of the model. As demonstrated in Section 2, the policy-induced shift from indeterminacy to determinacy reduces both persistence and volatility of the endogenous variables. Policy shifts that maintain a determinate equilibrium, on the other hand, have a less dramatic effect on the reduced-form error variance.[15] Our experiment is to change the inflation coefficient from 1.5 to 2.5. The other structural and policy parameter are as in Table 1. Two observations emerge: first, the within-regime shift is less successful in replicating the stylized facts, in particular the decline in output volatility and persistence. While obviously far from conclusive, we regard this as a simple plausibility check on the driving forces behind the Great Moderation. Secondly, tests for error variance constancy, reported in Table 8, show that the null of homoskedasticity would have to be accepted at typical confidence levels with the exception of inflation. Differences between variances are in any case noticeably smaller than in the previous case. Not surprisingly, this

[15] Rudebusch (2005) rules out the possibility of a shift from indeterminacy to determinacy by adjusting the estimated inflation response coefficient upwards. This is done out of a concern for avoiding instability in the equation system. However, the issue with indeterminacy is that there is not enough instability in the system since there are infinitely many *stable* adjustment paths.

affects the improvement in power of the heteroskedasticity-adjusted test statistics in Table 9. The size-adjusted probabilities of rejecting parameter stability and superexogeneity are now more similar. This is clearly one possible explanation why the previous literature has not found any evidence of Lucas critique effects in the data.

A second possible reason are small-sample problems which Lindé (2001) documented in the case of the superexogeneity test. We thus return to our baseline calibration and simulate a shift from indeterminacy to determinacy, but this time impose a sub-sample size of 150 observations, twice as many as in the baseline. The results can be found in Table 10. Small-sample problems are clearly present for all three versions of the test for parameter stability. Interestingly, a doubling of sample size has only little effect on the probabilities of rejecting the null hypothesis for the output equation, while the power of heteroskedasticity-adjusted statistics improve by more than that of the simple OLS-based statistics. A similar conclusion can be drawn for the superexogeneity test.

6 Detecting the Lucas Critique in Practice: The Volcker Policy Shift

We conclude our analysis by an application to historical data. In fact, many contributions have shown that postwar U.S. data are characterized by a substantial amount of heteroskedasticity, for instance, Stock and Watson (2002), Ahmed, Levin and Wilson (2004), Kim, Nelson and Piger (2004), and Cogley and Sargent (2005). The conceptual background for our simulation analysis is the break in U.S. monetary policy behavior in the early 1980s, characterized by the tenure of Paul Volcker. We are interested in whether parameter stability tests adjusted for the presence of heteroskedasticity can detect this shift in reduced-form representations.

The full sample spans the period 1960:1 - 2005:3. Quarterly data were collected from the FRED database at the Federal Reserve Bank of St. Louis. We use data on output growth, measured as the quarter-to-quarter change in real GDP; inflation is the quarter-to-quarter change in the GDP deflator; and the federal funds rate as policy instrument.

Figure 1 reports the statistics of two recursive tests for parameter stability of the reduced-form model (7)-(9). The minimum length of a sub-sample is eight years, implying that in the first recursion Period I ends in 1968:4, whereas in the last recursion Period II begins in 1998:1. The left column reports results of a test based on the Chow statistics

(10) whereas the right column refers to Wald statistics of the form (11), which correct for heteroskedasticity using the Newey-West estimator of the variance-covariance matrix of the parameters. The dashed horizontal line represents the 5% critical value computed by Andrews (1993) for tests of parameter instability with unknown change point within the middle 70 percent of the sample. For the sake of comparison, the critical values for the same tests, but with known break date are reported as dotted horizontal lines.[16] The null hypothesis of parameter stability is rejected if the maximum value of the statistics is above the critical value.

In analogy with the Monte Carlo simulations in Section 5, accounting for heteroskedasticity dramatically changes the conclusions of the stability tests. The *sup* of the Wald statistics in the right column are clearly above the 5% critical values for all equations. According to these statistics the point estimates of the break dates are 1980:2 for the output and interest rate equations, and 1980:1 for the inflation equation.[17] The recursive Chow tests in the left column uniformly fail to reject the null hypothesis of parameter stability. The only possible exception is the interest rate equation conditional on a priori knowledge that a break occurred in 1980:2. However, even if the econometrician were equipped with information, say, on the basis of the empirical literature on monetary policy rules (Clarida, Galí and Gertler, 2000), the statistics of the Chow tests would not detect any structural change in the output and inflation equations, being far below the relevant 5% critical values. As the decline in the error variance of the equations for output, inflation and interest rate has been widely documented in the literature on U.S. data, we conclude that the neglect of heteroskedasticity accounts for the lack of support received by the Lucas critique using the Chow and super-exogeneity tests.

7 Conclusions

We present two main arguments in this paper. First, we emphasize that within the framework of structural DSGE-models a change in policy parameters affects both reduced-form coefficients and the variance of the errors in these equations. We deduce from this that tests for parameter instability across sub-samples that are based on the assumption of ho-

[16] Andrews and Fair (1988) show that under the null hypothesis of parameter stability, $\hat{\theta}_1 = \hat{\theta}_2$, the Wald statistic is asymptotically distributed as a χ^2 random variable with k degrees of freedom.

[17] Using the minimum value of the sum of RSS_1 and RSS_2 for the two sub-samples, we select the same break dates for output and interest rate whereas the break for inflation is now dated $1980Q3$.

moskedasticity have low power. We consequently argue that this is behind the tendency in the literature to reject the empirical relevance of the Lucas critique. We suggest adjustments for heteroskedasticity in commonly used parameter stability tests, and show by means of a Monte Carlo analysis that the power of these tests is improved and Lucas critique effects are detected in simulated data.

As an empirical example we test for the presence of Lucas critique effects on reduced-form specifications of output and inflation using post-war U.S. data. We do, indeed, find evidence of a break in the behavior of the Federal Reserve in 1980, and of parameter instability across the sub-samples using our heteroskedasticity-adjusted test statistic. We believe this is a new finding in the literature. Obviously, our results are model-dependent, and further investigation in a richer framework is certainly warranted.

We want to conclude on a somewhat critical note. Given a fully-specified, structural DSGE-model as a DGP, the question arises why researchers should bother with reduced-form specifications that are subject to the Lucas critique. One answer is certainly ease of implementation. Furthermore, only very little is known about structural break tests within the context of estimated DSGE-models.[18] However, a deeper issue is whether DSGE-models that are used for policy analysis are not themselves subject to the Lucas critique. Implicitly, Lucas' argument rests on the notion that the information set of economic agents and their decision problems were not fully specified in traditional macroeconometric models. Yet, with the use of *ad hoc* monetary policy rules that same issue surely comes up in DSGE-models that do not include optimizing policy-makers.

[18] A notable exception is Ireland (2001). Initial steps in this direction have been also made by Canova (2005) and Justiniano and Primiceri (2006).

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Table 1: Model Parameters

Panel A: Structural Parameters

Subsample	β	κ	τ^{-1}	ρ_g	ρ_z	σ_g	σ_z
I and II	0.99	0.7	1.4	0.7	0.8	0.3	1.1

Panel B: Monetary Policy Rule

Subsample	ψ_π	ψ_y	ρ_R	σ_R
I: Indeterminacy	0.4	0.2	0.6	0.3
II: Determinacy	1.5	0.5	0.8	0.3

Table 2: Good Policy Hypothesis: Descriptive Statistics

Panel A: Standard Deviations

Equation	I: Indeterminacy	II: Determinacy	% Change
Output	2.52	1.47	-41.58
Inflation	2.17	0.72	-66.91
Interest rate	0.81	0.47	-41.85

Panel B: Persistence

Equation	I: Indeterminacy	II: Determinacy	% Change
Output	0.69	0.75	8.96
Inflation	0.69	0.29	-58.07
Interest rate	0.85	0.77	- 9.33

Table 3: Good Luck Hypothesis: Descriptive Statistics

Panel A: Standard Deviations

Equation	I: High Variance	II: Low Variance	%Change
Output	1.50	0.80	-46.57
Inflation	0.72	0.26	-64.65
Interest rate	0.48	0.21	-57.65

Panel B: Persistence

Equation	I: High Variance	II: Low Variance	%change
Output	0.78	0.83	6.50
Inflation	0.30	0.33	9.27
Interest rate	0.76	0.80	5.08

Table 4: Innovation Standard Errors: Descriptive Statistics and Stability Tests

Equation	I: Indeterminacy	II: Determinacy	Statistics (p-value)
Output	1.79	0.88	4.11 (0.00)
Inflation	1.55	0.65	5.64 (0.00)
Interest rate	0.25	0.23	1.22 (0.23)

Table 5: Parameter Stability Tests

Panel A: Size-Adjusted Probability of Rejecting Parameter Stability

Equation	OLS-based	GLS-based	Newey-West corrected
Output	0.03	0.17	0.11
Inflation	0.09	0.45	0.38
Interest rate	0.91	0.94	0.92

Panel B: Size-Adjusted Probability of Rejecting Super-Exogeneity

Equation	OLS-based	GLS-based	Newey-West corrected
Output	0.02	0.11	0.11
Inflation	0.09	0.46	0.37

Table 6: Parameter Stability Tests
-Sensitivity Analysis-

	Parameter Stability		Superexogeneity	
	OLS-based	GLS-based	OLS-based	GLS-based
$\kappa = 0.2$				
Output	0.03	0.18	0.03	0.13
Inflation	0.08	0.60	0.22	0.61
Interest Rate	0.30	0.44	-	-
$\tau^{-1} = 3$				
Output	0.05	0.20	0.04	0.14
Inflation	0.18	0.53	0.16	0.54
Interest Rate	0.97	0.99	-	-
$\sigma_\zeta = 0.2$				
Output	0.03	0.16	0.03	0.11
Inflation	0.08	0.45	0.08	0.45
Interest Rate	0.84	0.89	-	-

Table 7: Good Policy Hypothesis: Descriptive Statistics
- from Determinacy to Determinacy -

Panel A: Standard Deviations

Equation	I: Determinacy	II: Determinacy	% Change
Output	1.49	1.50	0.35
Inflation	0.69	0.52	-24.57
Interest rate	0.73	0.50	-31.75

Panel B: Persistence

Equation	I: Determinacy	II: Determinacy	% Change
Output	0.77	0.75	1.98
Inflation	0.47	0.25	-45.66
Interest rate	0.78	0.77	- 1.38

Table 8: Innovation Standard Errors: Descriptive Statistics and Stability Tests
- from Determinacy to Determinacy -

Equation	I: Determinacy	II: Determinacy	Statistics (p-value)
Output	0.89	0.90	0.97 (0.55)
Inflation	0.59	0.48	1.55 (0.05)
Interest rate	0.26	0.23	1.28 (0.18)

Table 9: Parameter Stability Tests
- from Determinacy to Determinacy -

Panel A: Size-Adjusted Probability of Rejecting Parameter Stability

Equation	OLS-based	GLS-based	Newey-West corrected
Output	0.06	0.16	0.10
Inflation	0.10	0.23	0.20
Interest rate	0.48	0.61	0.61

Panel B: Size-Adjusted Probability of Rejecting Super-Exogeneity

Equation	OLS-based	GLS-based	Newey-West corrected
Output	0.06	0.11	0.11
Inflation	0.21	0.23	0.20

Table 10: Parameter Stability Tests
-from Indeterminacy to Determinacy: sub-samples of 150 obs-

Panel A: Size-Adjusted Probability of Rejecting Parameter Stability

Equation	OLS-based	GLS-based	Newey-West corrected
Output	0.03	0.12	0.12
Inflation	0.19	0.78	0.75
Interest rate	0.99	0.99	0.99

Panel B: Size-Adjusted Probability of Rejecting Super-Exogeneity

Equation	OLS-based	GLS-based	Newey-West corrected
Output	0.03	0.11	0.09
Inflation	0.15	0.76	0.71

Figure 1: Andrews Test of Parameter Instability with Unknown Break Date