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A Small-Sample Study of the New-Keynesian Macro Model

This paper presents a small-sample study of the three-equation-three variable New-Keynesian macro model. While the point estimates imply that the Fed has been stabilizing inflation fluctuations since 1980, our econometric analysis suggests considerable uncertainty regarding the stance of the Fed against inflation. The canonical New-Keynesian macro model is strongly rejected by the likelihood ratio test, but we propose the direction in which it needs to be modified in order to fit the data.

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NEW-KEYNESIAN MODELS have become the benchmark of much of the recent monetary policy literature. A wide variety of this class of models has been formulated, solved, and estimated. One common feature of these works is the high importance attributed to the model's structural parameters in the dynamics of the macro variables. While a number of papers have estimated New-Keynesian models with different methodologies, surprisingly very little has been said about the small-sample properties of the structural parameters. The present paper covers this gap in the literature.

Virtually all New-Keynesian systems include a monetary policy rule. At the same time, many theoretical monetary policy studies show that a coefficient larger than 1

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in the interest rate response to inflation is required for policy optimality. As a result, estimates of monetary policy rules have received a lot of attention recently. Clarida, Galí, and Gertler (1999), for instance, estimate a policy rule across sample periods with U.S. data and conclude that the post-Volcker policy rule is consistent with an optimal monetary policy strategy, unlike its pre-Volcker counterpart. One difficulty faced by these studies, however, is that only a small number of observations is typically available in each subsample estimation.

In order to conduct more precise inference about the monetary policy and private sector parameters of the structural model, we perform a bootstrap exercise, which yields the empirical probability distribution of the structural parameters. Two main empirical findings emerge from this small-sample analysis. First, it shows that standard estimates of the monetary authority response to expected inflation are upwardly biased, implying that inference on the stance of the monetary authority based on standard asymptotic theory can be misleading. Second, the empirical distributions of both the Phillips curve parameter and the coefficient relating the output gap and the real interest rate in the IS equation are very different from their asymptotic distributions.

Our structural model is a linearized Rational Expectations model consisting of AS, IS, and monetary policy rule equations with endogenous persistence. The AS equation is a generalization of the Calvo (1983) pricing model. The IS equation can be derived through representative agent optimization with external habit persistence, as in Fuhrer (2000). The monetary policy rule in our model is the forward looking Taylor rule proposed by Clarida, Galí, and Gertler (2000). Our model, though parsimonious, is rich enough to capture the macro dynamics implied by recently developed New-Keynesian models.

We estimate the model by full information maximum likelihood (FIML).[1] Even though the original model is strongly rejected using the likelihood ratio (LR) test, our analysis shows that when the error terms of the model are allowed to be serially correlated, the model is only marginally rejected at the 5% level using the small-sample distribution of the LR test statistic. In contrast, allowing cross-correlation of the errors terms does not result in a nearly similar improvement.

Several authors have estimated New-Keynesian macro models. Smets and Wouters (2003) and Lubik and Schorfheide (2004) estimate different versions of this class of models using Bayesian techniques. McCallum and Nelson (1998) and Ireland (2001) obtain instrumental variable and maximum likelihood estimates, respectively. Finally, Rotemberg and Woodford (1998), Christiano, Eichenbaum, and Evans (2001), and Boivin and Giannoni (2003) estimate structural New-Keynesian models by minimizing a measure of distance between empirical VARs and their models. None of these studies, however, analyzes the small-sample properties of the full New-Keynesian model.[2]

1. In order to avoid the potential problem of parameter instability, we select a sample period, 1980:4Q-2000:1Q, which does not include the most likely structural break in all the reduced form parameters of the model. This choice is based on the sup-Wald statistic derived by Bai, Lumsdaine, and Stock (1998).

2. Fuhrer and Rudebusch (2004) have recently analyzed the small-sample properties of the parameters in the IS equation through a Monte-Carlo exercise.

The paper is organized as follows. Section 1 formulates the structural model, which we consider, and discusses the model solution. Section 2 describes our estimation procedure. Section 3 discusses the data and the selection of the sample period based on the sup-Wald break date test statistic. In Section 4 we present our empirical results. First we show the estimates of the structural model and implied solution. Then we perform a small-sample study of the structural model. Finally we carry out model diagnostics using the asymptotic and small-sample LR tests. Section 5 concludes.

1. MODEL AND SOLUTION

Our structural model contains three equations: the AS or supply equation, the IS or demand equation, and a monetary policy rule. As Woodford (2003) shows, this set of equations can be formulated with explicit micro-foundations as a general equilibrium model. Each of the equations exhibits endogenous persistence, which allows for more realistic macro dynamics, and a forward-looking part. We assume that there is no informational difference between the private sector (firms and households) and the Central Bank.

1.1 A New-Keynesian Macro Model

The aggregate supply equation is a generalization of the supply specification originally developed by Calvo (1983):

$$\pi_t = \delta E_t \pi_{t+1} + (1 - \delta)\pi_{t-1} + \lambda y_t + \epsilon_{AS,t} \tag{1}$$

π_t is inflation between $t - 1$ and t, and y_t stands for the output gap between $t - 1$ and t. $\varepsilon_{AS,t}$ is the aggregate supply structural shock, assumed to be independently and identically distributed with homoskedastic variance σ^2_{AS}. It can be interpreted as a cost-push shock which makes real wages deviate from their equilibrium value or simply as a pricing error. E_t is the Rational Expectations operator conditional on the information set at time t, which comprises π_t, y_t, r_t (the nominal interest rate at time t), and all the lags of these variables. λ is the Phillips curve parameter. We assume a constant real wage markup, so that the output gap is proportional to the marginal cost, the original variable in the Calvo (1983) model. As Galí and Gertler (1999) and Woodford (2003) make clear, the endogenous persistence arises due to the existence of price setters who do not adjust optimally and index their prices with respect to past inflation.

The IS or demand equation is based on representative agent intertemporal utility maximization with external habit persistence, as proposed by Fuhrer (2000):

$$y_t = \mu E_t y_{t+1} + (1 - \mu)y_{t-1} - \phi(r_t - E_t \pi_{t+1}) + \epsilon_{IS,t} \tag{2}$$

where $\epsilon_{IS,t}$ is the IS or demand shock, assumed to be independently and identically distributed with homoskedastic variance σ^2_{IS}. In our specification, it is the habit formation specification in the utility function which imparts endogenous persistence to the output gap. The forward-looking parameter, μ, depends inversely on the level of habit persistence. The monetary policy channel in the IS equation is captured by the contemporaneous output gap dependence on the ex ante real rate of interest. Finally, the monetary transmission mechanism depends negatively on the curvature parameter in the utility function.

We close the model with the monetary policy rule formulated by Clarida, Galí, and Gertler (2000):

$$r_t = \alpha_{MP} + \rho r_{t-1} + (1 - \rho)[\beta E_t \pi_{t+1} + \gamma y_t] + \epsilon_{MP,t} \tag{3}$$

α_{MP} is a constant, and $\epsilon_{MP,t}$ is the monetary policy shock, assumed to be independently and identically distributed with homoskedastic variance σ^2_{MP}. The policy rule exhibits interest rate smoothing, placing a weight of ρ on the past interest rate. The Fed reacts to high-expected inflation and to deviations of output from its trend. The parameter β measures the long run response of the Central Bank to expected inflation, whereas γ describes its reaction to output gap fluctuations. We assume that the Federal funds rate is the monetary policy instrument, as much of the previous literature does.

1.2 Equilibrium

In this section, we follow the framework laid out in Cho and Moreno (2002) to derive the Rational Expectations equilibrium of the model. Our macroeconomic system of equations (1)–(3) can be expressed in matrix form as follows:

$$B_{11} X_t = \alpha + A_{11} E_t X_{t+1} + B_{12} X_{t-1} + \epsilon_t, \quad \epsilon_t \sim (0,D) \tag{4}$$

where $X_t = (\pi_t \; y_t \; r_t)'$, B_{11}, A_{11}, and B_{12} are the coefficient matrices of structural parameters, and α is a vector of constants. ϵ_t is the vector of structural errors, D is the diagonal structural error variance matrix, and 0 denotes a 3×1 vector of zeros. The Rational Expectations equilibrium to the system in Equation (4) can be expressed as:

$$X_{t+1} = c + \Omega X_t + \Gamma \epsilon_{t+1} \tag{5}$$

where c is a 3×1 vector of constants, and Ω and Γ are 3×3 matrices. The implied reduced-form of our structural model is thus a VAR of order 1 with highly nonlinear parameter restrictions.

The matrices Ω and Γ can be computed numerically using the generalized Schur matrix decomposition method (QZ) developed in Klein (2000) and Sims (2001). One limitation of the QZ method is that it does not indicate what solution to choose in the presence of multiple equilibria. When indeterminacy of equilibrium arises, we employ the recursive method derived by Cho and Moreno (2002). Appendix A provides a summary of the recursive method.

2. FULL INFORMATION MAXIMUM LIKELIHOOD ESTIMATION

We estimate the structural parameters using FIML by assuming normality of the structural errors. Galí and Gertler (1999) and Clarida, Galí, and Gertler (2000) estimate separately some of the equations of the model that we study. It seems adequate to estimate the whole model jointly, given the simultaneity between the private sector and the Central Bank behavior, as explained by Leeper and Zha (2000).

The log likelihood function can be written as:

$$\ln L(\theta \,|\, \bar{X}_T, \bar{X}_{T-1}, \ldots, \bar{X}_1) = \sum_{t=2}^{T} \Big[-\frac{3}{2} \ln 2\pi - \frac{1}{2} \ln |\Sigma| \\ - \frac{1}{2} (\bar{X}_t - \Omega \bar{X}_{t-1})' \Sigma^{-1} (\bar{X}_t - \Omega \bar{X}_{t-1}) \Big] \quad (6)$$

where θ is the vector of the structural parameters $\bar{X} = X_t - E(X)$ and $\Sigma = \Gamma D \Gamma'$. We check whether there is a unique, real-valued stationary solution at each iteration. Whenever there are multiple solutions at the i-th iteration, we apply the recursive method to select one solution. We choose the initial parameters from the values used in the literature. We found that the estimates obtained through our recursive method converge to the c, Ω, and Γ matrices obtained through the QZ method.

3. DATA DESCRIPTION AND SAMPLE SELECTION

We estimate the model with U.S. quarterly data from 1980:4Q to 2000:1Q. Implicit GDP deflator data are used for inflation. The inflation rate is computed as the log difference of the GDP deflator between the end and the beginning of each quarter. The Federal funds rate is the monetary policy instrument: we use the average of the Federal funds rate over the previous quarter. Our results are by and large robust to the use of the Consumer Price Index (CPI) for inflation and the three-month T-Bill rate for the short-term interest rate. We use three different measures for the output gap: output detrended with the Congressional Budget Office (CBO) Measure of Potential GDP, linearly detrended real GDP, and quadratically detrended real GDP.[3] The data are annualized and in percentages. Federal funds rate data were collected from the Board of Governors of the Federal Reserve website. Real GDP and the GDP deflator were obtained from the National Income and Product Accounts (NIPA).

Clarida, Galí, and Gertler (1999), Boivin and Watson (1999), and others have shown evidence of parameter instability across sample periods. We select our sample period based on the sup-Wald statistic for parameter instability, derived by Bai, Lumsdaine, and Stock (1998). This statistic detects the most likely date for a

3. The Hodrick-Prescott filter, linear filter, quadratic filter, and the CBO Measure of Potential GDP have been used extensively in the literature. There seems to be no consensus about the choice of filter to generate the output gap, since all of them seem to contain some measurement error.

TABLE 1

SUP-WALD BREAK DATE STATISTICS

Sample Period	VAR	Sup-Wald	Break Date	90% Confidence Interval
1954:3Q-2000:1Q	1	72.02	1980:4Q	1980:3Q-1981:1Q
1954:3Q-2000:1Q	2	103.33	1980:4Q	1980:3Q-1981:1Q
1954:3Q-2000:1Q	3	116.86	1980:4Q	1980:3Q-1981:1Q

NOTES: This table lists the sup-Wald values of the break date test derived by Bai, Lumsdaine, and Stock 1998). The test detects the most likely break date of a break in all of the parameters of unconstrained VARs of orders 1–3. The table shows the results of the test using the GDP deflator, linearly detrended output gap, and the Federal funds rate.

break in all the parameters of a reduced form VAR. We run the sup-Wald statistic for unconstrained VARs of orders 1–3. As shown in Table 1, the beginning of the fourth quarter of 1980, one year after Paul Volcker's beginning of his tenure as Federal Reserve chairman, is clearly identified as the most likely break date for the parameters of the reduced form relation. In all three cases, the value of the Sup-Wald statistic is significant at the 1% level,[4] and the 90% confidence interval is very tight, including only three quarters. The break date test is also robust across output gap measures. This date coincides with the biggest increase, between two quarters, in the average Federal funds rate during the whole sample: from 9.83% on the third quarter of 1980 to 15.85% on the fourth. This severe contraction engineered by the Federal Reserve lies at the root of the early 1980s disinflation.[5] We start the sample right after the break date occurs.[6]

4. EMPIRICAL RESULTS

In this section, we present our empirical findings. First, we report the structural parameter estimates and derive the Rational Expectations model solution. Then we provide the parameters' small sample distributions based on a bootstrap exercise. Finally, we perform specification tests of the structural model based on the asymptotic and small-sample LR test statistic.

4.1 Parameter estimates

FIML estimates are shown in Table 2. Asymptotic standard errors are obtained as the inverse of the Hessian Matrix. We present three sets of estimates in columns (1), (2), and (3): the first one is obtained using linearly detrended output, the second one uses quadratically detrended output, and the third one uses output detrended with the CBO measure of potential output. As is clear from Table 2, the estimates are reasonably robust across output gap specifications.

4. The associated asymptotic critical values can be found in Bakaert, Harvey, and Lumsdaine (2002).

5. Right after Volcker's arrival, the Federal Reserve also increased the Federal funds rate sharply, but it was decreased shortly thereafter. Feldstein (1994) dubs this episode the unsuccessful disinflation.

6. Empirical results are similar if we start the sample outside the 90% confidence interval.

TABLE 2

FIML Estimates and Small Sample Distribution of the Structural Parameters of the Model

Parameters	(1)	(2)	(3)	(4)	(5)	(6)
δ	0.5586 (0.0167)	0.5585 (0.0168)	0.5681 (0.0275)	[0.5257 0.5914]	0.5764	[0.5239 0.6565]
λ	0.0011 (0.0011)	0.0011 (0.0010)	−0.0002 (0.0015)	[−0.0010 0.0032]	0.0028	[−0.0034 0.0131]
μ	0.4859 (0.0327)	0.4810 (0.0350)	0.4801 (0.0338)	[0.4218 0.5500]	0.4826	[0.2386 0.5728]
ϕ	0.0045 (0.0046)	0.0054 (0.0054)	0.0065 (0.0057)	[−0.0044 0.0135]	0.0140	[−0.0065 0.0760]
ρ	0.8458 (0.0411)	0.8419 (0.0407)	0.8767 (0.0479)	[0.7652 0.9264]	0.8148	[0.6629 0.9211]
β	1.6409 (0.4630)	1.6413 (0.4453)	2.1506 (0.5705)	[0.7334 2.5483]	1.9027	[0.3983 5.0267]
γ	0.6038 (0.3243)	0.6126 (0.3019)	1.0079 (0.5957)	[−0.0319 1.2395]	0.6214	[−0.3402 1.7680]
σ_{AS}	0.4585 (0.0377)	0.4588 (0.0371)	0.4661 (0.0396)	[0.3846 0.5324]	0.4635	[0.3956 0.5344]
σ_{IS}	0.3734 (0.0332)	0.3766 (0.0344)	0.3570 (0.0337)	[0.3083 0.4384]	0.3841	[0.2996 0.5553]
σ_{MP}	0.7327 (0.0591)	0.7305 (0.0596)	0.7281 (0.0577)	[0.6169 0.8486]	0.7105	[0.5399 0.8818]

Notes: This table shows the FIML parameter estimates of the structural macro model in Equation (4), using the GDP deflator, the output gap, and the Federal funds rate. Standard errors are in parentheses below the estimates. The parameter sets in columns (1), (2), and (3) correspond to the estimations with linearly detrended output, quadratically detrended output, and output detrended using the CBO measure of potential output, respectively. Column (4) shows the 95% confidence interval of the asymptotic parameter estimates. Column (5) shows the sample means of the 1000 bootstrap parameter estimates. Column (6) shows the 95% interval of the empirical distribution of the parameter estimates. These last three columns are based on the estimates in Equation (1). The sample period is 1980:4Q-2000:1Q. The model's equations in demeaned form are:

$$\pi_t = \delta E_t\pi_{t+1} + (1-\delta)\pi_{t-1} + \lambda y_t + \varepsilon_{AS,t}$$

$$y_t = \mu E_t y_{t+1} + (1-\mu)y_{t-1} - \phi(r_t - E_t\pi_{t+1}) + \varepsilon_{IS,t}$$

$$r_t = \rho r_{t-1} + (1-\rho)[\beta E_t\pi_{t+1} + \gamma y_t] + \epsilon_{MP,t}$$

The parameter estimates are by and large consistent with previous findings in the literature. In the AS equation, δ is significantly greater than 0.5, implying that agents place a larger weight on expected inflation than on past inflation. Galí and Gertler (1999) found similar estimates. The Philips Curve parameter, λ, has the right sign in two of the three specifications, but it is not statistically different from 0 in any of the three cases. Fuhrer and Moore (1995) and Ireland (2001) obtained estimates of similar magnitude using a similar pricing specifications. Rudebusch (2002) obtains larger and significant estimates of the Phillips curve parameter. His approach, however, differs from ours since he includes several lags of inflation in the AS equation. In the IS equation, μ is statistically indistinguishable from 0.5, implying that agents place similar weights on expected and past output gap. The estimates of the implied inverse of ϕ, the coefficient on the real rate in the IS equation, are around 0.005. This value is considerably smaller than the ones usually employed in calibration (see McCallum 2001), but similar to the ones found in MLE or GMM estimation of the linearized IS equation (see Estrella and Fuhrer, 1999, Smets, 2000, Nelson and Nikolov, 2002).

In the monetary policy equation, the smoothing parameter, ρ, is around 0.85, reflecting the persistence in the short-term interest rate. β, the coefficient on expected inflation, is larger than 1, but only significantly above unity at the 5% level when the output gap is detrended with the CBO measure of potential output. γ, the coefficient on output gap, is also positive and only significantly different from 0 in the specification which uses the CBO measure of potential output. While these estimates are similar to the ones found by Clarida, Galí, and Gertler (1999) for the same monetary policy rule, our standard errors are considerably larger.

4.2 Model solution

For the first two specifications, the sets of FIML estimates imply a unique stationary solution, as we describe in Appendix A. For the remainder of our discussion, we will focus on the parameter estimates obtained when output is linearly detrended since their signs are fully in agreement with the theoretical model. The estimates of the implied reduced form matrices, Ω and Γ, which drive the dynamics of the model, are:[7]

$$\begin{bmatrix} \pi_t \\ y_t \\ i_t \end{bmatrix} = \begin{bmatrix} 0.670 \\ 0.258 \\ 0.579 \end{bmatrix} + \begin{bmatrix} 0.782^* & 0.056 & -0.011 \\ -0.002 & 0.961^* & -0.031 \\ 0.154^* & 0.114^* & 0.838^* \end{bmatrix} \begin{bmatrix} \pi_{t-1} \\ y_{t-1} \\ i_{t-1} \end{bmatrix} + \begin{bmatrix} 1.772^* & 0.106 & -0.013 \\ -0.004 & 1.870^* & -0.037 \\ 0.350^* & 0.221^* & 0.991^* \end{bmatrix} \begin{bmatrix} \epsilon_{AS_t} \\ \epsilon_{IS_t} \\ \epsilon_{MP_t} \end{bmatrix} \quad (7)$$

Panels A and B of Table 3 show the autocorrelation and cross correlation patterns exhibited by the structural errors, respectively. Panels C and D report some diagnostic tests of the residuals. The diagnostic tests give mixed results. Even though the Jarque-Bera test cannot reject the hypothesis of normality for the AS and IS residuals, the Ljung-Box Q-statistic rejects the hypothesis that their first five autocorrelations are zero. Under the null of the model, there should not be significant autocorrelations or cross-correlations, but this is a very difficult test to pass, given our parsimonious VAR(1) specification. The cross correlations of the error terms reveal nonzero contemporaneous correlations among the structural shocks.

The top row of Figure 1 compares the one step ahead predicted values of the model with the actual values of inflation, the output gap, and the interest rate. The predicted values generated by the model track the real values very closely. The bottom row of Figure 1 graphs the structural errors of the model. The IS shocks exhibit some persistence, as reported in Panel A of Table 3. It can also be seen that the monetary policy shocks were of very small magnitude after 1983. This corroborates

7. The stars denote the parameters that are significantly different from zero at the 5% level. The standard errors can be calculated using delta-method. Even though Ω and Γ cannot be expressed analytically in terms of structural parameters, we can derive numerical derivatives of Ω and Γ with respect to the structural parameters.

TABLE 3

Residuals Diagnostic Tests

Panel A. Autocorrelations of the Structural Errors

Lag $= i$	$\epsilon_{AS,t}, \epsilon_{AS,t-i}$	$\epsilon_{IS,t}, \epsilon_{IS,t-i}$	$\epsilon_{MP,t}, \epsilon_{MP,t-i}$
1	−0.3213	0.3555	0.1138
2	−0.1596	0.3798	−0.3057
3	0.1894	0.1860	0.2251
4	0.1356	−0.0029	0.2055

Panel B. Contemporaneous Crosscorrelations of the Structural Errors

$\epsilon_{AS,t}, \epsilon_{IS,t}$	$\epsilon_{AS,t}, \epsilon_{MP,t}$	$\epsilon_{IS,t}, \epsilon_{MP,t}$
0.0736	−0.2306	0.3027

Panel C. Ljung-Box Q-statistics

Lag	*Q*(*AS*,*t*)	pval(*AS*,*t*)	*Q*(*IS*,*t*)	pval(*IS*,*t*)	*Q*(*MP*,*t*)	pval(*MP*,*t*)
1	5.6600	(0.0174)	10.1481	(0.0014)	0.9574	(0.3278)
2	8.6191	(0.0134)	22.1299	(0.0000)	8.1270	(0.0172)
3	11.6441	(0.0087)	24.4316	(0.0000)	11.8207	(0.0080)
4	13.0648	(0.0110)	24.4745	(0.0001)	15.0232	(0.0047)

Panel D. Jarque-Bera Tests

	$JB(\epsilon_{AS,t})$	pval($\epsilon_{AS,t}$)	$JB(\epsilon_{IS,t})$	pval($\epsilon_{IS,t}$)	$JB(\epsilon_{MP,t})$	pval($\epsilon_{MP,t}$)
	3.6277	(0.1630)	5.0576	(0.0798)	55.5700	(0.0000)

Notes: Panel A reports the serial correlation of the AS, IS, and monetary policy shocks. Panel B lists the contemporaneous cross-correlations among the structural shocks. Panel C shows the Ljung-Box Q-statistics for autocorrelation of the error terms, with their corresponding probability values. Panel D reports the Jarque-Bera tests for normality of the residuals, with their corresponding probability values.

the analysis in Taylor (1999) and Leeper and Zha (2000) showing that monetary policy shocks during the 1990s were small.

4.3 Small-Sample Distributions of the Structural Parameters

Because our sample is relatively short, inference based on asymptotic distribution may be misleading. In order to draw a more precise inference on the validity of the structural parameters, we perform a bootstrap analysis. We bootstrap 1000 samples under the null and re-estimate the structural model to obtain an empirical probability distribution of the structural parameters. Appendix B details the bootstrap procedure. In the last two columns of Table 2, we report the small-sample means of the parameters and their associated 95% confidence intervals, respectively.

The coefficient on expected inflation in the monetary policy rule, β, appears significantly upwardly biased. Its small-sample 95% confidence interval includes 1 and is clearly wider than its asymptotic counterpart. Our finding implies then that inference on the stance of the monetary authority based on the asymptotic distribution can be misleading.

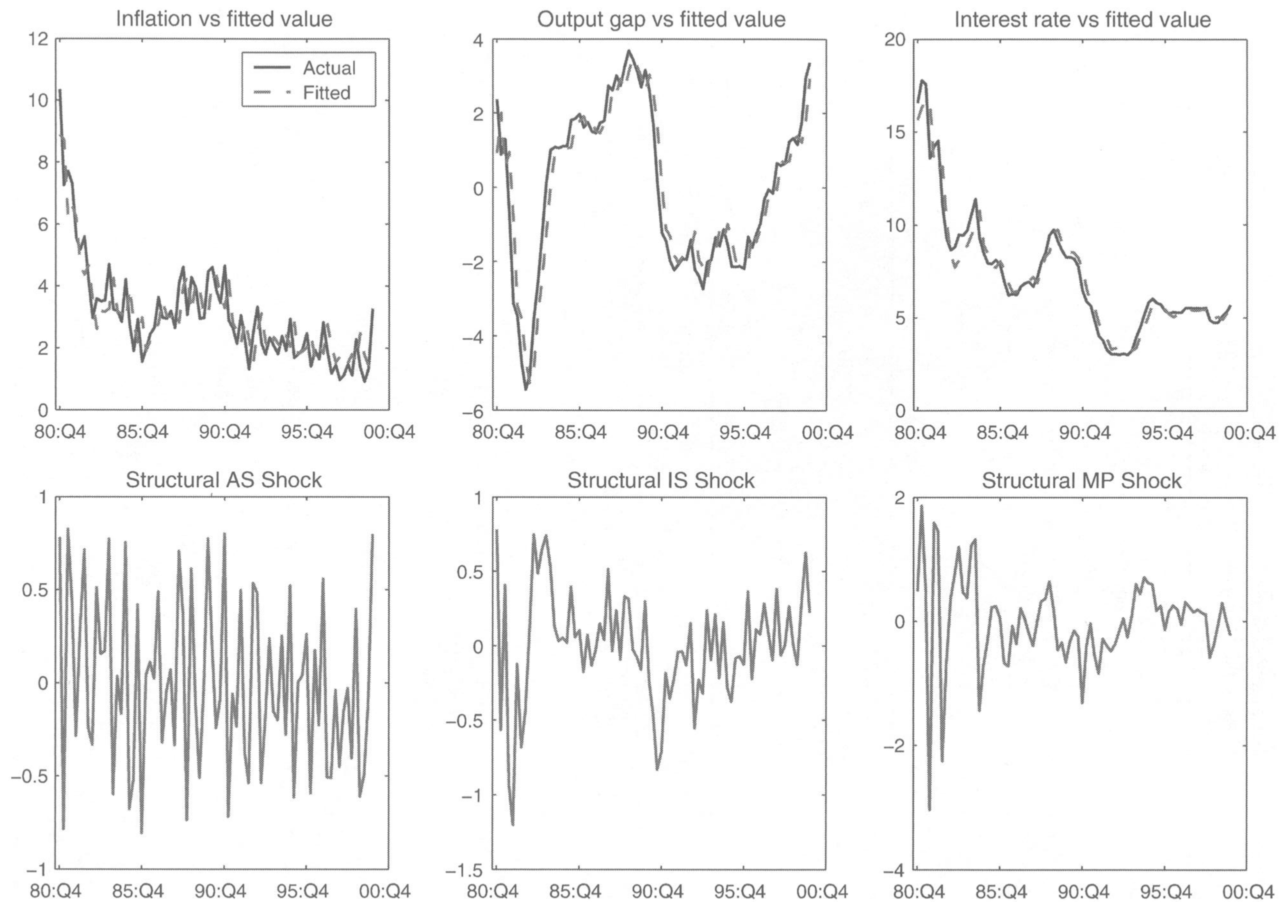

FIG. 1. Predicted values and structural shocks: 1980:Q4-2000:Q1. The top row of this figure shows the predicted (dashed lines) and actual (solid lines) values for inflation, the output gap, and the Federal funds rate associated with the FIML estimation of the structural model in Equation (5). The bottom row shows the structural errors estimates ($\epsilon_{AS,t}$, $\epsilon_{IS,t}$, and $\epsilon_{MP,t}$).

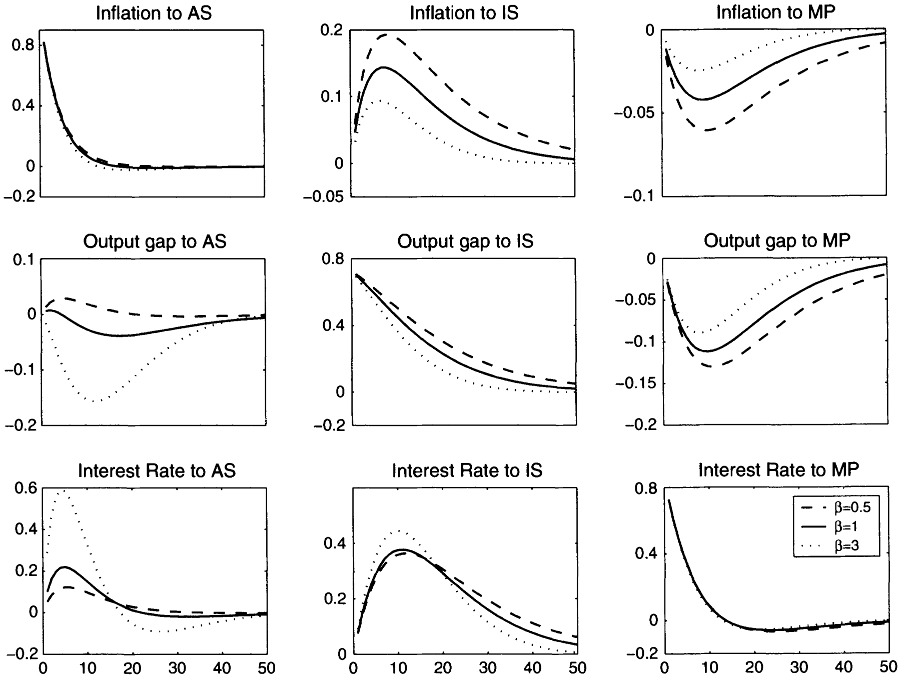

FIG. 2. Sensitivity analysis for different values of β This figure presents the impulse responses which arise under different values of β chosen from the 95% interval of their empirical distribution. The remaining parameter values are held fixed at their corresponding estimates, as shown in Table 2, column (1).

A larger uncertainty about β is transmitted into the dynamics of the macro variables following the structural shocks. Figure 2 shows the impulse response functions under different values of β, which belong to both its asymptotic and empirical distributions. As β increases, the Fed responds more aggressively to supply and demand shocks. As it is well known, the magnitude of β plays a pivotal role in the output gap response to the AS shock. When β is larger than 1 (the Taylor principle holds), the output gap decreases for a long time.[8] Therefore, a monetary policy which is very responsive to inflationary pressures may result, under an AS shock, in costly recessionary effects. A higher β also makes the private sector's responses to the monetary policy shock less pronounced. This is due to the fact that the contractionary policy shock lowers expected inflation below the steady state in the future. Larger values of β partially offset the impact of the monetary policy shock, since a stronger reaction from the Fed to lower expected inflation moves the interest rate in the opposite direction to the one implied by the shock. Conversely, if the Fed is not very responsive ($\beta = 0.5$), the impact of the policy shock is magnified.

8. In our standard New-Keynesian model, a $\beta > 1$ is required for monetary policy optimality. However, Christiano and Gust (1999), Benhabib, Schmitt-Grohe, and Uribe (2001), and Galí, López-Salido, and Vallí (2004) point out the limitations of the Taylor principle as a criterion of monetary policy optimality in different frameworks to the standard New-Keynesian model.

TABLE 4

Robustness Analysis: FIML Estimates and Small Sample Distribution of the Structural Parameters of the Model: 1979:4Q-2000:1Q

Parameters	(1)	(2)	(3)	(4)
δ	0.5241 (0.0185)	[0.4877 0.5604]	0.5362	[0.4805 0.5976]
λ	0.0023 (0.0021)	[−0.0018 0.0064]	0.0053	[−0.0027 0.0208]
μ	0.5026 (0.0194)	[0.4645 0.5407]	0.5050	[0.3682 0.5798]
ϕ	0.0058 (0.0044)	[−0.0029 0.0144]	0.0124	[−0.0033 0.0538]
ρ	0.7761 (0.0502)	[0.6777 0.8744]	0.7524	[0.6149 0.8618]
β	1.7012 (0.2784)	[1.1555 2.2469]	1.8164	[0.8854 3.0319]
γ	0.4007 (0.2830)	[−0.1540 0.9554]	0.3645	[−0.4941 1.3754]
σ_{AS}	0.5123 (0.0420)	[0.4301 0.5946]	0.5127	[0.4127 0.6204]
σ_{IS}	0.3951 (0.0324)	[0.3316 0.4587]	0.4011	[0.3216 0.5006]
σ_{MP}	0.9918 (0.0775)	[0.8399 1.1437]	0.9490	[0.7068 1.2212]

Notes: This table shows the FIML parameter estimates of the structural macro model in Equation (5), using the GDP deflator, linearly detrended output, and the Federal funds rate. Standard errors are in parenthesis below the estimates. Column (2) shows the 95% confidence interval of the asymptotic parameter estimates. Column (3) shows the average of the empirical probability distribution of the parameter estimates. Column (4) shows the 95% interval of the empirical distribution of the parameter estimates. The sample period is 1979:4Q-2000-1Q.

The empirical distributions of δ and ρ are mildly positively and negatively skewed, respectively. This bias is related to the well-known small-sample downward bias of the first order autocorrelation coefficients, as reported in Bekaert, Hodrick, and Marshall (1997). The most severe small-sample problems are the strong positive skewness exhibited by the empirical distribution of the Phillips curve parameter, λ, and that of the coefficient on the real interest rate in the IS equation, ϕ. This finding may be related to output gap mis-measurement, as these two parameters were not significantly different from zero in the FIML estimation. Finally, the averages of the empirical distributions of μ, γ, and those of the three structural shocks standard deviations are very similar to the FIML parameter estimates.

4.4 Robustness Analysis

Our motivation for choosing the baseline sample was to avoid structural breaks in the parameters. It is, however, interesting to analyze how results change depending on the sample chosen. In particular, more observations may result in a sharper inference on the parameter space, reducing the uncertainty reported above.

We first start the sample, on the fourth quarter, upon the arrival of Paul Volcker as chairman of the Fed. The ending period is still the first quarter of 2000. Table 4 shows that while the asymptotic results are similar to the baseline subsample, it is noteworthy that β, the long-run response of the interest rate to expected inflation, is now significantly larger than one. This would imply that the Fed behaved optimally according to most monetary policy studies. Analogously to the baseline sample, we perform a bootstrap exercise on this new sample. We find the same direction and a similar size in the biases of most of the structural parameters. Interestingly, β is again upwardly biased and not significantly different from one looking at the

TABLE 5

ROBUSTNESS ANALYSIS: FIML ESTIMATES AND SMALL SAMPLE DISTRIBUTION OF THE STRUCTURAL PARAMETERS OF THE MODEL: 1980:4Q-2004:4Q

Parameters	(1)	(2)	(3)	(4)
δ	0.5752 (0.0174)	[0.5411 0.6093]	0.5925	[0.5440 0.6636]
λ	0.0006 (0.0012)	[−0.0017 0.0030]	0.0012	[−0.0061 0.0097]
μ	0.5023 (0.0210)	[0.4611 0.5435]	0.4933	[0.1463 0.5728]
ϕ	0.0017 (0.0023)	[−0.0029 0.0063]	0.0075	[−0.0062 0.0619]
ρ	0.9023 (0.0303)	[0.8430 0.9617]	0.8697	[0.7423 0.9529]
β	1.6608 (0.6694)	[0.3488 2.9728]	1.8563	[−0.0638 5.5339]
γ	1.1366 (0.5174)	[0.1226 2.1507]	1.1622	[−0.0428 2.8632]
σ_{AS}	0.5077 (0.0394)	[0.4304 0.5849]	0.5137	[0.4378 0.5937]
σ_{IS}	0.3460 (0.0244)	[0.2982 0.3937]	0.3608	[0.2872 0.5840]
σ_{MP}	0.7085 (0.0496)	[0.6113 0.8057]	0.6860	[0.5318 0.8368]

NOTES: This table shows the FIML parameter estimates of the structural macro model in Equation (5), using the GDP deflator, linearly detrended output, and the Federal funds rate. Standard errors are in parenthesis below the estimates. Column (2) shows the 95% confidence interval of the asymptotic parameter estimates. Column (3) shows the average of the empirical probability distribution of the parameter estimates. Column (4) shows the 95% interval of the empirical distribution of the parameter estimates. The sample period is 1980:4Q-2004:4Q.

small sample. As a result, our original finding that β is not significantly larger than 1 looking at the small sample, is reinforced.[9]

We perform another robustness exercise adding more observations at the end of the sample, so that the sample now covers the period 1980:fourth quarter to 2004:fourthth quarter. Table 5 shows that both the asymptotic and the small-sample results are similar to the original subsample.

4.5 Model Specification

In this section, we examine, both asymptotically and at the small-sample level, how our estimated model fits the actual U.S. economy for our sample period with respect to an unrestricted model. First we study our original model. Then we analyze two augmented models, which incorporate autocorrelation and cross-correlation of the error terms.

Baseline model. Since our model is nested in a VAR(1) with highly nonlinear parameter restrictions, we compare the model with an unrestricted VAR(1).[10] The cross-equation restrictions implied by the New-Keynesian model are rejected by an LR test: we have seven parameters in the structural model and three variances of structural shocks. The unrestricted VAR(1) has nine parameters in the coefficient matrix and six in the variance covariance matrix of innovations. Therefore, there are five over-identification restrictions. The likelihood of our model and the unrestricted VAR are –259.975 and –243.360, respectively. This implies an LR test

9. Overall, we found that when we started the sample between the fourth quarter of 1979 and the third quarter of 1980, β was asymptotically larger than one in statistical terms. This may be related to the large increase of the Fed funds rate on the third quarter of 1980.

10. Even though the optimal number of lags chosen by the Schwarz criterion is three among the unrestricted VARs, it seems appropriate to compare our model with the nested VAR(1) for the purpose of our study. The impulse responses of an unrestricted VAR(3) are similar to those of the unrestricted VAR(1).

TABLE 6

EMPIRICAL SIZE AND POWER FOR THE LIKELIHOOD RATIO TEST

	Mean	Median	Std. Dev.	90%	95%	99%	Sample LR	pval
Panel A. Model with uncorrelated residuals								
$\chi^2(5)$	5	4.35	3.16	9.24	11.07	15.09	33.23	0.000
MODEL LR	6.83	5.94	4.54	12.43	15.48	22.61		0.000
SIZE(%)				23.0	15.5	5.2		
POWER(%)				95.6	91.4	73.1		
Panel B. Model with autocorrelated and cross-correlated residuals (non-diagonal F)								
$\chi^2(9)$	9	8.34	4.25	14.68	16.92	21.66	20.60	0.015
MODEL LR	9.88	9.03	4.89	16.26	19.83	25.37		0.039
SIZE(%)				14.1	8.9	3.1		
POWER(%)				79.4	64.4	37.8		
Panel C. Model with autocorrelated residuals (diagonal F)								
$\chi^2(11)$	11	10.37	4.71	17.28	19.68	24.72	47.33	0.000
MODEL LR	13.07	12.30	5.67	20.55	23.77	28.09		0.000
SIZE(%)				21.1	11.8	4.2		
Panel D. Model with $F = 0$ and cross-correlated residuals								
POWER(%)				97.4	92.8	85.2		
χ^2 (2)	2	1.38	2	4.61	5.99	9.21	11.99	0.002
MODEL LR	2.58	1.83	2.45	5.87	7.50	16.98		0.006
SIZE(%)				16.2	9.5	2.5		
POWER(%)				95.6	91.4	73.1		

NOTES: This table provides summary statistics for the asymptotic and empirical distributions of the likelihood ratio (LR) test statistic. The statistics are the mean, median, standard deviation (Std. Dev.), and the 90%, 95%, and 99% quantiles. MODEL LR refers to the empirical distribution of the LR statistic under the null hypothesis (restricted model). The table also provides empirical sizes and powers from the empirical distributions of the LR test statistic. The empirical size is the percent of the bootstrap experiments generated under the null hypothesis, where the test statistic exceeds a given asymptotic critical value. The empirical power of the test is the percent of the bootstrap experiments generated under the alternative hypothesis (unrestricted VAR), where the test statistic exceeds the given empirical critical value. The structural model in demeaned matrix form is: $X_t = \Omega X_{t-1} + \Gamma\epsilon_t$, where $\epsilon_t = F\epsilon_{t-1} + w_t$. Panels A, B, C, and D show the statistics for the original model ($F = 0$, errors uncorrelated), the model with cross-correlated and serially correlated structural errors (non-diagonal F), the model with serially correlated errors (F diagonal), and the model with cross-correlated errors ($F = 0$, errors cross-correlated), respectively.

statistic of 33.230, rejecting the null that the restricted model comes from the same asymptotic distribution as the unrestricted one.

As shown by Bekaert and Hodrick (2001) in the context of the Expectation Hypothesis, asymptotic tests such as the LR test can be severely biased in small samples. With the data generated by our bootstrap exercise, we re-estimate the structural model and the unconstrained VAR(1). This yields the small-sample distribution of the LR test statistic. As we report in the Panel A of Table 6, there is a considerable size distortion in the LR test of our model. For instance, the 5% critical value is 15.48, instead of the 11.07 asymptotic value, and the empirical size is 15.5%. The top panel of Figure 3 shows that the empirical distribution of the LR test statistic has a higher mean and a fatter tail than the asymptotic distribution. Unfortunately, the structural model is still strongly rejected. We also bootstrap 1000 samples under the alternative hypothesis of an unrestricted VAR(1) to calculate the empirical power

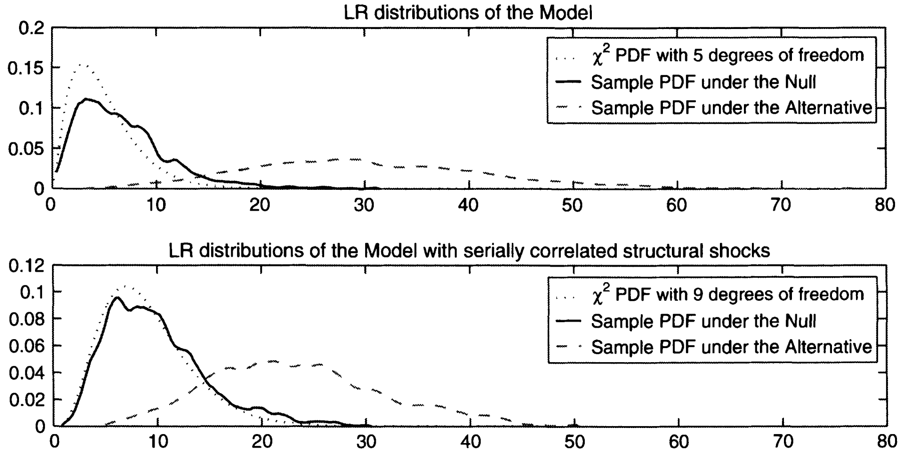

FIG. 3. Empirical distribution of the likelihood ratio. This figure compares the asymptotic probability density function (PDF) of the Likelihood Ratio test (dotted line) with its small sample counterpart (solid line) under the null of the structural model. It also graphs the Likelihood Ratio test under the alternative hypothesis, the unconstrained VAR(1) (dashed line). The PDF of the empirical LR test statistic is estimated using an Epanechnikov kernel. The top and bottom panels display the PDFs for the model with and without serially correlated structural errors, respectively.

of the LR test. The empirical power measures the probability of rejecting the null hypothesis when the alternative is true in a small sample. It is calculated as the percentage of LR tests obtained, under the alternative hypothesis, that are lower than a given empirical critical value. For a 5% significance level, the power of the test is 91.4%.

Extending the New-Keynesian model. The rejection of the baseline model is perhaps not very surprising given that the diagnostic tests had revealed significant autocorrelation and cross-correlation patterns in the error terms. In order to gauge how the New-Keynesian model could be improved, we first augment the model with autocorrelated residuals and analyze the associated LR test. Several authors have resorted to exogenous autocorrelation in order to produce realistic macro dynamics Rotemberg and Woodford (1999). In line with their work, we let the error terms follow a VAR(1) process:

$$\epsilon_{t+1} = F\epsilon_t + w_{t+1} \tag{8}$$

where F is a 3×3 stationary matrix that captures the structural shock serial correlation and w_{t+1} is independently and identically distributed with diagonal variance covariance matrix D. The reduced form solution of the model is still given by Equation (5). The same method of undetermined coefficients employed in Section 1.2 can be applied to solve for Ω, Γ, and c in terms of α, A_{11}, B_{11}, B_{12}, and F. It can be shown that the expressions for Ω and c are the same as the equations in the original model, and therefore the same methodology for solving the matrix quadratic form can be applied. However, Γ now depends on F:

$$\Gamma = (B_{11} - A_{11}\Omega)^{-1}(I + \mathrm{A}_{11}\Gamma F) \tag{9}$$

In order to estimate this model, we first express the model solution in terms of w_{t+1} as:

$$X_{t+1} = (I - \Gamma F \Gamma^{-1})c + (\Omega + \Gamma F \Gamma^{-1})X_t - \Gamma F \Gamma^{-1}\Omega X_{t-1} + \Gamma w_{t+1} \tag{10}$$

The dynamic effects of macro shocks could now be different from the baseline New-Keynesian model. As Equation (10) shows, the coefficient matrices of X_t, X_{t-1}, and Γ are also functions of F, so that the impulse response functions are not fully governed by the structural parameters. We first estimate the model by FIML without any restriction on F. Let F_{ij} be ij-th element of F. Then zero restrictions on F_{13}, F_{21}, F_{32}, and F_{33} are imposed because they are not significantly different from zero. Since the reduced form solution is VAR(2), a natural alternative is an unrestricted VAR(2). These four additional restrictions imply that the model has nine degrees of freedom in total.

Even though the asymptotic LR test still rejects the model at the 5% level, the rejection is marginal using the small-sample LR test (the p value is 0.039), as is shown in Panel B of Table 6 and the bottom panel of Figure 3. The empirical power of the test is much lower than the one in the original model: the power associated with empirical sizes of 5% and 1% are 64.4% and 37.8%, respectively. In contrast, the corresponding powers in the model without serial correlation are 91.4% and 73.1%. This evidence suggests that tests of models which imply restricted higher order VARs may suffer from low power against their unrestricted counterparts.

Our framework allows us to test other specifications which have been proposed in the literature for the autocorrelation of the error terms. Clarida, Galí, and Gertler (1999), for instance, let their structural errors follow AR(1) processes. In our setup, this is equivalent to a diagonal F matrix. Analogously to the case of the model with non-diagonal F, we estimate and test the model under the diagonal F specification. Panel C of Table 6 shows that this model is strongly rejected asymptotically and at the small sample.

We finally analyze an extension of the baseline New-Keynesian model where we allow contemporary cross-correlation in the error terms.[11] The model diagnostic tests did detect some cross-correlation among the disturbances. The reduced-form of this model is a VAR(1), but the variance-covariance matrix of the error terms differs from the baseline model and can be expressed as:

$$X_t = \Omega X_{t-1} + \Gamma \Psi w_t \tag{11}$$

where Ω and Γ are computed as in the original model. The matrix Ψ reflects the cross-correlation of the shocks and w_t is independently and identically distributed with diagonal variance covariance matrix D. An implication of this model solution is that a price puzzle could arise, because Ψ is estimated from the data. The monetary

11. Notice that the case serially autocorrelated error terms also implies contemporary cross-correlation in the errors as long as F is non-diagonal.

policy could thus increase inflation on impact, potentially affecting the subsequent dynamics of inflation.

Panel D of Table 6 shows the results of the LR test for the cross-correlated New-Keynesian model. Given that we estimate three more parameters than in the baseline model, we over-identify the model only by two parameters. The alternative model is again the unrestricted VAR(1). The model is rejected both asymptotically and at the small sample. Even though the rejection is less severe looking at the small sample, this model remains overall inconsistent with the data.

To summarize, adding persistence seems a fruitful strategy to improve the New-Keynesian model. This result highlights the need to produce different model specifications, in order to uncover the structural macro relations behind this significant autocorrelation of the residuals.

5. CONCLUSION

Policy parameters have qualitative and quantitative implications on the relation between macro dynamics and structural shocks. When the Fed reacts strongly to deviations of expected inflation from its target, two different effects take place: on the one hand, inflation returns faster to the target in response to AS and IS shocks. On the other hand, the economy enters into a longer recession in response to an AS shock. A number of authors have estimated a strong reaction of the Fed to deviations of expected inflation from the target since 1979. Our maximum likelihood estimation shows, however, that this result is not statistically significant using linearly and quadratically detrended output. Moreover, our small-sample study reveals that the coefficient on expected inflation is upwardly biased. One possibility is that the Taylor rule does not describe accurately the way the Fed conducts monetary policy and that the Fed reacts differently to the different shocks which buffet the economy.

Reconciling macro models with the data remains an important, if challenging, task for macroeconomists. This paper represents a step in this direction. We showed that adding persistence to a standard microfounded New-Keynesian model improves the fit of the model. Therefore, additional research efforts are needed to provide an economic interpretation to macro dynamics. Two examples of this line of research are Smets and Wouters (2003) and Bekaert, Cho, and Moreno (2005). They incorporate additional variables to the New-Keynesian model and show that the joint reduced-form for inflation, output, and the interest rate displays richer and more realistic dynamics.

APPENDIX A. UNIQUENESS OF THE SOLUTION

Whenever indeterminacy of equilibrium arises, we use the recursive method Cho and Moreno (2002) to pin down a solution. They solve the model forward recursively and propose a selection criterion which is stationary and real-valued by construction.

TABLE 7

GENERALIZED EIGENVALUES

Gen. Eig.	(1)	(2)	(3)
ξ_1	0.7845	0.7837	0.7608
ξ_2	0.8986 − 0.0348i	0.8973 − 0.0385i	0.9110 − 0.0593i
ξ_3	0.8986 + 0.0348i	0.8973 + 0.0385i	0.9110 + 0.0593i
ξ_4	1.0148	1.0148	0.9970
ξ_5	1.0987	1.1192	1.1419
ξ_6	∞	∞	∞

NOTES: This table reports the generalized eigenvalues which determine the stability of the structural macro model. The sets of eigenvalues in columns (1), (2), and (3) correspond to the estimations of the systems with linearly detrended output, quadratically detrended output, and output detrended using the CBO measure of potential output, respectively.

They postulate the existence of a unique vector of self-fulfilling expectations where agents coordinate in equilibrium. In practice, this equilibrium is reached by imposing a transversality condition where distant future expectations converge to their long run mean. The remaining expectations are discarded, since agents deem them incapable of being satisfied.

Table 7 shows the generalized eigenvalues associated with the three FIML estimated sets of parameters. As explained in Section 3.2, in the first two specifications (with output linearly and quadratically detrended), we have a unique solution, since there are exactly three eigenvalues less than unity, the same number as predetermined state variables in the model. We also verified that the recursive solution coincides with the one obtained through the QZ method.

For the third specification (with output detrended using the CBO measure), we have multiple solutions, since there are four eigenvalues less than one in moduli. Our recursive method converges to the QZ solution with the smallest three eigenvalues. In general, we found that, holding the remaining parameters at their estimated values in column (1) of Table 2, when λ is positive, the solution is unique. For negative values of λ, large in absolute value, there is no real valued solution. For small negative values of λ, as estimated with the CBO measure, there are multiple solutions.

APPENDIX B. BOOTSTRAP ANALYSIS

Our structural model and the unrestricted VAR(1) can be expressed respectively as:

$$X_t = c + \Omega X_{t-1} + \Gamma\epsilon_t \tag{12}$$

$$X_t = d + \Theta X_{t-1} + u_t \tag{13}$$

where $Var(\Gamma\epsilon_t) = \Gamma D\Gamma'$ and $Var(u_t) = \Upsilon$. If the structural model is true, it should be the case that $\Gamma D\Gamma' = \Upsilon$. We orthogonalize the unrestricted VAR(1) error terms through a Choleski decomposition, so that $Var(u_t) = E(u_t u'_t) = \Upsilon = CC'$, where C is lower triangular. Therefore, $u_t = C\zeta_t$, where ζ_t has mean zero and ones in the diagonal of its variance covariance matrix. The unrestricted VAR(1) can then be expressed as:

$$X_t = d + \Theta X_{t-1} + C\zeta_t \tag{14}$$

Under the null of the model $\epsilon_t = \sqrt{D}\xi_t$, where ξ_t has mean zero and ones in the diagonal of its variance covariance matrix. The model can then be expressed as:

$$X_t = c + \Omega X_{t-1} + \Gamma\sqrt{D}\xi_t \tag{15}$$

Therefore, if the model is true, it should be the case that $\Gamma\sqrt{D} = C$ and that $Var(\Gamma\sqrt{D}\xi_t) = Var(C\zeta_t)$. We perform a bootstrap analysis under the null of the structural model and under the alternative data generating process, the VAR(1). Under the null we proceed as follows:

1. We bootstrap the unconstrained errors, u_t, with replacement.
2. We reconstruct 1000 sample data sets of size 578 under the null hypothesis, using the estimated parameter matrices c, ς and D, and the historical initial values, along with the ζ_t disturbances, which are obtained by pre-multiplying the u_t errors by C^{-1}. For every sample we discard the first 500 data points and retain the last 78 observations to have the same size as the original data set.
3. We re-estimate both the model and the unrestricted VAR(1) 1000 times. This yields 1000 parameter sets and 1000 LR tests.

With the 1000 parameter sets, we obtain the small-sample distribution of the structural parameters under the null of the model. To compute the empirical critical values of the LR test statistic, we select the corresponding quantiles of the empirical distribution of the LR test statistic. The bootstrap simulations under the alternative hypothesis differ from the ones under the null in that, in step 2, the data sets are constructed conditional on d and Θ, instead of c, Ω, and D. The power of the test is calculated as the percentage of LR tests obtained, under the alternative hypothesis, which is lower than a given empirical significance level.

The case of the bootstrap of the model with autocorrelation, $F \neq 0$, is analogous to the one just presented. There are two differences with respect to the baseline case. First, the unconstrained residuals are bootstrapped from a VAR(2) model. Second, under the null hypothesis, Equation (10) is used to reconstruct the small-sample data sets.

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