An Estimated DSGE Model for the United Kingdom

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1. Introduction

In this paper we estimate a dynamic stochastic general equilibrium (DSGE) model with nominal rigidities for the U.K. economy. The model we estimate is due to Christiano, Eichenbaum, and Evans (2005) and has become a benchmark, matching important aspects of the U.S. data while also being derived from optimizing behavior.

Interest in DSGE modeling of the United Kingdom has been heightened in recent years with the introduction into the U.K. monetary policy process of the Bank of England Quarterly Model (BEQM), which is based to a considerable degree on explicit optimizing foundations; see Harrison, Nikolov, Quinn, Ramsay, Scott, and Thomas (2005) for the model, and Pagan (2005) for a discussion. Nevertheless, BEQM is dissimilar in important respects from the CEE model of the U.S. and the variant of the Christiano-Eichenbaum-Evans (CEE) model that Smets and Wouters (2003) estimate for the euro area, and these dissimilarities make it difficult to use BEQM to compare the structure of the U.K. economy with that of other economies. For example, the estimation procedure for BEQM is different from both that used by CEE or Smets-Wouters; portions of the BEQM model are estimated over a considerably shorter sample than CEE consider for the U.S.; and there are deviations from explicit optimization in the dynamics of the BEQM model.

All in all, it is probably fair to say that there has been considerably less work done for the U.K. in terms of DSGE modeling with systems estimation than there has been for the other economies. But U.K. data may contain information that is ideal for estimation of a DSGE model—namely, information on private sector responses to policy actions. As the present Governor of the Bank of England observed some 30 years ago,

> Maintenance of the existing order and existing rates produces no information, whereas more information can be obtained by making changes. In this respect the U.S.… is at a disadvantage by comparison with the U.K. A good illustration of this is afforded by the excitement generated amongst American economists in the 1960s by the investment tax credit and the attempts to assess its effects. A British economist would have shrugged this off as a mere trifle compared to the changes he had witnessed over the years. (King, 1977, p. 6).

This observation, though made with reference to the changes wrought in U.K. fiscal policy up to the 1960s, applies tenfold to monetary policy experience in the period since

the 1960s. Over that period the United Kingdom has undergone great variation in inflation, interest rates, and monetary regime.[1] It is true that for estimation this is a mixed blessing, because large regime changes make it problematic to estimate a structural model over a long sample. But Christiano, Eichenbaum, and Evans (1999) and Sims and Zha (2006) argue for the United States that constant-parameter policy reaction functions may be reasonable approximations over even long samples, a view also implicit in CEE's (2005) choice of a 1965–1995 estimation period. In modeling the U.K. using a DSGE model, we make a compromise between these positions by treating the period since 1979 as a single regime,[2] but also presenting results for pre-1979 and a long sample covering 1962–2005.

This paper proceeds as follows. Section 2 presents the model. Section 3 presents estimates for our main sample. Section 4 presents results for the longer sample and discusses other regime-change issues. Section 5 concludes.

2. Model

The model is the same as that in Christiano, Eichenbaum, and Evans (2005), by now standard in the DSGE literature. The model incorporates both nominal frictions (sticky prices and wages) and dynamics in preferences and production (habit formation in consumption, investment adjustment costs, and variable capital utilization). The pattern of timing in agents' decisions is consistent with the VAR identification restriction that we use in Section 3 below. In our outline here of the linearized version of the model, all variables are expressed in log-deviations from their steady-state values. For convenience, model parameters and variables are summarized in Tables 1 and 2, respectively.

Prices are governed by Calvo contracts, augmented by indexation to the previous period's inflation for those firms not allowed to reoptimize their pricing decision. The implied inflation dynamics are given by the following Phillips curve:

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[1] Various advantages of the U.K. data for testing macroeconomic hypotheses have been stressed by Ravn (1997) (evaluating the RBC model against the behavior of U.K. real aggregates), Nelson and Nikolov (2004) (using a small New Keynesian model to evaluate different U.K. policy regimes), and Benati (2004) (assessing the behavior of U.K. data moments over the postwar period).

[2] Some work for the U.K. (e.g. Castelnuovo and Surico, 2006) focuses on 1992 as the start of the present policy regime. But in commencing estimation in the late 1970s for our baseline results, we are in line with the choices implied by some of the BEQM equation estimation periods (see e.g. pp. 115–120 of Harrison *et al*, 2005).

Table 1. Parameters in model

PARAMETER	DESCRIPTION	PARAMETER	DESCRIPTION
β	discount factor	ξ_p	degree of price stickiness
α	output elasticity to capital	ξ_w	degree of wage stickiness
δ	depreciation rate	$1/\kappa$	elasticity of investment to $\hat{p}_{K'}$
μ	steady state money growth rate	$1/\sigma_a$	elasticity of utilization to $\hat{r}^k$
ψ_l	relative weight of leisure in utility	ρ	interest rate smoothing
ψ_q	relative weight of real balances in utility	r_π	int. rate response to inflation
λ_w	wage markup	r_y	int. rate response to output gap
$1/\sigma_q$	semielasticity of money demand to $\hat{R}$	$r_{\Delta\pi}$	int. rate response to changes in inflation
b	habits parameter	$r_{\Delta y}$	int. rate response to changes in output gap
λ_f	price markup	σ_ε	standard deviation of mon. policy shock

Table 2. Variable definitions

VARIABLE	DESCRIPTION	VARIABLE	DESCRIPTION
$\hat{m}$	real money supply	$\hat{\pi}$	inflation
$\hat{q}$	households' money demand	$\hat{w}$	real wage
$\hat{k}$	capital services	$\hat{r}^k$	rental rate on capital
$\hat{\bar{k}}$	physical capital	$\hat{\mu}$	growth rate of nominal money supply
$\hat{l}$	labor	$\hat{p}_{K'}$	price of capital
$\hat{c}$	consumption	$\hat{U}_C$	marginal utility of consumption
$\hat{i}$	investment	$\hat{R}$	nominal interest rate
$\hat{y}$	output		

$$E_{t-1}\Big(\frac{\beta\xi_p}{(1-\xi_p)(1-\beta\xi_p)}\hat{\pi}_{t+1}-\frac{\xi_p(1+\beta)}{(1-\xi_p)(1-\beta\xi_p)}\hat{\pi}_t+ \frac{\xi_p}{(1-\xi_p)(1-\beta\xi_p)}\hat{\pi}_{t-1}+\left[\alpha\hat{r}_t^k+(1-\alpha)\left(\hat{R}_t+\hat{w}_t\right)\right]\Big)=0 \quad (1)$$

Along similar lines, staggered contracts for nominal wages, with a clause for indexation to the preceding period's price inflation, give rise to the nominal wage equation:

$$E_{t-1}(\varpi_1\hat{w}_{t+1}+\varpi_2\hat{w}_t+\varpi_3\hat{w}_{t-1}+\varpi_4\hat{\pi}_{t+1}+\varpi_5\hat{\pi}_t+\varpi_6\hat{\pi}_{t-1}+\hat{U}_{C,t}-\hat{l}_t)=0 \quad (2)$$

where

$$\varpi_1=\frac{-\beta\xi_w}{(1-\beta\xi_w)(1-\xi_w)}\left(1+\frac{\lambda_w}{\lambda_w-1}\right)$$

$$\varpi_2=\frac{1+\beta\xi_w^2}{(1-\beta\xi_w)(1-\xi_w)}\left(1+\frac{\lambda_w}{\lambda_w-1}\right)-\frac{\lambda_w}{\lambda_w-1}$$

$$\varpi_3=\frac{-\xi_w}{(1-\beta\xi_w)(1-\xi_w)}\left(1+\frac{\lambda_w}{\lambda_w-1}\right)$$

$$\varpi_4=\varpi_1,\varpi_6=\gamma_w\varpi_3$$

$$\varpi_5=\frac{\xi_w(1+\beta)}{(1-\beta\xi_w)(1-\xi_w)}\left(1+\frac{\lambda_w}{\lambda_w-1}\right)$$

Firms' optimality conditions imply that total payments for capital services equal total labor cost:

$$\hat{r}_t^k+\hat{k}_t=\hat{w}_t+\hat{R}_t+\hat{l}_t \quad (3)$$

Underlying this condition is the assumption that firms finance the wage bill with funds borrowed one period in advance. Real unit labor costs are therefore (in log terms) equal to the sum of the real wage and the short-term nominal interest rate.

The household's intertemporal Euler equation for consumption and first-order condition for investment purchases are respectively:

$$\hat{U}_{C,t} = E_t(\hat{R}_{t+1} - \hat{\pi}_{t+1} + \hat{U}_{C,t+1}) \quad (4)$$

$$E_{t-1}\,(\widehat{p}_{\mathrm{K}',t} + \widehat{U}_{C,t}) = E_{t-1}(\widehat{U}_{C,t+1} + \frac{r^k}{r^k + 1 - \delta}\hat{r}^k_{t+1} + \frac{1-\delta}{r^k + 1 - \delta}\widehat{p}_{\mathrm{K}',t+1}) \quad (5)$$

With habit formation in preferences, the marginal utility of consumption depends on the current, prior and expected future levels of consumption:

$$\left(1 - b\beta\right)\left(1 - b\right)E_{t-1}\hat{U}_{C,t} - b\beta E_{t-1}\hat{c}_{t+1} + \left(1 + b^2\beta\right)E_{t-1}\hat{c}_t - b\hat{c}_{t-1} = 0 \quad (6)$$

The amount of capital utilization, i.e. the difference between capital services and physical capital, is chosen to equate the marginal cost to the marginal benefit:

$$E_{t-1}\left(\hat{k}_t - \hat{\bar{k}}_t\right) = \frac{1}{\sigma_a}E_{t-1}\hat{r}^k_t \quad (7)$$

where $1/\sigma_a$ is the elasticity of the utilization cost function.

The optimal condition for rental of capital services is:

$$E_{t-1}\widehat{p}_{\mathrm{K}',t} = \kappa E_{t-1}\left(\hat{\imath}_t - \hat{\imath}_{t-1}\right) - \beta\kappa E_{t-1}\left(\hat{\imath}_{t+1} - \hat{\imath}_t\right) \quad (8)$$

where β is the household's discount factor, and $1/\kappa$ is the elasticity of the adjustment cost function.

The stock of physical capital obeys the law of motion:

$$\hat{\bar{k}}_{t+1} = \left(1 - \delta\right)\hat{\bar{k}}_t + \delta\hat{\imath}_t \quad (9).$$

Here δ denotes the depreciation rate.

The money demand function is given by:

$$\hat{q}_t = -\frac{1}{\sigma_q}\left[\frac{R}{R-1}\hat{R}_t + \hat{U}_{C,t}\right] \quad (10)$$

The following identity gives the growth of nominal money supply in terms of inflation and changes in real money supply:

$$\hat{\mu}_{t-1} = \hat{m}_t - \hat{m}_{t-1} + \hat{\pi}_t \quad (11)$$

Loan market clearing requires that total money demand by firms (to finance their wage bill) and by households is equal to aggregate money supply:

$$q\hat{q}_t + wl\left(\hat{w}_t + \hat{l}_t\right) = \mu m\left(\hat{\mu}_t + \hat{m}_t\right) \quad (12)$$

The resource constraint and the aggregate production function can be written as:

$$\left(\frac{1}{\beta} + \delta - 1\right)\frac{K}{Y}\left(\hat{k}_t - \hat{\bar{k}}_t\right) + \frac{C}{Y}\hat{c}_t + \delta\frac{K}{Y}\hat{\imath}_t = \hat{y}_t \quad (13)$$

$$\hat{y} = \alpha\lambda_f \hat{k}_t + (1-\alpha)\lambda_f \hat{l}_t \quad (14)$$

Monetary policy follows a dynamic version of the Taylor rule:

$$\hat{R}_t = \rho\hat{R}_{t-1} + (1-\rho)\left[r_\pi \hat{\pi}_{t-1} + r_y \hat{y}_t\right] + r_{\Delta\pi}\left(\hat{\pi}_t - \hat{\pi}_{t-1}\right) + r_{\Delta y}\left(\hat{y}_t - \hat{y}_{t-1}\right) - \varepsilon_t \quad (15)$$

The short-term nominal interest rate is therefore a smoothed function of inflation, output, and changes in inflation and output.

Other than our use of an interest-rate rule, the model we use corresponds to the CEE benchmark. A limitation in our application to the U.K. data is that the CEE model is closed-economy. But there are several reasons for using a closed-economy model when analyzing the U.K; see Neiss and Nelson (2003) for a discussion. For the present paper, the main reasons why a closed-economy of the DSGE model might be suitable for the U.K. are: (*a*) openness makes it difficult to model capital formation endogenously, whereas the presence of endogenous capital is a key feature of the CEE model; (*b*) the simplest open-economy models give counterfactual weight to the exchange rate in CPI inflation dynamics; and once the exchange-rate channel is "tamed" by such approaches as assuming incomplete pass-through, imported intermediates, etc., the model's properties become more like those of a closed economy.

3. Estimation

To estimate the model, we first obtain data responses to a monetary policy shock from a VAR for the United Kingdom. Then, as in CEE (2005), we match these impulse responses as closely as possible with the CEE model, using a minimum-distance estimation procedure.[3] Our analysis here is limited to monetary policy shocks, but there is evidence for the U.S. that estimates of the CEE model are robust to incorporating technology shocks into the analysis (see DiCecio, 2005, and Altig, Christiano, Eichenbaum, and Linde, 2005).

VAR estimates. We estimate our VAR on U.K. data that are a subset of the series studied in the U.S. case by CEE. Our VAR contains the logs of real GDP, real consumption,[4] real investment, and labor productivity, as well as the nominal Treasury bill rate and the quarterly (retail) inflation rate.[5] These choices imply a focus upon the response to a monetary policy shock of the policy rate, inflation, and aggregate demand, as well as the split of aggregate demand among its components, and the division of the output response between labor and other inputs. The sample period is 1979 Q2–2005 Q4. The start date is the quarter corresponding to the period (May 1979) in which the Thatcher Government first took office, and so an important monetary policy regime change.[6] It also corresponds approximately to the date of some other important changes in government policy that are important for the VAR responses, as we discuss in Section 4.

Figure 1 plots the estimated VAR responses to a monetary policy shock and their standard errors, along with the match to each response made by our estimates of the CEE model; the model-based responses are the red lines. Parameters fixed in estimation are given in Table 3.

Parameter estimates. The parameter estimates resulting from this matching of impulse responses are given in Table 4.

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[3] This procedure was also used by Rotemberg and Woodford (1997) with a smaller VAR.
[4] We have not split consumption between durables and nondurables. VAR impulse responses in Aoki, Proudman, and Vlieghe (2002, Chart 2), using a different VAR specification and sample period from ours, found similar response functions for the two types of consumption.
[5] Our VAR does not include a time trend. Impulse responses look similar regardless of whether a linear trend is included in the VAR.
[6] See e.g. Goodhart (1989) for a perspective on this regime change.

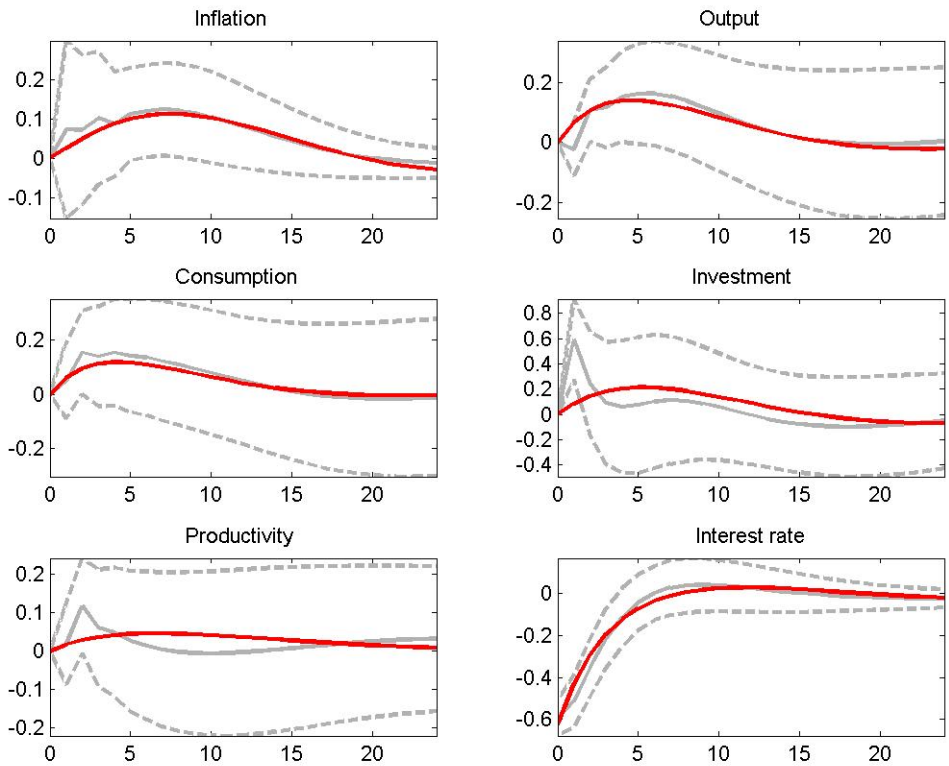

Figure 1. Estimated and model responses to a monetary policy shock

Table 3. Parameters fixed in estimation

Parameter	Value
β	$(1.03)^{-1/4}$
α	0.36
δ	0.025
μ	1.017
ψ_l	s.t. $l = 1$
ψ_q	s.t. $\frac{q}{m} = 0.25$
λ_w	1.05
σ_q	1

Table 4. Baseline model estimates Sample period 1979 Q2–2005 Q4	
Private sector parameters	
B	0.7739 (0.0007)
λ_f	2.0982 (0.0078)
ξ_p	0.9371 (0.0003)
ξ_w	0.1676 (0.0058)
κ	16.6225 (0.1001)
σ_a	∞ (—)
Monetary policy rule parameters	
ρ	0.8720 (0.0224)
r_π	1.269 (0.1228)
r_y	0.3482 (0.1426)
$r_{\Delta\pi}$	–0.6152 (0.1693)
$r_{\Delta y}$	0.4302 (0.0604)
σ_R	0.1564 (0.00001)

The parameter indexing habit formation in consumption is larger than that estimated by CEE (2005) and Altig, Christiano, Eichenbaum, and Linde (2005), but is basically in line with Fuhrer (2000). So the degree of habit formation in the U.K. appears similar to U.S. estimates.

The markup estimate is, at somewhat above 2, high by the standards of calibrated and estimated DSGE models. It is, however, in line with the estimate of the average U.K. gross markup (in manufacturing) by Haskel, Martin, and Small (1995, p. 30) of 2.0. Our high markup estimate is also made more standard by taking into account that it should be regarded as the wedge between consumer prices and (principally) nominal wages,[7] including the impact of cost elements we have not modeled explicitly.[8] At the back of one's mind it should be remembered that the model is an abstraction of a model with imported intermediate goods and indirect taxes. With these unmodelled elements built into the empirical price level series, the estimated markup of retail prices on nominal wages is increased.[9]

The estimated interest-rate policy rule has responses to both the level and growth rate of inflation as well as to the deviation of output from steady state. Because the interpretation of the inflation responses is affected by the output response, we deal with the latter response first. As technology shocks are held constant, any output movements reflect opening up of the output gap and so inflationary pressure. It is precisely this type of output variation that a monetary authority will have greatest interest in stabilizing. This may account for the output response being larger than is usual in estimated interest-rate rules, which typically do not remove from the output measure the variation that is due to technology shocks.

The inflation response consists of a standard level response and a negative response to the change in inflation. Under some parameter values, an estimated negative response to the change in inflation is equivalent to the policymaker having lagged inflation rather than current inflation in their rule. Our estimated $r_{\Delta\pi}$ response is, however, too large (in absolute value) for this to be the case. Instead, it is the case that the policymaker actually makes different-signed short-run responses to inflation, initially temporarily reducing the

[7] Interest on the nominal wage bill also enters the cost expression, with implications we discuss shortly.
[8] By contrast, Haskel, Martin, and Small's (1995) markup estimate allows for costs of materials, so our markup estimate should be higher than theirs, other things equal.
[9] Therefore, our high estimate may be consistent with Britton, Larsen, and Small (2000)'s setting the U.K. steady-state markup value closer to 1.0 when calibrating a model with explicit imported intermediates.

inflation rate in the short run. To understand this response, one has to keep the supply side of the CEE model in mind. In the standard sticky-price model the impulse response of output and the output gap to a monetary policy shock is identical, because potential output depends on real shocks only. In the CEE model, however, this is not the case because interest rates enter the production function, meaning that potential output depends on the nominal interest rate (see Ravenna and Walsh, 2006, for further discussion). Holding constant its other effects, a cut in the interest rate stimulates potential output and so helps inflation stabilization in the face of upward pressure on aggregate demand. In this light, in the wake of a monetary policy shock, policymaker stabilization of the output gap and inflation takes a three-pronged approach: a large response to output to rein in incipient excessive aggregate demand ($r_y > 0$, $r_{\Delta y} > 0$); a short-run cut in the interest rate as inflation rises to stimulate potential output ($r_{\Delta\pi} < 0$); and a sizable and durable positive response of the interest rate to the level of inflation relative to target ($r_\pi > 1$).

The estimates imply large investment adjustment costs, mainly driven by the matching of the smoother investment responses after the initial period; the model does not match the apparent initial spike in investment observed in the data.

Our model imposes rigidities in nominal wage adjustment (relative to a flexible price/flexible-wage baseline) in the form of automatic full indexation to prior inflation, but our estimates do not suggest much additional nominal wage stickiness beyond indexation. On the other hand, price stickiness is substantial. Because full indexation of prices is superimposed upon this price adjustment, it is not appropriate to infer from the low implied Calvo probability that prices are implausibly rigid; rather, the indexation implies substantial price movements every period even with the underlying price stickiness. Empirical support for lagged inflation terms in the Phillips curve, when this parameter is estimated freely, is not universal (see e.g. Ireland, 2001), and we intend to relax the assumption of full indexation. A lagged inflation term in the Phillips curve is, however, in line with the specification advocated by Blake and Westaway (1996) for U.K. monetary policy analysis.

Factor utilization is not found to be important, the relevant parameter being driven to the boundary of its admissible region. Our VAR productivity responses are not very precisely estimated, so the model can explain output variation in terms of input responses and so has little need to rely on the intensive margin to explain the data responses.

4. Estimates including pre-1979 data

In this section we present results for the long sample 1962–2005 as well as a sample using only pre-1979 data. The long-sample impulse responses and their matches are given in Figure 2, and those for the pre-1979 sample are given in Figure 3. Parameter estimates for each sample are given in Table 5.

Regarding the policy rule, the estimates deliver substantially lower inflation responses in the interest-rate rule pre-1979, consistent with the assignment of inflation control to nonmonetary devices in the U.K. before 1979. But the response is large enough even in this sample period to deliver determinacy (i.e. a single model equilibrium).[10] The output response is perversely signed pre-1979; this may be another reflection of the lack of monetary policy response to inflationary pressure since, as noted earlier, this coefficient denotes responses to the specific type of output increases that are likely to raise inflation.

Unlike our baseline estimates, both the sets of estimates that include pre-1979 data have a more standard (i.e., positive) interest-rate response to the change in inflation.

For the structural parameters, the main difference from our baseline estimates is that in the pre-1979 sample, the model cannot match the investment impulse response; the best the model can do to match is to suppress investment in the model (and so generate a flat investment model-response in Figure 3). Accordingly, the investment adjustment-cost parameter estimate becomes arbitrarily large.

It is tempting to suggest that the anomalous results regarding investment occur because the estimates including pre-1979 data are distorted by the existence of unmodelled breaks in monetary policy regime. But this does not really provide a satisfactory answer why we get these particular results. It is not obvious that estimated impulse responses over a sample that includes multiple regimes will be perverse in their shape; they are, more or less, an average of the responses observed across each regime, and so we should expect them to be of standard shape. Instead of this, we get model estimates that appear to extinguish the investment portion of aggregate demand.

[10] Our estimation routine only considers parameter combinations that deliver a single solution. An alternative procedure, which we have not pursued here, would be to consider both determinacy and indeterminacy regions and select a solution in the latter case using the minimum state variable procedure.

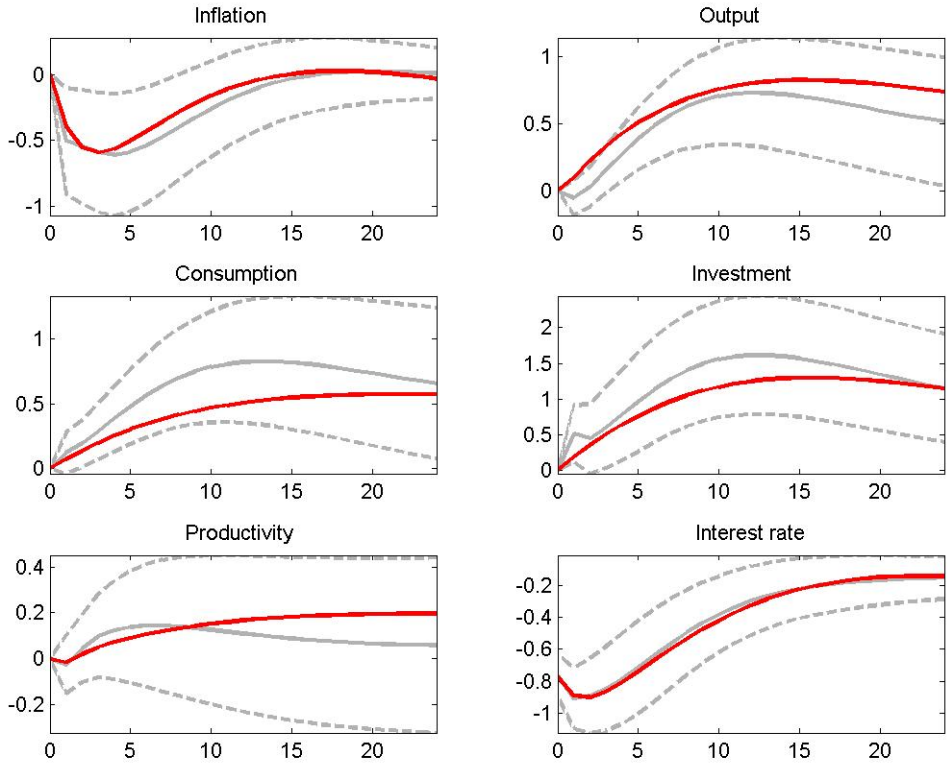

Figure 2. Estimated and model responses to a monetary policy shock
Sample 1962 Q3–2005 Q4

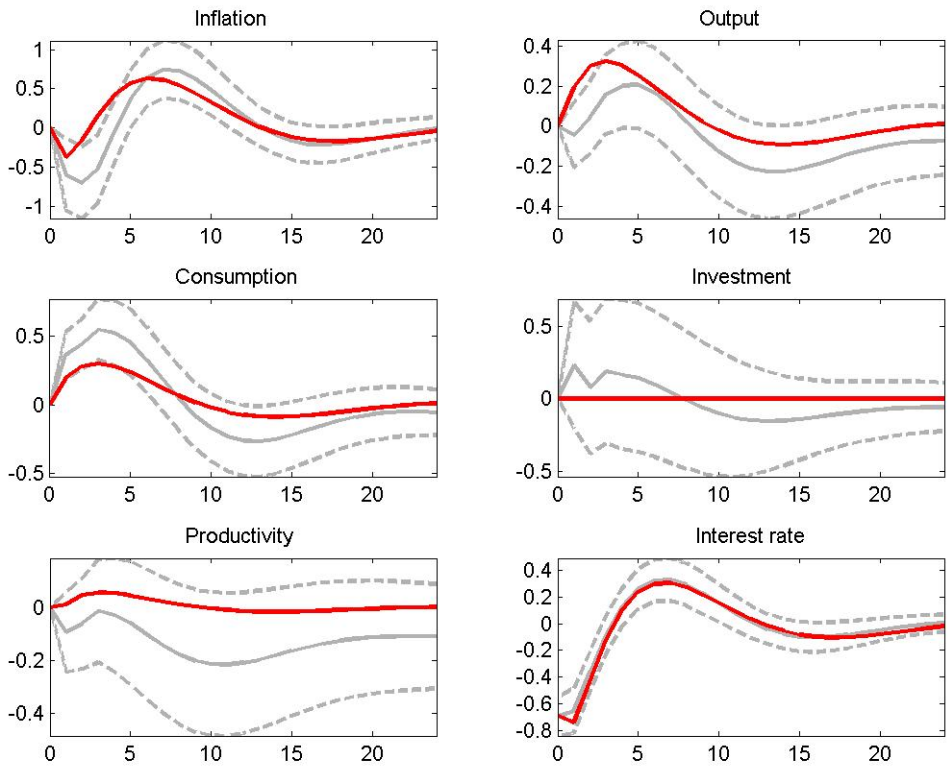

Figure 3. Estimated and model responses to a monetary policy shock
Sample 1962 Q3–1979 Q1

Table 5. Model estimates for samples that include pre-1979 data

	Sample period 1962 Q3–2005 Q4	Sample period 1962 Q3–1979 Q1
Private sector parameters		
b	0.9401 (0.0001)	0.5807 (0.0011)
λ_f	1.3839 (0.0011)	1.2018 (0.0024)
ξ_p	0.4366 (0.0016)	0.1696 (0.0015)
ξ_w	0.9882 (0.0001)	0.6071 (0.0004)
κ	52.0904 (0.1902)	∞ (—)
σ_a	0.3598 (0.0018)	0 (—)
Monetary policy rule parameters		
ρ	0.9726 (0.0109)	0.7011 (0.0013)
r_π	1.9491 (0.5981)	0.9430 (0.0380)
r_y	0.1296 (0.0793)	–0.1883 (0.0174)
$r_{\Delta\pi}$	0.4300 (0.0029)	0.5995 (0.0129)
$r_{\Delta y}$	0.1251 (0.0148)	0.1534 (0.0148)
σ_R	0.1941 (0.0002)	0.1720 (0.0002)

It is likely instead that government policy is indeed the culprit for the anomalies in the pre-1979 results, but that the driving factor is not monetary policy but U.K. governments' microeconomic interventions in the economy. Before the 1980s, many large industries (e.g. steel and telecommunications) were principally government-owned. What is more, in a misguided effort to control inflation by nonmonetary means, governments frequently intervened in the pricing decisions of their enterprises. For example, George Brown, then Secretary of State for Economic Affairs, said in 1965 said that the government was operating a price-control policy "in the field of government responsibility so far as charges for which they are responsible, prices which are there responsibility…"[11] Ted Heath, Prime Minister 1970–74, said shortly before being elected: "we are going to see to it that the State does not put up its prices and charges with gay abandon."[12] The attempt to enforce this policy led to considerable interference in government enterprises' operations, so much so that Anthony Crosland, a leading Labour Party figure, cited 1970–71 as a period that displayed a poor "balance between Ministerial control and entrepreneurial freedom."[13]

From around 1978, however, it became much more standard for government-owned enterprises to base their pricing and investment decisions on market signals, with a government report on the subject in 1978 stating: "The Government intends that the nationalized industries will not be forced into deficits by restraints on their prices."[14] The stepping-away by government from management of investment decisions was cemented by the privatization of many government enterprises in the 1980s.

Since the pre-1978 government interventions blocked investment from responding to market signals, including those from monetary policy shocks, one can understand why investment responses might deviate greatly from those predicted by our model, in which investment behavior is based on optimal firm choices. Government prohibitions on a firm's ability to raise prices might cut off funds to the firm, thus distorting investment decisions. On the other hand, for given monetary policy, government intervention in investment decisions might merely transfer aggregate demand pressure from investment to other categories of spending, rather than affect total demand. So impulse responses

[11] "Brown's Pay Policy Approved," *Glasgow Herald*, October 1, 1965, page 5.
[12] Quoted in William Russell, " 'Pushover' Jenkins Surrendered to Unions, Says Heath," *Glasgow Herald*, June 3, 1970, page 8.
[13] Crosland (1974, p. 39).
[14] U.K. White Paper on Nationalized Industries, 1978, quoted in *House of Commons Debates*, December 5, 1979, page 563.

other than those for investment might still be compatible with the model, which is essentially what we find.

5. Extensions

In this section we report further results regarding the robustness of our results to alternative data definitions (Section 5.1), and our choice of regime dates (Section 5.2).

5.1 Alternative investment series

The BEQM model and such sources as the Bank of England *Inflation Report* use a slightly narrower definition of investment than that we have employed. This narrower definition is known as "Business Investment" (though, like our series, investment by government enterprises). In Figure 4 we show that using this series in our VAR has little effect on the empirical responses to a monetary policy shock. The parameter estimates using this series are given in Table 6. These are little changed from the baseline parameter estimates, with a couple of important exceptions. Most notably, the estimates attribute some movements in production to varying utilization, whereas the previous estimates precluded a role for utilization; and the alternative estimates give roughly equal importance to price and wage stickiness, whereas in the baseline estimates prices were much stickier than wages.

5.2 VAR stability and regime breaks

As noted above, our data cover the whole period 1962–2005, but our baseline structural estimates are based on a sample covering only the period since the late 1970s, reflecting the changes in U.K. industrial and monetary policies that took place around that time. To investigate further the issue of regime break dates, in Table 7 we follow Boivin and Giannoni (2002, p. 99) by investigating the stability of the VAR when it is estimated for the long sample 1962 Q1–2005 Q4. We report the *p*-values for the constancy of the coefficients associated with each group of regressors in the VAR. The break date suggested by the test is also reported, and those with 10% significance or better are highlighted.

The results for the baseline VAR specification—that is, the VAR specification underlying the Figure 2 impulse responses—occupy the top half of the table. The results suggest a

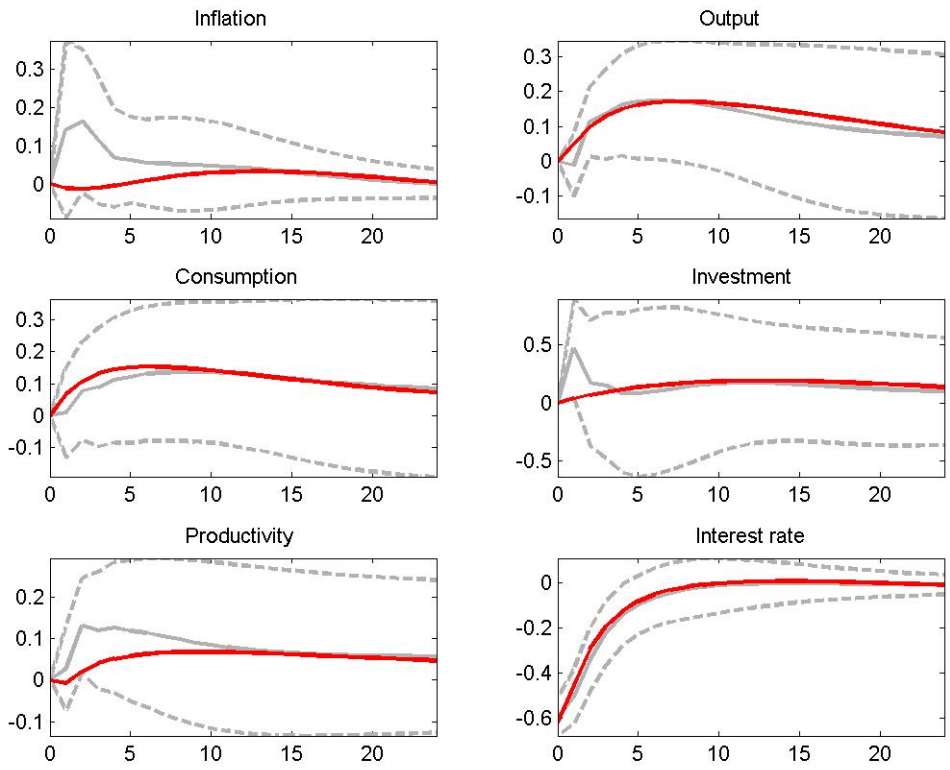

Figure 4. Estimated and model responses to a monetary policy shock, with alternative investment series

Sample 1979 Q2–2005 Q4

Table 6. Estimates using alternative investment series Sample period 1979 Q2–2005 Q4	
Private sector parameters	
b	0.7486 (0.0009)
λ_f	1.9238 (0.0059)
ξ_p	0.8486 (0.0016)
ξ_w	0.9620 (0.0007)
κ	35.2587 (0.3656)
σ_a	0.8482 (0.0142)
Monetary policy rule parameters	
ρ	0.9720 (0.0084)
r_π	1.5985 (0.2202)
r_y	0.9077 (0.3088)
$r_{\Delta\pi}$	−1.1072 (0.0563)
$r_{\Delta y}$	0.6744 (0.0182)
σ_R	0.1551 (0.0001)

Table 7. VAR stability tests (sample period: 1962 Q3–2005 Q4)

Dependent variable	Regressor					
	π	y	c	i	$y-h$	r
1. Baseline VAR specification						
π	**0.002** **1975 Q2**	**0.001** **1977 Q1**	0.713 1975 Q3	0.153 1971 Q1	0.510 1970 Q4	0.574 1973 Q2
y	0.127 1972 Q3	0.682 1975 Q3	0.373 1990 Q2	0.358 1976 Q1	0.948 1975 Q3	0.347 1977 Q4
c	0.469 1990 Q2	0.444 1990 Q2	**0.001** **1977 Q1**	**0.012** **1980 Q2**	**0.001** **1977 Q1**	**0.008** **1980 Q2**
i	0.553 1973 Q1	0.844 1968 Q4	0.788 1976 Q1	0.195 1979 Q2	0.524 1975 Q3	0.711 1975 Q3
$y-h$	0.382 1973 Q2	0.943 1975 Q3	0.922 1973 Q2	0.869 1973 Q2	0.361 1990 Q2	0.199 1974 Q1
r	0.275 1977 Q4	**0.000** **1981 Q2**	**0.000** **1981 Q2**	**0.006** **1977 Q4**	**0.000** **1981 Q2**	0.471 1973 Q1
2. Baseline VAR specification with y, c, i, $y-h$ detrended						
π	**0.002** **1975 Q2**	**0.072** **1975 Q2**	**0.040** **1981 Q4**	**0.004** **1976 Q1**	0.567 1978 Q4	0.741 1988 Q4
y	0.110 1972 Q3	**0.036** **1984 Q1**	0.121 1981 Q2	0.224 1987 Q2	0.296 1981 Q4	0.867 1977 Q4
c	0.738 1970 Q1	**0.063** **1981 Q3**	**0.065** **1980 Q3**	**0.095** **1980 Q2**	0.105 1975 Q2	**0.001** **1980 Q2**
i	0.812 1968 Q4	0.269 1987 Q1	0.239 1985 Q1	**0.014** **1976 Q4**	**0.049** **1980 Q1**	0.662 1990 Q2
$y-h$	0.325 1972 Q3	0.147 1984 Q1	**0.078** **1984 Q1**	0.187 1976 Q4	0.476 1990 Q2	0.665 1974 Q1
r	0.425 1977 Q4	0.813 1980 Q1	0.714 1986 Q1	0.801 1990 Q3	0.865 1998 Q3	0.460 1976 Q3

Note: Values reported are the *p*-values for the Andrews (1993) *sup*-Wald test, computed using the procedure of Diebold and Chen (1996). The null hypothesis is no structural breaks, while the alternative hypothesis is break in the constant and group of lag coefficients on the indicated regressor. Each panel also gives the break-date associated with the *p*-value.

significant break in the inflation equation around 1975 Q2. This date, however, does not constitute a monetary policy regime break; instead, the mid-1975 instability reflects a one-time shock to industrial policy. U.K. governments' policy, described above, of holding down nationalized industries' prices, underwent an adjustment in this period, with prices allowed to rise to eliminate the discrepancy with costs that had accumulated.

The large impact effect on consumer prices has sometimes been categorized as tantamount to a substantial increase in indirect taxes (Wilson, 1984, p. 50). The test statistics mechanically record this large shock as evidence of regime change. The remaining stability rejections are spread over 1977–81 and so are roughly in line with our focus in estimation on a 1979 break date.

Recall that our VAR does not use detrended real data, nor does it include a trend as a regressor. Some work on U.S. data, e.g. Rotemberg and Woodford (1997), Boivin and Giannoni (2002), and Giannoni and Woodford (2005), detrends real variables before putting them in the VAR. We report in the bottom half of Table 1 the stability results for our VAR when our specification is modified by replacing the four real variables with their detrended counterparts. The detrending assumes a broken linear trend with constant and trend breaks in both 1973 Q4 and 1981 Q4.

Beside continuing to show a break in 1975 in some of the inflation coefficients, these stability results largely reaffirm a focus on a regime break around the early 1980s (specifically, 1980 or 1981). Two of the significant rejections of stability do suggest a break in 1984 in GDP and productivity behavior, but these rejections can be discounted as reflecting the temporary disturbances to output from the coalmining strike of that year.

One puzzling aspect of the results with detrended variables is that the interest-rate equation no longer registers any significant regime break. This, however, is not decisive evidence against the importance of monetary policy regime change. First, relatively minor and statistically insignificant changes in the VAR coefficients can imply large changes in the implied "long-run response" of the interest rate to endogenous variables. This is the case here, since, despite the lack of rejection of stability: the VAR equation for the interest rate underlying the final row of the table has a long-run solution with an interest-rate response to (annualized) inflation of about 0.3, but this rises to 1.0 upon restricting the sample to 1979–2005.[15] Secondly, the inflation VAR equation now exhibits a significant early 1980s break, which is indirect support for a monetary policy regime change around that time.

[15] In some contexts (see e.g. Rotemberg and Woodford, 1997, and Rudebusch, 1998) the VAR equation for the interest rate coincides with the interest-rate policy rule. This is not the case in our analysis, as the policy rule that we use in estimation differs from the VAR equation. But solving the reduced-form VAR interest-rate equation for its long-run solution nevertheless provides a means of cross-checking the stability test results.

6. Conclusions

In this paper we have estimated the Christiano, Eichenbaum, and Evans (2005) model using U.K. data. While CEE found plausible estimates on U.S. data when treating the period since the 1960s as a single regime, for the United Kingdom it appears that more satisfactory estimates emerge if pre-1979 data are excluded; otherwise, the estimates imply degenerate behavior of investment. This result is consistent with policy regime changes being an important factor in the postwar U.K. economy. These regime changes include not only changes in the role assigned to monetary policy, but shifts toward making investment decisions more closely related to market forces. Another important implication of our estimates is that price stickiness, rather than wage stickiness, is the major source of nominal rigidity in the U.K.

Appendix 1. Data sources and definitions

Nominal interest rate: U.K. Treasury bill rate, quarterly average; source: Haver-*IFS*, quarterly average series.

Output: Real GDP, s.a., quarterly, series abmi.q; source: U.K. Office of National Statistics (ons.gov.uk), download date May 2006.

Private household consumption, s.a., quarterly, series abjr.q; source: ons.gov.uk, download date May 2006.

Investment: Gross fixed capital formation, s.a., quarterly, series npqt.q; source: ons.gov.uk, download date May 2006.

Alternative investment series: Business investment at 2003 prices, s.a., quarterly, series npel.q; source: ons.gov.uk, download date August 2006.

Productivity: Y/H, where H = hours worked. Source for H: series ybus.q (source: ONS); with splice into Ravn (1997) U.K. hours worked series to obtain pre-1971 Q1 data. Inflation: log difference of P, where P is a seasonally adjusted consumer price series. P was constructed as follows A quarterly average of RPI was spliced into a quarterly average of RPIX after 1973, the series was seasonally adjusted, then tax-related spikes of 4% (in 1979 Q3) and 2% (1990 Q2) were removed from the series. The seasonal regressions underlying the seasonal adjustment used the log-change as the dependent variable, and seasonal patterns were allowed to differ across 1955–1976, 1976–1986, and 1987–2005. The source for the monthly RPI underlying the quarterly averages was the U.K. Office of National Statistics (ONS) (ons.gov.uk). The ONS, however, provides RPIX data only from January 1987. An unofficial RPIX series starting in 1974 has, however, been constructed at the Bank of England, and this series underlies studies such as Nelson and Nikolov (2004) and Benati (2004). The OECD-Haver service also provides an RPIX series (though beginning only in 1975) that closely matches this series. We used the quarterly average of the unofficial RPIX series for 1974–87 and spliced it into quarterly average of the official RPIX series that begins in 1987. Splicing this RPIX series at 1974 Q1 with RPI delivered the RPI/RPIX quarterly average series underpinning P.

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