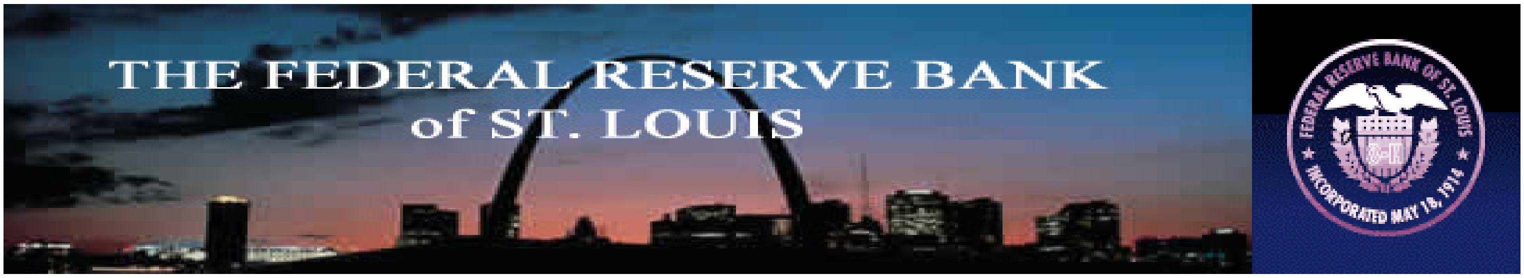

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International Evidence on the Stability of the Optimizing IS Equation

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International Evidence on the Stability of the Optimizing IS Equation

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Abstract

In this paper we provide international evidence on the issue of whether the optimizing IS equation is more stable than a backward-looking alternative. The international evidence consist of estimates of IS equations on quarterly data for the UK and Australia, both for the full sample of the last 40 years and for the period following major monetary policy shifts in 1979–80. Our results suggest that the parameters in the optimizing IS equations are more empirically stable than those of the backward-looking alternative. The use of dynamic general equilibrium modelling in empirical work does deliver material benefits, in the form of equations that are more suitable for monetary policy analysis.

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1. Introduction

Dynamic stochastic general equilibrium (DSGE) models have become prevalent in monetary policy analysis; see Walsh (2003) and Woodford (2003) for book-length treatments. A well-known property of these models is that, because they include first-order conditions that arise from dynamic constrained optimization by private agents, forward-looking terms—i.e. the rational expectation of future values of endogenous variables—typically appear in the structural equations describing spending and pricing decisions. This feature of DSGE models means that dynamics arising from expectations about policy behaviour are explicitly separated from dynamics that spring from the private sector's preference and production functions. Moreover, the preference and production functions are summarized by behavioural parameters that should not change with monetary policy regime. This feature suggests that parameters in DSGE models should be constant in the face of the shifts in monetary policy regime that have been observed empirically in many countries. By contrast, backward-looking representations of endogenous variables, such as vector error-correction or vector autoregressive characterizations of the data, do not distinguish dynamics from the different sources. The coefficients in these backward-looking equations are, in other words, vulnerable to the Lucas (1976) critique.

In a recent paper, Estrella and Fuhrer (EF) (2003) argue that for models intended for analyzing monetary policy, the advantages claimed for forward-looking optimization models over backward-looking models, do not hold in practice. Specifically, EF apply a variety of statistical tests to evaluate the relative stability of the fully backward-looking Rudebusch-Svensson (1999) model and one version of McCallum and Nelson's (1999) forward-looking model, where the latter consists of the optimizing IS equation combined with the standard New Keynesian Phillips curve.[1] EF conclude that "of the current crop of small macroeconomic models, it seems that backward-looking formulations… are somewhat more stable than their forward-looking counterparts" (EF, 2003, p. 102).[2]

In this paper, we present some arguments and results that suggest that EF's conclusion is unwarranted. We focus on the stability of the IS equation; as discussed below, EF's results shed little light on the stability of the New Keynesian Phillips curve. Our reinterpretation of the evidence on IS function stability takes two parts. First, we argue

[1] Earlier work (e.g. Favero, 1993; Ericsson and Irons, 1995) tested the Lucas critique on empirical consumption and money demand functions.
[2] This is the fifth of EF's conclusions, and the one most relevant for the present paper. We have little disagreement with EF's four other conclusions.

that EF's concentration on test statistics and forecast performance means they neglect the type of evaluation of the Lucas critique stressed in Lucas (1988). Lucas' approach to constancy compares directly the estimates of structural parameters across regimes; and by this criterion, as we show, EF's work is consistent with stability of the McCallum-Nelson IS equation. Secondly, EF reach their conclusions on US data; but as stressed in Lucas (1973) and Kormendi and Meguire (1984), international comparisons provide powerful evidence on the policy invariance of macroeconomic models. Our international evidence takes the form of estimated IS equations for Australia and the United Kingdom. We show that the estimated optimizing IS equations for these countries are considerably more stable and interpretable than the backward-looking alternatives.

Our paper proceeds as follows. Section 2 discusses and reinterprets EF's evidence for the US. Section 3 presents estimates of monetary policy reaction functions for the UK and Australia, and establishes that these countries exhibit the same kind of regime break in 1979–80 found by EF (2003, Section II) for the US. Section 4 reports full-sample and post-1979 estimates of the optimizing IS equation for the UK and Australia. The comparison shows that these estimates are similar to those for the US, and that the key structural parameter does not appear to change with policy regime. Section 5 finds that, by contrast, there is evidence of parameter instability in the backward-looking Rudebusch-Svensson IS equation for both the UK and Australia. Section 6 concludes.

2. Estrella and Fuhrer's evidence for the US

2.1 The forward-looking model

The forward-looking, optimization-based model for which EF report unfavourable empirical results is the following familiar New Keynesian model.[3]

$$y_t = \mathrm{E}_t y_{t+1} - \sigma(R_t - \mathrm{E}_t \pi_{t+1}) + v_t \qquad (1)$$

$$\pi_t = \mathrm{E}_t \pi_{t+1} + \lambda(y_t - y_t^*) \qquad (2)$$

Here y_t is the log of output, y_t^* is the log of potential output (whose definition in this model we discuss further below), R_t is the nominal interest rate, π_t the quarterly inflation rate, $\sigma > 0$, $\lambda > 0$, and v_t is an IS shock. The growing variables y and y^* only

[3] Many studies have used versions of this model; EF focus on McCallum and Nelson (1999) because that paper used timing assumptions that imply a model of exactly the form of equations (1)–(2). More complicated timing assumptions are used in (e.g.) Rotemberg and Woodford (1997).

appear in the system in stationary combinations with other variables—i.e. as the forward-looking difference term $y_t - E_t y_{t+1}$ in the optimizing IS equation (1) and as the output gap term $y_t - y_t^*$ in the New Keynesian Phillips curve (2). Therefore, the model could equivalently be written with detrended log output y_t^d replacing y_t and detrended log potential output y_t^{d*} replacing y_t^*. The influence of optimizing behaviour is felt in the presence of forward-looking terms ($E_t y_{t+1}$ and $E_t \pi_{t+1}$) in each equation, as well as the use of rational expectations in the computation of the ex-ante real short-term interest rate, $R_t - E_t \pi_{t+1}$.

To complete the model, EF specify a monetary policy rule (a reaction function for the short-term nominal interest rate) which they estimate on US data. The estimated policy rule exhibits a break in response coefficients in late 1979. EF seek to evaluate whether equations (1) and (2) are structural—in the sense that they are constant in the face of the 1979 policy regime change. EF's tests include evaluating the stability of the parameters of forward and backward-looking models via static simulations and test statistics for stability with known and unknown breakpoints. They estimate two versions of the forward-looking model, one with *i.i.d.* errors in the IS and Phillips curve equations, and an alternative with serially correlated shocks. Graphical evidence indicates that the one-period static forecast generated by the backward-looking model is superior to the forward-looking model for the post-1979 period. The formal stability tests also support a preference for the backward-looking model.

We contend that, on closer inspection, EF's findings have little to say about the constancy of the New Keynesian system (1)–(2). This reconsideration takes two parts. First, their analysis of the constancy of the New Keynesian Phillips curve is flawed by the failure to use a measure of the output gap ($y_t - y_t^*$) that is consistent with the underlying theory. Secondly, EF's results on the IS equation add to existing evidence that this equation is indeed policy-invariant in US data. We discuss each of these issues in the remainder of this section.

2.2 Phillips curve evidence

Quite apart from Lucas-critique issues, there is an important feature of optimization-based models that distinguishes them from backward-looking models. This feature is that they imply a disparity between the behaviour of detrended output, on the one hand, and the output gap, on the other. This disparity applies in principle regardless of the detrending procedure applied in obtaining a detrended output series. The source of the divergence is that the output-gap concept implied by optimization-based models is the percentage discrepancy between two series—detrended output and detrended potential output—while detrended output records variation only in the first of these series. Detrended output will coincide with the output gap only in the case of a smooth (in the limit, nonstochastic) potential output series. But stochastic DSGE models provide few *a priori* grounds for potential GDP (relative to its steady-state growth path) to be smooth. On the contrary, potential output in these models corresponds to the value that output takes under flexible prices; and output under flexible prices is typically volatile and cyclical in flexible-price DSGE models (e.g. Prescott, 1986). The time-variation in potential GDP reflects not only the presence of productivity shocks in the production function, but also variations in the optimal responses (under price flexibility) of labour and other endogenous productive inputs, to movements in real shocks (which include shocks to both production and utility functions).[4]

This distinction is relevant for the interpretation of EF's results. The real-activity variable that appears in their empirical IS equation *is the same* as that in their estimated Phillips curve. The Congressional Budget Office (CBO) output-gap series is used in both equations—as the aggregate-demand variable in the IS function and the output-gap variable in the Phillips curve. From a DSGE modelling perspective, however, output or detrended output is the appropriate variable to use in the IS equation, while the theory-consistent output gap should be used in the Phillips curve. The CBO gap series closely resembles quadratically detrended log output in US data; so, despite its label, the CBO series appears a better proxy for detrended output than the output gap. By our criteria, therefore, EF used an appropriate real-activity measure in their IS equation, but not an appropriate measure of the gap in their Phillips curve. It is therefore very difficult to interpret either EF's test statistics for stability of the Phillips curve, or their joint tests of the stability and forecasting properties of the IS and Phillips curves. In particular, it is impossible to make conclusions about the stability of the McCallum-Nelson structural model, because an omitted variable problem—the movements in potential GDP over the cycle—may be responsible for the unfavourable test results reported by EF.[5]

Existing results suggest that this omitted-variable problem is far from negligible. McCallum and Nelson (1999) found that detrended output can be a very poor proxy for

[4] This feature of optimization-based models was an important element of the discussion in McCallum and Nelson (1999), and has also been emphasized in McCallum (2001), Neiss and Nelson (2002, 2003) Woodford (2001, 2003), and others.

[5] The only (implicit) allowance for this omitted variable in EF's specification is their specification of a serially correlated shock process in the Phillips curve. Estimates of this shock series may be a valid proxy for potential-output movements when there is not a distinct disturbance term in equation (2); otherwise, however, the shock estimates are likely to confound movements in potential GDP, the output gap, and the Phillips curve disturbance. Galí, Gertler, and López-Salido (2001) and Neiss and Nelson (2002) offer alternative rationalizations for a Phillips curve disturbance in equation (2).

the output gap in policy rules, a result confirmed by Neiss and Nelson (2003) in a richer model with capital and habit formation. In addition, Neiss and Nelson (2002) construct a theory-consistent potential output series for Australia, the UK, and the US, and estimate New Keynesian Phillips curves using the resulting output-gap series. The results are uniformly better than NKPC estimates using detrended output as a proxy for the gap, and resolve many of the empirical difficulties with NKPCs reported in the literature. On the basis of these results and the argument above, we conclude that EF's findings shed little light on the stability over time of the New Keynesian Phillips curve.[6] We therefore concentrate in the remainder of this paper on the other structural equation in the forward-looking model that EF critique: the optimizing IS equation.

2.3 Stability of the IS equation for the US

EF's evaluation of the robustness of the forward-looking IS equation to the Lucas critique is based on out-of-sample forecasting power and on statistical tests for model stability. This contrasts with the approach that Lucas himself took when evaluating the constancy of an empirical optimization-based specification (Lucas, 1988). Lucas refrained from deploying 'standard statistical tests for the stability of parameter estimates across different sample periods' (1988, pp. 160–162). Discussing his application (which was to money demand), Lucas went on,

> We begin with a simple economic model that suggests a two-parameter description of money demand. When we hypothesize that this relationship is 'stable', we mean that we expect these two parameters to reflect relatively stable features of consumer preferences and the way in which business is carried out, and we expect them not to shift around as monetary or other policies are altered over time. This theory does not suggest that the residuals can be characterized in a simple, elegant fashion over a given time period, or even that the stochastic structure of the residuals should be stable over time. (1988, p. 162).

While our interest is in the stability of the IS equation, not money demand, the same principle applies. The prediction of the optimizing model is that the preference parameter(s) in the equation studied should be constant, though the shock processes need not be. For the optimizing IS equation (1), this boils down to the prediction that σ is constant over time, even though both the monetary policy regime and the IS shock process (v_t) may undergo shifts. In light of his interpretation of stability, Lucas

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[6] EF's misspecification, in other words, is sufficient to conclude that one cannot infer from their results that the New Keynesian Phillips curve is truly non-constant—just as one cannot draw inferences about the constancy of the parameter c in the relationship $y = c.x + e$ from econometricians' estimates of $y = d.z + u$, where z is a variable that theory provides no reason to be closely related to x.

eschewed statistical tests in favour of direct examination of estimates of the household preference parameters across regimes.

In their evaluation of the IS equation, EF do not provide evidence of this kind. Indeed, even their full-sample parameter estimates of the model are not given. EF (2003) refer to EF (2002) for a 'complete empirical assessment of the relative performance of the backward- and forward-looking models in this paper'. However, the model in EF (2003) is estimated, while that in EF (2002) is calibrated from prior studies. We presume EF (2003) meant to refer instead to EF (2000). The latter working paper presents the parameter estimates for the models tested in EF (2003). EF (2000) only report full-sample estimates, with no direct evidence reported on how resilient the key slope parameters are to estimation on sub-samples or across regimes. Instead, as in EF (2003), test statistics alone are used to evaluate stability.

Table 1. Existing estimates of the IS curve slope coefficient

Study	*Data*	*Estimation method*	*Estimate of interest elasticity of aggregate demand*
McCallum and Nelson (1999)	US, quarterly data, 1955 Q1–1996 Q4	IV	0.16
Estrella and Fuhrer (2000)	US, quarterly data, 1966 Q1–1997 Q4	IV	0.22
Fuhrer (2000)	US, quarterly data, 1966 Q1–1995 Q4	Maximum likelihood	0.16
Ireland (2001)	US, quarterly data, 1980 Q1–1999 Q2	Maximum likelihood	0.22

Note: This table gives estimates of b in econometric estimates of IS equations of the type $y_t = E_t y_{t+1} - br_t + e_t$ where r_t is the real interest rate ($r_t = R_t - E_t\pi_{t+1}$) and e_t is a stochastic term. Reported estimates have been expressed in comparable units to allow for the different conventions across papers in measuring the interest rate (i.e., as annualized or quarterly). Fuhrer's estimates use consumption as the activity variable and allow for nonseparable utility.

Given our focus on the IS equation, there is only one parameter whose stability is of interest: the interest elasticity of aggregate demand, σ. While EF do not report subsample results, their (2000) results can be included with estimates in other studies for other sample periods to give an impression of the stability of the IS parameter σ over time. We report in Table 1 a collection of estimates of the σ parameter in the IS equation from studies on quarterly US data that have used either instrumental variables or maximum likelihood estimation. Despite some overlap of sample periods, these studies encompass a variety of approaches. For example, McCallum and Nelson use a decade of data not used in other studies; Ireland (2001) and McCallum and Nelson, but not EF or Fuhrer (2000), exploit cross-equation restrictions from other optimizing equations; Ireland and Fuhrer use maximum likelihood procedures; and EF and McCallum and Nelson, while both using instrumental variables estimation, employ different instrument sets.

Despite this diversity of approaches, the impression from Table 1 is of close agreement on the value of σ, including from a study (Ireland's) that uses data only from the post-1979 monetary regime. Accordingly, claims of instability of the optimizing IS equation seem unwarranted to us for the United States, provided that stability is interpreted appropriately. Our principal contribution in this paper, however, is to examine international evidence on the issue, and we concentrate on that issue henceforth.

3. Policy rule estimates for the UK and Australia

We now move to the empirical evidence we contribute in this paper to the debate. The data we have are 40 years of quarterly observations for both the UK and Australia. Following EF, prior to examining the stability issue, we first report estimates of interest-rate policy rules for these economies. As in EF, the specification of the interest rate policy rule we use is based on Clarida, Galí, and Gertler (1998, 2000) and Judd and Rudebusch (1998), according to which the short-term nominal interest rate is moved in partial-adjustment fashion to expected future (quarterly annualized) CPI inflation and levels and first differences of detrended (here, HP-filtered) log output:

$$R_t = (1-\rho_R)c + (1-\rho_R)\phi_\pi \mathrm{E}_t\pi_{t+1}{}^a + (1-\rho_R)\phi_y y_t{}^{hp} + (1-\rho_R)\phi_{\Delta y}\, \Delta y_t{}^{hp} + \rho_R R_{t-1}. \qquad (3)$$

As in the prior studies, we distinguish between pre- and post-1979 monetary policy behaviour by estimating this specification separately over the two subsamples. Before we present our estimates, some general remarks about the appropriateness of representing pre-1979 policy by an interest-rate rule are in order. It is true that in the UK and Australia prior to the 1980s, policymakers adopted a plethora of methods to influence monetary conditions: reserve requirements, marginal reserve requirements, limits on interest rates paid by banks, lending controls, operations in the market for long-term government debt—beside operating on short-term rates. However, this is *not* a distinguishing feature of the UK or Australia relative to the US: the US had its own versions of these controls (e.g. the Operation Twist program of intervention in the long-term debt market during the Kennedy Administration; Regulation Q on housing market loans; prohibition of payment of interest on demand deposits; heavy reserve requirements on both banks and non-banks). Insofar as these measures invalidate representing pre-1979 policy by interest-rate rules, they do so for the US as well as the UK and Australia. Bernanke and Blinder (1992) argue for the US that, notwithstanding the multitude of monetary policy tools, it is legitimate to look at the short-term interest rate as an index of policy actions.[7] The same principle extends to other countries. To take an example, an increase in reserve requirements by the central bank while keeping the short rate constant is ineffective, because the central bank then supplies the extra required reserves at the same cost as previously supplied reserves (the contractionary effect of higher required reserves are offset by a higher quantity of total reserves).[8] Consistent with this, the Governor of the Reserve Bank of Australia said as early as 1969 that 'we would nowadays tend to look at the more pervasive instruments of open market operations and interest rates' rather than direct controls (*Banker*, 1969, p. 1289). There is one important respect, furthermore, in which it is *more* appropriate for the UK and Australia to model policy as an interest-rate rule, than it is for the US. In the US, it was not officially acknowledged until 1994 that the authorities had an interest-rate instrument; by contrast, in the UK and Australia this was always acknowledged.[9] Therefore, there are, if anything, stronger grounds for the UK and Australia than for the US to characterize pre-1980s monetary policy as a rule for the short rate.

Results for each country and sample are reported in Tables 2 and 3. Three-month Treasury rates are the dependent variable for both countries.[10] We now comment on the parameter estimates, and provide comparisons with some prior studies for each country.

We first consider our UK estimates. As is well known, UK monetary policy has undergone several changes in regime. We concentrate, however, on the break in behaviour in 1979, specifically that corresponding to the election of the Thatcher Government in 1979 Q2. While inflation targeting was not introduced until 1992, the

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[7] The formal condition for this is that different types of loan and security be a perfect substitute for one another. The perfect-substitution assumption is standard in monetary analysis; in particular, the assumption is common to both the backward and forward-looking IS equations that EF (and we) study.

[8] See e.g. Meltzer (2001, p. 25).

[9] The following quotation from Foster (1979, p. 152) illustrates this state of affairs: 'It is worth emphasizing that short-term securities in the United Kingdom differ from those in the United States in that the latter are generally viewed as market determined whereas the former have been administered through Bank Rate policy…' By contrast, studies today routinely characterize 1970s US short-term rates as policy-determined.

[10] Throughout our sample, the three-month Treasury note rate in both the UK and Australia has been closely related with policy actions on short-term interest rates. For the UK, the various official policy rates (Bank Rate, Minimum Lending Rate, repo rate, etc.) have always been highly correlated with the Treasury bill rate. In Australia, the Deputy Governor of the Reserve Bank said in 1964 that 'it is an object of policy to ensure that there is always a ready market for any government security' (Phillips, 1964, p. 81), and he mentioned the Treasury note rate as a representative short rate. Our estimation of the IS curve will allow explicitly for the occasion when private rates in Australia differed most sharply from the Treasury rate.

Table 2. Policy rule estimates: United Kingdom						
Specification: $R_t = (1-\rho_R)c + (1-\rho_R)\phi_\pi E_t\pi_{t+1}^a + (1-\rho_R)\phi_y y_t^{hp} + (1-\rho_R)\phi_{\Delta y}\Delta y_t^{hp} + \rho_R R_{t-1}$ + price dummies						
	c	ϕ_π	ϕ_y	$\phi_{\Delta y}$	ρ_R	SEE
Sample period 1957 Q2–1979 Q1	0.037 (t = 6.66)	0.416 (t = 6.75)	0.809 (t = 2.04)	0.565 (t = 0.85)	0.651 (t = 6.40)	0.0108
Sample period 1979 Q2–2002 Q3	0.014 (t = 0.57)	1.555 (t = 3.11)	2.861 (t = 1.78)	6.178 (t = 1.47)	0.894 (t = 18.90)	0.0086
Instruments: lags 1–5 of inflation, interest rates, y^{hp}; also a constant and dummies for tax changes in 1979 Q3 and 1990 Q2.						

Table 3. Policy rule estimates: Australia					
Specification: $R_t = (1-\rho_R)c + (1-\rho_R)\phi_\pi E_t\pi_{t+1}^a + (1-\rho_R)\phi_y y_t^{hp} + \rho_R R_{t-1}$ + price dummies					
	c	ϕ_π	ϕ_y	ρ_R	SEE
Sample period 1963 Q1–1979 Q4	0.026 (t = 2.03)	0.561 (t = 3.05)	0.757 (t = 0.81)	0.891 (t = 15.40)	0.0055
Sample period 1980 Q1–2002 Q1	0.035 (t = 2.64)	1.151 (t = 5.22)	1.131 (t = 1.78)	0.827 (t = 16.82)	0.0111
Instruments: lags 1–5 of inflation, interest rates, y^{hp}; also a constant and dummies for tax changes in 1975 Q3, 1975 Q4, 1976 Q4, and 2000 Q3.					

1979 shift was the quantitatively most important break in monetary policy regime, whether judged by inflation outcomes (the break in the mean of inflation between the 1970s and the 1980s being greater than the break observed pre- and post-1992)[11] or the visual relation of nominal interest rates to inflation (real rates were permitted to be positive from 1979 onward).

Estimates of the UK monetary policy reaction function have been presented for the pre-1979 period by Nelson (2003), and for post-1979 by Clarida, Galí, and Gertler (1998)

[11] UK inflation did rise above 20% in 1980, but this is not in itself evidence against a monetary policy regime break in 1979, because one must allow for a realistic lag in the effect of monetary policy actions on inflation. If we allow for a three-year lag to allow monetary changes to be have most of their effect manifested on inflation, then a policy regime in force from 1979–80 should be judged by inflation outcomes from 1983. Annualized RPIX inflation in the UK averaged 5.2% from 1983 Q1 to 1992 Q4, down 9.0 percentage points on the prior decade's mean and consistent with a major change in regime.

and Nelson. As for the US, there is a much lower response to inflation in the policy rule before 1979 than afterward, consistent (in the UK at least) with inflation control being delegated to other devices such as incomes policy. We find that the whole post-1979 period can be characterized as a policy with a coefficient on expected future inflation of about 1.6, similar to that found in Nelson for the 1992–97 inflation-targeting period. Relative to Nelson's results, we find here that the 1979–92 period is more homogeneous with the inflation-targeting epoch. In part, this reflects our controlling for indirect-tax and other distortions to inflation in our estimation: our empirical implementation of equation (3) specifies that policymakers do not respond to rises in inflation due to one-time events such as the increase in VAT in 1979.[12] We also find strong policy responses to developments in output.[13]

Our estimates for Australia indicate that the response to inflation increases from 0.56 in the early period to a non-accommodating 1.15 in the second period, which we take to be 1980 onward. These rule estimates seem more plausible and interpretable than those for Australia in two previous studies, Gagnon and Ihrig (2002) and Gerlach and Smets (2000, pp. 1693–94). Both studies focus on the 1990s, and neither obtains a rule that satisfies a basic version of the Taylor principle (i.e., $\phi_\pi > 1$; see Woodford, 2001). We have several differences in data and specification from these studies, including our use of one-period-ahead rather than four-period-ahead inflation in the estimated rule; the former, which is CGG's (1998) baseline specification for several G7 countries, evidently produces more plausible results. The different results are not due to sample period choices: our estimates for 1993 Q1–2001 Q4, like our 1980–2002 estimates, deliver a rule that satisfies $\phi_\pi > 1$.

This brings us to some further examination of our choice for the regime-break date for Australia. We emphasize again that our aim is to concentrate on a single, quantitatively important regime break, and are conscious that further changes in regime occurred in the 1990s in both Australia and the UK. Our choice of a new regime starting in 1980 Q1 is in line with Gruen and Shuetrim's (1994, p. 333) nomination of this date on the grounds that 'the [December 1983 exchange rate] float was not an important regime change because, between 1980 and 1983, the exchange rate was still fairly flexible.' Beside the evidence on policy rules presented here, other evidence also suggests that a more anti-inflationary policy took place from around 1980. For example, once an allowance is made for a lag of around three years from monetary developments to inflation, mean inflation over 1983 Q1–1992 Q4 was

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[12] We allow for this in estimation by including dummy variables for the tax changes in the policy rule.
[13] Our UK estimates are consistent with UK policymakers reacting to a form of (annualized) nominal income growth, $\pi^a + 4 \cdot \Delta y^{hp}$.

6.3%, down 5.4 percentage points from the prior decade. The 1980-82 period also had a real interest rate of 2.3% on average, compared to −1.0% in 1961−79 and 3.0% since 1993. In other words, even prior to the onset of inflation targeting in Australia in 1993–94, there was a substantial fall in inflation relative to the 1970s and early 1980s, consistent with an important break in policy behaviour relative to the 1970s.

Beechey *et al*'s (2000) work on Australia commences estimation in 1985 Q1, and so as a check on our estimates, we reestimated our policy rule estimates against that break. Dropping the 1980–84 observations from the later policy regime does not change the estimated long-run responses significantly, but leads to decreased precision of the estimates (the *t*-value for the inflation response falls from 5.2 to 4.5, that for the output response from 1.8 to 1.0). And adding the 1980–84 observations to the earlier sample leads to little improvement in the *t*-value on the output term (this rises from 0.8 to 0.9), while sharply reducing the significance of the expected-inflation term (its *t*-value falls from 3.1 to 1.3). These results suggest that our emphasis on a 1980 break is appropriate.

The evidence in this section supports the notion that, just as in the US, a very substantial monetary policy regime shift occurred in 1979–80 in the UK and Australia. We therefore zero in on this period when we estimate the IS equation across regimes.

4. Estimates of the IS equation for the UK and Australia

With policy rule estimates in hand, we turn to estimation of the IS equation and the effect of the 1979–80 regime change on our estimates of structural parameters. In this section we estimate the optimizing IS equation for the UK and Australia for the full sample (Section 4.1) and post-1979 (Section 4.2).

4.1 Specification and estimation

The optimizing IS equation that we are interested in estimating takes the form:

$$y_t = b_0 + E_t y_{t+1} - \sigma r_t + s_g(g_t - E_t g_{t+1}) + v_t\,, \qquad (4)$$

where y_t is log real GDP, r_t is the ex ante short-term real interest rate (defined as $R_t - E_t\pi_{t+1}$) , g_t is log real government purchases, and v_t is a preference shock, $\sigma > 0$, and s_g is the steady-state share of government purchases in GDP. Relative to the IS equation (1), this formulation allows explicitly for the presence of a public sector. We discuss the inclusion of government spending in the specification further below.

The underlying source of this equation deserves attention before we proceed to estimation. Formally, equation (4), like (1), is a semi-logarithmic linear approximation of a nonlinear Euler equation, where the linearization is conditional on constant steady-state values of ($y_t - E_t y_{t+1}$) and r_t. In practice, the steady-state growth rate of output, $E(\Delta y_t)$, appears to have undergone discrete changes during our 40-year sample period; in particular, it appears to be lower in both the UK and Australia for the period after 1973 than before 1973. The linear approximation (4) could continue to be valid provided the mean of r_t shifted along with that of Δy_t. Neoclassical growth theory does suggest that the means of r_t and Δy_t should shift together; in that case, the intercept term in equation (4) would be unchanged even in the face of changes in $E(\Delta y)$. Specifically, the nonlinear Euler equation underlying equation (4) takes means into account in the sense that it implies the following relation between the means (letting Y_t denote the unlogged level of output and assuming a constant s_g):

$$E[(Y_t/Y_{t+1})] = E[(1+r_t)^{-\sigma}] \Rightarrow E[1-\Delta y_{t+1}] = E[(1+r_t)^{-\sigma}] \Rightarrow E[\Delta y] \approx \sigma E(r), \qquad (5)$$

so DSGE models imply that changes in the steady-state growth of output should be proportional to changes in the steady-state real interest rate. See Laubach and Williams (2001) for a recent study of US data that exploits condition (5) to measure changes in the steady-state natural interest rate.[14]

In practice, however, there are grounds for believing that changes in the steady-state real interest rate and economic growth rate have not been closely related in the last 40 years. Financial liberalization and deregulation during the 1980s were probably associated with a one-time rise in the measured value of r_t that simply reflected the removal of regulations favouring the holding of government debt. Thus, the effective real interest rate relevant to household spending decisions in the 1970s was probably higher than suggested by government security rates, and, correspondingly, the rise in effective real rates in the 1980s was more muted than the rise in recorded real rates. And much of the fall in the average value of r_t in the 1990s in the UK, Australia, and other inflation-targeting countries may have reflected a decrease in the risk premium arising from the change to a more stable monetary policy environment, rather than a decrease in the risk-free ex-ante real interest rate, which appears in the model above.[15] Moreover, Prescott (1986, p. 30) expresses doubt that the sample mean of the real Treasury bill rate is a good approximation of the steady-state real return in a neoclassical growth model—reducing the likelihood that the means of Δy_t and r_t will be

[14] In econometric terms, the Δy_t and r_t series would be co-breaking (Hendry and Mizon, 2000).

[15] A similar property has been found elsewhere. Examining euro area data for 1982–97, Gerlach and Schnabel (2000, p. 166) observe that ‘countries where the credibility of monetary policy is low have had relatively high ex-post real interest rates’.

linked. All in all, it seems appropriate to decouple the means of the two series in empirical work. We therefore estimate an interest-elastic, forward looking aggregate demand function, while breaking the link between the steady-state growth rate and steady-state real rate.[16] We do so by estimating equation (4) as follows:

$$y_t = \mathrm{E}_t y_{t+1} + b_0 + b_1(R_t - \mathrm{E}_t\pi_{t+1}{}^a) + b_2 D7302_t + s_g(g_t - \mathrm{E}_t g_{t+1}), \tag{6}$$

where $D7302_t$ is an intercept dummy equal to 1.0 from 1973 Q4 onwards. For estimation, we have also substituted in the Fisher relation, and annualized both the nominal interest rate and the quarterly inflation series. These nominal variables therefore correspond to those used in estimating the policy rules in the previous section. The expected signs in (6) are $b_1 < 0$ and $b_2 > 0$. We note that equation (6) remains valid in an open-economy environment, provided that the interest elasticity is interpreted as capturing the net-trade channel of monetary policy, and the IS disturbance term is interpreted as including external shocks (Neiss and Nelson, 2003).

Our estimates of (6) by instrumental variables for the UK and Australia are given in Table 4. The instrument set consists of lagged values of the endogenous variables as well as dummies for various 'events' that shifted the price level (mainly changes in indirect taxes).[17] The use of lagged values as instruments involves the assumption of white-noise shocks, an assumption not made for the US in McCallum and Nelson (1999). The white-noise assumption seemed more appropriate for the two countries we study, in view of the lower serial correlation of Δy_t in the UK and Australia.[18] For both countries, in instrumenting for the nominal interest rate, we allow the coefficient on inflation to change after 1979–80. This captures an important aspect of the policy regime change of that period, and improves the efficiency of the estimates, while staying within an instrumental-variables framework.

The interest elasticity for the UK is highly significant[19] and in close agreement with Nelson and Nikolov (2002), who did not allow for an IS intercept-shift in 1973. The

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[16] For Δy this is implicitly done in studies, such as those of EF (2000, 2003) and Rotemberg and Woodford (1997), that detrend the y_t series prior to estimating the Euler relationship.

[17] These dummies are those used earlier in our policy rule estimates. See Neiss and Nelson (2002) for more discussion of the price-level dummies for each country.

[18] For example, quarterly OLS regressions of Δy_t on a constant produced Durbin Watson statistics of 2.08 for Australia and 2.04 for the UK (and corresponding first-order autocorrelations of Δy of −0.04 and −0.03 respectively). These statistics will be indicative of the serial correlation of the IS shock when the value of b_1 is low, which Table 1 suggests is realistic.

[19] We report conventional *t*-ratios and use them to evaluate significance. The estimation programmes we used did not generate automatic adjustments of the *t*-statistics for heteroscedasticity; but we compared the conventional *t*-ratios to approximately valid adjusted ratios and found little difference.

Table 4. IV estimates of the optimizing IS equation			
	United Kingdom Sample period 1957 Q1–2002 Q4	Australia Sample period 1962 Q1–2002 Q2	Australia Sample period 1962 Q1–2002 Q2
b_0	–0.0057 (t = 4.36)	–0.0103 (t = 5.21)	–0.0103 (t = 5.15)
b_1	–0.0844 (t = 3.92)	–0.0455 (t = 1.66)	–0.0508 (t = 1.84)
b_2	0.0029 (t = 1.77)	0.0058 (t = 2.38)	0.0062 (t = 2.49)
Dummy for 1974 Q2	—	—	–0.0246 (t = 2.66)
s_g	0.22[a]	0.23[a]	0.23[a]

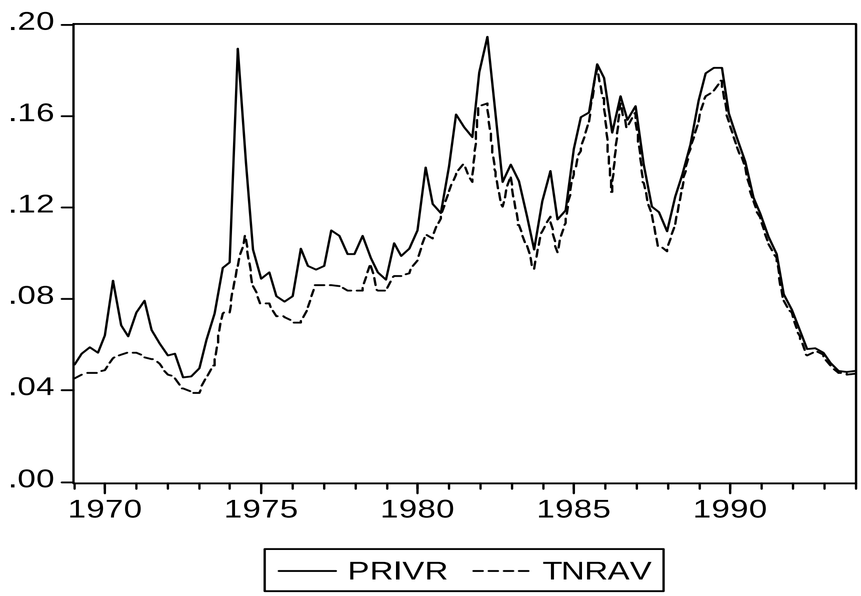
a Parameter imposed based on first-moment information (the sample mean of the government consumption/GDP share).

Instruments:
UK: $\pi_{t-1}, \pi_{t-2}, \pi_{t-3}, R_{t-1}, R_{t-2}, R_{t-3}, y_{t-1}, y_{t-2}, y_{t-3}, g_{t-1}, g_{t-2}, g_{t-3}$, dummies, constant.
Australia: $\pi_{t-1}, R_{t-1}, y_{t-1}, g_{t-1}$, lag 1 of log US output, dummies, constant.
The coefficient on π_{t-1} in the instrumenting for R_t is allowed to change after 1979 Q1 (UK) or 1979 Q4 (Australia).

shift term approaches significance for the UK and is highly significant for Australia.[20] The estimated value of the interest elasticity is imprecisely estimated (t = 1.7) for Australia, but similar in magnitude to the values for the US in Table 1, noting that the Table 1 estimates should be divided by four in making comparisons. Some of the imprecision of the interest elasticity reflects the influence of a single observation—the second quarter of 1974, when output fell sharply. During this episode, the government's monetary squeeze was felt more strongly in monetary aggregates and returns on private loans than on government paper, as Figure 1, which plots the Treasury note rate against a private rate of the same maturity, shows. The model underlying equation (1), in which different non-money assets are perfect substitutes, does not allow for the character of this squeeze.[21] As Table 4 shows, including a

[20] The lower significance of the 1973 shift in the UK partly reflects the fact that this shift was partially reversed in the 1980s. For this reason, when we estimate backward-looking IS equations for the UK using detrended output as the dependent variable, some of our results will allow for trend-breaks in both 1973 and 1981.
[21] Note that backward-looking models of the Rudebusch-Svensson type share this limitation.

Figure 1: Treasury note rate and private bill rate in Australia



dummy for the 1974 Q2 observation has little effect on the estimated interest elasticity in the IS function, but increases slightly the precision of the estimate.

4.2 Estimates following 1979–80 regime break

We now consider the issue of regime break by the method suggested in Section 2. Specifically, we re-estimate the IS equations on data following the monetary policy regime break of 1979–80, and compare the full-sample and single-regime estimates. Our post-1979 estimates are given in Table 5. For the UK, the estimated interest elasticity is again negative and highly significant. The point estimate is larger in absolute value, but the difference with the full-sample estimate is not statistically significant.[22] This direct examination is therefore supportive of stability of the IS equation for the UK. A similar conclusion holds for Australia: the interest-elasticity estimate remains negative and is not different statistically from the full-sample value, although, in this case, the precision of the estimate is low.

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[22] The confidence intervals for the full-sample and post-1979 interest elasticities substantially overlap. In addition, we tested the significance of a post-1979 slope-dummy variable, defined as the real interest rate series generated by the full-sample estimates of equation (4) multiplied by a dummy equal to 1.0 from 1979 Q2. The slope-dummy variable had a reported *t*-statistic of 1.68, which overstates its significance on account of the variable being a generated regressor.

Table 5. IV estimates of the optimizing IS equation after 1979		
	United Kingdom Sample period 1979 Q2–2002 Q4	Australia Sample period 1980 Q1–2002 Q2
b_0	0.0002 (t = 0.29)	–0.0052 (t = 2.74)
b_1	–0.1319 (t = 5.17)	–0.0264 (t = 0.70)
b_2	—	—
s_g	0.22[a]	0.23[a]

a Imposed parameter.

Instruments:

UK: π_{t-1}, π_{t-2}, π_{t-3}, R_{t-1}, R_{t-2}, R_{t-3}, y_{t-1}, y_{t-2}, y_{t-3}, g_{t-1}, g_{t-2}, g_{t-3}, dummies, constant.

Australia: π_{t-1}, R_{t-1}, y_{t-1}, g_{t-1}, lag 1 of log US output, dummies, constant.

In Table 6 we present some robustness results. First, we investigate the implications of omitting the government-expenditure term. The formal derivation of the IS equation (4) from optimizing behaviour suggests that this term should be included (see e.g. Walsh, 2003, Woodford, 2003, and references therein), but estimates for the U.S. typically leave it in the disturbance term. We have explicitly included it in our estimates and, in line with the theory, imposed its value at the steady-state share of government spending to output (obtained in practice by the sample mean of this ratio). When we instead drop this term from the estimated specification, Table 6 shows that our estimated IS equations remain very similar for both countries. For the UK, the estimated interest elasticity of aggregate demand drops slightly in significance; for Australia, it actually increases in significance, so that now the *t*-statistic for the interest elasticity does exceed 2. On the other hand, as Table 6 also shows, suppressing the government-expenditure term delivers less satisfactory results for Australia after 1980 because now the point estimate for the interest elasticity is wrongly signed. Explicitly controlling for the government expenditure term therefore does seem useful for both countries, and our preferred estimates remain those in Table 5.

Also in Table 6, we allow for a second break in the intercept of the IS equation. We include an intercept dummy equal to 1.0 from 1994 Q4, to capture the apparent break in productivity and output growth that seems to have occurred around then. The point estimate on the dummy is consistent with much of the post-1973 growth slowdown

Table 6. Robustness checks on optimizing IS equation estimates				
	Government expenditure term omitted			*Intercept change from 1994 Q4*
	United Kingdom Sample period 1957 Q1–2002 Q4	Australia Sample period 1962 Q1–2002 Q2	Australia Sample period 1980 Q1–2002 Q2	Australia Sample period 1962 Q1–2002 Q2
b_0	–0.0068 (t = 5.28)	–0.0132 (t = 7.20)	–0.0084 (t = 4.42)	–0.0103 (t = 5.15)
b_1	–0.0789 (t = 3.70)	–0.0534 (t = 2.08)	0.0045 (t = 0.12)	–0.0509 (t = 1.84)
b_2	0.0031 (t = 1.92)	0.0072 (t = 3.18)	—	0.0070 (t = 2.70)
Dummy for 1974 Q2	—	–0.0240 (t = 2.57)	—	–0.0247 (t = 2.66)
s_g	0^a	0^a	0^a	0.23^a
Intercept shift, 1994 Q4 on	—	—		–0.0030 (t = 1.05)

a Imposed parameter.

Instruments:
UK: π_{t-1}, π_{t-2}, π_{t-3}, R_{t-1}, R_{t-2}, R_{t-3}, y_{t-1}, y_{t-2}, y_{t-3}, g_{t-1}, g_{t-2}, g_{t-3}, dummies, constant.
Australia: π_{t-1}, R_{t-1}, y_{t-1}, g_{t-1}, lag 1 of log US output, dummies, constant.
The coefficient on π_{t-1} in the instrumenting for R_t is allowed to change after 1979 Q1 (UK) or 1979 Q4 (Australia).

being reversed; however, inclusion of the dummy has no effect on our estimate of the interest elasticity of the IS equation.

5. Comparison with backward-looking specifications

The backward-looking IS equation that EF study is that of Rudebusch and Svensson (1999), which may be written

$$y^d_t = a_0 + \alpha_1 y^d_{t-1} + \alpha_2 y^d_{t-2} + \beta\, r^b_{t-1} + u_t, \qquad (7)$$

where $r^b_t = \frac{1}{4}(\Sigma_{i=0}^3 [R_{t-i} - \pi^a_{t-i}])$. Specification (7) is 'backward-looking': there are no terms in expected future output (which appear when the IS equation is derived from optimizing behaviour, as in equation (1)); nor are the lags of the interest rate or output justified from an optimization problem. The nominal interest rate appears as a lagged moving average, with the current nominal rate (which appears in the optimizing IS equation) absent. Furthermore, the nominal interest rate is converted into a real interest rate r_t^b by subtracting current quarterly inflation, rather than the rational expectation of next quarter's inflation, as occurs in the optimizing case.

Because equation (7) recognizes no forward-looking terms and has unjustified lagged terms, a DSGE-based approach to modelling would classify estimates of equation (7) as reduced-form estimates. The coefficients estimated, therefore, would not correspond to underlying preference parameters, and would be predicted by DSGE theory to shift with regime changes. According to EF's stability tests, on the other hand, equation (7) exhibits a degree of constancy over time on US data that makes it more suitable for policy analysis than formal theory would suggest.

As we have emphasized, however, an added dimension on which to determine whether an IS equation is structural is to examine estimates from other countries. To that end, we estimate the Rudebusch-Svensson IS equation on UK and Australian data. The specifications we consider are (7) plus a less restricted version,

$$y^d_t = a_0 + \Sigma_{i=1}^{j} \alpha_i \, y^d_{t-i} + \Sigma_{i=1}^{k} \beta_i \, (R_{t-i} - \pi^a_{t-i}) + u_t. \qquad (8)$$

Equation (8) allows richer lagged dynamics than (7), and also lets the lagged rates appear in a less constrained way (i.e. with different weights over a year) while still appearing as a type of 'real rate'. As it is detrended log output that appears in these IS equations,[23] we present estimates for three different empirical counterparts to y^d: those arising from Hodrick-Prescott filtering, quadratic detrending, and a broken linear trend.[24] The HP filter uses a filtering parameter of 1600; the broken trend uses break-dates of 1973 Q4 for both countries, with an additional break in 1981 Q4 for the UK.

For the UK, the estimates of the interest elasticity (the sum of coefficients on r^b_t) are almost all wrongly signed—positive—across sample periods, detrending methods for output, and whether equation (7) or (8) is estimated. The only exception is the

[23] The optimizing IS equation, on the other hand, refers to the log-level of output, or rather a forward-looking first difference of log output, hence our use of y_t as the dependent variable in the previous section.

[24] The original Rudebusch-Svensson specification used the Congressional Budget Office's estimate of the output gap as the dependent variable. As noted above, this series resembles US quadratically detrended log output, and does not correspond to the output gap concept stressed by DSGE models.

Table 7. Estimates of backward-looking IS equation for the UK

$$y^d_t = a_0 + \Sigma_{i=1}^{j} \alpha_i \, y^d_{t-i} + \Sigma_{i=1}^{k} \beta_i \, (R_{t-i} - \pi^a_{t-i}) + u_t$$

1. y^d_t = HP-filtered log output

Sample	Lags of y^d	Sum of coefficients on lags of y^d	Lags of interest-rate term	Sum of coefficients on real rate
1957 Q1–2002 Q4	1–2	0.790 (t = 16.29)	1–4 with equal weights	0.009 (t = 0.53)
1979 Q2–2002 Q4	1–2	0.888 (t = 16.75)	1–4 with equal weights	0.039 (t = 0.03)
1957 Q1–2002 Q4	1–4	0.717 (t = 13.71)	1–4	0.003 (t = 0.21)
1979 Q2–2002 Q4	1–4	0.819 (t = 15.40)	1–4	0.029 (t = 0.95)

2. y^d_t = Quadratically detrended log output

Sample	Lags of y^d	Sum of coefficients on lags of y^d	Lags of interest-rate term	Sum of coefficients on real rate
1957 Q1–2002 Q4	1–2	0.955 (t = 38.52)	1–4 with equal weights	0.012 (t = 0.59)
1979 Q2–2002 Q4	1–2	0.976 (t = 38.11)	1–4 with equal weights	0.082 (t = 2.17)
1957 Q1–2002 Q4	1–4	0.928 (t = 36.27)	1–4	–0.008 (t = 0.39)
1979 Q2–2002 Q4	1–4	0.944 (t = 37.76)	1–4	0.055 (t = 1.62)

3. y^d_t = Broken-trend log output[a]

Sample	Lags of y^d	Sum of coefficients on lags of y^d	Lags of interest-rate term	Sum of coefficients on real rate
1957 Q1–2002 Q4	1–2	0.886 (t = 24.74)	1–4 with equal weights	0.005 (t = 0.29)
1979 Q2–2002 Q4	1–2	0.935 (t = 27.31)	1–4 with equal weights	0.060 (t = 1.67)
1957 Q1–2002 Q4	1–4	0.847 (t = 22.99)	1–4	0.005 (t = 0.29)
1979 Q2–2002 Q4	1–4	0.896 (t = 27.29)	1–4	0.056 (t = 1.72)

a. y^d is obtained by regressions of log real GDP on a constant, a time trend, and each of these terms multiplied by dummy variables equal to 1.0 after 1973 Q3 and 1981 Q3.

Table 8. Estimates of backward-looking IS equation for Australia:

$$y^d_t = a_0 + \Sigma_{i=1}^{j} \alpha_i \, y^d_{t-i} + \Sigma_{i=1}^{k} \beta_i \, (R_{t-i} - \pi^a_{t-i}) + \Sigma_{i=1}^{l} \gamma_i \, y^{usd}_{t-i} + u_t$$

Sample	Lags of y^d	Lagged y^d coefficient sum	Lags of interest rate term	Coefficient sum on real rate	Lags of US output	Sum on US output
1. y^d_t , y^{usd}_t measured by Hodrick-Prescott filtered log output						
1962 Q4–2002 Q3	1–2	0.644 (t = 9.70)	1–4 with equal weights	0.008 (t = 0.37)	—	—
1980 Q1–2002 Q3	1–2	0.775 (t = 13.94)	1–4 with equal weights	–0.001 (t = 0.02)	—	—
1962 Q4–2002 Q3	1–4	0.539 (t = 7.09)	1–4	0.008 (t = 0.38)	—	—
1980 Q1–2002 Q3	1–4	0.718 (t = 11.66)	1–4	–0.000 (t = 0.004)	—	—
1962 Q4–2002 Q3	1–4	0.519 (t = 6.12)	1–4	0.004 (t = 0.21)	1–4	0.133 (t = 1.93)
1980 Q1–2002 Q3	1–4	0.617 (t = 6.72)	1–4	–0.001 (t = 0.03)	1–4	0.186 (t = 2.15)
2. y^d_t , y^{usd}_t measured by quadratically detrended log output						
1962 Q4–2002 Q3	1–2	0.950 (t = 30.13)	1–4 with equal weights	–0.002 (t = 0.07)	—	—
1980 Q1–2002 Q3	1–2	0.954 (t = 33.45)	1–4 with equal weights	–0.038 (t = 1.07)	—	—
1962 Q4–2002 Q3	1–4	0.938 (t = 28.35)	1–4	–0.010 (t = 0.33)	—	—
1980 Q1–2002 Q3	1–4	0.946 (t = 31.65)	1–4	–0.045 (t = 1.26)	—	—
1962 Q4–2002 Q3	1–4	0.897 (t = 19.81)	1–4	–0.014 (t = 0.47)	1–4	0.107 (t = 2.09)
1980 Q1–2002 Q3	1–4	0.889 (t = 14.95)	1–4	–0.069 (t = 1.94)	1–4	0.111 (t = 1.61)

Note: The sample for the detrending is 1954 Q1–2002 Q4 for US real GDP, 1959 Q3–2002 Q4 for Australia real GDP.

estimates for equation (8) for the full sample when quadratically detrended log GDP is the dependent variable. Here, however, the interest elasticity estimate is highly insignificant statistically, in contrast to the corresponding full-sample estimates for the UK optimizing IS equation. Moreover, as Table 7 shows, the interest elasticity coefficients obtained with this dependent variable break into sizeable and quite statistically significant *positive* values when estimated on post-1979 data, hardly the behaviour of a structural parameter estimate.

For Australia, estimates of equation (7) and (8) generally deliver interest elasticity estimates that are either wrongly signed (the full-sample estimates) or highly statistically insignificant (the post-1979 estimates). Somewhat better results, though still far from statistical significance, arise with y^d measured by quadratically detrended output.

We have so far omitted explicit open-economy terms from our estimated specifications. For the optimizing IS equation, this neglect can be justified on theoretical grounds, as noted earlier. For the backward-looking IS equation, the specification choice reflects prior findings for the UK that neither the real exchange rate nor world output enters significantly when included (see e.g. Goodhart and Hofmann, 2003, and references therein). For Australia, the omission of terms-of-trade or real exchange rate variables reflects the fact that, again, these variables are rarely close to statistical significance, especially when long sample periods are considered (see e.g. Robinson, Stone, and van Zyl, 2003). Beechey *et al* (2000) report, however, that US output enters significantly in a backward-looking IS equation. We add detrended US output to equation (8), and report our estimates for Australia in Tables 8 and 9.

The foreign-output term generally enters significantly and positively across samples. With this term included, and with y^d measured as quadratically detrended log output, the Rudebusch-Svensson specification does have some respectable properties: quite a significant post-1979 interest elasticity estimate (t = 1.9) and little change in the parameter estimates across samples—although the point estimate of the interest elasticity diminishes fairly sharply when pre-1980 data are included. Since this specification is the only backward-looking IS equation that we have found that appears to exhibit parameter estimates that are both interpretable and constant, we further explored it by adding the dummy for 1974 Q2 that we found important in Table 4. When included, the coefficient estimate on this dummy is negative and highly significant (t = 3.1), while the estimated short-run interest elasticity increases in absolute value from –0.014 to –0.029, closer to the post-1979 value of –0.069. However, the t-statistic for the interest-elasticity estimate is still only 1.0, compared to the 1.8 we found for the forward-looking specification.

Table 9. Estimates of backward-looking IS equation for Australia:

$$y^d_t = a_0 + \Sigma_{i=1}^{j}\ \alpha_i\ y^d_{t-i} + \Sigma_{i=1}^{k}\ \beta_i\ (R_{t-i} - \pi^a_{t-i}) + \Sigma_{i=1}^{l}\ \gamma_i\ y^{usd}_t + u_t$$

y^d_t, y^{usd}_t measured by broken-trend log output

Sample	Lags of y^d	Lagged y^d coefficient sum	Lags of interest rate term	Coefficient sum on real rate	Lags of US output	Sum on US output
1962 Q4–2002 Q3	1–2	0.750 (t = 13.37)	1–4 with equal weights	0.005 (t = 0.18)	—	—
1980 Q1–2002 Q3	1–2	0.857 (t = 18.96)	1–4 with equal weights	0.023 (t = 0.66)	—	—
1962 Q4–2002 Q3	1–4	0.713 (t = 11.56)	1–4	0.005 (t = 0.19)	—	—
1980 Q1–2002 Q3	1–4	0.836 (t = 17.27)	1–4	0.023 (t = 0.69)	—	—
1962 Q4–2002 Q3	1–4	0.655 (t = 8.91)	1–4	–0.001 (t = 0.06)	1–4	0.134 (t = 2.48)
1980 Q1–2002 Q3	1–4	0.662 (t = 7.38)	1–4	0.019 (t = 0.57)	1–4	0.211 (t = 2.84)

Note: For both the US and Australia, y^d is obtained by regressions of log real GDP on a constant, a time trend, and each of these terms multiplied by a dummy variable equal to 1.0 after 1973 Q3. The sample for the detrending is 1954 Q1–2002 Q4 for US real GDP, 1959 Q3–2002 Q4 for Australia real GDP.

Perhaps more important, the reasonable results we obtain for Australia when GDP is quadratically detrended are not robust to alternative detrending procedures. The backward-looking equations uniformly deliver positive or virtually-zero interest elasticity estimates when output is detrended by HP filtering or (as in Table 8) by broken trends.

We therefore conclude that when full-sample and post-1979 estimates are compared directly, backward-looking IS equations seem to be less satisfactory and less constant than the forward-looking alternative. For both Australia and the UK, the key IS equation parameter (the interest elasticity) appears to exhibit better properties—full-sample statistical significance, robustness to including pre-1980 observations, and correct signs—when the optimizing IS equation is estimated.

While most of the interest elasticity estimates for the backward-looking IS equations are wrongly signed, few of them are *significantly* different from negative values. Why then do we place so much emphasis on the wrong signs of the point estimates? We do so for two reasons. First, IS equations of the type we discuss in this paper are principally employed for policy analysis—especially stochastic simulation exercises where they are used in conjunction with a Phillips curve and alternative policy rules—and, in these exercises, it is the point estimates of the parameters that are used. Wrongly-signed coefficient estimates then render the IS equation ‘unusable’. The incorrect signs in the empirical IS equations in this section therefore undermine the usefulness of these equations for policy analysis. Secondly, non-optimizing equations are often advocated on the grounds that empirical estimates of the key elasticities in optimizing models are wrongly signed. By contrast, we find for the UK and Australia that it is the backward-looking equations that have severe problems with wrong signs.

It should also be noted that, notwithstanding the wrong signs, most of the IS equations in this section ‘fit’ better, in the sense of lower residual standard deviation, than the optimizing IS equations estimated in Section 4.[25] But, in contrast to EF, and in the spirit of Lucas (1988), we do not regard this concept of fit as an appropriate criterion for judging the validity of an IS equation. The backward-looking IS equations have a built-in good fit by their specification of generous, unrestricted lags of the dependent variable. What should be used as a criterion is something that does not automatically arise from the specification: the constancy across sub-samples of the key structural parameter, namely the real interest elasticity of aggregate demand. By this criterion, we have found more support for the optimizing IS equation than the backward-looking alternative.

6. Conclusion

In this paper, we have offered an alternative perspective on the constancy of empirical optimization-based models to the judgement of Estrella and Fuhrer (2003). In the spirit of Lucas (1988), we have evaluated the constancy of both optimizing and backward-looking IS equations in terms of the robustness of a key parameter estimate to the inclusion of observations from a different monetary policy regime. We have also concentrated on international evidence—quarterly estimates for the UK and Australia, instead of the US. The result of this comparison suggests that empirical optimizing IS equations are more stable and interpretable than are the backward-looking alternatives. The use of current DSGE models in empirical work does deliver material benefits, in the form of equations that are more suitable for monetary policy analysis.

[25] Direct comparisons of residual standard deviations are difficult as the forward-looking IS equation pertains to the log-level or log-difference of output, while the backward-looking equation refers to a detrended or filtered version of log output.

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