Available online at www.sciencedirect.com
ScienceDirect
Journal of International Money and Finance 27 (2008) 40–53
Journal of International Money and Finance
www.elsevier.com/locate/jimf

Regime shifts and the stability of backward-looking Phillips curves in open economies

Efrem Castelnuovo*

Department of Economics, University of Padua, Via del Santo 33, I-35123 Padova (PD), Italy

Abstract

We assess the stability of open-economy backward-looking Phillips curves estimated over two different exchange rate regimes. We calibrate a new-Keynesian monetary policy model and employ it for producing artificial data. A monetary policy break replicating the move from a Target-Zone regime to a Free-Floating regime implemented in Sweden in 1992 is modeled. We employ two different, plausibly calibrated Taylor rules to describe the Swedish monetary policy conduct, and fit a reduced-form Phillips curve to the artificial data. While not rejecting the statistical relevance of the Lucas critique, we find that its economic importance does not seem to be overwhelming.

JEL classification: E17; E52; F41

Keywords: Lucas critique; Forwardness; Backward-looking Phillips curves; Exchange rates; Chow test

1. Introduction

Since the publication of the seminal paper by Lucas (1976), many researchers have explicitly embedded forward-looking expectations in their policy models. One of the fields that has been intensely affected by this push towards micro-foundation is the monetary one (e.g. Woodford, 2003). Interestingly, a different strand of this literature (e.g. Rudebusch and Svensson, 1999, 2002; Ball, 1999, 2000; Onatski and Stock, 2002; Laubach and Williams, 2003; Fagan

* Tel.: +39 049 827 4257; fax: +39 049 827 4211.
E-mail address: efrem.castelnuovo@unipd.it

0261-5606/$ - see front matter © 2007 Published by Elsevier Ltd.
doi:10.1016/j.jimonfin.2007.09.004

et al., 2005) has relied on ad-hoc backward-looking frameworks. In fact, backward-looking models tend to offer quite a good fit of the data. Moreover, their impulse responses closely resemble those stemming from structural VARs, an issue that pure forward-looking models have some troubles in dealing with (Estrella and Fuhrer, 2002).

Evidently, at a theoretical level backward-looking models are affected by the Lucas (1976) critique. The argument goes as follows. If agents are forward-looking they will adjust their expectations once a policy change is credibly announced. As a consequence, *reduced-form* coefficients will be unstable under a change in the policy regime. Then, a policy analysis performed with reduced-form coefficients may be severely mis-leading. This is true in principle: but how important is this change in practice?

While some researchers have undertaken empirical efforts to answer this question in a closed-economy set-up (e.g. Lindè, 2001; Estrella and Fuhrer, 2003; Rudebusch, 2005), to the best of our knowledge the only contribution dealing with this issue in an open-economy framework dates back to Taylor (1989). This is somewhat surprising, given the increasing openness in terms of trade and flows of resources conveyed in the international financial markets observed in several countries in the last decades (Lane, 2001).[1]

The aim of this paper is that of 'updating' Taylor's (1989) contribution. Taylor (1989) employs an estimated macro-model for simulating the shift from a 'fixed' to a 'flexible' nominal exchange rate in some industrialized countries. Once done so, he fits some reduced-form schedules (mainly demand and supply curves) to such simulated data, and compares the estimated parameters under the first regime to those estimated under the second one. Taylor (1989) observes that the differences in magnitude between those parameters are not really large, and concludes that the Lucas critique does not find a large support in the data. Notice that, in performing his analysis, Taylor (1989) does not use any statistical tool for assessing the stability of the estimated coefficients.

We refine the contribution by Taylor (1989) along two main dimensions. First, we employ a modern new-Keynesian DSGE open-economy monetary policy model in the spirit of the one proposed by Svensson (2000). In this model, the monetary policy makers follow a Taylor-type rule and aim at minimizing the oscillations of inflation, real GDP, and nominal exchange rate around their targets. As a second difference with respect to Taylor (1989), we assess both the *statistical* and the *economic* relevance of the Lucas critique. In this sense, we line up with Rudebusch (2005) in acknowledging the importance of employing a formal test for assessing the instability of the estimated schedules from a *statistical* perspective. Importantly, as underlined by Rudebusch (2005) this is not the only dimension one may want to consider in this kind of exercise. In fact, the macroeconometrician is also called to evaluate the pros and cons of employing a reduced-form model estimated under a policy break from an *economic* perspective. To do so, we compare the mean values and the distributions of the estimated coefficients in the two different subsamples we focus on.

For our exercise to be interesting, we need to calibrate our model in a plausible fashion. Interestingly, Sweden experienced a dramatic change in regime in 1992, when it moved from an exchange rate target-zone (TZ hereafter) to a flexible exchange rate regime (FF henceforth). Cúrdia and Finocchiaro (2005) split the Swedish sample 1980Q1–2003Q3 into two subsamples, i.e. 1980Q1–1992Q4 (TZ) and 1993Q1–2003Q3 (FF), estimate regime-specific Taylor

[1] Here we refer to contributions that are very closely related to our object of investigation, i.e. the empirical relevance of the Lucas critique for *backward-looking monetary policy* models. In general, the quantitative importance of the Lucas critique has been subject to wide attention since 1976. For a survey, see Ericsson and Irons (1995).

rules, and find that there are remarkable differences between the two sub-sets of estimated values.[2] We calibrate our Taylor rules with their point estimates in order to simulate a 'plausible' regime shift, i.e. a regime shift historically occurred in an open-economy such as Sweden. We calibrate the rest of our new-Keynesian model by borrowing parameter values from the recent empirical literature on monetary policy models for the Swedish economy, i.e. Leitemo and Söderström (2005), Lindé et al. (2004), and Cúrdia and Finocchiaro (2005).

In our exercise, we focus our attention on OLS estimated reduced-form Phillips curve. We do so to contribute to the recent discussion on inflation dynamics and its formalization, discussion that has led some authors to prefer the 'accelerationist' version of the Phillips curve (e.g. Mankiw, 2001; Estrella and Fuhrer, 2002) to the micro-founded, expectations-equipped new-Keynesian schedule (e.g. Woodford, 2003).

Our results acknowledge the statistical importance of the Lucas critique, i.e. the reduced-form Phillips curve we estimate tend to be unstable. However, the magnitude of the changes in the estimated coefficients leads us to think that the Lucas critique might not necessarily be of overwhelming importance for policy analysis purposes, at least as far as the Swedish case is concerned.

The paper is structured as follows. Section 2 outlines the small macro-model we employ to produce the simulated time-series of interest. Section 3 contains an explanation of the steps we implement to perform our econometric exercise. In Section 4 we present our findings, whose robustness is discussed in Section 5. Section 6 concludes, and References follow.

2. A simple open-economy macro-model

The open-economy framework we employ is basically the one put forward by Svensson (2000). In this set-up, the paths of the domestic inflation rate and the output gap are defined as follows:

$$\pi_{t+1} = \mu_\pi E_t \pi_{t+2} + (1-\mu_\pi)\pi_t + \alpha_y y_t + \alpha_q E_t q_{t+1} + u_{t+1}, \tag{1}$$

$$y_{t+1} = \mu_y E_t y_{t+2} + \left(1-\mu_y\right) y_t - \beta_r (i_t - E_t \pi_{t+1}) + \beta_q q_t + \beta_y y_t^* + v_{t+1}, \tag{2}$$

where π_t is the annualized quarterly inflation, y_t is the output gap (i.e. the log-difference between the real GDP and a measure of potential output), q_t is the real exchange rate, i_t is the short-term nominal interest rate controlled by the Central Bank, u_t and v_t are iid processes with zero mean and standard deviations σ_u and σ_v, and y_t^* is the foreign output gap (as, in general, starred variables refer to foreign variables).

Eq. (1) is an open-economy version of a stochastic hybrid Phillips curve, in which the inflation rate is pre-determined one period, it is endogenously inertial (e.g. due to price indexation, Christiano et al., 2005), it takes into account the effect of expected costs of imported intermediate inputs via the real exchange rate fluctuations, and it allows inflation to be hit by a 'cost push shock' u_{t+1}.[3] Eq. (2) defines the path of the output gap, which is caused by expectations on future output gap's realizations as well as past values (the latter finding their rationale, e.g. habit formation, as in Fuhrer, 2000), the ex-ante real interest rate, the real exchange rate

[2] The instability of a Taylor rule estimated over a comparable sample for Sweden is also certified by the empirical exercise conducted by Sgherri (2005), who finds a policy break at the beginning of the 1990s.

[3] In this model an increase of the nominal/real exchange rate stands for *depreciation.*

(which approximates the increased demand for domestic goods driven by exchange rate depreciation) and the foreign output gap, which captures the increased demand for domestic goods due to the expansions of the foreign business cycle.[4]

The evolution of the nominal exchange rate s_t is described by the following hybrid stochastic version of the uncovered interest parity (UIP) condition:

$$i_t = i_t^* + \mu_s E_t s_{t+1} + (1-\mu_s)s_{t-1} - s_t + \varphi_t, \tag{3}$$

where the risk-premium φ_t is shaped as an AR(1) process with root ρ_ψ and a zero-mean stochastic error ψ_t whose standard deviation is identified by σ_ψ.[5] We capture backward-looking exchange rate expectations (Frankel and Froot, 1987) by allowing for the parameter μ_s to assume a value smaller than 1. Clearly, when $\mu_s = 1$ we go back to the textbook UIP condition.

As indicated above, one of the arguments (potentially) of interest for the central banker is the CPI inflation rate π_t^{CPI}, which is defined as

$$\pi_t^{\mathrm{CPI}} = (1-\chi)\pi_t + \chi\pi_t^{\mathrm{M}}, \tag{4}$$

where χ is the weight of imported goods in the aggregate consumption basket, and π_t^{M} stands for imported inflation. Following Leitemo and Söderström (2005), we define the imported price level p_t^{M} as follows:

$$p_t^{\mathrm{M}} = (1-\theta)p_{t-1}^{\mathrm{M}} + \theta\left(p_t^* + s_t\right). \tag{5}$$

Importantly, the parameter θ allows for the possibility of deviating from the law of one price *in the short-run*. In fact, if $0 \le \theta < 1$, then the imported price level does not immediately fully adjust after a shock has hit the foreign inflation rate or the nominal exchange rate. This price stickiness is intended to capture the imperfection of the exchange rate pass-through observed in the real world, imperfection that tends to be much less important in the long run, as shown by Campa and Goldberg (2005).

Since the real exchange rate q_t is defined as

$$q_t = s_t + p_t^* - p_t, \tag{6}$$

Eqs. (4)–(6) suggest the following link between real exchange rate and CPI inflation:

$$\pi_t^{\mathrm{CPI}} = (1-\chi)\pi_t + \chi\left[(1-\theta)\pi_{t-1}^{\mathrm{M}} + \theta(\pi_t + \Delta q_t)\right], \tag{7}$$

which makes it clear that (the change of) the real exchange rate exerts an impact over CPI inflation.

As far as the Rest-Of-the-World (ROW henceforth) is concerned, in this framework the monetary authorities follow a Taylor rule, i.e.

[4] Note that the steady-state value of the real exchange rate q_t in this model is equal to zero, hence the model is consistent with the natural rate hypothesis. The lagged impact of the real exchange rate on the domestic output gap is due to our willingness of avoiding the contemporaneous presence of the current and the expected domestic policy rate in the IS equation, which would render the set-up of problem non-standard.

[5] We shape the stochastic component φ_t as an AR(1) process to capture the commonly observed persistence of the risk-premium, as in Svensson (2000), and Leitemo and Söderström (2005).

$$i_t^* = \left(1 - \rho_*^i\right)\left(\psi_*^\pi \pi_t^* + \psi_*^y y_t^*\right) + \rho_*^i i_{t-1}^* + \zeta_t^*, \tag{8}$$

where f_π^* and f_y^* are the coefficients, respectively, associated to foreign inflation and foreign output gap, ρ_*^i is the interest rate smoothing coefficient, while ζ_t^* is a zero-mean white noise process with variance σ_ζ^*. To catch the persistence typically observed in macro-data, π_t^* and y_t^* are defined as AR(1) processes, i.e.

$$\pi_{t+1}^* = \rho_\pi^* \pi_t^* + u_{t+1}^*, \tag{9}$$

$$y_{t+1}^* = \rho_y^* y_t^* + v_{t+1}^*, \tag{10}$$

with u_{t+1}^* and v_{t+1}^* being iid processes whose variances are, respectively, σ_u^* and σ_v^*.

2.1. Monetary policy

The monetary authorities' behavior closes the model. We aim at simulating the shift from the 'Target-Zone' exchange rate regime experienced by Sweden in the pre-1992 period to the 'Free-Floating' regime adopted later on. In our framework, the different monetary policy regimes are identified by different Taylor rules. In particular, under the TZ regime we represent the monetary policy conduct as follows:

$$i_t = (1 - \rho_{\mathrm{TZ}})\left(\psi_{\mathrm{TZ}}^\pi \pi_t^{\mathrm{CPI}} + \psi_{\mathrm{TZ}}^y y_t + \psi_{\mathrm{TZ}}^s \Delta s_t\right) + \rho_{\mathrm{TZ}} i_{t-1}. \tag{11}$$

By contrast, under FF the policy rule reads as follows:

$$i_t = (1 - \rho_{\mathrm{FF}})\left(\psi_{\mathrm{FF}}^\pi \pi_t^{\mathrm{CPI}} + \psi_{\mathrm{FF}}^y y_t\right) + \rho_{\mathrm{FF}} i_{t-1}. \tag{12}$$

Cúrdia and Finocchiaro (2005) fit the schedules (11) and (12) to Swedish data. In particular, the Taylor rule under TZ is estimated for the subsample 1980Q1–1992Q4, while the Taylor rule under FF is estimated for the subsample 1993Q1–2003Q3. They find remarkable differences in terms of point estimates, so documenting a regime-break in the Swedish monetary policy conduct. This is the historically plausible, empirically relevant break we exploit in performing our Montecarlo exercise.

2.2. Model parameterization

As already discussed, to perform an interesting exercise from a policy perspective we need a plausible parameterization of the structural model we employ. First of all, the calibration of the policy break must rely on an observed natural experiment, in order to capture an economically meaningful regime shift. As already pointed out, the Swedish case is informative in such sense. We then rely on Cúrdia and Finocchiaro's (2005) estimates of the two different Taylor rules (11) and (12). In particular, we set $\psi_{\mathrm{TZ}}^\pi = 1.52$, $\psi_{\mathrm{TZ}}^y = 0.13$, $\psi_{\mathrm{TZ}}^s = 3.89$, $\rho_{\mathrm{TZ}} = 0.94$ to represent the Swedish monetary policy conduct under the TZ regime, and $\psi_{\mathrm{FF}}^\pi = 2.20$, $\psi_{\mathrm{FF}}^y = 0.03$, $\rho_{\mathrm{FF}} = 0.76$ under FF.[6] The benchmark parameterization used in our exercise is borrowed from

[6] The estimated coefficients for the output gap presented by Cúrdia and Finocchiaro (2005) were rescaled (divided by 4) in order to adapt them to the quarterly model in use in this paper. Cúrdia and Finocchiaro (2005) estimate a Taylor rule with the nominal exchange rate in deviations with respect to its steady-state level: we adopt their estimate for the nominal exchange rate in first differences.

some of the existing literature that concentrates on the Swedish case. The domestic economy is almost fully parameterized on the basis of the contributions by Leitemo and Söderström (2005), Cúrdia and Finocchiaro (2005), and Lindé et al. (2004). In particular, we parameterize the Phillips curve as in Leitemo and Söderström (2005) and assign a value of 0.2 to α_y and of 0.04 to α_q.[7] As far as the IS curve is concerned, we set the intertemporal elasticity of substitution β_r to 0.15 as in Leitemo and Söderström (2005), a value that seems to represent a good compromise between the low point estimate − 0.03 − proposed by Cúrdia and Finocchiaro (2005) and the relatively large point estimate − 0.32 − obtained by Lindé et al. (2004).[8] Given the evidence in favor of the impact of the ROW business cycle on the domestic one (Lindé et al., 2004), we set β_y to be equal to 0.12 as in Leitemo and Söderström (2005). We acknowledge the role played by the real exchange rate in influencing the domestic business cycle by fixing β_q to 0.05, in order to allow for a positive but moderate exchange rate channel as in Leitemo and Söderström (2005). As far as the degree of forwardness of the UIP condition is concerned, we set $\mu_s = 0.7$ to acknowledge to the nominal exchange rate its feature of 'forward-looking determined asset price' (Svensson, 2000). The risk-premium autoregressive coefficient ρ_φ is fixed to 0.3, and its volatility is set to 0.844 as in Leitemo and Söderström (2005). We set the weight of the imported goods in the aggregate consumption basket χ to 0.35, in line with the measure employed by Cúrdia and Finocchiaro (2005), and the exchange rate pass-through coefficient θ to 0.5 as in Campa and Goldberg (2005).

We take the US economy as a proxy for the Rest-Of-the-World, and OLS estimate equations (8)−(10) on the sample 1980Q1−2003Q3, i.e. the same sample considered by Cúrdia and Finocchiaro (2005) in their empirical work on the Swedish economy.[9] The benchmark calibration we employ in our exercise is reported in Table 1.

It is well known in the empirical monetary policy literature that it is very difficult to quantify the magnitudes of the coefficients related to the importance of the forward-looking expectations terms (see e.g. Canova and Sala, 2005). Therefore, we consider three different pairs of 'degrees of forwardness', identifying, respectively,

- a moderately forward-looking model, i.e. $[\mu_\pi, \mu_y] = [0.5, 0.5]$;
- a forward-looking model with highly forward-looking entrepreneurs, i.e. $[\mu_\pi, \mu_y] = [0.8, 0.5]$;
- a highly forward-looking model, i.e. $[\mu_\pi, \mu_y] = [0.8, 0.8]$.

[7] Leitemo and Söderström (2005) work with a *quarterly* inflation rate. By contrast, we work with an *annualized* inflation rate. For consistency, we rescaled their Phillips curve coefficients by multiplying them by 4.

[8] Cúrdia and Finocchiaro (2005) estimate a DSGE open-economy model for Sweden with Bayesian techniques for the sample 1980Q1−2003Q3. When referring to their 'point estimates', we actually mean their 'posterior mean estimates'. Lindé et al. (2004) match the impulse responses of a DSGE open-economy model to those of a VAR(5) with dummies, a time-trend, and exogenous ROW variables such as short-term nominal interest rate, inflation rate, and the foreign trade-weighted GDP at market prices fit to the sample 1986Q1−2002Q4. When referring to both Cúrdia and Finocchiaro (2005) and Lindé et al.'s (2004) point estimates, we mean the convolutions − of their estimated structural parameters − that correspond to our structural parameters.

[9] The database on the US economy was constructed by downloading the time-series concerning the federal funds rate (quarterly average), the real GDP, the potential output level, and the GDP deflator from the Federal Reserve Bank of St. Louis. We computed the output gap as percentage deviation of the real GDP with respect to the potential output level, and the annualized inflation rate by multiplying the quarterly growth rate of the GDP deflator by 400. The full set of OLS estimates is available upon request.

Table 1
Benchmark parameterization

Domestic economy									
Phillips curve		IS curve		UIP condition		CPI equation		Policy rules	
α_y	0.2	β_r	0.15	μ_s	0.7	χ	0.35	ψ_{TZ}^{π}	1.52
α_q	0.04	β_q	0.05	ρ_φ	0.3	θ	0.5	ψ_{TZ}^{y}	0.13
σ_u^2	1.556	β_y	0.12	σ_ψ^2	0.844			ψ_{TZ}^{s}	3.89
		σ_v^2	0.656					ρ_{TZ}	0.94
								ψ_{FF}^{π}	2.20
								ψ_{FF}^{y}	0.03
								ρ_{FF}	0.76

Foreign economy					
Phillips curve		IS curve		Taylor rule	
$\rho_{\pi*}$	0.89	ρ_{y*}	0.95	ψ_*^{π}	1.94
σ_{u*}^2	0.97	σ_{v*}^2	0.74	ψ_*^{y}	1.12
				ρ_*^{i}	0.89
				$\sigma_{\xi*}^2$	0.99

Sources of the parameters indicated in the text.

Some support for this modeling choices is provided by Cúrdia and Finocchiaro (2005), who provide evidence of a very forward-looking domestic Phillips curve and fairly forward-looking IS schedule for Sweden in the last two decades. We now move to the description of our empirical exercise.

3. Assessing the importance of the Lucas critique

In our 'in-lab' exercise we concentrate on the stability of the following reduced-form open-economy Phillips curves:

$$\pi_t = \sum_{j=1}^{4} \left(\gamma_{\pi j} \pi_{t-j} + \gamma_{yj} y_{t-j} + \gamma_{qj} q_{t-j} \right) + \xi_t^{\pi}. \tag{13}$$

Eq. (13) embeds all and no more than the variables present in the structural Phillips curve (1), and it is intended to capture its dynamics in a backward-looking fashion.[10] This is nothing but an open-economy version of the one proposed by Rudebusch and Svensson (1999, 2002) for the US case. Notably, with adequate restrictions on the coefficients γ_s, this reduced-form equation collapses to the one in Ball (1999, 2000).

To assess the empirical relevance of the Lucas critique, we proceed as follows. We estimate Eq. (13) on simulated data drawn from repeated samples with regime-specific Taylor rules, and examine the estimated autoregressive coefficients. The data are drawn from samples created by alternatively imposing the different 'degrees of forwardness' μ_π and μ_y presented in the previous section. For each model, we simulate 5000 samples. In each data sample, 100 observations

[10] It would be interesting to write (and estimate) the exact reduced-form of the structural inflation equation (1). Unfortunately, given the complicated structure of the economic model at hand, this is not feasible. In fact, that of estimating a reduced-form Phillips curve whose coefficients are complicated (and unknown) convolutions of the structural parameters of the economy is nothing but what an econometrician working with backward-looking models typically does.

Table 2
Parameters' stability: rejection rates

True model	Policy rule	Phillips curve
[μ_π, μ_y]	[1st, 2nd]	Rej. rate
Panel A: empirical size of the Chow test		
[0.5, 0.5]	TZ, TZ	0.052
	FF, FF	0.047
[0.8, 0.5]	TZ, TZ	0.047
	FF, FF	0.054
[0.8, 0.8]	TZ, TZ	0.055
	FF, FF	0.055
Panel B: stability of the reduced-forms		
[0.5, 0.5]	TZ, FF	0.695
[0.8, 0.5]	TZ, FF	0.292
[0.8, 0.8]	TZ, FF	0.213

Panel A: empirical size of the test; panel B: test on the Lucas critique. Sample size: $T = 200$; number of Montecarlo iterations: 5000. Further details on the Montecarlo exercise are presented in the main text.

are generated from a particular structural model with the policy rule TZ, then 100 observations are generated from the same model with the policy rule FF. [11] Notice that the switch from one policy rule to another is the *only* variation in the structural model from the first to the second sample, i.e. the structural equations of the data-generating process (1)–(10) are left unchanged across policy regimes. As in Taylor (1989) and Rudebusch (2005), the policy rule is assumed to be perfectly credible, i.e. we assume 'commitment' by the monetary authorities.[12] This implies that the change in the policy rule is unexpected but immediately recognized by the agents in the economy. Notice that this aspect, which unfortunately neglects the role of learning or uncertainty about the monetary policy regime, maximizes the ability to detect structural shifts.

4. Findings

We now turn to the analysis of our results. Table 2 provides some statistical evidence on the stability of the estimated reduced-form Phillips curve under the simulated policy regime shift.

Panel A displays the outcome of our investigation regarding the empirical size of the test, and indicates the 'Rejection Rates', i.e. the ratios between the number of times the Chow (1960) test statistic exceeded the 5% theoretical critical value, which in our analysis is 1.8075. Interestingly, when the null of stability is true – i.e. there is no regime shift and the same policy rule is followed in the entire sample – the proportions of rejections of the such null are very close to the theoretical 5%. This implies that the Chow test is well sized – i.e. the empirical size is close to the nominal one – and that can be used as a reliable statistical tool for detecting a break in the estimated coefficients.

Panel B displays the rejection rates computed in case of policy shift in the data-generating process. The rejection rates remarkably increase, and support the *statistical* relevance of the

[11] The initial conditions of the two samples (per each model) are generated as random draws from their unconditional distributions. The results are unaffected if the last few observations of the first sample are employed as initial conditions for the second sample.

[12] In our exercise we compute the unique and stable solution of the linear model under rational expectations as explained in Söderlind (1999).

Table 3
Parameters' stability: rejection rates

True model	Policy rule	Phillips curve
$[\mu_\pi, \mu_y]$	[1st, 2nd]	Critical values
Panel A: empirical 5% critical values		
[0.5, 0.5]	TZ, TZ	1.821
	FF, FF	1.838
[0.8, 0.5]	TZ, TZ	1.833
	FF, FF	1.809
[0.8, 0.8]	TZ, TZ	1.778
	FF, FF	1.871
$[\mu_\pi, \mu_y]$	[1st, 2nd]	Rej. rate
Panel B: stability of the reduced-forms		
[0.5, 0.5]	TZ, FF	0.681
[0.8, 0.5]	TZ, FF	0.292
[0.8, 0.8]	TZ, FF	0.190

Panel A: empirical size of the test; panel B: test on the Lucas critique. Sample size: $T = 200$; number of Montecarlo iterations: 5000. Chow critical values adjusted for the sample-size. Details on the Montecarlo exercise are presented in the main text.

Lucas critique in the context under analysis, i.e. given the reduced-form and the data-generating process we are working with. In fact, the lowest value recorded is 0.213, well above the (approximately) 5% previously encountered.[13]

Is the result we just commented due to the fact that we employed the theoretical − i.e. not-adjusted (per sample size) − Chow critical value? The fact that the rejection rates shown in Table 2 are quite close to 5% tends to suggest a negative answer. However, we computed the sample-size adjusted Chow critical values and repeated our Montecarlo simulation with the latter.

Table 3 shows the results of our investigation. Panel A displays the empirical 5% critical values computed by imposing the null of stability of the policy regime. Indeed, we have the confirmation that the empirical critical values are all very similar to the 5% theoretical critical value we referred to when testing for the null of stability of Eq. (13). Consequently, the rejection rates computed in presence of a regime shift and presented in Table 3 (Panel B) are very close to those displayed in Table 2 (Panel B). Therefore, also these simulations confirm the statistical relevance of the Lucas critique in this open-economy context.[14]

Does this *statistical* evidence necessarily support the *economic* importance of the Lucas critique? To answer this question, we analyze the point estimates we obtained with our Montecarlo exercise. The magnitude of the estimated coefficients in a Phillips curve is economically important because it is often seen as a way of supporting (or rejecting) a theoretical framework (e.g. in certain situations, the sum of the estimated coefficients of the inflation regressors suggests whether the natural rate hypothesis is true or not), as well as a mean of evaluating the cost

[13] Notice that in this analysis, as well as in Rudebusch's (2005), we cannot state that 'the more forward-looking the structural model is, the more unstable the estimated reduced-form coefficients turn out to be'.

[14] This conclusion is in contrast with the one put forward by Rudebusch (2005) in his analysis of the reduced-form closed-economy model for the US economy. This contrast is likely to be mainly due to the different (in terms of magnitude and variables involved) policy break considered, as well as to the different structure of the economy considered.

Table 4
Phillips curve, parameters estimates

Estimated coefficients					
$[\mu_\pi, \mu_y]$	Subsample	$\sum_{j=1}^{4} \widehat{\gamma}_{\pi j}$	$\sum_{j=1}^{4} \widehat{\gamma}_{yj}$	$\sum_{j=1}^{4} \widehat{\gamma}_{qj}$	$\overline{R}^2$
[0.5, 0.5]	1st	0.715 (0.131)	0.375 (0.125)	0.065 (0.041)	0.912 (0.045)
	2nd	0.829 (0.156)	0.372 (0.113)	0.055 (0.079)	0.725 (0.079)
[0.8, 0.5]	1st	0.490 (0.205)	0.312 (0.163)	0.063 (0.057)	0.565 (0.092)
	2nd	0.719 (0.199)	0.265 (0.138)	0.032 (0.049)	0.439 (0.119)
[0.8, 0.8]	1st	0.520 (0.186)	0.499 (0.210)	0.090 (0.053)	0.496 (0.109)
	2nd	0.718 (0.191)	0.371 (0.192)	0.054 (0.046)	0.387 (0.136)

Standard deviations – computed over 5000 point estimates – in brackets.

of some policy actions (e.g. the sacrifice ratio is a function of the sum of the output gap coefficients in the Phillips curve).

Table 4 displays the average values of the sum of the estimated coefficients of the three regressors (inflation rate, the output gap, and the real exchange rate) of the Phillips curve, along with the adjusted R^2. The observable pattern in the point estimates suggests larger (sums of) point estimates for the inflation and real exchange rate regressors in the first subsample, while the 'relative' importance of the output gap across subsamples depends on which data-generating process one considers. The adjusted R^2 tends to be lower for the 2nd subsample, probably because of the lower significance of the real exchange rate under the FF regime. Nevertheless, there is not an overwhelming difference between the estimated coefficients in the two regimes under investigation. Indeed, while supporting the statistical importance of the Lucas critique, our (average) point estimates seem not to support its economic relevance.

Of course, two very different distributions may display the same mean. In fact, simple sample-averaging may cover differences in the distributions that may be of interest for undertaking economic analysis. This suggests that one may want to have a look also to the empirical distributions obtained with our Montecarlo exercise. Figs. 1–3 plot the empirical distributions of the (sums of the) estimated coefficients. In fact, the differences among the displayed distributions are far from appearing dramatic, with the distributions of the inflation coefficients just slighly more tilted rightwards in the 2nd subsample with respect to those referring to the 1st one, and those of the output gap and the real exchange rate slightly tilded leftwards. By looking at the aforementioned figures we can confirm that the economic importance of the Lucas critique is far from being remarkable.

5. Robustness checks

As in all the empirical exercises, our findings may be affected by some of the choices we made when setting up the 'true' model of the economy. Therefore, we performed some robustness checks. First, we allowed for a higher intertemporal elasticity of substitution, so that admitting a larger influence of the monetary policy moves on the aggregate demand. In particular, we set $\beta_r = 0.20$ and re-ran the experiment.[15] Our results turn out to be fairly robust to this perturbation. Another key-parameter in our new-Keynesian model is β_y, i.e. the one regulating the

[15] For higher values of β_r, given the rest of the parameterization as indicated in Table 1, we verified that the system is unstable, i.e. there are too few stable roots compared to the number of pre-determined variables of the system.

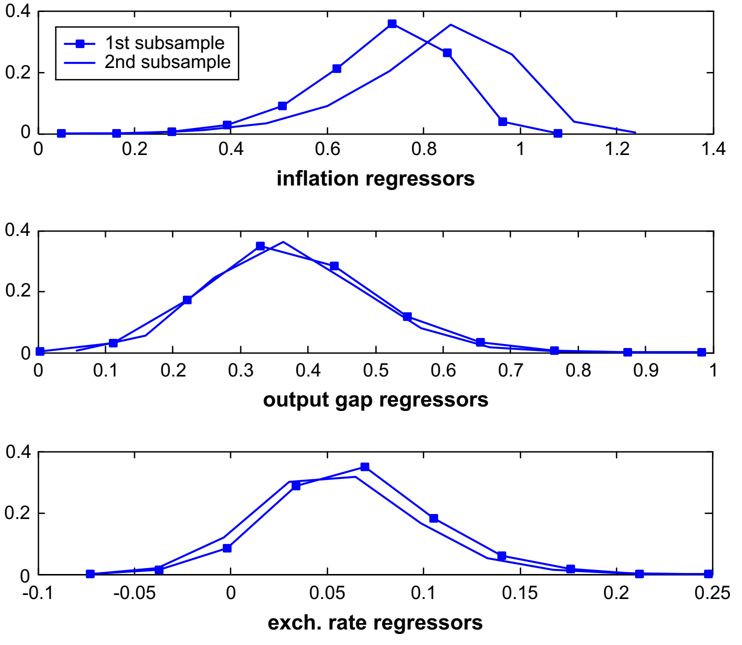

Fig. 1. Empirical distributions of the estimated coefficients, model identified by $\mu_\pi = 0.5$, $\mu_y = 0.5$. For the remaining parameterization, see Table 1.

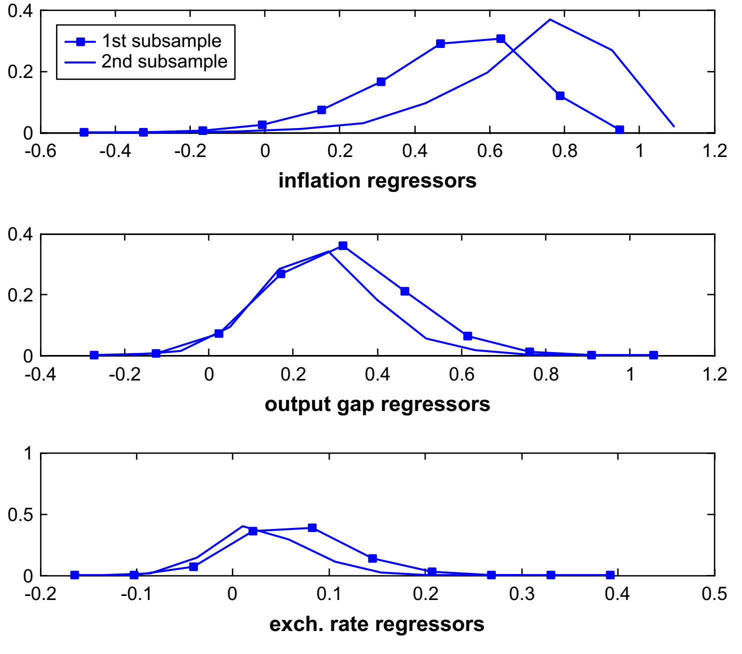

Fig. 2. Empirical distributions of the estimated coefficients, model identified by $\mu_\pi = 0.8$, $\mu_y = 0.5$. For the remaining parameterization, see Table 1.

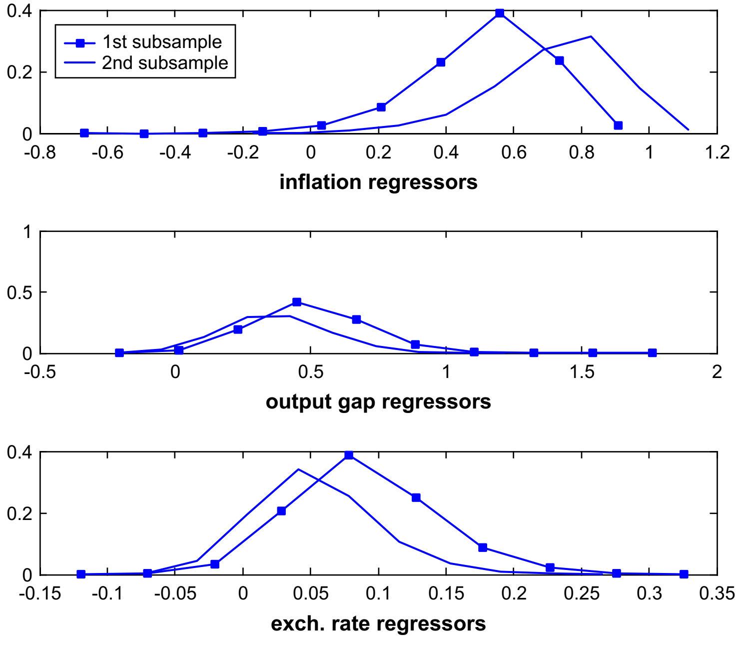

Fig. 3. Empirical distributions of the estimated coefficients, model identified by $\mu_\pi = 0.8$, $\mu_y = 0.8$. For the remaining parameterization, see Table 1.

direct impact of the business cycle on the domestic inflation rate. We doubled it – i.e. we set $\beta_y = 0.24$ – and re-run the exercise. Once more, the results presented in the previous section turn out to be robust. Another key-parameter is the degree of 'forwardness' μ_e in the UIP equation. We performed some checks also along this dimension, and allowed both for a value $\mu_e = 0.4$, lower with respect to the benchmark one, and then for a higher one, i.e. $\mu_e = 0.9$. Our results turn out to be robust also to these variations of the benchmark model.[16]

6. Conclusions

This paper aims at assessing the stability of reduced-form Phillips curves in presence of a policy break in the nominal exchange rate regime. We employ a new-Keynesian small scale open-economy dynamic stochastic model allowing for imperfect exchange rate pass-through and endogenous persistence in inflation, the output gap, and the nominal exchange rate for simulating such policy break. The model is carefully calibrated in order to plausibly represent the dynamics of an economy that experienced a dramatic policy regime change in the 1990s, i.e. Sweden. To assess the relevance of the Lucas critique in this context we estimate a reduced-form Phillips curve and assess its stability.

Our results tend to support the statistical importance of the Lucas critique. Indeed, some signs in instability in the estimated Phillips curve seem to emerge. However, when looking at the estimated coefficients of the inflation curve, the differences under the two regimes are far from being large, both in terms of average realizations and from a distributional perspective.

[16] The outcome of our robustness checks is available upon request.

Therefore, our exercise does not tend to support the economic importance of the critique, and corroborates the conclusions by Taylor (1989) and Rudebusch (2005).

Acknowledgements

We thank Michael Melvin (co-Editor) and an anonymous referee for useful comments. We also thank Nunzio Cappuccio, Carlo Favero, Giovanni Ferri, Alfred Haug, Alessandro Missale, Roberto Perotti, Luca Sessa, Saverio Simonelli, Frank Smets, Ulf Söderström, Paolo Surico, Giovanni Urga, Guglielmo Weber, and participants at EcoMod 2004 (Paris), SCE 2004 (Amsterdam), North American Winter Meeting of the Econometric Society 2005 (Philadelphia), and seminars held at Ente L. Einaudi, University of Padua, and University of Bari for discussions and suggestions. The hospitality of Ente L. Einaudi, where part of this research was conducted, is gratefully acknowledged. All remaining errors are ours.

References

Ball, L., April 1999. Efficient rules for monetary policy. International Finance, 63–83.

Ball, L., 2000. Policy rules and external shocks. NBER Working Paper No. 7910.

Campa, J.M., Goldberg, L.S., November 2005. Exchange-rate pass-through into import prices. Review of Economics and Statistics 87 (4).

Canova, F., Sala, L., 2005. Back to square one? Identification issues in DSGE models. IGIER, mimeo.

Chow, G.C., 1960. Tests of equality between sets of coefficients in two linear regressions. Econometrica 28, 591–605.

Christiano, L., Eichenbaum, M., Evans, C., February 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. Journal of Political Economy 113, 1–45.

Cúrdia, V., Finocchiaro, D., October 2005. An estimated DSGE model for Sweden with a monetary regime change. IIES Seminar Paper No. 740.

Ericsson, N.R., Irons, J.S., 1995. The Lucas critique in practice: theory without measurement. Board of Governors of the Federal Reserve System, International Finance Discussion Paper No. 506.

Estrella, A., Fuhrer, J.C., September 2002. Dynamic inconsistencies: counterfactual implications of a class of rational-expectations models. American Economic Review 92 (4), 1013–1028.

Estrella, A., Fuhrer, J., February 2003. Are 'deep' parameters stable? The Lucas critique as an empirical hypothesis. Review of Economics and Statistics 85 (1), 94–104.

Fagan, G., Henry, J., Mestre, R., January 2005. An area-wide model for the Euro area. Economic Modelling 22 (1), 39–59.

Frankel, J.A., Froot, K.A., March 1987. Using survey data to test standard propositions regarding exchange rate expectations. American Economic Review 77 (1), 133–153.

Fuhrer, J.C., September 2000. Habit formation in consumption and its implications for monetary policy. American Economic Review 90 (3), 367–390.

Lane, P., August 2001. The new open economy macroeconomics: a survey. Journal of International Economics 54 (2), 235–266.

Laubach, T., Williams, J.C., 2003. Measuring the natural rate of interest. Review of Economics and Statistics 85, 1063–1070.

Leitemo, K., Söderström, U., April 2005. Simple monetary policy rules and exchange rate uncertainty. Journal of International Money and Finance 24 (3), 481–507.

Lindé, J., 2001. Testing for the Lucas critique: a quantitative investigation. American Economic Review 91, 986–1005.

Lindé, J., Nessén, M., Söderström, U., June 2004. Monetary policy in an estimated open-economy model with imperfect pass-through. IGIER Working Paper No. 263.

Lucas Jr., R.E., 1976. Econometric policy evaluation: a critique. Carnegie-Rochester Conference Series on Public Policy 1, 19–46.

Mankiw, G.N., May 2001. The inexorable and mysterious tradeoff between inflation and unemployment. The Economic Journal 111, C45–C61.

Onatski, A., Stock, J.H., 2002. Robust monetary policy under model uncertainty in a small model for the U.S. economy. Macroeconomic Dynamics 6, 85–110.

Rudebusch, G.D., April 2005. Assessing the Lucas critique in monetary policy models. Journal of Money, Credit and Banking 37 (2), 245–272.

Rudebusch, G.D., Svensson, L.E.O., 1999. Policy rules for inflation targeting. In: Taylor, J.B. (Ed.), Monetary Policy Rules, NBER Conference Report Series. University of Chicago Press, Chicago and London.

Rudebusch, G.D., Svensson, L.E.O., March 2002. Eurosystem monetary targeting: lessons from U.S. data. European Economic Review 46, 417–442.

Sgherri, S., 2005. Explicit and implicit targets in open economies. IMF Working Paper No. 05-176.

Söderlind, P., 1999. Solution and estimation of RE macromodels with optimal policy. European Economic Review 43, 813–823.

Svensson, L.E.O., February 2000. Open-economy inflation targeting. Journal of International Economics 50 (1), 155–184.

Taylor, J.B., 1989. Monetary policy and the stability of macroeconomic relationships. Journal of Applied Econometrics 4, S161–S178.

Woodford, M., 2003. Interest and Prices: Foundations of a Theory of Monetary Policy. Princeton University Press.