Regime Shifts and the Stability of Backward Looking Phillips Curves in Open Economies

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Abstract

In this paper we assess the stability of open economy backward-looking Phillips curves estimated over two different exchange rate regimes. The pseudo-data employed in our econometric exercise come from the simulation of a New-Keynesian hybrid model suited for performing monetary policy analysis. Two main results arise: i) in most of the simulated scenarios the estimated reduced-form Phillips curves turn out to be unstable. However, if the structural new-keynesian model is predominantly - even if not fully - backward-looking, the estimated reduced-form parameters are stable; ii) the Chow-breakpoint test tends to underestimate the importance of regime-shifts in small samples.

Keywords: Lucas Critique, forwardness, backward looking Phillips curves, exchange rates, Chow test.

JEL Classification: E17, E52, F41

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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 microfoundation 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]) has relied on ad-hoc backward-looking frameworks. In fact, backward-looking models tend to offer a quite 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, backward-looking models are affected by the Lucas [1976] critique. The theoretical 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 theoretically 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 this change is from a quantitative perspective?

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 [2003]), to the best of our knowledge the only contribution dealing with an open-economy framework dates back to Taylor [1989]. This is somewhat surprising, given the incresing openess 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 [1989]'s contribution. Taylor [1989] employs an estimated macro-model for simulating the shift from a 'fixed' to a 'flexible' nominal exchange rate in some industrial-

[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 in this sense, see Ericsson and Irons [1995].

ized countries. Once done so, he estimates with such simulated data some reduced-form schedules (mainly demand and supply curves), 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 differences among the estimated parameters.

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 manage the short-term interest rate to minimize a penalty function representing the loss the Society bears because of the fluctuations of the main economic macro-variables. This minimization problem is complicated by the complex transmission mechanism present in the economy. We employ this 'targeting-rule' framework a la Svensson [1999] because it represents the workhorse approach in modern monetary policy analysis. As a second difference with respect to Taylor [1989], to evalutate the impact of the regime shift on the reduced form coefficients we rely on a statistical tool, i.e. the popular Chow [1960] breakpoint test.

The reduced-form schedules we concentrate on are two different versions of the 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]) over the micro-founded, expectations-equipped new-Keynesian schedule (e.g. Woodford [2003]). Since the former is probably a reduced-form schedule capturing dynamics stemming from a different structural model, it is interesting to gauge the stability of such curve once a policy shift is implemented.

Our results highlight the relationship existing between the instability of our estimated reduced-form Phillips curves and the relative importance of agents' forward-looking expectations in the structural model at hand. Furthermore, we provide evidence against the power of the Chow-breakpoint test in small samples, so offering a possible explanation for the commonly rejected empirical relevance of the Lucas critique under a regime shift.

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:

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

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

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).

Equation (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 (as in 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}.[2] Equation (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 in e.g. habit formation, as in Fuhrer [2000]), the ex-ante real interest rate, the real exchange rate (which approximates

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

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.[3] Notice that equations (1) and (2) allows for explicit lags in the transmission mechanism; in fact, it is hard to derive these lags from micro-foundations, but they are quite useful to match the gradual response of inflation and output to monetary policy shocks observed in the data (Christiano et al [2005]).

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.[4] 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

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

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:

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

[3] 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 regulator problem non-standard.

[4] 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].

Importantly, the parameter θ allows for the possibility of deviating from the law of one price *in the short-run.* In fact, if $0 \leq \theta < 1$, 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 tend to be much less important in the long run, as shown in Campa and Goldberg [2002].

Since the real exchange rate q_t is defined as

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

equations (4), (5), and (6) suggest the following link between real exchange rate and CPI inflation:

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

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.

$$i_t^* = (1-\rho_{i*})(f_\pi^* \pi_t^* + f_y^* y_t^*) + \rho_{i*} i_{t-1}^* + \zeta_t^* \tag{8}$$

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 i.i.d. processes whose variances are respectively σ_u^* and σ_v^*.

2.1 Optimal monetary policy

The monetary authorities' behavior closes the model. We aim at simulating the shift from a 'controlled exchange rate regime' to a 'flexible' one, a change implemented by several countries in the last decades (Reinhart and Rogoff [2004]). In our framework, the different monetary policy regimes are identified by different penalty functions. In particular, under the 'controlled exchange rate regime' the penalty function reads as follows:[5]

$$E(L_t) = \lambda_{\Delta s} Var(\Delta s_t) + \lambda_{\Delta i} Var(\Delta i_t) \tag{11}$$

where $Var(\Delta s_t)$ stands for the volatiliy of the nominal exchange rate in first differences, while the interest rate smoothing argument $Var(\Delta i_t)$ is mainly introduced to avoid counterfactual extreme fluctuations of the short-term nominal interest rate.[6] By contrast, the 'flexible' exchange rate regime is formalized by allowing for a 'CPI inflation targeting', i.e.

$$E(L_t) = \lambda_{\pi^{CPI}} Var(\pi_t^{CPI}) + \lambda_y Var(y_t) + \lambda_{\Delta i} Var(\Delta i_t) \tag{12}$$

where $Var(\pi_t^{CPI})$ captures the concern the monetary authorities have for CPI inflation fluctuations, while the weight λ_y measures the relative importance of business-cycle fluctuations in the penalty function.

The monetary authorities aim at minimizing either (11) or (12) subject to the constraints (1)-(10). We assume that the central bank conducts monetary policy under discretion, an assumption supported by the analysis by Bernanke and Mishkin [1997]. Söderlind [1999] shows that the solution of the optimal control problem links the vector of state variables $x_{1t} = \left[\pi_t, y_t, \varphi_t, q_{t-1}, \pi_{t-1}^M, i_t^*, \pi_t^*, y_t^*, i_{t-1}\right]'$ to the policy rate i_t, i.e.

[5] In fact, the CB solves an *intertemporal* problem featured by the following loss function: $E_t \sum_{\tau=0}^{\infty} \delta^\tau \left(\sum_{i=1}^{n} x_{i,t+\tau}^2 \right)$, where x_i is one of the n arguments targeted by the monetary authorities. As shown by Rudebusch and Svensson [1999], the conditional expectation presented here tends to the unconditional expectation discussed in the text as $\delta \rightharpoonup 1$. In this study, we fix the discount factor δ to be equal to .99, a standard choice given the quarterly frequency assumed for our model.

[6] Discussions on some possible alternative ways of formalizing a 'controlled' exchange rate regime and on some technicalities on the optimal-control problem solved by the central bank are offered in the Technical Appendix of this paper available upon request.

$i_t = \widetilde{f}^i(\lambda_s^i)x_{1t}$. The $(1x9)$ vector $\widetilde{f}^i(\lambda_s^i)$ is the numerically-computed solution of the optimal control problem solved by the monetary policy authorities having the vector of preferences λ_s^i, with i indicating either 'regime1' [identified by the penalty function (11)] or 'regime2' [identified by (12)].

2.2 Model parameterization

Our 'in-lab' exercise allows us to perfectly identify the source of the instability of the reduced-form Phillips curves (if any). Evidently, to perform an interesting exercise from a policy perspective we need a sensible parameterization of the structural model we employ. The benchmark parameterization used in our exercise is largely borrowed from the existing literature. The domestic economy is almost fully parameterized on the basis of the contributions by Svensson [2000] and Leitemo and Söderström [2005].[7] In particular, we think of $\mu_\pi = .5$ and $\mu_\pi = .3$ as plausible benchmark values for the degrees of 'forwardness' related to - respectively - the Phillips curve and the IS schedule. However, given the huge uncertainty surrounding such values, we also investigate the dynamics stemming from other two different models identified by $[\mu_\pi, \mu_y] = [.3, .1]$ and $[\mu_\pi, \mu_y] = [.8, .8]$. As far as the degree of forwardness of the UIP condition is concerned, we set $\mu_s = .7$ to acknowledge to the nominal exchange rate its feature of 'forward-looking determined asset price' (Svensson [2000]). We set the exchange rate pass-through coefficient $\theta = .5$ to be in line with the indications coming from the empirical work by Campa and Goldberg [2002]. We parameterize the foreign economy as in Svensson [2000], the sole exception being the interest rate smoothing parameter ρ_{i*} that we set to .75, in line with the point-estimate by Clarida et al [2000] for the US. All the parameters identifying our benchmark model are collected in Table 1.

To close the model, the two different regimes we simulate are identified by the following central bank's sets of preferences: $\lambda_{\Delta s} = 1, \lambda_{\Delta i} = .2$ for the 'controlled nominal exchange rate volatility' regime (Loss function [11]) vs. $\lambda_{\pi^{CPI}} = 1, \lambda_y = .5, \lambda_{\Delta i} = .2$ for the 'CPI Quasi-Strict Inflation Targeting'

[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.

regime (Loss function [12]).[8]

Table 2 collects the numerically computed coefficients of the optimal 'targeting' rules $i_t = \widetilde{f}^{regime1}x_{1t}$ and $i_t = \widetilde{f}^{regime2}x_{1t}$. As expected, huge differences between the two optimal rules arise. In particular, the central bank attributes a large importance to the elements entering the UIP condition under the first regime; by contrast, domestic elements play a key-role when a 'CPI inflation targeting' is implemented. Notably, the regime-shift also leads to a higher optimal interest rate smoothing, given the higher concern toward inflation stabilization under discretion (Woodford [2003]).

[Tables 1-2 here]

3 Assessing the importance of the Lucas critique

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

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

$$\pi_t = \sum_{j=1}^{4}(\gamma_{\pi j}\pi_{t-j} + \gamma_{yj}y_{t-j} + \gamma_{qj}q_{t-j} + \gamma_{ij}i_{t-j} + \gamma_{i^*j}i^*_{t-j}) + \widetilde{\xi_t^{\pi}} \tag{14}$$

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.[9] This is nothing but an open-economy version

[8] The models is thought for 'replicating' quarterly dynamics. All the variables are in log-deviations with respect to their steady states, which are normalized to zero. The timing of the model goes as follows: at the beginning of the t^{th}-period, shocks strike the economy; then, private agents form their expectations; finally, CB sets the policy rate. The model's impulse response functions - not presented in the text, but available upon request - confirm the dynamic sensibility of the model.

[9] 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.

of the one proposed by Rudebusch and Svensson [1999,2002] for the US. Notably, with adequate restrictions on the coefficients γ_s, this reduced-form equation collapses to the one in Ball [1999,2000]. We label this curve as 'Phillips [1]'.

Eq. (14) enriches the former by adding both the domestic and the foreign policy rate. This is done in order to approximate the attempt an econometrician might perform for capturing important and possibly omitted dynamics for the inflation rate path, or to acknowledge an explicit role to the policy rates in the inflation formation as in the 'cost-channel' literature (e.g. Ravenna and Walsh [2004]). We label the latter curve as 'Phillips [2]'.

Steps for assessing the statistical relevance of the Critique

We now turn to the description of the algorithm we employ for assessing the statistical importance of the Lucas critique in such an open-economy context. In particular, we implement the following steps:

1. We simulate the structural model described in Section 2 for $I + T$ periods *under the null of absence of regime shifts.* For each period of the simulated sample we draw a realization of the vector $[u_t, v_t, \psi_t, u_t^*, v_t^*, \zeta_t^*]' \sim N(0_{6x1}, diag[\sigma_u^2, \sigma_v^2, \sigma_\psi^2, \sigma_u^{*2}, \sigma_v^{*2}, \sigma_\zeta^{*2}])$.[10] The first $I = 100$ periods are simulated to get a stochastic vector of initial values for the model, and are discarded before implementing Step 2.

2. With this sample of pseudo-data (sample whose size is equal to T), we OLS estimate the 'reduced form' coefficients of the backward-looking Phillips curves (13) and (14). Then, we split the samples in two subsamples of equal size $T_1 = T_2 = \frac{T}{2}$, and compute the F-statistic of the Chow [1960]-breakpoint test.[11] We perform our exercise with samples features by different sizes: a 'small' one ($T = 200$) and a 'large' one ($T = 1,000$).

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[10] In our notation, $diag[\sigma_1, ..., \sigma_n]$ stands for the diagonal matrix $\begin{bmatrix} \sigma_1^2 & \cdots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \cdots & \sigma_n^2 \end{bmatrix}$.

[11] To compute the F-statistic, we adopt the following formula (k stands for the number of estimated coefficients): $\frac{(\widehat{\sigma}_T^2 - \widehat{\sigma}_{T1}^2 - \widehat{\sigma}_{T2}^2)/k}{(\widehat{\sigma}_{T1}^2 + \widehat{\sigma}_{T2}^2)/(T-2k)} \sim F(k, T-2k)$ under the null of stability.

3. We repeat Steps 1-2 $N = 3,000$ times; each time, we store the F-statistic computed according to Step 2.

4. The $3,000$ values of the F-statistic are employed for calculating the F-critical value for the Chow test, so obtaining the corrected-per-sample size critical value of the test. In our exercise, we concentrate on the statistical significance at the 5% level;

5. We implement Steps 1-3 *allowing for the monetary policy regime shift* at $t = \frac{T}{2}$.[12]

6. We compare the 3,000 F-statistical values obtained in Step 5 with the F-critical value computed in Step 3. In particular, if the statistical value is larger/smaller than the critical one, the null of stability is rejected/non rejected.[13] According to this criterium, we compute a *rejection rate* per each different parameterization and sample.

The next Section presents and comments on our results.

4 Findings

To assess the stability of 'Phillips [1]' and 'Phillips [2]', we first concentrate on the rejection rates we obtained with our simulations. Two main interesting considerations might be done. First, the fact that in the 'true' model rational expectations play a direct and indirect role in shaping the inflation path does not necessarily call for the instability of our reduced-form Phillips curves under the regime-shift we simulate. In fact, Table 3 suggests that as long as the impact of such 'forwardness' is positive but minor - i.e.

[12] Notice that in moving from the first regime to the second one we are assuming that agents are not concerned with any learning issue; this is a limitation of our approach, and probably renders our 'in-lab' exercise less close to reality than a study performed with actual data. On the other hand, this approach amplify the power of the Chow test, so rendering its suggestions (above all those coming from large samples) more reliable.

[13] In fact, the 'superexogeneity test' by Engle and Hendry [1993] would suggest to compute the rejection rates conditional on the rejection of the null of stability for the monetary policy rule. However, in this simulated exercise the break of the policy rule is a certain and known event. This is the reason why we perform an unconditional calculation of the rejection rates.

when $[\mu_\pi, \mu_y] = [.3, .1]$ - the stability of 'Phillips [1]' and 'Phillips [2]' cannot be rejected with a large statistical confidence. But reduced-forms are not stable in general; in fact, when the economy is quite forward-looking, their stability is not supported by the data. Second, the Chow test is potentially misleading in 'small' samples. This statement is supported by the rejection rates computed for the intermediate scenario, i.e. the one identified by $[\mu_\pi, \mu_y] = [.5, .3]$. In particular, while the stability of the reduced-form schedules is supported according to the rejection-rates when $T = 200$, the opposite is suggested when the larger sample is considered. This is an interesting finding, because if offers a rationale for the huge amount of empirical contributions which did not offer any statistical support to the Lucas critique (see the long list of papers surveyed by Ericsson and Irons [1995], and the recent contribution by Rudebusch [2003]). Notably, this is the very same conclusion reached by Lindé [2001], who performed an exercise similar in spirit but limited to a closed-economy set up. Given that the Chow [1960] test is still widely applied in econometric applications, our evidence seems to provide a serious warning against its use, even when the breakpoint is known.

Our results differ with respect to those in Taylor [1989]. This might be due to the different structure of the two 'true' models, as well as the different parameterization, the magnitude of the regime-shift, and so on. We stress here that the plus of our contribution is that of going over the simple observation of two different sets of estimated parameters. Indeed, we rely on a statistical test, and we clearly put in evidence its limits in small samples and its ability to detect instability in large samples.

[Table 3 here]

5 Robustness checks

Of course, our qualitative findings may be affected by some of the choices we made when setting up the 'true' model of the economy. In particular, the literature is scant on the value that the degree of 'forwardness' μ_e in the UIP equation should take. Therefore, we performed some checks along this dimension to verify the robustness of our results. In particular, we allowed for the value $\mu_e = .4$, lower with respect to the benchmark one, and for a higher one, i.e. $\mu_e = .9$. Figures 1 and 2 display the sensitivity of the rejection rates to such variations. A few considerations may be put forward. First, the function linking the rejection rates to the parameter μ_e seems to be quite complex and non-linear. In particular, conditioning to a given pair of values $[\mu_\pi, \mu_y]$, we can either find a monotonically increasing function, or a monotonically decreasing, or a peak in correspondence to the benchmark value for μ_e. Then, from a qualitative perspective, the above mentioned function deserves further investigation. Second, from a quantitative perspective the parameter μ_e does not really affect any of the conclusions as far as the pairs $[\mu_\pi, \mu_y] = [.3, .1]$ and $[\mu_\pi, \mu_y] = [.8, .8]$ are concerned. By contrast, such a parameter plays an important role for the stability of the reduced-form schedules in the intermediate case, i.e. $[\mu_\pi, \mu_y] = [.5, .3]$. Somewhat counter-intuitively, the higher the importance of the forward-looking term in the UIP condition, the lower the instability of the curves 'Phillips [1]' and 'Phillips [2]'. This is an interesting result which calls for a even more intense effort towards a better understanding of the relative importance of explicitely formalized endogenous persistence vs. forward-looking components in both the structural version of the Phillips curve and in the IS schedule, along the lines recently developed by Estrella and Fuhrer [2003].

[Figures 1-2 here]

A relevant issue here is that regarding the precision of the estimated parameters in (13) and (14). In fact, the Chow test relies upon the estimated residuals of such schedules, then the more consistently and efficiently the parameters of such schedule are estimated the more reliable the test is. Given that we cannot compute the exact reduced-form of the structural Phillips curve (1), we are not aware of the exact figures the estimated reduced-form coefficients should take. Nevertheless, we can at least judge the sign and 'significance' coming out of our econometric exercise.

Tables 4 and 5 collect the estimated parameters and their standard deviations.[14] As expected, the most important regressors - namely, lagged inflation rates, lagged output gap observations, and lagged real exchange rate realizations - turn out to be significant and to have the expected positive sign. Not surprisingly, the higher the importance of the forward-looking component in the structural Phillips curve, the lower the estimated-mean values of the autoregressive coefficients in the reduced-form schedules. Overall, our point-estimates seem to support the reliability of the test we employed.

[**Tables 4-5 here**]

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 modern 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. Then, we estimate two reduced-from Phillips curves and assess their stability. In most cases, their stability is rejected by a standard Chow-test, above all when the test is run over a large sample. However, if forward-looking expectations play a limited role in the structural model of the economy, the stability of the reduced-from schedules is hardly rejected. This finding seems to re-qualify the discussion on the importance of the Lucas [1976] critique and its consequences for the use of backward-looking schedules in monetary policy analysis. It is worth stressing that, according to our econometric exercise, the Chow-test is not able to detect coefficients' instability in small samples. This finding offers a rationale for the long list of contributions rejecting the importance of the Lucas critique.

[14] As already described when we presented our algorithm, for each reduced-form coefficient j we estimated $3,000$ values - with $3,000$ different samples extracted from the same population - under the policy regime shift. Tables 4 and 5 present the mean-value and the standard deviation of each estimated parameter computed over the $3,000$ point estimates available.

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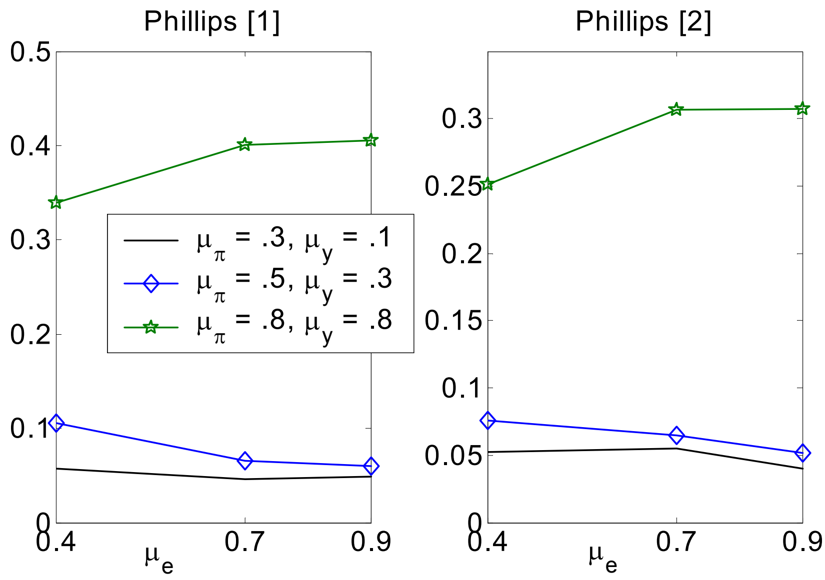

Figure 1: REJECTION RATES: SAMPLE SIZE T = 200.

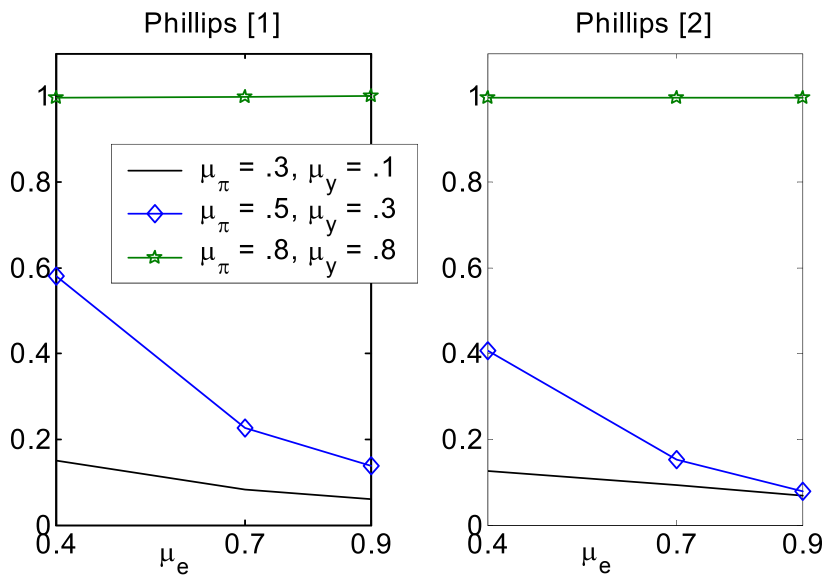

Figure 2: REJECTION RATES: SAMPLE SIZE T = 1,000.

Domestic economy							
Phillips curve		IS curve		UIP condition		CPI equation	
μ_π	.5	μ_y	.3	μ_s	.7	χ	.35
α_y	.2	β_r	.15	ρ_φ	.3	θ	.5
α_q	.04	β_q	.05	σ^2_ψ	.844		
σ^2_u	1.556	β_y	.12				
		σ^2_v	.656				
Foreign economy						*Central Bank*	
Phillips curve		IS curve		Taylor rule		Key-parameters	
$\rho_{\pi*}$	.8	ρ_{y*}	.8	$f_{\pi*}$	1.5	δ	.99
σ^2_{u*}	.5	σ^2_{v*}	.5	f_{y*}	.5	$\lambda_{\Delta i}$	.2
				ρ_{i*}	.75		
				$\sigma^2_{\xi*}$	.5		

Table 1: BENCHMARK PARAMETRIZATION. Sources of the parameters indicated in the text.

Optimal policy rule	π_t	y_t	φ_t	q_{t-1}	π^M_{t-1}	i^*_t	π^*_t	y^*_t	i_{t-1}
Regime 1	.27	.22	.85	-.06	.00	1.00	-.10	-.01	.06
Regime 2	.41	.65	.23	.01	.04	.53	.22	.11	.34

Table 2: OPTIMAL REACTION FUNCTIONS UNDER ALTERNATIVE REGIMES. Model parameter as in the benchmark case, see Table 1.

True model	*Phillips [1]*		*Phillips [2]*	
$[\mu_\pi, \mu_y]$	$\frac{Rej.rate}{T=200}$	$\frac{Rej.rate}{T=1{,}000}$	$\frac{Rej.rate}{T=200}$	$\frac{Rej.rate}{T=1{,}000}$
$[.3, .1]$	.0460	.0933	.0553	.0823
$[.5, .3]$	.0660	.2270	.0650	.1533
$[.8, .8]$	.4010	.9993	.3067	.9997

Table 3: PARAMETERS STABILITY: REJECTION RATES.

Sample size T = 200				
$[\mu_\pi, \mu_y]$	$\sum_{j=1}^{4}\widehat{\gamma_{\pi j}}$	$\sum_{j=1}^{4}\widehat{\gamma_{yj}}$	$\sum_{j=1}^{4}\widehat{\gamma_{qj}}$	$\overline{R}^2$
$[.3, .1]$	.9082 (.1224)	.2714 (.2636)	.0404 (.0627)	.6987 (.0651)
$[.5, .3]$	.7450 (.1049)	.4449 (.2695)	.0724 (.0670)	.4851 (.0722)
$[.8, .8]$	.1740 (.1813)	.3602 (.1850)	.1562 (.0650)	.1559 (.0542)
Sample size T = 1,000				
$[\mu_\pi, \mu_y]$	$\sum_{j=1}^{4}\widehat{\gamma_{\pi j}}$	$\sum_{j=1}^{4}\widehat{\gamma_{yj}}$	$\sum_{j=1}^{4}\widehat{\gamma_{qj}}$	$\overline{R}^2$
$[.3, .1]$	.9170 (.0503)	.2678 (.1076)	.0400 (.0258)	.7176 (.0263)
$[.5, .3]$	.7607 (.0440)	.4386 (.1128)	.0721 (.0280)	.4996 (.0305)
$[.8, .8]$	.2031 (.0749)	.3503 (.0767)	.1526 (.0281)	.1596 (.0224)

Table 4: PHILLIPS CURVE [1], PARAMETERS ESTIMATES. Note: Standard deviations - computed over 3,000 point-estimates and adjusted R2 - in brackets.

Sample size T = 200						
$[\mu_\pi, \mu_y]$	$\sum_{j=1}^{4}\widehat{\gamma_{\pi j}}$	$\sum_{j=1}^{4}\widehat{\gamma_{yj}}$	$\sum_{j=1}^{4}\widehat{\gamma_{qj}}$	$\sum_{j=1}^{4}\widehat{\gamma_{ij}}$	$\sum_{j=1}^{4}\widehat{\gamma_{i^*j}}$	$\overline{R}^2$
$[.3, .1]$	.8749 (.1783)	.2284 (.3242)	.0261 (.0892)	.0023 (.2520)	.0641 (.3030)	.6994 (.0653)
$[.5, .3]$	.7231 (.1782)	.4168 (.3308)	.0440 (.0989)	−.0666 (.2806)	.1823 (.3243)	.4884 (.0736)
$[.8, .8]$	.1113 (.2154)	.3959 (.2393)	.1489 (.0916)	.1188 (.2788)	−.0499 (.3141)	.1596 (.0577)
Sample size T = 1,000						
$[\mu_\pi, \mu_y]$	$\sum_{j=1}^{4}\widehat{\gamma_{\pi j}}$	$\sum_{j=1}^{4}\widehat{\gamma_{yj}}$	$\sum_{j=1}^{4}\widehat{\gamma_{qj}}$	$\sum_{j=1}^{4}\widehat{\gamma_{ij}}$	$\sum_{j=1}^{4}\widehat{\gamma_{i^*j}}$	$\overline{R}^2$
$[.3, .1]$	.8939 (.0694)	.2280 (.1282)	.0261 (.0357)	−.0083 (.0993)	.0677 (.1218)	.7181 (.0264)
$[.5, .3]$	.7426 (.0700)	.4028 (.1321)	.0446 (.0387)	−.0546 (.1110)	.1617 (.1277)	.5029 (.0308)
$[.8, .8]$	.1588 (.0873)	.3797 (.0954)	.1448 (.0381)	.1044 (.1107)	−.0452 (.1264)	.1626 (.0227)

Table 5: PHILLIPS CURVE [2], PARAMETERS ESTIMATES. Note: Standard deviations - computed over 3,000 point-estimates and adjusted R2 - in brackets.