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Monetary Policy in a Monetary Union: What Role for Regional Information?*

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Can the central bank of a monetary union, whose objectives are exclusively defined in terms of union-wide variables, improve its performance by reacting to regional variables rather than to union-wide variables only? Our answer is not clear-cut. We find the improvement to be large when we use a backward-looking model of the economy and negligible when we use a hybrid model. The main determinant of this finding seems to be the different degree of inertia (or its opposite, the forward-lookingness) characterizing the two models, rather than other model features.

JEL Code: E52.

1. Introduction

The launch of the euro in January 1999 has rekindled an interest in monetary unions. Several papers have focused on the issue of regional heterogeneity and its implications for monetary policy. For the euro area, relevant differences have often been detected in the reactivity of key national macro variables to monetary policy (Angeloni, Kashyap, and Mojon 2003). Even in long-established

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monetary unions such as the United States, significant regional heterogeneity in the transmission of monetary policy has been documented (see, e.g., Carlino and DeFina 1998, 1999; Meade and Sheets 2002; and Owyang and Wall 2003).

In spite of this, little attention has thus far been devoted to the implications of regional heterogeneity for the conduct of monetary policy. The recent literature on monetary policy rules has focused on a large number of factors potentially affecting the design of policy, including the uncertainty on the structure or the parameters of the model, the presence of measurement errors in variables, the impact of data revisions, and the features of the reaction function (should simple rules be preferred to fully optimal ones? should the central bank react to current or expected inflation? should it react to the exchange rate?). A necessarily incomplete list of recent contributions dealing with these themes includes Peersman and Smets (1999), Taylor (1999a, 1999b), Orphanides (2001), Rudebusch (2001, 2002), Levin, Wieland, and Williams (2003), and Lubik and Schorfheide (2007). However, with a few notable exceptions, which we discuss below, the potential role of regional disaggregation is neglected.

Thus, the typical policy rule includes only union-wide variables among its arguments. Yet the presence of heterogeneity in a monetary union implies that indicators pertaining to the constituent regions should appear among the state variables of the model representing the economy of the union, and hence in the policy rule, by standard optimal control arguments, even when the central bank is exclusively interested in union-wide developments. In other words, the fact that the *objectives* of policy (the loss function) are defined exclusively in terms of union-wide variables does not imply that the *decision-making process* (the policy rule) should also rely exclusively on union-wide variables. While obvious, this distinction is often overlooked.[1]

[1] The European Central Bank (ECB) has repeatedly argued in the past that to pursue area-wide objectives—as unambiguously required by the Maastricht Treaty and the ECB Statute—only area-wide economic developments mattered (see, e.g., President Duisenberg's statement during the press conference of September 9, 1999: "Our decisions today, again and as always, were based on a euro area-wide analysis of economic and financial developments—and nothing else"). More recently, however, ECB officials have increasingly stressed that national

There is, thus, a surprising gap between the widespread recognition of significant heterogeneities within important monetary unions and their virtual neglect in the design of monetary policy. The present paper addresses this issue. Specifically, we ask whether and to what extent the monetary authority of a currency union, exclusively interested in union-wide variables, should react to regional economic developments within the union.

We address this issue, adopting the linear quadratic framework commonly used in the literature on policy rules, and provide some answers for the euro area. We first estimate a simple multicountry model of the three main national economies of the area—France, Germany, and Italy. The model, in line with much of the existing literature, suggests that statistically significant differences among these economies do exist. Next, to assess whether these differences are of practical relevance to the policymaker, we compute optimal Taylor-type simple rules that minimize the policymaker's loss function under the constraint provided by the model. We focus on two classes of simple rules: one allows the policy rate to react to region-specific variables; in the other, the arguments are union-wide variables only. We compare the minimized loss under the two alternative rules, interpreting the difference as the cost of neglecting the heterogeneities inside the union. As the loss function depends on the unconditional variances of union-wide inflation and output gap, this difference can be viewed as a clear, policy-oriented metric to decide whether regional heterogeneities deserve practical consideration. This metric can usefully complement a statistical approach at assessing the relevance of heterogeneities, such as that proposed in the literature on aggregation (see, e.g., Pesaran, Pierse, and Kumar 1989 and Marcellino, Stock, and Watson 2003).

We focus on the euro area and on its component countries mainly because national differences in the economic structures (as well as in language, culture, institutional features, and legal systems) are at present pronounced, and idiosyncratic shocks are still likely to have a relatively important role. One may therefore conjecture that the

information is instrumental in gaining a better assessment of euro-area developments (see, e.g., the remarks by Otmar Issing, former member of the Executive Board, at the ECB workshop "Monetary Policy Implications of Heterogeneity in a Currency Area," Frankfurt, December 13–14, 2004).

potential loss associated with the neglect of country-specific information might be large. However, our approach has a wider methodological import. The above questions seem relevant, at least in principle, for the United States and for other monetary unions, such as Canada and China, where regional differences appear to be major.

Our main result is twofold. On the one hand, the loss reduction that can be attained when policy is allowed to react to regional variables can be large, and therefore the issue is potentially relevant. On the other hand, the size of the loss reduction is heavily dependent on the model of the area adopted. Specifically, using an estimated backward-looking model, we find that the loss reduction typically exceeds 30 percent, reflecting a similarly sized decline in inflation and output-gap variability. By contrast, when we replicate the analysis using an estimated hybrid model, featuring backward- as well as forward-looking components, we find that the gain associated with the use of regional information is almost negligible. A sensitivity analysis suggests that most of the difference in the results can be traced back to the degree of forward-lookingness of the model. Unfortunately, this is precisely the feature that seems most difficult to pin down empirically with a reasonable degree of confidence (see the papers collected in the September 2005 issue of the *Journal of Monetary Economics*, which we briefly discuss below).

A few recent papers deal with the role of regional variables in a monetary union. Using a two-region dynamic stochastic general equilibrium (DSGE) model, Benigno (2004) shows that the central bank should target an average price indicator in which the weight assigned to each region is proportional to its degree of price stickiness. Lombardo (2006) extends this analysis to the role of unequal degrees of competition. Our paper can be viewed as complementary to these studies, as it explores a similar issue using a completely different approach. The methodology in De Grauwe (2000), De Grauwe and Piskorski (2001), Aksoy, De Grauwe, and Dewachter (2002), and, to some extent, Wyplosz (1999) is similar to ours. However, these works associate the use of information on the individual countries with the (undue) nationalistic attitude of the Governing Council members, which implies loss functions defined over national variables. Their primary focus is therefore on the decision procedures within the ECB Governing Council rather than on the optimal use of regional information. Monteforte and Siviero (2008) define both the central bank's objective and policy rule over area-wide aggregates, and compute two alternative optimal rules under the constraint provided by either an area-wide or a multicountry model. Their main focus is on the consequences of following aggregate and disaggregate modeling strategies. Jondeau and Sahuc (2008) analyze the same issue but also address our main question using a DSGE model. Brissimis and Skotida (2007) compare the results achieved by the monetary policymaker using a disaggregate model and policy rule versus those attainable with an aggregate model and rule.

The organization of the paper is as follows. Section 2 describes the basic setup of our exercises. Section 3 illustrates the simple euro-area multicountry models used in the analysis—a fully backward-looking and a hybrid backward-forward-looking version. Section 4 reports the main results. Section 5 concludes.

2. Design of Experiments

In the standard optimal control of a linear system under a quadratic loss function, the solution of a matrix Riccati equation yields an optimal policy rule whose arguments are all the state variables of the system. Most of the recent literature on optimal monetary policy focuses instead on simple optimal policy rules, obtained by imposing some constraint on the functional form of the optimal rule, e.g., reducing the number of (lagged) variables appearing among its arguments. The underlying idea is to weigh the underperformance of the simple rules against their simplicity, which can make them easier to use for the monetary authorities and a more useful tool for communication with the public. Furthermore, simple rules may be more robust, as compared with more model-dependent optimal rules, thus implying a trade-off between performance in the context of a specific model and robustness.

The analytical framework adopted in the present paper is borrowed from that literature. However, we are not interested in the functional form of the policy rules, nor in their robustness as such. Rather, we focus on comparing the performance of rules that include national variables among their arguments *vis-à-vis* rules that only react to area-wide variables. We restrict our attention to Taylor-type rules, in which only contemporaneous inflation and output gap appear among the arguments, along with a lagged interest rate term.[2] The first rule we consider, which we dub the multicountry information-based rule (MCIBR), has the following form:

$$i_t = \alpha_M^G \pi_t^G + \alpha_M^F \pi_t^F + \alpha_M^I \pi_t^I + \beta_M^G y_t^G + \beta_M^F y_t^F + \beta_M^I y_t^I + \rho_M i_{t-1}, \quad (1)$$

where y denotes the output gap, π denotes annualized quarter-on-quarter inflation, i denotes a policy interest rate, and α_M^j, β_M^j, and ρ_M are coefficients to be determined. Superscripts $j = G, F, I$ stand, respectively, for Germany, France, and Italy, the three economies in the simple multicountry models of the euro area presented in section 3. The second rule, dubbed the area-wide information-based rule (AWIBR), is of the type

$$i_t = \alpha_A \pi_t + \beta_A y_t + \rho_A i_{t-1}, \quad (2)$$

where π_t and y_t are area-wide variables and α_A, β_A, and ρ_A are coefficients to be determined. Following the methodology adopted by Eurostat, we compute area-wide variables as $\pi_t = \sum_j w_\pi^j \pi_t^j$ and $y_t = \sum_j w_y^j y_t^j$, $j = G, F, I$, where w_x^j, $x = \pi, y$ represent the appropriate country weights.[3] In the case of the AWIBR, the policymaker is assumed to react to area-wide inflation and output gap only; in other words, the AWIBR constrains policy to depend on national variables with coefficients proportional to country weights. In the case of the MCIBR, instead, the parameters on the individual countries' inflation and output gap do not obey any proportionality constraint. The restriction imposed on the area-wide rule could be interpreted as a "political economy" constraint: to avoid the suspicion that country-specific concerns may affect its choices for the area, the monetary authority ties its own hands and decides not to react to country variables separately, but only to their aggregation.

[2]One advantage of reaction functions of this family, besides their simplicity, is that they have been shown to have good global stabilization properties (Benhabib, Schmitt-Grohé, and Uribe 2003).

[3]Specifically, w_y^j are 1999 GDP weights and w_π^j are 1999 consumer spending weights (under purchasing power parity). The weights are as follows: Germany: 0.43, 0.44; France: 0.29, 0.27; Italy: 0.28, 0.29.

To assess the performance of these rules, we adopt a standard quadratic, time-separable specification for the policymaker's loss function:

$$L_t = (1-\delta)E_t \sum_{\tau=0}^{\infty} \delta^{\tau} \left[\pi_{t+\tau}^2 + \lambda y_{t+\tau}^2 + \mu(\Delta i_{t+\tau})^2\right], \tag{3}$$

where δ is a discount factor, and λ and μ are parameters that reflect the weights attached by the policymaker to the variability of the output gap and of the policy interest rate changes relative to the variability of inflation around a target, assumed to be zero for simplicity. Note that both inflation and the output gap are area-wide variables, in line with the basic tenet that the monetary policy authority is solely interested in area-wide developments.

For $\delta \longrightarrow 1$ the intertemporal loss function can be interpreted as the unconditional mean of the period loss functions, which in turn is given by the weighted sum of the unconditional variances of the target variables (see Rudebusch and Svensson 1999):

$$L_t = \mathrm{var}(\pi_t) + \lambda \mathrm{var}(y_t) + \mu \mathrm{var}(\Delta i_t). \tag{4}$$

The optimal AWIBR (MCIBR) is found by searching for values of the three parameters in (2) (the seven parameters in (1)) that minimize the loss function (4) subject to the constraints given by the estimated multicountry models described in the next section. As a benchmark, we also compute the fully optimal rule (FOR), i.e., the rule derived from the unconstrained solution to the dynamic problem of minimizing (4) subject to the model economy, which depends on all state variables of the models.

3. Two Simple Models

In this section we present the two main models that we use throughout the paper. Both consist of an aggregate supply (AS) equation or Phillips curve and an aggregate demand (AD) equation or IS curve for each of the three main economies in the euro area—Germany, France, and Italy—which jointly account for over 70 percent of the area GDP. The first is a backward-looking model, popularized by Rudebusch and Svensson (1999) and Rudebusch (2002). It displays remarkable parameter stability and tracks the data well; furthermore, it is simple to estimate and use. However, its theoretical grounding is weak. The second model, presented in section 3.2, features backward- as well as forward-looking right-hand-side variables. It belongs to a class of models that has become increasingly popular, as the models tend to fit the data better than the purely forward-looking version initially suggested by the theoretical New Keynesian literature (see Goodfriend and King 1997, Rotemberg and Woodford 1997, and Clarida, Galí, and Gertler 1999) and can at the same time be reconciled with theory.[4] However, these models have two main drawbacks. First, the estimated parameters tend to be fragile: as we document below, our model is very sensitive to the choice of the sample period. Second, the quantitative importance of forward-looking elements, which turns out to be crucial for our results, is particularly hard to pin down with precision. Indeed, this issue is the subject of an ongoing debate. Among the latest contributions, Galí, Gertler, and López-Salido (2005) and Sbordone (2005) argue that forward-looking behavior is dominant, whereas others, including Lindé (2005) and Rudd and Whelan (2005a, 2005b), criticize this viewpoint. All these authors concur that the estimation of the degree of forward-lookingness of the economy is fraught with econometric difficulties, especially in the context of a relatively large system of equations.

In our view, these reasons warrant an assessment of the sensitivity of our results to the model choice.

The models were estimated with quarterly data over the period 1978:Q1–2004:Q4. Inflation is measured as the quarter-on-quarter rate of change of the households' consumption deflator. Potential output was estimated by applying the band-pass filter (Baxter and King 1999) to the (log) real GDP for each country.

[4] Galí and Gertler (1999) show that a hybrid model for inflation can be derived under the assumption that a fraction of firms act as backward-looking price setters. Smets and Wouters (2003) and Christiano, Eichenbaum, and Evans (2005) show that lagged inflation and output terms may appear in the model as the result of Calvo-pricing firms with indexation to last-period inflation and habit persistence in consumers' behavior.

3.1 Backward-Looking Model

We started from a general specification involving up to six lags of each variable on the right-hand side of each equation. After eliminating statistically insignificant terms, the following specification was selected (standard errors are given in parentheses below each coefficient):[5]

$$
\begin{aligned}
\pi_t^G &= \underset{(.086)}{0.29}\,\pi_{t-1}^G + \underset{(.066)}{0.58}\,\pi_{t-4}^G + \underset{(.132)}{0.36}\,y_{t-1}^G + \underset{(restr.)}{0.13}\,\pi_t^F + \varepsilon_{\pi t}^G \\
y_t^G &= \underset{(.058)}{0.77}\,y_{t-1}^G - \underset{(.034)}{0.08}\,(i_{t-2} - \pi_{t-2}^G) + \varepsilon_{yt}^G \\
\pi_t^F &= \underset{(.049)}{0.92}\,\pi_{t-1}^F + \underset{(.168)}{0.32}\,\frac{1}{4}\sum_{i=2}^{5} y_{t-i}^F + \underset{(restr.)}{0.08}\,\pi_t^G + \varepsilon_{\pi t}^F \\
y_t^F &= \underset{(.047)}{0.86}\,y_{t-1}^F - \underset{(.015)}{0.04}\,(i_{t-2} - \pi_{t-2}^F) + \varepsilon_{yt}^F \\
\pi_t^I &= \underset{(.010)}{0.96}\,\pi_{t-1}^I + \underset{(.108)}{0.24}\,y_t^I + \underset{(restr.)}{0.04}\,\pi_t^G + \varepsilon_{\pi t}^I \\
y_t^I &= \underset{(.065)}{0.70}\,y_{t-1}^I + \underset{(.067)}{0.11}\,y_t^G - \underset{(.014)}{0.03}\,(i_{t-1} - \pi_{t-i}^I) + \varepsilon_{yt}^I,
\end{aligned}
\tag{5}
$$

where ε_k^j $(j = G, F, I; k = \pi, y)$ are serially uncorrelated disturbances. Due to the presence of simultaneous terms, the model was estimated with 3SLS.[6]

[5]The values of some coefficients were restricted so as to impose long-run verticality of the national Phillips curves. The restriction was tested and accepted in all cases (see Monteforte and Siviero 2008). The adjusted R^2 for the equations are, respectively, 0.49, 0.64, 0.89, 0.78, 0.96, and 0.73.

[6]Constant terms were used in estimation but are not displayed in equations (5), which we use in the optimization exercises. This is in line with the approach followed in similar literature and amounts to interpreting equation systems such as (5) as a description of small deviations from the equilibrium. Since for most of the estimation period the exchange rates among the German, French, and Italian currencies were not fixed, though constrained by the Exchange Rate Mechanism of the European Monetary System, in estimation the inflation “imported” in country j from country i was constructed as the sum of the inflation rate in country i and the percentage variation of the exchange rate between the two countries (lagged values of all variables were used as instruments for the latter). In equations (5), to be used for the experiments presented below, the percentage change of the exchange rate was set identically equal to zero, consistent with the introduction of the single currency as of January 1, 1999.

The main nonstandard ingredient of model (5) is represented by the interactions among the various countries: to capture trade links, we let the inflation and output gap of each country have an effect on the corresponding variables in the other two countries. From a theoretical viewpoint, this choice is grounded in the two-country model of a monetary union in Benigno (2004). He shows that each country's Phillips curve depends, inter alia, on the terms of trade but that, via repeated substitutions, a specification can be obtained in which inflation in one country depends on its own lagged values and on the other country's current and lagged inflation.

The main properties of specification (5), as summarized by the impulse responses obtained by simulating the model augmented with a Taylor-type stabilizing policy rule, are roughly in line with well-established stylized facts about the euro-area economy.[7] The model estimates support the view that a significant degree of heterogeneity exists among the three economies analyzed. This result is in line with some of the most recent literature, although the issue is far from settled.[8] Further, heterogeneities within the euro area do not seem to have faded in recent times. Indeed, model (5) displays remarkable parameter stability: truncating the sample before the start of the monetary union (1978:Q1–1998:Q4) yields negligible changes in the parameter values (see the estimates in Angelini et al. 2002); also, we failed to detect appreciable signs of convergence in the shocks hitting the three economies considered.[9]

[7] In particular, the shocks hitting the AS and AD equations in France tend to have more persistent effects on area-wide inflation than the corresponding shocks hitting the Italian or the German economies; the effects of an AD shock are smallest and less volatile if they originate in Italy, while the effects on the output gap of an AS shock originating in the same country are largest and most volatile. Monetary policy takes longer to affect inflation in France than in either Italy or Germany; the time pattern in the latter two countries is similar, but the effect is markedly more pronounced in Italy.

[8] A large research effort devoted by the Eurosystem to the understanding of the monetary transmission mechanism in the euro area (see Angeloni, Kashyap, and Mojon 2003) shows that significant cross-country differences in the transmission mechanism exist but that alternative methods, data sets, and sample periods yield different, sometimes conflicting results.

[9] We compared the variance-covariance matrix of the model estimated with data up to 1998:Q4 with that of the model estimated with data up to 2004:Q4. The correlation of the shocks rises somewhat in the more recent period. However,

3.2 Hybrid Model

The model was estimated with GMM, using three lags of the variables as instruments. As in the case of the backward-looking model, we started from a general specification, featuring several leads and lags of the variables on the right-hand side, and gradually eliminated nonsignificant terms. The final version of the model is the following:

$$
\begin{aligned}
\pi_t^G &= \underset{(0.140)}{0.59}\, E_t\pi_{t+4}^G + \underset{(0.140)}{0.36}\, \pi_{t-4}^G + \underset{(restr.)}{0.05}\, \pi_t^F + \underset{(0.096)}{0.28}\, y_{t-1}^G + e_{\pi t}^G \\
y_t^G &= \underset{(0.039)}{0.69}\, y_{t-1}^G - \underset{(0.015)}{0.03}\left(i_{t-1} - E_{t-1}\widetilde{\pi}_{t+3}^G\right) + e_{yt}^G \\
\pi_t^F &= \underset{(0.131)}{0.69}\, E_t\pi_{t+3}^F + \underset{(0.125)}{0.27}\, \pi_{t-1}^F + \underset{(restr.)}{0.04}\, \pi_t^G \\
&\quad + \underset{(0.14)}{0.16}\left(\frac{1}{2}\left(y_{t-4}^F + y_{t-5}^F\right)\right) + e_{\pi t}^F \qquad (6) \\
y_t^F &= \underset{(0.036)}{0.83}\, y_{t-1}^F - \underset{(0.012)}{0.03}\left(i_{t-1} - E_{t-1}\widetilde{\pi}_{t+3}^F\right) + e_{yt}^F \\
\pi_t^I &= \underset{(0.196)}{0.49}\, E_t\pi_{t+2}^I + \underset{(0.200)}{0.46}\, \pi_{t-1}^I + \underset{(restr.)}{0.05}\, \pi_t^G \\
&\quad + \underset{(0.14)}{0.28}\left(\frac{1}{2}\left(y_{t-2}^I + y_{t-3}^I\right)\right) + e_{\pi t}^I \\
y_t^I &= \underset{(0.046)}{0.75}\, y_{t-1}^I + \underset{(0.047)}{0.11}\, y_t^G - \underset{(0.11)}{0.03}\left(i_t - E_t\widetilde{\pi}_{t+4}^I\right) + e_{yt}^I,
\end{aligned}
$$

where $\widetilde{\pi}_t^j = (1/4)\sum_{i=0}^{3}\pi_{t-i}^j$. Altogether, the model fits the data well.[10] However, whereas the backward-looking model (5) displays remarkable parameter stability over time, the hybrid model estimates turn out to be very sensitive to the choice of the sample period and of the instruments. For instance, if the first two years of data are dropped from the estimation sample, all the real interest rate terms in the aggregate demand equations become insignificant, whereas the size of the output-gap coefficients in the aggregate supply equations declines dramatically (as much as halving for both

using the model estimated on the shorter data set to replicate the experiments conducted in section 4 below yields very similar results.

[10]The adjusted R^2 for the equations are, respectively, 0.48, 0.57, 0.87, 0.70, 0.94, and 0.72. As for model (5), long-run verticality of the national Phillips curves was imposed.

Germany and Italy). We also attempted measuring the output gap with a proxy for real marginal cost (arguably a better measure from a theoretical viewpoint), as proposed by Clarida, Galí, and Gertler (1999). However, such alternative proxy was not found to improve the fit of the model nor the robustness of the estimates.

From a methodological viewpoint, the only difference when using this model instead of model (5) is that the iteration process needed to compute the optimal coefficients of the AWIBR and MCIBR requires an additional step to solve the model (augmented with the policy rule) for the rational-expectation equilibrium.[11]

4. How Large Are the Costs of Overlooking National Information?

In this section we assess the importance of national information using the two models estimated in the previous section.

4.1 Exercises Using the Backward-Looking Model

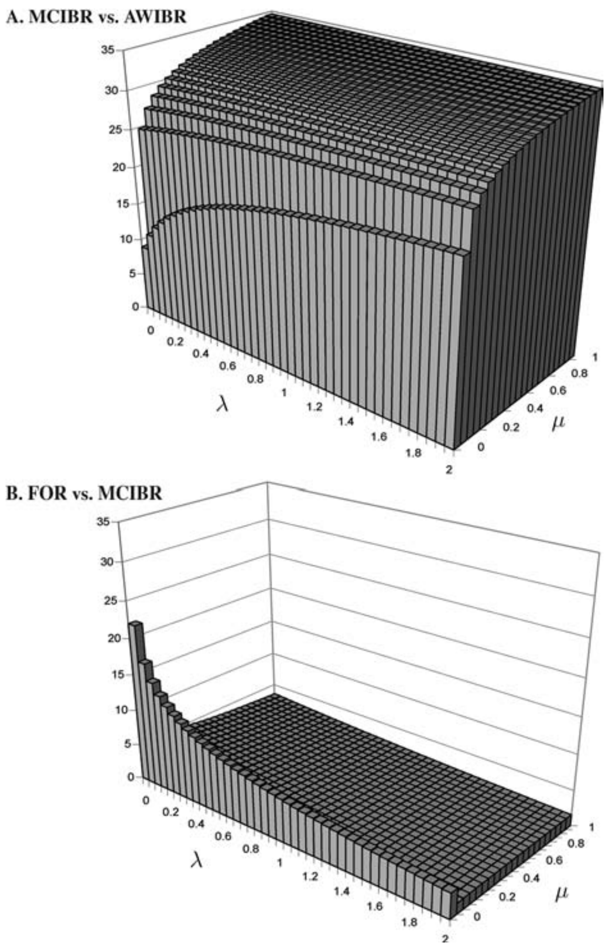
Using model (5), we computed the optimal losses under the optimized AWIBR (equation (2)), the optimized MCIBR (equation (1)), and the fully optimal rule (FOR) for different choices of the weights λ and μ assigned to output-gap and interest rate variability in the loss function (4). Since (2) is a restricted version of (1), the loss under the former is by construction at least as large as under the latter. The issue is, are the welfare gains under the MCIBR large enough to conclude that national information is important?

The results are reported in figure 1. Panel A shows that the answer is positive. Relative to the AWIBR, the MCIBR yields a loss reduction in the 25 to 35 percent range for most combinations of the preference parameters. Only when $\mu = 0$ (i.e., when there is no concern for interest rate variability) does the loss reduction shrink

[11] The solution found corresponds to what Clarida, Galí, and Gertler (1999) call "constrained commitment." It involves a form of commitment, since the central bank follows a rule decided in advance, rather than reoptimizing every period. It is, however, constrained to a class of rules that does not include the "full commitment" solution, which would also require the central bank to react to lagged values of the state variables.

Figure 1. Loss Reduction in the Backward-Looking Model

(in percent)



Note: The figure reports the percentage reduction in the value of the loss function attained under the multicountry information-based rule (MCIBR) relative to the area-wide information-based rule (AWIBR) under the constraint represented by the backward-looking model in section 3.1. The parameters λ and μ reflect the weight attached by the policymaker to the output gap and the interest rate variability, respectively (see equation (3) in the text). Panel A reports the percentage reduction in the value of the loss function attained under the MCIBR relative to the AWIBR. Panel B reports the percentage reduction in the value of the loss function attained under the fully optimal rule (FOR) relative to the MCIBR.

to 10–20 percent, and only when both λ and μ are zero, corresponding to an extreme form of inflation targeting, does it reduce to just below 10 percent. Clearly, in terms of "welfare units" the size of this improvement is arbitrary, as the central bank loss function is only identified up to a positive linear transformation. However, the improvement can be directly interpreted as a summary measure of the reductions of the unconditional variances of inflation, output gap, and interest rate changes. The variance reduction is between 10 and 30 percent for inflation, 20 and 35 percent for the output gap, and 20 and 45 percent for interest rate changes. Summing up, under model (5) the welfare gains associated with the use of national information are undisputedly large. Panel B of the figure shows that the MCIBR does a good job relative to the fully optimal rule. The latter improves upon the former by less than 4 percent on average across the various combinations of the preference parameters (the only sizable improvement occurring when no weight is assigned to interest rate variability).

To check the sensitivity of these results to parameter uncertainty, we ran one-tailed tests of the hypothesis that the average loss associated with the MCIBR and the AWIBR are the same.[12] The tests allow for rejection of the null hypothesis for most combinations of preference parameters, suggesting that the results in figure 1 are fairly robust to parameter uncertainty.

A second issue we explore is, what are the consequences of the adoption of the MCIBR at the country level? Since policy is by construction aimed at minimizing the aggregate loss (4), the adoption of the MCIBR need not yield a Pareto improvement for the three economies comprising our model. Indeed, in our exercises the optimal MCIBR systematically entails, vis-à-vis the optimal AWIBR, a reduction in the national loss function for Germany and

[12]To this end, for each combination of (λ, μ) we extracted one realization from the empirical distribution of the estimated model coefficients and, without recomputing the AWIBR and MCIBR, we simulated the model for 800 periods under either one or the other of the two competing rules. The process was repeated 5,000 times. We then computed (using the last 400 simulated values) the average loss associated with each rule, as well as the corresponding variances and covariances, and applied the standard test for equality of means of normal distributions.

France.[13] By contrast, Italy always fares better under the AWIBR. For instance, when $\lambda = \mu = 1$, Germany and France experience a loss reduction of roughly 40 percent, as opposed to an increase of almost 30 percent for Italy (see panel A of figure 4 in the appendix). We tried to assess how these results are affected by country size, but failed to detect clear patterns, as country size interacts with the specific features of the estimated country structures and shocks.[14] The comparison of the two rules, reported in table 1, also fails to display clear patterns. Overall, the MCIBR seems less aggressive with respect to inflation: the impact coefficient is 0.93, as opposed to 1.24 in the case of the AWIBR. This reflects reactivity that is lower for Italy and for Germany, and equal for France. In the case of the output gap, the MCIBR still underweighs Italy but overweighs France and leaves the German weight unchanged.

Finally, we analyze the issue of convergence. Clearly, if the national economies were identical, the different performance of the two rules would vanish. Therefore, it seems reasonable to conjecture that the potential relevance of regional information for monetary policy would gradually disappear were the regional asymmetries to fade away. As argued in section 3.1, no clear signs of recent convergence among the euro-area economies can yet be detected. However, the possibility of a gradual reduction in asymmetries in the future clearly cannot be ruled out. Therefore, we assess how our results would be affected should more symmetry among the regional economies gradually prevail in the future. Our counterfactual exercises, described in the appendix, provide two main results. First,

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[13]For each of the three economies, the national loss function is defined as in equation (4), except that the inflation and output-gap terms are replaced by their respective national counterparts.

[14]Specifically, we performed the following counterfactual exercises. We gradually increased the size of each country w^j, one at a time, from 1 to 99 percent; correspondingly, the size of the other two countries was reduced from 49.5 to 0.5 percent. Every time, the optimization exercise was rerun and the national and area-wide losses were recomputed. When the weight of Italy is increased, the area-wide benefits of the MCIBR (versus the AWIBR) monotonically increase (to above 60 percent), as do those for France and Germany, whereas the outcome for Italy initially worsens and then improves. However, when the same exercise is run for either Germany or France, the area-wide as well as their national benefits decline (to about 20 percent), while the benefits for Italy go from negative to positive.

Table 1. Optimal Rule Coefficients ($\lambda = 1, \mu = 0.5$)

	Inflation				**Output Gap**				**Lagged Interest Rate**
	Area	Fra	Ger	Ita	Area	Fra	Ger	Ita	
Backward-Looking Model									
Area-Wide Rule	1.24	**0.34**	**0.55**	**0.35**	1.23	**0.36**	**0.56**	**0.31**	0.78
Multicountry Rule	**0.93**	0.34	0.43	0.16	**1.23**	0.51	0.55	0.16	0.76
Hybrid Model									
Area-Wide Rule	0.25	**0.07**	**0.11**	**0.07**	0.42	**0.19**	**0.12**	**0.11**	0.95
Multicountry Rule	**0.23**	0.05	0.12	0.06	**0.44**	0.11	0.17	0.16	0.95

Note: The coefficients in bold are obtained by aggregating (in the case of the multicountry rule) or disaggregating (for the area-wide rule) the computed coefficients, using country weights.

heterogeneity of *both* country models and country shocks is needed for national information to be valuable. Under full convergence of either type, the relative gain associated with the MCIBR is negligible. Second, convergence of shocks or models has a different effect on the usefulness of national information. A moderate degree of convergence of models alone is sufficient for the welfare gains associated with the MCIBR to fall below 10 percent. By contrast, almost perfect convergence of shocks is needed for national information to become irrelevant.

4.2 Exercises Using the Hybrid Model

In this section we replicate the basic exercise underlying figure 1 using the hybrid model (6). The optimization exercise was run for the two rules (1) and (2) and various (λ, μ) combinations, and the percentage loss reduction was computed. Figure 2—the analogue of figure 1, panel A—reports the results. The magnitude of the loss reduction is now negligible, reaching a maximum of 3 percent for the pure inflation-targeting case.

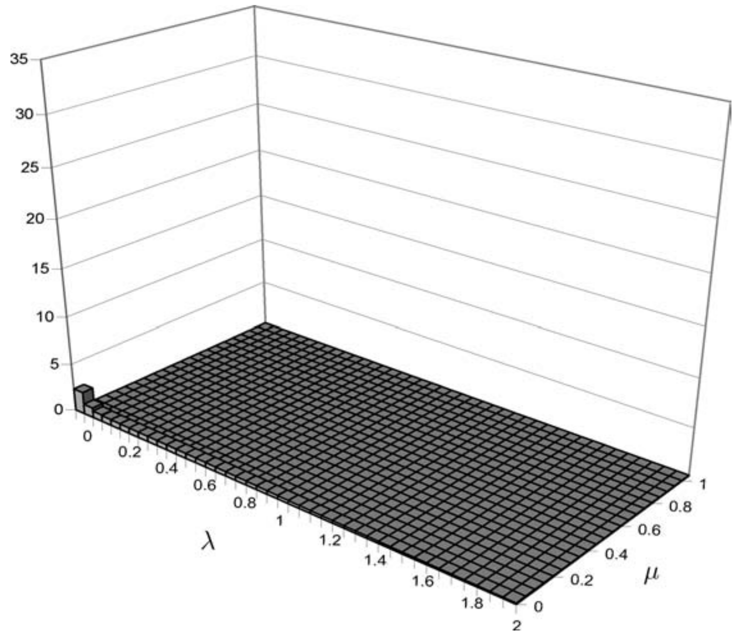
This clearly raises the question, what drives this change in the results? We first checked whether the change can be attributed to any particular parameter of the hybrid model (6). Setting $\lambda = 1$ and $\mu = 0.5$ (similar conclusions hold for all combinations of preference parameters) we perturbed, one at a time, each of the following groups of coefficients, starting from their respective estimated values: (i) the sensitivity of inflation to the output gap (AS equations); (ii) the sensitivity of the output gap to its lagged value or (iii) to the real interest rate (AD equations); and (iv) the degree of forward-lookingness (AS and AD equations). After each perturbation, we recomputed the loss reduction.[15]

The results are shown in figure 3. The horizontal axis reports the percentage perturbation applied to each group of coefficients, up to 60 percent. The vertical axis measures the loss reduction. Small

[15]Specifically, for groups (i) through (iii) we reduced the coefficients' value; for group (iv) we reduced the coefficient on expected variables and increased those of the corresponding lagged variables, so as to leave constant the sum of the two, hence preserving the long-run verticality of the Phillips curve.

Figure 2. Loss Reduction in the Hybrid Model

(in percent; MCIBR vs. AWIBR)

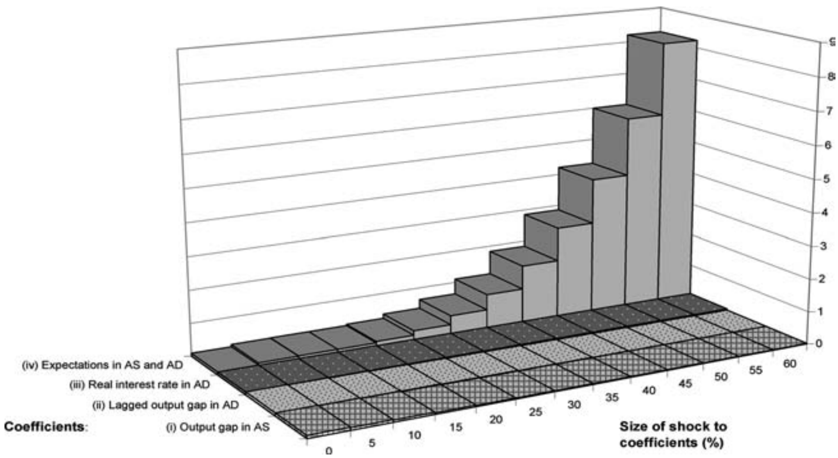


Note: The figure reports the percentage reduction in the value of the loss function attained under the multicountry information-based rule (MCIBR) relative to the area-wide information-based rule (AWIBR) under the constraint represented by the hybrid backward-forward-looking model in section 3.2. The parameters λ and μ reflect the weight attached by the policymaker to the output gap and the interest rate variability, respectively (see equation (3) in the text).

shocks—up to 15–20 percent of the estimated coefficient value—have no appreciable effect on the results, regardless of the coefficient group: the MCIBR still delivers negligible gains. For families of coefficients (i) to (iii), this also holds true for large shocks. However, the loss reduction is sensibly affected when the degree of forward-lookingness is modified. Namely, as the model becomes predominantly backward looking, the loss reduction goes back toward the high values documented in the previous section: applying a 60 percent shock to this family of parameters, the loss reduction increases

Figure 3. Sensitivity of the Loss Reduction in the Hybrid Model

(in percent; MCIBR vs. AWIBR; $\lambda = 1$, $\mu = 0.5$)



Note: The figure reports the sensitivity of the loss reduction in figure 2 to shocks to the coefficients of the hybrid model. Shocks are defined as a percentage reduction in the estimated coefficient value. See section 4.2 for a detailed description of the exercise.

to around 8 percent.[16] The pattern is highly nonlinear: a 40 percent shock yields a loss reduction of only 2 percent.

As an additional check that the change in results relative to those of section 4.1 is mainly driven by the forward-looking nature of model (6), we constructed an artificial hybrid model using the estimated backward equations (5). We parameterized the degree of forward-lookingness by $\phi \in [0, 1]$ so that $\phi = 0$ would yield model (5), whereas $\phi = 1$ would yield a fully forward-looking model with no autoregressive terms.[17] We replicated our optimization exercises using this model, letting ϕ gradually increase from zero to one. The

[16] Using shocks as large as 60 percent of the coefficient size clearly raises doubts about the validity of the approximation. However, we believe the experiment is still helpful in understanding what drives the difference in the results observed with the two models.

[17] For instance, we adopted the following version of the AS equation for Italy: $\pi_t^I = \phi 0.96 E_t \pi_{t+1}^I + (1 - \phi) 0.96 \pi_{t-1}^I + 0.24 y_t^I + 0.04 \pi_t^G + \varepsilon_{\pi t}^I$.

results of the exercises, not reported, confirm the importance of the degree of inertia of the economy: for values of $\phi < 0.4$ the results of section 4.1 are replicated; as ϕ approaches 0.5 the importance of national information declines, becoming negligible for $\phi > 0.7$.

These results are in line with those obtained in studies using dynamic stochastic general equilibrium models incorporating forward-looking terms. Jondeau and Sahuc (2008) show that significant welfare improvements can be obtained using a multicountry model, while those stemming from a multicountry rule are limited.

The near irrelevance of national information under the hybrid economy (6) seems to reflect a general reduced sensitivity of the loss to the precise specification of the policy in a forward-looking environment, as suggested by the following experiment. Setting $\lambda = 1$ and $\mu = 0.5$ we shocked, one at a time, the six country-specific coefficients of the optimized MCIBR and computed the increase in the loss function. The results obtained using the two models are dramatically different. For instance, using the hybrid model, a 50 percent reduction in any of the three coefficients of inflation yields a negligible loss deterioration, always below 1 percent; in the backward-looking model, by contrast, the deterioration ranges between 20 and 40 percent.

This feature of the forward-looking model need not be surprising, since the economy is less heavily dependent on its past and the behavior of the endogenous variables is more directly driven by that of the exogenous shocks. Policy, in these circumstances, can do relatively little. In particular, as noted by Clarida, Galí, and Gertler (1999), in the pure forward-looking case even the fully optimal policy (under discretion or "constrained commitment") cannot affect the rate at which inflation converges back to its target value, when displaced by a cost-push shock. This is not true in a model with inflation inertia. Therefore, in a model exhibiting non-negligible inertial behavior, policy appears to be more important. Results with a similar flavor are obtained by Adalid et al. (2005).

5. Concluding Remarks

In this paper we address an important question for the design of monetary policy for a monetary union: what is the cost of neglecting regional information in the decision-making process? Our answer is not clear-cut. We find that the cost is large using a backward-looking model of the economy; it is negligible using a hybrid model, featuring expected variables on the right-hand side. The main determinant of this finding seems to be the different degree of inertia (or its opposite, the forward-lookingness) characterizing our two models. Ideally, one should be able to identify the “best” model and pick one of the two answers. However, as we discussed above, both models have pros and cons: the backward-looking estimates are econometrically more reliable but theoretically unappealing, whereas the opposite holds true for the hybrid model. Indeed, the issue of the quantitative importance of forward-looking elements in New Keynesian Phillips curves, which turns out to be crucial for our results, is the subject of an ongoing debate. In this paper we do not take a stand in this debate, and therefore we remain somewhat agnostic about the answer to our key question. However, we believe that our results warrant a close scrutiny of the potential benefits of regional information for monetary policy purposes, certainly closer than the virtual neglect this issue has received so far.

The exercises performed in the paper yield several other interesting insights. For instance, in the backward-looking framework, the reduction of area-wide inflation and output variability associated with the use of national information does not reflect similar improvements in all participating countries; indeed, some countries may experience an increase in the variability of their economies. In the hybrid framework, the near irrelevance of national information seems to reflect a general reduced sensitivity of the loss function to the precise specification of the policy in a forward-looking environment: halving any of the three inflation coefficients of the optimal policy rule yields a negligible loss increase, always less than 1 percent; replicating the exercise in the backward-looking economy yields a loss increase of 20 to 40 percent.

Our concluding caveat is that none of our results should be interpreted as providing precise indications on the appropriate reaction of monetary policy to the actual asymmetries prevailing in the euro area, as the models we use are much too simple. Rather, we would stress the methodological case for paying attention to the potential costs of neglecting regional information in the design of the monetary policy in a monetary union, and for pursuing a disaggregate modeling strategy. Obvious extensions of this line of research would include further checking the robustness of the results to alternative modeling choices, exploring the monetary policy implications of the well-documented heterogeneities at the sector level within the monetary union (Dedola and Lippi 2005), and looking at other monetary unions.

Appendix. Relevance of Regional Information as the Economies Converge

In this appendix we assess in detail how the comparison between the AWIBR and the MCIBR reported in section 4.1 would be affected should more symmetry among the regional economies gradually prevail in the future.

We consider the two cases of convergence in shocks and structures separately. Regarding the convergence of stochastic disturbances, we take full AD (or AS) convergence to mean that the shocks in the corresponding equation become exactly the same in all countries (hence, the cross-country correlations equal 1—unlike fully optimal rules, optimal simple rules are affected by changes in the covariance matrix of the stochastic disturbances; see Currie and Levine 1985, 1987). As in De Grauwe and Piskorski (2001), we define the variance of the common AD shock under full convergence as

$$\overline{\sigma}_y^2 \equiv \left(w_y^G \sigma_y^G + w_y^F \sigma_y^F + w_y^I \sigma_y^I\right)^2, \tag{7}$$

where σ_y^i is the estimated standard deviation of ε_y^i in country $i (i = G, F, I)$, and w_y^i is the corresponding GDP weight.

We also consider the possibility of partial convergence, which we parameterize by $\xi_{AD}^s \in [0, 1]$ (with the extrema of the interval corresponding to zero and full convergence, respectively). The elements of the variance-covariance matrix of the shocks under partial convergence are defined as

$$\left(\widetilde{\sigma}_y^i\right)^2 \equiv \xi_{AD}^s \overline{\sigma}_y^2 + \left(1 - \xi_{AD}^s\right)\left(\sigma_y^i\right)^2 \tag{8}$$

$$\widetilde{\sigma}_y^{ij} \equiv \xi_{AD}^s \widetilde{\sigma}_y^i \widetilde{\sigma}_y^j \tag{9}$$

for all countries i, j, so that the correlation of shocks among countries is given by ξ_{AD}^s itself. Convergence of AS shocks is defined in an analogous way and parameterized by ξ_{AS}^s.

Full and partial convergence of economic structures (i.e., the parameters of the model) is modeled in a similar fashion. In particular, it is assumed that, under full convergence, inflation and output gap in each country respond to their determinants in the same way, except for the effects of country size. For each AS and AD equations, the coefficients under full convergence are defined as a weighted average of the original coefficients in all countries. The average cross-country effects at full convergence are computed as a weighted average of the original ones in equations (5), so that each country is influenced by the other two. Partial convergence is modeled as in the case of disturbances, with the parameters controlling the degree of convergence for the AD and AS equations denoted ξ^m_{AD} and ξ^m_{AS}, respectively.

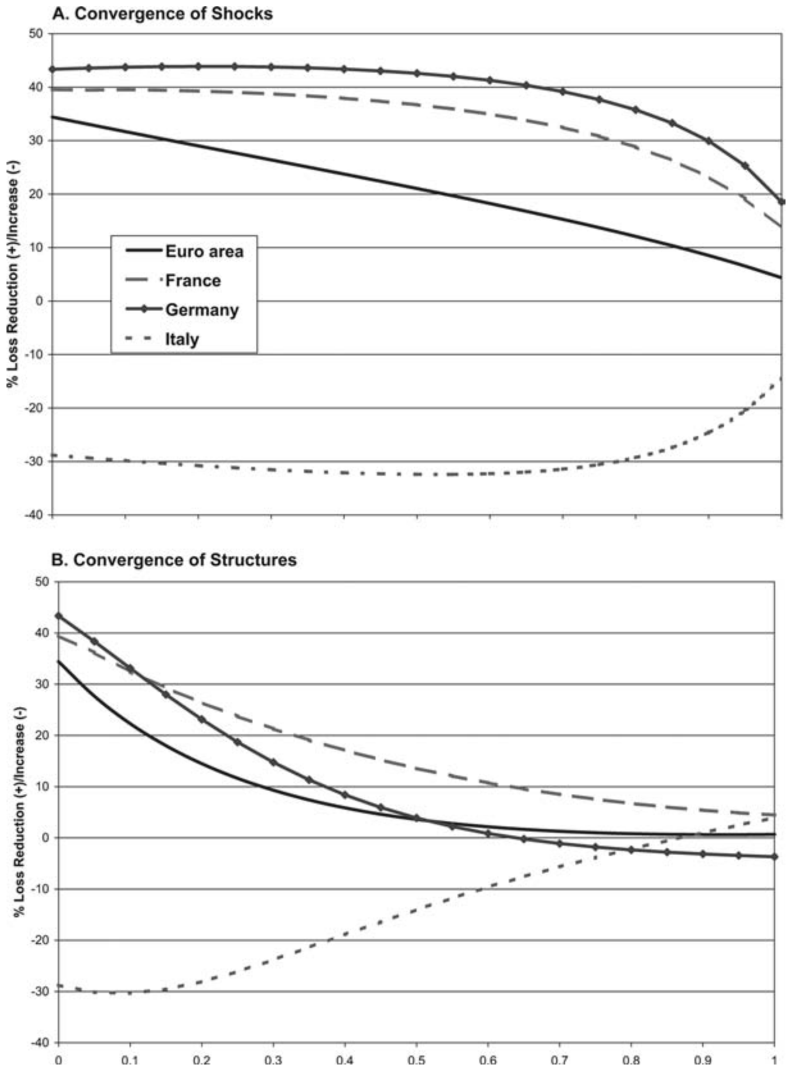
Figure 4 plots the reduction of the loss attained with the MCIBR, relative to that achievable with the AWIBR, as a function of the degree of convergence, for the case $\lambda = \mu = 1$ (similar results are obtained with alternative (λ, μ) combinations). In panel A we assume symmetric convergence in the shocks (i.e., $\xi^s_{AS} = \xi^s_{AD} \in [0, 1]$) and no convergence in the parameters of the models (i.e., $\xi^m_{AS} = \xi^m_{AD} = 0$). Conversely, in panel B the lines refer to the case of symmetric convergence of the economic structures, keeping the shocks unchanged. Two main results emerge from the figure.

First, heterogeneity of *both* country models and country shocks is needed for national information to be valuable. Under full convergence of either type, the relative gain associated with the MCIBR is negligible.

Second, convergence of shocks or models has a different effect on the relevance of national information. A moderate degree of (symmetric) convergence of models alone ($\xi^m_{AS} = \xi^m_{AD} > 0.3$) is sufficient for the welfare gains associated with the MCIBR to fall below 10 percent. By contrast, for $\xi^s_{AS} = \xi^s_{AD} = 0.7$ (without any convergence of models), such gains are still above 15 percent; almost perfect convergence of shocks is needed for national information to become irrelevant.

The figure also suggests that the country-specific implications of convergence of economic structures are roughly the same as for the area as a whole: a moderate degree of convergence of models suffices to reduce the difference in performance between the MCIBR and the AWIBR by a significant amount. By contrast, only with a

Figure 4. Usefulness of National Information as the Economies Converge
($\lambda = \mu = 1$)



Note: We simulate an increase in the symmetry of both AS and AD equations of the three countries in the estimated backward-looking model (5). On the vertical axis we measure the loss reduction (+) or increase (−) attained with the multi-country rule (MCIBR) relative to the area-wide rule (AWIBR). The horizontal axis measures the degree of convergence, increasing from left to right, so that zero corresponds to the case of no convergence.

rather high degree of shock convergence does the difference in performance between the two rules decline sensibly. For values of ξ^s_{AS} and ξ^s_{AD} below 0.7, in Germany, France, and Italy it is basically the same as in the case of no convergence, while it halves for the area as a whole.

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Capital Requirements and Bank Behavior in the Early 1990s: Cross-Country Evidence*

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This paper uses a simultaneous-equations model to investigate how banks from six G-10 countries adjusted their capital and their risk-weighted assets after the passage of the 1988 Basel Accord. In particular, the analysis tests whether weakly capitalized banks increased their capital or decreased their risk-weighted assets more rapidly than did well-capitalized banks. If so, did market discipline play a significant role? The results suggest that only in the United States were weakly capitalized banks observed to increase their capital ratios faster than well-capitalized banks; however, the weakly capitalized U.S. banks did not modify their risk-weighted assets at different rates from other U.S. banks. In addition, market discipline appears to have played an essential role: weakly capitalized U.S. banks that did not also face market pressure did not increase their capital ratios faster than other U.S. banks. This suggests that market pressure was an important factor in the capital build-up of the early 1990s.

JEL Codes: G21, G28.

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1. Introduction

One of the major developments undergone by the banking industry in the 1990s has been the worldwide implementation of the 1988 Basel Accord that set minimum capital standards for internationally active banks. The Basel guidelines were initially adopted by the central banking authorities from the G-10 countries. Their implementation started in 1989 and was completed four years later, in 1993. The purpose of the Accord was twofold. First, it aimed at creating a level playing field for banks by raising capital ratios, which were generally perceived as too low in some G-10 countries. Second, and connected to this, it aimed at promoting financial stability by linking the required amount of capital to a measure of the bank's risk-weighted assets. However, the relatively simple approach to calculating risk-weighted assets had the potential for distorting incentives for bank risk taking.

Twenty years after the adoption of the 1988 Basel standards, though still at the beginning of the implementation of the Basel II framework, it is fair to say that empirical research has not fully answered the following questions: Was the 1988 agreement effective in raising capital ratios among banking institutions, especially those whose initial ratios fell close to the minimum of the requirements? For banks that increased their capital adequacy ratios, did regulatory pressure play a greater role than market discipline? Analysis of how G-10 banks respond to capital standards is important given that some parts of the Basel II framework—e.g., the standardized approach to credit risk—represent a refinement of the 1988 standards.

The lack of answers to the questions raised above is mainly due to data limitations. Indeed, data on capital and credit risk of G-10 banks are often confidential or hard to obtain on a standardized cross-country basis. Existing studies on the impact of the 1988 Basel Accord focus on Japan, Switzerland, the United Kingdom, and the United States, while evidence remains scarce for other countries that were part of the Accord. Therefore, an important contribution of this paper is to shed further light on the impact of bank capital requirements in a number of G-10 countries for which studies have not yet been undertaken.

More precisely, this paper uses the simultaneous-equations model developed by Shrieves and Dahl (1992) to analyze adjustments in capital and credit risk at banks from six G-10 countries (Canada, France, Italy, Japan, the United Kingdom, and the United States) between 1988 and 1995. Credit risk, which is defined as the ratio of risk-weighted assets to total assets, is the only type of risk analyzed here since it was the main focus of the 1988 Basel Accord.[1]

Shrieves and Dahl's model has been used by several other studies documenting the impact of capital requirements on bank capital and credit risk, including Aggarwal and Jacques (1997, 2001) and Jacques and Nigro (1997) for the United States, and Rime (2001) for Switzerland.[2] These studies find little evidence that weakly capitalized banks adjust their ratio of risk-weighted assets to total assets following the introduction of bank capital requirements, but they find support for the hypothesis that these banks increase their capital-to-assets ratios faster than well-capitalized banks. The latter result is consistent with increased pressure from regulators or market participants following the introduction of bank capital requirements (Basel Committee on Banking Supervision 1999). However, the above-mentioned studies do not distinguish between both types of pressures, and they interpret their results as a sign of increased regulatory pressure.

In addition to focusing on a different set of countries, this paper contributes to the existing literature on the effects of capital requirements by disentangling the impact of regulatory and market pressures on bank capital and credit risk taking. In the analysis, regulatory pressure is measured by a dummy variable equal to one if a bank's capital ratio falls below some threshold and zero otherwise, while market pressure is measured by a dummy variable equal to one if a bank is listed or rated and zero otherwise.

[1]It is possible that G-10 banks modified other risks, such as interest or market risk, following the introduction of the Basel capital requirements. However, there is little empirical evidence on this (Basel Committee on Banking Supervision 1999).

[2]In the United States, it is difficult to distinguish between the effects of the 1988 Basel standards and the effects of the Federal Deposit Insurance Corporation Improvement Act (FDICIA), which was passed three years later.

Identifying the impact of regulatory and market pressures on bank capital and credit risk is not only important in the context of the 1988 capital adequacy rules but is also relevant for the Basel II framework. Indeed, the first pillar of the New Accord (minimum capital requirements) is supplemented by two other pillars, where the third (market discipline) is intended to promote higher disclosure standards and reinforce market pressure on banks to hold adequate capital ratios. However, very little is known about the effectiveness of market discipline in complementing regulatory pressure in order to increase capital and decrease risk taking among banks.

Consistent with the existing literature, this paper finds that, ceteris paribus, weakly capitalized U.S. banks increased their total capital ratio faster than did well-capitalized U.S. banks in the early 1990s. However, and contrary to previous studies, the analysis suggests that this increase was due to both regulatory and market pressures rather than regulatory pressure alone.

As regards the other G-10 countries included in the study, little evidence is found that weakly capitalized banks raised their capital-to-assets ratios at a faster rate than well-capitalized banks. In addition, no evidence is found that U.S. or non-U.S. weakly capitalized banks modified their ratio of risk-weighted assets to total assets differently from well-capitalized banks.

Taken as a whole, these results suggest that the effectiveness of the 1988 bank capital requirements to increase capital and/or reduce credit risk was rather limited outside the United States, where it reflected both regulatory and market pressures.

The remainder of the paper is organized as follows. Section 2 briefly reviews the literature on the impact of capital requirements on bank capital and credit risk and summarizes the 1988 capital standards. Section 3 presents the data used in the analysis, while section 4 describes the methodology. Results are presented in section 5 and conclusions are drawn in section 6.

2. Capital Requirements and Bank Behavior

2.1 Review of the Theoretical Literature

One of the main justifications for regulating bank capital is the need to avoid the risk-shifting incentive generated by improperly priced deposit insurance. Although it may promote financial stability in the short run, risk-insensitive deposit insurance tends indeed to reduce banks' incentives to maintain adequate capital and may thus endanger stability in the long run. The ability of capital standards to successfully eliminate this moral hazard problem has been at the heart of a theoretical debate for more than twenty-five years.

A first strand of the literature focuses on utility-maximizing banks using the portfolio approach of Pyle (1971) and Hart and Jaffee (1974), which explains the existence of financial intermediaries within a mean-variance framework. In this setting, Koehn and Santomero (1980) show that the introduction of higher capital-to-assets ratios will lead banks to shift their portfolio to riskier assets and that this reshuffling effect will be larger for institutions that initially held relatively more risky assets per unit of capital.

This conclusion has been challenged by Furlong and Keeley (1989) and Keeley and Furlong (1990), who use an option model and find that a higher capital ratio does not lead banks to increase asset risk. Both papers contend that the mean-variance framework, which reaches opposite conclusions, is inappropriate because it does not adequately describe the bank's investment opportunity set by neglecting the option value of deposit insurance and the possibility of bank failure.

One way to eliminate the risk-shifting incentive is to require banks to meet risk-related capital ratios, as suggested by Kim and Santomero (1988). However, Rochet (1992) shows that when the objective of banks is to maximize the market value of their future profits, risk-related capital ratios cannot prevent them from choosing very specialized and very risky portfolios.

In a nutshell, theoretical contributions do not agree on whether imposing harsher capital requirements leads banks to increase the risk structure of their portfolios. However, these studies suggest that the impact of capital requirements on bank capital and credit risk depends on the extent to which such requirements are binding. Moreover, the degree of response of capital and credit risk to capital requirements may be affected by the presence of market discipline (Basel Committee on Banking Supervision 1999).

The next section attempts to clarify further the relation between capital and credit risk taking by briefly restating the key rules of the 1988 Basel Accord and analyzing how banks can comply with them.

2.2 The 1988 Basel Accord

As mentioned above, the 1988 Basel standards were entirely focused on credit risk. An amendment to incorporate market risk was included in 1996, and the Basel Committee on Banking Supervision issued a revised capital adequacy framework in June 2004. This new framework, which replaces the 1988 standards, is based on three mutually reinforcing pillars that allow banks and supervisors to evaluate additional types of risks such as operational risk (Basel Committee on Banking Supervision 2004). The implementation of the Basel II framework began in 2007 in Europe and in 2008 in the United States.

Under the 1988 Basel Accord, internationally active banks were required to meet two capital adequacy ratios: the tier 1 and total capital ratios.

The tier 1 ratio is equal to tier 1 capital divided by risk-weighted assets. Tier 1 capital consists mainly of stockholder equity capital and disclosed reserves, while risk-weighted assets are calculated by assigning each asset and off-balance-sheet item to one of four broad risk categories. These categories receive risk weights of 0 percent, 20 percent, 50 percent, and 100 percent, with riskier assets being placed in the higher-percentage categories. For example, the 0 percent category consists of assets with zero default risk (e.g., cash, government bonds/securities), the 20 percent category consists of assets with a low rate of default (e.g., loans to OECD banks), the 50 percent category consists of medium-risk assets (essentially residential mortgage loans), and the 100 percent category consists of the remaining assets (in particular, loans to nonbanks).

The total capital ratio is the sum of tier 1 and tier 2 capital divided by risk-weighted assets. Tier 2 capital includes elements like undisclosed reserves and subordinated term debt instruments, provided that their original fixed term to maturity does not exceed five years.

The 1988 Basel standards required banks to have a tier 1 ratio of at least 4 percent and a total capital ratio of at least 8 percent, with the contribution of tier 2 capital to total capital not exceeding 50 percent.[3]

As shown in the appendix, banks that wish to raise their capital adequacy ratio (for regulatory or nonregulatory reasons) can use three types of balance-sheet adjustments: they can increase their capital level, decrease their risk-weighted assets, or sell off their assets. This is summarized in equation (1), which decomposes the growth rate of the capital adequacy ratio into three terms: the growth rate of capital, the growth rate of the credit-risk ratio, and the growth rate of total assets:

$$\frac{\dot{\mathrm{CAR}}}{\mathrm{CAR}} = \frac{\dot{\mathrm{K}}}{\mathrm{K}} - \frac{\dot{\mathrm{RISK}}}{\mathrm{RISK}} - \frac{\dot{\mathrm{A}}}{\mathrm{A}}, \tag{1}$$

where CAR = K/RWA = capital adequacy ratio (tier 1 ratio or total capital ratio); K = capital (tier 1 capital or total capital); RISK = RWA/A = risk-weighted assets/total assets = credit-risk ratio; and A = total assets. The dots denote time derivatives.

Thus, banks can increase their capital adequacy ratio (CAR) by raising their capital level (K), lowering their credit-risk ratio (RISK), or lowering their total assets (A). In a nutshell, the impact of an increase in capital requirements on bank capital and risk choices is not clear a priori. This paper attempts to clarify this impact by focusing on the behavior of weakly capitalized banks, which are under regulatory pressure to increase their capital adequacy ratios.

In the analysis, capital is defined as the capital-to-assets ratio (K/A) and risk as the credit-risk ratio (RWA/A) of banks. I adopt these definitions for the purpose of understanding how G-10 banks adjusted the numerator of their capital adequacy ratio following changes in its denominator, and vice-versa.[4] However, it is well

[3]Following the passage of FDICIA in 1991, U.S. banks were also required to comply with a third ratio—namely, a tier 1 leverage ratio of at least 4 percent. Under FDICIA, banks are classified in three main categories: (i) well-capitalized (total capital ratio $\geq$ 10 percent, tier 1 ratio $\geq$ 6 percent, and tier 1 leverage ratio $\geq$ 5 percent), (ii) adequately capitalized (total capital ratio $\geq$ 8 percent, tier 1 ratio $\geq$ 4 percent, and tier 1 leverage ratio $\geq$ 4 percent), and (iii) undercapitalized (total capital ratio $<$ 8 percent, tier 1 ratio $<$ 4 percent, or tier 1 leverage ratio $<$ 4 percent).

[4]Alternative measures of risk taking such as value-at-risk (VaR) or the volatility of the market price of bank assets were not available for the period considered.

known that RWA/A is a very crude measure of credit risk and that the four risk categories specified by the 1988 Basel Accord only imperfectly reflect the actual credit risk taking of banks (Jones 2000). One may therefore view RWA/A more as a measure of portfolio composition (regulatory risk) than of "true" credit risk (economic risk). The latter interpretation is independent of whether RWA/A is a correct measure of credit risk.

3. Data

The variables used in this study are obtained from Bankscope. The sample consists of an unbalanced panel containing yearly data on 576 G-10 commercial banks (but no holding companies) with assets of more than $100 million. Consistent with most studies on the impact of the 1988 capital standards, the sample is restricted to the 1988–95 period.[5]

The analysis is further restricted to six G-10 countries (Canada, France, Italy, Japan, the United Kingdom, and the United States) because capital adequacy data were not available for the other G-10 countries over the period of interest (although data were available for Sweden, this country was excluded from the sample because of the banking crisis it experienced in the early 1990s). In addition, banks with a total capital ratio above 50 percent or a credit-risk ratio above 200 percent were treated as outliers and excluded from the sample.

Table 1 shows the distribution of banks by country. Although the sample contains mostly banks located in the United States and Japan, it is also representative of the banking sector in the other four countries. Indeed, the sample always includes at least six of the ten biggest banks in terms of assets of each country, and the sample banks' assets always exceed half of the total banking assets of each country.

[5]Data on capital adequacy are not available for years prior to 1988, preventing any comparison with the pre-Basel period. The choice of 1995 is somewhat arbitrary but quite standard given that most studies on the impact of the Basel guidelines focus on the first half of the 1990s. In the case of the United States, Flannery and Rangan (2002) show that none of the 100 largest banks appear to have been constrained by regulatory capital requirements since 1995.

Table 1. Representativeness of the Sample

Country	Number of Banks	Number of Banks from the National Top-Ten	Sample Bank Assets/Total National Banking Assets (%)
Canada	7	7	92.19
France	9	7	54.18
Italy	16	10	86.06
Japan	76	9	83.98
United Kingdom	9	6	69.66
United States	459	10	91.74

Note: The figures in the table are for year-end 1995. The whole sample consists of 576 commercial banks with assets of more than $100 million. The analysis is restricted to six G-10 countries because data on capital adequacy were not available for other G-10 countries over the period of interest. Sweden was excluded from the sample because of the banking crisis it experienced in the early 1990s.

Panels A–C of table 2 show the average total capital-to-assets ratio, tier 1 capital-to-assets ratio, and credit-risk ratio of banks in each country over the period surveyed. Figures are slightly difficult to compare, as the number of observations is increasing over time.[6] Nevertheless, some tentative remarks can be made. First, looking at panels A and B, the total capital-to-assets and tier 1 capital-to-assets ratios of banks are upward trending in each country over the period surveyed, except in Canada. Second, looking at panel C, some countries (Canada, France, the United Kingdom, and perhaps Italy) appear to have experienced a decrease in credit risk, whereas others (Japan and the United States) have seen credit risk remaining fairly constant.

The remainder of table 2 and table 3 report additional descriptive statistics on the relation between capital and credit risk. Panels D and E of table 2 show the total capital and tier 1 ratios of banks over the period surveyed. Both series are increasing in each country

[6]The Basel standards were implemented gradually, which explains the low number of observations in 1988 and 1989. Results of logit regressions (not reported here) indicate that banks with high capital-to-assets ratios were not more likely to join the sample between 1989 and 1995.

Table 2. Summary Statistics (Capital-to-Assets Ratios, Credit-Risk Ratio, and Capital Adequacy Ratios)

Country	1988		1989		1990		1991		1992		1993		1994		1995	
	Avg.	Obs.	Avg.	Obs.	Avg.	Obs.	Avg.	Obs.	Avg.	Obs.	Avg.	Obs.	Avg.	Obs.	Avg.	Obs.
Panel A: Total Capital-to-Assets Ratio																
Canada	6.94	6	7.02	7	7.03	7	7.38	7	7.19	7	7.35	7	7.11	7	6.05	7
France	4.27	1	3.93	5	4.24	8	4.58	9	4.76	9	5.03	9	5.25	7	5.10	7
Italy	—	—	5.65	1	5.79	2	5.79	6	5.63	10	5.58	14	6.28	16	6.36	14
Japan	—	—	—	—	6.58	5	6.08	11	7.02	11	7.12	11	6.60	11	6.65	11
United Kingdom	8.90	1	8.01	5	8.70	6	8.73	7	7.98	8	8.43	8	8.43	9	8.20	9
United States	—	—	8.54	1	7.50	156	7.74	153	8.35	450	8.85	435	8.78	422	9.27	392
All Countries	6.85	8	6.48	19	7.33	184	7.46	193	8.18	495	8.62	484	8.56	472	8.97	440
Panel B: Tier 1 Capital-to-Assets Ratio																
Canada	4.74	6	4.88	7	4.91	7	5.09	7	4.96	7	4.96	7	4.84	7	4.77	7
France	2.29	1	2.72	4	2.51	8	2.62	9	2.86	9	2.94	9	3.04	7	3.06	7
Italy	—	—	4.22	1	4.93	2	4.70	6	4.48	9	4.40	14	4.97	16	5.18	13
Japan	3.25	37	3.62	51	3.78	74	3.84	75	4.08	76	4.13	76	4.20	76	4.11	72
United Kingdom	4.89	1	4.69	5	5.11	6	5.31	7	4.59	7	5.17	8	5.42	9	5.32	9
United States	—	—	5.07	1	5.54	3	6.33	153	7.15	450	7.58	435	7.41	422	7.82	392
All Countries	3.46	45	3.80	69	3.91	100	5.37	257	6.56	558	6.87	549	6.76	537	7.07	500
Panel C: Credit-Risk Ratio																
Canada	97.65	6	93.27	7	89.40	7	83.59	7	79.78	7	74.58	7	71.64	7	68.37	7
France	71.70	1	55.59	5	58.57	8	58.82	9	58.64	9	57.16	9	54.59	7	53.06	7
Italy	—	—	58.08	1	55.89	2	57.45	6	60.13	9	55.48	14	54.60	16	57.48	13
Japan	66.78	37	68.45	51	67.81	74	68.23	75	69.78	76	68.72	76	68.36	76	69.95	72
United Kingdom	75.17	1	79.33	4	76.30	6	74.02	7	65.62	8	65.00	8	62.72	9	61.57	9
United States	—	—	92.12	1	85.80	3	79.14	152	69.92	415	69.70	422	70.88	420	72.87	390
All Countries	71.19	45	70.86	69	69.39	100	74.71	256	69.61	524	68.97	536	69.70	535	71.21	498

(continued)

Table 2. (Continued)

Country	1988		1989		1990		1991		1992		1993		1994		1995	
	Avg.	**Obs.**	**Avg.**	**Obs.**	**Avg.**	**Obs.**	**Avg.**	**Obs.**	**Avg.**	**Obs.**	**Avg.**	**Obs.**	**Avg.**	**Obs.**	**Avg.**	**Obs.**
Panel D: Total Capital Ratio																
Canada	7.14	6	7.59	7	7.89	7	8.86	7	9.01	7	9.84	7	9.94	7	9.91	7
France	6.00	1	7.03	7	7.34	9	7.72	9	8.26	9	8.80	9	9.65	9	9.41	9
Italy	10.02	1	8.84	3	8.97	6	9.75	8	9.32	11	10.15	14	11.69	16	11.33	16
Japan	9.24	37	8.61	51	8.93	74	8.49	75	9.23	76	9.60	76	9.24	76	9.58	72
United Kingdom	10.95	2	9.91	6	11.31	7	12.16	7	12.22	9	12.59	9	13.71	9	13.55	9
United States	—	—	9.30	1	8.90	3	10.09	152	12.30	415	13.05	422	12.67	421	13.04	392
All Countries	8.99	47	8.49	75	8.88	106	9.55	258	11.68	527	12.36	537	12.09	538	12.39	505
Panel E: Tier 1 Ratio																
Canada	4.84	6	5.33	7	5.54	7	6.07	7	6.23	7	6.67	7	6.76	7	6.97	7
France	3.20	1	4.34	5	4.41	8	4.58	8	5.11	8	5.41	8	5.95	8	5.89	8
Italy	8.88	1	7.31	3	7.76	4	7.85	8	7.37	10	8.00	14	9.28	16	8.96	14
Japan	4.91	37	5.30	51	5.61	74	5.66	75	5.89	76	6.07	76	6.21	76	6.12	72
United Kingdom	6.05	2	5.72	6	6.68	7	7.47	7	7.14	8	7.72	9	8.85	9	8.89	9
United States	—	—	5.50	1	6.47	3	8.31	152	10.58	415	11.28	422	10.82	421	11.17	392
All Countries	5.00	47	5.36	73	5.69	103	7.32	257	9.64	524	10.24	536	9.97	537	10.20	502

Note: The whole sample consists of 576 commercial banks with assets of more than $100 million. The analysis is restricted to six G-10 countries because data on capital adequacy were not available for other G-10 countries over the period of interest. Sweden was excluded from the sample because of the banking crisis it experienced in the early 1990s. The Basel standards were implemented gradually, which explains the low number of observations in 1988 and 1989. Results of logit regressions (not reported here) indicate that banks with high capital-to-assets ratios were not more likely to join the sample between 1989 and 1995.

Table 3. Decomposition of the Average Annual Growth Rate of CAR (%), 1988–95

Country	$\frac{\dot{CAR}}{CAR}$	$\frac{\dot{K}}{K}$	$\frac{\dot{RISK}}{RISK}$	$\frac{\dot{A}}{A}$	Obs.
CAR = Total Capital Ratio					
Canada	4.56	8.66	−4.23	8.33	48
France	3.62	6.66	−2.08	5.12	44
Italy	−1.60	5.85	0.61	3.65	45
Japan	1.55	5.51	0.32	3.64	49
United Kingdom	3.16	5.09	−4.04	5.98	43
United States	3.33	11.43	0.83	7.27	1,348
All Countries	3.10	10.35	0.44	6.81	1,577
CAR = Tier 1 Capital Ratio					
Canada	4.56	8.66	−4.23	8.33	48
France	5.25	8.36	−2.31	5.42	39
Italy	−0.03	4.24	0.61	3.65	45
Japan	1.80	9.20	0.23	7.17	436
United Kingdom	3.98	5.98	−3.91	5.90	42
United States	2.96	11.06	0.83	7.27	1,348
All Countries	2.57	9.96	0.40	6.98	1,958

Note: This table decomposes the annual growth rate of the capital adequacy ratio (CAR) into three terms: the annual growth rate of capital (K), the annual growth rate of the credit-risk ratio (RISK), and the annual growth rate of total assets (A). The dots denote time derivatives. A proof is given in the appendix.

across the years 1988–93, with no significant increase afterward. On average, G-10 banks already met the minimum requirements of 8 percent for the total capital ratio and 4 percent for the tier 1 ratio as early as 1989, except in Canada and France.

Table 3 further decomposes the average annual growth rate of both capital adequacy ratios into three terms, as in equation (1). The growth rate of both ratios is roughly similar and is mainly driven by a rise in capital levels (Italy, Japan, and the United States) or by a rise in capital levels and a decrease in risk-weighted assets (Canada, France, and the United Kingdom), which offset the rise in total assets.

On the whole, however, tables 2 and 3 do not tell us whether changes in the capital-to-assets ratio and changes in the credit-risk ratio of banks were related, nor whether the increase in capital-to-assets ratios that took place between 1988 and 1995 was due to the introduction of capital adequacy rules.[7] Determining whether the Basel standards caused changes in the capital-to-assets and credit-risk ratios of banks and whether these changes were related requires a more sophisticated analysis than just looking at descriptive statistics.

The following section presents a model that aims at assessing the empirical determinants of observed changes in the capital-to-assets and credit-risk ratios, with a particular emphasis on the role played by regulatory and market pressures.

4. Methodology

4.1 The Model

In order to acknowledge that capital and risk decisions are determined together, I use the simultaneous-equations model developed by Shrieves and Dahl (1992). In this model, observed changes in banks' capital and credit risk taking consist of two components—a discretionary adjustment and a change caused by factors exogenous to the bank:[8]

$$\Delta \mathrm{CAP}_{i,t} = \Delta^{\mathrm{d}}\mathrm{CAP}_{i,t} + E_{i,t}, \tag{2}$$

$$\Delta \mathrm{RISK}_{i,t} = \Delta^{\mathrm{d}}\mathrm{RISK}_{i,t} + S_{i,t}, \tag{3}$$

where $\Delta \mathrm{CAP}_{i,t}$ and $\Delta \mathrm{RISK}_{i,t}$ are the observed changes in capital and risk levels, respectively, for bank i in period t. The $\Delta^{\mathrm{d}}\mathrm{CAP}_{i,t}$ and

[7] For instance, in the case of the United States, an alternative explanation for the capital build-up observed in table 2 may be that banks were recapitalizing following the 1990–91 recession. The regression analysis in section 4 therefore controls for the state of the business cycle in each country via the lagged rate of GDP growth.

[8] The model analyzes the relation between changes in capital and changes in risk rather than the relation between capital and risk levels because the objective of this study is to understand how banks adjust their risk to changes in capital, and vice-versa.

$\Delta^{\mathrm{d}}\mathrm{RISK}_{i,t}$ variables represent discretionary adjustments in capital and risk, while $E_{i,t}$ and $S_{i,t}$ are random error terms.

Following Shrieves and Dahl (1992), I model the discretionary changes in capital and risk using a partial-adjustment framework such that

$$\Delta^{\mathrm{d}}\mathrm{CAP}_{i,t} = \alpha\big(\mathrm{CAP}^{*}_{i,t} - \mathrm{CAP}_{i,t-1}\big), \tag{4}$$

$$\Delta^{\mathrm{d}}\mathrm{RISK}_{i,t} = \beta\big(\mathrm{RISK}^{*}_{i,t} - \mathrm{RISK}_{i,t-1}\big), \tag{5}$$

where $\mathrm{CAP}^{*}_{i,t}$ and $\mathrm{RISK}^{*}_{i,t}$ are bank i's target capital and risk levels, respectively. Thus, the discretionary changes in capital and risk for bank i are proportional to the difference between the target level in period t and the observed level in period $t-1$.

Substituting equations (4) and (5) into equations (2) and (3), the changes in capital and risk can be written as

$$\Delta\mathrm{CAP}_{i,t} = \alpha\big(\mathrm{CAP}^{*}_{i,t} - \mathrm{CAP}_{i,t-1}\big) + E_{i,t}, \tag{6}$$

$$\Delta\mathrm{RISK}_{i,t} = \beta\big(\mathrm{RISK}^{*}_{i,t} - \mathrm{RISK}_{i,t-1}\big) + S_{i,t}. \tag{7}$$

This means that observed changes in capital and risk are a function of the target capital and risk levels, the lagged capital and risk levels, and any random shocks. As mentioned earlier, capital (CAP) is defined as the capital-to-assets ratio (K/A)—either the total capital-to-assets ratio or the tier 1 capital-to-assets ratio—while risk (RISK) is defined as the credit-risk ratio (RWA/A).

4.2 Variables Affecting Changes in Banks' Capital and Risk

Although the target capital and risk levels of banks are not observable, they are assumed to depend on a set of observable variables describing the banks' financial condition and the state of the economy in each country. In this paper, the variables used to approximate the target capital-to-assets ratio (CAP^{*}) are the size of the bank (SIZE), a measure of its liquidity (LOANS), a measure of its asset quality (LLOSS), a measure of its profitability (ROA), the rate of GDP growth (GROWTH), regulatory pressure interacted with market pressure (REG $\times$ MARKET), regulatory pressure interacted with the inverse of market pressure (REG $\times$ (1$-$MARKET)), changes in the credit-risk ratio (ΔRISK), changes in the credit-risk ratio interacted with regulatory pressure (ΔRISK × REG), and year dummies (YEAR). The variables used to proxy the target credit-risk ratio (RISK*) are SIZE, LOANS, LLOSS, GROWTH, REG × MARKET, REG × (1−MARKET), changes in the capital-to-assets ratio (ΔCAP), changes in the capital-to-assets ratio interacted with regulatory pressure (ΔCAP × REG), and YEAR. Table 4 gives the definition of each variable and shows summary statistics for three subsamples: European and Canadian banks, U.S. banks, and Japanese banks.

4.2.1 Bank-Specific Variables

Bank size (SIZE) is measured as the natural log of total assets. It is included as a control variable because large banks have easier access to equity capital markets and are thus expected to have lower capital-to-assets ratios than smaller banks. In addition, large banks carry out a wider range of activities, which should increase their ability to diversify their portfolio and, hence, decrease their credit risk. The percentage of total assets tied up in loans (LOANS) is included both in the capital and in the risk equations because higher LOANS values correspond to higher investment in risk-weighted assets and should therefore lead to higher credit risk and a greater need for capital. Following Rime (2001), loan losses (LLOSS) are approximated with the ratio of provisions to total assets and are included in the system of equations with an expected negative effect on credit risk and capital. Indeed, loan losses affect risk, as they are deducted from outstanding loans and should therefore lead to a decrease in the ratio of risk-weighted assets to total assets. In addition, banks with higher loan losses are forced to make higher provisions, thereby reducing net earnings and, ultimately, capital. The return on assets (ROA) is included in the capital equation with an expected positive effect on capital, as banks may prefer to increase capital through retained earnings rather than through equity issues in the presence of asymmetric information in capital markets.

The regulatory pressure variable (REG) describes the behavior of banks close to or below the Basel minimum capital requirements. These banks are expected to have increased their regulatory

Table 4. Summary Statistics (All Variables)

Variable	European and Canadian Banks			U.S. Banks			Japanese Banks		
	Avg.	**SD**	**Obs.**	**Avg.**	**SD**	**Obs.**	**Avg.**	**SD**	**Obs.**
1990 Dummy	0.09	0.29	180	0.00	0.03	1,348	0.11	0.32	436
1991 Dummy	0.12	0.33	180	0.00	0.05	1,348	0.16	0.36	436
1992 Dummy	0.16	0.37	180	0.11	0.32	1,348	0.17	0.38	436
1993 Dummy	0.18	0.39	180	0.29	0.46	1,348	0.17	0.38	436
1994 Dummy	0.20	0.40	180	0.30	0.46	1,348	0.17	0.38	436
1995 Dummy	0.20	0.40	180	0.29	0.45	1,348	0.17	0.37	436
SIZE_{t-1}	10.63	1.34	180	8.09	1.15	1,348	10.38	1.07	436
LOANS_{t-1}	56.46	15.52	180	60.15	14.71	1,348	65.86	6.49	436
LLOSS_{t-1}	0.66	0.51	180	0.63	1.21	1,348	0.13	0.20	436
ROA_{t-1}	0.39	0.51	180	1.05	1.02	1,348	0.19	0.08	436
GROWTH_{t-1}	0.70	2.20	180	1.95	1.60	1,348	2.29	2.27	436
$\Delta\mathrm{RISK}_{t}$	−1.83	5.10	180	0.55	6.08	1,348	0.17	1.91	436
RISK_{t-1}	67.61	15.56	180	71.15	15.89	1,348	68.45	6.48	436
CAP = Total Capital-to-Assets Ratio									
$\mathrm{REG}_{t-1} \times (1\text{–MARKET})$	0.26	0.44	180	0.07	0.26	1,348	0.10	0.31	49
$\mathrm{REG}_{t-1} \times \mathrm{MARKET}$	0.39	0.49	180	0.06	0.24	1,348	0.84	0.37	49
$\Delta\mathrm{CAP}_{t}$	0.00	0.90	180	0.38	1.40	1,348	0.06	0.94	49
CAP_{t-1}	6.55	2.07	180	8.58	2.11	1,348	6.69	0.92	49

(continued)

Table 4. (Continued)

Variable	European and Canadian Banks			U.S. Banks			Japanese Banks		
	Avg.	**SD**	**Obs.**	**Avg.**	**SD**	**Obs.**	**Avg.**	**SD**	**Obs.**
CAP = Tier 1 Capital-to-Assets Ratio									
$REG_{t-1} \times$ (1–MARKET)	0.18	0.38	174	0.03	0.16	1,348	0.14	0.35	436
$REG_{t-1} \times$ MARKET	0.28	0.45	174	0.02	0.13	1,348	0.50	0.50	436
ΔCAP_t	0.04	0.63	174	0.29	1.30	1,348	0.08	0.23	436
CAP_{t-1}	4.44	1.62	174	7.24	1.81	1,348	3.93	0.65	436

Note: The variables in the table are year dummy variables, SIZE (log of total assets in millions of dollars), LOANS (loans/total assets), LLOSS (loan loss provisions/total assets), ROA (net income/total assets), GROWTH (GDP growth rate), RISK (risk-weighted assets/total assets), CAP (total capital/total assets or tier 1 capital/total assets), REG (regulatory pressure), and MARKET (market pressure). REG is a dummy variable equal to one if the total capital adequacy ratio falls below 10 percent (regressions with CAP = total capital-to-assets ratio) or if the tier 1 capital adequacy ratio falls below 6 percent (regressions with CAP = tier 1 capital-to-assets ratio), and zero otherwise. MARKET is a dummy variable equal to one if banks had a credit rating from Moody's or S&P or were listed on a stock exchange between 1988 and 1995, and zero otherwise. All variables are in percent except year dummies, SIZE, REG, and MARKET. Statistics include average (Avg.), standard deviation (SD), and number of observations (Obs.) of each variable.

capital and/or decreased their risk-weighted assets more than well-capitalized banks because not meeting the Basel standards could trigger exclusion from international banking business.

The studies mentioned in section 1 generally measure regulatory pressure by a dummy variable equal to one if the capital adequacy ratio falls below the regulatory minimum (4 percent for tier 1 ratio and 8 percent for the total capital ratio) plus one standard deviation of the bank's capital adequacy ratio series, and zero otherwise. The rationale for this definition of regulatory pressure is that the regulatory minimum capital constraint was not binding for a majority of G-10 banks at the beginning of the 1990s (cf. table 2). At the same time, it seems reasonable to assume that the size of a bank's capital buffer partially depends on the volatility of its capital adequacy ratio.[9]

This definition of regulatory pressure is not used here because the data are unbalanced and, hence, computing the standard deviation of the capital adequacy ratio would require using a different number of observations for each bank, which does not make sense. In addition, this definition implies that regulatory pressure is influenced by bank behavior and, as a result, is endogenous.

For these reasons, I rely on a much simpler definition of regulatory pressure: banks are under regulatory pressure if their total capital ratio falls below 10 percent (regressions with CAP = total capital-to-assets ratio) or if their tier 1 ratio falls below 6 percent (regressions with CAP = tier 1 capital-to-assets ratio). These thresholds, which are similar to those imposed by FDICIA on U.S. banks to be recognized as well capitalized, produce sensible percentages of observations with REG equal to one in each subsample (see table 4).[10]

The regulatory pressure variable is nevertheless difficult to interpret, as the behavior of banks for which REG is equal to one is

[9]See Bauman and Nier (2003) and Lindquist (2004) for an investigation of the determinants of banks' capital buffers in the United Kingdom and in Norway, respectively.

[10]Robustness checks (not reported here) show that the results are not affected by the choice of alternative thresholds for the tier 1 ratio (5 percent or 7 percent) and for the total capital ratio (9 percent or 11 percent). In the case of U.S. banks, the regulatory pressure variable also includes the 4 percent tier 1 leverage requirement set by FDICIA for banks to be considered well capitalized.

likely to reflect not only regulatory pressure from prudential authorities but also pressure from market participants such as investors or credit-rating agencies (cf. section 1). In other words, it may be hard to disentangle the effects of regulatory pressure from increased market discipline when REG is used alone in the regressions.

For this reason, I introduce a market pressure variable (MARKET) in the analysis. This variable is equal to unity if banks had a credit rating from Moody's or S&P or were listed on a stock exchange over the period surveyed, and zero otherwise.[11] Since I am primarily interested in the impact of regulatory pressure, I interact REG with MARKET and with its inverse to create two new variables: REG $\times$ MARKET and REG $\times$ (1$-$MARKET). The former variable reflects the behavior of banks under both types of pressures, while the latter captures the behavior of banks under regulatory pressure but under no market pressure. Banks under no regulatory pressure act as a comparison group.

Finally, since previous sections indicate that banks' capital and credit-risk choices are interdependent, ΔCAP and ΔRISK are included on the right-hand side of equations (7) and (6), respectively. The sign of the relationship between both variables is not clear a priori. A positive and significant relation between ΔCAP and ΔRISK would be consistent with the unintended effects of more stringent bank capital requirements on credit risk (section 2.1) or with the fact that banks want to maintain their capital adequacy ratios (CAP/RISK) constant following a change in capital and credit risk. A negative and significant relation between ΔCAP and ΔRISK could indicate either an increase or a decrease in bank capital adequacy ratios, depending on which variable is increasing or decreasing and at what rate.

4.2.2 Country-Specific Variable

The rate of GDP growth (GROWTH) is included in the capital and the risk equations in order to take account of country-specific macroeconomic shocks—such as changes in the volume or in the structure

[11] Data on the ownership structure of banks could have been useful to refine the definition of market pressure but were not available for the period of interest. Note that market pressure does not show any significant correlation with bank size.

of loan demand—that may have affected banks' capital and credit-risk choices. There are reasons to believe that this variable may be significant, since several papers (e.g., Ayuso, Pérez, and Saurina 2004 and Jiménez and Saurina 2006) show that capital and credit risk tend to be driven by cyclical factors.

4.2.3 Year Dummy Variables

Year dummy variables (YEAR) are added to the specification in order to take account of common country shocks that may have affected banks' capital and credit-risk choices (e.g., end of the implementation period of the Basel Accord in 1992).

4.2.4 Specification and Estimation Technique

In order to avoid potential endogeneity problems, the variables selected to explain target capital and risk ratios are lagged once in the regressions. The model defined by equations (6) and (7) is thus written as follows:

$$\begin{aligned}\Delta\mathrm{CAP}_{i,t} = {} & a_0 + \sum_t a_{1t}\mathrm{YEAR}_t + a_2\mathrm{SIZE}_{i,t-1} + a_3\mathrm{LOANS}_{i,t-1}\\ & + a_4\mathrm{LLOSS}_{i,t-1} + a_5\mathrm{ROA}_{i,t-1}\\ & + a_6\mathrm{GROWTH}_{j,t-1} + a_7(\mathrm{REG}_{i,t-1} \times (1\text{–}\mathrm{MARKET}_i))\\ & + a_8(\mathrm{REG}_{i,t-1} \times \mathrm{MARKET}_i)\\ & + a_9\mathrm{CAP}_{i,t-1} + a_{10}\Delta\mathrm{RISK}_{i,t}\\ & + a_{11}(\Delta\mathrm{RISK}_{i,t} \times \mathrm{REG}_{i,t-1}) + E_{i,t}, \qquad (8)\end{aligned}$$

$$\begin{aligned}\Delta\mathrm{RISK}_{i,t} = {} & b_0 + \sum_t b_{1t}\mathrm{YEAR}_t + b_2\mathrm{SIZE}_{i,t-1} + b_3\mathrm{LOANS}_{i,t-1}\\ & + b_4\mathrm{LLOSS}_{i,t-1} + b_5\mathrm{GROWTH}_{j,t-1}\\ & + b_6(\mathrm{REG}_{i,t-1} \times (1\text{–}\mathrm{MARKET}_i))\\ & + b_7(\mathrm{REG}_{i,t-1} \times \mathrm{MARKET}_i) + b_8\mathrm{RISK}_{i,t-1}\\ & + b_9\Delta\mathrm{CAP}_{i,t} + b_{10}(\Delta\mathrm{CAP}_{i,t} \times \mathrm{REG}_{i,t-1}) + S_{i,t}, \qquad (9)\end{aligned}$$

where i is a bank index and t is a time index.

The system formed by equations (8) and (9) is estimated separately for three different subsamples of banks: European and Canadian banks, U.S. banks, and Japanese banks.[12] Regressions are run separately for the United States and Japan because these two countries have enough observations to allow estimation of the model at the country level. Canadian, French, Italian, and UK banks are included together in the estimated system of equations because table 2 shows that their capital-to-assets ratios and credit-risk ratio had relatively similar patterns (increasing for CAP and decreasing for RISK) between 1988 and 1995.[13]

For each subsample of banks, the system of equations is estimated by three-state least squares (3SLS) with bank fixed effects. The use of 3SLS is motivated by the fact that the right-hand side of each equation includes an endogenous variable that is the dependent variable from the other equation in the system. Bank fixed effects are added to the specification because Chow tests reject the null hypothesis of absence of bank fixed effects in each equation, while Hausman tests reject the null hypothesis of no correlation between the bank fixed effects and the explanatory variables.[14]

5. Results

5.1 Preliminary Results

Tables 5, 6, and 7 present the results for European and Canadian banks, U.S. banks, and Japanese banks, respectively. CAP is defined as the total capital-to-assets ratio in the first system of equations and as the tier 1 capital-to-assets ratio in the second system of equations of each table (in the case of Japanese banks, I only present results for the system of equations where CAP is equal to the tier 1

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[12] Country dummies were also added in equations (8) and (9) in the European and Canadian subsample. Their coefficient is not reported due to the estimation procedure chosen (fixed effects).

[13] Since the capital-to-assets ratios of Canadian banks slightly decreased after 1991, I also estimated the system of equations for European banks only as a robustness check. The results found for European banks only are qualitatively similar to those that include Canadian banks and that are reported in table 5.

[14] The studies mentioned in section 1 do not test for the presence of fixed or random effects and systematically rely on pooled 3SLS for estimation purposes.

capital-to-assets ratio because too few banks report a total capital-to-assets ratio). Before analyzing the role played by regulatory and market pressures, as well as the relation between changes in capital and changes in credit risk, I briefly discuss the sign of the most important control variables.

Consistent with previous studies (e.g., Jacques and Nigro 1997 and Aggarwal and Jacques 2001), I find that bank size (SIZE) has a negative effect on ΔCAP in table 6, a result which suggests that large U.S. banks have easier access to capital markets and can therefore operate with lower amounts of capital. In addition, SIZE has a significant and highly positive impact on ΔRISK in table 7, reflecting large Japanese banks' disengagement from high-quality borrowers (risk weight equal to 0 or 20 percent) and increased exposure to the real-estate sector (risk weight equal to 100 percent) in the late 1980s and the early 1990s. The impact of loans as a percentage of total assets (LOANS) on changes in capital and credit risk, though often statistically significant, is not economically significant (less than 0.1 percentage point). Loan losses (LLOSS) exhibit little significance except in table 5, where they have a negative impact on ΔRISK, and in table 7, where they have a negative impact on ΔCAP, as expected. The return on assets (ROA) has a significantly positive effect on changes in capital in tables 5 and 7, a result consistent with the hypothesis that banks with higher earnings can improve more easily their capital position.

Interestingly, the rate of GDP growth (GROWTH) appears to have a somewhat negative and significant impact on the capital adequacy ratio of non-U.S. banks but no impact on the capital adequacy ratio of U.S. banks. Indeed, a 1-percentage-point change in GDP growth has a negative though very small (-0.09 percentage point) impact on capital changes in table 5 (CAP = total capital ratio), no impact on capital and risk changes in table 6, and a negative effect on both variables in table 7, with the overall effect on capital adequacy ratios (CAP/RISK) being slightly negative.[15]

The results for non-U.S. banks tend to confirm those of Ayuso, Pérez, and Saurina (2004), who find a negative relation between the

[15]The overall effect on capital adequacy ratios (-0.13) is obtained by applying the point estimates for ΔCAP and ΔRISK on Japanese banks' average tier 1 capital-to-assets and credit-risk ratios, respectively.

Table 5. Determinants of Changes in Capital and Credit-Risk Ratios (European and Canadian Banks)

Independent Variables	CAP = Total Capital-to-Assets Ratio		CAP = Tier 1 Capital-to-Assets Ratio	
	ΔCAP	ΔRISK	ΔCAP	ΔRISK
Intercept	−9.739	0.253	−8.113	4.352
	(1.00)	(0.01)	(1.27)	(0.11)
1990 Dummy	−0.496	-4.849^{***}	0.032	-5.561^{***}
	(1.39)	(3.07)	(0.14)	(3.69)
1991 Dummy	−0.768	-10.617^{***}	−0.171	-10.379^{***}
	(1.42)	(5.02)	(0.49)	(5.01)
1992 Dummy	-1.263^{**}	-12.336^{***}	−0.434	-12.473^{***}
	(1.98)	(4.72)	(1.04)	(4.85)
1993 Dummy	−0.795	-13.519^{***}	−0.150	-14.073^{***}
	(1.39)	(5.77)	(0.39)	(6.05)
1994 Dummy	−0.904	-14.434^{***}	−0.121	-14.741^{***}
	(1.56)	(6.04)	(0.32)	(6.28)
1995 Dummy	-0.974^{*}	-13.375^{***}	−0.181	-13.588^{***}
	(1.78)	(5.88)	(0.51)	(6.02)
SIZE_{t-1}	1.433	3.420	0.945	3.294
	(1.47)	(0.86)	(1.49)	(0.84)
LOANS_{t-1}	0.012	0.247^{***}	0.014	0.103
	(0.66)	(2.81)	(1.15)	(1.17)
LLOSS_{t-1}	0.465^{*}	-2.151^{***}	0.065	-1.408^{**}
	(1.83)	(3.25)	(0.40)	(2.05)
ROA_{t-1}	1.240^{***}	—	0.634^{***}	—
	(4.40)		(3.34)	
GROWTH_{t-1}	-0.091^{*}	−0.112	−0.037	−0.205
	(1.92)	(0.51)	(1.18)	(0.99)
REG_{t-1} × (1–MARKET)	−0.085	−1.242	0.352	−2.008
	(0.31)	(1.06)	(1.50)	(1.27)
REG_{t-1} × MARKET	−0.386	−1.083	−0.065	−0.262
	(1.39)	(0.95)	(0.46)	(0.29)
CAP_{t-1}	-0.901^{***}	—	-0.644^{***}	—
	(6.69)		(5.98)	—
ΔRISK_t	−0.015	—	0.024	—
	(0.54)		(1.31)	
RISK_{t-1}	—	-0.568^{***}	—	-0.501^{***}
		(8.36)		(6.99)
ΔCAP_t	—	2.416^{***}	—	3.570^{***}
		(5.17)		(4.74)
Observations	180	180	174	174
R-squared	0.41	0.59	0.47	0.61

Note: The dependent variables in the first system of equations are CAP (total capital/total assets) and RISK (risk-weighted assets/total assets). The dependent variables in the second system of equations are CAP (tier 1 capital/total assets) and RISK (risk-weighted assets/total assets). Each system of equations is estimated by 3SLS with bank fixed effects. Absolute t-statistics are in parentheses; *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 6. Determinants of Changes in Capital and Credit-Risk Ratios (U.S. Banks)

Independent Variables	CAP = Total Capital-to-Assets Ratio		CAP = Tier 1 Capital-to-Assets Ratio	
	ΔCAP	ΔRISK	ΔCAP	ΔRISK
Intercept	15.034***	48.922***	13.494***	51.342***
	(6.69)	(6.13)	(6.30)	(6.45)
1990 Dummy	—	—	—	—
1991 Dummy	—	—	—	—
1992 Dummy	2.025	10.974	0.871	8.773
	(0.98)	(1.41)	(0.45)	(1.13)
1993 Dummy	1.408	2.383	1.616*	6.227*
	(1.45)	(0.65)	(1.73)	(1.67)
1994 Dummy	1.494*	5.366*	1.402*	8.572***
	(1.95)	(1.87)	(1.90)	(2.95)
1995 Dummy	1.627	2.750	2.104	8.309
	(1.03)	(0.46)	(1.39)	(1.38)
SIZE_{t-1}	−1.095***	−0.769	−1.036***	−1.013
	(4.45)	(0.87)	(4.41)	(1.15)
LOANS_{t-1}	−0.022**	−0.075**	−0.019**	−0.086**
	(2.32)	(2.25)	(2.16)	(2.57)
LLOSS_{t-1}	−0.085	−0.077	−0.069	−0.050
	(1.46)	(0.46)	(1.23)	(0.30)
ROA_{t-1}	−0.085	—	−0.083	—
	(1.27)		(1.26)	
GROWTH_{t-1}	0.250	2.332	−0.108	0.970
	(0.41)	(1.00)	(0.18)	(0.42)
REG_{t-1} ×	0.115	−0.621	0.134	−0.871
(1–MARKET)	(0.70)	(0.99)	(0.63)	(1.01)
REG_{t-1} ×	0.486***	1.128	0.344	6.125***
MARKET	(2.80)	(1.63)	(1.33)	(5.99)
CAP_{t-1}	−0.750***	—	−0.684***	—
	(21.25)		(19.60)	
ΔRISK_t	0.000	—	0.015	—
	(0.02)		(1.36)	
RISK_{t-1}	—	−0.654***	—	−0.663***
		(20.82)		(22.01)
ΔCAP_t	—	0.026	—	−0.059
		(0.15)		(0.30)
Observations	1,348	1,348	1,348	1,348
R-squared	0.36	0.47	0.33	0.48

Note: The dependent variables in the first system of equations are CAP (total capital/total assets) and RISK (risk-weighted assets/total assets). The dependent variables in the second system of equations are CAP (tier 1 capital/total assets) and RISK (risk-weighted assets/total assets). Each system of equations is estimated by 3SLS with bank fixed effects. Absolute *t*-statistics are in parentheses; *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 7. Determinants of Changes in Capital and Credit-Risk Ratios (Japanese Banks)

	CAP = Tier 1 Capital-to-Assets Ratio	
Independent Variables	**ΔCAP**	**ΔRISK**
Intercept	−6.331*	−54.581
	(1.90)	(1.44)
1990 Dummy	—	—
1991 Dummy	−0.063	−1.148*
	(1.02)	(1.78)
1992 Dummy	−0.204*	−1.656
	(1.74)	(1.26)
1993 Dummy	−0.676***	−7.117***
	(3.17)	(2.80)
1994 Dummy	−0.739***	−8.337***
	(2.94)	(2.76)
1995 Dummy	−0.907***	−9.595***
	(3.24)	(2.75)
$SIZE_{t-1}$	0.752**	8.823**
	(2.42)	(2.46)
$LOANS_{t-1}$	0.018**	0.064
	(2.47)	(0.77)
$LLOSS_{t-1}$	−0.280***	−0.206
	(4.14)	(0.28)
ROA_{t-1}	0.532**	—
	(2.47)	
$GROWTH_{t-1}$	−0.128***	−0.741**
	(4.01)	(2.26)
$REG_{t-1} \times$ (1–MARKET)	0.114*	−0.279
	(1.80)	(0.42)
$REG_{t-1} \times$ MARKET	−0.024	0.256
	(0.59)	(0.62)
CAP_{t-1}	−0.498***	—
	(7.92)	
$\Delta RISK_t$	0.016	—
	(1.02)	
$RISK_{t-1}$	—	−0.503***
		(6.29)
ΔCAP_t	—	−2.827**
		(2.15)
Observations	436	436
R-squared	0.51	0.34

Note: The dependent variables in the system of equations are CAP (tier 1 capital/total assets) and RISK (risk-weighted assets/total assets). The system of equations is estimated by 3SLS with bank fixed effects. Absolute t-statistics are in parentheses; *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

business cycle and the capital buffers of Spanish banks. The results for U.S. and non-U.S. banks do, however, contrast with those of Jiménez and Saurina (2006), who find a positive association between the business cycle and credit risk in Spain. A possible explanation, aside from the difference in sample, is that these authors use a different measure of credit risk (nonperforming loans) than the one employed in this paper (risk-weighted assets to total assets).

Finally, the parameter estimates on lagged capital and credit risk are negative and significant in each subsample, with values lying in the range $[-0.901, -0.501]$ in table 5, $[-0.750, -0.654]$ in table 6, and $[-0.503, -0.498]$ in table 7. These figures indicate that G-10 banks were adjusting their capital and credit-risk ratios very rapidly to desired levels in the first half of the 1990s.

5.2 Impact of Regulatory and Market Pressures on Changes in Capital and Credit Risk

Looking at banks under regulatory pressure, a distinction must again be made between U.S. and non-U.S. banks.

In the case of U.S. banks, regulatory pressure without market pressure (REG $\times$ (1$-$MARKET)) has no effect on capital-to-assets ratios, while the combination of both regulatory and market pressures (REG $\times$ MARKET) has a positive and significant impact on the total capital-to-assets ratio but no impact on the tier 1 capital-to-assets ratio. This result suggests that pressure from both regulators and market participants was effective in raising U.S. bank capital ratios in the early 1990s. This finding contrasts with existing studies on the impact of bank capital requirements in the United States (e.g., Aggarwal and Jacques 1997, 2001 and Jacques and Nigro 1997), which do not estimate the impact of market pressure and find that regulatory pressure alone had a positive and significant impact on U.S. bank capital ratios.[16] The results in table 6 suggest rather that it is the pressure exerted by both regulators and market participants which contributed to increasing the capital ratios of U.S. banks. The magnitude of the increase in banks' capital due

[16]Interestingly, and similar to these papers, I find that REG becomes significant in the capital equation when it is used alone, i.e., when it is not interacted with MARKET and with (1$-$MARKET).

to regulatory and market pressures (0.49 percentage points on an annual basis) is somewhat lower than the one attributed to regulatory pressure alone in the above-mentioned papers. The results for the system of equations where CAP is equal to the tier 1 capital-to-assets ratio further seem to suggest that U.S. banks under both regulatory and market pressures increased their credit risk taking in the early 1990s. However, the small percentage of observations for which REG and MARKET are both equal to one in this system of equations (2 percent, cf. table 4) more than probably reduces the reliability of this estimate.

In the case of non-U.S. banks (tables 5 and 7), regulatory pressure—either with or without market pressure—is insignificant at the 5 percent level both in the capital and in the risk equations. This result, which indicates that weakly capitalized banks located outside the United States did not significantly modify their capital-to-assets ratios and their ratio of risk-weighted assets to total assets more rapidly than well-capitalized banks, represents new evidence on the impact of the 1988 bank capital requirements.

The insignificance of regulatory pressure in the capital equations of table 5 is rather surprising given the widespread belief that weakly capitalized banks had a stronger capital response than well-capitalized banks in all countries following passage of the 1988 Basel standards (see, e.g., Basel Committee on Banking Supervision 1999). This result might be explained by the behavior of several Canadian and European banks that adjusted only slowly to the Basel standards and had their capital adequacy ratios on the edge or below the required minimum during most of the period studied.[17]

The results in table 7 are also interesting given that one of the goals of the 1988 Basel Accord was to create a level playing field by eliminating the funding-cost advantage enjoyed by Japanese banks, which operated with significantly lower capital ratios than their competitors (Wagster 1996). The fact that REG $\times$ (1$-$MARKET) has a small and only weakly significant impact and REG $\times$ MARKET has no impact in the capital equation suggests that the pressure

[17] Additional results (not reported here) show that when the sample is restricted to the 1988–93 or 1988–94 periods, regulatory pressure interacted with market pressure does have a negative and significant impact on ΔRISK when CAP is equal to the total capital-to-assets ratio.

exerted by regulators was not really effective in raising the tier 1 capital-to-assets ratio of weakly capitalized banks in Japan. This result is in line with Ito and Sasaki (1998) and Montgomery (2005), who show that undercapitalized Japanese banks tended to issue more subordinated debt (i.e., an increase in tier 2 capital but not in tier 1 capital) after the passage of the new capital adequacy rules. In addition, Montgomery (2005) also finds that banks with low capital ratios tended to shift their asset portfolio out of heavily weighted risky assets such as corporate bonds and into zero-weighted riskless assets such as government bonds. This effect is not observed here, probably because risk-weighted assets to total assets is a broader measure of credit risk than those used in that paper and, therefore, simultaneous changes may counterbalance one another.

The results so far indicate that U.S. banks experiencing both regulatory and market pressures increased their total capital-to-assets ratio faster than well-capitalized banks in the early 1990s. However, regulatory pressure—either with or without market pressure—was not effective in raising the capital-to-assets ratios of banks in the other G-10 countries analyzed here. Also, there is no strong evidence that weakly capitalized G-10 banks modified their credit risk taking over the period of interest.[18]

5.3 Relation between Changes in Capital and Credit Risk

The relation between changes in capital (ΔCAP) and changes in credit risk (ΔRISK) also appears to depend on the country or group of countries considered.

As shown in table 5 (European and Canadian banks), changes in capital and credit risk are positively and significantly related to each other in the ΔRISK equation of each system. Although this result

[18]The low within-variability of REG $\times$ MARKET and REG $\times$ (1$-$MARKET) suggests comparing the fixed effect estimates of both parameters with their pooled estimates. The latter show the same level of significance as the former, except for REG $\times$ (1$-$MARKET), which has a weakly positive impact on ΔCAP in the European and Canadian subsample when CAP is defined as the tier 1 capital-to-assets ratio.

is consistent with the unintended effects of higher capital requirements on credit risk mentioned in section 2.1, it does not imply a decrease in banks' capital adequacy ratios (CAP/RISK), as the response of ΔRISK to a 1-percentage-point increase in CAP (2.42 percentage points when CAP = total capital-to-assets ratio and 3.57 percentage points when CAP = tier 1 capital-to-assets ratio) is not large enough. The remainder of table 5 shows that changes in capital and credit risk are not significantly related to each other in the ΔCAP equation of each system, meaning that European and Canadian banks did not alter significantly their capital-to-assets ratios in reaction to changes in the composition of their ratio of risk-weighted assets to total assets.

In table 6 (U.S. banks), changes in capital and credit risk are not significantly related to each other, while in table 7 (Japanese banks), both changes are unrelated in the ΔCAP equation and are negatively and significantly related in the ΔRISK equation. In the latter case, an increase of 1 percentage point in the tier 1 capital-to-assets ratio of banks leads to a decrease of 2.83 percentage points in their credit-risk ratio, all other things being equal. This result indicates that Japanese banks improved their tier 1 ratio by simultaneously increasing their tier 1 capital-to-assets ratio and lowering their ratio of risk-weighted assets to total assets.

6. Conclusion

This paper analyzes adjustments in capital and risk-weighted assets at banks from six G-10 countries between 1988 and 1995 using the simultaneous-equations model developed by Shrieves and Dahl (1992). In particular, the paper tests whether weakly capitalized banks increased their capital or decreased their risk-weighted assets more rapidly than did well-capitalized banks.

The analysis distinguishes between changes in capital and credit risk brought about by regulatory and market pressures. This distinction is important, as little is known about the effectiveness of market pressure in complementing regulatory pressure in order to increase capital and/or decrease risk taking among weakly capitalized banks. It is also important in light of the fact that minimum capital requirements and market discipline constitute two of the three pillars of the New Basel Accord.

The results suggest that only in the United States did weakly capitalized banks increase their total capital ratio faster than well-capitalized banks; moreover, this increase appears to have been due to both regulatory and market pressures. In the other G-10 countries, little evidence is found that weakly capitalized banks increased their capital ratios at a faster rate than well-capitalized banks. Finally, no evidence is found that U.S. or non-U.S. weakly capitalized banks modified their ratio of risk-weighted assets to total assets differently from well-capitalized banks.

Taken as a whole, these results suggest that the effectiveness of the 1988 bank capital requirements to increase capital and/or reduce credit risk was rather limited outside the United States. The results also highlight the role of market discipline in influencing U.S. bank capital choices.

Appendix. Proof of Equation (1)

As $\mathrm{CAR} = \frac{\mathrm{K}}{\mathrm{RWA}}$ and $\mathrm{RISK} = \frac{\mathrm{RWA}}{\mathrm{A}}$, we have that $\mathrm{CAR} = \frac{\mathrm{K}}{\mathrm{RISK.A}}$.

Taking logs and differentiating with respect to time yields,

$$\frac{d\log \mathrm{CAR}}{dt} = \frac{d\log \mathrm{K}}{dt} - \left[\frac{d\log \mathrm{RISK}}{dt} + \frac{d\log \mathrm{A}}{dt}\right].$$

We obtain easily that $\frac{\dot{\mathrm{CAR}}}{\mathrm{CAR}} = \frac{\dot{\mathrm{K}}}{\mathrm{K}} - \frac{\dot{\mathrm{RISK}}}{\mathrm{RISK}} - \frac{\dot{\mathrm{A}}}{\mathrm{A}}$.

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Inflation Thresholds and Relative Price Variability: Evidence from U.S. Cities*

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The impact of inflation on relative price variability (RPV) is an important channel for real effects of inflation. With a view to the recent debate on the Federal Reserve's implicit lower and upper bounds of its inflation objective, we introduce a modified version of Hansen's panel threshold model to explore the inflation-RPV linkage in U.S. cities. We find two significant inflation thresholds and both positive and negative effects of inflation on RPV. The smallest effect of inflation on RPV is ensured if inflation is low but well above zero. If monetary policy aims at minimizing inflation's impact on relative prices, our estimates suggest that U.S. inflation should range between 1.8 percent and 2.8 percent.

JEL Codes: E31, C23.

1. Introduction

There is a growing consensus that inflation affects the economy through its impact on relative price variability (RPV). In theoretical models, the link between inflation and RPV is typically generated by menu costs or imperfect information about the price level.[1] In both cases, inflation increases RPV and, thus, distorts the information content of nominal prices. The resulting real effects of inflation

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[1]Given inflation, price adjustments or menu costs increase RPV by making it optimal for (heterogenous) firms to change their prices infrequently even if their real prices erode (see, e.g., Rotemberg 1983). In incomplete-information models introduced by Lucas (1973), noisy price information leads to misperceptions of relative price changes and inefficient supply.

play a predominant role for recent macroeconomics. In particular, in standard New Keynesian dynamic stochastic general equilibrium (DSGE) models, increased relative price variability is "the root of all evil" caused by inflation (see Green 2005, 132).

In line with these theoretical predictions, several studies have provided evidence in favor of a positive impact of inflation on RPV for various countries; see, e.g., Parsley (1996), Debelle and Lamont (1997), Aarstol (1999), Jaramillo (1999), Chang and Cheng (2000), Konieczny and Skrzypacz (2005), and Nautz and Scharff (2005). Yet, there are notable exceptions: Following Lastrapes (2006), the established relationship between U.S. inflation and RPV breaks down in the mid-eighties, while Reinsdorf (1994) found that the relation is even negative during the disinflationary early 1980s. In the same vein, Fielding and Mizen (2000) and Silver and Ioannidis (2001) show that RPV decreases in inflation for several European countries.

A common feature of empirical contributions on the inflation-RPV linkage is that they restrict the attention to *linear* relationships. However, the mixed evidence provided by the empirical literature suggests that the relationship between inflation and RPV is more complex. In particular, the marginal impact of inflation on RPV may differ for high- and low-inflation regimes. For example, Jaramillo (1999) finds that U.S. inflation's impact on RPV is stronger when it is below zero. Further evidence in favor of threshold effects of inflation on RPV is provided by Caglayan and Filiztekin (2003) for Turkey and Caraballo, Dabús, and Usabiaga (2006) for Spain and Argentina. In all these contributions, however, both the number and the location of inflation thresholds are not estimated but are imposed exogenously.

This paper sheds more light on the empirical relevance of inflation thresholds for RPV by applying a modified version of the panel threshold model introduced by Hansen (1999) to recent price data from U.S. cities. This enables us to estimate the number of inflation thresholds, the threshold levels, and the marginal impact of inflation on RPV in the various regimes. Although our sample focuses on the recent low-inflation period, the panel data provides us with a sufficient variation of inflation rates in a range that should be of particular interest for assessing the current low-inflation environment.

Threshold models nest the linear case, such that they can be viewed as a first, natural step to generalize the standard inflation-RPV equations. Of course, one may think of alternative nonlinear specifications—see, e.g., Fielding and Mizen (2008), who investigate the inflation-RPV linkage with nonparametric methods. Our particular interest in the empirical relevance of inflation thresholds is stirred by the recent discussion about the acceptable range of inflation. Although the Federal Reserve has never officially stated a target range of inflation, most analysts believe that the Federal Reserve has *implicit* upper and lower limits of its inflation objective; see, e.g., Thornton (2006). Threshold effects of inflation are not required for explaining the increasing role of inflation targets.[2] However, in view of the important role of the inflation-RPV linkage for the inflation transmission mechanism, the identification of inflation thresholds could provide useful information about the appropriate location and width of an inflation-targeting band.

The remainder of the paper is structured as follows. Section 2 introduces the data and presents results from a linear panel regression. Section 3 applies the threshold model to the inflation-RPV linkage, revealing regime-dependent effects of inflation. Section 4 summarizes our main results and offers some conclusions.

2. Inflation and RPV in U.S. Cities

2.1 The Data Set

Our empirical analysis uses price data of the eight major CPI subcategories published by the Bureau of Labor Statistics (BLS) for a panel of fourteen U.S. cities from January 1998 through August 2005.[3] As a consequence, our sample of yearly inflation rates starts in January 1999.

[2]Inflation targets may help anchor inflation expectations and increase the transparency and accountability of the central bank. Mishkin and Westelius (2006) show that the announcement of an inflation-targeting band can be interpreted as an inflation contract ameliorating the inflation bias of discretionary policy. Explicit inflation-targeting bands or critical values of inflation are used by many central banks, including the Bank of England and the European Central Bank, to facilitate the communication of monetary policy.

[3]We use the CPI-U index representing the expenditures by all urban consumers, which can be downloaded from http://www.bls.gov/cpi/home.htm. The

The frequency and timing of the CPI publication differ across cities: for eleven cities data are released every second month. Only for three cities (Chicago, Los Angeles, and New York) are price data available on a monthly basis. Since the estimation of Hansen's (1999) panel threshold model requires a balanced panel, we took only the data of every second month for these three cities. Specifically, we selected the observations of the odd months because this choice implied that the number of observations in our sample from odd and even months is exactly the same.[4] After these data adjustments, we are left with $40 \times 14 = 560$ observations of yearly inflation rates, a sufficient sample size for applying panel threshold models.

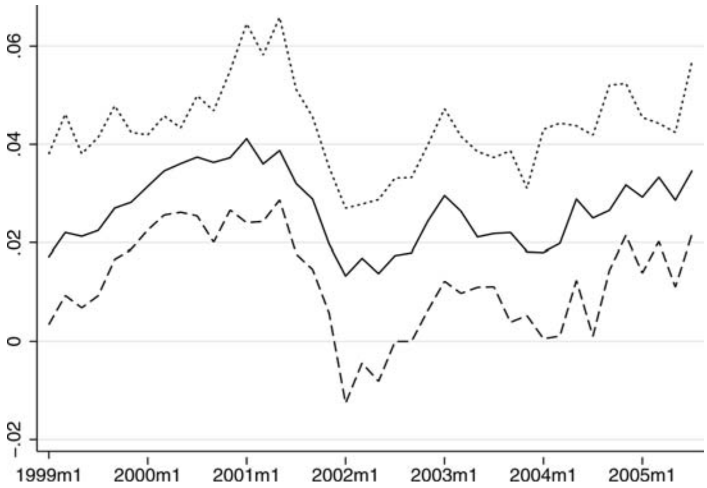
U.S. inflation has been low and stable over the last years. Since 1999 the average inflation rate across U.S. cities has fluctuated around 2.7 percent. Figure 1 further displays the minimum and the maximum of the city-specific inflation rates, indicating that inflation in U.S. cities exceeded 6 percent and even went below zero, at least for some cities in some periods. This illustrates that inflation differentials between U.S. cities have been modest but far from negligible. Typically, inflation rates varied in a range of 3 to 4 percentage points.[5] Figure 2 reveals more information about the distribution of city inflation rates from a timeless perspective. Note that our sample provides us with a sufficient variation of inflation rates. In particular, 25 percent of the observed inflation rates were below 1.88 percent or above 3.50 percent.

eight subcategories are food and beverages, housing, apparel, transportation, medical care, recreation, education and communication, and other goods and services. We selected January 1998 as a starting point for two reasons: (i) before then, data for Atlanta, Seattle, and Washington were only published twice a year and (ii) two new major groups were introduced in the CPI-U in January 1998.

[4]Note that the lack of synchronicity in the data is not a problem, because both the traditional linear-equation model and the threshold model will contain no lagged variables. Data for Atlanta, Detroit, Houston, Miami, Philadelphia, San Francisco, and Seattle are released only in even months; data for Boston, Cleveland, Dallas, and Washington are released only in odd months.

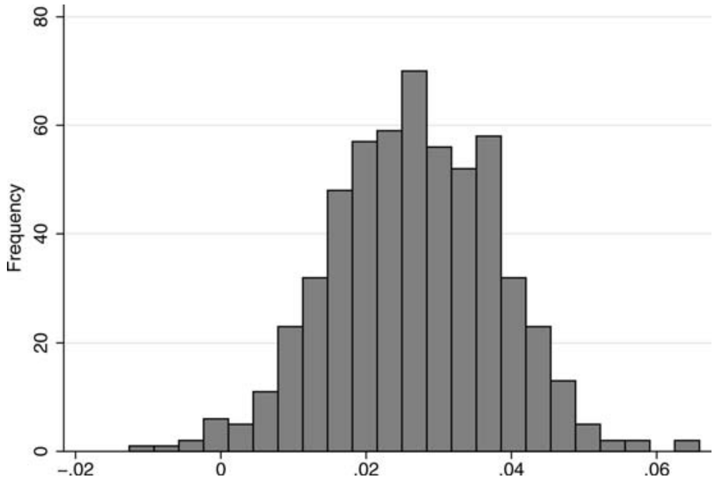
[5]The persistence of inflation differentials between U.S. cities has been relatively low. Therefore, inflation differentials between cities did not lead to significant price-level divergence.

Figure 1. Inflation Rates across U.S. Cities



Note: Minimum (dashed line), mean (solid line), and maximum (dotted line) of yearly CPI-U inflation rates of fourteen U.S. cities from January 1999 through August 2005. Source: Bureau of Labor Statistics.

Figure 2. Distribution of City Inflation Rates in the United States



Note: Yearly CPI-U inflation rates of fourteen U.S. cities from January 1999 through August 2005. Source: Bureau of Labor Statistics.

2.2 Relative Price Variability

Following the empirical literature, we define relative price variability (RPV_{it}) for city $i = 1, \ldots, 14$ in period $t = 1, \ldots, 40$ as

$$RPV_{it} = \sqrt{\sum_{j=1}^{8} w_j (\pi_{ijt} - \pi_{it})^2}, \qquad (1)$$

where $\pi_{ijt} = \ln P_{ijt} - \ln P_{ijt-6}$ is the yearly inflation rate for subcategory $j = 1, \ldots, 8$ and P_{ijt} is the level of the corresponding price index. $\pi_{it} = \sum_{j=1}^{8} w_j \pi_{ijt}$ denotes the inflation rate for city i, and w_j refers to the weight of the j-th subcategory in the aggregate index such that $\sum_{j=1}^{8} w_j = 1$.[6] Silver and Ioannidis (2001) introduce the coefficient of variation as an alternative measure of relative price variability. However, this RPV measure is not applicable in our sample because it includes inflation rates below zero.

2.3 The Linear Relation between Inflation and RPV

Following the empirical literature on inflation and RPV, we begin our analysis with a linear panel regression of RPV on aggregate inflation with city-specific fixed effects α_i:

$$RPV_{it} = \alpha_i + \beta \pi_{it} + \varepsilon_{it}. \qquad (2)$$

In line with Lastrapes (2006), the coefficient of inflation in the linear RPV equation (2) is clearly insignificant for recent U.S. data.[7] In a widely used alternative specification, RPV is regressed on $|\pi_{it}|$, the absolute value of inflation; see, e.g., Parks (1978) and Jaramillo (1999). In a low-inflation environment, where some inflation rates are actually below zero (compare figures 1 and 2), this could make a difference. For our data, however, the results presented in table 1

[6]In the following, the RPV measure will take into account that subcategory weights are adjusted on a yearly basis. The subcategory weights can also be downloaded from http://www.bls.gov/cpi/home.htm. They are only available as averages over all cities covered in the CPI.

[7]As a consequence, Lastrapes (2006) suggests including all individual prices in a linear VAR to estimate the cross-sectional distribution of impulse responses of these prices to, e.g., monetary shocks.

Table 1. The Linear Relation between Inflation and RPV

	$\hat{\boldsymbol{\beta}}$	$\boldsymbol{R^2}$
$RPV_{it} = \alpha_i + \beta\pi_{it} + \varepsilon_{it}$ (2)	-0.025 (0.03)	0.00
$RPV_{it} = \alpha_i + \beta\|\pi_{it}\| + \varepsilon_{it}$ (2a)	-0.016 (0.02)	0.00

Note: Standard errors are given in parentheses. Inflation and relative price variability (RPV) for fourteen U.S. cities; sample period is January 1999 through August 2005. Data source: Bureau of Labor Statistics.

(equation (2a)) reveal that this plausible nonlinearity in inflation's impact on RPV is not supported by the data.

3. Inflation Thresholds and the Inflation-RPV Linkage

3.1 The Threshold Model

In this section, we investigate whether the linear model (2) is misspecified because the marginal impact of inflation on RPV depends on the inflation level. In contrast to recent work by, e.g., Caglayan and Filiztekin (2003) for Turkish provinces and Caraballo, Dabús, and Usabiaga (2006) for Spain and Argentina, we do not impose the number and the locations of the different inflation regimes a priori. Rather, we employ a modified version of Hansen's (1999) panel threshold model that enables us to test for the number of thresholds and to estimate the threshold values—i.e., the critical inflation levels where the impact of inflation on RPV changes.

Specifically, we consider the following threshold model for the inflation-RPV linkage:

$$\begin{aligned} RPV_{it} = \alpha_i &+ \sum_{k=0}^{K-1} (\delta_{k+1} + \beta_{k+1}\pi_{it}) I(\gamma_k < \pi_{it} \leq \gamma_{k+1}) \\ &+ \beta_{K+1}\pi_{it} I(\gamma_K < \pi_{it} \leq \gamma_{K+1}) + \varepsilon_{it}, \end{aligned} \tag{3}$$

where $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$, and I is the indicator function. Equation (3) allows for K inflation thresholds and, thus, $K + 1$ regimes. In each regime, the marginal impact of inflation (β_k) on RPV may differ. Given the observed inflation differentials across cities, it is an additional feature of the panel threshold model that different cities are allowed to be in different inflation regimes.

It is worth noting that the panel threshold model (3) generalizes the original setup in Hansen (1999) by allowing for regime-dependent intercepts (δ_k). According to Bick (2007), ignoring intercepts can lead to biased estimates of both the thresholds and the corresponding marginal impacts.

3.2 The Number of Inflation Thresholds

In a first step, we applied Hansen's (1999) sequential testing procedure for determining the number of inflation thresholds. Following Hansen (1999), we require that each regime contains a minimum number of observations. Column 1 of table 2 shows the results obtained for the 5 percent rule predominantly applied in empirical applications of the threshold model. For our data set, the 5 percent rule implies that inflation thresholds may range from 0.81 percent to 4.44 percent. The results indicate a clear rejection of a linear relation ($K = 0$) between RPV and inflation in favor of a double-threshold model. Specifically, the null hypothesis of a single inflation threshold ($K = 1$) in the inflation-RPV equation can be rejected at the 1 percent significance level, while the hypothesis of a double threshold ($K = 2$) cannot be rejected at the 10 percent significance level.

This conclusion appears very robust with respect to different assumptions concerning the minimum number of observations in each regime. In particular, we found that the 5 percent constraint is not binding, implying that adopting the less restrictive 1 percent rule (where feasible inflation thresholds range from 0 percent to 5.24 percent) leads to identical results. This already indicates that the evidence in favor of a regime-dependent influence of inflation on RPV is not driven by a few outliers. We also performed the test adopting the unusually restrictive 10 percent rule. In this case, the range of feasible inflation thresholds shrinks to $[1.22\%, 4.12\%]$, which leads to slightly different values of the test statistics; see column 2 of table 2. Yet, the main result of the test remains unaffected. In particular, regardless of the minimum number of observations contained in each regime, table 2 strongly suggests that the inflation-RPV linkage is characterized by two inflation thresholds and, thus, three different inflation regimes.

Table 2. Test Procedure Establishing the Number of Thresholds

$$RPV_{it} = \alpha_i + \sum_{k=0}^{K-1}(\delta_{k+1} + \beta_{k+1}\pi_{it})I(\gamma_k < \pi_{it} \leq \gamma_{k+1}) + \beta_{K+1}\pi_{it}I(\gamma_K < \pi_{it} \leq \gamma_{K+1}) + \varepsilon_{it}$$

	5% Rule	**10% Rule**
No Threshold (H_0: $K = 0$)		
F_1	37.75	36.48
p-value	0.00	0.00
(10%, 5%, 1% critical values)	(11.48, 13.13, 16.17)	(11.16, 12.78, 16.30)
One Threshold (H_0: $K = 1$)		
F_2	18.61	18.31
p-value	0.00	0.01
(10%, 5%, 1% critical values)	(11.24, 12.53, 16.46)	(10.61, 12.48, 16.72)
Two Thresholds (H_0: $K = 2$)		
F_3	10.64	11.68
p-value	0.15	0.07
(10%, 5%, 1% critical values)	(11.48, 12.99, 15.94)	(10.82, 12.16, 14.78)

Note: $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$. The sequential test procedure indicates that the number of thresholds is $K = 2$. One thousand bootstrap replications were used to obtain the p-values. Following Hansen (1999), each regime is required to contain at least 5% or 10% of all observations.

3.3 A Double-Threshold Model for the Relation between Inflation and RPV

In view of the evidence in favor of two inflation thresholds, we estimated the following double-threshold model:

$$RPV_{it} = \alpha_i + (\delta_1 + \beta_1\pi_{it})I(\pi_{it} \leq \gamma_1) + (\delta_2 + \beta_2\pi_{it})I(\gamma_1 < \pi_{it} \leq \gamma_2) + \beta_3\pi_{it}I(\gamma_2 < \pi_{it}) + \varepsilon_{it}. \tag{4}$$

Table 3. A Double-Threshold Model for the Inflation-RPV Linkage

$RPV_{it} = \alpha_i + (\delta_1 + \beta_1\pi_{it})I(\pi_{it} \leq \gamma_1) + (\delta_2 + \beta_2\pi_{it}) \times I(\gamma_1 < \pi_{it} \leq \gamma_2) + \beta_3\pi_{it}I(\gamma_2 < \pi_{it}) + \varepsilon_{it}$		
	5% Rule	**10% Rule**
Threshold Estimates		
$\hat{\gamma}_1$	1.672	1.672
95% confidence interval	[1.586, 1.803]	[1.586, 1.820]
$\hat{\gamma}_2$	4.274	3.648
95% confidence interval	[2.852, 4.385]	[2.824, 4.102]
Regime-Dependent Inflation Coefficients:		
$\hat{\beta}_1$	−0.595**	−0.593**
	(0.13)	(0.13)
$\hat{\beta}_2$	−0.198**	−0.189**
	(0.05)	(0.07)
$\hat{\beta}_3$	0.548*	0.712**
	(0.23)	(0.14)
Regime-Dependent Intercepts:		
$\hat{\delta}_1$	0.028**	0.036**
	(0.01)	(0.01)
$\hat{\delta}_2$	0.026**	0.035**
	(0.02)	(0.01)
R^2	0.095	0.091
Observations in Regime 1	105	105
Observations in Regime 2	415	344
Observations in Regime 3	40	111

Note: ** and * indicate significance at the 1% and 5% level, respectively. Standard errors are in parentheses. Each regime consists of at least 5% and 10% of all observations.

Table 3 reports the estimates obtained under the 5 percent rule and the 10 percent rule. Results for the 1 percent rule are not presented since they are identical to those received for the 5 percent rule.

The upper panel of the table shows the results for the two inflation thresholds. The results for the lower inflation threshold are virtually unaffected by the applied rule. Both the point estimate for the threshold (1.672 percent) and the corresponding 95 percent confidence intervals are very similar. Note that the confidence interval does not contain 2 percent, probably the most popular number for inflation targets. The second inflation threshold estimated for the 5 percent rule (4.274 percent) exceeds the upper limit for feasible thresholds under the 10 percent rule (4.12 percent). As a consequence, the point estimate for the second threshold decreases under the 10 percent rule. Yet, the main conclusions about the threshold's location are very robust: according to the 95 percent confidence intervals, the upper threshold is clearly above 2.8 percent and certainly below 4.4 percent. Note that observations of all three regimes do not belong exclusively to a small subset of cities; see table 4. Therefore, the established nonlinearity of the inflation-RPV linkage does indeed result from a *regime*-dependent marginal impact of inflation and cannot be captured by city-specific inflation coefficients.

The estimates $(\hat{\beta}_1, \hat{\beta}_2, \hat{\beta}_3)$ for the marginal impact of inflation in the three inflation regimes are shown in the middle panel of table 3. In contrast to the results obtained for the linear specification, the threshold model reveals that inflation has a significant impact on RPV. However, both magnitude and sign of the inflation coefficient depend on the level of inflation. In the low-inflation regime—i.e., when inflation is below 1.672 percent—the marginal impact of inflation on RPV is significantly negative (–0.595). Thus, a further decline of inflation would increase RPV significantly. According to, e.g., Akerlof, Dickens, and Perry (1996) or Jaramillo (1999), this effect of inflation rates close to zero may point to the presence of nominal downward wage and price rigidities. In fact, in the intermediate-inflation regime, when inflation is low but well above zero, the impact of inflation on RPV is significantly weaker. In the high-inflation regime, the marginal impact of inflation is positive under both the 5 percent ($\widehat{\beta}_3 = 0.548$) and the 10 percent ($\widehat{\beta}_3 = 0.712$) specification. When inflation exceeds an upper threshold, it seems that RPV-increasing aspects of inflation (including, e.g., menu costs and imperfect information about the price level) become eventually dominant, while RPV-decreasing aspects of inflation have faded out.

Table 4. U.S. Cities and Inflation Regimes

	Low-Inflation Regime	Medium-Inflation Regime	High-Inflation Regime
Atlanta	14 (14)	25 (20)	1 (6)
Boston	2 (2)	28 (21)	10 (17)
Chicago	11 (11)	28 (27)	1 (2)
Cleveland	14 (14)	26 (23)	0 (3)
Dallas	9 (9)	27 (23)	4 (8)
Detroit	7 (7)	32 (29)	1 (4)
Houston	10 (10)	27 (19)	3 (11)
Los Angeles	0 (0)	39 (28)	1 (12)
Miami	8 (8)	30 (25)	2 (7)
New York	1 (1)	38 (34)	1 (5)
Philadelphia	5 (5)	30 (26)	5 (9)
San Francisco	14 (14)	16 (9)	10 (17)
Seattle	9 (9)	30 (25)	1 (6)
Washington	1 (1)	39 (35)	0 (4)
Total	105 (105)	415 (344)	40 (111)

Note: The table shows how often a city appears in the various inflation regimes estimated for the inflation-RPV linkage under the 5 percent rule; compare with table 3. The respective numbers under the 10 percent rule are given in parentheses.

4. Concluding Remarks

The impact of inflation on relative price variability is a major channel for real effects of inflation. This paper focused on the recent low-inflation period using price data from a panel of U.S. cities from 1998 through 2005. For this sample, we found that the common linear inflation-RPV equation has to be rejected in favor of a double-threshold model with surprisingly small 95 percent confidence intervals for both inflation thresholds. Partly reconciling the mixed evidence provided by the empirical literature, the estimated inflation coefficients reveal that there are both positive and negative effects of inflation on RPV. Inflation increases RPV only if it exceeds a critical value, which is estimated to range from about 2.8 percent to 4.4 percent. By contrast, inflation decreases RPV for inflation rates close to zero or, more precisely, below 1.67 percent. The weakest impact of inflation on RPV is found for the intermediate regime, when inflation is still low but well above zero.

Even central banks with a strong commitment to price stability are not really interested in zero inflation rates.[8] Typically, central banks prefer a more sophisticated notion of price stability. According to, e.g., Blinder et al. (1998, 98), "one prominent definition of 'price stability' is inflation so low that it ceases to be a factor in influencing people's decisions." Therefore, given the crucial importance of relative prices for economic decisions, an acceptable band of inflation rates should ensure the smallest impact of inflation on the variability of relative prices.

The recent literature on the importance of price rigidities revealed that there are notable differences in the frequency of price adjustments and implied durations between sectors. Golosov and Lucas (2007), and Klenow and Kryvtsov (2007) demonstrated that idiosyncratic shocks are an important factor for the price setting of firms. In accordance with Woodford (2003), the efficient level of RPV in an economy with multiple sectors is typically not zero, since it reflects relative price changes driven by fundamentals. Therefore, the optimal rate of inflation need not drive the level of RPV to zero. It is the marginal effect of inflation on RPV that has to be minimized. From this perspective, our empirical results may shed light on the location and width of an appropriate inflation-targeting band. In particular, the inflation thresholds in the inflation-RPV nexus suggest that U.S. inflation should range between 1.8 percent and 2.8 percent.

The repercussions of the introduction of an explicit targeting band on inflation's impact on relative prices are not obvious. The analysis of highly disaggregated price data indicates that the relation between inflation and the price setting of firms has been underresearched; see, e.g., Golosov and Lucas (2007). In particular, the Calvo and Taylor sticky-price models predominantly used in current New Keynesian DSGE models generate only poor predictions for the persistence and volatility of inflation; see Bils and Klenow

[8] Several arguments point to the difficulties implied by inflation rates too close to zero. For example, positive inflation rates may ameliorate problems caused by the zero bound for nominal interest rates; see, e.g., Adam and Billi (2006).

(2004). Recent evidence on the price setting of firms seems to support the relevance of inflation thresholds. According to Nakamura and Steinsson (2007), aggregate inflation plays no role in the frequency and the size of price changes during the low-inflation period 1998 through 2005. However, from 1988 to 1997, when average inflation exceeded 2.8 percent, the impact of aggregate inflation is significant and plausibly signed.

Generalizing the traditional linear inflation-RPV regressions, the current paper employed Hansen's (1999) panel threshold model to allow for a more complex relation between inflation and RPV. In the current model, the marginal impact of inflation jumps to a new value whenever inflation exceeds a threshold. Of course, allowing for a more gradual change of the inflation coefficient might lead to a more realistic view on the inflation-RPV linkage. Therefore, following Strickholm and Teräsvirta (2006), incorporating elements of smooth transition into a threshold model could be a natural extension and is left for future research.

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Incomplete Interest Rate Pass-Through and Optimal Monetary Policy*

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Many recent empirical studies have reported that the pass-through from money-market rates to retail lending rates is far from complete in the euro area. This paper formally shows that when only a fraction of all the loan rates is adjusted in response to a shift in the policy rate, fluctuations in the average loan rate lead to welfare costs. Accordingly, the central bank is required to stabilize the rate of change in the average loan rate in addition to inflation and output. It turns out that the requirement for loan rate stabilization justifies, to some extent, the idea of policy rate smoothing in the face of a productivity shock and/or a preference shock. However, a drastic policy reaction is needed in response to a shock that directly shifts retail loan rates, such as an unexpected shift in the loan rate premium.

JEL Codes: E44, E52, E58.

1. Introduction

Many empirical studies have shown that in the majority of industrialized countries, a cost channel plays an important role in the

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transmission of monetary policy.[1] Along with this, many authors have attempted to incorporate a cost channel in formal models of monetary policy. For example, Christiano, Eichenbaum, and Evans (2005) introduce a cost channel into the New Keynesian framework in accounting for the actual dynamics of inflation and output in the United States, while Ravenna and Walsh (2006) explore optimal monetary policy in the presence of a cost channel.

However, a huge number of recent studies have also reported that, especially in the euro area, shifts in money-market rates, including the policy rate, are not completely passed through to retail lending rates.[2] Naturally, since loan rates are determined by commercial banks, to what extent shifts in money-market rates affect loan rates and thereby the behavior of firms depends on how commercial banks react to the shifts in the money-market rates. If not all of the commercial banks promptly respond to a change in the money-market rates, then a policy shift will not affect the whole economy equally.[3] Given this situation, it is natural to ask whether or not the presence of loan rate sluggishness alters the desirable monetary policy compared with the case in which a shift in the policy rate is immediately followed by changes in retail lending rates. Nevertheless, to the best of my knowledge, little attention has been paid to such a normative issue since the main purpose of the previous studies was to estimate the degree of pass-through.

The principal aim of this paper is to formally explore optimal monetary policy in an economy with imperfect interest rate pass-through, where retail lending rates are allowed to differ across regions. Following Christiano and Eichenbaum (1992), Christiano, Eichenbaum, and Evans (2005), and Ravenna and Walsh (2006),

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[1] See, for example, Barth and Ramey (2001), Angeloni, Kashyap, and Mojon (2003), Christiano, Eichenbaum, and Evans (2005), Chowdhury, Hoffmann, and Schabert (2006), and Ravenna and Walsh (2006).

[2] Some recent studies, to name a few, are Mojon (2000), Weth (2002), Angeloni, Kashyap, and Mojon (2003), Gambacorta (2004), de Bondt, Mojon, and Valla (2005), Kok Sørensen and Werner (2006), and Gropp, Kok Sørensen, and Lichtenberger (2007). A brief review of the literature on interest rate pass-through is provided in the next section.

[3] Possible explanations for the existence of loan rate stickiness have been continuously discussed in the literature. Some of those explanations are introduced in the next section.

it is assumed in our model that the marginal cost of each production firm depends on a borrowing rate, since the owner of each firm needs to borrow funds from a commercial bank in order to compensate for wage bills that have to be paid in advance. A novel feature of our model is that there is only one commercial bank in each region, and each commercial bank does business only in the region where it is located. Since loan markets are assumed to be geographically segmented, each firm owner can borrow funds only from the corresponding regional bank. In this environment, retail loan rates are not necessarily the same across firms. The commercial banks' problem for loan rate determination is specified as Calvo-type pricing.

It is shown that the approximated utility function takes a form similar to the objective function that frequently appears in the literature on "interest rate smoothing." An important difference, however, is that the central bank is now required to stabilize the rate of change in the average loan rate, not the rate of change in the policy rate. The necessity for the stabilization of the average loan rate can be understood by analogy with the requirement for inflation stabilization, which has been widely discussed within the standard Calvo-type staggered-price model. Under staggered pricing, the rate of inflation should be stabilized because price dispersion would otherwise take place. Under staggered loan rates, changes in the average loan rate must be dampened because loan rate dispersion would otherwise take place. Since loan rate dispersion inevitably causes price dispersion through the cost channel, it consequently leads to an inefficient dispersion in hours worked.

It turns out that the introduction of a loan rate stabilization term in the central bank's loss function causes the optimal policy rate to become more inertial in the face of a productivity shock and a preference shock. This implies that the optimal policy based on a loss function with a loan rate stabilization term is quite consistent with that based on the conventionally used loss function that involves a policy rate stabilization term. Yet, this smoothing effect appears to be limited quantitatively.

On the other hand, the presence of a loan rate stabilization term requires a drastic policy response in the face of an exogenous shock that *directly* shifts retail loan rates, such as an unexpected change in the loan rate premium. For example, an immediate reduction in the policy rate is needed in response to a positive loan premium shock since it can partially offset the rise in loan rates. This is in stark contrast to the policy suggested by conventional policy rate smoothing. The case of a loan premium shock is an example for which it is crucial for the central bank to clearly distinguish between policy rate smoothing and loan rate smoothing.

The rest of the paper is organized as follows. The next section briefly reviews recent empirical studies on interest rate pass-through. Section 3 presents a baseline model, and section 4 summarizes the equilibrium dynamics of the economy. Section 5 derives a utility-based objective function of the central bank, and optimal monetary policy is explored in section 6. Section 7 concludes the paper.

2. A Review of Recent Studies on Interest Rate Pass-Through

Over the past decade, a huge number of empirical studies have been conducted in an attempt to estimate the degree of interest rate pass-through in the euro area. In the literature, the terminology "interest rate pass-through" generally has two meanings: loan rate pass-through and deposit rate pass-through. In this paper, we focus on the former since the general equilibrium model described below treats only the case of loan rate stickiness. Although it is said that deposit rates are also sticky in the euro area, constructing a formal general equilibrium model that includes loan rate stickiness is a reasonable first step to a richer model that could also take into account the sluggishness in deposit rates. This section briefly reviews recent studies on loan rate pass-through in the euro area.[4]

Although recent studies on loan rate pass-through differ in terms of the estimation methods and the data used, a certain amount of broad consensus has been established. First, at the euro-area aggregated level, the policy rate is only partially passed through to retail loan rates in the short run, while the estimates of the degree of

[4] de Bondt, Mojon, and Valla (2005) and Kok Sørensen and Werner (2006) also provide a survey of the literature on the empirical study of interest rate pass-through, including deposit rate pass-through.

pass-through differ among researchers. For example, according to table 1 of de Bondt, Mojon, and Valla (2005), the estimated degree of short-run (i.e., monthly) pass-through of changes in the market interest rates to the loan rate on short-term loans to firms varies from .25 (Sander and Kleimeier 2002; Hofmann 2003) to .76 (Heinemann and Schüler 2002). Gropp, Kok Sørensen, and Lichtenberger (2007) argued that interest rate pass-through in the euro area is incomplete even after controlling for differences in bank soundness, credit risk, and the slope of the yield curve. On the other hand, there is no general consensus about whether the long-run interest rate pass-through is perfect or not.[5]

Second, although the degree of interest rate pass-through significantly differs across countries, the extent of heterogeneity has been reduced since the introduction of the euro (de Bondt 2002; Toolsema, Sturm, and de Haan 2001; Sander and Kleimeier 2004). At this point, it also seems to be widely admitted that the speed of loan rate adjustment has, to some extent, been improved (de Bondt 2002; de Bondt, Mojon, and Valla 2005).

While there is little doubt about the existence of sluggishness in loan rates, there is still much debate as to why it exists and why the extent of pass-through differs across countries. For instance, Gropp, Kok Sørensen, and Lichtenberger (2007) insisted that the competitiveness of the financial market is a key to understanding the degree of pass-through. They showed that a larger degree of loan rate pass-through would be attained as financial markets become more competitive. Schwarzbauer (2006) pointed out that differences in financial structure, measured by the ratio of bank deposits to GDP and the ratio of market capitalization to GDP, have a significant influence on the heterogeneity among euro-area countries in the speed of pass-through. de Bondt, Mojon, and Valla (2005) argued that retail bank rates are not completely responsive to money-market rates since bank rates are tied to long-term market interest rates even in the case of short-term bank rates. From a different point

[5]For instance, Mojon (2000), Heinemann and Schüler (2002), Hofmann (2003), and Sander and Kleimeier (2004) reported that the long-run pass-through of market rates to interest rates on short-term loans to firms is complete. On the other hand, Donnay and Degryse (2001) and Toolsema, Sturm, and de Haan (2001) argued that the loan rate pass-through is incomplete even in the long run.

of view, Kleimeier and Sander (2006) emphasized the role of monetary policymaking by central banks as a determinant of the degree of pass-through. They argued that better-anticipated policy changes tend to result in a quicker response of retail interest rates.[6]

In the theoretical model presented in the next section, we consider a situation where financial markets are segmented and thus each regional bank has a monopolistic power. While the well-known Calvo-type staggered pricing is applied to banks' loan rate settings, it turns out that the degree of pass-through depends largely on the central bank's policy rate setting. Moreover, a newly charged loan rate can be interpreted as a weighted average of short- and long-term market rates, where the size of each weight is dependent on the degree of stickiness. Thus, although our way of introducing loan rate stickiness into the general equilibrium model is fairly simple, the model's implications for the relationship between loan rates and the policy rate seem quite consistent with what some of the previous studies have suggested.

3. The Model

The economy consists of a representative household, intermediate-goods firms, final-goods firms, financial intermediary, and the central bank. The representative household consumes a variety of final consumption goods while supplying labor service in the intermediate-goods sector. Each intermediate-goods firm produces a differentiated intermediate good and sells it to final-goods firms. Following Christiano and Eichenbaum (1992), Christiano, Eichenbaum, and Evans (2005), and Ravenna and Walsh (2006), we consider a situation in which the owner of each intermediate-goods firm has to pay wages in advance to workers at the beginning of each period. The owner thereby needs to borrow funds from a commercial bank since they cannot receive revenue until the end of the period. Final-goods firms produce differentiated consumption goods by using a composite of intermediate goods.

[6]For a more concrete discussion about the source of imperfect pass-through, see Gropp, Kok Sørensen, and Lichtenberger (2007). As for the heterogeneity in the degree of pass-through, see Kok Sørensen and Werner (2006).

3.1 Households

The one-period utility function of a representative household is given as

$$\begin{aligned} U_t &= u(C_t;\xi_t) - \int_0^1 v(L_t(i))di \\ &= \frac{(\xi_t C_t)^{1-\sigma}}{1-\sigma} - \int_0^1 \frac{L_t(i)^{1+\omega}}{1+\omega} di, \end{aligned}$$

where $C_t \equiv \left[\int_0^1 C_t(j)^{\frac{\theta-1}{\theta}} dj\right]^{\frac{\theta}{\theta-1}}$, and $C_t(j)$ and $L_t(i)$ are the consumption of differentiated good j and hours worked at intermediate-goods firm in region i, respectively. Henceforth, index i is used to denote a specific region as well as the variety of intermediate goods. Since there is only one intermediate-goods firm in each region, this usage is innocuous. ξ_t represents a preference shock with mean unity, and $\theta(>1)$ denotes the elasticity of substitution between the variety of goods. It can be shown that the optimization of the allocation of consumption goods yields the aggregate price index $P_t \equiv \left[\int_0^1 P_t(j)^{1-\theta} dj\right]^{\frac{1}{1-\theta}}$.

Assume that the household is required to use cash in purchasing consumption goods. At the beginning of period t, the amount of cash available for the purchase of consumption goods is $M_{t-1} + \int_0^1 W_t(i)L_t(i)di - \int_0^1 D_t(i)di$, where M_{t-1} is the nominal balance held from period $t-1$ to t, and $\int_0^1 W_t(i)L_t(i)di$ represents the total wage income paid in advance by intermediate-goods firms. The household also makes a one-period deposit $D_t(i)$ in commercial bank i, the interest on which (R_t) is paid at the end of the period. It is assumed that the household has deposits in all of the commercial banks. Accordingly, the following cash-in-advance constraint must be satisfied at the beginning of period t:[7]

$$\int_0^1 P_t(j)C_t(j)dj \leq M_{t-1} + \int_0^1 W_t(i)L_t(i)di - \int_0^1 D_t(i)di.$$

[7]With this specification, it is implicitly assumed that financial markets open before the goods market.

The household's budget constraint is given by

$$M_t = M_{t-1} + \int_0^1 W_t(i)L_t(i)di - \int_0^1 D_t(i)di - \int_0^1 P_t(j)C_t(j)dj \\ + R_t \int_0^1 D_t(i)di + \Pi_t - T_t,$$

where Π_t denotes the sum of profits transferred from firms and commercial banks, and T_t is a lump-sum tax.

The demand for good j is expressed as

$$C_t(j) = \left(\frac{P_t(j)}{P_t}\right)^{-\theta} C_t. \tag{1}$$

The budget constraint can then be rewritten as

$$M_t = M_{t-1} + \int_0^1 W_t(i)L_t(i)di - \int_0^1 D_t(i)di - P_tC_t \\ + R_t \int_0^1 D_t(i)di + \Pi_t - T_t.$$

In an equilibrium with a positive interest rate, the following equality must hold:

$$P_tC_t = M_{t-1} + \int_0^1 W_t(i)L_t(i)di - \int_0^1 D_t(i)di. \tag{2}$$

This implies that the amount of total consumption expenditure is equal to cash holdings as long as there is an opportunity cost of holding cash. Then, the budget constraint leads to $M_t = R_t \int_0^1 D_t(i)di + \Pi_t - T_t$. Eliminating the money term from equation (2) yields an alternative expression of the budget constraint:

$$P_tC_t = R_{t-1}\int_0^1 D_{t-1}(i)di + \int_0^1 W_t(i)L_t(i)di - \int_0^1 D_t(i)di + \Pi_{t-1} - T_{t-1}.$$

The first-order conditions for the household's optimization problem are

$$\frac{\xi_t^{1-\sigma} C_t^{-\sigma}}{P_t} = \beta R_t E_t \left[\frac{\xi_{t+1}^{1-\sigma} C_{t+1}^{-\sigma}}{P_{t+1}} \right], \tag{3}$$

$$\frac{W_t(i)}{P_t} = \frac{L_t(i)^{\omega}}{\xi_t^{1-\sigma} C_t^{-\sigma}}, \tag{4}$$

where β and E_t are the subjective discount factor and the expectations operator conditional on information in period t, respectively.

3.2 Intermediate-Goods Firms

Intermediate-goods firm $i \in (0, 1)$ produces a differentiated intermediate good, $Z_t(i)$, by using the labor force of type i as the sole input. The production function is simply given by

$$Z_t(i) = A_t L_t(i), \tag{5}$$

where A_t is a countrywide productivity shock with mean unity. The owners of intermediate-goods firms must pay wage bills before goods markets open. Specifically, the owner of firm i borrows funds, $W_t(i)L_t(i)$, from commercial bank i at the beginning of period t at a gross nominal interest rate R_t^i. At the end of the period, intermediate-goods firm i must repay $R_t^i W_t(i) L_t(i)$ to bank i, so that the nominal marginal cost for firm i leads to $MC_t(i) = R_t^i W_t(i)/A_t$. Here, it is assumed that firm i can borrow funds only from the regional bank i since loan markets are geographically segmented. This assumption prohibits arbitrages, and thereby lending rates are allowed to differ across regional banks. Although such a situation might overly emphasize the role of the financial market's segmentation, a number of studies have found evidence of lending rate dispersion across intranational and international regions that cannot be explained by differences in riskiness.[8]

[8] For instance, see Berger, Kashyap, and Scalise (1995), Davis (1995), and Driscoll (2004) for the United States and Buch (2001) for the euro area. Buch (2000) provides a survey of the literature on lending-market segmentation in the United States.

It is assumed for simplicity that intermediate-goods firms are able to set prices flexibly. The price of $Z_t(i)$ will then be given by

$$P_t^z(i) = \frac{\theta_z}{(\theta_z - 1)(1 + \tau^m)} \frac{R_t^i W_t(i)}{A_t}, \tag{6}$$

where τ^m is a subsidy rate imposed by the government in such a way that $\theta_z \bar{R}/[(\theta_z - 1)(1 + \tau^m)] = 1$. It should be noted that since intermediate-goods firms borrow funds, the borrowing rates become an additional production cost. Thus, a rise in borrowing rates has a direct effect of increasing intermediate-goods prices.[9] Note also that since borrowing rates are allowed to differ across firms, it would become a source of price dispersion.

3.3 Final-Goods Firms

Each final-goods firm uses a composite of intermediate goods as the input for production. The production function is given by

$$Y_t(j) = \left[\int_0^1 Z_t^j(i)^{\frac{\theta_z - 1}{\theta_z}} di\right]^{\frac{\theta_z}{\theta_z - 1}}, \theta_z > 1,$$

where $Y_t(j)$ and $Z_t^j(i)$ represent a differentiated consumption good and the firm j's demand for individual intermediate good i, respectively. Optimization regarding the allocation of inputs yields the price index $P_t^z \equiv \left[\int_0^1 P_t^z(i)^{1-\theta_z} di\right]^{\frac{1}{1-\theta_z}}$. Accordingly, the firm j's demand for intermediate good i is expressed as follows:

[9] τ^m eliminates the distortions stemming both from monopolistic power and a positive steady-state interest rate ($\bar{R}$). Here, a positive steady-state interest rate is distortionary since the marginal cost would no longer be equal to v'/u'.

$$Z_t^j(i) = \left(\frac{P_t^z(i)}{P_t^z}\right)^{-\theta_z} Y_t(j).$$

Since $Z_t(i) = \int_0^1 Z_t^j(i)dj$ must hold in equilibrium, the demand function leads to

$$\begin{aligned} Z_t(i) &= \left(\frac{P_t^z(i)}{P_t^z}\right)^{-\theta_z} \int_0^1 Y_t(j)dj \\ &= \left(\frac{P_t^z(i)}{P_t^z}\right)^{-\theta_z} Y_t V_t^y, \end{aligned} \tag{7}$$

where $Y_t \equiv \left[\int_0^1 Y_t(j)^{\frac{\theta-1}{\theta}} dj\right]^{\frac{\theta}{\theta-1}}$ and $V_t^y \equiv \int_0^1 (Y_t(j)/Y_t)dj$. Note that V_t^y becomes larger than unity if $Y_t \neq Y_t(j)$ for some j.

It is assumed that final-goods firms are unable to adjust prices freely. Following Calvo (1983), we consider a situation in which a fraction $1-\phi$ of firms can change their prices, while the remaining fraction ϕ cannot. The price-setting problem of final-goods firms leads to

$$\max_{\tilde{P}_t} E_t \sum_{s=0}^{\infty} \phi^s \Gamma_{t,t+s} \left[(1+\tau^f)\tilde{P}_t - P_{t+s}^z\right] \left(\frac{\tilde{P}_t}{P_{t+s}}\right)^{-\theta} C_{t+s}, \tag{8}$$

where $\tilde{P}_t$ is the price of final goods set by firms that can adjust prices in period t, and $\Gamma_{t,t+s} \equiv \beta^s \frac{u'(C_{t+s};\xi_{t+s})P_t}{u'(C_t;\xi_t)P_{t+s}}$ denotes the stochastic discount factor up to period $t+s$. τ^f represents a subsidy rate, where $\tau^f = 1/(\theta-1)$.[10]

Log-linearizing the resultant first-order condition leads to

$$\pi_t = \beta E_t \pi_{t+1} + \lambda_F \left(p_t^z - p_t\right), \tag{9}$$

[10]Note that firms that can adjust prices in the same period set an identical price. Although different intermediate-goods firms may set different prices, marginal costs for final-goods firms are identical since the allocations of intermediate inputs are the same.

where $\lambda_F \equiv (1-\phi)(1-\beta\phi)/\phi$ and $\pi_t \equiv p_t - p_{t-1}$. Henceforth, for an arbitrary variable X_t, $x_t \equiv \log(X_t/\bar{X})$, where $\bar{X}$ denotes the steady-state value. Equation (9) is a version of the New Keynesian Phillips curve that has been used in numerous recent studies. Note that the term $p_t^z - p_t$ is equivalent to the real marginal cost of producing a final good, which is common across firms. Evidently, $p_t^z - p_t$ becomes zero if final-goods prices are fully flexible.

3.4 Financial Intermediary

Intermediate-goods firm i needs to borrow funds from commercial bank i at the start of each period in order to compensate for wage bills that must be paid in advance. At the beginning of period t, commercial bank i receives deposit $D_t(i)$ and money injection $M_t - M_{t-1} \equiv \Delta M_t$ from the household and the central bank, respectively. The former becomes the liability of the commercial bank, while the latter corresponds to its net worth. On the other hand, commercial bank i lends funds, $W_t(i)L_t(i)$, to intermediate-goods firm i. Therefore, the following equality must hold in equilibrium:

$$D_t(i) + \Delta M_t = W_t(i)L_t(i), \forall\ i \in (0,1). \tag{10}$$

The left-hand side and right-hand side can also be interpreted as representing the supply and the demand for funds, respectively. At the end of the period, commercial bank i repays its principle plus interest, $R_t(W_t(i)L_t(i) - \Delta M_t)$, to the household. The household also indirectly receives the money injection from the central bank through the profit transfer from commercial banks.

As is shown in appendix 1, firm i's demand for funds can be expressed as

$$W_t(i)L_t(i) = (R_t^i)^{\frac{-(1+\omega)\theta_z}{1+\omega\theta_z}} \Lambda_t \equiv \Psi\big(R_t^i; \Lambda_t\big),$$

where Λ_t is a function of aggregate variables that individual firms and commercial banks take as given. Obviously, firm i's demand for funds, $\Psi(R_t^i; \Lambda_t)$, decreases in R_t^i since an increase in R_t^i raises the marginal costs and thereby reduces its production.

Now let us specify the profit-maximization problem of commercial banks. It is assumed here that in each period, each commercial bank can adjust its loan rate with probability $1-q$. The probability of adjustment is independent of the time between adjustments. The problem for commercial bank i is then given by

$$\max_{R_t^i} E_t \sum_{s=0}^{\infty} q^s \Gamma_{t,t+s} \left[(1+\tau^b) R_t^i \Psi\left(R_t^i; \Lambda_{t+s}\right) - R_{t+s} \Psi\left(R_t^i; \Lambda_{t+s}\right)\right], \tag{11}$$

where τ^b represents a subsidy rate such that $(1+\omega)\theta_z/[(\theta_z - 1)(1+\tau^b)] = 1$. The commercial bank in region i takes into account the effect of a change in R_t^i on $W_t(i)L_t(i)$, while taking as given P_t, P_t^z, Y_t, C_t, V_t^y, ΔM_t, and R_t. The second term in the square bracket is according to the equilibrium condition (10), which implies that, given the value of ΔM_t, a change in $W_t(i)L_t(i)$ must be followed by the same amount of change in $D_t(i)$.

It can be shown that the first-order condition for this problem is given by

$$E_t \sum_{s=0}^{\infty} (q\beta)^s \frac{C_{t+s}^{-\sigma} \xi_{t+s}^{1-\sigma} \Lambda_{t+s}}{P_{t+s}} \left(R_t^i - R_{t+s}\right) = 0.$$

Log-linearizing this condition yields

$$r_t^i = \tilde{r}_t = (1 - q\beta) E_t \sum_{s=0}^{\infty} (q\beta)^s r_{t+s}. \tag{12}$$

This optimality condition implies that all the commercial banks that adjust in the same period impose an identical loan rate, $\tilde{r}_t$. It should be pointed out that the newly adjusted loan rates depend largely on the expectations of future policy rate as well as the current policy rate. The weight on the current policy rate is only $1 - q\beta$, while the weights on future policy rates sum up to $q\beta$. This is the well-known forward-looking property stemming from staggered pricing. If one interprets the banks' problem as price determination under conventional Calvo pricing, the value of q is simply considered as representing the degree of stickiness. From a different point of view, however, the newly adjusted loan rates expressed as (12) could be regarded as an outcome of a long-term contract, where commercial banks lend funds by charging a fixed interest rate with the proviso that there is a possibility of revaluation with probability $1 - q$. In this case, the length of maturity is expressed as a random variable that has a geometric distribution with parameter $1-q$. In fact, as is shown below, there is a close relation between the newly adjusted loan rates and long-term "market" interest rates.

In order to obtain model-consistent long-term interest rates, suppose for the moment that the length of maturity is known with certainty. The representative commercial bank's problem for the determination of an n-period loan rate will be given by[11]

$$\max_{R_{n,t}} E_t \sum_{s=0}^{n-1} \Gamma_{t,t+s}[(1+\tau^b)R_{n,t}\Psi(R_{n,t};\Lambda_{t+s}) - R_{t+s}\Psi(R_{n,t};\Lambda_{t+s})].$$

The first-order condition is

$$E_t \sum_{s=0}^{n-1} \beta^s \frac{C_{t+s}^{-\sigma}\xi_{t+s}^{1-\sigma}\Lambda_{t+s}}{P_{t+s}}(R_{n,t} - R_{t+s}) = 0.$$

It follows that

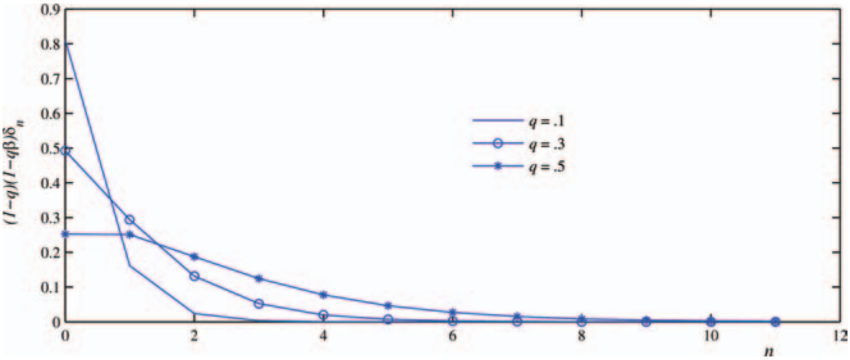
$$r_{n,t} = \left(\sum_{s=0}^{n-1} \beta^s\right)^{-1} E_t \sum_{s=0}^{n-1} \beta^s r_{t+s}. \tag{13}$$

While this is an expression of loan rates of maturity-n, this can also be interpreted as the n-period market interest rates since the bank will set $r_{n,t}$ in such a way that the expected return equals the expected cost as long as there are neither adjustment costs nor default risk.[12] Because the bank faces no uncertainty in regard to the length of periods between adjustments, $r_{n,t}$ must be an efficient estimate of the per-period cost of funds from period t to $t+n$. Unsurprisingly, this endogenously derived relation, (13), takes a form known as the expectations theory of the term structure. Here, the consumer's subjective discount factor, β, is used as the discount factor on expected future short-term rates.

[11] Index i is now dropped for brevity since the following hypothetical problem is common to all banks.

[12] In order to consider a competitive equilibrium of market interest rates, distortion stemming from the monopolistic market power is removed by government subsidies.

Figure 1. Weights on Long-Term Interest Rates



Using expression (13), we can present the following proposition.

PROPOSITION 1. *If the n-period market interest rate is written as (13), then the newly adjusted loan rates, $\tilde{r}_t$, can be expressed as*

$$\tilde{r}_t = (1 - q\beta)(1 - q)(r_t + \delta_1 r_{2,t} + \delta_2 r_{3,t} + \ldots),$$

where $\delta_k = \frac{q^k(1-\beta^{k+1})}{1-\beta}$ *for* $k \geq 0$*, and* $\sum_{k=0}^{\infty} \delta_k = ((1-q\beta)(1-q))^{-1}$*. Moreover,* $\delta_{k+1} < \delta_k$ *holds for all* $k \geq 0$ *if and only if* $q < (1+\beta)^{-1}$*.*

Proof. See appendix 2.

This proposition states that the newly adjusted loan rate can be expressed as a weighted average of long-term market interest rates of various maturities. It turns out that the weights on long-term rates are largely dependent on the probability of revaluation.[13] Figure 1 illustrates examples of δs. As is clear from the figure, the weights on short-term rates decrease with larger q. This reflects the fact that the currently adjusted loan rates will be expected to live for longer periods as the revaluation probability becomes lower.

[13]Interestingly, if one interprets δ as the time-varying discount factor, it takes a form of hyperbolic discounting, where the discount rate itself decreases as the maturity increases.

4. Equilibrium Dynamics

Before proceeding, let us summarize the key equilibrium relations in preparation for succeeding analyses. Appendix 3 shows that the real marginal cost of final-goods firms, $p_t^z - p_t$, can be expressed as $p_t^z - p_t = r_t^l + (\sigma + \omega)x_t$, where r_t^l and x_t denote an average loan rate and an output gap, respectively. The New Keynesian Phillips curve (NKPC) can thus be written as

$$\pi_t = \beta E_t \pi_{t+1} + \lambda_F(\sigma + \omega)x_t + \lambda_F r_t^l. \tag{14}$$

As was pointed out by Ravenna and Walsh (2006), the difference between the standard NKPC and the NKPC with the cost-channel effect lies in the presence of an additional interest rate term. Yet, our expression differs from theirs in that the interest term in (14) is expressed by the average loan rate, not by the policy rate. Since our model incorporates profit-maximization behavior of commercial banks, retail loan rates are distinguished from the policy instrument in an endogenous manner.[14] It turns out that the average loan rate, r_t^l, becomes a determinant of inflation because a rise in the average loan rate leads to a higher marginal cost for final-goods' production.

An obvious outcome of this modification is that as long as $q > 0$, the cost-channel effect is weakened compared with the case of perfect pass-through. This is not only because only a fraction $(1 - q)$ of commercial banks reset their loan rates each period, but also because a newly charged loan rate differs from the policy rate in that period. Since the correlation between the policy rate and the marginal cost of intermediate-goods firms becomes weaker as q increases, the influence of a policy shift on final-goods prices will be reduced accordingly.[15]

[14] Chowdhury, Hoffmann, and Schabert (2006) also make distinctions between a money-market rate and a lending rate in a model similar to ours, but their distinction depends fully on the assumption that there exists a proportional relationship between the two interest rates.

[15] Recently, Tillmann (2007) estimated the NKPC of the form (14) using the data for the United States, the United Kingdom, and the euro area. He showed that inflation dynamics can be better explained if the short-term rate that appeared in the Ravenna-Walsh NKPC is replaced with lending rates.

The standard aggregate demand equation can be obtained by log-linearizing the Euler equation (3):

$$x_t = E_t x_{t+1} - \frac{1}{\sigma}\left(r_t - E_t \pi_{t+1} - rr_t^n\right), \tag{15}$$

where $rr_t^n \equiv \sigma((1+\omega)/(\sigma+\omega))E_t\Delta a_{t+1} + \omega((\sigma-1)/(\sigma+\omega))E_t\Delta\hat{\xi}_{t+1})$ denotes the natural rate of real interest, where $\hat{\xi}_t \equiv \log(\xi_t)$.

Now we turn our attention to the determination of the average loan rate. By the nature of commercial banks' loan rate setting, the average loan rate is given by

$$r_t^l = q r_{t-1}^l + (1-q)\tilde{r}_t.$$

The current average loan rate can be expressed as a weighted average of the newly adjusted loan rate and the previous average loan rate. Eliminating $\tilde{r}_t$ from (12) yields

$$\Delta r_t^l = \beta E_t \Delta r_{t+1}^l + \lambda_B\left(r_t - r_t^l\right), \tag{16}$$

where $\Delta r_t^l \equiv r_t^l - r_{t-1}^l$ and $\lambda_B \equiv (1-q)(1-q\beta)/q$. Equation (16) says that a shift in the average loan rate will be caused by a discrepancy between the policy rate and the average loan rate as well as a change in the expectation of future loan rate. This equation can also be written as

$$r_t^l = \frac{\beta}{1+\beta+\lambda_B}E_t r_{t+1}^l + \frac{\lambda_B}{1+\beta+\lambda_B}r_t + \frac{1}{1+\beta+\lambda_B}r_{t-1}^l.$$

Intuitively, the average loan rate is expressed as a weighted average of the expected loan rate, the current policy rate, and the previous loan rate.[16] It states that the relative weights on the expected loan rate and the previous loan rate increase as the sluggishness of loan rates deteriorates. Conversely, the current loan rate approaches the current policy rate as q goes to zero.

In an environment where the central bank controls r_t, equations (14), (15), and (16) and a policy rule describe the behavior of π, x, r^l, and r. We next explore the central bank's optimal policy rate setting in the following sections.

[16] After I finished writing this paper, I found that Teranishi (2008) also obtained similar results in a different setting. We arrived at the similar results completely independently of each other.

5. Social Welfare

This section attempts to obtain a welfare-based objective function for monetary policy by approximating the household's utility function up to a second order. Appendix 4 shows that the one-period utility function can be approximated as

$$U_t = -\frac{\bar{L}^{1+\omega}}{2}(\sigma + \omega)\left\{x_t^2 + \left(\frac{\theta}{\sigma+\omega}\right) var_j p_t(j) + \left[\frac{\theta_z}{(1+\omega\theta_z)(\sigma+\omega)}\right] var_i r_t^i\right\} + t.i.p., \tag{17}$$

where an upper bar means that the variable denotes the corresponding steady-state value, and $t.i.p.$ represents terms that are independent of policy, including terms higher than or equal to third order. A notable feature of equation (17) is the presence of the variance of loan rates. This result is quite intuitive given that the determination of loan rates is specified as Calvo-type pricing. Equation (17) reveals that the variance of lending rates reduces social welfare in the same manner as the variance of final-goods prices does.

Woodford (2001, 22–23) shows that the present discounted value of the variance of prices can be expressed in terms of inflation squared. That is,

$$\sum_{s=0}^{\infty} \beta^s var_j p_{t+s}(j) = \lambda_F^{-1} \sum_{s=0}^{\infty} \beta^s \pi_{t+s}^2.$$

It is straightforward to apply this result to rewriting the present discounted value of the variance of lending rates. It follows that

$$\sum_{s=0}^{\infty} \beta^s var_i r_{t+s}^i = \lambda_B^{-1} \sum_{s=0}^{\infty} \beta^s \left(\Delta r_{t+s}^l\right)^2.$$

It turns out that the present discounted value of the variance of lending rates can be expressed in terms of a change in the average loan rate.

Consequently, the social welfare function can be rewritten as

$$E_t \sum_{s=0}^{\infty} \beta^s U_{t+s} = -\frac{\bar{L}^{1+\omega}}{2}(\sigma+\omega) E_t \sum_{s=0}^{\infty} \beta^s \{x_{t+s}^2 + \psi_\pi \pi_{t+s}^2 + \psi_r \left(\Delta r_{t+s}^l\right)^2\} + t.i.p., \tag{18}$$

where $\psi_\pi \equiv \theta/[\lambda_F(\sigma+\omega)]$ and $\psi_r \equiv \theta_z/[\lambda_B(1+\omega\theta_z)(\sigma+\omega)]$ represent the relative weights on inflation and the rate of change in the average loan rate, respectively. Equation (18) states that fluctuations in the average loan rate will reduce social welfare when commercial banks adjust loan rates only infrequently. This finding is closely parallel to a well-known result obtained under staggered goods prices. Under staggered goods prices, the rate of inflation enters into the welfare function because a nonzero inflation gives rise to price dispersion. Under staggered loan rate contracts, the rate of change in the average loan rate enters into the welfare function because changes in the average loan rate inevitably entail loan rate dispersion.

It might also be noted that equation (18) closely resembles a conventional loss function that has been frequently employed in the recent literature on monetary policy for the purpose of capturing actual central banks' interest rate smoothing (i.e., policy rate smoothing) behavior. Specifically, in many previous studies it has been *assumed* that a monetary authority tries to minimize a loss function of the form[17]

$$Loss_t^c = x_t^2 + \lambda \pi_t^2 + \nu(\Delta r_t)^2.$$

This expression essentially differs from ours in that the third term is expressed in terms of the policy instrument rather than the average loan rate. Here, the relation between Δr_t and Δr_t^l can be written from proposition 1 as

$$\Delta r_t^l = (1-q\beta)(1-q)^2[\Delta r_t + \delta_1 \Delta r_{2,t} + \ldots + q(\Delta r_{t-1} + \delta_1 \Delta r_{2,t-1} + \ldots) + \ldots].$$

[17]See, for example, Rudebusch and Svensson (1999), Rudebusch (2002a, 2002b), Levin and Williams (2003), and Ellingsen and Söderström (2004). See Sack and Wieland (2000) and Rudebusch (2006) for a survey of studies on interest rate smoothing.

Thus, Δr_t constitutes only a fraction $(1-q\beta)(1-q)^2$ of Δr_t^l. The rest of the components of Δr_t^l are expressed by the past policy shifts and the current and past changes in long-term rates. Notice that equation (18) and the conventional loss function never coincide since the loan rate smoothing term will disappear in the limiting case of $q = 0$, where $r_t^l = r_t$ holds. Nevertheless, the desirability of policy rate smoothing might be retained in that it contributes to the stabilization of loan rates through the stabilization of long-term rates. A further discussion about the relationship between loan rate stabilization and the central bank's policy rate smoothing will be given in the next section.

6. Monetary Policy in the Presence of Loan Rate Stickiness

This section attempts to explore desirable monetary policy in the presence of incomplete interest rate pass-through, focusing on the question of how the desirable path of the policy rate will be modified once loan rate stickiness is taken into account. Provided that the central bank tries to maximize social welfare function (18), the presence of loan rate stickiness affects inflation and output through two channels. On one hand, the presence of loan rate stickiness mitigates the cost-channel effect of a policy shift on inflation. On the other hand, the central bank has to put some weight on loan rate stabilization in the face of loan rate stickiness. It is shown below that the former effect tends to reduce the desirability of policy rate smoothing since there is less need for the central bank to pay attention to the undesirable effect that a policy shift has on inflation. In contrast, the latter effect increases the desirability of policy rate smoothing since the stabilization of the policy rate leads, at least to some extent, to loan rate stability. These two aspects are thoroughly examined in the succeeding subsections.

In the following, we consider two alternative policy regimes: standard Taylor rule and commitment under a timeless perspective. In addition, we also investigate optimal policy in the face of a loan premium shock, which directly alters the markup in loan rate pricing. It is shown that the role of the loan rate stability term depends largely on the underlying nature of shocks.

6.1 Baseline Parameters

The baseline parameters used in the analysis are as follows: $\beta = .99$, $\sigma = 1.5$, $\omega = 1$ (Ravenna and Walsh 2006), and $\theta = 7.88$ (Rotemberg and Woodford 1997). We set the elasticity of substitution for intermediate goods at the value equal to θ, thus $\theta_z = 7.88$. Following Galí and Monacelli (2005), we specify the process of productivity shock as $a_t = .66a_{t-1} + \zeta_t^a$, where the standard deviation of ζ_t^a is set at .007. As for the preference shock, we specify the process as $\hat{\xi}_t = .5\hat{\xi}_{t-1} + \zeta_t^\xi$, where the standard deviation of ζ_t^ξ is set at .005.[18] The degree of price stickiness, ϕ, is chosen such that the slope of the Phillips curve is equal to .58, the value reported by Lubik and Schorfheide (2004). It follows that $\phi = .623$ $((1-\phi)(1-\beta\phi)(\sigma+\omega)/\phi = .58)$, which leads to $\psi_\pi = 13.582$.

As mentioned in section 2, recent studies reported different estimates of the degree of loan rate pass-through at the euro-area aggregated level. Here, three alternative values are considered: q_L, q_M, and q_H. According to table 1 of de Bondt, Mojon, and Valla (2005), the lowest value of the estimated degree of loan rate pass-through for short-term loans to enterprises is .25 (Sander and Kleimeier 2002; Hofmann 2003), while the largest one is .76 (Heinemann and Schüler 2002). Since these estimates are obtained from monthly data, we have to convert them to their quarterly counterparts. For example, in the case of the largest degree of pass-through, q_L is set such that $1 - q_L = .76 + (1 - .76).76 + (1 - .76)^2.76$, which leads to $q_L = .014$. Likewise, q_H is set at .422. Finally, q_M is set at .177, the average of all the estimates reported by thirteen studies cited in table 1 of de Bondt, Mojon, and Valla (2005). This implies that the relative weight on the loan rate, ψ_r, is .445, .092, and .005 if $q = q_H$, q_M, and q_L, respectively.

6.2 Policy Rate Smoothing and the Degree of Interest Rate Pass-Through

Before investigating optimal policy, it should be pointed out that the degree of interest rate pass-through is heavily dependent on the

[18]The essential results shown below will never change in the absence of the preference shock.

policy rate behavior. The impact of a policy shift on retail loan rates can vary not only with the frequency of loan rate adjustments but also with the expectation of future policies. It is useful to gain a better understanding of the relation between policy rate smoothness and the degree of interest rate pass-through.

Suppose for exposition that the policy rule is expressed as

$$r_{t+1} = \rho r_t + \eta_{t+1},$$

where $\rho \in [0, 1)$ describes the degree of policy rate inertia. η_{t+1} is a white noise, which represents an unpredictable component of the policy rate. Then, the average loan rate is given as

$$r_t^l = \frac{(1-q)(1-q\beta)}{1-\rho q\beta} r_t + q r_{t-1}^l.$$

In this case, the degree of *instantaneous* interest rate pass-through can be expressed as

$$\frac{\partial r_t^l}{\partial r_t} = \frac{(1-q)(1-q\beta)}{1-\rho q\beta}.$$

This implies that the impact of a policy shift on the current average loan rate will become larger as the degree of policy inertia increases.

More generally, it can be shown that the impact of a current policy shift on the s-period-ahead average loan rate leads to

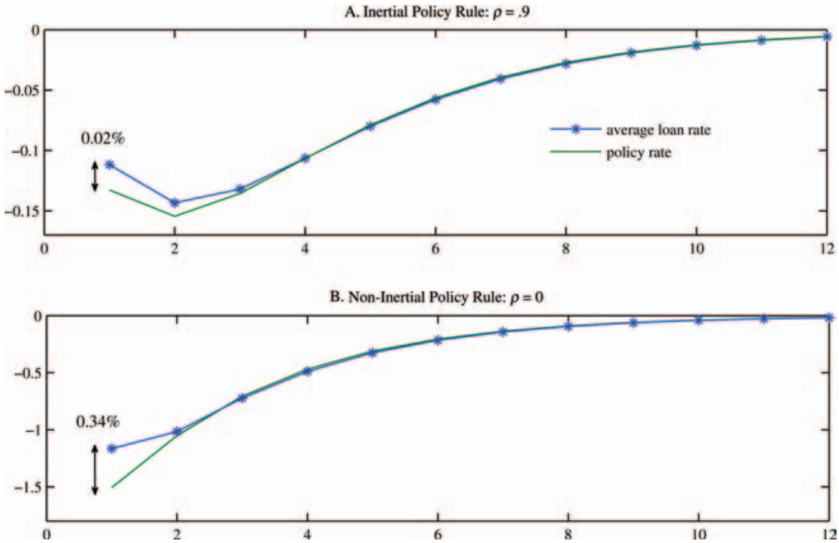
$$\frac{\partial r_{t+s}^l}{\partial r_t} = \frac{(1-q)(1-q\beta)(\rho^{s+1} - q^{s+1})}{(1-\rho q\beta)(\rho - q)}.$$

Accordingly, the degree of *cumulative* interest rate pass-through can be given as

$$\sum_{s=0}^{\infty} \frac{\partial r_{t+s}^l}{\partial r_t} = \frac{1-q\beta}{(1-\rho q\beta)(1-\rho)}.$$

Notice that under an inertial policy rule, current policy rate has an impact on future average loan rate not only through the persistent dynamics in the average loan rate but also through the policy rate dynamics itself. Since commercial banks' loan rate determination is made in a forward-looking manner, a policy shift can have a larger impact on loan rates as the shift becomes more persistent.

Figure 2. Responses of the Average Loan Rate and the Policy Rate under Alternative Policy Rules



Now, let us reexamine the above implications by using a more general policy rule. We employ the following standard Taylor rule:

$$r_t = \rho r_{t-1} + (1-\rho)\big(rr_t^* + \phi_\pi \pi_t + (\phi_x/4)x_t\big), \qquad (19)$$

where ϕ_π and ϕ_x are set at 1.5 and .5, respectively. ρ is set at .9 under an "inertial policy," while $\rho = 0$ under a "non-inertial policy." Figure 2 illustrates impulse responses to a one-standard-deviation productivity shock.[19] The figure shows that there is an appreciable difference between the two cases in the reaction of the average loan rate to the policy rate. Under the inertial policy rule, the paths of r_t and r_t^l are shown to be very close. Specifically, the spread between the two paths on impact is only .02 percent, where the instantaneous interest rate pass-through turns out to be 84.2 percent. Under the non-inertial policy rule, on the other hand, the initial spread amounts to .34 percent, where the instantaneous interest rate pass-through is 77.3 percent. Thus, the property that a lagged policy rate

[19] Henceforth, interest rate responses are illustrated in annual rate.

term plays a key role as a determinant of the degree of interest rate pass-through still holds under the standard Taylor rule.

6.2.1 Some Intuitions into the Desirability of Policy Rate Smoothing

In order to obtain some intuitions into the relationship between loan rate smoothing and policy rate smoothing, let us first express the current loan rate solely in terms of the policy rate.

$$\begin{aligned} r_t^l &= (1-q)\tilde{r}_t + q r_{t-1}^l \\ &= (1-q)(1-q\beta)[r_t + \beta q E_t r_{t+1} + (\beta q)^2 E_t r_{t+2} + \ldots] \\ &\quad + q(1-q)(1-q\beta)[r_{t-1} + \beta q E_{t-1} r_t + (\beta q)^2 E_{t-1} r_{t+1} + \ldots] + \ldots . \end{aligned}$$

It follows that

$$\begin{aligned} \Delta r_t^l = {} & (1-q)(1-q\beta)[\Delta r_t + \beta q(E_t \Delta r_{t+1} + r_t - E_{t-1} r_t) + \ldots] \\ & + q(1-q)(1-q\beta)[\Delta r_{t-1} + \beta q(E_{t-1}\Delta r_t + r_{t-1} \\ & - E_{t-2} r_{t-1}) + \ldots] + \ldots . \end{aligned}$$

This expression shows that the growth rate of the current average loan rate is determined not only by the current and the past policy rate increments but also by the expectations of future increments and policy surprises. This reveals two important implications for loan rate stabilization. First, the presence of increment terms implies that the policy rate should be continuously smoothed. This is simply because any policy rate changes inevitably give rise to a shift in the newly adjusted loan rates. The average loan rate becomes more stable as the policy rate in any given period becomes closer to the previous period's level. This necessarily requires the policy rate to be inertial or history dependent.

Second, the "surprise" terms, $r_t - E_{t-1} r_t$, $r_{t-1} - E_{t-2} r_{t-1}$, and so on, state that the central bank should avoid causing a policy surprise, for a revision of commercial banks' policy rate expectations will entail a shift in the newly adjusted loan rates. This is quite natural in that the commercial banks' loan rate determination is based on the expectation of future policy rates conditional on information available at that time.[20] It should be noted that not only expectation errors in the current period but also expectation errors made in the past cause a change in the current average loan rate. The reason for this is as follows: suppose that the policy rate has not been changed since m periods ago, and the last policy shift had not been anticipated at that time. In the current period, given that the policy rate is still expected to be constant in the future, loan rates between the ages of 1 and m need not be changed even if they have a chance of adjustment, because the last policy shift is already incorporated. In contrast, loan rates that have not been adjusted for the past m periods need to be readjusted in the current period since they have not yet incorporated the unexpected policy shift that occurred m periods ago. Since a certain fraction of all the loan rates is necessarily over the age of m, their readjustments inevitably occur and cause a shift in the current average loan rate.

It is evident that once the central bank changes the policy rate, the resultant loan rate readjustments will persist forever. These loan rate readjustments will never end, even if the policy rate is (and is expected to be) kept unchanged from then on. However, such persistent effects would be alleviated if the policy shift was correctly anticipated in advance. Of course, even if a policy shift is incorporated in advance, a revision of expectation necessarily occurs at least to some extent (unless the entire policy rate path was fully incorporated at the initial period). Nevertheless, the extent of an expectation revision can be made smaller as the timing of incorporation becomes earlier since the corresponding adjustments of loan rates will be dispersed over some periods. In this sense, it could be said that the forecastability of future policy rates becomes another key to loan rate stability. Policy rate smoothing will contribute to loan rate stability by revealing some information regarding future policy rates.[21]

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[20] In fact, Svensson (2003) notes that the central bank should minimize the *surprise* in the policy rate. He proposed the (ad hoc) central bank's loss function of the form $L_t = Var(x_t) + \lambda Var(\pi_t) + \nu Var(E_{t-1}r_t - r_t)$.

[21] The necessity of the central bank's communicability is stressed by Kleimeier and Sander (2006). They argue that the impact of policy rate shifts on retail lending rates tends to be large in countries in which the central bank communicates well with the public. See also Woodford (2005) for a discussion of central bank communication.

6.3 The Role of the Loan Rate Stabilization Term

In this section, we address the issue of how the presence of a loan rate stabilization term, $\psi_r(\Delta r_t^l)^2$, affects the desirable policy. To this end, we especially focus on its relation with conventional policy rate smoothing or policy inertia. In investigating policy inertia under various degrees of loan rate stickiness, we have to explicitly distinguish between *policy* inertia and *intrinsic* inertia. To isolate policy inertia, we need to know to what extent current policymaking depends on the previous policy decision. Thus, we measure here the desirable degree of policy inertia by the size of the coefficient on the lagged policy rate in the case of a simple rule, and by the size of the relative weight on the policy rate smoothing term in the case of commitment.

6.3.1 Simple Rule

As a policy rule, we again employ (19). Here, optimal combinations of (ρ, ϕ_π, ϕ_x) are searched for under alternative values of q, ψ_π, and ψ_r. Since we want to know the effects of introducing the loan rate stabilization term on the optimal value of ρ, the relative weight on the loan rate, ψ_r, is set either at the endogenously determined value or at zero.

Optimal combinations of (ρ, ϕ_π, ϕ_x) are given in table 1.[22] It shows that, given the values of q and ψ_π, the optimal value of ρ

Table 1. Optimal Combinations of (ρ, ϕ_π, ϕ_x)

(ψ_π, ψ_r)	$(13.58, \psi_r)$	(13.58, 0)	$(4, \psi_r)$	(4, 0)
$q = q_L$	(.35, 3, 0)	(.35, 3, 0)	(.35, 1, 0)	(.35, 1, 0)
Relative Loss[a]	1.225	1.225	1.004	1.004
$q = q_M$	(.35, 3, 0)	(.35, 3, 0)	(.4, 1.65, 0)	(.35, 1.15, 0)
Relative Loss	1.116	1.116	1.004	1.008
$q = q_H$	(.3, 3, 0)	(.25, 3, 0)	(.4, 2.6, 0)	(.35, 3, 0)
Relative Loss	1.027	1.054	1.021	1.046

[a]The denominator of “relative loss” is the value of social loss that would be attained under commitment policy with the appropriate loan rate smoothing objective.

[22]The examined parameter ranges are as follows: [0, .95] for ρ, [0, 3] for ϕ_π, and [0, 1] for ϕ_x. The increment size of the grid is .05.

under loan rate stabilization is greater than or equal to that under $\psi_r = 0$. This implies that the introduction of a loan rate smoothing term into the welfare function tends to strengthen the desirability of policy rate smoothing, although this smoothing effect seems to be limited, especially when loan rate stickiness is not severe. This argument is also confirmed by the value of social loss under an optimal policy rule relative to the loss under timeless commitment. It turns out that the cost of putting a "too-small" weight on the previous policy rate is at most 2.6 percent.

6.3.2 Commitment

Next, let us turn to the case of commitment. Under commitment, the central bank is assumed to minimize a given loss function. In order to investigate the desirability of policy inertia, we specify here the central bank's one-period loss function as follows:

$$L_t^{smooth} = x_t^2 + \psi_\pi \pi_t^2 + \tilde{\psi}_r (\Delta r_t)^2.$$

In this case, the value of $\tilde{\psi}_r$ that minimizes social loss can be interpreted as a proxy for the optimal degree of policy inertia.

The optimal values of $\tilde{\psi}_r$ under alternative values of q are shown in table 2. It shows that as long as the value of q is not that small, the optimal value of $\tilde{\psi}_r$ tends to be larger under loan rate stabilization than under $\psi_r = 0$. This reconfirms the previous result that the introduction of a loan rate stabilization term justifies conventional policy rate smoothing. However, the quantitative importance of this smoothing effect is again not very large. The cost of ignoring policy rate stability is at most 3.6 percent.

Table 2. Optimal Weights on $(\Delta r_t)^2$

(ψ_π, ψ_r)	$(13.58, \psi_r)$	(13.58, 0)	$(4, \psi_r)$	(4, 0)
$q = q_L$	0	0	0	0
Relative Loss	1.000	1.000	1.000	1.000
$q = q_M$	.05	0	.05	0
Relative Loss	1.001	1.014	1.001	1.026
$q = q_H$	.05	0	.05	0
Relative Loss	1.007	1.026	1.014	1.051

Figure 3. Policy Rate Responses under Commitment

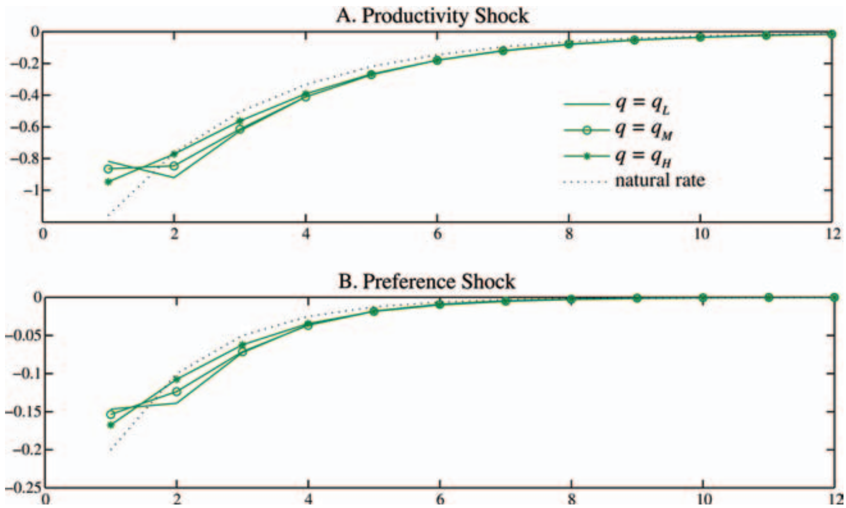
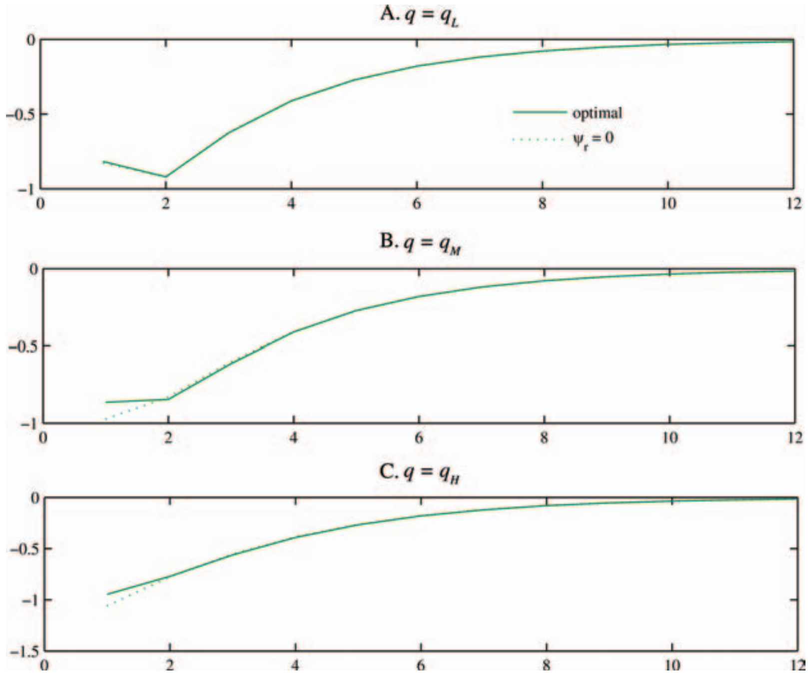


Figure 3 illustrates the policy rate responses under timeless commitment. It turns out that the optimal policy rate responses under alternative values of q differ only slightly. Nevertheless, it can be said from the figure that the initial reduction of the policy rate is largest (smallest) when $q = q_H$ (q_L). This is because an increase in q mitigates the cost-channel effect, the direct impact of a policy change on inflation. A reduction in the policy rate is needed in the face of a positive productivity or preference shock, but such an expansionary policy necessarily entails a negative effect on inflation due to the presence of the cost channel. As q increases, however, such an undesirable aspect becomes less important, and thereby the central bank can set the policy rate at a level closer to the natural rate of interest. The figures show that the cost-channel effect is quantitatively more influential than the policy rate smoothing effect, which stems from the change in the value of ψ_r.

In order to clarify the strength of the policy rate smoothing effect, figure 4 illustrates policy rate responses under commitment with and without loan rate smoothing.[23] It can be confirmed that the initial reduction in the policy rate is smaller in the presence of loan rate

[23]The obtained results are essentially the same in the case of a preference shock as well.

Figure 4. Policy Rate Responses to a Productivity Shock



stabilization. As expected, however, the difference between the two cases is not significantly different.

In summary, in the face of a productivity shock and/or a preference shock, the presence of a loan rate stabilization term itself supports, to some extent, the idea of conventional policy rate smoothing. However, it appears that the optimal policy is more strongly influenced by the cost-channel effect than the policy rate smoothing effect that stems from the presence of a loan rate stabilization term. In the next section, we reexamine the role of loan rate stabilization by introducing a loan rate premium shock, which directly changes the markup in loan rates.

6.4 Undesirability of Policy Rate Smoothing: The Case of a Loan Rate Premium Shock

In the above analysis, we investigated the optimal policy response in the face of a productivity shock or a preference shock. In such an environment, retail loan rates are determined solely by the policy rate, although those shocks have an indirect influence through the policy rate. In practice, however, it is usual for loan rates to fluctuate for reasons that are not directly linked to the policy rate behavior. One possible case is a shift in the loan rate premium triggered by changes in financial market conditions. While here we do not emphasize a particular cause of loan premium fluctuations, optimal policy in the face of such kinds of shocks is worth considering. A loan rate premium shock can be introduced by modifying the first-order condition of the commercial banks' problem as follows:

$$E_t \sum_{s=0}^{\infty} (q\beta)^s \frac{C_{t+s}^{-\sigma} \xi_{t+s}^{1-\sigma} \Lambda_{t+s}}{P_{t+s}} \left(R_t^i - \varphi_{t+s} R_{t+s}\right) = 0,$$

where $\varphi_{t+s} > 0$ and $E[\varphi_{t+s}] = 1$. It follows that

$$\Delta r_t^l = \beta E_t \Delta r_{t+1}^l + \lambda_B \left(r_t - r_t^l\right) + \lambda_B \hat{\varphi}_t,$$

where $\hat{\varphi}_t \equiv \log(\varphi_t)$. $\hat{\varphi}_t$ is actually a shock to the change in the loan rate but can also be interpreted as a shock to the level of the loan rate once redefined as $\tilde{\varphi}_t \equiv \frac{\lambda_B}{1+\lambda_B+\beta} \hat{\varphi}_t$.[24]

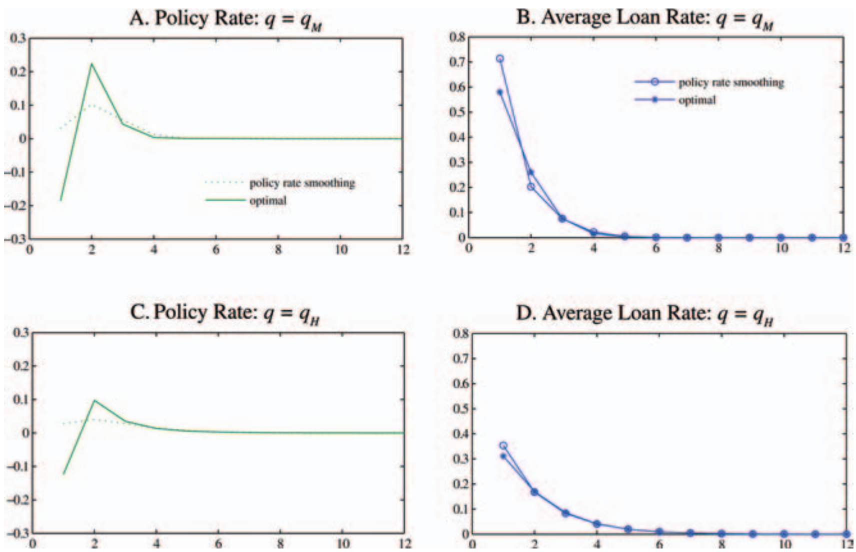
Figure 5 illustrates optimal policy rate responses under commitment to a 1 percent (at the annual rate) positive loan rate premium shock. It is assumed that this shock is a temporary one. It clearly shows that the presence of a loan rate stabilization term now plays a critical role in the conduct of monetary policy. Under optimal policies with a loan rate stabilization objective, the policy rate needs to be drastically reduced in the face of a rise in the loan rate premium. This is because a reduction in r_t can partially offset a rise in $\hat{\varphi}_t$, as is evident from the above equation. On the other hand, such a drastic policy rate reduction cannot be observed when the central

[24] Another way of introducing an exogenous loan rate shock is to assume a time-varying subsidy rate, τ_t^b. Provided that $E[\tau_{t+s}^b] = \tau^b$, then the modified loan rate adjustment equation leads to

$$\Delta r_t^l = \beta E_t \Delta r_{t+1}^l + \lambda_B \left(r_t - r_t^l\right) - \lambda_B \hat{\tau_t^b},$$

where $\hat{\tau_t^b} = \log((1+\tau_t^b)/(1+\tau^b))$. In this case, a below-the-average subsidy rate will act as a positive loan rate shock. See also Woodford (2003, ch. 6) for a discussion of time-varying markups in goods prices.

Figure 5. Impulse Responses to a 1 Percent Loan Premium Shock



bank conducts "optimal" policy rate smoothing, which is incorrect from the point of view of welfare maximization. In fact, since the optimal weight on the policy rate is very small (.05), this result will also hold even in the case where the central bank pays no attention to policy rate smoothing.

Figure 5 also illustrates the behavior of the average loan rate. As is clear from the figure, the response of the average loan rate is mitigated as q increases, since the fraction of newly adjusted loan rates declines. Along with this, the required amount of policy rate reduction turns out to be smaller under $q = q_H$ than under $q = q_M$. Although a rise in q has the effect of requiring more drastic policy shifts by increasing the size of the relative weight on the loan rate, such an effect is relatively small.

To sum up, the role of a loan rate stabilization term fundamentally alters according to the underlying nature of shocks. In the face of a shock that would directly shift inflation and output, the presence of the loan rate stabilization objective itself requires inertial policy. This is because loan rates are determined based only on the policy rate, in which case the only way to avoid fluctuations in loan rates is to avoid fluctuations in the policy rate. However, a shock that would directly give rise to loan rate fluctuations should be dampened by a drastic policy shift. Policy rate smoothing is not needed (and in fact is even harmful) when there is requirement for loan rate stabilization. From this point of view, it can be said that conventional policy rate smoothing is no longer a panacea. The case of a loan rate premium shock is an example for which policy rate smoothing should be abandoned.

7. Concluding Remarks

The main findings of this paper can be summarized as follows. First, when the pass-through from the policy rate to retail loan rates is incomplete, fluctuations in the average loan rate will reduce social welfare. This is because shifts in the average loan rate immediately give rise to a loan rate dispersion across firms, which ultimately yields an inefficient dispersion in hours worked. Accordingly, the central bank faces a policy trade-off in stabilizing inflation, an output gap, and the rate of change in the average loan rate.

Second, the introduction of a loan rate stabilization term in the central bank's loss function causes the optimal policy rate to become more inertial in the face of a productivity shock and a preference shock. In this sense, loan rate smoothing is closely parallel to conventional policy rate smoothing. However, such a smoothing effect turned out to be less influential than the cost-channel effect.

Third, the presence of a loan rate stabilization term requires a drastic policy reaction in the face of an exogenous shock that directly shifts retail lending rates, such as a shift in the loan rate premium. This result is counter to the conventional wisdom that the policy rate must be adjusted gradually in short steps. However, given the fact that the standard dynamic stochastic general equilibrium model usually ignored the cost of loan rate dispersion, this disagreement is not so surprising. The case of a loan premium shock is an example for which the central bank has to clearly distinguish between policy rate smoothing and loan rate smoothing.

We conclude by noting several points that should be addressed in future research. First, a more realistic framework for long-term interest contracts should be introduced. In the present paper, loan rate determination is specified as Calvo-type pricing. However, a more plausible situation would be that the length of maturity is determined at the time of contract and is allowed to differ across borrowers. Second, although our model treats the frequency of loan rate adjustments as exogenous, there is a possibility that the frequency of loan rate adjustments depends on the policy rate behavior. Finally, stickiness in deposit rates as well as in loan rates should also be considered, since many previous studies have reported that deposit rates are also sticky. Although this paper treats deposit rates as equivalent to the policy rate, the relaxation of this assumption may affect the desirability of policy rate smoothing.

Appendix 1. Derivation of the Demand for Funds

From (4), (5), and (7), labor wage $W_t(i)$ can be expressed as

$$\begin{aligned} W_t(i) &= \xi_t^{\sigma-1} P_t C_t^{\sigma} L_t^{\omega}(i) \\ &= \xi_t^{\sigma-1} P_t C_t^{\sigma} \left(\frac{Z_t(i)}{A_t}\right)^{\omega} \\ &= \xi_t^{\sigma-1} P_t C_t^{\sigma} \left(\frac{P_t^z(i)}{P_t^z}\right)^{-\omega\theta_z} \frac{Y_t^{\omega}(V_t^y)^{\omega}}{A_t^{\omega}} \\ &\equiv \Xi_t P_t^z(i)^{-\omega\theta_z}. \end{aligned} \tag{20}$$

Inserting this equality into (6) gives

$$\begin{aligned} P_t^z(i) &= \frac{\Xi_t R_t^i P_t^z(i)^{-\omega\theta_z}}{\bar{R} A_t} \\ &= \left(\frac{\Xi_t R_t^i}{\bar{R} A_t}\right)^{\frac{1}{1+\omega\theta_z}}. \end{aligned} \tag{21}$$

Therefore, the amount of funds demanded by intermediate-goods firm i, $W_t(i)L_t(i)$, leads to

$$\begin{aligned} W_t(i)L_t(i) &= \xi_t^{\sigma-1} P_t C_t^{\sigma} L_t(i)^{1+\omega} \\ &= \xi_t^{\sigma-1} P_t C_t^{\sigma} \left(\frac{P_t^z(i)}{P_t^z}\right)^{-\theta_z(1+\omega)} \frac{Y_t^{1+\omega}\left(V_t^y\right)^{1+\omega}}{A_t^{1+\omega}} \\ &\equiv \left(R_t^i\right)^{\frac{-(1+\omega)\theta_z}{1+\omega\theta_z}} \Lambda_t, \end{aligned}$$

where $\Lambda_t \equiv \xi_t^{\sigma-1} P_t C_t^{\sigma} A_t^{\frac{(1+\omega)(\theta_z-1)}{1+\omega\theta_z}} (P_t^z)^{(1+\omega)\theta_z} (Y_t V_t^y)^{1+\omega} \Xi_t^{\frac{-(1+\omega)\theta_z}{1+\omega\theta_z}} \times \bar{R}^{\frac{(1+\omega)\theta_z}{1+\omega\theta_z}}$.

Appendix 2. Proof of Proposition 1

Given the definition of long-term interest rates, the newly adjusted loan rate, $\tilde{r}_t$, should be expressed as

$$\begin{aligned}\tilde{r}_t &= (1-q\beta)E_t[r_t + q\beta r_{t+1} + (q\beta)^2 r_{t+2} + \ldots] \\ &= \left(\sum_{s=0}^{\infty}\delta_s\right)^{-1} E_t\left[r_t + \delta_1\left(\frac{r_t+\beta r_{t+1}}{1+\beta}\right)\right. \\ &\quad \left. + \delta_2\left(\frac{r_t+\beta r_{t+1}+\beta^2 r_{t+2}}{1+\beta+\beta^2}\right) + \ldots\right].\end{aligned}$$

Accordingly, comparing the coefficients on r_t and on $E_t r_{t+1}$, respectively, yields

$$1-q\beta = \left(\sum_{s=0}^{\infty}\delta_s\right)^{-1}\left(1 + \frac{\delta_1}{1+\beta} + \frac{\delta_2}{1+\beta+\beta^2} + \ldots\right)$$

and

$$(1-q\beta)q\beta = \left(\sum_{s=0}^{\infty}\delta_s\right)^{-1}\left(\frac{\delta_1\beta}{1+\beta} + \frac{\delta_2\beta}{1+\beta+\beta^2} + \ldots\right).$$

Summarizing these two equations leads to

$$\left(\sum_{s=0}^{\infty}\delta_s\right)^{-1} = (1-q)(1-q\beta).$$

Therefore, comparing the coefficients on $E_t r_{t+k}$ and on $E_t r_{t+k+1}$, respectively, yields

$$(1-q\beta)(q\beta)^k = (1-q)(1-q\beta)\left(\frac{\delta_k\beta^k}{\sum_{s=0}^{k}\beta^s} + \frac{\delta_{k+1}\beta^k}{\sum_{s=0}^{k+1}\beta^s} + \ldots\right)$$

and

$$(1-q\beta)(q\beta)^{k+1} = (1-q)(1-q\beta)\left(\frac{\delta_{k+1}\beta^{k+1}}{\sum_{s=0}^{k+1}\beta^s} + \frac{\delta_{k+2}\beta^{k+1}}{\sum_{s=0}^{k+2}\beta^s} + \dots\right)$$

for all $k \geq 1$. Summarizing these two equations, we have

$$\delta_k = q^k \sum_{s=0}^{k} \beta^s,$$

which is the desired result.

Next, let us derive a condition that attains $\delta_{k+1} < \delta_k$ for all $k \geq 0$, where $\delta_0 \equiv 1$. From the expression of δ_k, we have

$$\begin{aligned}\delta_k - \delta_{k+1} &= q^k \sum_{s=0}^{k} \beta^s - q^k \sum_{s=0}^{k} \beta^s \\ &= \frac{q^k}{1-\beta}[1-\beta^{k+1} - q(1-\beta^{k+2})].\end{aligned}$$

Then, the following condition has to be satisfied for this to be positive for all $k \geq 0$:

$$q < \frac{1-\beta^{k+1}}{1-\beta^{k+2}} \equiv \varpi(k), \text{for all } k \geq 0.$$

At this point, note that $\partial\varpi(k)/\partial k = -\beta^{k+1}(1-\beta)\ln\beta/(1-\beta^{k+2})^2 > 0$, and $\varpi(0) = (1+\beta)^{-1}$. Therefore, the condition $\delta_{k+1} < \delta_k$ is satisfied for all $k \geq 0$ if and only if $q(1+\beta) < 1$.

Appendix 3. Derivation of $p_t^z - p_t = r_t^l + (\sigma + \omega)x_t$

From the household's optimality condition (4), it is obvious that

$$w_t(i) - p_t = \sigma y_t + \omega l_t(i) - (1-\sigma)\hat{\xi}_t.$$

Using this equality and the pricing rule of intermediate-goods firms, (6), we have

$$p_t^z - p_t - r_t^l + a_t = \sigma y_t + \omega l_t - (1-\sigma)\hat{\xi}_t. \tag{22}$$

A linear approximation of (7) leads to $z_t = y_t$, and the production function (5) implies $l_t = z_t - a_t$. Notice that, as is shown in Galí and Monacelli (2005), the term $v_t^y = d \log \int_0^1 \left(\frac{P_t(j)}{P_t}\right)^{-\theta} dj$ is of second order. It follows that

$$l_t = y_t - a_t.$$

Inserting this condition into equation (22) yields

$$p_t^z - p_t = r_t^l + (\sigma + \omega)\left[y_t - \left(\frac{1+\omega}{\sigma+\omega}\right)a_t - \left(\frac{1-\sigma}{\sigma+\omega}\right)\hat{\xi}_t\right]. \qquad (23)$$

Let us define z_t^f as the flexible-price equilibrium of an arbitrary variable z_t. It follows from (23) that

$$y_t^f + \left(\frac{1}{\sigma+\omega}\right)r_t^{lf} = \left(\frac{1+\omega}{\sigma+\omega}\right)a_t + \left(\frac{1-\sigma}{\sigma+\omega}\right)\hat{\xi}_t \equiv \tilde{y_t}^f.$$

Let us call $\tilde{y_t}^f$ the quasi-flexible-equilibrium output. This relation states that the sum of the flexible-equilibrium output and the flexible-equilibrium loan rate can be expressed in terms of a productivity shock and a preference shock. By defining $x_t \equiv y_t - \tilde{y_t}^f$, equation (23) can be rewritten as

$$p_t^z - p_t = r_t^l + (\sigma + \omega)x_t.$$

Appendix 4. Derivation of Equation (17)

A second-order approximation of $u(C_t)$ and $v(L_t(i))$, respectively, leads to

$$u(C_t;\xi_t) = u'\bar{C}\left[c_t + \frac{1}{2}(1-\sigma)c_t^2 + \frac{u'_{c\xi}}{u'_c}c_t\hat{\xi}_t\right] + t.i.p. \qquad (24)$$

$$v(L_t(i)) = v'\bar{L}\left[l_t(i) + \frac{1}{2}(1+\omega)l_t^2(i)\right] + t.i.p.$$

From the relation $l_t(i) = z_t(i) - a_t$, the latter can be written as

$$v(L_t(i)) = v'\bar{L}\left[z_t(i) + \frac{1}{2}(1+\omega)z_t^2(i) - (1+\omega)a_t z_t(i)\right] + t.i.p.$$

It immediately follows that

$$\int_0^1 v(L_t(i))di = v'\bar{L}\left\{[1-(1+\omega)a_t]\int_0^1 z_t(i)di \right.$$
$$\left. +\frac{1}{2}(1+\omega)\int_0^1 z_t^2(i)di\right\} + t.i.p. \qquad (25)$$

It turns out that the disutility of labor depends on $\int_0^1 z_t(i)di$ and $\int_0^1 z_t^2(i)di$. We focus on these expressions in turn.

From a second-order approximation of the definition of intermediate-goods price index, we have

$$\int_0^1 p_t^z(i)di = p_t^z - \left(\frac{1-\theta_z}{2}\right) var_i p_t^z(i).$$

Inserting this into a linearized version of equation (7) yields

$$\int_0^1 z_t(i)di = \frac{\theta_z(1-\theta_z)}{2} var_i p_t^z(i) + y_t + v_t^y. \qquad (26)$$

Thus, the total intermediate goods can be expressed as a function of the variance of individual prices, $var_i p_t^z(i)$.

Next, we show that the variance of intermediate-goods price can be written in terms of the variance of loan rates. Based on equation (6), we can establish that

$$\begin{aligned} p_t^z(i) &= r_t^i + w_t(i) - a_t \\ &= r_t^i - a_t - (1-\sigma)\hat{\xi}_t + p_t + \sigma c_t + \omega l_t(i) \\ &= r_t^i - a_t - (1-\sigma)\hat{\xi}_t + p_t + \sigma c_t + \omega\big[-\theta_z\big(p_t^z(i) - p_t^z\big) + y_t - a_t\big], \end{aligned}$$

where the second and the third equalities follow from (4) and (7), respectively. Since only $p_t^z(i)$ and r_t^i are dependent on index i, the variance of $p_t^z(i)$ leads to

$$var_i p_t^z(i) = \left(\frac{1}{1+\omega\theta_z}\right)^2 var_i r_t^i. \qquad (27)$$

Meanwhile, the term $\int_0^1 z_t^2(i)di$ can be rewritten as follows:

$$\begin{aligned}\int_0^1 z_t^2(i)di &= var_i z_t(i) + \left[\int_0^1 z_t(i)di\right]^2 \\ &= var_i z_t(i) + y_t^2 \\ &= \theta_z^2 var_i p_t^z(i) + y_t^2 \\ &= \left(\frac{\theta_z}{1+\omega\theta_z}\right)^2 var_i r_t^i + y_t^2,\end{aligned} \tag{28}$$

where the last line comes from (27).

Therefore, from equations (25)–(28), the disutility of labor leads to

$$\begin{aligned}\int_0^1 v(L_t(i))di = \frac{v'\bar{L}}{2}\Big\{&(1+\omega)\big(y_t^2 - 2a_t y_t\big) + 2y_t + 2v_t^y \\ &+ \left(\frac{\theta_z}{1+\omega\theta_z}\right) var_i r_t^i\Big\} + t.i.p.\end{aligned}$$

Since $u'\bar{C} = v'\bar{L}$ holds in the efficient steady state, the utility of the representative household can be expressed as

$$\begin{aligned}U_t &= u(C_t;\xi_t) - \int_0^1 v(L_t(i))di \\ &= \frac{v'\bar{L}}{2}\Big\{(1-\sigma)y_t^2 + \frac{2u'_{c\xi}}{u'_c}\hat{\xi}_t y_t - (1+\omega)\big(y_t^2 - 2a_t y_t\big) \\ &\quad - 2v_t^y - \left(\frac{\theta_z}{1+\omega\theta_z}\right) var_i r_t^i\Big\} + t.i.p. \\ &= -\frac{v'\bar{L}}{2}\Big\{(\sigma+\omega)\left[y_t^2 - 2\left(\frac{1+\omega}{\sigma+\omega}\right)a_t y_t - 2\left(\frac{u'_{c\xi}/u'_c}{\sigma+\omega}\right)\hat{\xi}_t y_t\right] \\ &\quad + 2v_t^y + \left(\frac{\theta_z}{1+\omega\theta_z}\right) var_i r_t^i\Big\} + t.i.p.\end{aligned}$$

Note that v_t^y can be approximated as $(\theta/2)var_j p_t(j)$. In addition, the specification of the total utility function yields $u'_{c\xi}/u'_c = 1-\sigma$ and $v'\bar{L} = \bar{L}^{1+\omega}$, which establishes equation (17).

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The Role of the Chairman in Setting Monetary Policy: Individualistic vs. Autocratically Collegial MPCs*

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This paper models the role of the Chairman in the decision making of individualistic and autocratically collegial monetary policy committees, assuming that uncertainty about the optimal interest rate causes policymakers' views to differ and that they are unable to communicate their opinions perfectly. The Chairman's ability to moderate the discussion and his economic skills—and, in an autocratically collegial committee, the authority arising from his position—impact the path of interest rates and the distribution of votes. Simulations suggest that his influence on the quality of policy itself is limited and that interest rate setting is only slightly worse in an autocratically collegial setup. The Chairman's main impact is to help build consensus in the committee, which enhances the credibility of monetary policy.

JEL Codes: D81, E52.

1. Introduction

In a growing number of countries, monetary policy is set by committee.[1] It is sometimes argued, certainly in the case of the Federal Open Market Committee (FOMC) in the United States, that the Chairman of the committee exerts a disproportionate influence on the

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[1]Fry et al. (2000) study the policy frameworks of eighty-eight central banks across the world and report that in seventy-nine, policy is set by committee.

interest rate decision (see, e.g., former Vice Chairman of the Board of Governors Blinder 1999 and former Governor Meyer 2004). There are at least three potential reasons for this. First, the Chairman may be particularly capable in the monetary policy area—as, e.g., Paul Volcker or Alan Greenspan have been argued to be (see Blinder and Reis 2005 and Goodfriend and King 2005). Of course, it would be optimal for such Chairmen to exert a large influence on policy decisions. Second, the Chairman guides the discussion and summarizes the different views in the committee, and may thereby help shape the outcome of the deliberations. We refer to these two sets of skills as the Chairman's "economic" and "moderating" abilities. Third, the Chairman's position per se may in certain institutional setups lend his views extra weight.[2]

Blinder (2004, 2007) distinguishes between three types of monetary policy committees (MPCs): individualistic, genuinely collegial, and autocratically collegial. In an individualistic committee, members vote according to their views and communicate their opinions to the public, so that the Chairman's views carry no special weight. Blinder argues that the Bank of England's MPC fits this description. A genuinely collegial committee, of which Blinder argues the ECB's Governing Council is an example, reaches a consensus decision and talks with one—the Chairman's—voice to the public.[3] While the Chairman's position does not make him more influential than any other committee member in the interest rate decision, he gets the most public attention. In an autocratically collegial committee, such as (according to Blinder) the FOMC under Chairman Greenspan, the committee as a rule adopts the level of interest rates preferred by the Chairman.[4] However, the committee "serve[s] as a kind of check on the [C]hairman" (Blinder 2007, 115) in that it may block interest rate proposals that deviate much from the majority's view.

[2]Kuttner and Posen (2007) show that financial markets react to the appointment of new central bank heads, which suggests that it matters for policy who the Chairman is.

[3]The Statute of the European System of Central Banks and of the European Central Bank allows the Governing Council to reach decisions by vote. (The Statute is available at http://www.ecb.int/ecb/pdf/orga/escbstatutes_en.pdf.)

[4]The *International Herald Tribune* refers to Chairman Greenspan's style of decision making as "strong-arm[ing] his colleagues" (Uchitelle 2008, 1).

The Chairman tries to avoid this outcome, since it could be damaging to his and the committee's credibility, and therefore only makes proposals that can gain a clear majority of committee members.[5] Searching to achieve consensus in this way matters for monetary policy: Janet Yellen, former Governor and currently President of the Federal Reserve Bank of San Francisco, emphasizes that consensus in the committee is crucial for "the credibility, legitimacy, and likely effectiveness of monetary policy" (Yellen 2005, 2–3).[6]

The present paper provides a first model for the Chairman's impact on interest rate setting by committee. Section 2 discusses the empirical, experimental, and theoretical literature on MPCs. Section 3 describes the model, which assumes that policymakers are uncertain about the level of the optimal interest rate. During their deliberations, committee members signal their private information on the optimal rate, subject to communication errors. We assume that the importance of these depends on the Chairman's moderating skills.[7] Section 4 simulates the model and assesses the Chairman's impact on policy decisions in individualistic and autocratically collegial committees. We show that the baseline specifications for these two types of committee fit the outcomes of the MPC's voting at the Bank of England and of the FOMC, respectively. We then compare the two committees and vary the key parameters in the model to assess their effect on the path of interest rates and the distribution of votes. The Chairman's impact on the path of interest rates is found to be smaller in an individualistic committee than in an autocratically collegial committee, which renders the latter's interest rate setting inferior. However, while the exact path of monetary

[5]Meyer (2004, 50) refers to the phenomenon of "musical chairs" in the FOMC—i.e., the committee members' understanding that more than two dissents represent a rebellion against the Chairman—and states that "the Chairman is expected to resign if the Committee rejects his policy recommendation." For simplicity we do not model any strategic behavior arising from this tradition, nor do we take into account that there are governors and Bank presidents on the FOMC and that the former are arguably more likely to agree with the Chairman.

[6]See also Blinder and Wyplosz (2004) and Chappell, McGregor, and Vermilyea (2004). Sager and Gastil (2006) discuss the advantages of consensus over majority decisions from a psychological perspective.

[7]Since genuinely collegial committees make decisions by consensus, there is no voting record that would allow us to test the performance of our model. We therefore do not consider this type of MPC.

policy chosen by the two kinds of committee differs frequently, the difference in quality as measured by the deviation from the optimal interest rate is small. Independently of the type of committee, the more skilled as a moderator and as a monetary policy expert the Chairman is, the more there is agreement in the committee. Agreement is most common in autocratically collegial committees in which the Chairman attaches great importance to consensus and in which policymakers are keen to adopt his views. Section 5 concludes.

We note from the outset that the paper abstracts from strategic behavior that may arise if the policy goals differ between individuals or if information is costly. Under such circumstances, coalitions may be formed, the composition of which may depend on political considerations or the state of the economy. However, while there is a growing literature on strategic behavior in MPCs, policymakers themselves do not seem to believe that such strategic considerations matter much. To quote Yellen again: "In fact, I think FOMC members behave far less individualistically and strategically than assumed in some of the models" (Yellen 2005, 1).[8]

2. Brief Survey of the Literature

Next we provide a brief and highly selective overview of the empirical, experimental, and theoretical literature on MPCs. The empirical literature focuses on evidence from the FOMC. Belden (1989) analyzes the voting pattern of the twelve FOMC members and finds that Bank presidents dissent more frequently than Board members by favoring tighter policy.[9] Chappell, Havrilesky, and McGregor (1997) estimate reaction functions for individual FOMC members and also find that some systematically voted for tighter (looser) monetary policy than others. This evidence suggests that policymakers may hold different views of the optimal level of inflation and/or potential output that make them disagree about the desirable level of

[8] In private communication with the author, Alan Blinder stated that there is very little strategic behavior in the FOMC. Bank of England Governor Mervyn King argues that the MPC members in the United Kingdom share the same policy objective and that "it makes no sense" to label them "doves" or "hawks" (King 2002, 5). He moreover sees little evidence for the formation of voting blocks.

[9] See also Havrilesky and Gildea (1991).

interest rates. Meade and Sheets (2005) establish that regional economic conditions impact the voting pattern of members, which indicates that policymakers display a "home bias" that may arise either because they care more about local than national economic conditions or because they rely overly on regional data in interpreting the countrywide economy.

The empirical literature on other MPCs is limited. Bhattacharjee and Holly (2006), Gerlach-Kristen (2004), and Harris and Spencer (2007) study the voting record of the committee at the Bank of England and detect differences in the voting patterns of internal and external members. Andersson, Dillén, and Sillen (2001) examine the voting record of the Bank of Sweden and find that the votes of dissenting policymakers are informative about future changes in policy, whereas Fujiki (2005) shows that the voting record in Japan does not forecast policy changes well.

A number of descriptive empirical papers study the role of the Chairman. Blinder (1999) states that the Chairman is "more equal" than the other FOMC members, and Chappell, McGregor, and Vermilyea (2004) estimate that his impact on policy decisions corresponds to a voting weight of 40 to 50 percent. Romer and Romer (2003) analyze minutes and transcripts of FOMC meetings and find that even Chairmen who are not necessarily seen as especially able monetary policymakers are influential. Chappell, McGregor, and Vermilyea (2005) document that Chairman Greenspan spoke longer in FOMC meetings than the other members, and Meyer (2004) describes Chairman Greenspan's habit of summarizing the staff's analysis and giving his policy recommendation at the beginning of the policy discussion.[10]

The experimental literature on interest rate setting by committee is limited. Blinder and Morgan (2005) and Lombardelli, Proudman, and Talbot (2002) find that committees make better decisions than individual policymakers. They also present evidence indicating that deliberation before voting improves monetary policy. More interestingly for the paper at hand, Blinder and Morgan (2006) study the impact of committee size and the role of the Chairman in interest rate setting. They find that larger groups perform slightly better

[10] Chairman Bernanke, by contrast, tends to speak last (*Wall Street Journal* 2006).

than smaller groups and that monetary policy does not appear to improve if there is a designated leader. The latter result clearly surprises the authors, but it is compatible with the finding below that the Chairman, both in an individualistic and an autocratically collegial committee, hardly affects the quality of monetary policy. Unfortunately, Blinder and Morgan did not record whether there was a difference in the frequency of dissents, which this paper argues should be the case.

The theoretical literature on MPCs focuses on why policymakers disagree about the level of interest rates. Four arguments have been advanced. First, policymakers may disagree because they have different views about the inflation objective (see Mihov and Sibert 2002 and Waller 1989 and the related popular discussion on "hawks" and "doves"). Second, members might hold different views because some are more skilled than others and therefore have a better sense of the appropriate level of interest rates (see Gersbach and Hahn 2001). Gerlach-Kristen (2006) shows that in this situation it is optimal to attach a greater weight to the more skilled members' views. Third, different opinions regarding the appropriate stance of policy may arise because members rely on different data sets. Aksoy, De Grauwe, and Dewachter (2002) and von Hagen and Süppel (1994) discuss how a national perspective in analyzing data in the Governing Council of the ECB can lead to disagreement about optimal policy. Fourth, members may disagree for strategic reasons, such as to raise their profile with the public. Sibert (2006) discusses the fact that information cascades, in which the views offered early on in the deliberation have a disproportionate impact on the final decision, occur only if the exchange of information is costly or if members hold different preferences.[11]

Policy discussions in the committee and the Chairman's role in them have only recently started to attract attention in the theoretical literature. Gerlach-Kristen (2006), Spencer (2005), and Weber (2007) model policy deliberations in an MPC but disregard the role of the Chairman. Berk and Bierut (2007) consider a Chairman whose economic skills are greater than those of the rest of the committee. Favarque, Matsueda, and Méon (2007) and Riboni and Ruge-Murcia

[11] The literature on strategic voting has been growing rapidly; Fujiki (2005) and Gerling, et al. (2005) provide comprehensive surveys.

(2007) present a strategic voting game in which the Chairman makes interest rate proposals. A committee member accepts this proposal if it is closer to his own view of optimal policy than the current level of the interest rate. A major weakness of these models is that they do not capture the idea that in an autocratically collegial committee policymakers go along with the Chairman's proposal—rather than voting for the rate they personally favor—as long as it is not too different from their own view.

In sum, the empirical literature shows that MPC members hold different views regarding the appropriate path of monetary policy and that the Chairman is seen as having a disproportionate influence on the interest rate decision. The experimental literature suggests that committee decisions are superior to those of a single policymaker and that discussions in the committee prior to the vote improve policy outcomes, but that appointing a Chairman does not. The theoretical literature, finally, provides explanations for differences in views in the committee and explores the agenda-setting power of the Chairman. However, it does not capture the committee's role as a "check on the Chairman." This is the gap we attempt to fill.

3. The Model

3.1 Basic Assumptions

We assume that there is an optimal interest rate, i_t^*, which is observed only imprecisely. The goal of monetary policy is to set the interest rate, i_t, as close as possible to i_t^*. We let the optimal rate follow an autoregressive process of second order since empirical estimates of reaction functions of the Taylor type suggest that monetary policy reacts to inflation and the output gap, and since there is broad evidence that the output gap follows an autoregressive process of second order (see, e.g., Watson 1986). Thus,

$$i_t^* = c + \rho_1 i_{t-1}^* + \rho_2 i_{t-2}^* + u_t, \tag{1}$$

with $u_t \sim N(0, \sigma^2)$. As is common in the empirical literature estimating output gaps using state-space models, we let $\rho_1 > 1$, $\rho_2 < 0$, and $\rho_1 + \rho_2 < 1$.

The model assumes that there are n policymakers who are rational, share the same goals, and do not behave strategically. We furthermore assume that policymakers have difficulties observing i_t^*—for instance, because inflation and output-gap data are measured imprecisely (see Orphanides 2001) or because members interpret the same data differently. To capture these difficulties in a simple way, we assume that policymaker j's perception or "observation" of i_t^* is given by

$$i_{j,t} = i_t^* + v_{j,t}, \tag{2}$$

with $v_{j,t} \sim N(0, s^2)$. Each policymaker thus receives the sum of a common signal, i_t^*, and private information, $v_{j,t}$, which is white noise. We assume that these observation errors are uncorrelated across policymakers and over time, and that the variance s^2 is identical for all policymakers except possibly for the Chairman.

3.2 Interest Rate Setting by Committee

We now turn to the question of how the committee sets policy on the basis of the individual observations of i_t^*. We first examine an individualistic committee and then turn to an autocratically collegial committee.

3.2.1 An Individualistic Committee

In all committees, policymakers deliberate before deciding on the level of interest rates and thereby pool the private information available to them. As in practice, we assume that in the deliberations members reveal their general impression of economic conditions but do not state their views of the exact, numerical level of the optimal interest rate. Since there is consequently some uncertainty about committee member j's view of the optimal rate, we assume that his signal of i_t^* is received by policymaker k subject to noise, so that policymaker k's understanding of $i_{j,t}$ is given by

$$\widetilde{i}_{kj,t} = i_t^* + v_{j,t} + \widetilde{w}_{kj,t},$$

where $\widetilde{w}_{kj,t}$ is the communication error, with $\widetilde{w}_{kj,t} \sim N(0, \alpha s^2)$ and $\alpha > 0$. We assume that $v_{j,t}$ and $\widetilde{w}_{kj,t}$ are uncorrelated, so that the variance of their sum is given as $(1 + \alpha)s^2$. Thus, policymaker k's perception of $i_{j,t}$ is distributed around the optimal interest rate with a variance that is by α larger than that of his own. Below we refer to α as a measure of the committee's "communication difficulties."

Since the optimal interest rate is serially correlated, policymaker j's best assessment of i_t^* draws not only on his own $i_{j,t}$ and his understanding of his colleagues' current observations of i_t^* but also on past observations. The optimal way to assess the current value of i_t^* involves a signal extraction mechanism based on the Kalman filter (see, e.g., Hamilton 1994). We derive policymaker j's assessment of i_t^* formally in the appendix.

Differences in view persist even after deliberation since information aggregation in the committee is imperfect.[12] We assume that the committee sets interest rates in steps of 25 basis points, so that each policymaker j votes for the step of the policy interest rate i_t that is closest to his assessment of the optimal rate. The level of i_t favored by the majority of members is adopted as policy.

We assume that the Chairman's special status in an individualistic committee arises from two sources. The first is his role as a moderator of the policy discussions: he may be able to reduce the communication difficulties within the committee by structuring the discussions, by asking clarifying questions if a member's statements are difficult to interpret, and by summarizing the different views in the committee. Formally, the Chairman reduces the variance of the communication error, so that policymaker k's understanding of policymaker j's observation changes to

$$i_{kj,t} = i_t^* + v_{j,t} + w_{kj,t}, \tag{3}$$

with $w_{kj,t} \sim N[0, (\alpha/\beta)s^2]$ and $\beta \geq 1$, where β captures the Chairman's "moderating skills."

The second source of the Chairman's influence is his ability as a monetary policymaker. We model this by assuming that his observation error of the optimal interest rate is smaller than s^2. Thus,

$$i_{C,t} = i_t^* + v_{C,t},$$

where C denotes the Chairman and where $v_{C,t} \sim N(0, s^2/\gamma)$, with $\gamma \geq 1$. We let γ capture the Chairman's "economic skills." The larger

[12]This notion is supported by the empirical fact that there are dissents in monetary policy votes.

γ is, the better the Chairman is at judging i_t^*, which increases the optimal weight attached to his view on i_t^* (Gerlach-Kristen 2006).

3.2.2 An Autocratically Collegial Committee

To model an autocratically collegial committee, we assume that the Chairman discusses (perhaps individually, perhaps in premeetings) with the other committee members before the policy decision their views regarding the optimal stance of policy. This allows him to form a view of what level of the interest rate the individual members prefer and how large a deviation from that interest rate they are willing to accept.

Since the Chairman would like to set the policy rate equal to $i_{C,t}$ (rounded to the closest 25 basis points) but does not have the statutory power to do so, he assesses how many committee members would vote for or against him if he proposed $i_{C,t}$ as the policy rate. We assume that policymakers display a certain tolerance toward the Chairman's views in the sense that they dissent only if their own view differs considerably from his proposal (see Blinder 2007 and Meyer 2004). More technically, we assume that policymaker j dissents if the Chairman's proposal lies outside a tolerance interval around j's own assessment of i_t^* that is given by

$$\pm\Phi^{-1}_{1-(1-\tau)/2} * \sqrt{P_j(1,1)},$$

where Φ^{-1} denotes the inverse cumulative normal distribution, τ denotes policymaker j's tolerance level, and $P_j(1,1)$ denotes the first element of his forecast error variance matrix, which is derived in the appendix and which in the steady state corresponds to the mean squared error of policymaker j's assessment of i_t^*. We assume that all committee members have the same tolerance level τ. In the baseline simulation we set $\tau = 99$ percent, so that, given the other parameter assumptions, policymaker j votes with the Chairman if the latter's proposal lies within a range of ± 23 basis points of j's assessment of i_t^*. Committee members' tendency to go along with the Chairman's proposals implies that his influence on interest rate decisions is larger than optimal.

We furthermore assume that the Chairman would like to achieve a majority $\mu > 0.5$ to capture the fact that he values the "appearance of unity" (Blinder 2007). If he expects a majority of μ or larger to support the rate he personally thinks optimal, he proposes this rate. If he expects a smaller majority, he adjusts his proposal toward, but not necessarily all the way to, the rate preferred by the majority.

To clarify this mechanism, assume that there are three MPC members: the Chairman and members J and K. Suppose that the Chairman thinks i_t^* equals 1.70 percent, that J thinks it is 2.15 percent, and that K thinks it is 2.30 percent. Let J's and K's tolerance interval be equal to ± 20 basis points and assume for simplicity that the Chairman knows all this with certainty. In an individualistic committee, the Chairman votes for 1.75 percent, J and K vote for 2.25 percent, and the policy rate is set equal to 2.25 percent. In an autocratically collegial committee, if the Chairman proposes a rate of 1.75 percent, J and K vote for 2.25 percent, so that the policy rate is again set equal to 2.25 percent. If the Chairman proposes 2.00 percent instead, this value lies within J's tolerance interval of [1.95%, 2.35%] but outside K's. Thus, the Chairman and J vote for 2.00 percent and K votes for 2.25 percent. The policy rate is set equal to 2.00 percent.

We next assess the impact on monetary policy of the Chairman's moderating and economic skills and the effect of an individualistic versus an autocratically collegial setup by simulating the model.

4. Simulations

The assumption that interest rates are set in steps renders the model difficult to handle mathematically. We therefore simulate it using 10,000 draws. To choose realistic parameters, we attempt first to match the pattern of the monthly UK repo rate between June 1997, when the MPC began setting interest rates, and August 2007. Over that period, the policy repo rate followed an autoregressive process of second order with coefficients 1.35 and −0.37. The first column of table 1 presents the voting pattern for the Bank of England's MPC and the FOMC over the period. In 66.1 percent of all meetings, at least one of the nine MPC members in the United Kingdom cast a dissenting vote. On average, the majority was 85.5 percent, which corresponds to roughly eight MPC members. In the United States

Table 1. Characteristics of Interest Rate Decisions

	Actual Voting		**Baseline Simulation**	
	MPC	**FOMC**	**Individualistic**	**Autocratically Collegial**
Fraction of Meetings Ending with Disagreement	66.1%	23.2%	66.1%	23.3%
Average Size of Majority	85.5%	97.7%	84.2%	95.4%

Note: UK and U.S. voting patterns for June 1997 to August 2007. Results take fluctuations in committee size into account. Simulations (10,000 draws) assuming $\rho_1 = 1.95$, $\rho_2 = -0.98$, $\sigma^2 = 0.001$, $s^2 = 0.05$, $n = 9$, $\alpha = 2.7$, $\beta = 1$, $\gamma = 1$, $\mu = 0.6$, and $\tau = 0.99$.

in the same period, only 23.2 percent of all FOMC decisions were not unanimous, and the average majority was 97.7 percent—much larger than in the United Kingdom.

In the baseline simulation, we set the autoregressive coefficients of the optimal interest rate equal to $\rho_1 = 1.95$ and $\rho_2 = -0.98$.[13] The variance σ^2 of the shocks affecting i_t^* is assumed to be 0.001, which implies that 95 percent of all monthly innovations lie in a range of ± 6 basis points. The variance s^2 of the observation error is given as 0.05, so that policymaker j's observation of the optimal rate lies with 95 percent probability in a range of ± 45 basis points of the true value of i_t^*. Discussions in the committee reduce this range to ± 18 basis points. We assume that $n = 9$ and that the committee's communication difficulties are captured by $\alpha = 2.7$. The Chairman is assumed not to possess special moderating or economics skills ($\beta = \gamma = 1$) and to aim for a majority of 60 percent ($\mu = 0.6$). Finally, we let committee member j cast a dissenting vote only if the interest rate proposed by the Chairman lies outside j's 99 percent tolerance interval.

[13] We choose (in absolute terms) large autoregressive coefficients because i_t^* is observed with an error, which biases the coefficient estimates for the policy rate toward zero. The autoregressive coefficients of the simulated policy interest rate are 1.31 and -0.34 for the baseline setup of an individualistic committee.

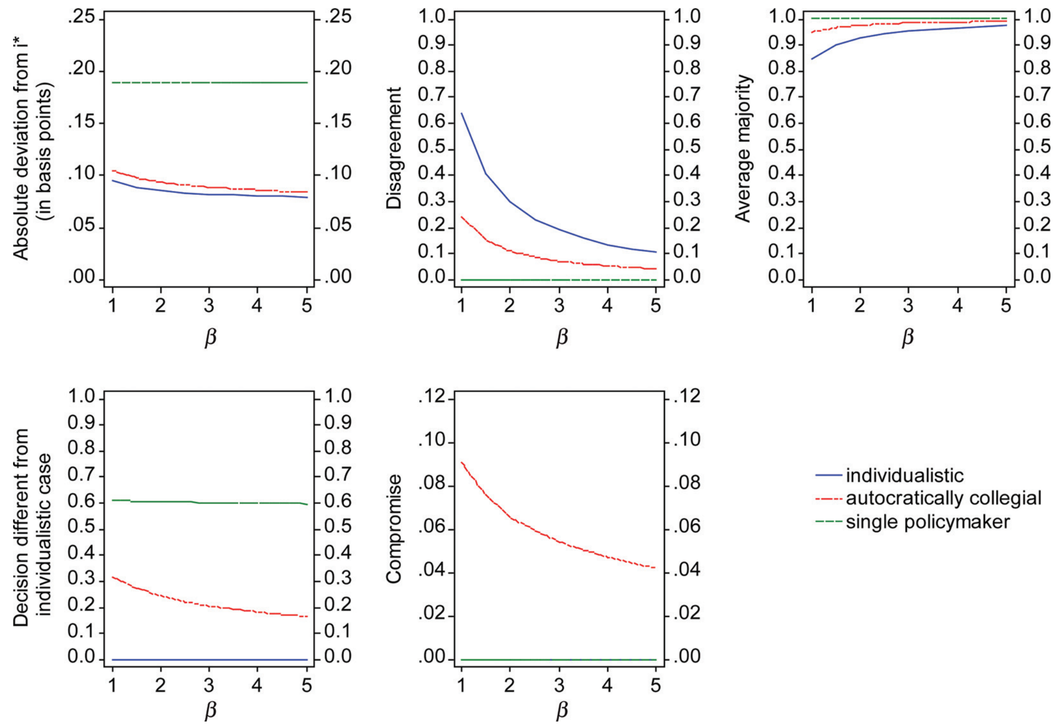
The second column of table 1 reports that under these assumptions, policy meetings of an individualistic committee end with at least one dissent in 66.1 percent of all cases and that the average majority is 84.2 percent. For an autocratically collegial committee, there is disagreement only in 23.3 percent of all cases, and the average majority is 95.4 percent. These values are very close to their empirical counterparts.

Next we study the impact of the Chairman's moderating and economics skills, his preferred majority, and policymaker j's tolerance toward the Chairman's proposals on five characteristics of interest rate setting: the quality of monetary policy, as measured by the average absolute deviation (in basis points) of the interest rate set by the committee from i_t^*; the frequency of dissents; the average majority by which decisions are taken; the frequency with which an autocratically collegial committee chooses a different interest rate than an individualistic committee; and the frequency with which the Chairman proposes a compromise, i.e., a rate that lies between his and the committee's view.

Figure 1 assesses the effect of the Chairman's moderating skills on these characteristics by varying β between 1 and 5. The first plot indicates that the interest rate deviates from i_t^* on average by roughly 10 basis points. The deviation decreases as the Chairman's moderating skills improve, since he helps committee members communicate their views more clearly. Consequently, their decision is based on better data and is therefore superior. It should be noted that the average absolute deviation from i_t^* is only about 1 basis point smaller for an individualistic than for an autocratically collegial committee, which is compatible with the experiments discussed in Blinder and Morgan (2006), who report no significant improvement in policy if a Chairman is appointed.[14] The second plot shows the frequency with which there is disagreement in the committee. The better the Chairman moderates the discussion, the less disagreement there is in both types of committee, and the third plot correspondingly indicates that the average majority increases. Since policymakers in an autocratically collegial committee cast dissenting

[14]We also plot the average absolute deviation of policy from i_t^* if policy is set by a single policymaker. This deviation equals 19 basis points and thus is roughly twice as large as for a committee.

Figure 1. Impact of Chairman's Moderating Skills (β between 1 and 5)



Note: Simulations (10,000 draws) assuming $\rho_1 = 1.95$, $\rho_2 = -0.98$, $\sigma^2 = 0.001$, $s^2 = 0.05$, $n = 9$, $\alpha = 2.7$, $\gamma = 1$, $\mu = 0.6$, and $\tau = 0.99$.

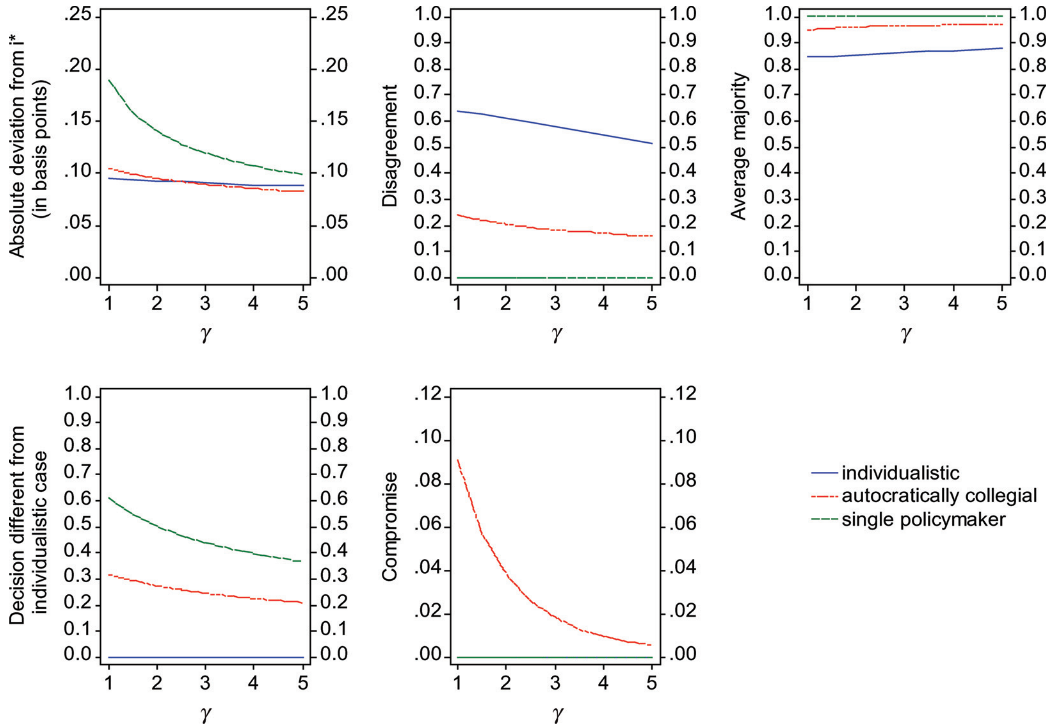
votes only if they disagree strongly with the Chairman, unanimity is, in this framework, more common. The fourth plot shows that the better a moderator the Chairman is, the more the decisions of an autocratically collegial MPC become similar to those of an individualistic committee. The reason for this is that the larger β is, the more clearly committee members understand one another's views and trust them. They are therefore less willing to accept any extreme position the Chairman may take, which forces him to propose more frequently the rate an individualistic MPC would set. The fifth plot shows that the Chairman's move toward the majority view also reduces the frequency with which he proposes a compromise.

Figure 2 assesses how monetary policy and the voting pattern change if the Chairman's economics skills improve. As we increase γ from 1 to 5, interest rate setting by an autocratically collegial committee comes closer to i_t^* than that by an individualistic committee. It thus appears that it is optimal to choose a framework in which the Chairman dominates if he is a much better monetary policymaker than the other committee members.[15] Nevertheless, the difference in performance between the two committee types is rather small for any γ. As γ rises and thus uncertainty about i_t^* diminishes, fewer policy meetings end with disagreement and the average majority increases. Decisions by an autocratically collegial committee become more similar to those of an individualistic MPC because an especially skilled Chairman also exerts a great influence in an individualistic committee. With his rising influence, there are also fewer compromises since the committee reaches virtually the same assessment of the optimal interest rate as the Chairman.

Figures 3 and 4 concentrate on the impact of μ and τ, which are parameters that matter only in an autocratically collegial committee. Figure 3 shows that as the Chairman's preferred majority μ is varied from 0.5 to 0.9, the quality of monetary policy, again, hardly is affected. As μ rises, policy by an autocratically collegial committee approaches that of an individualistic MPC, and there is less disagreement and a larger average majority since the Chairman attempts to meet the committee's preferences. Initially, there are, for the same reason, fewer compromises. For $\mu = 0.9$, however, the

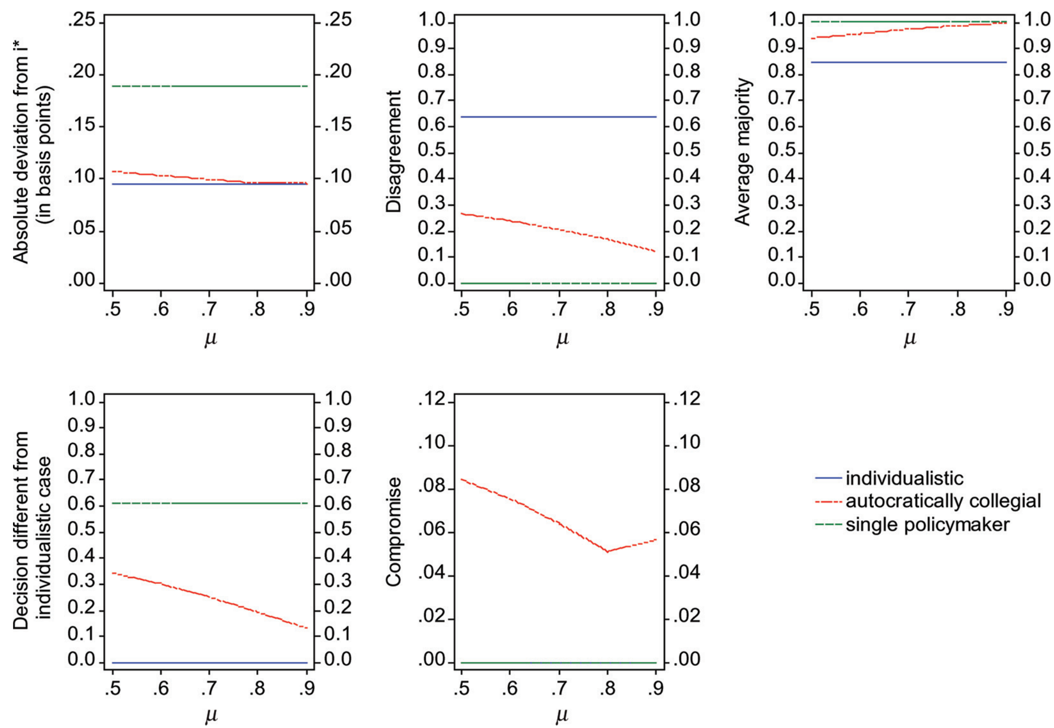
[15] This finding disappears if there are no communication difficulties, i.e., if $\alpha = 0$.

Figure 2. Impact of Chairman's Economics Skills (γ between 1 and 5)



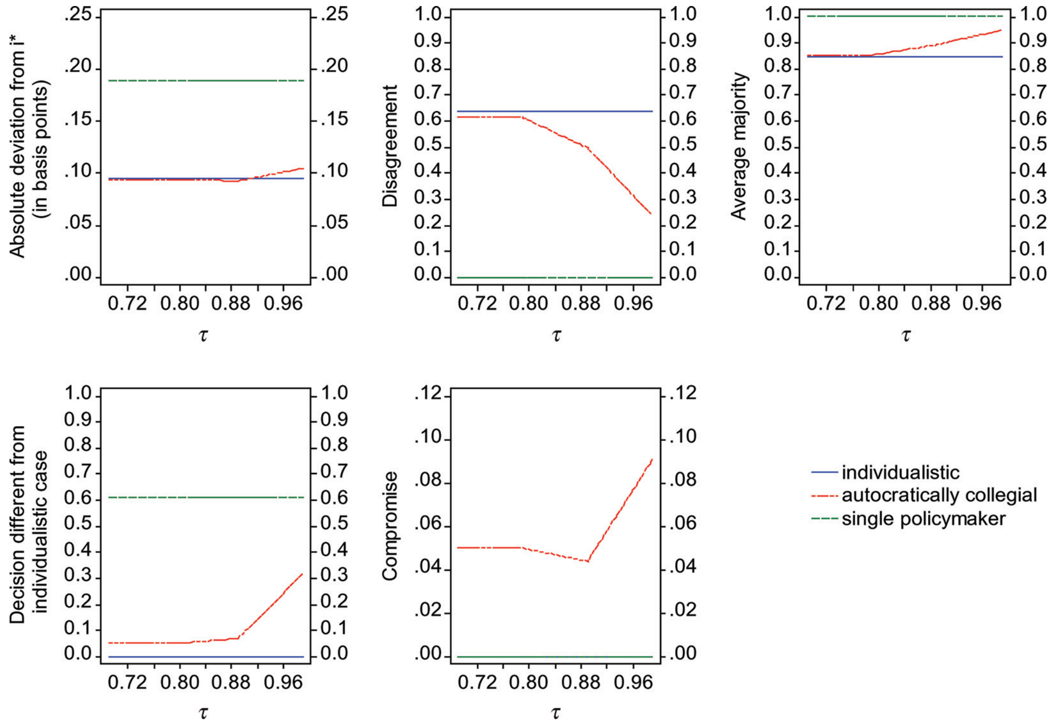
Note: Simulations (10,000 draws) assuming $\rho_1 = 1.95$, $\rho_2 = -0.98$, $\sigma^2 = 0.001$, $s^2 = 0.05$, $n = 9$, $\alpha = 2.7$, $\beta = 1$, $\mu = 0.6$, and $\tau = 0.99$.

Figure 3. Impact of Chairman's Preferred Majority (μ between 0.5 and 0.9)



Note: Simulations (10,000 draws) assuming $\rho_1 = 1.95$, $\rho_2 = -0.98$, $\sigma^2 = 0.001$, $s^2 = 0.05$, $n = 9$, $\alpha = 2.7$, $\beta = 1$, $\gamma = 1$, and $\tau = 0.99$.

Figure 4. Impact of Chairman's Preferred Majority (τ between 0.5 and 0.9)



Note: Simulations (10,000 draws) assuming $\rho_1 = 1.95$, $\rho_2 = -0.98$, $\sigma^2 = 0.001$, $s^2 = 0.05$, $n = 9$, $\alpha = 2.7$, $\beta = 1$, $\gamma = 1$, and $\mu = 0.6$.

number of compromises increases since a majority of more than 90 percent implies a committee with $n = 9$ unanimity. In this case, the Chairman also has to accommodate individuals with extreme views, which occasionally forces him to deviate from the rate preferred by the committee as a whole.

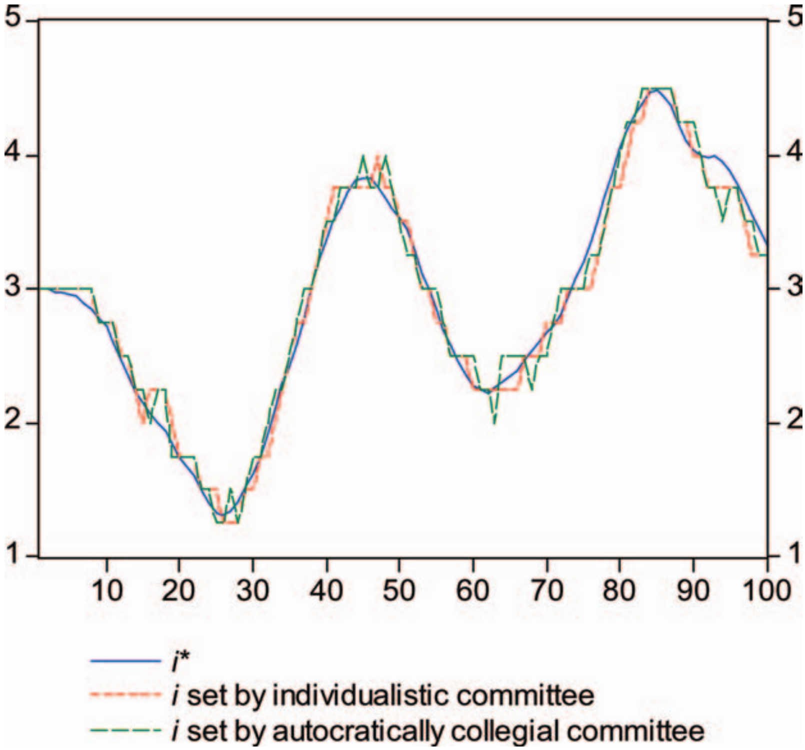
Figure 4 studies how committee members' willingness to accept extreme interest rate proposals affects interest rate setting. We vary τ between 0.69 and 0.99. If policymakers' tolerance level is low, they tend to vote against the Chairman if he proposes a rate that deviates from their view. This situation forces the Chairman to make compromises. If the committee becomes more tolerant toward the Chairman's opinions, there are fewer dissents and he starts proposing rates that he thinks are appropriate. As a consequence, there are fewer compromises at first. As the committee becomes even more tolerant, the number of compromises increases again since the Chairman deviates more often from the rate the committee would like to set. Policy therefore worsens.

Figures 1–4 suggest that the quality of monetary policy, as measured by the deviation from the optimal interest rate i_t^*, is similar under the two types of MPCs but that the rates set differ frequently. At first glance, this may appear contradictory. To gain an understanding of this finding, we plot in figure 5 the first 100 draws of the two baseline simulations. It can be seen that the rates set by an individualistic and an autocratically collegial committee differ frequently by 25 basis points and that the optimal interest rate tends to lie between these two rates, thus accounting for the fact that the deviations from i_t^* are comparable for both types of committee.

5. Conclusions

This paper models the role of the Chairman in the decision process of the MPCs. We argue that he may help reduce the uncertainty about the optimal level of interest rates and thus bring about larger majorities. We find that interest rate setting is generally worse in an autocratically collegial committee than in an individualistic committee, and that it is the worse, the less able a monetary policymaker the Chairman is, the less he is concerned with gaining the committee's support, and the more tolerant the committee is toward his views. However, the difference in quality is quantitatively small. The main

Figure 5. Simulated Interest Rate Paths



Note: Simulations for an individualistic and an autocratically collegial committee, assuming a mean of i_t^* of 3 percent, $\rho_1 = 1.95$, $\rho_2 = -0.98$, $\sigma^2 = 0.001$, $s^2 = 0.05$, $n = 9$, $\alpha = 2.7$, $\beta = 1$, $\gamma = 1$, $\mu = 0.6$, and $\tau = 0.99$.

advantage of an autocratically collegial committee is that decisions are made by larger majorities, which may contribute to the central bank's credibility.

The model presented in this paper does not address the issue of strategic behavior in MPCs. It is plausible that the voting strategies of inflation "hawks" and "doves" impact policy decisions, as may shifting coalitions of committee members. We also leave for future research the impact of political pressure on committee members' voting and the possibility that the information exchange during deliberation might be costly.

Appendix

To derive policymaker j's assessment of the optimal interest rate, it is useful to rewrite equations (1)–(3) in state-space form. Equation (1) can be written as

$$\begin{bmatrix} i_t^* \\ i_{t-1}^* \end{bmatrix} = \begin{bmatrix} c \\ 0 \end{bmatrix} + \begin{bmatrix} \rho_1 & \rho_2 \\ 1 & 0 \end{bmatrix} \begin{bmatrix} i_{t-1}^* \\ i_{t-2}^* \end{bmatrix} + \begin{bmatrix} u_t \\ 0 \end{bmatrix},$$

with

$$Q = \begin{bmatrix} u_t \\ 0 \end{bmatrix} \begin{bmatrix} u_t & 0 \end{bmatrix} = \begin{bmatrix} \sigma^2 & 0 \\ 0 & 0 \end{bmatrix}.$$

Equations (2) and (3) can be combined to yield member j's observation equation,

$$\begin{bmatrix} i_{j1,t} \\ \dots \\ i_{j,t} \\ \dots \\ i_{j(n-1),t} \\ i_{jC,t} \end{bmatrix} = H' \begin{bmatrix} i_t^* \\ i_{t-1}^* \end{bmatrix} + \begin{bmatrix} v_{1,t} \\ \dots \\ v_{j,t} \\ \dots \\ v_{n-1,t} \\ v_{C,t} \end{bmatrix} + \begin{bmatrix} w_{j1,t} \\ \dots \\ 0 \\ \dots \\ w_{j(n-1),t} \\ w_{jC,t} \end{bmatrix}, \quad (4)$$

where we assume for convenience that the Chairman's view of the optimal interest rate is listed last. H denotes a $2 \times n$ vector with ones in the first row and zeros in the second. The variance-covariance matrix of the last two vectors in equation (4), which we denote by $V_{j,t}$ and $W_{j,t}$, is given by

$$\begin{aligned} R_j &= (V_{j,t} + W_{j,t})(V_{j,t} + W_{j,t})' \\ &= \begin{bmatrix} (1+\alpha/\beta)s^2 & \dots & 0 & \dots & 0 & 0 \\ \dots & \dots & \dots & \dots & \dots & \dots \\ 0 & \dots & s^2 & \dots & 0 & 0 \\ \dots & \dots & \dots & \dots & \dots & \dots \\ 0 & \dots & 0 & \dots & (1+\alpha/\beta)s^2 & 0 \\ 0 & \dots & 0 & \dots & 0 & (1+\alpha/\beta)s^2/\gamma \end{bmatrix}. \end{aligned}$$

(For the Chairman, the last element of R_C equals s^2/γ, whereas the other elements on the diagonal are given by $(1+\alpha/\beta)s^2$.)

Following Hamilton (1994), it can be shown that policymaker j's optimal assessment $i_{j,t|t}$ of i_t^* is computed as

$$\begin{bmatrix} i_{j,t|t} \\ i_{j,t-1|t} \end{bmatrix} = (I - K_j H') \left(\begin{bmatrix} c \\ 0 \end{bmatrix} + \begin{bmatrix} \rho_1 & \rho_2 \\ 1 & 0 \end{bmatrix} \begin{bmatrix} i_{j,t-1|t-1} \\ i_{j,t-2|t-1} \end{bmatrix} \right) + K_j \begin{bmatrix} i_{j1,t} \\ \dots \\ i_{jC,t} \end{bmatrix},$$

with I a 2×2 identity matrix, $K_j = P_j H(H' P_j H + R_j)^{-1}$, and

$$P_j = \begin{bmatrix} \rho_1 & \rho_2 \\ 1 & 0 \end{bmatrix} (P_j - K_j H' P_j) \begin{bmatrix} \rho_1 & \rho_2 \\ 1 & 0 \end{bmatrix}' + Q.$$

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The Expected Interest Rate Path: Alignment of Expectations vs. Creative Opacity*

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We examine the effects of the release by a central bank of its expected future interest rate in a simple two-period model with heterogeneous information between the central bank and the private sector. The model is designed to rule out common-knowledge and time-inconsistency effects. Transparency—when the central bank publishes its interest rate path—fully aligns central bank and private-sector expectations about the future inflation rate. The private sector fully trusts the central bank to eliminate future inflation and sets the long-term interest rate accordingly, leaving only the unavoidable central bank forecast error as a source of inflation volatility. Under opacity—when the central bank does not publish its interest rate forecast—current-period inflation differs from its target not just because of the unavoidable central bank expectation error but also because central bank and private-sector expectations about future inflation and interest rates are no longer aligned. Opacity may be creative and raise welfare if the private sector's interpretation of the current interest rate leads it to form a view of expected inflation and to set the long-term rate in a way that systematically offsets the effect of the central bank forecast error on inflation volatility. Conditions that favor the case for transparency are a high degree of precision of central bank information relative to private-sector information, a high precision of early information, and a high elasticity of current to expected inflation.

JEL Codes: D78, D82, E52, E58.

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1. Introduction

A number of central banks—the Reserve Bank of New Zealand, the Bank of Norway, the Central Bank of Iceland, and the Swedish Riksbank—now announce their expected interest rate paths, in addition to their inflation and output-gap forecasts. One reason for this practice is purely logical. Inflation-targeting central banks publish the expected inflation rate and the output gap, typically over a two- or three-year horizon, but what assumptions underlie their forecasts? Obviously, they make a large number of assumptions about the likely evolution of exogenous variables. One of these is the policy interest rate. Most banks used to assume a constant policy interest rate. If, however, the resulting expected rate of inflation exceeds the inflation target, the central bank is bound to raise the policy rate, which implies that the inflation forecast does not really reflect what the central bank expects. This is why many central banks now report that their inflation-forecasting procedure relies on the interest rate implicit in the yield curve set by the market. As long as the central bank agrees with the market forecasts, this might seem to be an acceptable procedure. But what if the market forecasts do not lead, in the central bank's view, to the desirable outcome? Then the inflation forecasts are not what the central bank expects to see and, therefore, the market interest forecasts must differ from those of the central bank. As noted by Woodford (2006), consistency requires that the central bank report the expected path of the policy rate along with its inflation and output-gap forecasts.

Why then do most central banks conceal their conditional inflation forecasts by not revealing their expected interest rate paths? Would it not be preferable for central banks to reveal their own expectations of what they anticipate to do? Most central banks reject this idea. Goodhart (2006) offers a number of reasons of why they do so:

Carl Walsh, and John Williams, as well as from participants in seminars at the University of California, Berkeley; the Federal Reserve Bank of San Francisco; the Bank of Korea; the Bank of Norway; the Riksbank; and the Third Banca d'Italia-CEPR Conference on Money, Banking and Finance. All errors are our own.

> If, as I suggest, the central bank has very little extra (private, unpublished) information beyond that in the market, [releasing the expected interest rate path forces the bank to choose between] the Scilla of the market attaching excess credibility to the central bank's forecast (the argument advanced by Stephen Morris and Hyun Song Shin), or the Charybdis of losing credibility from erroneous forecasts.

The first concern is that the central bank could become unwillingly committed to earlier announcements even though the state of the economy has changed in ways that were then unpredictable. The risk is that either the central bank validates the pre-announced path, and enacts suboptimal policies, or it chooses a previously unexpected path and loses credibility since it does not do what it earlier said it would be doing. This argument is a reminder of the familiar debate on time inconsistency. The debate has shown that full discretion is not desirable. Blinder et al. (2001) and Woodford (2005) argue instead in favor of a strategy that is clearly explained and shown to the public to guide policy decisions.

The second concern is related to the result by Morris and Shin (2002) that the public tends to attribute too much weight to central bank announcements—not because central banks are better informed, but because these announcements are common knowledge. This argument is far from convincing. It is based on the doubtful assumption that the central bank is poorly informed relative to the private sector (Svensson 2005a). It also ignores the fact that central banks must reveal at least the current interest rate (Gosselin, Lotz, and Wyplosz 2008).

The third, related, concern is that revealing future interest rates might create a potential credibility problem. The central bank's announcement is bound to shape the market-set yield curve, but what if the implied short-term rates do not accord with those announced by the central bank? Since it is the long end of the yield curve that affects the economy, and therefore acts as a key transmission channel of monetary policy, it could force the central bank to take more abrupt actions to move the yield curve to match its own interest rate forecasts. Would this note be countereffective?

Finally, central bank decisions are normally made by committees—the Reserve Bank of New Zealand is an exception among inflation-targeting central banks—which, it is asserted, are unlikely to be able to agree on future interest rates. The Bank of Norway and the Riksbank show that this is not really the case. Quite to the contrary, these central banks not only explain that committees can think about the expected interest rate path, but they also report that doing so improves the quality of analyses carried out by both the decision makers and the staff.[1]

We deal with some, not all, of these questions. Because they have been extensively studied, we deliberately ignore the time-consistency issue and the Morris-Shin effect. Instead, we focus on the information role of interest rate forecasts with two aims. First, we examine how the publication of the expected interest rate path affects private-sector expectations in a simple model characterized by information heterogeneity—the central bank and the private sector receive different information about a random shock. Second, we ask whether revealing the forecasted policy rates is desirable.

In our model, full central bank transparency is not necessarily desirable because an imperfectly informed central bank policy inevitably makes forecast errors; this is indeed one argument put forward against the publication of the interest rate path. The private sector recognizes that the central bank's forecast errors result in misguided policy choices, but it fully trusts the central bank to do the best that it can given its information set. With no further information about this information set, the private sector does not fully understand the policy choice about the current interest rate and therefore draws wrong conclusions about this choice. When it publishes its interest rate forecast, the central bank reveals its information set, which helps the private sector to more accurately interpret the current interest rate decision; yet, this is not always optimal. In a typical second-best fashion, it may be that the private sector's erroneous inference of the central bank's erroneous policy choice delivers a welfare-superior outcome. For the publication of the expected interest rate path to be desirable, the central bank

[1]This information was obtained via private communication from Anders Vredin.

information must be precise relative to that of the private sector and early signals must be precise relative to subsequent updates.[2]

Two other results are worth mentioning at the outset. First, because they receive different signals, the central bank and the private sector do not generally agree on expected future inflation. In our model, the publication of the interest rate path forecast fully aligns expectations, not because the information sets become identical but because expectations coincide. Second, the publication of the interest rate path forecast leads to a process of information swapping between the central bank and the private sector: we call this a mirror effect. The central bank initially provides information about its signals and subsequently recovers information about the private-sector signals.

The literature on the revelation of expected future policy interest rates is limited so far. Archer (2005) and Qvigstad (2005) present, respectively, the approach followed by the Reserve Bank of New Zealand and the Bank of Norway. Svensson (2005b) presents a detailed discussion of the shortcomings of central bank forecasts based on the constant interest rate assumption or on market rates to build up the case for using and revealing the policy interest rate path. Faust and Leeper (2005) emphasize the distinction between conditional and unconditional forecasts. They assume that the central bank holds an information advantage over the private sector, which in their model implies that sharing that information is welfare enhancing. They show that conditional forecasts—i.e., not revealing the policy interest rate path—provide little information on the more valuable unconditional forecasts, for which they find some supporting empirical evidence.

Similarly, Rudebusch and Williams (2006) assume an information asymmetry between the central bank and the private sector regarding both policy preferences and targets.[3] The private sector

[2]This second-best result is related to the demonstration by Hellwig (2005) that the reason why nontransparency may be desirable in Morris and Shin (2002) is the existence of a market failure due to the combination of asymmetric information and incomplete markets.

[3]Rudebusch and Williams (2006) also offer an excellent overview of the policy debate about how central banks signal their intentions regarding future policy actions.

learns about these factors by running regressions on past information, which may include the expected interest rate path. The paper also allows for a "transmission noise" that distorts its communication. Through simulations, they find that revealing the expected path improves the estimation process and welfare, with a gain that declines as the transmission noise increases. Additionally, they explore the case when the accuracy of the central bank signals is not known by the public. They find that accuracy underestimation limits the gains from releasing the expected interest rate path, while overestimation may be counterproductive. This result is not of the Morris-Shin variety, however, because what is at stake is not the precision of information but the size of the transmission noise, a very different phenomenon.

Walsh (2007) considers a model where the central bank and individual firms receive different signals about aggregate demand and firm-level cost shocks. As a consequence, as in Morris and Shin (2002), the publication by the central bank of its output-gap forecasts—which is equivalent in his model to revealing expected inflation—has a large effect on individual firm forecasts, which can be welfare reducing if the central bank is poorly informed. Walsh examines the possibility that the central bank information is not received by all firms. Partial transparency may offset the common-knowledge effect. The optimal degree of transparency—the proportion of firms that receive the central bank's information—depends on the relative accuracy of the central bank's information about demand and supply shocks.

Our contribution differs from Faust and Leeper (2005) and Rudebusch and Williams (2006). They assume the existence of an information *asymmetry*, which makes transparency always desirable as long as the central bank is credible. Instead, we assume that the central bank is credible with known preferences—which fully accord with social preferences—and we focus on information *heterogeneity* between the central bank and the private sector. Walsh (2007) too deals with information heterogeneity but, as we consider a single representative private agent, we eliminate the common-knowledge effect that is at the center of his analysis.

The next section presents the model, a simple two-period version of the standard New Keynesian log-linear model. Section 3 looks at the case when the central bank optimally chooses the interest rate and announces its expected future interest rate. In section 4, the central bank follows the same rule as in section 3 but does not reveal its expected future interest rate. Section 5 compares the welfare outcomes of the two policy regimes, and the last section concludes with a discussion of arguments frequently presented to reject the release of interest rate expectations by central banks.

2. The Model

2.1 Macroeconomic Structure

We adopt the now-standard New Keynesian log-linear model, as in Woodford (2003). It includes a Phillips curve:

$$\pi_t = \beta E_t^P \pi_{t+1} + \kappa_1 y_t + \varepsilon_t, \tag{1}$$

where y_t is the output gap and ε_t is a random disturbance, which is assumed to be uniformly distributed over the real line, therefore with an improper distribution and a zero unconditional mean. In what follows, without loss of generality, we assume a zero rate of time preference so that $\beta = 1$. The output gap is given by the forward-looking IS curve:

$$y_t = E_t^P y_{t+1} - \kappa_2 \big(r_t - E_t^P \pi_{t+1} - r^*\big), \tag{2}$$

where r_t is the nominal interest rate. We do not allow for a demand disturbance because allowing for two sources of uncertainty would greatly complicate the model.[4] We assume that the natural real interest rate $r^* = 0$. Note that all expectations E^P are those of the private sector, which sets prices and decides on output after the central bank has decided on the contemporaneous interest rate.

We limit our horizon to two periods by assuming that the economy is in steady state at $t = 0$ and $t \geq 3$, i.e., when inflation, output gap, and the shocks are nil. This simplifying assumption is meant to describe a situation where past disturbances have been absorbed so that today's central bank action is looked upon as dealing with the current situation ($t = 1$) given expectations about the near future

[4] A generalization to both demand and supply disturbances, which could preclude obtaining closed-form solutions, is left for future work. Walsh (2007) examines the different roles of these disturbances.

($t = 2$)—say two to three years ahead—while too little is known about the very long run ($t \geq 3$) to be taken into consideration. Consequently, (1) and (2) imply

$$\pi_1 = E_1^P \pi_2 - \kappa\big(r_1 - E_1^P \pi_2 + E_1^P r_2 - E_1^P \pi_3\big) + \kappa_1 E_1^P y_3 + \varepsilon_1,$$

where $\kappa = \kappa_1\kappa_2$. Note that the channel of monetary policy is the real long-term interest rate, the second term in the above expression. This long-term rate is decided partly by the central bank—it chooses r_1—and partly by the private sector, which sets the longer end of the yield curve $E_1^P r_2$ and the relevant expected inflation rates $E_1^P \pi_2$ and $E_1^P \pi_3$. This implies that, when it sets the interest rate r_1, the central bank must take into account the effect of its decision on market expectations. Put differently, the central bank must forecast how private-sector forecasts will react to the choice of r_1.

Since the economy is known to return to steady state in period 3, $E_1^P \pi_3 = 0$ and $E_1^P y_3 = 0$ and the previous equation simplifies to

$$\pi_1 = (1 + \kappa)E_1^P \pi_2 - \kappa\big(r_1 + E_1^P r_2\big) + \varepsilon_1, \tag{3}$$

where $r_1 + E_1^P r_2$ is the long-run (two-period) nominal interest rate. Similarly,

$$\pi_2 = -\kappa r_2 + \varepsilon_2, \tag{4}$$

where we also assume that the central bank sets $r_t = r^*$ for $t \geq 3$, which is indeed optimal, as will soon be clear.

The loss function usually assumes that society is concerned with stabilizing both inflation and the output gap around some target levels, which allows for a well-known inflation-output trade-off. Much of the literature on central bank transparency additionally focuses on the idea that the public at large may not know how the central bank weighs these two objectives. This assumption creates an information asymmetry, which makes transparency generally desirable, as shown in Rudebusch and Williams (2006). Here, instead, we ignore this issue by assuming that the weight on the output gap is zero and that the target inflation rate is also nil. Since the rate of time preference is zero, the loss function is, therefore, evaluated as the unconditional expectation:

$$L = E\big(\pi_1^2 + \pi_2^2\big) \tag{5}$$

and this is known to everyone.

2.2 Information Structure

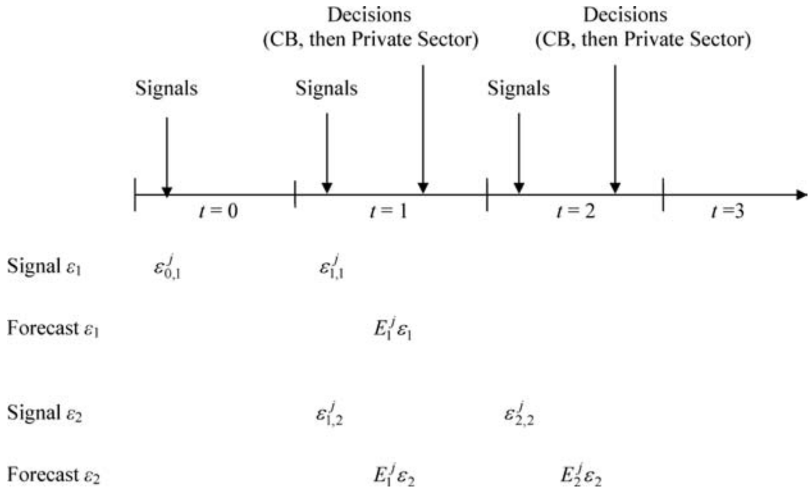
The information structure is crucial. Information asymmetry requires that the central bank and the private sector receive different signals about the shock ε_t. In addition, in order to meaningfully discuss the publication of interest rate forecasts, we allow for the central bank to discover new information between the release of its forecast and the decision on the corresponding interest rate. To that effect, we assume that two signals are received for each shock ε_t, both of which are centered around the shock: (i) an early signal $\varepsilon^j_{t-1,t}$ obtained in the previous period, which leads to the forecast $E^j_{t-1}\varepsilon_t$, and (ii) a contemporaneous signal $\varepsilon^j_{t,t}$, where $j = CB, P$ denotes the recipient of the signals—the central bank and the private sector, respectively. Both of them then combine the early and updated signals to form new forecasts $E^{CB}_t\varepsilon_t$ and $E^{P'}_t\varepsilon_t$.[5] Note that the private-sector forecast based on its own signals is denoted with a prime to distinguish it from the forecasts made subsequently, after the central bank has decided on the interest rate, which is instantly revealed. Thus the operator E^P_1 in (3) combines $E^{P'}_1$ with the information content of r_1.

Figure 1 presents the information structure and the timing of decisions. At the beginning of period 0, the central bank and the private sector receive an early signal $\varepsilon^j_{0,1}$ on the shock ε_1. These signals have known variances $(k\alpha)^{-1}$ and $(k\beta)^{-1}$ for the central bank and the private sector, respectively. Equivalently, the signal precisions are $k\alpha$ and $k\beta$. At the beginning of period 1, updated signals on $\varepsilon^{CB}_{1,1}$ and $\varepsilon^P_{1,1}$—with variances $[(1-k)\alpha]^{-1}$ and $[(1-k)\beta]^{-1}$, respectively—are received by the central bank and the private sector. Using Bayes's rule to exploit both signals, the central bank and the private sector infer expectations $E^{CB}_1\varepsilon_1$ and $E^{P'}_1\varepsilon_1$, respectively, with variances α^{-1} and β^{-1} or, equivalently, precisions α and β. The parameter k measures the relative precision of early signals vis à vis the updated signals, and we assume that $0 \leq k \leq 1$.

Much the same occurs concerning the period 2 disturbance ε_2, with a slight but importance difference. At the beginning of period 1,

[5]The assumption that ε_t is uniformally dsitributed implies that Bayes's rule is only applied to the signals. Note that $cor(\varepsilon_t, \varepsilon^j_t) = 1$.

Figure 1. Timing of Information and Decisions



the central bank and the private sector receive, respectively, the early signals $\varepsilon^{CB}_{1,2}$ and $\varepsilon^{P}_{1,2}$ with variances $(k\alpha)^{-1}$ and $(k\beta)^{-1}$. The central bank then forms $E^{CB}_1\varepsilon_2 = \varepsilon^{CB}_{1,2}$ and sets r_1 to minimize $E^{CB}_1 L$. The private sector waits until r_1 is set and announced to form $E^{P}_{1,2}\varepsilon_2$, using both its early signal $\varepsilon^{P}_{1,2}$ and whatever information it can extract from r_1. Thus, as previously noted, $E^{CB}_1\varepsilon_2$ and $E^{P}_1\varepsilon_2$ are formed at different times during period 1: $E^{CB}_1\varepsilon_2$ before r_1 is known and $E^{P}_1\varepsilon_2$ afterwards. The reason is that r_1 conveys new information to the private sector, not to the central bank.

At the beginning of period 2, the central bank and the private sector receive contemporaneous signals $\varepsilon^{CB}_{2,2}$ and $\varepsilon^{P}_{2,2}$, with variances $[(1-k)\alpha]^{-1}$ and $[(1-k)\beta]^{-1}$, respectively. We further assume that, at the beginning of period 2, the realized values of π_1 and ε_1 become known to both the central bank and the private sector. The central bank uses all information available—the early and contemporaneous signals $\varepsilon^{CB}_{1,2}$ and $\varepsilon^{CB}_{2,2}$ as well as π_1 and ε_1—to form its forecast $E^{CB}_2\varepsilon_2$ and sets r_2 to minimize $E^{CB}_2 L$. After the central bank decision, the private sector observes r_2, forms its expectations, and decides on output and prices.

The focus of the paper is whether, in addition to choosing and announcing r_t, the central bank should also reveal its expectation of the interest rates in the following periods r_{t+i}. This issue is made simpler once we recognize that $r_t = 0$ for all $t \geq 3$, so that we will only need to consider the choice of r_1 and r_2 and whether the central bank reveals $E_1^{CB} r_2$.

2.3 Comments

The model combines some highly stylized features with a rather complicated information structure. The objective is to work with the simplest possible model that can meaningfully explore the role of interest rate forecasts. Two periods allow us to distinguish between the current and the future interest rate. Two signals—early and updated—make it possible for the actually chosen interest rate to differ from its forecast. Information heterogeneity provides a channel for central bank release of information to affect private-sector decisions, raising a few issues along the way.

We intentionally shut down two prominent channels that provide arguments against full central bank transparency: creative ambiguity and the common-knowledge effect.[6] Creative ambiguity emerges in the presence of time inconsistency due to uncertainty about the central bank preferences, presumed to differ from those of the private sector. It is desirable in a model where only unanticipated money matters so that central banks need to preserve some secrecy margin, a dubious assumption rejected by the New Keynesian Phillips curve. Here, instead, we assume that the central bank and the private sector only care about inflation, a special—simple—case of identical preferences further discussed below. The beauty-contest effect arises when the private sector includes a large number of agents who each receive a different signal and pay excessive attention to central bank signals simply because these signals are seen, and are known to be seen, by all. We previously voiced doubts about the practical relevance of this effect. Here we assume that there is a single representative private-sector agent.

We also rule out information asymmetry between the central bank and the private sector and focus instead on information heterogeneity. Information asymmetry generally provides support for

[6]The seminal contribution for creative ambiguity is Cukierman and Meltzer (1986); the common-knowledge effect is due to Morris and Shin (2002).

transparency. However, except for its own preferences, it is hard to imagine what information advantage is enjoyed by central banks.[7] Information heterogeneity arises when central banks and the private sector have different (non-nested) information about the economy or the "right model," a highly plausible assumption.

In our model, information heterogeneity occurs because the central bank does not observe the long-term rate before it sets the interest rate. This may seem unrealistic. Central banks, which move at discrete times, can and do observe the continuously updated yield curve and other financial variables whenever they make decisions. But if we allow the private sector to set the long-term interest rates before the central bank makes its decision, the result is information asymmetry, not heterogeneity. Indeed, the central bank would know both the private-sector signals and its own signals. As we explain below, this would eliminate the welfare difference between transparency and opacity. We could allow the private sector to move first and yet preserve information heterogeneity by allowing for more than one source of uncertainty. In this case, observing the yield curve would provide the central bank with information on the combination of shocks, not on the individual shocks.[8] However, adding one more shock would make our model intractable analytically.

More generally, in this kind of linear model, the fact that the central bank observes the variables (interest rate, asset prices, etc.) set by the private sector does not lead to information asymmetry as long as the number of these variables is smaller than the number of private-sector signals. Allowing for a single private-sector signal that is not observed by the central bank is a parable meant to capture the idea that the central bank cannot uncover all private signals. This is the simplest possible framework that gives rise to information heterogeneity.

Finally, our loss function (5) implies that the central bank pursues a strict inflation-targeting strategy. In practice, however, the commonly adopted strategy is flexible inflation targeting. This issue is, again, related to the general issue of the number of shocks and signals. Meaningfully adding the output gap to the loss function would

[7]We ignore confidential central bank information about the situation of banks during financial crises, a different phenomenon from the one at hand.

[8]This is the modeling strategy adopted by Walsh (2007).

require allowing for demand shocks. Two variables and their signals would make the model considerably more complicated, most likely analytically intractable. We believe that our generic results would be qualitatively preserved.

3. The Central Bank Reveals Its Interest Rate Forecast

We first look at the case where the central bank reveals $E_1^{CB} r_2$, which we refer to as the transparency case. In period 2, the central bank sets the interest rate in order to minimize $E_2^{CB}(\pi_2)^2$ conditional on the information available at the beginning of this period, i.e., after it has received the signal $\varepsilon_{2,2}^{CB}$.[9] The central bank seeks to offset the perceived shock and sets

$$r_2 = \frac{1}{\kappa} E_2^{CB} \varepsilon_2. \tag{6}$$

The simplicity of this choice is a consequence of our assumption that the economy will return to the steady state in period $t = 3$. It can be viewed either as a rule or as discretionary action given the new information received at the beginning of the period.

Moving backward to period 1, the central bank publishes $E_1^{CB} r_2 = \frac{1}{\kappa} E_1^{CB} \varepsilon_2 = \frac{1}{\kappa} \varepsilon_{1,2}^{CB}$. This shows that publishing the interest rate is equivalent to fully revealing the central bank signal $\varepsilon_{1,2}^{CB}$.[10] As a consequence, in period 1 the private sector receives two signals about ε_2: its own signal $\varepsilon_{1,2}^{P}$ with precision $k\beta$ and, as just noted, the central bank signal $\varepsilon_{1,2}^{CB}$ with precision $k\alpha$. Denoting the relative precision of the central bank and private-sector signals as $z = \frac{\alpha}{\beta}$, the private sector uses Bayes's rule in period 1 to optimally forecast ε_2:

$$E_1^P \varepsilon_2 = \gamma_1^{tr} \varepsilon_{1,2}^{P} + \left(1 - \gamma_1^{tr}\right) \varepsilon_{1,2}^{CB} = \frac{1}{1+z} \varepsilon_{1,2}^{P} + \frac{z}{1+z} \varepsilon_{1,2}^{CB}. \tag{7}$$

[9] Note that we do not allow for the private sector to use newly received information $\varepsilon_{2,2}^{P}$, which arrives too late to be of any use.

[10] This is so because the model allows for one signal and one policy instrument. If there were more signals than instruments, publishing the expected future value of one instrument would not be fully revealing.

In order to set the long-term interest rate, the private sector also needs to forecast the future short-term interest rate given by (6) and therefore the central bank's own forecast of the future shock. Conjecture that, similarly to (7), the optimal forecast is

$$E_1^P E_2^{CB} \varepsilon_2 = \gamma_2^{tr} \varepsilon_{1,2}^P + \left(1 - \gamma_2^{tr}\right) \varepsilon_{1,2}^{CB} \tag{8}$$

with unknown coefficient γ_2^{tr} to be determined.

When period 2 starts, π_1 and ε_1 become known. As a consequence, (3) and (6) show that $\pi_1 + \kappa r_1 - \varepsilon_1 = (1+\kappa)(E_1^P \varepsilon_2 - E_1^P E_2^{CB} \varepsilon_2) - E_1^P E_2^{CB} \varepsilon_2$ is known to both the central bank and the private sector. Using (7) and (8) we have

$$\pi_1 + \kappa r_1 - \varepsilon_1 = \left[(1+\kappa)\left(\gamma_1^{tr} - \gamma_2^{tr}\right) - \gamma_2^{tr}\right]\left(\varepsilon_{1,2}^P - \varepsilon_{1,2}^{CB}\right) + \gamma_2^{tr} \varepsilon_{1,2}^{CB}.$$

This implies that, at the beginning of period 2, when π_1 and ε_1 become known, the central bank can recover the private signal $\varepsilon_{1,2}^P$. We have a delayed mirror effect: by revealing the expected future interest rate, the central bank gives out its period 1 information $\varepsilon_{1,2}^{CB}$ and gets in return, in period 2, the private information $\varepsilon_{1,2}^P$. Put differently, by observing how its own information was previously interpreted, the central bank now recovers the signal previously received by the private sector. Importantly, the mirror image is not identical to the original; it provides the central bank with useful information when it decides on the interest rate r_2. Indeed, it can use three signals about ε_2: $\varepsilon_{1,2}^{CB}$ received in period 1 with precision $k\alpha$, $\varepsilon_{2,2}^{CB}$ received in period 2 with precision $(1-k)\alpha$, and now $\varepsilon_{1,2}^P$ with precision $k\beta$. Applying Bayes's rule we have

$$E_2^{CB} \varepsilon_2 = \frac{z\left[k\varepsilon_{1,2}^{CB} + (1-k)\varepsilon_{2,2}^{CB}\right] + k\varepsilon_{1,2}^P}{z+k}.$$

Noting that $E_1^P \varepsilon_{2,2}^{CB} = E_1^P \varepsilon_2$, it follows that $\gamma_1^{tr} = \gamma_2^{tr}$ and therefore[11]

$$E_1^P E_2^{CB} \varepsilon_2 = \frac{1}{1+z}\varepsilon_{1,2}^P + \frac{z}{1+z}\varepsilon_{1,2}^{CB} = E_1^P \varepsilon_2.$$

The private sector's own forecast of the future shock is perfectly aligned with its perception of the future central bank estimate of this shock, which it knows will lead to the choice of the future interest rate. As they swap signals, both the central bank and the private sector learn from each other. As a consequence, the private sector knows that its own forecast will be taken into account by the central bank when it applies Bayes's rule before deciding on r_2.

PROPOSITION 1. *When the central bank reveals its expected future interest rate, the private sector and the central bank exchange information about their signals received in period 1 about the period 2 shock:*

- *In period 1, the central bank fully reveals its early signal about the period 2 shock, which is then used by the private sector to improve its own forecast.*
- *In period 2, the central bank can identify the corresponding early signal previously received by the private sector.*
- *As a result, central bank and private-sector expectations are fully aligned and, in period 1, both expect future inflation to be zero.*

The last statement in the proposition is readily established. In period 2, the interest rate r_2 is set by the central bank according to

[11]Proof:

$$\begin{aligned} E_1^P E_2^{CB} \varepsilon_2 &= \frac{zE_1^P\left[k\varepsilon_{1,2}^{CB} + (1-k)\varepsilon_{2,2}^{CB}\right] + k\varepsilon_{1,2}^P}{z+k} \\ &= \frac{z\left[k\varepsilon_{1,2}^{CB} + (1-k)\left(\dfrac{\varepsilon_{1,2}^P}{1+z} + \dfrac{z\varepsilon_{1,2}^{CB}}{1+z}\right)\right] + k\varepsilon_{1,2}^P}{z+k} \\ &= \frac{\varepsilon_{1,2}^P}{1+z} + \frac{z\varepsilon_{1,2}^{CB}}{1+z}. \end{aligned}$$

(6), which fully reveals $E_2^{CB}\varepsilon_2$, the central bank updated information about the shock ε_2. Using (4), it follows that

$$\pi_2 = \varepsilon_2 - E_2^{CB}\varepsilon_2. \tag{9}$$

As a consequence, $E_1^P\pi_2 = E_1^P\varepsilon_2 - E_1^P E_2^{CB}\varepsilon_2 = 0 = E_1^{CB}\pi_2$. The publication in period 1 by the central bank of its inflation forecast $E_1^{CB}\pi_2$ is uninformative: it simply restates that the central bank aims at bringing inflation to its target level. This is similar to forecasts of inflation-targeting central banks, which are invariably on target at the chosen horizon, typically two to three years ahead.[12]

We can characterize the optimal monetary policy. In period 2, it is described by (6). In period 1, the central bank sets the interest rates to minimize $E_1^{CB}(\pi_1^2 + \pi_2^2)$ conditional on available information. Since (9) shows that r_1 does not affect π_2, in period 1 the central bank can simply minimize $E_1^{CB}\pi_1^2$. Since $E_1^P\pi_2 = 0$, from (3) we see that the central bank chooses the short-term interest rate r_1 such that, in expectation, the long-term interest rate—which is what matters for aggregate demand—fully offsets the current shock:

$$r_1 + E_1^{CB}E_1^P r_2 = \frac{1}{\kappa}E_1^{CB}\varepsilon_1.$$

Since $E_1^P\varepsilon_2 = E_1^P E_2^{CB}\varepsilon_2$, $E_1^P r_2 = E_1^{CB} r_2$ and, using (6), we find the optimal policy decision in period 1:

$$r_1 = \frac{1}{\kappa}\left(E_1^{CB}\varepsilon_1 - E_1^{CB}\varepsilon_2\right). \tag{10}$$

In period 1, having observed r_1, the private sector uses (6) to set the long-term interest rate:

$$r_1 + E_1^P r_2 = r_1 + \frac{1}{\kappa}E_1^P E_1^{CB}\varepsilon_2 = r_1 + \frac{1}{\kappa}E_1^{CB}\varepsilon_2,$$

[12] On the other hand, evidence so far by Archer (2005) and Ferrero and Secchi (2007) suggests that market expectations only partially adjust following the publication of the interest rate path. We find that expectations are fully aligned because we assume that there is only one source of uncertainty. Allowing for more shocks would mean that the central bank revelation of its expected interest rate path would not fully reveal all its information, as noted in section 2.3 above. We are grateful to the anonymous referee for attracting our attention to this point.

which is the same as the central bank's own forecast. Thus the yield curve exactly matches the interest rate path published by the central bank.

Collecting the previous results, we obtain

$$\pi_1 = \left(\varepsilon_1 - E_1^{CB}\varepsilon_1\right) + \frac{1}{1+z}\left(\varepsilon_{1,2}^{CB} - \varepsilon_{1,2}^{P}\right).$$

Period 1 inflation depends on two forecasting errors: the period 1 central bank forecasting error and the discrepancy between the central bank and the private-sector signals regarding period 2 shock.[13] Note that the impact of this last discrepancy is less than one for one ($\frac{1}{1+z} < 1$) because the revelation of $\varepsilon_{1,2}^{CB}$ by the central bank leads the private sector to discount its own signal $\varepsilon_{1,2}^{P}$ and to bring its forecast $E_1^P\varepsilon_2$ in the direction of $\varepsilon_{1,2}^{CB}$. Note also that $E_1^{CB}\pi_1 = 0$: the central bank always forecasts inflation rate to be on target because its objective does not call for any trade-off with other objectives. That forecast is also uninformative.

The private sector is well aware that the central bank's interest rate forecast is bound to be inaccurate. Indeed, in general, there is no reason for $E_1^P E_2^{CB}\varepsilon_2$ to be equal to ε_2, but the eventual realization of this difference is irrelevant. The private sector fully understands that the future interest rate will usually differ from what was announced, since the central bank will then respond to newly received information $\varepsilon_{2,2}^{CB}$; see (6). This eventual discrepancy is fully anticipated by the private sector because the central bank strategy—its loss function—is public knowledge, so credibility is not an issue here. The difference between the pre-announced rate $E_1^{CB}r_2$ and the actually chosen rate r_2 is purely random and therefore uninformative. Importantly, this result holds independently of the degree of precision of the signals received by the central bank and the private sector. What matters is that signal precision be known.[14]

[13]More precisely, inflation is the result of three forecasting errors since $\pi_1 = (\varepsilon_1 - E_1^{CB}\varepsilon_1) + \frac{1}{1+z}[(\varepsilon_{1,2}^{CB} - \varepsilon_2) - (\varepsilon_{1,2}^{CB} - \varepsilon_2)]$, which includes the central bank and private-sector early forecast errors about ε_2.

[14]The case when the signal precisions are not known is left for further research. For a study of this case in a different setting, see Gosselin, Lotz, and Wyplosz (2008).

Finally, for future reference, in this case of transparency the unconditional loss function is

$$L^{tr} = E(\pi_1)^2 + E(\pi_2)^2 = \frac{1}{\beta}\left[\frac{1}{z} + \frac{1}{k}\left(\frac{1}{1+z}\right)^2\left(\frac{1}{z}+1\right) + \frac{1}{z+k}\right].$$

4. The Central Bank Does Not Reveal Its Interest Rate Forecast

We consider now the case when the central bank does not announce its expectation of the future interest rate. We call this the opacity case. The optimal interest rate in period 2 remains given by (6). The resulting inflation rate is also the same as in (9), although the information available to the central bank is different from that in the previous case, as will be emphasized below.

In period 1, the central bank still reveals the current interest rate, which is set on the basis of its available information, i.e., $E_1^{CB}\varepsilon_1$ and $E_1^{CB}\varepsilon_2$. We restrict our attention to the following policy linear rule, which optimally uses all available information:[15]

$$r_1 = \mu E_1^{CB}\varepsilon_1 + \nu E_1^{CB}\varepsilon_2, \tag{11}$$

where μ and ν are unknown parameters to be determined by the optimal policy.

Having observed r_1, the private sector sets the inflation rate according to (3). To that effect, it needs to forecast future inflation, which by (9) depends on $E_2^{CB}\varepsilon_2$, the central bank's forecast. In forming this forecast, the central bank uses its signals $\varepsilon_{1,2}^{CB}$ and $\varepsilon_{2,2}^{CB}$ as well as period 1 inflation, which has now become known. In contrast to the previous case, $\varepsilon_{1,2}^{CB}$ is now unknown to the private sector. As a consequence, $E_1^P\varepsilon_2$ no longer coincides with $E_1^P E_2^{CB}\varepsilon_2$. In order to form its forecast $E_1^P E_2^{CB}\varepsilon_2$, following Bayes's rule, the

[15]There is no reason to presume that a linear rule is optimal. This restrictive assumption, required to carry through the calculations that follow, can be seen as a linear approximation of the optimal policy. This introduces some asymmetry between the transparency and opacity cases: in the former, the rule is optimal; in the latter, it may not be. Unfortunately, we are not able to derive the optimal policy choice under opacity.

private sector uses its three available signals $E_1^P \varepsilon_1$, $\varepsilon_{1,2}^P$, and r_1.[16] It can use $\varepsilon_{1,2}^P$ directly. In addition, the interest rate rule (11) implies that $E_1^{CB}\varepsilon_2 = (r_1 - \mu E_1^{CB}\varepsilon_1)/\nu$, so r_1 can be used to make inference about $E_2^{CB}\varepsilon_2$. The optimal forecast is necessarily of the form

$$\begin{aligned} E_1^P E_2^{CB}\varepsilon_2 &= \gamma_2^{op}\varepsilon_{1,2}^P + \left(1-\gamma_2^{op}\right)\left(\frac{r_1 - \mu E_1^{P'}\varepsilon_1}{\nu}\right) \\ &= \gamma_2^{op}\varepsilon_{1,2}^P + \left(1-\gamma_2^{op}\right)\left[\varepsilon_{1,2}^{CB} - \frac{\mu}{\nu}\left(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right] \end{aligned} \quad (12)$$

with γ_2^{op} to be determined. The same reasoning can be applied to $E_1^P\varepsilon_2$ to obtain

$$E_1^P\varepsilon_2 = \gamma_1^{op}\varepsilon_{1,2}^P + \left(1-\gamma_1^{op}\right)\left[\varepsilon_{1,2}^{CB} - \frac{\mu}{\nu}\left(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right], \quad (13)$$

where $\gamma_1^{op} = \frac{k(1+z)+\left(\frac{\nu}{\mu}\right)^2}{(1+z)\left[k+\left(\frac{\nu}{\mu}\right)^2\right]}$.

As in the transparency case, the unknown weighting coefficient γ_2^{op} can be found by identification. In this case, there is no simple analytical solution. The appendix shows that $\gamma_1^{op} - \gamma_2^{op}$ is the solution to a third-order equation that satisfies the following relation:

$$\gamma_1^{op} - \gamma_2^{op} = \frac{z\theta k\left[\theta k - \left(\frac{\nu}{\mu}\right)^2\right]}{(1+z)\left[k+\left(\frac{\nu}{\mu}\right)^2\right]\left[\left(\frac{\mu}{\nu}\right)^2\theta^2 zk + z + k\right]}, \quad (14)$$

where θ is defined as

$$\theta = 1 + \frac{1}{(1+\kappa)\left(\gamma_1 - \gamma_2^{op}\right) - \gamma_2^{op}}.$$

In comparison with the case where the central bank publishes its expected future interest rate, (14) implies that, in general, $\gamma_1^{op} \neq \gamma_2^{op}$ so that $E_1^P\varepsilon_2 \neq E_1^P E_2^{CB}\varepsilon_2$. From (4) and (6), it follows that

$$E_1^P\pi_2 = E_1^P\varepsilon_2 - E_1^P E_2^{CB}\varepsilon_2 \neq 0. \quad (15)$$

[16] More precisely, $E_1^P\varepsilon_1$ is not a signal but the expectation formed on the basis of signals $\varepsilon_{0,1}$ and $\varepsilon_{1,1}$.

Well aware that its own period 1 forecast of the disturbance ε_2 differs from that of the central bank, the private sector is no longer sure that the central bank can achieve its aim. This is the key difference between transparency and opacity. Private-sector doubt is reflected in the discrepancy between central bank and private-sector expectations, which is captured by $\gamma_1^{op} - \gamma_2^{op}$.

The appendix shows that the optimum interest rate rule in period 1 requires $\mu = -\nu = \kappa^{-1}$. The monetary policy rule is formally identical to (10) in the transparency case. As before, the reason is that, in order to minimize the volatility of π_1, the central bank seeks to set the nominal long-term interest rate to offset the first-period shock, which it expects to be $E_1^{CB}\varepsilon_1$; to do so, it must take into account its future interest rate, which it expects to choose so as to offset the future shock, which is expected to be $E_2^{CB}\varepsilon_2$. Thus, even if the central bank is transparent, it must still form a view of its future action.[17] The above results can be summarized as follows.

PROPOSITION 2. *Private-sector and central bank expectations are no longer aligned under opacity. While the interest rate rule is the same as when the central bank announces its expected future interest rate, the yield curve no longer matches the central bank forecast of the interest rate path.*

The resulting inflation rate in period 1 is

$$\pi_1 = \frac{1}{\theta - 1}\left[\left(\varepsilon_{1,2}^{P} - \varepsilon_{1,2}^{CB}\right) - \theta\left(E_1^{P'}\varepsilon_1 - \varepsilon_1\right) + \left(E_1^{CB}\varepsilon_1 - \varepsilon_1\right)\right], \quad (16)$$

which combines the forecast errors of both the private sector and the central bank. It follows that

$$E_1^{P}\pi_2 = \left(\gamma_1^{op} - \gamma_2^{op}\right)\left[\left(\varepsilon_{1,2}^{P} - \varepsilon_{1,2}^{CB}\right) - \left(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right], \quad (17)$$

which shows the role of the doubt factor $\gamma_1^{op} - \gamma_2^{op}$: the private sector will not expect the central bank to eliminate inflation in period 2 unless $\gamma_1^{op} - \gamma_2^{op} = 0$.

[17]Note that even though the interest rate rules are formally the same under both transparency regimes, this does not imply the same interest and inflation rates. Indeed, the information sets of the central bank and of the private sector change with the transparency regime.

Using the expression for $E(\pi_2)^2$ provided in the appendix, we find the loss function under central bank opacity:

$$L^{op} = E(\pi_1)^2 + E(\pi_2)^2$$
$$= \frac{1}{\beta}\left[\left(\frac{1}{\theta-1}\right)^2\left(\frac{1}{z}+\frac{1}{k}+\frac{1}{kz}\right)+\left(\frac{\theta}{\theta-1}\right)^2+\frac{1+\theta^2 k}{\theta^2 zk+z+k}\right].$$

We mentioned in section 2.3 that there would be no welfare difference between transparency and opacity if the central bank could observe in period 1 the yield curve before making its decision. Indeed, in this case, observing $E_1^P r_1$ and $E_1^P r_2$ fully reveals the two private-sector signals $\varepsilon_{1,1}^P$ and $\varepsilon_{1,2}^P$. Similarly, observing $E_2^P r_2$ fully reveals $\varepsilon_{2,2}^P$. It follows that, in period 2, independently of the transparency regime, the central bank knows everything that the private sector knows, there is no mirror effect, and L_2 is the same. In period 1, independently of the regime, we have $E_1^P \pi_2 = E_1^P \varepsilon_2 - E_1^P E_2^{CB} \varepsilon_2 = E_1^P \varepsilon_2 - E_1^P \varepsilon_2 = 0$, i.e., inflation expectations are always aligned. It follows that $\pi_1 = -\kappa(r_1 + E_1^P r_2) + \varepsilon_1$ even though $E_1^P r_2$ is not the same under transparency and opacity. The central bank optimal decision then is $r_1 = -E_1^P r_2 + \frac{1}{\kappa} E_1^{CB} \varepsilon_1$, which implies $\pi_1 = \varepsilon_1 - E_1^{CB} \varepsilon_1$.[18] This shows that period 1 inflation, and therefore welfare, does not depend on the transparency regime.

5. Welfare Analysis

We now compare welfare when the central bank reveals its expected interest rate—labeled transparency—and when it does not—labeled opacity. To do so we study the difference of welfare losses under the two regimes: $\Delta L = L^{op} - L^{tr}$. In spite of the model's extreme simplicity, we cannot derive an explicit condition that determines the sign of ΔL. Consequently, we proceed in three steps. In section 5.1, we derive a sufficient condition for period 1 loss difference ΔL_1 to be positive; since $\Delta L_2 > 0$, this is also a sufficient condition for

[18] Proof: Note that $r_2 = \frac{1}{\kappa} E_2^{CB} \varepsilon_2$ so $E_1^P r_2 = \frac{1}{\kappa} E_1^P E_2^{CB} \varepsilon_2 = \frac{1}{\kappa} E_1^P \varepsilon_2$. This, in turn, implies that $E_1^{CB} E_1^P r_2 = \frac{1}{\kappa} E_1^{CB} E_1^P \varepsilon_2 = \frac{1}{\kappa} E_1^P \varepsilon_2$. Optimal monetary policy in period 1 is $r_1 = -E_1^{CB} E_1^P r_2 + \frac{1}{\kappa} E_1^{CB} \varepsilon_1$. With the previous result, this means $r_1 = -\frac{1}{\kappa} E_1^P \varepsilon_2 + \frac{1}{\kappa} E_1^{CB} \varepsilon_1$.

transparency to dominate opacity. Then, in section 5.2, we provide a necessary condition for $\Delta L_1 < 0$. Finally, we present in section 5.3 the results from the formal analysis of ΔL that is described in the appendix.

5.1 Preliminary Observation

We first compare the welfare losses separately period by period. Starting with period 2, we have

$$\beta \Delta L_2 = L_2^{op} - L_2^{tr} = \frac{1}{\beta} \frac{\theta^2 k^2}{(\theta^2 zk + z + k)(z + k)} > 0. \tag{18}$$

PROPOSITION 3. *Transparency is always welfare increasing in period 2.*

The reason is that the central bank is better informed when it can recover the private-sector signal $\varepsilon_{1,2}^P$; see (9).

Thus, a sufficient condition for transparency to be welfare improving is that the period 1 welfare difference $\Delta L_1 = L_1^{op} - L_1^{tr} \geq 0$. In the appendix we show that

$$\beta \Delta L_1(\theta) = \left(\frac{1}{\theta - 1}\right)^2 \frac{1 + k + z}{kz} + \left(\frac{\theta}{\theta - 1}\right)^2 - \frac{1 + k(1 + z)}{kz(1 + z)}, \tag{19}$$

and we study this expression as a function of θ. This analysis yields the following sufficient condition for transparency to be welfare improving.

PROPOSITION 4. *A sufficient condition for the release by the central bank of its expected future interest rate to be welfare improving is that $z > \frac{1+k}{\sqrt{k}}$.*

The more precise is the central bank signal α relative to the private-sector signal β—the higher is z—the more likely it is that transparency pays off. Conversely, if central bank information is of poor quality—i.e., when $z < \frac{1+k}{\sqrt{k}}$—the situation becomes ambiguous.[19]

[19] Over the relevant range of $\not{k}$, from zero to one, the function $(1 + k)/\sqrt{k}$ is decreasing from ∞ when $k = 0$ to 2 when $k = 1$.

The intuition is as follows. Transparency allows for the exchange of early signals between the central bank and the private sector: in period 1, the central bank reveals $\varepsilon_{1,2}^{CB}$; in period 2, it discovers $\varepsilon_{1,2}^{P}$. The ambiguous period 1 welfare effect of transparency, therefore, depends on the precision of the central bank early signal $\varepsilon_{1,2}^{CB}$, i.e., on k and z. Thus k and z act as complementary factors favoring transparency. A higher k means that transparency is achieved with a lower z, and conversely.

5.2 Why May Opacity Raise Welfare?

We now ask why opacity could ever raise welfare. It might seem that more information is always better than less. This is not necessarily true here since we have two agents—the central bank and the private sector—who strategically interact under heterogeneous information.[20]

The appendix provides a formal explanation of why less information may be welfare increasing. Here we use a very simple example to provide an intuitive interpretation. Consider the case where it will turn out that the two shocks ε_1 and ε_2 are nil, but the signals received by the central bank lead it to mistakenly infer an inflationary shock in period 1 ($E_1^{CB}\varepsilon_1 > 0$) and, correctly, no shock in period 2 ($E_1^{CB}\varepsilon_2 = 0$). Assume also that the private-sector signals turn out to be accurate, so $E_1^{P}\varepsilon_1 = E_1^{P}\varepsilon_2 = 0$. From (10) we know that, expecting an inflationary shock, the central bank raises the interest rate ($r_1 > 0$). This will turn out to be a policy mistake—optimal policy would call for $r_1 = 0$. When it observes the positive interest rate, knowing that it is set according to (10), the private sector can infer either that $E_1^{CB}\varepsilon_1 > 0$ or that $E_1^{CB}\varepsilon 2 < 0$, or a suitable combination of both. Under transparency, the central bank reveals $E_1^{CB}\varepsilon_2 = 0$, so the private sector understands that the central bank has raised the period 1 interest rate because of an inflationary signal. The private sector correctly expects the central bank to bring inflation to target in period 2 ($E_1^{P}\pi_2 = 0$) by keeping $r_2 = 0$. Then (3)

[20] In period 1, the central bank acts as a Stackelberg leader in setting r_1 and then the private sector reacts, setting π_1 and the long-term interest rate. Then, in period 2, the central bank reacts and sets r_2.

shows that $\pi_1 < 0$ because the central bank policy is too restrictive in period 1.

When the central bank does not publish its interest rate forecast, the private sector no longer knows for sure why the interest rate has been raised. For the sake of reasoning, let us consider two possible extreme assumptions about private-sector inference in this situation. If the private sector correctly guesses that $E_1^{CB}\varepsilon_1 > 0$ and $E_1^{CB}\varepsilon_2 = 0$, the situation is the same as under transparency. If instead the private sector incorrectly infers from the interest rate increase that the central bank expects a deflationary shock in period 2 ($E_1^{CB}\varepsilon_1 = 0$ and $E_1^{CB}\varepsilon_2 < 0$), it will conclude that the central bank plans to lower r_2 and has raised r_1 to keep the long-term rate unchanged in shockless period 1. Its expectation of a lower interest rate ($E_1^P r_2 < 0$) leads the private sector to raise its inflation forecast ($E_1^P \pi_2 > 0$). Both terms tend to offset in (3) the effect on π_1 of the contractionary policy actually carried out by the central bank.

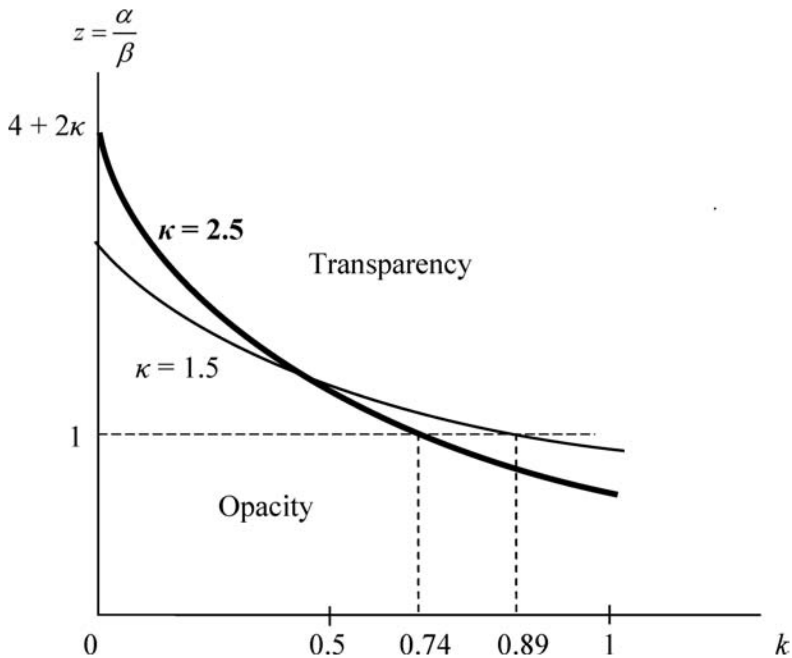
This example illustrates how, under opacity, the private sector's misinterpretation of the central bank action mitigates the effect of a policy mistake and possibly raises welfare. This is a special case of the more general result, developed in the appendix, that opacity is desirable when it leads the private sector to systematically draw inference from the central bank action in a way that offsets policy mistakes due to imperfect signals. This leads to the following proposition.

Proposition 5. *(Creative Opacity) A necessary condition for opacity to welfare dominate transparency is that the private sector's own forecasts systematically offset the impact on inflation volatility of the central bank forecast errors.*

5.3 Welfare Ranking

We have derived a *sufficient* condition for transparency to be desirable and a *necessary* condition for opacity to dominate. We now study how the necessary and sufficient sign condition for ΔL relates to the three model parameters z, κ, and k, with $z \geq 0$, $k \in [0, 1]$, and $\kappa \geq 0$. Figure 2 summarizes the results established in the appendix. It displays two curves that correspond to two values of κ. The area below each curve corresponds to $\Delta L = L^{op} - L^{tr} < 0$, i.e., to the

Figure 2. Welfare Outcomes



case where welfare is higher when the central bank does not reveal its interest rate path forecast.

The following proposition summarizes the results of this analysis.

PROPOSITION 6. *When the central bank follows the optimal linear interest rate rule (11), ceteris paribus, transparency dominates when z is large and when k is large. The role of κ is ambiguous: when k is small, an increase in κ favors opacity, while it favors transparency when k is large.*

We interpret these results below.

The Role of Relative Signal Precision. We first look at the role of $z = \alpha/\beta$, the ratio of central bank signal precision α to private-sector signal precision β. The higher is z, the more likely it is that transparency is desirable. The reason is clear: the publication of the interest rate path provides the private sector with a central bank signal that is more useful the more precise it is relative to its own signals.[21] As z becomes smaller, the benefit from information disclosure declines because the private sector increasingly doubts any signal from the central bank, for good reason. Opacity increases welfare when, having observed the current interest rate r_1, the private sector sets prices and the long-term interest rate so as to systematically offset the effects of potential large central bank forecast errors. Note that this is not the same result as in Morris and Shin (2002), which deals with a common-knowledge effect that is not considered here.

The Role of Early Information Precision. The parameter k, which ranges from zero to one, represents that precision of early signals relative to updated signals. By releasing in period 1 its forecast of the interest rate that it expects to set in period 2, the central bank reveals its early signal $\varepsilon_{1,2}^{CB}$ of the shock ε_2 expected in period 2. Then, in period 2, the central bank can decipher the early signal $\varepsilon_{1,2}^{P}$ received by the private sector in period 1 concerning the same shock ε_2. Since transparency makes this exchange of early signals possible, it is more desirable the more precise are these signals. Indeed, when $k = 0$, these signals become nearly useless.

Yet, figure 2 shows that there always exists a high enough z to make transparency desirable.[22] The reason is that, even when when $k = 0$, in period 1 the private sector still needs to set inflation and the long end of the yield curve. Under opacity, it must rely on its infinitely imprecise early signal $\varepsilon_{1,2}^{P}$ as well as on the interest rate r_1 announced by the central bank, on the basis of its own infinitely imprecise signal $\varepsilon_{1,2}^{CB}$. When the central bank is generally better informed than the private sector—when z is large—it therefore helps the private sector to know $\varepsilon_{1,2}^{CB}$. Put differently, with $k = 0$, the private sector "buys" whatever information it gets from the central bank when z is large. Under opacity, it will not assume that the central bank is misled and will still conclude that $E_1^P \pi_2 \simeq 0$. As expectations are aligned, there is no room for creative

[21] Under opacity, as $z \to \infty$, $\gamma_1^{op} - \gamma_2^{op} \to 0$, so $E_1^P \pi_2 = 0$: opacity converges to transparency. From (18) we know that $\lim_{z\to\infty} \Delta L_2 = 0$ and, from (19), $\lim_{z\to\infty} \beta \Delta L_1 = \frac{1/k+\theta^2}{(\theta-1)^2} > 0$, so transparency dominates when z is large enough.

[22] The intercept of the curve on the vertical axis is $4 + 2\kappa$.

opacity to trigger the kind of welfare-improving correction described in section 4.[23]

As k increases, more attention is paid by both the central bank and the private sector to their own early signals, not just to the other agents' early signals. Under opacity, this heightened attention increases the expectation discrepancy, which is a source of welfare loss. At the same time, because it interprets the current interest rate as conveying information on the central bank's early signal when it sets the long-term rate, the private sector may offset the central bank forecast error, which improves welfare. The expectation discrepancy, which rises with k, directly hurts welfare but may be exploited to raise it indirectly.

Put differently, when it is welfare improving under opacity, the private-sector correction of the mistaken central bank policy decision is more effective the lower is the precision k of early signals. This is because early information swapping under transparency is less effective when k is low, which makes the private-sector correction relatively more helpful.

The Role of the Elasticity of Current to Expected Inflation. Parameter κ represents the channel through which private forecasts of inflation and the long-term interest rate affect current inflation; see (3). As κ increases, the curve that marks the frontier between transparency and opacity in figure 2 shifts up on the left where k is small (the intercept with the horizontal axis is $4 + 2\kappa$) and down on the right where k is large.

To understand why, we need to consider two different effects. First, remember that, due to the non-alignment of central bank and private-sector expectations, opacity tends to increase the volatility of private-sector forecasts and therefore inflation volatility. This effect increases when κ rises, which tends to make transparency more desirable, i.e., to shift the curve down. Second, we have previously noted that the private-sector correction of the central bank error is more likely to stabilize inflation, and therefore to be welfare increasing, the lower is k; we also noted that this effect is reinforced when κ rises. This explains why an increase in κ favors opacity for low values of k. When, instead, k is large, the exchange of noisy early signals under

[23]The appendix shows that the non-alignment of inflation expectations is proportional to the doubt factor $\gamma_1^{op} - \gamma_2^{op}$, which becomes nil as $k \to 0$.

transparency stabilizes expectations and, through κ, the inflation rate.

Lessons from Calibrated Models. The model that we use to derive this result has been calibrated in the literature. In this section, we look at the welfare implications of the parameter values suggested by Galí and Gertler (2007). In their quarterly model, they set $\kappa = 0.167$. In our model, a period is better thought of as lasting two or three years, which approximately implies that κ should range between 1.33 and 2. Svensson (2005a) argues that z cannot be lower than unity and is probably larger. Estimates of z from Clark and McCrackin (2006b) range from 0.83 to 1.55. Taking $z = 1$ as a reference, we ask what the minimum value of k must be in order for transparency to welfare dominate opacity. As shown in figure 2, the critical values of k are 0.74 for $\kappa = 2.5$ and 0.89 for $\kappa = 1.5$. Estimates of k by Clark and McCrackin (2006a) range from 0.43 to 1.

On this basis, the conclusion is that the desirability of publishing the interest rate path is a close call, with most parameter values falling in the no-transparency zone. Of course, these parameter values are to be taken with considerable precaution and, even more importantly, the model is far too simple to be taken at face value. Our purpose is emphatically not to reach normative conclusion but to explore what mechanisms come into play when a central bank publishes its interest rate forecast.

6. Conclusions

The general presumption in the (so far limited) academic literature is that transparency is welfare superior. With few exceptions, most central banks take the opposite view. This paper is a first step to breach the gap. For opacity to welfare dominate transparency, we must identify a market failure. This paper is based on the view that there exists an important degree of information heterogeneity between central banks and the private sector. The results imply that neither side can ever fully recover the information of the other side by simply observing its actions—the private sector observes the interest rate set by the central bank and the central bank observes financial prices. This makes it possible for opacity to welfare dominate transparency. In the simple setup adopted here, opacity is desirable when the private sector misinterprets the central bank decisions and sets its forecasts in a way that offsets the effect of central bank forecasting errors. This double coincidence of forecast errors makes the case for opacity quite weak.

In contrast, the case made by central banks against transparency relies on private-sector confusion between forecasts and commitments. We show that, when it is assumed that both the central bank and the private sector act optimally on the basis of optimal signal extraction and in the absence of a time-inconsistency problem, this is a non-issue. This is so because the private sector has no reason to doubt that the interest rate path announced by the central bank is optimal given its information set. As the information set changes, so must the optimal path. Put differently, the standard case against transparency relies either on suboptimal private behavior—the private sector does not form its expectations on the basis of available information—or on the dubious assumption that central banks act strategically in a way that gives rise to time inconsistency. Narrowing down the policy debate is, we hope, a relevant contribution.

Another insight is that the case for transparency is enhanced when early signals are precise relative to contemporaneous signals. Put differently, the interest rate path becomes less useful when the outlook becomes more uncertain. This resembles the Morris-Shin result, but the mechanism is completely different: it pits early against contemporaneous information precision. This aspect does not seem to have been noted in the literature so far. It suggests that the benefits from transparency can change over time, depending on the prevailing situation. For instance, revealing the expected interest rate path may be undesirable when longer-run uncertainty rises.

Transparency does not just allow a central bank to better (i.e., more credibly) share its information with the private sector; it also gives rise to the mirror effect whereby the central bank also obtains some information back. In our simple model, this means that inflation expectations of the private sector and the central bank are perfectly aligned. The realistic version of this result, which would follow from allowing for more signals, is that transparency lowers the volatility of expected future inflation and therefore the volatility of current inflation. This is a testable proposition, which could narrow down the policy debate when enough observations from the current experiments become available.

A last, fairly obvious result is that transparency is more desirable the better informed is the central bank and the more elastic is output to the long-term interest rate, i.e., the more effective is monetary policy. In other words, the case for transparency is stronger when the central bank is well informed and powerful.

Obviously, we do not address all the arguments against the publication of the interest rate path. Consider, for example, the articulate presentation of the case against transparency by Goodhart (2005):

> If an MPC's non-constant forecast was to be published, there is a widespread view, in most central banks, that it would be taken by the public as more of a commitment, and less of a rather uncertain forecast than should be the case (though that could be mitigated by producing a fan chart of possible interest rate paths, rather than a point estimate: no doubt, though, measuring rulers and magnifying glasses would be used to extract the central tendency). Once there was a published central tendency, then this might easily influence the private sector's own forecasts more than its own inherent uncertainty warranted, along lines analyzed by Morris and Shin (1998, 2002, 2004). Likewise when new, and unpredicted, events occurred, and made the MPC want to adjust the prior forecast path for interest rates, this might give rise to criticisms, ranging from claims that the MPC had made forecasting errors to accusations that they had reneged on a (partial) commitment.

Part of the argument directly refers to Morris and Shin's common-knowledge effect. We do not address this issue here because it has been shown to rest on highly unlikely assumptions. Indeed, it assumes that the central bank is relatively poorly informed (z is low) and that the central bank does not even reveal the current interest rate.[24] Another part of the argument is that releasing the expected interest rate might lock the central bank into setting its interest rate in the future at forecasted level, even though it is no longer desirable given newly available information. This is the classic rules-versus-discretion argument in the presence of time inconsistency, as discussed in Woodford (2005). In our model, time inconsistency is

[24]See Hellwig (2005), Svensson (2005a), and Gosselin, Lotz, and Wyplosz (2008).

eliminated because we do not allow for its two constituent ingredients, the presence of an inflation bias and unknown central bank preferences—two assumptions that we consider unrealistic.[25]

Of course, we too make a large number of assumptions. Some of them, discussed in section 2.3, are simplifications that can be generalized to be more realistic without, we believe, affecting the policy conclusions. We assume that all signal precisions are known. In Gosselin, Lotz, and Wyplosz (2008), in a different setup that focuses on the common-knowledge effect, we show that uncertainty about signal precision carries subtle changes, most of which tend to favor opacity.

Our assumption that the economy starts from and ends at the steady state is not innocuous. In particular, it implies that inflationary expectations are perfectly anchored. Along with a loss function that focuses only on inflation, it implies that the central bank does not aim at a gradual path guiding inflation to its target. Preliminary investigation of an extension of our model to an arbitrary number of periods suggests the following tentative observations. The sharp distinction between periods 1 and 2 would disappear. With it, the result that transparency is always welfare superior in period 2 would be lost. This would work against transparency. On the other hand, in each period the central bank would benefit from the mirror effect as the result of previous publication of the expected interest rate path. This is likely to strengthen the welfare case for transparency.

Appendix

Proof of Equation (14)

Using (11), note that $E_1^{CB}\varepsilon_1 = (r_1 - \nu\varepsilon_{1,2}^{CB})/\mu$ is a signal about ε_1. In period 1, the private sector observes $\frac{r_1 - \nu\varepsilon_{1,2}^{P}}{\mu} = E_1^{CB}\varepsilon_1 + \frac{\nu}{\mu}(\varepsilon_{1,2}^{CB} - \varepsilon_{1,2}^{P})$, which is therefore also a signal about ε_1 available for the private sector with variance $\frac{1}{\alpha} + (\frac{\nu}{\mu})^2(\frac{1}{k\alpha} + \frac{1}{k\beta})$. Similarly, in period 1,

[25] The experience of the Bank of Norway is particularly interesting in this respect. Realizing that credibility is necessary to avoid misinterpretations of the difference between the forecasted and actual interest rate, the Bank of Norway is actively engaged in describing its preferences.

the private sector observes $\frac{r_1-\mu\varepsilon_1^P}{\nu} = \varepsilon_{1,2}^{CB} + \frac{\mu}{\nu}(E_1^{CB}\varepsilon_1 - E_1^{P'}\varepsilon_1)$, which is a signal about ε_2 with variance $\frac{1}{\beta}\left[\frac{1}{kz} + \left(\frac{\mu}{\nu}\right)^2\left(1+\frac{1}{z}\right)\right]$. Using these signals, we can apply Bayes's theorem to obtain

$$E_1^P\varepsilon_1 = \frac{\left[k+(1+z)\left(\dfrac{\nu}{\mu}\right)^2\right]E_1^{P'}\varepsilon_1 + kz\left[E_1^{CB}\varepsilon_1 - \dfrac{\nu}{\mu}\left(\varepsilon_{1,2}^P - \varepsilon_{1,2}^{CB}\right)\right]}{k(1+z)+(1+z)\left(\dfrac{\nu}{\mu}\right)^2}$$

$$E_1^P\varepsilon_1^{CB} = \frac{\left(\dfrac{\nu}{\mu}\right)^2 E_1^{P'}\varepsilon_1 + k\left[E_1^{CB}\varepsilon_1 - \dfrac{\nu}{\mu}\left(\varepsilon_{1,2}^P - \varepsilon_{1,2}^{CB}\right)\right]}{k+\left(\dfrac{\nu}{\mu}\right)^2}$$

$$\begin{aligned}E_1^P\varepsilon_2 &= \frac{\left[k(1+z)+\left(\dfrac{\nu}{\mu}\right)^2\right]\varepsilon_{1,2}^P + z\left(\dfrac{\nu}{\mu}\right)^2\left[\varepsilon_{1,2}^{CB} - \dfrac{\mu}{\nu}\left(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right]}{(1+z)\left(k+\left(\dfrac{\nu}{\mu}\right)^2\right)}\\ &= \gamma_1^{op}\varepsilon_{1,2}^P + \left(1-\gamma_1^{op}\right)\left[\varepsilon_{1,2}^{CB} - \frac{\mu}{\nu}\left(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right],\end{aligned}$$

which defines $\gamma_1^{op} = \frac{k(1+z)+\left(\frac{\nu}{\mu}\right)^2}{(1+z)\left(k+\left(\frac{\nu}{\mu}\right)^2\right)}$.

$$E_1^P\varepsilon_2^{CB} = \frac{k\varepsilon_{1,2}^P + \left(\dfrac{\nu}{\mu}\right)^2\left[\varepsilon_{1,2}^{CB} - \dfrac{\mu}{\nu}\left(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right]}{k+\left(\dfrac{\nu}{\mu}\right)^2}.$$

It follows that

$$E_1^P\varepsilon_1 - E_1^P\varepsilon_1^{CB} = \frac{E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1 + \dfrac{\nu}{\mu}\left(\varepsilon_{1,2}^P - \varepsilon_{1,2}^{CB}\right)}{(1+z)\left(k+\left(\dfrac{\nu}{\mu}\right)^2\right)}.$$

Recalling (12) and using (3), (6), and (9), we can now compute π_1, which is necessary to obtain the signal extracted by the central bank at time 2:

$$\begin{aligned}\pi_1 &= (1+\kappa)\big(E_1^P\varepsilon_2 - E_1^P E_2^{CB}\varepsilon_2\big) - E_1^P E_2^{CB}\varepsilon_2 - \kappa r_1 + \varepsilon_1\\ &= (1+\kappa)\big(\gamma_1^{op} - \gamma_2^{op}\big)\left[\varepsilon_{1,2}^P - \left(\varepsilon_{1,2}^{CB} - \frac{\mu}{\nu}\big(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\big)\right)\right]\\ &\quad - \kappa r_1 - \left[\gamma_2^{op}\varepsilon_{1,2}^P + \big(1-\gamma_2^{op}\big)\left(\varepsilon_{1,2}^{CB} - \frac{\mu}{\nu}\big(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\big)\right)\right] + \varepsilon_1.\end{aligned}$$

This expression can be rewritten as

$$\frac{\pi_1 + \kappa r_1 - \varepsilon_1}{(1+\kappa)\big(\gamma_1^{op} - \gamma_2^{op}\big) - \gamma_2^{op}} + \theta\varepsilon_{1,2}^{CB} = \varepsilon_{1,2}^P + \theta\frac{\mu}{\nu}\big(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1\big),$$

where we have introduced an auxiliary variable $\theta = 1 + \frac{1}{(1+\kappa)(\gamma_1^{op}-\gamma_2^{op})-\gamma_2^{op}}$. Now note that π_1 and ε_1 become known in period 2 (and r_1 is always known). It follows that the right-hand side in the previous expression is known to the central bank when period 2 starts and it can be used as a signal about ε_2. However, the central bank can improve this signal by replacing $E_1^{CB}\varepsilon_1$ with ε_1 so that the signal about ε_2 is now $\varepsilon_{1,2}^P + \theta\frac{\mu}{\nu}(E_1^{P'}\varepsilon_1 - \varepsilon_1)$, with variance $\frac{1}{\beta}\left(\frac{1}{k} + \theta^2\left(\frac{\mu}{\nu}\right)^2\right)$.

We next use Bayes's rule to find $E_1^P E_2^{CB}\varepsilon_2$. The relevant computation leads to

$$E_2^{CB}\varepsilon_2 = \frac{\left[\left(\frac{\mu}{\nu}\right)^2\theta^2 zk + z\right]\left[k\varepsilon_{1,2}^{CB} + (1-k)\varepsilon_{2,2}^{CB}\right] + k\left[\varepsilon_{1,2}^P + \frac{\mu}{\nu}\theta\big(E_1^{P'}\varepsilon_1 - \varepsilon_1\big)\right]}{\left(\frac{\mu}{\nu}\right)^2\theta^2 zk + z + k}$$

so that the compounded expectation is given by

$$E_1^P E_2^{CB}\varepsilon_2 = \frac{\left[\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z\right]\left[kE_1^P\varepsilon_{1,2}^{CB} + (1-k)E_1^P\varepsilon_{2,2}^{CB}\right]}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k}$$
$$+ k\frac{\varepsilon_{1,2}^P + \frac{\mu}{\nu}\theta(1-\gamma_1)\left[\left(E_1^{P'}\varepsilon_1 - \varepsilon_1^{CB}\right) + \left(\frac{\nu}{\mu}\right)\left(\varepsilon_{1,2}^P - \varepsilon_{1,2}^{CB}\right)\right]}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k}.$$

Using the expressions for the various private-sector expectations, we can deduce by identification

$$\gamma_2^{op} = \left[\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z\right]\frac{\frac{k^2}{k+\left(\frac{\nu}{\mu}\right)^2} + (1-k)\frac{k(z+1)+\left(\frac{\nu}{\mu}\right)^2}{k(z+1)+(z+1)\left(\frac{\nu}{\mu}\right)^2}}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k}$$
$$+ k\frac{1+\theta\left(\frac{x\left(\frac{\nu}{\mu}\right)^2}{k(z+1)+(z+1)\left(\frac{\nu}{\mu}\right)^2}\right)}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k}$$

from which we find (14).

Proof of Proposition 2

The parameters for r_1 are found by minimizing the unconditional loss function $E(\pi_1)^2 + E(\pi_2)^2$.[26] Using (14), the previous expression for π_1 can be rewritten as

$$\pi_1 = \frac{1}{\theta-1}\left(\varepsilon_{1,2}^P - \varepsilon_2\right) + \frac{\mu}{\nu}\frac{\theta}{\theta-1}\left(E_1^{P'}\varepsilon_1 - \varepsilon_1\right)$$
$$+ \left(\kappa\nu + \frac{\theta}{\theta-1}\right)\left[\left(\varepsilon_2 - \varepsilon_{1,2}^{CB}\right) + \frac{\mu}{\nu}\left(\varepsilon_1 - E_1^{CB}\varepsilon_1\right)\right]$$
$$+ (1-\kappa\mu)\varepsilon_1 - (1+\kappa\nu)\varepsilon_2,$$

[26]It is unconditional because, if it were conditional on central bank information, the coefficients μ and ν would be nonlinear functions of $E_1^{CB}\varepsilon_1$ and $E_1^{CB}\varepsilon_2$, so the rule would not be linear—and impossible to derive in closed form.

which implies that

$$E(\pi_1)^2 = (1-\kappa\mu)^2 E(\varepsilon_1)^2 + (1+\kappa\nu)^2 E(\varepsilon_2)^2 + \text{other terms},$$

where the other terms depend on k, $z = \frac{\alpha}{\beta}$, μ, and ν.

Similarly, note that $\pi_2 = \varepsilon_2 - E_2^{CB}\varepsilon_2$ and that $E_2^{CB}\varepsilon_2$ is optimally found by the central bank by using the signals $\varepsilon_{1,2}^{CB}$, $\varepsilon_{2,2}^{CB}$, and $\varepsilon_{1,2}^{P} + \theta\frac{\mu}{\nu}(E_1^{P'}\varepsilon_1 - E_1^{CB}\varepsilon_1)$ as indicated above, which gives

$$\begin{aligned} E_2^{CB}\varepsilon_2 = {} & \left[\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z\right] \frac{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z\left[k\varepsilon_{1,2}^{CB} + (1-k)\varepsilon_{2,2}^{CB}\right]}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k} \\ & + k\frac{\varepsilon_{1,2}^{P} + \frac{\mu}{\nu}\theta\left(E_1^{P'}\varepsilon_1 - \varepsilon_1\right)}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k} \end{aligned}$$

so that

$$\begin{aligned} \pi_2 = {} & \left[\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z\right] \frac{k\left(\varepsilon_2 - \varepsilon_{1,2}^{CB}\right) + (1-k)\left(\varepsilon_2 - \varepsilon_{2,2}^{CB}\right)}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k} \\ & + k\frac{\left(\varepsilon_2 - \varepsilon_{1,2}^{P}\right) - \dfrac{\mu}{\nu}\theta\left(E_1^{P'}\varepsilon_1 - \varepsilon_1\right)}{\left(\frac{\mu}{\nu}\right)^2 \theta^2 zk + z + k} \end{aligned}$$

and $E(\pi_2)^2$ only includes terms in k, z, μ, and ν. It follows that the total unconditionally expected loss under opacity can be written as

$$L^{op} = (1-\kappa\mu)^2 E(\varepsilon_1)^2 + (1+\kappa\nu)^2 E(\varepsilon_2)^2 + \text{other terms}.$$

Since both ε_1 and ε_2 are assumed to be uniformly distributed, $E(\varepsilon_1)^2$ and $E(\varepsilon_2)^2$ are arbitrarily large relative to the other terms—in particular, the variances α^{-2} and β^{-2}. It follows that the rule that minimizes L^{op} sets these terms equal to zero. Using the expression for π_2, we find the unconditional expectation $E(\pi_2)^2$, which measures the second period loss:

$$E(\pi_2)^2 = \frac{1+\theta^2 k}{\theta^2 zk + z + k}.$$

Proof of Proposition 4

The study of (19) shows that $\beta\Delta L_1(\theta)$ reaches a minimum of $\frac{kz^2-(1+k)^2}{kz(1+k)(1+z)}$ when $\theta = -\frac{1+k+z}{kz}$. This minimum is positive when $z > \frac{1+k}{\sqrt{k}}$.

The Sign of $\gamma_1^{op} - \gamma_2^{op}$

Using the optimality condition $\frac{\mu}{\nu} = -1$, the parameters γ_2^{op} and θ are jointly determined by the two following equations:

$$\theta = 1 + \frac{1}{(1+\kappa)\left(\gamma_1^{op} - \gamma_2^{op}\right) - \gamma_2^{op}} = 1 + \frac{1}{(2+\kappa)\left(\gamma_1^{op} - \gamma_2^{op}\right) - \gamma_1^{op}}$$

$$\gamma_1^{op} - \gamma_2^{op} = \frac{z\theta k(\theta k - 1)}{(1+z)(1+k)(\theta^2 zk + z + kt)}.$$

Defining $x = (1+z)(1+k)(\gamma_1^{op} - \gamma_2^{op})$, we can rewrite these two equations as

$$\theta = 1 + \frac{1}{(2+\kappa)\dfrac{x}{(1+z)(1+k)} - \gamma_1^{op}} = \frac{((2+\kappa)x + z)}{(2+\kappa)x - [k(1+z)+1]}$$

$$x(\theta^2 zk + z + k) = z\theta k(\theta k - 1).$$

They can be combined to yield the following third-order equation in x:

$$A_3x^3 + A_2x^2 + A_1x + A_0 = 0,$$

where

$$\begin{aligned}
A_3 &= kz(2+\kappa)^2 + (k+z)(2+\kappa)^2 \\
A_2 &= 2kz^2(2+\kappa) - 2(k+z)(2+\kappa)[1 + k(z+1)] \\
&\quad - kz(2+\kappa)[k(2+\kappa) - 2 - \kappa] \\
A_1 &= kz^3 + (k+z)(-k(z+1)-1)^2 - z^2k[k(2+\kappa) - 2 - \kappa] \\
&\quad - kz(2+\kappa)[zk + k(z+1) + 1] \\
A_0 &= -z^2k(zk + k(z+1) + 1).
\end{aligned}$$

The graphical study of the solutions of this equation shows that, for κ not too large (the threshold exceeds any realistic value of κ), this equation admits a single noncomplex solution x, which is always positive.[27] This establishes our claim that $\gamma_1^{op} - \gamma_2^{op} > 0$.

For further reference, we have the following limit conditions:

- When $z \to 0$, $x \to 0$ (as $\frac{z^2}{(k+1)}$).
- When $z \to \infty$, $x \to 2\frac{k}{k+1}$ and $\gamma_1^{op} - \gamma_2^{op} \to 0$.
- When $k \to 0$, $x \to 0$ (as kz).
- When $\kappa \to \infty$, there are three possibilities, all of which imply that $\gamma_1^{op} - \gamma_2^{op} < 0$: $x \to 0$ (as $\frac{-z}{k}$), $x \to \frac{-zk(1-k)}{zk+z+k}$, or $x \to -\frac{k+2kz+1}{-2-\kappa+2k+k\kappa}$.

Proof of Proposition 5

Remember first that welfare in period 2 is always higher under transparency because it provides the central bank with more information and therefore a more precise estimate of the period shock. For opacity to welfare dominate transparency, therefore, it must reduce inflation volatility in period 1 by enough to offset the welfare loss of period 2. From (3) we know that period 1 inflation is driven by expected inflation and the extent to which the central bank fails to stabilize output in period 1. Using (10), (6), and (15), (3) can be rewritten as

$$\pi_1 = (1+\kappa)E_1^P\pi_2 - \kappa\big(r_1 + E_1^P r_2\big) + \varepsilon_1 = (2+\kappa)E_1^P\pi_2 - \psi. \quad (20)$$

The "policy miss" term $\psi = \kappa r_1 - (\varepsilon_1 - E_1^P\varepsilon_2)$ measures the private sector's perception of the extent to which the central bank fails to achieve its period 1 objective when it optimally chooses r_1. Without any information about the private signals, the central bank's best forecast of $E_1^P\pi_2$ is $E_1^{CB}E_1^P\pi_2 = 0$, which explains (10). The policy miss term can be rewritten as

$$\psi = \big(E_1^{CB}\varepsilon_1 - \varepsilon_1\big) + \big(E_1^P\varepsilon_2 - E_1^{CB}\varepsilon_2\big). \quad (21)$$

[27] When κ is above this threshold and z is not too large, we have three real solutions for x, two of which are negative. When z becomes large, again, there is a unique real solution, which is positive.

Let us now compare (20) under the two regimes. Under transparency, but not under opacity, $E_1^P \pi_2 = 0$, so opacity tends to add volatility to period 1 inflation, the more so the higher is κ. Furthermore, we can show that $Var^{op}(\psi) > Var^{tr}(\psi)$. This is quite intuitive: the first term in (21), the period 1 signal error, is regime invariant, while the second term reflects disagreements between the central bank and the private sector. When it announces $E_1^{CB} r_2$, the central bank fully reveals $E_1^{CB} \varepsilon_2$ and therefore moves $E_1^P \varepsilon_2$ toward $E_1^{CB} \varepsilon_2$.

It follows that opacity always raises period 1 inflation as well, unless $cov(E_1^P \pi_2, \psi)$ under opacity is positive and large enough to offset the other two effects. Thus $cov(E_1^P \pi_2, \psi) > 0$ is a necessary, but not sufficient, condition for opacity to raise welfare. Since $cov(E_1^P \pi_2, \psi) = (\gamma_1^{op} - \gamma_2^{op}) \frac{k+1}{\alpha k} > 0$, as shown above, this condition is always satisfied. What is needed, therefore, is that $cov(E_1^P \pi_2, \psi) > 0$ be large enough. This is the case when z, k, and κ are small.

Proof of Proposition 6

Using the loss functions given in the text and using $\gamma_1^{op} = \frac{k(1+z)+1}{(1+z)(k+1)}$, we have

$$\begin{aligned}\beta \Delta L = {} & \frac{1}{((z+1)(k+1))^2} \left(((2+\kappa)x - (k(z+1)+1))^2 \frac{k+1+z}{kz} \right. \\ & \left. + ((2+\kappa)x + z)^2 \right) - \frac{1+k(1+z)}{kz(1+z)} \\ & + \frac{k}{z(z+k)} \frac{((2+\kappa)x+z)}{(2k + k\kappa - 2 - \kappa)x + k + 2zk + 1} x,\end{aligned}$$

where x, defined in this appendix (in the study of the sign of $\gamma_1^{op} - \gamma_2^{op}$), is the solution of a polynomial of degree 3. We can also write

$$\beta \Delta L = \frac{P(x)}{-(1-k)(2+\kappa)x + k + 2zk + 1},$$

with $P(x)$ of degree 2 where the coefficient of x^2 is positive $\forall$ k, z, κ. We do not specify the form of $P(x)$, since we will only need some of properties of this function, which we study graphically using Mapple. Note that x is a solution of the third-degree polynomial shown above in this appendix. For any such x, the sign of $-(1-k)(2+\kappa)x+k+2zk+1$ is always positive (the hypersurfaces defined by the third-degree equation for x and by $-(1-k)(2+\kappa)x+k+2zk+1=0$ do not intersect). As a consequence, the sign of ΔL reduces to the sign of the second-degree polynomial $P(x)$. Denoting $\Delta(k,z,\kappa)$ the discriminant of $P(x)$, we draw the following conclusions:

- If $\Delta < 0$, then $\Delta L > 0$ whatever x, the solution of the third-order equation, and transparency dominates.
- If $\Delta > 0$ and x, the solution of the third-order equation, lies outside the roots of P, then $\Delta L > 0$ and transparency dominates.
- If $\Delta > 0$ and x, the solution of the third-order equation, lies inside the roots of P, then $\Delta L < 0$ and opacity dominates.

The study of ΔL consists then in checking whether or not the roots of the third-order equation lie between the roots of $P(x)$. The results are presented graphically in the (k,z) plane in figure 2, which displays $\Delta(k,z,\kappa)=0$. Above this curve, $\Delta(k,z,\kappa) < 0$ and transparency dominates. The shape of the opacity zone below $\Delta(k,z,\kappa)=0$ has been determined from a graphical three-dimensional analysis using Mapple and is therefore not precisely known. The figures are also informed by the analytic study of the following limit cases:

- For $z \to 0$, $\Delta L < 0$.
- For $z \to \infty$, $\Delta L > 0$.
- For $k \to 0$, $\Delta L > 0$ when $z - 2\kappa - 4 > 0$, and $\Delta L < 0$ otherwise.

Proposition 6 states that, as κ increases, the curve $\Delta L = 0$ shifts up when k is small and down when k is large. This is not entirely accurate. The graphical analysis indicates that when κ is small, the curve shifts upward $\forall k$. For economically relevant values of κ, however, the statement in proposition 6 is accurate.

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