



INTERNATIONAL JOURNAL OF CENTRAL BANKING

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Communication, Decision Making, and the Optimal Degree of Transparency of Monetary Policy Committees*

Anke Weber
International Monetary Fund

This paper develops a theoretical model of a monetary policy committee with heterogeneous members whose decisions and public communications are observed by the financial markets. It thereby provides a link between the literatures on monetary policy committees and central bank communication. The results show that transparency about the different views among committee members surrounding the economic outlook is beneficial. However, communicating the diversity of views about the monetary policy decision may not be welfare enhancing, at least in the short term. These results support previous empirical findings and have strong implications for how committees should communicate.

JEL Codes: E50, E52, E58.

1. Introduction

The conduct of monetary policy has changed markedly since the 1990s. Over the past two decades, a number of central banks have shifted responsibility for interest rate setting to a monetary policy

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committee (MPC).¹ In addition, central banks have become more independent and now pay close attention to explaining what they do and what underlies their decisions. Greater transparency and increased communication are consequences of these developments (de Haan, Eijffinger, and Rybinski 2007).²

This paper provides a link between the literatures on monetary policy committees and central bank communication, both of which are important elements in modern central bank design. It develops a theoretical model of a monetary policy committee whose decisions and public communications are observed by the financial markets, and thereby analyzes how committees should communicate. The model contrasts in several ways with previous work on central bank transparency and communication. First, decisions are set by a monetary policy committee with heterogeneous members rather than by a representative central banker. It is assumed that policymakers face uncertainty in assessing the state of the economy.³ They receive different signals on the state of the economy and face uncertainty about the precision of these observations. They also have different preferences regarding the optimal level of the output gap. In addition, two sources of asymmetric information between the committee and market participants are introduced. It is assumed that market participants have imperfect information about the precision of public information on the state of the economy and about the policy preferences of committee members.

Several interesting insights are obtained. It is shown that there is a decreasing marginal benefit of increasing the committee size. It is also found that communication matters for financial markets' expectations of inflation and future short-term interest rates. Committee members should communicate their views on the state of the economy without error, as this reduces inflation and output-gap

¹A recent survey by Pollard (2004) shows that seventy-nine out of eighty-eight central banks conduct monetary policy by committee.

²Eijffinger and Geraats (2006) provide an index of transparency for a set of developed countries that includes some inflation targeters (United Kingdom, Sweden, Australia, and Canada) as well as non-targeters (Japan, United States, and Switzerland). They show that between 1998 and 2002 transparency increased for virtually all of the central banks they studied.

³The fact that central banks do not know the current state of the economy with certainty has been stressed by Orphanides (2003) and Aoki (2006).

variability. Communication of a divergence of views by committee members regarding the monetary policy decision (as, for example, the publication of voting records) may lead to greater variability in inflation and the output gap in the short run. In the long run, welfare is enhanced, as financial markets are able to infer members' preferences from voting records once they have a better estimate of the precision of public information.

This paper contributes to the previous theoretical and empirical literature on monetary policy committees and central bank communication.⁴ Blinder (2007) argues that delegating monetary policy to a committee leads to superior policy for a number of reasons, such as the ability to pool judgments of different members and the possibility of learning from other members. The finding that committees make superior decisions to individual policymakers has been confirmed by the experimental literature (Blinder and Morgan 2005; Lombardelli, Proudman, and Talbot 2005). Furthermore, the theoretical literature has concluded that a committee of limited size is optimal due to the costs of acquiring information about different members and coordination costs (Sibert 2006). This paper argues that committee members are not efficient at pooling different information sets when they do not know their precision and that this implies that there are decreasing marginal benefits of increasing the committee size.

An increasing number of papers have examined the optimal communication strategy of central banks. Some economists stress the importance of communication in providing central banks with the means to influence key asset prices in the economy (Blinder 1998; Bernanke 2004). However, others, such as Morris and Shin (2002), argue that too much communication may not be desirable. Morris and Shin (2002) develop a model in which welfare decreases when private agents rely on a sufficiently noisy public signal to coordinate their actions and thus disregard their own private information.

Ehrmann and Fratzscher (2005) investigate communication by three major central banks (the Bank of England, the Federal Reserve, and the European Central Bank) empirically and evaluate the effectiveness of central bank communication on financial

⁴Vandenbussche (2006) and Blinder (2007) provide comprehensive overviews of the literature on central bank committees whilst Blinder et al. (2008) provide a survey of the literature on central bank communication.

markets in terms of their ability to anticipate monetary policy decisions. They conclude that a higher degree of communication dispersion among committee members about monetary policy worsens the ability of financial markets to anticipate future monetary policy decisions. They also find that communication of the risks and diversity of views of the committee surrounding the economic outlook enhances the ability of the public to anticipate future monetary policy decisions, at least in the case of the Federal Reserve. Gerlach-Kristen (2004), on the other hand, investigates the voting records of committee members at the Bank of England and finds that those records can convey information about the views of individual members. Their publication may hence lead to greater policy predictability. The present paper is best seen as providing some theoretical underpinning to explain those previous empirical findings.

The remainder of this paper is structured as follows. Section 2 sets out the basic model. The solution of this model is discussed in section 3. Section 4 presents the results, and conclusions are summarized in section 5.

2. The Model

2.1 *The General Setup*

Inflation is determined by a New Keynesian Phillips-curve relationship:

$$\pi_t = E_t \pi_{t+1} + y_t + s_t, \quad (1)$$

where π_t denotes inflation and y_t is the output gap defined as the difference between actual and potential output, where the latter is conveniently normalized to 0. For convenience, the coefficients of the New Keynesian Phillips curve are set to 1. This does not affect the qualitative results. In addition, s_t is a persistent supply shock, which follows a simple AR(1) process:

$$s_t = \alpha s_{t-1} + v_t, \quad (2)$$

where $0 < \alpha < 1$ and v_t is i.i.d. white noise with variance σ_v^2 .

Following Kozicki and Tinsley (2008), the output gap is determined by the following simple relationship:⁵

$$y_t = -R_t, \quad (3)$$

where R_t denotes the long-term real interest rate. The assumption that the output gap is decreasing in the long-term real interest rate can be justified by the fact that the interest-rate-sensitive components of aggregate demand generally depend on the yield on financial assets with longer maturity (Eijffinger, Schaling, and Tesfaselassie 2004).

To facilitate the derivation of closed-form results, we replace the n -period real interest rate in (3) with a two-period rate. The two-period real interest rate, R_t , and the current and future short-term real interest rates, $i_t - E_t\pi_{t+1}$ and $E_t i_{t+1} - E_t\pi_{t+2}$, are related by the expectations theory of the term structure:⁶

$$R_t = \frac{1}{2}(i_t - E_t\pi_{t+1}) + \frac{1}{2}(E_t i_{t+1} - E_t\pi_{t+2}). \quad (4)$$

Replacing the n -period bond rate with a two-period bond rate could be interpreted as describing a situation in which the central bank deals with the current situation given expectations about the near future (two periods ahead) while too little is known about the long run (Gosselin, Lots, and Wyplosz 2008).

The central bank controls the short-term nominal interest rate, i_t , and has exogenously given inflation and output-gap targets, π^* and y^* . It also wants to minimize deviations of inflation and the output gap from those assigned targets. The central bank can observe and respond directly to private-sector expectations, and it sets its interest rate after private expectations are set in every period. This timing implies that monetary policy is set under discretion. Since i_t affects only y_t and π_t and there are no endogenous state variables, the central bank will set interest rates to minimize the following loss function:

⁵This equation can be derived from a forward-looking IS curve of the form $y_t = E_t y_{t+1} - (i_t - E_t\pi_{t+1})$. After recursive forward substitution, the relationship can be approximated as $y_t \simeq -E_t[\frac{1}{n} \sum_{i=1}^n i_{t+i} - \frac{1}{n} \sum_{i=1}^n \pi_{t+i+1}]$, where i_t denotes the short-term nominal interest rate.

⁶Without loss of generality, the unobserved term premium is set to 0.

$$L = \frac{1}{2}[(\pi_t - \pi^*)^2 + \mu(y_t - y^*)^2], \quad (5)$$

where $0 < \mu < 1$ is the relative weight on output-gap stabilization. In order to find the optimal reaction function, (5) is minimized subject to (1), (2), (3), and (4). The central bank's reaction function can be written as

$$\begin{aligned} i_t = & \left(1 + \frac{2}{1 + \mu}\right) E_t \pi_{t+1} + E_t \pi_{t+2} - E_t i_{t+1} \\ & + \frac{2}{1 + \mu}(s_t - \pi^*) - \frac{2\mu}{1 + \mu} y^*. \end{aligned} \quad (6)$$

Hence, i_t is increasing in the current supply shock because it determines current inflation. Optimal policy is also inversely related to financial markets' forecast of the future short-term real rate, since the higher $E_t i_{t+1} - E_t \pi_{t+2}$, the lower the optimal short-term real interest rate can be. The higher the relative weight that the central bank gives to output-gap stabilization, μ , the smaller is the response of i_t to changes in expected inflation and the aggregate supply shock.

The two-period real rate can be expressed as

$$R_t = \frac{1}{1 + \mu}(E_t \pi_{t+1} + s_t - \pi^*) - \frac{\mu}{1 + \mu} y^*,$$

where equations (4) and (6) have been used. Intuitively, the two-period real interest rate is a positive function of expected inflation and the supply shock and a negative function of the inflation and output-gap targets.

The method of undetermined coefficients is used to obtain the public's rational expectation of inflation as⁷

$$E_t \pi_{t+1} = \frac{\mu}{1 + \mu(1 - \alpha)} \alpha s_t + \pi^* + \mu y^*. \quad (7)$$

Private-sector inflation expectations are increasing in the current supply shock and the inflation and output-gap targets of the central

⁷The solution method is outlined in the appendix.

bank. Substituting (4) and (7) into (1) and (3) yields expressions for inflation and the output gap:

$$\pi_t = \frac{\mu}{[1 + \mu(1 - \alpha)]} s_t + \pi^* + \mu y^* \quad \text{and} \quad y_t = -\frac{1}{[1 + \mu(1 - \alpha)]} s_t.$$

Hence inflation is increasing in the supply shock and the central bank's targets for inflation and the output gap. Moreover, it is a positive function of μ . Therefore, the greater the preference of the central bank for output-gap stabilization, the stronger is the response of inflation to supply shocks. If $s_t > 0$, the output gap is also a positive function of the central bank's preference for output-gap stabilization. Furthermore, it is a negative function of the supply shock.

2.2 *Imperfect Information about the State of the Economy*

The central bank and the public are assumed to face uncertainty about the current supply shock but to have their own best estimate of s_t . Sections 3.1 and 3.2 show how such an estimate can be derived. The central bank and the private sector know the structural equations of the economy, and the central bank continues to observe perfectly the public's inflation expectations, which are set before the central bank makes an interest rate decision. This assumption is made in order to be able to derive closed-form solutions of the model under uncertainty about the state of the economy and asymmetric information between the central bank committee and the financial markets. If policy is set under discretion, then the committee does not need to consider the effect of its short-term interest rate decision on financial markets' expectations of future inflation and interest rates, since these are already set. As Svensson and Woodford (2004) show, under asymmetric information between the central bank and the public and uncertainty about the state of the economy, certainty equivalence continues to hold under discretion; that is, the committee can continue to set interest rates as it would under certainty about the supply shock but use its best estimate of the supply shock instead. Furthermore, the signal-extraction problem of finding an optimal estimate of the supply shock can be treated separately from the optimization problem of finding an optimal response coefficient for the reaction function of the committee. This is because

the best estimate of the supply shock does not depend on the current nominal interest rate, inflation, or output and thus a separation principle applies (Svensson and Woodford 2004). Furthermore, financial markets use their best estimate of the current supply shock and ignore the uncertainty surrounding this estimate when forming expectations about future inflation and interest rates.

It is assumed that in the final stage of each period t , the supply shock is realized and observed both by the central bank and the public. This assumption is made in order to simplify the analysis. Whilst it would be more reasonable to assume that a new and better estimate of the supply shock becomes available at the end of each period, this would make the signal-extraction procedure significantly more complicated.

Using (6), the short-term nominal interest rate under uncertainty is given by

$$i_t = \left(1 + \frac{2}{1 + \mu}\right) E_t \pi_{t+1} + E_t \pi_{t+2} - E_t i_{t+1} + \frac{2}{1 + \mu} (s_{tit}^{(CB)} - \pi^*) - \frac{2\mu}{1 + \mu} y^*, \quad (8)$$

where s_{tit}^{CB} denotes the optimal estimate of the central bank for the supply shock in period t .

Inserting the central bank's reaction function (8) into (4) and using (7), the two-period real interest rate under uncertainty can be written as

$$R_t = \frac{\mu\alpha}{(1 + \mu)[1 + \mu(1 - \alpha)]} s_{tit}^{(F)} + \frac{1}{1 + \mu} s_{tit}^{(CB)}, \quad (9)$$

where s_{tit}^F denotes the optimal estimate of the financial markets for the supply shock in period t .

Using (7) and (9), inflation and the output gap in period t will equal

$$\pi_t = \frac{\mu^2\alpha}{(1 + \mu)[1 + \mu(1 - \alpha)]} s_{tit}^{(F)} - \frac{1}{1 + \mu} s_{tit}^{(CB)} + \pi^* + \mu y^* + s_t \quad (10)$$

and

$$y_t = -\frac{\mu\alpha}{(1 + \mu)[1 + \mu(1 - \alpha)]} s_{tit}^{(F)} - \frac{1}{1 + \mu} s_{tit}^{(CB)}. \quad (11)$$

Thus, inflation continues to increase with higher central bank targets for inflation and the output gap. It is a positive function of the true supply shock and the estimated supply shock by the public. The higher the perceived supply shock by the central bank, the lower will be inflation. This is because the higher the current estimate of the supply shock by the central bank, the higher will be the short-term nominal interest rate that it sets. The output gap, on the other hand, is decreasing in both the central bank's and the public's estimates of the supply shock. This paper assumes that $\alpha > 0$ and $\mu > 0$. From equations (10) and (11) it can be seen why these assumptions are made. If $\alpha = 0$, then the private sector's estimate of the supply shock has no effect on inflation or the output gap. In this case, if the inflation and output-gap targets of the central bank are known to the public, there is no role for communication by committee members regarding their views on the innovation to the supply shock. The same is true for $\mu = 0$. In this case, there is also no effect of the output-gap target on inflation. Thus, for the results in this paper to hold, the supply shock should be persistent and the central bank should not be a strict inflation targeter. Empirical studies typically confirm that supply shocks are persistent (Ireland 2004). In addition, whether they explicitly target inflation or not, the central banks of most developed economies act as flexible inflation targeters (Cukierman 2007).

2.3 *Transparency and the Monetary Policy Committee*

This section introduces a model of a monetary policy committee with N members whose decisions and public communications are observed by the financial markets. Each committee member, j , receives a signal on the innovation, v_t , to the supply shock, s_t :

$$v_t^{(j)} = v_t + \varepsilon_t^{(j)}, \quad (12)$$

where $\varepsilon_t^{(j)}$ is i.i.d. with variance $\sigma_{\varepsilon,j}^2$ for $j = 1, 2, \dots, N$. Thus $E(v_t^{(j)} - v_t)^2 = \sigma_{\varepsilon,j}^2$.

It is assumed that the error terms of those signals are uncorrelated among committee members. Furthermore, their true variance is unknown. The precision of committee member j 's observation is estimated in period t by member k as $\tilde{\sigma}_{\varepsilon,t}^{2(k,j)}$ for $j = 1, \dots, N$ and

$k = 1, \dots, N$. Whilst committee members may receive the same information on the state of the economy, they interpret and process this information differently, resulting in different observations on the state of the economy. Furthermore, committee members are likely to be uncertain about the correct underlying model of the economy, and this is reflected in different variances of observation errors and uncertainty about the precision of information.

In addition, committee members are modeled as having different preferences regarding the level of the output-gap target, y^* . Even though the objectives of a central bank are often assigned by law, these are usually defined in terms of a specific inflation target, whereas the objectives regarding the optimal level of the output target are less precisely defined (Cukierman 2007). Thus, it is likely that committee members have different views about the optimal level of the output-gap target, and this is confirmed by the empirical literature (Chappell, McGregor, and Vermilyea 2005). The paper assumes that these output-gap preferences are constant and denoted by $\theta^{(j)}$ for member j . This may not be the case in reality. Cukierman (2007) argues that these preferences fluctuate due to changes in the intensity of political pressures on members. However, the assumption that the output-gap target preferences are constant greatly simplifies the analysis of the paper. If they were following an autoregressive process as in Faust and Svensson (2002), this would add another signal-extraction problem to the paper. However, the basic conclusions should still hold.

There are two important information asymmetries between the monetary policy committee and financial markets. First, financial markets do not observe the signals that individual committee members receive on the innovation to the supply shock. Instead, financial markets receive public information on the state of the economy from committee members—for instance, in the form of speeches. This communication takes the following form for each committee member j :

$$x_t^{(j)} = v_t^{(j)} + \varpi_t^{(j)}, \quad (13)$$

where $\varpi_t^{(j)}$ is i.i.d. with variance $\sigma_{\varpi,j}^2$ for $j = 1, \dots, N$. It is assumed that $E_t(\varpi_t^{(F,j)}, \varpi_t^{(F,k)}) = 0$ for $j \neq k$. Thus, each committee member, j , communicates his view on the innovation to the supply shock

to the financial markets, but may do so imperfectly, which is reflected in the error term $\varpi_t^{(j)}$. When $\sigma_{\varpi,j}^2 = 0$, the signals $x_t^{(j)}$ communicate $v_t^{(j)}$ without noise and there is perfect actual transparency about committee member j 's observation on the innovation to the supply shock. In reality, it will be difficult for the private sector to establish how transparent the central bank is, and it is unlikely that the private sector will know the variance of the communication error, $\sigma_{\varpi,j}^2$. Thus, the paper follows Geraats (2007) and introduces the notion of perceived transparency. Committee member j is perceived to be perfectly transparent about his signal on the innovation to the supply shock if $\tilde{\sigma}_{\varpi,t}^{2(j)} = 0$.

Equation (13) can alternatively be written as

$$x_t^{(j)} = v_t + \varepsilon_t^{(j)} + \varpi_t^{(j)} = v_t + \zeta_t^{(j)}, \quad (14)$$

where the variance of $\zeta_t^{(j)}$ equals $Var(\zeta_t^{(j)}) = \sigma_{\varepsilon,j}^2 + \sigma_{\varpi,j}^2$ for $j = 1, 2, \dots, N$. There is no communication about the estimated precision of signals by committee members, and financial markets are uncertain about the variance of $\varpi_t^{(j)}$. The variance of $\zeta_t^{(j)}$ is estimated by financial markets as $\tilde{\sigma}_{\zeta,t}^{2(j)} = \tilde{\sigma}_{\varepsilon,t}^{2(F,j)} + \tilde{\sigma}_{\varpi,t}^{2(j)}$.

Financial markets also have imperfect knowledge about the policy preferences of committee members. As Beetsma and Jensen (1998) argue, policymakers are likely to have private information about their preferences. Financial markets have a prior on the preferred output-gap targets of policymakers.⁸ Denote this prior for the preferred output-gap target of committee member j as $z^{(j)} = \theta^{(j)} + \psi^{(j)}$, where $\psi^{(j)}$ is i.i.d. with variance $\sigma_{z,j}^2$ for $j = 1, 2, \dots, N$. The variance of this prior is estimated by the financial markets as $\tilde{\sigma}_{z,j}^2$.

It should be noted that the paper follows the definition of Geraats (2002) and defines transparency as the absence of asymmetric information between policymakers and financial markets. The paper focuses exclusively on what Geraats (2006) refers to as the information effect of transparency—that is, the effect of giving the private sector new information to act upon. It does not investigate

⁸This prior could have been derived as a result of some previous communication by committee members about their preferences.

the indirect incentive effects of transparency, whereby the committee members would adjust their behavior according to which information is disclosed. These incentive effects are particularly relevant for the question of whether voting records should be published. There certainly is a concern that if the European Central Bank were to publish voting records, this would induce members of the Governing Council to vote for the optimal interest rate of the country they represent (Issing 1999; Gersbach and Hahn 2005). Strategic voting behavior has received considerable attention in the theoretical literature.⁹ However, policymakers themselves do not believe that strategic considerations play an important role in policy meetings. For instance, Yellen (2005) claims that “in fact, I think FOMC members behave far less individualistically and strategically than assumed in some of these models.”

The timing of events in any period t is as follows: First, each committee member receives a signal on the innovation to the current supply shock (denote this stage as t_{first}). Subsequently, the financial markets receive the public information $x_t^{(j)}$ and rationally form expectations of inflation and short-term nominal interest rates. The committee then meets, deliberates, votes on an interest rate by majority, and publishes its interest rate decision (t_{third}). The committee may also publish its voting records (t_{fourth}).¹⁰ Finally, and before the beginning of period $t + 1$, the supply shock, s_t , is realized and observed by both the committee and the financial markets.

3. Solving the Model

3.1 Committee Decision Making

When the committee meets and deliberates, each committee member combines his signal on the innovation to the supply shock with those

⁹An extensive overview of the game-theoretic literature that includes an analysis of strategic voting behavior of committee members and its relevance for monetary policy committees has been provided by Gerling et al. (2003).

¹⁰We hence assume that the publication of voting records occurs in between policy meetings. In the case of the Bank of England, for example, voting records and minutes are published at 9:30 a.m. on the Wednesday thirteen days after the monthly committee decision, so that they are available to the public before the committee next meets.

of the other members in order to form an assessment of the state of the economy in period t .¹¹ The optimal combination of signals for each committee member j is a linear combination of all signals with the weights determined by the precision of signals:

$$\tilde{v}_t^{(j)} = v_t + \tilde{\mathbf{B}}_{j,t} [\varepsilon_t^{(1)} \ \varepsilon_t^{(2)} \ \dots \ \varepsilon_t^{(N)}]' \quad (15)$$

for $j = 1, 2, \dots, N$. The weights given to the different signals, the vector $\tilde{\mathbf{B}}_{j,t}$, is given by (44). It follows that $E_t(\tilde{v}_t^{(j)}) = v_t$ and that $Var(\tilde{v}_t^{(j)}) = \sigma_v^2 + \tilde{\mathbf{B}}_{j,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}_{j,t}'$.

Committee members can form an optimal assessment of the innovation to the supply shock for period t using (15) and the fact that the innovation to the supply shock has zero mean. This assessment is given by¹²

$$v_{tit}^{(j)} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}_{j,t}'} \tilde{v}_t^{(j)}, \quad (16)$$

where it is assumed that σ_v^2 is known to all committee members. An important characteristic of equation (16) is that the more precise the optimal combination of signals is perceived to be, the more weight it will be given by committee member j . In particular, the signal-extraction parameter, $\frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}_{j,t}'}$, will approach 0 as $\tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}_{j,t}' \rightarrow \infty$. If $\tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}_{j,t}' = 0$ and the optimally combined signal is perceived to be fully accurate, the signal-extraction parameter will be equal to 1.

The optimal prediction for the supply shock by member j in period t is given by

$$s_{tit}^{(j)} = \alpha s_{t-1} + v_{tit}^{(j)}, \quad (17)$$

where $v_{tit}^{(j)}$ is given by (16). As long as $\tilde{\mathbf{B}}_{j,t} \neq \tilde{\mathbf{B}}_{k,t}$ for $j = 1, \dots, N$, $k = 1, \dots, N$, and $j \neq k$, it is true that $v_{tit}^{(j)} \neq v_{tit}^{(k)}$. Thus, there will

¹¹Signals are combined using the methods developed by Bates and Granger (1969) and Dickinson (1973) on the optimal combination of forecasts.

¹²The basic Kalman filtering formulae have been used with (15) as the observation equation. For a derivation of the Kalman filtering formulae, see, for example, Hamilton (1994).

be disagreement between committee members after the deliberation process when members vote on the interest rate. Actual voting data of MPCs (Gerlach-Kristen 2003, 2009) confirm this feature of the model.

The interest rate that will be voted for by each committee member j will equal

$$i_{t|t}^{(j)} = \left(1 + \frac{2}{1 + \mu}\right) E_t \pi_{t+1} + E_t \pi_{t+2} - E_t i_{t+1} + \frac{2}{1 + \mu} (s_{t|t}^{(j)} - \pi^*) - \frac{2\mu}{1 + \mu} \theta^{(j)} \quad (18)$$

for $j = 1, 2, \dots, N$.

Since monetary policy decisions are made by majority voting, the interest rate set by the committee is determined by the median voter:

$$i_{t|t}^{(c)} = \text{median}(i_{t|t}^{(1)}, i_{t|t}^{(2)}, \dots, i_{t|t}^{(N)}). \quad (19)$$

3.2 Financial Markets

This section derives expressions for inflation and interest rate expectations of financial markets. In order to fully understand the role of the two information asymmetries in the model, in section 3.2.1 it is assumed that policy preferences of committee members are identical (and equal to y^*) and that financial markets face only uncertainty about the supply shock and the precision of public information. In section 3.2.2, uncertainty about committee members' preferences regarding the level of the output-gap target, $\theta^{(j)}$, is added.

3.2.1 Imperfect Common Knowledge about Signals and Their Precision

Financial markets solve a signal-extraction problem similar to that faced by committee members. The optimally combined public signal can be written as

$$\tilde{v}_t^F = v_t + \tilde{\mathbf{B}}_{F,t} [\zeta_t^{(1)} \quad \zeta_t^{(2)} \quad \dots \quad \zeta_t^{(N)}]', \quad (20)$$

where the vector $\tilde{\mathbf{B}}_{F,t}$ is defined in equation (49). It follows that $E_t(\tilde{v}_t^F) = v_t$ and that $\text{Var}(\tilde{v}_t^F) = \sigma_v^2 + \tilde{\mathbf{B}}_{F,t} \Omega_F \tilde{\mathbf{B}}_{F,t}'$. Such a

combination of the public information implies that the more accurate the public signal of a specific committee member is perceived to be by the markets, the more weight it will be given. This can provide an explanation for the finding by Ehrmann and Fratzscher (2007) that financial markets reacted significantly stronger to statements by Alan Greenspan than to statements by other FOMC members. Alan Greenspan might have been viewed by the markets as a particularly able policymaker with a precise observation on the state of the economy.¹³

Financial markets use (20) to form an optimal assessment of the innovation to the supply shock in period t :

$$v_{tit}^{(F)} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{F,t} \tilde{\mathbf{\Omega}}_{F,t} \tilde{\mathbf{B}}'_{F,t}} \tilde{v}_t^F. \quad (21)$$

The assessment of financial markets of the current supply shock equals

$$s_{tit}^{(F)} = \alpha s_{t-1} + v_{tit}^{(F)}. \quad (22)$$

The interest rate set by the median committee member is given by (8), where

$$s_{tit}^{(CB)} = \text{median}(s_{tit}^{(1)}, s_{tit}^{(2)}, \dots, s_{tit}^{(N)}).$$

Inflation and the output gap are given by equations (10) and (11).

3.2.2 *Imperfect Common Knowledge about Committee Members' Preferences*

Financial markets continue to form an optimal estimate of the current supply shock as in section 3.2.1. However, financial markets also need to derive optimal estimates of committee members' output-gap target preferences. It will be shown how voting records provide financial markets with partial information about those preferences. It is first investigated how private-sector expectations are formed when voting records are not published.

¹³An alternative explanation for this empirical finding lies in the institutional power of the chairman.

Voting Records Are Not Published. In this case financial markets receive a signal on the output-gap target of the median voter in period t once the policy decision is published. However, in the next period, $t + 1$, the assessments of the innovation to the current supply shock by committee members have changed and the median voter is likely to be a different committee member than in period t . Thus the usefulness of the policy decision for deducing preferences of committee members is limited. Financial markets realize this and use their priors on committee members' preferences. Private-sector expectations of the median member's preferences are thus given by $z^C = \text{median}(z^{(1)}, z^{(2)}, \dots, z^{(N)})$.

Voting Records Are Published. Voting records provide financial markets with information on the preferences of committee members. Financial markets know that the interest rate voted for by committee member j in period $t - 1$ follows

$$\begin{aligned} i_{t-1|t-1}^{(j)} = & \left(1 + \frac{2}{1 + \mu}\right) E_{t-1}\pi_t + E_{t-1}\pi_{t+1} - E_{t-1}i_t \\ & + \frac{2}{1 + \mu} (s_{t-1|t-1}^{(j)} - \pi^*) - \frac{2\mu}{1 + \mu} \theta^{(j)}. \end{aligned} \quad (23)$$

Thus, if market agents were able to observe $s_{t-1|t-1}^{(j)}$, they would be able to perfectly infer $\theta^{(j)}$ from the individual voting records. Unless committee members communicate their optimal assessment of the economy after deliberation has taken place, financial markets observe $s_{t-1|t-1}^{(j)}$ imperfectly. The estimate by financial markets of $s_{t-1|t-1}^{(j)}$ for all $j = 1, 2, \dots, N$ is equal to $s_{t-1|t-1}^{(F)}$. Therefore the signal financial markets receive on $\theta^{(j)}$ in period $t - 1$ is given by

$$\xi_{t-1}^{(j)} = \theta^{(j)} + \frac{1}{\mu} [s_{t-1|t-1}^{(F)} - s_{t-1|t-1}^{(j)}], \quad (24)$$

where $E_{t-1}(\xi_{t-1}^{(j)}) = \theta^{(j)}$ and $E_{t-1}(\xi_{t-1}^{(j)} - \theta^{(j)})^2 = (\frac{1}{\mu})^2 (E_{t-1}(s_{t-1|t-1}^{(F)} - s_{t-1|t-1}^{(j)})^2)$. It should be noted that the true variance of $\xi_{t-1}^{(j)}$ is unknown to financial markets. This variance is a function of the precision of committee members' observations on the innovation to the supply shock and the precision of the public information.

Financial markets can update their initial prior and derive a new assessment of $\theta^{(j)}$. In period $t - 1$ after voting records have been published, the best assessment of $\theta^{(j)}$ equals

$$\theta_{t-1|t-1}^{(F,j)} = \theta_{t-1|t-2}^{(F,j)} + \frac{\widetilde{Var}(\theta_{t-1|t-2}^{(F,j)})}{\widetilde{Var}(\theta_{t-1|t-2}^{(F,j)}) + \widetilde{Var}(\xi_{t-1}^{(j)})} [\xi_{t-1}^{(j)} - \theta_{t-1|t-2}^{(F,j)}], \quad (25)$$

where the weights are written in perceived terms because the true variances are unknown to financial markets. Equation (25) shows that the best assessment of $\theta^{(j)}$ in period $t - 1$, $\theta_{t-1|t-1}^{(F,j)}$, is a linear combination of the previous best assessment, $\theta_{t-1|t-2}^{(F,j)}$, and the signal on preferences contained in voting records, $\xi_{t-1}^{(j)}$. The more precise the signal on committee member j 's preference, $\xi_{t-1}^{(j)}$, is perceived to be, the more weight it will be given in the signal-extraction procedure. If this signal is perceived to be perfectly accurate, $\widetilde{Var}(\xi_{t-1}^{(j)}) = 0$, then from (25) the best assessment of $\theta^{(j)}$ will equal $\xi_{t-1}^{(j)}$. If, on the other hand, $\widetilde{Var}(\xi_{t-1}^{(j)}) \rightarrow \infty$, then $\theta_{t-1|t-1}^{(F,j)} \rightarrow \theta_{t-1|t-2}^{(F,j)}$ and thus the weight given to the signal on preferences contained in voting records will turn to 0.

The best assessment of the median member's output-gap target in period t before voting records in that period are published is given by

$$z_{t|t}^C = median(\theta_{t-1|t-1}^{(F,1)}, \theta_{t-1|t-1}^{(F,2)}, \dots, \theta_{t-1|t-1}^{(F,N)}). \quad (26)$$

It is assumed that the public ignores the uncertainty surrounding its estimates of the supply shock and the preference of the median policymaker and instead uses its own best estimates to form expectations of inflation and short-term nominal interest rates. In order to be able to obtain a closed-form expression for inflation expectations, the public is modeled as making the implicit assumption that the median voter also has the median output-gap target preference. This clearly need not be the case, since committee members' interest rate recommendations are determined both by their best assessment of the supply shock and their output-gap target preference.

Using (7), (22), and (26), we can derive private-sector expectations of inflation in period $t + 1$ as

$$E_t \pi_{t+1} = \frac{\mu \alpha}{1 + \mu(1 - \alpha)} s_{tit}^{(F)} + \pi^* + \mu z_{tit}^C. \quad (27)$$

The public's expectation of the future short-term nominal interest rate can be written as

$$E_t i_{t+1} = \frac{2\alpha + \mu(1 + \alpha)\alpha^2}{(1 + \alpha)[1 + \mu(1 - \alpha)]} s_{tit}^{(F)} + \pi^* + \mu z_{tit}^C. \quad (28)$$

Therefore, the two-period real interest rate is given by

$$R_t = \frac{1}{2} (i_{tit}^{(c)} - E_t \pi_{t+1}) + \frac{1}{2} (E_t i_{t+1} - E_t \pi_{t+2}), \quad (29)$$

where $i_{tit}^{(c)}$ is given by (19). Substituting for R_t and $E_t \pi_{t+1}$ in (1) and (3) then yields inflation and the output gap under the publication of voting records. It is straightforward to derive inflation and the output gap when voting records are not published, the only difference being the financial markets' estimate of the median member's output-gap target.

4. Results

Sections 4.1 and 4.2 solve a version of the model with a single central banker under both perfect and imperfect knowledge of the central bank's output-gap target. Section 4.3 then turns to the committee case and investigates welfare when there is only uncertainty about the state of the economy and the precision of public information. Section 4.4 introduces imperfect information about preferences of committee members, and the desirability of publishing voting records is analyzed when committee members meet repeatedly over time.

The parameters and variances are set as follows: $\alpha = 0.8$, $\mu = 0.5$, and $\sigma_v^2 = 1$ and $\sigma_{z,j}^2 = 1$ for all $j = 1, 2, \dots, N$. In addition, π^* and y^* are set to 2. For all simulations, the true variances, $\sigma_{\varepsilon,j}^2$ and $\sigma_{\varpi,j}^2$, and perceived variances, $\tilde{\sigma}_{\varepsilon,t}^{2(k,j)}$ and $\tilde{\sigma}_{\varpi,t}^{2(j)}$, are randomly drawn for each committee member $j = 1, \dots, N$. The perceived and true variances are drawn as follows for each committee member: $\sigma^2 = (u^{(j)})^2$,

where $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$ and σ^2 can stand for the true variances, $\sigma_{\varepsilon,j}^2$ and $\sigma_{\varpi,j}^2$, or the perceived variances, $\tilde{\sigma}_{\varepsilon,t}^{2(k,j)}$ and $\tilde{\sigma}_{\varpi,t}^{2(j)}$. This implies that σ^2 is basically drawn from a χ_1^2 distribution, which has a variance of 1.¹⁴ The covariance matrices— $\mathbf{\Omega}$, $\mathbf{\Omega}_F$, $\tilde{\mathbf{\Omega}}_{j,t}$, and $\tilde{\mathbf{\Omega}}_{F,t}$ —are then computed. Given these covariance matrices, it is possible to draw random shocks with the covariance properties of the system being modeled. In all simulations, those random shocks are generated using 10,000 draws.

The baseline parameters are chosen so that the perceived and true variances on the signals received on the innovation to the supply shock are drawn from a distribution with a variance of 1. Given that the variance of the innovation to the supply shock, σ_v^2 , is also set to 1, this seems a sensible baseline parameter specification. Section 4.5 considers the effects of changing the variance of $u^{(j)}$.

4.1 A Single Central Banker: Perfect Common Knowledge of Policy Preferences

It follows directly from (12) and section 3.1 that the best estimate of the innovation to the supply shock by the central banker is given by

$$v_{tit}^{(CB)} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\sigma}_{\varepsilon,t}^{2(CB)}} v_t^{(CB)} = \tilde{\tau}_{CB} v_t^{(CB)}. \quad (30)$$

Similarly for financial markets, it can be deduced from (14) and section 3.2 that

$$v_{tit}^{(F)} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\sigma}_{\varepsilon,t}^{2(F,CB)} + \tilde{\sigma}_{\varpi,t}^{2(CB)}} x_t^{(CB)} = \tilde{\tau}_F x_t^{(CB)}. \quad (31)$$

In what follows, the optimal weights that should be given to the observation on the innovation to the supply shock by the central

¹⁴It is possible to convert $u^{(j)} \sim N(0, \sigma_u^2)$ into a standard normal random variable, $Z \sim N(0, 1)$, where $Z = \frac{u^{(j)}}{\sigma_u}$. If a standard normal random variable is squared, it follows a χ_1^2 distribution with variance 2. Thus, in order to find the variance of $(u^{(j)})^2$, the variance of Z^2 needs to be multiplied by $(\sigma_u^2)^2$.

bank and the public information are defined as $\tau_{CB} = \frac{\sigma_v^2}{\sigma_v^2 + \sigma_{\varepsilon, CB}^2}$ and $\tau_F = \frac{\sigma_v^2}{\sigma_v^2 + \sigma_{\varepsilon, CB}^2 + \sigma_{\varpi, CB}^2}$, respectively. These are the weights that would be given to the signal by the central bank and the public information if their precision were perfectly known. Furthermore, let $A_1 = \frac{\mu^2 \alpha}{(1+\mu)[1+\mu(1-\alpha)]}$ and $A_2 = \frac{1}{1+\mu}$.

Using equations (10) and (11) as well as (12), (14), (17), and (22), the following expressions for inflation and the output gap are obtained:

$$\begin{aligned} \pi_t = & (A_1 - A_2 + 1)\alpha s_{t-1} + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB} + 1]v_t \\ & + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB}] \varepsilon_t^{(CB)} + A_1 \tilde{\tau}_F \varpi_t^{(CB)} + \pi^* + \mu y^* \end{aligned} \quad (32)$$

and

$$\begin{aligned} y_t = & - \left(\frac{1}{\mu} A_1 + A_2 \right) \alpha s_{t-1} \\ & - \left(\frac{1}{\mu} A_1 \tilde{\tau}_F + A_2 \tilde{\tau}_{CB} \right) (v_t + \varepsilon_t^{(CB)}) - \tilde{\tau}_F \frac{1}{\mu} A_1 \varpi_t^{(CB)}. \end{aligned} \quad (33)$$

In the appendix it is shown the expected loss can be written as

$$E(L) = \frac{1}{2} \sigma_v^2 \left[\begin{aligned} & \left(A_1^2 \left(1 + \frac{1}{\mu} \right) + 2A_1 - A_2 + 1 \right) \frac{\alpha^2}{1-\alpha^2} \\ & + (A_1 \tilde{\tau}_F)^2 \left[1 + \frac{1}{\mu} \right] \left[\frac{1}{\tau_F} \right] + A_2 \tilde{\tau}_{CB}^2 \left[\frac{1}{\tau_{CB}} \right] \\ & + 2A_1 \tilde{\tau}_F - 2A_2 \tilde{\tau}_{CB} + 1 \end{aligned} \right]. \quad (34)$$

It is possible to evaluate how the overall loss depends on τ_{CB} and $\tilde{\tau}_{CB}$ as well as τ_F and $\tilde{\tau}_F$. It is straightforward to deduce that $\frac{\partial E(L)}{\partial \tau_{CB}} < 0$. Thus it is optimal for $\tau_{CB} = 1$. This is intuitive, as in this case the true variance of the signal received by the central bank is 0. Furthermore,

$$\frac{\partial E(L)}{\partial \tilde{\tau}_{CB}} = 2A_2 \left(\frac{\tilde{\tau}_{CB}}{\tau_{CB}} - 1 \right).$$

The sign of this derivative is ambiguous. As long as $\tilde{\tau}_{CB} \leq \tau_{CB}$, the loss will be a decreasing function of $\tilde{\tau}_{CB}$. Thus if the central bank receives a signal on the innovation that is fully accurate, $\tau_{CB} = 1$,

it should understand that this is the case and use this signal as an estimate of the innovation to the supply shock.

It can easily be derived that $\frac{\partial E(L)}{\partial \tau_F} < 0$. The derivative of the loss with respect to $\tilde{\tau}_F$ equals

$$\frac{\partial E(L)}{\partial \tilde{\tau}_F} = 2A_1 \left(1 + A_1 \frac{\tilde{\tau}_F}{\tau_F} \left(1 + \frac{1}{\mu} \right) \right) > 0.$$

Thus it is optimal for the central bank to provide perfect transparency, $\sigma_{\varpi, CB}^2 = 0$, with regard to its information on the supply shock and therefore maximize τ_F , which is decreasing in $\sigma_{\varpi, CB}^2$. However, since the expected loss is increasing in $\tilde{\tau}_F$, perceived transparency should be minimal in order to maximize $\tilde{\sigma}_{\varpi, t}^{2(CB)}$ and thus minimize $\tilde{\tau}_F$. These results are summarized in the following proposition.

PROPOSITION 1. *When the committee consists of a single central banker and there is perfect common knowledge of central bank output-gap preferences,*

- (i) *it is optimal for the central banker to be perfectly transparent about its observation on the innovation to the supply shock, and*
- (ii) *transparency about the innovation to the supply shock as perceived by the financial markets should be minimal.*

These results correspond to the results by Geraats (2007), who in contrast to this paper uses a model with no long-term interest rate and no uncertainty about the supply shock by the central bank. As Geraats (2007) explains, the intuition for this result is as follows: Actual transparency about the innovation to the supply shock reduces the noise of the public information on the supply shock and thus makes inflation expectations more stable, thereby reducing the variability in the output gap and inflation. Lower perceived transparency, on the other hand, reduces the response of private-sector expectations to the supply shock (as the less precise the signal is perceived to be, the less weight it will be given). This will lead to a smaller variability in inflation, and thus the central bank needs to adjust the short-term nominal interest rate by less to offset this

increased variability in inflation, thereby muting the variability of the output gap.

4.2 *A Single Central Banker: Asymmetric Information about Policy Preferences*

The best estimates of the innovation to the supply shock by the central bank and the financial markets continue to be given by (30) and (31), respectively. However, in contrast to section 4.1, financial markets now need to derive an optimal estimate of the level of the output-gap target of the central bank, $\theta^{(CB)}$, in order to form expectations of inflation and nominal short-term interest rates.

Using equation (25), and assuming that voting records are published for the first time in period $t - 1$, financial markets form the following estimate of the output-gap target for period t given their prior $z^{(CB)}$:

$$\theta_{t|t-1}^{(F,CB)} = (1 - \tilde{\tau}_\xi)z^{(CB)} + \tilde{\tau}_\xi \xi_{t-1}^{(CB)},$$

where $\tilde{\tau}_\xi = \frac{\tilde{\sigma}_z^2}{\tilde{\sigma}_z^2 + \widetilde{Var}(\xi_{t-1}^{(CB)})}$. Thus, the more accurate the information on the central bank's policy preference contained in voting records is perceived to be, the more weight it will be given by the financial markets. When the perceived variance of the signal on the central bank's output-gap target contained in voting records is 0, $\widetilde{Var}(\xi_{t-1}^{(CB)}) = 0$, then $\tilde{\tau}_\xi = 1$, so that the estimate of the output-gap target in period $t - 1$ will be equal to the signal contained in voting records in that period. If, on the other hand, $\widetilde{Var}(\xi_{t-1}^{(CB)}) \rightarrow \infty$, then $\tilde{\tau}_\xi \rightarrow 0$, so that the weight given to the prior, $z^{(CB)}$, will turn to 1.

Using equations (17), (22), and (24), and taking into account that there is only a single central banker, the variance of the signal on policy preferences, $\xi_{t-1}^{(CB)}$, can be written as follows:

$$Var(\xi_{t-1}^{(CB)}) = \left(\frac{1}{\mu}\right)^2 [E_{t-1}(v_{t-1|t-1}^{(F)} - v_{t-1|t-1}^{(CB)})]^2.$$

Using equations (30) and (31), this expression can be simplified to

$$Var(\xi_{t-1}^{(CB)}) = \left(\frac{1}{\mu}\right)^2 \sigma_v^2 \left[\frac{\tilde{\tau}_F^2}{\tau_F} + \frac{\tilde{\tau}_{CB}}{\tau_{CB}} [\tilde{\tau}_{CB} - 2\tilde{\tau}_F] \right],$$

where the fact that $Var(\varepsilon_t^{(CB)}) = \frac{1-\tau_{CB}}{\tau_{CB}}\sigma_v^2$ and $Var(\varpi_t^{(CB)}) = (\frac{1}{\tau_F} - \frac{1}{\tau_{CB}})\sigma_v^2$ was used. Intuitively, when $\tilde{\tau}_F = \tilde{\tau}_{CB}$ and $\tau_F = \tau_{CB}$, the variance of the signal on policy preferences, $Var(\xi_{t-1}^{(CB)})$, is equal to 0. In this case, it follows from the definitions of τ_F and τ_{CB} in section 4.1 that $\sigma_{\varpi, CB}^2 = 0$ and that the central bank communicates its signal on the innovation to the supply shock without any noise. In addition, since $\tilde{\tau}_F = \tilde{\tau}_{CB}$, the central bank and the private sector attach the same weight to their observations on the innovation to the supply shock. Thus, the information on preferences contained in voting records is perfectly accurate in this case.

Given the estimate of the output-gap target, $\theta_{tit-1}^{(F,CB)}$, and the resulting private-sector expectations, inflation and the output gap can be derived using equations (1), (3), (18), (27), (28), and (29). Simplifying the resulting expression and substituting using the definitions of section 4.1, inflation can be written as follows:

$$\begin{aligned} \pi_t = & (A_1 - A_2 + 1)\alpha s_{t-1} + [A_1\tilde{\tau}_F - A_2\tilde{\tau}_{CB} + 1]v_t \\ & + [A_1\tilde{\tau}_F - A_2\tilde{\tau}_{CB}]\varepsilon_t^{(CB)} + A_1\tilde{\tau}_F\varpi_t^{(CB)} + \pi^* + \mu A_2\theta^{(CB)} \\ & + A_2\mu^2\theta_{tit-1}^{(F,CB)}. \end{aligned} \quad (35)$$

Similarly, the output gap will be given by

$$\begin{aligned} y_t = & -\left(\frac{1}{\mu}A_1 + A_2\right)\alpha s_{t-1} - \left(\frac{1}{\mu}A_1\tilde{\tau}_F + A_2\tilde{\tau}_{CB}\right)(v_t + \varepsilon_t^{(CB)}) \\ & - \tilde{\tau}_F\frac{1}{\mu}A_1\varpi_t^{(CB)} - \mu A_2[\theta_{tit-1}^{(F,CB)} - \theta^{(CB)}]. \end{aligned} \quad (36)$$

As shown in the appendix, the expected loss function can now be written as

$$\begin{aligned} E(L_{asym}) = & E(L) + \frac{1}{2}\sigma_v^2\frac{\mu}{1+\mu}\left[\tilde{\tau}_\xi^2\left(\frac{\tilde{\tau}_F^2}{\tau_F} + \frac{\tilde{\tau}_{CB}}{\tau_{CB}}[\tilde{\tau}_{CB} - 2\tilde{\tau}_F]\right)\right. \\ & \left.+ \tilde{\tau}_\xi(\tilde{\tau}_F - \tilde{\tau}_{CB})\frac{\mu\alpha^2}{1+\mu(1-\alpha)}\right] + \frac{\mu^3}{1+\mu}(1 - \tilde{\tau}_\xi)^2\sigma_z^2. \end{aligned}$$

This is clearly decreasing in τ_F , making actual transparency again optimal. Intuitively, if the central bank communicates its observation on the innovation to the supply shock without error, then this improves the estimate of the supply shock of financial markets, making voting records a less noisy signal on the central bank's desired level of the output-gap target. Because the sign of $\frac{\partial Var(\xi_{t-1}^{(CB)})}{\partial \tilde{\tau}_F}$ will depend on the sizes of $\tilde{\tau}_F$, $\tilde{\tau}_{CB}$, τ_F , and τ_{CB} , how welfare depends on $\tilde{\tau}_F$ is ambiguous.

When the vote is not communicated to financial markets, financial markets will use their prior on the output-gap target of the central bank and thus $\tilde{\tau}_\xi = 0$. The expected loss will then equal

$$E(L_{asym/nc}) = E(L) + \frac{\mu^3}{1 + \mu} \sigma_z^2.$$

When voting records are not published, the findings of section 4.1 apply, and thus the expected loss is decreasing in τ_F and increasing in $\tilde{\tau}_F$. Therefore, the central bank should be transparent about the supply shock and minimize $\sigma_{\varpi, CB}^2$. However, since the expected loss is increasing in $\tilde{\tau}_F$, perceived transparency should be minimal.

It is straightforward to deduce that if $\tilde{\tau}_F = \tilde{\tau}_{CB}$ and $\tau_F = \tau_{CB}$, the expected loss when voting records are published becomes

$$E(L_{asym, \tilde{\tau}_F = \tilde{\tau}_{CB}, \tau_F = \tau_{CB}}) = E(L) + \frac{\mu^3}{1 + \mu} (1 - \tilde{\tau}_\xi)^2 \sigma_z^2.$$

In this special case, publishing the vote is always preferable as long as $\sigma_z^2 > 0$. If $\sigma_z^2 = 0$, then the variance of the prior on preferences is 0, and thus the expected loss when voting records are published will be identical to the expected loss when they are not published.

From the above analysis it follows that as long as the difference between the perceived variance of the innovation to the supply shock of the central bank and the financial markets is sufficiently small and the central bank is transparent about its view on the innovation to the supply shock, publishing voting records will lead to a smaller expected loss compared with the case when voting records are not published.

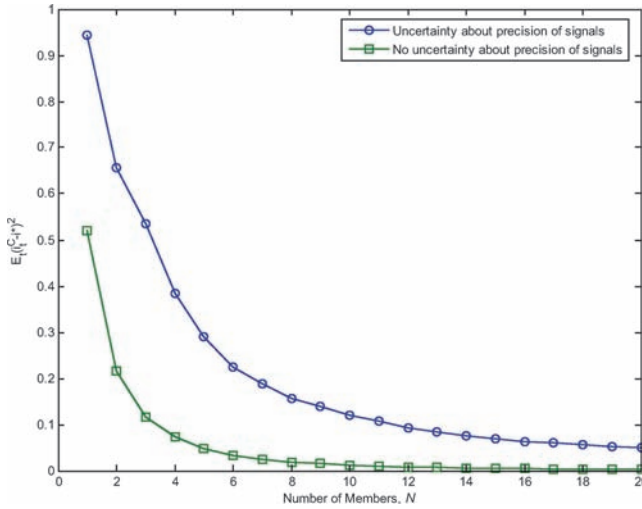
The following proposition summarizes these findings.

PROPOSITION 2. *When the committee consists of a single central banker and there is imperfect common knowledge of the central banker's preferred level of the output gap, $\theta^{(CB)}$,*

- (i) it is optimal for the central banker to be perfectly transparent about its observation on the innovation to the supply shock;*
- (ii) if voting records are not published, transparency about the innovation to the supply shock as perceived by the financial markets should be minimal; and*
- (iii) publishing voting records will always be beneficial if the central bank is perfectly transparent and the central bank and the financial markets share the same perceived variance of the signal on the innovation to the supply shock.*

4.3 The Committee Case: Perfect Common Knowledge of Policy Preferences

An interesting aspect of committee decision making is to analyze how the committee size influences welfare. One issue that can be evaluated is the average difference between the interest rate set by a committee of size N under uncertainty about the state of the economy and the interest rate set by the committee if the supply shock were perfectly known. This is depicted in figure 1 for $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$. Figure 1 assumes that there is no communication by the committee with the financial markets, so that financial market expectations are independent of the committee size. The figure shows that a larger committee sets interest rates under uncertainty closer to the level that would be set if the supply shock were perfectly known. However, when there is imperfect knowledge about the true variances of observations, then a committee of size N will set an interest rate further away from the optimal interest rate than a committee that is not facing such imperfect knowledge. This is because when committee members are uncertain about the precision of members' observations on the innovation to the supply shock, they are not efficient at pooling those different information sets, resulting in a welfare loss compared with the case when the underlying variances of signals are known. In this case the policy error is not converging toward 0. Figure 1 also shows that the marginal benefit of having a larger committee is decreasing with the number of

Figure 1. Policy Errors under Different Committee Sizes

Notes: This figure depicts policy errors for different committee sizes, under both uncertainty and perfect knowledge of the precision of signals. Perceived and true variances are randomly drawn for $\sigma^2 = (u^{(j)})^2$, where $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$. $E_t(i_t^C - i^*)^2$ denotes the average squared deviation of the interest rate set by the committee under uncertainty about the supply shock from the interest rate set if there was no such uncertainty.

members. Given that coordination costs are likely to be increasing with committee size (Sibert 2006), this may provide a partial explanation for the empirical fact that the size of committees is limited (Ehrhart, Lehment, and Vasquez-Paz 2007; Berger and Nitsch 2008; Berger, Nitsch, and Lybeck 2008). Whilst real-world MPCs vary greatly in size—ranging from three members at the Swiss National Bank to twenty-one at the European Central Bank—Berger, Nitsch, and Lybeck (2008) show that the median size is around seven to nine members.

It is possible to derive a general expression for the expected loss in the committee case. The optimal estimates of the innovation to the supply shock of committee members and financial markets are given by (16) and (21). Thus $\tilde{\tau}_{CB}$ and $\tilde{\tau}_F$ can be written as

$$\tilde{\tau}_{CB} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{CB,t} \tilde{\boldsymbol{\Omega}}_{CB} \tilde{\mathbf{B}}'_{CB,t}} \text{ and } \tilde{\tau}_F = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{F,t} \tilde{\boldsymbol{\Omega}}_F \tilde{\mathbf{B}}'_{F,t}},$$

where $\tilde{\tau}_{CB}$ denotes the weight given to the optimally combined observations by the median member and $\tilde{\tau}_{CB}$ and $\tilde{\tau}_F$ are the signal-extraction parameters in equations (16) and (21). The more accurate the optimally combined signal, $\tilde{v}_t^{(CB)}$, is perceived to be by the median member, and thus the smaller is $\tilde{\mathbf{B}}_{CB,t} \tilde{\boldsymbol{\Omega}}_{CB} \tilde{\mathbf{B}}'_{CB,t}$, the larger will be $\tilde{\tau}_{CB}$. The same is true for the signal-extraction parameter of the financial markets, $\tilde{\tau}_F$. In addition, define

$$\tau_{CB} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t}} \text{ and } \tau_F = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}}.$$

The parameters τ_{CB} and τ_F are similar to the signal-extraction parameters of equations (16) and (21). However, their size will depend on the true underlying variance of the combined signals on the innovation to the supply shock, $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t}$ and $\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}$, rather than the perceived variances. Therefore, τ_{CB} will be decreasing in $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t}$. If $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t} = 0$, then $\tau_{CB} = 1$. If $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t} \rightarrow \infty$, $\tau_{CB} \rightarrow 0$. The same applies to τ_F , which is decreasing in $\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}$.

The expected loss under the committee case is derived in the appendix, and it is shown that the expression is equivalent to the expected loss in the single central banker case, given by equation (34). The only difference is that $\tilde{\tau}_{CB}$ is a function of all the perceived weights given to committee members by the median given to committee members by the median policymaker. Similarly, $\tilde{\tau}_F$ will be a function of the perceived weights given by the public to the communication of each member. Thus, the loss is decreasing in τ_F and τ_{CB} and increasing in $\tilde{\tau}_F$. Furthermore, if $\tau_{CB} \geq \tilde{\tau}_{CB}$, the expected loss will be a decreasing function of $\tilde{\tau}_{CB}$. Therefore, all committee members should be actually transparent and communicate their signals on the innovation to the supply shock without error (thereby minimizing $\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}$ and maximizing τ_F) but perceived

transparency should be minimized.¹⁵ These theoretical findings are summarized in the following proposition.

PROPOSITION 3. *When the output-gap targets of committee members are identical and known to the public but there exists uncertainty about the precision of committee members' signals and communication,*

- (i) *actual transparency about committee members' observations on the innovation to the supply shock leads to lower variability in both inflation and the output gap and thus greater welfare, and*
- (ii) *transparency as perceived by the financial markets should be minimal.*

The intuition for these results is as follows. When committee members are perfectly transparent about their observations on the innovation to the supply shock, this reduces the noise in the public information on the supply shock. It can be easily inferred from equation (21) that the variance of the public's estimate of the innovation to the supply shock is a positive function of the variance of the combined public signals, i.e., $\tilde{\mathbf{B}}_{F,t}\mathbf{\Omega}_F\tilde{\mathbf{B}}'_{F,t}$. The variance of this optimally combined signal is given by (47). This is minimized when $\sigma_{\varpi,j}^2 = 0$ for $j = 1, 2, \dots, N$. When the noise in the public information on the supply shock is smaller, this makes inflation expectations more stable, thereby reducing the variability of inflation and the output gap. Lower perceived transparency, on the other hand, reduces the response of the private-sector expectations to the innovation to the supply shock. This again directly follows from equation (21). The larger is $\tilde{\sigma}_{\varpi,t}^{2(j)}$, the smaller is $\tilde{\tau}_F$, which is the weight given to $\tilde{v}_t^F$ in (21). This will lead to a smaller variability in inflation, and thus the committee needs to adjust the short-term nominal interest rate by less, also muting the variability of the output gap.

In order to evaluate how the expected loss depends on the committee size, it needs to be investigated how τ_F , τ_{CB} , $\tilde{\tau}_F$, and $\tilde{\tau}_{CB}$ vary

¹⁵The fact that $\tilde{\mathbf{B}}_{F,t}\mathbf{\Omega}_F\tilde{\mathbf{B}}'_{F,t}$ is minimized for $\sigma_{\varpi,j}^2 = 0$ where $j = 1, 2, \dots, N$ directly follows from (47). Similarly, from expression (48) it can be deduced that $\tilde{\mathbf{B}}_{F,t}\tilde{\mathbf{\Omega}}_{F,t}\tilde{\mathbf{B}}'_{F,t}$ is increasing in $\tilde{\sigma}_{\varpi,t}^{2(j)}$.

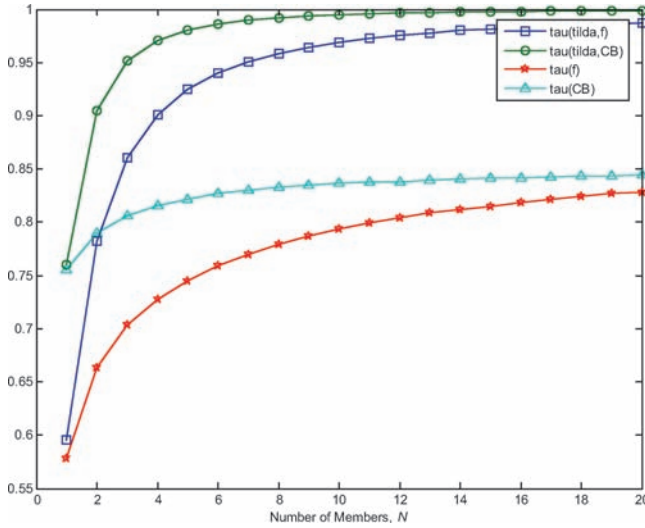
with committee size. It is straightforward to deduce from expressions (46) and (48) that as $N \rightarrow \infty$, $\tilde{\mathbf{B}}_{CB,t} \tilde{\boldsymbol{\Omega}}_{CB,t} \tilde{\mathbf{B}}'_{CB,t} \rightarrow 0$ and $\tilde{\mathbf{B}}_{F,t} \tilde{\boldsymbol{\Omega}}_{F,t} \tilde{\mathbf{B}}'_{F,t} \rightarrow 0$. Thus, it follows from the discussion above that $\tilde{\tau}_F$ and $\tilde{\tau}_{CB}$ are increasing in N . How the true variances of the combined information, (45) and (47), vary with committee size is less obvious. Therefore, it is not clear whether τ_F and τ_{CB} are increasing or decreasing with committee size. However, when simulating τ_F and τ_{CB} under different committee sizes for $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$, it is found that they are increasing with committee size. This is shown in figure 2. Thus, it follows that as N increases, $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t}$ and $\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}$ decrease. In order to provide some intuition for this numerical finding, it is interesting to evaluate the expression for the true underlying variance of the combined signals, (45), when $N = 1$ and $N = 2$. It is straightforward to deduce that the true underlying variance is greater when $N = 1$ than when $N = 2$, if

$$\frac{1}{(\tilde{\sigma}_{\varepsilon,t}^{2(j,2)})} (\sigma_{\varepsilon,1}^2 - \sigma_{\varepsilon,2}^2) + 2 \frac{\sigma_{\varepsilon,1}^2}{\tilde{\sigma}_{\varepsilon,t}^{2(j,1)}} > 0.$$

This inequality will generally hold unless $\sigma_{\varepsilon,2}^2$ exceeds $\sigma_{\varepsilon,1}^2$ significantly. Figure 2 also shows that $\tilde{\tau}_F$ and $\tilde{\tau}_{CB}$ exceed τ_F and τ_{CB} . This is because as N increases, the perceived variances of the combined signals, (46) and (48), decrease more rapidly than the true variances, (45) and (47).

It is now possible to evaluate numerically how welfare varies with committee size. The model is simulated assuming that $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$. As figure 3 shows, as N increases, the expected loss decreases but there is a decreasing marginal benefit of adding further committee members. Given that there are costs associated with larger committees, such as coordination costs, this supports the conclusion of the previous literature that a committee of limited size is optimal.

The intuition for these results is as follows: As N increases, both $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}'_{CB,t}$ and $\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}$ decrease whilst τ_{CB} and τ_F increase. This reduces the variability of inflation and the output gap and thus increases welfare. However, $\tilde{\mathbf{B}}_{F,t} \tilde{\boldsymbol{\Omega}}_{F,t} \tilde{\mathbf{B}}'_{F,t}$ also decreases. Thus, $\tilde{\tau}_F$ increases, which reduces welfare. The numerical results show that the first effect dominates. Therefore, at least for the

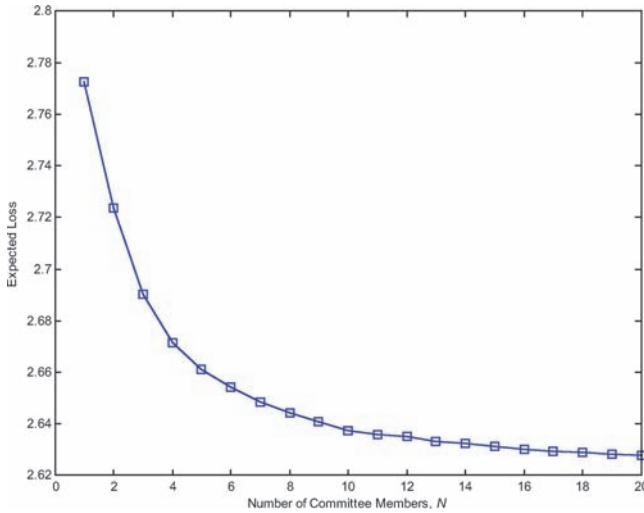
Figure 2. Signal-Extraction Parameters

Notes: The figure depicts the weights given to the combined signals by committee members and financial markets for $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$. It should be noted that $\tau(\text{tilda}, f)$ denotes $\tilde{\tau}_F$ in the text. Similarly, $\tau(\text{tilda}, CB) = \tilde{\tau}_{CB}$, $\tau(f) = \tau_F$, and $\tau(CB) = \tau_{CB}$.

parameter configuration considered here, when N increases, the benefit of reducing the noise in both the median member's and the public's assessment of the innovation to the supply shock outweighs the costs of increasing the response of the private-sector expectations to the innovation to the supply shock.

4.4 *The Committee Case: Asymmetric Information about Policy Preferences*

This section investigates how the expected loss varies over time for a given committee size, if committee members not only face uncertainty about the supply shock but also have different preferences regarding the desired level of the output gap. Financial markets have imperfect information about both the state of the economy and the policy preferences of committee members. It is assumed furthermore that committee members are appointed in the same period and meet repeatedly over time.

Figure 3. Welfare under Different Committee Sizes

Notes: This figure depicts welfare under different committee sizes when output-gap targets by committee members are identical and there is no uncertainty about them and for $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$.

After the committee has met, the true supply shock in that period is revealed to all committee members. It is hence possible for each member to observe $v_t^{(j)} - v_t$, and this difference can be used to update the previous estimate of the variance of that committee member's observation. As shown in the appendix, estimates of the variance of observations converge over time as the asymmetry in assessing precision matrices disappears. Similarly, financial markets observe $x_t^{(j)} - v_t$ after the policy decision has been made in period t . This means that they can calculate their estimate of $\tilde{\sigma}_{\zeta,t}^{2(j)}$ using (50). Over time, the initial prior will be given less and less weight and hence the estimated variance of financial markets, $\tilde{\sigma}_{\zeta,t}^{2(j)}$, will converge to the true variance, $\sigma_{\zeta,j}^2$.

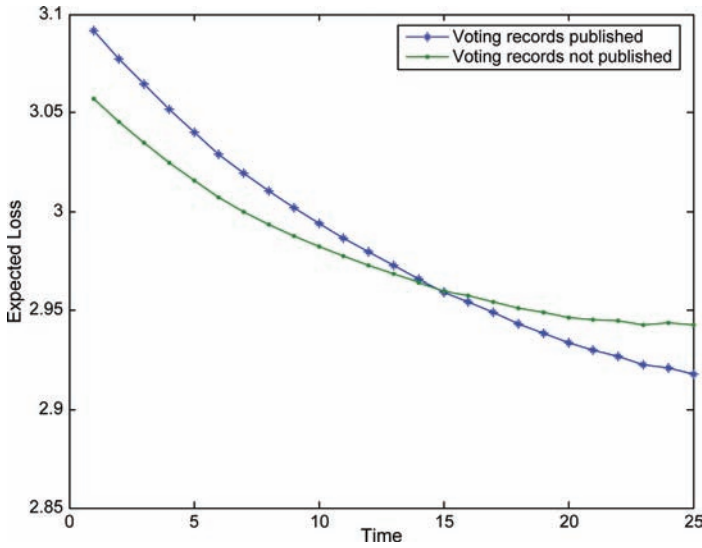
Given these results on the convergence of perceived and actual variances over time, it is possible to evaluate under which circumstances it is beneficial to publish voting records. In the simple version with just one central banker, it was shown that if the optimal and perceived weights given to the information on the supply shocks are

identical between the central bank and the financial markets, then publishing voting records will be beneficial. In this case, the signal on the preference of the central bank regarding the level of the output-gap target is perfectly accurate. Thus as long as the difference between the optimal and perceived weights is not too significant, publishing voting records is welfare enhancing.

Whether or not this result holds in the committee case has to be numerically evaluated. In the committee case, when $N > 1$, it is no longer possible to find closed-form solutions for the expected loss when there is imperfect information regarding the output-gap targets of committee members. However, we can simulate the model for N members and evaluate how the expected loss varies over time, comparing the case when voting records are published with the case when they are not. As is illustrated in figure 4 for $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$ and $N = 9$, the expected loss is generally decreasing over time as policymakers and financial markets learn the optimal weights, which they should use in the signal-extraction process. Initially, welfare is greater when voting records are not published. This changes over time, when the publication of voting records eventually becomes beneficial.

The intuition for these results is as follows: It can be seen from equation (25) that for each committee member j , the weight given to the noisy signal of j 's output-gap target preference in period t depends on the perceived precision of the information on θ^j contained in the voting records in period $t - 1$, i.e., $\widetilde{Var}(\xi_{t-1}^{(j)})$. This is a function of the perceived expected squared deviation of the best assessment of the supply shock of financial markets from the best assessment of committee member j , i.e., $\widetilde{E}_{t-1}(s_{t-1|t-1}^{(F)} - s_{t-1|t-1}^{(j)})^2$. Because of the uncertainty about the true variances of the public information, initially there may be too much or too little weight given to the information contained in voting records, relative to what is optimal. However, over time the true variances of the public information, $\sigma_{\zeta,j}^2$, will be learned by the financial markets. Thus, the weight given to the information on j 's output-gap target contained in voting records will become more accurate. In addition, voting records become a less noisy signal on the true preferences of committee members as the estimates of the supply shock by the financial markets and committee members are converging to each other. Thus, whilst there may be losses in welfare in the short term

Figure 4. Welfare under Different Communication Strategies of the Committee



Notes: Time denotes the number of meetings of a committee in which all members are appointed at the same time. The model is simulated for a committee with $N = 9$ members and $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$.

when voting records are published, eventually, publishing voting records becomes beneficial.

From equation (25) it can be deduced that the variance of the updated estimates of policymakers' preferences, $Var(\theta_{t-1|t-1}^{(F,j)})$, is a function of $Var(\xi_{t-1}^{(F,j)})$. This in turn is a function of $E_t[(v_{t-1|t-1}^{(F)} - v_{t-1|t-1}^{(j)})^2]$ and hence the actual communication error of the committee. Using equations (15), (16), (20), and (21), it is straightforward to show that $E_t[(v_{t-1|t-1}^{(F)} - v_{t-1|t-1}^{(j)})^2] = \tilde{B}_{F,t}\Omega_F\tilde{B}'_{F,t} + \tilde{B}_{j,t}\Omega\tilde{B}'_{j,t} - 2\tilde{B}_{F,t}\Omega\tilde{B}'_{j,t}$. From (47), it can be seen that $\tilde{B}_{F,t}\Omega_F\tilde{B}'_{F,t}$ is minimized when $\sigma_{\omega,j}^2 = 0$ for $j = 1, 2, \dots, N$. Therefore, the more transparent the committee is about the signals it receives on the innovation to the supply shock (i.e., the smaller $\sigma_{\omega,j}^2$), the smaller will be the possible short-term loss of publishing voting records.

Thus, when there are two sources of information asymmetry between the committee and financial markets, it is still optimal for

committee members to be actually transparent about the state of the economy. Even if there is full transparency about the economic outlook, depending on the differences between perceived and actual variances, there may be a short-term welfare loss from publishing voting records resulting from the uncertainty about the precision of public information. Morris and Shin's (2002) finding that financial markets may overreact to the public information (here the information contained in voting records) hence may hold in the short run.

Ehrmann and Fratzscher (2005) analyze the communication strategies pursued by the Bank of England, the European Central Bank, and the Federal Reserve between January 1999 and May 2004 and show that the predictability of policy decisions decreases when the committee communicates its diverging views about policy decisions. Our theoretical model provides an explanation for this result in the short term.

Empirical findings by Geraats (2009) using a sample of ninety-eight central banks show that in 2006 only ten of those central banks published their voting records, whereas fifty-six published numerical economic forecasts. In addition, only five central banks released individual voting records. The other five only published non-attributed voting patterns. This paper provides a partial explanation of this finding because it shows that transparency about the state of the economy is always beneficial, whereas there may be short-term losses from the publication of voting records.

It should be noted that these results are based on the simulation of a committee in which committee members that are appointed in the same period do not change over time. In practice, committee members are not appointed at the same time and their terms are staggered. This means that the overall uncertainty about policymakers' observation errors should be roughly constant over time. This procedure might be able to ensure that the uncertainty about the precision of the public information is never large enough to make the publication of voting records undesirable and could be beneficial in this respect.¹⁶

¹⁶On the other hand, short staggered terms, in which membership changes quite frequently, could ensure that uncertainty remains sufficiently high to make the publication of voting records detrimental.

Table 1. Sensitivity Analysis of Figure 1

N	Unknown Signal Precision $E_t(i_t^C - i^*)^2$		Known Signal Precision $E_t(i_t^C - i^*)^2$	
	$u^{(j)} \sim N(0, 0.25)$	$u^{(j)} \sim N(0, 1)$	$u^{(j)} \sim N(0, 0.25)$	$u^{(j)} \sim N(0, 1)$
1	0.4494	1.2358	0.2842	0.6075
2	0.2596	0.8944	0.0965	0.2882
3	0.2136	0.7733	0.0455	0.1571
4	0.1517	0.5256	0.0260	0.0953
5	0.1134	0.4101	0.0167	0.0608
6	0.0849	0.3150	0.0119	0.0439
7	0.0693	0.2589	0.0088	0.0329
8	0.0596	0.2166	0.0067	0.0252
9	0.0499	0.1899	0.0053	0.0207
10	0.0424	0.1649	0.0043	0.0169
11	0.0379	0.1500	0.0037	0.0145
12	0.0338	0.1299	0.0031	0.0122
13	0.0307	0.1175	0.0027	0.0107
14	0.0275	0.1074	0.0023	0.0091
15	0.0253	0.0992	0.0020	0.0078
Notes: This table depicts policy errors for different committee sizes. $E_t(i_t^C - i^*)^2$ denotes the average squared deviation of the interest rate set by the committee under uncertainty about the supply shock from the interest rate set if there was no such uncertainty.				

4.5 Sensitivity Analysis

The model considered in this paper is highly stylized and, as such, the numerical results are likely to depend on its precise parameter specification. This section evaluates the sensitivity of the numerical results in sections 4.3 and 4.4 to the distribution of $u^{(j)}$. Two alternative specifications are considered— $u^{(j)} \sim N(0, 1)$ and $u^{(j)} \sim N(0, 0.25)$ —compared with the benchmark case of $u^{(j)} \sim N(0, \sqrt{\frac{1}{2}})$. When $u^{(j)} \sim N(0, 1)$, the variance of the χ_1^2 distribution from which σ^2 is drawn is 2. When $u^{(j)} \sim N(0, 0.25)$, this variance will equal 0.125.

Tables 1 and 2 show that the qualitative results are broadly similar to those discussed in section 4.3 when $u^{(j)} \sim N(0, 1)$ and $u^{(j)} \sim N(0, 0.25)$. However, it is interesting to note that when $u^{(j)} \sim N(0, 0.25)$, the expected loss is generally smaller for different

Table 2. Sensitivity Analysis of Figures 2 and 3

Signal-Extraction Parameters and Expected Loss, Perfect Common Knowledge of y^*						
N	$u^{(j)} \sim N(0, 1)$			$u^{(j)} \sim N(0, 0.25)$		
	τ_F	τ_{CB}	EL	τ_F	τ_{CB}	EL
1	0.4614	0.6533	3.0163	0.7720	0.8450	2.6053
2	0.5204	0.6899	3.0030	0.7827	0.8715	2.5830
3	0.5604	0.7063	2.9932	0.8131	0.8812	2.5698
4	0.5874	0.7164	2.9821	0.8306	0.8873	2.5591
5	0.6103	0.7244	2.9741	0.8433	0.8914	2.5525
6	0.6269	0.7301	2.9704	0.8531	0.8940	2.5498
7	0.6393	0.7355	2.9685	0.8611	0.8962	2.5471
8	0.6516	0.7396	2.9648	0.8668	0.8971	2.5457
9	0.6629	0.7436	2.9628	0.8719	0.8984	2.5440
10	0.6698	0.7446	2.9607	0.8762	0.8999	2.5435
11	0.6773	0.7467	2.9595	0.8793	0.9009	2.5427
12	0.6841	0.7481	2.9590	0.8825	0.9012	2.5421
13	0.6902	0.7490	2.9584	0.8851	0.9018	2.5412
14	0.6965	0.7496	2.9578	0.8874	0.9027	2.5410
15	0.7018	0.7508	2.9575	0.8894	0.9031	2.5401
Notes: This table depicts the signal-extraction parameters attached to the combined signals by committee members and financial markets and the expected loss (EL) under different committee sizes when the output-gap targets of committee members are identical and known to the markets.						

committee sizes than in the baseline simulation shown in figure 3. On the other hand, when $u^{(j)} \sim N(0, 1)$, the expected loss is generally greater than in the benchmark case. This is because the larger the variance of $u^{(j)}$, the more different are the perceived and actual variances on average. This decreases welfare, as committee members are less efficient at pooling the different information sets and financial markets are less able to attach a correct weight to the public information.

The sensitivity of the results in section 4.4 to different parameter specifications is also evaluated. In addition to the two alternative specifications considered above, $u^{(j)} \sim N(0, 0.5)$ is also considered. In this case, $(u^{(j)})^2$ will be drawn from a χ_1^2 distribution

Table 3. Sensitivity Analysis of Figure 4

Expected Loss, Imperfect Common Knowledge of $\theta^{(j)}$							
Time	$u^{(j)} \sim N(0, 0.25)$		$u^{(j)} \sim N(0, 0.5)$		Time	$u^{(j)} \sim N(0, 1)$	
	VR	NR	VR	NR		VR	NR
1	2.4671	2.5548	2.7035	2.6835	1	4.2670	3.4198
2	2.3639	2.4745	2.6884	2.6690	3	3.9072	3.2562
3	2.2842	2.4181	2.6699	2.6533	5	3.6061	3.1143
4	2.2102	2.3650	2.6516	2.6381	7	3.3764	2.0050
5	2.1533	2.3260	2.6323	2.6233	9	3.2053	2.9263
6	2.1112	2.2954	2.6137	2.6098	11	3.0810	2.8726
7	2.0818	2.2718	2.5964	2.5981	13	2.9960	2.8363
8	2.0626	2.2507	2.5816	2.5887	15	2.9314	2.8059
9	2.0482	2.2347	2.5695	2.5814	17	2.8676	2.7735
10	2.0393	2.2228	2.5535	2.5758	19	2.8035	2.7436
11	2.0391	2.2143	2.5486	2.5715	21	2.7517	2.7377
12	2.0388	2.2095	2.5450	2.5678	23	2.7130	2.7316
13	2.0387	2.2093	2.5421	2.5643	25	2.6928	2.7217
14	2.0484	2.2091	2.5395	2.5608	27	2.6674	2.7173
15	2.0383	2.2089	2.5322	2.5571	29	2.6466	2.7116
Notes: This table shows the expected loss when voting records are published (VR) and when they are not published (NR) over time. Time denotes the number of meetings of a committee in which all members are appointed at the same time. The model is simulated for a committee with $N = 9$ members.							

with variance 0.5. The qualitative results of section 4.4 are sensitive to the distribution of $u^{(j)}$ and thus it is interesting to evaluate a larger number of parameter specifications and to analyze for which specifications the results hold. Table 3 shows the expected loss over time for $N = 9$. The simulation results in table 3 for $u^{(j)} \sim N(0, 1)$ and $u^{(j)} \sim N(0, 0.5)$ confirm the results of section 4.4 shown in figure 4. For these parameter specifications, initially the publication of voting records is not beneficial. It can be seen that the publication of voting records is detrimental for a longer period of time than in the benchmark case when $u^{(j)} \sim N(0, 1)$. This is intuitive, as the differences between perceived and actual variances will be larger in this case. When $u^{(j)} \sim N(0, 0.5)$, the period of time during which the publication of voting records is detrimental is shorter

than in the benchmark case since the initial difference between perceived and actual variance is smaller than in the baseline simulations. However, if the initial difference between perceived and actual variances is sufficiently small, then the publication of voting records is always welfare improving, as the results for $u^{(j)} \sim N(0, 0.25)$ show. Therefore, whether or not there is a short-term loss from publishing voting records is not independent of the parameters used in the simulation. The publication of voting records will only be detrimental in the short run if the initial difference between perceived and actual variances is sufficiently large.

5. Conclusion

Recently, there has been a growing number of theoretical and empirical papers investigating the optimal degree of transparency and communication of central banks. This paper addresses these questions within a model of a monetary policy committee with heterogeneous members whose communications and public decisions are observed by the financial markets.

The results show that it is optimal for the committee to be transparent about information regarding the current state of the economy. However, perceived transparency should be minimal. Moreover, there may be losses from publishing voting records and thereby revealing possible disagreement within the committee regarding the policy decision. This is because financial markets might overreact to the information contained in voting records, depending on their ability to estimate the true precision of the information on preferences contained in voting records. Financial markets can learn about the precision of the public information, which will make voting records less noisy and financial markets become better at attaching the correct weight to this information. Thus, in the long run, the publication of voting records is always beneficial. If committee members' terms are staggered, this might ensure that uncertainty is never large enough to make the publication of voting records undesirable, and such an institutional setup could prove useful in this respect.

There are also implications for the optimal design of monetary policy committees. It is found that a committee of limited size is optimal since there are decreasing marginal benefits of adding additional

members and coordination costs are likely to increase as the committee size gets larger. Furthermore, uncertainty about the precision of members' observations on the economy is welfare reducing. Such uncertainty should be kept at a minimum, and therefore it would be detrimental to replace all members at the same time. Staggering might again be beneficial.

Some useful directions for further research should be noted. Whilst the paper builds a model of individualistic decision making, in practice, monetary policy committees around the world use different decision-making strategies. Decisions could be made by consensus or by a committee with a chairman that dominates the proceedings. It would be interesting to make the model more applicable to such different decision-making processes and to investigate how the optimal communication strategy of a central bank depends on the decision-making procedure of its monetary policy committee.

Appendix

Finding a Solution for Private-Sector Expectations under Certainty

To find a rational-expectations equilibrium, the technique of undetermined coefficients is employed, where the bubble-free solution is obtained through a minimal-state-variable procedure as outlined by McCallum (1983). The best guess is that the nominal interest rate will be of the form

$$i_t = a_0 s_t + a_1 \pi^* + a_2 y^*. \quad (37)$$

Taking expectations and shifting the time index one period ahead gives

$$E_t i_{t+1} = a_0 \alpha s_t + a_1 \pi^* + a_2 y^*. \quad (38)$$

Using (1) together with (3) and (4) yields the following expression for inflation:

$$\pi_t = E_t \pi_{t+1} - \frac{1}{2} [i_t - E_t \pi_{t+1} + E_t i_{t+1} - E_t \pi_{t+2}] + s_t. \quad (39)$$

It is apparent that π_t will be of the form

$$\pi_t = b_0 s_t + b_1 \pi^* + b_2 y^*. \quad (40)$$

Thus

$$E_t \pi_{t+1} = b_0 \alpha s_t + b_1 \pi^* + b_2 y^* \quad (41)$$

and

$$E_t \pi_{t+2} = b_0 \alpha^2 s_t + b_1 \pi^* + b_2 y^*. \quad (42)$$

Substituting (37), (38), (41), and (42) into (39) and comparing the resulting expression with (40) yields the following:

$$b_0 = \frac{2 - a_0(1 + \alpha)}{2 - \alpha(3 + \alpha)}, b_1 = a_1 \text{ and } b_2 = a_2.$$

The above solution can be substituted into (40) to obtain an expression for inflation as a function of the coefficients a_i . Similarly, expressions for $E_t \pi_{t+1}$ and $E_t \pi_{t+2}$ in terms of a_i can be obtained. Together with (37) and (38), an expression for the output gap as a function of a_i can be found. To deduce the coefficients a_i , the optimality condition is used, substituting for y_t and π_t to determine the values for a_i such that (6) is satisfied. After some algebra, it is found that

$$a_0 = \frac{2 + \mu(1 + \alpha)\alpha}{(1 + \alpha)[1 + \mu(1 - \alpha)]}, a_1 = 1, \text{ and } a_2 = \mu. \text{ Thus}$$

$$b_0 = \frac{\mu}{[1 + \mu(1 - \alpha)]}.$$

The Signal-Extraction Problem of Committee Members

The optimally combined signal is a linear combination of all observations for committee member j :

$$\tilde{v}_t^{(j)} = b_{j,1} v_t^{(1)} + b_{j,2} v_t^{(2)} + b_{j,3} v_t^{(3)} + \dots + b_{j,N} v_t^{(N)},$$

where the weights must add up to 1. Let $\tilde{\mathbf{B}}_{j,t} = [b_{j,1} \ b_{j,2} \ \dots \ b_{j,N}]$; then the true variance of $\tilde{v}_t^{(j)}$ is

$$\begin{aligned}
E_t \left(\tilde{v}_t^{(j)} - v_t \right)^2 &= E \left(\left(\tilde{\mathbf{B}}_{j,t} \begin{bmatrix} \varepsilon_t^{(1)} & \varepsilon_t^{(2)} & \dots & \varepsilon_t^{(N)} \end{bmatrix}' \right) \right. \\
&\quad \left. \times \left(\tilde{\mathbf{B}}_{j,t} \begin{bmatrix} \varepsilon_t^{(1)} & \varepsilon_t^{(2)} & \dots & \varepsilon_t^{(N)} \end{bmatrix}' \right)' \right) \\
&= \tilde{\mathbf{B}}_{j,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}_{j,t}', \tag{43}
\end{aligned}$$

where $\boldsymbol{\Omega}$ is an $N \times N$ diagonal covariance matrix where the diagonal elements are given by the true variances of members' signals.

Since the true variance is unknown, committee member j needs to use the matrix of perceived variances, $\tilde{\boldsymbol{\Omega}}_{j,t}$. This is an $N \times N$ diagonal matrix where the diagonal elements are given by member j 's estimated variances for members 1 to N . In order to optimally combine signals, members minimize $\tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}_{j,t}'$ subject to the constraint that $\mathbf{iB}'_j = 1$, where $\mathbf{i} = [1 \ 1 \ \dots \ 1]$. The Lagrangian of this problem can be written as

$$L = \tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}_{j,t}' - \lambda [\mathbf{iB}'_j - 1].$$

Two first-order conditions are obtained:

$$2\tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} - \lambda \mathbf{i} = 0 \text{ and } [\mathbf{iB}'_j - 1] = 0.$$

These can be solved to obtain that

$$\tilde{\mathbf{B}}_{j,t} = \frac{\mathbf{i} \tilde{\boldsymbol{\Omega}}_{j,t}^{-1}}{\mathbf{i} \tilde{\boldsymbol{\Omega}}_{j,t}^{-1} \mathbf{i}'}. \tag{44}$$

The true underlying variance of the combined signals equals

$$\begin{aligned}
\tilde{\mathbf{B}}_{j,t} \boldsymbol{\Omega} \tilde{\mathbf{B}}_{j,t}' &= \frac{\mathbf{i} \tilde{\boldsymbol{\Omega}}_{j,t}^{-1}}{\mathbf{i} \tilde{\boldsymbol{\Omega}}_{j,t}^{-1} \mathbf{i}'} \boldsymbol{\Omega} \left[\frac{\mathbf{i} \tilde{\boldsymbol{\Omega}}_{j,t}^{-1}}{\mathbf{i} \tilde{\boldsymbol{\Omega}}_{j,t}^{-1} \mathbf{i}'} \right]' \\
&= \frac{\frac{1}{(\tilde{\sigma}_{\varepsilon,t}^{2(j,1)})^2} \sigma_{\varepsilon,1}^2 + \frac{1}{(\tilde{\sigma}_{\varepsilon,t}^{2(j,2)})^2} \sigma_{\varepsilon,2}^2 + \dots + \frac{1}{(\tilde{\sigma}_{\varepsilon,t}^{2(j,N)})^2} \sigma_{\varepsilon,N}^2}{\left(\frac{1}{\tilde{\sigma}_{\varepsilon,t}^{2(j,1)}} + \frac{1}{\tilde{\sigma}_{\varepsilon,t}^{2(j,2)}} + \dots + \frac{1}{\tilde{\sigma}_{\varepsilon,t}^{2(j,N)}} \right)^2}. \tag{45}
\end{aligned}$$

The perceived variance of the combined signal is equal to

$$\tilde{\mathbf{B}}_{j,t} \tilde{\boldsymbol{\Omega}}_{j,t} \tilde{\mathbf{B}}'_{j,t} = \frac{1}{\frac{1}{\tilde{\sigma}_{\varepsilon,t}^{2(j,1)}} + \frac{1}{\tilde{\sigma}_{\varepsilon,t}^{2(j,2)}} + \cdots + \frac{1}{\tilde{\sigma}_{\varepsilon,t}^{2(j,N)}}}. \quad (46)$$

The Signal-Extraction Problem of Financial Markets

The solution is identical to section 5. The true variance of the combined public signal equals

$$\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t} = \frac{\frac{1}{(\tilde{\sigma}_{\zeta,t}^{2(1)})^2} \sigma_{\zeta,1}^2 + \frac{1}{(\tilde{\sigma}_{\zeta,t}^{2(2)})^2} \sigma_{\zeta,2}^2 + \cdots + \frac{1}{(\tilde{\sigma}_{\zeta,t}^{2(N)})^2} \sigma_{\zeta,N}^2}{\left(\frac{1}{\tilde{\sigma}_{\zeta,t}^{2(1)}} + \frac{1}{\tilde{\sigma}_{\zeta,t}^{2(2)}} + \cdots + \frac{1}{\tilde{\sigma}_{\zeta,t}^{2(N)}} \right)^2}, \quad (47)$$

where $\tilde{\sigma}_{\zeta,t}^{2(j)} = \tilde{\sigma}_{\varepsilon,t}^{2(F,j)} + \tilde{\sigma}_{\varpi,t}^{2(j)}$ and $\sigma_{\zeta,j}^2 = \sigma_{\varepsilon,j}^2 + \sigma_{\varpi,j}^2$, and the perceived variance of the combined public signal is equal to

$$\tilde{\mathbf{B}}_{F,t} \tilde{\boldsymbol{\Omega}}_{F,t} \tilde{\mathbf{B}}'_{F,t} = \frac{1}{\frac{1}{\tilde{\sigma}_{\zeta,t}^{2(1)}} + \frac{1}{\tilde{\sigma}_{\zeta,t}^{2(2)}} + \cdots + \frac{1}{\tilde{\sigma}_{\zeta,t}^{2(N)}}}, \quad (48)$$

where

$$\tilde{\mathbf{B}}_{F,t} = \frac{\mathbf{i} \tilde{\boldsymbol{\Omega}}_F^{-1}}{\mathbf{i} \tilde{\boldsymbol{\Omega}}_F^{-1} \mathbf{i}'}. \quad (49)$$

The Expected Loss in Section 4.1

Given the expressions for inflation and the output gap—i.e., (32) and (33)—inflation and output-gap variability can be written as

$$\text{Var}(\pi_t) = \sigma_v^2 \left[\frac{((A_1 - A_2 + 1)\alpha)^2}{1 - \alpha^2} + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB} + 1]^2 + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB}]^2 \frac{1 - \tau_{CB}}{\tau_{CB}} + (A_1 \tilde{\tau}_F)^2 \left(\frac{1}{\tau_F} - \frac{1}{\tau_{CB}} \right) \right]$$

and

$$\begin{aligned} Var(y_t) = \sigma_v^2 & \left[\frac{\left(\left(\frac{1}{\mu} A_1 + A_2 \right) \alpha \right)^2}{1 - \alpha^2} + \left(\frac{1}{\mu} A_1 \tilde{\tau}_F + A_2 \tilde{\tau}_{CB} \right)^2 \frac{1}{\tau_{CB}} \right. \\ & \left. + \left(\frac{A_1}{\mu} \tilde{\tau}_F \right)^2 \left(\frac{1}{\tau_F} - \frac{1}{\tau_{CB}} \right) \right], \end{aligned}$$

where the fact that $Var(s_t) = \frac{1}{1-\alpha^2} \sigma_v^2$, $Var(\varepsilon_t^{(CB)}) = \frac{1-\tau_{CB}}{\tau_{CB}} \sigma_v^2$, and $Var(\varpi_t^{(CB)}) = \left(\frac{1}{\tau_F} - \frac{1}{\tau_{CB}} \right) \sigma_v^2$ was used. It can easily be seen that the variability of both inflation and the output gap are decreasing in τ_F . Using the loss function, (5), together with the expressions for inflation and output-gap variability yields the following expression for the expected loss:

$$E(L) = \frac{1}{2} \sigma_v^2 \left[\begin{aligned} & \left(A_1^2 \left(1 + \frac{1}{\mu} \right) + 2A_1 - A_2 + 1 \right) \frac{\alpha^2}{1-\alpha^2} \\ & + (A_1 \tilde{\tau}_F)^2 \left[1 + \frac{1}{\mu} \right] \left[\frac{1}{\tau_F} \right] \\ & + A_2 \tilde{\tau}_{CB}^2 \left[\frac{1}{\tau_{CB}} \right] + 2A_1 \tilde{\tau}_F - 2A_2 \tilde{\tau}_{CB} + 1 \end{aligned} \right].$$

The Expected Loss in Section 4.2

Inflation and the output gap are given by equations (35) and (36). These expressions are similar to the ones in section 4.1, with the exception that the public's estimate of the output-gap target now enters the above equations. But $\theta_{tit-1}^{(F,CB)}$ is a function of the difference between $s_{t-1|t-1}^{(F)}$ and $s_{t-1|t-1}^{(CB)}$. This difference can be written as $(s_{t-1|t-1}^{(F)} - s_{t-1|t-1}^{(CB)}) = (\tilde{\tau}_F - \tilde{\tau}_{CB})v_{t-1} + (\tilde{\tau}_F - \tilde{\tau}_{CB})\varepsilon_{t-1}^{(CB)} + \tilde{\tau}_F \varpi_t^{(CB)}$. Thus $\theta_{tit-1}^{(F,CB)}$ will be correlated with s_{t-1} . The variance of $\theta_{tit-1}^{(F,CB)}$ equals

$$Var(\theta_{tit-1}^{(F,CB)}) = \tilde{\tau}_\xi^2 \frac{1}{\mu^2} \sigma_v^2 \left(\frac{\tilde{\tau}_F^2}{\tau_F} + \frac{\tilde{\tau}_{CB}}{\tau_{CB}} [\tilde{\tau}_{CB} - 2\tilde{\tau}_F] \right) + (1 - \tilde{\tau}_\xi)^2 \sigma_z^2.$$

The variance of the output gap and inflation can thus be written as

$$\begin{aligned}
& Var_{asym}(\pi_t) \\
&= \left[Var(\pi_t) + \sigma_v^2 \left[A_2^2 \tilde{\tau}_\xi^2 \mu^2 \left(\frac{\tilde{\tau}_F^2}{\tau_F} + \frac{\tilde{\tau}_{CB}}{\tau_{CB}} [\tilde{\tau}_{CB} - 2\tilde{\tau}_F] \right) \right. \right. \\
&\quad \left. \left. + A_2 \mu \tilde{\tau}_\xi (\tilde{\tau}_F - \tilde{\tau}_{CB}) (A_1 - A_2 + 1) \alpha \right] \right] \\
& Var_{asym}(y_t) \\
&= \left[Var(y_t) + \sigma_v^2 \left[A_2^2 \tilde{\tau}_\xi^2 \left(\frac{\tilde{\tau}_F^2}{\tau_F} + \frac{\tilde{\tau}_{CB}}{\tau_{CB}} [\tilde{\tau}_{CB} - 2\tilde{\tau}_F] \right) \right. \right. \\
&\quad \left. \left. + A_2 \tilde{\tau}_\xi (\tilde{\tau}_F - \tilde{\tau}_{CB}) \left(\frac{1}{\mu} A_1 + A_2 \right) \alpha \right] \right] \\
&\quad \left. + \mu^2 A_2^2 (1 - \tilde{\tau}_\xi)^2 \sigma_z^2 \right].
\end{aligned}$$

Using these expressions together with (5) yields the expected loss:

$$E(L_{asym}) = \left[E(L) + \frac{1}{2} \sigma_v^2 \frac{\mu}{1+\mu} \left[\tilde{\tau}_\xi^2 \left(\frac{\tilde{\tau}_F^2}{\tau_F} + \frac{\tilde{\tau}_{CB}}{\tau_{CB}} [\tilde{\tau}_{CB} - 2\tilde{\tau}_F] \right) \right. \right. \\
\left. \left. + \tilde{\tau}_\xi (\tilde{\tau}_F - \tilde{\tau}_{CB}) \frac{\mu \alpha^2}{1+\mu(1-\alpha)} \right] \right. \\
\left. + \frac{\mu^3}{1+\mu} (1 - \tilde{\tau}_\xi)^2 \sigma_z^2 \right].$$

The Expected Loss in Section 4.3

With identical output-gap targets and no uncertainty by the public about them, inflation and the output gap continue to be given by (10) and (11). Inserting (16), (17), (21), and (22) into these expressions and simplifying yields

$$\begin{aligned}
\pi_t &= (A_1 - A_2 + 1) \alpha s_{t-1} + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB} + 1] v_t \\
&\quad + A_1 \tilde{\tau}_F \tilde{\mathbf{B}}_{F,t} [\zeta_t^{(1)} \quad \zeta_t^{(2)} \quad \dots \quad \zeta_t^{(j)}]' \\
&\quad - A_2 \tilde{\tau}_{CB} \tilde{\mathbf{B}}_{CB,t} [\varepsilon_t^{(1)} \quad \varepsilon_t^{(2)} \quad \dots \quad \varepsilon_t^{(N)}]' + \pi^* + \mu y^*.
\end{aligned}$$

Similarly for the output gap,

$$\begin{aligned}
y_t &= - \left(\frac{1}{\mu} A_1 + A_2 \right) \alpha s_{t-1} - \left[\frac{1}{\mu} A_1 \tilde{\tau}_F + A_2 \tilde{\tau}_{CB} \right] v_t \\
&\quad - \frac{1}{\mu} A_1 \tilde{\tau}_F \tilde{\mathbf{B}}_{F,t} [\zeta_t^{(1)} \quad \zeta_t^{(2)} \quad \dots \quad \zeta_t^{(j)}]' \\
&\quad - A_2 \tilde{\tau}_{CB} \tilde{\mathbf{B}}_{CB,t} [\varepsilon_t^{(1)} \quad \varepsilon_t^{(2)} \quad \dots \quad \varepsilon_t^{(N)}]'.
\end{aligned}$$

where $\tilde{\tau}_{CB}$ and $\tilde{\mathbf{B}}_{CB,t}$ correspond to the median policymaker. Thus,

$$Var(\pi_t) = \begin{bmatrix} \sigma_v^2 \left(\frac{((A_1 - A_2 + 1)\alpha)^2}{1 - \alpha^2} + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB} + 1]^2 \right) \\ + (A_1 \tilde{\tau}_F)^2 \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t} + (A_2 \tilde{\tau}_{CB})^2 \tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega}_{CB} \tilde{\mathbf{B}}'_{CB,t} \\ - 2A_1 \tilde{\tau}_F A_2 \tilde{\tau}_{CB} \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{CB,t} \end{bmatrix}$$

and

$$Var(y_t) = \begin{bmatrix} \sigma_v^2 \left(\frac{((\frac{1}{\mu} A_1 + A_2)\alpha)^2}{1 - \alpha^2} + \left[\frac{1}{\mu} A_1 \tilde{\tau}_F + A_2 \tilde{\tau}_{CB} \right]^2 \right) \\ + \frac{1}{\mu^2} (A_1 \tilde{\tau}_F)^2 \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t} + (A_2 \tilde{\tau}_{CB})^2 \tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega}_{CB} \tilde{\mathbf{B}}'_{CB,t} \\ - 2 \frac{1}{\mu} A_1 \tilde{\tau}_F A_2 \tilde{\tau}_{CB} \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{CB,t} \end{bmatrix}.$$

The expected loss equals

$$E(L) = \frac{1}{2} \sigma_v^2 \begin{bmatrix} \left(A_1^2 \left(1 + \frac{1}{\mu} \right) + 2A_1 - A_2 + 1 \right) \frac{\alpha^2}{1 - \alpha^2} \\ + [A_1 \tilde{\tau}_F - A_2 \tilde{\tau}_{CB} + 1]^2 + \mu \left[\frac{1}{\mu} A_1 \tilde{\tau}_F + A_2 \tilde{\tau}_{CB} \right]^2 \\ + \frac{1 + \mu}{\mu} (A_1 \tilde{\tau}_F)^2 \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t} + (1 + \mu) \tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega}_{CB} \tilde{\mathbf{B}}'_{CB,t} \end{bmatrix}.$$

If $\tau_{CB} = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega}_{CB} \tilde{\mathbf{B}}'_{CB,t}}$ and $\tau_F = \frac{\sigma_v^2}{\sigma_v^2 + \tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t}}$, it can easily be seen that the above is equivalent to (34) since $\tilde{\mathbf{B}}_{F,t} \boldsymbol{\Omega}_F \tilde{\mathbf{B}}'_{F,t} = \frac{1 - \tau_F}{\tau_F}$ and $\tilde{\mathbf{B}}_{CB,t} \boldsymbol{\Omega}_{CB} \tilde{\mathbf{B}}'_{CB,t} = \frac{1 - \tau_{CB}}{\tau_{CB}}$.

Learning about the Precision of Signals

Committee members are assumed to recursively update their estimates of the precision of signals on the supply shock. Thus, committee member j forms the following estimate about the precision of member k 's signal in period t :

$$\tilde{\sigma}_{t,\varepsilon}^{2(j,k)} = \frac{t-1}{t} \tilde{\sigma}_{t-1,\varepsilon}^{2(j,k)} + \frac{1}{t-1} (v_{t-1}^{(k)} - v_{t-1})^2$$

for $j = 1, 2, \dots, N$ and $k = 1, 2, \dots, N$. If committee members initially have different priors, then k 's evaluation will not equal j 's evaluation of the variance of k 's signal in period t unless $t = \infty$.

This intuitive result can be shown mathematically. When the committee meets for the first time at $t = 1$, members j and k have different priors on k 's variance of the signal; i.e., $\tilde{\sigma}_{1,\varepsilon}^{2(j,k)} \neq \tilde{\sigma}_{1,\varepsilon}^{2(k,k)}$. When the true supply shock becomes known, member j can update his prior of k 's variance:

$$\tilde{\sigma}_{2,\varepsilon}^{2(j,k)} = \frac{1}{2}\tilde{\sigma}_{1,\varepsilon}^{2(j,k)} + (v_1^{(k)} - v_1)^2.$$

Hence for period t ,

$$\tilde{\sigma}_{t,\varepsilon}^{2(j,k)} = \frac{1}{t}\tilde{\sigma}_{1,\varepsilon}^{2(j,k)} + \sum_{l=0}^{t-1} \left[\frac{t-l}{t} \frac{1}{t-l-1} \right] (v_{t-l-1}^{(k)} - v_{t-l-1})^2,$$

where we repeatedly substituted for $\tilde{\sigma}_{t-1,\varepsilon}^{2(j,k)}$. The first term, $\frac{1}{t}\tilde{\sigma}_{1,\varepsilon}^{2(j,k)}$, converges to 0 as $t \rightarrow \infty$. Hence, estimates of committee members converge over time as the different initial priors are given less weight. For financial markets, the estimate of the precision of the public information received from member j equals

$$\tilde{\sigma}_{\zeta,t}^{2(j)} = \frac{t-1}{t}\tilde{\sigma}_{\zeta,t-1}^{2(j)} + \frac{1}{t-1}(x_{t-1}^{(j)} - v_{t-1})^2. \quad (50)$$

If committee members are actually transparent about their information on the innovation to the supply shock, $x_{t-1}^{(j)} = v_{t-1}^{(j)}$, estimates of the variance of public information converge to their true values as $t \rightarrow \infty$.

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Canada's Pioneering Experience with a Flexible Exchange Rate in the 1950s: (Hard) Lessons Learned for Monetary Policy in a Small Open Economy*

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This paper revisits Canada's pioneering experience with a flexible exchange rate over the period 1950–62. It examines whether the floating rate was the best option for Canada in the 1950s by developing and estimating a New Keynesian small open-economy model of the Canadian economy. The model is then used to conduct a counterfactual analysis of the impact of different monetary policies and exchange rate regimes. The main finding is that the flexible exchange rate helped reduce the volatility of key macroeconomic variables. The Canadian monetary authorities, however, did not understand all of the implications of conducting monetary policy under a flexible exchange rate and a high degree of capital mobility. The paper confirms that monetary policy was more volatile in the post-1957 period and Canada's macroeconomic performance suffered as a result.

JEL Codes: E32, E37, F31, F32, N1.

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Reflections on Canada's Exchange Rate and Monetary Policies: 1950–62

Milton Friedman:

"Floating rates are not a guarantee of sensible internal monetary policy. . . . All floating rates do is make it possible for you to have a sensible internal monetary policy without considering the rest of the world."

"The reason Canada went off floating rates [in 1962] was because they were working so well, and their internal monetary policy was so bad."

(Friedman and Roosa 1967, p. 122)

Robert Mundell:

"Whether insulation is achieved or not depends on the precise behaviour of the monetary authorities."

"The tight monetary policy. . . . suggests a faulty understanding of how the advantages of a flexible exchange rate system can be exploited."

(Mundell 1964, pp. 82, 85)

James Coyne (Governor, Bank of Canada, 1955–61):

On the definition of "tight monetary policy":

"To the extent that the phrase might be taken to imply a contraction in the availability of money, it is not applicable. In this sense of the phrase there has never been a 'tight monetary policy' in Canada."

(Bank of Canada, *Annual Report 1957*, p. 15)

"[I] have always felt a special responsibility as Governor. . . to protect the value of the Canadian dollar."

(July 10, 1961 testimony before the Canadian Senate, p. 340)¹

¹The Canadian Senate defeated the Government bill declaring the position of Governor of the Bank of Canada vacant on July 14, 1961; nonetheless, Coyne resigned immediately thereafter.

1. Introduction

On September 30, 1950, the Canadian government took the bold step of allowing the Canadian dollar to float, which meant leaving the Bretton Woods par-value system less than five years after its establishment.² This action marked the beginning of the first policy experiment with a flexible rate by an industrialized country in the post-war period.³ Canada's decision to adopt a flexible exchange rate in 1950 was bitterly criticized by the International Monetary Fund (IMF) because Canada was a founding and integral member of the IMF and the Bretton Woods system, and IMF officials were concerned that other member countries might follow suit. Milton Friedman, in contrast, cited the Canadian situation in the late 1940s as one highly amenable to floating in his famous 1953 article "The Case for Floating Exchange Rates."⁴ In this article, he argued that a floating exchange rate provides two key benefits—insulation from external shocks and monetary independence. Canadian officials subsequently used Friedman's article and these arguments as an intellectual justification for their choice of exchange rate regime.

Despite the fact that Canada's flexible exchange rate remained remarkably stable for more than ten years (indeed, perhaps it was too stable), the experiment ended in May 1962 amidst a storm of controversy over the recent management of domestic monetary, exchange rate, and fiscal policies. This political uproar forced the resignation of James Coyne, the Governor of the Bank of Canada. His dismissal over his conduct of monetary policy was unprecedented among industrial countries and it initiated far-reaching reforms to clarify the relationship between central banks and governments, their

²The floating rate period actually began the following Monday on October 3, 1950. Canadian commercial banks established an interbank market for foreign exchange over the weekend.

³Earlier experiences with floating include the U.S. Greenback period (1862–78); Austria-Hungary (1896–1914); and the United Kingdom (1931–39).

⁴Although the essay was written in 1950, Friedman had in 1948 engaged in a radio debate at the University at Chicago with Deputy Governor Donald Gordon and Professor W. A. Mackintosh (Queen's University) in which he strongly advocated that Canada float rather than impose currency controls to protect reserves. Friedman, Gordon, and Mackintosh (1948) and Powell (2005) provide further details.

political masters.⁵ Furthermore, the relatively poor performance of the Canadian economy and the controversy over monetary and fiscal policies served as the inspiration for Mundell's contribution to the Mundell-Fleming model.^{6, 7}

Friedman's view of the Canadian experience with a floating exchange rate over this period is neatly summarized by the second quote given above: the flexible rate was successful to the extent that it behaved much as Friedman had expected given the fundamentals, but its ultimate demise was due to misguided monetary policy. James Coyne, on the other hand, felt his monetary policies were anything but misguided; he was committed to low inflation and maintaining the value of the dollar. Although these two goals could be simultaneously achieved with a tight monetary policy, such a policy was often inconsistent, especially in the short run, with achieving a full employment level of output when the economy experienced export and fiscal demand shocks.

The purpose of this paper is to revisit Canada's experience with a floating rate from 1950–62 to reevaluate the conduct of monetary policy under a flexible exchange rate and a high degree of capital mobility. Such reconsideration is timely because the implementation of monetary policy under a flexible exchange rate and mobile capital is an important issue in a world where more and more small open economies are adopting floating rates and global financial markets are becoming more integrated. In particular, the role of the exchange rate in the transmission of monetary policy, the appropriate response of monetary policy to exchange rate movements, and the choice of a nominal anchor are critical monetary policy questions.

⁵The Bank of Canada Act was amended so that in the event of a disagreement over monetary policy, the government would have to issue and publish a directive to the Governor and the Governor would be forced to resign.

⁶See Gordon (1961) for a detailed analysis of the Bank of Canada's controversial monetary policies during the Coyne era. Mundell's research on the small open-economy model and the assignment problem were motivated by Canada's monetary and fiscal problems. Mundell's research clearly showed that under a flexible exchange rate, tight monetary policy and expansionary fiscal policy were inimical to a full employment level of output in a small open economy. See Mundell (1964) and Bordo, Gomes, and Schembri (2010) for further details.

⁷The ruling Diefenbaker Conservative government was elected in 1958 in a landslide; it was defeated in the 1963 election primarily because of Canada's poor economic performance stemming from mistaken monetary and fiscal policies.

In this regard, useful lessons can be drawn from Canada's early experience with floating rates. In addition, Canada's decisions in 1950 and 1962 to change exchange rate regimes provide useful insights in the debate on the choice of exchange rate regimes, which has been revived recently in the international policy arena for emerging-market countries, for potential euro-zone members, and also with respect to Canada's current adherence to a floating exchange rate coupled with explicit inflation targets versus some form of credibly fixed exchange rate with the U.S. dollar.⁸

The paper makes an important contribution to the earlier literature on Canada's flexible exchange rate experience in the 1950s by developing and estimating a dynamic stochastic general equilibrium (DSGE) model of the Canadian economy in order to conduct a counterfactual analysis of the impact of different exchange rate and monetary policies—in particular, the monetary policy of James Coyne—on the performance of the Canadian economy. The theoretical model is largely based on the recent work of Galí and Monacelli (2005) and is estimated using Bayesian techniques as in Smets and Wouters (2003) and Lubik and Schorfheide (2007).

Two counterfactual experiments are conducted. The first assumes that the monetary policy of the period, until 1956 (the "pre-Coyne" period) is maintained throughout the sample period and the second assumes that Canada remained on a fixed exchange rate over the 1950–62 period rather than adopting a floating rate. The adoption of the pre-Coyne monetary policy would have reduced the volatility of output and interest rates by allowing the inflation and exchange rates to adjust more fully to exogenous shocks. Had the pre-Coyne monetary policy been in place throughout the sample period, it is likely that Canada would not have returned to the Bretton Woods par-value system in 1962, only to abandon it for good in 1970. The second experiment shows that had Canada remained on a fixed exchange rate and essentially followed U.S. monetary policy over this period, the volatility of output and inflation could have increased significantly. Thus, even though the monetary policy in the post-1957 ("Coyne") period is found to be highly volatile, it was still better than having a fixed exchange rate.

⁸For references on the recent debate in Canada, see Lafrance and Schembri (2003).

The unique Canadian experiment with a flexible exchange rate sparked extensive research, most of which was published in the 1960s and early 1970s in an era when the Bretton Woods par-value system was undergoing severe strains and debates raged about viable alternatives.⁹ This earlier literature on the Canadian experience focused on the merits of floating exchange rates by analyzing whether the regime's performance was satisfactory in terms of the stability of the exchange rate, the overall macroeconomic performance of the Canadian economy, the ability of the flexible rate to provide insulation from external shocks, and the amount of monetary independence actually achieved under the flexible rate regime. Following Friedman's arguments on stabilizing speculation, much of the research focused on the first issue, the stability of the exchange rate, and the role of short-term capital movements. The consensus opinion from this research was that, on the whole, the Canadian experience was successful, and this finding contributed to the case for generalized floating and the dismantling of the par-value system between 1971 and 1973.¹⁰

Our general conclusion is similar to that of the older literature in the qualitative sense that our research indicates that Canada's flexible rate performed reasonably well in terms of helping the Canadian economy adjust to shocks. Its effectiveness was, however, hindered by inappropriate monetary and fiscal policies—policies that did not fully incorporate an understanding of the response of a flexible exchange rate in an environment of capital mobility. Our results are generally consistent with Friedman's and Mundell's opening quotes: namely, that the Canadian floating rate experiment demonstrated that a flexible rate combined with a sensible monetary policy allowed the exchange rate to play a stabilizing role in the face of external asymmetric shocks; as a result, Canada's economic performance over approximately the first half of the period, 1950–56, was reasonably good. Over the second half of the period, 1957–62, however, the evidence indicates monetary policy under Governor Coyne was sometimes too tight and the government reacted to the growing output gap by expanding fiscal policy, which only made matters worse.

⁹See Yeager (1976) for a useful critical survey of the older literature.

¹⁰The United Kingdom and West Germany also considered the adoption of a floating rate in the 1950s, but, in the end, both decided against it.

The resultant high unemployment set in motion the political unrest that led to Coyne's departure. Coyne and other Canadian officials, however, were not alone in their mistaken beliefs about the effects of monetary and fiscal policies under a flexible exchange rate and capital mobility because they were not well understood until the pioneering theoretical work of Mundell and Fleming.¹¹

In the next section of the paper, we review the history of Canada's exchange rate and monetary policy experience in the 1950s and the debates over the floating exchange rate regime and the conduct of monetary policy. In section 3, the behavior of the floating Canadian dollar is examined in more detail. In section 4, a DSGE model of the Canadian economy is developed, then estimated and used to conduct counterfactual exercises. The final section provides some concluding remarks on lessons learned from Canada's floating rate experience.

2. Historical Background

In this section, we provide an overview of Canada's exchange rate and monetary policy experience from 1950 to 1962 by summarizing briefly the historical experience and highlighting the key issues. Monetary policy in Canada over this period was primarily directed by the Bank of Canada, which was established in 1935 by the Bank of Canada Act. The Act, however, did not clearly define the independence of the Bank of Canada over the conduct of monetary policy. This lack of clarity would become a significant political issue toward the end of the floating rate period as pressure mounted on the government for a change in the direction of monetary policy. Graham Towers served as Governor of the Bank of Canada for the first twenty years of its existence from 1935 through 1954. He was succeeded by James Coyne, who was eventually forced to resign by the government in July 1961, six months before the end of his seven-year term, essentially over differences in views concerning domestic economic policy, including fiscal and commercial policy as well as monetary policy (Powell 2009).

¹¹Haberler (1937) and Meade (1951) developed many of the ideas that were later expanded and formalized by Mundell and Fleming. See Bordo and James (2001) for more details.

Like the U.S. Federal Reserve, the Bank of Canada has a relatively diffuse legal mandate, but in the years immediately after World War II, the Bank was much less transparent. The responsibility for Canadian monetary policy rested solely with the Governor and so, in contrast to the Federal Reserve, there are no minutes from meetings to explain the process by which monetary policy decisions were reached.¹² Information was provided via the Bank's annual report, speeches by the Governor, parliamentary testimony, and a limited number of other publications. Moreover, unlike the Federal Reserve, the Bank did not consistently employ a small set of well-defined indicators/intermediate targets for monetary policy such as free reserves in the banking system.¹³

In our review of the Bank's publications and public pronouncements as well as secondary sources, it appears that monetary policy over the period was guided by the principles of maintaining low inflation and relatively stable output growth. In addition to measures of inflation and economic activity, the Bank also considered a variety of other indicators such as commercial bank cash reserves, market interest rates, and measures of aggregate liquidity such as bank loans. The concept of an intermediate target is not used initially, but later the Bank Rate, the rate at which chartered banks can borrow from the Bank of Canada, is linked to the short-term Treasury-bill (T-bill) rate. The implementation of monetary policy changed over the period as financial markets developed. Initially, moral suasion and quantitative restrictions on commercial bank lending were used and then open-market operations and changes in the Bank Rate became the preferred instruments to influence short-term market interest rates.

The Canadian federal government was responsible for exchange rate policy and made the decisions to change the exchange rate parity rates and the exchange rate regime after World War II.¹⁴ These

¹²Although the monetary policy at the Bank of Canada has become more transparent, the Governor remains solely responsible for its conduct.

¹³See Romer and Romer (2002).

¹⁴The federal government is also responsible for international reserves, which are held in the Exchange Fund Account, and thus, for foreign exchange market intervention as well. The Bank of Canada, as the financial agent for the government, actually performs the intervention.

decisions were somewhat more transparent than the Bank's monetary policy decisions because they were debated in Parliament. Although the government (Department of Finance) and the Bank of Canada had different responsibilities in terms of economic policy, there was regular consultation between them, especially on exchange rate and monetary policy.

2.1 Prelude: 1946–50

Canada allowed its currency to float in October 1950 after two unsuccessful attempts to establish a sustainable Bretton Woods par value. In July 1946, the Canadian dollar was revalued from a wartime discount with the U.S. dollar to parity. It soon became evident that this rate was too high and, beginning in 1948, official reserves began to decline, creating deflationary pressure. In September 1949, Canada followed the United Kingdom and thirty other countries and devalued its currency back to the pre-July 1946 level of US\$0.909. International economic conditions, however, soon changed in favor of Canada's exports and this rate became too low; there were strong capital inflows beginning in 1949 and continuing into 1950 (reflecting the demand for Canadian resources for the Korean War, which began in June 1950) and these led to a significant increase in official reserves (driven, in part, by speculation of another revaluation) and the money supply. To offset this inflationary pressure, the Canadian authorities decided to float the dollar rather than try to pick another par value only to find out, as in 1946 and 1948, that it would soon prove to be inappropriate. The decision to float was presented as a temporary move, presumably with a return to the par-value system once a new "fundamental equilibrium" had been attained.¹⁵

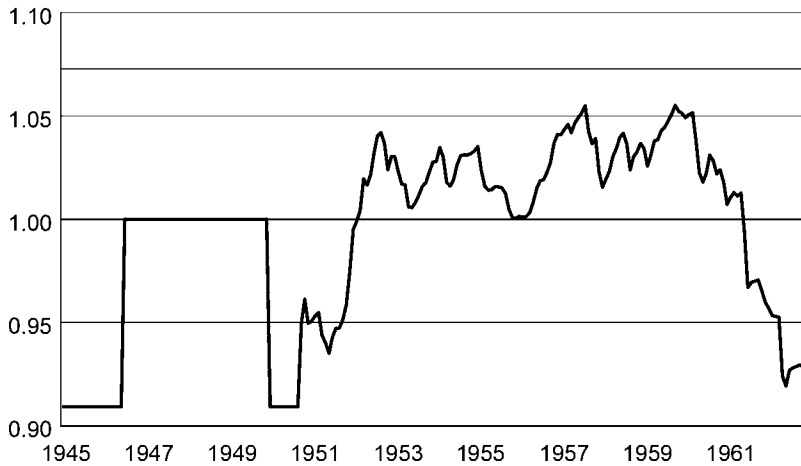
2.2 1950–51: Transition to a Free Float

Figure 1 shows that over the next eighteen months the Canadian dollar appreciated markedly from US\$0.909 to US\$1.02, a 12

¹⁵In a speech on October 20, 1952, the Conservative Minister of Finance, Douglas Abbott, said, "At some future time conditions may develop in Canada in which it would be appropriate to establish a fixed rate of exchange for the Canadian dollar." See Binhammer (1964, p. 639) and Yeager (1976, p. 544) for further details.

Figure 1. Exchange Rate

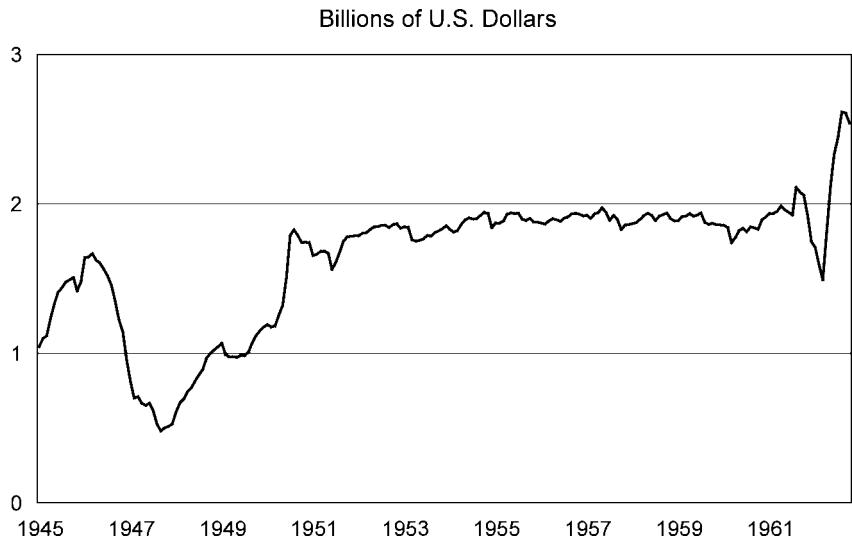
Monthly Average Noon Rates, U.S. Dollars Per Unit

**Source:** Bank of Canada.

percent increase. This rapid appreciation has largely been explained by the massive capital inflows—largely foreign direct investment (FDI) from the United States to develop Canada's natural resources (Yeager 1976, p. 544), as shown in figure 2. The Korean War boom was associated with very rapid real growth, which put upward pressure on Canadian interest rates and provided further support for the appreciating currency.

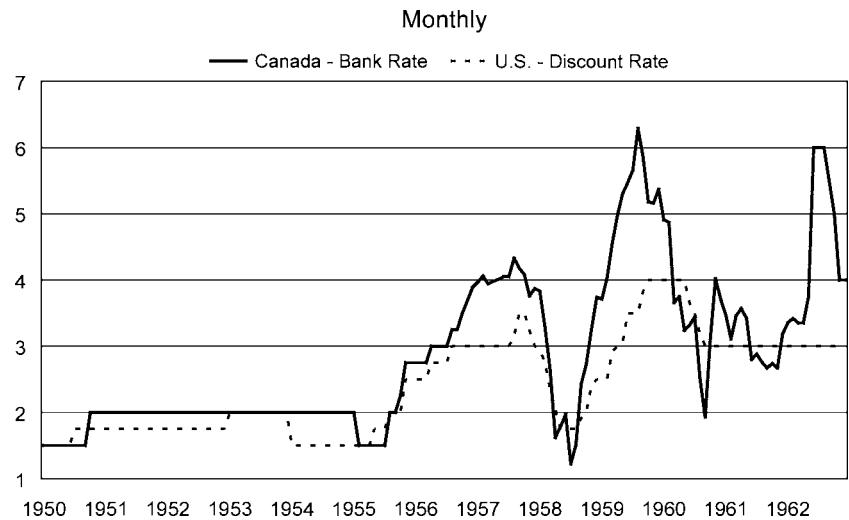
When the floating rate period started in October 1950, foreign exchange controls were still in place in Canada and an active market for short-term government securities did not exist; neither did an overnight market for reserves. Chartered banks rarely borrowed from the Bank of Canada and so the Bank Rate, which was perhaps the most visible instrument of monetary policy, was largely ineffective in influencing monetary conditions. For example, in October 1950, it was raised to 2 percent from 1.5 percent and it remained at that level until February 1955, when it was reduced back to 1.5 percent (figure 3). As a consequence, monetary policy was also conducted through various forms of open-market operations involving government securities and government deposits held by the

Figure 2. Canadian Official Holdings of Gold and U.S. Dollars

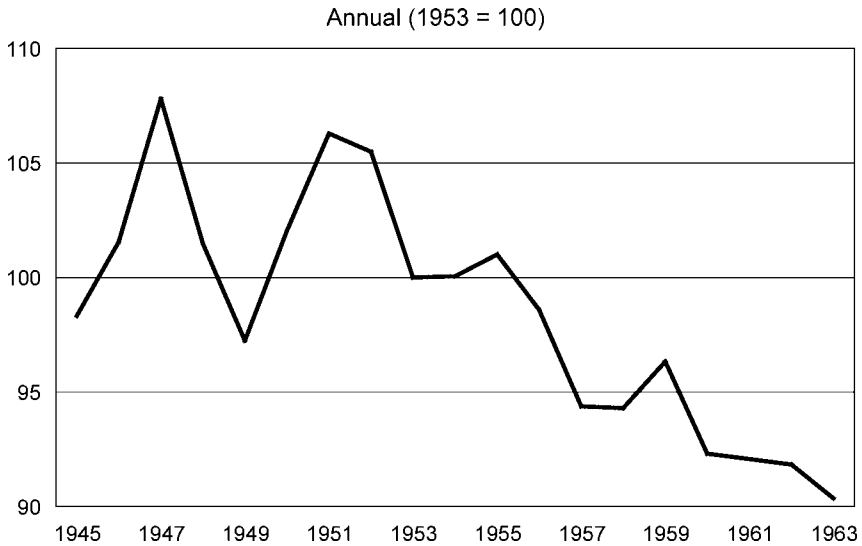


Source: Bank of Canada.

Figure 3. Interest Rates



Source: Statistics Canada and the Bank for International Settlements.

Figure 4. Canadian Real Commodity Price Index

Source: Bank of Canada.

chartered banks, and by moral suasion and direct regulation to influence the volume of chartered bank lending.

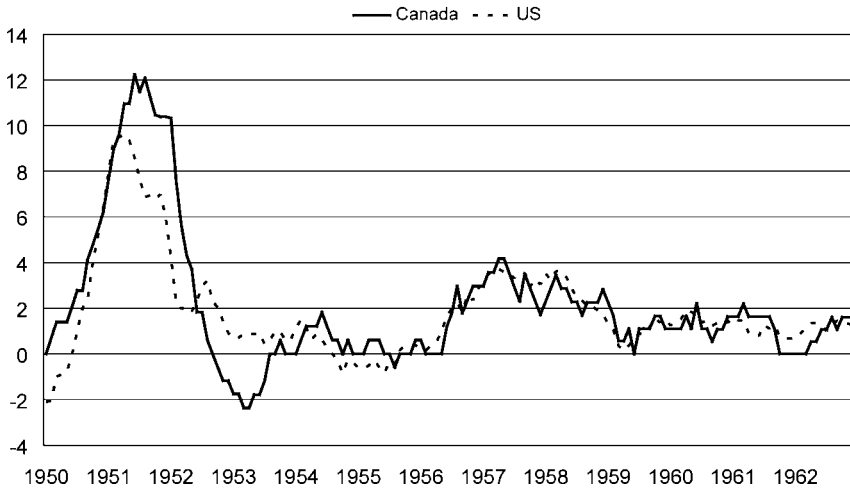
In this vein, special direct restrictions on consumer and bank credit were implemented in 1950 and 1951 to help manage the abrupt inflationary pressure stemming from the Korean War expansion in the United States. The U.S. expansion also caused commodity prices to rise sharply (see figure 4). Although the Canadian dollar appreciated over this period, the appreciation did not completely offset the inflationary impact of the increase in reserves and the money supply and the rise in commodity prices. CPI inflation was 6 percent in 1950 and rose to over 10 percent in 1951 (figure 5), and much of it was driven by food prices.

2.3 1952–56: The Textbook Case

The 1952–56 period was the heyday of the floating regime; the Canadian dollar traded at a substantial premium relative to the U.S. dollar (figure 1), and FDI-driven capital inflows continued (figure 3). With the exception of the brief post-Korean War recession in

Figure 5. Consumer Price Index

Monthly (1997 = 100), Year-Over-Year Growth Rate



Source: U.S. Bureau of Labor Statistics and Statistics Canada.

1953–54, growth was robust and inflation remained relatively low. Monetary policy was somewhat countercyclical; the actions were modest and slow to respond to cyclical output movements. Consequently, exchange rate adjustment was hindered and the full insulating value of a flexible exchange rate was not realized.

By the end of 1951, inflationary pressure was waning as commodity prices started to decline. Exchange controls were lifted in December 1951, and direct restrictions on consumer and bank credit were removed in 1952.¹⁶ The Bank of Canada took several important steps in 1953 toward encouraging the development of a broad and active market in Treasury bills by shifting from a biweekly to a weekly auction, increasing the range of maturities to 273 days, and by entering into purchase and resale agreements with dealers of government securities. This latter innovation—along with the change to the Bank Act in 1954, which raised the primary (i.e., non-interest

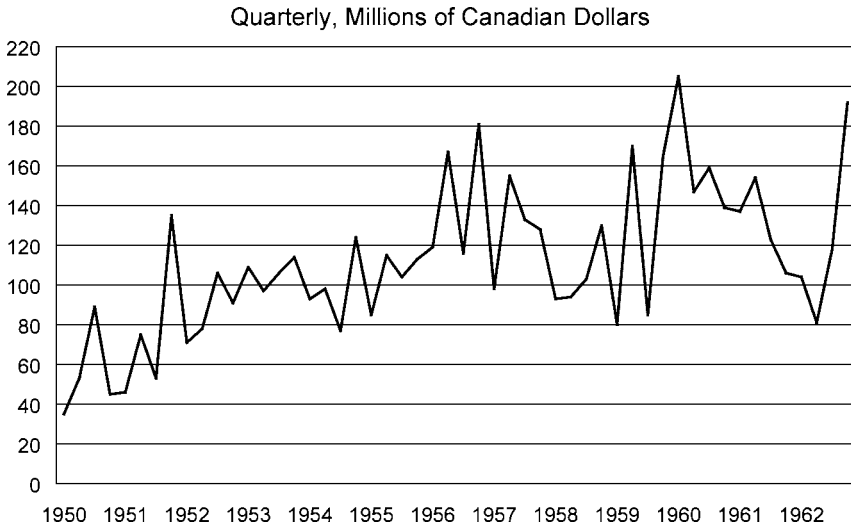
¹⁶Canada was the second country after the United States to remove exchange and capital controls after WWII. Indeed, the removal in 1951 restored the situation to what it was before WWII. See Powell (2005) for further details.

bearing) reserve requirement from 5 percent to 8 percent—spurred the establishment of a day-to-day loan market among the bank and investment dealers as banks became more interested in managing their reserves and the investment dealers were able to use the purchase and resale agreements to obtain cash from the Bank of Canada.

The period of slow growth in Canada, which began in the second half of 1953, in part as a result of reduced defense expenditures on both sides of the border, continued through most of 1954 and inflation remained close to zero. Interest rates also remained relatively low and so in February 1955 the Bank Rate was reduced to 1.5 percent because the Bank felt that this rate should be more “flexible and bear a closer (but not fixed) relation to other short-term interest rates.”¹⁷ This change marked the beginning of more frequent use of the Bank Rate as an instrument of monetary policy.

The Canadian and (to a lesser extent) U.S. economies grew strongly through the rest of 1955, 1956, and into 1957. Investment boomed in both countries, and in Canada it was centered on the development of natural resources. As aggregate demand grew, inflation pressures began to mount, and for the first time in the floating rate period, the inflationary pressure was domestic in origin. Although inflation was again almost zero in 1955, it jumped to 3 percent in 1956. The Bank reacted by offering increased resistance to the expansion in bank credit through open-market sales. As a result of the higher demand for credit for investment, market interest rates rose and the Bank Rate was increased three times in 1955 to 2 percent until it reached 2.75 percent by November. Because the banks were making more use of the lending window to meet reserve requirements, the increases were affecting interest rates at the short term. The banks also agreed to maintain voluntarily—starting May 31, 1956—a minimum liquid asset ratio (cash, Bank of Canada deposits, day-to-day loans, and Treasury bills) of 15 percent to deposits in addition to the 8 percent primary reserve requirement. In 1956, the Bank Rate was again raised on three occasions

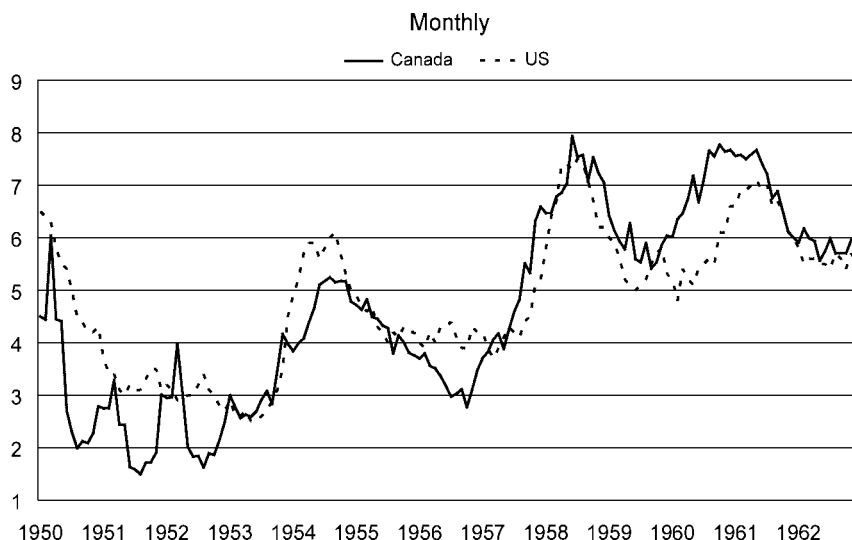
¹⁷This quote was taken from a public statement by the Bank of Canada on February 14, 1955, and published in the 1955 *Annual Report* (p. 7).

Figure 6. Direct Investment in Canada

Source: Statistics Canada.

until it reached 3.5 percent by October. In November, the decision was made that the Bank Rate would float and be equal to 25 basis points more than the Treasury-bill rate established at the weekly auction, and at the end of 1956, the Bank Rate closed at 3.92 percent, which represented an increase of almost 250 basis points in less than twenty-three months.

Over the 1950–56 period, monetary policy and exchange rate movements seemed to have been reasonably effective in managing inflationary pressures, especially those stemming from the Korean War expansion, but there is less evidence of a significant counter-cyclical impact on output growth (see figures 1 and 6.) Graham Towers retired as Governor of the Bank of Canada at the end of 1954 and James Coyne began his tenure in January 1955. Although the economy continued to grow from 1956 into 1957, higher interest rates and a stronger dollar—which had appreciated by almost 7 percent over 1955 and 1956 to a premium of US 4¢ by the end of 1956—were starting to have an impact. These were the seeds of the later controversy surrounding the conduct of Canadian monetary policy.

Figure 7. Unemployment Rate

Source: Statistics Canada and U.S. Bureau of Labor Statistics.

2.4 1957–60 Deteriorating Performance and Questionable Policies

In 1957, the economy began slowing after more than two years of rapid growth. Because this deceleration was marked by a sharp increase in the rate of unemployment, from 4 percent to 7 percent (see figure 7), observers began to question the wisdom of Canadian monetary policy, especially as the Bank continued to tighten monetary conditions until August 1957 with the Bank Rate rising to 4.33 percent and the Canadian dollar appreciating to a peak of US\$1.06 at the same time. This further tightening seemed unwarranted since the signs of a slowdown were apparent as the inflation rate started to decline early in 1957 and ended up the year at 2.2 percent, down from 3 percent in 1956.

These criticisms were further supported by the fact that the Canadian economy seemed to slow well before the U.S. economy, indicating that the source of the adverse shock was not foreign but domestic (i.e., tight monetary policy). Indeed, Canada's economic performance during the 1957–58 recession was probably worse than that of the United States, as growth was slower and

unemployment higher. Controversy also swirled around the money-supply figures provided in the 1957 *Annual Report*.¹⁸ Indeed, the tone of the 1957 *Annual Report* (and subsequent *Annual Reports* while Coyne was Governor) changed; they became more defensive and responded directly to criticisms, as illustrated by the opening quote from Coyne concerning the money supply.

The Bank also argued that the increase in the unemployment rate was due to structural factors—in particular, the rapid growth of the labor force because of massive immigration and a higher labor force participation rate. The structural argument, which can be interpreted as an increase in the natural rate of unemployment, should not be immediately dismissed. The unemployment rate in the United States also increased by almost the same amount as the recession deepened into 1958. Although U.S. immigration rates were much lower, the two countries shared a similar demographic profile and were undergoing comparable socioeconomic changes that likely had similar labor market effects. Hence, it would not be unreasonable for the natural rate of unemployment to have increased in both countries.

The trough in the recession in both countries was reached in the spring of 1958 and large-scale monetary expansions helped both economies recover quickly; interest rates in Canada fell as the Bank Rate declined from 3.92 percent at the end of 1957 to a low of 1.91 percent in July 1958. This monetary easing was part of an effort in both countries to convert or roll over into longer maturities government bonds that were issued to finance WWII defense expenditures and were coming due shortly.

As both economies rebounded in the second half of 1958, interest rates rose sharply to levels that existed at the beginning of the year, especially in Canada, where the recovery was much stronger. In the 1958 *Annual Report*, the Bank blamed the higher interest rates on higher expected future inflation (investor “inflation psychosis,” p. 3). The critics, however, blamed the higher rates on monetary policy that was too tight. Since neither side had the benefit of Mundell's later work, neither recognized that under a floating

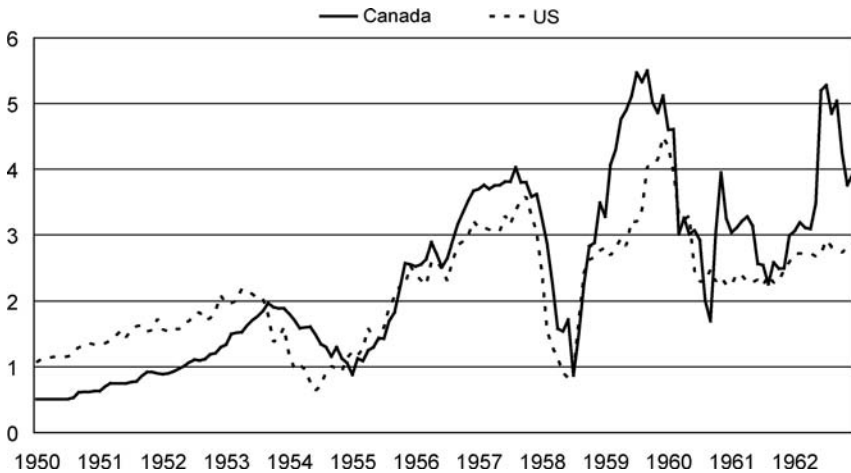
¹⁸Gordon (1961, p. 11) argued that the Bank broke with IMF data-reporting guidelines and used a “shallow statistical trick” to hide the fact that money supply had, in fact, increased over the last four months of the year by four times less than the Bank reported in the 1957 *Annual Report*.

rate, the government's countercyclical fiscal expansion would also contribute to higher interest rates and a stronger Canadian dollar. The currency appreciated by roughly 2 percent in 1958 and remained at a substantial premium to the U.S. dollar.

Inflation in 1959 fell to 2 percent from 2.5 percent in 1958 despite the fact that the strong recovery continued in Canada through to the end of 1959. Nonetheless, the Bank continued to push up short-term interest rates over the first eight months of 1959 (the Bank Rate increased by 257 basis points from 3.85 percent to 6.41 percent over this short period), without much evidence of any significant inflationary pressure. A sizable spread developed in Canada-U.S. interest rates, especially in the first half of 1959, and the dollar appreciated in 1959 by another 1 percent. The unemployment rate fell during 1959 but remained higher than the U.S. level.

The Federal Reserve also had concerns about higher future inflation in 1959 and 1960; it increased its discount rate as well, but less dramatically than the Bank of Canada. The impact of this tightening was felt in 1960 as both economies grew more slowly and inflation fell to 1.3 percent in Canada. The unemployment rate increased sharply from 6.5 percent at the beginning of 1960 to 8.7 percent by the end of the year. In the 1960 *Annual Report* (pp. 15–16), the Bank attempted to shift the blame for the slower growth and higher unemployment from its own policies to “serious structural distortions and inadequacies in the Canadian economy,” namely, current account and fiscal deficits. For Coyne, the most significant inadequacy was Canada’s “undue” dependence on foreign capital that arose, in his view, because of excessive domestic spending (1958 *Annual Report*, p. 7), and thus he felt that higher interest rates were appropriate to curb this spending. He, however, failed to recognize that higher interest rates made Canada a more attractive destination for foreign capital. As a result, Canadian short-term rates rose well above U.S. levels (see figure 8) and this differential caused portfolio capital inflows to increase and put upward pressure on the Canadian dollar.¹⁹

¹⁹Coyne’s views on Canada’s economic inadequacies and the tight monetary policy that he implemented in order to address them were heavily criticized by the academic community (Gordon 1961). His concerns about Canada’s dependence on foreign capital were, however, consistent with the growing political backlash against FDI because of popular anxieties about the perceived U.S. domination of the Canadian economy (Powell 2009).

Figure 8. Three-Month Treasury-Bill Market Yield

The rising unemployment rate and Coyne's contentious arguments, which were repeated in several speeches in 1960 and 1961, embarrassed the government and created substantial political pressure to remove Coyne. In May 1961, the government reacted by introducing legislation to declare the position of the Governor of the Bank of Canada vacant.

In summary, the Bank of Canada's tight monetary policy in the last half of the floating rate period was excessive. The overarching fears of higher future inflation were not realized because of a failure to understand the effectiveness of monetary policy under a floating exchange rate and a high degree of capital mobility. High interest rates served to appreciate the exchange rate and lower inflation by reducing both economic activity and the cost of imported goods.

2.5 1961–62 Transition to a Pegged Exchange Rate

With the resignation of James Coyne, Louis Rasminsky became Governor. Rasminsky accepted the position conditional on a clarification of the responsibility for monetary policy between the central bank and the government. Drafted by Rasminsky, the directive principle, as it came to be known, stated that under normal circumstances

the Bank of Canada was responsible for monetary policy, but if a conflict between the Bank and the government arose, the government was required to issue a specific directive to the Governor which would be published in the *Canada Gazette* (the record of the Parliament). Under these circumstances, the Governor would likely resign. The principle was incorporated into the 1967 revision to the Bank of Canada Act, and thus represented an important achievement because it clearly specified the independence of the central bank for monetary policy and the procedure to be followed in the event of a conflict between the central bank and the government. The extraordinary forced resignation of a central bank governor and the adoption of the directive principle had an impact on policies governing other central banks, including Australia.²⁰

Unfortunately, Rasminsky's accomplishment with the directive principle was overshadowed in his first year of office by the government's bungled attempts to depreciate the Canadian dollar in an effort to reflate the economy; this misguided effort eventually precipitated an exchange rate crisis, which required IMF intervention. In response to the weak economy and high unemployment rate, the government's 1961 budget promised a host of expansionary fiscal policy measures. In his budget speech of June 19, 1961, the Minister also declared his desire to see a depreciation of the Canadian dollar (not comprehending that the expansionary fiscal policy would have the opposite effect). To this end, the Bank of Canada was instructed to intervene in the foreign exchange market with unsterilized sales of Canadian dollars, which soon declined from a premium of approximately 1 percent on the U.S. dollar in July 1961 to a discount of US 3¢ less than two months later (see figure 1). Additional intervention in September 1961 led to a further drop in the Canadian dollar to a US 5¢ discount. Several months later, in April 1962, a speculative attack on the dollar occurred which forced the government to repeg the dollar at US\$0.925 in an effort to stem the free fall. Speculation continued, however, and it took an austerity program and an IMF/U.S./UK rescue package of slightly more than US \$1 billion in June 1962 to restore stability.

²⁰For a survey of the experience in other similar central banks, see Tuladhar (2005).

3. The Behavior of the Exchange Rate

This section focuses on two issues: the remarkable stability of the Canadian dollar over the floating period, 1950–62, and the related issue of whether this relatively stable exchange rate actually helped insulate the Canadian economy from external shocks. Over the twelve-year period, the dollar fluctuated in a narrow range of US 13¢, from a low of US\$0.93 in early 1950 to a peak of US\$1.06 in August 1957. If we focus on the core period, 1952–60, the range was much smaller, only US 6¢, from US\$1.00 in early 1952 to US\$1.06 in August 1957. Moreover, high-frequency fluctuations were very mild and orderly. Over the whole period, the average daily change was 0.08 percent, and only 5 percent of the daily changes over the whole floating rate period exceeded one-quarter of a percent (Poole 1967).

Several explanations have been put forward to rationalize the dollar's stability. Many attributed it to stabilizing speculation by agents who believed that movements in the rate were temporary (Poole 1967; Marsh 1969; Yeager 1976). This evidence was seen as being consistent with the original assertion by Friedman (1953) that speculation under a floating exchange rate would necessarily be stabilizing in order to be profitable. Others attributed it to the coincidence of Canadian with U.S. cyclical positions and monetary policies (Hawkins 1968, p. 31) (see figures 5, 6, and 8).

It has also been argued that official intervention operations served to reduce the volatility of the Canadian dollar exchange rate, but the literature has concluded that official intervention did not play a significant role in stabilizing the nominal exchange rate because the extent of intervention was limited to simply offsetting short-run fluctuations in order to maintain an orderly market (Plumptre 1970, p. 4). Indeed, net monthly changes in official reserves were less than US \$20 million in the majority of months when intervention occurred.²¹

Several observers, including Plumptre (1970, p. 6), argue that the relative stability of the floating Canadian dollar was due, in part, to

²¹See Wonnacott (1965), Yeager (1976), Binhammer (1964), and Royal Commission on Banking and Finance (1964) for more information.

the absence of large shocks during the 1950s.²² This is a plausible argument because this was a period of relatively stable growth driven by post-war reconstruction and increasing prosperity. Evidence provided by the estimated DSGE model developed in this paper indicates that the structural shocks experienced by the Canadian economy during the floating rate period in the 1950s were generally smaller than those found in subsequent decades.²³

In summary, the Canadian dollar was stable over this period—not only because shocks were relatively small and to some degree common to both the Canadian and U.S. economies, as evinced by the close correlation of their business cycles, but also because capital was not very mobile globally. Apart from Canada and the United States, capital controls were widespread. Finally, it is important to recognize that Canada was the only major industrialized country that was floating at that time—all other major countries had pegged rates to the U.S. dollar.

In addition to giving the domestic authorities control over monetary policy, the other main benefit of a floating rate is its ability to shelter the domestic economy from external shocks. The observation that the Canadian flexible exchange rate was fairly stable over this period despite the volatility in output growth has led some observers to conclude that Canada's experience did not provide overwhelming evidence on the postulated insulation properties of a floating rate. Wonnacott (1965) compares exchange rate movements with price-level changes and shifts in unemployment and concludes that the evidence that the flexible rate provided insulation from external shocks is mixed at best. McLeod (1965) reaches a similar conclusion for price movements; he argues that Canadian inflation performance was about the same as that of the United States over the floating rate period.²⁴

Unfortunately, the qualitative bivariate analysis that these authors conducted is incomplete and does not provide an adequate counterfactual analysis. In particular, it is likely that exchange rate

²²Plumtre (1970) also notes that when the Canadian dollar floated in the 1930s, its movements were relatively stable as well.

²³See an earlier version of this paper—Bordo, Dib, and Schembri (2007)—for more details.

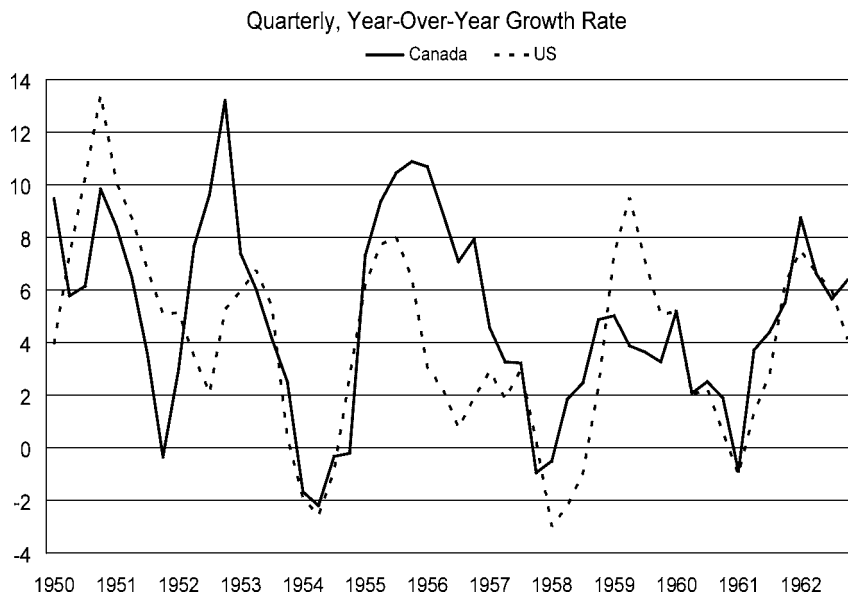
²⁴Rhomberg (1960), using an econometric model, finds the opposite result: that is, the flexible Canadian dollar did reduce the import of foreign inflation.

adjustment to movements in U.S. export demand was hindered by weakly countercyclical domestic monetary policy. Mundell (1964), McLeod (1965), and Dunn (1971) argue that Canadian monetary policy was much less countercyclical than U.S. monetary policy in the two coincident recessions of 1953–54 and 1957–58 (figure 9). Not only was the response of Canadian monetary policy and short-term rates weaker, but it was also slower (figure 8). The fact that the peak-to-trough movements in output over these two business cycles were larger in Canada than in the United States (figure 9) is consistent with this view. Consequently, the Canadian dollar tended to appreciate when monetary policy was eased earlier and more aggressively in the United States than in Canada in response to weaker output growth, and thus did not appear to provide much insulation for the Canadian economy in the face of U.S. economic slowdowns and reductions in export demand.²⁵

4. Counterfactual Analysis

In this section, we conduct two counterfactual experiments to examine the economic impact of Canadian monetary and exchange rate policies in the 1950s. The first involves eliminating the apparent shift in monetary policy that took place over the second half of the sample by maintaining the monetary policy that prevailed in the first half throughout the floating rate period. The second experiment consists of assuming that the fixed exchange rate parity of 1950 is not abandoned in favor of a flexible exchange rate. To perform these experiments, we proceed in the following steps: firstly, we develop a small theoretical DSGE model of the Canadian economy; secondly, we estimate the key parameters using a Bayesian technique; thirdly, we econometrically extract the shocks that prevailed during this period; fourthly, we conduct the experiments by modifying the monetary and exchange rate policies as required and then simulating the model, adding back the shocks that actually took

²⁵The interest rate differential had a significant impact on the Canadian dollar because of the high degree of capital mobility between Canada and the United States. Caves and Reuber (1971), for example, provide empirical evidence of the sensitivity of the Canadian dollar to the short-term interest differential in the 1950s.

Figure 9. Real Gross Domestic Product (1997 Prices)

Source: Statistics Canada.

place; and finally, we calculate and compare the variances of the key macroeconomic variables in the data with those generated by the model with the counterfactual policies.

4.1 *The Model*

In this paper, we use a small theoretical New Keynesian model for a small open economy with Calvo-type staggered price setting. The model is similar to the one used in Lubik and Schorfheide (2007), which is a simplified version of Galí and Monacelli (2005).²⁶ It is assumed that (i) financial markets are complete; (ii) the law of one price holds; (iii) prices of domestically produced goods are sticky because they are set with Calvo (1983) contracts; and (iv) domestically produced goods are either exported or combined with imported goods to produce final domestic consumption goods using a

²⁶See Galí and Monacelli (2005) for a detailed derivation of the model. The full specification of our model is given in appendix 2.

Cobb-Douglas aggregation function.²⁷ Canada is assumed to be a small open economy that takes import and export prices as well as the foreign nominal interest rate as given.

The log-linear model consists of an IS equation for a small open economy, a New Keynesian Philips-curve equation, an equation determining the nominal exchange rate, and a Taylor-type monetary policy rule. We assume that the evolution of the terms of trade is exogenous and the economy is disturbed by four additional domestic and world shocks.²⁸ The exogeneity of the terms of trade is motivated by the fact that Canada has limited influence over its import prices and its exports are mainly composed of commodities whose prices are determined exogenously on world markets in U.S. dollars. As shown in Galí and Monacelli (2005), the consumption Euler equation leads to the following IS-curve equation for a small open economy:

$$y_t = E_t y_{t+1} - [\tau + \alpha(2 - \alpha)(1 - \tau)](R_t - E_t \pi_{t+1}) + \rho_A dA_t + \alpha[\tau + \alpha(2 - \alpha)(1 - \tau)]E_t \Delta q_{t+1} + [\alpha(2 - \alpha)(1 - \tau)/\tau]E_t \Delta y_{t+1}^*, \quad (1)$$

where $0 < \alpha < 1$ is the share of imports in domestic consumption, which is a natural measure of the degree of openness of the economy, and $\tau > 0$ is the intertemporal elasticity of substitution. Note that if $\alpha = 0$, equation (1) reduces to its closed-economy variant. The endogenous variables, y_t , R_t , and π_t , denote aggregate output (normalized for technological growth),²⁹ the nominal interest rate, and the consumer price index (CPI) inflation rate, respectively; the exogenous variables, dA_t , Δq_t , and y_t^* , denote the technological growth rate, changes in terms of trade, and world output, respectively. The terms of trade, q_t , is defined as the relative

²⁷The composite consumption index is defined by $c_t = c_{h,t}^{1-\alpha} c_{f,t}^\alpha$, with $c_{h,t}$ and $c_{f,t}$ being indices of consumption of domestic and foreign goods, and α denoting the share of imports in total consumption. Since there is no investment, $c_{h,t}$ is equal to domestic output, y_t , net of exports.

²⁸All the variables in the model are expressed in terms of their log-deviations from steady-state values.

²⁹Aggregate output is normalized for trend productivity growth to ensure stationarity.

price of exports in terms of imports and we assume that its first-difference form, Δq_t , evolves exogenously, because Canada is a small open economy and most of the variations in its terms of trade were driven largely by commodity prices.³⁰

These exogenous variables evolve according to the following first-order autoregressive processes:

$$dA_t = \rho_A dA_{t-1} + \varepsilon_{At}, \quad (2)$$

$$\Delta q_t = \rho_q \Delta q_{t-1} + \varepsilon_{qt}, \quad (3)$$

and

$$y_t^* = \rho_{y^*} y_{t-1}^* + \varepsilon_{y^*t}, \quad (4)$$

where ρ_A , ρ_q , and ρ_{y^*} are autoregressive coefficients; ε_{At} , ε_{qt} , and ε_{y^*t} are i.i.d. random shocks normally distributed with mean zero and standard deviations σ_A , σ_q , and σ_{y^*} , respectively. The output growth rate computed from the data is defined as $\Delta Y_t = Y_t - Y_{t-1} = y_t - y_{t-1} + dA_t$, where Y_t is the logarithm of real aggregate output per capita.

We introduce price stickiness by assuming Taylor price setting, in which only a fraction of firms producing domestic goods are allowed to change their prices each period with an exogenous probability. Under this assumption, the New Keynesian Phillips curve is given by

$$\pi_t^d = \beta E_t \pi_{t+1}^d + \frac{\kappa}{[\tau + \alpha(2 - \alpha)(1 - \tau)]} (y_t - y_t^p), \quad (5)$$

where $0 < \beta < 1$ is the discount factor; π_t^d is producer price index (PPI) inflation—defined as the rate of change in the index of domestic goods prices—and $(y_t - y_t^p)$ is the output gap; and y_t^p is potential output when prices are fully flexible. The parameter $\kappa > 0$ depends on the model's other structural parameters, such as the price-stickiness parameter and labor supply and demand elasticities.

³⁰We have estimated the model with an endogenous terms of trade defined as $[\tau + \alpha(2 - \alpha)(1 - \tau)]\Delta q_t = \Delta y_t^* - \Delta y_t$. Some of the resulting parameter estimates have implausible values, and the model performs very poorly in terms of the in-sample fit; a good fit is critical for the counterfactual analysis. We have therefore decided to maintain the exogenous terms of trade assumption. These results are similar to those found by Lubik and Schorfheide (2007).

In the estimation, this parameter is treated as a structural parameter of the model. In the context of a small open economy, Lubik and Schorfheide (2007) show that potential output can be derived as a linear function of foreign output:

$$y_t^p = -\frac{\alpha(2-\alpha)(1-\tau)}{\tau}y_t^* \quad (6)$$

Galí and Monacelli (2005) derive the following equation that relates CPI inflation, π_t , to PPI inflation, π_t^d , and changes of the terms of trade, Δq_t :

$$\pi_t = \pi_t^d - \alpha \Delta q_t \quad (7)$$

Therefore, the equation for the exchange rate, e_t , can be derived from the definition of the CPI, assuming that the uncovered interest rate parity (UIP) and relative PPP always hold:

$$\Delta e_t = \pi_t - (1-\alpha)\Delta q_t - \pi_t^*, \quad (8)$$

where π_t^* is the world inflation rate that evolves exogenously according to the following first-order autoregressive process,

$$\pi_t^* = \rho_\pi \pi_{t-1}^* + \varepsilon_{\pi^*t}, \quad (9)$$

where ρ_π is an autoregressive coefficient and ε_{π^*t} is a random shock normally distributed with mean zero and standard error σ_{π^*} .³¹

The terms of trade enters in first-difference form because only changes in relative prices affect inflation and real variables via the definition of the CPI. This specification is consistent with our assumption that the Canadian firms had no market power in the international markets; thus the prices of internationally traded goods can be treated as exogenous to the Canadian economy.

For the convenience of the analysis, it is assumed that the Bank of Canada followed a Taylor-type rule to conduct its monetary policy. Thus, it managed the short-term nominal interest rate to respond to deviations of CPI inflation, output, and changes of the nominal

³¹Under a fixed exchange rate regime, $\Delta e_t = 0$, the CPI inflation is fully determined by the PPI inflation rate and the terms of trade, as in equation (7). Nevertheless, equation (8) disappears from the model.

exchange rate from their long-term equilibrium values. This policy rule is the same one used in Lubik and Schorfheide (2007) to examine if the Bank of Canada actively responded to exchange rate movements.³² This rule allows us to simply compare the reaction of the monetary authority with inflation, output, and exchange rate variations during the pre- and post-Coyne periods. The monetary policy rule is described by

$$R_t = \rho_R R_{t-1} + (1 - \rho_R)[\psi_\pi \pi_t + \psi_y y_t + \psi_e \Delta e_t] + \varepsilon_{Rt}, \quad (10)$$

with all policy coefficients greater than zero. The first variable on the right-hand side of (10) is the smoothing term introduced to match the persistence in the nominal interest rate, where ρ_R is the smoothing coefficient; the term ε_{Rt} is a zero-mean, serially uncorrelated monetary policy shock with standard deviation σ_R . This exogenous monetary policy shock is interpreted as the unsystematic component of monetary policy. The parameters ψ_π , ψ_y , and ψ_e measure the responses of the monetary authority to inflation, output, and nominal exchange rate variations. A unique equilibrium exists as long as the sum of ρ_R , $(1 - \rho_R)\psi_\pi$, and $(1 - \rho_R)\psi_e$ is larger than unity. The model assumes the following UIP condition between the domestic nominal interest rate, R_t , the foreign nominal interest rate, R_t^* , and the expected exchange rate movement, $E_t \Delta e_{t+1}$:

$$R_t = R_t^* + E_t \Delta e_{t+1}. \quad (11)$$

The model of a small open economy given by equations (1), (5), (7), (8), and (10) can be solved using the Blanchard and Kahn (1980) procedure to obtain the following state-space representation:

$$\begin{aligned} s_t &= \Phi_1 s_{t-1} + \Phi_2 \varepsilon_t, \\ d_t &= \Phi_3 s_t, \end{aligned} \quad (12)$$

where the state variable vector s_t includes predetermined and exogenous variables, d_t is a vector of control variables, and the vector ε_t contains the random innovations. The coefficients matrices, Φ_1 , Φ_2 ,

³²As in Lubik and Schorfheide (2007), we assume that monetary policy is specified in terms of deviations of current output from its steady-state value.

and Φ_3 , have elements that depend on the structural parameters of the model. The state-space solution (12) is used to estimate and simulate the model.

4.2 *Estimation Procedure and Results*

The model's parameters are estimated using Bayesian estimation techniques, as in Lubik and Schorfheide (2007). These techniques update prior distributions for the deep parameters of the model using the actual data.³³ The estimation is done using recursive simulation methods, particularly the Metropolis-Hastings algorithm. This approach—which is described in Schorfheide (2000) and widely used in estimating DSGE models, such as Smets and Wouters (2003)—is system based and fits the solved DSGE model to a vector time series.

The model has five shocks, so we use five aggregate time series in the estimation to avoid the singularity problem. The data series are monthly and consist of the growth rate of the Canadian industrial production index as a measure of real output growth, the nominal ninety-day T-bill interest rate, the annualized monthly CPI inflation rate, the first difference of the nominal exchange rate, and the first difference of the terms of trade. The terms of trade is defined as the relative price of exports in terms of imports. All of the series, including the output growth, are demeaned before estimation, because the model is log-linearized around steady-state equilibrium values.

The vector of observable variables, z_t , is composed of real output growth, the nominal interest rate, the CPI inflation rate, and the first differences in the nominal exchange rate and the terms of trade. Therefore, z_t is given by

$$z_t = [\Delta Y_t, R_t, \pi_t, \Delta e_t, \Delta q_t]',$$

which is a 5×1 vector of observable variables, and $Z^T = \{z_1, \dots, z_T\}$ represents a $5 \times T$ data matrix of T observations on z_t . The structural parameters of the model are in a 16×1 vector θ . The linear structural

³³The discount factor, β , is set at 0.998 prior to estimation, because there is insufficient information in the data to estimate it accurately. This calibration implies an annual steady-state real interest rate of 2.5 percent, which matches the historical average for the period 1952–61.

model is reduced to a state-space representation for z_t . Under the assumption that all the structural shocks are normally distributed and uncorrelated, we obtain a likelihood function $L(\theta|Z^T)$ that can be evaluated using the Kalman filter described in detail in Hamilton (1994, ch. 13).

To implement the Bayesian approach, a prior distribution of the parameters is assumed with a density $p(\theta)$. The advantage of the Bayesian technique over other estimation methods is that, by setting prior densities, we can attach implicit weights to the prior information on the structural parameters. Then the data given by Z^T are used to update the prior distribution using the likelihood function. The posterior distribution based on Bayes's theorem is

$$p(\theta|Z^T) = \frac{L(\theta|Z^T)p(\theta)}{\int L(\theta|Z^T)p(\theta)d\theta}.$$

The Bayesian simulation technique, described in Schorfheide (2000), generates random posterior draws of the parameters that are used in the simulation of the model. This approach allows us to estimate structural parameters of the model and at the same time extract the smoothed structural shocks.

Canada had a flexible exchange rate for the period from 1950:M10 to 1962:M5. However, exchange controls remained in place until the end of 1951, and an active money market was not established until 1953 (Watts 1993). Therefore, we estimate the model over the relevant flexible exchange rate period, 1952:M1 to 1961:M12, and over two sub-samples: the pre-1957 period from 1952:M1 to 1956:M12 and the post-1957 period from 1957:M1 to 1961:M12. The second sub-sample corresponds to the Coyne period.³⁴ The estimates of monetary policy parameters in the two sub-samples allow us to compare the reaction of the monetary authority with inflation, output, and nominal exchange rate variations.

³⁴The estimation sample starts in 1952:M1 because the Bank of Canada abandoned exchange control in December 1951. The sample is divided at the end of 1956 because it roughly corresponds to the change of monetary policy that occurred in the mid-1950s with the change in Governors. With monthly data, there are sixty observations in each sub-sample.

Table 1. Prior and Posterior Distributions
of the Parameters

Param.	Priors			Posteriors		
	Density	Mean	Std.	Entire Period Posterior Means [95% Interval]	Pre-1957 Posterior Means [95% Interval]	Post-1957 Posterior Means [95% Interval]
ρ_R	Beta	0.9	0.05	0.852 [0.81–0.89]	0.935 [0.91–0.96]	0.742 [0.67–0.82]
ψ_π	Gamma	1.5	0.3	1.326 [0.92–1.70]	1.141 [0.75–1.51]	1.778 [1.30–2.28]
ψ_y	Gamma	0.5	0.3	0.692 [0.39–0.99]	0.503 [0.23–0.78]	0.627 [0.32–1.01]
ψ_e	Gamma	0.5	0.3	1.716 [1.00–2.39]	1.4915 [0.62–2.30]	0.874 [0.38–1.44]
α	Beta	0.25	0.05	0.217 [0.15–0.28]	0.232 [0.16–0.29]	0.261 [0.18–0.33]
τ	Gamma	0.5	0.05	0.426 [0.34–0.50]	0.458 [0.38–0.53]	0.4590 [0.38–0.54]
κ	Beta	0.5	0.05	0.421 [0.34–0.49]	0.492 [0.41–0.56]	0.433 [0.36–0.50]
ρ_A	Beta	0.5	0.05	0.547 [0.44–0.64]	0.480 [0.41–0.56]	0.567 [0.46–0.67]
ρ_q	Beta	0.2	0.05	0.242 [0.15–0.34]	0.2340 [0.13–0.33]	0.206 [0.13–0.27]
ρ_{y^*}	Beta	0.9	0.05	0.971 [0.95–0.99]	0.977 [0.96–0.99]	0.928 [0.88–0.97]
ρ_{π^*}	Beta	0.7	0.05	0.568 [0.50–0.63]	0.580 [0.50–0.66]	0.656 [0.58–0.73]
σ_R	InvGam	0.01	0.05	0.67 [0.55–0.79]	0.29 [0.22–0.35]	0.91 [0.68–1.13]
σ_A	InvGam	0.01	Inf.	0.73 [0.55–0.90]	0.77 [0.61–0.92]	0.72 [0.54–0.91]
σ_q	InvGam	0.01	Inf.	0.93 [0.83–1.03]	1.07 [0.91–1.22]	0.76 [0.64–0.85]
σ_{y^*}	InvGam	0.01	Inf.	1.85 [1.00–2.68]	1.53 [0.69–2.51]	1.97 [1.03–2.76]
σ_{π^*}	InvGam	0.01	Inf.	1.34 [1.19–1.49]	1.63 [1.38–1.89]	0.90 [0.78–1.02]
Log Data Density				1,946.63	985.40	994.62
Notes: In columns 5–7, posterior means and 95 percent posterior probability intervals [in brackets] are reported.						

Table 1 displays the prior and posterior distributions for the structural parameters of the model. The priors used for the monetary policy parameters are the instrumental variable estimates that do not impose the model restrictions. The priors for the autoregressive coefficient

and standard deviations in the exogenous processes are recovered from the least-squared estimation of different laws of motion using the actual data. These priors are estimated using the relevant floating rate period that covers 1952:M1–1961:M12. For the priors of the remaining structural parameters α , τ , and κ , we employ values similar to those used in Lubik and Schorfheide (2007) for Canada. Table 1 also reports the estimation results of the model's structural parameters: the posterior means and 95 percent posterior probability intervals. We evaluate the robustness of the estimation of our model by comparing the prior and posterior distributions (density) for the estimated parameters. Figure 10 plots the prior distributions together with the posterior means and distributions. It shows that actual data are very informative about almost all of the estimated parameters. This is because posterior distributions are significantly different from the prior ones.³⁵

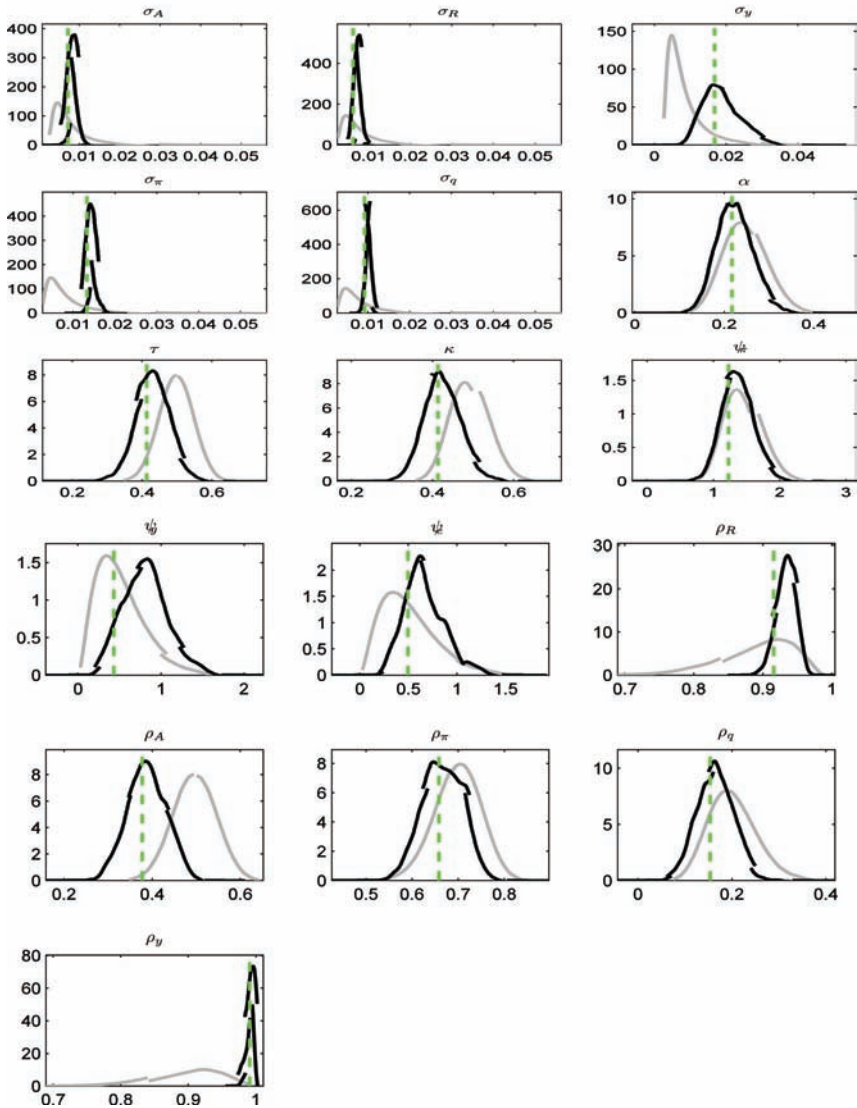
We focus our empirical analysis on the estimates of the monetary policy parameters in the pre- and post-1957 periods. The priors of the monetary policy parameters are assumed to be the same for the three samples; that is, they are treated symmetrically a priori. The prior of equal monetary policy pre- and post-1957 does not hold. Indeed, we find significant heterogeneity in the reaction of the monetary authority across the two sub-samples.

The estimates of the policy rule coefficients show that the Bank of Canada did not strongly respond to inflation, output, and exchange rate fluctuations during the pre-1957 period, as the estimates of the coefficients ψ_π , ψ_y , and ψ_e are 1.14, 0.50, and 1.49, respectively. Nevertheless, the responses to inflation, output, and the nominal exchange rate were more aggressive in the post-1957 period. The estimates of the coefficients ψ_π and ψ_y are larger at 1.78 and 0.63, respectively, while the estimate of ψ_e is smaller at 0.87. The estimates of the smoothing coefficient, ρ_R , are 0.94 and 0.74 in the pre- and post-1957 sub-samples. This implies higher interest rate inertia in the pre-1957 period compared with the post-1957 period. It is important to note that the estimated standard error of unsystematic monetary policy shocks, σ_R , is 0.29 percent in the pre-1957 period, whereas it is 0.91 percent during the post-1957 period, indicating

³⁵Prior and posterior distribution plots are similar for the three sample estimations, so we report only those for the entire floating-period estimation.

Figure 10. Prior and Posterior Distributions

(Sample: 1952:M1–1961:M12)



Note: Posterior distribution = bold line; prior distribution = grey line; posterior means = dashed line.

that monetary policy shocks, in general, were much more volatile in the post-1957 period than in the pre-1957 period.³⁶ Thus, the Bank of Canada appears to have been more strongly smoothing the nominal interest rate during the pre-1957 period, given the greater estimated interest rate inertia and the lower volatility of policy shocks.

Table 1 also reports the estimates of the remaining non-monetary-policy parameters. The estimate values of α , τ , and κ are very similar in the three samples, indicating some stability in the values of these parameters across the three periods. The estimates of these parameters are also close to those estimated in Lubik and Schorfheide (2007). Technology and world output shocks are similarly persistent and more volatile in the two periods, while terms of trade and world inflation were more volatile in the pre-1957 period.

To evaluate the performance of the model, we report volatilities, autocorrelations, forecast-error-variance decompositions, and the model's fit for output, the nominal interest rate, CPI inflation, and changes of the nominal exchange rate using the estimates for the entire floating and post-1957 periods. These exercises represent different ways of evaluating the performance of our model. Table 2 reports volatilities and autocorrelation functions of key simulated variables (output, the nominal interest rate, inflation, and the first difference of the nominal exchange rate) and compares them with those observed in the data. For the two estimated periods, the model generates unconditional volatilities and autocorrelations that are close to those observed in the data, except for the exchange rate. Overall, these results indicate that the model largely captures the salient features of the Canadian data.

We also consider the forecast-error-variance decompositions. Table 3 shows the variance decompositions of the variables attributed to each of the five shocks for a one-month-ahead horizon. In both periods, foreign output shocks account for the bulk of the fluctuations in these variables. Monetary policy shocks also account

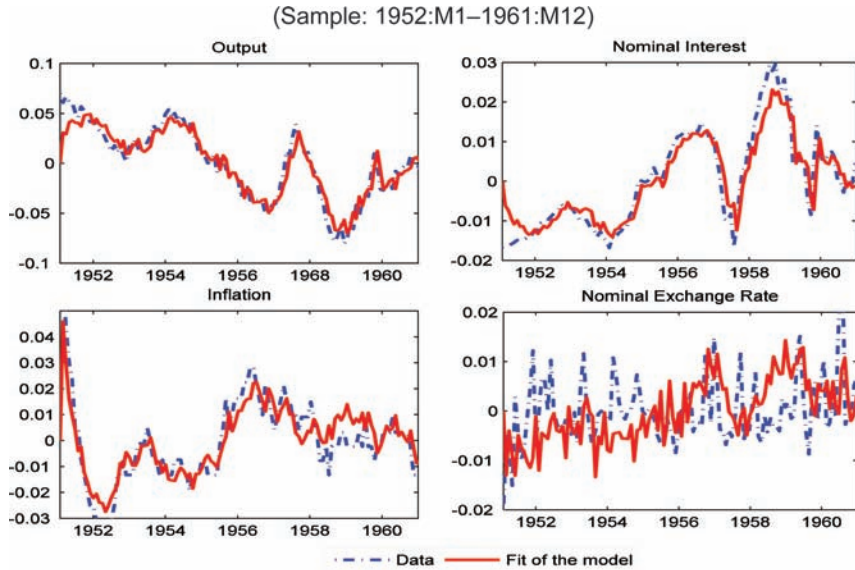
³⁶The estimated monetary policy is $R_t = 0.94R_{t-1} + 0.07\pi_t + 0.03y_t + 0.09\Delta e_t$ in the pre-1957 period, while it is $R_t = 0.74R_{t-1} + 0.46\pi_t + 0.16y_t + 0.23\Delta e_t$ in the post-1957 period. Therefore, responses to inflation, output, and changes of the nominal exchange rate were, respectively, seven, five, and three times more aggressive in the post-1957 period than in the pre-1957 period.

Table 2. Volatilities and Autocorrelations: Data and Estimated Model

Variables	Volatilities		Autocorrelations	
	Data	Model	Data	Model
A. Post-1957 Period				
Output	2.54	2.76	0.90	0.89
Nominal Interest Rate	1.07	0.95	0.90	0.79
CPI Inflation	1.02	1.25	0.84	0.67
Dif. (Nominal Exchange Rate)	0.69	1.53	0.28	0.61
B. Entire Floating Period				
Output	3.68	4.07	0.95	0.94
Nominal Interest Rate	1.22	1.25	0.96	0.93
CPI Inflation	1.77	1.78	0.93	0.72
Dif. (Nominal Exchange Rate)	0.65	1.85	0.33	0.68
Notes: Volatilities are measured by standard deviations in percentages. Simulated volatilities and autocorrelations are based on the posterior means for the post-1957 period and entire floating period as displayed in table 1.				

Table 3. Forecast-Error-Variance Decompositions

Variables	Percentage Owing To:				
	dA_t	ϵ_{Rt}	Δq_t	y_t^*	π_t^*
A. Post-1957 Period					
Output	1.92	5.93	0.67	90.6	0.86
Nominal Interest Rate	15.5	20.6	1.08	62.1	0.73
CPI Inflation	4.66	24.7	0.17	60.5	9.93
Dif. (Nominal Exchange Rate)	3.08	16.4	13.1	40.1	27.4
B. Entire Floating Period					
Output	0.98	2.71	0.69	93.6	2.05
Nominal Interest Rate	4.27	7.40	0.74	87.5	0.09
CPI Inflation	2.41	16.2	1.18	58.6	21.5
Dif. (Nominal Exchange Rate)	2.23	15.0	9.88	54.1	18.8

Figure 11. Data and Model's Fit: Entire Floating Period

for a significant fraction of the variance in the nominal interest rate, inflation, and the nominal exchange rate, as found in recent studies that have estimated DSGE models for Canada; for example, see Dib (2006, 2010).

To assess the ability of the model to explain the actual data, we plot the data against the model's predictions (at the posterior means) of output, the nominal interest rate, CPI inflation, and changes of the nominal exchange rate. Figure 11 shows that the model does remarkably well tracking output, the nominal interest rate, and inflation. It does, however, less well predicting changes of the nominal exchange rate, which is a variable hardly predicted in standard DSGE models. The model underpredicts the nominal exchange rate before 1957 and overpredicts it after 1957. Nevertheless, overall, the model is successful in fitting the data and it is useful for the counterfactual monetary policy analysis.

4.3 Counterfactual Simulations

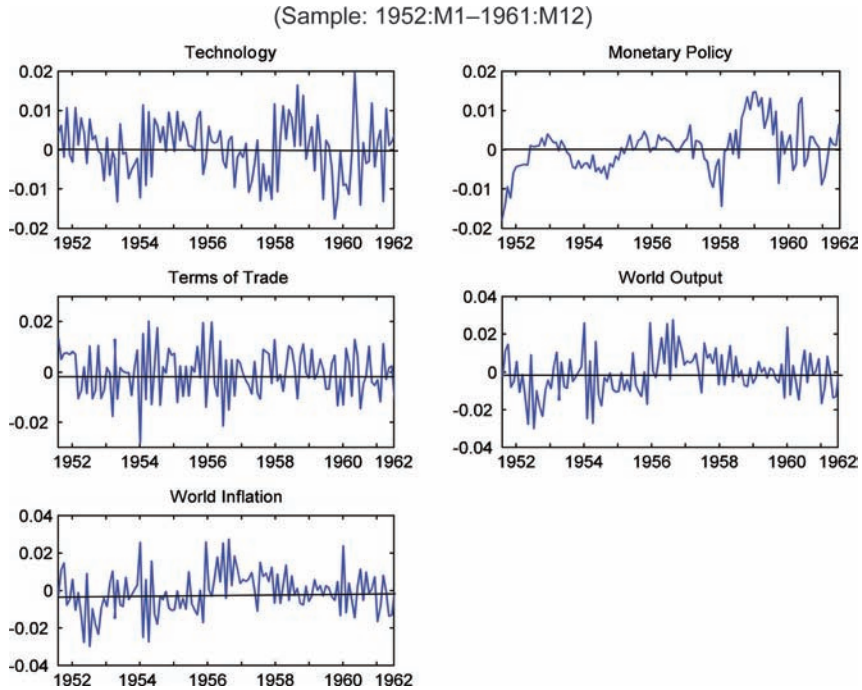
In order to perform counterfactual simulations of the model, we need to identify and extract the historical shocks that disturbed

the economy over the counterfactual periods, so that they can be reintroduced when the counterfactual model is simulated. For our rational-expectation model, we use the model's state-space representation, equation (12), to extract the smoothed shocks identified by a record of structural innovations $\{\varepsilon_{At}, \varepsilon_{Rt}, \varepsilon_{qt}, \varepsilon_{yt}, \varepsilon_{\pi t}\}; t = 1, \dots, T$, using a Kalman filter. We extract the historical shocks for the three estimated periods: the entire floating period, the pre-1957 period, and the post-1957 period. In this procedure, we use the estimates of the model's parameters for the three periods, as reported in table 1. Thus, our approach extends the methodologies used by Leigh (2004) and Boivin and Giannoni (2006) in their counterfactual analysis done for the Japanese and U.S. monetary policies, respectively. Our model is more complex and includes structural (as opposed to non-structural) shocks. All the generated shocks are i.i.d. and serially uncorrelated.

Figure 12 displays the extracted smoothed shocks for the entire floating rate period. What is noteworthy from this figure is that the volatility of the extracted monetary policy shocks is much higher in the second half of the floating period (which is the post-1957 period), reflecting the instability in monetary policy noted earlier. The volatilities of the other shocks are relatively stable throughout the period.

Before examining the results of the counterfactual simulations, it is important to note that conducting a counterfactual experiment in economics is problematic because a controlled experiment is impossible. To conduct the simulation, we set the model's parameters at their posterior mean values reported in table 1. Thus, by leaving the parameters unchanged, we are biasing the volatilities upward in the counterfactual simulations because agents would have likely acted in a way to mitigate welfare-reducing volatility.³⁷ In conducting the counterfactual experiments, we simulate the estimated model with the new assumptions regarding the conduct of monetary and/or exchange rate policies. We also incorporate the historical smoothed structural shocks, extracted via the Kalman filter, in an attempt to

³⁷In fact, we know that some parameters would likely have changed had a different monetary policy or exchange rate regime been in place because agents would have reoptimized.

Figure 12. Smoothed Shocks: Entire Floating Period

duplicate the actual economic circumstances that occurred during the sample periods.

We perform two types of counterfactual experiments. The first one, called the “pre-1957” monetary policy, assumes that the monetary policy of the pre-1957 (“pre-Coyne”) period is maintained through the post-1957 (“Coyne”) period. This experiment imposes the pre-1957 monetary policy over the post-1957 period and is conducted using the model’s structural parameters as estimated over the post-1957 period, except for the monetary policy coefficients. With this counterfactual experiment, we investigate what would have happened to the Canadian economy had the pre-1957 monetary policy been in place during the Coyne period.

In our model, the monetary policy consists of two components: the systematic component represented by the monetary policy rule and the unsystematic component represented by the smoothed monetary policy shocks. Therefore, in the monetary policy counterfactual

experiment, we conduct *three* simulations for the post-1957 period. We then compare the results across simulations and relative to the actual data. In the first or benchmark simulation for the post-1957 period, we simulate the estimated model using the post-1957 monetary policy rule but *without* any monetary policy shocks (monetary policy shocks are set equal to zero). In the second counterfactual exercise, we simulate the estimated model with the pre-1957 monetary policy rule but also *without* monetary policy shocks.³⁸ Hence the difference between the benchmark and second simulations is simply the pre- and post-1957 monetary policy rules (that is, the systematic component of monetary policy). Finally, we simulate the estimated model with both the pre-1957 monetary policy rule *and* the monetary policy shocks as extracted for the pre-1957 period. This simulation is interesting because the volatility of monetary policy shocks was greater in the post-1957 period than in the pre-1957 period, reflecting the more erratic nature of monetary policy over the latter period. Thus, this simulation captures the impact of this difference in volatility as well as the difference in the monetary rule across the two periods.

The second counterfactual experiment, called the fixed nominal exchange rate (FNER), assumes that instead of having a flexible nominal exchange rate, Canada had remained on the fixed nominal exchange rate over the floating period. Thus, this counterfactual experiment incorporates the commitment to maintain a fixed nominal exchange rate over the 1952–61 period. This is an interesting experiment because it could be argued that given the observed instability of monetary policy during the Coyne period, Canada might have been better off with a fixed nominal exchange rate. In this simulation, we assume that the goal of the monetary authority is to keep the exchange rate fixed, and the Canadian interest rate is simply assumed to follow the U.S. nominal interest rate.

For this experiment, we use the model's structural parameters as estimated over the entire floating period, 1952–61, except for Canadian monetary policy, which is assumed to follow U.S. monetary policy. Therefore, to perform this counterfactual experiment, changes of the nominal exchange rate, Δe_t , and all monetary policy

³⁸The unsystematic monetary policy shocks for the post-1957 period are set equal to zero in the first two simulations.

**Table 4. Volatilities from Counterfactual Experiments
(Standard Deviations in %)**

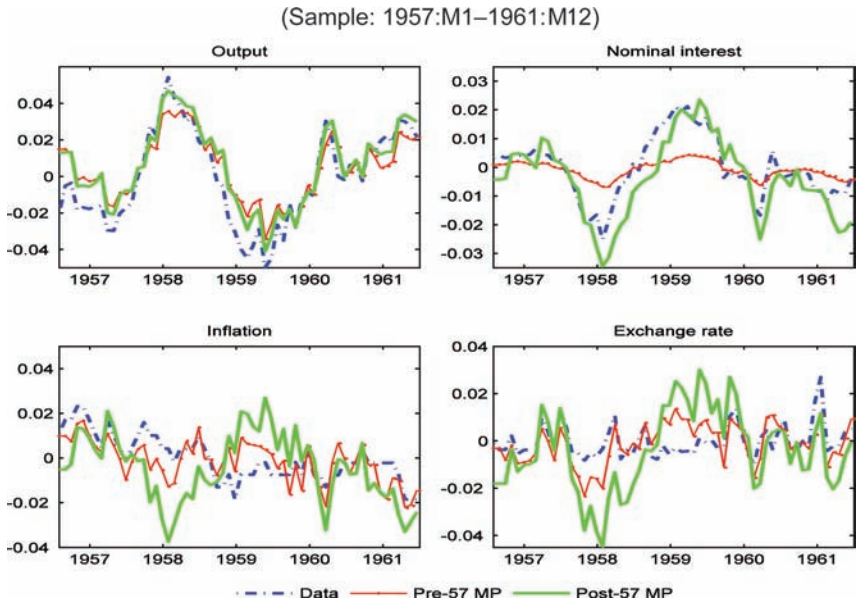
Variables	Monetary Policy (1957:M1–1961:M12)				Fixed NER (1952:M1–1961:M12)	
	Data	1	2	3	Data	FNER
Output	2.54	2.16	1.75	1.50	3.51	4.41
Nominal Interest Rate	1.07	1.35	0.29	0.20	1.22	1.01
CPI Inflation	1.02	1.52	0.95	1.65	1.77	5.24
Δ (Nominal Exchange Rate)	0.69	1.72	0.86	1.34	0.65	0
Notes: Results in column 1 are for the post-1957 period monetary policy rule <i>but</i> without monetary policy shocks (monetary policy shocks are set to zero). Results in column 2 are for the pre-1957 monetary policy rule <i>but</i> without policy shocks (monetary policy shocks are set to zero). Results in column 3 are for the pre-1957 monetary policy rule <i>and</i> the pre-1957 smoothed monetary policy shocks. In the FNER column, the monetary policy parameters, ρ_R , ψ_π , ψ_y , and ψ_e are set equal to zero, while the domestic nominal interest rate, R_t , is set equal to the U.S. interest rate, R_t^* .						

parameters— ρ_R , ψ_π , ψ_y , and ψ_e —are set equal to zero, while the Canadian nominal interest rate, R_t , is set equal to the U.S. nominal interest rate, R_t^* , as implied by the UIP condition given by equation (11).

Because the model is log-linearized around steady-state equilibrium values, our comparison between simulated counterfactual series and the actual data is based on second moments measured by the standard deviations in percentage. Table 4 reports the second moments of the data and those simulated for key counterfactual series: output, the nominal interest rate, CPI inflation, and changes of the nominal exchange rate. Also, the comparison between counterfactual series and the actual data is illustrated by the plots in figures 13 and 14 for the pre-1957 monetary policy and the fixed nominal exchange rate counterfactual experiments, respectively.

The second moments of the key variables calculated from the three counterfactual monetary policy simulations for the “pre-1957” counterfactual experiment are reported in table 4, columns 1, 2, and 3. First, we note that, under the pre-1957 monetary policy *but*

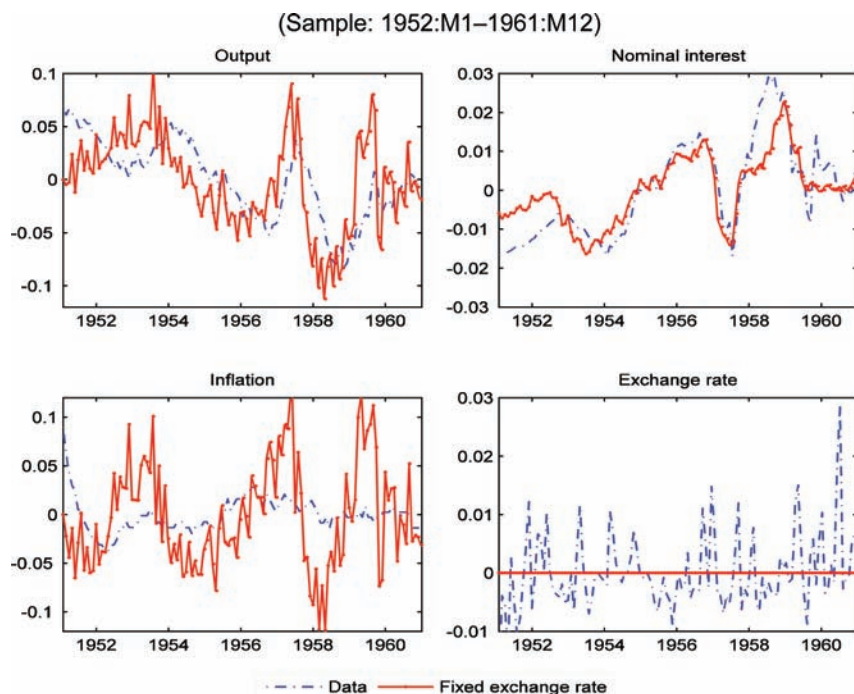
Figure 13. Data and Counterfactual Series: Pre-1957 Monetary Policy (MP) Experiment



without monetary policy shocks, the volatilities of output, the nominal interest rate, and inflation (column 2) are substantially *lower* relative to those of the actual data and of the benchmark case where the post-1957 monetary policy is maintained *but* the monetary policy shocks are set at zero (column 1). This finding is significant because the assumption of unchanged parameters *biases upward* the volatility of the macro variables in the experimental results. We also note that volatility of the nominal exchange rate is slightly higher in the pre-1957 monetary policy simulation without policy shocks, compared with the data, but significantly lower than in the benchmark case. Figure 13 plots the data and simulated series from the two counterfactual simulations with the pre- and post-1957 monetary policy rule but without policy shocks. This figure confirms that the variability of output, the nominal interest rate, and inflation would have been lower under the pre-1957 monetary policy relative to the actual data and the benchmark case.

Column 3 in table 4 indicates that, relative to the data, the volatilities of output and the nominal interest rate are much lower

Figure 14. Data and Counterfactual Series: Fixed Exchange Rate Experiment



in the simulation with the pre-1957 monetary policy *and* the unsystematic policy shocks, which are extracted from the pre-1957 period. In contrast, the volatilities of inflation and the nominal exchange rate are higher.

The counterfactual simulation results suggest that the pre-1957 monetary policy would have dramatically lowered the volatility of the nominal interest rate. This finding can be explained by the higher degree of monetary policy smoothness in the pre-1957 period relative to that in the post-1957 period.³⁹ It also reflects the less aggressive monetary policy response to inflation fluctuations during the pre-1957 period. In addition, the volatility of the exchange rate is higher

³⁹The estimated interest rate smoothing coefficient in the policy rule is 0.94 in the pre-1957 period, but is only 0.74 in the post-1957 period. Therefore, policy rates were much less volatile in the pre-1957 period and thus contribute to more stable output.

in the counterfactual simulation with the pre-1957 monetary policy relative to the data. These results suggest that under the pre-1957 monetary policy there was more scope for the flexible exchange rate to adjust to structural shocks, thereby helping to stabilize output.

In summary, the two sets of counterfactual simulation results displayed in columns 2 and 3 in table 4 indicate that the pre-1957 monetary policy would have stabilized output because simulated output volatility is lower than what was observed in the data and what was generated using the post-1957 monetary policy but without policy shocks. Therefore, changes to the monetary policy rule and monetary policy shocks were the main resources of large output disturbances that occurred in the second half of the 1950s. Thus, we claim that the pre-1957 monetary policy would have reduced the macroeconomic fluctuations in the Canadian economy.

Table 4 also reports the simulation results of the second counterfactual experiment of a fixed nominal exchange rate (FNER). Under this experiment, the volatilities of output and inflation drastically increase, while the volatility of the domestic nominal interest rate is, by assumption, the same as that for the U.S. nominal interest rate.⁴⁰ See figure 14. Although these volatilities are biased upward given the assumption of unchanged parameters, they suggest that the flexible exchange rate regime was successful in stabilizing the Canadian economy in the 1950s, even during the post-1957 period when monetary policy was more volatile.

5. Concluding Remarks

The purpose of this paper has been to revisit Canada's pioneering experience with floating exchange rate over the period 1950–62 and to reconsider the evidence and extensive commentary with the benefit of forty to fifty years of improvements in the tools of economic analysis and in our understanding of how monetary policy works in an open economy. Our goal is to draw useful lessons for the conduct of monetary policy today. One key observation is that Canada's floating exchange rate in the 1950s behaved very much like the exchange rate today; it was driven largely by interest rate

⁴⁰The volatility of the changes of the nominal exchange rate is equal to zero under the fixed exchange rate regime.

differentials and by commodity prices. The interesting question is whether the floating rate was the best option for Canada in the 1950s. There is little doubt that in September 1950, when the Canadian dollar was under incredible pressure to appreciate, the float was the correct choice because it removed the speculative pressure that pegged rates often and it sheltered Canada from U.S. inflationary pressures.⁴¹ Our counterfactual analysis indicates that relative to a fixed exchange rate, the flexible rate helped reduce the volatility of key macroeconomic variables, despite the fact the Canadian monetary authorities clearly did not understand all of the implications of conducting monetary policy under a flexible exchange rate and a high degree of capital mobility.

The other counterfactual experiment suggests that had monetary policy not changed during the Coyne period, the Canadian economy would likely have performed much better. The policy was not only more volatile but also produced higher interest rates. Consequently, output was more volatile and growth was likely slower because higher interest rates also generated a more appreciated exchange rate. This misunderstanding of the conduct of a monetary policy under a floating rate unfortunately led to the demise of this regime.

Nonetheless, Canada's floating exchange rate experience in the 1950s was a very useful policy experiment; it demonstrated (and inspired Mundell to show theoretically) that monetary policy is a relatively powerful policy instrument under a floating exchange rate; it also showed that for a flexible rate to work effectively as a stabilizing mechanism, as Friedman argued, monetary policy should not be the source of economic disturbances, but should work in tandem with the exchange rate to facilitate adjustment.

Appendix 1. Data for DSGE Model

Nominal Exchange Rate: Calculated as the monthly average of the daily CDN\$-US\$ noon spot rate as reported by the Bank for International Settlements.

⁴¹Mexico, in contrast to Canada, maintained a pegged rate and the inflation rate increased sharply to double-digit levels. See Murray, Schembri, and St-Amant (2003) for more details.

Real Exchange Rate: Calculated as the nominal CDN\$-US\$ exchange rate deflated by Canadian and U.S. CPI data.

Nominal Ninety-Day T-Bill Interest Rate: Prior to 1962, the Canadian T-bill interest rate is calculated as the average yield determined by auctions for ninety-day Treasury bills. From 1962 onward, the variable is simply the ninety-day T-bill rate. All Canadian data are taken from Statistics Canada. The U.S. T-bill rate is the three-month market yield as reported by the Bank for International Settlements.

CPI Inflation Rate: The year-over-year percentage change in the headline consumer price index as measured by Statistics Canada and the U.S. Department of Labor. The indices are rebased to the year 1997.

Terms of Trade: The price of exports over the price of imports, taken from Statistics Canada.

Industrial Production Index: The Canadian data consist of two series that are spliced together. For months prior to 1971, data are taken from Statistics Canada. After 1971, all data are obtained from the IMF's International Financial Statistics. U.S. data for all periods are taken from the Bank for International Settlements. The indices are rebased to the year 1997.

Commodity Terms of Trade: The price of commodity exports over the price of commodity imports, taken from Statistics Canada.

Unemployment Rate: Canadian data are obtained from hard copies of the "Labour Survey" published by Statistics Canada. U.S. data are taken from hard copies of the Bureau of Labor Statistics' household data publication (the "Current Population Survey").

Appendix 2. Log-Linearized DSGE Equilibrium System

$$y_t = E_t y_{t+1} - [\tau + \alpha(2 - \alpha)(1 - \tau)](R_t - E_t \pi_{t+1}) + \rho_A d A_t \\ + \alpha[\tau + \alpha(2 - \alpha)(1 - \tau)]\rho_q \Delta q_t + [\alpha(2 - \alpha)(1 - \tau)/\tau](\rho_y - 1)y_t^*, \quad (13)$$

$$\pi_t^d = \beta E_t \pi_{t+1}^d + \frac{\kappa}{[\tau + \alpha(2 - \alpha)(1 - \tau)]}(y_t - y_t^p), \quad (14)$$

$$y_t^p = -\frac{\alpha(2-\alpha)(1-\tau)}{\tau}y_t^*, \quad (15)$$

$$R_t = \rho_R R_{t-1} + (1-\rho_R) [\psi_\pi \pi_t + \psi_y y_t + \psi_e \Delta e_t] + \varepsilon_{Rt}, \quad (16)$$

$$\Delta e_t = \pi_t - (1-\alpha)\Delta q_t - \pi_t^*, \quad (17)$$

$$\pi_t = \pi_t^d - \alpha\Delta q_t, \quad (18)$$

$$\Delta q_t = \rho_q \Delta q_{t-1} + \varepsilon_{qt}, \quad (19)$$

$$\pi_t^* = \rho_\pi \pi_{t-1}^* + \varepsilon_{\pi^*t}, \quad (20)$$

$$dA_t = \rho_A dA_{t-1} + \varepsilon_{At}, \quad (21)$$

$$y_t^* = \rho_{y^*} y_{t-1}^* + \varepsilon_{y^*t}, \quad (22)$$

$$\Delta Y_t = y_t - y_{t-1} + dA_t. \quad (23)$$

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The Impact of Banks' Cumulative Reserve Position on Federal Funds Rate Behavior*

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We analyze the impact that reserve levels accumulated through the preceding day in a reserve maintenance period have on the level of the federal funds rate each morning prior to when open-market operations are arranged. Our empirical results and other evidence provided about intraday patterns of the federal funds rate demonstrate that the pace at which reserves are supplied over a maintenance period to meet banks' total reserve requirements is an important determinant of federal funds rate behavior.

JEL Codes: E5, G21.

1. Introduction

In this study we examine the factors that determine the level of the federal funds rate each morning, prior to when any open-market operations (OMOs) would typically be arranged by the Federal Reserve.¹ In particular, we focus on the impact from the quantity of reserves accumulated through the previous day in the same two-week

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¹The open-market operations that this statement refers to are the final operations arranged each day that would be intended to adjust the supply of reserves that day.

reserve maintenance period. The level of the morning federal funds rate is of interest because this rate is indicative not only of trading conditions at that time but also of the general level of rates that prevails through much of the day, often until near the close of the trading session and well after the time when any new OMOs might have been arranged. Thus, knowing how the cumulative level of reserves acquired over previous days of a maintenance period may influence the morning level of rates is essential for developing a reserve management strategy designed to control the federal funds rate over a multiday maintenance period.

The federal funds rate refers to the interest rate banks pay to borrow reserve balances held by other banks at the Federal Reserve in the federal funds market.² In this market, the vast majority of trades settle on the trade date and mature the next business day. For years, the Federal Open Market Committee (FOMC) has set a target level of the federal funds rate as its operating objective for OMOs arranged by the Open Market Trading Desk (the Desk) of the Federal Reserve Bank of New York (FRBNY). The OMOs used by the Desk typically take the form of purchases or sales of government securities, with purchases adding to reserve supply and sales decreasing reserves. Each day, the Desk estimates the aggregate level of reserve supply for that day and for upcoming days in the maintenance period, assesses those estimates against estimates of reserve levels needed to maintain the funds rate around the target, and arranges OMOs accordingly. In making its daily decisions about the level of reserves to provide each day through OMOs, the Desk will consider current trading conditions in the federal funds market, measured by the deviation of the rate in the morning from the target. But importantly, on occasions when this morning rate deviates from the target, the Desk will not necessarily just seek to provide the level of reserves that it would expect to foster an average, or “effective,” rate for that entire day that is equal to the target. Rather, in making a decision, it will also take into account the level of reserves that banks will have accumulated through that day in the

²Descriptions of the federal funds market can be found in Bartolini et al. (2005), Board of Governors of the Federal Reserve System (2005), and Hilton (2005). For simplicity, institutions in the funds market are referred to as “banks” in this study, although several other types of institutions actively participate in this market.

maintenance period and the expected impact of this reserve position on the federal funds rate the next morning.

Many earlier studies have looked for evidence of a liquidity effect by investigating the impact of exogenous daily reserve shocks to bank reserves and various calendar effects on the daily average level of the federal funds rate, measured relative to the target. This study represents a departure from previous efforts by examining determinants of the federal funds rate prior to the execution of any OMOs (the AM rate, defined below, measured relative to the federal funds rate target). We find that reserves accumulated through the previous day in the maintenance period generally have a negative relationship with the AM minus target rate spread, and that this relationship grows in magnitude and significance over the course of a two-week maintenance period. This result is consistent with banks' having a diminishing capacity to "average" daily reserve holdings for purposes of meeting reserve requirements as the end of the two-week maintenance period approaches. Overall, this analysis highlights the need for the Desk, and any central bank that operates with multiday reserve maintenance periods, to take into account the pace at which it provides reserves over an entire maintenance period in its daily reserve decisions.

The paper is structured as follows. Section 2 contains background information on the federal funds market, reserves and open-market operations, and the Desk's approach for controlling the federal funds rate. A literature review is found in section 3. Section 4 contains a description of the data, and empirical results can be found in section 5. Section 6 concludes.

2. Background³

In the portion of the federal funds market examined in this study, reserve balances in the accounts of commercial banks at the Federal

³The framework and procedures described in this section detail arrangements used by the Federal Reserve over the sample period of the study. Since September 2008, however, the Federal Reserve has adopted a very different approach to implementing monetary policy, one associated with interest rate objectives near zero, use of non-traditional operations, and the exercise of recently acquired authority to pay interest on reserve balances. Our sample period therefore ends with the last full maintenance period in August 2008. In the future, it is possible that the Federal Reserve will adopt operating procedures very different from the ones described and analyzed in this study.

Reserve are borrowed or loaned for one business day (overnight) and settle on the trade date. In volume, transactions with these characteristics account for the vast majority of all trading in the federal funds market. By convention, loans in this market are unsecured. Trading ends when Fedwire closes, scheduled for 6:30 p.m.⁴ The rate in this market, the federal funds rate, is sensitive to the supply of reserve balances, and for many years the FOMC has implemented monetary policy by setting a target for this rate and manipulating reserve levels to achieve this rate objective.

2.1 Determinants of Reserve Supply⁵

Using a simple balance-sheet identity, the supply of reserve liabilities of the Federal Reserve can be calculated by taking the difference between the level of domestic financial assets (less liabilities) held by the Federal Reserve plus discount-window loans outstanding, plus the net value (assets less liabilities) of all other factors on the Federal Reserve balance sheet—so-called “autonomous factors” over which the Desk has little or no direct control.⁶

Changes in autonomous factors, in domestic financial assets via OMOs, or in discount-window lending will add to or decrease the supply of reserves. For example, increases in Federal Reserve notes outstanding (currency), which are liabilities of the Federal Reserve and by far the largest of these autonomous factors, reduce reserves. When a bank requests a shipment of Federal Reserve notes from the Federal Reserve to satisfy their depositors’ demands for cash, the

⁴Fedwire is the wholesale electronic funds transfer network owned and operated by the Federal Reserve System. Federal funds transactions typically settle when the lending institution instructs the Federal Reserve via Fedwire to transfer reserves out of its account and into the account of the borrowing bank.

⁵In this study the term “reserves” is used to refer to the actual balances held by depository institutions at the Federal Reserve. As used in this study, the term does not include applied vault cash (which technically are reserves) and does include balances used to meet clearing-balance requirements (which technically are not reserves). Applied vault cash and clearing-balance requirements are defined below.

⁶Examples of major autonomous factors include balances in the Treasury’s account at the Federal Reserve and Federal Reserve notes outstanding. Weekly averages of the main autonomous factors and other items on the Federal Reserve’s balance sheet are published each Thursday by the Federal Reserve, in Publication H.4.1.

Federal Reserve debits the reserve account of the bank in payment for the currency that it ships. Although currency is the largest of these autonomous factors, some other factors are more volatile and have a less predictable impact on reserve supply.

Open-market operations—the purchase or sale of government securities by the Desk with securities dealers—are a discretionary tool that is used by the Desk to affect the supply of reserves. When the Desk purchases government securities, it pays for the securities it acquires by crediting the reserve balance of its counterparty, or of the correspondent bank of a non-bank counterparty. This action increases the supply of reserves. Sales of securities reduce reserve supply. And when the Federal Reserve extends a discount window loan to a bank, it credits the reserve balance of the borrowing bank for the amount of the loan and books the loan as an asset on its balance sheet.

2.2 *Reserve Accounting: Requirements, Penalties, and Carryover Privileges*⁷

Over each two-week *reserve maintenance* period beginning every other Thursday, each bank is required to hold a quantity of reserves in proportion to the quantity of transactions deposit liabilities on its balance sheet which, under lagged reserve accounting procedures in place since 1998, are based on amounts taken from a preceding two-week *reserve computation* period. This reserve requirement can be satisfied either by holding balances at the Federal Reserve during the reserve maintenance period or with cash held on a bank's premises during the reserve computation period (called "applied vault cash"). To meet the portion of reserve requirements not satisfied with vault cash, a bank must hold in total a sufficient amount of reserves over the entire maintenance period, based on its end-of-day levels. But it

⁷Several changes to the operating environment occurred over the course of the sample period, including (i) a shift from a contemporaneous to a lagged reserve accounting framework in 1998, (ii) termination of the adjustment credit facility and creation of the primary credit facility in 2003, and (iii) widespread adoption of retail sweep programs by depository institutions, which significantly reduced their deposit liabilities subject to reserve requirements. None of these developments materially affected the approach used by the Federal Reserve to control short-term interest rates. In this section, only the latest arrangements in place for the sample period are described.

may accumulate these reserves in any daily pattern it desires, subject to overdraft policies that apply to reserve positions at the end of each day.

Clearing-balance requirements represent obligations to hold reserves that are set at the discretion of a bank before each reserve maintenance period begins. Only balances held at the Federal Reserve during the two-week reserve maintenance period are eligible to satisfy clearing-balance requirements.

A bank is penalized for ending any day overdrawn on its account at the Federal Reserve, as well as for failing to meet its combined reserve and clearing-balance requirements at the end of the maintenance period. To obtain the necessary reserves to avoid these fees if it is unable to borrow the necessary amount in the market, a qualifying bank may borrow reserves directly from the Federal Reserve at the primary credit facility (commonly called the discount window) at a rate typically set 100 basis points above the target federal funds rate.⁸ This spread between the primary credit facility rate and the target funds rate is generally viewed as representing a *de facto* penalty associated with being reserve deficient. During our sample period, the Federal Reserve did not pay interest on reserves held in excess of requirements.⁹ Thus, the opportunity cost of holding excess reserves is a bank's marginal funding cost, which is represented in most studies by the federal funds rate.

To provide banks with some flexibility in meeting their requirements to avoid these penalties and costs, the Federal Reserve allows banks to apply excess reserve balances held in one maintenance period to meet reserve requirements in the following period, in an amount up to 4 percent of reserve requirements in the following period. Similarly, a bank may end a period up to 4 percent short of its

⁸During the period of market turmoil that began in the fall of 2007, the spread between the primary credit facility rate and the target funds rate was narrowed.

⁹Banks also did not earn interest on reserves used to meet reserve requirements. Banks did earn income credits, at a rate based on market rates, on the balances held to meet clearing-balance requirements. Credits earned can be used only to pay for priced services offered by the Federal Reserve, such as check-clearing fees and Fedwire transfer charges. In 2006, the Federal Reserve Act was amended, granting the Federal Reserve authority to pay interest on balances held by depository institutions, including balances used to meet reserve requirements, effective October 1, 2011. The Emergency Economic Stabilization Act of 2008 accelerated the effective date to October 1, 2008, and the Federal Reserve began paying interest on reserves on October 9, 2008.

reserve requirements and pay no penalty, so long as it holds sufficient excess reserves in the following period to offset this deficiency.¹⁰

2.3 Link between the Supply of Reserves and the Federal Funds Rate

Participation in the brokered segment of the federal funds market, from which the rate data used in this study is taken, is dominated by larger banking institutions. The rate in this segment of the market is influenced by the probabilities and costs these banks associate either with ending any day overdrawn or maintenance period deficient of their requirements, or with accumulating excess reserves. The probability of each of these outcomes will be influenced by the aggregate supply of reserves available on any given day and for a maintenance period as a whole.

The supply of reserves needed to maintain the federal funds rate around its target is closely linked to the level that allows banks to meet all their requirements, but without accumulating excess reserves, while simultaneously avoiding daily overdrafts.¹¹ When the supply of reserves is so low that some banks have no alternative but to borrow at the Federal Reserve's discount window to avoid ending a maintenance period short of their requirements or ending any day overdrawn, upward pressure on the funds rate will build, until

¹⁰Instead of having carryover flexibility, clearing-balance requirements are considered to have been satisfied so long as a bank accumulates an amount of reserves anywhere within a 2 percent band around its specified requirement.

¹¹Smaller-sized banks that don't have access to funding markets which would allow for a quick adjustment in their reserve balance typically demand some level of excess reserves as a source of liquidity, to guard against reserve-draining shocks. As a group, these smaller banks historically have held about \$1.5 billion of reserves in excess of their requirements each day. Reflecting their inability to access broad funding markets, this "frictional" demand amongst smaller banks has proven to be largely insensitive to both current trading conditions in the funds market and to the level of the funds target. In describing the link between the supply of reserves and the federal funds rate, the excess reserve demands of these banks are ignored with no loss of generality. In effect, their holdings of excess reserves are akin to an autonomous factor that reduces reserve supply, which the Desk must offset in its reserve provision to ensure that the larger banks can accumulate sufficient reserves to meet their requirements. Demands for excess reserves by small and large banking institutions are discussed in Federal Reserve Bank of New York (2006) and have been described in earlier annual reports in this series.

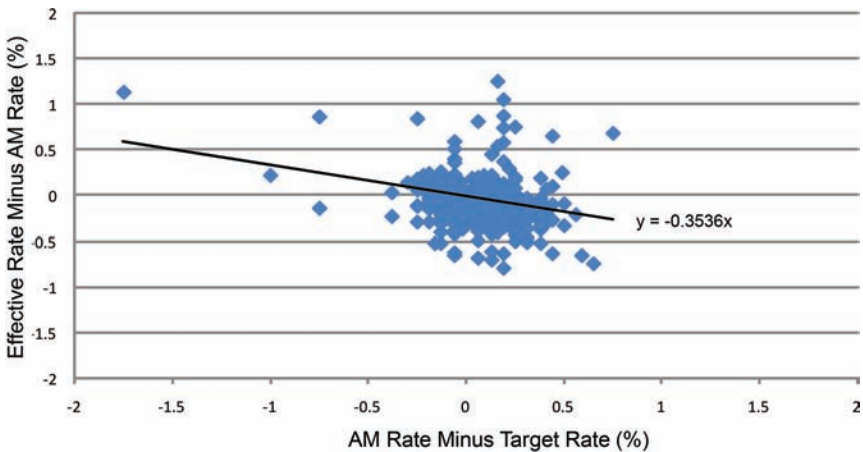
market rates reach a level at which a bank would prefer to pay the primary credit facility rate for borrowing directly from the Federal Reserve. And when reserve supply is so high that some banks face the prospect of accumulating unwanted excess reserves for the period, rates will fall off. In effect, at least for larger banks that actively manage their reserve balance by trading in the federal funds market, requirements are binding and are the principal determinant of reserve demand over an entire maintenance period.

Despite the ability of banks to freely substitute reserve holdings between days within a maintenance period for purposes of meeting their requirements, the funds rate can be influenced by the daily distribution of reserve supply, even if the level of reserves provided for the period as a whole is consistent with banks just meeting their requirements and sufficient each day to allow banks to avoid daily overdrafts. For example, relatively high levels of reserves provided by the Desk early in a period could put downward pressure on rates at that point in the period by fostering a perception among banks that they face a heightened risk of ending the maintenance period with excess reserves. Conversely, a low pace of reserve provision in the early part of a maintenance period could place upward pressure on rates early in a period, by raising fears among banks that they are at greater risk of failing to meet their requirements before the end of the period. Counterparty credit limits that all banks impose as lenders and face as borrowers reinforce these perceptions, as they constrain any bank's ability to make large adjustments to its daily or cumulative reserve position over short intervals without potentially having to make severe rate concessions on trades in the market, irrespective of the aggregate supply of reserves that subsequently might be available.

2.4 Daily Desk Operating Procedures and Intraday Rate Behavior

In making a decision each day about the level of reserves to provide that same day through OMOs, the Desk will consider current trading conditions in the federal funds market, measured by the deviation of the rate in the morning from target, with a goal of pushing the rate in the desired direction in later trading. However, when this morning rate differs from the target, the Desk will not just simply seek to provide the level of reserves that it estimates would yield an average rate

Figure 1. Scatter Plot for AM Rate Minus Target Rate



Note: Observations between $-.01\%$ and $.01\%$ and the first, middle, and last business day of each month are removed.

for the day equal to the target. Doing so would require a deliberate engineering of above-target (below-target) levels of the rate later in the day to offset below-target (above-target) rates that might have prevailed earlier in the day. Moreover, empirical evidence indicates that the federal funds rate tends to be relatively stable over much of the trading session, from the opening often until very late in the day, well after the time any new OMOs might have been arranged (Bartolini et al. 2005). Thus, seeking to fully offset morning deviations in rate levels from target could entail a very substantial overshooting of rates relative to the target late in the session.

As evidence of this proposition, we examine the relationship between the spread between the morning rate and the federal funds rate target ($AM - Tar$) and the spread between the daily effective federal funds rate and the AM rate ($Eff - AM$) from 1994 to 2008 on days when the morning rate is more than 1 basis point either above or below the target.¹² The following simple relationship is estimated and the underlying data presented in figure 1:

$$(Eff - AM)_t = \beta'(AM - Tar)_t + \varepsilon_t.$$

¹²The morning rate and the effective rate will be more thoroughly defined in section 4.

Consistent with the Desk's intention to push rates toward the target late in the day when the rate deviates from the target in the morning, this relationship is negative. But the estimate of β ($-.3536$) is less than -1 , indicating that the Desk does not generally seek to achieve an *average* daily rate equal to the target when making its reserve decisions on days when the morning rate deviates substantially from the target.¹³

In choosing the quantity of reserves to leave in place on any given day, the Desk will also take into account the cumulative level of reserves that banks will have acquired through that day in the maintenance period as a result of any new OMOs it arranges, and the expected impact of this cumulative reserve position on the federal funds rate the next morning (which, based on the findings in Bartolini et al. 2005 cited earlier, would also be expected to prevail over much of the trading session the next day). For this reason, the relationship between cumulative reserve levels and morning rates is a critical factor that helps determine the pace at which the Desk will provide reserves over an entire two-week reserve maintenance period consistent with its operating objective for targeting the federal funds rate.

3. Literature Review

The earlier literature up through the mid-1990s examining the impact of monetary policy focused on the relationship between the monetary aggregates (e.g., M1 and M2) and interest rates, and used low-frequency data (monthly or quarterly). The goal was to investigate the existence of a liquidity effect, basically a negative relationship between the money supply and interest rates. The interest rate examined was either the federal funds rate or a short-term interest rate such as the rate on three-month Treasury bills (see table 2 in Pagan and Robertson 1995 for a summary of selected studies). Studies employing a single equation, typically regressing the interest rate on the money supply and other economic variables, tended to find no liquidity effect. Later studies employed simultaneous-equation estimation techniques and were more likely to find evidence of a liquidity effect (for example, Galí 1992 and Sims 1992).

¹³The coefficient has a t -statistic of -14.98 and the adjusted- R^2 value is $.104$.

The work of Hamilton (1997) was among the first that focused on the federal funds market at the daily frequency. Hamilton examined the change in the daily average federal funds rate and its relationship with a supply shock to reserves, namely the forecast miss for Treasury's account held at the Federal Reserve. His results suggest the presence of a liquidity effect, although only on certain days of a maintenance period.

Thornton (2001) reexamines Hamilton's study and finds that Hamilton's results are driven by a few outliers. In addition, Hamilton's results do not appear to hold for other sample periods. Thornton then examines the relationship between non-borrowed reserves (total reserves less borrowing from the discount window) and changes in the target federal funds rate. Thornton does not find a statistically significant relationship and concludes that a liquidity effect does not exist at the daily frequency.

Carpenter and Demiralp (henceforth C&D) (2006b) examine the relationship between the effective federal funds rate minus the target rate and forecast misses for all the autonomous factors, not just the Treasury balance. This approach provides a more direct measure of a shock to the supply of bank reserves. They find that the liquidity effect exists at the daily frequency, and also that the liquidity effect is non-linear in that only large forecast misses (\$1 billion or more in absolute value) have a significant effect in the federal funds market. More recently, Thornton (2006) has extended C&D's analysis in a number of dimensions. For example, Thornton examines the change in the effective federal funds rate, examines a longer sample period, and deletes outlier observations. Thornton's results confirm C&D's findings of a liquidity effect, but the magnitude of the relationship is smaller.

Taking a different perspective, Bartolini, Bertola, and Prati (2002) examine the biweekly maintenance period patterns of the federal funds rate for evidence of a "martingale effect," but they do not explicitly consider how cumulative reserve positions might influence the daily pattern in rates over a maintenance period. Carpenter and Demiralp (2006a) do include a cumulative measure of excess reserves in their analysis of the effective federal funds rate, but the basis of this relationship is not explicitly explored. In contrast, our study focuses on the direct link between cumulative reserve levels

and the formation of rates in the federal funds market before new OMOs are arranged, and considers how this relationship influences the strategy for controlling the federal funds rate over an entire reserve maintenance period.

4. Data

The daily data used in our study cover the first full maintenance period starting in March 1994, roughly corresponding to when the Federal Reserve first began to publicly announce changes in the target federal funds rate, and run through the last full maintenance period in August 2008, when long-standing operating procedures and arrangements for implementing monetary policy were upended amid the financial crisis. The maintenance period that includes September 11, 2001, is omitted from the sample.

Data on the effective federal funds rate (Eff) and the target federal funds rate (Tar) is available from the FRBNY public web site (www.newyorkfed.org). The effective federal funds rate is a trade-weighted average of federal funds transactions provided to the Desk by a group of the major brokers in this market. The morning (AM) federal funds rate is a representative rate quote taken from an informal “reading” of the market conducted by the Desk, which is made by calling the major brokers, typically close to 9 a.m. The dispersion in rate quotes provided by the individual brokers is almost always extremely small. The morning federal funds rate and data on the supply of bank reserves are obtained from FRBNY’s proprietary database.

The focus of this study is on the determinants of the morning rate minus the target rate $(AM - Tar)_t$.¹⁴ The key explanatory variable is reserves accumulated through the previous day $t - 1$ (CR_{t-1}), measured relative to total maintenance period requirements (TMPR)—the quantity of reserves banks must accumulate over the course of an

¹⁴Note that we now include all observations during the sample period, aside from the observations in the maintenance period that includes September 11, 2001, as opposed to the analysis for figure 1.

Table 1. Summary Statistics

	Mean	Std. Dev.	Min.	Max.
Effective Minus Target (%)	0.013	0.16	-1.51	2.55
AM Minus Target (%)	0.034	0.12	-1.75	1.50
Effective Minus AM (%)	-0.021	0.14	-1.51	2.36
Cumulated Reserves/TMPR	0.463	0.30	0	1
Obs. = 3,634				

entire period to meet all requirements.¹⁵ We expect that the measure of reserves accumulated through the previous day (CR_{t-1})/TMPR will have a negative influence on the morning funds rate $(AM - Tar)_t$.

We control for a variety of calendar events that influence the behavior of the funds rate by including separate dummy variables for each day of a two-week maintenance period and for several calendar events (such as month start and month end dates, mid-month dates, and days after holidays). We also include a dummy variable for the day after the Northeast blackout on August 14, 2003, which had a pronounced impact on trading and rates.

To control for market expectations of policy changes in the target federal funds rate, we allow for anticipation effects up to five days ahead of each scheduled FOMC meeting date, but only for those days ahead of a meeting that fell within the same maintenance period as the meeting. Table 1 presents summary statistics for the main variables of interest in this study.

5. Empirical Framework and Results

A Lagrange multiplier test for no ARCH effects is rejected at the 95 percent confidence level, so we will present results from

¹⁵Note that TMPR has the same value on all days in a given maintenance period but may change from period to period. The calculation of TMPR was adjusted upward to include the Desk's estimate of frictional excess reserve demand of smaller banks. It also captures carryover positions larger banks brought into a maintenance period that could have generated some positive or negative excess demand. TMPR was constructed by taking actual total reserves provided during an entire maintenance period, less the size of any reserve supply projection miss made on the last day of the period and less primary credit discount borrowing on the final day. Carpenter and Demiralp's (2006a) measure of cumulative reserves is not standardized by a TMPR measure.

the Threshold GARCH (TARCH) model based on Glosten, Jagannathan, and Runkle (1993). The model is as follows:

$$(\text{AM} - \text{Tar})_t = \alpha + \beta' X_t + \varepsilon_t$$

$$\text{Var}(\varepsilon_t) = \sigma_t^2 = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 \varepsilon_{t-1}^2 (1(\varepsilon_{t-1} < 0)) + \gamma_3 \sigma_{t-1}^2.$$

The γ_1 coefficient will capture ARCH effects, and the γ_3 coefficient will capture GARCH effects. The γ_2 coefficient will capture whether or not negative shocks ($\varepsilon_{t-1} < 0$) have an asymmetric effect on the conditional variance (TARCH effects).

The results in table 2 show that all but three of the cumulated excess coefficients are negative, as expected. The three positive coefficients are all found in the first week of the maintenance period and are statistically insignificant. These results indicate that there is generally a negative relationship between cumulated excess reserves ($\text{CR}_{t-1}/\text{TMPR}$) and the $(\text{AM} - \text{Tar})_t$ spread for day t . Furthermore, the magnitude of the coefficients increases noticeably for the last two days of the maintenance period compared with the preceding days in the period. The estimated coefficients are not statistically significant on most of the days within the first week of a maintenance period, but are significant over most of the second week. These findings suggest that the sensitivity of the morning funds rate (relative to its target) to variations in the level of reserves cumulated through the previous day tends to rise as a maintenance period progresses. Overall, the results suggest that banks have a diminishing capacity to “average” reserve holdings as a maintenance period unfolds.¹⁶

Using the estimated constant terms and coefficients on the cumulative reserves variable ($\text{CR}_{t-1}/\text{TMPR}$) from table 2, for each day within a maintenance period, we calculate the level of cumulated reserves through the previous day associated with the morning funds rate being at the target (i.e., $(\text{AM} - \text{Tar})_t = 0$). For example, for the second Wednesday (W2) of the maintenance period, this

¹⁶The estimated values of the coefficients on the reserve variables and the day-in-period dummy variables are difficult to interpret directly. However, the expected deviation of the morning rate from the target associated with a given level of reserves accumulated through the previous day in the maintenance period can be calculated by summing the day-in-period dummy term for that day with the product of the cumulated reserves coefficient and the level of reserves cumulated through the previous day in the period.

Table 2. TARCH Results

Dependent Variable: AM Minus Target	Coeff.	t-stat.
AM minus Target _{t-1}	0.536	41.3
f1-Cumulated Reserves _{t-1} /TMPR	0.159	0.9
m1-Cumulated Reserves _{t-1} /TMPR	-0.071	-1.0
t1-Cumulated Reserves _{t-1} /TMPR	0.008	0.2
w1-Cumulated Reserves _{t-1} /TMPR	0.019	0.6
r2-Cumulated Reserves _{t-1} /TMPR	-0.137	-2.5
f2-Cumulated Reserves _{t-1} /TMPR	-0.101	-2.0
m2-Cumulated Reserves _{t-1} /TMPR	-0.094	-1.7
t2-Cumulated Reserves _{t-1} /TMPR	-0.280	-4.8
w2-Cumulated Reserves _{t-1} /TMPR	-0.681	-7.2
r1	0.012	7.2
f1	-0.022	-1.7
m1	0.035	1.9
t1	-0.003	-0.2
w1	-0.011	-0.9
r2	0.072	2.8
f2	0.047	1.7
m2	0.084	2.1
t2	0.230	4.8
w2	0.623	7.3
month start/end or mid-month date	0.057	38.3
day after month start/end or mid-month date	-0.033	-14.9
day before quarter end (excluding year end)	-0.008	-3.6
quarter end (excluding year end)	0.163	40.0
quarter start (excluding year start)	0.011	1.1
day after quarter start (excluding year start)	-0.026	-3.0
day before year end	-0.078	-28.1
year end	0.019	1.8
year start	0.032	1.6
day after year start	-0.004	-0.2
day after Monday holiday	0.031	9.4
second day after Monday holiday	0.006	1.6
day after other holiday	0.024	4.5
second day after other holiday	0.010	2.3
FOMC _{t-5} *target change	0.001	4.4
FOMC _{t-4} *target change	0.0005	3.0
FOMC _{t-3} *target change	0.002	17.5
FOMC _{t-2} *target change	0.002	19.4

(continued)

Table 2. (Continued)

Dependent Variable: AM Minus Target	Coeff.	t-stat.
FOMC _{t-1} *target change	0.003	19.7
FOMC date*target change	-0.004	-20.1
day after FOMC date*target change	-0.000005	-0.01
day after Blackout	0.195	4.9
Conditional Variance:		
arch	0.239	18.0
tarch	0.296	13.3
garch	0.744	147.7
constant	-10.246	-65.5

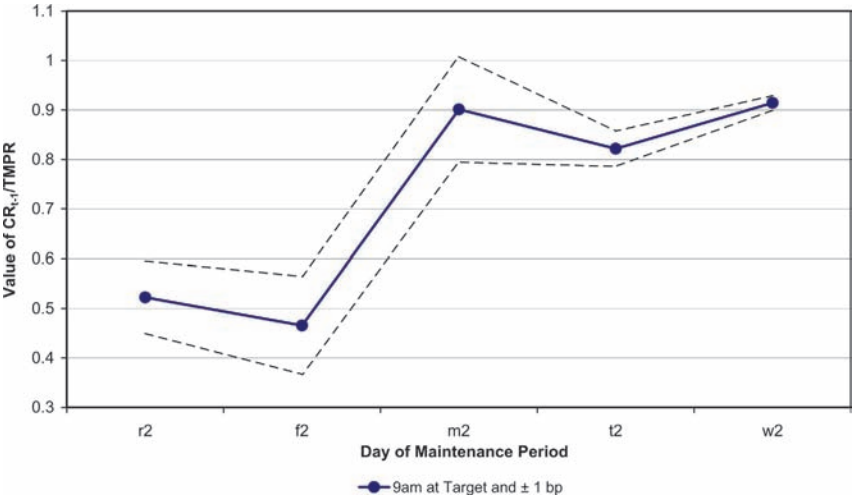
value is $(.623/.681) = .914$.¹⁷ That is, the funds rate would be expected to be at the target on the morning of the second Wednesday so long as banks accumulated about 90 percent of their total requirements for the maintenance period through the second Tuesday. These calculated values for each day over the second week of a maintenance period are presented in figure 2, along with the values for $(CR_{t-1}/TMPR)$ for each day of the period associated with the funds rate in the morning being 1 basis point above and 1 basis point below the target. As expected, the values of $(CR_{t-1}/TMPR)$ associated with keeping the funds rate around the target in the morning generally rise and approach a level of “1” in value. Moreover, the range of values of cumulative reserve levels associated with the morning funds rate being within 1 basis point of the target narrows as the period progresses.

6. Discussion and Conclusion

Our empirical analysis is differentiated from previous work in two important ways. First, it examines the federal funds rate prior to the execution of any daily OMOs. Second, accumulated reserve levels

¹⁷These calculations are based on an assumption that the values for all other variables in the equation, such as the dummy variables for high-payment-flow days, are zero.

Figure 2. Estimated Relation between $CR_{t-1}/TMPR$ and 9 a.m. Rate Relative to Federal Funds Target: Second Week of Maintenance Period



through the preceding day in a maintenance period are studied as a key variable of interest and determinant of this rate.

We cite previous empirical work and provide new evidence to support the argument that when the Desk makes its reserve management decisions each day, it may not simply seek to achieve an *average* daily rate for that day that is equal to the target on days when the morning rate deviates from the target. We then show that the level of accumulated reserves through the preceding day has a significant impact on the level of the morning rate. This impact strengthens as a maintenance period progresses, a finding that is consistent with banks' having a diminishing capacity to "average" reserve holdings as the number of days remaining in the period falls. These results highlight the need for the Desk, and analogously for other central banks that operate with multiday reserve maintenance periods, to consider the impact of cumulative reserve positions on the behavior of the funds rate on different days of a maintenance period when developing a reserve management strategy designed to control the federal funds rate for an entire maintenance period.

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The Sub-Prime Crisis and UK Monetary Policy*

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The “sub-prime” crisis, which led to major turbulence in global financial markets beginning in mid-2007, has posed major challenges for monetary policymakers. We analyze the impact on monetary policy of the widening differential between policy rates and the three-month LIBOR rate, the benchmark for private-sector interest rates. We show that the optimal monetary policy rule should include the determinants of this differential, adding an extra layer of complexity to the problems facing policymakers. Our estimates reveal significant effects of risk and liquidity measures, suggesting that the widening differential between base rates and LIBOR was largely driven by a sharp increase in unsecured lending risk. We calculate that the crisis increased LIBOR by up to 60 basis points; in response, base rates fell further and more quickly than would otherwise have happened as policymakers sought to offset some of the contractionary effects of the sub-prime crisis.

JEL Codes: C51, C52, E52, E58.

1. Introduction

The “sub-prime” crisis, which led to major turbulence in global financial markets beginning in mid-2007, has posed major challenges for monetary policymakers. Most prominence has been given to the

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attempts by policymakers to avoid systemic failures in financial institutions by means of liquidity injections and proposed regulatory reforms. But the crisis also posed new problems for policymakers in setting interest rates in order to steer the economy toward stable inflation and output levels. One of the main symptoms of the sub-prime crisis has been the widening differential between medium-term interest rates such as the three-month LIBOR rate and the short-term base rate set by policymakers. This differential is important since aggregate demand is more responsive to the LIBOR rate than to the base rate, as it is the benchmark interest rate that influences the interest rate at which the private sector, both corporate and personal, can borrow. A changing relationship between base and LIBOR rates implies that a given base rate implies a different level of aggregate demand and hence different levels of inflation and output. This added an extra layer of complexity to the problems facing policymakers.

This paper analyzes the effects of this changing relationship on the behavior of monetary policymakers in the United Kingdom. We begin by extending a prominent model of optimal monetary policy to introduce the distinction between the interest rate set by the policymaker and the interest rate that affects aggregate demand, something that is neglected by existing models. Doing so, we obtain an optimal monetary policy rule for the base rate that includes not just inflation and the output gap but also the determinants of the differential between the base rate and the rate at which the private sector can borrow. The emerging literature on the sub-prime crisis has identified risk and liquidity factors as being central to the changing relationship between LIBOR and the base rate; our model implies that these factors should also be components of the optimal policy rule. The augmented policy rule suggests two new insights into monetary policy: First, a rising differential implies that the level of aggregate demand can contract even if the base rate is constant or even falling, a situation that arguably occurred in the United Kingdom in early 2008 (e.g., Lomax 2007). This also implies that in the medium term a higher differential between LIBOR and base rates will imply lower base rates on average (Smith 2007). Second, any factor that affects the slope of the relationship between the base rate and LIBOR will affect the optimal response of interest rates to inflation and output. For example, we shall argue below that a

deterioration in market liquidity after mid-2007 made base rates more responsive to inflation and output.

We estimate an empirical version of the monetary policy rule, augmented by an equation for the yield-curve relationship between base rates and LIBOR. We use monthly data, as the interest-rate-setting Monetary Policy Committee (MPC) meets monthly. For inflation, we use the rate targeted by the MPC, the annual change in the Retail Prices Index excluding mortgage interest payments (RPIX) until December 2003 and the annual change in the Consumer Price Index (CPI) thereafter. Correspondingly, the inflation target is 2.5 percent until December 2003 and 2 percent thereafter. For output we use monthly GDP data. Following the literature on the sub-prime crisis (e.g., Michaud and Upper 2008, and Taylor and Williams 2008), we use the difference between rates on secured and unsecured borrowing in the interbank market as our main measure of risk in the relationship between the three-month LIBOR rate and the base rate. This has been used to capture the perceived risk in lending between banks when there is concern that the counterparty may default, a prominent issue during the crisis. For liquidity, we use the composite index published in the biannual *Financial Stability Report*. This index reflects bid-ask spreads, return-to-volume ratios, and liquidity premia using data for the United States, the euro zone, and the United Kingdom. For further details, see Bank of England (2007).

We find considerable empirical support for our model; the exclusion of risk and liquidity measures from the policy rule is rejected, as is the assumption of a constant response of interest rates to inflation and output. Unsecured lending risk and liquidity are significant determinants of the differential between base and LIBOR rates; however, the increase in the differential since mid-2007 is largely driven by increases in unsecured lending risk. Our evidence therefore further supports the argument that the sub-prime crisis was largely the result of the unwillingness of banks to enter the interbank market because of uncertainty of the value of assets on offer and, at times, because of fears of the solvency of their counterparties.

The remainder of the paper is structured as follows. The theoretical model is developed in section 2. Section 3 describes the variables used in our empirical model. Sections 4 and 5 contain our estimates. Section 6 concludes.

2. Theory

We use a slightly modified version of the canonical model of Svensson (1997) in which a standard Taylor rule emerges as the result of a theoretical framework in which policymakers adjust interest rates in order to pursue an inflation target. The model is

$$\pi_{t+1} = \pi_t + \alpha_y y_t + v_{t+1} \quad (1)$$

$$y_{t+1} = \beta_y y_t - \alpha_r (E_t i_t^{borrow} - E_t \pi_{t+1} - \bar{r}) + \eta_{t+1} \quad (2)$$

$$i_t^{borrow} = \omega_{0t} + \omega_{1t} i_t^{base} + \varepsilon_t. \quad (3)$$

Equation (1) is a Phillips curve in which inflation depends on inflation and the output gap in the previous period and on a supply shock (v). Equation (2) is an aggregate demand relationship in which the output gap (y) depends on the lagged output gap, the real interest rate at which the private sector can borrow relative to its equilibrium value, and on a demand shock (η). This differs from the standard formulation in using the nominal interest rate at which the private sector can borrow (i^{borrow}) rather than the base rate (i^{base}). Since the borrowing rate is closely linked to medium-term interest rates—most prominently, the three-month LIBOR rate—equation (3) relates the nominal borrowing rate to the base rate using a yield-curve relationship. We allow both the intercept and slope of this relationship to vary over time, to reflect the pronounced movements in the differential between medium-term rates and base rates that have been observed since mid-2007.

The model is a simple extension of Svensson (1997), which is obtained if $i_t^{borrow} = i_t^{base}$. To solve for the optimal policy rule, we follow the approach of Svensson. We assume that at time t policymakers choose current and future base rates to minimize the loss function

$$L_t = E_t \sum_{i=0}^{\infty} \delta^i \left\{ \frac{1}{2} (\pi_{t+i} - \pi^T)^2 \right\}. \quad (4)$$

Equation (4) specifies the policymakers' loss function as the discounted sum of expected quadratic deviations of inflation π from the inflation target π^T , where δ is the discount factor. We assume policymakers know the value of ε_t but not the other errors, which

become apparent at the end of the period. Since the base rate chosen at time t affects the inflation rate two periods ahead, the policymakers' problem is equivalent to minimizing $L_t = \delta^2 E_t \frac{1}{2} (\pi_{t+2} - \pi^T)^2$. The first-order condition is¹

$$E_{t-1} \pi_{t+2} = \pi^T, \quad (5)$$

where we assume that policymakers choose interest rates at the start of period t based on information available up to the end of period $t - 1$. Using (1)–(3), we can express the optimal policy rule as

$$\hat{i}_t^{base} = \frac{\bar{r} + \pi^T - \omega_{0t}}{\omega_{1t}} + \frac{1 + \alpha_r \alpha_y}{\omega_{1t} \alpha_r \alpha_y} (E_t \pi_t - \pi^T) + \frac{1 + \beta_y + \alpha_r \alpha_y}{\omega_{1t} \alpha_r} E_t y_t. \quad (6)$$

This extends the familiar Taylor rule by including the yield-curve factors, ω_{0t} and ω_{1t} . Intuitively, because aggregate demand and inflation depend on the borrowing rate, the optimal base rate will be a function of factors that affect the relationship between the base rate and the borrowing rate. The inclusion of these factors is consistent with evidence that yield-curve determinants improve the performance of Taylor rules in the United States (e.g., Piazzesi 2005).

3. Empirical Specification

In this section we develop an estimable model consisting of empirical versions of the yield-curve relationship in (3) and the optimal policy rule in (6). Beginning with the yield-curve relationship, the recent literature on the “sub-prime crisis” has focused on the widening differential between overnight and medium-term rates. Using daily data, Michaud and Upper (2008) consider the differential between the overnight indexed swap (OIS) rate and the three-month LIBOR rate and investigate the role of risk and liquidity factors in explaining the widening of this differential from mid-2007. They consider two measures of risk—the spread between secured and unsecured interbank rates and premia on credit default swaps—and argue that movements in the secured-unsecured spread are more closely related

¹This argument closely follows Svensson (1997); see equation (2.8) of that paper and the preceding discussion for a more extensive analysis.

to movements in the OIS-LIBOR differential. Liquidity is measured using indicators of trading volume, bid-ask spreads, and the impact of trades on prices, derived using data from the e-MID electronic trading platform. These measures have no clear relationship with movements in the LIBOR/base rate spread.

Taylor and Williams (2008) mainly consider the United States and also use daily data. They focus on the role of risk factors in explaining the spread between one- and three-month interbank rates and OIS rates. As with Michaud and Upper (2008), risk is measured using credit default swap premia and the spread between secured and unsecured interbank rates, although the former measure is preferred as a less noisy measure of risk. The effects of liquidity are confined to allowing for effects of the Term Auction Facility (TAF) introduced in December 2007. Both measures of risk and the effects of the TAF are significant in regressions explaining both the one- and three-month spreads.

Schwarz (2009) uses daily data from January 2007 to April 2008 to explain the spread between the OIS and one- and three-month LIBOR rates in the euro zone, again investigating risk and liquidity effects. Risk is measured using the spread between the highest and lowest decile of overnight borrowing rates for banks on the e-MID trading platform. Liquidity is measured using the differential in yields between bonds issued by a German state-backed development bank and government bonds of equal maturity, using the principle that the difference in yields between assets with similar maturity and risk characteristics must reflect liquidity (Longstaff 2004 uses a similar approach in measuring liquidity in the United States). The effect of liquidity is found to be stronger than in the other studies.

Our econometric model will further test some of these ideas and integrate them into a model that also allows us to analyze the response of monetary policymakers to the financial crisis. We use monthly data since the Monetary Policy Committee meets monthly. This forces us to use somewhat different explanatory variables than in the recent literature. We use the base rate rather than the overnight indexed swap rate, as this is the rate directly set by the Monetary Policy Committee and since the OIS rate is only available for a relatively short period of time. In practice this has little effect; for the period for which they are available monthly, OIS rates

are highly correlated with the base rate. We are also unable to use credit default swap premia, as data on these are only available from mid-2004, and so we use the spread between secured and unsecured interbank rates as our main measure of risk.

To capture liquidity effects,² we use the index of liquidity calculated by the Bank of England; unlike data from the e-MID platform, this is available on a monthly basis since 1992. This index reflects three factors: (i) bid-ask spreads (for gilt repos, the FTSE100 and major currencies) as a measure of the “tightness” of markets (Kyle 1985); (ii) the return-to-volume ratio (for gilts, the FTSE100 and equity options) as a measure of the impact of volumes of prices (Amihud 2002); and (iii) liquidity premia, measured as the spread between corporate bonds and a credit spread and between bond and LIBOR rates in the United States, the euro zone, and the United Kingdom. For further details, see Bank of England (2007).

On the basis of this discussion, the empirical version of the yield-curve equation in (3) is

$$\begin{aligned} i_t^{lib,3} = & \omega_{00} + \omega_{01}(i_t^{lib,1} - i_t^{repo,1}) + \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) \\ & + \omega_{03}liq_t^{FSR} + (\omega_{11} + \omega_{12}liq_t^{FSR})i_t^{base} + \varepsilon_t. \end{aligned} \quad (7)$$

Comparing (7) with (3),³ $\omega_{0t} = \omega_0 + \omega_{01}(i_t^{lib,1} - i_t^{repo,1}) + \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) + \omega_{03}liq_t^{FSR}$ and $\omega_{1t} = \omega_{11} + \omega_{12}liq_t^{FSR}$. $i_t^{lib,3}$ is the average three-month LIBOR rate. $(i_t^{lib,1} - i_t^{repo,1})$ is the differential between one-month LIBOR and gilt repo rates. This measures the differential between unsecured and secured lending rates.⁴ We also include the differential between the three-month LIBOR and one-month gilt repo rates to capture term-structure effects. liq_t^{FSR} is the liquidity index published by the Bank of England in the *Financial Stability Report*. We include an interaction term, whereby the effect of liquidity on $(i_t^{lib,3} - i_t^{base})$ varies with the base rate, to allow for changes in the slope of the yield curve.

²For more detailed analyses of the impact of liquidity, see papers by Goodhart, Crocker, and Tirole in Banque de France (2008).

³The interest rates in (7) are monthly averages of daily observations, obtained from the Bank of England web site.

⁴Borrowing on the LIBOR market is unsecured, while borrowing on the gilt repo rate market is secured.

For the optimal policy rule in (6), the main modeling choices are how to interpret the timing of events and which measures of inflation and the output gap to use. In our preferred specification, policymakers respond to forecasts of inflation and the output gap over the coming three months. Although this may appear to conflict with the specification of the optimal policy rule in (6), it has been argued that policymakers in effect review decisions every three months as forecasts of inflation, output, and the time profile of interest rates are updated in the quarterly *Inflation Report*. Using (7) to express ω_{0t} and ω_{1t} in terms of risk and liquidity factors, our empirical policy rule is then

$$\begin{aligned} \hat{i}_t^{base} = & \frac{\bar{r} + \pi^T - \omega_{00} - \omega_{01}(i_t^{lib,1} - i_t^{repo,1}) - \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) - \omega_{03}liq_t^{FSR}}{\omega_{11} + \omega_{12}liq_t^{FSR}} \\ & + \frac{\phi_\pi}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 (\pi_{t+k} - \pi^T) + \frac{\phi_y}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 y_{t+k} + \zeta_t, \end{aligned} \quad (8)$$

where $\phi_\pi = \frac{1+\alpha_r\alpha_y}{\alpha_r\alpha_y}$, $\phi_y = \frac{1+\beta_y+\alpha_r\alpha_y}{\alpha_r}$, and the error term reflects the errors induced by replacing expected values of inflation and the output gap with the realized ex post values. We use the inflation rate targeted by monetary policy, namely the annual change in the RPIX until December 2003 and the annual change in the CPI thereafter. Correspondingly, the inflation target is 2.5 percent until December 2003 and 2 percent thereafter. For output we use monthly GDP data (kindly provided by the National Institute of Economics and Social Research) and derive the output gap as the proportional difference between GDP and its Hodrick-Prescott trend. Finally, we allow for the effects of interest rate smoothing⁵ by expressing the observed

⁵Interest rate smoothing is difficult to model in the context of the Svensson approach to deriving the optimal monetary policy rule. As an alternative approach, we could derive the optimal policy rule from an amended model in which the loss function has quadratic terms in inflation, output, and interest rate changes. The resultant rule would be similar to (6) above but with an interest rate smoothing term.

base rate as a weighted average of the current optimal and previous base rates:

$$i_t^{base} = \rho i_{t-1}^{base} + (1 - \rho) \hat{i}_t^{base}, \quad (9)$$

where i_t^{base} is the observed base rate and the optimal base rate, $\hat{i}_t^{base}$, is given by (8). Combining (8) and (9),

$$\begin{aligned} i_t^{base} = & \rho i_{t-1}^{base} + (1 - \rho) \\ & \times \left\{ \frac{\bar{r} + \pi^T - \omega_{00} - \omega_{01}(i_t^{lib,1} - i_t^{repo,1}) - \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) - \omega_{03}liq_t^{FSR}}{\omega_{11} + \omega_{12}liq_t^{FSR}} \right. \\ & \left. + \frac{\phi_\pi}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 (\pi_{t+k} - \pi^T) + \frac{\phi_y}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 y_{t+k} + \zeta_t \right\}. \end{aligned} \quad (10)$$

Our empirical model comprises equations (7) and (10).

There are two sets of testable restrictions that simplify our model to models estimated elsewhere in the literature. First, the yield-curve model in (7) simplifies to models estimated in the existing literature if $\omega_{11} = 1$ and $\omega_{12} = 0$, in which case

$$\begin{aligned} i_t^{lib,3} - i_t^{base} = & \omega_{00} + \omega_{01}(i_t^{lib,1} - i_t^{repo,1}) + \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) \\ & + \omega_{03}liq_t^{FSR} + \varepsilon_t. \end{aligned} \quad (11)$$

This simplified model for the LIBOR/base rate differential is similar to the model in Michaud and Upper (2008) for the LIBOR-OIS differential, where, as they discuss, variations in liquidity between financial institutions and market microstructure effects are captured in the error term. Although the theoretical framework in Taylor and Williams (2008)—based on the Ang and Piazzesi (2003) model of the term structure—is rather different, the model they estimate is consistent with (11). With these restrictions, (10) simplifies to

$$\begin{aligned}
i_t^{base} = & \rho i_{t-1}^{base} + (1 - \rho) \left\{ \bar{r} + \pi^T - \omega_{00} - \omega_{01} (i_t^{lib,1} - i_t^{repo,1}) \right. \\
& - \omega_{02} (i_t^{repo,3} - i_t^{repo,1}) - \omega_{03} liq_t^{FSR} \\
& \left. + \phi_\pi \sum_{k=0}^2 (\pi_{t+k} - \pi^T) + \phi_y \sum_{k=0}^2 y_{t+k} + \zeta_t \right\}. \quad (12)
\end{aligned}$$

In this policy rule, risk and liquidity continue to affect base rates, but the optimal response to inflation and output is now independent of these factors.

A second simplified model is obtained if $\omega_{01} = \omega_{02} = \omega_{03} = \omega_{12} = 0$, in which case our model simplifies to

$$i_t^{lib,3} - i_t^{base} = \omega_{00} + \varepsilon_t \quad (13)$$

and

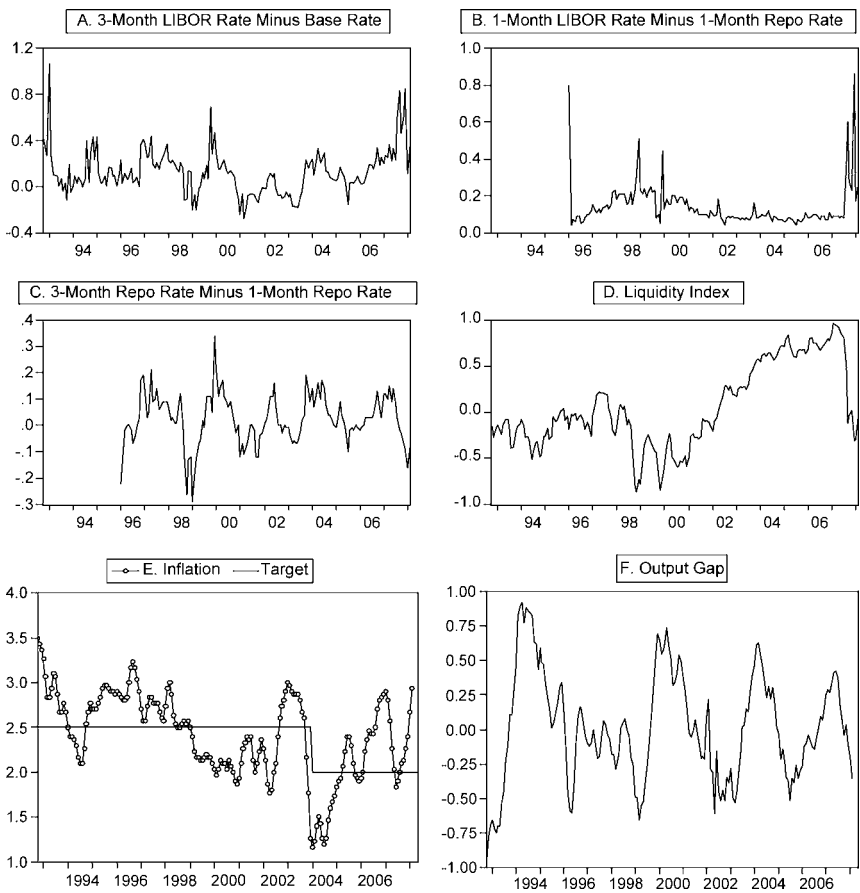
$$\begin{aligned}
i_t^{base} = & \rho i_{t-1}^{base} + (1 - \rho) \\
& \times \left\{ \bar{r} + \pi^T - \omega_{00} + \phi_\pi \sum_{k=0}^2 (\pi_{t+k} - \pi^T) + \phi_y \sum_{k=0}^2 y_{t+k} + \zeta_t \right\}. \quad (14)
\end{aligned}$$

In this case there is a fixed proportional relationship between the LIBOR and base rates, and our policy rule simplifies to the familiar Taylor rule.

Our empirical strategy will be to compare estimates of the system comprising (7) and (10) with the simplified models in (11)–(12) and (13)–(14). We also note that the model in (7) and (10) implies that estimates of each of (11)–(14) will suffer from parameter instability. We will use this as an additional test of the specification of our model.

The main features of our data are depicted in figure 1. Figure 1A shows the differential between the three-month LIBOR and base rates; this exhibits a sharp increase in mid-2007, followed by fluctuations around an elevated level until the end of our sample. Figure 1B shows $(i_t^{lib,1} - i_t^{repo,1})$, where we see a sharp jump in mid-2007 that matches the sharp rise in $(i_t^{lib,3} - i_t^{base})$. Movements in $(i_t^{repo,3} - i_t^{repo,1})$, depicted in figure 1C, are less dramatic and are

Figure 1. Interest Rate Spreads, Liquidity, Inflation, and Output Gap in the United Kingdom



less correlated with $(i_t^{lib,3} - i_t^{base})$. Movements in this liquidity index are depicted in figure 1D, where it is apparent that liquidity fell sharply in mid-2007, having increased steadily over the previous five years.⁶ Figures 1E and 1F depict the inflation rate relative to the target and the output gap.

⁶The index is expressed in standardized form, relative to the mean value of the mid-1990s and where the vertical scale measures deviations in terms of standard deviations; data on the liquidity index are available since 1992, whereas data on the LIBOR and repo rates are only available since 1996.

4. Estimates

Our main estimates are presented in table 1. We estimated our system using GMM. We treat all variables as endogenous, using the first four lags of each as instruments. Column 1 presents estimates of (7) and (10), column 2 has estimates of (11)–(12), and column 3 has estimates of (13)–(14). There is considerable support for our model in equations (7) and (10). We reject $\omega_{01} = \omega_{02} = \omega_{03} = \omega_{12} = 0$, which simplify the model to that in column 2, and reject $\omega_{11} = 1$ and $\omega_{12} = 0$, which simplify the model to that in column 3. We also find parameter instability on each of (11)–(14), which further supports our preferred specification.

The estimated parameters of our preferred model are all significant and of the expected sign (this is not true of the simplified model in (11)–(12)). In the policy rule, the responses of interest rates to inflation and output are comparable to those estimated in the existing literature, although the estimated equilibrium real interest rate is perhaps a little low. In the LIBOR equation, increases in unsecured lending risk and medium-term risk are associated with a larger differential. Liquidity has two offsetting effects on the differential between LIBOR and the base rate; an increase in liquidity reduces the differential, via the negative estimate on ω_{03} , but also increases it via the positive estimate ω_{12} .

Our estimates imply a more complex policy rule than is usually considered, including measures of the risk and liquidity that affect the relationship between the base rate set by policymakers and the rate at which the private sector can borrow. To illustrate this, we conducted a simple counterfactual analysis by calculating the implied predicted value of the base rate, assuming that our risk and liquidity measures were fixed at their 2007:Q1 values for the remainder of our sample.⁷ As figure 2 shows, base rates would have risen by 50 basis points in response to rising inflation in the spring of 2007 before falling by 25 basis points late in the year. With risk and liquidity factors unchanged, movements in the three-month LIBOR rate would have mirrored these changes. The impact of increasing

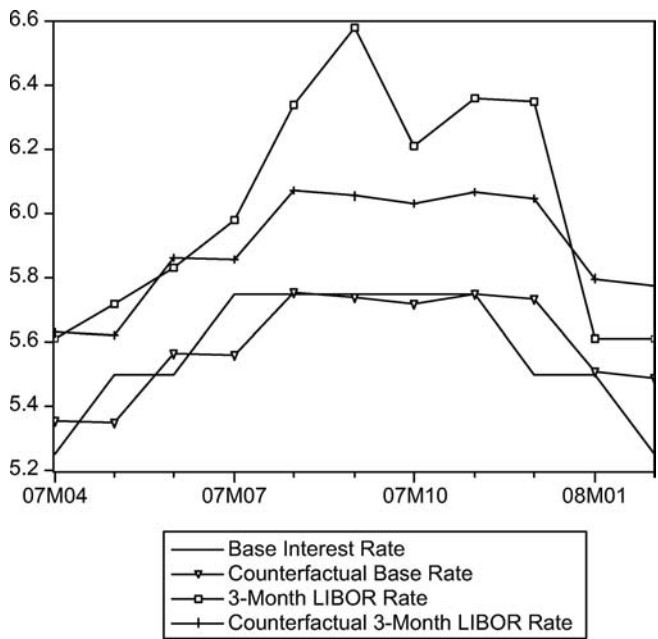
⁷We use realized values of inflation and output; this can, of course, only be justified in the very short run and in the context of a limited illustration of the implications of our results.

Table 1. Main Estimates

	1	2	3	4	5	6
Eqs. Estimated	(7) and (10)	(11)–(12)	(13)–(14)	(15)–(16)	(17)–(18)	(7) and (10)
	1996:M1–2008:M2	1996:M1–2008:M2	1996:M1–2008:M2	1992:M10–2008:M2	2000:M12–2008:M2	1996:M1–2008:M2
Policy Rule						
ρ	0.900 (0.040)	0.894 (0.030)	0.911 (0.042)	0.891 (0.041)	0.901 (0.012)	0.910 (0.042)
$\bar{\tau}$	1.109 (0.174)	3.076 (0.056)	3.053 (0.057)	2.466 (0.161)	−0.672 (0.129)	1.134 (0.199)
ϕ_{π} (Inflation)	1.130 (0.161)	1.572 (0.173)	1.628 (0.150)	1.537 (0.128)	0.567 (0.085)	1.082 (0.184)
ϕ_y (Output Gap)	2.276 (0.136)	2.780 (0.173)	2.524 (0.137)	1.721 (0.116)	1.942 (0.107)	1.130 (0.125)
Regression	0.15	0.17	0.17	0.17	0.10	0.16
Standard Error						
$\bar{R}^2$	0.98	0.97	0.97	0.98	0.97	0.97
Parameter Stability (p-value)	0.27	0.00	0.00	0.28	0.11	0.21
LIBOR Equation						
ω_{00}	0.159 (0.024)	−0.150 (0.011)	0.120 (0.009)	0.080 (0.039)	0.321 (0.016)	0.155 (0.050)
ω_{01} (Unsecured Lending Risk)	1.688 (0.087)	1.689 (0.063)			1.677 (0.045)	1.175 (0.130)
ω_{02} (Medium- Term Risk)	1.196 (0.080)	1.102 (0.060)			−0.535 (0.067)	1.033 (0.118)

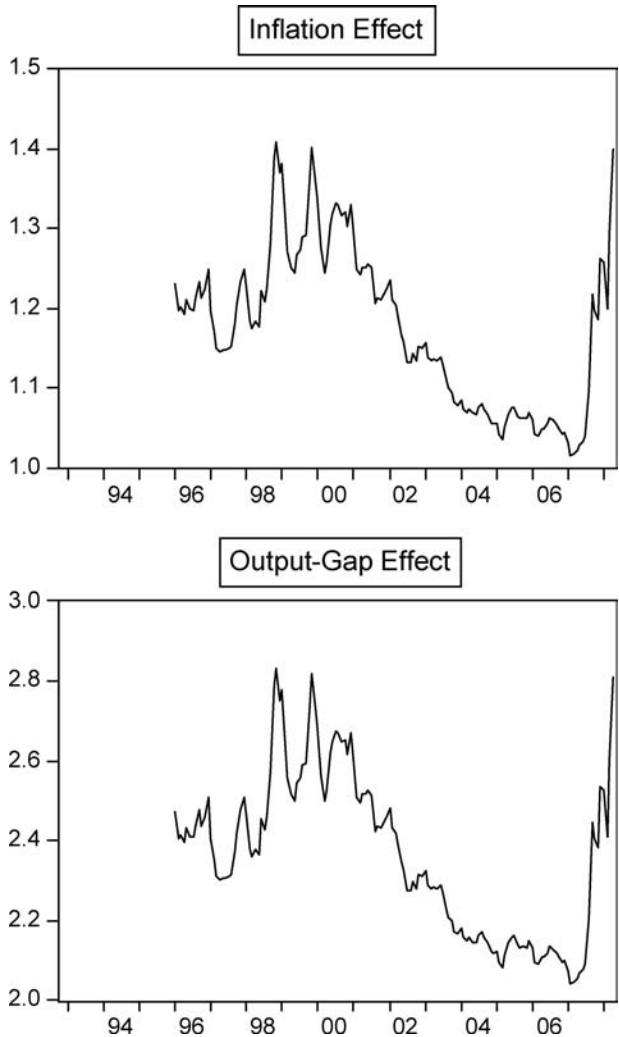
(continued)

Figure 2. Counterfactual Experiment



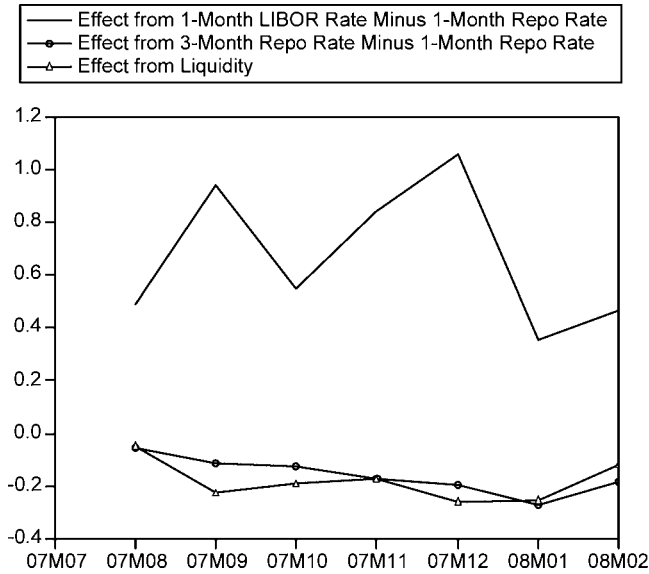
risk and deteriorating market liquidity is apparent in the divergence between the actual and counterfactual LIBOR rates, leading to an additional increase of up to 60 basis points in the LIBOR rate, representing a sharp tightening in monetary policy in the summer and autumn of 2007. The impact of risk and liquidity on the base rate is also apparent; the base rate was cut by 25 basis points in December 2007 compared with a counterfactual cut in January 2008, and the further 25-basis-point reduction in February 2008 is not predicted by this counterfactual experiment. This illustrates the impact of the sub-prime crisis on monetary policy, both in terms of increasing the LIBOR rate relative to the base rate and in terms of lowering the base rate itself. The impact of the sub-prime crisis on the response of interest rates to inflation and output is depicted in figure 3; this shows that the expansion in liquidity in the early years of this century reduced these responses, which then returned to the higher levels of the late 1990s following the sharp deterioration in liquidity in 2007–8.

Figure 3. Time-Varying Inflation and Output-Gap Effects



Note: Figure 3 plots the time-varying inflation effect $(\frac{\phi_{\pi}}{\omega_{11}+\omega_{12}liq_t^{FSR}})$ and the time-varying output-gap effect $(\frac{\phi_y}{\omega_{11}+\omega_{12}liq_t^{FSR}})$ using estimates of equation (10) reported in table 1.

Figure 4. Decomposing the Differential between the Three-Month LIBOR Rate and Base Rate



Note: The decomposition is based on the estimates of equations (7) and (10) reported in table 1.

Our estimates suggest that the key indicator of the “sub-prime” crisis, the rise in the differential between LIBOR and base rates, is largely explained by increases in unsecured lending risk. Figure 4 shows the results of using our model to decompose the rise in this differential since July 2007 (this is also discussed in the appendix). The offsetting effects of liquidity largely cancel themselves out, while, as expected, the yield-curve effects of the differential between the three- and one-month gilt repo rates are small. The effect of the increase in the differential between one-month LIBOR and gilt repo rates is dominant. These estimates support the view that the rise in the perceived risk of unsecured lending on the interbank market was the main driving force behind the sub-prime crisis.

The remainder of the paper considers the robustness of our findings. First, we use the differential between corporate bonds and ten-year government bonds, both to assess this alternative measure of

risk and to assess the effects of a longer sample, since we have data on this from 1992. We amend our model to be

$$\begin{aligned} i_t^{lib,3} = & \omega_{00} + \omega_{03}liq_t^{FSR} + \omega_{04}(i_t^{corp} - i_t^{gb}) \\ & + (\omega_{11} + \omega_{12}liq_t^{FSR})i_t^{base} + \varepsilon_t \end{aligned} \quad (15)$$

$$\begin{aligned} i_t^{base} = & \rho i_{t-1}^{base} + (1 - \rho) \left\{ \frac{\bar{r} + \pi^T - \omega_{00} - \omega_{03}liq_t^{FSR} - \omega_{04}(i_t^{corp} - i_t^{gb})}{\omega_{11} + \omega_{12}liq_t^{FSR}} \right. \\ & + \frac{\phi_\pi}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 (\pi_{t+k} - \pi^T) \\ & \left. + \frac{\phi_y}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 y_{t+k} + \zeta_t \right\}, \end{aligned} \quad (16)$$

where i^{corp} is the corporate bond rate and i_t^{gb} is the ten-year government bond rate. Estimates of this model, presented in column 4 of table 1, are similar to those in column 1. The restrictions that would simplify this model to (11)–(12) or (13)–(14) are again rejected. Second, we use the overnight interest rate swap rate in place of the base rate. The model in this case is

$$\begin{aligned} i_t^{lib,3} = & \omega_{00} + \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) + \omega_{03}liq_t^{FSR} \\ & + (\omega_{11} + \omega_{12}liq_t^{FSR})i_t^{ois} + \varepsilon_t \end{aligned} \quad (17)$$

$$\begin{aligned} i_t^{ois} = & \rho i_{t-1}^{ois} \\ & + (1 - \rho) \left\{ \frac{\bar{r} + \pi^T - \omega_{00} - \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) - \omega_{03}liq_t^{FSR}}{\omega_{11} + \omega_{12}liq_t^{FSR}} \right. \\ & + \frac{\phi_\pi}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 (\pi_{t+k} - \pi^T) \\ & \left. + \frac{\phi_y}{\omega_{11} + \omega_{12}liq_t^{FSR}} \sum_{k=0}^2 y_{t+k} + \zeta_t \right\}, \end{aligned} \quad (18)$$

where i^{ois} is the monthly average overnight interest rate swap rate. We are limited by lack of data here, since data on OIS rates are only available from late 2000. This seems to have particularly affected the

estimates on the inflation term in the policy rule and on medium-term risk in the LIBOR equation. This aside, the main features of our estimates are unchanged and we are unable to simplify to either (11)–(12) or (13)–(14).

We also tried estimating (7)–(10) using the inflation forecasts provided by the Bank's *Inflation Report* and using real-time data on output. These experiments are hampered by the fact that inflation forecasts and real-time data are only available on a quarterly frequency; to overcome this we assumed a constant inflation forecast for each month within the same quarter and a constant growth rate of real-time output data within a quarter. Doing so makes no qualitative difference to our empirical results, although estimates of the policy rule become less reliable. Estimates of models which used a quadratic trend filter and the Baxter and King (1999) filter to detrend output data were similar to those reported in table 1.

Column 6 of table 1 reports estimates of equations (7)–(10) which use an alternative measure of liquidity—the difference between the one-year government bond rate and the one-year local authority rate (an increase in the spread denotes improved liquidity). This measure is similar in spirit to the liquidity measure used by Schwarz (2009), as these assets have the same maturity and are both essentially risk free. Estimates of our model using this liquidity measure are similar to those reported in column 1 of table 1 and are significant; the remaining estimates are in line with the earlier ones.

Finally, we note that our estimates suffer from some residual ARCH effects, which seem to be associated with the liquidity measure, which is highly volatile toward the end of the sample. Estimating (7) and (10) with a correction for these ARCH effects has little effect on the estimates. Alternatively, we obtain a model free of ARCH effects if the liquidity variable is dropped from the model, estimates of the other parameters being relatively unaffected. We prefer not to omit liquidity effects, as these are important factors in the sub-prime crisis.

5. A Counterfactual Experiment over the Most Recent Period

In this section we address the robustness of our estimates by assessing their post-sample forecasting ability over the period from March

2008 to February 2009. This is a formidable test of our model, as this period includes the collapse of Lehman Brothers and the rescue of AIG and other institutions in September 2008, the banking crisis of October 2008, and the rapid sequence of deep cuts to central bank policy rates that left them at unprecedented lows by early 2009.

We construct the counterfactual values of the base rate and the three-month LIBOR rate over this more recent period using the estimates of equations (7) and (10) reported in column 1 of table 1. These are depicted in figure 5, alongside the actual values of these variables. Our model provides a good account of the rapid reduction in the Bank of England policy rate and also picks up most, but not all, of the fall in the LIBOR rate. Table 2 quantifies these counterfactual experiments. The average value of the base rate is lower than the counterfactual, with a lower standard deviation. The average value of LIBOR is lower than the counterfactual, with a higher standard deviation. However, these differences are not statistically significant. On balance, these experiments suggest that our estimates are robust and are able to account for the most recent period despite the extreme turbulence of that period.

6. Conclusions

We have analyzed the impact of the sub-prime crisis that began in mid-2007 on the interest-rate-setting behavior of UK monetary policymakers. Our focus is on the widening differential between base rates and the three-month LIBOR rate, the latter being a key determinant of aggregate demand, as it is the baseline against which many interest rates relevant to the private sector are set. In order to do so, we extend a familiar model of optimal monetary policy, due to Svensson (1997), to allow for the distinction between the interest rate set by the Central Bank and the interest rate relevant to private-sector expenditure decisions. We show that the resulting optimal policy rule includes the determinants of the differential between the two interest rates and that factors which affect the slope of the relationship between the interest rates affect the optimal response of interest rates to inflation and output.

We estimate our model using UK data using the three-month LIBOR rate to measure the interest rate relevant to aggregate demand and following the literature on the “sub-prime” crisis in

Figure 5. Actual and Counterfactual Values of Base and Three-Month LIBOR Rate, 2008:M3–2009:M2

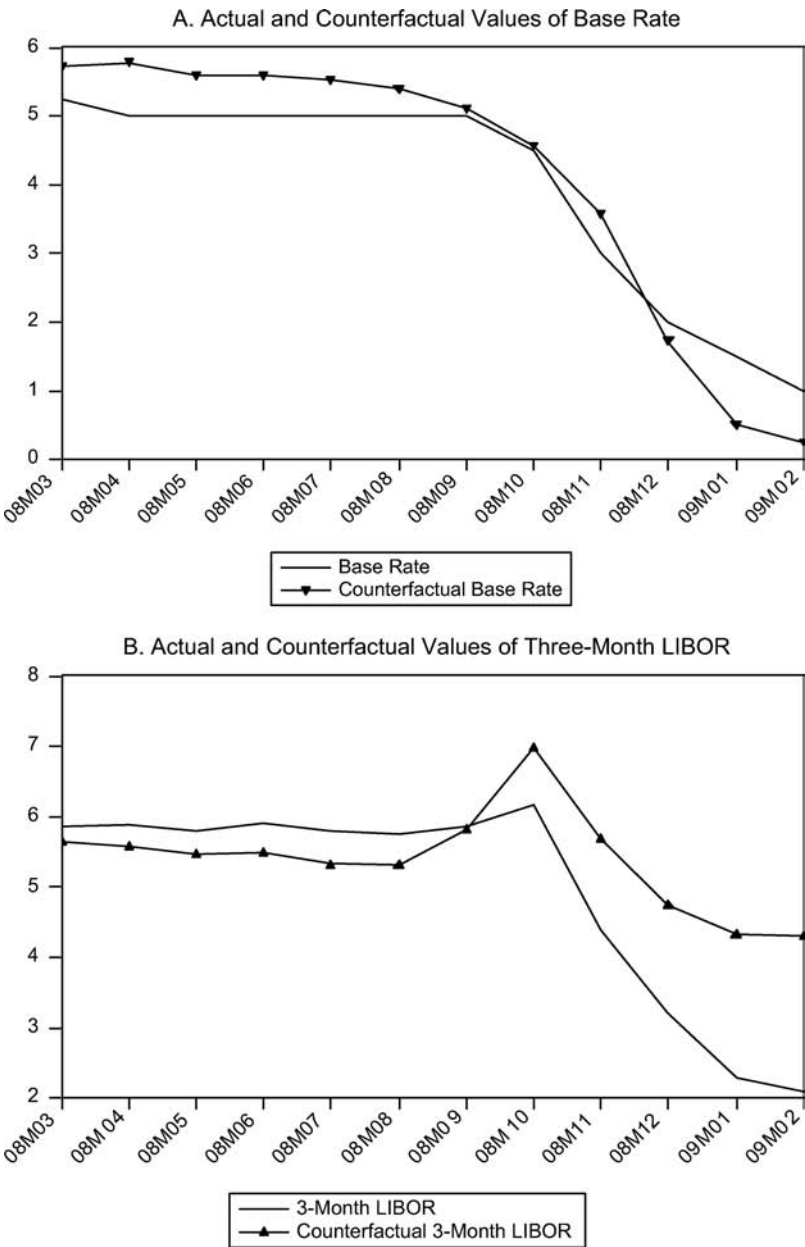


Table 2. Mean Values and Standard Deviations of Actual and Counterfactual Base and Three-Month LIBOR Rate, 2008:M3–2009:M2

	Mean	Standard Deviation
i_t^{base}	3.94	1.59
$i_t^{base,counter}$	4.10	2.10
Equality Test: F-test and p-value in Brackets	0.03 [0.86]	1.74 [0.37]
$i_t^{lib,3}$	4.92	1.49
$i_t^{lib,3,counter}$	5.39	0.74
Equality Test: F-test and p-value in Brackets	0.95 [0.33]	2.22 [0.03]

using measures of risk and liquidity as determinants of the differential between this and the base rate. Our estimates support our model. We find strong effects from both unsecured lending risk, measured by the difference between the one-month LIBOR and gilt repo rates, and liquidity; exclusion of these factors from the policy rule is strongly rejected by the data. We use our model to investigate the effects of the sub-prime crisis. We calculate that the effects of the sub-prime crisis increased the three-month LIBOR rate by up to 60 basis points in the summer and autumn of 2007, representing a significant tightening in monetary policy. They also affected the base rate, which fell further and more quickly than would otherwise have happened as policymakers sought to offset some of the contractionary effects of the sub-prime crisis on aggregate demand. We also establish that the rise in the differential between the three-month LIBOR rate and base rates was largely driven by unsecured lending risk, supporting the view that the perceived risk of unsecured lending on the interbank market was the main driving force behind the sub-prime crisis, perhaps because of the unwillingness of banks to enter the interbank market in view of the uncertain value of the assets on offer and fears of the solvency of counterparties.

Although suggestive, our work is necessarily preliminary, as the sub-prime crisis is still ongoing at the time of writing, in summer 2009. A more definitive analysis must wait for the end of the crisis

and for its macroeconomics effects to have unwound. We intend to return to this in future work.

Appendix. Decomposing Changes in the Relationship between the Borrowing Rate and the Base Rate

The relationship between the borrowing rate and the base rate is

$$i_t^{borrow} = \omega_{0t} + \omega_{1t}i_t^{base}. \quad (19)$$

So the differential between the borrowing rate and base rate is

$$diff_t = i_t^{borrow} - i_t^{base} = \omega_{0t} + \omega_{1t}i_t^{base} - i_t^{base}, \quad (20)$$

where $\omega_{0t} = \omega_0 + \omega_{01}(i_t^{lib,1} - i_t^{repo,1}) + \omega_{02}(i_t^{repo,3} - i_t^{repo,1}) + \omega_{03}liq_t^{FSR}$ and $\omega_{1t} = \omega_{11} + \omega_{12}liq_t^{FSR}$.

The impact of a change in a variable x on the differential is then

$$\frac{\partial diff_t}{\partial x_t} = i_t^{borrow} - i_t^{base} = \frac{\partial \omega_{0t}}{\partial x_t} + \frac{\partial \omega_{1t}i_t^{base}}{\partial x_t} - \frac{\partial i_t^{base}}{\partial x_t}. \quad (21)$$

Since

$$\begin{aligned} \omega_{1t}i_t^{base} &= \rho i_{t-1} \\ &+ (1 - \rho) \left\{ \bar{r} + \pi^T - \omega_{0t} + \phi_\pi \sum_{k=0}^2 (\pi_{t+k} - \pi^T) + \phi_y \sum_{k=0}^2 y_{t+k} \right\}, \end{aligned} \quad (22)$$

then

$$\frac{\partial \omega_{1t}i_t^{base}}{\partial x_t} = -(1 - \rho) \frac{\partial \omega_{0t}}{\partial x_t}. \quad (23)$$

Also since

$$\begin{aligned} i_t^{base} &= \frac{\rho}{\omega_{1t}} i_{t-1} \\ &+ (1 - \rho) \left\{ \frac{\bar{r} + \pi^T - \omega_{0t}}{\omega_{1t}} + \frac{\phi_\pi}{\omega_{1t}} \sum_{k=0}^2 (\pi_{t+k} - \pi^T) + \frac{\phi_y}{\omega_{1t}} \sum_{k=0}^2 y_{t+k} \right\}, \end{aligned}$$

then

$$\begin{aligned}
\frac{\partial i_t^{base}}{\partial x_t} = & -(1 - \rho) \left(\frac{1}{\omega_{1t}} \frac{\partial \omega_{0t}}{\partial x_t} - \frac{\omega_{0t}}{\omega_{1t}^2} \frac{\partial \omega_{1t}}{\partial x_t} \right) - \frac{\rho}{\omega_{1t}^2} \frac{\partial \omega_{1t}}{\partial x_t} i_{t-1} \\
& - (1 - \rho)(\bar{r} + \pi^T) \left(\frac{1}{\omega_{1t}^2} \frac{\partial \omega_{1t}}{\partial x_t} \right) \\
& - (1 - \rho) \left(\frac{1}{\omega_{1t}^2} \frac{\partial \omega_{1t}}{\partial x_t} \right) \phi_\pi \sum_{k=0}^2 (\pi_{t+k} - \pi^T) \\
& - (1 - \rho) \left(\frac{1}{\omega_{1t}^2} \frac{\partial \omega_{1t}}{\partial x_t} \right) \phi_y \sum_{k=0}^2 y_{t+k},
\end{aligned}$$

which simplifies to

$$\frac{\partial i_t^{base}}{\partial x_t} = -(1 - \rho) \frac{1}{\omega_{1t}} \frac{\partial \omega_{0t}}{\partial x_t} - \frac{i_t^{base}}{\omega_{1t}} \frac{\partial \omega_{1t}}{\partial x_t}. \quad (24)$$

Combining (21)–(24),

$$\frac{\partial diff_t}{\partial x_t} = \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right) \frac{\partial \omega_{0t}}{\partial x_t} + \frac{i_t^{base}}{\omega_{1t}} \frac{\partial \omega_{1t}}{\partial x_t}. \quad (25)$$

Next we calculate individual effects.

Unsecured Lending Risk ($i_t^{lib,1} - i_t^{repo,1}$)

$$\frac{\partial \omega_{0t}}{\partial x_t} = 1.688 \text{ and } \frac{\partial \omega_{1t}}{\partial x_t} = 0.$$

Therefore,

$$\frac{\partial diff_t}{\partial (i_t^{lib,1} - i_t^{repo,1})} = 1.688 \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right),$$

so change due to changes in unsecured lending risk is

$$1.688 \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right) \{ (i_t^{lib,1} - i_t^{repo,1}) - (i_{july\ 2007}^{lib,1} - i_{july\ 2007}^{repo,1}) \}.$$

Medium-Term Lending Risk ($i_t^{repo,3} - i_t^{repo,1}$)

$$\frac{\partial \omega_{0t}}{\partial x_t} = 1.196 \text{ and } \frac{\partial \omega_{1t}}{\partial x_t} = 0,$$

so

$$\frac{\partial diff_t}{\partial (i_t^{repo,3} - i_t^{repo,1})} = 1.196 \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right),$$

so change due to changes in medium-term lending risk is

$$1.196 \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right) \{ (i_t^{repo,3} - i_t^{repo,1}) - (i_{july\ 2007}^{repo,3} - i_{july\ 2007}^{repo,1}) \}.$$

Liquidity (liq_t^{FSR})

$$\frac{\partial \omega_{0t}}{\partial x_t} = -0.804 \text{ and } \frac{\partial \omega_{1t}}{\partial x_t} = 0.171,$$

so

$$\frac{\partial diff_t}{\partial liq_t^{FSR}} = -0.804 \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right) + 0.171 \frac{i_t^{base}}{\omega_{1t}},$$

so change due to changes in liquidity is

$$\left\{ -0.804 \left(\rho + \frac{1 - \rho}{\omega_{1t}} \right) + 0.171 \frac{i_t^{base}}{\omega_{1t}} \right\} (liq_t^{FSR} - liq_{july\ 2007}^{FSR}).$$

The two effects of liquidity in (25) go in opposite directions, as the reduction in liquidity will

- increase ω_{0t} and thus widen the differential or
- reduce ω_{1t} and thus reduce the differential.

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Central Bank Communication and Multiple Equilibria*

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In this paper, we construct a simple model for communication between a central bank and money-market traders. It is demonstrated that there are multiple equilibria. In one equilibrium, traders truthfully reveal their own information, and by learning this, the central bank can make better forecasts. Another equilibrium is a “dog-chasing-its-tail” equilibrium described by Blinder (1998). Traders mimic the central bank’s forecast, so the central bank simply observes its own forecast from traders. The latter equilibrium is socially worse, as inflation variability becomes larger. As policy implications, we find that too-high transparency of central banks is bad because it yields the “dog-chasing-its-tail” equilibrium, and central banks should conduct continuous monitoring or emphasize that their forecasts are conditional because doing so eliminates the “dog-chasing-its-tail” equilibrium. We also consider the possibility of the existence of an optimal degree of transparency.

JEL Codes: C72, D83, E52.

1. Introduction

Monetary policy transparency plays an essential role in conducting monetary policy. It not only serves legitimacy but also enhances the effectiveness of monetary policy. High transparency helps anchor

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inflation expectations (e.g., Faust and Svensson 2001) and influences the path of future interest rates (e.g., Woodford 2005). For these reasons, central banks around the globe have greatly increased monetary policy transparency in recent years.

However, too-high transparency is costly, and no central bank discloses all the information it has. This may partly reflect the fear of losing credibility (e.g., Mishkin 2007) or of political pressure in a federal state like Canada (Chant 2003). From a different perspective, and in a theoretical manner, Morris and Shin (2002) illustrate that a coordination problem makes too-high transparency harmful. They argue that public information disclosed by central banks is focal, which plays an excessive coordination role among private agents. Furthermore, as Blinder (1998) points out, too-high transparency prevents the efficient feedback from private agents to the central bank. Blinder (1998) calls this the “dog-chasing-its-tail” problem. He argues that if a central bank tries to “please the markets” too much, the markets stop functioning and the central bank only observes its own image in the markets. However, Blinder (1998) does not specify the meaning of “please the markets.”

In this paper, to formulate Blinder’s (1998) idea, we construct a simple model for communication between a central bank and private agents.¹ Communications are made in both directions: from the central bank to private agents and vice versa. As private agents, we consider an infinite number of money-market traders in the financial market. In order to finance their portfolios, hedge their positions, and square liquidity imbalances, the traders take position in relation to short-term interest rate expectations. This paper therefore assumes that the money-market traders aim to forecast interest rates set by a central bank as accurately as possible. The central bank aims to stabilize inflation by accurately forecasting an incoming shock. Following Grossman and Stiglitz (1976, 1980) and Hellwig and Veldkamp (2009), we assume that traders must pay an information acquisition cost to gather private information.

By constructing a simple model and solving it analytically, we demonstrate that there are multiple equilibria regarding the action of

¹Woodford (1994), Bernanke and Woodford (1997), Morris and Shin (2005), and Gosselin, Lotz, and Wyplosz (2008) also formulate Blinder’s idea, as is discussed below.

the central bank and the traders. In one equilibrium, traders truthfully reveal their own information, and by learning this, the central bank can make a better forecast. This equilibrium represents the good-coordination outcome as in Hellwig (2005). Transparency helps traders accurately forecast the interest rate set by the central bank, and owing to traders' truthful revelation, the central bank can also make a better forecast in setting the interest rate.

What is notable in this paper is the existence of another equilibrium. This equilibrium corresponds to the bad-coordination outcome as in Morris and Shin (2002) and more closely to the "dog-chasing-its-tail" equilibrium described by Blinder (1998). At the equilibrium, traders mimic the forecast disclosed by the central bank. The central bank therefore cannot improve its own forecast through communication and cannot optimally set the interest rate to stabilize inflation. The reason for the existence of the "dog-chasing-its-tail" equilibrium is as follows. Suppose all other traders mimic the central bank's forecast. Since there are an infinite number of traders in the market, even if one trader truthfully reveals its information, the central bank can obtain little information by communicating with the traders. This makes the central bank set an interest rate close to its original forecast. Therefore, by mimicking the central bank, the trader can make a good forecast of the central bank's action, yielding the "dog-chasing-its-tail" equilibrium.

Our analyses yield several policy implications. First, too-high transparency of central banks is bad because it yields the "dog-chasing-its-tail" equilibrium. Higher transparency helps traders forecast the central bank's action more accurately. However, if traders reveal their information truthfully, it makes the central bank's actions uncertain because truthfully revealed information by other traders is uncertain. Therefore, high transparency increases the incentive to mimic the central bank's forecast, which yields the "dog-chasing-its-tail" equilibrium. This implies that pleasing the markets may be good in the short run but detrimental in the long run because it prevents the markets from functioning, which induces instability in the economy. Second, the cost of transparency depends positively on the information acquisition cost and negatively on traders' forecast precision and the central bank's forecast precision after disclosure. In this respect, for central banks, it is important to evaluate traders' ability to judge whether the "dog-chasing-its-tail"

equilibrium is likely to arise. Furthermore, a central bank should conduct continuous monitoring or emphasize that its forecasts are conditional because this eliminates the “dog-chasing-its-tail” equilibrium by making a central bank’s disclosed information less valuable for the guidance of future monetary policy. Third, we consider the possibility that there exists an optimal degree of transparency. It is the highest level of transparency that eliminates the “dog-chasing-its-tail” equilibrium.

There is a sizable amount of literature on central bank transparency. Morris and Shin’s (2002) paper has attracted much attention and has been challenged, for example, by Angeletos and Pavan (2004), Hellwig (2005), Svensson (2006), Demertzis and Viegi (2008), and Hellwig and Veldkamp (2009). Svensson (2006) argues that the coordination induced by high transparency is good, provided that the public signal is more precise than the private signal. In reply to Svensson (2006), Morris, Shin, and Tong (2006) argue that if the public signal is correlated with the private signal, then quantitative evaluation supports Morris and Shin (2002). Morris and Shin (2005) and Gosselin, Lotz, and Wyplosz (2008) formulate Blinder’s idea in the context of central bank communication. With a simple theoretical model, they demonstrate the risk of informational inefficiency. According to their model, transparency prevents the central bank from obtaining potentially valuable market information and reduces the precision of its forecasts. However, in Morris and Shin (2005) and Gosselin, Lotz, and Wyplosz (2008), private agents’ disclosed information is still valuable, while in our “dog-chasing-its-tail” equilibrium, a central bank simply observes its own forecast, so private agents’ disclosed information has no value. Blinder’s idea is also partly formulated by Woodford (1994) and Bernanke and Woodford (1997), who point out the risk of indeterminacy arising from using not a structural model but private forecasts.

As the communication counterpart of a social planner, we focus on money-market traders in the financial market. All the above literature looks at private firms that engage in capital investment (Angeletos and Pavan 2004), price-setting (Hellwig 2005, Demertzis and Viegi 2008, and Gosselin, Lotz, and Wyplosz 2008), or zero-sum trades (Morris and Shin 2002, 2005) and regards these firms as consumers ultimately, so the sum of their utility equals social welfare, which a social planner aims to maximize. We admit that it is very

important to study central bank communication with consumers, particularly in terms of anchoring inflation expectations (Demertzis and Viegi 2008). However, regarding central bank communication, traders in the financial market should not be neglected. Traders are keen observers of central bank disclosure and also providers of information on financial market conditions. Further, these traders account for only a fraction of the total population, so the sum of the traders' utility does not represent social welfare. The "dog-chasing-its-tail" problem raised by Blinder (1998) seems to be related to traders in the financial market. In this paper, we consider more explicitly what can happen in central bank communication with money-market traders in the financial market. In doing so, we assume that a central bank's action is focal in that the interest rate set by the central bank directly affects a trader's loss. A coordination motive appears not explicitly but implicitly. In contrast, in Morris and Shin (2002, 2005), a coordination motive is introduced as an assumption. A common feature is, however, that a central bank's disclosed information is focal due to the coordination motive.

The existence of multiple equilibria we demonstrate in this paper differs from Morris and Shin (2002) and Hellwig (2005), where the equilibrium is unique. Morris and Shin (2002) and Hellwig (2005) disagree as to whether transparency yields better or worse coordination, but because of multiple equilibria, our paper integrates both the bad-coordination outcome and the good-coordination outcome. In our model, whether there are multiple equilibria (in particular, the "dog-chasing-its-tail" equilibrium) depends crucially on various parameters such as those of the information acquisition cost, forecast precision, and transparency. In particular, the assumption of the information acquisition cost is essential to yield multiple equilibria and differs from Morris and Shin (2002) and Hellwig (2005). Another modeling difference between our model and that of Morris and Shin (2002) and Hellwig (2005) is that we incorporate money-market traders in the financial market. In Angeletos and Pavan (2004) and Demertzis and Viegi (2008), the possibility of multiple equilibria is demonstrated, but their characteristics are very different from ours. In particular, the "dog-chasing-its-tail" equilibrium in Blinder (1998) does not appear.

The coexistence of the good and the bad equilibria provides us with implications on the optimal degree of transparency. In this

paper, we examine how the central bank's and traders' utility change depending on the transparency parameter and ask whether there is an optimal degree of transparency. Regarding the optimal degree of transparency, the above papers (with the exception of Angeletos and Pavan 2004) assume either full transparency or perfect secrecy. However, this assumption is unrealistic and may lead to an extreme conclusion that supports perfect secrecy. Angeletos and Pavan (2004) argue that there is a transparency parameter, but it is merely a central bank's forecast accuracy that is not controllable. In this paper, we explicitly introduce the measure of central bank transparency that is controllable, continuous, and representative of the clearness of central bank disclosure.² We then examine how the central bank's and traders' utility change depending on the transparency parameter and ask whether there is an optimal degree of transparency.

The structure of this paper is as follows. Section 2 introduces our model. Section 3 solves the model and shows that there arise multiple equilibria. Section 4 concludes the paper.

2. Model

We assume that inflation π is given by

$$\pi = i + u, \quad (1)$$

where i is an interest rate set by a central bank (hereafter CB) and u is a shock. Neither CB nor money-market traders (hereafter T) know the amount of the shock u in advance. By setting an appropriate interest rate, CB aims to minimize inflation variability,

$$L^{CB} = \pi^2, \quad (2)$$

which results in ultimately minimizing the loss of consumers.

²There have been relatively few papers on the optimal extent of transparency to date. Furthermore, their measures of transparency are not necessarily controllable and representative of the clearness of central bank disclosure. For example, Morris and Shin (2007) investigate the optimal number of signals to be disclosed. Walsh (2007) and Cornand and Heinemann (2008) investigate how many private agents should be informed by a central bank. Myatt and Wallace (2008) interpret the correlation between public signal and pure private signal as the measure of publicity.

As to CB's counterpart, we consider a sufficiently large number of money-market traders. The traders take positions in relation to short-term interest rate expectations, to finance their portfolios, to hedge their positions, and to square liquidity imbalances, so they aim to forecast i set by CB as accurately as possible. In other words, trader j 's loss function is written as

$$L_j^T = (i_j^T - i)^2, \quad (3)$$

where i_j^T is trader j 's forecast about i . Two remarks are worth making. First, CB's action is, by construction, focal because the interest rate set by CB directly affects T's loss. Instead, a coordination motive is not explicit in the above setup. However, as will become clear, traders have a coordination motive because CB's action is influenced by the average of traders' forecasts. Second, we explicitly model money-market traders who are different from consumers, so the sum of the traders' utility does not represent social welfare.

In the communication between CB and T, we consider the following four steps. At step 1, CB receives some information about the shock u —that is, u_1^{CB} —and discloses an imperfectly transparent forecast about the interest rate i —that is, i^{CB} . CB does not fully commit to the disclosed interest rate at step 1 because CB gathers new information by step 3, at which CB sets the actual interest rate i . In other words, CB's disclosed forecast is conditional. At step 2, T gathers information and, using its information set, tries to forecast the interest rate set by CB at step 3. As in Hellwig and Veldkamp (2009), we assume that trader j can gather its own information about the shock u_j^T by paying a fixed cost, $cE(u^2)$. Public information i^{CB} disclosed by CB is free. This assumption is motivated by Grossman and Stiglitz (1976, 1980), who argue that when there is a cost of information, a price system cannot perfectly convey information to the public. In our model, as we will soon see, this assumption yields an incentive for T to mimic CB's disclosed information.³ At step 3, CB observes the average of T's forecasts i^T and also receives new

³The appendix considers whether multiple equilibria arise without introducing fixed costs. It reveals that if there are a limited number of money-market traders but CB cannot observe each trader's strategy, then similar multiple equilibria can arise.

information about the shock u_2^{CB} . Using CB's information set, CB refines its forecast about the shock and sets the optimal interest rate i so as to stabilize inflation. Finally, at step 4, the inflation rate π is realized, by which the true amount of the shock u is known to everyone.

Forecast precisions are formulated in a standard way following Morris and Shin (2002, 2005). CB and T cannot observe an incoming u but observe u_1^{CB} , u_j^T , and u_2^{CB} given by

$$\begin{aligned} u_1^{CB} &= u + e_1^{CB}, & E(e_1^{CB^2})/E(u^2) &= 1/\alpha_1^{CB}, \\ u_j^T &= u + e_j^T + e^T, & E(e_j^{T^2})/E(u^2) &= 1/\alpha_j \text{ and} \\ & & E(e^{T^2})/E(u^2) &= 1/\alpha^T, \\ u_2^{CB} &= u + e_2^{CB}, & E(e_2^{CB^2})/E(u^2) &= 1/\alpha_2^{CB}, \end{aligned} \quad (4)$$

where α_1^{CB} and α_2^{CB} represent CB's forecast precisions at step 1 and at step 3, respectively. At step 2, trader j encounters idiosyncratic uncertainty e_j^T and aggregate uncertainty e^T , and the inverse of its amplitude is given by α_j^T and α^T . It is convenient to transform these parameters into

$$\begin{aligned} x_1^{CB} &= \frac{\alpha_1^{CB}}{1 + \alpha_1^{CB}}, \quad x_2^{CB} = \frac{\alpha_2^{CB}}{1 + \alpha_1^{CB} + \alpha_2^{CB}}, \quad x_j^T = \frac{\alpha_j^T}{1 + \alpha_j^T}, \\ x^T &= \frac{\alpha^T}{1 + \alpha^T}. \end{aligned} \quad (5)$$

Moreover, different from the earlier literature discussed in the introduction, we explicitly introduce a continuous transparency parameter τ so that it is controllable, continuous, and representative of the clearness of CB's disclosure. CB's disclosure is described as

$$i^{CB} = -E(u|u_1^{CB}) + v, \quad E(v^2) = E[(u - E(u|u_1^{CB}))^2]/\tau. \quad (6)$$

In our model, it is optimal for CB to set an interest rate so as to cancel a coming shock, so if CB is perfectly transparent, the error term v becomes zero. In this sense, the variance of v represents the obscurity of CB's disclosure. Note, however, that the variance of the error term v depends on CB's forecast precision before disclosure x_1^{CB} , which

cannot be easily controlled by CB, although early literature such as Angeletos and Pavan (2004) interprets CB's forecast precision as the measure of transparency. In order to draw policy implications using a controllable measure of transparency, we focus on a continuous transparency parameter τ as the measure of transparency.⁴ The variance of v is decreasing with the transparency parameter τ as well as CB's forecast precision. As τ becomes larger, the variance of v becomes smaller, which suggests higher transparency.

3. Equilibrium

This section solves the above model and shows that multiple equilibria can exist. The solution of this model is derived by answering the following two questions. The first question is how T discloses its own forecast about the interest rate at step 2. More precisely, we ask if T tries to make the best guess about the shock u and truthfully reveals its own forecast, or if T mimics CB's forecast made at step 1. The second question is how CB evaluates the shock u and sets the optimal interest rate i . More precisely, we ask how CB refines its forecast by combining its own information and T's forecast, and if CB can learn anything valuable from T's forecasts.

Since the model is linear quadratic, the actions of CB and T are written in the following linear form:

$$\begin{aligned} \text{Step 2} \quad i_j^T &= E(i|I_j^T) = a_1^j i^{CB} + a_2^j u_j^T, \\ \text{Step 3} \quad E(u|I^{CB}) &= -i = b_1 i^T + b_2 u_2^{CB} + b_3 u_1^{CB} + b_4 v. \end{aligned} \quad (7)$$

Here we use the fact that trader j 's information set at step 2 is $I_j^T = \{i^{CB}, u_j^T\}$ and that CB's information set at step 3 is $I^{CB} = \{i^T, u_2^{CB}, u_1^{CB}, v\}$. The coefficients a and b are obtained from

$$\begin{aligned} (a_1^j, a_2^j) &= \arg \min E[i - E(i|I_j^T)]^2 \quad \text{given } (a_1^k, a_2^k, b_1, b_2, b_3, b_4) \\ (b_1, b_2, b_3, b_4) &= \arg \min E[u - E(u|I^{CB})]^2 \quad \text{given } (a_1, a_2), \end{aligned} \quad (8)$$

where $k \neq j$.

⁴As we will show later in figure 1, CB's forecast precision before disclosure x_1^{CB} appears to be irrelevant to the condition of multiple equilibria. Therefore, in fact, our results appear to hold, even though we focus on v .

The above problem can be solved analytically, which leads to our first proposition. Hereafter, we limit our attention to the equilibrium where all traders have the same strategy.

PROPOSITION 1. *There exist multiple equilibria, “M” and “R,” when fixed costs c satisfies*

$$c > F(x_1^{CB}, x_2^{CB}, x^T, x_i^T, \tau). \quad (9)$$

Otherwise, only equilibrium “R” exists. In equilibrium “M” (or “dog-chasing-its-tail”), T mimics CB ’s forecast. In equilibrium “R,” T truthfully reveals its own forecast.

Proof. We can show that there can exist two equilibria, where the coefficients a and b are respectively written as

$$\begin{aligned} a_1 &= \left\{ 1 - x_1^{CB} + \frac{1 - x_1^{CB}}{\tau x_1^{CB}} + \left(\frac{1}{\tau} \frac{1 - x_1^{CB}}{x_1^{CB}} + 1 \right) \right. \\ &\quad \times \left. \left(\frac{1 - x_i^T}{x_i^T} + \frac{1 - x^T}{x^T} \right) \right\}^{-1} \left(\frac{1 - x_i^T}{x_i^T} + \frac{1 - x^T}{x^T} \right), \\ a_2 &= -1 + a_1 \left(\frac{1}{\tau} \frac{1 - x_1^{CB}}{x_1^{CB}} + 1 \right), \\ b_1 &= \frac{1}{a_2} \frac{(1 - x_1^{CB})(1 - x_2^{CB})}{(1 - x_1^{CB})(1 - x_2^{CB}) + \frac{1 - x^T}{x^T}}, \quad b_2 = x_2^{CB}(1 - a_2 b_1), \\ b_3 &= x_1^{CB}(1 - a_2 b_1 - x_2^{CB} + x_2^{CB} a_2 b_1 + a_1 b_1), \quad b_4 = -a_1 b_1, \end{aligned} \quad (10)$$

or

$$\begin{aligned} a_1 &= \left(\frac{1 - x_1^{CB}}{\tau x_1^{CB}} + 1 \right)^{-1}, \quad a_2 = 0, \\ b_1 &= \text{arbitrary}, \quad b_2 = x_2^{CB}, \quad b_3 = x_1^{CB}(1 - x_2^{CB} + a_1 b_1), \quad b_4 = -a_1 b_1. \end{aligned} \quad (11)$$

In equation (11), a_2 is zero, which suggests that T does not reveal its own information and mimics CB ’s forecast. CB therefore sees its own image through communication with T . This is the “M” or “dog-chasing-its-tail” equilibrium. The parameter b_1 is arbitrary. This is

because, in “M,” CB’s reaction function in equation (7) becomes redundant. Using equation (11), we can easily show that equation (7) is transformed into

$$\begin{aligned} E(u|I^{CB}) &= b_1 i^T + b_2 u_2^{CB} + b_3 u_1^{CB} + b_4 v \\ &= x_2^{CB} u_2^{CB} + x_1^{CB} (1 - x_2^{CB}) u_1^{CB}. \end{aligned} \quad (12)$$

In other words, in “M,” CB has to rely on its own information. On the other hand, in equation (10), a_2 is non-zero, which suggests that T truthfully reveals its own forecast.

We next examine the condition that no trader has an incentive to deviate from “M.” If we neglect fixed costs, and when all the other traders choose “M,” a trader j who deviates from “M” can minimize its expected loss by choosing

$$\begin{aligned} a_2^j &= \left(\frac{1}{x_i^T} + \frac{1 - x_1^{CB}}{\tau x_1^{CB} x_i^T} - x_1^{CB} + \frac{1 - x^T}{x^T} \left(1 + \frac{1 - x_1^{CB}}{\tau x_1^{CB}} \right) \right)^{-1} \\ &\quad \times \left(x_2^{CB} - x_1^{CB} x_2^{CB} + \frac{1 - x_1^{CB}}{\tau x_1^{CB}} (x_1^{CB} + x_2^{CB} - x_1^{CB} x_2^{CB}) \right), \end{aligned} \quad (13)$$

$$a_1^j = \frac{1}{x_1^{CB}} \left(a_2^j \left(\frac{1}{x_i^T} + \frac{1 - x^T}{x^T} \right) + x_1^{CB} + x_2^{CB} - x_1^{CB} x_2^{CB} \right).$$

Comparing the expected loss including fixed costs in this case with the expected loss in “M” suggests that the equilibrium “M” exists when

$$\begin{aligned} c &> F(x_1^{CB}, x_2^{CB}, x^T, x_i^T, \tau) \\ &\equiv -x_1^{CB} \left\{ \left(1 + \frac{1 - x_1^{CB}}{\tau x_1^{CB}} \right)^{-1} - 1 + x_2^{CB} \right\} \\ &\quad - 2x_1^{CB} x_2^{CB} \left\{ \left(1 + \frac{1 - x_1^{CB}}{\tau x_1^{CB}} \right)^{-1} - 1 + x_2^{CB} \right\} - x_2^{CB^2} \\ &\quad - \frac{1 - x_1^{CB}}{\tau} \left(1 + \frac{1 - x_1^{CB}}{\tau x_1^{CB}} \right)^{-2} \end{aligned}$$

$$\begin{aligned}
& + \frac{1 - x_1^{CB}}{x_1^{CB}} x_2^{CB^2} + \frac{1 - x_1^{CB}}{\tau} \left(\frac{x_1^{CB} + x_2^{CB} - x_1^{CB} x_2^{CB}}{x_1^{CB}} \right)^2 \\
& - \frac{(1 - x_1^{CB})^2}{x_1^{CB}} \left(\frac{1}{x_i^T} + \frac{1 - x^T}{x^T} \right) \left(x_2^{CB} + \frac{x_1^{CB} + x_2^{CB} - x_1^{CB} x_2^{CB}}{\tau x_1^{CB}} \right)^2 \\
& \times \left\{ \frac{1}{x_i^T} - x_1^{CB} + \frac{1 - x_1^{CB}}{x_1^{CB}} \frac{1}{x_i^T} + \frac{1 - x^T}{x^T} \left(1 + \frac{1 - x_1^{CB}}{x_1^{CB}} \right) \right\}^{-1}.
\end{aligned} \tag{14}$$

In equilibrium “R,” T truthfully reveals its own information, and by learning this, CB can make a better forecast. This equilibrium represents the good-coordination outcome as in Hellwig (2005). Transparency helps traders accurately forecast the interest rate set by the central bank, and due to traders’ truthful revelation, the central bank can also make a better forecast in setting the interest rate.

On the other hand, the equilibrium “M” corresponds to the bad-coordination outcome as in Morris and Shin (2002) and more closely to the “dog-chasing-its-tail” equilibrium in Blinder (1998). In this equilibrium, T mimics CB’s forecast. CB therefore cannot improve its own forecast through communication, preventing CB from optimally setting the interest rate to stabilize inflation. The reason for the existence of the “dog-chasing-its-tail” equilibrium is as follows. Suppose all other traders mimic CB’s forecast. Since there are an infinite number of traders in the market, even if one T truthfully reveals its information, CB can obtain little information by communicating with T. This makes CB set an interest rate close to its original forecast. Therefore, by mimicking CB, T can make a good forecast of CB’s action, yielding the “dog-chasing-its-tail” equilibrium.

The equilibrium “M” in our model is much closer to the “dog-chasing-its-tail” problem raised by Blinder (1998) than that of the earlier literature. For example, the problem pointed out by Woodford (1994) and Bernanke and Woodford (1997) is the risk of indeterminacy arising from using not a structural model but private forecasts. Their papers do not necessarily explain Blinder’s argument that if a central bank tries to “please the markets” too much, the markets stop functioning and the central bank only observes its own image

from the markets. In Morris and Shin (2005), transparency makes private agents' disclosed information less valuable to a central bank, but it is still valuable. On the other hand, our paper demonstrates that in the "dog-chasing-its-tail" equilibrium, CB simply observes its own forecast, so private agents' disclosed information has no value.

Morris and Shin (2002) and Hellwig (2005) disagree as to whether transparency yields better or worse coordination, and their disagreement is not easily resolved because the equilibrium is unique and their models are different. Our model helps resolve their disagreement because it integrates both the bad-coordination outcome and the good-coordination outcome in a unified framework. Compared with Morris and Shin (2002) and Hellwig (2005), there are two key features in our model that lead to multiple equilibria. First, we introduce the information acquisition cost. The appendix suggests that we need an alternative assumption to yield multiple equilibria without introducing the information acquisition cost. Second, as the communication counterpart of a social planner, we focus on money-market traders in the financial market, while Morris and Shin (2002) and Hellwig (2005) focus on private firms that engage in zero-sum trades and price setting, respectively.

As we noted in the previous section, CB's action is, by construction, focal because the interest rate set by CB directly affects T's loss. A coordination motive appears not explicitly but implicitly because CB's action is influenced by the average of traders' forecasts. This is understood by rearranging equation (3) using equation (7):

$$L_j^T = (i_j^T - i)^2 = \left\{ i_j^T - (-b_1) \lim_{N \rightarrow \infty} \frac{1}{N} \sum_k i_k^T \right\}^2 + \dots \quad (15)$$

Omitted terms (...) in the right-hand side of equation (15) are the function of the shock u .⁵ From this, it is clear that trader j who

⁵We can show that even though T's loss depends on a fundamental described by inflation variability as well as T's forecast error, the same multiple equilibria can arise. This implies that this result is robust even though traders are not of the money market but are of long-term bond or equity markets. However, the following discussion of traders' welfare will be modified.

aims to forecast CB's action attempts to evaluate other traders' forecasts. This induces a coordination motive among agents. In this sense, T's loss function is similar to that in Morris and Shin (2002, 2005). However, a difference is that in Morris and Shin (2002, 2005), a coordination motive is introduced as an assumption. A common feature is, however, that CB's disclosed information is focal due to the coordination motive.

Which equilibrium is better for CB? The second proposition answers this question.

PROPOSITION 2. *For CB, "R" is better than "M."*

Proof. CB's loss, $L^{CB} = \pi^2$, is given by

$$\begin{aligned} \text{(M)} \quad L^{CB(M)} &= (1 - x_1^{CB})(1 - x_2^{CB}), \\ \text{(R)} \quad L^{CB(R)} &= \left(\frac{1}{(1 - x_1^{CB})(1 - x_2^{CB})} + \frac{x^T}{1 - x^T} \right)^{-1} < L^{CB(M)}. \end{aligned} \tag{16}$$

As simple New Keynesian economics states, if we consider that minimizing inflation variability leads to maximizing social welfare, then this proposition suggests that "R" is socially more desirable than "M." This proposition is intuitive. In "R," CB can learn T's information about the shock u , so CB can improve its forecast accuracy for u and set a better interest rate to stabilize inflation than in "M," where CB has to rely only on its own information. T's information, even if its precision is low, is always useful for CB if CB extracts the signal appropriately. It is also worth pointing out that equation (16) does not depend on the transparency parameter τ . The degree of transparency does not make any difference to CB's loss.

To find a way to eliminate the undesirable equilibrium "M," we next examine the condition of multiple equilibria in more detail. A condition on the fixed costs to acquire private information c is obvious. As the fixed costs become lower, profits from saving the fixed costs become smaller. Therefore, the equilibrium "M" is more likely

to be eliminated. Regarding transparency and forecast precision parameters, the following lemma provides a way to eliminate “M.”

LEMMA 3.

$$\begin{aligned} dF(x_1^{CB}, x_2^{CB}, x^T, x_i^T, \tau)/dx_2^{CB} &\geq 0, dF(x_1^{CB}, x_2^{CB}, x^T, x_i^T, \tau)/dx^T > 0, \\ dF(x_1^{CB}, x_2^{CB}, x^T, x_i^T, \tau)/dx_i^T &> 0, dF(x_1^{CB}, x_2^{CB}, x^T, x_i^T, \tau)/d\tau \leq 0. \end{aligned}$$

Many valuable policy implications can be obtained from this lemma. Firstly, too-high transparency is bad because transparency that is too high increases the parameter space over which the socially bad equilibrium “M” is possible. The fact that $dF/d\tau$ is non-positive suggests that as CB’s transparency becomes higher, “M” becomes more likely to arise. The reason why $dF/d\tau$ is non-positive is simple. As transparency increases, T can infer more accurately CB’s valuation. However, if traders reveal their information truthfully, it makes CB’s action uncertain because truthfully revealed information by other traders is uncertain. Therefore, high transparency increases an incentive to mimic CB’s forecast, which yields the “dog-chasing-its-tail” equilibrium. The cost of transparency depends on the fixed costs to acquire private information c . As c becomes lower, the bad equilibrium “M” is more likely to be eliminated, so high transparency becomes less detrimental.

Secondly, CB should conduct continuous monitoring or emphasize that CB’s forecasts are conditional. The fact that dF/dx_2^{CB} is non-negative suggests that as CB’s forecast after disclosure becomes more accurate, “M” becomes less likely to arise. The reason why dF/dx_2^{CB} is non-negative is as follows. If x_2^{CB} is high, CB weighs its own information after disclosure more than that before disclosure, so CB’s disclosed information does not provide T with a good sign for future interest rate decisions. Therefore, mimicking CB becomes less valuable to T than the case of truthfully revealing. By conducting continuous monitoring or emphasizing that CB’s forecasts are conditional, CB can increase its forecast precision after the disclosure and make itself less committed to future policy, which can reduce the incentive for T to choose “M.”

Thirdly, we find that both dF/dx^T and dF/dx_i^T are positive. This suggests that as T’s forecast precision is lower, “M” becomes more likely to arise. This is because if T’s forecast precision is low,

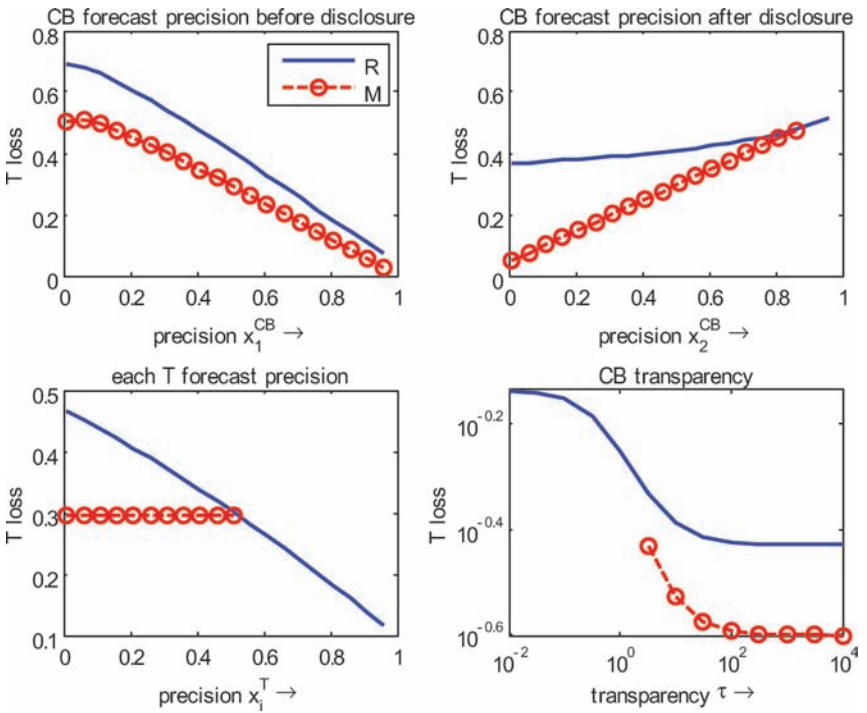
then T's payoff by truthful revelation becomes small, which lowers the incentive for T to deviate from "M." Since policymakers cannot adjust the parameters x^T and x_1^T , this finding does not necessarily give them a direct policy implication. However, it gives them a warning of "M" when T's ability is low. In this respect, it is important for CB to evaluate T's ability.

To check these results visually, we implement numerical evaluations by changing some parameters. As a baseline, we assume the values of $x_1^{CB} = 0.5$, $x_2^{CB} = 0.5$, $x_1^T = 0.2$, and $\tau = 10$. The forecast precision of each trader is worse than that of CB. The aggregate forecast accuracy of traders, x^T , is chosen to be 0.8 so that traders' aggregate information has the same accuracy as CB's information. This corresponds to the conservative benchmark in Svensson (2006). Fixed costs, c , are chosen to be 0.05 so that our numerical calculation clearly demonstrates this model's characteristics. By changing one parameter while keeping the other parameters fixed, we examine how T's loss changes depending on forecast precision and transparency.

Figure 1 shows the results. The solid line shows T's loss in "R," while the dotted line with circles shows T's loss in "M." We find that the dotted line disappears in some regions. This suggests that, as proposition 1 and lemma 3 state, the bad equilibrium "M" is eliminated when x_2^{CB} is high, x_1^T (so as x^T) is high, or τ is low.

Finally, we consider the possibility that there exists an optimal degree of transparency. A candidate for this is the highest τ^* that leads to the unique equilibrium. In our simulation, the bottom-right panel of figure 1 suggests that τ^* is approximately one. If the transparency parameter becomes lower than τ^* , the undesirable equilibrium "M" is eliminated but T suffers more losses, as the bottom-right panel of figure 1 demonstrates. In this sense, as in Hellwig (2005), it is good to be transparent—for traders as well as a central bank—so as to promote good coordination. On the other hand, if the transparency parameter exceeds τ^* , it produces the socially bad equilibrium "M" as in Morris and Shin (2002, 2005). Therefore, too much transparency is risky. Of course, higher transparency than τ^* does not necessarily lead to the selection of "M" because, as long as all other T believes that the equilibrium "R" is achieved, the equilibrium "R" becomes stable. However, T prefers "M" to "R," as the bottom-right panel suggests, so T may be inclined to coordinate to achieve "M."

Figure 1. Trader’s Loss



4. Concluding Remarks

This paper constructed a very simple model for communication between a central bank and money-market traders in the financial market. We assume that a central bank aims to stabilize inflation, while traders aim to forecast the interest rate set by the central bank, and that traders have to pay a fixed cost when gathering information. With these assumptions, we demonstrate that there may be multiple equilibria. In one equilibrium, “R,” traders truthfully reveal their own information, and by learning this, the central bank can make a better forecast. Another equilibrium, “M,” corresponds to a “dog-chasing-its-tail” equilibrium described by Blinder (1998): traders mimic the central bank’s forecast, so the central bank simply observes its own forecast from traders. The latter equilibrium is worse for the central bank and ultimately for consumers because inflation variability increases.

Policy implications are, first, that too-high transparency of central banks is bad because it yields the socially bad equilibrium “M.” Pleasing the markets may be good in the short run but detrimental in the long run because it prevents the markets from functioning, which induces instability in the economy. Second, the cost of transparency depends positively on the fixed costs to acquire private information c and negatively on traders’ forecast precision x_i^T and x_j^T and a central bank’s forecast precision after disclosure x_2^{CB} . In this respect, for central banks, it is important to evaluate traders’ ability to judge whether the equilibrium “M” is likely to arise. Furthermore, central banks should conduct continuous monitoring or emphasize that their forecasts are conditional because it increases x_2^{CB} and eliminates “M.” Third, we consider the possibility that there exists an optimal degree of transparency. The highest τ^* that eliminates “M” may be interpreted as the optimal degree of transparency.

There are remaining tasks, however. In particular, it is important to extend this model into a dynamic model with consideration of the inflation target and inflation expectations. In this model, we assume that traders in the financial market do not influence inflation except for their expectation formation, but it is natural to think that traders’ expected interest rates or traders’ reactions to a central bank’s action directly influence inflation by affecting inflation expectations and long-term interest rates. Our paper neglects this aspect, although many people emphasize its importance (e.g., Faust and Svensson 2001, Woodford 2005, Aoki and Kimura 2008, and Demertzis and Viegi 2008). This may well reestablish the importance of greater transparency.

Appendix. Model without Fixed Costs

In this paper, we assume that T has to pay a fixed cost when gathering information. This is clearly one of the key assumptions used to demonstrate the existence of multiple equilibria—in particular, “M.” This appendix considers another way of obtaining similar multiple equilibria without introducing fixed costs.

We make four modifications in our model. Firstly, we do not assume a fixed cost in gathering information. Secondly, we consider a finite number of money-market traders given by N . Thirdly and

most importantly, we assume that CB cannot observe the deviation of each strategy of T (a_1^j, a_2^j) from a certain aggregate equilibrium strategy (a_1, a_2). At step 3, CB observes i^T , but CB considers that its change comes not from the deviation of each strategy of T (a_1^j, a_2^j) but from a change in the shock traders receive, u_j^T . Therefore, each trader's strategy does not influence CB's strategy, and the trader optimizes its strategy with CB's strategy fixed. Fourthly, for simplicity we neglect aggregate uncertainty e^T . In other respects, our model setup is the same as before.

With this setup, we can obtain the following two propositions.

PROPOSITION A1. *There always exist multiple equilibria. In equilibrium "M" (or "dog-chasing-its-tail"), T mimics CB's forecast. In equilibrium "R," T truthfully reveals its own forecast.*

Proof. We can show that there are two equilibria, where the coefficients a and b are written as

$$\begin{aligned} a_1 &= - \left\{ x_1^{CB} - \frac{1}{x^T} \left(\frac{1}{\tau} \frac{1 - x_1^{CB}}{x_1^{CB}} + 1 \right) \right\}^{-1} \left(\frac{1 - x_j^T}{x_j^T} \right), \\ a_2 &= -1 + a_1 \left(\frac{1}{\tau} \frac{1 - x_1^{CB}}{x_1^{CB}} + 1 \right), \\ b_1 &= \frac{1}{a_2} \frac{(1 - x_1^{CB})(1 - x_2^{CB})}{(1 - x_1^{CB})(1 - x_2^{CB}) + \frac{1 - x_j^T}{Nx_j^T}}, \quad b_2 = x_2^{CB}(1 - a_2 b_1), \\ b_3 &= x_1^{CB}(1 - a_2 b_1 - x_2^{CB} + x_2^{CB} a_2 b_1 + a_1 b_1), \quad b_4 = -a_1 b_1, \end{aligned} \quad (17)$$

or

$$\begin{aligned} a_1 &= -b_4/b_1, \quad a_2 = 0, \\ b_1 &= -N, \quad b_2 = x_2^{CB}, \quad b_3 = x_1^{CB}(1 - x_2^{CB} + a_1 b_1), \quad b_4 = \text{arbitrary}. \end{aligned} \quad (18)$$

In equation (18), $a_2 = 0$, which suggests that T mimics CB's forecast. A parameter b_4 is arbitrary, but we can easily show that equation (7) is transformed into

$$\begin{aligned}
 E(u|I^{CB}) &= b_1 i^T + b_2 u_2^{CB} + b_3 u_1^{CB} + b_4 v \\
 &= x_2^{CB} u_2^{CB} + x_1^{CB} (1 - x_2^{CB}) u_1^{CB}.
 \end{aligned} \tag{19}$$

On the other hand, in equation (17), a_2 is non-zero, which suggests that T truthfully reveals its own forecast.

PROPOSITION A2. *For CB, “R” is better than “M.”*

Proof. CB’s loss, $L^{CB} = \pi^2$, is given by

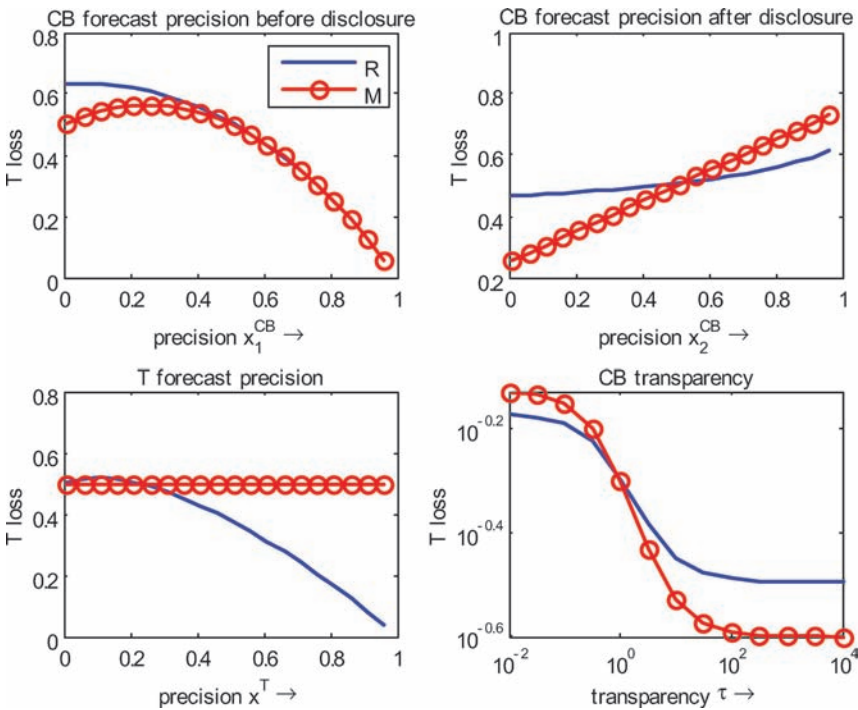
$$\begin{aligned}
 \text{(M)} \quad L^{CB(M)} &= (1 - x_1^{CB})(1 - x_2^{CB}), \\
 \text{(R)} \quad L^{CB(R)} &= \left(\frac{1}{(1 - x_1^{CB})(1 - x_2^{CB})} + \frac{N x_j^T}{1 - x_j^T} \right)^{-1} < L^{CB(M)}.
 \end{aligned} \tag{20}$$

These multiple equilibria are highly similar to those in our main model. A notable difference is that in this modified model, multiple equilibria always exist irrespective of the forecast precision parameters α_1^{CB} , α^T , and α_2^{CB} and the transparency parameter τ .

As this model demonstrates, fixed costs are not necessarily needed to yield multiple equilibria—in particular, “M.” For T, deviating from “M” is not worthwhile because CB can respond to the deviation of a single trader’s strategy by choosing $b_1 = -N$, which lowers T’s profit by deviating. However, as we noted at the beginning of this appendix, this result crucially depends on the assumption that CB observes i^T but CB considers that its change comes not from the deviation of each T’s strategy (a_1^j, a_2^j) but from a change in the shock traders receive, u_j^T .

Finally, in order to evaluate T’s loss, we assume the values of $x_1^{CB} = 0.5$, $x_2^{CB} = 0.5$, $x_j^T = 0.2$, $\tau = 10$, and $N = 16$. The number of traders, N , is set so that T’s pooled forecast precision is the same as CB’s forecast precision. Figure 2 demonstrates the results. Although multiple equilibria always exist, these panels closely resemble figure 1. Comparing the losses in “M” and “R” implies that high transparency increases the attractiveness of “M” over “R” for T. On the other hand, CB’s high forecast precision after disclosure decreases the attractiveness of “M” for T.

Figure 2. Trader's Loss



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Estimation of Monetary Policy Preferences in a Forward-Looking Model: A Bayesian Approach*

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In this paper we adopt a Bayesian approach toward the estimation of monetary policy preference parameters in a general equilibrium framework for the euro area. We assume that monetary policy authorities optimize an intertemporal quadratic loss function under commitment and study two alternative specifications of the loss function. The first specification includes inflation, output gap, and the interest rate differential as targets. The second loss function includes an additional wage-inflation target. The weights assigned to target variables—i.e., monetary policy preferences—are estimated jointly with the structural model parameters. The results imply that inflation variability remains the main concern of optimal monetary policy. Interest rate smoothing and the output gap appear to be, to a lesser extent, important target variables as well. Due to the time-inconsistency problem under commitment, we propose to initialize the estimates by a pre-sample period of forty quarters. This allows us to approach, empirically, the timeless-perspective framework.

JEL Codes: E42, E52, E58, E61.

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1. Introduction

Correct knowledge of the variables that are of main concern for monetary policy is an important asset, since knowing the alternative monetary policy targets will have an effect on the formation of private-sector expectations. Given the role these expectations play in stabilizing the economy, it is desirable for monetary policymakers to provide the private agents with sufficient information concerning the relative importance of each target variable. The value attached to a particular target—for example, the inflation target—can be described by the relative weight assigned to this target in the loss function that the central bank aims to minimize over the infinite horizon. The relative weights therefore generally reflect the preferences of monetary policymakers with respect to the corresponding target variables.

In order to infer the monetary policy preferences, one could analyze empirical monetary policy reaction functions and study the behavior of monetary policymakers. This kind of approach, however, has often been criticized (Svensson 2002b, 2003, and Dennis 2000, 2002, 2004a, 2006, and 2007). While an estimated reaction function gives a good description of monetary policy behavior, the intertemporal loss function is a more appropriate measure of (changes in) monetary policy objectives. The variables entering the reaction function mainly provide monetary policy with information needed to achieve the policy objectives. These variables are therefore not necessarily equal to the target variables in the loss function and cannot be attributed directly to the monetary policy objectives. A more theoretical justification for assuming a representative monetary policymaker that systematically optimizes an intertemporal loss function is that this approach will bring monetary policy behavior in line with the behavior of private agents (Svensson 1999 and Woodford 2003). In this paper, we adopt a general equilibrium framework with rational and optimizing agents, where all structural equations result from optimal decisions made by private agents as well as monetary policymakers.

An extensive number of studies in the literature have recently focused attention on the estimation or calibration of preferences of optimizing monetary policy authorities, which is analogous to estimating the weights assigned to the target variables in the loss

function of the central bank. Many of these estimation exercises consider the case of discretionary monetary policy, as in Dennis (2000, 2004a), Söderström, Söderlind, and Vredin (2005), and Castelnuovo (2006) for the U.S. economy and Lippi and Neri (2007) for the euro-area economy.¹ The case of full commitment as in Söderlind (1999) and Ilbas (forthcoming) or commitment to a simple rule of the kind adopted by Salemi (2006) and Givens and Salemi (2008) for the U.S. economy has, to our knowledge, not been applied to the euro-area economy. The aim of this paper is to study the case of monetary policy that systematically minimizes an intertemporal quadratic loss function under full commitment in a forward-looking model for the euro area. A commitment strategy, if credible, enables the central bank to control the expectations of private agents and provides it with an additional stabilization tool. We consider the Smets and Wouters (henceforth SW) (2003) model for the euro area as the benchmark model, where we replace the estimated Taylor rule with optimal monetary policy under commitment. This enables us to estimate the preference parameters of the monetary policy objective function jointly with the structural parameters of the model economy. We further show that the time-inconsistency problem under commitment can be dealt with empirically by considering an initialization period that is long enough to reduce the effect of the initial values on the estimation results. This enables us to implement the timeless-perspective framework of Woodford (2003) in empirical exercises.

This paper is organized as follows. In the next section, we outline the theoretical framework. We start from the SW (2003) model and describe the assumed structural behavior of the private agents in the economy, followed by the introduction of optimal monetary policy, which leads to a set of Euler equations that can be estimated accordingly. In introducing optimal monetary policy, we consider two types of loss functions. The first loss function includes inflation, the model-consistent output gap, and the interest rate differential as target variables, whereas the second loss function considers an additional wage-inflation target. The third section explains the data

¹Ozla (2003) and Dennis (2006) perform a similar exercise for the United States in the context of the Rudebusch and Svensson (1999) model, which is a purely backward-looking model and hence avoids time-inconsistency issues.

set used and the methodology adopted in the Bayesian estimation procedure, followed by a discussion of the results. We compare the results for the structural parameters obtained from the modified model with optimal monetary policy with the results of the original SW (2003) model. In section 4, we compare the alternative models based on their marginal-likelihood values and discuss the impulse responses obtained under the best-performing model that is characterized by optimal policy with respect to the benchmark impulse responses of SW (2003). Finally, section 5 concludes, followed by the appendices.

2. Theoretical Framework

The structural behavior of the euro-area economy is described by the model developed in Smets and Wouters (2003). However, instead of capturing the behavior of the monetary policy authorities by an empirical Taylor rule as is done in the original setup of the SW (2003) model, we will assume that monetary policy is performed optimally under commitment. In the following, we present a brief summary of the linearized SW (2003) model and introduce the optimizing monetary policy authorities.

2.1 The Smets and Wouters Model for the Euro Area

The dynamic stochastic general equilibrium (DSGE) framework presented in SW (2003) for the euro area consists of a household sector that supplies a differentiated type of labor, leading to nominal rigidities in the labor markets. The goods markets are characterized by intermediate and final goods producers. The former type of agents are monopolistically competitive and produce a differentiated type of intermediate goods, leading to nominal rigidities in the goods markets. The latter type of agents operate in a perfectly competitive market and produce one final good.

The model also features real rigidities like habit formation, costs of adjustment in capital accumulation, and variable capital utilization. There are many similarities with the model presented in Christiano, Eichenbaum, and Evans (2005). However, the SW (2003) model includes an additional number of structural shocks, includes

partial indexation to past inflation in the labor and the goods markets, and is estimated using (Bayesian) estimation methods. The linearized rational-expectations equations that result from private-sector optimizing behavior are summarized next, where the same notation as in SW (2003) is adopted and where all variables are expressed as log-deviations from their steady-state levels denoted by $\hat{\cdot}$, i.e. $\hat{x} = \log(\frac{x}{x^*})$.²

Each household supplies a differentiated type of labor and maximizes an intertemporal utility function of the following type:

$$E_0 \sum_{t=0}^{\infty} \beta^t U_t^\tau, \quad (1)$$

where E is the rational-expectations operator and β is the discount factor. The utility function U_t^τ is separable in consumption, C_t^τ , and labor supply, l_t^τ , defined as follows (see linearized equations below for a description of the parameters):

$$U_t^\tau = \varepsilon_t^b \left(\frac{1}{1 - \sigma_c} (C_t^\tau - H_t)^{1 - \sigma_c} - \frac{\varepsilon_t^L}{1 + \sigma_L} (l_t^\tau)^{1 + \sigma_L} \right). \quad (2)$$

This setup leads to the following consumption Euler equation, which includes an external habit variable that leads to a backward-looking component to capture habit persistence:

$$\begin{aligned} \hat{C}_t = & \frac{h}{1 + h} \hat{C}_{t-1} + \frac{1}{1 + h} E_t \hat{C}_{t+1} - \frac{1 - h}{(1 + h)\sigma_c} (\hat{R}_t - E_t \hat{\pi}_{t+1}) \\ & + \frac{1 - h}{(1 + h)\sigma_c} (\hat{\varepsilon}_t^b - E_t \hat{\varepsilon}_{t+1}^b), \end{aligned} \quad (3)$$

where h is the degree of habit persistence in the external habit variable $H_t = hC_{t-1}$. The symbol σ_c represents the inverse of the intertemporal elasticity of substitution between consumption at different dates. $\hat{R}_t$ is the nominal interest rate, $\hat{\pi}_t$ is the inflation rate, and $\hat{\varepsilon}_t^b$ is an AR(1) preference shock to the discount rate with an

²For a detailed description of the optimizing behavior of the agents that led to the linearized version of the model, we refer to the original SW (2003) paper.

i.i.d. normal error term: $\varepsilon_t^b = \rho_b \varepsilon_{t-1}^b + \eta_t^b$. Nominal wages are set by households according to a Calvo (1983) type of scheme. Households that cannot reoptimize will adjust their nominal wages partially to past inflation with a degree $0 \leq \gamma_w \leq 1$, leading to the following real-wage (w_t) equation:

$$\begin{aligned} \hat{w}_t = & \frac{\beta}{1+\beta} E_t \hat{w}_{t+1} + \frac{1}{1+\beta} \hat{w}_{t-1} + \frac{\beta}{1+\beta} E_t \hat{\pi}_{t+1} - \frac{1+\beta\gamma_w}{1+\beta} \hat{\pi}_t \\ & + \frac{\gamma_w}{1+\beta} \hat{\pi}_{t-1} - \frac{1}{1+\beta} \frac{(1-\beta\xi_w)(1-\xi_w)}{\left(\frac{\lambda_w+(1+\lambda_w)\sigma_L}{\lambda_w}\right)} \xi_w \\ & \times \left[\hat{w}_t - \sigma_L \hat{L}_t - \frac{\sigma_c}{1-h} (\hat{C}_t - h\hat{C}_{t-1}) - \hat{\varepsilon}_t^L - \eta_t^w \right], \end{aligned} \quad (4)$$

where ξ_w is the probability that the wage will not be reoptimized in the next period. λ_w is a constant in $\lambda_{w,t} = \lambda_w + \eta_t^w$, with $\lambda_{w,t}$ a shock to the wage markup assumed to be i.i.d. normal around the constant term. σ_L is the inverse of the elasticity of work with respect to the real wage, $\hat{L}_t$ is the aggregate labor, and $\hat{\varepsilon}_t^L$ is a shock to labor supply assumed to follow an AR(1) process with an i.i.d. normal error term: $\varepsilon_t^L = \rho_L \varepsilon_{t-1}^L + \eta_t^L$. Households also make the investment and capital accumulation decision; they are owners of the capital stock and rent it out to the producers of intermediate goods at the rental rate of capital r_t^k . The supply of rental services can be adjusted by increasing investments I_t or by increasing the utilization rate z_t of installed capital. The linearized investment equation, which is characterized by adjustment costs depending on the size of investments, is given by the following equation:

$$\hat{I}_t = \frac{1}{1+\beta} \hat{I}_{t-1} + \frac{\beta}{1+\beta} E_t \hat{I}_{t+1} + \frac{\varphi}{1+\beta} \hat{Q}_t + \frac{\beta E_t \hat{\varepsilon}_{t+1}^I - \hat{\varepsilon}_t^I}{1+\beta}, \quad (5)$$

with φ the inverse of the investment adjustment costs (S''), which captures the investment dynamics, and $\hat{Q}_t$ the real value of capital expressed in units of consumption goods. $\hat{\varepsilon}_t^I$ is an AR(1) shock to investment costs with an i.i.d. normal error term: $\varepsilon_t^I = \rho_I \varepsilon_{t-1}^I + \eta_t^I$. The real value of capital is represented by

$$\hat{Q}_t = -(\hat{R}_t - \hat{\pi}_{t+1}) + \frac{1 - \tau}{1 - \tau + \bar{r}^k} E_t \hat{Q}_{t+1} + \frac{\bar{r}^k}{1 - \tau + \bar{r}^k} E_t \hat{r}_{t+1}^k + \eta_t^Q, \quad (6)$$

with τ the depreciation rate, $\bar{r}^k$ the steady-state real rental rate of capital, and η_t^Q an i.i.d. normal shock that captures changes in the external finance premium due to informational frictions. However, this shock is not related to the structure of the economy, and it is further assumed that distortions caused by informational frictions are zero in the steady state. By definition, $\beta = 1/(1 - \tau + \bar{r}^k)$. The capital-accumulation equation fulfills³

$$\hat{K}_t = (1 - \tau) \hat{K}_{t-1} + \tau \hat{I}_t, \quad (7)$$

where K_t stands for capital. There is a final-goods sector and an intermediate-goods sector. The final-goods sector is perfectly competitive and produces one final good used for consumption and investment by the households. The intermediate-goods sector is characterized by monopolistic competition where each firm is the only producer of the intermediate good j . In the intermediate-goods sector each monopolist sets its price, given the total demand function for its differentiated good in line with Calvo (1983). As in the case of wage setting by households, each firm can only reoptimize its price after receiving a price-change signal with a constant probability $1 - \xi_p$. Firms that cannot reoptimize will adjust their price partially to past inflation with a degree $0 \leq \gamma_p \leq 1$. This leads to the following New Keynesian Phillips curve:

$$\begin{aligned} \hat{\pi}_t = & \frac{\beta}{1 + \beta\gamma_p} E_t \hat{\pi}_{t+1} + \frac{\gamma_p}{1 + \beta\gamma_p} \hat{\pi}_{t-1} \\ & + \frac{1}{1 + \beta\gamma_p} \frac{(1 - \beta\xi_p)(1 - \xi_p)}{\xi_p} [\alpha \hat{r}_t^k + (1 - \alpha) \hat{w}_t - \hat{\varepsilon}_t^a + \eta_t^p], \quad (8) \end{aligned}$$

with α the share of capital in total output in the Cobb-Douglas production function in the intermediate-goods sector with constant

³Note that we correct for a typo in SW (2003) by replacing lagged investment with its current value. In addition, we include in equation (10) the cost-of-capital-adjustment component (last term) in the first line, which is omitted in SW (2003).

returns to scale. $\hat{\varepsilon}_t^a$ is an AR(1) productivity shock with an i.i.d. normal error term: $\varepsilon_t^a = \rho_a \varepsilon_{t-1}^a + \eta_t^a$. The final term in square brackets represents the real marginal costs augmented with an i.i.d. (cost-push) shock η_t^p to the markup in the goods market: $\lambda_{p,t} = \lambda_p + \eta_t^p$. The labor-demand equation for a given capital stock is given by

$$\hat{L}_t = -\hat{w}_t + (1 + \Psi)\hat{r}_t^k + \hat{K}_{t-1}, \quad (9)$$

where Ψ is the inverse of the elasticity of the capital-utilization cost function, which comes from the condition $\hat{r}_t^k = \frac{1}{\Psi}\hat{z}_t$ and determines the dynamics. Finally, the goods-market equilibrium condition is represented by the following expression, where total output, $\hat{Y}_t$, equals total demand for output by households and the government (first line) and total supply of output by the firms (second line):

$$\begin{aligned} \hat{Y}_t &= c_y \hat{C}_t + g_y \varepsilon_t^G + \tau k_y \hat{I}_t + \bar{r}^k k_y \Psi \hat{r}_t^k \\ &= \phi(\hat{\varepsilon}_t^a + \alpha \hat{K}_t + \alpha \Psi \hat{r}_t^k + (1 - \alpha) \hat{L}_t), \end{aligned} \quad (10)$$

where ε_t^G is an AR(1) government spending shock with an i.i.d. normal error term: $\varepsilon_t^G = \rho_G \varepsilon_{t-1}^G + \eta_t^G$. c_y , g_y , and k_y stand for the steady-state ratio of consumption to output, government spending to output, and capital to output, respectively. ϕ is equal to $1 + \frac{\Phi}{\bar{Y}}$, where the second term is the ratio of fixed production cost (Φ) to output in the steady state ($\bar{Y}$).

In order to complete the model, we introduce optimizing monetary policy authorities, discussed in the next section, which distinguishes our model from the original SW (2003) specification where monetary policy is described by a generalized empirical Taylor rule of the following type⁴:

$$\begin{aligned} \hat{R}_t &= \rho \hat{R}_{t-1} + (1 - \rho) \{ \bar{\pi}_t + r_\pi (\hat{\pi}_{t-1} - \bar{\pi}_t) + r_y (\hat{Y}_{t-1} - \hat{Y}_{t-1}^p) \} \\ &\quad + r_{\Delta\pi} (\hat{\pi}_t - \hat{\pi}_{t-1}) + r_{\Delta y} ((\hat{Y}_t - \hat{Y}_t^p) - (\hat{Y}_{t-1} - \hat{Y}_{t-1}^p)) + \eta_t^R. \end{aligned} \quad (11)$$

⁴Note that we correct a typo in SW (2003) in the first line of the rule where the lag of the output gap appears instead of the current output gap.

The interest rate $\hat{R}_t$ is the monetary policy instrument; $\bar{\pi}_t$ is the inflation target that follows an AR(1) process with i.i.d. normal error term and is normalized to zero: $\bar{\pi}_t = \rho_\pi \bar{\pi}_{t-1} + \eta_t^\pi$. ρ is the monetary authorities' smoothing parameter, η_t^R is an i.i.d. normal monetary policy shock, and $(\hat{Y}_t - \hat{Y}_t^p)$ is the output gap defined as the difference between the actual output level and the level of potential output. The latter is assumed to be the output level under the absence of nominal rigidities (i.e., full price and wage flexibility) and the three i.i.d. cost-push shocks $(\eta_t^Q, \eta_t^p, \eta_t^w)$. In addition to the one-period difference in inflation and the output gap as measured by $r_{\Delta\pi}$ and $r_{\Delta y}$, respectively, a gradual response to lagged inflation (r_π) and lagged output gap (r_y) is assumed.

2.2 Optimal Monetary Policy

Monetary policy authorities are assumed to minimize a discounted intertemporal quadratic loss function of the following type:

$$E_t \sum_{i=0}^{\infty} \delta^i [y'_{t+i} W y_{t+i}], \quad 0 < \delta < 1, \quad (12)$$

with δ the monetary policy discount factor, which in the estimations will be fixed to the value chosen for the households' discount factor: $\delta = \beta = 0.99$. E_t is the expectations operator conditional on information available at time t . This type of quadratic loss function is commonly adopted in the literature, as in Rudebusch and Svensson (1999), Söderlind (1999), Giannoni and Woodford (2003), Svensson (2003), and Dennis (2007). The vector $y_t = [x'_t u'_t]'$ contains the $n \times 1$ endogenous variables and AR(1) exogenous variables in the model included in x_t and the $p \times 1$ vector of control variables included in u_t . Since we assume only one control variable—i.e., the interest rate— u_t is a scalar with $u_t = \hat{R}_t$. W is a time-invariant symmetric, positive semi-definite matrix of policy weights which reflect the monetary policy preferences of the central bank over the target variables. An alternative and theoretically more justifiable approach toward monetary policy is to derive the approximated welfare-based loss function. Levin et al. (2006) derive a welfare-based loss function in a DSGE model similar to the one of SW (2003) and consider the

effects of uncertainty on the conduct of monetary policy. The authors show that presence of uncertainty about model specification and the parameter estimates has an effect on the specification of the welfare-based loss function. However, adopting a welfare-based loss function would make our results rely more heavily on the specific assumptions about the underlying model of the economy than would be the case under an ad hoc loss function. Angeloni, Coenen, and Smets (2003) claim that the welfare loss from not using a welfare-based loss function is small enough to justify the use of an ad hoc loss function in the SW (2003) model. As shown by Woodford (2003), in more simple models this type of loss function can be obtained from the micro-foundations.⁵ Recent developments in the literature, however, have made it possible to consider linear-quadratic approximate problems that yield a solution representing a correct local linear approximation to the actual non-linear optimal policy. Benigno and Woodford (2008), among others, show that in a large number of cases, under the linear-quadratic approximation to the exact optimal policy problem, it is possible to correctly rank alternative simple rules as long as the size of the shocks is small enough. Altissimo, Cúrdia, and Rodríguez Palenzuela (2005) provide a standard algorithm that makes it possible to apply the theoretical framework of Benigno and Woodford (2008) to various type of models, including the SW (2003) model. Debortoli and Nunes (2006) in addition prove that the methodology of Benigno and Woodford (2006) can be generalized, and that in the case of a distorted steady state the timeless-perspective assumption is crucial in obtaining a linear-quadratic approximation to the welfare measure.

We follow the literature⁶ in assuming that the one-period ad hoc loss function for (12) includes inflation, a measure of the output gap, and a smoothing component for the policy instrument:

$$Loss_t = \hat{\pi}_t^2 + q_y(\hat{Y}_t - \hat{Y}_t^p)^2 + q_r(\hat{R}_t - \hat{R}_{t-1})^2. \quad (13)$$

⁵See also Rotemberg and Woodford (1999) and Giannoni and Woodford (2003), among others, for the same exercise in DSGE models.

⁶See, e.g., Rudebusch and Svensson (1999), Giannoni and Woodford (2003), Dennis (2004a), Söderström, Söderlind, and Vredin (2005), Castelnuevo (2006), and Lippi and Neri (2007).

As explained in the next section, the data set used in the estimation procedure contains series on (detrended) inflation considered in deviation from its sample mean. Since we normalize the inflation target in (13) to zero, this implies that monetary policy aims to stabilize inflation around the sample mean, which is a known constant.⁷ The inclusion of the term $\hat{R}_t - \hat{R}_{t-1}$ in the loss function can be justified by concerns about financial stability or in order to take into account the observed inertial behavior in the policy instrument, which suggests a gradualist monetary policy approach. The output gap considered here is the one implied by the model—i.e., the deviation of output from its natural level.⁸ As is done in most applications in the literature, the weight assigned to inflation in the above loss function is normalized to one.⁹ Hence, the weights corresponding to the output gap and the interest rate smoothing component—i.e., q_y and q_r , respectively—are relative weights with respect to the inflation target variable. We will estimate the preferences of monetary policy reflected by these parameters along with the structural parameters of the economy.

Since the structural model of the economy outlined above is characterized by nominal wage rigidities, we investigate also the case where monetary policy is concerned about stabilizing nominal wage inflation in addition to the target variables in (13). A case for a nominal wage-inflation target is provided by Erceg, Henderson, and Levin (1998, 2000) and Woodford (2003). In a dynamic general equilibrium model featuring wage and price inertia, monetary policy faces a trade-off between stabilizing wage and price inflation (in addition to the output gap) since rigidities in the two markets at the same

⁷One could criticize this approach and alternatively consider the inflation target as an additional parameter to be estimated. This approach is considered in Ilbas (forthcoming). Conditions under which the inflation target can be identified and estimated are provided in Dennis (2004a, 2004b).

⁸Although it is more common in empirical applications to use deviations of output from a linear trend as an approximation to the output gap, we prefer to adopt the theoretical concept of the output gap. To support this choice, we experiment with alternative definitions of the output gap in the next section which yield less favorable results in terms of the marginal likelihood.

⁹Examples can be found in Rotemberg and Woodford (1999), Rudebusch and Svensson (1999), Woodford (2003), and Dennis (2004a).

time make the Pareto optimal equilibrium unattainable for monetary policy. Moreover, Erceg, Henderson, and Levin (2000) derive a social welfare function under the circumstances of nominal wage and price rigidities and show that there are high welfare costs attached to targeting price inflation only and ignoring wage-inflation stabilization. Therefore, we analyze the following alternative specification of the loss function:

$$\begin{aligned} Loss_t = & \hat{\pi}_t^2 + q_y(\hat{Y}_t - \hat{Y}_t^p)^2 + q_r(\hat{R}_t - \hat{R}_{t-1})^2 \\ & + q_w(\hat{W}_t - \hat{W}_{t-1} + \hat{\pi}_t)^2, \end{aligned} \quad (14)$$

where q_w is the weight attached to the wage-inflation target. The specification of the one-period loss function in (13) can be considered as a special case of (14) where q_w is set to zero. We will estimate the model under both specifications of the loss function, by treating the case under each one of the two specifications as a separate model. Therefore, we will refer to the case where $q_w = 0$ —i.e., period loss function (13)—as model *M1* and consider the case where nominal wage inflation is an additional target variable as model *M2*. We then use the corresponding marginal-likelihood values in order to assess to what extent monetary policy has been concerned about nominal wage-inflation stabilization over the sample period.¹⁰

The central bank minimizes the intertemporal loss function (12), the one-period loss function of which is given by either (13) or (14), under commitment subject to the structural equations of the economy (3)–(10) augmented by their flexible-price versions, written and represented by the following second-order form:

$$Ax_t = BE_t x_{t+1} + Fx_{t-1} + Gu_t + Dz_t, \quad z_t \sim iid[0, \Sigma_{zz}], \quad (15)$$

with z_t an $n \times 1$ vector of stochastic innovations to the variables in x_t , having mean zero and variance-covariance matrix Σ_{zz} , and the

¹⁰We admit that in this study we make the strong assumption that the monetary policy regime has been unchanged in the euro area throughout the sample period. This might be doubtful given the separate monetary policy strategies adopted in the individual countries before the introduction of the euro. However, the transition period toward the euro and the restrictions imposed by the Maastricht Treaty justify our assumption that the policy regimes were more likely to have been in line.

matrices A , B , F , G , and D containing the structural parameters. We follow the optimization routine for commitment suggested by Dennis (2007). For details concerning this procedure, we therefore refer to Dennis (2007). The Euler equations of the monetary policy optimization problem are represented in the following second-order form:

$$A1^* \Upsilon_t = B1^* E_t \Upsilon_{t+1} + C1^* \Upsilon_{t-1} + D1^* z_t, \quad (16)$$

with $\Upsilon_t = \begin{bmatrix} x_t \\ u_t \\ \rho_t \end{bmatrix} = \begin{bmatrix} y_t \\ \rho_t \end{bmatrix}$ and the final term in Υ_t , ρ_t , the vector of

Lagrange multipliers. The matrices $A1^*$, $B1^*$, and $C1^*$ have dimension $(2n + p) \times (2n + p)$. In the next section, we will estimate the Euler equations (16) resulting from the optimization procedure.

3. Estimation

In this section we discuss the data set used and the methodology followed in estimating the system (16), which yields estimates of the structural parameters resulting from optimizing private agents and policy preferences of optimizing monetary policy authorities. Next, we present the results under $M1$ and $M2$ and compare them with each other and with the estimates obtained from the benchmark SW (2003) model.

3.1 Data

We use the updated version of the euro-area data set used in SW (2003), i.e., constructed by Fagan, Henry, and Mestre (2005). The data set contains observed series on real GDP, real consumption, real investment, GDP deflator, real wages, employment, and the nominal interest rate. The series range over the period 1980:Q2–2005:Q4, preceded by an initialization period of forty quarters. In appendix 2, we show that this pre-sample method in addition makes it possible to overcome the time-inconsistency problem in the estimation procedure by yielding non-zero initial values for the Lagrange multipliers of forward-looking variables, which is in line with the timeless-perspective approach of Woodford (2003).

As in SW (2003), all variables are considered in deviations from their sample means. Inflation and nominal interest rates are detrended by the inflation trend, whereas the remaining variables in the data set are detrended separately by a linear trend. Following SW (2003), we introduce an additional equation for employment to correct for the use of data on employment instead of data on aggregate hours worked in the euro area, since the latter is not available for the euro area:

$$\hat{E}_t - \hat{E}_{t-1} = \hat{E}_{t+1} - \hat{E}_t + \frac{(1 - \beta\xi_e)(1 - \xi_e)}{\xi_e}(\hat{L}_t - \hat{E}_t). \quad (17)$$

This equation reflects the fact that employment moves more slowly in response to shocks than aggregate hours worked. $\hat{E}_t$ is the amount of people employed and ξ_e is the fraction of firms that is able to respond to shocks in a given period by adjusting the level of employment. The labor demand in a given period is satisfied by adjusting the number of hours worked per employer, which is flexible and not observed. In the line of Hansen and Sargent (1980), we further assume that the econometrician does not necessarily have access to the same (i.e., more complete) data set possessed, and responded to, by the monetary authorities. In order to capture the orthogonal information possessed by the monetary authorities, which we are not able to observe, we therefore introduce an i.i.d. error ϵ_t^R to account for this discrepancy in the monetary policy instrument, i.e., the interest rate. Hence, the relation between the observed series of the nominal interest rate ($\hat{R}_t^{obs}$) and the non-observed policy instrument rate ($\hat{R}_t^{nobs}$) is expressed as follows:

$$\hat{R}_t^{obs} = \hat{R}_t^{nobs} + \epsilon_t^R. \quad (18)$$

As opposed to SW (2003), the two monetary policy shocks that appear in the generalized Taylor rule—i.e., a shock to the inflation objective and an interest rate shock—are absent from the models *M1* and *M2*. Given that there are seven observables only, we need a sufficient amount of restrictions to enable identification. As in SW (2003), identification is obtained through the assumption that all

shocks are uncorrelated, that the three cost-push shocks together with the interest rate error term follow a white noise process, and that the remaining shocks related to preferences and technology are AR(1). We also follow SW (2003) in fixing parameters that are related to the steady state of the state variables and cannot be estimated because the data set is demeaned. The discount factors β and δ are set equal to 0.99, implying an annual real interest rate of 4 percent in the steady state. The annual depreciation rate on capital is assumed to be 10 percent, i.e., $\tau = 0.025$. The income share of labor in total output is assumed to be 0.7 in the steady state, i.e., $\alpha = 0.3$. The share of consumption and investment in total output is 0.6 and 0.22 in the steady state, respectively. Finally, there is one parameter that needs to be fixed since it is not identified: the constant wage markup parameter λ_w is calibrated to be 0.5.

3.2 Methodology

We apply Bayesian methods in order to estimate the parameters of the alternative models.¹¹ The posterior density distribution is derived by combining the prior distribution with the likelihood function. After maximizing the posterior mode, we use the Metropolis-Hastings algorithm to generate draws from the posterior distribution in order to approximate the moments of the distribution, calculate the modified harmonic mean, and construct the Bayesian impulse responses.¹² In discussing the estimation results, we will focus on the maximized posterior mode and the Hessian-based standard errors.

Panel A in table 1 shows the details of the prior distributions for the shock processes, i.e., the standard errors σ of all nine shocks and the AR(1) coefficients ρ of the five preference shocks. The type of the prior distributions, the prior means, and the prior standard errors are identical to the assumptions made in SW (2003) and are kept constant throughout the estimation processes for both models *M1* and *M2*. All variances of the shocks are assumed to have an

¹¹For a detailed discussion on Bayesian estimation of DSGE models, we refer to SW (2003, 2005), Schorfheide (2006), and An and Schorfheide (2007).

¹²All estimations are performed using Michel Juillard's software Dynare, which can be downloaded from the web site www.dsge.net.

Table 1. Prior Specifications of the Shocks and the Structural Parameters

A. Parameters of Shock Processes	Prior Specification		
	Distribution Type	Prior Mean	Prior se
σ Productivity Shock	Inverse Gamma	0.4	2
σ Preference Shock	Inverse Gamma	0.2	2
σ Government Spending Shock	Inverse Gamma	0.3	2
σ Labor Supply Shock	Inverse Gamma	1	2
σ Investment Shock	Inverse Gamma	0.1	2
σ Equity Premium Shock	Inverse Gamma	0.4	2
σ Price Markup Shock	Inverse Gamma	0.15	2
σ Wage Markup Shock	Inverse Gamma	0.25	2
σ Inflation Objective Shock (Only SW (2003))	Inverse Gamma	0.02	2
σ Monetary Policy Shock (Only SW (2003))	Inverse Gamma	0.1	2
σ Interest Rate Error (Only M1 and M2)	Gamma	0.05	0.025
ρ Inflation Objective Shock (Only SW (2003))	Beta	0.85	0.1
ρ Productivity Shock	Beta	0.85	0.1
ρ Preference Shock	Beta	0.85	0.1
ρ Government Spending Shock	Beta	0.85	0.1
ρ Labor Supply Shock	Beta	0.85	0.1
ρ Investment Shock	Beta	0.85	0.1

(continued)

Table 1. (Continued)

B. Structural Parameters	Prior Specification		
	Distribution Type	Prior Mean	Prior se
$S''(.)$ Investment Adjustment Cost	Normal	4	1.5
σ_c Consumption Utility	Normal	1	0.375
h Consumption Habit	Beta	0.7	0.1
ξ_w Calvo Wages	Beta	0.75	0.05
σ_L Labor Utility	Normal	2	0.75
ξ_p Calvo Prices	Beta	0.75	0.05
ξ_e Calvo Employment	Beta	0.5	0.15
γ_w Indexation Wages	Beta	0.75	0.15
γ_p Indexation Prices	Beta	0.75	0.15
Capital-Utilization Adjustment Cost	Normal	0.2	0.075
ϕ Fixed Cost	Normal	1.45	0.125
ρ Smoothing Parameter Empirical Taylor Rule	Beta	0.8	0.1
r_π Lagged Inflation Parameter Empirical Taylor Rule	Normal	1.7	0.1
r_y Lagged Output-Gap Parameter Empirical Taylor Rule	Normal	0.125	0.05
$r_{\Delta\pi}$ Inflation Differential Empirical Taylor Rule	Normal	0.3	0.1
$r_{\Delta y}$ Output-Gap Differential Empirical Taylor Rule	Normal	0.0625	0.05
q_r Interest Smoothing Preference	Gamma	0.5	0.4
q_y Output-Gap Preference	Gamma	0.5	0.4
q_w Wage-Inflation Preference	Gamma	0.5	0.4
Notes: Panel A summarizes the assumed prior distributions, the prior means, and the prior standard errors in the estimation of the standard errors of the shocks (σ), followed by the AR(1) coefficients of the persistent shocks (ρ). Panel B in the table shows the assumed prior distributions, the prior means, and the prior standard errors in the estimation of the structural parameters and the policy preference parameters.			

inverted gamma distribution with two degrees of freedom, except for the interest rate error term, which we assume to be gamma distributed with a prior mean of 0.05 and standard error of 0.025. The five AR(1) coefficients are assumed to have a beta distribution with a prior mean of 0.85 and a strict prior standard error of 0.1 in order to distinguish the persistent shocks clearly from the i.i.d. shocks.

The prior specifications of the structural and the monetary policy preference parameters—i.e., the weights assigned to the target variables q_y , q_r , and q_w —are reported in panel B of table 1. The monetary policy preference parameters are assumed to be gamma distributed, in order to rule out negative values, with 0.5 prior mean. We further assign a loose prior standard error of 0.4 given that our prior knowledge about these parameters is rather restricted.

3.3 Results

In this section, we report and discuss the estimation results obtained for $M1$ and $M2$.¹³ We also reestimate the SW (2003) version of the model, which will serve as a benchmark for the results obtained under optimal monetary policy.

3.3.1 Structural Shocks and Private-Sector Parameters

The results obtained from the posterior maximization—i.e., the posterior mode and the (Hessian-based) standard errors—for the models $M1$, $M2$, and the benchmark SW (2003) model are reported in table 2.

In panel A of the table, the results for the shock processes are shown. The symbol σ stands for the standard error of each shock and ρ for the AR(1) parameter of the persistent shocks. The estimated

¹³In our experiments we consider alternative loss functions where we replace the interest rate smoothing term by the interest rate level (as in Giannoni and Woodford 2002a, 2002b; Woodford 2003; and Onatski and Williams 2004), or the output gap by the differences in output. We also studied the case where we used simply output deviations from a linear trend instead of the model-consistent output gap. We examined loss functions including a difference in the output gap, a difference in the inflation rate, or of the welfare-approximated type presented by Onatski and Williams (2004) as well. None of these cases, however, could yield better outcomes in terms of their corresponding marginal-likelihood values.

Table 2. Estimates of the Shock Processes and the Structural Parameters for the Period 1980:Q2–2005:Q4

A. Parameters of Shock Processes	Results from Posterior Maximization $M1(q_w = 0)$		Results from Posterior Maximization $M2(q_w \neq 0)$		Results SW (2003) (Reestimated)	
	Mode	se (Hessian)	Mode	se (Hessian)	Mode	se (Hessian)
σ Productivity Shock	0.5767	0.0754	0.5769	0.0755	0.5705	0.0738
σ Preference Shock	0.1799	0.0497	0.1749	0.0494	0.3053	0.0750
σ Government Spending Shock	0.2961	0.0205	0.2961	0.0205	0.2948	0.0203
σ Labor Supply Shock	2.0571	0.4004	2.0676	0.4047	3.1778	0.7346
σ Investment Shock	0.0392	0.0120	0.0393	0.0121	0.0506	0.0153
σ Equity Premium Shock	0.5735	0.0611	0.5745	0.0617	0.6192	0.0503
σ Price Markup Shock	0.1601	0.0134	0.1597	0.0134	0.1611	0.0136
σ Wage Markup Shock	0.2854	0.0211	0.2858	0.0212	0.2014	0.0165
σ Inflation Objective Shock (Only SW (2003))	—	—	—	—	0.0092	0.0038
σ Monetary Policy Shock (Only SW (2003))	—	—	—	—	0.0734	0.0158
σ Interest Rate Error (Only M1 and M2)	0.0439	0.0069	0.0444	0.0071	—	—
ρ Inflation Objective Shock (Only SW (2003))	—	—	—	—	0.9219	0.0859
ρ Productivity Shock	0.9867	0.0096	0.9867	0.0095	0.9835	0.0104
ρ Preference Shock	0.8752	0.0322	0.8772	0.0323	0.8673	0.0294
ρ Government Spending Shock	0.9342	0.0216	0.9342	0.0216	0.9466	0.0229
ρ Labor Supply Shock	0.9892	0.0071	0.9898	0.0070	0.9336	0.0212
ρ Investment Shock	0.9262	0.0520	0.9251	0.0532	0.9197	0.0331

(continued)

Table 2. (Continued)

B. Structural Parameters	Results from Posterior Maximization $M1(q_w = 0)$		Results from Posterior Maximization $M2(q_w \neq 0)$		Results SW (2003) (Reestimated)	
	Mode	se (Hessian)	Mode	se (Hessian)	Mode	se (Hessian)
$S''(.)$	4.4435	1.0543	4.4222	1.0508	6.8295	1.0566
σ_c	1.5026	0.3015	1.5080	0.3018	1.3687	0.2646
h	0.5655	0.0642	0.5651	0.0641	0.6281	0.0607
ξ_w	0.8930	0.0161	0.8953	0.0161	0.7273	0.0364
σ_L	1.4510	0.4848	1.4605	0.4894	2.4772	0.5757
ξ_p	0.9038	0.0097	0.9040	0.0098	0.9279	0.0082
ξ_e	0.7438	0.0289	0.7438	0.0287	0.7410	0.0261
γ_w	0.7704	0.1265	0.7684	0.1269	0.3578	0.1257
γ_p	0.4379	0.0726	0.4476	0.0722	0.4531	0.0883
Capital Utilization	0.2153	0.0730	0.2146	0.0731	0.2076	0.0734
ϕ	1.3437	0.1169	1.3454	0.1170	1.3239	0.1138
ρ	—	—	—	—	0.9642	0.0104
r_π	—	—	—	—	1.6920	0.0999
r_y	—	—	—	—	0.1271	0.0404
$r_{\Delta\pi}$	—	—	—	—	0.1338	0.0415
$r_{\Delta\pi}$	—	—	—	—	0.1815	0.0215
q_r	0.9267	0.4128	0.9258	0.4159	—	—
q_y	0.0410	0.0191	0.0406	0.0192	—	—
q_w	—	—	0.1092	0.1427	—	—

Notes: Panel A shows the posterior mode and the corresponding Hessian-based standard errors for the shock processes (with standard errors σ and AR(1) coefficients of the persistent shocks ρ). Panel B in the table shows the estimated posterior mode of the structural parameters and the policy preference parameters (in bold), followed by the Hessian-based standard errors for each estimate in M1 and M2, and in the SW (2003) specification of the model for the sample period 1980:Q2–2005:Q4. A pre-sample period of forty quarters is used. M1 corresponds to the loss function specification that does not include the wage-inflation target, i.e., $Loss_t = \hat{\pi}_t^2 + q_y(\hat{Y}_t - \hat{Y}_t^p)^2 + q_r(\hat{R}_t - \hat{R}_{t-1})^2$. M2 is the case where an additional wage-inflation target is considered in the estimations, i.e., $Loss_t = \hat{\pi}_t^2 + q_y(\hat{Y}_t - \hat{Y}_t^p)^2 + q_r(\hat{R}_t - \hat{R}_{t-1})^2 + q_w(\hat{W}_t - \hat{W}_{t-1} + \hat{\pi}_t)^2$. Both of these specifications are compared with the results obtained for the benchmark SW (2003) specification of the model where monetary policy is described by the empirical Taylor rule $\hat{R}_t = \rho\hat{R}_{t-1} + (1 - \rho)\{\bar{\pi}_t + r_\pi(\hat{\pi}_{t-1} - \bar{\pi}_t) + r_y(\hat{Y}_{t-1} - \hat{Y}_{t-1}^p) + r_{\Delta\pi}(\hat{\pi}_t - \hat{\pi}_{t-1}) + r_{\Delta y}(\hat{Y}_t - \hat{Y}_t^p) - (\hat{Y}_{t-1} - \hat{Y}_{t-1}^p) + \eta_t^R\}$.

parameters and their corresponding standard errors under our specification $M1$ of the model are similar to those under $M2$. The preference, labor supply, and investment shocks are estimated to be lower under the models $M1$ and $M2$ compared with SW (2003). Furthermore, a more persistent labor supply shock is estimated under $M1$ and $M2$ than under the SW (2003) specification.

Comparing the SW (2003) estimates of the structural parameters with these obtained under $M1$ and $M2$, which are reported in panel B of table 2, yields the following conclusions. The investment adjustment cost parameter is estimated to be lower than the baseline estimated value. This suggests a higher elasticity of investment with respect to an increase in the current price of installed capital of 1 percent under $M1$ and $M2$. A strikingly higher value is obtained under $M1$ and $M2$ for the Calvo wage parameter ξ_w than in the baseline case. Exactly the same conclusion holds for the wage-indexation parameter γ_w . Hence, a significantly higher wage stickiness and wage indexation is present whenever monetary policy is assumed to behave optimally, suggesting an average duration of wage contracts of slightly more than two years. The reverse conclusion can be drawn from estimates for the Calvo price ξ_p and the price-indexation parameter γ_p . These parameters are lower under the models characterized by optimal monetary policy. This implies a lower degree of price stickiness in the goods markets under $M1$ and $M2$ compared with SW (2003), with an average duration of price contracts of more than two years, which is very similar to the average duration of the wage contracts. The estimates of γ_p imply that price indexation is lower under $M1$ and $M2$, and is in line with the findings of Galí, Gertler, and López-Salido (2001) for the euro area, where a low degree of backward-looking behavior in the goods market is estimated. Salemi (2006) likewise finds support for forward-looking behavior in a New Keynesian model when monetary policy is assumed to be committed to a simple rule; in particular, the coefficient on the forward-looking term in the Phillips curve turns out to be significant, which is not the case when the optimality assumption is not imposed. Hence, the degree of forward-lookingness is more precisely estimated under the assumption of monetary policy optimality in Salemi (2006). This finding is supported by our estimation results for the inflation equation. Finally, the estimate of the

labor-utility parameter σ_L is considerably lower under the models $M1$ and $M2$ than under SW (2003).¹⁴

3.3.2 Monetary Policy Preference Parameters

Table 2 also shows the parameter estimates of our main interest, i.e., the monetary policy preferences in the two models $M1$ and $M2$.¹⁵ The preference for interest rate smoothing is estimated to be higher than the preference for output-gap stabilization, while overall the main concern is still the inflation target. However, although the importance of the output gap as a target variable is estimated to be small, the posterior probability interval for this parameter q_y ranges above zero, as can be deduced from the posterior distribution for this parameter in model $M1$ displayed in figure 5 in appendix 1. The 90 percent posterior lower and upper bound for this parameter is $[0.0209 - 0.1249]$ in model $M1$. This implies that although the weight on the output gap is estimated to be small, it is significantly different from zero.¹⁶ The estimated weight assigned to both the output gap q_y and the interest rate smoothing preference parameter q_r hardly changes over varying specifications of the loss function, with values around 0.04 and 0.92, respectively. The high weight on interest rate smoothing suggests that although inflation remains the main target, monetary policy in the euro area has been prudent due to the initial concerns of establishing a credible single monetary policy regime in the euro area. In addition, financial-stability concerns

¹⁴Note that, as was the case in the original SW (2003) paper, our estimates of this parameter did not appear to be robust across specifications.

¹⁵Since inflation and the interest rates, both target variables, are measured on a quarterly basis and the literature occasionally considers target variables on a yearly frequency, the weights obtained from the estimates have to be adjusted in order to make the results comparable to those in the literature that base their results on yearly data. Taking this into account boils down to multiplying q_y by a factor of 16 and converting the inflation and the interest rate in the model to a yearly frequency.

¹⁶It should also be noted that in order to evaluate the relative importance of the components in the loss function, their corresponding weights should be combined with the relative volatility of the related variables. Hence the weights per se are not sufficient to evaluate the importance of the alternative target variables, due to the fact that the output-gap concept we use in this study is a theoretical one.

might have played an important role in attaching a high weight on smoothing the interest rate.

In general, estimates of a small role for the output gap seem to find support in the literature. Lippi and Neri (2007) estimate a very low value for q_y for the euro area, although they use a different output-gap concept¹⁷ and assume that monetary policy is conducted under discretion, which requires some caution in comparing the results. Dennis (2004a) likewise finds an ignorable weight for the output gap for the United States under discretionary monetary policy.¹⁸ Salemi (2006) and Givens and Salemi (2008), who consider commitment to a simple rule, estimate a close-to-zero and insignificant weight on the output gap in the context of a forward-looking model for the United States. Söderlind (1999) and Ilbas (forthcoming), considering the case of full commitment, however, estimate a relatively high value for q_y in the framework of a standard loss function similar to (13). However, it is important to keep in mind that, as in Ilbas (forthcoming), our output-gap concept differs from the one used in the other studies, which makes comparison of the results a bit troublesome. While Svensson (1999, 2002a) argues for a case of gradual monetary policy where some weight should be given to stabilizing the output gap, findings in the literature mentioned above for q_y generally do not support this concept of flexible inflation targeting. On the other hand, the computed posterior interval for q_y in this paper suggests that monetary policy in the euro area has considered the output gap as an important target variable.

Dennis (2004a) and Lippi and Neri (2007) estimate a weight on the interest rate smoothing component that is higher than the weight on the inflation target.¹⁹ This is not the case in our study. Although our estimates of q_r show that interest rate smoothing is a relatively important target, inflation remains the main policy goal. From an

¹⁷Lippi and Neri (2007) describe the output gap as the deviation of output from a linear trend. On the contrary, we assume that the output gap is the deviation of output from the natural output level in the absence of nominal rigidities and the three i.i.d. cost-push shocks.

¹⁸Söderström, Söderlind, and Vredin (2005) show in their calibration exercise under discretion analogously a low concern for output-gap stabilization based on U.S. data. See also Favero and Rovelli (2003).

¹⁹Söderström, Söderlind, and Vredin (2005) show in their calibration exercise under discretion also a high importance for interest rate smoothing based on U.S. data.

economic point of view, this finding is plausible and in line with the statements that inflation should be the main target variable in monetary policy's objective function. Castelnovo (2006) finds, through a calibration exercise in the framework of discretionary monetary policy, a value for the interest rate smoothing weight close to ours whenever forward-looking agents are added to the model. When agents are assumed to be backward looking only, this weight increases considerably up to a point where interest rate smoothing becomes twice as important as the inflation target. This leads Castelnovo (2006) to conclude that finding an economically difficult justifiable high value for q_r is probably due to model misspecification by the omission of factors like forward-looking behavior. Salemi (2006) confirms the empirical importance of forward-looking behavior by estimating a relatively low weight on interest rate stability in the loss function, suggesting inflation to be the main target, as in Givens and Salemi (2008) and Ilbas (forthcoming), while finding a high coefficient on the lagged interest rate in monetary policy's reaction function. Since the model we consider includes forward-looking agents, our values for q_r are therefore in line with these results.²⁰ An intuitive explanation for this moderating effect of the presence of forward-looking agents on the estimated weight assigned to the interest rate smoothing component in the loss function can be given as follows.²¹ Whenever (rational) agents are forward looking, their expectations will play a key role in the stabilization process of monetary policy and the law of motion of the target variables. If the economy is hit by a shock in the current period, requiring a change in the policy instrument rate in order to stabilize the target variables, expectations will adjust accordingly, and since agents are rational, they will take into account the fact that interest rate smoothing is a

²⁰In order to assess this positive link between the degree of backward-lookingness and the estimates of the preference for interest rate smoothing in our model, we look at the correlation between the series on the inflation-indexation parameter γ_p and the interest rate smoothing preference parameter q_r obtained from the Markov chain Monte Carlo (MCMC) draws. Based on these draws, we detect a positive correlation of around 0.3, which is in line with Castelnovo (2006).

²¹Castelnovo (2006) provides a detailed explanation on this issue and quantifies the role played by forward-looking agents in lowering the calibrated values of the weights on the interest rate smoothing component. See also Sack and Wieland (2000) as a key reference on this issue.

target variable. Therefore, a slow and persistent move in the interest rates is anticipated. Hence, expectations will have a stabilizing effect on current inflation and output gap, which in turn results in a slow and inertial behavior in the interest rates. If agents were assumed to be backward looking, like in the case of Ozlale (2003) and Dennis (2006), this inertial behavior in the interest rates could only be taken into account by the assumption that smoothing receives a high weight in the loss function of the central bank. On the other hand, if agents are forward looking, interest rate inertia is attributed to the stabilizing effect of expectations, which results in lower concern for the interest rate smoothing target. Also note that the commitment framework assumed in this paper enforces this history dependence more than would be the case under discretion like in most studies previously mentioned. This suggests that if we would perform a similar exercise under discretionary monetary policy, the estimated values for q_r would probably be higher. Finally, our estimates of q_r do not seem to support the argument of Svensson (2002a, 2003) that an interest rate stabilization or smoothing component should not enter the loss function, since the values obtained for the smoothing target are significantly higher than zero.

When nominal wage inflation is introduced in the loss function of monetary policy—i.e., the case of $M2$ —this additional target receives a weight of 0.1092. However, on the basis of comparison of marginal-likelihood values in table 3, the $M1$ specification of the model where $q_w = 0$ is preferred over $M2$.²² Hence, we can conclude that inflation, interest rate smoothing, and output-gap stabilization are the only targets of monetary policy.

Figures 5–7 in appendix 1 plot the prior and the posterior distributions of the parameters for model $M1$. We apply the Metropolis-Hastings sampling algorithm based on 100,000 draws in order to derive the posterior distributions.²³

²²Given that the $M2$ specification of the model includes an additional parameter to $M1$, the worse performance of $M2$ is mainly driven by the prior. We experimented with different prior assumptions for the monetary policy preference parameters but found that this observation does not change with the alternative prior assumptions.

²³Convergence is assessed graphically by the Brooks and Gelman (1998) MCMC univariate diagnostics for each individual parameter and the MCMC multivariate diagnostics for all parameters simultaneously. These graphs are available upon request.

Table 3. Model Comparison Based on the Marginal Likelihood for the Period 1980:Q2–2005:Q4

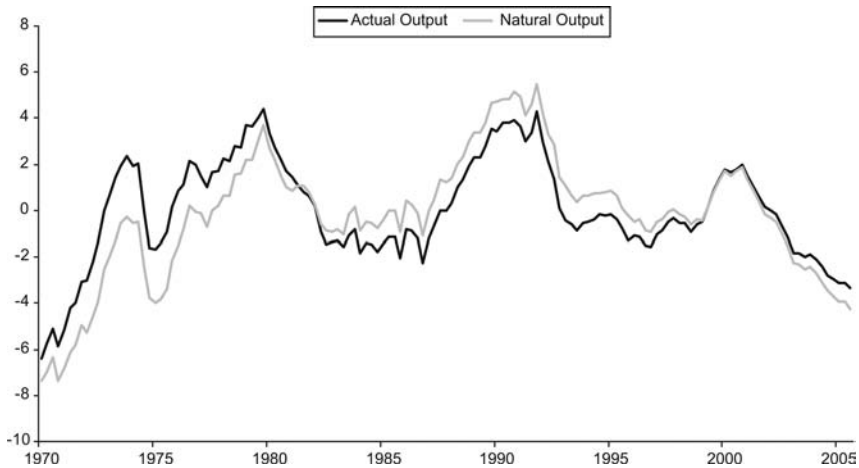
	<i>M1</i> ($q_w = 0$)	<i>M2</i> ($q_w \neq 0$)	SW (2003)
Laplace Approximation	−275.35	−276.26	−245.21
$\log \frac{p(Q_T M_i)}{p(Q_T SW(2003))} \quad i = 1, 2$	−30.14	−31.05	—
Modified Harmonic Mean	−274.65	−275.18	−244.65
$\log \frac{p(Q_T M_i)(1970:2-2005:4)}{p(Q_T M_i)(1970:2-1980:1)}$ $i = 1, 2, SW(2003)$	−300.61	−1331.4	−253.89
Notes: All estimations are performed with the same pre-sample period 1970:Q2–1980:Q1. The table shows the comparison for the period 1980:Q2–2005:Q4. The first row reports the Laplace approximation to approximate the marginal likelihood through the evaluation at the posterior mode. The table also reports the modified harmonic mean for each model (third row), obtained through the Markov chain Monte Carlo simulations, which does not differ much from the Laplace approximation. The second row in the table computes the performance of the models <i>M1</i> and <i>M2</i> with respect to the benchmark SW (2003) model. The fourth row computes the marginal likelihood by correcting for the effect of common priors, since same priors are used over the alternative model specifications, which can bias the model selection process.			

Finally, given that the model-consistent output-gap framework adopted in this paper differs from the more commonly used output-gap definition, we plot the natural output level implied by the estimated model *M1* against the historical path of the actual output level. Figure 1 illustrates a positive output-gap pattern until the late seventies, i.e., the period during which inflation was high. From the early eighties until the end of the nineties, the output gap is negative but turns slightly positive again around 2003.

4. Model Comparison

In this section we compare the models characterized by optimal monetary policy—i.e., the models *M1* and *M2*—and the benchmark SW (2003) model, based on their marginal-likelihood values²⁴ reported in table 3. In a next step, we compare Bayesian impulse responses for

²⁴See Geweke (1999) and Schorfheide (2006) for a detailed discussion on the marginal-likelihood function in Bayesian estimation.

Figure 1. Actual vs. Natural Output Levels in Model $M1$ 

Notes: The dark line in the figure describes the historical path of the actual output level; the shaded line represents the estimate of the natural output level in model $M1$. The latter is the output level in the absence of nominal rigidities (i.e., full price and wage flexibility) and the three i.i.d. cost-push shocks $(\eta_t^Q, \eta_t^p, \eta_t^w)$.

selected shocks under optimizing monetary policy authorities ($M1$) with those under SW (2003).

4.1 *Marginal-Likelihood Comparison*

The marginal likelihood of a model can be represented as follows:

$$p(Q_T | M_i) = \int_{\omega} p(Q_T | \omega, M_i) p(\omega | M_i) d\omega, \quad (19)$$

where Q_T contains the observable data series, ω the vector of parameters, and M_i the model under consideration—in our case of three models, $i = 1, 2$, or SW (2003). The likelihood function $p(Q_T | \omega, M_i)$ of the data series is conditional on the parameter vector ω and the model M_i . $p(\omega | M_i)$ is the prior density of the parameters conditional on the model. As in Schorfheide (2000), we use the Laplace approximation to approximate the marginal likelihood through the evaluation at the posterior mode. Table 3 also reports the modified harmonic mean for each model, obtained through the

Markov chain Monte Carlo simulations, which does not differ much from the Laplace approximation. As pointed out by Sims (2003) and Del Negro and Schorfheide (2008), using the same priors for alternative specifications of a model can bias our choice toward one type of specification. Given this potential pitfall in model comparison, we correct for the effect of common priors by estimating and evaluating the models over the training sample 1970:Q2–1980:Q1 as well, and subtract accordingly the corresponding marginal likelihood from the one obtained by estimation over the whole sample period 1970:Q2–2005:Q4. The conclusions based on this comparison, however, are in line with the comparisons reported in the first three rows in table 3.

Although from the table we can conclude that model $M1$, where $q_w = 0$, fits the data better than model $M2$, both models perform relatively worse than the benchmark SW (2003) specification of the model. This result indicates that monetary policy was not optimal (under commitment) during the sample period, which is not very surprising due to the restrictive assumption of optimizing monetary policy in the models $M1$ and $M2$. A similar comparison exercise is performed by Salemi (2006), who constructs a likelihood-ratio statistic to test the restrictions imposed by the optimal policy hypothesis under which the parameters of the policy rule are over-identified. Based on this formal test, Salemi (2006) shows that the estimates based on the structural estimation strategy with an optimal reaction function (i.e., the optimal monetary policy restriction) provide a worse fit to the data than in the case where the optimal monetary policy restriction is not imposed, which leads to the rejection of the hypothesis that monetary policy was optimal in the United States over the period 1965–2001. Moreover, Salemi (2006) and Givens and Salemi (2008) show that imposing the restrictions associated with the optimal monetary policy hypothesis leads to an alteration and sharpening of the estimates of some of the structural parameters, which is, as mentioned in the previous section, in line with our findings for the structural parameters.

Also note that in the SW (2003) description of monetary policy behavior, the Taylor rule includes five parameters, while there are only two monetary policy preference parameters to be estimated in $M1$. Hence, it is not very surprising that the SW (2003) model with more free parameters to be estimated performs better than $M1$. We therefore test whether this conclusion would change if the Taylor

rule (11) in the benchmark SW (2003) is replaced by two alternative specifications that contain fewer free parameters, and compare accordingly their corresponding marginal-likelihood values with $M1$. In the first alternative case, the SW (2003) model is reestimated with a Taylor rule that responds to only lagged inflation and lagged output gap:²⁵

$$\hat{R}_t = \rho \hat{R}_{t-1} + (1 - \rho) \{ \bar{\pi}_t + r_\pi (\hat{\pi}_{t-1} - \bar{\pi}_t) + r_y (\hat{Y}_{t-1} - \hat{Y}_{t-1}^p) \} + \eta_t^R. \quad (20)$$

The corresponding value of the marginal likelihood is -292.23 , which worsens the fit of the SW (2003) benchmark against $M1$ characterized by optimal policy. In the second alternative case, the SW (2003) model is reestimated with a forward-looking version of the rule (20), where monetary policy reacts to a four-quarter change in the inflation rate and a four-quarter change in the output gap:

$$\begin{aligned} \hat{R}_t = \rho \hat{R}_{t-1} + (1 - \rho) \{ & \bar{\pi}_t + r_\pi (E_t \hat{\pi}_{t+4} - \bar{\pi}_t) \\ & + r_y E_t (\hat{Y}_{t+4} - \hat{Y}_{t+4}^p) \} + \eta_t^R. \end{aligned} \quad (21)$$

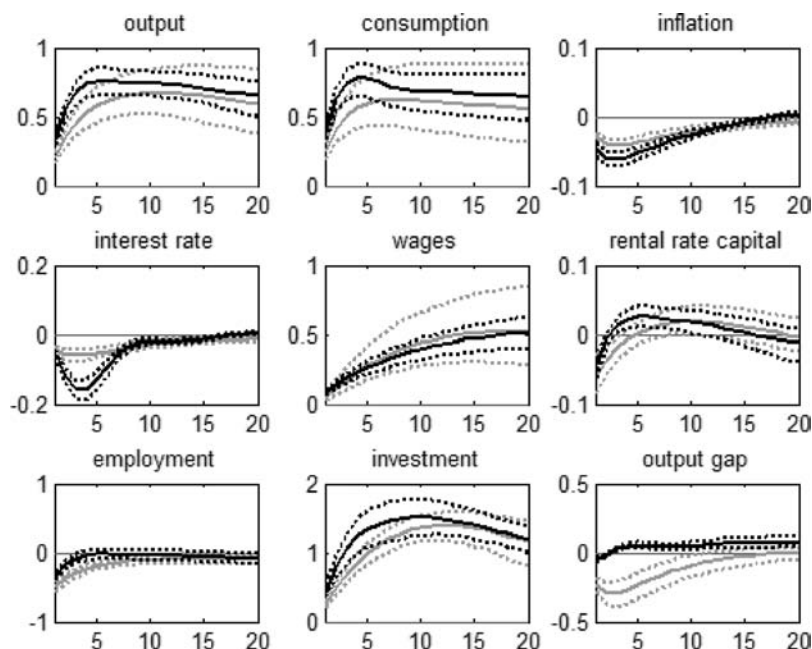
This version of the Taylor rule increases the value of the marginal likelihood to -244.69 , i.e., improving the fit of the SW (2003) benchmark against the $M1$ specification of the model.²⁶ Therefore, the performance of the model when monetary policy is assumed to be optimal against the benchmark where policy is simply described by a Taylor rule largely depends on how the latter is defined.

4.2 Bayesian Impulse-Response Analysis

In this section, we visualize the consequences of assuming optimizing monetary policy authorities on the impact and the dynamics of the variables in the case of a supply shock (productivity shock), a demand shock (equity premium shock), and a cost-push shock (price markup shock) over a period of twenty quarters in figures 2–4. We

²⁵Note that this rule is a special case of the original Taylor rule (11) where $r_{\Delta\pi} = r_{\Delta y} = 0$.

²⁶Note that when we assume a four-quarter change in the inflation rate only, i.e., $\hat{R}_t = \rho \hat{R}_{t-1} + (1 - \rho) \{ \bar{\pi}_t + r_\pi (E_t \hat{\pi}_{t+4} - \bar{\pi}_t) + r_y (\hat{Y}_t - \hat{Y}_t^p) \} + \eta_t^R$, the marginal likelihood worsens to -357.06 .

Figure 2. Productivity Shock

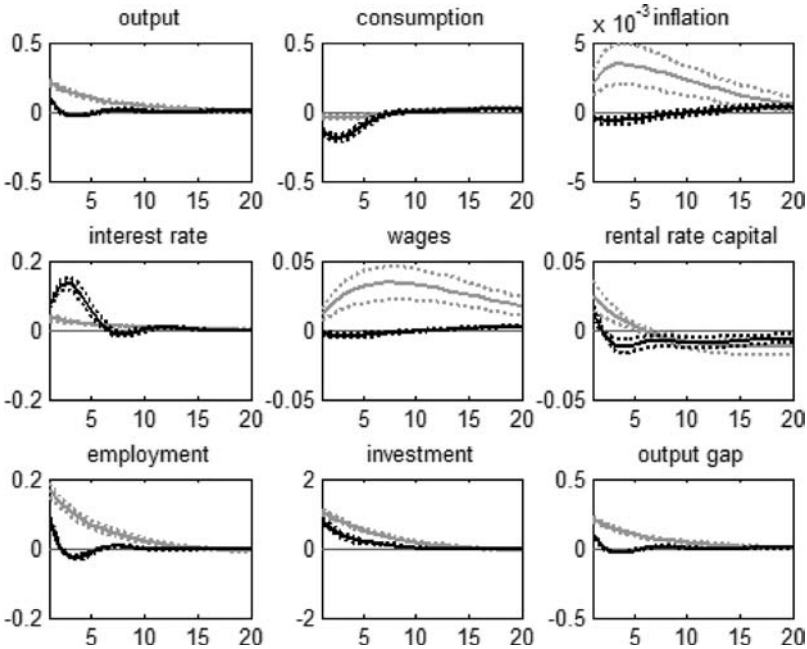
Notes: The figure shows the impulse responses of selected variables to a productivity shock in the SW (2003) model (shaded lines) and in the model *M1* (dark lines) estimated for the period 1980:Q2–2005:Q4. The solid lines are the mean impulse responses, whereas the dotted lines are the 10 percent and the 90 percent posterior intervals. The horizon of twenty quarters is given on the horizontal axis; the vertical axes show the percentage deviations of the variables from the steady state.

take the SW (2003) model, which includes an estimated policy reaction function as the benchmark case (shaded lines), and assess to which extent the reactions of the variables differ when monetary policy minimizes an intertemporal loss function with one-period loss as specified under model *M1* (dark lines).²⁷

Figure 2 shows the responses of the variables to a productivity shock. The interest rate shows a slightly lower impact and becomes more negative (accommodative) around the third quarter under *M1*

²⁷We focus only on the impulse responses obtained under *M1*. However, the impulse responses under *M2* are, with the exception of responses to the equity premium shock, very similar to those under *M1*.

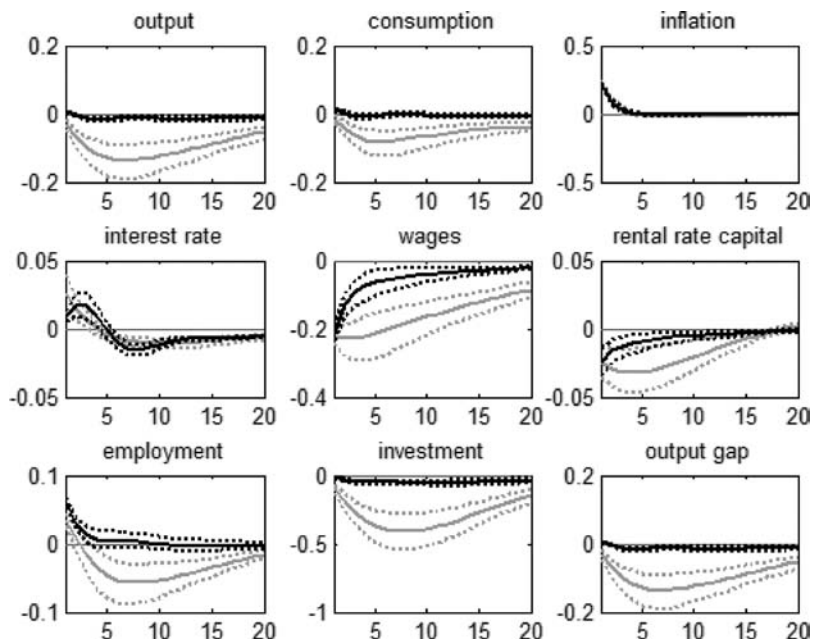
Figure 3. Equity Premium Shock



Notes: The figure shows the impulse responses of selected variables to an equity premium shock in the SW (2003) model (shaded lines) and in the model *M1* (dark lines) estimated for the period 1980:Q2–2005:Q4. The solid lines are the mean impulse responses, whereas the dotted lines are the 10 percent and the 90 percent posterior intervals. The horizon of twenty quarters is given on the horizontal axis; the vertical axes show the percentage deviations of the variables from the steady state.

as opposed to the benchmark SW (2003) case. Hence, consumption and investment both increase to a greater extent, resulting in a higher increase in output and lower decrease in employment. The output gap becomes only slightly negative, in contrast to the SW (2003) benchmark case, since monetary policy accommodates the productivity shock more strongly. While the impact on wages is similar in both models, the initial decline in the rental rate of capital is slightly smaller under *M1*. Finally, the initial effect on inflation is slightly more negative.

Figure 3 shows the impulse responses to an equity premium shock. The interest rate responds more strongly to the equity premium shock and gets more positive (more activist policy) around

Figure 4. Price Markup Shock

Notes: The figure shows the impulse responses of selected variables to a price markup shock in the SW (2003) model (shaded lines) and in the model *M1* (dark lines) estimated for the period 1980:Q2–2005:Q4. The solid lines are the mean impulse responses, whereas the dotted lines are the 10 percent and the 90 percent posterior intervals. The horizon of twenty quarters is given on the horizontal axis; the vertical axes show the percentage deviations of the variables from the steady state.

the third quarter, in contrast to the baseline model, which explains the stronger initial decline in consumption and the weaker response of investment and hence employment, output, and the output gap. The impact on both wages and rental rate of capital is much weaker and even turns slightly negative under *M1*. Therefore, inflation responds negatively; however, the effect is very small.

Comparing the impulse responses to a price markup shock in figure 4 with the benchmark case yields the following conclusions. The impact and dynamics of the interest rate are similar to the benchmark case. The initial decrease in investment is lower than the benchmark model, followed by weak dynamics that are close to zero. In contrast to what we observe for SW (2003), the impact of

the shock on consumption is slightly positive under $M1$, followed by smaller dynamics also in this case. Hence, it is not surprising that output and employment are affected very weakly in the initial period. The former reaches its lowest value around the tenth quarter, which is still higher than in SW (2003). Interestingly, and in line with the findings of Salemi (2006), inflation shows very similar patterns in both models.

5. Conclusion

In this paper, we estimate the preferences of optimizing monetary policy authorities in a DSGE framework. Taking the Smets and Wouters (2003) model for the euro area as the benchmark case, we look at the effects of the assumption that monetary policy is conducted optimally under commitment. Accordingly, we estimate the preferences in the monetary policy's objective function, which is assumed to take a standard quadratic form, using Bayesian estimation techniques. The optimizing monetary authorities are modeled in two alternative ways. The first model adopts a one-period loss function that includes inflation, the model-consistent output gap, and the one-period difference in the policy instrument. The second model includes an additional wage-inflation component. The estimation results for the structural parameters under these two model specifications are compared with the results of the benchmark Smets and Wouters (2003) model.

The estimate of the wage-inflation preference parameter in the monetary policy's loss function in our second alternative model does not turn out to be significant. This model performs relatively worse by yielding a lower marginal likelihood than our first alternative model with three target variables defining the loss function. The results for the first model suggest a significant value assigned to both the output gap and interest rate smoothing, in addition to the main inflation objective, in monetary policy's loss function.

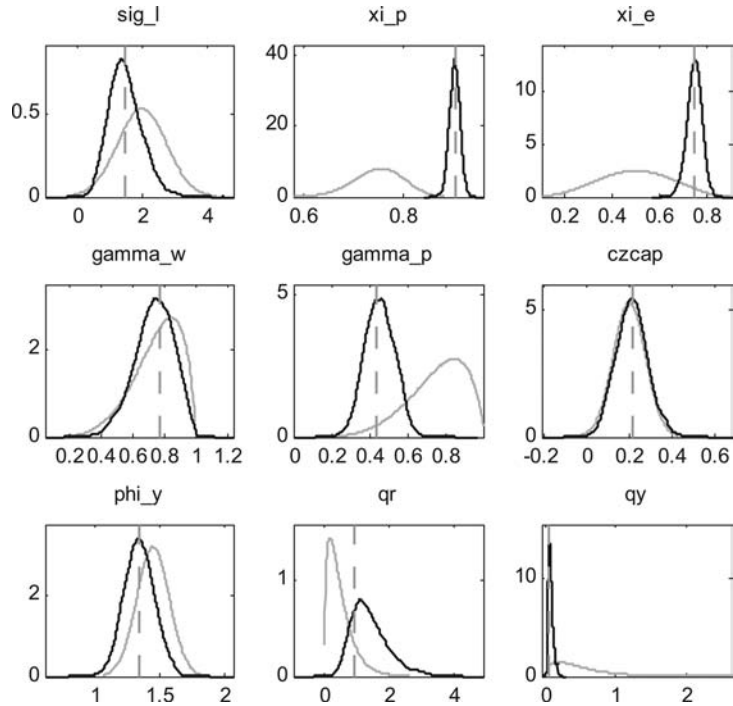
We show that by using an initialization sample in the estimations, we are able to approach the framework of optimal monetary policy under the timeless perspective of Woodford (2003) very closely.

There are a number of points on which we would like to improve and extend this work in the future. The results presented in this paper are conditional on the assumption that monetary policy performs optimally under commitment. Therefore, these results may

change and lead to different conclusions if one considers the alternative case of discretion. A simplifying and perhaps naive assumption throughout our estimations is that of a constant loss function over the sample period, implying stable preferences over the target variables. Moreover, we have restricted our analysis to a case where the inflation target is assumed to be a known constant. In Ilbas (forthcoming), these assumptions are relaxed and the out-of-sample forecasting performance of the model under the optimal monetary policy assumption is examined for the U.S. economy.

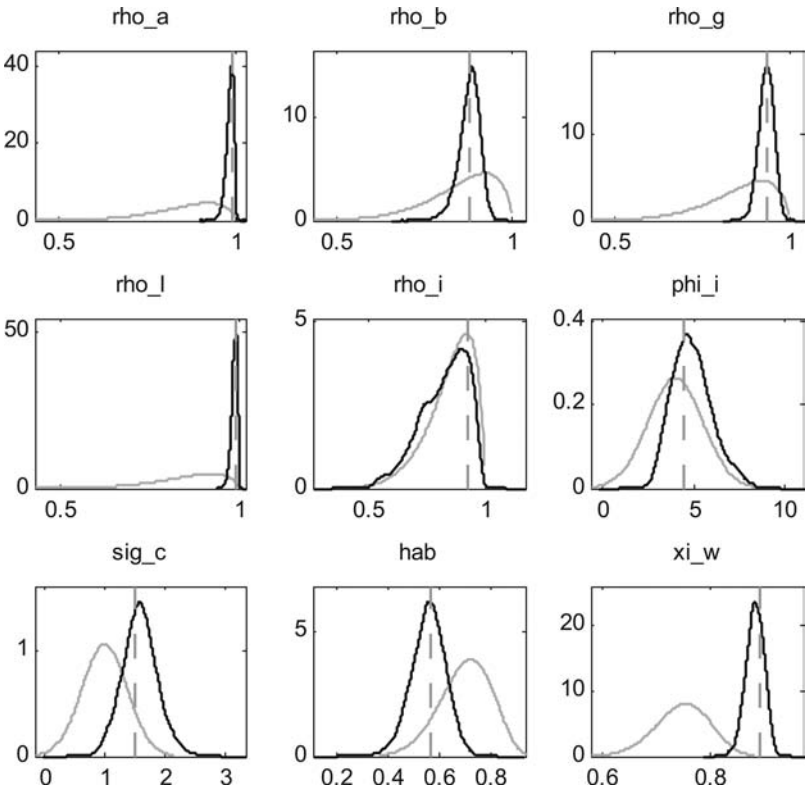
Appendix 1. Prior and Posterior Distributions of Parameters in Model *M1*

Figure 5. Posterior Distributions of the Parameters of *M1*



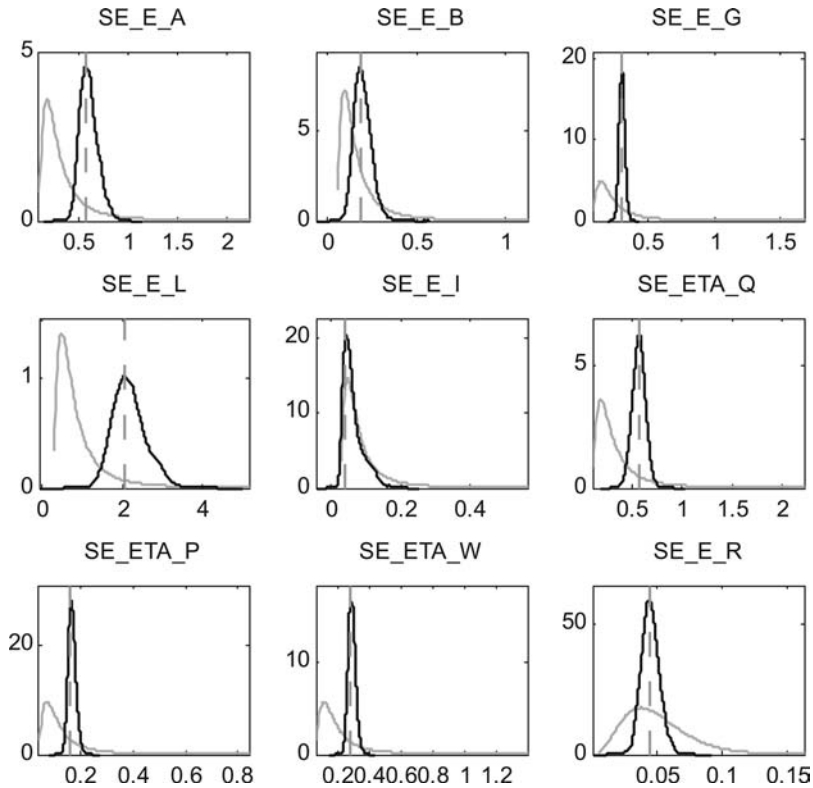
Notes: Prior distributions are indicated by shaded lines, and estimated posterior distributions are indicated by dark lines. The posterior mode is indicated by the vertical lines. The Metropolis-Hastings sampling algorithm is based on 100,000 draws using the sample 1980:Q2–2005:Q4.

Figure 6. Posterior Distributions of the Parameters of $M1$



Notes: Prior distributions are indicated by shaded lines, and estimated posterior distributions are indicated by dark lines. The posterior mode is indicated by the vertical lines. The Metropolis-Hastings sampling algorithm is based on 100,000 draws using the sample 1980:Q2–2005:Q4.

Figure 7. Posterior Distributions of the Parameters of $M1$



Notes: Prior distributions are indicated by shaded lines, and estimated posterior distributions are indicated by dark lines. The posterior mode is indicated by the vertical lines. The Metropolis-Hastings sampling algorithm is based on 100,000 draws using the sample 1980:Q2–2005:Q4.

Appendix 2. Use of a Pre-Sample and the Timeless-Perspective Approach

Due to our assumption of optimal monetary policy under commitment, we are dealing with a policy that is not time consistent. Woodford (2003) addresses this potential time-inconsistency problem by considering the concept of optimal policy from a timeless perspective (see also Svensson and Woodford 1999). In this paper, we perform the estimations by adopting a pre-sample period; i.e., the first forty

quarters (1970:Q2–1980:Q1) of our sample are treated as the initialization period preceding the estimations. This implies that although the Lagrange multipliers are set equal to zero at the start of the pre-sample (the initial period), at the start of the estimation the Lagrange multipliers will be derived from their past values. Hence, the use of a pre-sample has implications for the time consistency of monetary policy since it determines the way the lagged Lagrange multipliers are treated and the values they are assigned at the start of the estimation.

In order to test the effect of the initial values chosen for the Lagrange multipliers on the estimation results, we experiment with different initial values at the start of the initialization period (i.e., at time 1970:Q2) for the Lagrange multipliers of the forward-looking variables.²⁸ The estimation results are almost unaffected by changes in initial values thanks to a relatively long initialization period.²⁹ Therefore, we check for sensitivity of the estimation results to the length of the initialization period. For this purpose, we decrease the size of the pre-sample gradually below forty and reestimate the model *M1* twice, i.e., once by setting all initial values equal to zero and once by setting the Lagrange multipliers of forward-looking variables equal to non-zero, for a given pre-sample size. While decreasing the sample size, the estimation results start to show higher differences once the pre-sample reaches the size of around twenty quarters and below.³⁰ Therefore, given our pre-sample choice of forty quarters, our estimation results appear to be robust to the initial values of the Lagrange multipliers of forward-looking variables and the pre-sample length. Hence, our results hold in the context of optimal policy approached from a timeless perspective.³¹

²⁸Unfortunately, we do not know the exact initial values of the Lagrange multipliers of the forward-looking variables because these are not observable. Therefore, we experiment with arbitrary values.

²⁹We perform this experiment for model *M1* only. The reasoning that the optimal time-inconsistent policy approaches closely the timeless-perspective policy after a sufficiently long period following the initial period of the optimization is similar to the one followed by Juillard and Pelgrin (2007).

³⁰A detailed report of this sensitivity exercise is available on request.

³¹There is an alternative empirical application to the timeless-perspective solution through the pre-sample approach proposed in this paper. Juillard and Pelgrin (2007) show a two-step method to estimate policy from a timeless perspective in a Bayesian framework.

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Bank of England Interest Rate Announcements and the Foreign Exchange Market*

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Since 1997, the Bank of England Monetary Policy Committee (MPC) has met monthly to set the UK policy interest rate. Using a Markov-switching framework that incorporates endogenous transition probabilities, we examine intraday, five-minute return data for evidence of systematic patterns in exchange rate movements on MPC policy announcement days. We find evidence for non-linear regime switching between a high-volatility, informed trading state and a low-volatility, liquidity trading state. MPC surprise announcements are shown to significantly affect the probability that the market enters and remains within the informed trading regime, with some limited evidence of market positioning just prior to the announcement.

JEL Codes: E42, E44, F31.

1. Introduction

The Bank of England (BoE) was granted operational independence to set its key interest rate in May 1997, with the goal of implementing

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policy consistent with stable inflation and economic growth.¹ Interest rate decisions are made by the Bank's Monetary Policy Committee (MPC), which meets for two days each month—as well as an additional pre-meeting briefing day—and issues a statement regarding interest rate decisions at noon on the second meeting day. This framework allows a natural laboratory setting for examining the impact of monetary policy decisions around a known time and date. Since market participants know that interest rate announcements arrive at noon on the second meeting day, there may be positioning prior to the announcement and news effects after the announcement that result in systematic differences in the behavior of financial market variables on MPC meeting days compared with other, non-MPC days. In this paper, we concentrate on the pattern of exchange rate volatility surrounding the MPC's interest rate decisions as well as the role played by the surprise content of these announcements.

Although activities directly related to each MPC meeting are spread over three different days, our empirical analysis will focus upon the second MPC meeting day, when the policy decision is made and announced. We use high-frequency, intraday data and a Markov-switching econometric model where exchange rate returns switch between a high-volatility, informed trading state and a low-volatility, uninformed or liquidity trading state. This framework allows for a characterization of macroeconomic news effects on the foreign exchange market that differs from the traditional approach. Thus, we hypothesize that macroeconomic news does not simply affect the market as shocks to otherwise continuous processes, but instead may change, temporarily, the entire data-generating process of exchange rates. One reason is that “hot-potato” trades are likely to dominate market turnover to an unusual degree around news events as individual dealers adjust inventory and offload onto others, effectively generating a multiplier effect on trades (Lyons 1996).

An econometric specification that allows for regime switches therefore appears appropriate, particularly as it facilitates a plausible interpretation of observed non-linearities. Moreover, and in contrast to the deterministic models typically employed in similar

¹From the creation of the MPC until July 2006, policy decisions were framed in terms of the repurchase, or repo, rate and after that date in terms of the Bank rate. We use the two names interchangeably.

analyses, we allow for a probabilistic and thus flexible characterization of the data. In particular, by modeling switching probabilities endogenously, our approach allows the probability of regime switching to vary during MPC meeting days. Given the notoriously capricious nature of financial markets, our approach therefore provides an interesting alternative perspective on the impact of news effects on asset prices. This is the first important contribution of our research to the existing empirical literature. The second contribution is the size of our data set—28,556 high-frequency observations spanning ten years—which to the best of our knowledge is far longer than employed by any existing study and is important in ensuring that our results are robust.

Adopting this approach, we find evidence for non-linear regime switching between a high-volatility, informed trading state and a low-volatility, liquidity trading state. MPC surprise announcements are shown to significantly affect the probability that the market enters and remains within the informed trading regime, with some limited evidence of market positioning just prior to the policy announcement.

The next section provides a brief review of the literature on the impact of macroeconomic news announcements on financial markets. In section 3 we provide some background institutional details on the MPC and the UK monetary-policy-setting process. Section 4 contains a discussion of our econometric methodology and the various hypotheses to be tested. Section 5 describes our data sets and contains our main empirical findings. Finally, section 6 summarizes our conclusions and discusses directions for future research.

2. Exchange Rate and Asset-Price Effects of Monetary Policy Announcements: A Brief Review of the Literature

Early intraday studies of the impact of macroeconomic news effects on exchange rates—for instance, Hakkio and Pearce (1985) and Ito and Roley (1987)—report mixed results in terms of statistical significance. This may reflect the coarseness of sampling intervals, with observations of exchange rates taken at opening, noon, and closing. If news effects work themselves out within periods less than several hours, observing the market at three equally spaced points over the

trading day will miss much of the action. The increased availability of high-frequency, intraday foreign exchange rate data considerably advanced research in this area.

High-frequency, intraday exchange rate volatility effects of news announcements were first documented by Ederington and Lee (1993, 1995, 1996).² Ederington and Lee (1993) use five-minute tick data from November 1988 to November 1991 for mark-dollar, as well as various interest rate futures, and define their variable of interest as the deviation of the absolute value of exchange or interest rate returns in a given five-minute period on day j from the average return during that period across the whole sample. Ederington and Lee (1993) regress this variable on a series of dummy variables that designate the publication schedule of U.S. macroeconomic data series. They conclude in favor of a significant change in intraday exchange and interest rate volatility upon publication of various series, including the monthly employment report, producer price inflation, and trade data. They find that the standard deviation of five-minute returns immediately after publication is at least five times higher on announcement days than on non-announcement, or control, days. In addition, although the largest volatility impact occurs within one minute of publication, the standard deviation of returns remains significantly above normal for up to forty-five minutes after publication for a number of macroeconomic series.

In an extension to their original paper, Ederington and Lee (1995) perform a similar analysis using ten-second data, and conclude that the price reaction to macroeconomic news is largely completed after only forty seconds. They also find evidence of a significant change in volatility immediately ahead of key macroeconomic data releases, suggesting that market participants act to square positions in advance of key event risk. Ahn and Melvin (2007) also report evidence of switching to a high-volatility, informed trading state during U.S. Federal Reserve (Fed) policy meetings but prior to the announcement of decisions. An extensive search of public news suggests that this informed trading state cannot be explained as the response to public information, and instead is suggestive of informed

²Taylor (1987, 1989) provides early high-frequency studies of the foreign exchange market and finds some evidence of the impact of news on deviations from covered interest rate parity.

traders taking positions in advance of the meeting conclusion based upon their expectations of the outcome. This is a theme to which we return below.

A number of other papers have also found significant evidence of policy and macroeconomic news effects upon exchange rates, as well as other asset prices. Goodhart et al. (1993) report the most persistent impact upon exchange rate volatility—four to five days—in a GARCH-M analysis of U.S. monetary policy announcements and publication of U.S. trade data. Other studies find the persistence of news effects to be more fleeting, consistent with Ederington and Lee (1993). This includes Andersen and Bollerslev (1998) in the context of a wider study of the determinants of mark-dollar volatility, and Almeida, Goodhart, and Payne (1998), who find that the volatility impact of U.S. and German macroeconomic data releases generally dissipates within fifteen minutes of publication for U.S. data releases and within approximately three hours for German releases. In addition, Almeida, Goodhart, and Payne (1998) report that relatively few German data releases have a significant impact upon exchange rate volatility, although the number does increase when the authors account for the proximity of the next Bundesbank policy meeting; the closer the meeting, the more likely was the Bundesbank to react to data surprises. Andersen et al. (2003) similarly find that relatively few German data releases exert a statistically significant effect on exchange rates—in this case, the conditional mean. Their study also considers the impact of Federal Reserve policy announcements and various U.S. macroeconomic data series, and finds in favor of a significant, asymmetric jump effect associated with both types of news; interestingly, negative U.S. data surprises often exhibit a larger impact upon exchange rates than positive surprises.

Faust et al. (2003) use intraday, daily, and monthly data from 1994 to 2001 to estimate structural vector autoregressions (SVARs), incorporating current and future U.S. and foreign short-term interest rates, and exchange rates in order to assess the effect of U.S. monetary policy shocks on other variables in the SVARs. Although the results for interest rates are mixed, the impact of policy shocks upon exchange rates using intraday data is statistically significant. In a similar vein, Harvey and Huang (2002) examine the impact of Federal Reserve open-market operations on a range of interest and exchange rates using GMM estimation and both two-minute and

hourly returns, over the sample 1982 to 1988. They find in favor of a significant increase in intraday interest rate futures volatility associated with so-called Fed time, but against any significant, generalized increase in exchange rate return volatility.

3. The Monetary Policy Committee

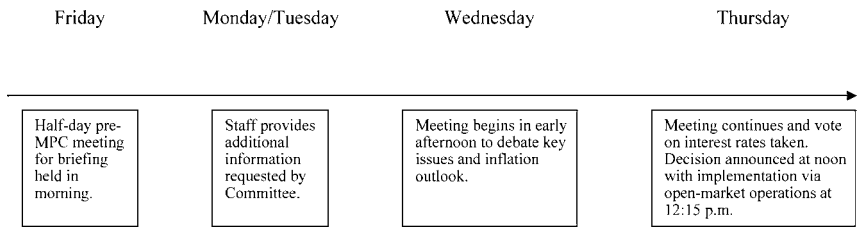
In May 1997, the UK Chancellor of the Exchequer announced that the BoE was to be given operational responsibility for setting interest rates via the newly created MPC. The MPC was to focus on ensuring that inflation was in line with the government-set target of 2.5 percent for the Retail Prices Index excluding mortgage interest payments “within a reasonable time period without creating undue instability in the economy.” Although not made explicit, this language was widely interpreted as indicating a policy horizon of two years. The policy goal was subsequently changed to 2.0 percent in December 2003, and is now defined in terms of the harmonized consumer price index.³ Conditional on achieving its inflation target, the MPC can also address fluctuations in economic growth and employment.

The MPC meets monthly, normally on the Wednesday and Thursday following the first Monday of each month. Meeting dates for each year are made available at the end of the previous year.⁴ The timetable for a representative meeting is given in figure 1. On the Friday morning prior to each meeting, the Committee receives a briefing from senior BoE staff on important news and data trends. The monthly MPC meeting typically begins at 15:00 on the following Wednesday afternoon (that is, the first meeting day) with a review of the state of the UK and world economy. The BoE Chief Economist starts the meeting with a short summary of any major events since the Friday briefing. On Thursday morning (the second meeting day), the MPC reconvenes and the Governor begins with a summary of the major issues. Members are then invited to state

³The UK government retains responsibility for establishing the goal of monetary policy. The inflation target is reconfirmed in the government’s annual budget statement. For institutional background on the MPC and the UK monetary policy process, see Bean (2001) and www.bankofengland.co.uk/monetarypolicy/framework.htm.

⁴These are published at www.bankofengland.co.uk.

Figure 1. Timeline for a Representative Monetary Policy Committee Meeting



their views on the appropriate policy action. The Deputy Governor responsible for monetary policy will usually speak first, with the Governor speaking last. Ultimately, the Governor offers a motion that he suspects will result in a majority and then calls for a vote, on the basis of a one-member, one-vote rule. Those in the minority are asked to state their preferred level of Bank rate. Lastly, the press statement is developed. If the decision is to change interest rates or follow a policy that was not expected by the market, the press statement will include the reasons for the action taken. In other cases, simply the decision is reported. This decision is announced at noon, London time, and policy is implemented with open-market operations beginning at 12:15 p.m. on the same day.

4. Methodology

The focus of this paper is on inference regarding movements in the dollar-sterling exchange rate around the time of the monthly MPC policy announcement, which occurs at noon on the second meeting day, as discussed above. As foreign exchange market participants know in advance when MPC decisions are announced, we examine five-minute dollar-sterling exchange rate returns for evidence of changes in market positioning during the meeting and whether such changes are driven by the news content of the policy announcement.

It is usual to think of high-frequency exchange rate data on any given day as bounded within a fairly narrow band and exhibiting first-order autocorrelation. By contrast, on MPC meeting days we may expect important news to be received by the market. We find it convincing to think of these news effects as changing,

temporarily, the entire data-generating process of exchange rates—and other financial variables—rather than simply introducing a one-time shock to an otherwise continuous process. Intuitively, so-called “hot-potato” trades are likely to dominate the market to an unusual degree in the immediate aftermath of the news as dealers adjust their inventory and offload onto other dealers, effectively generating a multiplier effect on trades (Lyons 1996).

An econometric specification allowing for regime switches is therefore appropriate. We adopt the Markov-switching framework associated with Hamilton (1990, 1994). An important advantage of this framework is that it facilitates a plausible interpretation of observed non-linearities and allows, in our application, for probabilistic rather than deterministic switching between regimes. A Markov-switching first-order autoregressive model can be written as

$$\begin{aligned}\Delta e_t &= \mu(S_t) + \rho(S_t)[\Delta e_{t-1} - \mu(S_{t-1})] + \varepsilon_t \\ \varepsilon_t &\sim N[0, \sigma^2(S_t)],\end{aligned}\tag{1}$$

where Δe_t is the change in the logarithm of the exchange rate at time t . The mean of the exchange rate returns process, μ , the autocorrelation coefficient, ρ , and the variance of the innovation, ε_t , are allowed to take on one of two values depending on the realization of an unobserved state variable $S_t \in \{1, 2\}$. In our application, we assume a two-state Markov process. One of the states (say, state 2) may be thought of as reflecting the usual pattern of exchange rate returns with negative autocorrelation and a relatively small variance. This tranquil state is associated with liquidity trading when no important information arrives in the market. The other state (state 1) may be thought of as the informed trading state when volatility is high and realized returns much larger than normal (Easley and O’Hara 1992; Lyons 2001).

Thus far, our proposed methodology is similar to that employed, *inter alia*, in Engel and Hamilton (1990). However, we diverge from the traditional Markov approach by modeling the probability of switching from one regime to another endogenously. Denoting the transition probability of switching from regime j to regime i at time t as P_t^{ij} for $i, j \in \{1, 2\}$, we can write the postulated functions for the transition probabilities, conditional upon information at time t , I_t , and the previous state, as

$$P_t^{ii} = \Pr[S_t = i | S_{t-1} = i, I_t] = \Phi[\alpha_{ii} + \beta'_{ii} X_t] \quad (2)$$

for $i \in \{1, 2\}$, where $\Phi[\]$ denotes the cumulative normal density function (in order to ensure that the probabilities lie in the unit interval) and where $X_t \in I_t$ is a vector of variables known at time t which may influence the transition probability according to the vector of loadings β_i . Given P_t^{11} , we implicitly have $P_t^{21} = 1 - P_t^{11}$; similarly, given an estimate of P_t^{22} , we implicitly have $P_t^{12} = 1 - P_t^{22}$.

The Markov-switching framework is applied to our data set to address several questions of interest. First, can we identify endogenous regime switching? Second, are the transition probabilities driven by the news component in the policy announcements? To test if the MPC policy announcement is price-relevant public news, we incorporate various dummy variables into the explanatory variable vector X_t . These dummies are set equal to one for a certain afternoon period on the second MPC meeting day, say noon to 13:00, and to zero otherwise. Third, is there evidence of positioning during the second meeting day prior to the noon policy announcement? To address this question, we incorporate dummy variables set equal to one for various time intervals prior to noon and zero otherwise.

5. Data and Empirical Findings⁵

Our data sample spans more than a decade, running from the inception of the MPC in June 1997 through October 2007, and incorporates 126 MPC meetings. Table 1 lists the MPC meeting days in our sample and the associated interest rate decisions. We classify an MPC decision as a surprise to the market if it differs from the median expectation taken from a Bloomberg survey of market economists.⁶ The standard deviation of analysts' expectations is reported as a measure of forecast dispersion. Table 1 also provides

⁵Unless otherwise stated, all references to MPC meeting days relate to second meeting days, when the policy announcement is made.

⁶This survey is carried out on the Friday before each MPC meeting and asks respondents for the magnitude—if any—of the interest rate change that they expect to result from the upcoming meeting. In its current guise, the survey collates the expectations of up to sixty financial economists. Although the sample of economists is not necessarily the same from one month to the next, a core subset ensures continuity.

Table 1. Monetary Policy Committee Meetings, Interest Rate Decisions, and Surprise Measures

The table contains interest rate decisions of the MPC for the period June 1997–October 2007. “Bloomberg Expectation” refers to the interest rate change predicted by the median expectation in a Bloomberg survey of market economists. “Forecast Dispersion” is the standard deviation calculated from individual analysts’ forecasts. The final five columns in the table indicate whether the interest rate announcement surprised the market according to the respective measure. The variable “IB10” (“IB15”) indicates whether the change in the three-month interbank rate from one day before the announcement to one day after the announcement was greater or equal to 10 (15) basis points. “LIFFE10” and “LIFFE15” are defined similarly but based on sterling three-month interest rate futures contracts.

Date	Interest Rate Decision	Bloomberg Expectation	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	LIFFE10 Surprise	LIFFE15 Surprise
6-Jun-97	6.5	6.25	Missing	Yes	No	No	No	No
10-Jul-97	6.75	6.75	Missing	No	No	No	No	No
7-Aug-97	7	6.75	Missing	Yes	No	No	No	No
11-Sep-97	7	7	Missing	No	Yes	No	No	No
9-Oct-97	7	7	Missing	No	No	No	No	No
6-Nov-97	7.25	7	Missing	Yes	No	No	Yes	Yes
4-Dec-97	7.25	7.25	Missing	No	No	No	No	No
8-Jan-98	7.25	7.25	Missing	No	No	No	No	No
5-Feb-98	7.25	7.25	Missing	No	No	No	No	No
5-Mar-98	7.25	7.25	Missing	No	No	No	No	No
9-Apr-98	7.25	7.25	Missing	No	Yes	No	No	No
7-May-98	7.25	7.25	Missing	No	No	No	No	No
4-Jun-98	7.5	7.25	Missing	Yes	Yes	Yes	Yes	Yes
9-Jul-98	7.5	7.5	Missing	No	Yes	No	Yes	No
6-Aug-98	7.5	7.5	Missing	No	No	No	No	No
10-Sep-98	7.5	7.5	Missing	No	Yes	No	No	No

(continued)

Table 1. (Continued)

Date	Interest Rate Decision	Bloomberg Expectation	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	LIFFE10 Surprise	LIFFE15 Surprise
8-Oct-98	7.25	7.5	0.127	Yes	Yes	Yes	Yes	Yes
5-Nov-98	6.75	7	0.059	Yes	Yes	No	No	No
10-Dec-98	6.25	6.5	0.145	Yes	No	No	No	No
7-Jan-99	6	6.25	0.000	Yes	Yes	Yes	No	No
4-Feb-99	5.5	5.75	0.128	Yes	Yes	Yes	Yes	Yes
3-Mar-99	5.5	5.25	0.128	No	No	No	Yes	No
8-Apr-99	5.25	5.25	0.079	No	No	No	No	No
6-May-99	5.25	5.25	0.107	No	No	No	No	No
10-Jun-99	5	5.25	0.109	Yes	Yes	No	Yes	No
8-Jul-99	5	5	0.000	No	No	No	No	No
5-Aug-99	5	5	0.000	No	No	No	No	No
8-Sep-99	5.25	5	0.000	Yes	Yes	Yes	Yes	Yes
7-Oct-99	5.25	5.25	0.101	No	Yes	No	Yes	No
4-Nov-99	5.5	5.5	0.106	No	Yes	Yes	Yes	Yes
9-Dec-99	5.5	5.5	0.056	No	No	No	No	No
13-Jan-00	5.75	5.75	Missing	No	No	No	No	No
10-Feb-00	6	6	0.074	No	No	No	No	No
9-Mar-00	6	6	0.086	No	No	No	No	No
6-Apr-00	6	6	0.127	No	No	No	No	No
4-May-00	6	6.25	0.000	Yes	No	No	No	No
7-Jun-00	6	6	0.072	No	No	No	No	No
6-Jul-00	6	6	0.115	No	No	No	No	No

(continued)

Table 1. (Continued)

Date	Interest Rate Decision	Bloomberg Expectation	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	LIFFE10 Surprise	LIFFE15 Surprise
3-Aug-00	6	6	0.101	No	No	No	No	No
7-Sep-00	6	6	0.067	No	No	No	No	No
5-Oct-00	6	6	0.000	No	No	No	No	No
9-Nov-00	6	6	0.000	No	No	No	No	No
7-Dec-00	6	6	0.394	No	No	No	No	No
11-Jan-01	6	6	0.083	No	No	No	No	No
8-Feb-01	5.75	5.75	0.067	No	No	No	No	No
8-Mar-01	5.75	5.75	0.125	No	No	No	No	No
5-Apr-01	5.5	5.5	0.049	No	No	No	No	No
10-May-01	5.25	5.25	0.000	No	No	No	No	No
6-Jun-01	5.25	5.25	Missing	No	No	No	No	No
5-Jul-01	5.25	5.25	0.000	No	No	No	Yes	No
2-Aug-01	5	5.25	0.000	Yes	Yes	Yes	Yes	Yes
6-Sep-01	5	5	0.131	No	No	No	No	No
18-Sep-01	4.75	Missing	0.111	Yes	Yes	No	No	No
4-Oct-01	4.5	4.5	0.096	No	No	No	No	No
8-Nov-01	4	4.25	0.000	Yes	Yes	Yes	Yes	No
5-Dec-01	4	4	0.000	No	No	No	Yes	No
10-Jan-02	4	4	0.069	No	No	No	No	No
7-Feb-02	4	4	0.000	No	No	No	No	No
7-Mar-02	4	4	0.000	No	No	No	No	No
4-Apr-02	4	4	0.000	No	No	No	No	No

(continued)

Table 1. (Continued)

Date	Interest Rate Decision	Bloomberg Expectation	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	LIFFE10 Surprise	LIFFE15 Surprise
9-May-02	4	4	0.073	No	No	No	No	No
6-Jun-02	4	4	0.045	No	No	No	No	No
4-Jul-02	4	4	0.000	No	No	No	No	No
1-Aug-02	4	4	0.000	No	No	No	No	No
5-Sep-02	4	4	0.071	No	No	No	No	No
10-Oct-02	4	4	0.127	No	No	No	Yes	No
7-Nov-02	4	4	0.000	No	No	No	No	No
5-Dec-02	4	4	0.000	No	No	No	No	No
9-Jan-03	4	4	0.046	No	No	No	No	No
6-Feb-03	3.75	4	0.000	Yes	Yes	Yes	Yes	Yes
6-Mar-03	3.75	3.75	0.080	No	No	No	No	No
10-Apr-03	3.75	3.75	0.126	No	No	No	No	No
8-May-03	3.75	3.75	0.097	No	No	No	No	No
5-Jun-03	3.75	3.75	0.107	No	No	No	No	No
10-Jul-03	3.5	3.75	0.000	Yes	Yes	Yes	No	No
7-Aug-03	3.5	3.5	0.000	No	No	No	No	No
4-Sep-03	3.5	3.5	0.000	No	No	No	No	No
9-Oct-03	3.5	3.5	0.037	No	Yes	Yes	No	No
6-Nov-03	3.75	3.75	0.039	No	No	No	No	No
9-Dec-03	3.75	3.75	0.041	No	No	No	No	No
8-Jan-04	3.75	3.75	0.041	No	No	No	No	No
5-Feb-04	4	4	0.000	No	No	No	No	No

(continued)

Table 1. (Continued)

Date	Interest Rate Decision	Bloomberg Expectation	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	LIFFE10 Surprise	LIFFE15 Surprise
4-Mar-04	4	4	0.124	No	No	No	No	No
8-Apr-04	4	4	0.053	No	No	No	No	No
6-May-04	4.25	4.25	0.126	No	No	No	No	No
10-Jun-04	4.5	4.5	0.055	No	No	No	No	No
8-Jul-04	4.5	4.5	0.000	No	No	No	No	No
5-Aug-04	4.75	4.75	0.038	No	No	No	No	No
9-Sep-04	4.75	4.75	0.000	No	No	No	No	No
7-Oct-04	4.75	4.75	0.053	No	No	No	No	No
4-Nov-04	4.75	4.75	0.000	No	No	No	No	No
9-Dec-04	4.75	4.75	0.000	No	No	No	No	No
12-Jan-05	4.75	4.75	0.000	No	No	No	No	No
10-Feb-05	4.75	4.75	0.000	No	No	No	No	No
10-Mar-05	4.75	4.75	0.000	No	No	No	No	No
7-Apr-05	4.75	4.75	0.042	No	No	No	No	No
9-May-05	4.75	4.75	0.000	No	No	No	No	No
9-Jun-05	4.75	4.75	0.065	No	No	No	No	No
7-Jul-05	4.75	4.75	0.084	No	No	No	Yes	No
4-Aug-05	4.5	4.5	0.000	No	No	No	No	No
8-Sep-05	4.5	4.5	0.000	No	No	No	No	No
6-Oct-05	4.5	4.5	0.000	No	No	No	No	No
10-Nov-05	4.5	4.5	0.000	No	No	No	No	No
8-Dec-05	4.5	4.5	0.000	No	No	No	No	No

(continued)

Table 1. (Continued)

[illegible]

two alternative surprise measures—based on three-month interbank interest rates (IB) and three-month sterling interest rate futures contracts traded on the London International Financial Futures Exchange (LIFFE)—to be used for robustness checks.⁷

Table 1 suggests that the Bank of England has succeeded in achieving its goal of improving monetary policy transparency (King 2000). All measures of market expectations show a clear trend lower in the frequency of policy surprises. The Bank rate was changed on thirty-six occasions during our sample: raised at nineteen meetings and lowered at seventeen meetings. One-half of these instances were fully expected by the market, as measured by the Bloomberg survey. For the other eighteen instances, the market was either surprised that the MPC changed the policy rate or was surprised by the extent of the change. There were no instances where the market expected a change in the policy rate in the opposite direction to the change actually announced, although in May 2000 the market expected a change whereas the MPC kept the Bank rate constant. Overall, therefore, we observe nineteen policy surprises during our sample according to the Bloomberg survey.⁸ Consequently, we separate our sample of 126 second MPC meeting days into two parts: (i) the 107 second meeting days when the change in the policy rate—including a change of 0 basis points—was in line with expectations (we term these “No Unexpected Change” days) and (ii) the nineteen days when the announced rate changed by an amount different to market expectations (we term these “Unexpected Change” days).

Tick data for the dollar-sterling exchange rate were obtained from a major international bank for each of our 126 MPC meeting days and a set of 126 control days, defined as the same day of the week as the MPC meeting seven days later. Insufficient exchange rate data were available for 14 out of the total of 252 days.⁹

⁷The period t policy announcement is classified as a surprise to the market if the difference between the rate in periods $t + 1$ (interbank or LIFFE) and $t - 1$ is greater than 10 (15) basis points, where t is the second MPC meeting day and the interest rate data are sampled daily.

⁸According to “IB10” (“IB15,” “LIFFE10,” “LIFFE15”), twenty-one (thirteen, nineteen, ten) policy surprises can be identified during the sample period.

⁹We also do not include the extraordinary, unscheduled meeting of September 18, 2001, and the respective control day on September 25, 2001.

We sample the last quotation of each five-minute interval over the hours 7:00–17:00 London time to create a series of exchange rate returns, defined as the change in the logarithm of the five-minute observations multiplied by 10,000.¹⁰ By way of example, the 12:05 observation on any given day is the last quotation from within the interval 12:00–12:05. The data for each day are stacked in serial order to create a data set with 28,556 observations.

The Markov model represented by the set of equations (1) above is used to estimate the effect of MPC announcements on the transition probabilities. Estimation of the model is carried out using a modified version of the EM algorithm due to Diebold, Lee, and Weinbach (1994). The two states are identified by significant shifts in the mean μ , the autocorrelation coefficient ρ , and the variance σ^2 . Recall that state 1 is the high-variance state associated with information-based trading and state 2 is the low-variance state associated with the normal market conditions of liquidity trading. Consistent with these definitions, the results in table 2 show that the estimated state 1 variance, $\sigma^2(1)$, is generally found to be about 3.5 times that of state 2, $\sigma^2(2)$. Statistically significant negative first-order autocorrelation, a common feature of high-frequency exchange rate returns, is also apparent.

In table 2, panel A we report estimates of the (restricted) constant transition probability model, and in panel B the (unrestricted and preferred) time-varying transition probability model. The reported likelihood-ratio statistic is statistically significant and justifies our decision to estimate the transition probabilities endogenously.¹¹ The transition probabilities estimated by the restricted model are reported in table 2; P^{11} is the probability of remaining in

¹⁰Danielsson and Payne (2002) compare one week of indicative quote data with firm quotes from an electronic FX brokerage and find that the properties of returns for each series become quite similar at a five-minute sampling frequency. At higher frequencies, the indicative quotes tend to lag firm quotes. We therefore choose five-minute sampling to ensure that our exchange rate returns are representative of market conditions. The raw data were referenced to Greenwich Mean Time (GMT), so time references were appropriately adjusted to account for British Summer Time.

¹¹From the log-likelihood values reported in table 2, this statistic is $-2(-73022 + 71082) = 3880$ (p-value = 0.00). Note that the means are not significantly different from zero in the specification with constant transition probabilities. This is the only specification for which this is the case.

Table 2. Markov-Switching Model of MPC News and News Anticipation Effects

The table reports estimates of a Markov-switching model for dollar-sterling exchange rate returns sampled at a frequency of five minutes over the London business day. The final, preferred specification was a first-order autoregression in returns with a regime-switching variance, a regime-switching mean, and a regime-switching autocorrelation coefficient. Transition probabilities are modeled as switching endogenously as a function of MPC-related events as in the following equation (where Φ denotes the cumulative normal density function):

$$P^{11} = \Phi(\alpha_{11} + \sum_k \beta_{11,k} dum_k) \text{ and } P^{22} = \Phi(\alpha_{22} + \sum_k \beta_{22,k} dum_k).$$

A. Constant Transition Probability Model						
Coefficient Estimates (p-values)						
$\rho(1)$	-0.13 (0.000)	$\sigma^2(1)$	5.51 (0.000)			
$\rho(2)$	-0.15 (0.000)	$\sigma^2(2)$	1.61 (0.000)			
Constant						
P^{11}	1.68 (0.000)					
P^{22}	1.74 (0.000)					
LogL = -73022						
B. Preferred Time-Varying Transition Probability Model						
B1. Surprise Measure Based on Bloomberg Survey of Market Economists						
Coefficient Estimates (p-values) of Regime-Switching AR(1) Model						
$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)	
$\mu(2)$	-0.28 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)	
Coefficient Estimates (p-values) for Endogenous Transition Probabilities						
	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day Time		Unexp. Ch. 12:05-13:00	All MPC 12:05-13:00	All Days 12:05-13:45	All MPC 11:30-11:55	All MPC 11:15-11:55
P^{11}	0.74 (0.000)	0.96 (0.000)	0.31 (0.000)	-0.01 (0.000)	0.31 (0.019)	-0.87 (0.002)
P^{22}	7.15 (0.000)			-2.34 (0.000)		
LogL = -71082						

(continued)

Table 2. (Continued)

B2. Surprise if Change in Three-Month Interbank Rate Is Greater Than 10 Basis Points						
Coefficient Estimates (p-values) of Regime-Switching AR(1) Model						
$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)	
$\mu(2)$	-0.29 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)	
Coefficient Estimates (p-values) for Endogenous Transition Probabilities						
	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day Time		Unexp. Ch. 12:05–13:00	All MPC 12:05–13:00	All Days 12:05–13:45	All MPC 11:30–11:55	All MPC 11:15–11:55
P^{11}	0.74 (0.000)	0.86 (0.000)	0.33 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.003)
LogL = -71084						
B3. Surprise if Change in Three-Month Interbank Rate Is Greater Than 15 Basis Points						
Coefficient Estimates (p-values) of Regime-Switching AR(1) Model						
$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)	
$\mu(2)$	-0.28 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)	
Coefficient Estimates (p-values) for Endogenous Transition Probabilities						
	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day Time		Unexp. Ch. 12:05–13:00	All MPC 12:05–13:00	All Days 12:05–13:45	All MPC 11:30–11:55	All MPC 11:15–11:55
P^{11}	0.74 (0.000)	0.51 (0.061)	0.43 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.14 (0.000)			-2.34 (0.000)		-0.86 (0.001)
LogL = -71082						

(continued)

Table 2. (Continued)

B4. Surprise if Change in Price of Three-Month Interest Rate Futures Is Greater Than 10 Basis Points						
Coefficient Estimates (p-values) of Regime-Switching AR(1) Model						
$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)	
$\mu(2)$	-0.28 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)	
Coefficient Estimates (p-values) for Endogenous Transition Probabilities						
	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day Time		Unexp. Ch. 12:05-13:00	All MPC 12:05-13:00	All Days 12:05-13:45	All MPC 11:30-11:55	All MPC 11:15-11:55
P^{11}	0.74 (0.000)	0.62 (0.011)	0.39 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.000)
LogL = -71082						
B5. Surprise if Change in Three-Month Interest Rate Futures Is Greater Than 15 Basis Points						
Coefficient Estimates (p-values) of Regime-Switching AR(1) Model						
$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)	
$\mu(2)$	-0.28 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)	
Coefficient Estimates (p-values) for Endogenous Transition Probabilities						
	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day Time		Unexp. Ch. 12:05-13:00	All MPC 12:05-13:00	All Days 12:05-13:45	All MPC 11:30-11:55	All MPC 11:15-11:55
P^{11}	0.74 (0.000)	0.67 (0.048)	0.44 (0.000)	-0.01 (0.000)	0.30 (0.020)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.001)

the high-volatility state and P^{22} is the probability of remaining in the low-volatility state. Normally, we would expect $P^{22} > P^{11}$ and, from table 2, this is what the data reveal. Estimating a Markov-switching model with fixed transition probabilities results in the following estimates: $P^{11} = \Phi(1.68) = 0.95$ and $P^{22} = \Phi(1.74) = 0.96$. The unconditional probability of being in state 2 associated with these transition probabilities is given as $\frac{(1-P^{11})}{(1-P^{11})+(1-P^{22})} = 0.556$, and the unconditional probability of being in state 1 is 0.444.

Moving to the time-varying probabilities model (table 2, panel B), it is first interesting to find that the mean return is significantly positive in the high-volatility state and significantly negative in the low-volatility state. This result might appear puzzling, as it suggests that return volatility is high when the pound appreciates against the dollar and low when it depreciates. Consequently, it is interesting to investigate whether this result is driven by MPC announcements or is simply an artifact of the data in the sample period considered.

To consider this issue, we include additional intercept terms into the mean equation of our preferred specification:

$$\begin{aligned} \Delta e_t = & \mu(S_t) + \mu_{pos}(S_t)I(\Delta i > 0) + \mu_{neg}(S_t)I(\Delta i < 0) \\ & + \rho(S_t)[\Delta e_{t-1} - \mu(S_{t-1})] + \varepsilon_t, \end{aligned} \quad (3)$$

where μ_{pos} and μ_{neg} are additional constant terms, $I(\Delta i > 0)$ is an indicator function that takes the value of one on Unexpected Change days between 12:05 and 13:00 if the announced interest rate is higher than expected, and $I(\Delta i < 0)$ is an indicator function that takes the value of one on Unexpected Change days between 12:05 and 13:00 if the announced interest rate is lower than expected.¹²

As table 3 reveals, the coefficients on μ_{pos} are significantly positive in both states and the coefficients on μ_{neg} significantly negative. These results indicate, as expected, that a higher UK policy rate than expected yields a positive mean return to holding sterling during the main impact period of the announcement, implying that the pound appreciates. A lower policy rate than expected yields a negative mean return during the main impact period of

¹²We select this definition for the indicator function, as our analysis below will demonstrate that the impact of interest rate announcements on the market is largest between 12:05 and 13:00 on Unexpected Change days.

the announcement, implying that the pound depreciates. Despite inclusion of these control dummies, however, none of our previous results are changed in any significant way. This indicates that our original finding is a general characteristic of the data during our sample, rather than a function of surprise policy announcements. Periods when the pound appreciates against the dollar tend to be more volatile than periods when the pound depreciates.

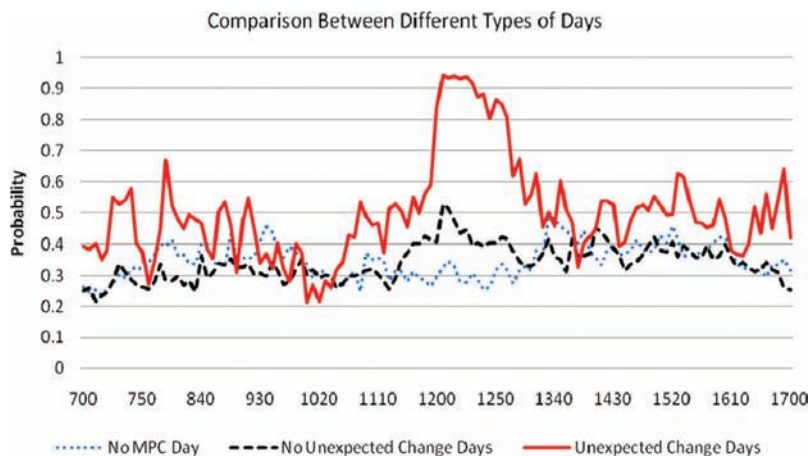
Returning to the core focus of our investigation, and table 2, panel B, transition probabilities in our preferred specification are modeled as varying with dummy variables that equal one at designated times of day and zero otherwise. Preliminary estimates suggest that the preferred model expresses P^{11} as a function of a constant, a dummy equal to one from 12:05 to 13:00 only on Unexpected Change days, a dummy equal to one from 12:05 to 13:00 on All MPC days, a dummy equal to one on All days between 12:05 and 13:45, and a dummy equal to one between 11:30 and 11:55 on All MPC days.¹³ P^{22} is a function of a constant, a dummy equal to one on All days between 12:05 and 13:45, and a dummy equal to one on All MPC meeting days from 11:15 to 11:55. All estimated coefficients differ significantly from zero with p-values of 0.02 or lower.

The results indicate that the probability of remaining in the informed trading state, P^{11} , is significantly higher from 12:05 to 13:00 on Unexpected Change days. This can be seen in figure 2, which plots the smoothed unconditional probability of being in state 1 for the three types of days in our sample, as generated by the preferred model reported in table 2. This probability is averaged across all observations for each type of day and for each five-minute interval.

As figure 2 highlights, there are dramatic differences across types of days and time of day. Based upon our preferred model specification, non-MPC meeting days are characterized by low-volatility, liquidity trading, as the probability of remaining in the informed trading state (that is, state 1) remains low throughout the day, fluctuating between 0.25 and 0.45. On No Unexpected Change days, the average unconditional probability of being in state 1 begins to rise just before 11:30 until 12:05, when it jumps to about 0.53. After

¹³Notice that the observation for 12:05 on any given day is the last quotation within the interval 12:00–12:05, i.e., the first observation in our data set after the interest rate announcement.

Figure 2. Smoothed Unconditional Probability of Informed Trading State



this peak, the probability quickly falls to 0.40 by 12:30, and then by 13:00 is quite similar to the afternoon pattern on non-MPC days.

On Unexpected Change days, however, there is a dramatic jump in the probability of being in state 1 at noon when the policy announcement is released, from about 0.55 to more than 0.90. The probability of being in the informed trading state remains above 0.70 until about 13:00. Thereafter it falls back and, by about 13:30, follows a pattern much like other days.

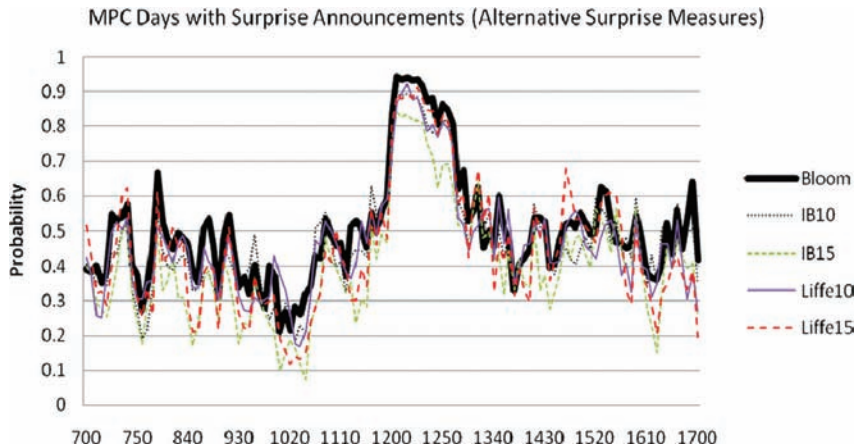
The probability of remaining in the tranquil state, P^{22} , falls between 11:15 and 11:55 on All MPC days and between 12:05 and 13:45 on All days. Although these changes in P^{22} are statistically significant, one may argue that they are not economically significant.

It is important to verify that our findings thus far are robust, and accordingly we now report and discuss a series of checks to these results.

5.1 Alternative Measures of Market Expectations

Thus far, we have measured market expectations using the Bloomberg survey of economists. Table 2, panels B2–B5 report estimated results for our preferred model above, albeit replacing

Figure 3. Smoothed Unconditional Probability of Informed Trading State



the Bloomberg measure with the alternative measures of market expectations based upon interest rates and futures, as detailed above. For each of the alternative surprise measures, the smoothed unconditional probability of being in the informed trading state on Unexpected Change days is plotted in figure 3. The similarity of the results in tables 2 and 3, and of figures 2 and 3, is striking and suggests that our findings are robust to changing the precise characterization of policy surprises. Accordingly, for the remainder of our analysis we concentrate on the Bloomberg survey surprise measure.

5.2 Alternative Dummy Variables

Results reported above concentrate on Unexpected Change days. In table 4, we augment our preferred model with an additional explanatory dummy variable, defined according to the same time divisions as previously, but also for No Unexpected Change days and All days. For instance, the dummy “Additional 1” is equal to one from 12:05 to 13:00 on All days. From table 4, none of the added dummy variables is statistically significant, indicating once more that our initial results are robust.

Table 4. Markov-Switching Model Estimated over Different Types of Days

The table reports estimates from the preferred specification (table 2, panel B) augmented to include dummy variables for different types of days (Unexpected Change days, No Unexpected Change days, and All days).

	Coefficient Estimates (p-values) for Additional Variables						
	Additional 1	Additional 2	Additional 3	Additional 4	Additional 5	Additional 6	Additional 7
Type of Day	All Days 12:05–13:00	All MPC 12:05–13:45	Unexp. Ch. 12:05–13:45	All Days 11:30–11:55	Unexp. Ch. 11:30–11:45	All Days 11:15–11:55	Unexp. Ch. 11:15–11:55
P^{11}	–0.12 (0.260)	–0.16 (0.120)	0.09 (0.69)	0.02 (0.878)	0.37 (0.560)		
P^{22}		–0.07 (0.782)	0.34 (0.688)			0.12 (0.770)	0.05 (0.964)

5.3 *Concurrent Central Bank Actions*

One difficulty with the analysis of macroeconomic news impacts upon exchange rates, or asset prices, is to ensure that the estimated intraday state probabilities truly reflect the impact of MPC policy announcements, rather than the effect of other news or shocks. One potential omitted variable in our analysis thus far is the announcement calendar of other central banks. In particular, there are twenty-eight meetings in our sample where MPC announcement days coincide with policy announcements by the Governing Council of the European Central Bank (ECB). Announcements by the ECB occur at 12:45 GMT, which coincides with the reported significant increase in the probability of remaining in the informed trading state P^{11} following MPC announcements. To test whether significant volatility shifts in dollar-sterling returns partly reflect a response to ECB interest rate decisions, we include a set of dummy variables to proxy for these announcements. This analysis is reported in table 5. Dummies take the value of one for time periods starting at 12:45 GMT on (i) all days on which MPC and ECB policy announcements coincide, (ii) those coincident days on which the ECB announced an interest rate change, or (iii) only those coincident days which involved an ECB policy surprise.¹⁴ As table 5 reports, only the dummy covering the period 12:45–17:00 on All coinciding days is significant when included in the specification for P^{11} .¹⁵ But our initial results are not altered substantively by the inclusion of ECB dummy variables.

Overall, we can conclude that the evidence in tables 2–5 represents a robust set of results. We have demonstrated significant evidence of a systematic regime switch to a high-volatility, informed trading state on MPC Unexpected Change days. This effect is significant for about an hour following the interest rate announcement, before the probability of remaining in the informed trading state

¹⁴We define ECB surprise announcement days according to changes in the short-term interbank rate (EURIBOR) using the same approach as for the BoE in table 1.

¹⁵A priori, one would expect the coefficient on this dummy variable to have a positive sign, indicating that ECB policy announcements raise the probability of higher dollar-sterling return volatility. The observed negative sign may indicate, then, that this dummy is capturing something other than the volatility impact of ECB policy announcements.

Table 5. Concurrent Central Bank Announcements

The table reports estimates from the preferred specification (table 2, panel B) augmented to include time dummy variables to capture days on which MPC and ECB policy announcements coincide. There are three types of coincident days: all days on which the MPC and ECB policy announcements coincide, those coincident days on which the ECB announced an interest rate change, and those coincident days on which the ECB policy announcement was unexpected by the median market expectation.

Equation	Day	Time	Coefficient (p-value)	Log Likelihood
P^{11}	ECB Coincides	12:45–14:00	−0.12 (0.255)	−71082
P^{11}	ECB Rate Change	12:45–14:00	−0.09 (0.636)	−71082
P^{11}	ECB Surprise Change	12:45–14:00	−0.54 (0.151)	−71081
P^{22}	ECB Coincides	12:45–14:00	−0.004 (0.992)	−71082
P^{22}	ECB Rate Change	12:45–14:00	−0.657 (0.402)	−71082
P^{22}	ECB Surprise Change	12:45–14:00	0.35 (0.777)	−71082
P^{11}	ECB Coincides	12:45–17:00	−0.21 (0.02)	−71078
P^{11}	ECB Rate Change	12:45–17:00	−0.06 (0.630)	−71082
P^{11}	ECB Surprise Change	12:45–17:00	−0.41 (0.056)	−71080
P^{22}	ECB Coincides	12:45–17:00	0.30 (0.182)	−71081
P^{22}	ECB Rate Change	12:45–17:00	−0.29 (0.473)	−71082
P^{22}	ECB Surprise Change	12:45–17:00	0.84 (0.330)	−71082

falls back to pre-announcement levels. The pattern of return volatility is clearly distinguished from other days in our sample and is not simply a “time-of-day” effect that is always present, regardless of policy announcements. These results are also robust to the inclusion of various alternative time-of-day dummies that account for concurrent actions of the ECB.

Consequently, in response to the first two questions posed above (Can we identify endogenous regime switching? Are the transition probabilities driven by the news component in the policy announcements?), we can answer with a strong affirmation.

5.4 Pre-Announcement Positioning Effects?

We now turn to the third question to be addressed: Is there evidence of positioning during MPC meetings prior to the policy announcement at noon on the second meeting day?

We capture this news anticipation effect in our preferred model (table 2, panel B) by the coefficients on the dummy variable for 11:30–11:55 on All MPC days in the P^{11} equation and the dummy variable for 11:15–11:55 on All MPC days in the P^{22} equation. Both dummy variables are significant, and the coefficients imply an increased probability of remaining in the informed trading state from 11:30 to 11:55 and a decreased probability of remaining in the normal liquidity trading state from 11:15 to 11:55.

Panels A, B, and C of table 6 incorporate alternative morning dummy variables into the preferred model as a further robustness check. This check proceeds much like the analysis associated with the post-announcement effect above for three different types of days: All days, All MPC days, and Unexpected Change days. For instance, the first row of table 6, panel A includes a dummy equal to one from 11:45 to 11:55 on All days in the P^{11} equation. None of the dummies for either All days or All MPC days is significant, and our preferred model results are not altered by the inclusion of these variables.

By contrast, in table 6, panel C, dummies for the periods 11:45–11:55, 11:30–11:55, and 11:00–11:55 on Unexpected Change days enter the P^{11} equation with significant p-values and positive coefficients. The dummy for 9:00–11:55 in the P^{22} equation is borderline significant as well and has a negative coefficient. This suggests that the probability of being in the high-volatility state increases during the morning of Unexpected Change days, whereas it does not on days when the announcement is anticipated. In addition, the results of our preferred model specification are once again not substantively changed by the inclusion of these variables.

Taken as a whole, therefore, there is evidence that exchange rate volatility switches between low and high regimes during the morning prior to the end of MPC meetings and, in particular, prior to those meetings that eventually deliver an unexpected change in the Bank rate. The evidence is strongest for the P^{11} equation for the 11:30–11:55 time period. Because the meetings always end prior to the noon announcement and the MPC's policy decision is known by insiders, this regime switching could be a result of signals read by market participants. This is not to claim that there are deliberate information leaks emanating from the Committee; it may be something much more subtle. Alternatively, it may be that traders simply close down positions during the morning in order to limit their

**Table 6. Markov-Switching Model of MPC News
Anticipation Effects**

The table reports estimates and tests of alternative specifications of the Markov-switching model for dollar-sterling exchange rate returns, using the model reported in table 2, panel B as the baseline model. Only the estimated coefficients (p-values) for the additional dummy in the transition probability equations are reported here.

A. Alternative Specifications (Morning Dummies for All Days) Additional Variables Added to the Preferred Model Specification			
Equation	Variable	Coefficient (p-value)	Log Likelihood
P^{11}	11:45–11:55	− 0.169 (0.300)	−71082
P^{11}	11:30–11:55	−0.08 (0.564)	−71082
P^{11}	11:15–11:55	−0.07 (0.500)	−71082
P^{11}	11:00–11:55	−0.000 (0.999)	−71082
P^{11}	9:00–11:55	−0.000 (0.999)	−71082
P^{22}	11:45–11:55	0.10 (0.862)	−71082
P^{22}	11:30–11:55	0.32 (0.459)	−71082
P^{22}	11:15–11:55	0.12 (0.769)	−71082
P^{22}	11:00–11:55	0.000 (0.999)	−71082
P^{22}	9:00–11:55	0.000 (0.999)	−71082
B. Alternative Specifications (Morning Dummies for All MPC Days) Additional Variables Added to the Preferred Model Specification			
P^{11}	11:45–11:55	−0.02 (0.933)	−71082
P^{11}	11:15–11:55	−0.03 (0.879)	−71082
P^{11}	11:00–11:55	0.01 (0.930)	−71082
P^{11}	9:00–11:55	0.04 (0.530)	−71082
P^{22}	11:45–11:55	−0.44 (0.51)	−71082
P^{22}	11:30–11:55	0.53 (0.445)	−71082
P^{22}	11:00–11:55	0.66 (0.291)	−71082
P^{22}	9:00–11:55	0.18 (0.460)	−71082
C. Alternative Specifications (Morning Dummies for Unexpected Change Days) Additional Variables Added to the Preferred Model Specification			
P^{11}	11:45–11:55	0.27 (0.047)	−71080
P^{11}	11:30–11:55	0.65 (0.038)	−71080
P^{11}	11:00–11:55	0.93 (0.041)	−71079
P^{11}	11:00–11:15	1.051 (0.161)	−71082
P^{11}	9:00–11:55	4.65 (0.953)	−71082
P^{22}	11:45–11:55	−0.28 (0.552)	−71082
P^{22}	11:30–11:55	−0.07 (0.935)	−71082
P^{22}	11:15–11:55	0.05 (0.964)	−71082
P^{22}	11:00–11:55	0.28 (0.828)	−71082
P^{22}	9:00–11:55	−1.88 (0.056)	−71082

risk exposure precisely because they are unsure about the upcoming policy announcement. Furthermore, the evidence presented here indicates no particularly large probability shifts prior to the conclusion of MPC meetings, especially when compared with the afternoon of Unexpected Change days. Consequently, the news impact of policy announcements appears to be much larger than any anticipation effect.

5.5 *Dispersion of Market Expectations*

An interesting feature of figure 2 is that the probability of moving into the high-volatility state rises even on No Unexpected Change days. This is worthy of further investigation, as one might expect *anticipated* announcements to be discounted into the exchange rate in advance. One explanation may be our use of median Bloomberg survey expectations with no regard for any dispersion in expectations among survey participants. There will in general be some people surveyed who are surprised by the announcement even when it coincides with the median view. If these people then react to the announcement, this may generate a series of “hot-potato” trades, although the multiplier effect is likely to be smaller than if a majority of the market is surprised.

In order to investigate the validity of this argument, we use the measure of forecast dispersion introduced in table 1 to distinguish MPC days when analysts were unanimous on the expected policy announcement from days when they were not. We construct a dummy that takes a value of one from 12:05 to 13:00 on No Unexpected Change days when the standard deviation of analysts’ expectations is zero, and another that takes the value of one from 12:05 to 13:00 on No Unexpected Change days when the standard deviation is greater than zero. We include both of these variables in the P^{11} equation of our preferred specification and exclude “Dummy 2” from the original specification; recall from table 2, panel B that dummy 2 takes a value of one during the period 12:05–13:00 on All MPC days. The data set is reduced to 24,683 observations due to a lack of data on analysts’ expectations prior to October 1998.

Results are reported in table 7. Only the dummy variable equal to one on No Unexpected Change days with dispersion in expectations exerts a significant influence on P^{11} , with a p-value of 0.001 and a coefficient of 0.4. This suggests that the rise in P^{11} at noon on

No Unexpected Change days is indeed significant only when at least one individual deviates from the median market expectation. This validates the argument above that our use of the median individual analyst's expectation as a proxy for the entire market's expectation explains why apparently anticipated policy actions always impact the probability of being in the high-volatility state.

Overall, the evidence in figure 2 indicates that MPC days are different from other days. The noon policy announcement appears to be price-relevant news—in particular, when the announcement comes as a surprise to the market. There is some modest evidence of positioning in advance of the announcement on all MPC days, but for days when interest rates are changed unexpectedly, it appears that the largest market response comes immediately at noon with the news. It also appears that the market takes around an hour to digest the news component of an unanticipated announcement.

It is interesting to contrast our results with those of Evans and Lyons (2007). We conclude in favor of a significant but relatively short-lived news impact upon the volatility of exchange rate returns from interest rate announcements. By contrast, Evans and Lyons' analysis of proprietary order flow data concludes in favor of a very persistent relationship between order flow and exchange rate returns, with the former exhibiting out-of-sample predictive power for returns as much as one quarter ahead but no significant impact in the short term. This contrast appears to reflect differences in the behavior of market participants in the various segments of the foreign exchange market. In this paper, we have isolated the impact of knee-jerk trading on the volatility of returns around the time of MPC interest rate announcements, as interdealer positioning adjusts to reflect the arrival of this new information. This is an important and quick process, as befits a market as liquid and relatively efficient as foreign exchange. But it is only part of the story. Evans and Lyons (2007) focus explicitly away from interdealers and on the customer segment of the market that accounts for more than 50 percent of market turnover.^{16,17} As Sager and Taylor (2006) discuss, other than

¹⁶For information on the share in foreign exchange market turnover of the various market segments, see Bank for International Settlements (2007).

¹⁷This segment includes asset management firms, hedge funds, corporations, and central banks.

smaller hedge funds, the majority of foreign exchange market customers typically do not exhibit knee-jerk reaction to news in the same way as the interdealer market. Although this behavior contradicts the rational-expectations hypothesis, it is rational—in the sense of being profit maximizing—and reflects both the size of assets under management, and associated transaction costs of trading, and that a large proportion of the trading activity of this market segment is not driven by news innovations, but rather by benchmark adjustments (Lyons 2001).

6. Conclusion

The Bank of England Monetary Policy Committee, created in 1997, was charged with implementing monetary policy consistent with stable inflation and economic growth. A stated aim of the new regime was that monetary policy should be more transparent than hitherto (King 2000).

Because the MPC meets at regularly scheduled, pre-announced times and the policy decision is always announced at noon on the second meeting day, it provides a rare opportunity to examine how the decisions of a key policy-setting committee are impounded into the foreign exchange market. Our particular interest is with respect to the news content of the policy announcement and also whether there is evidence of market positioning during meetings but prior to announcements.

We examined MPC policy announcement days using five-minute exchange rate returns and a Markov-switching framework. Our estimated model assumed the existence of a high-volatility state associated with informed trading and a low-volatility state associated with liquidity trading. We diverged from the usual non-linear regime-switching framework to model endogenous transition probabilities as a function of information regarding the meeting days. The transition probabilities were found to switch systematically and significantly on meeting days. In particular, the probability of remaining in the high-volatility state was estimated to increase significantly on MPC meeting days when interest rates were changed by an amount different from the *ex ante* median market expectation, or were not changed when the market expected a change.

The announcement day of MPC meetings can therefore be characterized as having a statistically and economically significant exchange rate reaction to the MPC news announcement at noon, with some evidence of pre-positioning during the morning of the meeting. These results are qualitatively similar to those reported by Sager and Taylor (2004) in their high-frequency study of the exchange rate effects of ECB interest rate announcements, suggesting that they are robust.¹⁸

An interesting extension of our results would be to empirically test the ability of market participants to profitably exploit these announcement effects—that is, to validate the economic significance of our findings—through a profit-loss analysis of trading strategies that, say, introduce short-lived option structures in dollar-sterling on the second day of MPC meetings around the time of the policy announcement. This is a task we leave to future research.

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¹⁸Likewise, Clare and Courtenay (2001) find in favor of a significant impact on exchange and interest rate volatility from UK monetary policy announcements and macroeconomic data releases using one-minute tick data. They also find that news effects have become more short lived since the BoE was granted operational independence.

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