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Anticipated Alternative Policy Rate Paths in Policy Simulations*

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This paper specifies a new convenient algorithm to construct policy projections conditional on alternative *anticipated* policy rate paths in linearized dynamic stochastic general equilibrium (DSGE) models, such as Ramses, the Riksbank's main DSGE model. Such projections with anticipated policy rate paths correspond to situations where the central bank transparently announces that it, conditional on current information, plans to implement a particular policy rate path and where this announced plan for the policy rate is believed and then anticipated by the private sector. The main idea of the algorithm is to include among the predetermined variables (the "state" of the economy) the vector of non-zero means of future shocks to a given policy rule that is required to satisfy the given anticipated policy rate path.

JEL Codes: E52, E58.

1. Introduction

This paper specifies a new convenient way to construct policy projections conditional on alternative *anticipated* policy rate paths in linearized dynamic stochastic general equilibrium (DSGE) models,

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such as Ramses, the Riksbank's main DSGE model.[1] Such projections with anticipated policy rate paths correspond to situations where the central bank transparently announces that it, conditional on current information, plans to implement a particular policy rate path and where this announced plan for the policy rate is believed and then anticipated by the private sector. Such projections are particularly relevant for central banks such as the Reserve Bank of New Zealand (RBNZ), Norges Bank, the Riksbank, and the Czech National Bank (CNB), where the policy announcement includes not only the current policy rate decision but also a forecast path for the future policy rate. They are also relevant in the discussion about the kind of "forward guidance" about the future policy rate that the Federal Reserve System and the Bank of Canada have recently given.

A common method to do policy simulations for alternative policy rate paths is to add *unanticipated* shocks to a given instrument rule (a rule that specifies the policy rate as a function of observed variables), as in the method of modest interventions by Leeper and Zha (2003) (see appendix 4). That method is designed to deal with policy simulations that involve "modest" unanticipated deviations from a given instrument rule. Such policy simulations correspond to a situation when the central bank would non-transparently and secretly plan to surprise the private sector by deviations from an announced instrument rule (or, alternatively, a situation when the central bank announces and follows a future path but the path is not believed by, and each period surprises, the private sector). Aside from corresponding to policy that is either non-transparent or lacks credibility, such deviations are in practical simulations often both serially correlated and large, which can be inconsistent with the assumption that they would remain unanticipated and interpreted as i.i.d. shocks by the private sector. In other words, they are in practice often not "modest" in the sense of Leeper and Zha. Projections with anticipated policy rate paths would in many cases seem more relevant for

[1]The policy rate (also called the instrument rate) is the short interest rate that the central bank uses as a (policy) instrument (control variable). For the Riksbank, the policy rate is the repo rate.

the transparent flexible inflation targeting that central banks such as the RBNZ, Norges Bank, the Riksbank, and the CNB conduct.[2]

A standard way to incorporate anticipated shocks (that is, shocks with non-zero time-varying means) in an economic model with forward-looking variables is to use a deterministic, perfect-foresight variant of the model where all future shocks are set equal to their means and are assumed to be known in the first period. Furthermore, a finite horizon is assumed, with a terminal condition where all variables equal their steady-state values. The problem can then be seen as a two-point boundary problem with an initial and a terminal condition. Stacking the model equations for the finite number of periods together with the initial and terminal condition gives rise to a finite-dimensional simultaneous equation system, non-linear for a non-linear model and linear for a linear model. The model can then be solved with the Fair-Taylor (1983) algorithm or the so-called stacked-time algorithm of Laffargue (1990), Boucekkine (1995), and Juillard (1996). The horizon is extended until it has a negligible effect on the solution.[3] The Dynare (2009) collection of Matlab and Octave routines uses the stacked-time algorithm for deterministic, perfect-foresight settings.

Assuming a linear model (a linearized DSGE model), we provide an alternative simple and convenient algorithm that allows a stochastic interpretation—more precisely, a standard state-space representation of a stochastic linear model with forward-looking variables, the solution of which can be expressed in a recursive form and found with standard algorithms for the solution of linear rational expectations systems, such as the Klein (2000), Sims (2000), or AIM algorithms (Anderson and Moore 1983, 1985). The main idea is to include among the predetermined variables (the "state" of the economy) the vector of non-zero means of future shocks to a given instrument rule. By modeling the shocks as a moving-average process—more precisely, the sum of zero-mean i.i.d. shocks—we allow a consistent stochastic interpretation of new information about the non-zero

[2]However, as noted in Svensson (2010), there are recent cases when the Riksbank's policy rate path has been far from credible and when projections with unanticipated shocks may be more relevant.

[3]That is, one need only extend the horizon until such a point that the extension no longer affects the simulated results over the horizon of interest. This is a "type III iteration" in the parlance of Fair and Taylor (1983).

means. The policy rate path can then be written as a function of the initial state of the economy, including the vector of anticipated shocks, and the vector of anticipated shocks can be chosen so as to result in any desired anticipated policy rate path. This is a special case of the more general analysis of judgment in monetary policy in Svensson (2005) and of optimal policy projections with judgment in Svensson and Tetlow (2005).

Our algorithm thus adds an anticipated sequence of shocks to a general but constant policy rule, including targeting rules (conditions on the target variables, the variables that are the arguments of the loss function) and explicit or implicit instrument rules (instrument rules where the policy rate depends on predetermined variables only or also on forward-looking variables). It very conveniently allows the construction of policy projections for alternative arbitrary nominal and real policy rate paths, whether or not these are optimal for a particular loss function.

We consider policy simulations where restrictions on the nominal or real policy rate path are eventually followed by an anticipated future switch to a given well-behaved policy rule, either optimal or arbitrary. With such a setup, there is a unique equilibrium for each specified set of restrictions on the nominal or real policy rate path. The equilibrium will, in a model with forward-looking variables, depend on which future policy rule is implemented, but for any given such policy rule, the equilibrium is unique. It is well known since Sargent and Wallace (1975) that an exogenous nominal policy rate path will normally lead to indeterminacy in a model with forward-looking variables (and to an explosive development in a backward-looking model), so at some future time the nominal policy rate must become endogenous for a well-behaved equilibrium to result (see also Gagnon and Henderson 1990). Such a setup with a switch to a well-behaved policy rule solves the problem with multiple equilibria for alternative policy rate projections that Galí (2010) has emphasized. On the other hand, consistent with Galí's results, the unique equilibrium depends on and is sensitive to both the time of the switch and the policy rule to which policy shifts.

We demonstrated our method for three different models, namely the small empirical backward-looking model of the U.S. economy of Rudebusch and Svensson (1999), the small empirical forward-looking model of the U.S. economy of Lindé (2005), and Ramses, the medium-sized model of the Swedish economy of Adolfson et al.

(2007a).[4] From the examples examined in this paper, we see that in a model without forward-looking variables such as the Rudebusch-Svensson model, there is no difference between policy simulations with anticipated and unanticipated restrictions on the policy rate path. In a model with forward-looking variables, such as the Lindé model or Ramses, there is such a difference, and the impact of anticipated restrictions would generally be larger than that of unanticipated restrictions. In a model with forward-looking variables, exogenous restrictions on the policy rate path are consistent with a unique equilibrium, if there is a switch to a well-behaved policy rule in the future. For given restrictions on the policy rate path, the equilibrium depends on that policy rule.

If inflation is sufficiently sensitive to the real policy rate, "unusual" equilibria may result from restrictions for sufficiently many quarters on the nominal policy rate. Such cases have the property that a shift up of the real interest rate path reduces inflation and inflation expectations so much that the nominal interest rate path (which by the Fisher equation equals the real interest rate path plus the path of inflation expectations) shifts down. Then, a shift up of the nominal interest rate path requires an equilibrium where the path of inflation and inflation expectations shifts up more and the real policy rate path shifts down. In the Rudebusch-Svensson model, which has no forward-looking variables, inflation is so sluggish and insensitive to changes in the real policy rate that there are only small differences between restrictions on the nominal and real policy rate. In the Lindé model, inflation is so sensitive to the real policy rate that restrictions for five to six quarters or more on the nominal policy rate result in unusual equilibria. In Ramses, unusual equilibria seem to require restrictions for ten quarters or more. In order to avoid unusual equilibria, restrictions should be imposed for fewer quarters than that.

The paper is organized as follows: Section 2 presents the state-space representation of a linear(ized) DSGE model and shows how to do policy simulations with an arbitrary constant (that is, time-invariant) policy rule, such as an instrument rule or a targeting rule. Section 3 shows our convenient way of constructing policy projections that satisfy arbitrary anticipated restrictions on the nominal or real policy rate by introducing anticipated time-varying deviations in

[4]See also Adolfson et al. (2007b, 2008).

the policy rule. Section 4 provides examples of restrictions on nominal and real policy rate paths for the Rudebusch-Svensson model, the Lindé model, and Ramses. Section 5 presents some conclusions.

A few appendices contain some technical details. Appendix 1 specifies the policy rule under optimal policy under commitment. Appendices 2 and 3 provide some details on the Rudebusch-Svensson and Lindé models, respectively. Appendix 4 demonstrates the Leeper and Zha (2003) method of modest interventions in this framework.

2. The Model

A linear model with forward-looking variables (such as a DSGE model like Ramses that is linearized around a steady state) can be written in the following practical state-space form:

$$\begin{bmatrix} X_{t+1} \\ Hx_{t+1|t} \end{bmatrix} = A \begin{bmatrix} X_t \\ x_t \end{bmatrix} + Bi_t + \begin{bmatrix} C \\ 0 \end{bmatrix} \varepsilon_{t+1} \qquad (1)$$

for $t = \ldots, -1, 0, 1, \ldots$. Here, X_t is an n_X-vector of *predetermined* variables in period t (where the period is a quarter); x_t is an n_x-vector of *forward-looking* variables; i_t is generally an n_i-vector of (policy) *instruments,* but in the cases examined here it is a scalar, the policy rate (in the Riksbank's case the repo rate), so $n_i = 1$; ε_t is an n_ε-vector of i.i.d. shocks with mean zero and covariance matrix I_{n_ε}; A, B, and C, and H are matrices of the appropriate dimension; and for any stochastic process y_t, $y_{t+\tau|t}$ denotes $\mathrm{E}_t y_{t+\tau}$, the rational expectation of $y_{t+\tau}$ conditional on information available in period t. The forward-looking variables and the instruments are the *non-predetermined* variables.[5]

The variables can be measured as differences from steady-state values, in which case their unconditional means are zero. Alternatively, one of the components of X_t can be unity, so as to allow the variables to have non-zero means. The elements of the matrices A, B, C, and H are normally estimated with Bayesian methods. Here they are considered fixed and known for the policy simulations.

[5]A variable is predetermined if its one-period-ahead prediction error is an exogenous stochastic process (Klein 2000). For (1), the one-period-ahead prediction error of the predetermined variables is the stochastic vector $C\varepsilon_{t+1}$.

More precisely, the matrices are considered structural—for instance, functions of the deep parameters in an underlying linearized DSGE model. Hence, with a linear model with additive uncertainty and a quadratic loss function as specified in appendix 1, the conditions for certainty equivalence are satisfied; that is, mean forecasts are sufficient for policy decisions.

The upper block of (1) provides n_X equations determining the n_X-vector X_{t+1} in period $t+1$ for given X_t, x_t, i_t, and ε_{t+1},

$$X_{t+1} = A_{11}X_t + A_{12}x_t + B_1 i_t + C\varepsilon_{t+1}, \tag{2}$$

where A and B are partitioned conformably with X_t and x_t as

$$A \equiv \begin{bmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{bmatrix}, \qquad B = \begin{bmatrix} B_1 \\ B_2 \end{bmatrix}. \tag{3}$$

The lower block provides n_x equations determining x_t in period t for given $x_{t+1|t}$, X_t, and i_t,

$$x_t = A_{22}^{-1}(Hx_{t+1|t} - A_{21}X_t - B_2 i_t). \tag{4}$$

Hence, we assume that the $n_x \times n_x$ submatrix A_{22} is non-singular, an assumption which must be satisfied by any reasonable model with forward-looking variables.[6]

In a backward-looking model—that is, a model without forward-looking variables—there is no vector x_t of forward-looking variables and no lower block of equations in (1).

With a constant (that is, time-invariant) arbitrary instrument rule, the policy rate satisfies

$$i_t = [f_X \quad f_x] \begin{bmatrix} X_t \\ x_t \end{bmatrix}, \tag{5}$$

where the $n_i \times (n_X + n_x)$ matrix $[f_X \quad f_x]$ is a given (linear) instrument rule and partitioned conformably with X_t and x_t. If $f_x \equiv 0$, the instrument rule is an *explicit* instrument rule; if $f_x \neq 0$, the instrument rule is an *implicit* instrument rule. In the latter case,

[6]Without loss of generality, we assume that the shocks ε_t only enter in the upper block of (1), since any shocks in the lower block of (1) can be redefined as additional predetermined variables and introduced in the upper block.

the instrument rule is actually an equilibrium condition, in the sense that in a real-time analogue the policy rate in period t and the forward-looking variables in period t would be simultaneously determined.

The instrument rule that is estimated for Ramses is of the following form (see the appendix of Adolfson et al. 2011 for the notation):

$$\begin{aligned} i_t = \rho_R i_{t-1} + (1-\rho_R)[\widehat{\bar{\pi}}_t^c + r_\pi(\hat{\pi}_{t-1}^c - \widehat{\bar{\pi}}_t^c) + r_y \hat{y}_{t-1} + r_x \widehat{\tilde{x}}_{t-1}] \\ + r_{\Delta\pi}\left(\hat{\pi}_t^c - \hat{\pi}_{t-1}^c\right) + r_{\Delta y}(\hat{y}_t - \hat{y}_{t-1}) + \varepsilon_{Rt}. \end{aligned} \tag{6}$$

Since $\hat{\pi}_t^c$ and $\hat{y}_t$, the deviation of CPI inflation and output from trend, are forward-looking variables in Ramses, this is an implicit instrument rule.

An arbitrary more general (linear) policy rule (G, f) can be written as

$$G_x x_{t+1|t} + G_i i_{t+1|t} = f_X X_t + f_x x_t + f_i i_t, \tag{7}$$

where the $n_i \times (n_x + n_i)$ matrix $G \equiv [G_x \quad G_i]$ is partitioned conformably with x_t and i_t and the $n_i \times (n_X + n_x + n_i)$ matrix $f \equiv [f_X \quad f_x \quad f_i]$ is partitioned conformably with X_t, x_t, and i_t. This general policy rule includes explicit, implicit, and forecast-based instrument rules (in the latter the policy rate depends on expectations of future forward-looking variables, $x_{t+1|t}$) as well as targeting rules (conditions on current, lagged, or expected future target variables).[7] When this general policy rule is an instrument rule, we require the $n_x \times n_i$ matrix f_i to be non-singular, so (7) determines i_t for given X_t, x_t, $x_{t+1|t}$, and $i_{t+1|t}$.

The optimal instrument rule under commitment (see appendix 1) can be written as

$$0 = F_{iX} X_t + F_{i\Xi} \Xi_{t-1} - i_t, \tag{8}$$

where the matrix F_i in (33) is partitioned conformably with X_t and Ξ_{t-1}. Here the n_x-vector of Lagrange multipliers Ξ_t in equilibrium follows

$$\Xi_t = M_{\Xi X} X_t + M_{\Xi\Xi} \Xi_{t-1}, \tag{9}$$

[7] A targeting rule can be expressed in terms of expected leads, current values, and lags of the target variables (the arguments of the loss function); see Svensson (1999), Svensson and Woodford (2005), and Giannoni and Woodford (2010).

where the matrix M in (32) has been partitioned conformably with X_t and Ξ_{t-1}. Thus, in order to include this optimal instrument rule in the set of policy rules (7) considered, the predetermined variables need to be augmented with Ξ_{t-1} and the equations for the predetermined variables with (9). For simplicity, the treatment below does not include this augmentation. Alternatively, below the vector of predetermined variables could consistently be augmented with the vector of Lagrange multipliers, so everywhere we would have $(X_t', \Xi_{t-1}')'$ instead of X_t, with corresponding augmentation of the relevant matrices.

The general policy rule can be added to the model equations (1) to form the new system to be solved. With the notation $\tilde{x}_t \equiv (x_t', i_t')'$, the new system can be written

$$\begin{bmatrix} X_{t+1} \\ \tilde{H}\tilde{x}_{t+1|t} \end{bmatrix} = \tilde{A} \begin{bmatrix} X_t \\ \tilde{x}_t \end{bmatrix} + \begin{bmatrix} C \\ 0_{(n_x+n_i)\times n_\varepsilon} \end{bmatrix} \varepsilon_{t+1}, \tag{10}$$

for $t = \ldots, -1, 0, 1, \ldots$, where

$$\tilde{H} \equiv \begin{bmatrix} H & 0 \\ G_x & G_i \end{bmatrix}, \quad \tilde{A} \equiv \begin{bmatrix} A_{11} & A_{12} & B_1 \\ A_{21} & A_{22} & B_2 \\ f_X & f_x & f_i \end{bmatrix},$$

and where $\tilde{H}$ is partitioned conformably with x_t and i_t and $\tilde{A}$ is partitioned conformably with X_t, x_t, and i_t.

Then, under the assumption that the policy rule gives rise to the saddlepoint property (that the number of eigenvalues with modulus greater than unity is equal to the number of non-predetermined variables), the system can be solved with the Klein (2000) algorithm or the other algorithms for the solution of linear rational expectations models mentioned in the introduction. The Klein algorithm generates the matrices M and F such that the resulting equilibrium satisfies

$$X_{t+1} = MX_t + C\varepsilon_{t+1}, \tag{11}$$

$$\tilde{x}_t \equiv \begin{bmatrix} x_t \\ i_t \end{bmatrix} = FX_t \equiv \begin{bmatrix} F_x \\ F_i \end{bmatrix} X_t \tag{12}$$

for $t = \ldots, -1, 0, 1, \ldots$, where the matrices M and F depend on $\tilde{A}$ and $\tilde{H}$, and thereby on A, B, H, G, and f.

In a backward-looking model, the time-invariant instrument rule depends on the vector of predetermined variables only, since there are no forward-looking variables, and the vector $\tilde{x}_t$ is identical to i_t.

Consider now *projections* in period t—that is, mean forecasts, conditional on information available in period t, of future realizations of the variables. For any stochastic vector process u_t, let $u^t \equiv \{u_{t+\tau,t}\}_{\tau=0}^{\infty}$ denote a *projection* in period t, where $u_{t+\tau,t}$ denotes the mean forecast of the realization of the vector in period $t+\tau$ conditional on information available in period t. We refer to τ as the horizon of the forecast $u_{t+\tau,t}$.

The projection (X^t, x^t, i^t) in period t is then given by (11) and (12) when we set the mean of future i.i.d. shocks equal to zero, $\varepsilon_{t+\tau,t} = \mathrm{E}_t \varepsilon_{t+\tau} = 0$ for $\tau > 0$. It then satisfies

$$X_{t+\tau,t} = M^\tau X_{t,t}, \tag{13}$$

$$\tilde{x}_{t+\tau,t} \equiv \begin{bmatrix} x_{t+\tau,t} \\ i_{t+\tau,t} \end{bmatrix} = F X_{t+\tau,t} \equiv \begin{bmatrix} F_x \\ F_i \end{bmatrix} X_{t+\tau,t} = \begin{bmatrix} F_x \\ F_i \end{bmatrix} M^\tau X_{t,t}, \tag{14}$$

$$X_{t,t} = X_{t|t}, \tag{15}$$

for $\tau \geq 0$, where $X_{t|t}$ is the estimate of predetermined variables in period t conditional on information available in the beginning of period t. Thus, "$,t$" and "t" in subindices refer to projections (forecasting) and estimates ("nowcasting" and "backcasting") in the beginning of period t, respectively.

3. Projections with Time-Varying Restrictions on the Policy Rate

The projection of the policy rate $i^t = \{i_{t+\tau,t}\}_{\tau=0}^{\infty}$ in period t is by (14) given by

$$i_{t+\tau,t} = F_i M^\tau X_{t+\tau,t}$$

for $\tau \geq 0$.[8]

[8] The projection of the policy rate and the other variables will satisfy the policy rule,

$$G_x x_{t+\tau+1,t} + G_i i_{t+\tau+1,t} = f_X X_{t+\tau,t} + f_x x_{t+\tau,t} + f_i i_{t+\tau,t},$$

for $\tau \geq 0$.

Suppose now that we consider imposing restrictions on the policy rate projection of the form

$$i_{t+\tau,t} = \bar{\imath}_{t+\tau,t}, \quad \tau = 0, \ldots, T, \tag{16}$$

where $\{\bar{\imath}_{t+\tau,t}\}_{\tau=0}^{T}$ is a sequence of $T+1$ given policy rate levels. Alternatively, we can have restriction on the real policy rate projection of the form

$$r_{t+\tau,t} = \bar{r}_{t+\tau,t}, \quad \tau = 0, \ldots, T, \tag{17}$$

where

$$r_t \equiv i_t - \pi_{t+1|t} \tag{18}$$

is the real policy rate and $\pi_{t+1|t}$ is expected inflation. With restrictions of this kind, the nominal or real policy rate is exogenous for period $t, t+1, \ldots, t+T$.

These restrictions are here assumed to be anticipated by both the central bank and the private sector, in contrast to Leeper and Zha (2003) where they are anticipated and planned by the central bank but not anticipated by the private sector. Thus, our case corresponds to a situation where the restriction is announced to the private sector by the central bank and believed by the private sector, whereas the Leeper and Zha case corresponds to a situation where the central bank either makes secret plans to implement the restriction or the restriction is announced but not believed by the private sector.

The restrictions make the nominal or real policy rate projection exogenous for the periods t, $t+1$, $\ldots$, $t+T$. We know from Sargent and Wallace (1975) that exogenous interest rates may cause indeterminacy when there are forward-looking variables. In order to ensure determinacy, we assume that there is an anticipated switch in period $t+T+1$ to the policy rule (G, f). Then the restrictions can be implemented by augmenting a stochastic deviation, z_t, to the policy rule (7),

$$G_x x_{t+1|t} + G_i i_{t+1|t} = f_X X_t + f_x x_t + f_i i_t + z_t. \tag{19}$$

The projection $\{z_{t+\tau,t}\}_{\tau=0}^{T}$ of the future deviations is then chosen such that (16) or (17) is satisfied. The projection of the future deviation from the horizon $T+1$ and beyond is zero, corresponding to the anticipated shift then to the policy rule (G, f).

More precisely, we let the $(T+1)$-vector $z^t \equiv (z_{t,t}, z_{t+1,t}, \ldots, z_{t+T,t})'$ (where $z_{t,t} = z_t$) denote a projection of the stochastic variable $z_{t+\tau}$ for $\tau = 0, \ldots, T$. As in the treatment of central bank judgment in Svensson (2005), the stochastic variable z_t is called the deviation. In particular, we assume that the deviation is a moving-average process that satisfies

$$z_t = \eta_{t,t} + \sum_{s=1}^{T} \eta_{t,t-s}$$

for a given $T \geq 0$, where $\eta^t \equiv (\eta'_{t,t}, \eta'_{t+1,t}, \ldots, \eta'_{t+T,t})'$ is a zero-mean i.i.d. random $(T+1)$-vector realized in the beginning of period t and called the innovation in period t. For $T = 0$, we have $z_t = \eta_{t,t}$, and the deviation is a simple i.i.d. disturbance. For $T > 0$, the deviation instead follows a moving-average process. Then we have

$$\begin{aligned} z_{t+\tau,t+1} &= z_{t+\tau,t} + \eta_{t+\tau,t+1}, \quad \tau = 1, \ldots, T, \\ z_{t+T+1,t+1} &= \eta_{t+T+1,t+1}. \end{aligned}$$

It follows that the dynamics of the deviation and the projection z^t can be written

$$z^{t+1} = A_z z^t + \eta^{t+1}, \tag{20}$$

where the $(T+1) \times (T+1)$ matrix A_z is defined as

$$A_z \equiv \begin{bmatrix} 0_{T\times 1} & I_T \\ 0 & 0_{1\times T} \end{bmatrix}.$$

Hence, z^t is the central bank's mean projection of current and future deviations, and η^t can be interpreted as the new information the central bank receives in the beginning of period t about those deviations.[9]

[9] In Svensson (2005) the deviation z_t is an n_z-vector of terms entering the different equations in the model, and the projection z^t of future z_t deviation is identified with central bank judgment. The graphs in Svensson (2005) can be seen as impulse responses to η^t, the new information about future deviations. (The notation here is slightly different from Svensson 2005 in that there the projection $z^t \equiv (z_{t+1,t}, \ldots, z_{t+T,t})'$ does not include the current deviation.)

Combining the model (1) with the augmented policy rule (19) gives the system

$$\begin{bmatrix} \tilde{X}_{t+1} \\ \tilde{H}\tilde{x}_{t+1|t} \end{bmatrix} = \tilde{A} \begin{bmatrix} \tilde{X}_t \\ \tilde{x}_t \end{bmatrix} + \begin{bmatrix} \tilde{C} \\ 0_{(n_x+n_i)\times(n_\varepsilon+T+1)} \end{bmatrix} \begin{bmatrix} \varepsilon_{t+1} \\ \eta^{t+1}, \end{bmatrix} \tag{21}$$

for $t = \ldots, -1, 0, 1, \ldots$, where

$$\tilde{X}_t \equiv \begin{bmatrix} X_t \\ z^t \end{bmatrix}, \quad \tilde{x}_t \equiv \begin{bmatrix} x_t \\ i_t \end{bmatrix}, \quad \tilde{H} \equiv \begin{bmatrix} H & 0 \\ G_x & G_i \end{bmatrix},$$

$$\tilde{A} \equiv \begin{bmatrix} A_{11} & 0_{n_X\times 1} & 0_{n_X\times T} & A_{12} & B_1 \\ 0_{T\times n_X} & 0_{T\times 1} & I_T & 0_{T\times n_x} & 0_{T\times 1} \\ 0_{1\times n_X} & 0 & 0_{1\times T} & 0_{1\times n_x} & 0 \\ A_{21} & 0_{n_x\times 1} & 0_{n_x\times T} & A_{22} & B_2 \\ f_X & 1 & 0_{1\times T} & f_x & f_i \end{bmatrix},$$

$$\tilde{C} \equiv \begin{bmatrix} C & 0_{n_X\times(T+1)} \\ 0_{(T+1)\times n_\varepsilon} & I_{T+1} \end{bmatrix}.$$

Under the assumption of the saddlepoint property, the system of difference equations (21) has a unique solution and there exist unique matrices M and F such that projection can be written

$$\tilde{X}_{t+\tau,t} = M^\tau \tilde{X}_{t,t},$$
$$\tilde{x}_{t+\tau,t} = F\tilde{X}_{t+\tau,t} = FM^\tau \tilde{X}_{t,t}$$

for $\tau \geq 0$, where $X_{t,t}$ in $\tilde{X}_{t,t} \equiv (X'_{t,t}, z^{t\prime})'$ is given but the $(T+1)$-vector z^t remains to be determined. Its elements are then determined by the restrictions (16) or (17).

In order to satisfy the restriction (16) on the nominal policy rate, we note that it can now be written

$$i_{t+\tau,t} = F_i M^\tau \begin{bmatrix} X_{t,t} \\ z^t \end{bmatrix} = \bar{\imath}_{t+\tau,t}, \quad \tau = 0, 1, \ldots, T.$$

This provides $T+1$ linear equations for the $T+1$ elements of z^t.

In order to instead satisfy the restriction (17) on the real policy rate, we note that inflation expectations in a DSGE model similar to Ramses generally satisfy

$$\pi_{t+1|t} \equiv \varphi \tilde{x}_{t+1|t} + \Phi \begin{bmatrix} \tilde{X}_t \\ \tilde{x}_t \end{bmatrix} \tag{22}$$

for some vectors φ and Φ. These vectors φ and Φ are structural, not reduced-form expressions. For instance, if π_t is one of the elements of x_t, the corresponding element of φ is unity, all other elements of φ are zero, and $\Phi \equiv 0$. If $\pi_{t+1|t}$ is one of the elements of $\tilde{x}_t$, the corresponding element of Φ is unity, all other elements of Φ are zero, and $\varphi \equiv 0$. Then the restriction (17) can be written

$$\begin{aligned} r_{t+\tau,t} \equiv i_{t+\tau,t} - \pi_{t+\tau+1,t} &= (F_i - \varphi F M - \Phi) M^\tau \begin{bmatrix} X_{t,t} \\ z^t \end{bmatrix} \\ &= \bar{r}_{t+\tau,t}, \quad \tau = 0, 1, \ldots, T. \end{aligned}$$

This again provides $T+1$ linear equations for the $T+1$ elements of z^t.

When the restriction is on the nominal policy rate, we can think of the equilibrium as being implemented by the central bank announcing the nominal policy rate path and the private sector incorporating this policy rate projection in their expectations, with the understanding that the policy rate will be set according to the given policy rule (G, f) from period $t+T+1$. When the restriction is on the real policy rate, we need to consider the fact that in practice central banks set nominal policy rates, not real ones. The restriction on the real policy rate will result in an endogenously determined nominal policy rate projection, which together with the endogenously determined inflation projection will be consistent with the real policy rate path. We can then think of the equilibrium as being implemented by the central bank calculating that nominal policy rate projection and then announcing it to the private sector.

3.1 Backward-Looking Model

In a backward-looking model, the projection of the instrument rule with the time-varying constraints can be written

$$i_{t+\tau,t} = f_X X_{t+\tau,t} + z_{t+\tau,t}, \tag{23}$$

so it is trivial to determine the projection z^t recursively so as to satisfy the restriction (16) on the nominal policy rate projection.

Inflation can be written

$$\pi_t = \Phi X_t$$

for some vector Φ, so expected inflation can be written

$$\pi_{t+1|t} = \Phi X_{t+1|t} = \Phi(AX_t + Bi_t). \tag{24}$$

By combining (23), (24), and (18), it is trivial to determine the projection z^t so as to satisfy the restriction (17) on the real policy rate projection.

4. Examples

In this section we examine restrictions on the nominal and real policy rate path for the backward-looking Rudebusch-Svensson model and the two forward-looking models, the Lindé model and Ramses. Appendices 2 and 3 provide some details on the Rudebusch-Svensson and Lindé models. We also show a simulation with Ramses with the method of modest interventions by Leeper and Zha. Appendix 4 provides some details on the Leeper-Zha method.

4.1 The Rudebusch-Svensson Model

The backward-looking empirical Rudebusch-Svensson (1999) model has two equations (with estimates rounded to two decimal points):

$$\pi_{t+1} = 0.70\,\pi_t - 0.10\,\pi_{t-1} + 0.28\,\pi_{t-2} + 0.12\,\pi_{t-3} + 0.14\,y_t + \varepsilon_{\pi,t+1}, \tag{25}$$

$$y_{t+1} = 1.16\,y_t - 0.25\,y_{t-1} - 0.10\left(\frac{1}{4}\Sigma_{j=0}^{3} i_{t-j} - \frac{1}{4}\Sigma_{j=0}^{3}\pi_{t-j}\right) + \varepsilon_{y,t+1}. \tag{26}$$

The period is a quarter, π_t is quarterly GDP inflation measured in percentage points at an annual rate, y_t is the output gap measured in percentage points, and i_t is the quarterly average of the federal funds rate, measured in percentage points at an annual rate. All variables are measured as differences from their means, their steady-state levels. The predetermined variables are

$X_t \equiv (\pi_t, \pi_{t-1}, \pi_{t-2,}\pi_{t-2}, y_t, y_{t-1}, i_{t-1}, i_{t-2,}i_{t-3})'$. See appendix 2 for details.

The target variables are inflation, the output gap, and the first difference of the federal funds rate. The period loss function is

$$L_t = \frac{1}{2}\big[\pi_t^2 + \lambda_y y_t^2 + \lambda_{\Delta i}(i_t - i_{t-1})^2\big], \tag{27}$$

where π_t is measured as the difference from the inflation target, which is equal to the steady-state level. The discount factor, δ, and the relative weights on output-gap stabilization, λ_y, and interest rate smoothing, $\lambda_{\Delta i}$, are set to satisfy $\delta = 1$, $\lambda_y = 1$, and $\lambda_{\Delta i} = 0.2$.

For the loss function (27) with the parameters $\delta = 1$, $\lambda_y = 1$, and $\lambda_{\Delta i} = 0.2$, and the case where ε_t is an i.i.d. shock with zero mean, the optimal instrument rule is as follows (the coefficients are rounded to two decimal points):

$$\begin{aligned} i_t = {} & 1.22\,\pi_t + 0.43\,\pi_{t-1} + 0.53\,\pi_{t-2} + 0.18\,\pi_{t-3} + 1.93\,y_t - 0.49\,y_{t-1} \\ & + 0.36\,i_{t-1} - 0.09\,i_{t-2} - 0.05\,i_{t-3}. \end{aligned}$$

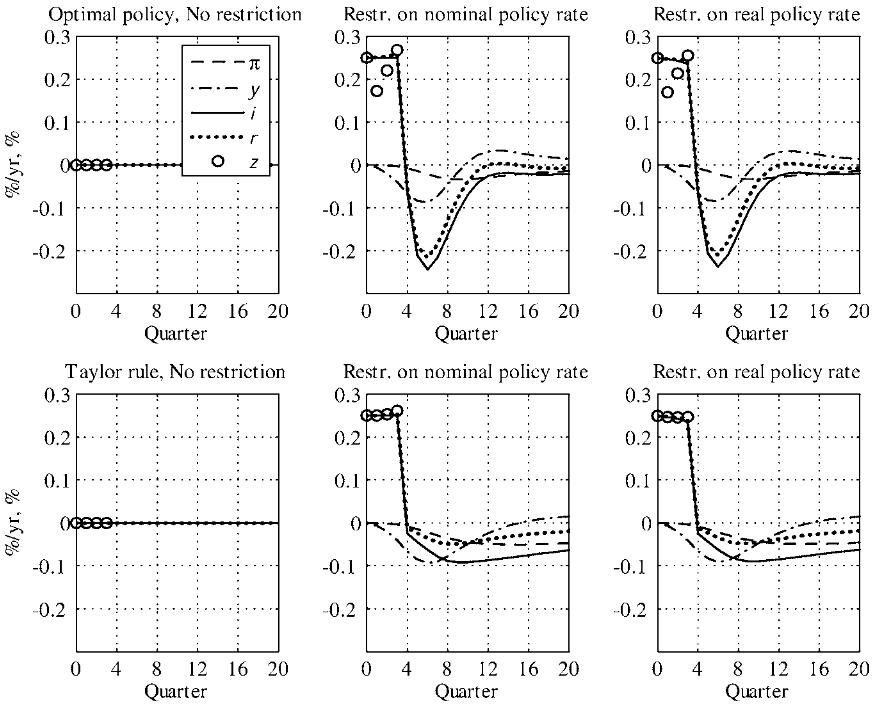
Figure 1 shows projections for the Rudebusch-Svensson model. The top row of panels shows projections under the optimal policy, whereas the bottom row of panels shows projections under a Taylor rule,

$$i_t = 1.5\,\pi_t + 0.5\,y_t,$$

where the policy rate responds to the predetermined inflation and output gap with the standard coefficients 1.5 and 0.5, respectively.

The projections start in quarter 0 from the steady state, when all the predetermined variables are zero. The left column of panels shows the projections when there is no restriction imposed on the nominal or real policy rate path. This corresponds to zero projected deviations $z_{t+\tau,t}$ in the optimal instrument rule and the Taylor rule. These are denoted by circles for the first four quarters, quarters 0–3. The economy remains in the steady state, and inflation (denoted by a dashed curve), the output gap (denoted by a dashed-dotted curve), the nominal policy rate (denoted by a solid curve), and the real policy rate (denoted by a dotted curve) all remain at zero.

Figure 1. Projections for Rudebusch-Svensson Model with Unrestricted and Restricted Nominal and Real Policy Rate for Optimal Policy (Top Row) and Taylor Rule (Bottom Row): Four-Quarter Restriction



The middle column shows projections when the nominal policy rate is restricted to equal 25 basis points for the first four quarters. For both the optimal policy and the Taylor rule, this requires positive and (except for quarter 1) increasing time-varying projected deviations in the instrument rule. The upward shift in quarters 0–3 in the nominal policy rate path reduces inflation and expected inflation somewhat, and the real policy rate path shifts up a bit more than the nominal policy rate path. The increased real policy rate also reduces the output gap. In the Rudebusch-Svensson model, inflation is very sluggish and the output gap responds more to the nominal and real policy rate than inflation. From quarter 4, there is no restriction on the policy rate path, and according to both the optimal policy and the Taylor rule, the nominal and real policy rate are reduced substantially so as to bring the negative inflation and output gap eventually back to zero. The optimal policy is more effective in bringing back inflation and the output gap than the Taylor rule, which is natural since the Taylor rule is not optimal.

The right column shows projections when the real policy rate is restricted to equal unity during quarters 0–3. Since there is so little movement in inflation and expected inflation, the projections for these restrictions on the real and the nominal policy rate are very similar.

Since there are no forward-looking variables in the Rudebusch-Svensson model, there would be no difference between these projections with anticipated restrictions on the policy rate path and simulations with unanticipated shocks as in Leeper and Zha (2003).

4.2 The Lindé Model

The empirical New Keynesian model of the U.S. economy due to Lindé (2005) also has two equations. We use the following parameter estimates:

$$\pi_t = 0.457\,\pi_{t+1|t} + (1-0.457)\pi_{t-1} + 0.048 y_t + \varepsilon_{\pi t},$$
$$y_t = 0.425\,y_{t+1|t} + (1-0.425)y_{t-1} - 0.156(i_t - \pi_{t+1|t}) + \varepsilon_{yt}.$$

The period is a quarter, and π_t is quarterly GDP inflation measured in percentage points at an annual rate, y_t is the output gap measured in percentage points, and i_t is the quarterly average of the federal funds rate, measured in percentage points at an annual rate. All variables are measured as differences from their means, their steady-state levels. The shock $\varepsilon_t \equiv (\varepsilon_{\pi t}, \varepsilon_{yt})'$ is i.i.d. with mean zero.

For the loss function (27), the predetermined variables are $X_t \equiv (\varepsilon_{\pi t}, \varepsilon_{yt}, \pi_{t-1}, y_{t-1}, i_{t-1})'$ (the lagged policy rate enters because it enters into the loss function, and the two shocks are included among the predetermined variables in order to write the model on the form (1) with no shocks in the equations for the forward-looking variables). The forward-looking variables are $x_t \equiv (\pi_t, y_t)'$. See appendix 3 for details.[10]

For the loss function (27) with the parameters $\delta = 1$, $\lambda_y = 1$, and $\lambda_{\Delta i} = 0.2$, the optimal policy function (8) is as follows (the coefficients are rounded to two decimal points):

$$\begin{aligned} i_t = {} & 1.06\,\varepsilon_{\pi t} + 1.38\,\varepsilon_{yt} + 0.58\,\pi_{t-1} + 0.78\,y_{t-1} + 0.40\;i_{t-1} \\ & + 0.02\,\Xi_{\pi,t-1,t-1} + 0.20\,\Xi_{y,t-1,t-1}, \end{aligned}$$

where $\Xi_{\pi,t-1,t-1}$ and $\Xi_{y,t-1,t-1}$ are the Lagrange multipliers for the two equations for the forward-looking variables in the decision problem in period $t-1$ (see appendix 1). The difference equation (9) for the Lagrange multipliers is

$$\begin{aligned} \begin{bmatrix} \Xi_{\pi t} \\ \Xi_{yt} \end{bmatrix} = {} & \begin{bmatrix} 10.20 & 0.74 & 5.54 & 0.43 & -0.21 \\ 0.74 & 1.48 & 0.40 & 0.85 & -0.28 \end{bmatrix} \begin{bmatrix} \varepsilon_{\pi t} \\ \varepsilon_{yt} \\ \pi_{t-1} \\ y_{t-1} \\ i_{t-1} \end{bmatrix} \\ & + \begin{bmatrix} 0.72 & 0.16 \\ 0.03 & 0.38 \end{bmatrix} \begin{bmatrix} \Xi_{\pi,t-1} \\ \Xi_{y,t-1} \end{bmatrix}. \end{aligned}$$

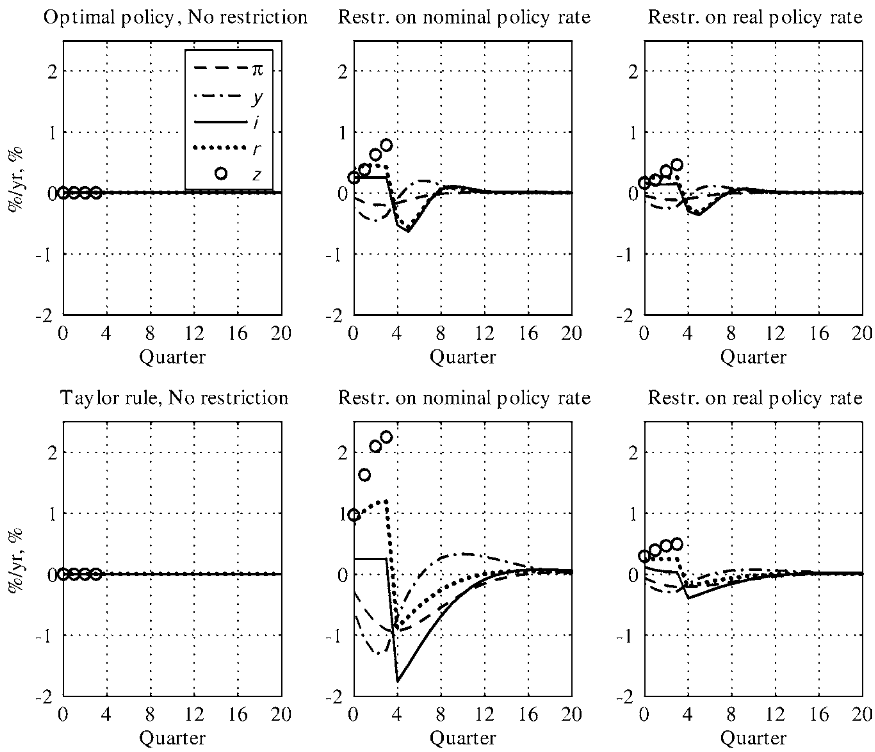
We also examine the projections for a Taylor rule for which the policy rate responds to current inflation and the output gap,

$$i_t = 1.5\,\pi_t + 0.5\,y_t.$$

Figure 2 shows projections for the optimal policy (top row) and Taylor rule (bottom row) when there is a restriction to equal 25 basis points for quarters 0–3 for the nominal policy rate (middle column) and the real policy rate (right column). In the middle column, we see that a restriction to a 25-basis-points-higher nominal policy rate reduces inflation and inflation expectations so the projection of the

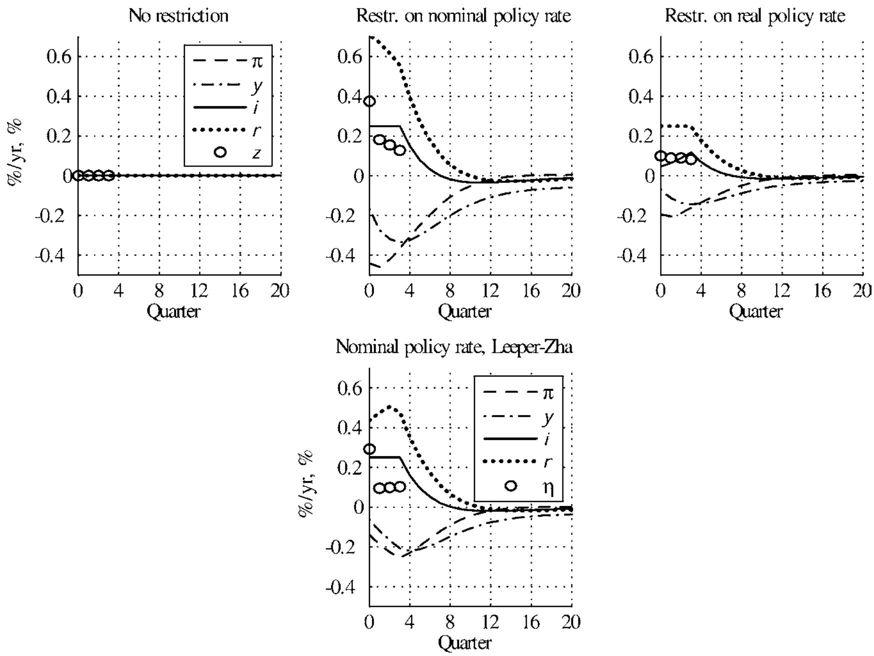
[10] It is arguably unrealistic to consider inflation and output in the current quarter as forward-looking variables. Alternatively, current inflation and the output gap could be treated as predetermined, and one-quarter-ahead plans for inflation, the output gap, and the policy rate could be determined by the model above. Such a variant of the New Keynesian model is used in Svensson and Woodford (2005).

Figure 2. Projections for the Lindé Model with Unrestricted and Restricted Nominal and Real Policy Rate for Optimal Policy (Top Row) and Taylor Rule (Bottom Row): Four-Quarter Restriction



real policy rate is above 25 basis points and higher than the policy rate for the first four quarters. In line with this, in the right column, the restriction on the real policy rate reduces inflation and inflation expectations so the corresponding nominal policy rate projection is below 25 basis points. We note that these restrictions require positive and rising time-varying projected deviations (denoted by the circles). The magnitude of the projected deviations is larger than those in figure 1 for the Rudebusch-Svensson model. Using the magnitude of the projected deviations as indicating the severity of the restriction, we conclude that the restriction to nominal or real policy rates equal to unity is more severe in the Lindé model.

Figure 3. Projections for Ramses with Anticipated Unrestricted and Restricted Nominal and Real Policy Rate (Top Row) and Unanticipated Restrictions on the Nominal Policy Rate (Bottom Row): Four-Quarter Restriction



Because inflation is more sensitive to movements in the real policy rate in the Lindé model than in the Rudebusch-Svensson model, there is a greater difference between restrictions on the nominal and the real policy rate. Also, from quarter 4, when there is no restriction on the policy rate, a fall in the real and nominal policy rate, according to both the optimal policy and the Taylor rule, more easily stabilizes inflation and the output gap back to the steady state than in the Rudebusch-Svensson model.

4.3 Ramses

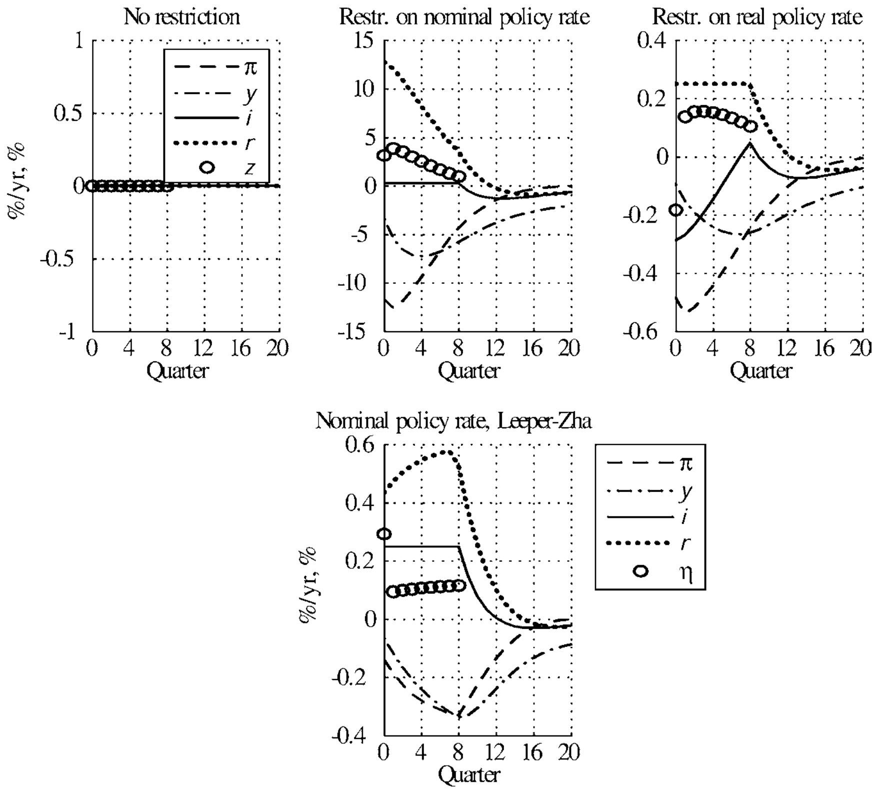
Adolfson et al. (2011) provide more details on Ramses, including the elements of the vectors X_t, x_t, i_t, and ε_t. Figure 3 shows projections with Ramses for the estimated instrument rule. The top row shows the result of restrictions on the nominal and real policy rate to equal 25 basis points for four quarters, quarters 0–3. We see that there is a substantial difference between restrictions on the nominal and the real policy rate, since inflation is quite sensitive to the real policy rate in Ramses. In the top-middle panel, we see that a restriction on the nominal policy rate projection to equal 25 basis points for quarters 0–3 corresponds to a very high and falling real policy rate projection. In the top-right panel we see that the restriction on the real policy rate to equal 25 basis points for quarters 0–3 corresponds to a nominal policy rate projection quite a bit below the real policy rate.

The bottom panel of figure 3 shows the result of a projection with the Leeper-Zha method of modest interventions to implement a restriction on the nominal policy rate to equal 25 basis points for quarters 0–3. There, positive unanticipated shocks (denoted by circles) are added to the estimated instrument rule to achieve the restriction on the nominal policy rate. Comparing the bottom panel with the top-right panel, we see that the impact on inflation, the output gap, and the real interest rate is smaller for the unanticipated shocks in the Leeper-Zha method than for the anticipated projected deviations in our method.

4.4 Unusual Equilibria

If restrictions are imposed on the nominal policy rate for many periods, "unusual" equilibria can occur. We can illustrate this for Ramses in figure 4, where in the top-middle panel the nominal policy rate is restricted to equal 25 basis points for nine quarters, quarters 0–8. This is a very contractionary policy, which shows in inflation and inflation expectations falling very much and the real policy rate becoming very high. (Note that the scale varies from panel to panel in figure 4.) If we look at the top-right panel, where the real policy rate is restricted to equal 25 basis points for nine quarters, we see that inflation and inflation expectations fall so much that the nominal policy rate becomes negative in quarter 0 (relative to when there is no restriction) and then rises to become positive only in quarters 7 and 8. We realize that if inflation and inflation expectations respond so much that nominal and real policy rates move in

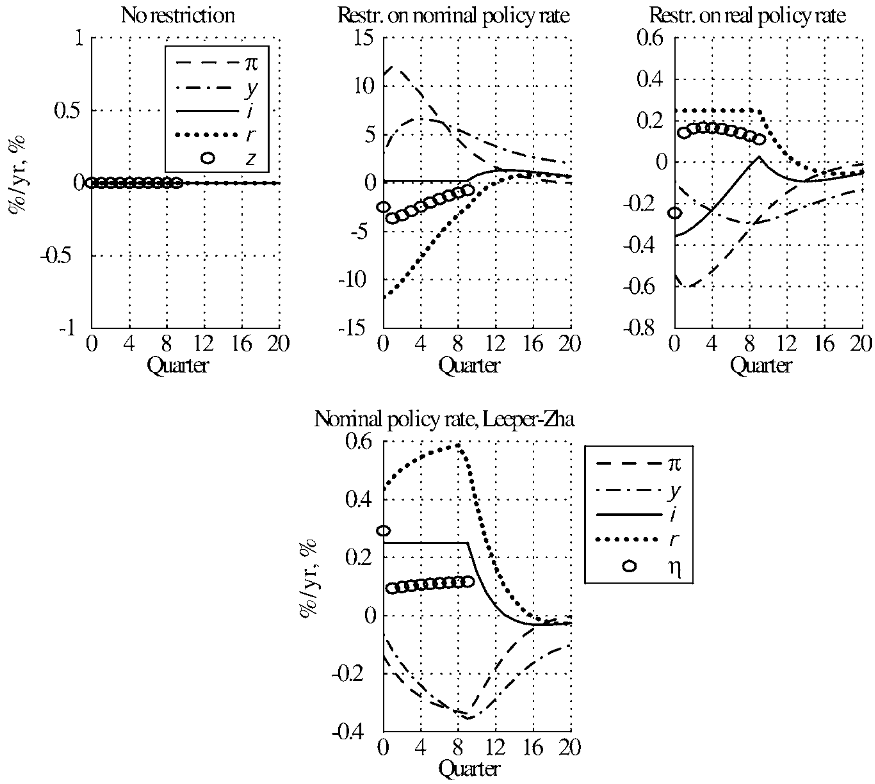
Figure 4. Projections for Ramses with Anticipated Unrestricted and Restricted Nominal and Real Policy Rate (Top Row) and Unanticipated Restrictions on the Nominal Policy Rate (Bottom Row): Nine-Quarter Restriction



opposite directions, some unusual equilibria may arise. This is confirmed in figure 5, where in the top-middle panel the nominal policy rate is restricted at 25 basis points for one more quarter, quarter 9. We see that then there is no longer an equilibrium where the real policy rate is positive and high. Instead the equilibrium is such that the real policy rate is negative, policy is very expansionary, and inflation and inflation expectations are high.

This phenomenon of unusual equilibria clearly requires that inflation and inflation expectations are quite sensitive to the real policy rate so that for multiple-period restrictions the nominal and real

Figure 5. Projections for Ramses with Anticipated Unrestricted and Restricted Nominal and Real Policy Rate (Top Row) and Unanticipated Restrictions on the Nominal Policy Rate (Bottom Row): Ten-Quarter Restriction



policy rates move in opposite directions. It requires as much as around ten-quarter restrictions to occur in Ramses. In the Lindé model, inflation is more sensitive to the real policy rate, so there it can occur already at six-quarter restrictions. We have not observed the phenomenon in the Rudebusch-Svensson model even for very long restrictions.

The phenomenon implies that restrictions for many quarters should be avoided in models where inflation and inflation expectations are sufficiently sensitive to the real policy rate.

5. Conclusions

We have presented a new convenient way to construct projections conditional on alternative *anticipated* policy rate paths in linearized dynamic stochastic general equilibrium (DSGE) models, such as Ramses, the Riksbank's main DSGE model. The main idea is to include the anticipated future time-varying deviations from a policy rule in the vector of predetermined variables, the "state" of the economy. This allows the formulation of the linear(ized) model on a standard state-space form, the application of standard algorithms for the solution of linear rational expectations models, and a recursive representation of the equilibrium projections. Projections for anticipated policy rate paths are especially relevant for central banks—such as the Reserve Bank of New Zealand, Norges Bank, the Riksbank, and the Czech National Bank—that publish a policy rate path, but they are also relevant for the discussion of the kind of "forward guidance" recently given by the Federal Reserve and Bank of Canada.

From the examples in this paper, we have seen that in a model without forward-looking variables such as the empirical model of the U.S. economy by Rudebusch and Svensson (1999), there is no difference between policy simulations with anticipated and unanticipated restrictions on the policy rate path. In a model with forward-looking variables, such as Ramses or the empirical New Keynesian model of the U.S. economy by Lindé (2005), there is such a difference, and the impact of anticipated deviations from a policy rule will generally be larger than that of unanticipated deviations. In a model with forward-looking variables, exogenous restrictions on the policy rate path are consistent with a unique equilibrium, if there is an anticipated switch to a well-behaved policy rule in the future. For given restrictions on the policy rate path, the equilibrium depends on that policy rule.

Furthermore, our analysis shows that if inflation is sufficiently sensitive to the real policy rate, "unusual" equilibria may result from restrictions on the nominal policy rate for sufficiently many periods. Such cases have the property that nominal and real policy rates move in opposite directions and nominal policy rates and inflation (expectations) move in the same direction. This phenomenon implies that restrictions on nominal policy rates for too many periods should be avoided.

Appendix 1. Optimal Policy

Let Y_t be an n_Y-vector of *target* variables, measured as the difference from an n_Y-vector Y^* of *target levels*. This is not restrictive, as long as we keep the target levels time invariant. If we would like to examine the consequences of different target levels, we can instead interpret Y_t as the absolute level of the target levels and replace Y_t with $Y_t - Y^*$ everywhere below. We assume that the target variables can be written as a linear function of the predetermined variables, the forward-looking variables, and the instruments,

$$Y_t = D \begin{bmatrix} X_t \\ x_t \\ i_t \end{bmatrix} \equiv [D_X \quad D_x \quad D_i] \begin{bmatrix} X_t \\ x_t \\ i_t \end{bmatrix}, \tag{28}$$

where D is an $n_Y \times (n_X + n_x + n_i)$ matrix and partitioned conformably with X_t, x_t, and i_t.

Let the intertemporal loss function in period t be

$$\mathrm{E}_t \sum_{\tau=0}^{\infty} \delta^{\tau} L_{t+\tau}, \tag{29}$$

where $0 < \delta < 1$ is a discount factor, L_t is the period loss given by

$$L_t \equiv Y_t' \Lambda Y_t, \tag{30}$$

and Λ is a symmetric positive semi-definite matrix containing the weights on the individual target variables.[11]

Optimization under commitment in a timeless perspective (Woodford 2003), combined with the model equations (1), results in a system of difference equations (see Söderlind 1999 and Svensson 2009). The system of difference equations can be solved with several alternative algorithms—for instance, those developed by Klein (2000) and Sims (2000) or the AIM algorithm of Anderson and Moore (1983, 1985) (see Svensson 2005, 2009 for details of the derivation and the application of the Klein algorithm). The equilibrium

[11] For plotting and other purposes, and to avoid unnecessary separate program code, it is convenient to expand the vector Y_t to include a number of variables of interest that are not necessary target variables or potential target variables. These will then have zero weight in the loss function.

under optimal policy under commitment can be described by the following difference equation:

$$\begin{bmatrix} x_t \\ i_t \end{bmatrix} = \begin{bmatrix} F_x \\ F_i \end{bmatrix} \begin{bmatrix} X_t \\ \Xi_{t-1} \end{bmatrix}, \tag{31}$$

$$\begin{bmatrix} X_{t+1} \\ \Xi_t \end{bmatrix} = M \begin{bmatrix} X_t \\ \Xi_{t-1} \end{bmatrix} + \begin{bmatrix} C \\ 0 \end{bmatrix} \varepsilon_{t+1}. \tag{32}$$

The Klein algorithm returns the matrices F_x, F_i, and M. The submatrix F_i in (32) represents the optimal instrument rule,

$$i_t = F_i \begin{bmatrix} X_t \\ \Xi_{t-1} \end{bmatrix}. \tag{33}$$

These matrices depend on A, B, H, D, Λ, and δ, but they are independent of C. That they are independent of C demonstrates the certainty equivalence of optimal projections (the certainty equivalence that holds when the model is linear, the loss function is quadratic, and the shocks and the uncertainty are additive); only probability means of current and future variables are needed to determine optimal policy and the optimal projection. The n_x-vector Ξ_t consists of the Lagrange multipliers of the lower block of (1), the block determining the projection of the forward-looking variables. The initial value for Ξ_{t-1} is discussed in Adolfson et al. (2011).

In a backward-looking model—that is, a model without forward-looking variables—there is no vector x_t of forward-looking variables, no lower block of equations in (1), no Lagrange multiplier Ξ_t, and the vector of target variables Y_t only depends on the vector of predetermined variables X_t and the (vector of) instrument(s) i_t.

Appendix 2. The Rudebusch-Svensson Model: An Empirical Backward-Looking Model

The two equations of the model of Rudebusch and Svensson (1999) are

$$\pi_{t+1} = \alpha_{\pi 1}\pi_t + \alpha_{\pi 2}\pi_{t-1} + \alpha_{\pi 3}\pi_{t-2} + \alpha_{\pi 4}\pi_{t-3} + \alpha_y y_t + z_{\pi,t+1} \tag{34}$$

$$y_{t+1} = \beta_{y1} y_t + \beta_{y2} y_{t-1} - \beta_r \left(\frac{1}{4} \Sigma_{j=0}^3 i_{t-j} - \frac{1}{4} \Sigma_{j=0}^3 \pi_{t-j} \right) + z_{y,t+1}, \tag{35}$$

where π_t is quarterly inflation in the GDP chain-weighted price index (P_t) in percentage points at an annual rate, i.e., $400(\ln P_t - \ln P_{t-1})$; i_t is the quarterly average federal funds rate in percentage points at an annual rate; and y_t is the relative gap between actual real GDP (Q_t) and potential GDP (Q_t^*) in percentage points, i.e., $100(Q_t - Q_t^*)/Q_t^*$. These five variables were demeaned prior to estimation, so no constants appear in the equations.

The estimated parameters, using the sample period 1961:Q1 to 1996:Q2, are shown in table 1.

Table 1. Estimated Parameters, Rudebusch-Svensson Model

$\alpha_{\pi 1}$	$\alpha_{\pi 2}$	$\alpha_{\pi 3}$	$\alpha_{\pi 4}$	α_y	β_{y1}	β_{y2}	β_r
0.70 (0.08)	−0.10 (0.10)	0.28 (0.10)	0.12 (0.08)	0.14 (0.03)	1.16 (0.08)	−0.25 (0.08)	0.10 (0.03)

The hypothesis that the sum of the lag coefficients of inflation equals one has a p-value of .16, so this restriction was imposed in the estimation.

The state-space form can be written

$$\begin{bmatrix} \pi_{t+1} \\ \pi_t \\ \pi_{t-1} \\ \pi_{t-2} \\ y_{t+1} \\ y_t \\ i_t \\ i_{t-1} \\ i_{t-2} \end{bmatrix} = \begin{bmatrix} \sum_{j=1}^4 \alpha_{\pi j} e_j + \alpha_y e_5 \\ e_1 \\ e_2 \\ e_3 \\ \beta_r e_{1:4} + \beta_{y1} e_5 + \beta_{y2} e_6 - \beta_r e_{7:9} \\ e_5 \\ e_0 \\ e_7 \\ e_8 \end{bmatrix} \begin{bmatrix} \pi_t \\ \pi_{t-1} \\ \pi_{t-2} \\ \pi_{t-3} \\ y_t \\ y_{t-1} \\ i_{t-1} \\ i_{t-2} \\ i_{t-3} \end{bmatrix} + \begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \\ -\frac{\beta_r}{4} \\ 0 \\ 1 \\ 0 \\ 0 \end{bmatrix} i_t + \begin{bmatrix} z_{\pi,t+1} \\ 0 \\ 0 \\ 0 \\ z_{y,t+1} \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix},$$

where e_j ($j = 0, 1, \ldots, 9$) denotes a 1×9 row vector, for $j = 0$ with all elements equal to zero, for $j = 1, \ldots, 9$ with element j equal to unity and all other elements equal to zero; and where $e_{j:k}$ ($j < k$)

denotes a 1×9 row vector with elements $j, j+1, \ldots, k$ equal to $\frac{1}{4}$ and all other elements equal to zero. The predetermined variables are π_t, π_{t-1}, π_{t-2}, π_{t-3}, y_t, y_{t-1}, i_{t-1}, i_{t-2}, i_{t-2}, and i_{t-3}. There are no forward-looking variables.

For a loss function (27) with $\delta = 1$, $\lambda = 1$, and $\nu = 0.2$, and the case where z_t is an i.i.d. zero-mean shock, the optimal instrument rule is as follows (the coefficients are rounded to two decimal points):

$$\begin{aligned} i_t &= 1.22\,\pi_t + 0.43\,\pi_{t-1} + 0.53\,\pi_{t-2} + 0.18\,\pi_{t-3} + 1.93\,y_t - 0.49\,y_{t-1} \\ &\quad + 0.36\,i_{t-1} - 0.09\,i_{t-2} - 0.05\,i_{t-3}. \end{aligned}$$

Appendix 3. The Lindé Model: An Empirical New Keynesian Model

An empirical New Keynesian model estimated by Lindé (2005) is

$$\begin{aligned} \pi_t &= \omega_f \pi_{t+1|t} + (1-\omega_f)\pi_{t-1} + \gamma y_t + \varepsilon_{\pi t}, \\ y_t &= \beta_f y_{t+1|t} + (1-\beta_f)(\beta_{y1} y_{t-1} + \beta_{y2} y_{t-2} + \beta_{y3} y_{t-3} + \beta_{y4} y_{t-4}) \\ &\quad - \beta_r (i_t - \pi_{t+1|t}) + \varepsilon_{yt}, \end{aligned}$$

where the restriction $\sum_{j=1}^{4} \beta_{yj} = 1$ is imposed and $\varepsilon_t \equiv (\varepsilon_{\pi t}, \varepsilon_{yt})'$ is an i.i.d. shock with mean zero. The estimated parameters (table 6a in Lindé 2005, non-farm business output) are shown in table 2.

Table 2. Estimated Parameters, Lindé Model

ω_f	γ	β_f	β_r	β_{y1}	β_{y2}	β_{y3}
0.457 (0.065)	0.048 (0.007)	0.425 (0.027)	0.156 (0.016)	1.310 (0.174)	−0.229 (0.279)	−0.011 (0.037)

For simplicity, we set $\beta_{y1} = 1$, $\beta_{y2} = \beta_{y3} = \beta_{y4} = 0$. Then the state-space form can be written as

$$\begin{bmatrix} \varepsilon_{\pi,t+1} \\ \varepsilon_{y,t+1} \\ \pi_t \\ y_t \\ i_t \\ \omega_f \pi_{t+1|t} \\ \beta_r \pi_{t+1|t} + \beta_f y_{t+1|t} \end{bmatrix}$$

$$= \begin{bmatrix} 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ -1 & 0 & -(1-\omega_f) & 0 & 0 & 1 & -\gamma \\ 0 & -1 & 0 & -(1-\beta_f) & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_{\pi t} \\ \varepsilon_{yt} \\ \pi_{t-1} \\ y_{t-1} \\ i_{t-1} \\ \pi_t \\ y_t \end{bmatrix} + \begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \\ 1 \\ 0 \\ \beta_r \end{bmatrix} i_t + \begin{bmatrix} \varepsilon_{\pi,t+1} \\ \varepsilon_{y,t+1} \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}.$$

The predetermined variables are $\varepsilon_{\pi t}$, ε_{yt}, π_{t-1}, y_{t-1}, and i_{t-1}, and the forward-looking variables are π_t and y_t.

For a loss function (27) with $\delta = 1$, $\lambda_y = 1$, and $\lambda_{\Delta i} = 0.2$, and the case where ε_t is an i.i.d. zero-mean shock, the optimal instrument rule is as follows (the coefficients are rounded to two decimal points):

$$\begin{aligned} i_t = {} & 1.06\,\varepsilon_{\pi t} + 1.38\,\varepsilon_{yt} + 0.58\,\pi_{t-1} + 0.78\,y_{t-1} + 0.40\;i_{t-1} \\ & + 0.02\,\Xi_{\pi,t-1,t-1} + 0.20\,\Xi_{y,t-1,t-1}, \end{aligned}$$

where $\Xi_{\pi,t-1,t-1}$ and $\Xi_{y,t-1,t-1}$ are the Lagrange multipliers for the two equations for the forward-looking variables in the decision problem in period $t-1$. The difference equation (9) for the Lagrange multipliers is

$$\begin{bmatrix} \Xi_{\pi t} \\ \Xi_{yt} \end{bmatrix} = \begin{bmatrix} 10.20 & 0.74 & 5.54 & 0.43 & -0.21 \\ 0.74 & 1.48 & 0.40 & 0.85 & -0.28 \end{bmatrix} \begin{bmatrix} \varepsilon_{\pi t} \\ \varepsilon_{yt} \\ \pi_{t-1} \\ y_{t-1} \\ i_{t-1} \end{bmatrix}$$

$$+ \begin{bmatrix} 0.72 & 0.16 \\ 0.03 & 0.38 \end{bmatrix} \begin{bmatrix} \Xi_{\pi,t-1} \\ \Xi_{y,t-1} \end{bmatrix}.$$

Appendix 4. Unanticipated Policy Rate Shocks: "Modest Interventions" as in Leeper and Zha (2003)

The method of "modest interventions" of Leeper and Zha (2003) can be interpreted as generating central bank projections that satisfy the restriction on the policy rate by adding a sequence of additive shocks to the instrument rule. These planned shocks are unanticipated by the private sector.

In order to illustrate the Leeper and Zha (2003) method of modest interventions, we set $T = 0$, in which case

$$z_t = \eta_{t,t}$$

and the deviation is a simple zero-mean i.i.d. disturbance. We can then write the projection model as perceived by the private sector as

$$\begin{bmatrix} \tilde{X}_{t+\tau+1,t} \\ \tilde{H}\tilde{x}_{t+\tau+1,t} \end{bmatrix} = \tilde{A} \begin{bmatrix} \tilde{X}_{t+\tau,t} \\ \tilde{x}_{t+\tau,t} \end{bmatrix} \tag{36}$$

for $\tau \geq 0$, where

$$\tilde{X}_t \equiv \begin{bmatrix} X_t \\ z_t \end{bmatrix}, \quad \tilde{x}_t \equiv \begin{bmatrix} x_t \\ i_t \end{bmatrix}, \quad \tilde{H} \equiv \begin{bmatrix} H & 0 \\ G_x & G_i \end{bmatrix},$$

$$\tilde{A} \equiv \begin{bmatrix} A_{11} & 0_{n_X \times 1} & A_{12} & B_1 \\ 0_{1 \times n_X} & 0_{1 \times 1} & 0_{1 \times n_x} & 0_{1 \times 1} \\ A_{21} & 0_{n_x \times 1} & A_{22} & B_2 \\ f_X & 1 & f_x & f_i \end{bmatrix}.$$

The solution to this system can be written

$$\begin{bmatrix} X^p_{t+\tau,t} \\ 0 \end{bmatrix} = M^\tau \tilde{X}_{t,t},$$

$$\tilde{x}^p_{t+\tau,t} \equiv \begin{bmatrix} x^p_{t+\tau,t} \\ i^p_{t+\tau,t} \end{bmatrix} = F \begin{bmatrix} X^p_{t+\tau,t} \\ 0 \end{bmatrix} = \begin{bmatrix} F_x \\ F_i \end{bmatrix} M^\tau \tilde{X}_{t,t}$$

for $\tau \geq 0$, where the superscript p denotes that this is the projection believed by the private sector in period t.

Let us demonstrate the method of modest interventions only for the restriction (16). The central bank plans to satisfy this restriction by a sequence of shocks $\{\tilde{\eta}_{t+\tau,t}\}_{\tau=0}^{T}$ that are unanticipated by the private sector. These shocks are chosen such that $\tilde{\eta}_{t,t}$ satisfies

$$i_{t,t} = F_i \begin{bmatrix} X_{t,t} \\ \tilde{\eta}_{t,t} \end{bmatrix} = \bar{\imath}_{t,t}.$$

Then the projection of the current forward-looking variables is given by

$$x_{t,t} = F_x \begin{bmatrix} X_{t,t} \\ \tilde{\eta}_{t,t} \end{bmatrix}.$$

For $\tau = 1, \ldots, T$, the projection of the predetermined variables is then given by

$$\begin{bmatrix} X_{t+\tau,t} \\ 0 \end{bmatrix} = M \begin{bmatrix} X_{t+\tau-1,t} \\ \tilde{\eta}_{t+\tau-1,t} \end{bmatrix},$$

the shock $\tilde{\eta}_{t+\tau,t}$ is chosen to satisfy

$$i_{t+\tau,t} = F_i \begin{bmatrix} X_{t+\tau,t} \\ \tilde{\eta}_{t+\tau,t} \end{bmatrix} = \bar{\imath}_{t+\tau,t},$$

and the projection of the forward-looking variables is given by

$$x_{t+\tau,t} = F_x \begin{bmatrix} X_{t+\tau,t} \\ \tilde{\eta}_{t+\tau,t} \end{bmatrix}.$$

There are some conceptual difficulties in a central bank announcing such a policy rate path and projection to the private sector. The projection is only relevant if the private sector does not believe that the central bank will actually implement the path but instead follow the instrument rule with zero expected shocks to the instrument rule. The method of modest interventions is instead perhaps more appropriate for secret policy simulations and plans that are not announced to the private sector, or for a situation when the announced policy rate path is not credible and the private sector is surprised each period when the path is implemented.

References

Adolfson, M., S. Laséen, J. Lindé, and L. E. O. Svensson. 2011. "Optimal Monetary Policy in an Operational Medium-Sized DSGE Model." Forthcoming in *Journal of Money, Credit, and Banking*.

Adolfson, M., S. Laséen, J. Lindé, and M. Villani. 2007a. "Bayesian Estimation of an Open Economy DSGE Model with Incomplete Pass-Through." *Journal of International Economics* 72 (2): 481–511.

———. 2007b. "RAMSES — A New General Equilibrium Model for Monetary Policy Analysis." *Economic Review* (Sveriges Riksbank) 2007 (2): 5–39.

———. 2008. "Evaluating an Estimated New Keynesian Small Open Economy Model." *Journal of Economic Dynamics and Control* 32 (8): 2690–2721.

Anderson, G. S., and G. Moore. 1983. "An Efficient Procedure for Solving Linear Perfect Foresight Models." Working Paper. Available at http://www.frb.gov.

———. 1985. "A Linear Algebraic Procedure for Solving Linear Perfect Foresight Models." *Economics Letters* 17 (3): 247–52.

Boucekkine, R. 1995. "An Alternative Method for Solving Nonlinear Forward-Looking Models." *Journal of Economic Dynamics and Control* 19 (4): 711–34.

Dynare. 2009. *Dynare Manual*, version 4.1.0. Available at http://www.dynare.org.

Fair, R. C., and J. B. Taylor. 1983. "Solution and Maximum Likelihood Estimation of Dynamic Nonlinear Rational Expectations Models." *Econometrica* 51 (4): 1169–85.

Gagnon, J. E., and D. W. Henderson. 1990. "Nominal Interest Rate Pegging under Alternative Expectations Hypotheses." In *Financial Sectors in Open Economies: Empirical Analysis and Policy Issues*, ed. P. Hopper, K. H. Johnson, D. L. Kohn, D. E. Lindsey, R. D. Porter, and R. Tryon, 437–73. Washington, DC: Board of Governors of the Federal Reserve System.

Galí, J. 2010. "Are Central Banks' Projections Meaningful?" Working Paper. Available at http://www.crei.cat/people/gali.

Giannoni, M. P., and M. Woodford. 2010. "Optimal Targeting Criteria for Stabilization Policy." Working Paper.

Juillard, M. 1996. "DYNARE: A Program for the Resolution and Simulation of Dynamic Models with Forward Variables Through the Use of a Relaxation Algorithm." CEPREMAP Working Paper No. 9602.

Klein, P. 2000. "Using the Generalized Schur Form to Solve a Multivariate Linear Rational Expectations Model." *Journal of Economic Dynamics and Control* 24 (10): 1405–23.

Laffargue, J.-P. 1990. "Résolution d'un modèle macroéconomique avec anticipations rationnelles." *Annales d'Économie et de Statistique* 17: 97–119.

Leeper, E. M., and T. Zha. 2003. "Modest Policy Interventions." *Journal of Monetary Economics* 50 (8): 1673–1700.

Lindé, J. 2005. "Estimating New-Keynesian Phillips Curves: A Full Information Maximum Likelihood Approach." *Journal of Monetary Economics* 52 (6): 1135–49.

Rudebusch, G. D., and L. E. O. Svensson. 1999. "Policy Rules for Inflation Targeting." In *Monetary Policy Rules*, ed. J. B. Taylor. Chicago: University of Chicago Press.

Sargent, T. J., and N. Wallace. 1975. "Rational Expectations, the Optimal Monetary Instrument and the Optimal Money Supply Rule." *Journal of Political Economy* 83 (2): 241–54.

Sims, C. A. 2000. "Solving Linear Rational Expectations Models." Working Paper.

Söderlind, P. 1999. "Solution and Estimation of RE Macromodels with Optimal Policy." *European Economic Review* 43 (4–6): 813–23. Related software available at http://home.tiscalinet.ch/paulsoderlind.

Svensson, L. E. O. 1999. "Inflation Targeting as a Monetary Policy Rule." *Journal of Monetary Economics* 43 (3): 607–54.

———. 2005. "Monetary Policy with Judgment: Forecast Targeting." *International Journal of Central Banking* 1 (1): 1–54.

———. 2009. "Optimization under Commitment and Discretion, the Recursive Saddlepoint Method, and Targeting Rules and Instrument Rules: Lecture Notes." Available at http://www.larseosvensson.net.

———. 2010. "Some Problems with Swedish Monetary Policy and Possible Solutions." Speech given in Stockholm, Sweden, November 24. Available at http://www.larseosvensson.net.

Svensson, L. E. O., and R. J. Tetlow. 2005. "Optimal Policy Projections." *International Journal of Central Banking* 1 (3): 177–207.

Svensson, L. E. O., and M. Woodford. 2005. "Implementing Optimal Policy through Inflation-Forecast Targeting." In *The Inflation-Targeting Debate*, ed. B. S. Bernanke and M. Woodford, 19–83. Chicago: University of Chicago Press.

Woodford, M. 2003. *Interest Rates and Prices.* Princeton, NJ: Princeton University Press.

A Bivariate Model of Federal Reserve and ECB Main Policy Rates

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This paper studies when and by how much the Federal Reserve and the European Central Bank change their target interest rates. I develop a new non-linear bivariate framework, which allows for elaborate dynamics and potential interdependence between the two countries, as opposed to linear feedback rules, such as a Taylor rule, and I use a novel real-time data set. Although the data sample is inherently small, through a Bayesian estimation approach, I find some evidence in favor of timing synchronization between central banks and against the hypothesis of follower behaviors. Results for the magnitude model support zero correlation in the size of the target rate changes. Institutional factors and inflation represent relevant variables for both timing and magnitude decisions, while output plays a secondary role.

JEL Codes: C11, C3, C52, E52.

1. Introduction

This paper focuses on the Federal Reserve (Fed) and the European Central Bank (ECB) interest rate feedback rules. In particular, it

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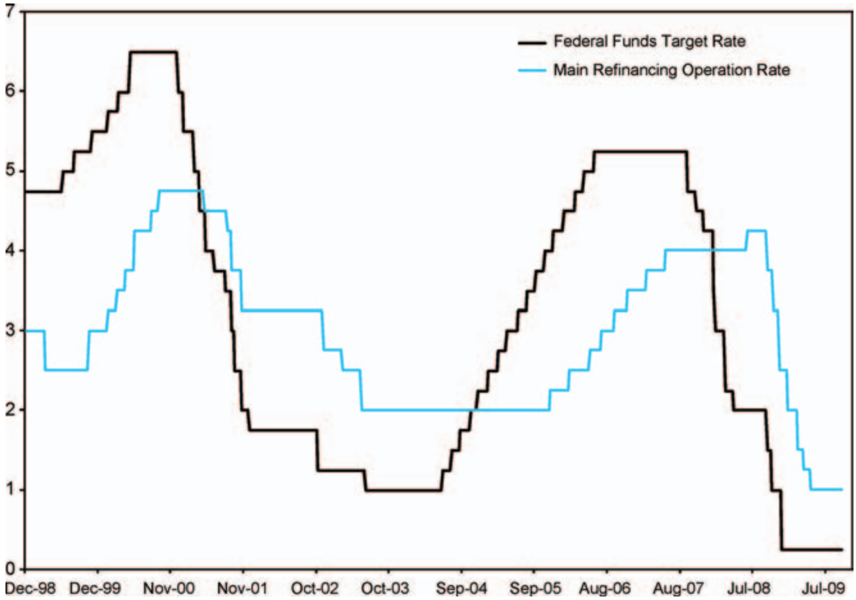
develops an econometric model that empirically analyzes when a target rate change is adopted (timing), as well as by how much the rate is changed (magnitude change). While central banks' behavior has typically been described with the use of univariate linear interest rate feedback rules, such as Taylor rules, I exploit a non-linear bivariate framework, which allows for elaborate dynamics and for potential interactions between the two central banks.

This study is appealing for several reasons. The most important way the Fed and the ECB achieve their monetary policy goals is by setting, respectively, the federal funds target rate (FFTR) and the main refinancing operation (MRO) rate.[1] These policy rates are important because they signal the stance of monetary policy, affect investment decisions, and often have considerable impact on financial markets. Understanding the way the two central banks set their target rates and identifying the variables taken into account in the process is of great interest. Since it started to operate at the beginning of 1999, the ECB, together with the Fed, has been meticulously scrutinized in the way it conducts monetary policy. Bartolini and Prati (2003) analyze the Fed and the ECB with particular attention to institutional structures, policy frameworks, and operational procedures. Cecchetti and O'Sullivan (2003) compare the central banks' approaches to the execution of monetary policy. U.S. and European Monetary Union (EMU) policy rates followed a roughly similar pattern over the period January 1999 to December 2009; see figure 1. The EMU rate fluctuates over a narrower range than the U.S. rate; both rates are characterized by frequent changes in the first half of the sample and sporadic changes in the middle part of the sample; and the FFTR displays frequent changes also in the last part of the sample. In addition, the size and sign of interest rate changes display some similarities. My analysis addresses these issues by studying

[1]In particular, the Federal Open Market Committee (FOMC) implements its monetary policy decisions by changing its target for the federal funds rate (FFR), which is the rate at which depository institutions borrow and lend reserves to and from each other overnight. Although the Federal Reserve does not control the FFR directly, it can do so indirectly by varying the supply of reserves available to be traded in the market. On the other hand, the key policy rate set by the Governing Council of the ECB is the rate applied to main refinancing operations, which provide the bulk of liquidity to the financial system.

Figure 1. Evolution of the Federal Funds Target Rate (FFTR) and of the Main Refinancing Operation (MRO) Rate from the Beginning of 1999 until the End of 2009



when interest rate changes are implemented and by examining the size and sign of those changes, with the idea that the two decisions could carry distinct information and might be triggered by different variables.[2]

Open questions in the monetary policy debate are whether in an open economy interest rate feedback rules should include the exchange rate, in addition to inflation and output, and, more generally, how optimal interest rate feedback rules should be designed within an international framework; see, among others, Benigno (2002), Clarida, Galí, and Gertler (2002), Pappa (2004), and Corsetti and Pesenti (2005). While the theoretical literature has focused on optimality

[2]In a traditional approach (e.g., VAR or Taylor rule), the same variables affect both timing and magnitude because the two decisions are normally analyzed together. That is, when there is a jump/change in the rate (magnitude decision), the event is considered to occur (timing decision).

issues, this paper introduces a new methodology to provide evidence about the interaction between the Fed and the ECB. It does not address issues related to cooperation or potential gains from cooperation. This paper only provides stylized facts about Fed and ECB interest rate feedback rules, exploring the possibility that interdependence could play a role in describing timing and magnitude of interest rate changes.

Moreover, it is not clear that conventional linear interest rate feedback rules are sufficient to explain the complexity of central banks' behavior, especially in the presence of potential interdependences. I therefore investigate the possibility that a non-linear model could better describe interest rate decisions.

Methodologically, one novelty of the paper is to provide a bivariate autoregressive conditional hazard (BACH) model to study the timing of interest rate changes and a conditional bivariate ordered (CBO) probit model to analyze the magnitude of interest rate changes. The BACH model extends the autoregressive conditional hazard (ACH) model of Hamilton and Jordà (2002) in order to account for interdependence between the two central banks. The timing/duration framework is based on the autoregressive conditional duration (ACD) model developed by Engle and Russell (1998, 2002, 2005) and Engle (2000). Bergin and Jordà (2004) analyze empirical evidence of monetary policy interdependence within a set of fourteen OECD countries, making use of the Hamilton and Jordà (2002) model, but they do not analyze the EMU. They study interdependence by investigating whether the probability of a change in the domestic target rate at time t depends on a similar decision by a "leading country" at time $t-1$ (United States, Germany, or Japan). This implies a hierarchy between banks and assumes that the leading central bank's decision is known by the other central bank. Moreover, their setup does not allow them to recover a joint hazard probability for the two central banks. My model differs from Bergin and Jordà (2004) because it is a truly bivariate model which converts the marginal hazards to a joint hazard probability through the use of a conditional discrete copula-type representation. This allows me to treat the Fed and ECB symmetrically and to study interdependence in the form of decision synchronization and follower behaviors. The CBO probit model represents a special case of a bivariate ordered probit model, where I rescale the probability mass to condition on the timing decision. It differs from Bergin and Jordà (2004) because, by rescaling, I am able to study the exact magnitude of the change (basis-points change) as opposed to merely the direction of change (strong increase, increase, decrease, strong decrease).

According to the Taylor-type rule literature, past interest rates, inflation, output gaps, and exchange rate movements are relevant factors in choosing the level of interest rate changes. I analyze whether these variables are important in determining when the Fed and ECB change their interest rates, as well as whether they play a role in explaining the magnitude of the change. Moreover, I study whether the Fed and the ECB synchronize their policies, whether one follows the other, and whether there exists a contemporaneous correlation in the magnitude of their interest rate changes. Timing synchronization of policies is analyzed with the odds ratio, which indicates how much the odds of one bank changing its target rate move when the other bank changes its target. Follower behavior is studied with dummy variables that capture the effect of one country's interest rate decision on the subsequent decisions of the other country. A test on the coefficient of this dummy variable can be interpreted as a test of one country (Granger) causing the other country's interest rate decision, and hence one country following the other country's decision. The correlation between interest rate changes captures the correlation which is left unexplained after traditional explanatory variables have been considered. This paper shows whether the traditional variables that have commonly been used in the literature are sufficient in explaining timing and magnitude changes of policy rates, and whether the interdependence could be a factor in explaining monetary policies. However, it does not provide a complete answer to the underlying problem about what is in fact the source of the interdependence and whether interdependence is optimal.

Another novel feature of the paper is the empirical application with the use of a real-time data set and the Bayesian estimation. Persuaded that the available information set that central banks observe is of great importance to the decisions they make, I construct and use a real-time data set that includes output and inflation measures, exchange rates, and data on target rates and duration between changes. This real-time approach to monetary policy has been studied, among others, by Orphanides (2001), who demonstrates that real-time policy recommendations differ from those derived with ex post revised data.[3] Bayesian estimation is not new to monetary policy studies (see Sims and Zha 1998, Schorfheide 2000, and Cogley and Sargent 2002), but, to the best of my knowledge, it has never been applied to ACD- or ACH-type models. The methods used in the paper generally require fairly large samples to produce results. Because of the youth of the ECB and the type of data, the sample used in the paper is small. The Bayesian framework is particularly well suited to this small-sample problem, because it allows me to incorporate pre-sample information to better evaluate the available information. I use ten years of data for the Fed and the Bundesbank to elicit the prior, following the view that the German central bank is, among the European central banks, the one that most closely resembles the ECB. Although the Bayesian approach facilitates the estimation, it does not completely eliminate the small-sample problem.

Estimation results for the timing model support the hypothesis that institutional factors, such as scheduled meetings of the FOMC and the Governing Council, as well as inflation rates, are important variables in determining timing decisions. I find some evidence of timing synchronization between the two central banks. On the other hand, follower behaviors are not supported. However, although there is strong evidence against the Fed following the ECB timing strategy, the evidence is slim against the reverse scenario of the ECB following the Fed's timing decisions. Estimation results for the magnitude model illustrate the importance of inflation rates as explanatory variables for both countries. Unemployment and exchange rate dynamics turn out to have a secondary role, confirming the idea that the ECB's primary objective is to maintain price stability. I find evidence supporting zero correlation between the magnitude shocks. The zero correlation suggests that although there is *synchronization* in the timing of interest rate changes, its own macroeconomic conditions are sufficient to explain the magnitude of each central bank's target rate change.

The paper is organized as follows. The next section describes the model. Section 3 describes the data. Section 4 describes the

[3]It could be interesting to compare baseline results based on real-time data with results obtained by estimating the model using revised data. However, this goes beyond the scope of this paper. Many papers in the literature have in fact already addressed this issue.

Bayesian implementation and presents empirical results. Section 5 presents results for posterior predictive checks. Section 6 concludes.

2. The Model

For simplicity I refer to the EMU and the United States as countries e and f (for ECB and Fed). The basic idea is to separate timing and magnitude of interest rate changes, and to derive a model capable of accounting for the specific features of both decisions. I describe timing by binary variables that take the value one when the target interest rate is changed. Consequently, magnitude variables take non-zero values only when the timing binary variable is one.

More precisely, let x_t^i be a binary variable that takes values $\{0,1\}$ according to whether the target rate of country $i \in \{e,f\}$ has changed at calendar time t:

$$x_t^i = \begin{cases} 1 & \text{target rate of country } i \text{ is changed} \\ 0 & \text{otherwise,} \end{cases} \tag{1}$$

and let y_t^i be the interest rate change that takes place whenever event x^i occurs (i.e., whenever $x^i = 1$).

I am interested in studying whether the two central banks decide to change their target rate (x) and by how much (y); hence I want to study the joint probability $f(x_t^e, x_t^f, y_t^e, y_t^f | \mathcal{F}_{t-1})$, where $\mathcal{F}_{t-1}$ is the information set available at time $t-1$. The joint probability can be rewritten as the product of the marginal distribution of (x^e, x^f) and the conditional distribution of $(y^e, y^f |\ x^e, x^f)$:

$$\begin{aligned} & f\big(x_t^e, x_t^f, y_t^e, y_t^f | \mathcal{F}_{t-1}; \theta\big) \\ & \quad = g\big(x_t^e, x_t^f | \mathcal{F}_{t-1}; \theta_1\big) \cdot q\big(y_t^e, y_t^f | x_t^e, x_t^f, \mathcal{F}_{t-1}; \theta_2\big) \end{aligned} \tag{2}$$

so that the resulting log-likelihood can be decomposed into two parts,

$$\mathcal{L}(\theta_1, \theta_2) = \mathcal{L}_1(\theta_1) + \mathcal{L}_2(\theta_2), \tag{3}$$

where

$$\mathcal{L}_1(\theta_1) = \sum_{t=1}^{T} \log g\big(x_t^e, x_t^f | \mathcal{F}_{t-1}; \theta_1\big) \quad \text{Timing Model} \tag{4}$$

$$\mathcal{L}_2(\theta_2) = \sum_{t=1}^{T} \log q\big(y_t^e, y_t^f | x_t^e, x_t^f, \mathcal{F}_{t-1}; \theta_2\big) \quad \text{Level Model.} \tag{5}$$

As pointed out by Engle (2000), if θ_1 and θ_2 have no parameters in common and are variation free, then the maximization of $\mathcal{L}(\theta_1, \theta_2)$ is equivalent to maximizing $\mathcal{L}_1(\theta_1)$ and $\mathcal{L}_2(\theta_2)$ separately. The parameters (θ_1, θ_2) are variation free, as in Engle, Hendry, and Richard (1983), if θ_1 and θ_2 are not subject to cross-restrictions, meaning that the range of admissible values for θ_1 does not vary with θ_2, and vice versa. In a maximum-likelihood environment, this means that the separate maximization of $\mathcal{L}_1(\theta_1)$ and $\mathcal{L}_2(\theta_2)$ is equivalent to maximizing $\mathcal{L}(\theta_1, \theta_2)$. The same decomposition would work in a Bayesian framework with independent priors since

$$\max\left\{\frac{\mathcal{L}_1(Y|\theta_1)\mathcal{L}_2(Y|\theta_2)\Pr(\theta_1, \theta_2)}{\Pr(Y)}\right\} = \max\left\{\frac{\mathcal{L}_1(Y|\theta_1)\Pr(\theta_1)}{\Pr(Y)}\right\}\max\left\{\frac{\mathcal{L}_2(Y|\theta_2)\Pr(\theta_2)}{\Pr(Y)}\right\}$$

if the priors are independent and (θ_1, θ_2) are variation free. I follow this strategy and I refer to the first part of the likelihood as the *timing model* and to the second part as the *level model.* The former is characterized as a bivariate autoregressive conditional hazard (BACH) model, while the latter is a conditional bivariate ordered (CBO) probit model.

I describe both models below.

2.1 Bivariate Autoregressive Conditional Hazard (BACH) Model

The *timing model* hinges on the joint probability of type e and f events occurring, where type i event occurring means that country i, $i \in \{e, f\}$, has decided to change its target rate. Marginal probability distributions for individual interest rate decisions have been modeled in the literature with autoregressive conditional hazard (ACH) models; see Hamilton and Jordà (2002). The ACH model is derived from the autoregressive conditional duration (ACD) model proposed by Engle and Russell (1998) and Engle (2000). The ACD model is developed in event time[4] and aims to explain the duration of spells between events (between two consecutive trades or quotes, for example). It is called autoregressive conditional duration because the conditional expectation of the duration depends upon past durations. Within this duration framework, bivariate models have been studied, but none of them is suitable to the present framework. Engle and Lunde (2003), for example, model the joint likelihood function for trade and quote arrivals, but they include the possibility that an intervening trade censors the time between a trade and the subsequent quote.[5] Thus their model does not serve my purpose. Moreover, unlike them, I adopt calendar time because it readily allows me to incorporate updated explanatory variables.

I develop a bivariate model which converts the marginal distribution information, modeled following Hamilton and Jordà (2002), into a joint distribution, by using a conditional discrete copula-type representation.

2.1.1 Marginal Hazard Rates

Define $N^e(t)$ and $N^f(t)$ to be, respectively, the cumulative number of country e and f events as of time t, i.e., the number of target rate changes of country e and f as of time t. Following Hamilton and Jordà (2002), I rewrite the $ACD(g, m)$ model from Engle and Russel (1998) as

$$\psi^i_{N^i(t)} = \sum_{j=1}^{m^i} \alpha^i_j \; u^i_{N^i(t)-j} + \sum_{j=1}^{g^i} \beta^i_j \; \psi^i_{N^i(t)-j}$$
$$i = e, f, \tag{6}$$

where α^i and β^i are country-specific parameters, $\psi^i_{N^i(t)}$ is the expected duration for country i at calendar time t when $N^i(t)$ events have occurred, and $u^i_{N^i(t)-j}$ is the duration for country i when

[4]Event time is defined by a sequence $\{t_0, t_1, .., t_n, \ldots\}$ with $t_0 < t_1 < \ldots < t_n < \ldots$ representing the arrival time of an event. Calendar time is simply $t = 0, 1, 2, 3, \ldots, n, \ldots$

[5]In particular, they analyze the elapsed times between two consecutive trades and between a trade and a quote.

$N^i(t) - j$ events have occurred; i.e., $u^i_{N^i(t)-j}$ is the time elapsed between event $N^i(t) - j - 1$ and event $N^i(t) - j$. Viewed as a function of time, $\psi^i_{N^i(t)}$ is a step function that changes only when a new event occurs, i.e., when $N^i(t) \neq N^i(t-1)$.[6]

Define the hazard rate h^i_t as the probability of a country i event occurring at time t (the probability that the central bank of country i decides to change its target rate), given the information available up until time $t-1$, i.e., $\Pr(x^i_t = 1|\mathcal{F}_{t-1})$. The country i marginal hazard rate can be written as

$$\begin{aligned} h^i_{t|t-1} &= \Pr[N^i(t) \neq N^i(t-1)|\mathcal{F}_{t-1}] = \Pr\left[x^i_t = 1\middle|\mathcal{F}_{t-1}\right] \\ &= \frac{1}{\psi^i_{N^i(t-1)} + \delta'_i z_{t-1}}, \end{aligned} \tag{7}$$

where

$$\begin{aligned} \psi^i_{N^i(t-1)} &= \alpha^i\ u^i_{N^i(t-1)-1} + \beta^i \psi^i_{N^i(t-1)-1} \\ g &= m = 1. \end{aligned} \tag{8}$$

z_{t-1} is a vector of variables known at $t-1$ and δ_i is a parameter vector. I need to ensure that $h^i_{t|t-1} \in (0,1)$. To do so I follow Hamilton and Jordà (2002) and I use a smooth function λ so that

$$h^i_{t|t-1} = \frac{1}{\lambda\big(\psi^i_{N^i(t-1)} + \delta'_i z_{t-1}\big)},$$

where

$$\lambda(\nu) = \begin{matrix} 1.0001 & \nu \leq 1 \\ 1.001 + 2\Delta_0(\nu-1)^2/\left[\Delta_0^2 + (\nu-1)^2\right] & 1 < \nu < 1 + \Delta_0 \\ 0.0001 + \nu & \nu \geq 1 + \Delta_0 \end{matrix}$$

and $\Delta_0 = 0.1$.

Using the above marginal distributions for the Bernoulli variables $x^i,\ i = e, f$, I want to construct a joint distribution for (x^e, x^f). The following section gives some theoretical background about the

[6]Equation (6) does not include a constant, which is instead included in the z vector.

discrete copula-type representation that will allow me to recover the joint hazard of countries e and f.

2.1.2 Conditional Discrete Copula-Type Representation and Joint Hazard Rates

When marginal distributions are continuous, a joint distribution can be constructed from the marginal distributions with the use of a copula. The beauty of a copula is that for bivariate (multivariate) distributions, the univariate marginals and the dependence structure can be separated, with the copula containing all the dependence information. Since the marginals considered here are discrete, problems arise since copulas are not unique in this case. The way I solve the problem follows Tajar, Denuit, and Lambert (2001). As the copula contains the dependence information, Tajar, Denuit, and Lambert (2001) disentangle the dependence structure from the Bernoulli marginals.

In addition, I extend the existing results to allow for conditioning variables. For the purpose of exposition, I will assume below that $\mathcal{F}$ represents the conditioning set (it might contain one or more variables).

Let (x^e, x^f) be random Bernoulli variables for which the marginal conditional distributions $\Pr[x_t^e|\mathcal{F}]$ and $\Pr[x_t^f|\mathcal{F}]$ are known. Associate $(x^e, x^f)|\mathcal{F}$ to a random couple $(u^e, u^f)|\mathcal{F}$ with discrete uniform marginals such that

$$\Pr[u^e = 0|\mathcal{F}] = \Pr[u^e = 1|\mathcal{F}] = \Pr[u^f = 0|\mathcal{F}] = \Pr[u^f = 1|\mathcal{F}] = \frac{1}{2}. \tag{9}$$

Therefore, the joint distributions of $(x^e, x^f)|\mathcal{F}$ and $(u^e, u^f)|\mathcal{F}$ can be written as follows:

$x^e_{\|\mathcal{F}} \diagdown x^f_{\|\mathcal{F}}$	0	1	
0	$h_{00\|\mathcal{F}}$	$h_{01\|\mathcal{F}}$	$1 - h^e_{\|\mathcal{F}}$
1	$h_{10\|\mathcal{F}}$	$h_{11\|\mathcal{F}}$	$h^e_{\|\mathcal{F}}$
	$1 - h^f_{\|\mathcal{F}}$	$h^f_{\|\mathcal{F}}$	

(10)

$$\begin{array}{|c|c|c|c|}\hline u^e_{|\mathcal{F}} \diagdown u^f_{|\mathcal{F}} & 0 & 1 & \\ \hline 0 & \gamma_{00|\mathcal{F}} & \gamma_{01|\mathcal{F}} & 1/2 \\ \hline 1 & \gamma_{10|\mathcal{F}} & \gamma_{11|\mathcal{F}} & 1/2 \\ \hline & 1/2 & 1/2 & \\ \hline \end{array}. \tag{11}$$

The joint probabilities of $(x^e, x^f)|\mathcal{F}$ and $(u^e, u^f)|\mathcal{F}$ are such that

$$h_{lk|\mathcal{F}} = \left(p^e_{l|\mathcal{F}}\right) \cdot \left(p^f_{k|\mathcal{F}}\right) \cdot \left(\gamma_{lk|\mathcal{F}}\right), \; l, k = 0, 1, \tag{12}$$

where $p^e_{l|\mathcal{F}}$ and $p^f_{k|\mathcal{F}}$ depend only on the marginals, while $\gamma_{lk|\mathcal{F}}$ contains the dependence information.

I compute the p's from the marginals as

$$p^e_{0|\mathcal{F}} = \frac{\left(1 - h^e_{|\mathcal{F}}\right)}{nf_1}, \quad p^e_{1|\mathcal{F}} = \frac{h^e_{|\mathcal{F}}}{nf_2} \tag{13}$$

$$p^f_{0|\mathcal{F}} = \frac{\left(1 - h^f_{|\mathcal{F}}\right)}{nf_3}, \quad p^f_{1|\mathcal{F}} = \frac{h^f_{|\mathcal{F}}}{nf_4}, \tag{14}$$

where nf_p, $p = 1, 2, 3, 4$ are normalizing factors that guarantee $h_{00}+h_{01} = 1-h^e$, $h_{10}+h_{11} = h^e$, $h_{00}+h_{10} = 1-h^f$, $h_{01}+h_{11} = h^f$.

Let the odds ratio $\eta \in R^+$ be

$$\eta = \frac{h_{00|\mathcal{F}} \cdot h_{11|\mathcal{F}}}{h_{01|\mathcal{F}} \cdot h_{10|\mathcal{F}}} = \frac{\gamma_{00|\mathcal{F}} \cdot \gamma_{11|\mathcal{F}}}{\gamma_{01|\mathcal{F}} \cdot \gamma_{10|\mathcal{F}}}. \tag{15}$$

The odds ratio is a measure of association for binary random variables. For ease of interpretation, it can be rewritten as $\eta = \frac{(\gamma_{11|\mathcal{F}}/\gamma_{10|\mathcal{F}})}{(\gamma_{01|\mathcal{F}}/\gamma_{00|\mathcal{F}})}$, where the numerator gives the "odds" of country f event occurring versus not occurring given that country e event occurs, while the denominator gives the "odds" of country f event occurring given that country e event does not occur. Thus the odds ratio indicates how much the odds of country f changing its target rate increase when country e changes its target. Independence is $\eta = 1$.

For a pair of binary random variables with uniform marginals, the following property holds:

$$\gamma_{01|\mathcal{F}} = \frac{1}{2} - \gamma_{00|\mathcal{F}}, \quad \gamma_{10|\mathcal{F}} = \frac{1}{2} - \gamma_{00|\mathcal{F}}, \quad \gamma_{11|\mathcal{F}} = \gamma_{00|\mathcal{F}}. \tag{16}$$

Therefore, $\gamma_{00|\mathcal{F}}$ can be obtained as the solution[7] to the following quadratic equation:

$$\gamma_{00|\mathcal{F}}^2\ (\eta - 1) - \eta\ \gamma_{00|\mathcal{F}} + \frac{\eta}{4} = 0. \tag{17}$$

Hence $h_{ij|\mathcal{F}}\ i, j = 0, 1$ can be computed.

This copula-type representation allows me to construct the joint hazard rates to be used in the likelihood.

Some assumptions must be made. First, in my setup, both marginals depend on t. According to equation (15), the odds ratio should also depend on t. I instead assume that $\eta_t = \eta,\ \forall t$. Second, I define the conditioning set $\mathcal{F}$ as the information available as of time $t-1$. The conditioning set must be the same for both marginal distributions.[8]

Therefore, the joint probability of events e and f occurring (given $\mathcal{F}_{t-1}$) is

$$\begin{aligned} g\big(x_t^e x_t^f\,|\mathcal{F}_{t-1};\theta_1\big) = {} & \big(h_{00,t|t-1}\big)^{\mathit{1}\,(x_t^e=0,x_t^f=0)}\big(h_{10,t|t-1}\big)^{\mathit{1}\,(x_t^e=1,x_t^f=0)} \\ & \times \big(h_{01,t|t-1}\big)^{\mathit{1}\,(x_t^e=0,x_t^f=1)}\big(h_{11,t|t-1}\big)^{\mathit{1}\,(x_t^e=1,x_t^f=1)}, \end{aligned} \tag{18}$$

where $\mathit{1}\,(\cdot,\cdot)$ is an indicator function. Equation (18) yields the following likelihood function:

$$\begin{aligned} \mathcal{L}_1(\theta_1) = \sum_{t=1}^{T} & \mathit{1}\big(x_t^e = 0, x_t^f = 0\big)\ \log(h_{00,t|t-1}) \\ & + \mathit{1}\big(x_t^e = 1, x_t^f = 0\big)\ \log(h_{10,t|t-1}) \\ & + \mathit{1}\big(x_t^e = 0, x_t^f = 1\big)\ \log(h_{01,t|t-1}) \\ & + \mathit{1}\big(x_t^e = 1, x_t^f = 1\big)\ \log(h_{11,t|t-1}). \end{aligned} \tag{19}$$

---

[7] Only one of the two roots belongs to $[0, 1/2)$ and is therefore admissible. The root $\gamma_{00,t}^{+} \notin [0, 1/2)$.

[8] Thus in principle the marginal hazard rate for process e, h^e, depends on all the conditioning variables (even the one from process f). I will impose the restriction that process f variables have no effect on the duration. The same applies to h^f.

2.2 Conditional Bivariate Ordered (CBO) Probit Model

I use a special bivariate ordered probit model to analyze the interest rate magnitude changes (y_t^e and y_t^f). My framework is a special case of the standard bivariate ordered probit model, because I am interested in the distribution of y_t^e and y_t^f conditioned on the information set $\mathcal{W}_{t-1}$ and conditioned on the timing decision (x_t^e, x_t^f).

Assume there are two latent variables, one for each country, representing the optimal (but unobserved) target change

$$\widetilde{y}_t^e = w_{t-1}^{e\prime}\pi^e + \varepsilon_t^e \tag{20}$$

$$\widetilde{y}_t^f = w_{t-1}^{f\prime}\pi^f + \varepsilon_t^f, \tag{21}$$

where π^e and π^f are parameter vectors, w_{t-1}^e and $w_{t-1}^f \in \mathcal{W}_{t-1}$ are vectors of variables observed as of time $t-1$, and $(\varepsilon_t^e, \varepsilon_t^f)|w_{t-1}, \sim \mathcal{N}(0, \Sigma)$ with $\Sigma = \begin{bmatrix} 1 & \rho \\ \rho & 1 \end{bmatrix}$.

If the observable target change y_t^i could assume the discrete values $s^i \in \{-50, -25, 0, 25, 50\}$ measured in basis points (bps), $i = e, f$, then it would be related to the unobservable optimal target change, so that[9]

$$y_t^i = \begin{cases} s_1 \leq -50 & \text{if } \widetilde{y}_t^i \in \left(-\infty = c_0^i, c_1^i\right] \\ s_2 = -25 & \text{if } \widetilde{y}_t^i \in \left(c_1^i, c_2^i\right] \\ s_3 = 0 & \text{if } \widetilde{y}_t^i \in \left(c_2^i, c_3^i\right] \\ s_4 = 25 & \text{if } \widetilde{y}_t^i \in \left(c_3^i, c_4^i\right] \\ s_5 = 50 & \text{if } \widetilde{y}_t^i \in \left(c_4^i, c_5^i = \infty\right). \end{cases} \tag{22}$$

I observe x_t^e and x_t^f and am interested in the conditional distribution of y_t^e and y_t^f given x_t^e and x_t^f. The questions I want to address are as follows: What is the joint probability of y_t^e and y_t^f taking values $s_m, s_n \in \{-50, -25, 25, 50\}$, respectively, given $x_t^e = x_t^f = 1$ and given $\mathcal{W}_{t-1}$? What is the probability of y_t^e

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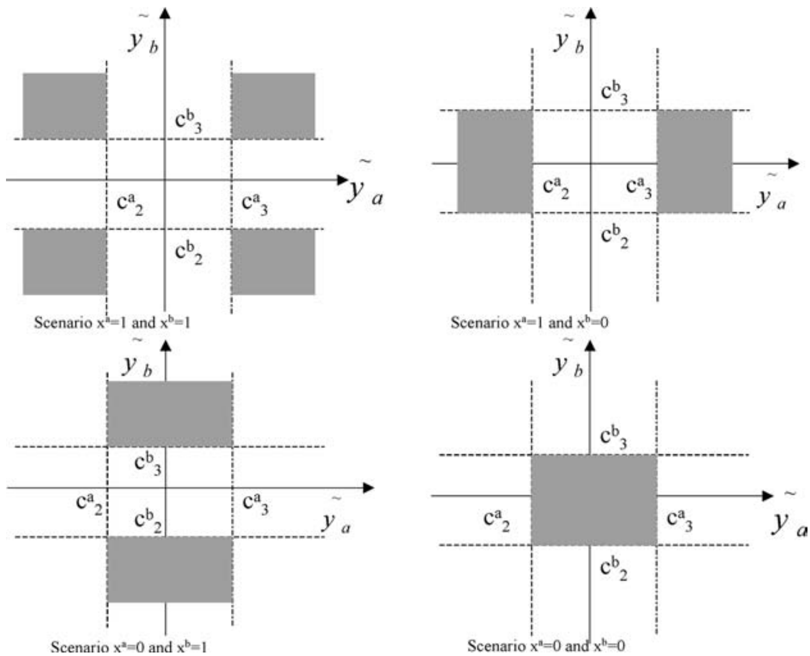
[9] In our framework, I also observe changes of −75 bps. Because these changes are not enough to identify a separate cut point, in the estimated model, I cluster the –75 bps changes together with the –50 bps changes.

Figure 2. Rescaling Necessary to Condition on (x_t^a, x_t^b)



Notes: The shaded areas of the top-left panel show the feasible regions for (y_t^a, y_t^b) when $(x_t^a = 1, x_t^b = 1)$; the top-right panel shows the feasible region for y^a when $(x_t^a = 1, x_t^b = 0)$, in which case $y^b = 0$ with probability 1; the bottom-left panel shows the feasible region for y^b when $(x_t^a = 0, x_t^b = 1)$, in which case $y^a = 0$ with probability 1; and the bottom-right panel shows the feasible region when $(x_t^a = 0, x_t^b = 0)$, in which case $y^a = y^b = 0$ with probability 1.

being equal to $s_m \in \{-50, -25, 25, 50\}$ when $y_t^f = 0$ (no change for country f occurs)? What is the probability of y_t^f being equal to $s_n \in \{-50, -25, 25, 50\}$ when $y_t^e = 0$?

Thus, starting from the bivariate normal distribution that characterizes $\binom{\widetilde{y}_t^e}{\widetilde{y}_t^f}|\mathcal{W}_{t-1}$, I want to retrieve $f(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1}, x_t^e, x_t^f)$. Conditioning on x_t^e and x_t^f, the bivariate normal density that characterizes the distributions of $(\varepsilon_t^e, \varepsilon_t^f)$ and $(\widetilde{y}_t^e, \widetilde{y}_t^f)$ is rewritten so as to redistribute the probability mass. Figure 2 contains a visual illustration of the necessary rescaling.

The log-likelihood relative to the magnitude decision can be written as

$$\mathcal{L}_2(\theta_2) = \sum_{t=1}^{T} \mathit{1}\left(x_t^e = 1, x_t^f = 0\right) \log P_{10} + \mathit{1}\left(x_t^e = 0, x_t^f = 1\right) \log P_{01} \\ + \mathit{1}\left(x_t^e = 1, x_t^f = 1\right) \log P_{11}, \tag{23}$$

where P_{10} denotes the probability of y_t^e being equal to $s_m \in \{-50, -25, 25, 50\}$ when $y_t^f = 0$ (no change for country f occurs), once I have rescaled to condition on the timing decision ($x_t^e = 1, x_t^f = 0$). A similar interpretation is given to P_{01} and P_{11}.[10] A detailed derivation of these probabilities is presented in the appendix.

3. Data

The raw data that I use to analyze Fed and ECB decisions are the dates and size of changes in the FFTR and the MRO rate. Table 1 displays the FFTR level, dates on which it was changed, and the size of the change. Table 2 displays similar data for the Eurosystem. Dummies for FOMC and Governing Council meetings have also been included. Due to the youth of the EMU, my sample spans the period January 1, 1999 to the last week of 2009 for a total of 575 weeks.

As is clear from both table 1 and table 2, the Fed has changed rates more frequently than the ECB. The average duration for the United States is about 80 days as opposed to 117 in the EMU. Figure 3 shows that there have been a total of forty-six changes in the United States and thirty-one in the EMU. In the United States, three were –75 bps changes, thirteen were –50 bps changes, seven were –25 bps changes, twenty-two were +25 bps changes, and one was a +50 bps change. In the EMU, one was a –75 bps change, nine were –50 bps changes, five were –25 bps changes, fourteen were +25 bps changes, and two were +50 bps changes. In the CBO probit estimation, I combined the –75 bps changes with –50 bps changes due to identificability issues of the cut points with so few observations.

[10]Note that, given $x_t^i = 0$, $i = e, f$, then $y^i = 0$ with probability 1. Thus $\log P_{00} = 0$.

Table 1. Calendar of Federal Funds Target Rate Changes

	FFTR	FFTR Change	Weekday	Duration in Days
17-Nov-98	4.75		Tue	
30-Jun-99	5	0.25	Wed	225
24-Aug-99	5.25	0.25	Tue	55
16-Nov-99	5.5	0.25	Tue	84
2-Feb-00	5.75	0.25	Wed	78
21-Mar-00	6	0.25	Tue	48
16-May-00	6.5	0.25	Tue	56
3-Jan-01	6	0.5	Wed	232
31-Jan-01	5.5	−0.5	Wed	28
20-Mar-01	5	−0.5	Tue	48
18-Apr-01	4.5	−0.5	Wed	29
15-May-01	4	−0.5	Tue	27
27-Jun-01	3.75	−0.25	Wed	43
21-Aug-01	3.5	−0.25	Tue	55
17-Sep-01	3	−0.5	Mon	27
2-Oct-01	2.5	−0.5	Tue	15
6-Nov-01	2	−0.5	Tue	35
11-Dec-01	1.75	−0.25	Tue	35
6-Nov-02	1.25	−0.5	Wed	330
25-Jun-03	1	−0.25	Wed	231
30-Jun-04	1.25	0.25	Wed	371
10-Aug-04	1.5	0.25	Tue	41
21-Sep-04	1.75	0.25	Wed	42
10-Nov-04	2	0.25	Wed	50
14-Dec-04	2.25	0.25	Wed	34
2-Feb-05	2.5	0.25	Wed	50
22-Mar-05	2.5	0.25	Wed	48
3-May-05	3	0.25	Tue	42
30-Jun-05	3.25	0.25	Tue	58
9-Aug-05	3.5	0.25	Thu	40
20-Sep-05	3.75	0.25	Tue	42
1-Nov-05	4	0.25	Tue	42
13-Dec-05	4.25	0.25	Tue	42
31-Jan-06	4.5	0.25	Tue	49
28-Mar-06	4.75	0.25	Tue	56
10-May-06	5	0.25	Wed	43
29-Jun-06	5.25	0.25	Thu	50

(continued)

Table 1. (Continued)

	FFTR	FFTR Change	Weekday	Duration in Days
18-Sep-07	4.75	−0.5	Tue	446
31-Oct-07	4.5	−0.25	Wed	43
11-Dec-07	4.25	−0.25	Tue	41
22-Jan-08	3.5	−0.75	Tue	42
30-Jan-08	3	−0.5	Wed	8
18-Mar-08	2.25	−0.75	Tue	48
30-Apr-08	2	−0.25	Wed	43
8-Oct-08	1.5	−0.5	Wed	161
29-Oct-08	1	−0.5	Wed	21
16-Dec-08	0.25	−0.75	Tue	48

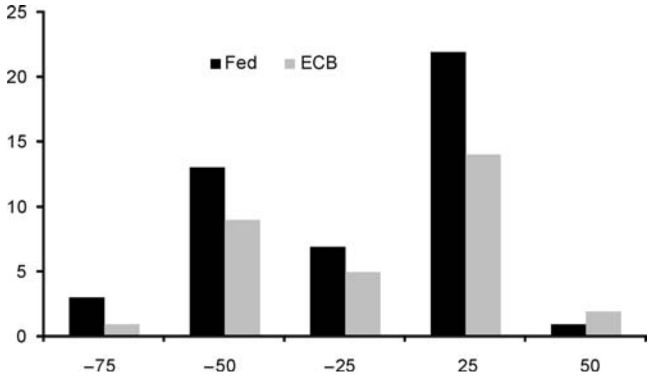
Table 2 deserves a few comments. Main refinancing operations which settled before June 28, 2000 were conducted on the basis of fixed-rate tenders, in which the ECB would specify the interest rate in advance and participating counterparties would bid the amount of money (volume) they were willing to transact at that rate. A side effect of the system was chronic overbidding by financial institutions. On June 8, 2000, the ECB announced that, starting with the operation to be settled on June 28, 2000, the main refinancing operations would be conducted as variable-rate tenders, in which counterparties would specify both the amount and the interest rate at which they want to transact. Starting from the operation to be settled on October 15, but announced on October 8, the ECB introduced a fixed-rate tender procedure with full allotment for all its refinancing operations.

Together with these key policy rates, I create dummy variables to control for the FOMC schedule in the United States and the Governing Council schedule in the EMU. These dummies are important since the majority of interest rate changes happen on these scheduled meeting dates. The Fed has made five intermeeting changes (January, April, and September 2001; January and October 2008), while the ECB has changed rates on a non-meeting day only twice, in the immediate aftermath of September 11, 2001 and in October 2008 at the height of the financial crisis.

Table 2. Calendar of MRO Rate Changes

	MRO		MRO Change	Weekday	Duration in Days
	Fix. Rate	Var. Rate			
1-Jan-99	3	—		Fri	
4-Jan-99	3	—		Mon	
21-Jan-99	3	—		Thu	
8-Apr-99	2.5	—	−0.5	Thu	77
4-Nov-99	3	—	0.5	Thu	210
3-Feb-00	3.25	—	0.25	Thu	91
16-Mar-00	3.5	—	0.25	Thu	42
27-Apr-00	3.75	—	0.25	Thu	42
8-Jun-00	4.25	—	0.5	Thu	42
28-Jun-00	—	4.25		Wed	20
31-Aug-00	—	4.5	0.25	Thu	64
5-Oct-00	—	4.75	0.25	Thu	35
10-May-01	—	4.5	−0.25	Thu	217
30-Aug-01	—	4.25	−0.25	Thu	112
17-Sep-01	—	3.75	−0.5	Mon	18
8-Nov-01	—	3.25	−0.5	Thu	52
5-Dec-02	—	2.75	−0.5	Thu	392
6-Mar-03	—	2.5	−0.25	Thu	91
5-Jun-03	—	2	−0.5	Thu	91
1-Dec-05	—	2.25	0.25	Thu	910
2-Mar-06	—	2.5	0.25	Thu	91
8-Jun-06	—	2.75	0.25	Thu	98
3-Aug-06	—	3	0.25	Thu	56
5-Oct-06	—	3.25	0.25	Thu	63
7-Dec-06	—	3.5	0.25	Thu	63
8-Mar-07	—	3.75	0.25	Thu	91
6-Jun-07	—	4	0.25	Wed	90
3-Jul-08	—	4.25	0.25	Thu	393
8-Oct-08	3.75	—	−0.50	Wed	97
6-Nov-08	3.25	—	−0.50	Thu	29
4-Dec-08	2.5	—	−0.75	Thu	28
15-Jan-09	2	—	−0.50	Thu	42
5-Mar-09	1.5	—	−0.50	Thu	49
2-Apr-09	1.25	—	−0.25	Thu	28
7-May-09	1	—	−0.25	Thu	35

Figure 3. Basis-Point-Change Distribution in the United States and in the EMU



Moreover, I construct a weekly real-time data set.[11] U.S. variables include the CPI and GDP deflator as inflation measures; GDP growth, industrial production (IP), and the unemployment rate as output measures; and the euro/dollar exchange rate. EMU variables include the euro-zone CPI[12] and GDP deflator as inflation measures; GDP growth, industrial production, and the unemployment rate as output measures; and the euro/dollar exchange rate. See table 3.[13] I take weekly average exchange rate data. Notice that some of these variables are released at a frequency which is lower than weekly, and therefore the latest number can potentially be quite old and stale. This might explain why variables that are updated more frequently—such as the CPI, IP, and the unemployment rate—will be preferred to GDP and GDP deflator in the estimation results. Evans (2005), among others, has focused on deriving daily or weekly

[11]The original data is available daily. That is, on a given day I observe whatever data is released, and for variables for which there is no release on that day, I consider the latest available number. To make this a weekly data set, I am forced to cut off the information on Fridays prior to the meetings. For the ECB, Governing Council meetings are only on Thursdays, so I disregard all the information that arrives thereafter. FOMC meetings are normally held on Tuesdays; thus considering information until the Friday of the previous week is not very restrictive.

[12]Euro-zone inflation is measured by the Harmonized Index of Consumer Prices (HICP).

[13]Inflation and output variables, together with their released dates, are taken from Bloomberg.

Table 3. Explanatory Variables Included in the Data Set

U.S. Variables		**EMU Variables**	
Meeting Dummies			
FOMC		Governing Council	
Inflation Measures			
CPI Excl. FE Index GDP Deflator	YOY% YOY%	MUCPI GDP Deflator	YOY% YOY%
Output Measures			
GDP Growth Industrial Production Unemployment Rate	YOY% MOM%	GDP Growth Industrial Production Unemployment Rate	YOY% MOM%
Exchange Rate			
Eurodollar Rate	Weekly Average	Eurodollar Rate	Weekly Average
Decision Dummies			
U.S. Decision		EMU Decision	

estimates of GDP and other macroeconomic variables. Including those "sophisticated" variables might be a way to overcome this problem, but it goes beyond the scope of this paper.

The aim of collecting real-time data is to consider all the available information that the ECB and the Fed have at the beginning of week t. I am interested in knowing all estimates, provisional, or final data released up until the end (Friday) of week $t-1$. In order to construct the euro-zone GDP and CPI series, I make use of actual released data as well as flash estimates. Euro-zone CPI[14] data for month t are released in the second half of month $t+1$. The same release schedule also applies to the United States. CPI flash estimates represent a considerable enhancement in the available information because they are released within five to ten days from the end of month t. Thus

[14]To compute the HICP flash estimates, Eurostat uses early price information for the reference month from member states for which data are available as well as early information about energy prices. The estimation procedure for the HICP flash estimate combines historical information with partial information on price developments in the most recent months to give a total index for the euro zone. No detailed breakdown is available.

I include flash estimates as soon as they become available and substitute those with final data when they are released. Whereas in the United States GDP data for the quarter ending in month t become available as early as the end of month $t+1$, in the euro area they used to become available at the beginning of month $t+3$. Flash estimates improve the available information because flash GDP estimates are now released as early as the middle of month $t+2$. Flash estimates for the CPI started being released in November 2001 for the October 2001 CPI. Flash estimates for GDP only began in May 2003, for 2003:Q1 GDP.[15] Unemployment data relative to month t are released in the first week of month $t+1$ in the United States and in the first week of month $t+2$ in the EMU.

I also construct two *decision* dummy variables that will be used to assess interdependence in timing decisions. The U.S. dummy variable takes the value one from the last EMU interest rate change until the first FOMC meeting. The EMU dummy variable takes the value one from the last U.S. interest rate change until the second subsequent Governing Council meeting. The asymmetry comes from the fact that Governing Council meetings are more frequent than FOMC meetings (especially in the first part of the sample, when the Governing Council was meeting every two weeks; the FOMC meets only eight times a year), and I want to allow sufficient time for both central banks to react to policy changes.

4. Estimation Strategy and Empirical Results

4.1 Bayesian Implementation[16]

I conduct the estimation in a Bayesian framework. A Bayesian model is characterized by the probability distribution of the data, $p(Y^T|\theta)$, and by the prior distribution $p(\theta)$. I look at the probability of θ given the realized Y^T:

$$p(\theta|Y^T) = \frac{p(Y^T|\theta)p(\theta)}{\int p(Y^T|\theta)p(\theta)d\theta} \tag{24}$$

[15] Thus I do not use flash estimates but only provisional and final estimates before those dates.

[16] I thank Frank Schorfheide for providing Gauss code for the Bayesian estimation, which can be found at www.ssc.upenn.edu/~schorf/programs/gauss-bayesdsge.zip.

so that the parameters θ are treated as random. Equation (24) simply shows how to recover the posterior distribution of θ by applying Bayes's Theorem. The prior belief, which is chosen by the researcher based on economic considerations, can be thought of as an augmentation of the data set which is particularly useful when there are many parameters to be estimated from a short data sample. Bayesian analysis is therefore particularly well suited to my framework because it allows me to augment the data set by including presample information, about the United States and Germany, into the prior.[17] ACD- and ACH-type models generally require fairly large data sets.[18] Because of the youth of the ECB and the type of events I analyze, the data set is small. Bayesian estimation helps in this respect, although it does not completely eliminate the problem.[19]

I use the posterior odds test to select between models. Let M_0 be the baseline model with prior probability $\pi_{0,0}$. The posterior odds of M_0 versus M_1 are

$$\frac{\pi_{0,T}}{\pi_{1,T}} = \left(\frac{\pi_{0,0}}{\pi_{1,0}}\right)\left(\frac{p(Y^T|M_0)}{p(Y^T|M_1)}\right), \tag{25}$$

where $\left(\frac{p(Y^T|M_0)}{p(Y^T|M_1)}\right)$ is the Bayes factor containing the sample evidence and $p(Y^T|M_i)$ is the data density, which I approximate with the modified harmonic mean estimation.[20]

[17]I use data on the German Bundesbank because it is the central bank in Europe that most closely resembles the ECB. Faust, Rogers, and Wright (2001) study the monetary policy of the ECB and compare it with a simple empirical representation of the monetary policy of the Bundesbank before 1999.

[18]Omitted simulation results show that the BACH model, as well as all ACD- and ACH-type models, needs a long data sample to identify the expected duration parameters (α and β) and the constants.

[19]This data problem could have been avoided by analyzing the Fed and the Bank of Japan, or the Fed and the Bank of Canada. However, I believe that investigating the Fed and the ECB is more interesting.

[20]Interpretation of the posterior odds is as follows: $log(\pi_{0,T}/\pi_{1,T}) > 2$ is decisive evidence against H_1; $1 < log(\pi_{0,T}/\pi_{1,T}) < 2$ is strong evidence against H_1; $1/2 < log(\pi_{0,T}/\pi_{1,T}) < 1$ is substantial evidence against H_1; and $0 < log(\pi_{0,T}/\pi_{1,T}) < 1/2$ is not worth more than a bare mentioning.

Table 4. Information on the BACH Parameter Priors

	BACH Model—Prior Distribution				
Parameter	**Range**	**Density**	**Mean**	**St. Dev.**	**90% Interval**
α^{US}	[0, 1)	Beta	0.14	0.08	[0.02, 0.25]
β^{US}	[0, 1)	Beta	0.56	0.14	[0.33, 0.79]
w^{US}	$R+$	Gamma	7.00	2.00	[3.71, 10.11]
$FOMC$	$R+$	Gamma	8.00	1.50	[5.48, 10.38]
CPI^{US}	$R+$	Gamma	1.50	0.50	[0.69, 2.27]
U^{US}	$R+$	Gamma	1.20	0.70	[0.16, 2.20]
d^{US}	$R+$	Gamma	2.00	1.00	[0.49, 3.48]
α^{EMU}	[0, 1)	Beta	0.14	0.08	[0.02, 0.25]
β^{EMU}	[0, 1)	Beta	0.56	0.14	[0.34, 0.80]
w^{EMU}	$R+$	Gamma	10.00	2.00	[6.71, 13.21]
GC	$R+$	Gamma	10.00	1.50	[7.53, 12.43]
CPI^{EMU}	$R+$	Gamma	1.50	0.40	[0.84, 2.11]
U^{EMU}	$R+$	Gamma	1.20	0.40	[0.55, 1.82]
d^{EMU}	$R+$	Gamma	2.00	1.00	[0.47, 3.46]
η	$R+$	Gamma	5.00	2.98	[0.63, 9.25]

4.1.1 Choice of Priors

I choose the priors for the parameters according to a number of considerations. I assume parameters are a priori independent of each other. Parameter restrictions are implemented by appropriately truncating the distribution or by redefining the parameters to be estimated.

Table 4 describes the distributional form, means, and 90 percent confidence intervals of the BACH model priors. According to the Taylor-rule literature, policy rates depend on their lagged values, on some measures of inflation and output deviations, and on exchange rate depreciation. Thus I assume that the probability of a rate change depends on the absolute deviation of inflation and output from a norm, and on the absolute exchange rate depreciation. I take absolute deviations because I want the probability of an interest rate change to increase with large deviations, regardless of their signs.

Let R_t, π_t, y_t, and er_t be the interest rate, inflation rate, a measure of output growth, and nominal exchange rate depreciation, respectively. Also, let π^* and y^* be the optimal level of inflation and output growth, and $0 < \rho^R < 1$ be the smoothing term. Then we can write the Taylor rule as[21]

$$\Delta R_t = (\rho^R - 1)R_{t-1} + (1-\rho^R)[b_1(\pi_t - \pi^*) + b_2(y_t - y^*) + b_3 er_t] + \varepsilon_t. \tag{26}$$

I should also include the lagged interest rate among the covariates. However, the duration in the model generates the dynamics that a lagged interest rate would normally generate. I will, however, verify that this is in fact the case. FOMC and Governing Council meeting dummies have also been included as covariates, following Hamilton and Jordà (2002).

I choose priors for α, β, and the constant w in equation (8), so that the marginal hazard rate, when all the covariates are at their average value,[22] matches the probability of an interest rate change over the ten-year pre-sample period January 1, 1989 to December 31, 1998. I approximate this probability by dividing the number of changes by the number of periods. Since I do not have any pre-sample data for the ECB, I use information about the German Lombard rate. I choose α, β, and w such that

$$h^{US} = \frac{1}{\alpha^{US}\ \overline{u}^{US} + \beta^{US}\ \overline{\psi}^{US} + w^{US}} = \frac{39}{552} = 0.07 \tag{27}$$

$$h^{EMU} = \frac{1}{\alpha^{G}\ \overline{u}^{G} + \beta^{G}\ \overline{\psi}^{G} + w^{G}} = \frac{19}{552} = 0.035, \tag{28}$$

where $\overline{u}$ and $\overline{\psi}$ represent the average values for duration and expected duration over the pre-sample period 1989–98. With this approach, of course, I cannot identify the three parameters involved in each equation. Thus I decide to fix α and β to values close to those that have been estimated in the literature (see Hamilton and Jordà 2002 for the United States) and I vary w to match the probability.

[21] See Lubik and Schorfheide (2007).

[22] I actually assume that π^* and y^* are the average inflation and output growth rates over the sample period.

I choose priors for the covariates z using the following relationships:

$$h^{US} = \frac{1}{\{\alpha^{US}\overline{u}^{US} + \beta^{US}\overline{\psi}^{US} + w^{US} - \delta^{US}|z^{US}|\}} \tag{29}$$

$$h^{EMU} = \frac{1}{\{\alpha^{G}\overline{u}^{G} + \beta^{G}\overline{\psi}^{G} + w^{G} - \delta^{EMU}|z^{EMU}|\}}. \tag{30}$$

In particular, assuming all the other variables are at their average value, the prior for U.S. inflation implies a 0.25 increase in the probability (from 0.07 to 0.32) when the inflation rate increases or decreases by 100 basis points and it is an FOMC day. Similarly, the prior on output/unemployment implies that, on FOMC weeks, a 100-basis-point change in targeted output growth/unemployment rate increases the probability by 0.22. The asymmetry in treating inflation and output is justified by the fact that inflation always has a greater coefficient in the Taylor-rule literature. Very similar priors are given to EMU inflation and output. The meeting dummy coefficients (d^i, $i = US,\ EMU$) are not treated equally in the two countries, due to the greater number of EMU meetings. Notice that, since I expect all the coefficients to be positive,[23] I parameterized them as gamma distributions.

I compute the mean of the odds ratio prior (η) by using the pre-sample proportions for scenarios $(0, 0)$, $(0, 1)$, $(1, 0)$, and $(1, 1)$, over the period January 1, 1989 to December 31, 1998. Again, I use data about the Bundesbank to construct priors about the ECB. The pre-sample odds ratio is about 5.

I also choose priors for the conditional bivariate ordered probit model based on pre-sample information. The means of the cut points are those that I would expect if I were to estimate an ordered probit with no covariates, based on the data 1989–98 for the United States and Germany. The first cut point $c_1 \in R$, hence the normal distribution. The other coefficients are appropriately redefined so as to guarantee that the cut points are ordered. Priors for U.S. inflation and output coefficients are centered at values that have been commonly estimated for Taylor-type rules (see equation (26)). Table 5 describes the distributional form, means, and 90 percent

[23] That is, I expect $-\delta < 0$.

Table 5. Information on the CBO Probit Parameter Priors

CBO Probit Model—Prior Distribution					
Parameter	**Range**	**Density**	**Mean**	**St.Dev.**	**90% Interval**
c_1^{US}	R	Normal	−2.15	2.00	[−7.00, 2.85]
$c_2^{US} - c_1^{US}$	R^+	Gamma	0.70	1.98	[0.00, 1.99]
$c_3^{US} - c_2^{US}$	R^+	Gamma	3.11	1.99	[0.30, 5.90]
$c_4^{US} - c_3^{US}$	R^+	Gamma	0.67	2.00	[0.00, 1.83]
CPI^{US}	R^+	Gamma	1.54	0.50	[0.74, 2.32]
U^{US}	R^+	Gamma	0.25	0.40	[0.00, 0.71]
Δer_t^{US}	R^+	Gamma	0.25	0.20	[0.00, 0.51]
c_1^{EMU}	R	Normal	−2.15	2.00	[−7.17, 2.76]
$c_2^{EMU} - c_1^{EMU}$	R^+	Gamma	0.92	2.01	[0.00, 2.78]
$c_3^{EMU} - c_2^{EMU}$	R^+	Gamma	4.28	2.00	[1.13, 7.23]
$c_4^{EMU} - c_3^{EMU}$	R^+	Gamma	0.81	1.97	[0.00, 2.46]
CPI^{EMU}	R^+	Gamma	1.54	0.50	[0.74, 2.33]
U^{EMU}	R^+	Gamma	0.25	0.20	[0.00, 0.52]
Δer_t^{UEMU}	R^+	Gamma	0.25	0.20	[0.00, 0.52]
ρ	$[-1, 1]$	Normal	0.00	0.40	[−0.66, 0.65]

confidence intervals of the priors of the CBO probit model. To better understand the meaning of these numbers, I consider two scenarios for the United States: a 100-basis-point increase in target inflation, and a 100-basis-point increase in target inflation with an additional 1-percentage-point decrease in the targeted unemployment rate. Conditioning on the Fed having decided to change its target rate, i.e., $x_t^f = 1$, the priors for the United States imply that a 100-basis-point increase in the target level of inflation will raise the probability of a 25-basis-point increase in the FFTR by approximately 0.20 and the probability of a 50-basis-point increase in the FFTR by approximately 0.40. A 100-basis-point increase in the target level of inflation and a 1-percentage-point decrease in targeted unemployment will raise the probability of a 25- and 50-basis-point increase in the FFTR by approximately 0.15 and 0.45, respectively, compared with the baseline scenario. A qualitatively similar analysis holds for the euro area, though in this case the priors imply a

Table 6. BACH Model Posterior Means and Intervals for (i) the Basic Specification and (ii) the Specification with the Odds Ratio Fixed at Its Independence Value (Odds Ratio = 1)

	BACH Parameter Estimation Results			
	Basic Specification		No Synchronization	
Parameter	Mean	90% Interval	Mean	90% Interval
α^{US}	0.05	[0.00, 0.09]	0.05	[0.00, 0.10]
β^{US}	0.44	[0.23, 0.66]	0.44	[0.22, 0.66]
w^{US}	15.05	[12.54, 17.55]	15.10	[12.59, 17.72]
$FOMC$	11.00	[8.80, 13.18]	10.82	[8.66, 13.07]
CPI^{US}	0.26	[0.14, 0.37]	0.28	[0.15, 0.40]
U^{US}	0.04	[0.01, 0.07]	0.04	[0.00, 0.08]
d^{US}	—	—	—	—
α^{EMU}	0.11	[0.00, 0.20]	0.12	[0.02, 0.10]
β^{EMU}	0.56	[0.35, 0.77]	0.57	[0.35, 0.78]
w^{EMU}	19.85	[16.17, 23.50]	19.94	[16.21, 23.66]
GC	9.24	[7.16, 11.27]	9.22	[7.15, 11.25]
CPI^{EMU}	0.28	[0.17, 0.38]	0.29	[0.18, 0.40]
U^{EMU}	0.35	[0.18, 0.53]	0.36	[0.18, 0.53]
d^{EMU}	—	—	—	—
η	4.27	[1.21, 7.14]	1	—

stronger reaction to inflation than unemployment, consistent with the sole mandate of price stability for the ECB.

4.2 BACH Estimates

I estimate a number of different specifications in order to assess which variables are in fact relevant for U.S. and EMU timing decisions. Table 3 shows the covariates I have considered. The basic specification I have selected includes meeting dummies, and inflation and unemployment absolute deviations.[24] Table 6 reports 90 percent posterior probabilities intervals and posterior means as point estimates. The constant parameters for both the United States and the

[24]Thus unemployment and the CPI dominate GDP and GDP deflator measures. Intuitively, unemployment and the CPI are monthly statistics and therefore more promptly incorporate new information.

EMU turn out to have a higher value compared with the prior means, possibly meaning a lower average probability over the sample. The increased value of the constant terms goes together with a smaller value for the coefficients of the other covariates, with the exception of the meeting dummies. The FOMC meeting dummy has a bigger coefficient than the Governing Council meeting dummy (GC): given that FOMC meetings are less frequent than Governing Council meetings, I expect a bigger increase in the probability of a rate change when the FOMC meets.[25] Inflation deviation seems to play a bigger role than unemployment in determining the timing of a rate change in the United States, while inflation and unemployment deviations have a similar weight in the ECB timing decisions. The estimated U.S. inflation parameter implies a 0.15 increase in the probability of an FFTR change when the inflation rate increases or decreases by 100 basis points and it is an FOMC day. Similarly, the estimated unemployment rate coefficient implies that, on FOMC weeks, a 100-basis-point change in the unemployment rate increases the probability of a rate change by 0.13. Consistent with the fact that the ECB is more resilient in changing its target rate, the estimate of the euro-area inflation parameter only implies a 0.04 increase in the probability of a target rate change when the inflation rate increases or decreases by 100 basis points and it is a meeting week for the ECB Governing Council. The odds ratio parameter η has a posterior mean of 4.27, suggesting interdependence between the two central banks.

An interesting by-product of the BACH model is that it generates persistence in the interest rate without including past interest rates. The basic specification with meeting dummies, inflation, and output has been tested against a specification that also includes lagged interest rates, and the former has been selected. I have also tested for a specification that includes exchange rate data in the covariates. Once again, the specification with meeting dummies, inflation, and output has been favored. Finally, unemployment is favored over industrial production and GDP growth as a measure of output, and the CPI is preferred to the GDP deflator as a measure of inflation (results omitted).

[25]The covariates enter with a negative sign in the denominator of the hazard rate; hence, a bigger coefficient on a covariate means a decrease in the denominator and an increase in the probability of a change.

Hu and Phillips (2004) apply a discrete choice approach to model the FFTR during the period 1994–2001, drawing information from the announced target rate as well as from explanatory variables that may be included in a Taylor-type rule. In their model, Hu and Phillips use a triple-choice specification to explain either a decrease, an increase, or no change in the FFTR. They find that four main economic variables contribute to explaining FFTR movements: M2, unemployment claims, consumer confidence, and new orders. While these variables are different from the one that I found significant in my model, their results also suggest that economic conditions trigger the timing of interest rate changes contrary to Hamilton and Jordà (2002), who found that only the absolute value of the spread between the effective federal funds rate and the six-month Treasury-bill rate is the main determinant of the timing, together with the FOMC dummy.

4.3 Conditional Bivariate Ordered Probit Estimates

The basic specification of the CBO probit model that I estimate includes inflation, unemployment, and exchange rates. Table 7 reports 90 percent posterior probabilities intervals and posterior means as point estimates. Inflation and unemployment results exhibit interesting features. While inflation continues to play the foremost role in explaining the size of changes in the U.S. and EMU policy rates, unemployment and exchange rate dynamics share a secondary role. Inflation and unemployment posterior means are, respectively, 0.53 and 0.09 for the United States in the basic model, whereas they are 0.47 and 0.11 for the EMU. The correlation coefficient has a posterior mean of 0.36, with a 90 percent interval equal to [0.06, 0.67].

Conditioning on the Fed having decided to change its target rate, i.e., $x_t^f = 1$, the parameter estimates for the United States imply that a 100-basis-point increase in the target level of inflation will increase the probability of a 25-basis-point rise in the FFTR by approximately 0.18. A 100-basis-point increase in the target level of inflation and a 1-percentage-point decrease in targeted unemployment will increase the probability of a 25- and 50-basis-point rise in the FFTR by approximately 0.21 and 0.05, respectively, compared with the baseline scenario. Results for the EMU suggest that,

Table 7. CBO Probit Model Posterior Means and Intervals for (i) the Basic Specification and (ii) the Specification with the Correlation Coefficient Fixed to Zero

CBO Probit Parameter Estimation Results				
	Basic Model		$\rho = 0$	
Parameter	Mean	90% Interval	Mean	90% Interval
c_1^{US}	1.35	[0.35, 2.78]	0.69	[−0.30, 1.69]
$c_2^{US} - c_1^{US}$	0.37	[0.16, 0.56]	0.29	[0.12, 0.46]
$c_3^{US} - c_2^{US}$	0.07	[0.01, 0.12]	0.82	[0.07, 1.56]
$c_4^{US} - c_3^{US}$	1.80	[1.17, 2.50]	1.75	[1.06, 2.42]
CPI^{US}	0.53	[0.25, 0.78]	0.52	[0.25, 0.78]
U^{US}	0.11	[0.00, 0.28]	0.04	[0.00, 0.12]
Δer_t^{US}	0.09	[0.00, 0.18]	0.09	[0.00, 0.18]
c_1^{EMU}	0.65	[0.08, 1.22]	1.96	[0.30, 3.80]
$c_2^{EMU} - c_1^{EMU}$	0.25	[0.07, 0.42]	0.27	[0.08, 0.46]
$c_3^{EMU} - c_2^{EMU}$	2.04	[0.84, 3.20]	1.90	[0.67, 3.12]
$c_4^{EMU} - c_3^{EMU}$	0.97	[0.50, 1.42]	1.14	[0.60, 1.65]
CPI^{EMU}	0.47	[0.28, 0.66]	0.51	[0.29, 0.72]
U^{EMU}	0.11	[0.03, 0.20]	0.25	[0.07, 0.43]
Δer_t^{UEMU}	0.13	[0.00, 0.26]	0.17	[0.01, 0.32]
ρ	0.36	[0.06, 0.67]	0	—

conditioning on the ECB having decided to change its target rate (i.e., $x_t^e = 1$), parameter estimates imply that a 100-basis-point increase in the target level of inflation will increase the probability of a 25-basis-point rise in the MRO rate by approximately 0.22. A 100-basis-point increase in the target level of inflation and a 1-percentage-point decrease in targeted unemployment will increase the probability of a 25- and 50-basis-point rise in the MRO rate by approximately 0.28 and 0.05, respectively, compared with the base scenario.

Magnitude results are therefore showing that, for both central banks, inflation has been crucial and unemployment has had a minor role during the period analyzed. This is expected in view of the fact that the ECB's primary objective is to maintain price stability; a policy of targeting output growth would probably be more problematic, given the intrinsic differences in the economies of the EMU countries. For the Fed, this result may hinge on the fact that most of the period analyzed in this paper is characterized by relatively low unemployment. Because of the non-linearities in my model, it is difficult to compare my results with the existing Taylor-rule literature, such as Evans (1998), Orphanides (2001), Clarida, Galí, and Gertler (2002), and Molodtsova, Nikolsko-Rzhevskyy, and Papell (2008), among others. However, my model seems to validate the idea that the inflation parameter tends to be bigger than the output/unemployment parameter. Moreover, in line with the real-time literature which suggests that real-time policy recommendations differ widely from those obtained with revised published data, my results for the United States differ from Hamilton and Jordà (2002), who employ a revised data set. In their paper, most of the candidate explanatory variables turn out to be insignificant, and they find that the spread between the effective federal funds rate and the six-month Treasury-bill rate is the main determinant of the size of the FFTR change as well as of the timing of the change. The importance of using real-time data is also confirmed by the fact that variables that are updated more frequently—such as CPI, IP, and the unemployment rate—are preferred, according to my estimation results, to variables that are updated less frequently, like GDP and GDP deflator. GDP information tends to be old and stale due to long reporting lags. Monetary policy committees might be more concerned about the current state of the economy and, as such, care more about the information contained in macroeconomic announcements that are more frequently and timely released.

4.4 Interdependence

The interdependence test of U.S. and EMU timing decisions is twofold: on the one hand, I am interested in assessing "contemporaneous" interdependence, after controlling for each country's macroeconomic conditions, which I refer to as *synchronization*; on the other hand, I investigate the possibility of *follower behaviors*, after controlling for each country's macroeconomic conditions. Assessing *synchronization* involves testing whether the odds ratio is different from 1 (1 meaning independence). The odds ratio indicates how

Table 8. BACH Model Posterior Means and Intervals for the Specifications with (i) the U.S. Dummy Variable in the EMU Decision Variables (EMU Follower), (ii) the EMU Dummy Variable in the U.S. Decision Variables (U.S. Follower), and (iii) Both Dummy Variables

	BACH Parameter Estimation Results					
	EMU Follower		U.S. Follower		Both Dummies	
Parameter	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval
α^{US}	0.05	[0.00, 0.09]	0.05	[0.00, 0.10]	0.05	[0.00, 0.09]
β^{US}	0.44	[0.21, 0.66]	0.44	[0.22, 0.66]	0.44	[0.21, 0.66]
w^{US}	15.12	[12.64, 17.63]	15.31	[12.75, 17.86]	15.35	[12.64, 17.63]
$FOMC$	11.03	[8.83, 13.20]	10.68	[8.45, 12.87]	10.68	[8.83, 13.20]
CPI^{US}	0.26	[0.14, 0.38]	0.25	[0.13, 0.37]	0.25	[0.14, 0.38]
U^{US}	0.04	[0.01, 0.08]	0.04	[0.00, 0.08]	0.04	[0.01, 0.08]
d^{US}	—	—	0.78	[0.20, 1.35]	0.78	[0.20, 1.34]
α^{EMU}	0.12	[0.02, 0.21]	0.11	[0.06, 0.20]	0.12	[0.02, 0.21]
β^{EMU}	0.57	[0.35, 0.78]	0.56	[0.36, 0.77]	0.57	[0.36, 0.78]
w^{EMU}	19.95	[16.24, 23.74]	19.91	[16.18, 23.55]	19.99	[16.27, 23.74]
GC	9.21	[7.18, 11.24]	9.26	[7.21, 11.27]	9.23	[7.14, 11.24]
CPI^{EMU}	0.29	[0.18, 0.40]	0.29	[0.17, 0.39]	0.29	[0.18, 0.40]
U^{EMU}	0.35	[0.18, 0.52]	0.36	[0.18, 0.53]	0.35	[0.18, 0.52]
d^{EMU}	1.82	[0.50, 3.09]	—	—	1.81	[0.49, 3.08]
η	4.38	[1.23, 7.34]	3.84	[1.12, 6.26]	4.20	[1.15, 7.02]

much the odds of one country changing its target rate increase when the other country changes its target. Columns 4 and 5 in table 6 display the estimation results for the independence setup. Setting the odds ratio to 1 does not significantly affect the other coefficients: both means and 90 percent probability intervals are very similar to the basic specification. However, as table 9 shows, the posterior odds of the model M_1 with the *odds ratio* $= 1$ versus the alternative basic model M_0 seems to support model M_0. Thus the BACH model tends to favor a setup with synchronization between the two central banks, after controlling for each country's macroeconomic conditions.

Follower behaviors are studied by including the two *decision* dummies to account for the effect of the other country's decisions (see section 3 for a more detailed explanation of the dummy variables). Results are shown in table 8. Table 9 suggests that the

Table 9. BACH Model Posterior Odds for the Synchronization and the Leader/Follower Models

	Log Marginal Data Densities	Posterior Odds
Basic Model	−286.61	
No Synchronization ($\eta = 1$)	−287.36	2.12 (M_0: Basic Model)
U.S. Leader/EMU Follower[a]	−286.80	1.21 (M_0: Basic Model)
EMU Leader/U.S. Follower[b]	−290.19	35.88 (M_0: Basic Model)
Both U.S. and EMU Dummies	−290.12	33.45 (M_0: Basic Model)

[a]Only dummy^{EMU}
[b]Only dummy^{US}

Table 10. CBO Probit Model Posterior Odds for the Correlation in Magnitude Hypothesis

	Log Marginal Data Densities	Posterior Odds
Basic Model	−130.05	
$\rho = 0$	−125.49	95.58 (M_0: $\rho = 0$)

posterior odds ratio supports the hypothesis of no follower behaviors when both dummies are included. I obtain similar results by testing for the Fed's follower behavior: there is strong evidence against the Fed following what the ECB does. On the other hand, though there is some evidence against the reverse scenario of the ECB following the Fed's timing decisions, this evidence is slim.

I analyze interdependence in the CBO probit framework by testing whether the correlation coefficient between the latent variables in equations (20) and (21) is different from zero. Table 7 presents the estimation results for this scenario. The posterior odds ratio in table 10 shows evidence in favor of the model specification in which the correlation is set to zero. The correlation coefficient measures the correlation between the shocks in the unobservable variable equations—the omitted factors.[26] The results seem to suggest that although there is *synchronization* in the timing of interest rate changes, each central bank then sets the target level based exclusively on its own macroeconomic conditions.

The BACH model with the odds ratio set to 1 and the CBO probit model with $\rho = 0$ can be thought of as the Hamilton and Jordà (2002) univariate model estimated for both the Fed and the ECB. Please note that results for the United States cannot be compared because the sample is different in term of length and included variables. Moreover, Hamilton and Jordà (2002) do not use a real-time data set.

5. Posterior Predictive Checks

To check the goodness of fit of the model described in the paper, I run some posterior predictive checks on the BACH and CBO probit models that were selected above.[27] The general idea is that if the model fits well enough, then the replicated data will look somehow similar to the observed data. I run predictive checks for the BACH model separately from the CBO probit model.

To generate data based on the BACH model, I use the following algorithm:

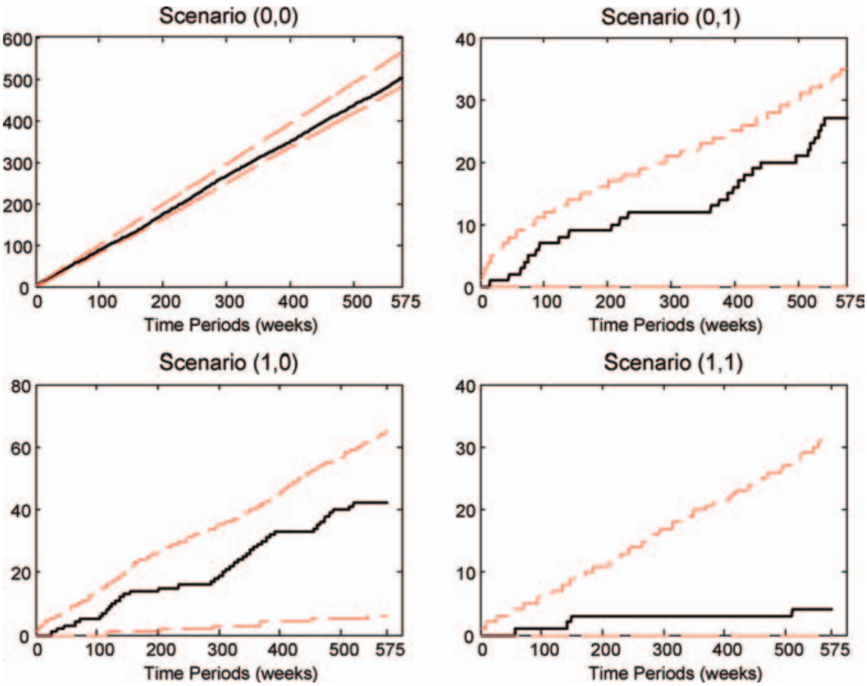
- Sample θ from the posterior distribution $p(\theta | x^{e,\ obs}, x^{f,\ obs}, z)$, where $x^{e,\ obs}, x^{f,\ obs}$ are defined as in equation (1) and z are the exogenous data (for $nrep = 100{,}000$).
- Using each one of the θ^i realizations above, sample $x^{i,rep}$ from the model $p(x|\theta, z)$, where each sample $x^{i,rep}$ has length $T = 575$ (length of my data set).
- Consider $(x^{e,rep}, x^{f,rep}) = \{(0,0),(0,1),(1,0),(1,1)\}$ and compare it with $(x^{e,\ obs}, x^{f,\ obs})$.

Figure 4 shows the results of the posterior predictive checks on the BACH model. The four panels each represent one of the four

[26] By relating these shocks to the VAR literature, it turns out that, given the assumption that interest rates only depend on past values of output and inflation, the disturbances in equations (20) and (21) are purely monetary shocks.

[27] See Bauwens, Lubrano, and Richard (1999).

Figure 4. Results of BACH Predictive Checks



Notes: The panels show the cumulative count of the (0,0), (0,1), (1,0), and (1,1) scenarios over the 575 weeks under consideration. The thick black lines represent the cumulative count of the observed data. The broken lines show the range of paths for the replicated data.

possible outcomes $(0,0)$, $(0,1)$, $(1,0)$, and $(1,1)$. The thick black line represents the cumulative count of the observed data. The upper-left panel shows the cumulative count of the $(0,0)$ scenarios over the 575 weeks under consideration. Similarly, the thick black lines in the other panels show the cumulative counts of scenarios $(0,1)$, $(1,0)$, and $(1,1)$. The value at the end of week 575 corresponds to the total amount of $(0,0)$, $(0,1)$, $(1,0)$, and $(1,1)$ scenarios that we observed from 1999 to 2009. The broken lines show the range of paths for the replicated data. As clearly shown by the figure, the observed path lies within the paths of the replicated data, endorsing the plausibility of the BACH model. Of note, because scenario $(1,1)$ is so little observed in reality, it is also perhaps the more difficult to simulate.

The algorithm for the predictive checking of the CBO probit model is similar to the one described above for the BACH model. After sampling from the parameters' posterior distribution, I sample $y^{i,rep}$ from the likelihood distribution $p(y^{obs}|\theta, z)$, where $y^{i,rep} = \{-50, -25, 25, 50\}$. For each repetition, I count the number of times $y^i = -50$, –25, 25, and 50. Results are shown in figure 5, where both the histogram and some statistics are shown. Although accuracy is not perfect, the mean and the median are not far off from the true realizations, endorsing the plausibility of the CBO probit model.

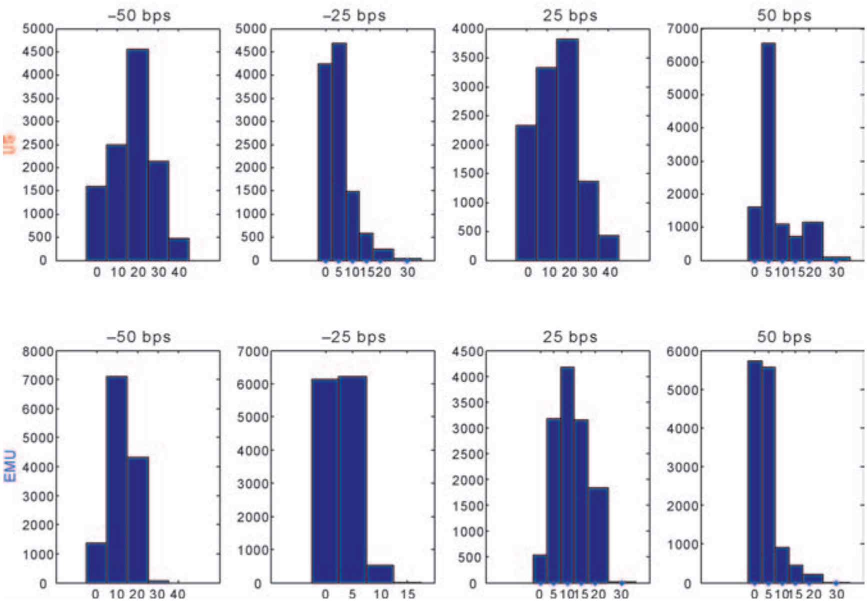
6. Conclusions

In this paper I have derived and estimated with Bayesian techniques a bivariate model to account for interdependence between Fed and ECB decisions. I have operationalized interest rate timing decisions with a bivariate autoregressive conditional hazard (BACH) model and magnitude decisions with a conditional bivariate ordered (CBO) probit model. The timing model yields evidence supporting the hypothesis that (i) institutional factors (scheduled FOMC and Governing Council meetings) and inflation rates are relevant variables for both central banks, (ii) output plays a minor role in the U.S. timing decisions, and (iii) there exists some evidence of synchronization but no follower behavior. The magnitude model illustrates that (i) inflation rates are the most important variables in determining interest rate levels, (ii) output and exchange rates play a secondary role in magnitude decisions, and (iii) the posterior odds ratio favors a model with zero correlation in magnitude changes. I also find that, based on posterior predictive checks, my model has a good fit and is able to generate data that are similar to the observed data.

My findings are necessarily based on a relatively small sample; however, there seems to be evidence suggesting that timing and magnitude changes are in fact quite interesting issues. The paper provides evidence in favor of interaction in U.S. and EMU interest rate timing and magnitude decisions, after controlling for traditional variables that have commonly been used in the literature. The paper offers a new methodology to analyze interdependence; however, it does not provide a complete answer to the underlying problem about what is in fact the source of the interdependence and

Figure 5. Results of CBO Probit Model Predictive Checks

		−50bps	−25bps	25bps	50bps
Fed	Observed	16	7	22	1
	Mean Realization	18	5	16	4
	Median Realization	19	4	15	4
ECB	Observed	10	5	14	2
	Mean Realization	13	3	11	4
	Median Realization	12	3	12	3



Notes: The top part of the figure shows the count of the times the Fed and ECB changed their respective target rate by −50, −25, 25, and 50 basis points. It also shows the mean and median of those counts for the simulated data. The bottom part of the figure shows the histogram of the simulated data for each one of the four outcomes (−50, −25, 25, and 50 basis points).

whether interdependence is optimal. Identifying where the interdependence comes from and analyzing whether results are robust to the inclusion of a larger set of explanatory variables remain important topics for further research.

Appendix

CBO Probit Model

The log-likelihood relative to the magnitude decision is

$$\mathcal{L}_2(\theta_2) = \sum_{t=1}^{T} \mathit{1}\left[x_t^e = 1, x_t^f = 0\right] \log P_{10} + \mathit{1}\left[x_t^e = 0, x_t^f = 1\right] \log P_{01} + \mathit{1}\left[x_t^e = 1, x_t^f = 1\right] \log P_{11} \tag{31}$$

with

$$\begin{aligned} P_{10} &= \Pr\left(y_t^e = s_m,\ y_t^f = 0 \mid w_{t-1}, x_t^e = 1, x_t^f = 0\right) \\ &= \Pr\left(c_{m-1}^e < \widetilde{y}_t^e \leq c_m^e,\ c_2^f < \widetilde{y}_t^f \leq c_3^f \mid w_{t-1}, x_t^e = 1, x_t^f = 0\right) \\ m &= 1, 2, 4, 5, \end{aligned} \tag{32}$$

where the probability is computed using the density

$$f\left(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1},\ x_t^e = 1,\ x_t^f = 0\right) = \frac{f\left(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1}\right)}{\Pr\left\{\begin{array}{c}\left[\left(\widetilde{y}_t^e \leq c_2^e\right) \vee \left(\widetilde{y}_t^e > c_3^e\right)\right] \\ \wedge\left(c_2^f < \widetilde{y}_t^f \leq c_3^f\right)\end{array}\right\}}; \tag{33}$$

with

$$\begin{aligned} P_{01} &= \Pr\left(y_t^e = 0,\ y_t^f = s_n \mid w_{t-1}, x_t^e = 0, x_t^f = 1\right) \\ &= \Pr\left(c_2^e < \widetilde{y}_t^e \leq c_3^e,\ c_{n-1}^f < \widetilde{y}_t^f \leq c_n^f \mid w_{t-1}, x_t^e = 0, x_t^f = 1\right) \\ n &= 1, 2, 4, 5, \end{aligned} \tag{34}$$

where the probability is computed using the density

$$f\left(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1},\ x_t^e = 0,\ x_t^f = 1\right) = \frac{f\left(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1}\right)}{\Pr\left\{\begin{array}{c}\left(c_2^e < \widetilde{y}_t^e \leq c_3^e\right) \\ \wedge\left[\left(\widetilde{y}_t^f \leq c_2^f\right) \vee \left(\widetilde{y}_t^f > c_3^f\right)\right]\end{array}\right\}}; \tag{35}$$

and with

$$
\begin{aligned}
P_{11} &= \Pr\left(y_t^e = s_m,\ y_t^f = s_n \mid w_{t-1}, x_t^e = x_t^f = 1\right) \\
&= \Pr(c_{m-1}^e < \widetilde{y}_t^e \leq c_m^e,\ c_{n-1}^f < \widetilde{y}_t^f \leq c_n^f \mid w_{t-1}, x_t^e = x_t^f = 1) \\
m, n &= 1, 2, 4, 5,
\end{aligned} \tag{36}
$$

where the probability is computed using the density

$$
f\left(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1},\ x_t^e = x_t^f = 1\right) = \frac{f\left(\widetilde{y}_t^e, \widetilde{y}_t^f | w_{t-1}\right)}{\Pr\left\{\begin{array}{c} \left[\left(\widetilde{y}_t^e \leq c_2^e\right) \vee \left(\widetilde{y}_t^e > c_3^e\right)\right] \\ \wedge\left[\left(\widetilde{y}_t^f \leq c_2^f\right) \vee \left(\widetilde{y}_t^f > c_3^f\right)\right] \end{array}\right\}}. \tag{37}
$$

Bayesian Implementation

Following Schorfheide (2000), I compute the mode $\widetilde{\theta}$ of the posterior density $p(\theta|Y^T)$ through a numerical optimization routine and then evaluate the inverse Hessian $\widetilde{\Sigma}$. I use a random-walk Metropolis algorithm to generate n_{sim} draws $\widetilde{\theta}^s$ from the posterior $p(\theta|Y^T)$.[28] At each iteration s, I draw a candidate parameter vector ϑ from a jumping distribution $J_s(\vartheta|\theta^{(s-1)})$ and I accept the jump from $\theta^{(s-1)}$ so that $\theta^{(s)} = \vartheta$ with probability $min(r, 1)$, where r is defined as

$$
r = \frac{p(Y^T|\vartheta)p(\vartheta)}{p(Y^T|\theta^{(s-1)})p(\theta^{(s-1)})}, \tag{38}
$$

and reject otherwise. The Markov chain sequence $\{\theta^{(s)}\}_{s=1}^{n_{sim}}$ converges to the posterior distribution as $n_{sim} \to \infty$. I use a Gaussian jumping distribution $J_s \sim \mathcal{N}(\theta^{(s-1)}, c^2\widetilde{\Sigma})$ with $c = 0.3$.

References

Bartolini, L., and A. Prati. 2003. "The Execution of Monetary Policy: A Tale of Two Central Banks." *Economic Policy* 18 (37): 437–67.

[28] I use n_{sim} = 150,000 for the BACH model and n_{sim} = 100,000 for the CBO probit model.

Bauwens, L., M. Lubrano, and J. F. Richard. 1999. *Bayesian Inference in Dynamic Econometric Models*. New York: Oxford University Press.

Benigno, P. 2002. "A Simple Approach to International Monetary Policy Coordination." *Journal of International Economics* 57 (1): 177–96.

Bergin, P. R., and O. Jordà. 2004. "Measuring Monetary Policy Interdependence." *Journal of International Money and Finance* 23 (5): 761–83.

Cecchetti, S., and R. O'Sullivan. 2003. "The European Central Bank and the Federal Reserve." *Oxford Review of Economic Policy* 19 (1): 30–43.

Clarida, R., J. Galí, and M. Gertler. 2002. "A Simple Framework for International Monetary Policy Analysis." *Journal of Monetary Economics* 49 (5): 879–904.

Cogley, T., and T. J. Sargent. 2002. "Evolving Post World War II U.S. Inflation Dynamics." In *NBER Macroeconomics Annual 2001*, ed. B. S. Bernanke and K. S. Rogoff. Cambridge, MA: MIT Press.

Corsetti, G., and P. Pesenti. 2005. "International Dimensions of Optimal Monetary Policy." *Journal of Monetary Economics* 52 (2): 281–305.

Engle, R. 2000. "The Econometrics of Ultra-High Frequency Data." *Econometrica* 68 (1): 1–22.

Engle, R., D. Hendry, and J. F. Richard. 1983. "Exogeneity." *Econometrica* 51 (2): 277–304.

Engle, R., and A. Lunde. 2003. "Trades and Quotes: A Bivariate Point Process." *Journal of Financial Econometrics* 1 (2): 159–88.

Engle, R. F., and J. Russell. 1998. "Autoregressive Conditional Duration: A New Model for Irregularly Spaced Transaction Data." *Econometrica* 66 (5): 1127–62.

———. 2002. "Forecasting the Frequency of Changes in Quoted Foreign Exchange Prices with the Autoregressive Conditional Duration Model." In *Forecasting Financial Markets*, Vol. 2, ed. T. C. Mills, 595–620. Cheltenham, UK and Northampton, MA.: Elgar.

———. 2005. "A Discrete-State Continuous-Time Model of Financial Transactions Prices and Times: The Autoregressive Conditional Multinomial-Autoregressive Conditional Duration Model." *Journal of Business and Economic Statistics* 23: 166–80.

Evans, C. L. 1998. "Real-Time Taylor Rules and the Federal Funds Futures Market." *Economic Perspectives* (Federal Reserve Bank of Chicago) (Third Quarter): 44–55.

Evans, M. D. D. 2005. "Where Are We Now? Real-Time Estimates of the Macroeconomy." *International Journal of Central Banking* 1 (2): 127–75.

Faust, J., J. H. Rogers, and J. H. Wright. 2001. "An Empirical Comparison of Bundesbank and ECB Monetary Policy Rules." International Finance Discussion Paper No. 705, Board of Governors of the Federal Reserve System.

Hamilton, J. D., and O. Jordà. 2002. "A Model of the Federal Reserve Rate Target." *Journal of Political Economy* 110 (5): 1135–67.

Hu, L., and P. Phillips. 2004. "Dynamics of Federal Funds Target Rate: A Nonstationary Discrete Choice Approach." *Journal of Applied Econometrics* 19 (7): 851–67.

Lubik, T., and F. Schorfheide. 2007. "Do Central Banks Respond to Exchange Rate Movements? A Structural Investigation." *Journal of Monetary Economics* 54 (4): 1069–87.

Molodtsova, T., A. Nikolsko-Rzhevskyy, and D. H. Papell. 2008. "Taylor Rules with Real-Time Data: A Tale of Two Countries and One Exchange Rate." *Journal of Monetary Economics* 55 (S1): S63–S79.

Orphanides, A. 2001. "Monetary Policy Rules Based on Real-Time Data." *American Economic Review* 91 (4): 964–85.

Pappa, E. 2004. "Do the ECB and the Fed Really Need to Cooperate? Monetary Policy in a Two-Country World." *Journal of Monetary Economics* 51 (4): 753–79.

Schorfheide, F. 2000. "Loss Function-Based Evaluation of DSGE Models." *Journal of Applied Econometrics* 15 (6): 645–70.

Sims, C., and T. Zha. 1998. "Bayesian Methods for Dynamic Multivariate Models." *International Economic Review* 39 (4): 949–68.

Tajar, A., M. Denuit, and P. Lambert. 2001. "Copula-Type Representation for Random Couples with Bernoulli Margins." Discussion Paper No. 0118, Universite Catholique de Louvain.

Interest Rate Smoothing and “Calvo-Type” Interest Rate Rules: A Comment on Levine, McAdam, and Pearlman (2007)*

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In a recent paper, Levine, McAdam, and Pearlman (2007) propose a new type of interest rate rule, which they denote a “Calvo-type” rule. The Calvo-type interest rate responds to the discounted sum of current and future rates of inflation. We show that a Calvo-type rule can be derived from a very different assumption than the one used by Levine, McAdam, and Pearlman (2007), namely a preference for interest rate smoothing. In addition to giving an alternative rationale for the Calvo-type rule, we provide additional empirical support for the specification.

JEL Codes: E52, E37, E58.

1. Introduction

Monetary policy is commonly assumed to be forward looking. A popular way to specify the forward-lookingness in monetary policy is to let the interest rate respond to the inflation forecast, as in forward-looking Taylor rules. Levine, McAdam, and Pearlman (2007), henceforth LMP, note that such rules may have poor stabilization properties and often give real indeterminacy. They propose an alternative representation of monetary policy, which they refer

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to as a "Calvo-type" rule, where the interest rate responds to the discounted sum of all expected future rates of inflation. They show that this type of rule has better stabilization properties than traditional forward-looking Taylor rules. Moreover, Gabriel, Levine, and Spencer (2009) find empirical support for this kind of rule using data for the United States.

In this comment, we show that the rule suggested by LMP can be derived from a very different assumption than that made by LMP, namely a preference for *interest rate smoothing*. The current interest rate decision affects both the change in the interest rate from the previous period to the current one and the expected change from the current period to the next one. The interest rate decision today should therefore take into account both the lagged interest rate and the expected future interest rate. When the interest rate responds to current inflation, this gives rise to a Calvo-type rule. The Calvo-type rule thus has a more general foundation than previously believed.

In addition to providing an alternative rationale for the Calvo-type rule, we provide additional empirical support for this specification. Gabriel, Levine, and Spencer (2009) estimate the interest rate rule using GMM. Since GMM estimates are known to suffer from small-sample bias (see e.g., Hall 2005, chapter 6 and the references therein), we analyze the robustness of Gabriel, Levine, and Spencer (2009)'s results using maximum-likelihood system estimation and single-equation estimation where the implied forward interest rate is used as a proxy for the expected future interest rate. The results in Gabriel, Levine, and Spencer (2009) are generally confirmed and strengthen the case for the Calvo-type specification of interest rate rules. Since the Calvo-type specification has a general theoretical foundation and strong empirical support, we argue that future work on simple rules should consider Calvo-type rules as an alternative to the more common but less general forward-looking Taylor rules.

2. A Model of Interest Rate Smoothing

In the empirical literature on simple interest rate rules, it is common to specify the rule as a *partial-adjustment* equation, i.e.,

$$r_t = \rho r_{t-1} + (1 - \rho) r_t^*, \tag{1}$$

where r_t is the nominal interest rate and r_t^* is the target interest rate. The main motivation for the partial-adjustment specification is empirical fit, but it is often interpreted as evidence of central banks' preference for interest rate smoothing.[1] With partial adjustment, the central bank moves the interest rate gradually to the target rate. In the empirical literature, the target rate r_t^* is commonly specified as a non-inertial rule, such as the classic Taylor rule. There are, however, theoretical reasons for having an inertial target rate, as argued by Woodford (2003). But as noted by Rudebusch (2002), the partial-adjustment specification does not distinguish between inertia in the target rate itself or gradual adjustment toward a non-inertial target rate.

LMP derive optimal rules adding a lagged interest rate term. Even if the loss function considered by LMP does not include a preference for interest rate smoothing,[2] the authors still find that current policy should respond to the lagged interest rate. Indeed, they find that the optimal coefficient is one, thereby implying an integral (or difference) rule. This is a common result when the coefficients are optimized subject to the type of forward-looking models considered by LMP. Since LMP do not have an interest rate smoothing term in the loss function, their result on the optimal coefficients can be interpreted as finding an optimal target rate r_t^*.

Our aim is to show that the Calvo-type specification does not hinge on a specific model, as long as the central bank has a preference for interest rate smoothing. Following the traditional literature on empirical policy rules, we assume that the target interest rate is a standard (non-inertial) Taylor rule, i.e.,

$$r_t^* = a\pi_t + by_t, \tag{2}$$

where we for simplicity abstract from constant terms and assume that the neutral interest rate is zero. Even if it can be argued that such a simple, non-forward-looking target rule is sub-optimal, we deliberately choose this specification to show that the forward-looking nature of the rule with interest rate smoothing does not hinge on a forward-looking target interest rate.

[1] See, e.g., Clarida, Galí, and Gertler (2000).

[2] They include a term with the interest rate *level*, but not the change in the interest rate.

Assume now that the central bank prefers to smooth the interest rate around the target rate. We model this by the following quadratic adjustment cost specification:

$$\Omega_t = \frac{1}{2} E_t \sum_{k=0}^{\infty} \delta^k \left[\left(r_{t+k} - r^*_{t+k} \right)^2 + \varphi (r_{t+k} - r_{t+k-1})^2 \right], \qquad (3)$$

where r^*_t is the target rate, δ is the discount factor, and φ is the cost of changing the interest rate. The first term represents the cost of deviating from the target interest rate, and the second term represents the cost of changing the interest rate. The first-order condition for minimization of (3) is

$$r_t - r^*_t + \varphi(r_t - r_{t-1}) - \delta\varphi(E_t r_{t+1} - r_t) - \sum_{k=0}^{\infty} \delta^k E_t \left(r_{t+k} - r^*_{t+k} \right) \frac{\partial E_t r^*_{t+k}}{\partial r_t} = 0. \qquad (4)$$

The term $-\sum_{k=0}^{\infty} \delta^k E_t (r_{t+k} - r^*_{t+k}) \frac{\partial E_t r^*_{t+k}}{\partial r_t}$ reflects that deviating from the target interest rate might affect the target rate itself, since the target rate depends on endogenous variables. We will, however, assume that interest rate smoothing has a negligible effect on the target interest rate in the near term. This is a reasonable assumption if the target rate depends on variables like inflation and the output gap that are affected by monetary policy with a time lag. Since the actual interest rate will only deviate significantly from the target rate in the first couple of periods, then for reasonable values of φ, one will tend to have that $|\frac{\partial E_t r^*_{t+k}}{\partial r_t}| \approx 0$ when $E_t(r_{t+k} - r^*_{t+k})$ is non-negligible and $E_t(r_{t+k} - r^*_{t+k}) \approx 0$ when $|\frac{\partial E_t r^*_{t+k}}{\partial r_t}|$ is non-negligible. The product $E_t(r_{t+k} - r^*_{t+k}) \frac{\partial E_t r^*_{t+k}}{\partial r_t} \approx 0$ for all $k = 1, 2, \ldots .T$, and the discounted sum of these products will be very small.[3] A close approximation to the optimal smoothing behavior given by (4) can then be written as

$$r_t = \gamma r_{t-1} + \gamma\delta E_t r_{t+1} + (1 - \gamma - \gamma\delta) r^*_t, \qquad (5)$$

[3]This is obviously not the case if the target interest rate depends on variables that display a significant contemporaneous response to changes in the interest rate such as, e.g., asset prices.

where $\gamma = \frac{\varphi}{1+\varphi(1+\delta)}$.[4] We see that optimal interest rate smoothing implies both forward- and backward-looking behavior, while partial adjustment implies only backward-looking behavior. More specifically, under partial adjustment the interest rate is set as a weighted average of the target rate and the lagged interest rate. Under optimal smoothing, the interest rate is a weighted average of the target rate, the lagged interest rate, and the expected next-period interest rate. Why does interest rate smoothing imply forward-looking behavior? The intuition is that a central bank that aims to smooth the interest rate is not only concerned about a smooth development in the interest rate from the previous period to the current period but also a smooth development from this period to the next. Since the interest rate set today has implications for both, a central bank with a preference for interest rate smoothing must be partly forward looking.

When the target rate r_t^* is given by (2), the rule with optimal interest rate smoothing can be written as

$$\begin{aligned} r_t &= \gamma r_{t-1} + \gamma\delta E_t r_{t+1} + (1-\gamma-\gamma\delta)(a\pi_t + by_t) \\ &= \hat{\rho} r_{t-1} + \varphi E_t \sum_{k=0}^{\infty} (\hat{\rho}\delta)^k (a\pi_{t+k} + by_{t+k}), \end{aligned} \tag{6}$$

where the last equality follows by solving the equation forward, which gives $\hat{\rho} = \frac{1}{2\delta\gamma}(\sqrt{1-4\delta\gamma^2}+1)$ and $\varphi = (1-(1+\delta)\gamma)\hat{\rho}\gamma^{-1}$. Note that for $b = 0$, the forward-solution specification is identical to the Calvo-type interest rate rule specified by equations (11) and (13) in LMP.[5] The key insight from our simple model is that a preference for interest rate smoothing is sufficient to make the central bank forward looking. This is in stark contrast to the sluggish backward-looking behavior implied by the standard partial-adjustment specifications in the empirical literature on interest rate rules. Rudebusch (2002) argued that the unreasonably high degree of inertia was due

[4] This specification is equal to the one in footnote 10 of LMP.

[5] Since our rule is derived from quadratic adjustment costs à la the Rotemberg (1982) approach of deriving the New Keynesian Phillips curve, it would perhaps be natural to call our specification a "Rotemberg-type" interest rate rule instead of a Calvo-type rule. But as with the New Keynesian Phillips curve, our Rotemberg foundation gives the same interest rate rule as LMP's Calvo type.

to omitted autocorrelated variables. Our model of optimal smoothing suggests that the omission of the expected future interest rate in (6) could be an important omitted variable.

3. Empirical Analysis

Gabriel, Levine, and Spencer (2009) find empirical support for the Calvo-type interest rate rule using single-equation GMM methods. Using U.S. data from 1960 to 2004, they report a positive and significant coefficient on the lagged interest rate term and the forward term in the interest rate rule.[6] It is well known that GMM estimators can exhibit substantial bias in small samples. In this section we examine the robustness of Gabriel, Levine, and Spencer (2009)'s results using two alternative approaches: maximum-likelihood system estimation and single-equation estimation where the implied forward interest rate is used as a proxy for the expected future interest rate.

We estimate the interest rule on quarterly U.S. data from 1987:Q3 to 2007:Q4.[7,8] Following the literature on estimated policy rules for the United States (e.g., Clarida, Galí, and Gertler 2000, Rudebusch 2002, and Jondeau, Le Bihan, and Galles 2004), we use the federal funds rate as the monetary policy instrument, r_t. Inflation is measured using the GDP deflator[9] (denoted P_t), so that $\pi_t = 400(\ln(P_t) - \ln(P_{t-1}))$.[10] The output gap is defined as the percentage deviation of real GDP from real potential GDP, i.e., $y_t = 100(\ln(GDP_t) - \ln(GDP_t^*))$, where real potential output is provided by the Congressional Budget Office.[11]

[6] Rewritten in comparable values, Gabriel, Levine, and Spencer (2009) find that (using the CBO output gap) $r_t = 0.56r_{t-1} + 0.4E_t r_{t+1} + 0.04(4.53E_t\overline{\pi}_{t+4} + 1.30E_t y_{t+1})$.

[7] The data series are obtained from the Federal Reserve Bank of St. Louis.

[8] The choice of estimation period is motivated by our desire to estimate the reaction function over a single monetary policy regime. Allowing for a structural break in the parameters of the reaction function in 1987:Q3, Jondeau, Le Bihan, and Galles (2004) strongly reject that the parameters are stable.

[9] The GDP deflator is seasonally adjusted.

[10] The results reported below are robust to using the GDP chain-weighted price index as the measure of inflation.

[11] The results are robust to replacing the output gap with a measure of the unemployment gap. Results are available upon request.

ML estimation requires that we specify an auxiliary model for the variables that determine the target rate (here, inflation and the output gap). We use a simple backward-looking model for inflation and output that has been shown to fit the data well. Specifically, we use a slightly modified version of the model proposed by Rudebusch and Svensson (1999). The model equations are

$$\pi_t = \alpha_1 \pi_{t-1} + \alpha_2 \pi_{t-2} + \alpha_3 \pi_{t-3} + (1 - \alpha_1 - \alpha_2 - \alpha_3)\pi_{t-4} + \alpha_y y_{t-1} + \varepsilon_{\pi,t}, \tag{7}$$

$$y_t = \beta_1 y_{t-1} + \beta_2 y_{t-2} + \beta_3 y_{t-3} + \beta_4 y_{t-4} - \beta_r(\overline{r}_t - \overline{\pi}_t) + \varepsilon_{y,t}, \tag{8}$$

where variables with a bar are defined as $\overline{x}_t = \frac{1}{4}\sum_{j=0}^{3} x_{t-j}$. We demean the variables prior to estimation; hence the equations do not contain any constant terms.[12] The baseline reaction function is

$$\begin{aligned} r_t &= \rho_1 r_{t-1} + \rho_2 E_t r_{t+1} + (1 - \rho_1 - \rho_2)(\gamma_\pi \overline{\pi}_t + \gamma_y y_t) + \varepsilon_{r,t}, \\ \varepsilon_{r,t} &= \lambda_r \varepsilon_{r,t-1} + \xi_{r,t}. \end{aligned} \tag{9}$$

The motivation for allowing for autocorrelation in the disturbance term is to guard against misspecification of the target rule: Rudebusch (2002) argues that the significance of the lagged interest rate term in estimated reaction functions is due to the erroneous omission of serially correlated variables. However, English, Nelson, and Sack (2003) find that partial adjustment plays an important role in describing the behavior of the federal funds rate, even if one allows for serially correlated errors.

The estimates of the parameters in the reaction function are reported in table 1.[13,14] The estimates of the coefficients on the

[12]Compared with the specification in Rudebusch and Svensson (1999), the IS curve includes two extra lags of the output gap. The extra lags improve the empirical fit of the model and are needed to eliminate the autocorrelation in the residuals.

[13]The maximum-likelihood estimates are obtained using the Matlab routines provided by Jeffrey Fuhrer. The closed-form solution is derived using the Anderson-Moore algorithm (see Anderson and Moore 1985), and the likelihood function is maximized using Matlab's sequential quadratic programming algorithm `constr`. The estimation procedure does not impose any restrictions on the variance-covariance matrix of the (structural) shocks.

[14]The estimates of the parameters in the auxiliary models for inflation and output are documented in the appendix.

Table 1. ML Estimates, 1987:Q3–2007:Q4

Parameter	Estimate	SE
ρ_1	0.4092	0.04376
ρ_2	0.4993	0.09019
γ_π	2.3781	0.68540
γ_y	2.565	1.33497
λ_i	0.8798	0.07516
	Statistic	***p*-value**
Ljung-Box Test for Autocorrelation Q(12)	6.3658	0.89653
Value of Likelihood Function	−189.419	

lagged interest term and the forward term are both positive and statistically significant, thus confirming the results in Gabriel, Levine, and Spencer (2009).[15] The estimate of the autoregressive coefficient in the process for the disturbance term is 0.9 and is statistically significant. Thus, the significance of the coefficients on the interest rate term should not reflect the omission of serially correlated variables in the specification of the target rate.

Table 2 reports the estimates of the reaction function when the target rate is assumed to depend on average inflation four periods ahead and the output gap one period ahead—that is,

$$\begin{aligned} r_t = {} & \rho_1 r_{t-1} + \rho_2 E_t r_{t+1} \\ & + (1 - \rho_1 - \rho_2)(\gamma_\pi E_t \overline{\pi}_{t+4} + \gamma_y E_t y_{t+1}) + \varepsilon_{r,t}. \end{aligned} \quad (10)$$

Following Rudebusch and Svensson (1999), we assume that the current (period t) state variables are included in the central bank's information set. As is evident from the table, the estimate of the coefficient on the forward term is now slightly smaller, but it is still statistically significant. The remaining parameters are not much affected.

[15]For comparison we also estimated the partial-adjustment version of the interest rate rule (i.e., excluding the forward interest rate term). The least-squares estimate of the coeffcient on the lagged interest rate is then 0.78.

Table 2. ML Estimates with Forward-Looking Target Rule, 1987:Q3–2007:Q4

Parameter	Estimate	SE
ρ_1	0.4582	0.0734
ρ_2	0.4234	0.1062
γ_π	2.3267	0.7405
γ_y	2.656	0.7349
λ_i	0.8603	0.0574
	Statistic	***p*-value**
Ljung-Box Test for Autocorrelation Q(12)	6.7673	0.8726
Value of Likelihood Function	−189.64	

We also estimated the reaction function using market expectations of the interest rate as a proxy for the expected policy rate interest rates one period ahead. We construct a measure of the expected future interest rate from the six-month and three-month LIBOR interest rates.[16] To guard against measurement error bias, we estimate the reaction function using GMM (see the discussion in Brissimis and Magginas 2008). The results are reported in table 3.[17,18] Again we find that the estimates of the coefficients on the interest rate terms are both positive and statistically significant. The J-statistic has a p-value of 0.71; hence, we cannot reject the validity of the over-identifying restrictions.

[16] We use the expectation hypothesis of the term structure of interest rates to compute the expected future interest rate: $(1+\frac{r_6}{100})^6 = (1+\frac{r_3}{100})^3(1+\frac{r_{impl,63}}{100})^{6-3}$, where r_3 and r_6 is three-month and six-month LIBOR, respectively, and $r_{impl,63}$ is the three-month forward rate to begin in three months.

[17] We use a heteroskedasticity and autocorrelation (HAC) consistent estimate of the variance-covariance matrix of the sample moments in the GMM estimator. The autocovariances are weighted using a Bartlett kernel with a bandwidth equal to 3 (selected using Newey West). The estimation results are obtained using EVIEWS.

[18] As discussed above, the monetary policy shock appears to be serially correlated, hence the first lag of the interest rate is not a valid instrument. Moreover, since forward rates are strongly correlated with the federal funds rate, we omit the first lag of forward rates from the instrument set.

Table 3. GMM Estimates with Forward Rates as Proxy for Expected Key Interest Rate, 1987:Q3–2007:Q4

Parameter	Estimate	SE
ρ_1	0.4625	0.0423
ρ_2	0.5000	0.0553
γ_π	1.9441	0.7742
γ_y	2.4927	0.8305
	Statistic	***p*-value**
J-test	6.2858	0.7110
Instrument set: $\{y_{t-j}, \pi_{t-j}\}_{j=0}^{3}\{r_{t-j}, r_{t-j}^{impl}\}_{j=2}^{3}$		

4. Conclusion

We show that the Calvo-type interest rate rule suggested by Levine, McAdam, and Pearlman (2007) can be derived from a preference for interest rate smoothing. Using both maximum-likelihood system estimation and single-equation estimation where the implied forward interest rate is used as a proxy for the expected future interest rate, we find additional empirical support for the Calvo-type rule.

Appendix. Estimated Auxiliary Models

Table 4. ML Estimates of Parameters in Auxiliary Model, 1987:Q3–2007:Q4 (Baseline Target Rate)

Phillips Curve			IS Curve		
Parameter	**Estimate**	**SE**	**Parameter**	**Estimate**	**SE**
α_1	0.2901	0.10625	β_1	1.1761	0.11660
α_2	0.1335	0.10634	β_2	0.0055	0.18072
α_3	0.1475	0.10655	β_3	−0.2445	0.14646
α_y	0.1221	0.05500	β_4	−0.0192	0.08687
			β_r	0.0108	0.02002

Table 5. ML Estimates of Auxiliary Model, 1987:Q3–2007:Q4 (Forward-Looking Target Rate)

Phillips Curve			IS Curve		
Parameter	**Estimate**	**SE**	**Parameter**	**Estimate**	**SE**
α_1	0.2735	0.1014	β_1	1.1358	0.1204
α_2	0.1303	0.1022	β_2	0.0738	0.1678
α_3	0.1733	0.1081	β_3	−0.2942	0.1377
α_y	0.1216	0.0561	β_4	−0.0010	0.0914
			β_r	0.0103	0.0182

References

Anderson, G., and G. Moore. 1985. "A Linear Algebraic Procedure for Solving Linear Perfect Foresight Models." *Economics Letters* 17 (3): 247–52.

Brissimis, S. N., and N. S. Magginas. 2008. "Inflation Forecasts and the New Keynesian Phillips Curve." *International Journal of Central Banking* 4 (2): 1–22.

Clarida, R., J. Galí, and M. Gertler. 2000. "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory." *Quarterly Journal of Economics* 115 (1): 147–80.

English, W. B., W. R. Nelson, and B. P. Sack. 2003. "Interpreting the Significance of the Lagged Interest Rate Term in Estimated Monetary Policy Rules." *Contributions to Macroeconomics* 3 (1): 1–16.

Gabriel, V. J., P. Levine, and C. Spencer. 2009. "How Forward-Looking Is the Fed? Direct Estimates from a 'Calvo-Type' Rule." *Economics Letters* 104 (2): 92–95.

Hall, A. R. 2005. *Generalized Method of Moments.* Oxford, UK: Oxford University Press.

Jondeau, E., H. Le Bihan, and C. Galles. 2004. "Assessing Generalized Method-of-Moments Estimates of the Federal Reserve Reaction Function." *Journal of Business and Economic Statistics* 22 (2): 225–39.

Levine, P., P. McAdam, and J. Pearlman. 2007. "Inflation-Forecast-Based Rules and Indeterminacy: A Puzzle and a Resolution." *International Journal of Central Banking* 3 (4): 77–110.

Rotemberg, J. J. 1982. “Monopolistic Price Adjustment and Aggregate Output.” *Review of Economic Studies* 49 (4): 517–31.

Rudebusch, G. D. 2002. “Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia.” *Journal of Monetary Economics* 49 (6): 1161–87.

Rudebusch, G. D., and L. E. O. Svensson. 1999. “Policy Rules for Inflation Targeting.” In *Monetary Policy Rules*, ed. J. B. Taylor, 203–46. Chicago: Chicago University Press.

Woodford, M. 2003. “Optimal Interest-Rate Smoothing.” *Review of Economic Studies* 70 (4): 861–86.

Does Monetary Policy React to Asset Prices? Some International Evidence*

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Previous estimates of the monetary policy response to stock market fluctuations in the United States are found to be sizable and significant once the simultaneous interdependence between stock prices and interest rates is properly taken into account. We show that this result is not confirmed when we apply the analysis to other countries and when we consider an extended sample period including the past decade for the United States. We do not find any response in the European Union and in six inflation-targeting countries, with the exception of Australia. Moreover, we find that the response in the United States declines over time and becomes not statistically significant during the housing bubble period (2003–07).

JEL Codes: E44, G10.

1. Introduction

The recent financial crisis and the previous series of boom-bust cycles in asset prices show that fluctuations in financial variables can have a potentially large impact on the macroeconomic outlook. In fact, large swings in asset prices are often associated with strains in the financial sector and in the real economy (see Borio, Kennedy, and

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Prowse 1994 and Borio and Lowe 2002, among many others). While the main concern of central bankers—the inflation rate—is now low and stable in the main developed economies, financial instability has become a central issue in the formulation of the monetary policy strategy.[1]

A small but growing empirical literature takes a positive perspective and attempts to estimate the monetary policy response to stock market fluctuations. Early contributions either impose a zero interest rate response on impact (see Thorbecke 1997 and Neri 2004, among others, for VAR studies) or find evidence of a very small response, always statistically insignificant and sometimes even negative (see Bernanke and Gertler 2000 for a GMM estimation of Taylor rules). More recent contributions argue that previous studies fail to find a significant interaction between monetary policy and asset prices because they do not properly consider the simultaneous interdependence between interest rates and asset prices. This endogeneity problem is ruled out by assumption in early VAR contributions and is addressed only partially in the estimation of monetary policy rules by using instruments that are likely to be weak. When the endogeneity issue is addressed carefully, a positive and significant response of monetary policy to stock prices emerges, at least in the United States. This is so in Rigobon and Sack (RS, henceforth) (2003) in a VAR identified through heteroskedasticity, in Castelnuovo and Nisticò (2010) in an estimated dynamic stochastic general equilibrium (DSGE) model where monetary policy is allowed to respond to fluctuations in the stock market, in Bjørnland and Leitemo (2009) in a VAR identified using a combination of short-run and long-run restrictions, and also in Chadha, Sarno, and Valente (2004) in a Taylor-rule estimation with GMM where the authors carefully check the quality of the instruments and use an adjusted labor share

[1]The conventional wisdom is that central banks react to asset prices only to the extent that they affect expected inflation and the output gap (Bernanke and Gertler 2001). According to this view, a more aggressive response to asset prices has destabilizing effects because it is difficult to know whether a change in asset prices is due to fundamental factors or not. These claims have not stood unchallenged. Cecchetti et al. (2000), Bordo and Jeanne (2002), and Akram and Eitrheim (2008), among others, provide examples where a more proactive policy has stabilizing effects on the economy.

as the appropriate, and theoretically grounded, proxy for the output gap. Importantly, all these studies argue that the size of the estimated response is compatible with statements made by central bank officials. Therefore, the estimated response would not reflect an attempt to target asset prices, but just an indirect response to the impact that asset prices have on aggregate demand, in particular on consumption and investment decisions.

In this paper we complement previous studies by extending the analysis to several other countries and by using data for the past decade in the United States. Our main result is that there are good reasons to believe that the response to stock prices in normal times has been substantially lower than that found in previous studies.

Two kinds of considerations motivate our analysis. First, previous studies concentrate their attention on the U.S. economy. To the best of our knowledge, only a few contributions consider other countries. Chadha, Sarno, and Valente (2004) also estimate Taylor rules for the United Kingdom and Japan, whereas Bohl, Siklos, and Werner (2007) extend the RS (2003) analysis to Germany over the period 1985–98. In this paper we provide evidence for the European Union (EU) and six inflation-targeting countries (Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom). Second, none of the studies quoted above consider recent data. Bjørnland and Leitemo (2009) estimate their VAR over the period 1983–2002 using monthly data. Chadha, Sarno, and Valente (2004) consider the period 1979–2000 using quarterly data. RS (2003) use a sample of daily data from January 1985 to December 1999.[2] In our paper, instead, we use data until July 2007, when the current financial crisis begins.

We follow the approach proposed by RS (2003) and we use a VAR identified through heteroskedasticity. This approach can obtain reliable estimates from a sample of just a few years of data, since it relies on data at daily frequency. This is especially important given our focus on the past decade, which was not included in previous studies.

Three sets of results bring us to the conclusion that the interest rate response to stock prices in the past two decades has been

[2] Castelnuovo and Nisticò (2010) use quarterly data until 2007, but their sample starts in 1954. Therefore, their results are not directly comparable with ours.

substantially lower than what was found in previous state-of-the-art studies. First, when we apply the RS (2003) procedure to the EU and six inflation-targeting economies, we are able to identify a significant, albeit small, response only in Australia. This highlights the peculiarity of the U.S. case when compared with the international evidence. Second, we show that the sizable response estimated by RS (2003) for the period 1985–99 in the United States relies heavily on the effects of the stock market crash of 1987. When we reestimate the same model over the period 1988–99, the estimated response is substantially reduced, although still statistically significant. Third, when we estimate the same model over the so-called bubble period (2003–07), the monetary policy response becomes statistically insignificant.

The fact that the response of monetary policy to stock prices has declined over time in the United States is a robust feature of our model. Interestingly, Taylor (2007) identifies a substantial deviation of U.S. monetary policy from the prescription dictated by a Taylor rule responding to inflation and the output gap over the period 2003–07. A lower response to asset prices is consistent with the Taylor view that U.S. monetary policy over the period 2003–07 was more expansionary than in the previous period.

This paper has the following structure. In section 2 we present the RS methodology and we replicate the RS main result. In section 3 we provide a cross-country analysis and we extend the evidence for the United States. In section 4 we discuss some critical issues related to the identification-through-heteroskedasticity approach. Concluding remarks are contained in section 5.

2. The Model

VAR models are the most frequently used tool to measure the interactions between macroeconomic variables. As we are interested in interest rates and stock prices, the structure of the simplest VAR is the following:

$$A \begin{bmatrix} i_t \\ s_t \end{bmatrix} = C(L) \begin{bmatrix} i_{t-1} \\ s_{t-1} \end{bmatrix} + B \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix},$$

where i_t is the three-month Treasury-bill interest rate, s_t is stock returns, A is a 2×2 matrix that describes contemporaneous relationships among the variables, $c(L)$ is a finite-order lag polynomial, and ε_t and η_t are structural disturbances. B is a 2×2 matrix in which non-zero off-diagonal elements allow some shocks to affect both endogenous variables.

We use the three-month Treasury-bill rate rather than the federal funds rate. While the federal funds rate is changed every six weeks or so, the three-month Treasury-bill rate adjusts daily, reflecting expectations of future variations in the federal funds rate, and is closely monitored by the central bank. RS (2003) and Bohl, Siklos, and Werner (2007) carefully motivate this choice. In this paper we follow the literature to facilitate results comparability.

The usual assumptions to achieve identification in this kind of model are to impose a triangular form to matrix A (Cholesky decomposition) and a diagonal structure to matrix B. In this way the model is exactly identified. However, a triangular matrix A implies that one of the two variables does not react contemporaneously to the other. This assumption, which is reasonable in other contexts, is clearly inappropriate in this case. In our application, each shock to one of the variables has an immediate effect on the other in the financial markets.

RS do not impose a triangular structure on matrix A, and they build an identification procedure relying on the heteroskedasticity that is present in the data and that is usually not considered in VAR studies. It is crucial to use daily data to exploit fully the heteroskedasticity present in the data. In fact, heteroskedasticity diminishes considerably in lower-frequency data.

Realizations of interest rates and stock returns can be seen as the intersection between two schedules. The first is the reaction function of asset prices to changes in the interest rate (supposed to be downward sloping because an increase in the interest rate lowers the discounted value of future dividends, i.e., the value of the asset). The second is the reaction of the interest rate to the evolution of the stock market. The objective of the procedure is to estimate the slope of this second schedule. Because of heteroskedasticity, endogeneity, and unobservability of a common shock z_t, introduced below, OLS estimates are biased. Thus, we look for a variable (an instrument) that shifts the stock market curve without affecting the monetary policy response. An increase in the variance of the stock market shock changes the covariance between stock returns and the interest rate, and this change plays the role of an instrument.

RS estimate the following VAR:

$$
\begin{aligned}
i_t &= \beta s_t + \theta x_t + \gamma z_t + \varepsilon_t \\
s_t &= \alpha i_t + \phi x_t + z_t + \eta_t,
\end{aligned}
$$

where i_t is the three-month interest rate (T-bill), s_t is the daily return on a stock market index, x_t includes five lags of the two endogenous variables and some macroeconomic shocks (measured as monthly releases of some macro indicators and subtracting the value expected by market participants; see RS 2003), and z_t represents some unobserved shocks affecting both i_t and s_t. The common shock takes into account any macroeconomic shock not included in x_t or shifts in agents' risk preferences.[3] ε_t is a monetary policy shock, and η_t is a stock market shock.

The structure of the model is quite rich, but our objective is very simple. We want to estimate the coefficient β that measures the response of the interest rate to the stock market return.

The assumption on the correlation structure of the shocks is the following: the shocks ε_t and η_t and the unobserved shock z_t are supposed to be orthogonal, and at this stage all three can be heteroskedastic. Note that orthogonality of ε_t and η_t does not imply that disturbances are uncorrelated: In fact, the presence of z_t induces correlation.

The system cannot be estimated directly, because of the endogeneity problem discussed above and because z_t is an unobservable variable, but we can write it in reduced form:

$$
\begin{bmatrix} i_t \\ s_t \end{bmatrix} = \Phi x_t + \begin{bmatrix} \upsilon_t^i \\ \upsilon_t^s \end{bmatrix},
$$

where

$$
\begin{aligned}
\upsilon_t^i &= \frac{1}{1-\alpha\beta}((\gamma+\beta)z_t + \beta\eta_t + \varepsilon_t) \\
\upsilon_t^s &= \frac{1}{1-\alpha\beta}((1+\alpha\gamma)z_t + \eta_t + \alpha\varepsilon_t).
\end{aligned}
$$

[3]The impact of z_t on s_t is normalized to one.

In the VAR literature, it is common practice to recover the estimates of structural-form parameters from reduced-form residuals. Given the structure of correlations specified above, the covariance matrix of reduced-form residuals is as follows:

$$\Omega = \frac{1}{(1-\alpha\beta)^2} \times \begin{bmatrix} (\beta+\gamma)^2\sigma_z^2 + \beta^2\sigma_\eta^2 + \sigma_\varepsilon^2 (1+\alpha\gamma)(\beta+\gamma)\sigma_z^2 + \beta\sigma_\eta^2 \\ \qquad + \alpha\sigma_\varepsilon^2 (1+\alpha\gamma)^2\sigma_z^2 + \sigma_\eta^2 + \alpha^2\sigma_\varepsilon^2 \end{bmatrix}.$$

By estimating the model in reduced form, we obtain a consistent estimate for the covariance matrix of reduced-form residuals. Unfortunately, the covariance matrix provides only three moments, which are not enough to achieve identification. The maximum number of parameters that can be identified is three, but in matrix Ω we have six unknowns: α, β, γ, σ_z^2, σ_η^2, σ_ε^2. Hence, we do not have enough restrictions to recover the structural-form parameters.

Still, heteroskedasticity can help in our task if we can identify different regimes for the covariance matrix of the reduced-form residuals. The additional regimes provide new restrictions and may enable us to identify the parameters of the structural form. Unfortunately, for each new regime indexed by the subscript i, we add three new equations but also three new unknowns: $\sigma_{i,z}^2$, $\sigma_{i,\eta}^2$, $\sigma_{i,\varepsilon}^2$. Nevertheless, if we assume that the monetary policy shock ε is homoskedastic (thus σ_ε^2 is constant across regimes), we add three equations and only two unknowns for each regime. With three regimes we have nine equations and ten unknowns $(\alpha, \beta, \gamma, \sigma_\varepsilon^2, \sigma_{1,z}^2, \sigma_{1,\eta}^2, \sigma_{2,z}^2, \sigma_{2,\eta}^2, \sigma_{3,z}^2, \sigma_{3,\eta}^2)$, but this is enough to achieve partial identification, and in particular we can estimate the parameter β.

The assumption that σ_ε^2 is constant is not very restrictive because it does not imply that i_t is homoskedastic. In fact, the variance of the interest rate is also composed of $\sigma_{i,z}^2$ and $\sigma_{i,\eta}^2$, which change through time.[4] The other essential assumption to achieve identification is

[4]The assumption of homoskedastic monetary shocks implies that we have to limit our attention to stable monetary regimes. For example, the RS procedure is not accurate (and delivers unstable results) for the case of Japan, where the central bank was forced to change its strategy many times in the past twenty years to fight against the zero lower bound.

that the parameters α, β, and γ are constant across regimes. This is common practice in the VAR literature, also when heteroskedasticity is not considered.

RS (2003) show that with three regimes, one solution of the following quadratic equation is a consistent estimator for β:

$$a\beta^2 - b\beta + c = 0,$$

where $a = \Delta\Omega_{31,22}\Delta\Omega_{21,12} - \Delta\Omega_{21,22}\Delta\Omega_{31,12}$, $b = \Delta\Omega_{31,22}\Delta\Omega_{21,11} - \Delta\Omega_{21,22}\Delta\Omega_{31,11}$, and $c = \Delta\Omega_{31,12}\Delta\Omega_{21,11} - \Delta\Omega_{21,12}\Delta\Omega_{31,11}$.[5]

With four regimes, we have overidentifying restrictions that allow us to estimate β by GMM.

An advantage with this model is that many assumptions are testable. In fact, if the model is correctly specified, we should find the same results for β under any three regimes, since the parameter β is supposed to be constant in the sample period. If not, the parameters are unstable across regimes, the assumption of homoskedasticity for the monetary policy shock is not correct, or there are non-linearities that are not captured in the RS specification.

Thus far, we have proved that with at least three regimes we are able to consistently estimate the parameter β. To determine the regimes, we estimate the VAR in reduced form and take the residuals. The heteroskedasticity of the shocks allows us to identify four regimes: regime 1 where both shocks have low volatility, regime 2 where the interest rate shock has low volatility and the stock market shock has high volatility, regime 3 where both shocks have high volatility, and regime 4 where the interest rate shock has high volatility and the stock market shock has low volatility. We split the observations into the four regimes according to the following criterion: one observation is considered to have high variance if the thirty-day rolling variance of the residual is more than one standard deviation over the average of the series. RS admit that this approach is arbitrary, but at least two arguments can justify this choice:

(i) As shown in Rigobon (2003), the estimates are consistent even if the regimes are badly specified. The estimates are

[5] $\Delta\Omega_{31,22}$ is the (2,2) element of matrix $\Delta\Omega_{31}$. $\Delta\Omega_{31} = \Omega_3 - \Omega_1$.

not consistent only if the misspecification is so large that the system fails to meet the following order condition:[6]

$$\Omega_{11,i}\Omega_{12,j} - \Omega_{11,j}\Omega_{12,i} \neq 0$$

for regimes i and j with $i \neq j$.

This condition has an intuitive explanation. It fails when two covariance matrices are proportional, i.e., relative variances are constant across regimes. In this case, some moment conditions are not independent and heteroskedasticity cannot be helpful (for a proof of this result and more details, see Rigobon 2003).[7]

(ii) The same criterion is largely used in the literature to identify periods of excessive volatility in asset markets (Bordo and Jeanne 2002).

The last step is to compute the distributions of the estimated coefficients. The distributions are calculated by bootstrap by using the asymptotic distribution of the covariance matrices. We simulate 1,000 draws for each Ω_i. For each covariance matrix, we estimate β using different subsets of regimes. In the end, we obtain 1,000 estimates and we are able to compute the distributions.

In our first experiment, we replicate the results of RS (2003). The sample period is January 1985–December 1999 and the data are daily. Data on three-month interest rates (T-bill rate for the United States) and stock market indexes (S&P 500 for the United States) are taken from Datastream and Bloomberg.

RS (2003) include in the variable x_t some observable macroeconomic shocks measured as the difference between the released value and the expected value of five monthly macroeconomic indicators.

[6] In this case, the parameters are not even identified because this condition is the equivalent of the rank condition that is tested in the identification literature once the order condition (number of equations equal to number of unknowns to achieve just-identification) is satisfied.

[7] Rigobon (2003) proves that the estimates are consistent also when the windows of the heteroskedasticity are wrongly specified (but the number of regimes is correct) and when there are more regimes in the data than the ones assumed in the model. Of course, if the regimes are badly specified, the differences across regimes are lower and the power of the model is reduced.

Table 1. Estimates for the United States, 1985–99

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0183	0.0180	0.0185	0.0177	0.0191
Std. Dev. of Distribution	0.0457	0.0093	0.0055	0.0055	0.0058
Mass Below Zero	0.7%	2.3%	0.1%	0.2%	0.2%

Table 2. Results of RS (2003), the United States, 1985–99

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0210	0.0214	0.0210	0.0273	0.1402
Std. Dev. of Distribution	0.0052	0.0058	0.0052	0.314	3.8122
Mass Below Zero	0.0%	0.0%	0.0%	1.4%	1.4%

The role of these shocks in the model is negligible, and in fact we are even able to reproduce the results for the United States without them. In our specification, the variable x_t only consists of five lags of the two endogenous variables.

The results (shown in table 1) are consistent with the assumptions of the model. The estimates of the coefficient β are almost identical across regimes and are very similar to the ones found by RS (2003), which we present in table 2.

The reader may notice that the quality of estimates involving both regimes 3 and 4 in the RS specification is significantly lower than the others. In regimes 3 and 4, the residuals in the interest rate equation exhibit high volatility. Apparently, regimes with high stock market volatility are more useful for identifying the parameters. This fact is confirmed in all our estimates throughout the paper.

The main result of the RS paper is that by employing an appropriate identification procedure, the reaction of monetary policy to stock price movements is positive and significant. A point estimate of 0.018 means that a 10 percent rise in the S&P 500 index increases the three-month interest rate by 18 basis points. RS argue that this result is very plausible and corresponds to the impact of stock prices on aggregate demand, mainly through the wealth effect on consumption.

Table 3. Estimates for the European Union, 1999–2006

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	−0.0003	0.0005	0.0005	0.0062	0.0063
Std. Dev. of Distribution	0.0193	0.0006	0.0006	0.0070	0.0086
Mass Below Zero	18.9%	19.5%	18.4%	17.6%	18%

3. Results

In this section we present our results, showing that the monetary policy response to stock prices in normal times has been substantially lower than that found in previous studies. First, we present evidence for the EU and six inflation-targeting countries. Second, we extend the analysis for the United States.

3.1 International Evidence

We estimate our model with data for the EU and six inflation-targeting countries (Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom).

We present the results for the EU in table 3. The sample period is January 1999–April 2006:

The European Central Bank (ECB) reaction turns out to be not significant. The size of the sample is smaller than for the other countries, due to the introduction of the euro in 1999, but is sufficient to derive some evidence on European monetary policy. In 2002, ECB President Trichet said explicitly that "it is clearly not opportune to introduce asset prices in the central bank's reaction function," and our result confirms this view.[8] A topic of discussion is whether we should find some sign of indirect reaction. Trichet recognizes that recent changes in asset prices have influenced private spending more than past swings because of the more widespread stock ownership in a number of industrialized countries. He states, however, that this

[8]President Trichet's speech is available at www.banque-france.fr/gb/instit/telechar/discours/sp230402.pdf.

Table 4. Estimates for Inflation-Targeting Countries Based on All Regimes

	Australia 93–07	Canada 95–07	New Zealand 94–07	Norway 01–07	Sweden 93–07	United Kingdom 93–07
Mean	0.0058	0.0000	−0.0035	−0.0046	0.0009	0.0010
Std. Dev.	0.0023	0.0017	0.0040	0.0039	0.0006	0.0009
Mass Below Zero	0.3%	51.3%	80.9%	88.9%	9.8%	13.9%

impact is still low compared with the United States. And in fact our results do not show any sign of indirect reaction.

We now turn our attention to several inflation-targeting countries: Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom. Since our procedure requires stability in the monetary policy framework, we decided to run estimates over a period of stable inflation, according to the classification in Roger (2009). New Zealand was the first country to adopt inflation targeting in 1990 and achieved a stable inflation rate in 1993. Canada adopted inflation targeting in February 1991 and was able to stabilize inflation in early 1995. Our estimation period starts in February 1994 for New Zealand (due to data availability) and in January 1995 for Canada. According to Roger (2009), the remaining four countries adopted inflation targeting when inflation was already under control. Therefore, we will start our estimation period at the beginning of the inflation-targeting regime in January 1993 for Australia, Sweden, and the United Kingdom and in March 2001 for Norway.[9]

The only country that has reacted significantly to stock prices under the inflation-targeting regime is Australia (table 4). Notice, however, that the size of the response is rather low and much smaller than in the United States. The response is even smaller, and significant only at the margin, in Sweden and the United Kingdom. We do not detect any response in Canada, whereas we find a somewhat negative response in New Zealand and Norway. We have not found

[9]Very few observations enter the high-volatility regimes in Norway, Sweden, and the United Kingdom. In these cases, we lower the window to one-half of a standard deviation instead of one standard deviation.

interesting subsample dynamics, as is the case for the United States (cf. below). On average, our results square well with previous estimates by Bohl, Siklos, and Werner (2007) for Germany and Chadha, Sarno, and Valente (2004) for the United Kingdom. It seems that stock prices played a limited role in the monetary policy strategy in inflation-targeting open economies.

According to the RS (2003) interpretation, the estimates across countries should be proportional to the financial wealth effect. Slacalek (2009) estimates the magnitude of the financial wealth effect in several countries. He finds significant effects in the United States, Australia, Canada, Sweden, and, to some extent, also in the United Kingdom. The estimates for other European countries are much lower and statistically not significant, in keeping with Altissimo et al. (2005) and Paiella (2007). In some cases our international evidence on the monetary policy response to stock prices is consistent with the above results: this is the case for the United States, Australia (although our estimated response is quantitatively very small), and the EU. However, we do not detect a meaningful response in Canada, Sweden, and the United Kingdom. Therefore, our results do not offer material for a systematic analysis of the cross-country variation in the monetary policy reaction. Instead, our results strongly highlight the peculiarity of the U.S. case in the international context, given that it is only in the United States that we find a non-negligible response, at least until 2003 (cf. below).[10] Our evidence seems to justify the fact that the literature focuses to such an extent on the United States and that only a few contributions consider other countries.

3.2 More Results for the United States

We now turn our attention to the United States and we extend the RS analysis until the beginning of the recent financial crisis.

A first legitimate question is whether the stock market crash of 1987 drove the RS result to some extent. It may very well be that such an unusual event required a large policy response. To analyze

[10] Assenmacher-Wesche and Gerlach (2008) study how the financial structure influences the impact of monetary policy on asset prices. They conclude that it plays a limited role. Our cross-country analysis for the opposite relationship (the impact of asset prices on monetary policy) suggests a similar result.

Table 5. Estimates for the United States, 1988–99

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0085	0.0090	0.0083	0.0362	0.0412
Std. Dev. of Distribution	0.0024	0.0024	0.0024	0.0145	0.0216
Mass Below Zero	0%	0%	0%	0.2%	2%

Table 6. Estimates for the United States, 1988–2003

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0080	0.0078	0.0080	0.0310	0.0157
Std. Dev. of Distribution	0.0016	0.0017	0.0017	0.0227	0.0524
Mass Below Zero	0%	0%	0%	6.5%	31.2%

this issue, we start our estimation in January 1988 instead of January 1985. The results shown in table 5 are considerably different from the evidence provided above (cf. table 1).

The estimated coefficient is still positive and strongly significantly different from zero. However, its magnitude is reduced from 0.018 to 0.0085 in the estimate using four regimes. Such a lower response is confirmed when we extend the sample until February 2003 (table 6). Results are rather stable and indicate an interest rate response much lower than the one estimated by RS (2003) for the period 1985–99.

So far we have disregarded the period from March 2003 to July 2007 because this is the period when stock and real estate prices surged, the so-called bubble period that preceded the current deep crisis. Taylor (2007) argues that during this period U.S. monetary policy was more expansionary than in the recent past: "During the period from 2003 to 2006 the federal funds rate was well below what experience during the previous two decades of good macroeconomic performance would have predicted."

On the basis of this argument, we believe that the period 2003–07 deserves a separate investigation to see whether we observe variations in the response to stock prices.

Table 7. Estimates for the United States, 2003-07

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0025	0.0018	0.0024	0.0162	0.0101
Std. Dev. of Distribution	0.0047	0.0056	0.0047	0.0641	0.0284
Mass Below Zero	28%	35%	28.7%	44.7%	30.2%

In fact, our analysis reveals results that are consistent with Taylor (2007). The size of the interest rate response is reduced even further over 2003–07 and, most importantly, becomes statistically insignificant (table 7).[11]

In his analysis, Taylor (2007) considers the policy response to inflation and the output gap. Our analysis identifies that at the same time there is also a lower response to asset prices, at least compared with the recent past. However, let us stress that our result does not necessarily imply that monetary policy was too expansionary over that period. As Taylor (2007) recognizes, at that time there were good reasons for a prolonged period of low interest rates—in particular, to avoid a deflationary spiral similar to that of Japan in the 1990s. More research is definitely needed on these issues.

All in all, our analysis shows that the policy response estimated using identification through heteroskedasticity declines over time and is not statistically different from zero during the "bubble period." Moreover, we show that sizable responses identified in previous studies rely heavily on the stock market crash of 1987. Therefore, our analysis offers sound evidence to suggest that the monetary policy response to stock prices in normal times has been substantially lower than that found in previous studies.

It is reassuring that some very recent papers find evidence supportive of our results. First, Ravn (2010) applies the RS procedure

[11] In the working paper version of this article (Furlanetto 2008) we show that the lower response to stock prices was not compensated by a higher response to a proxy for real estate prices given by the REIT index. In fact, the response to real estate prices becomes also statistically insignificant over the "bubble period." Interestingly, however, the response to real estate prices is increasing over time up to 2003, unlike for stock prices.

for the United States over the period 1998–2008, allowing for asymmetric responses to stock market increases and decreases. Consistent with our results, he finds that the interest rate response has declined over time. Interestingly, he shows that the response to stock market decreases is larger than the response to stock market increases. Second, Milani (2010) estimates a DSGE model with learning and meaningful asset price dynamics over the period 1960–2007 with monthly data. Interestingly, he reestimates the model for the post-1984 sample and identifies a significant reduction in the response of monetary policy to stock market fluctuations. Our evidence reinforces the Milani (2010) result, indicating that the response has declined even further when we consider the "bubble period" in isolation. Finally, Bjørnland and Leitemo (2009) estimate a monthly VAR identified through a combination of short-run and long-run restrictions. According to their estimates, a 10 percent increase in stock prices would lead to an impact increase in the interest rate of 40 basis points over the period 1983–2002. This value is somewhat larger than our estimate, although the comparison is only illustrative given that the sample is different. Moreover, given that stock market volatility (on an annualized basis) is higher at daily frequency, it is not surprising that we find a lower response using higher-frequency data. However, interestingly, Bjørnland and Leitemo (2009, online appendix) also find a sizable decline in the impact response when they use a shorter sample (starting in January 1987).

4. Identification through Heteroskedasticity: Discussion

Our paper questions the robustness and the stability of the main result in RS (2003), i.e., the positive and sizeable response of monetary policy to stock prices. However, we believe that our study confirms that the RS procedure is extremely powerful. Using identification through heteroskedasticity, it is possible to estimate the model in short samples (only a few years), whereas other econometric techniques need much larger samples. The use of daily data enables the RS approach to discover interesting subsample dynamics, as shown in the preceding sections. Using Taylor rules or other VAR, we could not detect the decline in reaction observed during the "bubble period."

Table 8. Estimates for the United States, 1988–2003, No Common Shock

	Regimes 1,2	Regimes 1,3	Regimes 1,4	Regimes All
Mean of Distribution	0.0083	−0.0031	0.0025	0.0005
Std. Dev. of Distribution	0.0018	0.0034	0.0021	0.0019
Mass Below Zero	0%	84.5%	10.6%	39%

A potential weakness in our approach is that the three-month interest rate, which is our measure of the monetary policy stance, can react for reasons that are unrelated to monetary policy. This is a relevant objection that can, however, be countered by the following considerations: The presence of the unobservable common shock z_t, in our opinion, takes into account, at least to some extent, the factors that induce variations in interest rates and stock prices that are unrelated to policy expectations (like shifts in investor risk preferences). Identification through heteroskedasticity relies heavily on the common shock and, in fact, when the common shock is excluded, the results worsen significantly. As an example, in table 8 we report the estimates for the United States over the period 1988–2003 excluding the common shock (to be compared with table 6).[12]

Moreover, RS (2003) use other measures for the interest rate variable and find that the results are confirmed. Furthermore, the fact that our estimates for the period 1985–99 are in the same ballpark as the ones in studies that use the policy rate is also reassuring (see Chadha, Sarno, and Valente 2004 and Bjørnland and Leitemo 2009).

Another potential weakness of our analysis is the use of daily data. One could argue that the response of monetary policy is extremely gradual and is not influenced by day-to-day movements in financial markets. On the other hand, the use of daily data is also the strength of our approach because it is at daily frequency that heteroskedasticity is more present and therefore more helpful in obtaining accurate estimates. Importantly, when we estimate our model for the United States with weekly and monthly data, we find similar results (table 9). As expected, identification through

[12] Only two regimes are enough to identify β in the absence of a common shock.

Table 9. Estimates for the United States, 1988–2003, Based on All Regimes

	Daily Data	Weekly Data	Monthly Data
Mean of Distribution	0.0080	0.0163	0.0146
Std. Dev. of Distribution	0.0016	0.0064	0.0108
Mass Below Zero	0%	0.2%	6.3%

heteroskedasticty achieves less-precise estimates using low-frequency data. However, estimates are still positive and significant.

A relevant question that is always present in the literature on monetary policy and asset prices is whether the identified response represents a direct or indirect reaction to asset prices. RS (2003) argue that the magnitude of their estimate is compatible with a central bank that is concerned only with the impact of the stock market on output and inflation. Fuhrer and Tootell (2008) and Milani (2010) provide strong supporting empirical evidence. According to their results, central banks respond to equity price returns not directly (the central bank is not targeting asset prices) but only indirectly, since variations in stock prices can have important implications for the macroeconomic outlook. The main result from our cross-country and cross-period analysis is that it is difficult to find even an indirect response in recent years. We believe that our results show that central banks have been extremely cautious in their response to asset price movements, in keeping with arguments in Bernanke and Gertler (2001).

5. Conclusion

In this paper, we show that previous estimates of the monetary policy response to stock market fluctuations in the United States are not confirmed when we apply the analysis to other countries and when the sample period for the United States is extended until the beginning of the recent financial crisis.

For the other countries, we find a significant, albeit small, response only in Australia. This highlights the peculiarity of the U.S. case in the international context.

For the United States, we find that the response declines over time and becomes statistically insignificant during the bubble period (2003–07). Moreover, we show that the inclusion (or not) of the stock market crash of 1987 in the sample period significantly affects the results.

As far as we know, this is the first paper that provides evidence of a significant decline in the response of monetary policy to stock prices. However, supportive evidence for our claim can be found in papers that are focused on broader issues (Bjørnland and Leitemo 2009, Milani 2010, and Ravn 2010).

We believe that our results need further investigation. An interesting extension of this paper would be the estimation of the interest rate response to the stock market in a model with time-varying parameters. We plan to work on this project in the future.

References

Akram, Q. F., and Ø. Eitrheim. 2008. "Flexible Inflation Targeting and Financial Stability: Is It Enough to Stabilize Inflation and Output?" *Journal of Banking and Finance* 32 (7): 1242–54.

Altissimo, F., E. Georgiou, T. Sastre, M. T. Valderrama, G. Sterne, M. Stocker, M. Weth, K. Whelan, and A. Willman. 2005. "Wealth and Asset Price Effects on Economic Activity." ECB Occasional Paper No. 29.

Assenmacher-Wesche, K., and S. Gerlach. 2008. "Financial Structure and the Impact of Monetary Policy on Asset Prices." Swiss National Bank, Working Paper No. 2008-16.

Bernanke, B., and M. Gertler. 2000. "Monetary Policy and Asset Price Volatility." NBER Working Paper No. 7559.

———. 2001. "Should Central Banks Respond to Movements in Asset Prices?" *American Economic Review* (Papers and Proceedings) 91 (2): 253–57.

Bjørnland, H., and K. Leitemo. 2009. "Identifying the Interdependence between US Monetary Policy and the Stock Market." *Journal of Monetary Economics* 56 (2): 275–82.

Bohl, M., P. Siklos, and T. Werner. 2007. "Do Central Banks React to the Stock Market? The Case of the Bundesbank." *Journal of Banking and Finance* 31 (3): 719–33.

Bordo, M., and O. Jeanne. 2002. "Monetary Policy and Asset Prices: Does 'Benign Neglect' Make Sense?" *International Finance* 5 (2): 139–64.

Borio, C., N. Kennedy, and S. Prowse. 1994. "Exploring Aggregate Asset Price Fluctuations across Countries: Measurement, Determinants and Monetary Policy Implications." BIS Economics Paper No. 40.

Borio, C., and P. Lowe. 2002. "Asset Prices, Financial and Monetary Stability: Exploring the Nexus." BIS Working Paper No. 114.

Castelnuovo, E., and S. Nisticò. 2010. "Stock Market Conditions and Monetary Policy in a DSGE model for the U.S." *Journal of Economic Dynamics and Control* 34 (9): 1700–1731.

Cecchetti, S. G., H. Genberg, J. Lipsky, and S. Wadhawani. 2000. *Asset Prices and Central Bank Policy.* Geneva Reports on the World Economy 2, CEPR.

Chadha, J. S., L. Sarno, and G. Valente. 2004. "Monetary Policy Rules, Asset Prices, and Exchange Rates." *IMF Staff Papers* 51 (3): 529–52.

Fuhrer, J., and G. Tootell. 2008. "Eyes on the Prize: How Did the Fed Respond to the Stock Market?" *Journal of Monetary Economics* 55 (4): 796–805.

Furlanetto, F. 2008. "Does Monetary Policy React to Asset Prices? Some International Evidence." Norges Bank Working Paper No. 2008/7.

Milani, F. 2010. "Learning about the Interdependence between the Macroeconomy and the Stock Market." Mimeo, University of California, Irvine.

Neri, S. 2004. "Monetary Policy and Stock Prices." Banca d'Italia, Working Paper No. 513.

Paiella, M. 2007. "Does Wealth Affect Consumption? Evidence for Italy." *Journal of Macroeconomics* 29 (1): 189–205.

Ravn, S. H. 2010. "Has the Fed Reacted Asymmetrically to Stock Prices?" Manuscript, University of Copenhagen and Danmarks Nationalbank.

Rigobon, R. 2003. "Identification through Heteroskedasticity." *Review of Economics and Statistics* 85 (4): 777–92.

Rigobon, R., and B. Sack. 2003. "Measuring the Reaction of Monetary Policy to the Stock Market." *Quarterly Journal of Economics* 118 (2): 639–69.

Roger, S. 2009. "Inflation Targeting at 20: Achievements and Challenges." IMF Working Paper No. 09-236.

Slacalek, J. 2009. "What Drives Personal Consumption? The Role of Housing and Financial Wealth." *The B.E. Journal of Macroeconomics* 9 (1) (Topics): Article 37.

Taylor, J. B. 2007. "Housing and Monetary Policy." In *Housing, Housing Finance, and Monetary Policy*, 463–76. Proceedings of the Federal Reserve Bank of Kansas City Annual Economic Symposium, Jackson Hole, Wyoming, August 30–September 1.

Thorbecke, W. 1997. "On Stock Market Returns and Monetary Policy." *Journal of Finance* 52 (2): 635–54.

The Financial Market Impact of Quantitative Easing in the United Kingdom*

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This paper investigates the impact of the Bank of England's quantitative easing policy on UK asset prices. Based on analysis of the reaction of financial market prices and model-based estimates, we find that asset purchases financed by the issuance of central bank reserves—which by February 2010 totalled £200 billion—may have depressed medium to long-term government bond yields by about 100 basis points, with the largest part of the impact coming through a portfolio balance effect. The wider impact on other asset prices is more difficult to disentangle from other influences: the initial impact was muted, but the overall effects were potentially much larger, though subject to considerable uncertainty.

JEL Codes: E43, E44, E52, E58.

1. Introduction

The intensification of the global financial crisis that followed the collapse of Lehman Brothers in September 2008 led to governments and central banks around the world introducing a variety

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of measures aimed at stabilizing financial conditions and supporting aggregate demand (see, e.g., Klyuev, de Imus, and Srinivasan 2009 for a review).

In the United Kingdom, a large monetary policy easing was accomplished using both conventional and unconventional measures.[1] The Bank of England's Monetary Policy Committee (MPC) cut the Bank Rate, the United Kingdom's policy rate, in a sequence of steps from 5 percent at the start of October 2008 to 0.5 percent in March 2009. But in reducing policy rates to their effective floor, the MPC also announced that, in view of the substantial downside risks to achieving the 2 percent CPI inflation target in the medium term, it would ease monetary conditions further through a program of asset purchases financed by the issuance of central bank reserves.

This policy of asset purchases has come to be known as quantitative easing (QE).[2] In general terms, QE is normally defined as a policy that expands the central bank's balance sheet, in order to increase the level of central bank money (in particular, bank reserves) in the economy (see Bernanke and Reinhart 2004). This is sometimes contrasted with a policy of changing the composition of the assets on the central bank's balance sheet (often referred to as credit easing); for example, by shifting between short and longer-maturity government bonds or by shifting into riskier private assets, such as corporate bonds or equities. The Bank of England's policy has elements of both, though the main emphasis was on expanding the balance sheet.[3] The MPC decided that it would purchase both private- and public-sector assets using central bank reserves, though the majority of purchases would be of UK government securities

[1]Though not identical, there are many similarities between the policies implemented by the main central banks during the financial crisis (see Miles 2010). D'Amico and King (2010) and Gagnon et al. (2011) review the impact of large-scale asset purchases by the U.S. Federal Reserve. Neely (2010) looks at the wider international effects of the Federal Reserve's asset purchases.

[2]The terminology was first used to describe the Bank of Japan's policy during 2001 to 2006 (see, e.g., Ugai 2007 and Shiratsuka 2009).

[3]The asset purchases were conducted through a separate legal entity, the Bank of England Asset Purchase Facility Fund, a limited company. The Fund and the Bank are fully indemnified by the Treasury from any losses arising out of or in connection with the asset purchase program. For a discussion of how asset purchases affected the Bank of England's accounts, see Bean (2009).

(gilts).[4] By purchasing financial assets from the private sector, the aim was to boost the amount of money in the economy, which would increase nominal spending and thereby ensure that inflation was on track to meet the CPI inflation target over the medium term.

By February 2010, the Bank of England had completed £200 billion of asset purchases as part of its QE policy, overwhelmingly comprising conventional gilts. Alongside separate liquidity support to the banking sector, these purchases expanded the Bank's balance sheet as a proportion of nominal GDP to three times its level before the onset of the crisis in the summer of 2007, as large as at any point in the past two centuries (see Cross, Fisher, and Weeken 2010). The Bank's gilt purchases represented 29 percent of the free float of gilts (the amount of non-official holdings of gilts) and were equivalent to around 14 percent of nominal GDP.

This paper examines the impact of these extraordinary measures on financial markets. Given their overwhelming importance, we will focus on the effects of the Bank's gilt purchases and will not directly discuss the impact of the other purchase facilities set up by the Bank. Our aim is to review how QE has affected gilt markets and how it has fed through more widely into other financial asset prices, like equities and corporate debt.

Since the motivation for the United Kingdom's QE purchases was to increase nominal spending on goods and services, in order to meet the MPC's inflation target, it might not be obvious why we should be concerned with the financial market impact per se. But judging the impact of QE in stimulating the macroeconomy is difficult, as the transmission mechanism may be subject to long lags, and it is hard to measure the specific contribution of the MPC's asset purchases, given the influence of other policy measures and other economic developments in the United Kingdom and internationally. The place where we might have expected to see the clearest direct impact of QE is in the reaction of financial markets. This in turn may provide the most timely and clearest read on the effectiveness of the policy and how it might be feeding through to the rest of the economy.

[4]The smaller purchases of corporate bonds and commercial paper were aimed at improving the functioning of those markets and therefore improving access to credit for firms (see Fisher 2010).

The rest of the paper is structured as follows. In section 2, we discuss the main channels through which QE asset purchases may affect financial markets and how we might attempt to estimate the relative importance of the various channels. Section 3 describes the evolution of the MPC's QE-related asset purchase program and how it has been implemented. In sections 4 and 5, we examine the immediate reaction of asset prices to the Bank's QE announcements, and allocate it into separate channels, using event-study analysis and survey data. Overall, our analysis suggests that the dominant effect has been through a portfolio balance channel. To provide a benchmark for the impact that might have been expected through this channel, section 6 uses two portfolio balance models estimated on pre-crisis data to quantify the impact on expected asset returns of changes in asset quantities. These results are broadly consistent with the observed initial reaction of asset prices to QE, although there is considerable uncertainty around the estimated effects, especially for equities. Section 7 draws overall conclusions.

2. QE and Asset Prices

By injecting money into the economy, in return for other assets, a central bank can increase the liquidity of private-sector balance sheets. As discussed in Benford et al. (2009), there are a number of ways through which this greater liquidity can have an impact on the economy. First, purchases of assets financed by central bank money should push up the prices of assets. This is the impact analyzed in this paper. If asset prices are higher, this reduces the cost of borrowing, encouraging higher consumption and investment spending. Higher asset prices also increase the wealth of asset holders, which should boost their spending. The other ways in which QE may potentially work—mainly, through expectations, by demonstrating that the MPC will do whatever it takes to meet the inflation target, and through influencing banks' lending ability—fall outside the scope of this paper.

2.1 Asset Price Channels

In our framework, there are three main channels through which QE might affect asset prices: macro/policy news, portfolio balance, and liquidity premia.

The *macro/policy news channel* refers to anything economic agents might learn from the Bank of England's QE announcements about the underlying state of the economy and the MPC's reaction function. This channel captures news about expected future policy rates—often referred to as the "signaling channel"[5]—but, if we define it more broadly to include perceptions of the risks around the path of future short-term interest rates, it should also include revisions to term premia. As well as affecting gilt yields, this channel will feed through into other asset prices to the extent that the relevant discount rates are affected. In principle, the overall sign of these effects on yields/prices might be either positive or negative. While QE might signal lower policy rates in the short term, it could also signal higher inflation in the future, leaving the impact on nominal gilt yields ambiguous.

The *portfolio balance channel* reflects the direct impact on asset prices of investors rebalancing their portfolios in response to the Bank of England's QE-related asset purchases. Tobin (1961, 1963, and 1969) and subsequently Brunner and Meltzer (1973) and Friedman (1978), amongst others, showed that if assets are not perfect substitutes, then a change in the quantity of a specific asset will lead, ceteris paribus, to a change in its relative expected rate of return. Thus imperfect substitutability provides a mechanism for QE-related asset purchases by the Bank to affect asset prices by inducing sellers to rebalance their asset portfolios. Provided long-term gilts and money are imperfect substitutes, QE-related gilt purchases would be expected to reduce bond yields and lead to investors increasing their demand for other long-term assets. The impact through this channel may occur both on announcement and over time as investors are able to adjust their portfolios. Since this channel depends on perceptions of the path of outstanding *stocks* of gilts and money, we would expect it to be persistent.

In conventional New Keynesian models, portfolio balance effects are not present and QE can only work through a signaling channel (see, e.g., Eggertsson and Woodford 2003). Asset purchases on their own do not change behavior because the assumptions typically

[5]Most of the related literature on QE refers to the signaling and portfolio balance channels. See, for example, Clouse et al. (2003), Bernanke, Reinhart, and Sack (2004), Ugai (2007), and Borio and Disyatat (2009).

made imply that the distinction between government and private asset holdings is unimportant, in a way reminiscent of Ricardian equivalence. In these models, QE can be effective only if it changes expectations regarding the path of future policy rates and/or inflation. This naturally leads to the conclusion that committing to a path for future interest rates may be more effective than undertaking asset purchases. But, in a model with financial frictions (e.g., credit constraints or distortionary taxes) or incomplete markets, and with imperfect substitutability between different assets, QE can also affect asset prices by changing the relative supplies of different assets.

The view that imperfect asset substitutability can be important is reflected in an emerging theoretical literature that builds microfoundations for these effects from the earlier contributions of Tobin and others. For example, Andrés, López-Salido, and Nelson (2004) introduce an adjustment to household preferences in a New Keynesian model to allow for imperfect asset substitutability between holdings of long-term bonds and money for certain households. Their framework can be thought of as a way of introducing "preferred habitat" investors (Modigliani and Sutch 1966) into a dynamic stochastic general equilibrium setting. More recently, using a partial equilibrium approach, Vayanos and Vila (2009) propose a theoretical model of preferred habitat, in which bond prices are determined through the activities of risk-averse arbitrageurs and preferred-habitat investors. In this setup, they demonstrate that shocks to bond supply are a determinant of bond prices, thus providing another rationale for expecting QE to have an effect on long-term bond yields.

In addition to the portfolio balance effect, the presence of the central bank in the market as a significant buyer of assets may improve market functioning and thereby reduce premia for illiquidity. This *liquidity premia channel* effect reflects the fact that the central bank's purchases may make it less costly for investors to sell assets when required. In normal times, markets may be deep and liquid, but in stressed conditions, premia for illiquidity could be significant. Since this channel depends on the *flow* of purchases for its effect, we would expect it to be temporary and limited to the duration of the asset purchase program.

How does the MPC's asset purchase program fit into this description? At a general level, the QE program seemed firmly based on a view that there would be significant portfolio rebalancing. The MPC's asset purchase program was directed toward large-scale purchases of conventional gilts: the impact was expected to be seen in gilt markets, but also across a broader range of asset prices and in real activity and inflation. The MPC did not explicitly use these purchases to signal future intentions, emphasizing instead its commitment to meeting the inflation target through the usual channels of monetary policy communications—including the MPC minutes and the quarterly Inflation Report. Nor were its actions focused on improving the functioning of gilt markets, where liquidity premia, even in stressed times, were considered to be small.[6]

Given the unusual character of the intervention, and the absence of a clear consensus on the exact impact of asset purchases generally, our approach is based on the notion that financial markets are incomplete or imperfect, while being agnostic on the exact source and size of any market frictions. That said, we do not want to rule out significant signaling or expectational effects, so we also investigate this channel in our empirical approach.

It is important to note here that though these channels are broadly defined compared with much of the literature on the topic, they do not capture the fact that asset purchases—with other macroeconomic policies—may have substantially changed the distribution of future macroeconomic outcomes, and thereby affected risk premia more broadly (e.g., equity risk premia). Dale (2010) discusses this in more detail.

2.2 Measuring the Asset Price Channels

In order to quantify the impact of QE purchases, we use several approaches: event-study methods are discussed in sections 4 and 5 and time-series econometric methods in section 6.

In attempting to quantify the role of the various channels in affecting gilt yields, we rely crucially on interest rates from overnight index swap (OIS) contracts. An OIS is a contract that involves the exchange of a predefined fixed interest rate (the OIS rate) with one linked to a compounded overnight interbank interest rate that has

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[6]The liquidity channel effect was nevertheless thought important for purchases of private-sector assets.

prevailed over the life of the contract. Since they settle on overnight interest rates and are collateralized, OIS rates should incorporate minimal credit risk. The OIS market has built up rapidly in recent years, and, at least at short maturities, these contracts are actively traded and should therefore also incorporate little liquidity risk.[7] On the assumption that OIS rates provide an accurate measure of default risk-free rates that are, as a derivative contract, less affected by supply constraints in the gilt market, movements in OIS rates should provide a measure of macro/policy news. Movements in the spread between gilt yields and OIS rates then represent the combined effect of the portfolio balance and liquidity channels.

To clarify our approach, it may help to start with the following well-known expression, which decomposes bond yields into expected future short-term interest rates and a term premium:

$$y(gilt)_t^n = (1/n)\sum_{i=0}^{n-1} E_t r_{t+i} + TP(gilt)_t^n, \tag{1}$$

where $y(gilt)_t^n$ is the n-period maturity yield on a government bond, r_{t+i} denotes the one-period (risk-free) short-term interest rate, and $TP(gilt)_t^n$ denotes the n-period term premium. In our framework, the term premium on gilts comprises two elements: $TP1(gilt)_t^n$, an instrument-specific effect that captures gilt-specific credit/liquidity premia and any effects from demand/supply imbalances, and $TP2(gilt)_t^n$, a term premium element that reflects uncertainty about future short-term interest rates:

$$TP(gilt)_t^n = TP1(gilt)_t^n + TP2(gilt)_t^n. \tag{2}$$

If we assume that credit risk premia on gilts are negligible, then movements in gilt-specific premia, $TP1(gilt)_t^n$, will reflect either changes in liquidity premia or demand/supply effects from QE that come through the portfolio balance channel. We examined separate evidence on market functioning (e.g., bid-ask spreads) to enable us to identify the role of the liquidity premia channel, but the importance of this channel appears to be small in the context of gilts,

[7] At longer maturities this may be less true, and it is possible that OIS rates may incorporate liquidity premia. See the discussion below.

so we place more emphasis on the relative importance of portfolio balance effects in driving gilt-specific premia around QE announcements.[8]

It is possible to write down a similar breakdown for yields implied by OIS contracts:

$$y(OIS)_t^n = (1/n)\sum_{i=0}^{n-1} E_t r_{t+i} + TP(OIS)_t^n, \tag{3}$$

where $y(OIS)_t^n$ is the n-period maturity OIS rate, r_{t+i} is the one-period short (risk-free) rate, and $TP(OIS)_t^n$ denotes the OIS n-period term premium. Again, in principle, the term premium implied by OIS rates can be broken down into two elements: $TP1(OIS)_t^n$, an instrument-specific premium, and $TP2(OIS)_t^n$, a conventional term premium.

$$TP(OIS)_t^n = TP1(OIS)_t^n + TP2(OIS)_t^n \tag{4}$$

The working assumption in our analysis is that the first $TP1(OIS)_t^n$ element is negligible, so that movements in OIS term premia reflect fundamentals to do with interest rate uncertainty rather than liquidity or credit risk premia or effects from demand/supply. A corollary of this is that the component of the gilt-yield term premium reflecting interest rate uncertainty (i.e., $TP2(gilt)_t^n$) will be the same as in the corresponding maturity-matched OIS rate:

$$TP(OIS)_t^n = TP2(OIS)_t^n = TP2(gilt)_t^n.$$

It follows that

$$\begin{aligned} y(gilt)_t^n - y(OIS)_t^n &= TP1(gilt)_t^n + TP2(gilt)_t^n - TP(OIS)_t^n \\ &= TP1(gilt)_t^n. \end{aligned} \tag{5}$$

Thus changes in the gilt-specific premia element, and the effects of the portfolio balance channel, should be proxied by changes in the spread between gilt yields and OIS rates. But to the extent

[8] See Joyce et al. (2010) for further details.

that OIS rates are driven by some of the same factors influencing gilt-specific premia (e.g., demand/supply imbalances), changes in gilt-OIS spreads will tend to underestimate the effects of portfolio rebalancing.

One implication of our approach is that QE can potentially affect the term premium through both the macro/policy news channel, as we have defined it, and through portfolio rebalancing. As we shall show in later sections, the evidence suggests on balance that the impact on gilt yields has been dominated by a portfolio balance effect, which would suggest that the term premium effect has broadly coincided with the portfolio balance effect.

3. The United Kingdom's Unconventional Policy Measures

In this section we describe the unconventional monetary policy measures that the Bank of England took in response to the financial crisis.

3.1 Initial Responses

The Bank's initial response to the financial crisis during 2007–08 included a range of measures aimed at providing liquidity insurance to the markets (see, e.g., Cross, Fisher, and Weeken 2010 for more details). The Bank's lending operations were extended beyond the amounts needed for banks to meet their pre-arranged reserves targets, which were themselves increased. The Bank conducted larger amounts of three-month repo operations and extended the collateral accepted. In April 2008, after the collapse of Bear Stearns, the Bank introduced a Special Liquidity Scheme (SLS) that allowed banks and building societies to swap high-quality, but temporarily illiquid, mortgage-backed and other securities for UK Treasury bills. Along with other central banks, in the wake of the collapse of Lehman Brothers in September 2008, the Bank established a swap facility with the Federal Reserve, providing an additional means whereby banks could borrow U.S. dollars. And, in October 2008, a Discount Window Facility was launched as a permanent liquidity insurance facility.

All these operations were aimed at providing liquidity support to the markets rather than changing the implementation of monetary policy. Towards the end of 2008, some of the extra liquidity introduced by these measures started to be drained with one-week Bank of England bills. The Bank's means of implementing monetary policy were largely unchanged until the start of the QE policy in March 2009.

3.2 The APF and QE

The Bank of England Asset Purchase Facility Fund was set up on January 30, 2009 as a subsidiary of the Bank of England. The Fund is fully indemnified by the Treasury from any losses arising out of or in connection with the Asset Purchase Facility (APF), ensuring that the Bank will not incur any losses arising from the asset purchase program (for further discussion, see Bean 2009). The APF was initially authorized to purchase up to £50 billion of private-sector assets—corporate bonds and commercial paper—financed by the issuance of Treasury bills and Debt Management Office (DMO) cash management operations, in order to improve liquidity in credit markets that were not functioning normally. The first purchases of commercial paper began on February 13, 2009.

The APF's remit was subsequently expanded to allow it to be used as a monetary policy tool ahead of the March 2009 MPC meeting. The Committee was given the option to finance purchases under the APF by issuing central bank reserves, and the range of eligible assets was expanded to include gilts. After the financial crisis worsened following the collapse of Lehman Brothers in September 2008, the MPC reduced the Bank Rate in a sequence of steps from 5 percent to 0.5 percent. When the final reduction of the Bank Rate from 1 percent to 0.5 percent was announced on March 5, 2009, the MPC also announced that it would undertake a program of asset purchases financed by the issuance of central bank reserves. The Sterling Monetary Framework was adjusted: among other changes, reserves targets were suspended and all reserves started being remunerated at the Bank Rate.[9]

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[9]For more details, see the consolidated notice at www.bankofengland.co.uk/markets/marketnotice090820smf-apf.pdf.

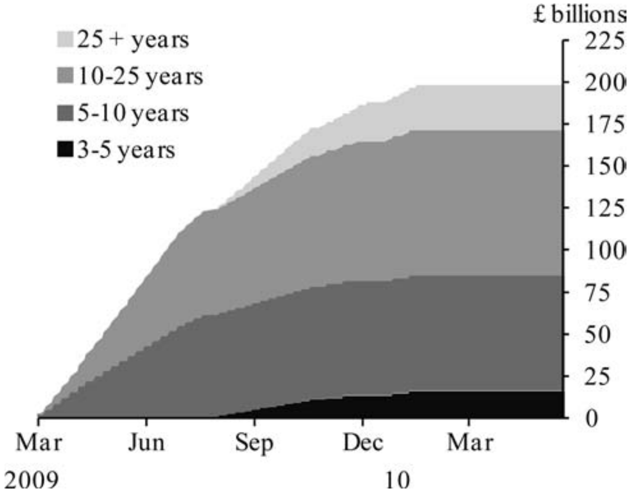
In order to meet the Committee's asset purchase objectives, the Bank announced that it would buy private and public assets, but that it was likely that the majority of overall purchases would be of gilts. The purchases of gilts were initially restricted to conventional gilts with a residual maturity between five and twenty-five years. Further extensions of the program were subsequently announced after the May, August, and November 2009 MPC meetings. After the August 2009 MPC meeting, the maturity range of gilt purchases was extended to three years and above. By February 2010, when the MPC announced that it would pause its program of purchases, the Bank had made £200 billion of asset purchases, of which £198 billion were gilts. Since January 2010, the Bank has been acting both as a buyer and a seller of corporate bonds, in order to improve liquidity in the market. From February 4, 2010, all purchases of corporate bonds and commercial paper have been financed by the issuance of Treasury bills and DMO cash management operations.

3.3 The Gilt Purchase Program

The Bank's gilt purchases were conducted through reverse auctions, whereby counterparties submitted prices at which they offered to sell specific quantities of individual gilts. These were held twice a week from March until August 2009 and three times a week after the August MPC meeting. The first gilt auction was conducted on March 14, 2009. At each auction the Bank accepted the cheapest offers (relative to market prices), up to the total amount to be purchased. The Bank bought widely across all maturities of available bonds (figure 1) but did not hold more than around 70 percent of the free float of any individual gilt. Although the counterparties in the auctions were banks and securities dealers, they could submit bids on behalf of their customers. And the auctions also allowed non-competitive bids to be made by other financial companies, whereby they agreed to sell gilts at the average successful price accepted in the competitive auction.

Since financial institutions may have bought up gilts in anticipation of selling them to the Bank, it is difficult to tell who the ultimate sellers were. But, as reported in Benford et al. (2009), the distribution of total gilt holdings at the end of 2008 suggests that

Figure 1. Cumulative Gilt Purchases by Maturity



Source: Bank of England.

banks held a comparatively small fraction of the total outstanding stock. Purchases of banks' gilts holdings will have shown up only in higher reserve balances at the Bank of England, and not in broad money aggregates (which includes deposits held by households and non-banks with commercial banks), unless the additional reserves led to increased bank lending or further purchases of assets from the non-bank private sector. But, other things equal, purchases from the non-bank private sector will have resulted in higher bank deposits and therefore will have been recorded as additional broad money. So to the extent that the purchases were ultimately from non-banks, we might have expected to see a large initial impact in the broad money data. (This motivates our approach in section 6, where we model the effect of QE as a swap between broad money and gilts.)

Table 1 sets out more details on the timetable of QE announcements. These are the events we will focus on in the next two sections, where we look at the reaction of financial markets to QE news. Although the first announcement of asset purchases was made in March, the publication of the February Inflation Report and the associated press conference on February 11 had given a strong indication that QE asset purchases were likely, which had an impact

Table 1. Key QE Announcement Dates

Announcement	Decision on QE	Other Information
February 11, 2009	The February Inflation Report and the associated press conference gave strong indication that QE asset purchases were likely.	
March 5, 2009	The MPC announced that it would purchase £75 billion of assets over three months financed by central bank reserves, with conventional bonds likely to constitute the majority of purchases. Gilt purchases were to be restricted to bonds with a residual maturity of between five and twenty-five years.	The Bank Rate was reduced from 1 percent to 0.5 percent.
May 7, 2009	The MPC announced that the amount of QE asset purchases would be extended by a further £50 billion to £125 billion.	
August 6, 2009	The MPC announced that the amount of QE asset purchases would be extended to £175 billion and that the buying range would be extended to gilts with a residual maturity greater than three years.	The Bank announced a gilt lending program, which allowed counterparties to borrow gilts from the APF's portfolio in return for a fee and alternative gilts as collateral.
November 5, 2009	The MPC announced that the amount of QE asset purchases would be extended to £200 billion.	
February 4, 2010	The MPC announced that the amount of QE asset purchases would be maintained at £200 billion.	The MPC's press statement said that the Committee would continue to monitor the appropriate scale of the asset purchase program and that further purchases would be made should the outlook warrant them.

on asset prices.[10] The next key dates were the further extensions of the program announced after the May, August, and November 2009 MPC meetings. At the August meeting, the Committee voted to raise the stock of assets purchased to £175 billion. Two additional decisions were also taken in August: the maturity range was increased from five to twenty-five years to three years and over, and some of the gilts purchased were made available for on-lending to the market through a gilt lending arrangement with the DMO.[11] The purchase program was further extended to £200 billion in November, maintaining the maturity range of three years and above. Finally, the decision in February 2010 to pause asset purchases, but to continue to monitor the appropriate scale of purchases, might have been expected to have an impact.

4. Gilt Market Reactions

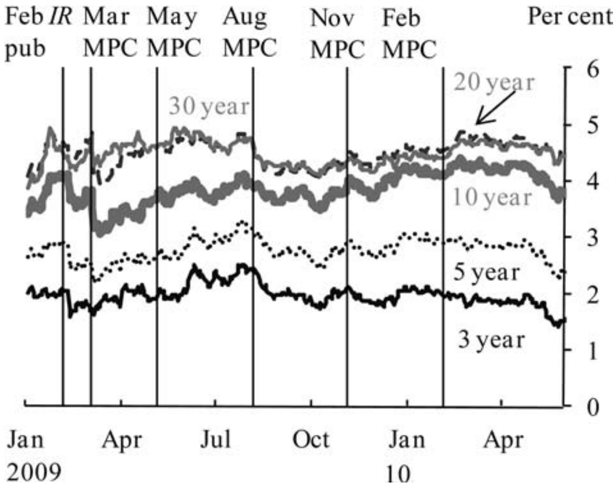
Since gilts made up the overwhelming majority of the Bank of England's asset purchases, it is natural to begin by first assessing the impact of QE on gilt yields.

Figures 2 and 3 show gilt yields and the spread between those yields and corresponding OIS rates at a number of maturities between January 2009 and May 2010. Both gilt yields and gilt-OIS spreads fell after the first announcements of QE in February and March 2009, consistent with a QE impact coming from both the macro/policy news and portfolio balance channels described in section 2. But comparing their levels at the end of May 2010 with where they were before the start of QE in February 2009 suggests little overall change at most maturities. However, net changes over the period are unlikely to provide a good measure of the overall

[10] Opening remarks at the press conference from the Bank of England Governor, Mervyn King, included the following statement: "The projections published by the Committee today imply that further easing in monetary policy may well be required. That is likely to include actions aimed at increasing the supply of money in order to stimulate nominal spending." (See www.bankofengland.co.uk/publications/inflationreport/irspnote110209.pdf). When answering questions from the press, he said that "we will be moving to a world in which we will be buying a range of assets, but certainly including gilts, in order to ensure that the supply of money will grow at an adequate rate to keep inflation at the target." (See www.bankofengland.co.uk/publications/inflationreport/conf090211.pdf).

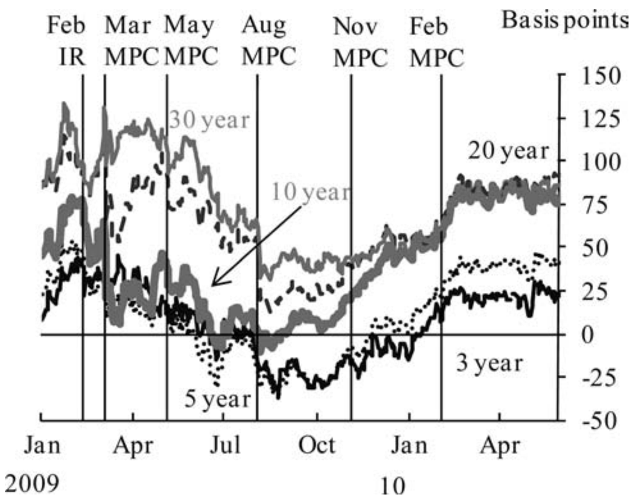
[11] See www.dmo.gov.uk/doc/gilts/press/sa060809b.pdf.

Figure 2. Gilt Yields[a]



[a]Estimated zero-coupon spot rates.
Sources: Bloomberg and Bank of England.

Figure 3. Gilt-OIS Spreads[a]



[a]Estimated zero-coupon gilt spot rates less corresponding zero-coupon OIS spot rates.
Sources: Bloomberg and Bank of England.

impact of QE on gilt yields, given the amount of other news there has been over the period, including on the likely scale of future gilt issuance by the UK government.

In the rest of this section we look at two different, but related, methods of quantifying the impact of QE on gilt yields. First, we use an event-study approach based on summing up the reactions of gilt yields and gilt-OIS spreads to announcements about QE. Second, we use a calibration based on scaling up reactions to the estimated news about total QE in those announcements, using the results of a survey of City economists conducted by Reuters.

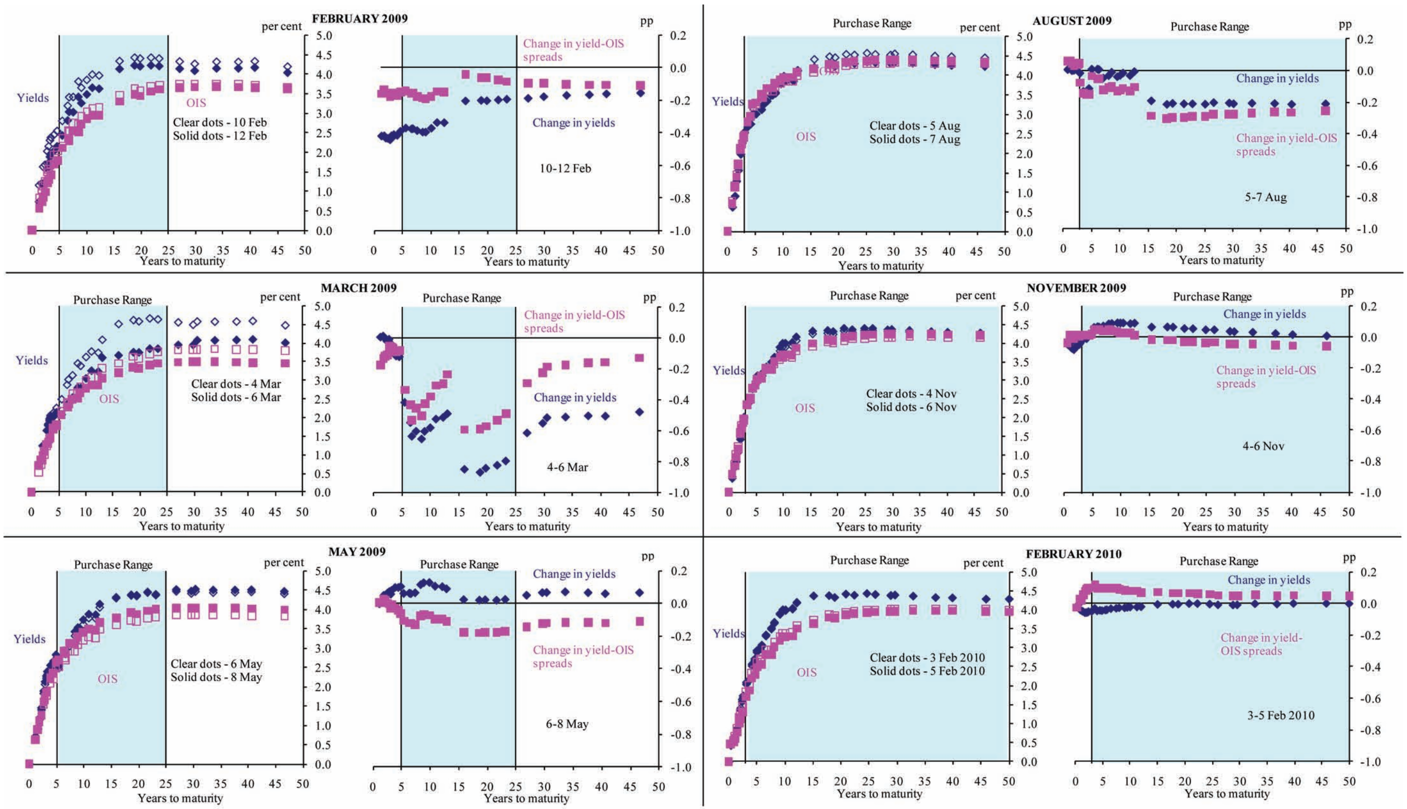
4.1 Event-Study Analysis

We might expect the majority of the impact of QE purchases on gilt yields to occur not when purchases are actually made but when expectations of those purchases are formed. One way, therefore, of quantifying the impact is to look at the immediate reaction of gilt yields and OIS rates to announcements relating to QE purchases (a similar approach is used in Bernanke, Reinhart, and Sack 2004 and Gagnon et al. 2011).

This event-study method involves focusing on the reaction of market prices over a fairly narrow interval after the QE-related news is released, with the aim of capturing the market's direct reaction to the news, abstracting from other factors that may also have been affecting asset prices. One judgment is how large to make the time interval (window) for comparison. Too short and we risk missing the full market reaction, as it may take time for the market to evaluate the news; too long and we risk the estimated reaction being contaminated by other news events. In what follows we use a two-day window, but for robustness we also examine the impact of using one- and three-day windows below. The relative novelty of QE in the United Kingdom, and the fact that market functioning may have been impaired, at least in early 2009, suggests that using a much shorter (intraday) window would not be appropriate.

Figure 4 shows the reaction of individual gilts to the six pieces of QE news discussed in section 3, as six pairs of figures. The left-hand figure in each pair shows yields-to-maturity at the end of the day before each announcement (clear diamonds) and on the day after the announcement (solid diamonds) corresponding to a two-day window. We also show equivalent OIS rates (clear and solid squares) for

Figure 4. Gilt Yield to Maturities and Corresponding Duration-Matched Zero-Coupon OIS Rates (Left Panel) and the Changes in Those Yields and the Yield-OIS Spread (Right Panel) Before and After Announcements Relating to QE Purchases



Sources: Bloomberg and Bank of England.

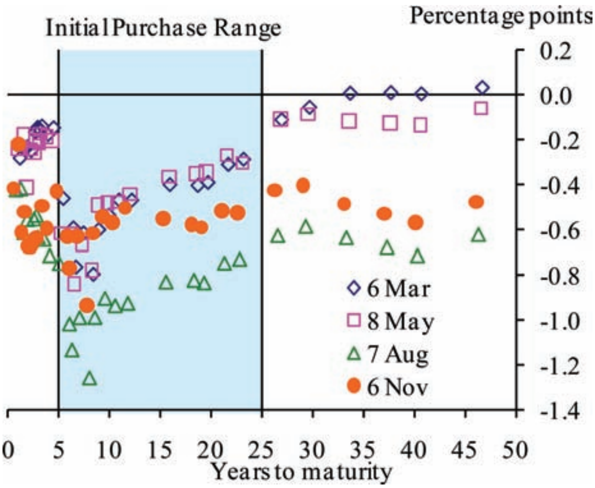
both days, where we have derived zero-coupon OIS rates from end-of-day prices to match the duration of each individual bond. The right-hand figure in each pair shows the corresponding change in gilt yields (diamonds) and the change in the spread between gilt yields and OIS rates (squares).

The largest two-day yield movements occurred following the publication of the Bank's Inflation Report and associated press conference in February 2009 and the announcement of the commencement of QE purchases after the March MPC meeting.

In February there was a reaction in both bond yields and gilt-OIS spreads, with yields on shorter-dated gilts falling by as much as 50 basis points (see figure 4, top-left panel). The reaction of yields on bonds with maturities above ten years was noticeably less. Intelligence gathered by the Bank of England from market participants suggested that some of this reflected perceptions that the Bank would target purchases on shorter-maturity bonds (see also Oakley 2009). The fact that both OIS rates and gilt-OIS spreads fell suggests that the news in the Inflation Report and the associated press conference comprised both macro/policy news and expected portfolio balance effects. Of course, not all of this macro/policy news reaction can be attributed to QE. Market intelligence and surveys suggest that the publication of the February Inflation Report was also associated with an increased expectation that the Bank Rate would be cut to 0.5 percent in March, though the impact of that on longer-term yields is likely to have been small.

When the MPC announced in March 2009 that the Bank would purchase up to £75 billion of gilts with residual maturities of between five and twenty-five years, there was a further significant reaction in yields and OIS rates (figure 4, middle-left panel). This effect was most pronounced in fifteen- to twenty-year maturities where yields fell by up to 80 basis points, perhaps reflecting a correction of previous expectations that purchases would be concentrated in gilts with shorter maturities. OIS rates also fell, though not as sharply, suggesting that the bulk of the fall reflected expected portfolio balance effects rather than changes in expected future short-term interest rates or the risks around those rates. Again the announcement accompanied other news, in that the Bank Rate was also reduced to 0.5 percent, but this change had been widely expected and any resulting reactions were likely to have been confined to the short end of the yield curve.

Figure 5. Cumulative Changes in Gilt-OIS Spreads Since February 10, 2009



Sources: Bloomberg and Bank of England.

The announcement in May 2009 of an extension of QE to £125 billion of purchases was widely anticipated and there was little reaction, with gilt yields and OIS rates actually rising by a small amount (figure 4, bottom-left panel). The August 2009 announcement of a further £50 billion extension was also largely expected, and the accompanying fall in yields of longer-maturity bonds seems more likely to have been caused by the extension of the purchase range to all bonds with a residual maturity of more than three years rather than news about the absolute size of purchases themselves (figure 4, top-right panel). Again the fact that this fall in yields was not reflected in OIS rates suggests that it was caused by a portfolio balance effect. The last two pieces of QE-related news appear to have had relatively little impact. The further extension of the program to £200 billion in November 2009 and the decision to pause purchases in February 2010 were both widely anticipated and so contained little news for prices (figure 4, middle and bottom-right panels).

The combined reaction to the February and March 2009 announcements was concentrated in those gilts within the five- to twenty-five-year purchase range. This changed the shape of the yield curve and introduced noticeable kinks around the five- and twenty-five-year points. Figure 5 shows the cumulative change in gilt-OIS spreads from before the February 2009 announcement to after the March, May, August, and November 2009 announcements. From this we can see that those differences in relative spreads were still present following the widening of the maturity range in August 2009. The fact that these differences were not arbitraged away by those who are broadly indifferent between gilts with similar maturities is indicative of increased segmentation in the gilt market and a lack of arbitrage activity in the first half of 2009. This suggests that, for those gilts in the initial purchase range, the downward pressure from QE purchases on their yields was greater than for other gilts. But figure 5 also shows that by November 2009 those differences had diminished. As described in section 3, the period between August and November saw the APF begin a scheme to lend out the gilts it had purchased via the DMO. The increased ability to borrow and short sell more easily those gilts held by the APF may have helped the arbitrage process, reducing segmentation in the gilt market. In so doing, the impact of QE on yields is likely to have been spread more evenly across gilts.[12]

In order to get an estimate of the effect of the QE announcements on gilt yields, we could simply sum over those reactions to QE news. But to get a more precise read of the overall impact on the term structure, we can examine the changes in the Bank of England's estimated zero-coupon yield curves, which strip out coupons from each gilt and allow us to construct continuous curves.[13] Using these yield curves, figure 6 shows a summary of how gilts reacted to each of the six announcements over a two-day window. It focuses for simplicity on the reaction averaged across five- to twenty-five-year spot rates, reflecting the maturity range of the initial purchases.[14]

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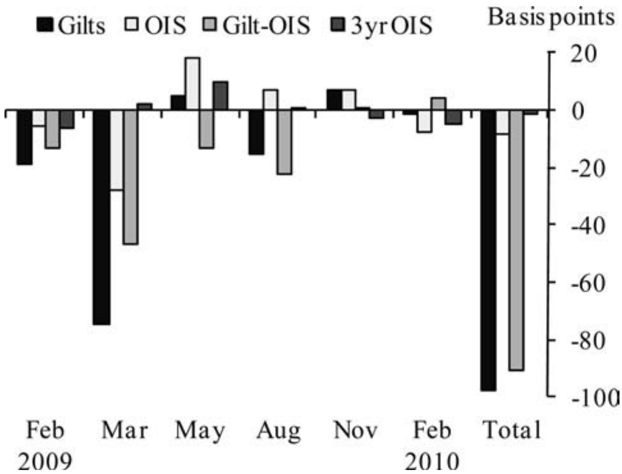
[12]Joyce et al. (2010) show that indicators of liquidity in the gilt market such as turnover and bid-ask spreads also improved over the period. This improvement in market liquidity may have been partly aided by APF gilt purchases and could also have contributed to the decrease in relative yield differences observed following August 2009.

[13]For data and more information, see www.bankofengland.co.uk/statistics/yieldcurve/index.htm.

[14]To the extent that the majority of the impact is likely to be concentrated at the duration of the gilt purchased, this could warrant focusing on the maturity range corresponding to the durations of the purchase range, or four to fifteen years. Here we attempt to capture the broader effects by using a five- to twenty-five-year range.

Figure 6. QE Announcement Impact on Gilt Yields, OIS, and Gilt-OIS Spreads: Average Change in 5- to 25-Year Spot Rates



Sources: Bloomberg and Bank of England.

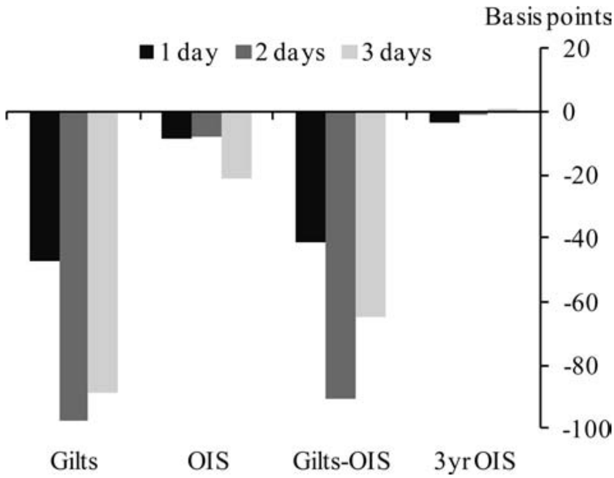
It also shows the reaction of gilt-OIS spreads and OIS rates for the same average maturities and the reaction of three-year OIS rates, in order to measure macro/policy news affecting just the short end of the yield curve. The publication of the Inflation Report in February 2009 appeared to have led markets to anticipate an additional 25-basis-point cut in the Bank Rate.[15] So to try and strip out that news from our measurement of the impact of QE, we make a simple adjustment to the reaction of gilts and OIS rates in February.[16]

Summing over the reactions in gilt yields to each of the QE news events gives an overall average fall of just under 100 basis points—with reactions ranging between 55 and 120 basis points across the five- to twenty-five-year segment of the yield curve (figure

[15] The mean expected level of the Bank Rate following the March MPC announcement, as measured by the Reuters poll of City economists, fell from 0.73 percent on February 5 to 0.53 percent on February 11.

[16] We subtract 25 basis points from instantaneous forward rates between zero and five years on a sliding scale (from 25 basis points at zero years to 0 basis points at five years) and then calculate the corresponding spot rates.

Figure 7. Total QE Announcement Impact and Sensitivity to Window Size



Sources: Bloomberg and Bank of England.

6).[17] Government bond yields in the United States, Germany, and France were largely unchanged over the same event windows, suggesting that these were UK-specific effects. The decomposition of the changes shows that the bulk of the effect came through changes in the gilt-OIS spread, which we expect to mainly reflect portfolio balance effects (as explained in section 2). The remaining change in OIS rates appears much smaller, at less than 10 basis points in total, and the overall reaction in shorter-maturity three-year OIS rates was close to zero. This suggests that the impact through the macro/policy news channel, as measured by changes in OIS rates, was much less important.

Figure 7 shows how sensitive these overall estimates are to changes in the size of the reaction window. Using a longer three-day window results in a similar overall impact, with a slightly smaller contribution from gilt-OIS spreads. Using a shorter one-day window

[17] On the basis of a very similar event-study approach, Meier (2009) suggested that the initial QE announcements reduced gilt yields by 35–60 basis points "at the very least" compared with where they would otherwise be. But his assessment only covered the period up to the middle of 2009.

reduces the overall impact to around 50 basis points, with the majority of the effect accounted for by movements in gilt-OIS spreads. So the overall impact varies between 50 basis points and 100 basis points according to the window size, but the conclusion that portfolio balance effects dominate remains robust to whatever window size is used.

4.2 News-Based Calibration

Figure 6 showed that the reactions in gilt yields were much larger for the February and March announcements than for later ones. One obvious explanation for these differences is that it reflects those first two events containing more news about QE for market participants.

An alternative way to estimate the impact on yields of QE purchases is to weight the announcement reactions by the amount of news each announcement contained. But in order to do so, it is necessary to calculate a measure of that news. Some partial information on market participants' expectations of QE is available from the Reuters poll of economists, which regularly surveys a panel of about fifty City economists on their future Bank Rate expectations. Between April 1, 2009 and February 25, 2010, Reuters also included a question in its poll on the total amount of QE purchases respondents expected. Bank of England market intelligence suggested that the responses to this survey provided a good proxy for market expectations of QE.

We can calculate a measure of the news in each announcement as the difference between the total QE purchase amount expected in the survey preceding the MPC's decision and the total QE amount expected in the survey released immediately after the MPC's decision. In the cases where there was no survey conducted immediately after the announcement, we use the difference between the amount announced and the previous survey expectation as our measure of news. There was no question on expectations of QE purchases in the Reuters surveys before April 2009, so any assumption about the news in the February and March 2009 announcements is necessarily arbitrary. But as most QE news appears to have occurred during this period, it is necessary to include it in our sample. Our baseline assumption is that the total amount of QE expected in the Reuters April 2009 survey represented genuine news, which was distributed equally between the February and March announcements. This is a conservative assumption as, to the extent that QE was anticipated before February and March, the amount of news will be overstated and hence the sensitivity of yields to that news understated. According to the Reuters survey, the February 2010 decision was broadly expected, as the mean of the Reuters survey was £204 billion before the announcement and £205 billion afterwards. For that reason, we do not include that announcement in the calibration.

To calibrate the impact of QE on the yield curve, we compare the two-day change in zero-coupon gilt and OIS rates across maturities of five to twenty-five years with our news measure for the QE events in February, March, May, August, and November 2009 and for the October 2009 Q3 GDP release.[18] Figure 8 shows there is a strong relationship between the size of the news and the average change in gilt yields across maturities after each event. A simple OLS regression of the two suggests a fall in gilt yields of around 0.6 basis points for each additional £1 billion of unanticipated QE purchases announced.[19]

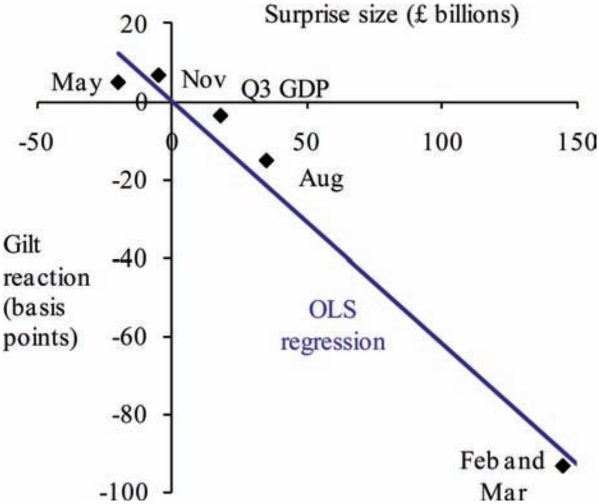
Scaling up the estimates from OLS regressions of QE news on gilt yields, OIS rates, and the gilt-OIS spread, figure 9 shows the total estimated impact of QE purchases averaged across maturities. The total impact on gilt yields from this news-based calibration is estimated to be around 125 basis points when a two-day window is used, with an impact on OIS rates (macro/policy news channel) of around 45 basis points and on gilt-OIS spreads (portfolio balance channel) of 80 basis points. This overall estimate is broadly similar to that estimated previously by summing up the announcement reactions, and the dominant effect is again estimated to come through the portfolio balance channel.

A sensitivity analysis of the results to the window length shows that, like before, the overall estimated impact is similar when we use two or three days, and smaller with a one-day window. The breakdown into changes in OIS rates and gilt-OIS spreads remains broadly

[18]The rise in expected purchases between the Reuters surveys on October 1 and October 28, 2009 appears to have been attributable to a lower-than-expected preliminary GDP release on October 23, which suggested more QE might be necessary.

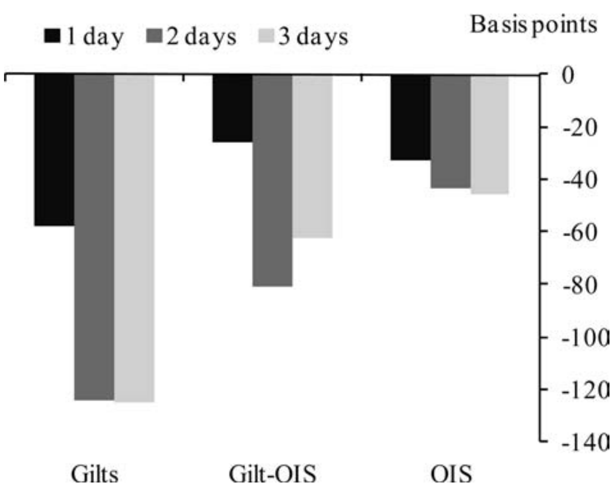
[19]The standard error for the coefficient is 0.04 and the R^2 is 0.98.

Figure 8. Size of Surprise and Average Gilt Movements



Sources: Thomson Reuters Datastream and Bank of England.

Figure 9. News-Based Calibration Impact and Sensitivity to Window Size



Sources: Thomson Reuters Datastream, Bloomberg, and Bank of England.

unchanged when we estimate the simple OLS regression using a two- or a three-day window. Using a one-day window, by contrast, results in a relatively larger impact on OIS rates than on gilt-OIS spreads.

5. The Reaction of Other Assets[20]

To the extent that investors do not regard money as a perfect substitute for gilts, we would expect them to reduce their money holdings associated with QE purchases by buying other sterling assets, such as corporate bonds and equities, and foreign assets. This will likely put upward pressure on the prices of those assets, and perhaps downward pressure on the sterling exchange rate. In addition, announcements about QE may contain information about the economy that has implications for perceptions of future corporate earnings and the uncertainty around them; and changes in the prices of gilts may affect the rate at which investors discount future cash flows. Both of these effects will also have an impact on asset prices. But all of these effects might be expected to take time to feed through, as it will take time for investors and asset managers to rebalance their portfolios, and asset prices are unlikely to anticipate fully this process, given the novelty of QE and uncertainty about the transmission mechanism.

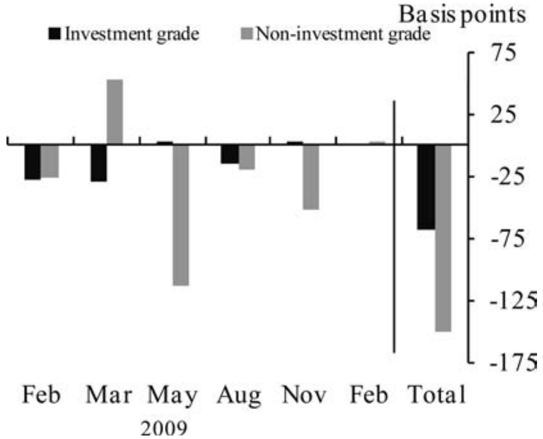
This section focuses on assessing the observed impact of QE on the two largest sterling asset classes in addition to gilts—corporate bonds and equities—and the impact on the exchange rate. Figures 10 and 11 summarize the immediate price reaction (over two days) following each of the six QE news announcements discussed earlier. These suggest that equity and corporate bond prices reacted in a less uniform way than gilts after the announcements. The rest of this section discusses each asset class in more detail.

5.1 Corporate Bonds

Lower gilt yields should lead to lower corporate bond yields for a given corporate bond spread (compensating for the risks of holding sterling corporate bonds relative to gilts). But, in addition, as

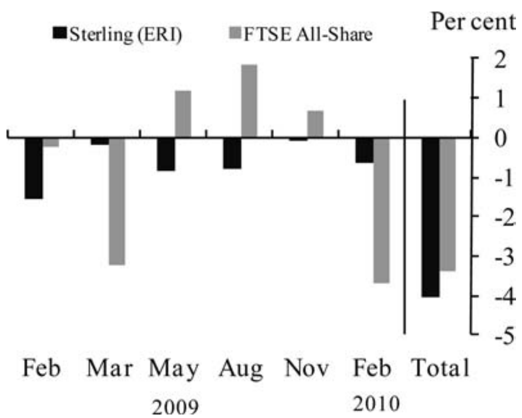
[20]A more detailed discussion of the reaction of other asset prices can be found in Joyce et al. (2010).

Figure 10. QE Impact on Corporate Bond Yields



Source: Merrill Lynch.

Figure 11. QE Impact on Sterling and FTSE All-Share



Source: Bloomberg.

investors attempt to rebalance their portfolios away from gilts and into corporate bonds, the component of that spread representing compensation for risk aversion and uncertainty (the so-called debt risk premium) should fall, reducing yields further, though the timing of this could depend on how long investors take to make portfolio decisions. But the announcement of QE may also give investors information about the outlook for the economy. This, if worse than expected, could affect the perceived risk of corporate default, putting upward pressure on yields. Over time, however, a successful QE policy would be expected to lead to lower corporate bond yields.

Summing over the immediate reaction to the six QE news announcements, sterling investment-grade corporate bond yields fell by 70 basis points, with spreads remaining broadly flat (figure 10). Sterling non-investment-grade corporate bond yields fell by 150 basis points, with spreads narrowing by 75 basis points.[21] The narrowing in non-investment-grade spreads is consistent with QE removing some of the perceived downside tail risks. Over the same announcement windows, U.S. dollar and euro-denominated investment-grade bond yields fell by 23 basis points and 11 basis points, respectively, around 50 basis points less than sterling-denominated bonds, suggesting that there was a UK-specific effect.

5.2 Equities

Lower gilt yields should, all else equal, increase the present value of future dividends, thus raising equity prices. In addition, as investors attempt to rebalance their portfolios away from gilts towards more risky assets, the additional compensation investors demand for the risk of holding equities (the so-called equity risk premium) should fall. This will put further upward pressure on equity prices. Again, the announcement of QE may also give investors information about the outlook for the economy. If worse than expected, this could lower their immediate expectations for future dividends and raise risk premia, thus putting downward pressure on equity prices in the short term. So, as for corporate bonds, it is therefore not clear what we would expect the immediate QE impact to be, although a successful QE policy would eventually be expected to lead to higher equity prices.

Equity prices did not react in a uniform way in response to QE news (figure 11). The FTSE All-Share Index fell slightly (−0.2 percent) following the publication of the February Inflation Report and more sharply (−3.2 percent) following the March MPC announcement. However, over the same period, international equity prices fell by even more, suggesting that there might have been a small positive

[21] These numbers imply gilt yields fell by 75 basis points. This is different from the estimate in section 4 because the average duration of corporate bonds is shorter than that for gilts.

UK-specific effect. UK equity prices increased somewhat following the next three QE announcements but fell sharply in February 2010, though this is unlikely to have been a QE effect, as the February decision was widely expected.

5.3 Sterling

Lower gilt yields should, all else equal, lead to a depreciation of sterling. A standard uncovered interest parity (UIP) decomposition would predict an 8 percent depreciation given the observed fall in ten-year spot gilt yields over the QE news events.[22] Summing over the immediate reactions to the six QE news announcements, the sterling exchange rate index (sterling ERI) depreciated by 4.0 percent overall (figure 11)—although the largest fall occurred after the publication of the February Inflation Report, which may not solely reflect QE news. If we instead perform a UIP decomposition using three-year OIS rates, in order to isolate the macro/policy news component, the implied fall in the exchange rate would be only 0.5 percent, which would imply that the initial reaction of sterling was slightly greater than expected.

5.4 Summary

Table 2 summarizes the movements in asset prices and yields around the main QE announcements and over a longer period up to mid-2010. Medium to long-term gilt yields appear to be 100–125 basis points lower than in the absence of QE, with most of the effect coming through the portfolio balance channel. Corporate bond yields also fell markedly around announcements, and there were modest falls in sterling. For equities, the impact of QE is harder to pinpoint, though equity prices rose strongly through 2009.

In addition to those immediate reactions, the impact on other asset prices through the portfolio balance channel may come through over a more prolonged period, as investors make decisions about how to rebalance their portfolios. Table 2 shows that between March 2009 and May 2010 sterling investment-grade bond spreads narrowed by 380 basis points and the FTSE All-Share Index rose by

[22] For an explanation of UIP see Brigden, Martin, and Salmon (1997).

Table 2. Summary of Movements for Different Assets

Asset	**Change Around Announcements**	**Change March 4, 2009–May 31, 2010**	**Comments**
Gilts	−100 bp (of which −90 in gilt-OIS spreads)	+30 bp (of which +15 in gilt-OIS spreads)	The portfolio balance channel dominates the macro/policy news channel.
Gilts (Surprise Calibration)	−125 bp (of which −80 in gilt-OIS spreads)	+30 bp (of which +15 in gilt-OIS spreads)	The portfolio balance channel also dominates when allowing for surprise component of announcements.
Corporate Bonds (Investment Grade)	−70 bp	−400 bp	Smaller fall than in gilts around announcements due to shorter average maturity; spreads flat around announcements but significantly down over the period.
Corporate Yields (High Yield)	−150 bp	−2,000 bp	Larger announcement effects, possibly reflecting the removal of tail risk.
FTSE All-Share	−3 percent	+50 percent	No announcement effects, but prices up during the period.
Sterling ERI	−4 percent	+1 percent	Hard to single out QE effect.

around 50 percent. All else being equal, higher equity and corporate bond prices are likely to encourage firms to raise finance through relatively higher capital market issuance, either in addition to or as a substitute for alternative means of raising funds. Net equity issuance by UK private non-financial corporations (PNFCs) was particularly strong in 2009, reversing the negative net issuance observed over 2003–08. Net corporate bond issuance by UK private non-financial corporations in 2009 was also stronger than over the 2003–08 period. It is not possible to know what would have happened in the absence of QE, but Bank of England market intelligence suggested there was strong institutional investor demand for corporate bonds during the second half of 2009 (see Bank of England 2009).

6. Portfolio Model Estimates

Our analysis of the reaction of asset prices to the MPC's QE announcements suggests that a large part of the effect came through a portfolio balance channel. But we have also noted that it is difficult to quantify the specific impact of QE, given the potential role of other policies and international factors. As an alternative approach, in this section we estimate two different portfolio balance models in order to quantify the possible effects of the MPC's asset purchases on asset prices.

6.1 The Portfolio Balance Model

A natural starting point for modeling the portfolio channel is the basic portfolio choice model arising from the "mean-variance" approach to portfolio allocation developed by Tobin and Markovitz in the 1950s (e.g., Tobin 1958) and set out in a number of papers, including Roley (1979, 1982), Walsh (1982), and Frankel (1985). In this model, expected returns on each asset are exogenous, from the perspective of each individual investor. An individual investor's problem is to choose the weight to allocate to each asset in his or her portfolio, in order to maximize expected utility from end-of-period wealth, subject to a wealth constraint. In aggregate, however, investors' total asset holdings are constrained to match the available (exogenous) asset supplies of each asset. In the case where investors' total desired asset holdings do not match the available asset supplies, investors will require additional returns on each asset to willingly hold the "excess" asset stocks, and vice versa. This provides a lever for a policy of asset purchases to affect asset prices by changing asset quantities (specifically, reducing the quantity of gilts) and thereby the excess returns (risk premia) investors require ex ante to hold the available stock of assets (in the case of QE purchases, reducing the required returns on gilts and assets that are substitutable for gilts).

The first-order conditions of the investor's maximization problem in the basic model generate a relationship between investors' asset demands, excess returns of each asset, and their covariances. By equating asset demands with exogenous asset supplies, it is then possible to derive the following equilibrium condition:

$$E_t(r_{t+1}) = \lambda \Omega \alpha_t, \tag{6}$$

where r_{t+1} is a vector of expected excess asset returns (where one of the assets performs the role of the numeraire asset), λ is the coefficient of constant relative risk aversion (CRRA), Ω is the covariance matrix of asset returns, and α_t is a vector of asset shares of the total portfolio. Equation (6) shows that expected returns on each asset in excess of the return on a benchmark asset are a function of risk aversion, the share of each asset in total wealth, and the asset return covariances.

In this simple model, given a set of asset shares, the expected excess returns are completely determined by the variance-covariance matrix of asset returns and the covariances capture relative substitutability between different assets. The model implies that the impact of a change in the relative stocks of assets—brought about by a swap of money for gilts, for example—is given by the covariance between asset returns together with the CRRA coefficient. This suggests that one might calibrate the impact of the Bank's asset purchases by estimating the return covariances and assuming a value for the coefficient of relative risk aversion. We follow this approach below.

It needs to be recognized, of course, that the model adopts a number of simplifying assumptions. There are a range of other important influences on asset returns, in addition to asset supplies, that are not captured by this model (e.g., the business cycle). Furthermore, the model is partial equilibrium in nature. Nevertheless, it seems surprisingly robust to various extensions (see Campbell 1999).

How do we implement this basic model empirically? We do not observe ex ante returns, so we shall assume in what follows that investors have rational expectations, so that the difference between ex post excess returns and ex ante excess returns is measured by a random error, orthogonal to the portfolio shares:[23]

$$r_{t+1} - E_t r_{t+1} = \varepsilon_{t+1}, \quad E_t(\varepsilon_{t+1}) = 0, \quad E_t(\varepsilon_{t+1}|\alpha_t) = 0.$$

Adding a constant term, we can therefore write the basic empirical model as (see, e.g., Engel et al. 1995 or Hess 1999 for a derivation):

$$r_{t+1} = A + \lambda\Omega_t\alpha_t + \varepsilon_{t+1}, \quad \Omega_t = E_t\varepsilon_{t+1}\varepsilon'_{t+1}. \tag{7}$$

We shall look at two different models: a basic vector autoregressive (VAR) model informed by the theory, but where we allow the data to speak, and a more sophisticated multivariate generalized autoregressive conditional heteroskedasticity (GARCH)-in-mean model (henceforth GARCH-M model), where we impose more structure using the theoretical restrictions implied by the basic theory.

6.2 A VAR Application

Our first approach is largely data driven. We estimate a VAR which includes both excess returns and asset shares and also allows for the influence of a set of exogenous variables, intended to capture other influences on asset demand and supply. The virtue of this approach is that it allows asset supplies to be treated as endogenous and to respond to movements in excess returns.[24]

Our VAR takes the following general form:

$$Y_t = \alpha_+ \sum_{i=1}^{p} \beta_i Y_{t-i} + \sum_{j=0}^{k} \gamma_j X_{t-j} + \varepsilon_t, \tag{8}$$

where Y_t is the vector of endogenous variables, which consists of both monthly excess returns and shares of total wealth held in these assets, and X_t is a vector of exogenous variables. In this model the

[23] If there are other information variables, then the errors would be orthogonal to the overall information set, which would include the portfolio shares.

[24] However, it is a reduced-form model and therefore subject to the usual caveats.

Table 3. Monthly Asset Returns and Asset Shares: Summary Statistics

	Mean	Standard Deviation	Min.	Max.
Excess Return on Equities (pp)	0.570	3.896	−12.743	10.053
Excess Return on Corporate Bonds (pp)	0.408	1.450	−3.759	4.648
Excess Return on Gilts (pp)	0.339	1.478	−4.127	4.933
Return on M4 (pp)	0.323	0.104	0.199	0.764
Equity Share	0.500	0.0442	0.411	0.600
Corporate Bond Share	0.0648	0.0290	0.020	0.109
Gilt Share	0.0927	0.0152	0.070	0.120
M4 Share	0.343	0.040	0.269	0.453

Notes: Sample is December 1990 to June 2007. Excess returns are calculated relative to the return on M4.

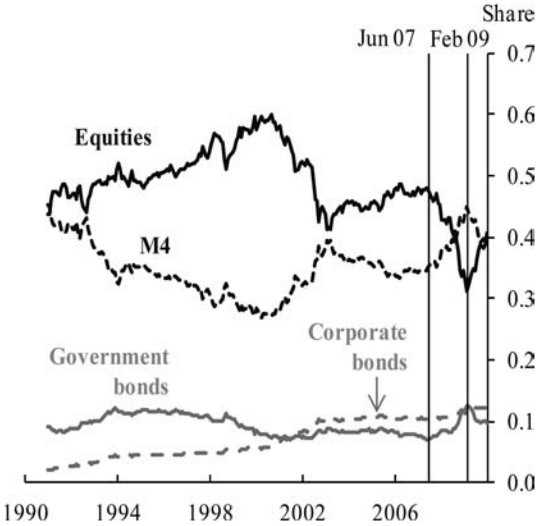
return covariances are implicit in the model estimates, rather than being explicitly modeled.

In our baseline model, we included monthly returns on gilts, sterling investment-grade corporate bonds, UK equities, and M4, with the latter defined as the numeraire asset. Details of the construction of the asset price and asset stock data are contained in the data appendix. For our exogenous variables we included variables attempting to pick up the state of the economic cycle: the growth rate of industrial production, (seasonally adjusted) RPI inflation, and the slope of the yield curve.[25]

Summary statistics for the asset price return and share data for the period December 1990 to June 2007 are shown in table 3; the asset shares are also plotted in figure 12. As we would expect, riskier assets tend to earn higher returns on average, so the average monthly return on equity is nearly three times as large as the return on holding M4. The volatility of corporate bond returns is slightly lower

[25]An extended version of the model including index-linked bonds produced similar results.

Figure 12. Asset Shares



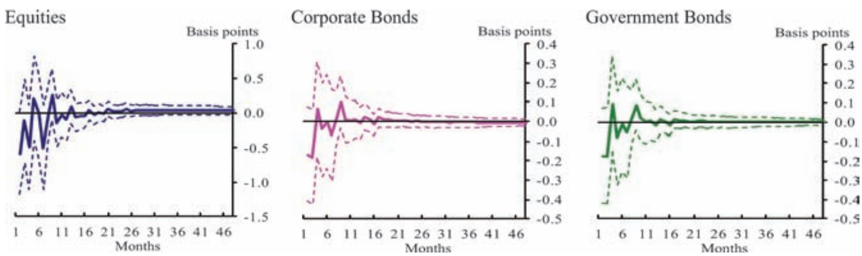
Sources: Thomson Reuters Datastream, Barclays Capital, and Bank calculations.

than gilt returns, at least for our sample, though the average return is slightly higher. One striking feature of the asset share data is the strong inverse relationship between the M4 share and the equity share (figure 12).

We estimated the model by OLS using monthly data on a sample from December 1991 to the middle of 2007, so before the onset of the current global financial crisis. We used seven lags of each endogenous variable, in line with the results from the normal Akaike and Schwarz lag selection criteria, and checked that post-estimation diagnostics including stability tests were satisfactory.[26] We then used the model to produce impulse responses, which allow us to summarize how excess asset returns and asset supplies are predicted to respond to a shock to the share of gilts in the aggregate portfolio. When conducting impulse response analysis, an important concern is the method used to identify shocks corresponding to each of the endogenous

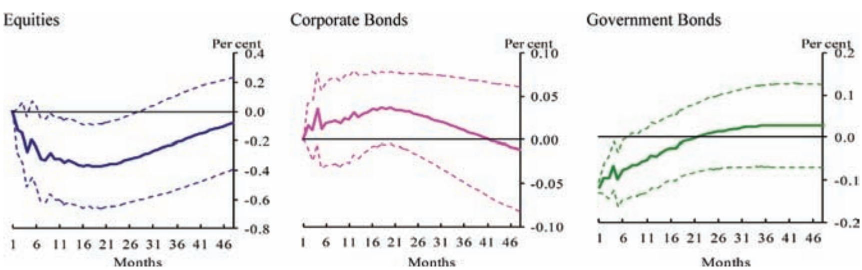
[26]The VAR was found to be stable with no roots outside the unit circle. Full estimation results are available upon request.

Figure 13. Impulse Responses of Excess Returns (One-Standard-Deviation Fall in Gilt Share)



Note: Excess returns are calculated relative to the return on M4.

Figure 14. Impulse Responses of Asset Shares (One-Standard-Deviation Fall in Gilt Share)



variables in the VAR. For our analysis, innovations to the gilt share are interpreted as the QE shock and this is identified in a standard recursive manner, by ordering the gilt share last in the VAR. We apply a Cholesky decomposition to compute the impulse responses.

Figures 13 and 14 show the impulse response functions for a one-standard-deviation fall in the share of gilts (offset by an increase in the share of M4). As the theory would suggest, the expected excess returns on gilts, corporate bonds, and equities all fall in response (although these responses are within the 95 percent confidence interval). This would be consistent with a rise in asset prices, as investors try to reallocate their portfolios away from gilts. The response of quantities to this shock is puzzling, however. While the corporate bond share increases slightly and the share of gilts falls, as might be expected, the share of equities also falls. This result is difficult to reconcile with the portfolio balance approach but might reflect the

Table 4. Estimated Impact of QE on Annualized Excess Returns (Basis Points)

	VAR Model		Multivariate GARCH-in-Mean (CCRA = 3)
	Immediate Impact	**Average Over Six Months**	**Effect**
Excess Returns on Gilts	−85	−32	−70
Excess Returns on Corporate Bonds	−81	−32	−66
Excess Returns on Equities	−282	−121	−34

Note: Excess returns are calculated relative to the return on M4.

fact that over our sample the share of M4 in wealth moved inversely with the share of equity.

The impulse responses are based on a one-standard-deviation shock, which translates roughly into a reduction of £5 billion of gilts using the gilt share sample average. In order to scale up these numbers to simulate the MPC's asset purchases, we assume for simplicity that all the purchases were from non-bank domestic investors (so that all the gilt purchases would have led to additional broad money holdings, at least initially) and were implemented at the start of the period.[27] The assumption that all the purchases come from the domestic non-bank private sector means that our estimates are likely to overestimate the effects, if anything.

To make the results more comparable with the changes in (annualized) yields shown earlier, table 4 shows the model-implied impact of QE in terms of annualized excess returns. Given uncertainty over the VAR dynamics, it is difficult to know which horizon to focus on. The second and third columns of the table therefore provide two measures of the implied impact on annualized monthly excess

[27] Actual QE announcements and purchases were staggered over a longer period, so we place less emphasis on the precise dynamics of the impulse responses.

returns: in the first period after the shock and on average over the first six months after the shock. The range of estimates for both excess gilt returns and excess corporate bond returns is broadly similar to the immediate market reactions discussed in sections 4 and 5. The range of estimates for excess equity returns is clearly much greater and is also more difficult to compare directly with the earlier analysis. Using a dividend discount model (as in Inkinen, Stringa, and Voutsinou 2010) to map the range of estimates into prices, however, implies a rise of between 20 percent and 70 percent. The upper estimate is clearly implausible. The main conclusion we draw is that the suggested impact on equity prices is potentially large but highly uncertain.

6.3 A Multivariate GARCH-in-Mean Model

One important caveat regarding our unrestricted VAR model is that it implicitly assumes that the covariance matrix between asset returns is constant. That is at odds with the empirical literature, which suggests that covariances can vary substantially over time and in particular at times of financial stress. So the model does not take into account the fact that the degree of substitutability of the different assets will have changed in response to evolving market conditions.

To allow explicitly for the possibility that the covariance matrix of asset returns may be changing over time, we also estimated the portfolio balance model in (7) using a multivariate GARCH-M framework (see Engel et al. 1995). This approach allows us to estimate a time-varying covariance structure but treats asset shares as exogenous. The estimated model takes the following form for an n-asset portfolio:

$$r_{t+1} = A + \lambda \Omega_t \alpha_t + \varepsilon_{t+1} \tag{9}$$

$$\Omega_t = C^{*'} C^* + A^{*'} \varepsilon_t \varepsilon_t' A^* + B^{*'} \Omega_{t-1} B^*. \tag{10}$$

The covariance structure given in (10) is the first-order BEKK model of Engle and Kroner (1995), where C^*, A^*, and B^* are $(N \times N)$ coefficient matrices with C^* upper triangular. The quadratic structure of the BEKK model ensures that the covariance matrix is positive

Table 5. Estimation Results for Multivariate GARCH-M Model, CRRA = 3

Mean Equation Estimates of the Constant Vector—*A*			
	Coefficient	**Robust Standard Error**	**Significance**
A (1)	0.00739	0.00277	0.00766
A (2)	0.00343	0.00089	0.00012
A (3)	0.00281	0.00092	0.00218
MVGARCH Equation Estimates of the Upper Triangular Matrix—C^*			
	Coefficient	**Robust Standard Error**	**Significance**
C^* (1,1)	−0.0037	0.00315	0.23923
C^* (2,1)	−0.00413	0.00105	0.00007
C^* (2,2)	−0.00005	0.00073	0.94001
C^* (3,1)	−0.00474	0.00117	0.00005
C^* (3,2)	−0.00006	0.00084	0.94155
C^* (3,3)	0.00000	0.00003	0.99946

(continued)

definite. The model is estimated by maximum likelihood assuming conditional normal errors.

We first estimated the model over the same pre-crisis sample period as the VAR model, in order to infer what the model would imply for the impact of a purchase of £200 billion of gilts. When the model was freely estimated, the CRRA parameter was negative, so following Hess (1999) we restricted this coefficient to three. The reported model fits the data reasonably well and there was no residual serial correlation.[28] Table 5 contains the estimation results. It needs to be borne in mind that a larger risk parameter would generate larger changes in expected returns.

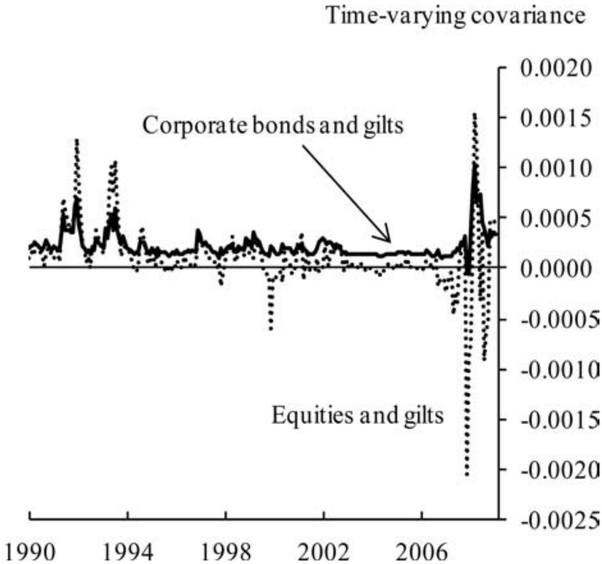
[28]In addition, we estimated a constant variance version of the model, by constraining the A^* and B^* matrices in (10) to be zero, but the null hypothesis that these parameters were zero was rejected at the 1 percent significance level.

Table 5. (Continued)

MVGARCH Equation							
Estimates of A^* Coefficient Matrix				**Estimates of B^* Coefficient Matrix**			
	Coefficient	**Robust Standard Error**	**Significance**		**Coefficient**	**Robust Standard Error**	**Significance**
A^* (1,1)	−0.43875	0.08090	0.00000	B^* (1,1)	−0.87862	0.04395	0.00000
A^* (2,1)	4.75812	1.22264	0.00010	B^* (2,1)	0.57410	0.40295	0.15423
A^* (3,1)	−4.72806	1.00346	0.00000	B^* (3,1)	−0.44677	0.48136	0.35333
A^* (1,2)	−0.01534	0.03121	0.62319	B^* (1,2)	−0.01359	0.01812	0.45320
A^* (2,2)	−0.40588	0.24215	0.09372	B^* (2,2)	−1.16749	0.05379	0.00000
A^* (3,2)	0.11543	0.24374	0.63579	B^* (3,2)	0.27857	0.06350	0.00001
A^* (1,3)	−0.01468	0.03706	0.69189	B^* (1,3)	−0.01263	0.02148	0.55659
A^* (2,3)	−0.84641	0.29500	0.00412	B^* (2,3)	−0.26368	0.08231	0.00136
A^* (3,3)	0.51189	0.30562	0.09395	B^* (3,3)	−0.61339	0.09044	0.00000
Log Likelihood	1821.48630						
Multivariate Q-stats							
Residuals from MVGARCH-in-Mean Model				**Standardized Residuals from MVGARCH-in-Mean Model**			
	Value	**Significance**	**χ^2 d.f.**		**Value**	**Significance**	**χ^2 d.f.**
Q(15)	139.50447	0.37767	135	Q(15)	142.43259	0.31395	135
Q(10)	106.71997	0.11020	90	Q(10)	88.73314	0.51799	90
Q(5)	60.46782	0.06147	45	Q(5)	53.36486	0.18362	45

Notes: Model was estimated using monthly data from 1990:01 to 2007:06. “Multivariate Q-stats” is the Hosking (1981) variant on the multivariate Q-stats.

Figure 15. Covariances between Equity and Gilt Excess Returns and between Corporate Bond and Gilt Excess Returns



Note: Excess returns are calculated relative to the return on M4.

To simulate the impact of QE, we make the same assumptions as before. We assume that all the gilt purchases were implemented at the start of the period and led one-for-one to additional broad money holdings, at least initially. The implications for annualized excess returns are shown in the final column of table 4, derived using the derivative of the asset demand relationship (using the average values of the estimated asset return covariances over the sample). These numbers are in the range implied by the VAR for gilts and corporate bonds but rather lower for equity returns. The fact that gilt and corporate yields move by similar amounts suggests that they are closer substitutes, which seems quite plausible.

We might expect that QE itself will have changed the covariance structure of returns. To try to examine this, we can reestimate the multivariate GARCH-M model over a longer sample up to the end of 2009. Figure 15 shows the estimated time-varying covariances between gilts and equities and gilts and corporate bonds from the model. The intensification of the financial crisis in late 2008 is clear from the large movements in both covariances over the same period. During 2009 there seems to be some reversion to normal conditions, though it is not possible to ascribe this directly to QE, given other developments over the same period.

In summary, these results lend some empirical support to the notion of imperfect asset substitutability in the portfolio choice models of Tobin and others. Given the considerable uncertainties involved, our empirical estimates seem reassuringly in line with the analysis reported in sections 4 and 5. Our estimates would suggest an effect on annualized excess gilt returns of 30 to 85 basis points, which is broadly similar to our estimates for the portfolio balance impact from our analysis of the announcement reactions. The major uncertainties concern the estimated impact on equities, where different approaches produce quite different estimates of the likely effect of QE on excess returns and the VAR-based analysis would imply a falling portfolio share.

7. Conclusions

As part of its response to the global financial crisis and a sharp downturn in economic prospects, the Bank of England's MPC began a program of quantitative easing in March 2009. Over a year, the Bank bought £200 billion of assets, most of them government securities. This paper attempts to evaluate the impact of these large purchases on financial markets.

Based on market reactions to news about QE purchases, we found that medium to long-term gilt yields were about 100 basis points lower than they would otherwise have been as a result of QE, which our estimates suggest mainly came through a portfolio balance effect. Separate econometric analysis suggests that these effects are broadly in the range that might have been expected. Analysis of announcement reactions is unlikely to capture the full effects of portfolio rebalancing on other assets, so it is difficult to disentangle the specific impact of QE purchases from other factors. But most other asset prices showed a marked recovery through 2009, suggesting that QE is likely to have had wider effects. Our econometric estimates suggested considerable uncertainty about the size of the impact, particularly regarding the impact on equity returns. Moreover, VAR-based analysis on its own would not have predicted the large pickup in equity issuance that occurred in 2009.

How do our findings compare with similar analysis of the Federal Reserve's asset purchases in the United States? Gagnon et al. (2011) estimate that the overall reduction in the ten-year term premium on U.S. Treasuries in response to the Federal Reserve's purchase program was "somewhere between 30 and 100 basis points." But in addition to this effect, they find an even more powerful effect on the yields on agency debt and agency mortgage-backed securities. Given the large range of uncertainty around these kinds of estimates, the effects of the Federal Reserve's purchases can be described as being of a similar order of magnitude to the Bank's for the United Kingdom.[29]

The effectiveness of QE asset purchases will ultimately be judged by their impact on the wider macroeconomy. Our analysis suggests that the purchases have had a significant impact on financial markets and particularly gilt yields, but there is clearly more to learn about the transmission of those effects to the wider economy.

Appendix. Data on Asset Returns and Stocks in Section 6

Our data consist of end-of-month realized returns and asset shares of four different assets: equities, corporate bonds, nominal gilts, and broad money from December 1990 to December 2009.

For *equities*, we use the total return index and market capitalization of the FTSE All-Share Index provided by Thomson Reuters Datastream. The return index includes an aggregate dividend as an incremental amount to the daily change in prices. For *corporate bonds*, we use the total return and market value of the Barclays Capital index corresponding to investment-grade corporate bonds of all maturities. The total index return includes coupon payments and paydowns in addition to changes in price. For *gilts*, we also use the

[29]The Federal Reserve's purchases of Treasuries, analyzed in Gagnon et al. (2011), were of a similar absolute size to those of the United Kingdom ($300 billion), albeit smaller compared with the overall size of the Treasuries market. Including the purchases of agency debt and agency mortgage-backed securities, however, gives a broadly similar figure as a percentage of GDP.

market value and returns from the Barclays Capital nominal gilts index, but we subtract holdings by the official sector using DMO data.[30]

We use an adjusted measure of M4 to capture the share of *broad money*[31] not held by financial institutions. M4 comprises the private sector's (i.e., the UK private sector other than monetary financial institutions (MFIs)) holdings of notes and coin, deposits, and other short-term instruments. Our adjusted M4 is constructed as M4 minus the sterling deposits of non-bank credit grantors, mortgage and housing credit corporations, bank holding companies, and other activities auxiliary to financial intermediation (intermediary offshore financial centers, or OFCs).[32] In addition, sterling deposits arising from transactions between banks or building societies and "other financial intermediaries" belonging to the same financial group are excluded from this measure of broad money.[33] Ideally, we would like to be able to exclude from our sample the equities, corporate bonds, and gilts held by MFIs and intermediary OFCs. This is not possible due to lack of available data.[34]

For the return on broad money, we construct an effective rate of return using rates and amounts from the Divisia money tables.[35] We first calculate separate retail and wholesale deposit rates from

[30]We only have data on official holdings since 2000. Since the proportion of gilts held by the official sector was small and relatively stable until 2008, we have deducted the percentage of average official holdings for 2000–01 from the pre-2000 figures.

[31]Detailed definitions of M4 and broad money are available at www.bankofengland.co.uk/mfsd/iadb/notesiadb/M4.htm. For a discussion of the economic meaning of M4, see Berry et al. (2007).

[32]A description of adjusted M4 can be found on www.bankofengland.co.uk/mfsd/iadb/notesiadb/m4adjusted.htm.

[33]Adjusted M4 is only available quarterly. We interpolate the adjustment linearly and deduct it from the monthly M4 data. Moreover, there are no adjusted data before December 1997. Given that the adjustment was stable at 10 percent of unadjusted M4 between 1998 and 2002, we deduct 10 percent from M4 for the pre-1997 period.

[34]There are data available on MFIs' holdings of some assets, but it is not possible to get their holdings of sterling investment-grade corporate bonds. No data on asset holdings by the other institutions excluded from the adjusted measure of M4 are available.

[35]This information is available from the interactive statistics database of the Bank of England.

several different deposit types in the Divisia tables.[36] We then weight those rates together into an overall deposit rate using the weightings of retail deposits, wholesale deposits, and notes and coin in M4. Notes and coin holdings yield no return.

The retail rates and weights are calculated by assuming that all deposits held by households, and non-financial corporates' sight deposits, are retail deposits. In turn, non-financial corporates' time deposits and all deposits by OFCs are considered wholesale. The weights obtained in this manner follow very closely the amounts of wholesale and retail deposits that make up M4 but for which no overall rates are available.

We only have quarterly deposit rate data prior to 1998. For those years, we interpolate the spreads to three-month sterling LIBOR rates linearly over each quarter and add them to the monthly LIBOR rates in order to construct monthly time series for retail and wholesale deposit rates.

References

Andrés, J., J. D. López-Salido, and E. Nelson. 2004. "Tobin's Imperfect Asset Substitution in Optimizing General Equilibrium." *Journal of Money, Credit, and Banking* 36 (4): 665–90.

Bank of England. 2009. "Markets and Operations." *Quarterly Bulletin* (Bank of England) 49 (4): 264.

Bean, C. 2009. "Quantitative Easing: An Interim Report." Speech to the London Society of Chartered Accountants, London, October 13.

Benford, J., S. Berry, K. Nikolov, M. Robson, and C. Young. 2009. "Quantitative Easing." *Quarterly Bulletin* (Bank of England) 49 (2): 90–100.

Bernanke, B., and V. R. Reinhart. 2004. "Conducting Monetary Policy at Very Low Short-Term Interest Rates." *American Economic Review, Papers and Proceedings* 94 (2): 85–90.

[36]This information is available on www.bankofengland.co.uk/statistics/bankstats/current/index.htm#1.

Bernanke, B., V. R. Reinhart, and B. Sack. 2004. "Monetary Policy Alternatives at the Zero Bound: An Empirical Assessment." *Brookings Papers on Economic Activity* 2: 1–78.

Berry, S., R. Harrison, R. Thomas, and I. de Weymarn. 2007. "Interpreting Movements in Broad Money." *Quarterly Bulletin* (Bank of England) 47 (3): 376–88.

Borio, C., and P. Disyatat. 2009. "Unconventional Monetary Policies: An Appraisal." BIS Working Paper No. 292.

Brigden, A., B. Martin, and C. Salmon. 1997. "Decomposing Exchange Rate Movements According to the Uncovered Interest Rate Parity Condition." *Quarterly Bulletin* (Bank of England) 4 (November): 377–89.

Brunner, K., and A. H. Meltzer. 1973. "Mr. Hicks and the 'Monetarists'." *Economica* 40 (157): 44–59.

Campbell, J. 1999. "Comment on Gregory D. Hess, 'The Maturity Structure of Government Debt and Asset Substitutability in the United Kingdom'." In *Government Debt Structure and Monetary Conditions*, ed. K. A. Chrystal, 156–61. (Proceedings of a conference organized by the Bank of England, June 18–19, 1998). London: Bank of England.

Clouse, J., D. Henderson, A. Orphanides, D. Small, and P. Tinsley. 2003. "Monetary Policy When the Nominal Short-Term Interest Rate Is Zero." *Topics in Macroeconomics* 3 (1): 1088–22.

Cross, M., P. Fisher, and O. Weeken. 2010. "The Bank's Balance Sheet During the Crisis." *Quarterly Bulletin* (Bank of England) 50 (1): 34–42.

Dale, S. 2010. "QE — One Year On." Remarks at the CIMF and MMF Conference on "New Instruments of Monetary Policy: The Challenges," Cambridge, March 12.

D'Amico, S., and T. B. King. 2010. "Flow and Stock Effects of Large-Scale Treasury Purchases." Federal Reserve Board Finance and Economics Discussion Series Paper No. 2010-52.

Eggertsson, G., and M. Woodford. 2003. "The Zero Bound on Interest Rates and Optimal Monetary Policy." *Brookings Papers on Economic Activity* 1: 139–211.

Engel, C., J. A. Frankel, K. A. Froot, and A. P. Rodrigues. 1995. "Tests of Conditional Mean-Variance Efficiency of the U.S. Stock Market." *Journal of Empirical Finance* 2 (1): 3–18.

Engle, R. F., and K. F. Kroner. 1995. "Multivariate Simultaneous Generalized ARCH." *Econometric Theory* 11 (1): 122–50.

Fisher, P. 2010. "The Corporate Sector and the Bank of England's Asset Purchases." Speech to the Association of Corporate Treasurers, London, February 18.

Frankel, J. 1985. "Portfolio Crowding-Out, Empirically Estimated." *Quarterly Journal of Economics* 100 (5): 1041–65.

Friedman, B. M. 1978. "Crowding Out or Crowding In? The Economic Consequences of Financing Government Deficits." *Brookings Papers on Economic Activity* 9 (3): 593–641.

Gagnon, J., M. Raskin, J. Remache, and B. Sack. 2011. "The Financial Market Effects of the Federal Reserve's Large-Scale Asset Purchases." *International Journal of Central Banking* 7 (1): 45–52.

Hess, G. 1999. "The Maturity Structure of Government Debt and Asset Substitutability in the United Kingdom." In *Government Debt Structure and Monetary Conditions*, ed. K. A. Chrystal, 131–55. (Proceedings of a conference organized by the Bank of England, June 18–19, 1998). London: Bank of England.

Hosking, R. M. 1981. "Equivalent Forms of the Multivariate Portmanteau Statistic." *Journal of the Royal Statistical Society* Series B (43): 261–62.

Inkinen, M., M. Stringa, and K. Voutsinou. 2010. "Interpreting Equity Price Movements Since the Start of the Financial Crisis." *Quarterly Bulletin* (Bank of England) 50 (1): 24–33.

Joyce, M., A. Lasaosa, I. Stevens, and M. Tong. 2010. "The Financial Market Impact of Quantitative Easing." Bank of England Working Paper No. 393.

Klyuev, V., P. de Imus, and K. Srinivasan. 2009. "Unconventional Choices for Unconventional Times: Credit and Quantitative Easing in Advanced Economies." IMF Staff Position Note No. 2009/27 (November).

Meier, A. 2009. "Panacea, Curse, or Nonevent? Unconventional Monetary Policy in the United Kingdom." IMF Working Paper No. 09/163.

Miles, D. 2010. "Interpreting Monetary Policy." Speech at the Imperial College Business School, London, February 25.

Modigliani, F., and R. C. Sutch. 1966. "Innovations in Interest Rate Policy." *American Economic Review, Papers and Proceedings* 56 (2): 178–97.

Neely, C. 2010. "The Large-Scale Asset Purchases Had Large International Effects." Federal Reserve Bank of St. Louis Working Paper No. 2010-018A.

Oakley, D. 2009. "Gilts Prices Jump as Bank Prepares to Buy Bonds." *Financial Times*, March 5. Available at http://www.ft.com.

Roley, V. 1979. "A Theory of Federal Debt Management." *American Economic Review* 69 (5): 915–26.

———. 1982. "The Effects of Federal Debt-Management Policy on Corporate Bond and Equity Yields." *Quarterly Journal of Economics* 97 (4): 645–68.

Shiratsuka, S. 2009. "Size and Composition of the Central Bank Balance Sheet: Revisiting Japan's Experience of the Quantitative Easing Policy." IMES Discussion Paper No. 2009-E-25.

Tobin, J. 1958. "Liquidity Preference as Behavior Towards Risk." *Review of Economic Studies* 25 (2): 65–86.

———. 1961. "Money, Capital and Other Stores of Value." *American Economic Review* 51 (2): 26–37.

———. 1963. "An Essay on the Principles of Debt Management." In *Fiscal and Debt Management Policies*, 143–218. Englewood Cliffs, NJ: Prentice Hall.

———. 1969. "A General Equilibrium Approach to Monetary Theory." *Journal of Money, Credit, and Banking* 1 (1): 15–29.

Ugai, H. 2007. "Effects of the Quantitative Easing Policy: A Survey of Empirical Analyses." *Monetary and Economic Studies* 25 (1): 1–48.

Vayanos, D., and J.-L. Vila. 2009. "A Preferred-Habitat Model of the Term Structure of Interest Rates." Working Paper, London School of Economics.

Walsh, C. 1982. "Asset Substitutability and Monetary Policy: An Alternative Characterization." *Journal of Monetary Economics* 9 (1): 59–71.

Determinants of House Prices in Nine Asia-Pacific Economies*

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The paper investigates the characteristics of house price dynamics and the role of institutional factors in nine Asia-Pacific economies during 1993–2006. On average, house prices tend to be more volatile in markets with lower supply elasticity and a more flexible business environment. At the national level, the current run-up in house prices mainly reflects adjustment to improved fundamentals rather than speculative housing bubbles. However, evidence of bubbles does exist in some market segments.

JEL Codes: G12, R31.

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1. Introduction

House price risk has attracted much attention in recent years. A number of industrialized economies, including the United States, the United Kingdom, and Spain, had witnessed a protracted period of significant increases in house prices in the mid-2000s. The perceived lower risk encouraged lax lending criteria in mortgage markets, which greatly contributed to the U.S. subprime crisis and the consequent global financial crisis. Just as suggested by previous studies, house price fluctuations have caused a major impact on household consumption,[1] the banking system, and the real economy.[2]

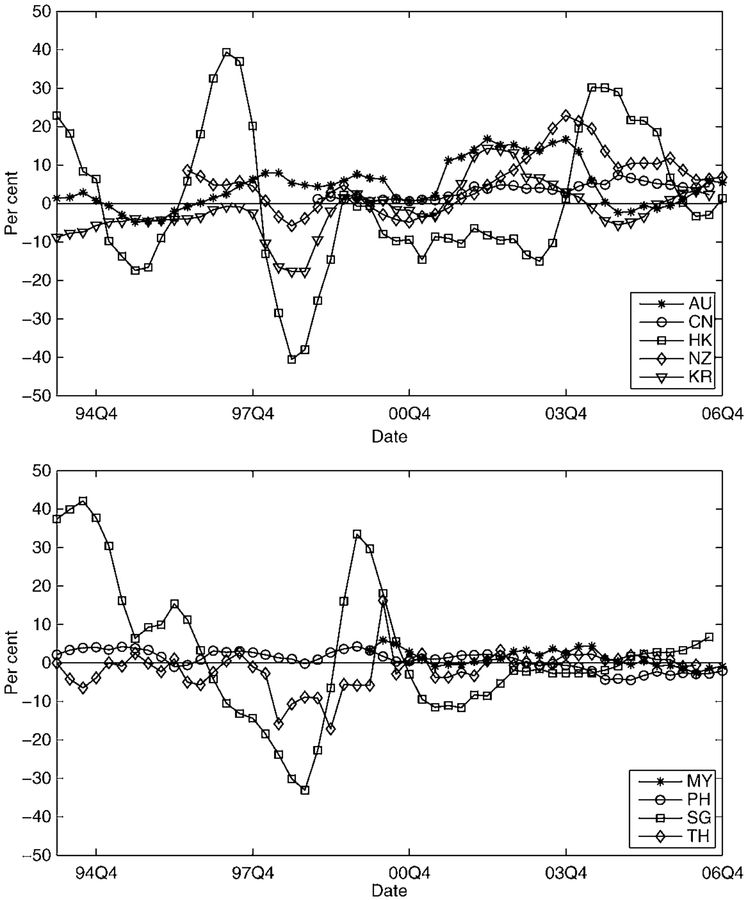
By comparison, housing markets in most Asian economies were relatively tranquil during the same period. In recent years, however, there have been growing concerns about the housing market in a few economies. China, Hong Kong SAR (Hong Kong hereafter), and South Korea (Korea hereafter) have witnessed very strong house price inflation in the past several years (see figure 1). Given the not-so-distant experience of financial crises in this region (such as the 1997 Asian crisis and the so-called lost decade in Japan), in which booms and busts in real estate markets played a crucial role, the question is whether the observed house price growth can potentially lead to bubble episodes.

There are two opposite views regarding the issue of possible housing bubbles in Asia. A pessimistic view argues that house prices have been overvalued in many countries and will face downward corrections in the near future. At the extreme, some consider it evidence of new speculative housing bubbles and call for supervisors and central banks to adopt prudential measures to contain them. By contrast, the optimistic view considers this round of house price growth as a manifestation of recovery from the previous crisis episode. The optimists argue that in the aftermath of previous crises, house prices were too low compared with their fundamental values. Therefore,

[1] See Girouard and Blöndal (2001), Cocco (2005), Yao and Zhang (2005), and Campbell and Cocco (2007). In addition, empirical findings (see Helbling and Terrones 2003; Case, Quigley, and Shiller 2005) also suggest that housing tends to have a bigger wealth effect than financial assets.

[2] Bernanke and Gertler (1995), Bernanke, Gertler, and Gilchrist (1996), Kiyotaki and Moore (1997), Aoki, Proudman, and Vlieghe (2004), and Gan (2007) show the strong linkages between the housing cycle and the credit cycle.

Figure 1. House Price Inflation (yoy) in Average Residential Markets, 1994–2006



Note: AU: Australia; CN: China; HK: Hong Kong SAR; KR: Korea; MY: Malaysia; NZ: New Zealand; PH: the Philippines; SG: Singapore; TH: Thailand.

the rebound of house prices from the very low levels is simply a consequence of the mean-reversion process. Moreover, the liberalization of housing markets and housing finance systems in the past decade, including a general trend towards more market-based housing markets, greater availability of mortgage products, and more liquid secondary mortgage markets, has arguably improved market efficiency, stimulated demand, and contributed to house price growth.

The paper sheds some light on this debate by examining house price developments in nine economies in the Asia-Pacific area, which include Australia, China, Hong Kong, Korea, Malaysia, New Zealand, the Philippines, Singapore, and Thailand.[3] Specifically, it attempts to address the following questions: What determines the fundamental values and short-term dynamics of house prices in Asia? What is the impact of institutional factors on house price movements? How can one gauge whether there is a housing bubble? To address these questions, we adopt the error-correction framework used by Capozza et al. (2002). The key results are as follows.

First, we find that the patterns of national house price dynamics exhibit significant cross-country heterogeneity. Moreover, institutional factors matter. In particular, distinction in house price dynamics can be largely attributable to cross-country differences in land supply and business environments.

Second, based on the econometric analysis, we characterize house price movements as the sum of three separate factors: (i) the fundamental value of housing (a trend term) that is determined by longer-term economic conditions and institutional arrangements, (ii) the deviation from fundamental values that is attributable to frictions in the housing market (a cyclical term), and (iii) an irrational or "bubble" component that is likely to be driven by overly optimistic expectations (an error term). Applying this approach to nine economies in Asia and the Pacific, we find that national house price movements before the onset of the global financial crisis mainly reflected changes in fundamental values or cyclical adjustments towards fundamentals. In other words, there is little evidence of housing bubbles in these economies, at least at the national levels.

The decomposition analysis has important policy implications. It allows for distinction between house price overvaluation (the sum of cyclical and error components) and a housing bubble. Policy recommendations are different accordingly. To mitigate house price overvaluation that is driven by cyclical movements related to market frictions, a policymaker should probably focus on measures that

[3]In this paper, we use the term "Asia" to represent the set of economies mentioned above.

aim at reducing the magnitude and frequency of house price cycles, such as loosening land use regulation, improving information availability and transparency, and enhancing the freedom in the business environment. By contrast, to contain a bubble, the policymaker should instead adopt measures that control unwarranted high expectations of capital gains or overconfidence of investors in the housing market.

The remainder of the paper is organized as follows. Section 2 provides an overview of the literature and highlights the contributions of this study. Section 3 describes the data and explains the empirical method used to examine the questions of interest, and section 4 discusses the empirical results. Finally, section 5 concludes and provides some policy perspectives.

2. A Review of the Literature and Our Contributions

2.1 Literature Review

To monitor the housing market, it is important to understand first the determinants of house prices. Housing is a special type of asset that has a dual role as a consumption and an investment good. From the long-term perspective, the equilibrium price a household is willing to pay for a house should be equal to the present discounted value of future services provided by the property, i.e., the present value of future rents and the discounted resale value of the house. From the short-term perspective, however, house prices can deviate from their fundamental values, on account of some unique characteristics of the real estate market (such as asset heterogeneity, downpayment requirements, short-sale restrictions, lack of information, and lags in supply). For instance, Leung and Chen (2006) show that land prices can exhibit cycles due to the role of intertemporal elasticity of substitution. Wheaton (1999) and Davis and Zhu (2004) develop a model in which there are lags in the supply of real estate and bank lending decisions depend on the property's current market value (labeled as historical dependence). They show that in response to a change in fundamental values, real estate prices can either converge to or exhibit oscillation around the new equilibrium values.

Existing literature suggest that house price movements are closely related to a common set of macroeconomic variables, market-specific conditions, and housing finance characteristics. Hofmann (2004) and Tsatsaronis and Zhu (2004) examine the determinants of house prices in a number of industrialized economies and find that economic growth, inflation, interest rates, bank lending, and equity prices have significant explanatory power. The linkage between property and bank lending is particularly remarkable, as also highlighted by Herring and Wachter (1999), Chen (2001), Hilbers, Lei, and Zacho (2001), and Gerlach and Peng (2005). This is not surprising given the heavy reliance on mortgage financing in the housing market. Moreover, housing markets are local in nature. Garmaise and Moskowitz (2004) find strong evidence that asymmetric information about local market conditions plays an important role in reshaping property transactions and determining the choice of financing. Green, Malpezzi, and Mayo (2005) find that house price dynamics differ across metropolitan areas with different degrees of supply elasticities.

On the important issue of detecting house price bubbles, there are several approaches adopted in the literature. Bubble episodes are sometimes assessed by market analysts in terms of the price-rent ratio or the price-income ratio. A bubble is typically identified if the current ratio is well above the historical average level. These measures, however, may be inadequate barometers for policy analysis because they ignore the variation in "equilibrium" price-rent (or price-income) ratios driven by fluctuations in economic fundamentals (e.g., rent growth, income growth, and the desired rate of return). To overcome these problems, two methods have been proposed. The first method compares observed price-rent ratios with time-varying discount factors that are determined by the user cost of owning a house, which consists of mortgage interest, property tax, maintenance cost, tax deductibility of mortgage interest payments, and an additional risk premium (see Himmelberg, Mayer, and Sinai 2005; Ayuso and Restoy 2006; Brunnermeier and Julliard 2008). The second method compares observed house prices with fundamental values that are predicted based on the long-run relationship between house prices and macroeconomic factors (see Abraham and Hendershott 1996; Kalra, Mihaljek, and Duenwald 2000; Capozza et al. 2002, and Holly, Pesaran, and Yamagata 2010, for example). In this paper, we adopt the second method because of data limitations and heterogeneity in what constitutes appropriate measurement of the user cost across countries.[4]

2.2 Contributions of This Study

This paper examines the determinants of house price fundamentals and short-term dynamics in nine Asia-Pacific economies and thirty-two cities/market segments in these economies, discusses the role of distinctive institutional arrangements, and explores the possible emergence of housing bubbles. The empirical framework used in this study follows Capozza et al. (2002), which examines the long-term and short-term dynamics of house prices in a three-step econometric analysis and characterizes the pattern of house price dynamics via a combination of serial correlation and mean-reversion coefficients (see section 3.2).[5] However, our study extends the previous analysis in three important ways.

First, previous studies have mainly focused on the lessons from industrialized economies. This study is one of the first papers to investigate the evidence in the Asia-Pacific, which has gained an increasing importance in the global economy. Given the remarkable experience of housing bubbles in many of the Asian economies in the 1990s, it is interesting to examine the house price movements after the crisis episode. In addition, Asia-Pacific housing markets differ substantially from those of industrialized economies in terms of the level of economic and financial market developments as well as institutional arrangements. In this regard, the results could provide complementary views to existing studies.

[4]Rent data in our sample economies are often not available or not comparable with the house price data (referring to different samples). It is also difficult to quantify some key components of the user cost, such as the tax deductibility and the risk premium in individual markets.

[5]The methodology is also similar to that of Holly, Pesaran, and Yamagata (2010), which determines the extent to which real house prices in forty-nine U.S. states (over a sample period of twenty-nine years) are driven by fundamentals such as real per capita disposable income and common shocks. Their study also looks into the speed of adjustment of real house prices to macroeconomic and local disturbances. However, it differs from ours in that it explicitly examines the role of spatial factors—in particular, the effect of contiguous states by use of a weighting matrix.

Second, our analysis emphasizes the impact of institutional factors on house price dynamics. The original paper by Capozza et al. (2002) analyzes the impact of a number of factors (such as population, income, and construction cost) on house price dynamics, but the role of institutions in not examined.[6] In this study, we construct a composite measure of institutional factors on the basis of four different aspects of market developments. This measure not only differs across countries but also varies over time. Indeed, our analysis shows that the institutional factor is important in explaining house price determination in the nine Asian economies.

Third, we extend the housing bubble literature by (i) distinguishing between house price growth and house price overvaluation, and (ii) decomposing house price overvaluation into cyclical and bubble components. The first distinction is quite obvious. House price growth may simply reflect the increase in the fundamental value of the property, which is driven by income, mortgage rates, and other factors. By contrast, house price overvaluation refers to the situation that current house prices are higher than the fundamental values. The second distinction is more subtle. A bubble is necessarily related to house price overvaluation, but not vice versa. This is because frictions in the housing market, including lags in supply and credit market imperfections, may cause house prices to deviate from their fundamental values in the short term. In this paper, this cyclical component of house price overvaluation is captured by the serial correlation and mean reversion of house price dynamics. The unexplained part is then defined as the bubble component that is more likely to be driven by overly optimistic expectations in the housing market. Understanding the source of house price overvaluation sheds important light on the appropriate policy actions.

3. Data Description and Empirical Methodology

In this section, we briefly describe the data used in this study and outline the empirical methodology adopted to characterize house

[6]One reason for the omission of institutional factors is that Capozza et al. examine the house price determination in a number of metropolitan areas in the United States. The institutional arrangements—such as business freedom, legal framework, and property rights protection—have little variation across areas.

price dynamics and to analyze the bubble component in house price overvaluation.

3.1 Data Description

Quarterly data for the residential property sector in nine economies and thirty-two cities/market segments in Asia[7] were used in the analysis. Where data are available, quarterly series spanning the period 1993–2006 were used.

The house price data have certain limitations. There are some subtle variations in the definition of house prices used in the estimation (see table 6 in the appendix). While some series are derived using a hedonic pricing method, some are simply based on floor-area prices collected by land registration authorities and the private sector, for which no quality adjustment was done. Moreover, the time series are relatively short. Except for Hong Kong, Korea, Singapore, and Thailand, quarterly house price data only cover the post-Asian-crisis period. However, longer time series of house price data may not necessarily improve the results in the sense that many Asian economies have experienced a regime shift in housing markets and housing finance systems, which has arguably led to discontinuities in the dynamics.

Apart from residential property prices, other series used in this study include real GDP, population, construction cost index, land supply index, mortgage credit-to-GDP ratios, real mortgage rates, real effective exchange rates, stock price index, and the first principal component of four institutional indices—the business freedom index, the financial freedom index, the corruption index, and the property

[7]Other emerging Asian economies are excluded from this study because house price data are not available. At the city level, Beijing, Chongqing, Guangzhou, Shanghai, Shenzhen, and Tianjin are included in China; Busan, Daegu, Daejon, Gwangju, Incheon, Seoul, and Ulsan are included in Korea; Johor, Kuala Lumpur, Pahang, Perak, and Pinang are included in Malaysia; and Caloocan, Makati, Manila, Pasay, Pasig, and Quezon are included in the Philippines. In addition, for Hong Kong, Singapore, Bangkok, Manila, and Kuala Lumpur, there are two separate sets of house prices for the average market and for the luxury market segments, respectively.

rights index[8]—which we believe to be the most relevant representation of the institutional factors that affect the housing market and, hence, house price movements. Table 1 reports summary statistics of key variables used in this study, for each country and for the whole sample.

3.2 Empirical Methodology: Characterizing House Price Dynamics

We follow the framework used by Capozza et al. (2002) to investigate the long-term and short-term determinants of house price movements. The approach can be divided into three steps. In the first step, the fundamental value of housing is calculated. In the second step, the short-term dynamics of house prices are characterized by a mean-reversion process to their fundamental values and by a serial correlation movement. In the third step, the degree of persistence in house prices and the relative speed by which house prices revert to their fundamental values are assessed by interacting the serial correlation and mean-reversion coefficients with macroeconomic and housing market variables as well as institutional factors.

3.2.1 The Fundamental Value of Housing

It is assumed that in each period and in each area (a country or a city), there is a fundamental value of housing that is largely determined by economic conditions and institutional arrangements:

$$P_{it}^{*} = f(X_{it}), \tag{1}$$

where P_{it}^{*} is the log of the real fundamental value of house prices in country i at time t, $f(\cdot)$ is a function, and X_{it} is a vector

[8]The business freedom index measures the ability to create, operate, and close an enterprise quickly and easily. Burdensome, redundant regulatory rules are the most harmful barriers to business freedom. The financial freedom index is a measure of banking security as well as independence from government control. The corruption index is a measure of the perception of corruption in the business environment, including levels of governmental legal, judicial, and administrative corruption. The property rights index measures the ability of individuals to accumulate private property, secured by clear laws that are fully enforced by the state (www.heritage.org).

Table 1. Summary Statistics

Variables	Total	AU	CN	HK	KR	MY	NZ	PH	SG	TH
RHP	109.07	109.05	108.35	114.28	116.87	102.29	116.87	105.95	95.73	109.50
	20.0	26.0	10.0	27.1	13.4	3.7	24.6	20.2	13.9	11.7
ΔRHP(%)	0.19	1.08	0.80	−0.25	−0.45	0.31	1.41	−0.93	0.60	−0.36
	5.5	1.8	0.9	6.3	2.2	1.1	2.0	12.5	4.1	5.0
ΔReal GDP(%)	5.12	3.72	9.08	4.33	5.26	5.66	3.51	4.36	6.183	4.01
	4.0	1.2	1.5	4.3	4.3	4.9	1.7	2.0	4.8	5.3
Population (mn)	161.41	19.09	1249.03	6.57	46.37	22.62	3.87	73.73	3.87	61.73
	380.5	0.9	39.5	0.2	1.3	2.1	0.2	5.6	0.3	2.4
RMR (%)	4.84	5.13	2.32	4.75	2.98	3.33	6.60	6.06	5.37	5.64
	3.3	1.7	6.1	3.9	0.7	2.1	1.3	2.4	1.3	2.4
Mort/GDP (%)	97.09	151.76	8.22	164.21	7.60	91.26	252.49	20.55	147.19	15.38
	82.1	40.6	1.7	34.5	7.6	15.1	37.5	5.9	31.3	1.4
LSI	147.05	105.95	108.47	91.74	123.18	87.94	119.26	115.26	138.75	440.68
	185.7	14.5	56.4	47.8	32.8	18.3	29.1	30.5	137.8	448.7
RCC	102.53	99.39	108.51	92.15	103.96	102.02	102.34	105.12	103.60	104.47
	7.7	3.1	11.1	5.9	4.9	3.7	3.5	10.1	4.4	5.9
EPI	104.16	110.89	94.48	93.24	103.46	106.14	120.41	102.72	100.31	105.83
	13.3	10.4	8.8	10.5	11.0	11.1	13.5	12.6	5.7	12.0
REER	110.32	93.82	73.67	74.13	110.41	99.76	108.94	130.99	90.00	106.04
	57.9	27.8	21.2	17.8	32.3	22.7	16.7	41.8	16.0	11.7

(continued)

Table 1. (Continued)

Variables	Total	AU	CN	HK	KR	MY	NZ	PH	SG	TH
BFI	60.64	60.37	31.74	89.78	52.80	61.73	72.55	35.35	90.36	52.12
	21.4	13.9	5.8	0.8	9.4	10.0	8.1	9.4	1.2	7.1
FFI	63.46	90	40	88.33	56.67	40	90	48.33	70	50
	21.0	0	10.1	5.6	9.5	10.1	0	5.6	0	0
CI	64.83	83.33	31.583	85.67	58.75	61.583	92.18	27	91.08	54.58
	25.1	8.1	2.1	5.3	13.5	10.1	2.5	5.5	1.5	18.5
PRI	72.80	90	30	90	83.33	60	90	53.33	90	70
	22.0	0	0	0	9.5	10.1	0	16.2	0	14.3

Notes: This table reports the summary statistics of key variables, in each country and in the whole sample (1993–2006). For each variable, the numbers in the first row represent sample mean and those in the second row represent the standard deviation. RHP: real house price index; ΔRHP: real house price growth (quarterly); RMR: real mortgage rate; Mort/GDP: mortgage credit/GDP ratio; LSI: land supply index; RCC: real construction cost index; EPI: equity price index; REER: real effective exchange rate; BFI: business freedom index; FFI: financial freedom index; CI: corruption index; PRI: property rights index.

of macroeconomic and institutional variables that determine house price fundamentals.

We adopt a general-to-specific approach in assessing the determinants of house price fundamentals. We start by including all the identified explanatory factors in our list to investigate their long-term relationship with house prices, using either single-equation ordinary least squares (OLS) or panel data techniques.[9] Only regressors found to be significant at the 5 percent level are retained in the final model specification. We choose four blocks of explanatory variables based on theoretic reasoning or previous empirical work.

The first block of explanatory variables consists of demand-side factors, including real GDP, population, the real mortgage rate, and the mortgage credit-to-GDP ratio. The inclusion of the real mortgage rate and mortgage credit is premised on the bank-dominated nature of financial systems across Asia. We posit that *higher income and higher population* tend to encourage greater demand for new housing and housing improvements. In addition, *mortgage rate* is expected to be negatively related to housing prices. A higher mortgage rate entails higher amortization, which, in turn, impinges on the cash flow of households. This reduces the affordability of new housing, dampens housing demand, and pushes down house prices. Similarly, the growth in *mortgage credit* increases the financing capacity of households and stimulates the demand for housing.

The second block of variables is made up of supply-side factors consisting of the land supply index and real construction cost. The *land supply index*, which refers to the building permit index in most countries, measures the flexibility of supply to demand conditions. In the long run, an increase in land supply tends to bring down house prices. By contrast, the burden of higher *real construction costs* will be shared by purchasers, and we expect a positive relationship between real construction costs and equilibrium house prices.

The third block of variables consists of prices of other types of assets such as equity prices and exchange rates. It is well documented that house prices tend to co-move with other asset prices. For instance, Sutton (2002) and Borio and McGuire (2004) find

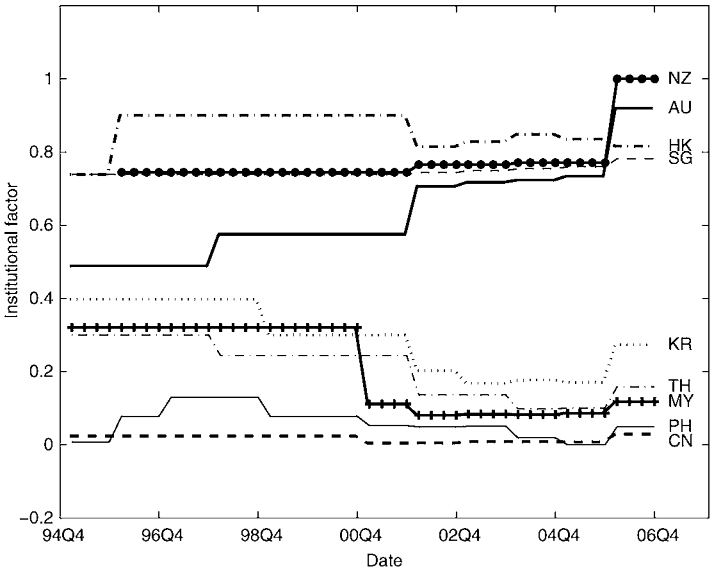
[9]To avoid simultaneity bias, contemporaneous variables are instrumented with own lags.

strong linkages between *equity price* and house price movements. The direction of such linkage, from a theoretical perspective, is not clear, as the substitution effect and wealth effect point in opposite directions.[10] Moreover, a *real effective exchange rate appreciation* is expected to exert positive influence on property market prices, particularly in markets where there is substantial demand from non-residents for investment purposes. In countries where foreign investment plays an important role in the economy, such as in Asia, an exchange rate appreciation is normally associated with housing booms.

Lastly, the fourth block consists of an institutional factor that attempts to account for the impact of market arrangements on equilibrium house prices. The institutional factor is constructed as the first principal component of four index variables compiled by the Heritage Foundation: the business freedom index, the corruption index, the financial sector index, and the property rights index. It is constructed so that we can examine the impact of business, regulatory, and financial conditions on the determination of house prices in a parsimonious way. The first principal component has approximately equal weights of the four indices and accounts for about 80 percent of the variability in the four-index series. A higher score in the institutional factor is associated with higher business freedom, better regulatory conditions, lower corruption, a greater range of intermediation functions by the financial sector, a higher degree of flexibility in acquiring land, and better legal protection to land/home owners. As shown in figure 2, the institutional factor exhibits substantial time variation and cross-country differences. The nine economies can be easily divided into two groups: Australia, Hong Kong, New Zealand, and Singapore are classified as more business friendly and the other five economies as less so. Over time, Australia and New Zealand experienced major improvements, while Malaysia and Thailand witnessed deterioration in their business environment during the period under review.

[10]A substitution effect predicts a negative relationship between the prices of the two assets, as the high return in one market tends to cause investors to leave the other market. A wealth effect, by contrast, predicts a positive relationship because the high return in one market will increase the total wealth of investors and their capability of investing in other assets.

Figure 2. Time Series of Institutional Factors



Notes: The figure plots the time series of the institutional factor in each of the nine economies under review. The institutional factor is defined as the first principal component of four index series: the business freedom index, the financial freedom index, the corruption index, and the property rights index. The institutional factor is rescaled into a range between 0 and 1.

Several remarks are worth mentioning here. First, we use the *trend* component of mortgage credit-to-GDP ratios and equity prices in explaining the long-run house price fundamentals. This modification recognizes that the original raw series may contain non-fundamental components and that a housing bubble often comes together with excessive growth in mortgage credit and sometimes interacts with extreme equity price movements. Using the trend series of the two variables can ensure that our estimates of house price fundamentals are not contaminated by the non-fundamental (or bubble) components and, by extension, minimize potential errors in the analysis.[11]

[11]We do recognize, nonetheless, that the trend needs to be estimated over a long enough sample period that is not dominated by bubble episodes. Data constraints, however, prevented us from estimating the trend series over a longer period of time.

Second, since the stochastic variables included in the long-run equation are mostly non-stationary, we check for cointegration by establishing first the stationarity of the residuals of the long-term equation before proceeding to the second stage. This is to address the concern of Gallin (2006) who, using the U.S. data, suggests that standard and more powerful panel data tests fail to reject the hypothesis of no cointegration, and therefore the error-correction specification in analyzing short-term dynamics may be inappropriate.

Third, there has been substantial evidence that mortgage finance system arrangements, including the terms of mortgage contract, lending practices, valuation method of collateral assets, real estate taxes, and innovations in the mortgage markets, have important implications on house price dynamics.[12] Ideally, we would like to also include a set of variables indicating the time variation and cross-country differences in housing finance systems. Nevertheless, information on housing finance systems is at best only available on a snapshot basis, and often with qualitative rather than quantitative features. Therefore, the impact of housing finance systems cannot be directly examined in this study. However, there is evidence that housing finance system arrangements may depend on the stage of economic development, the advances in credit information systems, and the strength of legal rights (Warnock and Warnock 2008). It is probably no accident that economies with a higher score in the composite institutional factor (including Australia, Hong Kong, New Zealand, and Singapore) coincide with those with more advanced housing financing systems and more active secondary mortgage markets (Zhu 2006). Therefore, the institutional factor in our study may also be interpreted as a proxy variable for the development in housing financing systems, although the link is very loose and at best an indirect one.

[12] See, for example, Estrella (2002), McCarthy and Peach (2002), Tsatsaronis and Zhu (2004), Peek and Wilcox (2006), and Égert and Mihaljek (2007). For descriptions of developments of housing finance systems in the past several decades, see Diamond and Lea (1992), European Central Bank (2003), Hegedüs and Struyk (2005), Organisation for Economic Co-operation and Development (2005), and Committee on the Global Financial System (2006).

3.2.2 Short-Run Dynamics of House Prices

Arguably, equilibrium is rarely observed in the short run due to the inability of economic agents to adjust instantaneously to new information. As suggested by Capozza et al. (2002), house price changes in the short run are governed by reversion to fundamental values and by serial correlation according to

$$\Delta P_{it} = \alpha \Delta P_{i,t-1} + \beta\big(P^*_{i,t-1} - P_{i,t-1}\big) + \gamma \Delta P^*_{it}, \qquad (2)$$

where P_{it} is the log of (observed) real house prices and Δ is the difference operator.

If housing markets are efficient, prices will adjust instantaneously such that $\gamma = 1$ and $\alpha = 0$. Considering that housing is a slow-clearing durable asset, it is reasonable to expect that current price changes are partly governed by previous changes in own price levels ($\alpha > 0$), by the deviation from the fundamental value ($0 < \beta < 1$), and partly by contemporaneous adjustment to changes in fundamentals ($0 < \gamma < 1$).

Capozza et al. (2002) shows that the above model specification allows for rich dynamics of house price movements, depending on the size of the coefficients α and β. The various patterns of house price dynamics can be summarized in figure 3.[13]

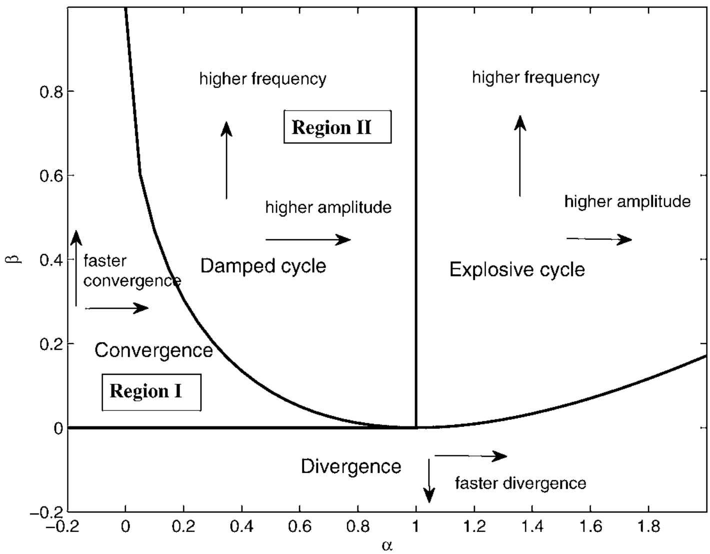
To summarize, the sufficient and necessary condition for a house price cycle to be stable is $\alpha < 1$ and $\beta > 0$. If satisfied, there are two possible types of house price movements:

(i) If $(1 + \alpha - \beta)^2 - 4\alpha \geq 0$ (region I in figure 3), the house price will converge monotonically to the equilibrium level. In this case, the transitory path itself does not generate house price cycles. In other words, house price cycles only reflect cyclical movements in their fundamental values. The speed of convergence depends on the magnitude of the two coefficients: the convergence rate is generally higher when α and β are larger.

(ii) If $(1 + \alpha - \beta)^2 - 4\alpha > 0$ (region II in figure 3), the transitory path in response to changes in equilibrium house price

[13]The strict proof is available upon request.

Figure 3. Characteristics of House Price Dynamics: Illustration



Note: The figure plots the characteristics of house price dynamics for different combinations of persistence (α) and mean-reversion (β) parameters.

values exhibits a damped fluctuation around the equilibrium level. The magnitude of the two coefficients, again, determines the property of the oscillation. Generally, a higher α implies a higher amplitude and a higher β implies a higher frequency of the fluctuation process.

If $\alpha \geq 1$ or $\beta \leq 0$, then the house price cycle is unstable. House prices may either diverge or exhibit an amplified fluctuation away from the equilibrium level, but such movements cannot be sustainable. In general, such features should not exist in any housing market for a prolonged period.

3.2.3 Endogenous Adjustment in Short-Run Dynamics

Given the importance of mean-reversion and serial correlation coefficients, the next step is to analyze what determines α and β.

Following Capozza et al. (2002), we introduce interactive terms in the mean-reversion and serial correlation coefficients:

$$\Delta P_{it} = \left[\alpha_0 + \sum_j \alpha_j Y_{ijt}\right] \Delta P_{i,t-1} + \left[\beta_0 + \sum_j \beta_j Y_{ijt}\right] \left(P^*_{i,t-1} - P_{i,t-1}\right) + \gamma \Delta P^*_{it}, \qquad (3)$$

where Y_{ijt} is a list of region-specific economic variables, housing market variables, and—what is new in this study—the composite institutional factor.[14] Introducing the interactive terms allows the two coefficients to differ across regions and to vary over time. For each country, the average serial correlation and mean-reversion coefficients are $\alpha_i = \alpha_0 + \sum_j \alpha_j \overline{Y_{ijt}}$ and $\beta_i = \beta_0 + \sum_j \beta_j \overline{Y_{ijt}}$, respectively, where $\overline{Y_{ijt}}$ represents the time average of Y_j in country i.

3.3 Detecting Housing Bubbles

We employ the above empirical results to investigate the issue of house price overvaluation and to quantify the two components of such overvaluation. One is the cyclical component that is attributable to the intrinsic house price cycles (related to supply and institutional frictions in the adjustment process) and the other is a bubble component that cannot be explained by these cyclical factors.

House price overvaluation is defined as observed house prices (P_t) being higher than predicted house price fundamentals (P^*_t) (see section 3.2.1, subscript i omitted). Intuitively, it is distinct from high house price inflation because the latter may simply reflect the increase in house price fundamentals.

More importantly, we also make a clear distinction between house price overvaluation and a house price bubble, which are often mixed in the existing literature. Throughout this paper, a housing bubble is defined via component analysis of house price overvaluation. As

[14] Similarly, we also adopt a general-to-specific approach, in that we start by including a list of possible factors but the final model specification only includes those variables with significant interactive effects.

suggested by Wheaton (1999) and Davis and Zhu (2004), frictions in housing markets can generate intrinsic house price cycles, causing house prices to deviate (sometimes substantially) from their fundamental values in the short term. We consider this cyclical component of house price overvaluation to be reflected in our estimates of short-term dynamics. The residual component that cannot be explained by the intrinsic adjustment process is what we define in this paper as the "bubble" component (also see Brunnermeier and Julliard 2008).

Specifically, for a given house price overvaluation $(P_t - P_t^*)$, the cyclical component is calculated as $P_{t-1} + E(\Delta P_t) - P_t^*$, where $E(\Delta P_t)$ is the predicted value from short-term dynamics (see equation (3)). Notice that the sum of the first two elements is the predicted house price based on short-term dynamics; its deviation from the fundamental value P_t^* is attributable to the short-run cyclical movement of house prices. By comparison, the residual component, labeled as the "bubble" component in this study, is defined as house price overvaluation minus this cyclical component. Hence, house price overvaluation is not equivalent to a house price bubble in our framework.

There are certain limitations in our definition of a housing bubble. For one, it is defined loosely. The definition of the bubble component is contingent on the accuracy of the model used to estimate house price dynamics. Strictly speaking, a house price bubble in our paper refers to the component that cannot be explained by the list of macrofinancial variables and institutional factors used in this study. If the list of variables is incomplete, then the bubble may mistakenly include a fundamental-related component. By contrast, if the estimates of house price fundamentals are not efficient and include a non-fundamental-driven component, they will introduce errors in the decomposition analysis. Certain aspects of the methodology are designed specifically to minimize the relevance of these concerns. As a reiteration, we use trend series of mortgage credit-to-GDP ratios and equity prices in examining the determination of house price fundamentals. Moreover, our analysis is constrained by the fact that the sample time series are not very long. To overcome this shortcoming, we adopt panel regressions (whenever data are available) to estimate house price fundamentals, in the hope of revealing the general relationship between house price fundamentals and macrofinancial factors. Nevertheless, these refinements are by no means perfect.

In addition, the above empirical methodology also provides another complementary evidence on the characteristics of house price cycles. If $\alpha \geq 1$ or $\beta \leq 0$, house prices are on a divergent path and their movement cannot be sustainable. Such evidence, although not directly related to the bubble component analysis, can shed light on irrational developments in the housing markets under review.

4. Empirical Findings

The empirical results consist of two parts: the characteristics of house price dynamics and the analysis on house price overvaluation and its bubble component.

To contextualize the findings, table 2 summarizes and compares the developments of housing markets in the nine Asia-Pacific economies. Culturally, there is a general trend towards encouraging homeownership in Asia during the period under review. The property sector is normally dominated by a few major developers. The banking system, alongside the government housing finance system, plays an important role in meeting the demand for housing in most sample economies. The national housing markets share certain similarities (e.g., the prevalent use of floating-rate mortgage contracts) but there exist important differences as well.

4.1 Characterizing House Price Dynamics

To investigate the characteristics of house price dynamics, we follow the Capozza et al. (2002) approach described in section 3.2. We run three sets of regressions, which are described below sequentially. The second regression is used as the benchmark for the bubble analysis discussed in section 4.2. The emphasis of analysis is based on the first and third steps of each regression, i.e., the determination of long-run fundamentals and endogenous adjustment in short-term dynamics.

The first regression relies on a panel data technique to estimate both the determinants of fundamental house prices and the short-run dynamics, with the results reported in table 3A and 3B, respectively. The regression attempts to capture the common picture, if any, of house price cycles for the nine economies during the sample period, i.e., 1993–2006.

Table 2. House Market Conditions in Selected Asia-Pacific Economies

Country	Mortgage Credit			Government Housing Finance Corporation	Homeownership Rates[a]
	LTV Ratio	Mortgage Rate	Loan Term		
Australia	60–70	Variable	25	—	72.0 (2002–04)
China	80	Variable	10–15 ($\leq$ 30)	HPF	59.0 (2000)
Hong Kong	70	Variable	20	HKMC	57.0 (2004)
Korea	70	Variable	3–20	KHFC	56.0 (2000)
Malaysia	80	Variable	30	Cagamas	85.0 (1998)
New Zealand	80–85	Variable	25–30	—	68.0 (2002–04)
Philippines	70	Variable	10–20	HDMF	71.1 (2000)
Singapore	80	Variable	30–35	HDB	92.0 (2005)
Thailand	80	Variable	10–20 ($\leq$ 30)	GHB	82.4 (2005)

[a]Various survey years reported in Cruz (2006) for Southeast Asian and East Asian countries and Ellis (2006) for Australia and New Zealand.
Sources: Global Property Guide (2007); Zhu (2006); national sources.

Table 3. Panel Regression Results

A. Determinants of House Price Fundamentals (Dependent Variable: Log of Real House Prices)		
Variables	**Coefficient**	**t-statistics**
Real GDP	0.36	2.0
Real Mortgage Rate	−0.033	6.4
Mort/GDP Trend	0.37	4.6
Land Supply Index	0.078	4.1
Real Effective Exchange Rate	0.55	3.8
EPI Trend	−0.22	3.6
Institutional Factor (IF)	0.14	3.4
Adjusted R^2	0.55	
B. Short-Run House Price Dynamics (Dependent Variable: Real House Price Growth)		
	Coefficient	**t-value**
Persistence Parameter (α)	0.24	5.1
Mean-Reversion Parameter (β)	0.22	7.8
Contemporaneous Adjustment Parameter (γ)	0.30	5.6
α^* (Change in Land Supply Index)	−0.42	3.9
α^* (Change in Construction Cost)	−10.95	2.9
α^* Institutional Factor	0.37	6.9
β^* (Change in Mortgage Rate)	0.14	4.4
β^* (Change in Land Supply Index)	−4.67	2.4
β^* Institutional Factor	−0.12	4.3
Adjusted R^2	0.36	

Notes: This table shows the regression results on the long-term determinants of house price fundamentals and short-term house price dynamics. Both regressions adopt the panel data regressions with fixed effects. "Mort/GDP Trend" and "EPI Trend" refer to the HP-filtered trend series of mortgage credit/GDP ratios and equity price indices, respectively. The institutional factor (IF) refers to the first principal component of four institutional variables: BFI, FFI, CI, and RPI as defined in table 1. In panel A, all variables (except for "Real Mortgage Rate" and "Mort/GDP Trend") are in logs. To avoid simultaneity bias, regressors are instrumented with own lags. Panel unit-root tests on the residuals reject null of unit-root process. Moreover, panel B uses the model as specified in equation (3).

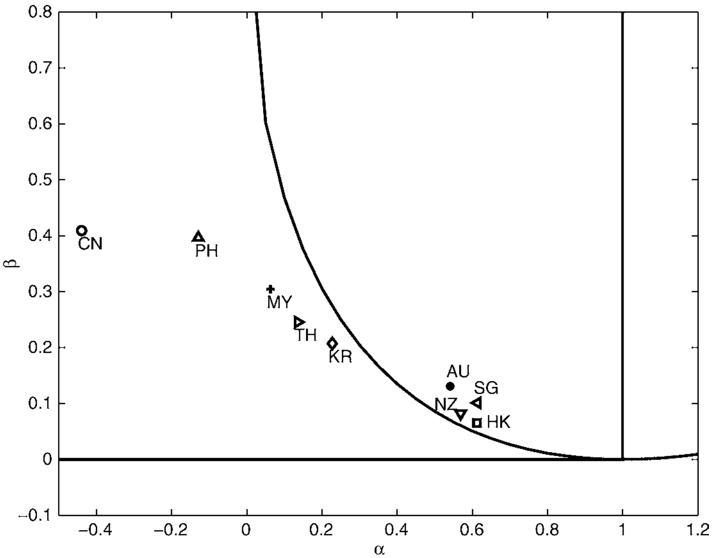
In the first stage, the determination of house price fundamentals yields results that are largely consistent with the theoretical predictions (table 3A). First, higher income, prospects of higher capital gains from real effective exchange rate appreciation, and greater credit availability (mortgage credit-to-GDP ratios) are associated with increases in house prices in Asia-Pacific economies. Second, increases in real mortgage rates have a dampening effect on house prices by raising the cost of housing purchase, but the magnitude is relatively small. Third, the coefficient of the land supply index is positive, which contradicts the theoretical prediction that increases in land supply have a dampening effect on house prices in the long run. This may, however, reflect a linkage in the opposite direction, i.e., higher house prices provide an incentive for developers to build up new residential property projects. Fourth, the institutional factor has a positive and significant effect, suggesting that the improvement in business environment (higher transparency in business regulations, lower corruption, a higher degree of financial sector development) facilitates greater transactions and exerts a positive impact on house prices. Lastly, equity prices are negatively related to house prices, suggesting that the substitution effect dominates the wealth effect during the sample period.

The results on the short-term dynamics, which embed the predicted house price fundamentals from stage 1 regression and the interactive terms to characterize the serial correlation and mean-reversion coefficients, are reported in table 3B. Figure 4 summarizes the characteristics of house price dynamics in each of the nine economies, by plotting the average persistence and mean-reversion coefficients using the time average of country-specific variables. They are separated into two groups. Australia, Hong Kong, New Zealand, and Singapore typically observe damped oscillation of house prices if the fundamental values change, whereas China, Korea, Malaysia, the Philippines, and Thailand observe a convergence to the fundamental values.[15]

The distinction in national house price dynamics as reflected in the persistence and mean-reversion coefficients can be explained by

[15] No country is in the zone of unstable divergence or amplified oscillation.

Figure 4. House Price Dynamics: Panel Regression Results



Note: The results are based on a panel regression on the determinants of house price fundamentals and a panel regression on the short-run dynamics (with fixed effects in both regressions).

differences in market arrangements, such as the supply elasticity embodied in the land supply index and real construction cost, mortgage rate adjustability, and the institutional factor (table 3B). The land supply index and real construction cost both have a negative interactive effect on the persistence coefficient. This means that increases in the land supply index and the construction cost index (which proxy for higher supply elasticity) temper the magnitude of house price cycles. As such, persistence of house prices is moderated in the process.

In addition, changes in mortgage rates have a positive interactive effect on the mean-reversion coefficient. This is probably because larger changes in mortgage rates may indicate a more liberalized mortgage market or higher flexibility in mortgage rate adjustment, thus reflecting faster speed of convergence to the equilibrium price (a higher mean-reversion coefficient).

Lastly, the institutional factor has a positive interactive effect on the persistence parameter and a negative interactive effect on the mean-reversion parameter. That is, a higher score in the institutional factor tends to increase the amplitude but lower the frequency of house price cycles. As the institutional factors become more favorable to growth, the price discovery function strengthens and the incentive to participate in the market improves. Thus, one would expect greater demand for housing. However, the housing market is unique because of inherent supply lags in the housing market. The processes of searching for a house and completing the transaction between sellers and buyers take longer than those in any other asset markets. Improved institutional environment, thus, causes house price growth to persist over a longer period of time.[16]

It is commonly known that housing is a local product and the determination of house prices tends to be market specific. To reflect this, we conduct a second regression that uses country-specific predicted fundamental values.[17] The results are reported in table 4.

Table 4A confirms that the driving factors of house price fundamentals are market specific; therefore, it is important to incorporate this heterogeneity in the analysis. Nevertheless, the results of short-run house price dynamics are quite robust, as reported in table 4B. The sign and significance of all coefficients, including the interactive terms, are retained. The cross-country differences in terms of the average persistence and mean-reversion coefficients do not change in the regression that uses country-specific fundamentals (figure 5 versus figure 4).

It is also worth reporting that when running the country-specific regressions on long-run fundamentals (reported in table 4A), the augmented Dickey-Fuller test confirms the stationarity of the

[16]This is shown in the plots of the persistence and mean-reversion parameters in figure 4. Along the same line, Zhu (2006) also suggests that house prices in Hong Kong and Singapore, the two economies with the most flexible housing finance arrangement, are much more volatile than those in a number of other Asian economies.

[17]For those countries with city-level data, the country-specific analysis is based on a panel regression within the country. This is to overcome major data limitations, i.e., the short time series and the quality difference in computing house price indices.

Table 4. Panel Regression Based on Country-Specific Models of House Price Fundamentals

A. Determinants of House Price Fundamentals (Dependent Variable: Log of Real House Prices)									
	AU (OLS)	**CN (Panel)**	**HK (OLS)**	**KR (Panel)**	**MY (Panel)**	**NZ (OLS)**	**PH (Panel)**	**SG (OLS)**	**TH (OLS)**
Constant	4.21	4.07	−8.39	5.60	2.42	−4.01	3.50	−4.82	4.76
Real GDP	0.38	0.18	0.022	—	0.41	0.56	—	—	−0.18
Mort/GDP Trend	0.92	—	—	—	0.24	—	1.08	−0.031	0.98
Real Mortgage Rate	—	—	−0.051	−0.034	0.010	—	0.017	—	—
Land Supply Index	0.23	−3.51	—	−0.16	—	—	0.16	—	0.074
Real Construction Cost	—	0.25	—	—	—	—	—	0.78	—
REER	—	—	0.99	—	—	0.32	—	1.30	—
Equity Price Trend	−0.84	—	2.22	—	—	0.98	—	—	—
Adjusted R^2	0.99	0.77	0.87	0.51	0.82	0.98	0.41	0.65	0.88

Notes: The results are based on country-specific regression results, by either using national-level data (OLS) or pooled city-level and national-level data (panel). All equations are cointegrated at 1 percent level of significance except for China. Regressors are expressed in logs except for mortgage credit-to-GDP ratio and real mortgage rate. A general-to-specific approach is adopted so that the final model specification in each economy only includes those explanatory variables with statistically significant coefficients. To avoid simultaneity bias, regressors are instrumented with own lags.

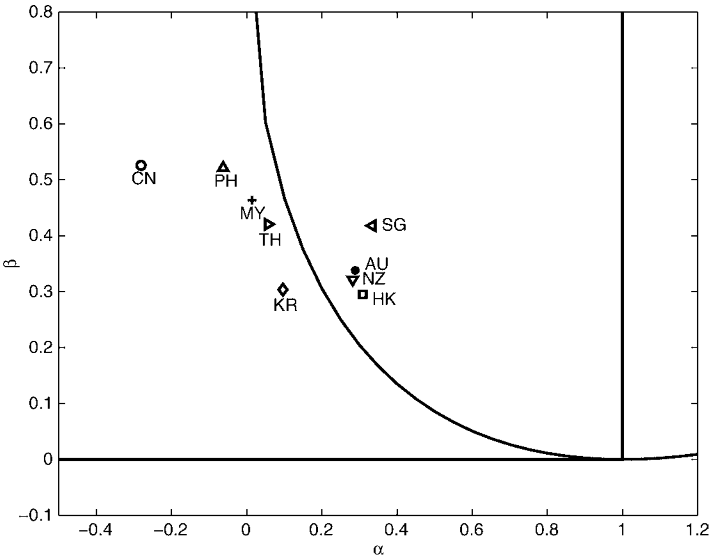
(continued)

Table 4. (Continued)

B. Short-Run House Price Dynamics (Dependent Variable: Real House Price Growth)		
	Coefficient	**t-value**
Persistence Parameter (α)	0.12	2.5
Mean-Reversion Parameter (β)	0.26	2.6
Contemporaneous Adjustment Parameter (γ)	0.68	10.9
α^* (Change in Land Supply Index)	−0.46	3.7
α^* (Change in Construction Cost)	−10.8	3.1
α^* Institutional Factor	0.20	4.1
β^* (Mortgage Rate)	0.018	1.8
β^* (Change in Land Supply Index)	−0.45	3.8
β^* Institutional Factor	−0.085	1.8
Adjusted R^2	0.51	

Notes: The regression is based on a panel data of the nine sample economies (with fixed effects). House price fundamentals are determined by the country-specific regression results as reported in table 4A. The institutional factor refers to the first principal component of four index variables: BFI, FFI, CI, and RPI as defined in table 1.

Figure 5. House Price Dynamics: Baseline Results



Note: The results are based on country-specific regressions on the determinants of house price fundamentals and a panel regression (with fixed effects) on the short-run dynamics.

residual terms in all markets except in China.[18] The evidence of cointegrating relationship justifies the validity of the error-correction specification in analyzing short-run dynamics.[19]

The third regression, instead, employs city-level data. As in the second regression, the fundamentals are determined on the basis of country-specific or market-specific analysis. The panel regression results of the endogenous adjustment equation, as reported in table 5, show significant and positive interactive effects of a dummy variable that defines the most important or high-end market segments in each economy, implying greater volatility in house price

[18]Similarly, in the first regression as described above, a panel unit-root test and the Kao residual cointegration test provide supporting evidence on the cointegration relationship among the variables included in table 3A.

[19]This is in contrast to the results in Gallin (2006), who reports little evidence of cointegration relationship in the U.S. market. The longer list of explanatory variables used in this study may contribute to the different findings.

Table 5. City-Level Endogenous Adjustment Panel Regression Results

	Coefficient	t-value
Persistence Parameter (α)	−0.14	5.7
Mean-Reversion Parameter (β)	0.54	11.8
Contemporaneous Adjustment Parameter (γ)	0.91	29.4
α^* (Change in Land Supply Index)	0.068	2.4
α^* (Dummy for Major Cities)	0.22	2.4
β^* (Change in Mortgage Rate)	0.084	2.6
β^* Institutional Factor	−0.086	3.0
β^* (Dummy for Major Cities)	0.084	2.6
Adjusted R^2	0.32	

Notes: The regression is based on a panel data of thirty-two cities (markets) in seven Asia-Pacific economies (Australia and New Zealand excluded), using the panel regression with fixed effects. House price fundamentals are determined by the country-specific panel regressions or market-specific regressions, which are not reported here. The institutional factor refers to the first principal component of four index variables: BRI, FFI, CI, and RPI as defined in table 1. The dummy for major cities (markets) equals one for the following cities (markets): Kuala Lumpur luxury, Bangkok luxury, Manila luxury, Hong Kong luxury, Singapore private, Beijing, Shanghai, and Seoul.

movements.[20] In addition, the negative (positive) interactive effect between the institutional factor (mortgage rate adjustment) and the mean-reversion parameter remains robust. However, the interactive effects of supply and construction cost indices are washed out.

The results suggest that the high-end markets or the leading markets are more likely to be associated with lower response of supply to market demand, which causes them to be more likely to face a higher volatility of house price movements. The low supply elasticity in these markets could be attributed to limited supply as well as high volatility in housing demand. The demand for new housing or house improvement tends to increase the most in the largest

[20]It equals to one for high-end markets (in Bangkok, Hong Kong, Kuala Lumpur, and Manila), the Singapore private housing market, and major commercial cities in the country (Beijing and Shanghai in China and Seoul in Korea).

cities during the urbanization process, and demand for investment purpose is often the most volatile in high-end markets.

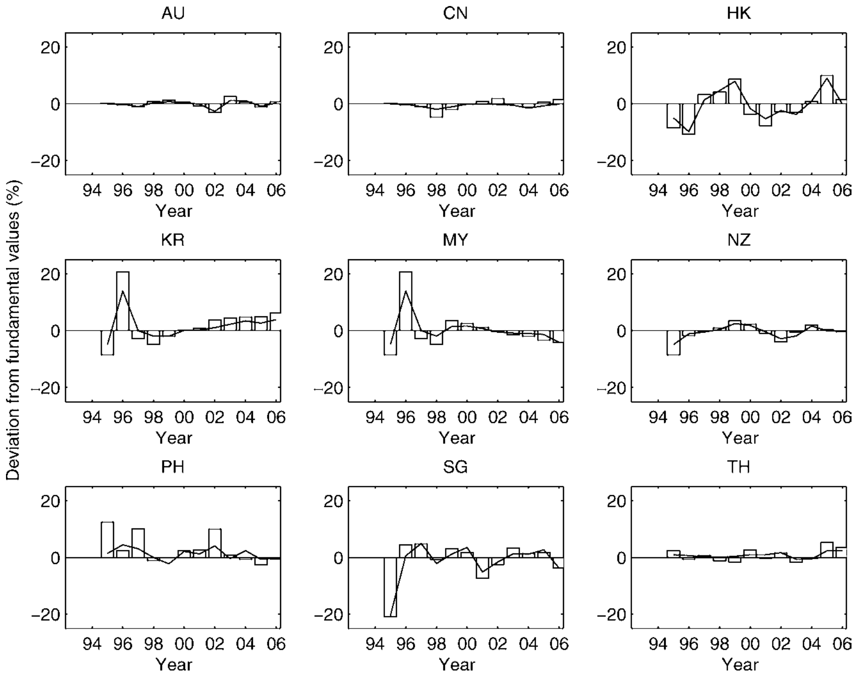
4.2 Detecting Housing Bubbles

Following the methodology described in section 3.3, we try to address the question of whether house prices in selected Asia-Pacific economies are overvalued and, if so, whether there is evidence of some bubble being formed in this region.

The analysis is based on the second regression described above, which treats the determination of house price fundamentals as country specific and relies on a panel data regression to analyze the patterns of short-run dynamics. In figure 6, we first plot the deviation of house prices from predicted fundamentals, represented in bars. At the national level, the evidence of house price overvaluation in recent years is rather weak. Except for Hong Kong (where the house price was 10 precent higher than predicted fundamentals in year 2005), the deviation of house prices from fundamental values is quite small. The result contrasts sharply with results before the Asian crisis, where house prices are about 20 percent higher than their fundamental values in Korea and Malaysia. It appears that the recent strong house price growth (e.g., in Australia, China, Hong Kong, and Korea; see figure 1) is mainly attributable to strong macroeconomic fundamentals.

When the cyclical component, depicted by lines in figure 6, is plotted against total house price overvaluation, the evidence of a house price bubble is even weaker. In Hong Kong, the modest house price overvaluation in year 2005 was mainly driven by the cyclical component, i.e., intrinsic house price adjustment due to house price frictions and other market factors. Only in Korea and Thailand is the bubble component positive, but at very low levels. Again, this contrasts with the findings before the Asian financial crisis, when the bubble component explains 7 percentage points of house price overvaluation in Korea and Malaysia and a double-digit bubble component in the Philippines. Therefore, a general conclusion is that, at least at the national level, there is little evidence of substantial house price overvaluation or house price bubbles in the selected economies in recent years.

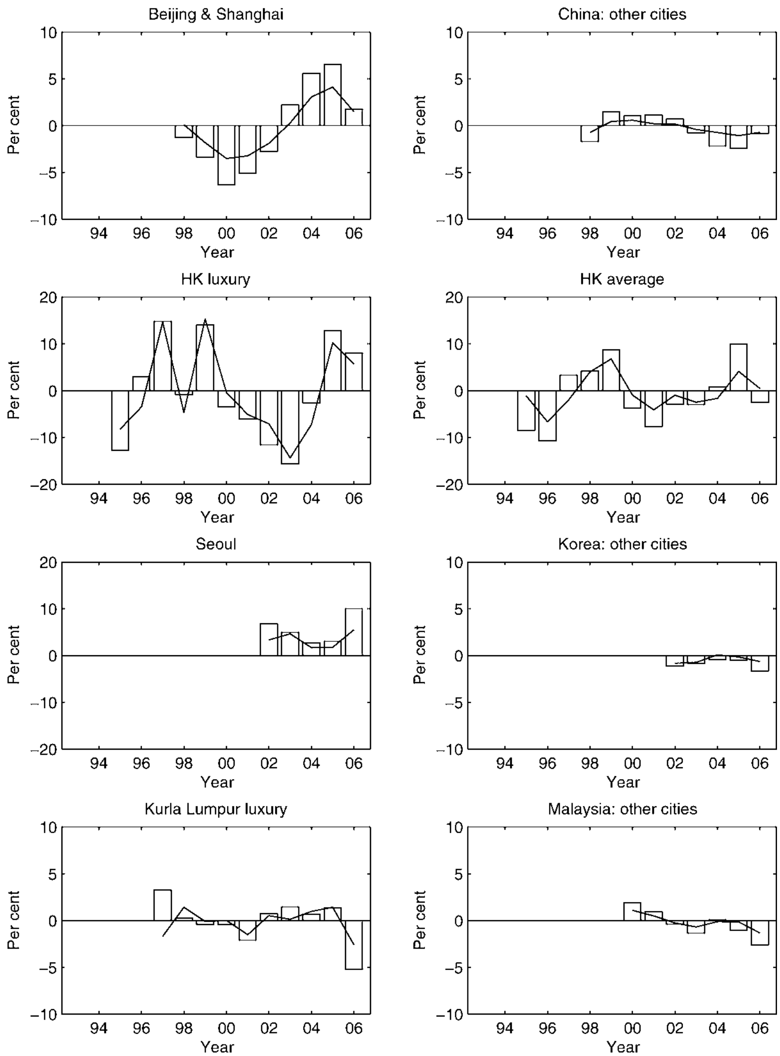
Figure 6. Deviation of Country-Level House Prices from Fundamental Values



Notes: The bars represent the average annual deviation of observed house prices from their fundamental values, and the lines represent the cyclical component of this average annual deviation, i.e., the component that can be explained by the short-term dynamics. The results are based on country-specific regressions on the determinants of house price fundamentals and a panel regression (with fixed effects) on the short-term dynamics (see table 4).

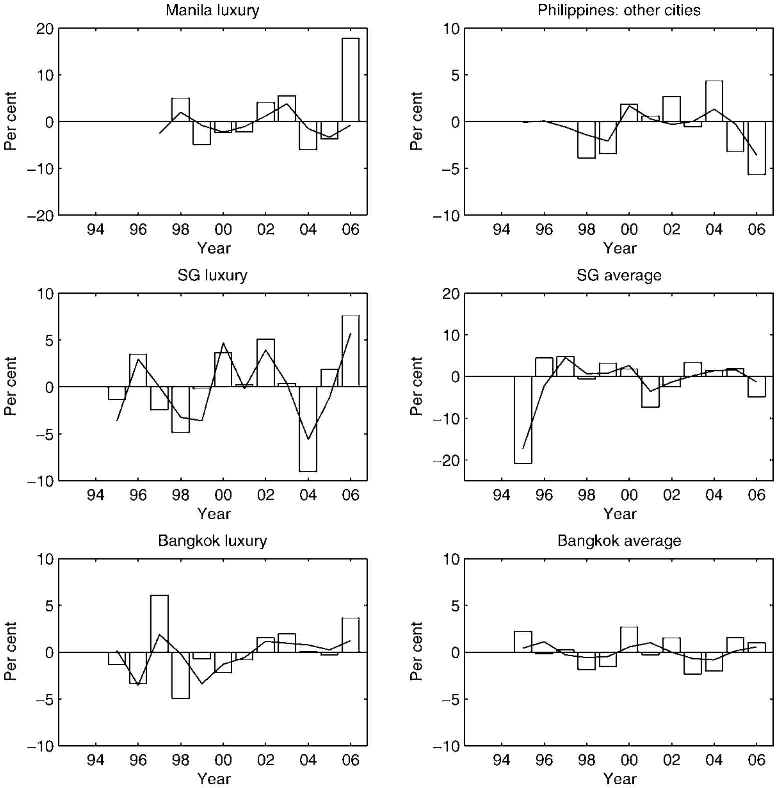
The analysis also extends to city-level (or market-level) house price dynamics. Figure 7 plots, in each economy, the house price deviation from fundamentals in the high-end market (or a leading market) versus the average market. There are two interesting findings. First, except for Malaysia, a more remarkable overvaluation has been detected in the leading market compared with the other markets in the current run-up of house prices. In other words, the house price overvaluation that is observed at the national level comes mainly from the leading market segment. Moreover, over the

Figure 7. Deviation of City-Level House Prices from Their Fundamentals



(continued)

Figure 7. (Continued)



Notes: The bars represent the average annual deviation of observed house prices from their fundamental values, and the lines represent the cyclical component of this average annual deviation, i.e., the component that can be explained by the short-term dynamics. The results are based on a city-level analysis. In China, "other cities" refers to the average of Chongqing, Guangzhou, Shenzhen, and Tianjin. In Korea, "other cities" refers to the average of Busan, Daegu, Daejon, Gwangju, Incheon, and Ulsan. In Malaysia, "other cities" refers to the average of Johor, Kuala Lumpur average market, Pahang, Perak, and Pinang. In the Philippines, "other cities" refers to the average of Caloocan, Makati, Manila average market, Pasay, Pasig, and Quezon.

whole sample period, house prices in the leading market are more likely to deviate substantially from their fundamental values. These results are consistent with the conventional view that the leading market is more volatile than the average market. Second, the breakdown analysis suggests that speculative housing bubbles may exist at particular market segments—for instance, the luxury market in Manila and to a lesser degree in Bangkok, Seoul, Beijing, and Shanghai. From a policy perspective, it is important for policymakers to implement market-specific diagnoses and to find the right policy instruments that can ideally distinguish between cyclical and bubble components.

5. Conclusion

The study documents evidence of serial correlation and mean reversion in nine Asia-Pacific economies and analyzes the patterns of house price dynamics in relation to local institutional features. Notwithstanding the nuances in each market, the regression results validate the hypothesis that the run-up in house prices up to 2006 reflects mainly an adjustment to more buoyant fundamentals rather than speculative housing bubbles. Looking back, it appears that property market developments in Asia and the Pacific were in line with our assessment of house price risk. Despite the spillover effect that hit the real economy, housing markets have only experienced mild adjustment in most Asia-Pacific economies without causing damage to the banking system.

Despite the relatively benign housing market environment in Asia, it remains crucial for regulators to understand the potential risks embedded in the evolving housing market structure. Whereas our study tries to investigate the determination of house price dynamics and evidence of house price bubbles, the answers are far from complete. Further exploration calls for improvement in data compilation and a better understanding of the mechanism of house price determination. For most of Asia, there appears to be a pressing need to improve the quality and timely availability of house price data if these are to aid in better analysis for policy decision-making purpose. Moreover, national average house prices mask the volatility in house price movements in leading cities/markets. Therefore, reliable information on the city level or across market segments is crucial to the understanding of possible local/market segment bubbles.

Appendix.

Table 6. House Prices: Definitions and Data Sources

Country	Series Definition	Sources	Remarks
Australia	Residential property price index	National source	Weighted average of eight capital cities in Australia, namely Sydney, Melbourne, Brisbane, Adelaide, Perth, Hobart, Darwin, and Canberra.
China	Property price index (both residential and commercial)	CEIC	Same source: city-level information is also available. Beijing, Chongqing, Guangzhou, Shanghai, Shenzhen, and Tianjin are included in this study.
Hong Kong	(i) Residential property price index (repeat sales); (ii) Capital value of luxury residential property	(i) CEIC; (ii) Jones Lang LaSalle (JLL)	(i) A composite index for all classes of private domestic, the most common official figures for property price measurement; (ii) Capital value for a prime-quality residential property in the best location.
Korea	Residential overall house price index (including detached house and apartment prices)	CEIC	Same source: city-level information is also available. Busan, Daegu, Daejon, Gwangju, Incheon, Seoul, and Ulsan are included in this study.

(continued)

Table 6. (Continued)

Country	Series Definition	Sources	Remarks
Malaysia	(i) Residential house price index; (ii) Capital value of luxury residential property in Kuala Lumpur	(i) National source; (ii) CEIC	(i) Nationwide house price index is from national source. City-level/state-level residential house prices are from CEIC, using hedonic method. Johor, Kuala Lumpur, Pahang, Perak, and Pinang are included in this study; (ii) Capital value for a prime-quality residential property in the best location in Kuala Lumpur.
New Zealand	Residential property price index	National source	Total New Zealand index is from current valuations of the relevant local authorities. These current valuations are used to calculate the average valuation and the price index in each quarter.
Philippines	(i) Residential property price index; (ii) Capital value of luxury residential property	(i) NSO; (ii) JLL/Colliers International	(i) Constructed from available value of building permits and corresponding floor area. City-level information is available for the national capital region (represented by Caloocan, Makati, Manila, Pasig, Pasay, and Quezon; 2000 = 100); (ii) Capital value for a prime-quality residential property in the best location in Manila, Makati, and Ortigas Center.

(continued)

Table 6. (Continued)

Country	Series Definition	Sources	Remarks
Singapore	(i) Residential property price index; (ii) Capital value of luxury residential property	(i) CEIC; (ii) JLL	(i) HDB resale price index, which is calculated from the quarterly average resale price of HDB flats by date of registration; (ii) Capital value for a prime-quality residential property in the best location.
Thailand	(i) Residential property price index; (ii) Capital value of luxury residential property in Bangkok	(i) BOT; (ii) JLL	(i) Bangkok and vicinities, single detached house and town house, including land (hedonic method); (ii) Capital value for a prime-quality residential property in the best location in Bangkok.

References

Abraham, J., and P. Hendershott. 1996. "Bubbles in Metropolitan Housing Markets." *Journal of Housing Research* 7 (2): 191–207.

Aoki, K., J. Proudman, and G. Vlieghe. 2004. "House Prices, Consumption and Monetary Policy: A Financial Accelerator Approach." *Journal of Financial Intermediation* 13 (4): 414–35.

Ayuso, J., and F. Restoy. 2006. "House Prices and Rents: An Equilibrium Asset Pricing Approach." *Journal of Empirical Finance* 13 (3): 371–88.

Bernanke, B., and M. Gertler. 1995. "Inside the Black Box: The Credit Channel of Monetary Policy Transmission." *Journal of Economic Perspectives* 9 (4): 27–48.

Bernanke, B., M. Gertler, and S. Gilchrist. 1996. "The Financial Accelerator and the Flight to Quality." *Review of Economics and Statistics* 78 (1): 1–15.

Borio, C., and P. McGuire. 2004. "Twin Peaks in Equity and Housing Prices?" *BIS Quarterly Review* (March): 79–93.

Brunnermeier, M., and C. Julliard. 2008. "Money Illusion and Housing Frenzies." *Review of Financial Studies* 21 (1): 135–80.

Campbell, J., and J. Cocco. 2007. "How Do House Prices Affect Consumption? Evidence from Micro Data." *Journal of Monetary Economics* 54 (3): 591–621.

Capozza, D., P. Hendershott, C. Mack, and C. Mayer. 2002. "Determinants of Real House Price Dynamics." NBER Working Paper No. 9262 (October).

Case, K., J. Quigley, and R. Shiller. 2005. "Comparing Wealth Effects: The Stock Market versus the Housing Market." Institute of Business and Economic Research Working Paper No. W01-004.

Chen, N. K. 2001. "Bank Net Worth, Asset Prices and Economic Activity." *Journal of Monetary Economics* 48 (2): 415–36.

Cocco, J. 2005. "Portfolio Choice in the Presence of Housing." *Review of Financial Studies* 18 (2): 535–67.

Committee on the Global Financial System. 2006. "Housing Finance in the Global Financial Market." CGFS Working Group Report No 26. Basel, Switzerland: Bank for International Settlements.

Cruz, P. C. 2006. "Inflation, Politics and Government Intervention: How Asian Residential Markets Have Suffered." Global Property

Guide (website and online newsletter) (November). Available at http://www.globalpropertyguide.com.

Davis, E. P., and H. Zhu. 2004. "Bank Lending and Commercial Property Cycles: Some Cross-Country Evidence." BIS Working Paper No. 150.

Diamond, D., and M. Lea. 1992. "Housing Finance in Developed Countries: An International Comparison of Efficiency." *Journal of Housing Research* 3 (1): 1–260.

Égert, B., and D. Mihaljek. 2007. "Determinants of House Prices in Central and Eastern Europe." BIS Working Paper No. 236.

Ellis, L. 2006. "Housing and Housing Finance: The View from Australia and Beyond." RBA Research Discussion Paper No. 2006-12.

Estrella, A. 2002. "Securitization and The Efficacy of Monetary Policy." *Economic Policy Review* (Federal Reserve Bank of New York) 8 (1): 243–55.

European Central Bank. 2003. "Structural Factors in the EU Housing Markets." Report (March).

Gallin, J. 2006. "The Long-Run Relationship between House Prices and Income: Evidence from Local Housing Markets." *Real Estate Economics* 34 (3): 417–38.

Gan, J. 2007. "Collateral, Debt Capacity, and Corporate Investment: Evidence from a Natural Experiment." *Journal of Financial Economics* 85 (3): 709–34.

Garmaise, M., and T. Moskowitz. 2004. "Confronting Information Asymmetries: Evidence from Real Estate Markets." *Review of Financial Studies* 17 (2): 405–37.

Gerlach, S., and W. Peng. 2005. "Bank Lending and Property Prices in Hong Kong." *Journal of Banking and Finance* 29 (2): 461–81.

Girouard, N., and S. Blöndal. 2001. "House Prices and Economic Activity." OECD Economics Department Working Paper No. 279.

Global Property Guide. 2007. Website and online newsletter. Available at www.globalpropertyguide.com.

Green, R., S. Malpezzi, and S. Mayo. 2005. "Metropolitan-Specific Estimates of the Price Elasticity of Supply of Housing, and Their Sources." *American Economic Review Papers and Proceedings* 95 (2): 334–39.

Hegedüs, J., and R. Struyk, eds. 2005. *Housing Finance: New and Old Models in Central Europe, Russia, and Kazakhstan.* Budapest: Local Government and Public Service Reform Initiative, Open Society Institute-Budapest.

Helbling, T., and M. Terrones. 2003. "When Bubbles Burst." In *World Economic Outlook*, 61–94. International Monetary Fund.

Herring, R., and S. Wachter. 1999. "Real Estate Booms and Banking Busts: An International Perspective." Center for Financial Institutions Working Paper No. 99-27, University of Pennsylvania.

Hilbers, P., Q. Lei, and L. Zacho. 2001. "Real Estate Market Developments and Financial Sector Soundness." IMF Working Paper No. 01/129.

Himmelberg, C., C. Mayer, and T. Sinai. 2005. "Assessing High House Prices: Bubbles, Fundamentals and Misperceptions." *Journal of Economic Perspectives* 19 (4): 67–92.

Hofmann, B. 2004. "Bank Lending and Property Prices: Some International Evidence." Working Paper.

Holly, S., M. H. Pesaran, and T. Yamagata. 2010. "A Spatio-Temporal Model of House Prices in the USA." *Journal of Econometrics* 158 (1): 160–73.

Kalra, S., D. Mihaljek, and C. Duenwald. 2000. "Property Prices and Speculative Bubbles: Evidence from Hong Kong SAR." IMF Working Paper No. 00/2.

Kiyotaki, N., and J. Moore. 1997. "Credit Cycles." *Journal of Political Economy* 105 (2): 211–48.

Leung, C., and N.-K. Chen. 2006. "Intrinsic Cycles of Land Price: A Simple Model." *Journal of Real Estate Research* 28 (3): 293–320.

McCarthy, J., and R. Peach. 2002. "Monetary Policy Transmission to Residential Investment." *Economic Policy Review* (Federal Reserve Bank of New York) 8 (1): 139–58.

Organisation for Economic Co-operation and Development. 2005. *Housing Finance Market in Transition Economies: Trends and Challenges.* OECD.

Peek, J., and J. Wilcox. 2006. "Housing, Credit Constraints, and Macro Stability: The Secondary Mortgage Market and Reduced Cyclicality of Residential Investment." *American Economic Review Papers and Proceedings* 96 (2): 135–40.

Sutton, G. 2002. "Explaining Changes in House Prices." *BIS Quarterly Review* (September): 46–55.

Tsatsaronis, K., and H. Zhu. 2004. "What Drives Housing Price Dynamics: Cross-Country Evidence." *BIS Quarterly Review* (March): 65–78.

Warnock, V., and F. Warnock. 2008. "Markets and Housing Finance." *Journal of Housing Economics* 17 (3): 239–51.

Wheaton, W. 1999. "Real Estate 'Cycles': Some Fundamentals." *Real Estate Economics* 27 (2): 209–30.

Yao, R., and H. Zhang. 2005. "Optimal Consumption and Portfolio Choices with Risky Housing and Borrowing Constraints." *Review of Financial Studies* 18 (1): 197–239.

Zhu, H. 2006. "The Structure of Housing Finance Markets and House Prices in Asia." *BIS Quarterly Review* (December): 55–69.