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Financial Market Implications of Monetary Policy Coincidences: Evidence from the UK and Euro Area Government-Bond Markets

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Abstract

Relatively little is known about the financial market impact of international monetary surprises arising on the same trading day. This paper estimates a suite of multi-security factor models, which captures international monetary surprise effects on UK and Euro Area government-bond markets over the period 1999-2014. In doing so, we shed light on the relative importance of coinciding, non-coinciding monetary surprises and non-monetary surprises across the yield curve. We find some support for the 'enrich-thy-neighbour' hypothesis of international monetary surprises, while our findings suggest that monetary policy cooperation during crises produces financial market effects that go above and beyond conventional policy.

JEL Classifications: E4, E5, F3, G1

Keywords: international; monetary policy; financial markets; factor model

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Abstract

Relatively little is known about the financial market impact of international monetary surprises arising on the same trading day. This paper estimates a suite of multi-security factor models, which captures international monetary surprise effects on UK and Euro Area government-bond markets over the period 1999-2014. In doing so, we shed light on the relative importance of coinciding, non-coinciding monetary surprises and non-monetary surprises across the yield curve. We find some support for the 'enrich-thy-neighbour' hypothesis of international monetary surprises, while our findings suggest that monetary policy cooperation during crises produces financial market effects that go above and beyond conventional policy.

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1. Introduction

Major central banks responded to the ‘great recession’ by cutting the base rate sharply; between October 2008 and the end of that same year the Bank of England (BoE), the European Central Bank (ECB), although the ECB, after increasing its ‘official’ rate as late as mid-2008, it started reducing it in May 2009, and the US Federal Reserve (FED) cut key interest rates by 300, 175 and 175 basis-points respectively.[1] The coinciding nature of monetary policy responses during the turmoil reflects the severe nature of events, which forced policy makers to act immediately and thus simultaneously. Additionally, or for both reasons at the same time, policy makers expected to make a bigger impact on the markets and boost confidence if they are seen to be cooperating (BIS, 2009; Lim et al., 2011; Leeper, 2016). Indeed, there is a growing policy debate about the role for greater international monetary policy cooperation (see, for example, Bernanke, 2008; Blanchard et al., 2013; Saccomanni, 2015; Draghi, 2016; Rajan, 2014, 2016). However, effective monetary policy cooperation requires a common understanding of the direction, size and significance of spillover effects. Yet there is much uncertainty as to whether and how coinciding policy responses would impact on the financial markets and promote expectations of recovery. Against a backdrop of globalisation, crisis and uncertainty, this paper exploits the existence of a number of coinciding BoE and ECB monetary policy announcements, to investigate whether international monetary policy surprises arising on the same trading day have been influential for UK and Euro Area (EA) government-bond markets; and, if so, whether their effects have been any different from other monetary and non-monetary surprises.

Whether or not international monetary surprises have ‘beggar-thy-neighbour’ or ‘enrich-thy-neighbour’ effects is an important and unresolved issue for academics and policy makers, particularly in view of the recent turmoil (Bernanke, 2013). Over more recent years, studies emphasise financial spillover channels, including international portfolio balance effects, which generate capital flows across countries and lower international bond yields (Ammer et al., 2016). However, the extent to which international monetary surprises pass-through to financial markets and promote expectations of economic recovery remains uncertain. For example, the portfolio balance effect has been established only in the case of non-coinciding international monetary surprises, whereas there are expenditure-switching and expenditure-enhancing channels embedded in open-economy models that offer differing conclusions about how a monetary surprise in one country affects economic prospects in another (Taylor, 2013; Blanchard et al., 2015).[2] Therefore, it remains very unclear what the ‘net’ effect of a

[1] The FED had previously cut its federal funds rate sharply from 4.25% to 2% between January and April of 2008 in response to the severe initial impact on US housing markets of the high federal funds rate over the period 2006-2007. While the federal funds rate remained at 2% at the start of October 2008, the BoE’s base interest rate and the ECB’s main financing rate were 5% and 4.25% respectively. More recently, unconventional monetary policies have been introduced, including asset purchasing programmes and, to an extent, forward guidance policies, (although by February 2014 in the case of the BoE and by December 2015 in the FED case, it was announced that no longer their policies would be linked to a particular economic indicator; more general economic conditions would dictate changes in interest rates), with the aim of promoting more broadly financial stability and economic recovery. The ECB has also introduced negative interest rates.

[2] Monetary policy coordination may reflect direct planning of policies across countries, coinciding best responses to shocks or perhaps some combination. Interestingly enough, the theoretical case for monetary

coinciding monetary-easing would be in terms of the financial market impact across the yield curve and more broadly for the economy. Thus, empirical investigation is now an essential next step.

Obtaining reliable estimates of monetary policy's impact on asset prices is an important, but challenging task to accomplish empirically, particularly due to the problem of simultaneity (Rigobon and Sack, 2004; Caporale et al., 2005). Moreover, the task is to identify policy's effects across the yield curve, while accounting for the fact that policy also responds to economic conditions. Other relevant factors may also affect asset prices at the same time as domestic policy, including financial conditions and macroeconomic outlook, as well as other common surprises affecting particular securities. That is, endogeneity makes more difficult the task of identifying monetary policy surprises on financial markets. Studies have typically adopted Kuttner's (2001) 'event-study' (ES) approach to overcome these problems (e.g. Bernanke and Kuttner, 2005; Bredin et al., 2007; Hussain, 2011; Rogers et al., 2014). Monetary surprises are identified using an observable change in a short-maturity yield, subsequent to the central bank policy committee changing the target rate.[3] Alternatively, Sentana and Fiorentini (2001), Rigobon and Sack (2004), Craine and Martin (2008) and Wright (2012) develop heteroskedasticity-based identification techniques (IH) using instrumental variables, which are generally valid under much weaker assumptions.[4]

This paper contributes relative to the literature by investigating empirically the incidence of monetary surprises from multiple sources and not just the home country. This paper's main contribution is to estimate the effect of coinciding monetary surprises on the UK and EA government-bond markets. By convention coinciding-event days have either been dropped from samples or not treated separately. In the latter case, a requirement is that both coinciding and non-coinciding monetary surprises are homoskedastic. However, this paper makes no such assumption and exploits the fact that a number of BoE and ECB monetary policy announcements occur on the same trading day, having effects across the yield curve. The case of the US is also accounted for in some instances, for relevant comparative purposes, but the existence of a sufficiently large number of monetary coincidences in the UK and EA renders

coordination stands at odds against much of the pre-crisis work in this area; for instance, 'New Open Economy' macroeconomic models found payoffs to be fairly small. Taylor (2013) provides a useful survey of the developments in this literature.

[3] The shorter the window, the more effective the ES approach is likely to be. Even so, monetary policy shocks may not be fully absorbed in asset prices during such short windows. Endogeneity problems cannot be entirely resolved since common shocks are likely to persist across shorter horizons. Another issue with the use of higher frequency data is that it may limit the timeframe for analysis.

[4] A growing literature investigates whether monetary policy surprises have international spillover effects (e.g. Andersen et al., 2007; Craine and Martin, 2008; Bredin et al., 2010; Ehrmann et al., 2011; Ammer et al., 2016). More recently, studies examine changes in the transmission of monetary surprises, during the 'great recession', and also the impact of quantitative easing, balance sheet and forward guidance programmes for government and corporate bond markets (e.g. Wright, 2012; Gilchrist and Zakrajsek, 2013; Rogers et al., 2014; Claus et al., 2014; Nyholm, 2016; Haldane et al., 2016). There has also been international analysis of monetary policy spillovers within and across advanced and emerging economies (e.g. Bowman et al., 2015; Fratzscher et al., 2016; Chen et al., 2016; Clark et al., 2016). However, the focus of these studies is on domestic and/or international monetary pass-through on non-coinciding monetary-event days.

these particular bond markets natural cases for analysis.[5] In doing so, this paper estimates a suite of econometric models, which enables comparison of the size and scope of spillover channels affecting international bond yields on different monetary and non-monetary-event days. Our preferred latent factor model incorporates coinciding and non-coinciding monetary surprises, in addition to non-monetary surprises, which significantly impact upon government-bond securities on the same trading day. To that end, we also provide some comparison against event-study approaches, which do not fully address the endogeneity issue despite its relevance in this context. Finally, this paper identifies monetary surprise level and rotation shocks, from which we infer the nature of the financial market impact of coinciding and non-coinciding monetary policy announcements across different maturity bonds. In this respect, our contribution sheds light on underlying theories about the term structure of interest rates.

The empirical strategy is based on estimation of a suite of relevant multi-security models, which accounts for different types of surprises arising from the UK, EA and the US, including up to five different event days: UK-EA coinciding monetary-event days, UK, EA, US non-coinciding monetary-event days and non-monetary-event days. To conduct the analysis, a suite of latent factor models is estimated over the period January 5 1999 to July 22 2014 using Generalised Method of Moments (GMM). An advantage of this paper's preferred IH methodology, which to our knowledge has never been utilised to investigate the financial market implications of monetary coincidences, is that it makes fuller use of non-monetary-event day data and enables inference about the relative importance of various monetary and non-monetary surprises; while crucially it addresses the inherent endogeneity issue.

This paper's findings indicate that monetary policy announcements on coinciding monetary-event days generate additional pass-through at the short-end of the yield curve. We attribute this to significant level and rotation effects, which under our preferred approach are quantitatively and qualitatively very different from the effects arising on non-coinciding monetary-event days. As such, the coinciding monetary policy effect appears to go above and beyond the non-coinciding announcement effect. Overall, our findings are more in line with the 'enrich-thy-neighbour' hypothesis and suggest that the swift and coordinated responses of major central banks must have helped to reduce the severity of the global crisis and promote expectations of a sustained recovery in the UK and EA. Several important policy considerations follow from this paper's findings.

The paper is organised as follows. Section 2 sets out the econometric framework. Section 3 discusses empirical implementation. Section 4 describes the data. Section 5 presents the main empirical results. Section 6 provides discussion. Section 7 summarises and concludes.

[5] Another consideration is that this paper's analysis relies on certain assumptions about the data in relation to heteroskedasticity; in this respect, the case for using the IH methodology appears to be stronger for UK and EA government-bond markets than for the US government-bond market.

2. Econometric framework

2.1. Linear factor model

This paper follows the basic linear factor model setup as detailed in the relevant literature (see, for example, Craine and Martin, 2008). Any given day in the economic calendar must be a coinciding monetary-event day (superscript p), a UK, EA or US non-coinciding monetary-event day (superscript q) or a non-monetary-event day. The model's matrix representation is as follows:

$$Y_t = \Phi^p \mu_t^p + \Lambda Z_t, \qquad t \in E^p \tag{1}$$

$$Y_t = \Phi^q \mu_t^q + \Lambda Z_t, \qquad t \in E^q \tag{2}$$

$$Y_t = \Lambda Z_t, \qquad t \in E^{NE} \tag{3}$$

$Z_t \sim WS(0,1)$ for all t.

$\mu_t^{p,q} \sim WS(0,1)$ for t on monetary-event days, 0 otherwise.

E^p denotes the set and number of coinciding monetary-event days; E^q is the set and number of non-coinciding monetary-event days; E^{NE} is the set and number of non-monetary-event days.

Y_t is a (N x 1) vector of N bond yield changes. Φ^p is a (N x 1) vector of factor loadings corresponding to a scalar coinciding monetary surprise, μ^p_t. Φ^q is a (N x 1) vector of factor loadings corresponding to a scalar non-coinciding monetary surprise, μ^q_t. UK, EA and US non-coinciding monetary surprises are orthogonal. Λ is a (N x (N + 1)) matrix of factor loadings corresponding to a ((N + 1) x 1) vector of non-monetary surprises, Z_t, which contains in row 1 a common surprise, $z_{1,t}$, that affects all bond yields. In the remaining N rows, there exist idiosyncratic surprises, $z_{4\ldots N+3,t}$, which are uncorrelated across securities.

Coinciding and non-coinciding monetary surprises are exogenously determined and occur only on the relevant monetary-event days with constant variance. That is, the variance of non-coinciding monetary surprises is zero on coinciding monetary-event days, while the variance of coinciding monetary surprises is zero on non-coinciding monetary-event days. Non-monetary surprises occur each day with constant variance, but are not identified a priori. Variances of surprises are drawn from a wide-sense (WS) distribution with zero mean and unit variance so that the factor loadings correspond to a one standard deviation surprise on the change in security yields. Moreover, in what follows below, this paper sets out the moment-conditions for monetary and non-monetary-event days.

2.2. Moment-conditions

2.2.1. Monetary-event days

The second moment-conditions for event day classifications are:

$$H_t^p = E[Y_t Y_t'] - \Omega - \Phi^p \Phi^{p\prime}, \qquad t \in E^p \tag{4}$$

$$H_t^q = E[Y_t Y_t'] - \Omega - \Phi^q \Phi^{q\prime}, \qquad t \in E^q \tag{5}$$

$$H_t^{NE} = E[Y_t Y_t'] - \Omega, \qquad t \in NE \tag{6}$$

The first set of moment-conditions in equation (4) is for coinciding monetary-event days. The second set in equation (5) is for non-coinciding UK, EA and US monetary-event days. The final set provided in equation (6) is for non-monetary-event days. $\Lambda\Lambda' = \Omega$ is the covariance matrix corresponding to non-monetary factors, Z_t; $\Phi^p\Phi^{p\prime}$ and $\Phi^q\Phi^{q\prime}$ are covariance matrices corresponding to coinciding- and non-coinciding monetary factors, respectively. These matrices contain the products of the relevant factor loadings of monetary surprises for securities ij.

2.2.2. Non-monetary-event days

This paper includes both common and idiosyncratic latent factors to account plausibly for non-monetary surprises.[6] This is a feasible choice for the implication of equation (6) is that on non-monetary-event days there must be at least as many non-monetary factors as securities, since the rank of the covariance matrix of securities is N, i.e. rank (Ω |t ϵ NE) = rank ($E[Y_tY_t']$ |t ϵ NE). Recall that Λ is a (N x (N + 1)) matrix of coefficients corresponding to a ((N + 1) x 1) vector of non-monetary surprises, Z_t. Next δ and γ are defined as (N x 1) and (N x N) matrices containing common and security specific factor loadings respectively. The relevant moment-conditions are as before, but now Ω, which is an (N x N) matrix, can be decomposed for element ij, where $\omega_{ij} \in \Omega$ reflects this structure for non-monetary factors:

$$\Omega = \delta\delta' + \gamma\gamma' \tag{7}$$

The covariance matrix for common surprises $\delta\delta'$ contains products of the relevant factor loadings. For the idiosyncratic surprise matrix $\gamma\gamma'$, only elements on the principal diagonal are non-zero. However, all off-diagonal elements take zero values, since idiosyncratic surprises are uncorrelated across securities. Summing each corresponding element across the matrices $\delta\delta'$ and $\gamma\gamma'$ yields the covariance matrix Ω for non-monetary-event days.

2.2.3. Summary of theoretical moments

Table 1 summarises the theoretical moments for coinciding monetary, non-coinciding-monetary and non-monetary-event days.

[6] Latent-factor models tend to fit the data well and often better than models based on pre-selected macroeconomic or financial variables to proxy for non-monetary influences (see, for example, Gurkaynak and Wright, 2012).

[TABLE 1]

For non-monetary-event days the sample variance and covariance matrices reflect elements in the covariance matrix Ω as discussed above. For coinciding and non-coinciding monetary-event days, in addition to ω_{ij} the theoretical moments include the product of the relevant monetary factor loadings in Φ^p and Φ^q respectively.

3. Empirical implementation

3.1. Introduction

There has been some debate in the literature over whether latent factor models ought to be preferred over more conventional regression-based approaches; this is for the purpose of identifying international monetary surprise effects on financial markets (see, for example, Borio and Zabai, 2016). The regression model under the ES approach specifies that an observable change in a short maturity rate, say Δr, on monetary policy event days is the monetary-surprise. Therefore, the estimate of a monetary-surprise impact on the yield change may be obtained by running a regression of the change in the short rate on the observable yield changes for other government-bond yields. The intuitive appeal of the conventional ES approach and its ease of implementation have facilitated a high practical application in the aforementioned empirical research.

However, biasedness and inconsistency are likely to arise if the observable change in the short rate also reflects the effects of other surprises arising on the same trading day. Indeed, policy makers are often forced to change the short rate due to events arising simultaneously in the macroeconomic and financial environment. The underlying problem of endogeneity bias may be especially severe in the case of coinciding monetary-surprises, since major events arising at a global level may require swift and sizeable responses from major central banks, such as the BoE, ECB and FED. As such, ES estimates of the monetary-surprise effects on bond-yields are likely to be biased and inconsistent due to a violation of the strict exogeneity assumption embedded within the classical linear regression model. Moreover, the nature of the bias is theoretically ambiguous and, in the case of omitted-variable bias, depends crucially on the correlation between the omitted variable and the included short rate. Additionally, the ES estimates may be biased towards zero when the monetary-surprise is measured with error; this is particularly relevant in the context of monetary coincidences, since there are multiple monetary-surprises arising on the same trading day. Therefore, a crucial assumption under the ES approach is that the change in the short rate is a suitable proxy variable for monetary-surprises; consequently, the errors-in-variables bias ought to be zero over the event window.

These concerns about the ES approach have resulted in the development and application of alternative approaches, particularly IH, which can be used to infer the financial market implications of monetary policy under much weaker assumptions related to the heteroskedasticity of monetary surprises, as discussed above. However, application of the latter remains limited in international economics and finance (see, for example, Rigobon, 2016).

3.2. GMM estimation

A minimum distance estimator is employed to estimate the IH factor model by matching the theoretically implied moments to those of the sample data. For the sample under study, the number of monetary-event days is much smaller than the number of non-monetary-event days. Therefore, the GMM approach of Hansen (1982) is implemented using an equal weighting matrix W^{-1}, which is generally optimal in smaller samples and yields unbiased estimates (Altonji and Segal, 1996).[7] Under this setup the parameter set θ is estimated via minimisation of the following loss function, where W^{-1} is an identity matrix with a dimension reflecting the total number of moments:

$$\min_{\theta} \mathrm{L}(\theta) = \mathrm{G}[\Phi,\Omega]'\mathrm{W}^{-1}\mathrm{G}[\Phi,\Omega] \tag{8}$$

The averaged moment matrices can be obtained by summing up over time each of the elements in H matrices for coinciding, non-coinciding, and non-monetary-event days; and dividing through by the number of days for each classification – these are denoted without time subscripts H^p, H^q, H^{NE}. For each event day classification there are N(N + 1)/2 unique elements, which can be stacked into vectors and combined to form vector G that contains averages of moments across p, q and NE.[8]

$$\mathrm{G} = \begin{bmatrix} \mathrm{vech}(\mathrm{H}^{p}) \\ \mathrm{vech}(\mathrm{H}^{q}) \\ \mathrm{vech}(\mathrm{H}^{NE}) \end{bmatrix} \tag{9}$$

The latent-factor approach set out above can also be related to the conventional IH 'differencing' approach of Rigobon and Sack (2004), based on the difference of moments on monetary and non-monetary-event days. Applying this differencing approach to equations (6)-(8) G can be restated as G^{RS}:

[7] Parameter estimates using the inverse weighting matrix are usually biased downwards in absolute value due to a correlation between sampling errors in the moments and those in the weighting matrix.

[8] We are interested only in the unique elements of symmetric matrices $\Lambda\Lambda'$, $\Phi^p\Phi^{p\prime}$ and $\Phi^q\Phi^{q\prime}$ and therefore use the upper triangular elements.

$$G^{RS} = \begin{bmatrix} \text{vech}(H^{p}) - \text{vech}(H^{NE}) \\ \text{vech}(H^{q}) - \text{vech}(H^{NE}) \end{bmatrix} = \begin{bmatrix} \text{vech}(E[YY']^{p} - E[YY']^{NE} - \Phi^{p}\Phi^{p}{}') \\ \text{vech}(E[YY']^{q} - E[YY']^{NE} - \Phi^{q}\Phi^{q}{}') \end{bmatrix} \quad (10)$$

where $E[YY']^{p}$, $E[YY']^{q}$ and $E[YY']^{NE}$ are covariance matrices of securities on coinciding, non-coinciding and non-monetary-event days with elements averaged across the respective event days. However, partial identification arises under Rigobon and Sack's (2004) approach because Ω does not differ across regimes of heteroskedasticity and thus non-monetary surprises are not separately identifiable.

3.3. Summary of models

This paper estimates three models, which are summarised in Table 2. First, to assess only coinciding monetary surprises, UK, EA and US non-coinciding-event days are excluded from the sample. Model 1 includes coinciding monetary and non-monetary surprises and is estimated in the spirit of Rigobon and Sack (2004) for $\theta = [\Phi^{p}]$, whereby $G(\Phi^{p}, \Omega)$ is based on the difference of moments on monetary and non-monetary-event days. Second, model 2 also includes the non-coinciding international monetary surprises, giving $\theta = [\Phi^{p}, \Phi^{q}]$, whereby the difference of moments is again stacked into $G(\Phi^{p}, \Phi^{q}, \Omega)$. Because moment-conditions in the Rigobon and Sack's (op. cit.) estimator are based on the difference of moments on monetary- and non-monetary-event days, the non-monetary factors are not identifiable (see Appendix A.1. for further information about identification in these models).[9] Third, model 3 permits full identification using the moments rather than moment-differences within vector G, along the lines of Craine and Martin (2008) and Claus and Dungey (2012). The model is fully identified for $\theta = [\Phi^{p}, \Phi^{q}, \delta, \gamma]$, by stacking coinciding, non-coinciding and non-monetary-event day moments into $G(\Phi^{p}, \Phi^{q}, \Omega(\delta, \gamma))$.

[TABLE 2]

Models 1, 2 and 3 enable estimation of N, 4N and 6N parameters, using N(N+1)/2, 2N(N+1) and 5N(N+1)/2 moment-conditions, respectively. Adding securities to the model increases the number of moment-conditions relative to parameters. In model 1, there are N coinciding monetary factor loadings and 2 x N non-monetary factor loadings (N common-surprise and N idiosyncratic surprise factor loadings), but only the former are identified. In model 2, there are an additional 3 x N monetary factor loadings that are identified for non-coinciding monetary surprises (N factor loadings each for UK, EA and US non-coinciding-surprises). In model 3, there are 4 x N monetary factor loadings and 2 x N non-monetary factor loadings that are identified. In contrast to models 1 and 2, model 3 identifies all monetary and non-monetary factor loadings.

[9] All Appendix material is available online at the following website: https://pphelps0.wixsite.com/appendix.

4. Measurement and data description

The BoE, ECB and the FED have typically changed key interest rates on meeting days and made announcements subsequently.[10] Most meetings pre-existed within respective financial calendars; however, some announcements were made in response to unscheduled events. Both scheduled and unscheduled announcements are used to identify monetary-event days. UK and EA monetary surprises affect securities on the same trading day. Because the FED announcement at 2pm (Eastern-Time) occurs after the UK trading day closes, US monetary surprises affect UK securities on the next trading day.[11] There are 61 non-coinciding BoE monetary policy announcements; 82 for the ECB; 101 for the FED. There are 86 coinciding UK-EA monetary policy announcements. A small number of other, coinciding monetary announcements are excluded.[12] The full sample comprises 3,257 observations over the period 5 January 1999 to 22 July 2014, excluding 2008-2010 due to instability associated with the global financial crisis.

This paper focuses on UK and EA government-bond yield responses arising from coinciding BoE and ECB monetary policy announcements, while it provides some subsequent discussion about the broader implications of monetary coordination. For the UK, we obtain treasury yields with 1, 3, 7, 10, 20 and 30-year maturities and use the 3-month LIFFE sterling futures yield as the short rate proxy. For the EA, we obtain bond-yields with 1, 3, 7, 10, 15 and 30-year maturities and use the 3-month EUREX euribor futures yield as the short rate proxy.[13] Yields are expressed as annualised changes and measured on basis-points. Data are sourced from DataStream.

There are five event day classifications; UK, EA and US monetary-event days (UK Money, EA Money and US Money, respectively in Table 3), coinciding monetary-event days (Joint Money in Table 3) and non-monetary-event days (NE in Table 3).[14] Table 3 presents for each classification the standard deviations of changes in the UK and EA government-bond yields over the sample period.

[TABLE 3]

[10] Monetary policy announcement dates are taken from publications available at BoE, ECB and FED websites. BoE announcements are taken from: www.bankofengland.co.uk/publications/minutes;
ECB announcements from: www.ecb.europa.eu/press/govcdec/mopo;
FED announcements from: www.federalreserve.gov/monetarypolicy/fomccalendars.

[11] We have also explored using a two-day window instead for US monetary announcements to capture monetary surprise effects on the actual announcement day and subsequent trading day; we have actually obtained very similar results. These results are not reported for brevity, but are available from the authors upon request.

[12] Other monetary coincidences sometimes occur on trading days, e.g. joint UK-US, EA-US or UK-EA-US monetary-events. Bracketing all monetary coincidences together would increase the total number of coinciding monetary-event days somewhat; but this might reasonably induce heteroskedasticity across different types of monetary coincidences. Furthermore, because the number of coinciding monetary announcements involving the FED is very small, non-UK-EA monetary coincidences are excluded from the sample.

[13] The 20-year maturity bond yield data for the EA are only available from 2002, whereas the 15-year maturity bond yield data are available for the full sample of our study.

[14] In what follows the terms coinciding and joint monetary-event days are used interchangeably.

The standard deviations reported in Table 3 typically differ on monetary-event days compared with non-monetary-event days. Standard deviations are higher on UK, EA and US monetary-event days than on non-event days in almost all cases. The standard deviations on coinciding monetary-event days are relatively large for short-medium maturity bonds, whereas on non-coinciding monetary-event days the variation tends to be distributed more broadly across the yield curve. Sometimes the standard deviations are considerably larger on coinciding monetary-event days than on non-monetary-event days. For example, the standard deviations of 1-year UK and EA bond-yields are 45% and 59% larger on coinciding monetary-event days than on non-monetary event days, respectively; this provides some support for utilising the IH methodology. Furthermore, the outcomes of a joint F-test for the equality of variances of bond yields on monetary-event and non-event days are reported in this table. The null hypothesis of equality is rejected in all cases at the 1% significance level.[15] As such, there is statistical evidence that bond yield variation on monetary-event days differs significantly from on non-monetary-event days.

5. Empirical results

5.1. Coinciding monetary-event days

Table 4 summarises the estimation output for model 1, which captures the coinciding monetary surprise effect within the coefficient-vector Φ^{p}. Estimation is conducted for UK and EA government-bond yields. The different securities are listed in the first column. IH GMM estimates are normalised (NGMM) by the short rate estimate of 5.25 for comparison against event-study estimations, whereby individual government-bond yields are regressed on the relevant 3-month short rate. Two event-study estimations are undertaken. First, regression analysis is carried out using the ordinary least squares estimator (OLS). Second, because outliers might make some observations more influential, Huber's (1964) iteratively re-weighted robust regression methodology is applied as it assigns a lower weight to less well-behaved observations. All estimates are signed so that a positive monetary surprise represents a monetary-easing, i.e. the surprise is the negative of the relevant yield change.

[TABLE 4]

Point estimates are statistically significant at conventional levels, although they are generally smaller for longer-maturity bonds, providing evidence of a duration effect on the yield curve. NGMM and event-study point estimates are similar in most cases, although the former are more precise. The implication of NGMM and event-study estimates is that a 25 basis-point monetary-easing reduces 1, 10 and 30-year UK government-bond yields by 14-21, 9-12 and 7-10 basis-points, respectively; the corresponding ranges for EA government-bond yields are

[15] The null hypothesis is also rejected at the 1% significance level for both short rate proxies, and the heteroskedasticity condition is met overall for all securities. Even so, some caution is needed when discussing the international spillover effects on securities for which the monetary-event day standard deviations are relatively low.

15-25, 9-16 and 7-11 basis-points, respectively.[16] These estimates are larger in absolute value in comparison to, for instance, Rigobon and Sack's (2004) factor model estimates for non-coinciding-event days, which imply that a 25 basis-point US monetary-tightening would reduce 1-year US treasury yields by 7 basis-points. As such, coinciding monetary announcements appear to have relatively large effects on bonds with shorter-term maturities.

[FIGURE 1]

[FIGURE 2]

If the event-study point estimates were unbiased, they would be identical to the normalised estimates obtained from the IH factor model (Craine and Martin, 2008). Moreover, NGMM and event-study estimates are similar in that they are always negative and significant for each security and exhibit a duration effect across the yield curve, for both UK and EA government-bond yields. However, there are some discrepancies. To formally investigate biasedness under the event-study approach, a version of Hausman's (1978) test is used as presented in Rigobon and Sack (2004). The test statistics given in Table 4 indicate that the event-study estimates are sometimes biased because the observed variables are endogenous and/or omitted variables are present. Additionally, the R-squared statistics from least squares estimations tend to be higher for shorter-term bond yields, whereas the proportion of variation explained for longer-term bond yields is much lower. Moreover, the estimates obtained under (ordinary and iteratively re-weighted) least squares tend to overstate the absolute impact of monetary announcements at the short-end of the yield curve (see Figures 1 and 2). This likely reflects the omission from the model of important news announcements occurring on coinciding monetary-event days, which shift expectations about the economy in the same direction as interest rates. The bias seems much larger than in the non-coinciding monetary surprise literature (see, for example Claus et al., 2014). Therefore, application of the IH factor model seems particularly useful in this context.

5.2. Coinciding- and non-coinciding monetary-event days

Table 5 summarises GMM-estimation results for model 2, which provides advancement by including both coinciding and non-coinciding monetary surprises. Estimation of the model is conducted separately for the UK and EA government-bond yields. The different securities are listed in the first column. The non-normalised estimates correspond to security responses to a one standard deviation surprise monetary-easing on coinciding and non-coinciding UK and EA monetary-event days; for comparison we also include spillovers corresponding to US FED announcements on non-coinciding monetary-event days.

[TABLE 5]

[16] We consider government-bond yield implications of a 25 basis-point change in the short rate under ES and NGMM, as it is conventional in the literature, and multiply the point estimates accordingly.

Under model 2 we observe that the pass-through on coinciding monetary-event days is again relatively high at the short-end of the yield curve for both UK and EA government-bond yields. Additionally, non-coinciding monetary surprises have negative and significant effects; however, a joint monetary-easing tends to reduce short-term bond yields by more than on other (non-coinciding) monetary-event days. The BoE-ECB pass-through is even higher at the short-end of the yield curve than for a monetary surprise corresponding to the FED, although US monetary surprises seem relatively influential for medium-long term bond yields. Our findings are indicative of strong international portfolio balance effects, which arise from a coinciding monetary surprise and generate capital outflows along with lower international bond yields (Ammer et al., 2016). International financial spillover channels seem to reinforce the domestic monetary policy pass-through to financial markets.

5.3. Monetary and non-monetary factors

Table 6 summarises GMM estimation for model 3, in the case of the UK and EA government-bond yields. This model includes non-monetary surprises in addition to coinciding and non-coinciding monetary surprises.

[TABLE 6]

Estimation output reported in Table 6 and the corresponding variance decompositions for different monetary-event days (see Tables A1-A4 and Tables A5-A8 in the Appendix for UK and EA securities, respectively) provide an indication of the relative importance of monetary and non-monetary surprises. The contribution of a given surprise to the total theoretical variance is computed as 100 multiplied by the squared coefficient for that particular surprise, weighted by the total theoretical variance on event days. The latter is computed as the sum of squared coefficients on both the monetary and non-monetary surprises as it is presented in Table 6.[17] Monetary surprises are generally very important in explaining movements in bond yields and account for approximately 72% and 75% of the variance in movements of UK and EA short rates on coinciding-event days, respectively. However, non-monetary surprises often account for the majority of the variance in movements of government-bond yields. The importance of both systematic and security specific surprises is consistent with the asset-pricing literature. Our findings differ from others in the literature, such as Craine and Martin (2008) and Claus and Dungey (2012), in that common rather than security specific non-

[17] For example, the contribution of a coinciding monetary surprise for security i is given by:

$$100\left[\frac{(\Phi_i^p)^2}{\delta_i^2+\gamma_i^2+(\Phi_i^p)^2}\right]$$

This expression includes the coinciding monetary surprise variance and variances of common and idiosyncratic surprises. The variances of UK, EA and US non-coinciding monetary surprises are zero by assumption on coinciding monetary-event days.

monetary surprises are typically much more influential and relevant in explaining movements in the short rate.[18]

5.4. Level and rotation effects

This sub-section advances on the previous analyses by identifying level and rotation effects on coinciding monetary-event days. To this end, we examine the product of the short rate and 10, 20 and 30-year government-bond yields for the UK and the product of the short rate and 10, 15 and 30-year government-bond yields for the EA. If the product is positive for each short-long maturity combination, this implies that short- and long-maturity yield changes are of the same sign and correspond to a level effect. If instead the product is negative for each short-long maturity combination, this implies that the signs differ and correspond to a rotation effect. Monetary announcements can have a rotation effect or a level effect on the yield curve, or it can be unclassified if the signs differ across the short and long-maturity yield products.[19] The precise nature of these effects is inferred from estimation of an extension of model 1, which in this context includes two coinciding monetary-event surprises, to account for level and rotation shocks. The estimation output for the UK and EA securities is reported in Table 7.

[TABLE 7]

Table 7 provides evidence of significant level and rotation effects on different monetary-event days. Regarding the level effects, all yields decrease following a monetary-easing according to the results presented in Table 7, which is suggestive of strong portfolio balance effects on bond yields. This is also consistent with the expectations hypothesis of the term structure, whereby a surprise monetary-easing implies lower short-term interest rates and lower expectations of future short-term interest rates (Cook and Hahn, 1989; Kuttner, 2001). Table 7 also presents evidence for rotation effects; yield changes following a coinciding monetary-easing tend to be much more negative at the short-end of the yield curve, whereas the opposite is true for long-maturity securities. In contrast to the expectations theory of the term structure, UK and EA long-maturity bond yields move in the opposite direction to the relevant short rates. This finding is consistent with monetary theory, as a reduction in the short rate should increase inflation and the level of sufficiently long rates (Romer and Romer, 2000). From a financial investment perspective, a surprise monetary-easing that promotes expectations of recovery (and higher inflation expectations), increases the required compensation in the form of yield over medium and longer-term horizons, thus generating a rotation of the yield curve (Gagnon et al., 2011). This finding corroborates with the expectation of major central banks; that by reacting promptly and in a cooperative manner following negative macroeconomic and financial shocks, it would broadly help to reduce the

[18] In Table 6 the point estimates that correspond to the common non-monetary surprises are always significant, whereas this is not the case for non-monetary idiosyncratic surprises. Furthermore, in most cases, point estimates are more sizeable for common non-monetary surprises.

[19] If an observation is unclassified as either a level or rotation monetary surprise, it is excluded for the purpose of this exercise.

severity of the global economic and financial crisis and increase expectations of recovery (Rogers et al., 2014).

The rotation effects on non-coinciding monetary-event days often, though not always, differ qualitatively. Yields of shorter-maturity bonds tend to increase on non-coinciding monetary-event days, whereas yields decline for longer-maturity bonds. As to why the rotations differ on non-coinciding monetary-event days, especially for international monetary surprises, it may be that a reduction in international interest rates and associated exchange rate appreciation of the domestic currency vis-à-vis the foreign currency and the US dollar (USD) currency signals weaker UK and EA economic prospects and lower inflationary expectations via the expenditure-switching effect (Bernanke, 2013). Perhaps one of the most compelling explanations for the coinciding interest rate cuts is that policy makers wish to be seen to be cooperating during a global crisis, thereby increasing confidence (BIS, 2009). In the absence of perceived cooperation, the confidence implications can be reversed. Therefore, coinciding monetary announcements seem to convey very strongly to financial markets that central bankers are acting in unity to promote recovery in terms of the longer-term economic prospects.[20] In this sense, the empirical coinciding monetary policy effect appears larger and more sustained than had been previously acknowledged within the theoretical literature (see, for example, Taylor, 2013).

To investigate the severity of endogeneity bias we also run OLS estimations of the UK and EA short rate on government-bond yields, along the lines of Table 4, accounting for both coinciding monetary-event day level and rotation surprises.[21] Qualitatively similar level and rotation effects are obtained under the ES approach (see Table A9 as in the Appendix for comparison). However, quantitative differences in the point estimates are apparent, with biasedness indicated formally through the Hausman test outcomes presented in Table 7. Sometimes OLS estimates are moderately or severely biased; this points to the presence of endogeneity bias and/or omitted variables. This supports further the use of the IH methodology adopted in this paper over more popular alternatives.

[20] We note that a domestic monetary-easing on non-coinciding monetary-event days sometimes does have a favourable rotation effect, although this is only apparent for non-coinciding EA monetary-event days. The rotation effect is small relative to the estimated effect on coinciding monetary-event days.

[21] To obtain ES estimates in this context we run the following regression of UK/EA government-bond security y_{jt} for $j \neq 1$ on a set of explanatory variables, including dummy variable interactions of security y_{1t}, where y_{1t} is the short rate, and a dummy variable D_t, which takes a value of 1 for a coinciding monetary level surprise and 0 for a coinciding monetary rotation surprise:

$$y_{j,t} = \phi_0 + \phi_1 D_t y_{1,t} + \phi_2 (1 - D_t) y_{1,t} + \phi_3 D_t + \varepsilon_t$$

where ε_t is a normally distributed disturbance term with the usual properties. Due to the interaction specification of the regression model, the coefficient estimates for ϕ_1 and ϕ_2 indicate the level and rotation monetary surprise effects on government-bond yield j, respectively. The monetary surprise is measured through an observable change in the short rate and is classified as a level or rotation surprise according to the criteria set out in the main text.

5.5. Model-implied and sample variance-covariance matrices

To evaluate the performance of the three factor models, the actual sample variance-covariance matrices are compared to the model-implied variance-covariance matrices on coinciding, non-coinciding monetary-event days and non-monetary-event days using estimation output presented in Tables 4, 5 and 6.

First, the model-implied variance-covariance matrix is computed using the GMM estimates presented in Table 4 for model 1, which only identifies the coinciding monetary-surprise effects on bond market yields. Using the first row from the N x N model-implied variance-covariance matrix on coinciding monetary-event days and the corresponding row of the N x N sample variance-covariance matrix, we compute the absolute percentage deviations of model-implied moments from the sample moments (i.e. for the variance of the security and its covariance with the other N – 1 bond yields). This process is repeated for all rows, allowing computation of absolute percentage deviations for each of the N securities from the sample moments on coinciding monetary-event days. Second, the model-implied variance-covariance matrices are computed using the GMM estimates presented in Table 5 for model 2, which only identifies the coinciding and non-coinciding monetary-surprise effects. Since in addition to the identical coinciding monetary-surprise effects there are three non-coinciding monetary-event days (UK, EA and US), there are three additional variance-covariance matrices to compare; therefore, three sets of absolute percentage deviations are obtained with respect to the sample variance-covariance matrices. However, model 1 and model 2 do not permit identification of non-monetary surprise effects. Therefore, we obtain another set of absolute percentage deviations using the GMM estimates presented in Table 6 for model 3, which is fully identified and permits computations of the model-implied moments and comparison with the sample counterparts on non-monetary-event days. In a nutshell, if the difference between model-implied and sample variance-covariance matrices is very small, this indicates that the model matches very well the key characteristics of the data.

Table 8 summarises the results from this exercise by subtracting from 100% the median absolute percentage deviation between the model-implied and sample variance-covariance matrices for individual securities. The results are presented for the three factor models. The values of approximately 79% and 83% on coinciding monetary-event days for UK and EA short rates indicate that model-implied and actual sample variances and covariances are typically very similar under model 3, the full-factor model. On average, around 80% of the sample-covariance matrix can be explained by the model-implied moments on coinciding monetary-event days compared to approximately 40% in models 1 and 2. Therefore, the full-factor model is better able to replicate the sample variance-covariance matrix on coinciding monetary-event days.

[TABLE 8]

Under model 3 the model-implied moments for UK securities on non-coinciding UK, EA and US monetary-event days account for approximately 83%, 74% and 54% of the observed variances and covariances, respectively. For EA securities the model-implied moments on non-coinciding UK, EA and US monetary-event days account for approximately 68%, 76% and 46% of the observed variances and covariances, respectively. Evidently, model 3 provides substantial improvement over model 2 in terms of its ability to replicate the sample moments on non-coinciding UK and EA monetary-event days for all securities. Additionally, the full-factor model explains up to 70% of sample moments of UK securities on non-monetary-event days and up to 80% for EA securities. Furthermore, the ability of the full-factor model to replicate sample moments is evident from Figures A1-A8 (as in the Appendix) on different monetary-event days, particularly on coinciding and non-coinciding UK and EA monetary-event days. Interestingly, model 2 is generally able to better replicate the sample moments on non-coinciding US monetary-event days than model 3, although this is not the case for all securities. In this light, application of our modelling suite is helpful in replicating key characteristics of the data for our sample.

5.6. Robustness

The following checks are conducted for robustness and to assess sample sensitivity.[22] First, the models are re-estimated for 1999-2007, excluding the period associated with the zero lower bound for interest rates and quantitative easing rounds. Second, to investigate robustness of the coinciding monetary surprise effect with respect to the business cycle, the first factor model is re-estimated to capture differences in monetary surprise effects when the short rate change is positive (associated with a monetary-tightening) and negative (associated with a monetary-easing). This entails including an additional set of monetary surprises for each monetary-event day; one vector of factor loadings when the change in the short rate is negative and another vector of factor loadings when the change in the short rate is positive.[23]

For the first check, after excluding the period associated with the zero lower bound, we find that the estimates are qualitatively and quantitatively similar to what is presented in Tables 4, 5 and 6 for UK and EA securities. Estimates for the level and rotation effects are qualitatively similar, although on coinciding monetary-event days point estimates are generally larger in absolute value and more negative at the short-end of the yield curve than in Table 7. For the second check, the pass-through at the short-end of the yield curve is relatively high during a monetary-easing and, for the securities considered, is slightly more pronounced for UK securities and even more so for EA securities. Therefore, the coinciding monetary policy effect appears somewhat sensitive to the economic cycle, although this may be favourable to policy makers attempting to stimulate recovery through a coordinated monetary-easing.[24] While some caution is needed when interpreting results based on refined samples, these checks do not undermine our main conclusions, which are set out further below following discussion.

[22] The complete estimation results from these checks are available from the authors upon request.

[23] Any remaining observations on coinciding monetary-event days are excluded from estimation. For $N = 7$ and with there being two types of coinciding monetary-events – level and rotation – the number of moment conditions, $2N(N+1)/2 = 56$, exceeds the number of parameters to estimate, $2N = 14$.

[24] Very similar results are obtained using the full-factor model.

6. Discussion

From a policy perspective, there is some debate as to whether the international responses of central bankers represent an optimal solution. On the one hand, we have observed a series of coinciding and non-coinciding monetary announcements, whereby central bank responses have closely followed one another. According to the 'beggar-thy-neighbour' proposition, exchange rate depreciations help first and foremost the economy whose currency weakened by making that country more internationally competitive. To that end, this paper identifies significant non-coinciding international monetary spillover effects arising from conventional monetary policy; although according to our preferred methodology, these monetary surprises have less favourable and weaker yield curve implications than domestic and coinciding monetary spillovers. In effect, our results are more in line with the old-Keynesian version of the Mundell-Fleming model, wherein the exchange rate channel plays a significant role. These findings do not contradict the empirical evidence, discussed by Taylor (2013) and Blanchard et al. (2015), which shows that the international growth effect of a domestic monetary-easing may only marginally offset the adverse effects of the exchange rate appreciation and capital inflows that arise from a monetary-easing from abroad.

Others argue that the pattern of international responses observed during the global financial crisis corresponds effectively to a joint monetary-easing, with monetary cooperation having broadly favourable international effects on financial markets and growth. For example, Bernanke (2013) compares recent monetary policy shifts to what happened during the Great Depression, as countries moved off the gold standard and sequentially engaged in so-called 'competitive devaluations'; however, he describes these as an appropriate move towards monetary-easing. According to Bernanke (op. cit.), the benefits of monetary accommodation, at least for the case of the G7 economies, are not significantly created by exchange rate movements, but come primarily from the stimulation of domestic aggregate demand. To an extent, this paper finds supportive evidence for this proposition, with a favourable rotation effect corresponding to ECB policy on non-coinciding monetary-event days; although this is not the case for the BoE. However, a closer form of cooperation may be more effective. Indeed, much more favourable are the financial market implications of coinciding monetary policy announcements, when the BoE and ECB are perceived by markets to be cooperating: on these days, the UK and EA bond market effect of monetary coincidences is positive, sizeable and significant.

The question remains as to whether or not greater monetary coordination among major central banks such as the BoE, ECB and FED, would be helpful to those countries that are not similarly aligned to the global financial cycle. There are also concerns that coordinated monetary-easing may over longer-horizons generate excessive inflation and the formation of asset price bubbles, in advanced and emerging economies (see, for example, Borio, 2014; Borio and Zabai, 2016). Moreover, the strength of coinciding monetary policy spillovers, evidenced in this paper, does suggest a need to consider the broader financial market implications and the international welfare consequences of greater monetary coordination among major central banks. In particular, there are growing calls from the international community to put in place appropriate safety mechanisms for other countries that might be adversely affected by the powerful monetary spillovers (see, for example, Ostry and Ghosh, 2016; Rajan, 2014, 2016).

7. Summary and conclusions

The coinciding nature of monetary policy responses during the turmoil reflects the severe nature of events, which forces policy makers to act immediately and thus simultaneously. Additionally, or for both reasons at the same time, policy makers expect to make a bigger impact on the markets and boost confidence if they are seen to be cooperating. Yet, considerable uncertainty remains as to whether and how the simultaneous responses of major central banks would impact on the financial markets and promote expectations of a sustained recovery. Against this background, our primary contribution is to investigate in a robust manner whether international monetary policy surprises arising on the same trading day have been influential for government-bond markets. To do this, we utilise a unique dataset of monetary coincidences corresponding to BoE and ECB monetary policy announcements and estimate a suite of multi-security factor models, which captures international spillover effects on government-bond markets over the period 1999-2014. In doing so, this paper establishes whether coinciding monetary surprise effects differ from other monetary and non-monetary surprise effects and whether the set of UK bond market reactions is unique or similar to EA reactions.

Estimation of the full-factor model, which is preferred in this context to the event-study approach, indicates that coinciding monetary surprises are relatively influential for short-maturity bonds and have a greater impact than other monetary and non-monetary surprises. Monetary policy announcements on coinciding monetary-event days generate additional pass-through at the short-end of the yield curve via a rotation effect, which is strikingly different from non-coinciding monetary-event days and contrasts with the expectations theory of the term structure of interest rates. As such, the overall UK and EA government-bond market effects of monetary coincidences are similarly positive, sizeable and significant, and go above and beyond the non-coinciding effects.

In conclusion this paper's findings suggest that the swift and cooperative responses of major central banks must have helped to reduce the severity of the global crisis and promote expectations of a sustained recovery in the UK and EA. Future cooperation in the form of coordinated monetary responses, which boost expectations of recovery in major economies, may therefore be an effective policy tool in the face of major global shocks, leading to 'enrich-thy-neighbour' outcomes, at least for the countries involved. Greater monetary policy coordination certainly appears favourable in this light, especially for countries similarly exposed to the global financial cycle. However, the strength of international monetary spillovers does suggest a need to investigate the broader financial market implications of monetary coordination and the global welfare consequences, which remain fruitful areas for future research.

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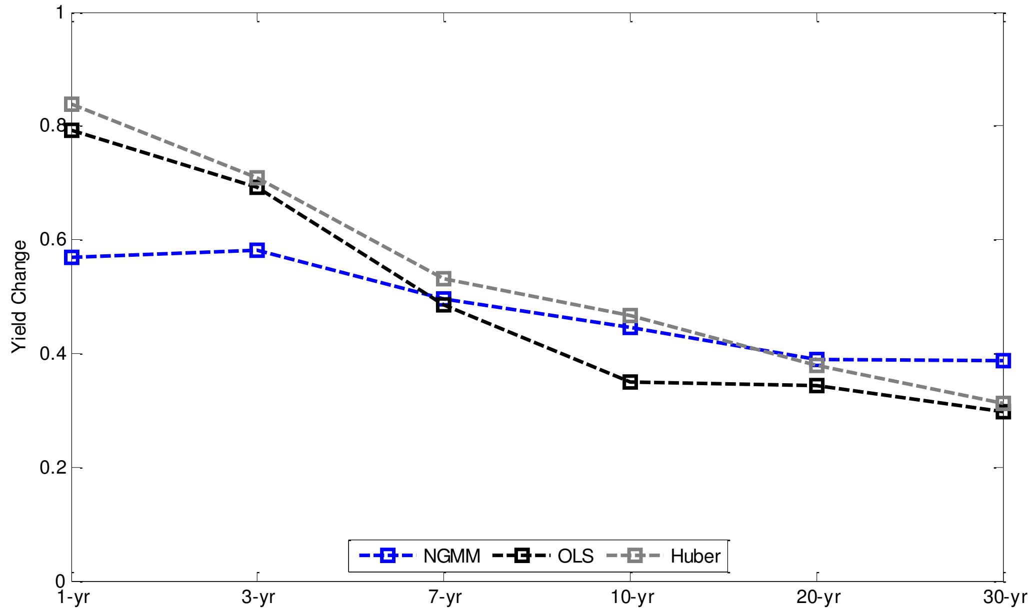
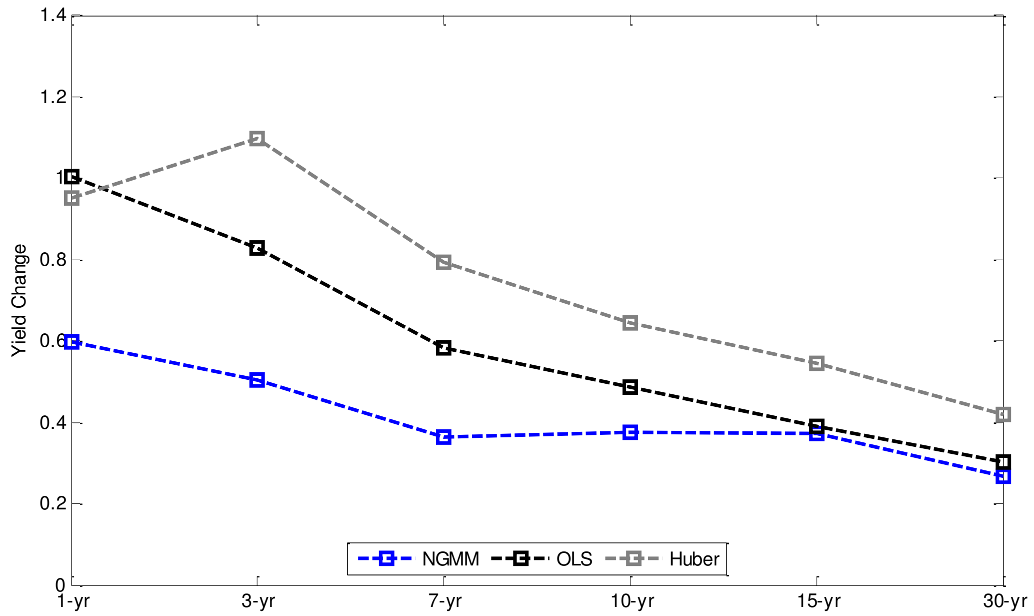
TABLES AND FIGURES

Figure 1. UK bond yield responses to a one basis-point monetary surprise



Source: Authors' own computation based on data from DataStream.
Notes: Yields are based on estimates of Φ's reported in Table 4 for 1, 3, 7, 10, 20 and 30-year UK government-bond securities. Black, grey and blue squares correspond to OLS, Huber (1964) and NGMM point estimates, respectively.

Figure 2. EA bond yield responses to a one basis-point monetary surprise



Source: Authors' own computation based on data from DataStream.
Notes: Yields are based on estimates of Φ's reported in Table 4 for 1, 3, 7, 10, 15 and 30-year EA government-bond securities. Black, grey and blue squares correspond to OLS, Huber (1964) and NGMM point estimates, respectively.

Table 1

Summary of theoretical moments

	Coinciding Monetary-Event Days	Non-Coinciding Monetary-Event Days	Non-Monetary-Event Days
Variance	$(\Phi_i^p)^2 + \delta_i^2 + \gamma_i^2$	$(\Phi_i^q)^2 + \delta_i^2 + \gamma_i^2$	$\delta_i^2 + \gamma_i^2$
Covariance	$\Phi_i^p \Phi_j^p + \delta_i \delta_j$	$\Phi_i^q \Phi_j^q + \delta_i \delta_j$	$\delta_i \delta_j$

Source: Authors' own construction.

Table 2

Summary of models

Model	Moments	Moment-Conditions	Parameters Estimated	Identification	Parameters Identified (θ)
1	N(N+1)	N(N+1)/2	N	Partial	Φ^p
2	5N(N+1)/2	2N(N+1)	4N	Partial	Φ^p, Φ^q
3	5N(N+1)/2	5N(N+1)/2	6N	Full	$\Phi^p, \Phi^q, \delta, \gamma$

Source: Authors' own construction.

Table 3

Standard deviations of government-bond yield changes

	Joint Money	UK Money	EA Money	US Money	NE	F-test
Panel A: UK						
1-year	6.06	4.90	4.64	4.62	4.19	1.47***
3-year	5.45	5.25	5.60	6.03	4.21	1.79***
7-year	4.97	5.27	4.85	6.53	4.32	1.64***
10-year	5.34	5.30	5.12	6.89	4.56	1.62***
20-year	4.63	4.73	4.32	6.10	4.02	1.59***
30-year	4.41	4.69	4.30	5.73	3.94	1.52***
Panel B: EA						
1-year	5.47	3.39	4.63	4.57	3.44	1.84***
3-year	5.57	4.62	5.56	6.31	4.18	1.82***
7-year	5.14	4.90	4.73	6.54	4.29	1.63***
10-year	4.73	5.17	5.09	6.44	4.21	1.68***
15-year	4.38	5.17	4.91	6.15	4.20	1.56***
30-year	4.12	4.83	4.48	5.91	4.04	1.49***
Observations	86	61	82	101	2,927	

Source: Authors' own computation based on data from DataStream.

Notes: The F-test statistic is based on the ratio of variances, whereby under the null hypothesis of equality of variances the ratio takes a value of unity. The F-statistic is distributed with $E^p + E^q - 1$ and NE - 1 degrees of freedom.

Table 4
GMM versus event-study estimation

	Joint Event GMM		Normalised GMM		Event-Study (OLS)		Event-Study (Huber)		Endogeneity Bias	Model Fit
	Φ^{GMM}	SE^{GMM}	Φ^{NGMM}	SE^{NGMM}	Φ^{OLS}	SE^{OLS}	Φ^{Huber}	SE^{Huber}	H-Test	R^2
Panel A: UK										
Short rate	-5.18***	(0.31)	-1.00***	(0.06)						
1-year	-2.94***	(0.24)	-0.57***	(0.05)	-0.79***	(0.07)	-0.84***	(0.05)	14.96***	0.57
3-year	-3.01***	(0.24)	-0.58***	(0.05)	-0.69***	(0.07)	-0.71***	(0.07)	4.68**	0.53
7-year	-2.56***	(0.23)	-0.50***	(0.04)	-0.49***	(0.08)	-0.53***	(0.07)	0.03	0.32
10-year	-2.31***	(0.23)	-0.45***	(0.04)	-0.35***	(0.09)	-0.47***	(0.08)	1.43	0.14
20-year	-2.02***	(0.22)	-0.39***	(0.04)	-0.34***	(0.08)	-0.38***	(0.07)	0.51	0.19
30-year	-2.00***	(0.22)	-0.39***	(0.04)	-0.30***	(0.08)	-0.31***	(0.07)	2.00	0.15
Observations		3,013		3,013		86		86	Average R^2 = 0.32	
Panel B: EA										
Short rate	-5.57***	(0.31)	-1.00***	(0.06)						
1-year	-3.33***	(0.25)	-0.60***	(0.04)	-1.00***	(0.07)	-0.95***	(0.06)	64.08***	0.72
3-year	-2.80***	(0.24)	-0.50***	(0.04)	-0.83***	(0.09)	-1.10***	(0.08)	14.85***	0.47
7-year	-2.02***	(0.22)	-0.36***	(0.04)	-0.58***	(0.10)	-0.79***	(0.09)	5.41**	0.28
10-year	-2.10***	(0.22)	-0.38***	(0.04)	-0.48***	(0.10)	-0.64***	(0.09)	1.57	0.23
15-year	-2.08***	(0.22)	-0.37***	(0.04)	-0.39***	(0.09)	-0.54***	(0.09)	0.04	0.16
30-year	-1.49***	(0.21)	-0.27***	(0.04)	-0.30***	(0.09)	-0.42***	(0.08)	0.18	0.09
Observations		3,013		3,013		86		86	Average R^2 = 0.33	

Source: Authors' own computation based on data from DataStream.
Notes: Estimates of Φ's and standard errors (SE's) with GMM, NGMM, OLS and Huber superscripts correspond to GMM, normalised GMM (both under model 1), OLS and Huber's (1964) robust event-study estimation on coinciding monetary-event days, respectively. Non-normalised GMM estimates correspond to security responses a one standard deviation monetary surprise on coinciding monetary-event days, whereas normalised GMM and event-study estimates correspond to responses to a one basis-point coinciding monetary surprise, measured through a change in the short rate. Values reported in column H-Test correspond to Hausman's (1978) test statistic, under the distribution $F(m, E^p - 1)$ where m is the number of coefficients being tested (m = 1 for tests are applied to individual securities) and E^p is the number of coinciding monetary-event days on which the event-study estimation is based (E^p = 86). For this test the 10%, 5% and 1% critical values are 2.77, 3.95 and 6.94, respectively. Model fit is determined from the R-squared statistic (R^2) based on individual OLS estimations. ***, **, * indicate significance at the 1%, 5% and 10% levels, respectively.

Table 5
Responses to coinciding and non-coinciding monetary surprises

	Joint Money		UK Money		EA Money		US Money	
	Φ^{p}	SE^{p}	Φ^{UK}	SE^{UK}	Φ^{EA}	SE^{EA}	Φ^{US}	SE^{US}
Panel A: UK								
Short rate	-5.18***	(0.31)	-4.24***	(0.26)	-3.68***	(0.23)	-1.73***	(0.26)
1-year	-2.94***	(0.24)	-2.92***	(0.23)	-2.36***	(0.19)	-3.71***	(0.30)
3-year	-3.01***	(0.24)	-2.36***	(0.21)	-2.19***	(0.18)	-4.57***	(0.31)
7-year	-2.56***	(0.23)	-2.99***	(0.24)	-2.18***	(0.19)	-4.81***	(0.33)
10-year	-2.31***	(0.23)	-2.53***	(0.22)	-2.01***	(0.19)	-2.97***	(0.28)
20-year	-2.02***	(0.22)	-2.34***	(0.21)	-2.08***	(0.19)	-4.27***	(0.30)
30-year	-2.00***	(0.22)	-2.75***	(0.23)	-1.83***	(0.18)	-3.76***	(0.29)
Observations								3,257
Panel B: EA								
Short rate	-5.57***	(0.31)	-0.90***	(0.18)	-4.81***	(0.27)	-2.15***	(0.28)
1-year	-3.33***	(0.25)	-2.09***	(0.17)	-2.37***	(0.22)	-3.88***	(0.31)
3-year	-2.80***	(0.24)	-2.33***	(0.18)	-2.56***	(0.22)	-4.40***	(0.31)
7-year	-2.02***	(0.22)	-2.74***	(0.20)	-2.61***	(0.22)	-4.57***	(0.32)
10-year	-2.10***	(0.22)	-2.31***	(0.18)	-1.81***	(0.21)	-3.33***	(0.30)
15-year	-2.08***	(0.22)	-2.84***	(0.20)	-2.35***	(0.21)	-4.29***	(0.31)
30-year	-1.49***	(0.21)	-2.49***	(0.18)	-2.22***	(0.21)	-4.02***	(0.30)
Observations								3,257

Source: Authors' own computation based on data from DataStream.

Notes: GMM estimates of Φ's and standard errors (SE's) with p, UK, EA and US superscripts correspond to coinciding and non-coinciding UK, EA and US monetary-event days, respectively. ***, **, * indicate significance at the 1% , 5% and 10% levels, respectively.

Table 6

Responses to monetary and non-monetary surprises

	Monetary-Event Days								Non-Monetary-Event Days			
	Joint Money		UK Money		EA Money		US Money		Common		Idiosyncratic	
	Φ^{p}	SE^{p}	Φ^{UK}	SE^{UK}	Φ^{EA}	SE^{EA}	Φ^{US}	SE^{US}	δ	SE(δ)	γ	SE(γ)
Panel A: UK												
Short rate	-5.18***	(0.31)	-4.24***	(0.26)	-3.68***	(0.23)	-1.73***	(0.26)	-3.24***	(0.26)	-0.24***	(0.01)
1-year	-2.94***	(0.24)	-2.92***	(0.23)	-2.36***	(0.19)	-3.71***	(0.30)	-3.31***	(0.27)	0.00	(0.04)
3-year	-3.01***	(0.24)	-2.36***	(0.21)	-2.19***	(0.18)	-4.57***	(0.31)	-3.89***	(0.28)	-1.54***	(0.08)
7-year	-2.56***	(0.23)	-2.99***	(0.24)	-2.18***	(0.19)	-4.81***	(0.33)	-3.98***	(0.29)	0.00	(0.05)
10-year	-2.31***	(0.23)	-2.53***	(0.22)	-2.01***	(0.19)	-2.97***	(0.28)	-2.77***	(0.26)	0.00	(0.04)
20-year	-2.02***	(0.22)	-2.34***	(0.21)	-2.08***	(0.19)	-4.27***	(0.30)	-3.70***	(0.28)	-0.97***	(0.05)
30-year	-2.00***	(0.22)	-2.75***	(0.23)	-1.83***	(0.18)	-3.76***	(0.29)	-2.97***	(0.26)	-2.60***	(0.13)
Observations												3,257
Panel B: EA												
Short rate	-5.57***	(0.31)	-0.90***	(0.18)	-4.81***	(0.27)	-2.15***	(0.28)	-2.56***	(0.24)	-1.89***	(0.09)
1-year	-3.33***	(0.25)	-2.09***	(0.17)	-2.37***	(0.22)	-3.88***	(0.31)	-3.02***	(0.26)	0.00	(0.05)
3-year	-2.80***	(0.24)	-2.33***	(0.18)	-2.56***	(0.22)	-4.40***	(0.31)	-3.43***	(0.26)	-2.14***	(0.11)
7-year	-2.02***	(0.22)	-2.74***	(0.20)	-2.61***	(0.22)	-4.57***	(0.32)	-4.07***	(0.28)	0.00	(0.04)
10-year	-2.10***	(0.22)	-2.31***	(0.18)	-1.81***	(0.21)	-3.33***	(0.30)	-2.74***	(0.25)	0.00	(0.06)
15-year	-2.08***	(0.22)	-2.84***	(0.20)	-2.35***	(0.21)	-4.29***	(0.31)	-3.80***	(0.27)	-0.95***	(0.05)
30-year	-1.49***	(0.21)	-2.49***	(0.18)	-2.22***	(0.21)	-4.02***	(0.30)	-3.14***	(0.25)	-2.54***	(0.13)
Observations												3,257

Source: Authors' own computation based on data sourced from DataStream.

Notes: GMM monetary estimates of Φ's and standard errors (SE's) with p, UK, EA and US superscripts correspond to coinciding and non-coinciding UK, EA and US monetary-event days, respectively. Non-monetary estimates (δ, γ) correspond to common and idiosyncratic surprises, respectively.

Table 7
Monetary-surprise level and rotation effects

	Joint Money		UK Money		EA Money		US Money		Endogeneity Bias
	Φ^{p}	SE^{p}	Φ^{UK}	SE^{UK}	Φ^{EA}	SE^{EA}	Φ^{US}	SE^{US}	H-Test
Panel A: UK									
Level Effects									
Short rate	-6.07***	(0.39)	-6.43***	(0.42)	-5.03***	(0.35)	-2.76***	(0.37)	
1-year	-4.57***	(0.35)	-5.01***	(0.39)	-4.22***	(0.33)	-5.27***	(0.42)	16.33***
3-year	-3.97***	(0.33)	-4.19***	(0.37)	-4.08***	(0.32)	-6.10***	(0.42)	9.26***
7-year	-3.96***	(0.34)	-4.74***	(0.38)	-3.72***	(0.31)	-6.10***	(0.43)	0.02
10-year	-3.95***	(0.34)	-4.58***	(0.38)	-3.76***	(0.32)	-4.56***	(0.39)	0.56
20-year	-3.65***	(0.33)	-3.91***	(0.36)	-3.72***	(0.32)	-5.26***	(0.40)	0.07
30-year	-3.68***	(0.33)	-4.77***	(0.39)	-3.44***	(0.31)	-5.05***	(0.41)	2.08
Rotation Effects									
Short rate	-6.69***	(0.38)	2.30***	(0.18)	2.35***	(0.18)	2.58***	(0.21)	
1-year	-3.15***	(0.32)	0.70***	(0.15)	1.17***	(0.22)	0.68***	(0.19)	25.01***
3-year	-3.97***	(0.32)	-2.06***	(0.15)	2.53***	(0.19)	0.95***	(0.15)	2.88*
7-year	-1.90***	(0.26)	-2.84***	(0.19)	0.86***	(0.13)	1.58***	(0.16)	6.16**
10-year	0.80***	(0.30)	-0.77***	(0.15)	-1.07***	(0.21)	-2.00***	(0.22)	5.61**
20-year	2.93***	(0.27)	-1.87***	(0.15)	-3.26***	(0.21)	-4.31***	(0.25)	0.30
30-year	2.70***	(0.27)	-0.07	(0.12)	-3.59***	(0.23)	-4.13***	(0.24)	0.04
Observations		2,985		2,991		2,981		2,986	
Panel B: EA									
Level Effects									
Short rate	-3.77***	(0.33)	-2.63***	(0.32)	-5.76***	(0.37)	-2.97***	(0.35)	
1-year	-4.48***	(0.33)	-4.13***	(0.34)	-3.89***	(0.34)	-4.97***	(0.39)	10.11***
3-year	-4.32***	(0.32)	-4.61***	(0.34)	-4.25***	(0.34)	-5.50***	(0.39)	27.23***
7-year	-3.81***	(0.31)	-4.92***	(0.36)	-4.41***	(0.35)	-5.41***	(0.39)	7.81***
10-year	-3.86***	(0.31)	-4.14***	(0.34)	-3.43***	(0.33)	-4.26***	(0.37)	3.61*
15-year	-3.92***	(0.32)	-4.88***	(0.35)	-4.01***	(0.33)	-4.99***	(0.38)	0.07
30-year	-3.61***	(0.31)	-4.73***	(0.35)	-3.99***	(0.34)	-4.81***	(0.38)	0.07
Rotation Effects									
Short rate	-10.54***	(0.53)	1.44***	(0.15)	-5.62***	(0.29)	1.00***	(0.24)	
1-year	-0.87***	(0.29)	0.98***	(0.13)	0.18	(0.17)	-1.15***	(0.24)	4.96***
3-year	0.40	(0.29)	-0.44***	(0.13)	0.08	(0.18)	-1.76***	(0.19)	4.30***
7-year	-0.67**	(0.29)	-0.75***	(0.12)	1.03***	(0.18)	-4.45***	(0.28)	2.29
10-year	0.37	(0.29)	0.44***	(0.10)	0.24	(0.17)	-2.71***	(0.28)	3.17*
15-year	0.29	(0.29)	-1.83***	(0.14)	1.24***	(0.18)	-5.34***	(0.33)	7.18***
30-year	0.85***	(0.29)	-0.84***	(0.11)	1.27***	(0.18)	-4.85***	(0.29)	9.21***
Observations		2,991		2,990		2,990		2,991	

Source: Authors' own computation based on data sourced from DataStream.

Notes: GMM estimates of Φ's and standard errors (SE's) with p, UK, EA and US superscripts correspond to coinciding and non-coinciding UK, EA and US monetary-event days, respectively. Estimation is based on an extension of model 1, where level and rotation effects are identified from two separate monetary surprises arising on coinciding monetary-event days. Values reported in column H-Test correspond to Hausman's (1978) test statistic, under the distribution $F(m, E^{p} - 1)$ where m is the number of coefficients being tested (m = 1 for tests are applied to individual securities) and E^{p} is the number of coinciding monetary-event days on which the event-study estimation is based ($E^{p} = 58$ for the UK and $E^{p} = 64$ for the EA, respectively). For this test the 10%, 5% and 1% critical values for the UK estimations are 2.80, 4.01 and 7.10, respectively; for the EA estimations, the corresponding critical values are 2.79, 3.99 and 7.06, respectively .. ***, **, * indicate significance at the 1%, 5% and 10% levels, respectively. ***, **, * indicate significance at the 1%, 5% and 10% levels, respectively.

Table 8
Model-implied versus actual sample moments

	Models 1 and 2	Model 2			Model 3				
	Joint Money	UK Money	EA Money	US Money	Joint Money	UK Money	EA Money	US Money	Non-Money
Panel A: UK									
Short rate	57.26	46.75	32.84	24.27	79.45	74.82	72.11	64.74	40.37
1-year	47.07	46.75	32.84	74.44	82.11	87.23	74.89	64.74	57.82
3-year	39.83	30.98	22.33	77.90	82.29	79.93	71.87	42.52	64.31
7-year	30.79	36.15	21.46	72.83	82.29	88.84	74.89	67.17	55.77
10-year	45.37	41.08	28.10	81.68	75.79	74.82	73.65	46.38	68.56
20-year	31.26	33.46	25.85	82.26	76.88	89.14	75.71	33.75	69.44
30-year	29.51	40.23	21.73	69.92	76.92	89.14	75.71	61.57	59.27
Average	40.16	39.34	26.45	69.04	79.39	83.42	74.12	54.41	59.36
Panel B: EA									
Short rate	60.32	8.34	50.63	33.30	83.23	42.21	73.12	64.34	32.56
1-year	48.50	23.97	33.76	71.69	72.73	68.82	77.87	71.39	46.97
3-year	36.92	28.72	31.90	67.46	82.05	60.16	70.86	40.44	62.98
7-year	25.55	32.80	32.11	78.84	84.00	72.15	77.01	46.56	57.96
10-year	45.59	42.00	34.09	77.20	76.46	72.69	77.87	41.89	66.69
15-year	27.98	42.00	32.11	91.42	89.93	78.52	83.18	17.16	80.04
30-year	22.83	33.30	31.65	77.20	67.80	83.05	73.07	40.44	60.48
Average	38.24	30.16	35.18	71.02	79.46	68.23	76.14	46.03	58.24

Source: Authors' own computation based on data sourced from DataStream.

Notes: Median percentage differences are computed using model-implied moments for models 1, 2 and 3 and the sample moments on joint, UK, EA and US monetary and non-monetary-event days. The average percentage difference across individual securities is also reported in the table.

Highlights

-We estimate international monetary surprise effects on government-bond markets

-We determine the relative importance of monetary and non-monetary surprises

-Financial market effects of monetary cooperation exceed that of conventional policy

-Our findings provide some support for the 'enrich-thy-neighbour' hypothesis