

An Empirical Analysis of the Relationship between US Monetary Policy and International Asset Prices

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Abstract

This article investigates the empirical relationship between monetary policy in the United States (US) and international equity, bond and real estate security markets for the sample period 01/1994 to 12/2007. The empirical results suggest that equity markets close to the US have a statistically significant response to US monetary policy shocks. The estimated impact of US monetary policy is heterogeneous across countries but statistically significant at the aggregate level in equity and bond markets. The aggregate impact (in absolute terms) and the ‘goodness of fit’ of US monetary policy on international equity and real estate security markets seems to be increasing over time.

JEL Classification: G12, G15, E44

Keywords: International asset pricing, monetary policy, Identification through Heteroskedasticity, recursive Mean Group estimation, bootstrap inference

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1 Introduction

The relationship between monetary policy in the United States (US) and asset prices has recently attracted the attention of economists. The relationship has already captured momentum at a national level since understanding the responsiveness of US asset prices to US monetary policy allows policymakers and market participants to build appropriate models for the distribution of risk bearing. In a world of financially interconnected markets, a relationship between US monetary policy and international asset prices may also exist. In this context, international asset prices may presumably respond not only to their domestic monetary policy shocks but also to US monetary policy shocks. This suggests that monetary policy authorities around the world would be able to take better decisions if they have an accurate understanding of the response of their asset markets to a monetary shock in the US. For instance, it would allow policy makers to devise coherent and coordinated risk management strategies in light of adverse contagion effects that may arise from international financial crises.

In this article we contribute to the understanding of the empirical relationship between US monetary policy and international equity, bond and real estate security markets for the sample period 01/1994 to 12/2007. Our empirical analysis is based on the recent methodology of Identification through Heteroskedasticity (IH) proposed by Rigobon and Sack (2004) as well the more time-honored Event Study (ES) approach in a heterogeneous panel framework. We argue that a US monetary policy shock might have different influences on international asset markets (e.g. due to financial integration, distance, distinct monetary policy frameworks, prudential rules, etc), thus, it is preferable to assume heterogeneity in the relationship and estimate the aggregate impact accordingly.

Previous studies have not focused on analyzing the impact of US monetary policy on international asset prices using recursive estimation at the panel dimension. In this study we propose recursive analysis to understand whether the aggregate impact and the ‘goodness of fit’ of US monetary policy is time-varying across international asset markets. The latter approach is in line with recent empirical and theoretical evidence that international market integration may be time-varying (Bekaert and Harvey, 1995; DeRoon and DeJong, 2005; Pavlova and Rigobon,

2009; Pachenko and Wu, 2009). Moreover, it is also plausible that monetary policy shocks may have asymmetric effects on asset markets due to a much larger impact during bear markets, recessions or times of tight credit market conditions (Chen, 2007; Basistha and Kurov, 2008). From an econometric perspective, recursive analysis via heterogeneous panels may also safeguard the analyst against structural changes present in the data and, thus, prevent her from arriving at spurious conclusions that may arise from single markets or fixed time periods.

It is also noteworthy that the empirical relationship between asset prices and monetary policy has been mostly analyzed with respect to equity and bond markets while leaving aside real estate security markets. The importance of understanding the behavior of real estate assets can be induced from the fact that total real estate accounts for about half of the world's wealth (Corgel et al., 2000). We partially try to overcome the general lack of attention to the latter asset class by accounting for indices of real estate securities at the country level. Moreover, studies that analyze the empirical relationship between monetary policy and bond markets have focused almost exclusively on bond *yields* as opposed to long term bond *returns* (Rigobon and Sack, 2004; Ehrmann et al., 2005; Hausman and Wongswan, 2006). In this study we use long term bond market indices to shed light on the empirical relationship between long term bond returns (i.e. changes in log bond prices) and US monetary policy shocks.

The empirical nature of our approach originates from the absence of a broad consensus on the stylized facts of the relationship between asset prices and monetary policy. The difficulty in arriving to such a consensus stems from the fact that, already at national level, the relationship may well depend on the specification of the model used, and in some cases, it is not transparent whether targeting asset prices may have desirable effects (Bernanke and Woodford, 1997; Bullard and Schaling, 2002; Geromichalos et al., 2007). Moreover, if the central bank's goal is price stability, and this is interpreted as stability of the price of current consumption, as opposed to stability of the price of current vs. future consumption, there might be no reason for a central bank to influence asset prices. Focusing on asset prices when setting monetary policy might be only relevant to the extent that they may signal inflationary or deflationary pressures (Bernanke and Gertler, 2000, 2004). On the other hand, other studies suggest that an active reaction of the monetary authority to asset price developments may help prevent bubbles and could even

be welfare improving (Cecchetti et al., 2000; Carlstrom and Fuerst, 2007).

Asset prices should theoretically play, nevertheless, a major role as providers of valuable information on expectations about future discount factors, which suggests that they may have an empirical relationship with the monetary policy rate (Campbell and Shiller, 1987, 1988; Vickers, 1999). Indeed, recent studies have identified a significant response of US asset markets to US monetary policy shocks (Cochrane and Piazzesi, 2002; Rigobon and Sack, 2004). The significant impact of monetary policy on equity prices could be explained by the effect of revisions on forecasted equity risk premia (Bernanke and Kuttner, 2005). It can also be argued that changes in asset prices and changes in the short-term interest rate can affect each other contemporaneously via channels such as through expectations of the future output path and inflation (Rigobon and Sack, 2002; Chadha et al., 2004).

The empirical evidence of the relationship between asset prices and monetary policy found for the US suggests that in a world of highly interconnected financial markets, monetary shocks that affect US asset prices could also affect international asset prices. For instance, foreign investors who hold US assets will be affected by US monetary policy shocks which could in turn affect their investment decisions with respect to foreign and domestic assets. However, while central banks still hold control over inflation, long term bond prices and equity prices are determined by global supply and demand forces (Rogoff, 2006; Bernanke, 2007). Thus, it is not obvious whether (growing) financial integration can strengthen or weaken the effects of monetary policy on asset prices at home or abroad. This issue is of particular importance since capital market rates are one of the most important channels through which monetary policy makers may influence the real economy and inflationary pressures. Therefore, a clearer understanding of the effect of US monetary policy on international asset prices can improve the view on how monetary authorities worldwide could coordinate actions to influence asset markets when necessary.

The relationship between US or domestic monetary policy and international asset prices has already been analyzed but following different empirical approaches. It has been found that international equity prices react negatively in response to interest rate surprises by the US monetary authority, and that the response's variation is mainly related to financial integration

with the US (Wongswan, 2005; Hausman and Wongswan, 2006). Moreover, it seems that asset prices react strongest to domestic monetary policy shocks but that there are also substantial international spillovers between money, bond, equity markets and exchange rates within and between the US and the Euro area (Ehrmann et al., 2005). The evidence on the relationship between monetary policy and asset prices in countries like the United Kingdom (UK), Japan and the European Union (EU) is tenuous particularly in periods after the 1990s (Furlanetto, 2008). Nevertheless, there is recent evidence that certain equity markets in the EU react significantly to monetary surprises of the European Central Bank (ECB) (Bohl et al., 2008). The latter findings suggest that the impact of US (domestic) monetary shocks across (within) border might be country dependent as well as time dependent.

In this study we shed light on the heterogeneity and time dependence of the impact of US monetary policy on international asset markets. To preview some of our main results we find that: (i) the equity markets of Mexico and Canada have a statistically significant response to US monetary shocks hinting at a proximity effect, (ii) the estimated impact of US monetary policy is heterogeneous across countries but statistically significant at the aggregate level in equity and bond markets and (iii) the aggregate impact (in absolute terms) and the ‘goodness of fit’ of US monetary policy on international equity and real estate security markets seems to be increasing over time.

The paper is organized as follows. Section 2 presents the empirical specification analyzed. Section 3 describes the dataset and the econometric methodology used. Section 4 discusses the results of the study. Section 5 concludes with some final remarks.

2 The model

This section presents a baseline model to analyze the empirical relationship between US monetary policy and international asset prices. US monetary policy is taken as the US Federal Reserve’s decisions to change the US short term interest rate (US STIR henceforth). In what follows let equity price $P_{s,t}$ be the price of a domestic all-share equity portfolio at time t . Bond

price $P_{b,t}$ is defined as the price of a domestic long-term government bond. Real estate price $P_{h,t}$ is defined as the price of a domestic portfolio of real estate securities. Moreover, let r_t^* denote the US STIR. To save on notation, we employ $P_t = P_{\bullet,t}$ henceforth to refer to asset $\bullet = s, h, b$ and $P_t^* = P_{\bullet,t}^*$ to refer to the US counterpart of asset $\bullet = s, h, b$. The empirical relationship between US monetary policy and international asset prices is expressed as

$$\Delta p_t = \alpha \Delta r_t^* + \phi' x_t + z_t + \eta_t, \quad (1)$$

where $\Delta p_t = \ln P_t - \ln P_{t-1}$ is the (log) return of the domestic asset, $\Delta r_t^* = r_t^* - r_{t-1}^*$ is the change in the US STIR, x_t is a $k \times 1$ vector of predetermined variables, z_t is an exogenous shock and η_t is the shock to the asset price. The vector x_t could contain, for instance, dynamics of Δp_t and Δr_t^* , equilibrium relationships, exchange return dynamics, inflation dynamics, etc. The variable z_t could capture conditions such as changes in risk preferences, changes in sentiment, liquidity shocks and macroeconomic shocks not captured by x_t (Rigobon and Sack, 2002). In addition, we could expect a-priori a negative value for α in the case of equity prices and real estate security prices since such asset classes should have a negative relationship with discount rates (Campbell and Shiller, 1987, 1988). However, the direction of the relationship between long term bond prices and interest rates is not clear-cut because revisions in expected risk-adjusted bond returns and expected inflation may have offsetting effects (Campbell and Ammer, 1993).

The estimation of (1) faces two major challenges. On the one hand, for certain countries, the variables Δp_t and Δr_t^* could be characterized as endogenous since (log) international asset price changes may be affected by changes in the US STIR and vice-versa. To see this, we may note that in a world of integrated financial markets, we could expect from the relationship $\Delta r_t^* \rightarrow \Delta p_t$, for instance, the causality $\Delta p_t \rightarrow \Delta r_t^*$ (Ehrmann et al., 2005) and thus $\Delta p_t^* \rightarrow \Delta r_t^*$ (Rigobon and Sack, 2002, 2004; Chadha et al., 2004) where $\Delta p_t^* = \ln P_t^* - \ln P_{t-1}^*$. In this set up, if Δr_t^* is endogenous then the OLS estimate of α will be biased. On the other hand, the omission of the variables in x_t or z_t constitutes a severe problem, given that these variables also carry important information that characterize the relationship between international asset prices and US monetary policy. Therefore, a careful identification strategy has to be employed.

Borrowing from Rigobon and Sack (2002, 2004), the following bivariate system of equations is considered to characterize the relationship between Δp_t and Δr_t^* as well as the macroeconomic variable(s) in x_t :

$$\Delta r_t^* = \beta \Delta p_t + \phi_1' x_t + \theta z_t + \epsilon_t, \quad (2)$$

$$\Delta p_t = \alpha \Delta r_t^* + \phi_2' x_t + z_t + \eta_t, \quad (3)$$

where ϵ_t is the monetary policy shock, and z_t is supposed as a common shock. The shocks ϵ_t and η_t are assumed to have no serial correlation and to be independent with each other and with the common shock z_t (Rigobon and Sack (2002, 2004)). Equations (2) and (3) cannot be estimated consistently using OLS due to the presence of simultaneous equations and omitted variables. For instance, in a simpler scenario of $\phi_1 = \phi_2 = 0$, if OLS is used to estimate (3), then it can be shown that one would obtain simultaneity bias if $\beta \neq 0$ and $\sigma_\eta^2 > 0$ and omitted variables bias if $\theta \neq 0$ and $\sigma_z^2 > 0$ where σ_\bullet^2 is used to denote the variance of $\bullet = \eta, z$. The system of equations in (2) and (3) can also be expressed in reduced form:

$$y_t = v_t + u_t, \quad (4)$$

where $y_t = (\Delta r_t^*, \Delta p_t)'$, $u_t = (u_{1t}, u_{2t})'$ and $v_t = \Gamma x_t$ defines the time-varying mean where Γ is a $2 \times k$ matrix of reduced form parameters attached to x_t . The reduced form residuals in u_t contain the contemporaneous parameters of interest, i.e. $u_{1t} = (1 - \alpha\beta)^{-1}[(\beta + \theta)z_t + \beta\eta_t + \epsilon_t]$ and $u_{2t} = (1 - \alpha\beta)^{-1}[(1 + \alpha\theta)z_t + \eta_t + \alpha\epsilon_t]$.

The main parameter under inspection is α which can be estimated consistently by the IH approach introduced in Rigobon (2003) and extended in other empirical studies of monetary policy (cf. Normandin and Phaneuf (2004), Lane and Lutkepohl (2009)). In our context, IH consists in looking at changes in the co-movements of the US STIR and international asset prices when the variance of one of the shocks in the system is known to shift while the parameters of equations (2) and (3) are assumed to remain stable. Following Rigobon and Sack (2004), we use two sub-samples, F ('policy dates') and \tilde{F} ('non-policy dates'). It can be shown that the

assumptions

$$\sigma_{\epsilon,F}^2 > \sigma_{\epsilon,\tilde{F}}^2, \sigma_{\eta,F}^2 = \sigma_{\eta,\tilde{F}}^2, \sigma_{z,F}^2 = \sigma_{z,\tilde{F}}^2, \quad (5)$$

on the variances of the shocks must hold for the IH estimator to be consistent. These conditions imply that the importance of policy shocks is larger in the subsample F . An important point is that the variance of the policy shock must not become infinitely large, but only increase relative to the variances of other periods preceding or following the shock. One may use, for instance, days of the US Federal Open Market Committee (FOMC) meeting and of the Chairman's semi-annual monetary policy testimony to the US Congress (so-called Humphrey Hawkins Report) to identify circumstances in which the conditions (5) are plausible (cf. Rigobon and Sack (2004)).

Let $\Omega_F = E[[u_{1t} \ u_{2t}]' \cdot [u_{1t} \ u_{2t}] | t \in F]$ and $\Omega_{\tilde{F}} = E[[u_{1t} \ u_{2t}]' \cdot [u_{1t} \ u_{2t}] | t \in \tilde{F}]$. The analytical expression for the difference in the latter covariance matrices under the assumptions in (5) can be shown to satisfy:

$$\Omega_D = \Omega_F - \Omega_{\tilde{F}} = \lambda \begin{bmatrix} 1 & \alpha \\ \alpha & \alpha^2 \end{bmatrix}, \quad (6)$$

with $\lambda = (\sigma_{\epsilon,F}^2 - \sigma_{\epsilon,\tilde{F}}^2)(1 - \alpha\beta)^{-2}$. Thus, α can be identified from the change in the covariance matrix since $\alpha = \Omega_{D,12}\Omega_{D,11}^{-1}$, $\alpha = \Omega_{D,22}\Omega_{D,12}^{-1}$ or $\alpha = (\Omega_{D,22}\Omega_{D,11}^{-1})^{-1/2}$. It is also noteworthy that the preceding IH approach is robust also in the case $\beta = 0$, i.e. under no simultaneity. Within the above framework, which is robust under heteroskedasticity, endogeneity and omitted variables, it is possible to devise a simple parameter stability test. More precisely, we can check whether the parameter α is stable within the sample of policy and non-policy dates by testing whether the determinant of the difference in *conditional* covariance of the two subsamples F and \tilde{F} is zero (*DCC* test) or simply by comparing the two estimates $\hat{\alpha} = \hat{\Omega}_{D,12}\hat{\Omega}_{D,11}^{-1}$ and $\hat{\alpha} = \hat{\Omega}_{D,22}\hat{\Omega}_{D,12}^{-1}$ (Rigobon (2000)).

3 Econometric methodology

This section consists of four subsections. In the first subsection the dataset used to carry out the analysis is described. In the second subsection the estimation strategy is discussed. The third and fourth subsections are devoted, respectively, to issues such as the selection of non-policy dates, comparison of estimated impacts, bootstrap inference and the recursive analysis.

3.1 Data

The empirical specification introduced in the previous section to study the response of international asset prices to US monetary policy (US Federal Reserve's decisions to change the US STIR) is examined from a panel perspective. Let $i = 1, \dots, N$ indicate in the following the respective cross sectional entities (countries). The sample covered for our analysis runs from 01/1994 to 12/2007 at the *daily* frequency. The asset price data consists of *daily* Datastream calculated all-share (total market) equity indices ($N = 29$), 10-year (benchmark) government bond indices ($N = 19$), real estate security indices ($N = 14$) and exchange rate data (currency in country i /US dollar) for all the different economies considered. Countries were selected upon data availability for the sample period covered.

Previous studies that analyze the relationship between US monetary and asset prices have used monetary policy rate measures such as the Federal Funds Futures rate (Rigobon and Sack, 2004), the Federal Funds rate (Bernanke and Kuttner, 2005; Cochrane and Piazzesi, 2002) and the 3-month Treasury rate (Rigobon and Sack, 2002). Similar to the Federal Funds Futures rate, the 3-month Treasury rate adjusts daily to capture changes in market expectations about monetary policy over the near term (Rigobon and Sack, 2002). In this study, the 3-month Treasury rate would capture changes in market expectations over the short term when inferring the response of international asset prices to changes in the US STIR. Moreover, as an interesting by-product, the interaction of international asset prices and the US Treasury rate allows us to analyze empirically the 'flight to quality' effect, i.e. the switch between risky (e.g. international equity) and riskless (e.g. US STIR) assets, which has recently attracted the attention of economists (Pavlova and Rigobon, 2009; Pachenko and Wu, 2009). Therefore,

for the purpose of this study, we employ the US 3-month Treasury rate at the *daily* frequency which is also collected from **Datastream**.

The policy dates are taken from the Federal Reserve’s web-site, where the meeting calendar, minutes and statements of the Federal Open Market Committee are published (<http://www.federalreserve.gov/monetarypolicy/fomc.htm>). In order to reduce the adverse effects of ‘outliers’ we concentrate on scheduled policy dates. Overall, a total number of 116 policy dates are employed.

3.2 Estimation methodology

As previously introduced, the impact of US monetary policy on international asset prices is studied via a heterogeneous panel framework. The panel version of the model presented in (4) is given by

$$y_{it} = v_{it} + u_{it}, \quad (7)$$

where $y_{it} = (\Delta r_t^*, \Delta p_{it})'$, contains the change in the US STIR Δr_t^* and the return $\Delta p_{it} = \ln P_{\bullet, it} - \ln P_{\bullet, it-1}$ of asset $\bullet = s, h, b$ for $i = 1, \dots, N$ countries and $t = 1, \dots, T$ daily time periods. The vector $u_{it} = (u_{1t}, u_{2it})'$ contains the reduced form residuals, more precisely, $u_{1t} = (1 - \alpha_i \beta_i)^{-1}[(\beta_i + \theta)z_t + \beta_i \eta_{it} + \epsilon_t]$, $u_{2it} = (1 - \alpha_i \beta_i)^{-1}[(1 + \alpha_i \theta)z_t + \eta_{it} + \alpha_i \epsilon_t]$ and α_i is the parameter of interest. The model for the time-varying mean $v_{it} = \Gamma_i x_{it}$ is specified as

$$v_{it} = v_i + A_{i,1} y_{it-1} + \dots + A_{i,p} y_{it-p} + B_{i,1} f_{it-1} + \dots + B_{i,p} f_{it-p}, \quad (8)$$

where $f_{it} = (\Delta s_{it}, \Delta p_t^*)'$ is a vector of exogenous components containing the country i /US dollar exchange rate return $\Delta s_{it} = \ln S_{it} - \ln S_{it-1}$ and the return on the US equity/bond/real estate index, i.e. $\Delta p_t^* = \ln P_{\bullet, t}^* - \ln P_{\bullet, t-1}^*$ for $\bullet = s, b, h$. Therefore, the lagged variables in y_{it} and f_{it} allow us to control for ‘conditional dependencies’ of international asset markets to (i) the country i equity/bond/real estate return, (ii) the change in the US STIR, (iii) the country i /US exchange rate return and (iv) the US equity/bond/real estate return. Model (7) coupled

with (8) reduces to a vector autoregression with exogenous components of order p for each cross-section member (henceforth VARX(p)). Note that the conditional mean v_{it} accounts for expectations (conditional on information available in time $t - 1$) of the variables contained in y_{it} and f_{it} . Thus, by using an estimate $\hat{u}_{it} = y_{it} - \hat{v}_{it}$ as opposed to the original data in y_{it} allows us to analyze the ‘surprise’ reaction of international asset prices to US monetary policy shocks.

The number of lags in (8) may be chosen as usual from a sequential test of the null $H_0 : p_{max}$ vs. $H_0 : p_{max} - 1$ or via Aikake and Schwarz information criteria. Given the optimal lag \hat{p} , model (7) is estimated and the reduced form residuals denoted \hat{u}_{it} are obtained. The response of the ‘conditionally centered’ (log) returns (\hat{u}_{2it}) of international equity, bond and real estate assets to US STIR changes (\hat{u}_{1t}) is estimated by employing the Instrumental Variable (IV) and Generalized Method of Moments (GMM) estimators proposed in Rigobon and Sack (2004) as well as the popular ES estimator. Specific issues not detailed here may be looked up in the article by Rigobon and Sack (2004). Our approach extends the latter estimation strategy to analyze the empirical relationship between US monetary policy and international asset prices within a heterogeneous panel framework.

In what follows let $\Delta r^* = [\hat{u}'_{1F} \hat{u}'_{1\tilde{F}_i}]'$ and $\Delta p_i = [\hat{u}'_{2iF} \hat{u}'_{2i\tilde{F}_i}]'$ be $(T_F + T_{\tilde{F}_i}) \times 1 = \mathcal{T}_i \times 1$ partitioned vectors containing (conditionally centered) US STIR changes and (log) asset price changes for country i , respectively, for the sample of policy dates $t \in F$ and the sample of non-policy dates $t \in \tilde{F}_i$. The latter $\tilde{F}_i = \{t_{\tilde{1}}, t_{\tilde{2}}, \dots, t_{\tilde{F}_i}\}$ may consist of the days preceding, following or surrounding those included in the former $F = \{t_1, t_2, \dots, t_F\}$. Once the sample of non-policy dates \tilde{F}_i is chosen, the coefficient α_i for $i = 1, \dots, N$ can be estimated via three-stage least squares (3SLS):

$$\hat{\alpha}_{IV,ij} = (\Delta r^{*'} q_j \Delta r^*)^{-1} \Delta r^{*'} q_j \Delta p_i, \quad (9)$$

where $q_j = w_j (w'_j \hat{\Sigma}_i w_j)^{-1} w'_j$ for $j = 1, 2$ with $w_1 = [\hat{u}'_{1F} \quad -\hat{u}'_{1\tilde{F}_i}]'$ or $w_2 = [\hat{u}'_{2iF} \quad -\hat{u}'_{2i\tilde{F}_i}]'$ the two possible vectors of instruments and $\hat{\Sigma}_i = \text{diag} [\hat{\varepsilon}_i^2]$ with $\hat{\varepsilon}_i$ the residuals of the two-stage least

squares regression (2SLS).¹ It is noteworthy that in the case that the number of observations in the sets F and \tilde{F}_i differ, each subsample in Δr^* , Δp_i and w_j has to be divided by the square root of the total number of dates in each particular set. The degree of explanation for the IV regression denoted $R_{IV,i}^2$ is computed as proposed by Pesaran and Smith (1994). The $R_{IV,i}^2$ provides a measure for the ‘goodness of fit’ in the response of international asset prices to US monetary policy.

In addition to IV estimation outlined above, GMM estimation may also be naturally employed given the analytical expression for the moment conditions in (6). The two parameters to be estimated via the GMM approach are α_i , which is the main parameter of interest, and $\lambda_i \equiv (\sigma_{\epsilon,F}^2 - \sigma_{\epsilon,\tilde{F}_i}^2)(1 - \alpha_i\beta_i)^{-2}$, which gives a measure of the heteroskedasticity in the data. Thus, in GMM estimation the parameter vector $\zeta_i = (\alpha_{GMM,i}, \lambda_i)'$ may be obtained. In our context, the GMM estimator for $i = 1, \dots, N$ is given by

$$\hat{\zeta}_{\mathcal{T}_i} = \arg \min_{\zeta_i \in \Phi} b_{\mathcal{T}_i}(\zeta_i)' A_{\mathcal{T}_i} b_{\mathcal{T}_i}(\zeta_i), \quad (10)$$

with Φ the parameter space, $b_{\mathcal{T}_i}(\zeta_i)$ the vector of differences between sample moments and analytical moments and $A_{\mathcal{T}_i}$ a positive definite and possibly random weighting matrix. Moreover, $\hat{\zeta}_{\mathcal{T}_i}$ is consistent and asymptotically Normal under suitable ‘regularity conditions’ (cf. Harris and Matyas (1999)). The J_i -statistic is available in GMM estimation as a measure of the goodness of fit of the model.

In contrast to the IV and GMM estimators, the popular ES estimator addresses the identification problem by focusing on periods immediately surrounding changes in the US STIR. In our context, the ES approach obtains the following estimator for $i = 1, \dots, N$:

$$\hat{\alpha}_{ES,i} = (\hat{u}'_{1F} \hat{u}_{1F})^{-1} \hat{u}'_{1F} \hat{u}_{2iF}. \quad (11)$$

Note that the heteroskedasticity-based estimator converges to the ES estimator if the shift in the

¹Note that the two stage least squares estimator (2SLS) estimator is numerically identical to $\hat{\alpha}_{IV,ij} = (w'_j \Delta r^*)^{-1} w'_j \Delta p_i$ which is the estimator provided in Rigobon and Sack (2004) since there is one independent variable and one instrument.

variance of the policy shocks is infinitely large. The degree of explanation of the ES regression denoted $R_{ES,i}^2$ may also be computed in the usual way from the OLS residuals of regression (11). A Hausmann test denoted \mathcal{H}_i can be used in order to test the null hypothesis $H_0 : \alpha_{IV,i} = \alpha_{ES,i}$. Under the null hypothesis, \mathcal{H}_i is F_{1, T_i-1} -distributed. Rejection of H_0 would indicate that the ES estimator is biased. Alternatively, it could indicate that the variance of the policy shock in the sample of policy dates F is not sufficiently large for near identification to hold (Rigobon and Sack, 2004).

In our panel approach we may regard heterogeneity in the estimators since we could expect different reactions to US monetary policy amongst the countries at hand and the distinct asset classes, i.e. we may assume that $\alpha_{\bullet,i} = \alpha_{\bullet} + \tau_{\bullet,i}$ for $\bullet = IV, ES, GMM$ where $\tau_{\bullet,i}$ is a cross-sectionally *iid* disturbance. Therefore, it is suitable under the latter assumption to aggregate the impact of US monetary policy via the Mean Group (MG) estimator proposed by Pesaran and Smith (1995). In order to account for heteroskedasticity across country estimates we perform a *standardized* MG estimation, i.e.

$$\bar{\alpha}_{\bullet} = N^{-1} \sum_{i=1}^N \hat{\alpha}_{\bullet,i} \hat{\sigma}_{\bullet,i}^{-1}, \quad (12)$$

for $\bullet = IV, ES, GMM$ where the standard errors $\hat{\sigma}_{\bullet,i} = \widehat{\text{Var}}[\hat{\alpha}_{\bullet,i}]^{1/2}$ are robust under heteroskedasticity. It is worthwhile noting that the above estimators rely on consistency of single market estimates and provide a guidance on the ‘average impact’ of US monetary policy on international asset prices under parameter heterogeneity. In addition, we estimate a ‘Mean Group Difference’ (MGD), i.e.

$$\bar{\alpha}_{MGD} = N^{-1} \sum_{i=1}^N (\hat{\alpha}_{IV,i} \hat{\sigma}_{IV,i}^{-1} - \hat{\alpha}_{ES,i} \hat{\sigma}_{ES,i}^{-1}), \quad (13)$$

in order to compare the heteroskedasticity based estimator and the ES estimator at the aggregate level.

3.3 Selection of non-policy dates and comparison of estimated impacts

In the international context the selection of the non-policy dates \tilde{F}_i is complicated in the sense that for certain countries (particularly those of other continents than America) it is difficult to know which set \tilde{F}_i leads to the most informative change in covariance $\Omega_{D,i}$ for IV and GMM estimation. As proposed in the econometrics literature, it is sometimes preferable to resort to data-driven methodologies for the identification and inference of key relationships as it reduces the complexity of empirical measurement (Friedman, 1987; Hendry, 1995; Hoover et al., 2008). Thus, in this study we ‘ask the data’ for the best window of non-policy dates \tilde{F}_i for each i around the policy dates F .

To motivate the selection of non-policy dates, note that the regression model in (1) could be transformed and estimated by OLS once an analyst has access to the true impact parameter α . This transformed regression is supposed to offer, conditional on x_t and z_t the highest accuracy in whitening the data. Transforming the regression model (1) with inefficient or even biased estimates of α is expected to worsen the model’s fitting accuracy. Therefore, competing estimates of α extracted from alternative sets of non-policy dates may be ranked according to an R^2 criterion derived from a transformed model representation. More precisely, the ‘algorithm’ to select the most appropriate set of instruments for each of the countries in the panel and for each of the asset markets consists of:

1. Obtaining a vector of non-policy dates $\tilde{F}_i = (F-1, F-2, \dots, F-h_{\bullet,i}^{max})'$ for $\bullet = IV, GMM$ constructed by stacking different sets of non-policy dates for $h_{\bullet,i}^{max}$ days preceding the policy dates. At each candidate horizon $h_{\bullet,i}$, the estimates $\hat{\alpha}_{\bullet,i}^h$ for $\bullet = IV, GMM$ are obtained and the model

$$\mu_{it} = \phi_i' x_{it} + \gamma_i t + \eta_{it}, \quad (14)$$

is estimated with the original (i.e. *non-centered*) data for $i = 1, \dots, N$ and $t = 1, \dots, T$ daily time periods via OLS where $\mu_{it} = \Delta p_{it} - \hat{\alpha}_{\bullet,i}^h \Delta r_t^*$, $x_{it} = (y'_{it-1}, \dots, y'_{it-p}, f'_{it-1}, \dots, f'_{it-p})'$ and t is a time trend which proxies the common shock.²

²We also constructed a common factor from international asset return data by means of Principal Component

2. Computing the R^2 of regression (14) and choosing the $\hat{h}_{\bullet,i}$ that corresponds to the maximum quantity from the set $\left\{R_{\bullet,i1}^2, \dots, R_{\bullet,ih^{max}}^2\right\}$ to obtain an optimal set of non-policy dates $\tilde{F}_i^* = (F - 1, F - 2, \dots, F - \hat{h}_{\bullet,i})'$.³

The above procedure is intuitive as a set of non-policy dates is chosen which, conditional on the candidate parameter $\hat{\alpha}_{\bullet,i}^h$, the vector x_{it} and the time trend t , minimizes the distance between the (*non-centered*) international (log) asset price change (Δp_{it}) and the change in the US STIR (Δr_t^*). We also experimented with alternative $h_{\bullet,i}^{max}$ in the sequential procedure but found $\hat{h}_{\bullet,i}$ to be mostly between [1,3] for all three asset markets. Therefore, $h_{\bullet,i}^{max} = 3$ was chosen. In order to test whether the set of selected non-policy dates \tilde{F}_i^* satisfies the heteroskedasticity requirement (6), the null hypothesis $H_0 : \Omega_{D,i} = 0$ versus the alternative hypothesis $H_1 : \Omega_{D,i} \neq 0$ is tested by means of a likelihood ratio statistic (Galeano and Pena, 2007). Rejection of the null hypothesis $H_0 : \Omega_{D,i} = 0$ would provide evidence of heteroskedasticity in the sub-samples F and \tilde{F}_i^* and, thus, supports the applied identification scheme.

The ‘auxiliary regression’ in (14) may also be applied with the ES estimator $\hat{\alpha}_{ES,i}$ to obtain a goodness of fit measure denoted $R_{ES,ih}^2$ for consistency of notation. In this way one may discriminate between estimators across the panel by comparing the quantities $R_{IV,ih}^2$, $R_{GMM,ih}^2$ (which correspond to the estimated impacts $\hat{\alpha}_{\bullet,i}$ and horizons $\hat{h}_{\bullet,i}$ for $\bullet = IV, GMM$) and $R_{ES,ih}^2$.⁴ It is also important to note that many exchange markets were closed in countries that are not in the American continent at the time of the monetary policy announcements (around 14:00 US Eastern Time). For this reason, one could argue that the (conditionally) centered returns \hat{u}_{2iF} and $\hat{u}_{2i\tilde{F}_i}$ of country i which is not in the American continent, should be modified to \hat{u}_{2iF+1} (IV, GMM, ES) and $\hat{u}_{2i\tilde{F}_i+1}$ (IV, GMM), respectively, in order to account for the timing effect of the shock. However, it is also plausible that the impact of FOMC announcements on asset markets occurred during overnight trading in European markets so that there might be a small overlap of trading times (Wongswan, 2005). To control for the

Analysis and used it to proxy the common shock. However, experimentations along these lines resulted in qualitatively similar results to the simple model with a time trend.

³Note that the set of non-policy dates \tilde{F}_i^* does not intersect with the set of policy dates F .

⁴Note that the quantities $R_{\bullet,ih}^2$ for $\bullet = IV, GMM, ES$ obtained from (14) are different from the quantities $R_{\bullet,i}^2$ for $\bullet = IV, ES$ which are obtained from the regressions in (9) and (11), respectively.

timing issue in non-American asset markets in our heterogeneous panel framework, we also used the ‘auxiliary regression’ in (14) above with the estimates $\hat{\alpha}_{\bullet,i}^s$ for $\bullet = IV, ES, GMM$, which were obtained by employing \hat{u}_{2iF+s} (IV, GMM, ES) and $\hat{u}_{2i\tilde{F}_i^*+s}$ (IV, GMM) for $s = 0, 1$ and chose the $\hat{\alpha}_{\bullet,i}^s$ which yielded the highest R^2 .

3.4 Bootstrap inference, parameter stability and recursive estimation

Studying the impact of US monetary policy shocks on international asset prices is complicated due to the large amount of factors that may affect the relationship. Nevertheless, in our set up, we control for (i) conditional effects of different explanatory variables in the time-varying mean v_{it} , (ii) a common shock z_t , (iii) regime changes and heteroskedasticity ($\Omega_{D,i}$) and (iv) selection of non-policy dates (\tilde{F}_i). Generally, our analysis is conditional on the set of policy dates and economic states and dynamics surrounding these time points governing the stochastic properties of model estimates and diagnostics. To disentangle random and structural features of the statistical outcomes we adopt resampling techniques to generate pseudo samples of policy dates. With them at hand, we build empirical distributions of the statistical tools employed. The resampling algorithm consists of:

1. Drawing with replacement from the set of policy dates $F = \{t_1, t_2, \dots, t_F\}$ to obtain a new set of bootstrap policy dates $F_b = \{t_{1b}, t_{2b}, \dots, t_{Fb}\}$.
2. Obtaining the ‘optimal’ set of non-policy dates \tilde{F}_{ib}^* for IV and GMM estimation and using \tilde{F}_{ib}^* and F_b to compute bootstrap quantities $\hat{\alpha}_{\bullet,ib}$, $\bar{\alpha}_{\bullet,b}$ for $\bullet = IV, ES, GMM, MGD$.
3. Repeating steps 1 to 2 for 500 times and obtaining bootstrap confidence intervals from the $\gamma/2$ and $(1 - \gamma/2)$ quantile of the bootstrap distribution.

As introduced previously, it is possible within the IH framework to inspect whether the impact of US monetary policy on international asset prices is stable over the sample period analyzed. One simply has to compare the $\hat{\alpha}_{IV,ij}$ estimators obtained with the two set of instruments $w_1 = [\hat{u}'_{1F} - \hat{u}'_{1\tilde{F}_i}]'$ or $w_2 = [\hat{u}'_{2iF} - \hat{u}'_{2i\tilde{F}_i}]'$. A significant difference between these estimates would hint at parameter instability. To study this issue in more detail across the panel,

a ‘Mean Group Difference’ in analogy to (13) is computed with the two different estimates $\hat{\alpha}_{IV,ij}$ obtained with the instruments w_1 or w_2 .

In order to safeguard the overall analysis from possible parameter instability, the impact of US monetary policy on international asset prices is also studied via a recursive *standardized* MG estimation procedure. The latter procedure consists on fixing a window of size $T_{F,min} = 72$ of policy dates which is moved over the subsample $\tau = T_{F,min}, \dots, \tau = T_F$. At each window, the parameter $\hat{\alpha}_{\bullet,i}$ for each i , and subsequently, the MG estimate $\bar{\alpha}_{\bullet}$ are computed for $\bullet = IV, ES, GMM$. The optimal set of non-policy dates \tilde{F}_i^* is also computed at each window as explained in the previous section to obtain the estimates $\hat{\alpha}_{\bullet,i}$ for $\bullet = IV, GMM$. Thus, a vector of MG impacts $\bar{\alpha}_{\bullet,rw} = (\bar{\alpha}_{\bullet,1}, \dots, \bar{\alpha}_{\bullet,RW})'$ is obtained which can be analyzed to understand the magnitude of the impact over different sub-samples. Since the new MG estimate at each window is based on past and new shocks, it also allows us to study the response of international asset markets to US monetary policy over time. Notably, the rolling window approach is consistent with the IH approach under the null hypothesis of a stable impact α_i in every sub-sample.

A rolling window ‘Mean Group R^2 ’ is also computed from $R_{\bullet,i}$ for $\bullet = IV, ES$ obtained from the regressions (9) and (11) along the latter lines i.e, $\bar{R}_{\bullet,rw}^2 = (\bar{R}_{\bullet,1}^2, \dots, \bar{R}_{\bullet,RW}^2)'$. Both rolling window estimates $\bar{\alpha}_{\bullet,rw}$ and $\bar{R}_{\bullet,rw}^2$ may uncover whether the actual degree of market integration is time-varying and whether there are asymmetric effects of monetary policy shocks over time as suggested in recent studies (DeRoos and DeJong, 2005; Pachenko and Wu, 2009; Chen, 2007; Basistha and Kurov, 2008).

4 Results

This section discusses the results of the impact of US monetary policy on international asset prices. Empirical results are reported in Tables 1 to 7. Figures 1, 2 and 3 display graphically the empirical densities of the estimates $\hat{\alpha}_{\bullet,i}/\hat{\sigma}_{\bullet,i}$ for $\bullet = IV, ES, GMM$ along with the recursive results. In what follows, we mostly refer to statistical significance at the 10% level.

insert Tables 1 and 2 around here

4.1 Single market results

Point estimates of equity markets for the sample of policy and non-policy dates between 01/1994 and 12/2007 are displayed in Table 1. The distinct lag selection tests performed (sequential likelihood ratio, AIC and SIC) for the VARX(p) pointed to three lags in most of the countries and the three asset markets so we have set $p = 3$ throughout. The VARX(3) estimation shows that the number of coefficients in the time-varying mean v_{it} that are significantly different from zero at a 10% level ranges from 3 to 10. The latter result puts forward that there are some significant conditional effects from the dynamic variables considered in the time-varying mean v_{it} and thus evidence of significant expectation effects for which one needs to control. Moreover, no serial correlation was diagnosed at the 5% level of significance in the residuals of the VARX(3) in any of the countries according to the Portmanteau statistic \mathcal{P}_i which supports the use of the estimated reduced form residuals. As expected, the GMM and IV approaches usually produce similar point estimates. In contrast, the ES approach yields estimates that are many times different from the heteroskedasticity based estimates in terms of magnitude and precision. Thus, in what follows, comparisons between estimators are mostly done between IV and ES.

The sample of international equity markets consists of 29 units. Amongst the 29 equity markets, 24 (5) show a negative (positive) impact from US monetary policy shocks with IV estimation. The number of positive point estimates increases to 10 when considering ES estimation. A very apparent feature of the results is the heterogeneity in the value of the estimates. Amongst the countries with a positive coefficient, the response to monetary policy estimated via IV (ES) ranges from 0.17 in Norway to 9.78 in Venezuela (0.01 in Switzerland to 3.57 in Venezuela). Amongst those countries with a negative coefficient obtained via IV (ES), the heterogeneity ranges from -0.08 in Germany to -6.65 in Finland (-0.1 in Japan to -2.88 in Mexico). Overall, we find that 13 (2) countries show a statistically significant negative impact at the 10% level according to the bootstrap confidence bands and/or robust standard errors in IV and GMM (ES) estimation, namely: Australia, Canada, Denmark, Finland, Greece, Ireland, Italy, Japan, Mexico, New Zealand, Spain, Sweden and the United Kingdom (Canada and Mexico).

Interestingly, US monetary policy has a statistically significant negative effect on the equity market of Mexico and Canada according to the bootstrap confidence bands of the heteroskedasticity and event study estimators. In fact, Mexico and Canada are the countries with the largest negative response to the Federal Reserve’s policy announcements in the American continent. This result speaks in favor of a proximity effect in the impact of US monetary policy on equity markets across border. The equity markets in Australia, Japan and the United Kingdom also show a significant response to US monetary policy shocks hinting at a size (market capitalization) effect. Another interesting finding is the significant response to US monetary policy in Scandinavian equity markets which is evident in both IV and GMM estimation. The only Scandinavian country with an insignificant impact to US monetary policy estimated via IV or GMM is that of Norway. However, the impact of the Norwegian equity market to US monetary policy is significantly positive at the 10% level with the ES estimator according to the bootstrap confidence bands. Another interesting case is the statistically significant response of the New Zealand equity market to US monetary policy since the latter was the first country in the world to introduce inflation targeting. In fact, most of the equity markets with a statistically significant response to US monetary policy follow inflation targeting. The latter findings show that the response of international equity markets to US monetary policy may be related to distance, financial integration and similar monetary policy frameworks as suggested by previous studies (Wongswan, 2005; Ehrmann et al., 2005; Ehrmann and Fratzscher, 2008).

Selected diagnostic statistics for equity markets are shown in Table 2. The null hypothesis $\Omega_{D,i} = 0$ is rejected at the 5% level in all markets which suggests heteroskedasticity between the sample of policy and non-policy dates and thus supports the IH scheme. The result on the Hausmann test \mathcal{H}_i , shows that we are able to formally reject the hypothesis $\alpha_{IV,i} = \alpha_{ES,i}$ at the 10% level in Denmark, Finland, Ireland, Japan, New Zealand and Sweden. The estimated horizon $\hat{h}_{IV,i}$ ($\hat{h}_{GMM,i}$) in equity markets is usually equal to 1 (3) in countries inside the American continent and equal to 2 (2) in countries located outside the American continent. Interestingly, the R_{ih}^2 of the ‘auxiliary regression’ in (14) is generally higher for IV and GMM estimates in comparison with ES estimates in countries of the American continent while the opposite is true when considering non-American countries. Moreover, the J_i -statistic is statistically insignificant

at the 5% level in 20 out of 29 international equity markets.

insert Tables 3 and 4 around here

As for equity markets, the results for the set of 19 international bond markets show that there are also conditional effects from the dynamic variables included in the time-varying mean v_{it} as shown by the number of coefficients significant at the 10% level (Table 3). Furthermore, there is no evidence of serial correlation in the residuals of the VARX(3). The IV and GMM estimates show that most bond markets, except for Canada and Australia, have a positive impact to US monetary policy. The ES estimator also obtains mostly positive impacts except for a few more negative cases than in IV or GMM. Out of the 17 countries with a positive coefficient 11 (9) show a statistically significant positive impact at the 10% level according to the bootstrap confidence bands and/or robust standard errors in IV (GMM) estimation. The only statistically significant positive coefficient at the 10% level in the case of ES estimation is that of Japan. However, both IV and ES estimators produce similar point estimates for Canada which are found to be negative and statistically significant at the 10% level according to the bootstrap confidence bands. Note that studies that use bond *yields* as opposed to bond *returns* (i.e. changes in log bond prices) have usually found a positive response of bond markets to US monetary policy shocks in a domestic and international setting (Rigobon and Sack, 2004; Ehrmann et al., 2005; Hausman and Wongswan, 2006). In our international setting, the positive response of (log) bond prices to changes in the US STIR might reflect market expectations of lower risk-premia and lower inflation (Campbell and Ammer, 1993).

Similar to equity markets, the null $\Omega_{D,i} = 0$ is rejected in all countries for bond markets at the 5% significance level (Table 4). The null hypothesis $\alpha_{IV,i} = \alpha_{ES,i}$ is rejected at the 10% level according to the Hausmann statistic \mathcal{H}_i in 9 out of the 19 bond markets. The estimated horizons $\hat{h}_{IV,i}$ and $\hat{h}_{GMM,i}$ are equal to 3 days prior to the shock in all countries except for Japan. The $R_{\bullet,ih}^2$ measures for $\bullet = IV, GMM$ from the ‘auxiliary regression’ are either greater than or equal to $R_{ES,ih}^2$ obtained from ES estimates. Moreover, the J_i -statistic is statistically insignificant at the 5% level in 12 bond markets.

insert Tables 5 and 6 around here

As for equity and bond markets, the panel of 14 international real estate security markets shows significant conditional effects in the time-varying mean v_{it} and no evidence of serial correlation (Table 5). Out of the 14 international real estate security markets considered, we obtain that 10 (11) of them show a negative effect to US monetary policy shocks via IV (GMM). The latter count is somewhat lower in ES estimation where we find 8 negative coefficients. There is a statistically significant negative impact in Australia, Denmark and Germany and a statistically significant positive impact in Belgium at the 10% level when considering the bootstrap confidence bands for the estimates obtained from the heteroskedasticity based estimators (IV and GMM). ES estimates are insignificant throughout. The insignificant response of real estate security markets to monetary policy found here confirms findings of previous studies (cf. Furlanetto (2008)). Results on selected diagnostics for real estate security markets are qualitatively similar to those of equity and bond markets (Table 6).

insert Figure 1 around here

4.2 Aggregate results

As can be noted from the preceding discussion, point estimates are generally similar between IV and GMM methods but in many instances dissimilar to those obtained from the ES method. This may be observed in the estimated densities of the standardized estimates $\hat{\alpha}_{\bullet,i}/\hat{\sigma}_{\bullet,i}$ for each of the international asset markets considered under the three different methodologies used (Figure 1). In general, the heteroskedasticity based estimators (IV and GMM) work ‘best’ in relation to the ES estimator in bond markets and similar in equity and real estate security markets according to the R_{ih}^2 measures and the Hausmann statistics \mathcal{H}_i . Overall, IV and GMM estimation produce estimates that are more frequently found statistically significant than ES estimation according to the bootstrap distribution and/or asymptotic (robust) standard errors. Nevertheless, there is one particularly interesting result for heteroskedasticity or event study estimators: Mexico and Canada have a statistically significant negative response to US monetary policy.

insert Table 7 around here

Another clear feature of the preceding findings is that the sensitiveness of the response of asset prices to US monetary policy is country specific in terms of magnitude and precision. The estimates for each country also vary with the considered asset. This confirms our a priori hypothesis of heterogeneity in the postulated panel relationship as well as empirical findings of recent studies (Ehrmann and Fratzscher, 2008). The results of the MG estimates in Table 7 put forward a negative average impact in equity markets, a positive average impact in bond markets and a negative average impact in real estate security markets with all three estimators considered. The average (standardized) impact is statistically significant at the 10% level in equity and bond markets with the IV and GMM estimators according to the bootstrap confidence bands and asymptotic standard errors. In the case of the ES estimator, the average impact is statistically significant at the 10% level in equity markets according to the asymptotic standard error. The negative and statistically significant response of international equity markets to changes in the US STIR provides evidence of the ‘flight to quality’ effect as proposed in recent theoretical and empirical models of the international propagation of shocks (Pavlova and Rigobon, 2009; Pachenko and Wu, 2009).

Interestingly, there is evidence of a statistically significant MGD at the 10% level between IV and ES estimates in all three asset markets considered (*MGD1*). Moreover, we diagnose a statistically significant MGD at the 10% level between the IV estimates $\hat{\alpha}_{IV,ij}$ obtained from the two different sets of instruments w_1 and w_2 in equity and bond markets (*MGD2*) according to the bootstrap confidence bands and asymptotic standard errors. The latter result hints at instability in the empirical relationship between US monetary policy and international equity and bond prices for the sample 01/1994 to 12/2007.

Thus, it seems preferable from the latter findings (high heterogeneity and possible parameter instability) to study the response of international asset prices to US monetary policy by taking different and smaller windows of policy dates to estimate the impact. In order to minimize parameter instability, one could in principle analyze different sub-samples based on information about events in the economies that might have caused structural breaks. However, this would be cumbersome to implement (and possibly misleading) given the high degree of heterogeneity found amongst cross-sectional members and across asset types. Alternatively, as previously

described, we aggregated the impact with a *standardized* rolling window MG procedure which allows us to have an aggregate view of the unstable vs. stable periods and the overall direction of the impact over time.

insert Figures 2, and 3 around here

Figure 2 shows graphically the results of the rolling window MG impact of US monetary policy on international asset prices. We present only the plots of the IV and ES estimators for space considerations, although similar pictures to those of IV can be obtained when employing the GMM estimator. Interestingly, the response of international asset prices to US monetary policy is significantly time-varying at the 10% level for most of the windows in IV estimation in the three asset markets. Moreover, we find that the impact of US monetary policy on international equity and real estate security markets is negative for most of the time windows and it shows a downward trend in IV and ES estimation. With respect to bond markets the relationship is positive for all time windows in IV and ES estimation.

The result on the stability test *MGD2* and the time variability of the (standardized) MG impact suggests that the relationship between US monetary policy and international asset prices might be sample specific. Thus, not considering recursive estimation will be misleading in the sense that one would not obtain a ‘true’ distribution of US monetary policy impacts in international asset markets when focusing only on particular samples. In addition, Figure 3 also shows the rolling window R^2 s computed from the $R_{\bullet,i}^2$, $\bullet = IV, ES$ measures obtained from regressions (9) and (11). The picture clearly shows that the ‘goodness of fit’ in the response of international equity and real estate security markets to US monetary policy has been increasing over time while it shows no clear trend for bond markets. From an econometric perspective, our findings on the time variability of the aggregate impacts and goodness of fit suggest that one should treat distributions over the entire sample as in Figure 1 with caution, as they might change over time. From an economic perspective, our results also confirm recent theoretical propositions and empirical findings of the time-varying dependencies of asset markets (DeRoos and DeJong, 2005; Pavlova and Rigobon, 2009; Pachenko and Wu, 2009). Moreover, the results also hint at the existence of asymmetric effects of US monetary policy on international asset

prices at the cross-sectional level which is in line with recent findings (Chen, 2007; Basistha and Kurov, 2008).

5 Conclusion

In this article we analyzed the empirical relationship between US monetary policy and international asset prices. The possible problems of endogeneity and omitted variables were addressed by using a robust identification strategy proposed by Rigobon and Sack (2004). We extend the latter approach to investigate the empirical relationship between US monetary policy and international asset prices in a heterogeneous panel framework. We consider issues such as controlling for conditional dynamic effects in the postulated relationship, instrument selection in the heterogeneous panel context, recursive estimation and bootstrap inference.

In general, IV and GMM estimation are more appropriate for identification of the impact of US monetary policy on international asset prices than the popular ES estimator at the aggregate level. The results also indicate that, for the entire sample period covered (01/1994 to 12/2007), the Federal Reserve's policy announcements had heterogeneous impacts in international equity, bond and real estate security markets. In international equity markets there is mostly a negative response to US monetary policy. The largest negative and statistically significant coefficients in the American continent are those of Mexico and Canada which suggests a proximity effect to shocks from the US. At the aggregate level, the MG impact of equity markets is negative and statistically significant. In the case of bond markets the relationship appears to be positive and statistically significant at the aggregate level and for most country estimates. Real estate security markets yield at the aggregate level a negative impact to US monetary policy shocks although not statistically significant.

We diagnose evidence of instability in the postulated empirical relationship between US monetary policy and international asset prices. A careful inspection of the MG estimates over time by means of a rolling window procedure shows that the (average) impact is negative for equity and real estate security markets and positive for bond markets. While the time-varying

MG impact appears to be increasing in absolute value over time for the latter two markets, it shows no clear trend for the former. Interestingly, we find that the ‘goodness of fit’ in the response of international equity and real estate security prices to US monetary policy seems to be increasing over time. These results also confirm recent findings of time-varying dependencies in international asset markets and recent evidence on the asymmetric effects of monetary policy shocks on asset prices.

The results of this study are important because they present a general pattern for the response of international asset prices to US monetary policy. The empirical approach allows to quantify and thus compare the effects across countries, asset markets, estimators and sample periods. In this sense, this paper contributes to the understanding of the empirical relationship between US monetary policy and international asset prices. Several opportunities are also raised for further research. For instance, a similar exercise could be done to study the response of international asset prices to ECB policy shocks and comparing the impact with that of the US Federal Reserve. Moreover, one could also analyze the impact of US monetary policy on exchange rates. We leave these issues for future exploration.

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Country	VARX(3)			IV			ES			GMM				
	v_{it}^{no}	$\hat{\alpha}_i$	\mathcal{P}_i	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$	$\hat{\alpha}_i$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$	$\hat{\alpha}_i$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$
Australia	9	44.175	-2.103	1.046	-5.301	-0.891	-0.754	0.597	-1.655	0.112	-1.948	0.842	-4.213	-0.787
Austria	5	28.491	-0.455	1.140	-2.863	0.859	0.245	0.629	-0.715	1.085	-0.386	0.810	-2.576	0.880
Belgium	5	44.530	0.320	1.286	-1.185	3.274	0.626	0.819	-0.310	2.098	0.562	1.002	-11.948	3.423
Canada	4	29.769	-3.371	2.509	-12.858	-0.291	-1.557	1.267	-3.546	-0.027	-4.762	2.086	-9.844	-1.762
Chile	5	36.238	0.755	1.967	-2.751	2.818	1.008	1.192	-0.819	2.218	0.387	1.360	-2.775	2.706
China	7	51.914	3.280	4.470	-3.653	16.377	-0.964	1.777	-3.674	0.757	0.645	2.212	-4.274	10.461
Colombia	5	20.803	-1.170	1.398	-5.353	1.342	-0.770	0.643	-1.672	0.046	-0.740	0.792	-3.731	1.087
Denmark	8	30.364	-2.672	1.283	-4.753	-0.933	-1.037	0.898	-2.019	0.703	-2.562	0.855	-4.284	-0.911
Finland	6	29.203	-6.646	3.079	-12.987	-2.779	-2.227	2.171	-4.381	1.361	-4.220	1.841	-8.279	-0.784
France	7	37.356	-0.422	1.335	-2.425	1.178	0.595	0.694	-0.360	1.610	-0.069	0.887	-1.845	1.383
Germany	9	42.715	-0.080	1.497	-2.405	2.015	1.042	0.824	-0.001	2.185	-0.100	2.445	-1.367	1.796
Greece	6	19.250	-2.542	1.892	-5.474	0.133	-0.815	1.401	-2.350	1.376	-2.722	1.179	-6.230	0.206
H. Kong	9	27.627	-0.598	2.651	-3.530	3.595	0.058	1.314	-1.352	2.230	-0.290	1.710	-1.245	5.724
Hungary	10	28.574	-2.471	4.731	-6.997	8.667	-2.138	3.126	-5.355	3.794	-1.635	3.015	-5.910	8.693
India	7	17.879	-0.978	3.600	-11.405	5.100	0.239	1.535	-2.169	2.183	0.360	1.563	-4.085	3.284
Indonesia	6	27.398	-2.686	1.583	-5.611	-0.301	-0.704	1.092	-1.879	1.158	-2.408	1.110	-4.616	-0.075
Italy	6	43.203	-3.073	1.860	-5.939	-0.675	-1.639	1.144	-3.137	0.092	-2.767	1.630	-5.660	-0.730
Japan	5	45.003	-2.085	1.755	-4.929	0.460	-0.117	1.307	-1.480	2.257	-2.610	1.197	-6.460	0.796
Luxembourg	8	28.759	-1.176	2.326	-3.150	4.164	-1.804	1.229	-3.036	0.286	-0.908	1.559	-3.101	4.665
Mexico	6	24.207	-6.010	4.234	-26.780	-0.126	-2.876	2.142	-6.373	-0.523	-5.486	3.118	-29.415	-0.128
N. Zealand	3	22.443	-2.246	1.246	-4.695	-0.631	-0.650	0.819	-1.539	0.601	-2.304	0.872	-5.084	-0.863
Norway	5	36.112	0.165	1.265	-1.860	1.780	0.993	0.686	0.221	2.105	0.392	0.844	-1.486	2.007
Peru	7	26.409	-0.435	2.506	-5.191	5.124	-0.783	1.230	-2.365	0.993	-2.235	1.469	-5.138	1.658
Spain	9	26.333	-2.625	1.771	-6.347	-0.326	-0.510	1.029	-1.742	1.032	-1.750	1.596	-4.722	0.370
S. Africa	5	14.842	-0.397	2.361	-2.504	5.837	-0.483	1.532	-1.911	2.344	-0.384	1.483	-2.484	4.240
Sweden	5	32.383	-3.284	1.657	-6.848	-1.329	-1.046	1.092	-2.421	0.553	-2.983	1.374	-6.725	-1.090
Switzerland	8	46.450	-0.734	1.438	-2.157	2.304	0.011	0.991	-1.062	1.724	-0.389	0.979	-1.201	2.274
U. Kingdom	8	38.752	-2.064	1.251	-4.095	-0.106	-0.680	0.785	-1.507	0.746	-1.009	1.969	-9.385	0.374
Venezuela	8	70.815	9.776	5.812	1.736	40.185	3.570	2.595	0.963	8.080	1.788	1.478	-0.235	39.352

Table 1: Country estimates for equity markets. v_{it}^{no} is the number of coefficients in the time-varying mean v_{it} of the VARX(3) which have robust t -ratios greater than 1.64 in absolute value, \mathcal{P}_i is the Portmanteau test of the null hypothesis of no 6th order serial correlation in the residuals of the VARX(3) which is $\chi^2(48)$ distributed, $\hat{\alpha}_{\bullet,i}$ ($\hat{\sigma}_{\bullet,i}$) is the point estimate (heteroskedasticity robust standard error) for country i with $\bullet = IV, ES, GMM$, $CI_{i,j}$ are the 90% bootstrap confidence bands. Note that the estimate $\hat{\alpha}_{IV,i1}$ is obtained with the set of instruments w_1 . Entries in **bold** are statistically significant at the 10% level according to the asymptotic and/or bootstrap distribution.

Country	$\Omega_{D,i}$			IV			ES			GMM		
	\mathcal{G}_i	\mathcal{H}_i	p -val	\hat{h}_i	$R_{\hat{h}_i}^2$	$R_{\hat{h}_i}^2$	\hat{h}_i	$R_{\hat{h}_i}^2$	$R_{\hat{h}_i}^2$	J_i -stat	p -val	
Australia	64.359	2.463	0.118	2	0.276	0.283	2	0.269	0.545	0.460		
Austria	64.275	0.543	0.462	2	0.092	0.094	2	0.089	0.795	0.373		
Belgium	71.616	0.095	0.758	2	0.097	0.100	2	0.095	3.926	0.048		
Canada	44.910	0.702	0.403	1	0.020	0.021	3	0.021	1.607	0.205		
Chile	60.793	0.026	0.872	2	0.068	0.068	2	0.067	1.648	0.199		
China	45.108	1.070	0.302	1	0.005	0.004	3	0.005	0.398	0.528		
Colombia	35.935	0.104	0.747	1	0.077	0.076	2	0.077	0.467	0.494		
Denmark	63.206	3.181	0.076	2	0.197	0.202	2	0.193	0.856	0.355		
Finland	82.584	4.098	0.044	2	0.144	0.146	2	0.143	3.694	0.055		
France	95.030	0.795	0.373	2	0.115	0.117	2	0.112	4.889	0.027		
Germany	94.409	0.806	0.370	2	0.109	0.111	2	0.107	5.976	0.015		
Greece	69.322	1.845	0.176	2	0.080	0.080	2	0.080	3.898	0.048		
H. Kong	79.136	0.081	0.776	2	0.184	0.185	2	0.183	2.833	0.092		
Hungary	63.107	0.009	0.925	2	0.094	0.094	2	0.094	0.357	0.550		
India	30.816	0.140	0.709	1	0.041	0.040	3	0.041	1.828	0.176		
Ireland	65.102	2.992	0.085	2	0.163	0.166	2	0.160	0.987	0.320		
Italy	81.862	0.957	0.329	2	0.054	0.055	2	0.053	4.745	0.029		
Japan	68.281	0.094	0.094	2	0.134	0.136	2	0.133	1.483	0.223		
Luxembourg	62.471	0.101	0.751	2	0.087	0.088	2	0.086	0.471	0.492		
Mexico	30.856	0.736	0.392	1	0.026	0.025	3	0.026	0.845	0.358		
N. Zealand	65.429	2.884	0.091	2	0.160	0.163	2	0.156	0.019	0.889		
Norway	65.779	0.608	0.436	2	0.116	0.118	2	0.114	2.604	0.107		
Peru	31.615	0.025	0.873	1	0.038	0.037	3	0.038	3.431	0.064		
Spain	83.652	2.153	0.144	2	0.074	0.074	2	0.073	7.539	0.006		
S. Africa	64.435	0.002	0.962	2	0.125	0.126	2	0.123	4.263	0.039		
Sweden	80.682	3.225	0.074	2	0.124	0.126	2	0.122	2.471	0.116		
Switzerland	81.809	0.511	0.476	2	0.099	0.100	2	0.097	4.248	0.039		
U. Kingdom	94.371	2.017	0.157	2	0.131	0.134	2	0.128	10.145	0.001		
Venezuela	46.867	1.424	0.234	1	0.052	0.052	3	0.052	2.039	0.153		

Table 2: Selected statistics for equity markets. \mathcal{G}_i are the statistics from the likelihood ratio test of the null $H_0 : \Omega_{D,i} = 0$ which is $\chi^2(3)$ distributed, \mathcal{H}_i is the Hausmann test of the null $H_0 : \alpha_{IV,i} = \alpha_{ES,i}$ and corresponding horizon of non-policy dates via the $\bullet = IV, GMM$ methodology, $R_{\hat{h}_i}^2$ is the degree of explanation from the ‘auxiliary regression’ in (14) with $\hat{\alpha}_{\bullet,i}$ for $\bullet = IV, GMM, ES$, J_i -stat is the statistic of the J -test in GMM estimation and corresponding p -value. Entries in **bold** are statistically significant at the 10% level.

Country	VARX(3)			IV			ES			GMM				
	v_{it}^{no}	\mathcal{P}_i	$\hat{\alpha}_{i,1}$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$	$\hat{\alpha}_i$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$	$\hat{\alpha}_i$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$
Australia	9	40.963	-0.495	0.642	-1.616	0.411	-0.242	0.375	-0.737	0.274	-0.454	0.450	-1.631	0.411
Austria	6	18.765	1.157	0.425	0.623	2.500	0.229	0.256	-0.240	0.475	0.860	0.294	0.445	1.580
Belgium	9	36.888	0.707	0.453	0.088	1.883	0.017	0.316	-0.478	0.343	0.709	0.377	-0.034	1.811
Canada	2	23.688	-1.411	1.148	-4.469	-0.241	-1.654	1.063	-3.425	-0.737	-0.694	0.837	-14.820	0.552
Denmark	9	35.595	0.934	0.523	0.287	2.785	0.078	0.243	-0.259	0.473	0.552	0.345	-0.126	1.443
Finland	7	44.514	0.612	0.542	-0.116	1.802	-0.038	0.395	-0.623	0.388	0.474	0.355	-0.211	1.349
France	9	18.712	0.794	0.547	0.061	2.125	0.103	0.364	-0.416	0.561	0.710	0.372	0.078	1.754
Germany	8	21.086	0.949	0.438	0.433	2.354	-0.008	0.277	-0.453	0.264	0.483	0.273	0.044	0.975
Ireland	8	28.199	0.612	0.584	-0.115	1.936	0.005	0.387	-0.548	0.469	0.368	0.371	-0.337	1.258
Italy	7	30.095	0.907	0.555	0.170	2.770	0.095	0.290	-0.363	0.507	0.445	0.360	-0.265	1.339
Japan	6	19.969	1.316	0.583	0.608	3.354	0.825	0.419	0.370	1.599	0.931	0.412	0.422	2.900
Netherlands	8	28.119	0.315	0.557	-0.438	1.199	-0.038	0.397	-0.600	0.382	0.071	0.341	-0.575	0.742
N. Zealand	6	23.836	1.370	1.039	-0.395	2.319	0.614	0.768	-0.769	1.311	1.291	0.565	-0.591	2.049
Norway	5	49.113	0.507	0.769	-0.384	2.727	-0.065	0.488	-0.663	0.709	0.562	0.594	-0.204	2.742
Portugal	6	18.741	0.981	0.520	0.391	2.540	0.300	0.301	-0.125	0.673	1.000	0.434	0.335	2.385
Spain	10	20.383	0.709	0.565	0.039	2.108	-0.148	0.352	-0.762	0.247	0.715	0.445	-0.103	1.940
Sweden	5	59.939	1.104	0.630	0.303	2.519	0.351	0.350	-0.208	0.804	0.822	0.440	0.176	1.632
Switzerland	7	27.049	0.848	0.450	0.354	2.371	0.225	0.251	-0.094	0.617	0.857	0.446	0.290	2.595
U. Kingdom	4	33.536	0.441	0.753	-0.447	1.758	-0.250	0.564	-1.011	0.338	0.826	0.543	-0.999	0.886

Table 3: Country estimates for bond markets. v_{it}^{no} is the number of coefficients in the time-varying mean v_{it} of the VARX(3) which have robust t -ratios greater than 1.64 in absolute value, \mathcal{P}_i is the Portmanteau test of the null hypothesis of no 6th order serial correlation in the residuals of the VARX(3) which is $\chi^2(48)$ distributed, $\hat{\alpha}_{\bullet,i}$ ($\hat{\sigma}_{\bullet,i}$) is the point estimate (heteroskedasticity robust standard error) for country i with $\bullet = IV, ES, GMM$, $CI_{i,j}$ are the 90% bootstrap confidence bands. Note that the estimate $\hat{\alpha}_{IV,i}$ is obtained with the set of instruments w_1 . Entries in **bold** are statistically significant at the 10% level according to the asymptotic and/or bootstrap distribution.

Country	$\Omega_{D,i}$			IV		ES		GMM			
	\mathcal{G}_i	\mathcal{H}_i	p -val	\hat{h}_i	R_{ih}^2	R_{ih}^2	R_{ih}^2	\hat{h}_i	R_{ih}^2	J_i -stat	p -val
Australia	48.331	0.237	0.627	3	0.387	0.387	0.387	3	0.387	0.249	0.617
Austria	62.732	7.473	0.007	3	0.150	0.149	0.128	3	0.150	2.311	0.128
Belgium	53.820	4.528	0.034	3	0.103	0.102	0.848	3	0.103	0.037	0.848
Canada	66.606	0.316	0.575	3	0.003	0.003	0.001	3	0.003	12.012	0.001
Denmark	69.246	3.413	0.066	3	0.244	0.243	0.041	3	0.244	4.197	0.041
Finland	54.771	3.068	0.081	3	0.077	0.077	0.409	3	0.077	0.682	0.409
France	53.655	2.848	0.093	3	0.068	0.068	0.694	3	0.068	0.155	0.694
Germany	65.849	7.984	0.005	3	0.164	0.163	0.009	3	0.164	6.767	0.009
Ireland	53.906	1.935	0.166	3	0.055	0.054	0.160	3	0.055	1.970	0.160
Italy	73.710	2.946	0.087	3	0.101	0.100	0.018	3	0.101	5.625	0.018
Japan	84.403	1.460	0.228	2	0.032	0.032	0.041	2	0.032	4.188	0.041
Netherlands	51.541	0.816	0.367	3	0.065	0.064	0.209	3	0.065	1.579	0.209
N. Zealand	53.533	1.171	0.280	3	0.337	0.337	0.696	3	0.337	0.153	0.696
Norway	53.324	0.929	0.336	3	0.121	0.120	0.746	3	0.121	0.105	0.746
Portugal	53.122	2.591	0.109	3	0.052	0.052	0.911	3	0.052	0.013	0.911
Spain	53.086	3.772	0.053	3	0.110	0.110	0.965	3	0.110	0.002	0.965
Sweden	71.576	2.069	0.152	3	0.066	0.066	0.019	3	0.066	5.510	0.019
Switzerland	60.358	2.784	0.097	3	0.077	0.003	0.193	3	0.077	1.696	0.193
U. Kingdom	58.781	1.918	0.167	3	0.033	0.033	0.002	3	0.033	9.958	0.002

Table 4: Selected statistics for bond markets. \mathcal{G}_i are the statistics from the likelihood ratio test of the null $H_0 : \Omega_{D,i} = 0$ which is $\chi^2(3)$ distributed, \mathcal{H}_i is the Hausmann test of the null $H_0 : \alpha_{IV,i} = \alpha_{ES,i}$ and corresponding p -value, $\hat{h}_{\bullet,i}$ is the estimated horizon of non-policy dates via the $\bullet = IV, GMM$ methodology, $R_{\bullet,ih}^2$ is the degree of explanation from the ‘auxiliary regression’ in (14) with $\hat{\alpha}_{\bullet,i}$ for $\bullet = IV, GMM, ES$, J_i -stat is the statistic of the J -test in GMM estimation and corresponding p -value. Entries in **bold** are statistically significant at the 10% level.

Country	VARX(3)			IV			ES			GMM				
	v_{it}^{no}	\mathcal{P}_i	$\hat{\alpha}_{i1}$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$	$\hat{\alpha}_i$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$	$\hat{\alpha}_i$	$\hat{\sigma}_i$	$CI_{i,1}$	$CI_{i,2}$
Australia	4	61.975	-1.398	1.179	-3.657	0.194	-0.830	0.780	-1.938	0.177	-1.873	0.866	-10.944	-0.418
Austria	5	38.457	-1.141	1.115	-3.695	0.077	0.052	0.445	-0.565	0.724	-0.247	0.474	-3.488	0.161
Belgium	5	38.356	2.000	1.283	0.335	6.228	0.769	0.829	-0.334	2.121	2.466	1.177	0.293	26.160
Canada	7	23.982	-1.500	2.806	-4.184	4.874	-0.748	1.893	-2.570	2.827	-1.776	1.737	-4.993	4.475
Denmark	6	20.081	-2.792	2.206	-9.216	-0.574	-1.579	1.567	-4.478	0.012	-0.527	0.599	-6.231	0.511
France	5	32.600	0.141	1.323	-1.674	2.723	0.412	0.841	-0.617	1.824	0.140	0.945	-1.922	2.551
Germany	4	50.743	-5.035	3.642	-15.849	-1.477	-1.942	2.107	-5.715	0.021	-4.783	3.185	-14.830	-0.855
Italy	7	22.119	-1.090	2.423	-4.958	2.681	-1.005	1.502	-3.295	0.929	-0.742	1.422	-4.945	4.291
Japan	7	30.001	-3.641	3.515	-8.615	2.405	0.875	2.819	-1.911	6.148	-2.369	2.432	-7.259	3.173
N. Zealand	10	33.459	2.252	1.497	-0.232	4.967	1.141	0.795	-0.223	1.890	3.205	1.802	-1.894	4.715
Norway	8	29.861	-0.844	1.933	-4.234	1.888	-0.233	1.127	-1.598	1.392	-0.777	1.393	-4.118	1.948
Spain	7	25.836	0.059	1.814	-2.580	3.417	0.197	1.149	-1.175	1.829	-0.110	1.238	-2.608	2.647
Sweden	5	27.595	-1.663	2.472	-4.003	2.791	-0.189	1.823	-2.081	3.038	-1.451	1.598	-5.502	2.629
U. Kingdom	3	49.139	-1.592	2.017	-4.445	2.271	-0.206	1.443	-1.743	2.651	-1.649	1.255	-4.167	2.076

Table 5: Country estimates for real estate security markets. v_{it}^{no} is the number of coefficients in the time-varying mean v_{it} of the VARX(3) which have robust t -ratios greater than 1.64 in absolute value, \mathcal{P}_i is the Portmanteau test of the null hypothesis of no 6th order serial correlation in the residuals of the VARX(3) which is $\chi^2(48)$ distributed, $\hat{\alpha}_{\bullet,i}$ ($\hat{\sigma}_{\bullet,i}$) is the point estimate (heteroskedasticity robust standard error) for country i with $\bullet = IV, ES, GMM$, $CI_{i,j}$ are the 90% bootstrap confidence bands. Note that the estimate $\hat{\alpha}_{IV,i1}$ is obtained with the set of instruments w_1 . Entries in **bold** are statistically significant at the 10% level according to the asymptotic and/or bootstrap distribution.

Country	$\Omega_{D,i}$			IV		ES			GMM		
	\mathcal{G}_i	\mathcal{H}_i	p -val	\hat{h}_i	R_{th}^2	R_{th}^2	\hat{h}_i	R_{th}^2	J_i -stat	p -val	
Australia	64.779	0.412	0.522	2	0.078	0.078	2	0.078	4.461	0.035	
Austria	323.682	1.362	0.244	2	0.011	0.010	2	0.011	1.943	0.163	
Belgium	67.136	1.582	0.210	3	0.021	0.021	3	0.021	9.185	0.002	
Canada	60.774	0.132	0.717	2	0.032	0.031	2	0.032	0.212	0.645	
Denmark	100.857	0.611	0.435	2	0.019	0.019	2	0.019	3.514	0.061	
France	63.674	0.070	0.791	2	0.032	0.032	2	0.032	0.019	0.892	
Germany	66.315	1.084	0.299	2	0.011	0.011	2	0.011	0.299	0.584	
Italy	51.425	0.002	0.964	3	0.018	0.018	3	0.018	2.309	0.129	
Japan	65.149	0.002	0.964	3	0.018	0.018	3	0.018	1.829	0.176	
N. Zealand	98.111	0.766	0.382	3	0.066	0.066	3	0.066	4.102	0.043	
Norway	48.690	0.151	0.698	3	0.033	0.033	3	0.033	0.431	0.512	
Spain	61.786	0.010	0.922	2	0.017	0.017	2	0.017	1.732	0.188	
Sweden	62.815	0.780	0.378	2	0.027	0.027	2	0.028	0.154	0.695	
U. Kingdom	61.797	0.968	0.326	2	0.035	0.035	2	0.035	0.008	0.930	

Table 6: Selected statistics for real estate security markets. \mathcal{G}_i are the statistics from the likelihood ratio test of the null $H_0 : \Omega_{D,i} = 0$ which is $\chi^2(3)$ distributed, \mathcal{H}_i is the Hausmann test of the null $H_0 : \alpha_{IV,i} = \alpha_{ES,i}$ and corresponding p -value, $\hat{h}_{\bullet,i}$ is the estimated horizon of non-policy dates via the $\bullet = IV, GMM$ methodology, $R_{\bullet,th}^2$ is the degree of explanation from the ‘auxiliary regression’ in (14) with $\hat{\alpha}_{\bullet,i}$ for $\bullet = IV, GMM, ES$, J_i -stat is the statistic of the J -test in GMM estimation and corresponding p -value. Entries in **bold** are statistically significant at the 10% level.

MG	Equity				Bonds				Real Estate						
	IV	ES	GMM	MGD1	MGD2	IV	ES	GMM	MGD1	MGD2	IV	ES	GMM	MGD1	MGD2
$\bar{\alpha}_\bullet$	-0.780	-0.330	-0.979	-0.450	-1.252	1.246	0.218	1.408	1.028	1.896	-0.398	-0.076	-0.459	-0.323	-1.065
$\hat{\sigma}_\bullet$	0.178	0.165	0.215	0.126	0.344	0.224	0.177	0.233	0.131	0.338	0.249	0.195	0.312	0.136	1.319
CI_1	-1.419	-1.120	-1.607	-0.759	-2.147	0.479	-0.655	0.285	0.636	0.775	-0.927	-0.535	-1.296	-0.583	-5.289
CI_2	-0.101	0.356	-0.151	-0.108	-0.085	2.136	1.208	2.312	1.576	3.269	0.063	0.433	0.129	-0.120	0.307

Table 7: Mean Group estimates per asset market. $\bar{\alpha}_\bullet$ ($\hat{\sigma}_\bullet$) is the *standardized* Mean Group estimator (asymptotic standard error) via $\bullet = IV, ES, GMM$ methodology and the ‘Mean Group Difference’ estimators $\bullet = MGD1, MGD2, CI_j$ are the 90% *standardized* Mean Group bootstrap confidence bands. Note that *MGD1* refers to the Mean Group Difference between the *IV* estimates and the *ES* estimates and *MGD2* refers to the Mean Group Difference between the *IV* estimates with the instruments w_1 and w_2 . Entries in **bold** are statistically significant at the 10% level according to the asymptotic and/or bootstrap distribution.

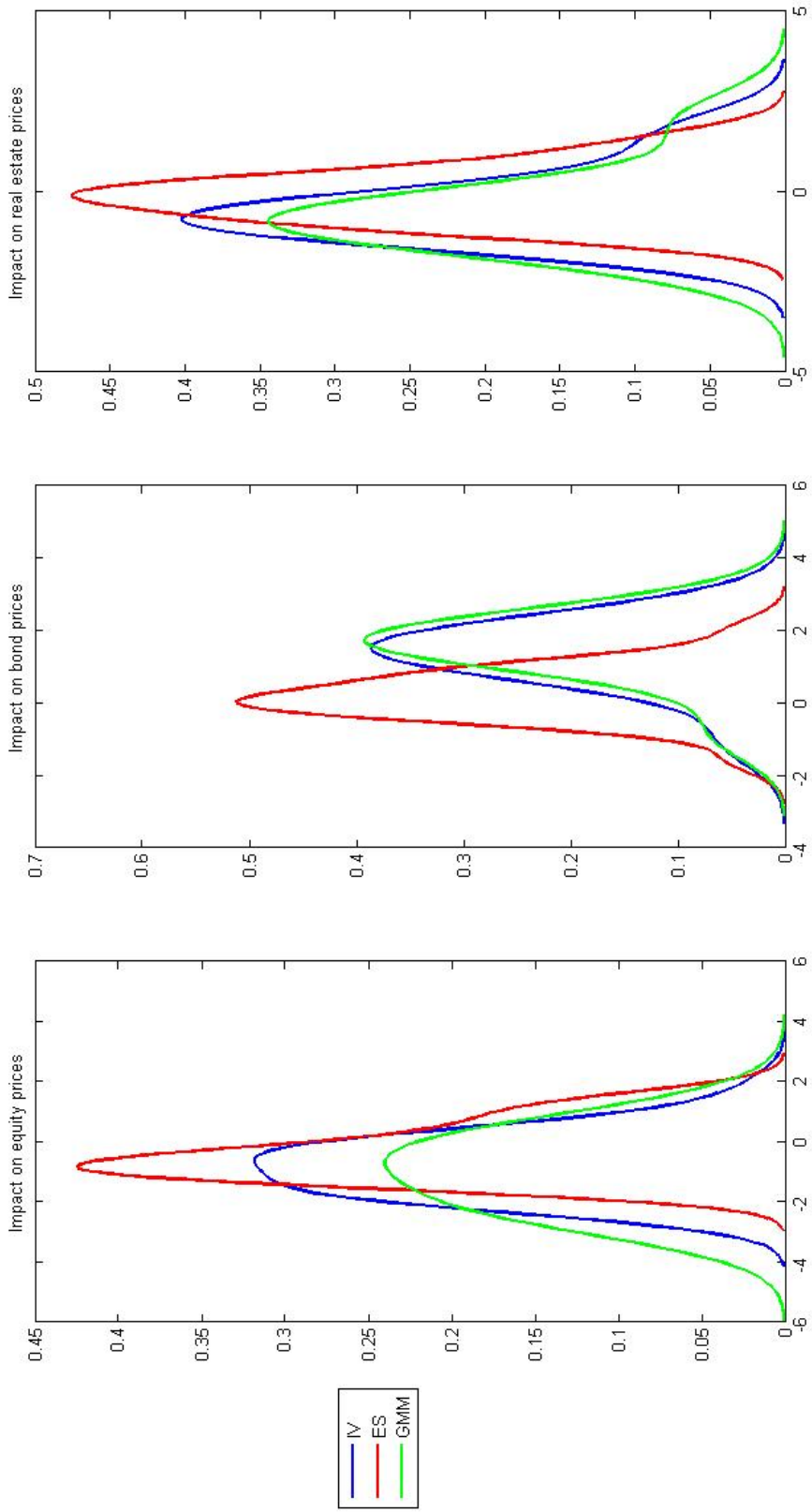


Figure 1: Densities of the impact of US monetary policy on international asset prices. Note: densities are estimated from the standardized estimates $\hat{\alpha}_{\bullet,i}/\hat{\sigma}_{\bullet,i}$ for $i = 1, \dots, N$ and $\bullet = IV, ES, GMM$ via a Kernel function.

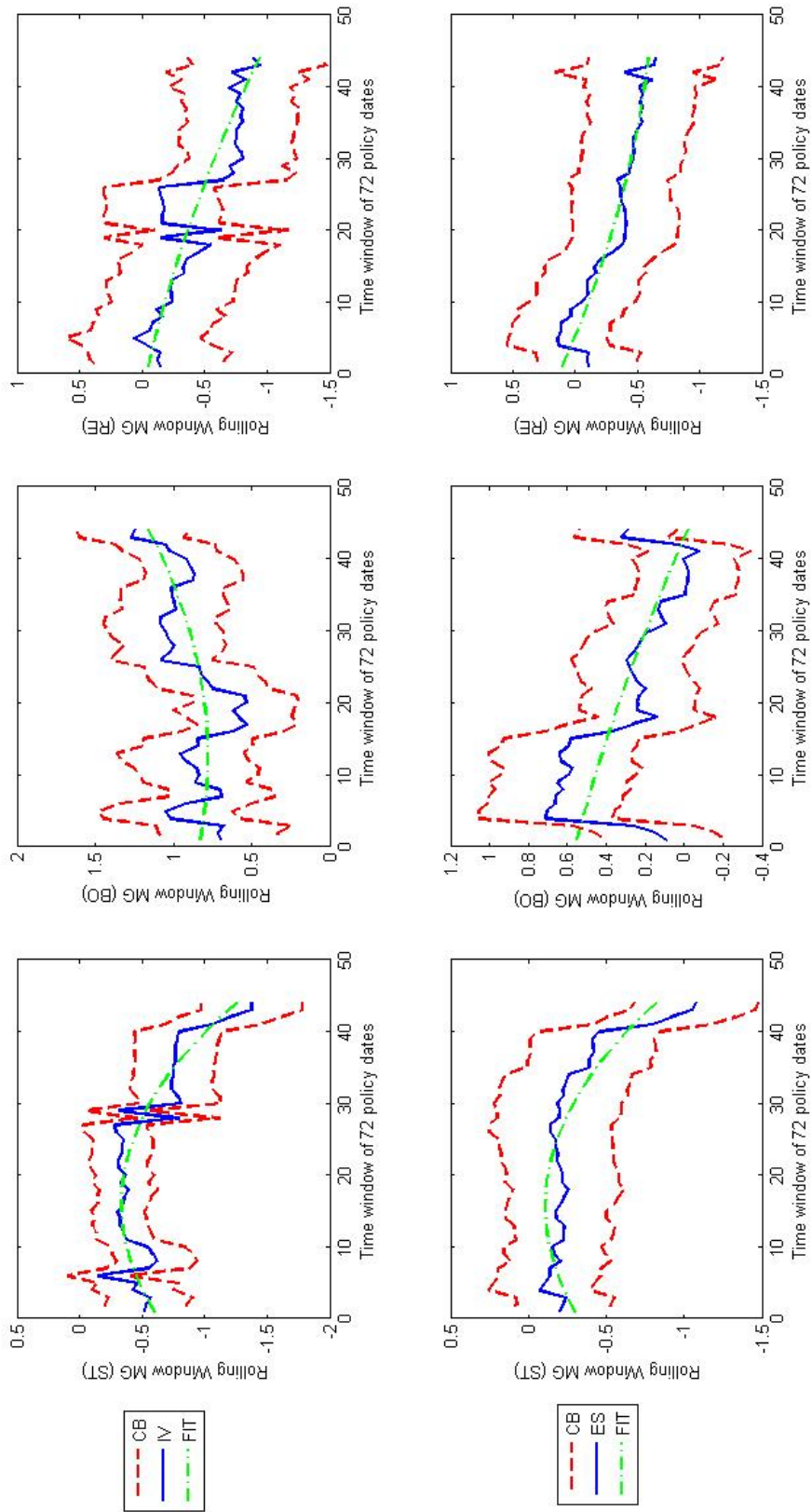


Figure 2: Rolling window Mean Group impact of US monetary policy on international asset prices. Note: *CB* (dashed red lines) denote the asymptotic 90% confidence bands, *FIT* (dashed and dotted green line) is a polynomial fit of the IV and ES rolling window Mean Group impacts.

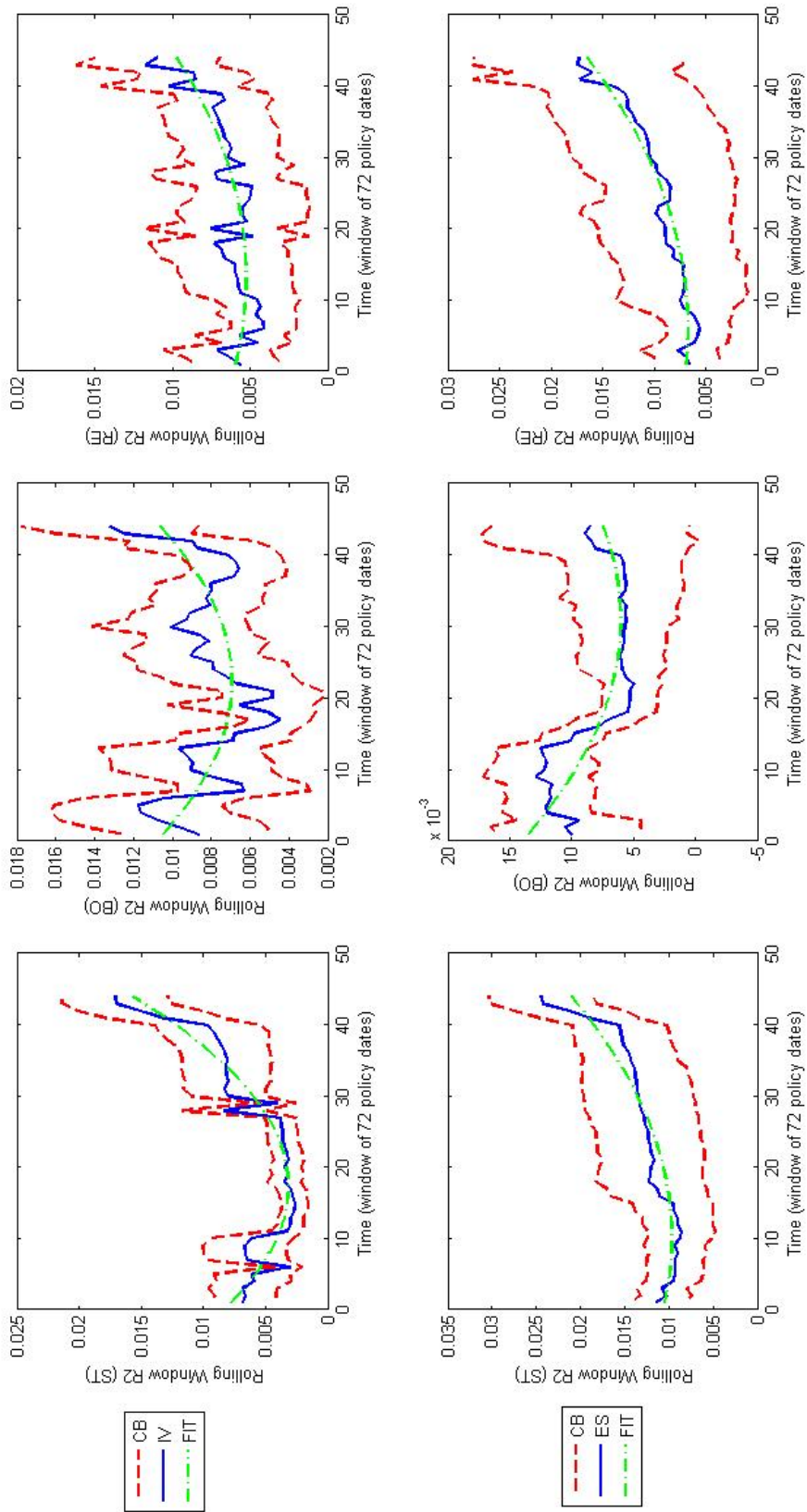


Figure 3: Rolling window Mean Group R^2 of the impact of US monetary policy on international asset prices. Note: *CB* (dashed red lines) denote the asymptotic 90% confidence bands, *FIT* (dashed and dotted green line) is a polynomial fit of the IV and ES rolling window Mean Group R^2 -s.