

# An empirical analysis of the relationship between US and Colombian long-term sovereign bond yields\*

Alexander Guarín, José Fernando Moreno and Hernando Vargast

## Abstract

We study two issues: (i) the relationship between interest rates on US and Colombian sovereign debt and (ii) the short-term response of the Colombian long-term bond yield and other asset prices to shocks to the US long-term Treasury rate. We use daily data between 2004 and 2013. Separating the period into three intervals (before, during and after the financial crisis), we consider the first issue with a moving window linear regression, and we address the second by estimating a VARX-MGARCH model. Our findings show that the link between sovereign bond yields has changed over time, and that the short-run responses of local asset prices to foreign financial shocks were qualitatively different in the three periods. The role of US Treasury securities as a safe haven asset during highly volatile periods seems to be at the root of these changes.

Keywords: Long-term bond yields, global financial crisis, emerging markets, moving window linear regression, VARX-MGARCH model

JEL classification: C30, E43, E58, F42, G15

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† Researcher (aguarilo@banrep.gov.co), Professional Economist (jmorengu@banrep.gov.co) and Technical Deputy Governor (hvargahe@banrep.gov.co) of the Bank of the Republic.

## 1. Introduction

During the recent global financial crisis and economic slowdown, central banks in advanced economies pushed down their monetary policy rates to boost the economy and to prevent a deeper recession. The policy rate reached the zero lower bound, preventing further monetary stimulus with that policy tool (Doh (2010) and Chen et al (2012)).

In consequence, central banks adopted unconventional measures to add further stimulus.<sup>1</sup> For instance, the Federal Reserve has since late 2008 issued “forward guidance” that informs the public about the future path of monetary policy to influence expectations. At that time it also implemented a quantitative easing (QE) program in which it buys long-term Treasury bonds and mortgage-backed securities.<sup>2</sup> The QE policy has led to a reduction of the net supply of long-term bonds, raising their prices and lowering their yields. The same measures also boosted other asset prices (eg commodities and equities) and increased market liquidity (Jones and Kulish (2013) and Cronin (2014)).

QE measures in the advanced economies have also spilled over to emerging market economies, lowering yields on their long-term sovereign bonds and generating other spillover effects such as appreciation of the local currency, rapid credit growth, inflationary pressures and booms in asset prices (see García-Cicco (2011); Chen et al (2012); Glick and Leduc (2012); Moore et al (2013); Fratzscher et al (2013); and Londoño and Sapriza (2014)).

The shifts in the yield on local currency long-term sovereign bonds are crucial for the financial market because that yield is a benchmark for the pricing of long-term assets. For example, a reduction encourages lengthening the maturity of credit obligations and undertaking long-term investment projects (see Turner (2014)). Nevertheless, if the benchmark long-term yield stays low for a prolonged period, financial stability risks could arise. For instance, excessive leverage could lead to a credit boom and the overvaluation of long-term assets such as houses and equities (Turner (2013) and Turner (2014)).

Therefore, the relationship between long-term bond yields of emerging market and advanced economies, its evolution over time and the effects of changes in these rates are crucial issues for macroprudential policy, financial stability, government debt management and monetary policy.

Our aim in this paper is two-fold. First, we study the changing relationship between the US and Colombian long-term sovereign bond yield over time. Second, we analyse the response of Colombian asset prices to shocks to the US Treasury yield and how these responses changed during the global financial crisis.

To study the first topic, we employ a moving window linear regression (MWLR) to examine the link between local asset prices and the US long-term Treasury rate. We also use the MWLR to study the relationship between Colombian and US bond

<sup>1</sup> See Bernanke and Reinhart (2004) for a detailed description of instruments used under this policy scheme.

<sup>2</sup> The Federal Reserve also performed the “Operation Twist” in 2011. In this action, the Fed sold short-term Treasury bonds and bought the same class of securities with long-term maturities.

yields while controlling for the sovereign risk premium and expected currency depreciation.

Second, we estimate a VARX-MGARCH model to compute the response of local asset prices to three distinct shocks to the US long-term Treasury yield. The sources of these shocks are changes in global volatility, the Treasury term premium and the stance of monetary policy in United States. The local asset prices we consider are the Colombian long-term sovereign interest rate, the foreign exchange rate, CDS spreads and the stock market index value.

The exercises employ daily data on US and Colombian financial variables between June 2004 and November 2013. For the second exercise, we divide the sample period into three segments – namely before, during and after the global financial crisis – and perform the estimations on each segment.

This paper contributes to the burgeoning literature on monetary policy spillovers in three ways. First, using daily data for financial variables, we model volatility using GARCH processes. Second, by computing results for the pre-crisis, crisis and post-crisis periods, we account for effects derived from periods with distinct economic and financial characteristics that otherwise could be missed. Third, we assess three distinct shocks to the US long-term bond yield to capture any differences in their effects on the response of Colombian asset prices.

Our findings show that the relationship between US and Colombian long-term sovereign bond yields has changed over time. In fact, the sign on this link turned negative between the second half of 2007 and the 1 May 2013 FOMC statements suggesting that the Federal Reserve would begin reducing (tapering) its bond purchases sooner than the markets were expecting. Our results also suggest that since 2008 the importance of the effects of movements of the US long-term Treasury rate on Colombian asset prices has increased. We also find that the short-run responses of both the Colombian interest rate and other local asset prices to shocks to the US long-term bond yield have been qualitatively different, depending on the sample period and the source of the shock. These changes seem to suggest, first, a special role for US Treasuries as a safe haven asset during the global financial crisis period, and second, a subsequent differentiation of local assets.

The remainder of the paper is organised as follows. In Section 2 we present the literature review, and in Section 3 we describe the main stylised facts on the recent evolution of interest rates on US and Colombia sovereign debt. In Section 4, we conduct a moving window linear regression analysis to study the relationship between long-term bond yields over time. Section 5 estimates the short-run responses on local asset prices to shocks to the US Treasury rate, and Section 6 concludes.

## 2. Literature review

A burgeoning segment of the literature has been addressing the effectiveness and spillover effects of the unconventional monetary policy actions taken after the September 2008 Lehman Brothers bankruptcy. One strand of this literature studies the impact of these actions on advanced economies. For instance, Doh (2010), Cúrdia and Woodford (2011), Gagnon et al (2011), Jones and Kulish (2013) and D'Amico and King (2013) examine the effects of the zero policy rate and QE

measures on the US long-term interest rates. They find that announcements about the future path of the monetary policy rate, and especially purchases of assets on a large scale, effectively reduced the long-term interest rate. This reduction reflected mainly a lower term premium.

Cúrdia et al (2012), Glick and Leduc (2012) and Cronin (2014) analyse spillover effects of unconventional measures on macroeconomic aggregates and asset prices of advanced economies. Particularly, Cúrdia et al (2012) explore the QE effects on both US GDP growth and inflation, while Glick and Leduc (2012) and Cronin (2014) examine the same effects on the foreign exchange rate, commodity prices, stocks and government bonds.

Likewise, Ugai (2007), Peersman (2011), Joyce et al (2011) and Schenkelberg and Watzka (2013) analyse the spillover effects of the unconventional monetary policy on advanced economies other than the United States. Ugai (2007) and Joyce et al (2011) highlight the impact of QE measures on the monetary base, aggregate demand and asset prices in Japan and England. Similarly, Lenza et al (2010) describe how the Federal Reserve, the Bank of England and the European Central Bank (ECB) have conducted monetary policy since 2007.

The other strand of this literature studies the spillover effects of QE measures on emerging market economies. For example, Neely (2010), Landau (2011), Chen et al (2012), Glick and Leduc (2012) and Londoño and Sapriza (2014) find that these policy measures not only stimulated the US domestic economy but also affected market expectations and extended global liquidity. The latter led to spillover effects on emerging market economies; the effects included large capital inflows, a boost to a broad range of asset prices, rapid credit growth, appreciation of local currencies and inflationary pressures. Similar results are found by Fratzscher et al (2013), who analyse the impact of each part of the QE program on the US market and across 65 other countries. In addition, Moore et al (2013) and Turner (2013) find that spillover effects also reduced long-term sovereign bond yields in emerging market economies.

This line of the literature also considers the spillover effects of unconventional monetary policy on specific countries. For example, Quispe and Rossini (2011) and Carrera et al (2013) evaluate the effects of QE measures on macroeconomic variables of Peru's economy (eg domestic growth, inflation, international reserves, liquidity and public debt). Carrera et al (2013) highlight that the response of macroeconomic variables is not uniform across each QE episode. García-Cicco (2011) and Barata et al (2013) measure spillover effects of unconventional policies in Chile and Brazil, respectively. They find that the most important consequences of these policies are capital inflows, the appreciation of the currency and a significant increase in credit.

### 3. Stylised facts

In the first part of this section, we illustrate the dynamics of long-term bond yields for the United States as well as for Colombia and other emerging market economies. We also compare the evolution of some financial variables for Colombia with net capital inflows into this economy during recent years.

In the second part, we analyse the changing relationship between US and Colombian interest rates. We divide our sample into three time spans and highlight the main financial characteristics of those periods.

### 3.1 Long-term sovereign bond yield dynamics

For the whole of the period from June 2004 to November 2013, both the 10-year US Treasury rate and the 10-year Colombian sovereign bond yield exhibit a negative trend (Figure 1, panel A). However, the trend varies for shorter intervals within that time span.

Between 2004 and the first half of 2007, the US Treasury yield exhibits a positive slope as a consequence of the increases in the federal funds rate (the US policy rate) to control inflation expectations. From there, bond yields have been decreasing as a result of the expansive monetary policy adopted by the Federal Reserve (ie the policy rate at the zero lower bound, the QE program and Operation Twist) to cope with the global financial crisis. Nonetheless, during short periods in 2009 and 2010, the US Treasury yield corrected upwards after reductions in global risk perceptions.

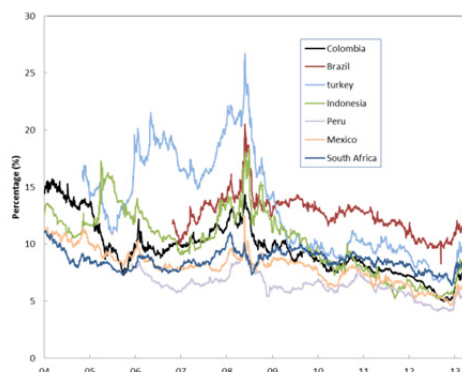
In the Colombian case, the long-term interest rate showed sharp variations through the sample period. Between June 2004 and February 2006 this rate dropped as a result of better fundamentals and external conditions in emerging market economies, and the decline in the Colombian risk spread.<sup>3</sup> From March 2006 to October 2008, the same interest rate rose mainly in response to two facts. First, the increase in Colombia's policy rate to cope with domestic inflationary pressures, and second, the rise of global risk since the mid-2007 beginnings of the global financial crisis. Since October 2008, the long-term sovereign bond yield has been decreasing as a consequence of the spillover effects of the unconventional measures adopted by the United States and other advanced economies.

<sup>3</sup> This reduction is associated with better terms of trade, fiscal consolidation, improvements in security conditions and the deepening of the market for local currency public debt .

(A) US and Colombia



(B) Emerging market economies



Source: Bloomberg.

Ten-year sovereign bond yields for Brazil, Colombia, Indonesia, Mexico, Peru, South Africa and Turkey between June 2004 and November 2013 exhibit similar dynamics and seem to be completely aligned after the Lehman Brothers bankruptcy (Figure 1, panel B). This behaviour appears to be related to the lower risk perception and better terms of trade of emerging market economies.

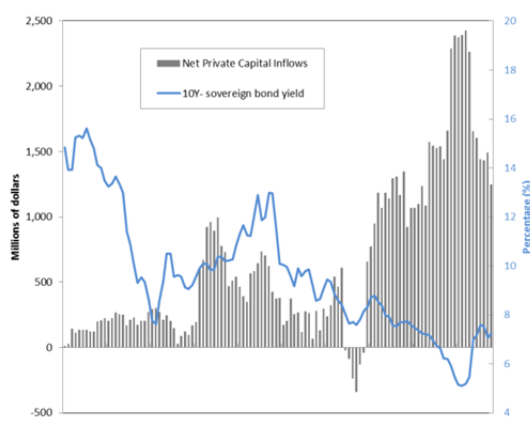
Colombia has recorded net capital inflows<sup>4</sup> during the past decade, with the exception of a short period at the end of 2010 (Figure 2, panels A–D).<sup>5</sup> Furthermore, since 2011 these capital inflows have been larger than in previous years due in part to positive net portfolio investments.

On the other hand, our four financial variables exhibited clear trends along the sample period. In particular, the long-term bond yield, the foreign exchange rate and CDS spreads dropped, while the stock market index rose. The behaviour of these variables before 2007 is explained mainly by local factors in the Colombian market. Between 2007 and 2008, there are specific breaks in the trend of these time series as a consequence of the risk and the economic uncertainty associated with the beginning of the global financial crisis. After this period, these financial variables resumed a decreasing trend. The tendency in this period is associated with the spillover effects of QE measures (see eg García-Cicco (2011), Chen et al (2012), Glick and Leduc (2012), Moore et al (2013), Fratzscher et al (2013) and Londoño and Sapriza (2014)).

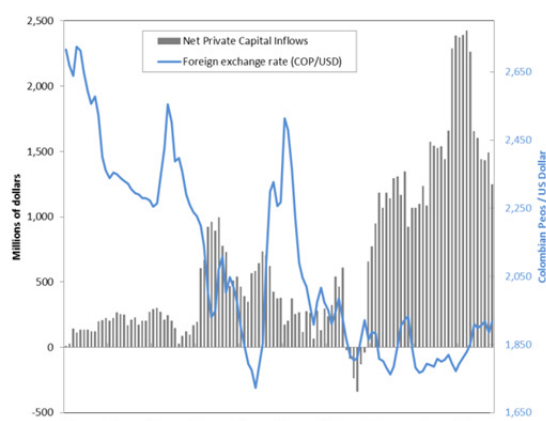
<sup>4</sup> The capital flows include both portfolio and foreign direct investment.

<sup>5</sup> Each panel also draws the sixth-order moving average of monthly net private capital inflows observed in the consolidated exchange balance. The sixth order captures the average of capital inflows during the last half year and hence allows us to avoid excessive volatility in our analysis. The consolidated exchange balance corresponds to current and capital account cash transactions conducted in the US dollar spot market.

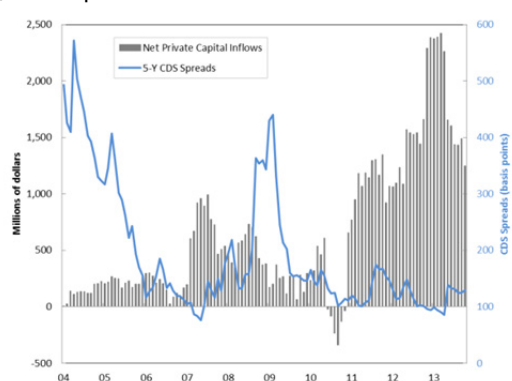
(A) Long-term sovereign bond yield



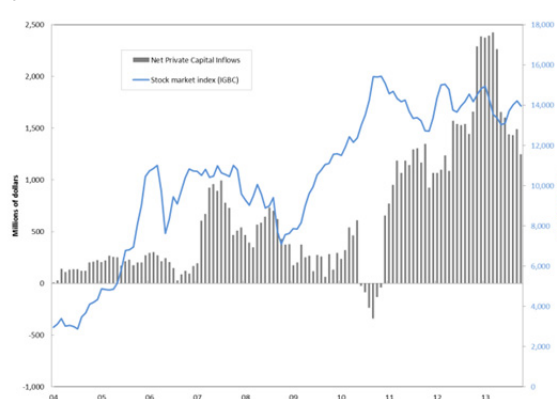
(B) Foreign exchange rate



(C) CDS spreads



(D) Stock market index



Sources: Bloomberg; and Bank of the Republic (central bank of Colombia).

### 3.2 The relationship between US and Colombian bond yields: before, during and after the global financial crisis

Here we divide the June 2004 to November 2013 period into three sub-periods: pre-crisis (June 2004 to February 2007), crisis (February 2007 to October 2009) and post-crisis (November 2009 to November 2013). The dates of each sub-period are chosen to reflect significant changes in the correlation between long-term bond yields arising from financial events or news with a high impact on the market.

#### US monetary policy tightening and developments in the Colombian bond market

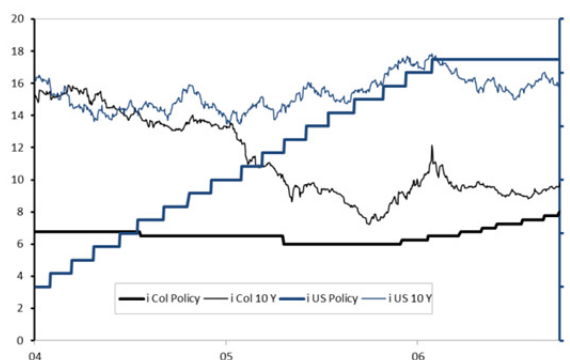
The pre-crisis period (June 2004 to February 2007) is characterised by a positive correlation between long-term bond yields and a non-significant relationship between monetary policy rates (Figure 3, panel A). In this period, the Federal Reserve sharply increased its policy rate to reduce inflationary pressures. Moreover, beginning in July 2006 the US market presented an inverted yield curve, giving an early warning signal of future recession. In the same time span, Colombian long-term bond yields dropped, while the policy rate remained relatively stable. In fact,

banks increased their fixed income portfolios in local sovereign debt and earned exceptional profits as the inflation rate fell.

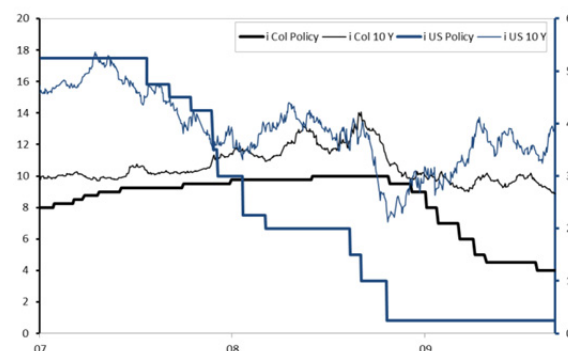
Evolution of US and Colombian sovereign interest rates

Figure 3

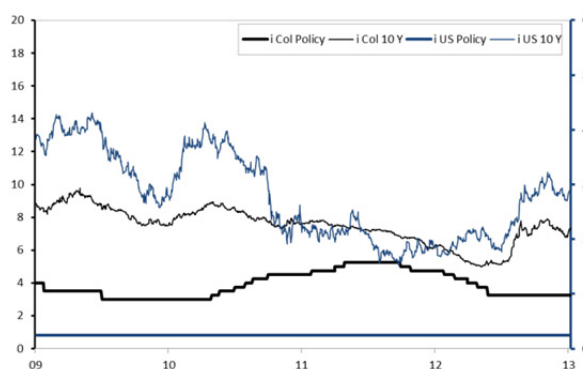
(A) Pre-crisis period



(B) Crisis period



(C) Post-crisis period



Sources: Bloomberg and Bank of the Republic (central bank of Colombia)

### The global financial crisis and the international recession

The crisis period (February 2007 to October 2009) was in general defined by a negative correlation between US and Colombian long-term bond yields, the reduction of the federal funds rate to nearly the zero lower bound, a decreasing trend of the long-term Treasury rate and a large demand for safe assets (Figure 3, panel B). Moreover, the period included several financial events and news with high impact in the market.<sup>6</sup> On the other hand, between 2006 and 2008, the central bank of Colombia raised its policy rate to arrest inflationary pressures and then began reducing the rate at the end of 2008. Since the second half of 2008, the Colombian long-term sovereign bond yield also exhibited a declining trend.

At the end of 2008 the correlation between long-term sovereign bond yields turned positive, in great part as a result of the 15 September collapse of Lehman

<sup>6</sup> For example, in February 2007 HSBC fires to the head of its US mortgage lending division and Freddie Mac announced that it would not buy risky mortgage securities. Furthermore, along 2007 and 2008 financial and real sector companies reported losses associated to the mortgage business, and the Fed warned on its negative effects on the economy.



Brothers. This event increased economic uncertainty as well as risk perceptions in the global financial market. To avoid a financial collapse, the Federal Reserve announced in November 2008 that it would initiate a QE programme (what became the first of three such programmes) in March 2009.

#### QE2, QE3, the Greek debt crisis and tapering

The post-crisis period (November 2009 to November 2013) was characterised by low interest rates in the US and Colombian markets as a result of the US QE programmes, particularly QE2 and QE3, Operation Twist and its spillover effects on emerging market economies (see for example García-Cicco (2011), Glick and Leduc (2012), Moore et al (2013), Fratzscher et al (2013) and Londoño and Sapriza (2014)). In general, the correlation between long-term bond yields after the global financial crisis is positive (Figure 3, panel C). The Greek debt crisis increased the demand for local safe assets in emerging market economies. In May 2013, statements by the Federal Reserve indicated to the market that the central bank's bond purchases would end sooner than expected, which led immediately to the rise of bond yields in both markets.

## 4. The changing relationship between US and Colombian sovereign bond yields

In this section we analyse the relationship between US and Colombian long-term interest rates over time. Our analysis is based on the MWLR.

### 4.1 Data

Our data set considers daily time series of local and foreign financial variables between June 2004 and November 2013. The domestic variables are the 10-year Colombian sovereign bond yield ( $i^{Col}$ ); the Colombian stock market index ( $igbc$ ); sovereign credit default swap (CDS) spreads ( $cds$ ) on five-year<sup>7</sup> Colombian sovereign bonds denominated in US dollars; and the foreign exchange rate denominated as Colombian pesos per US dollar ( $cop$ ) and its expected value for a horizon of 10 years ( $cope$ ).<sup>8</sup> Our foreign variable is the 10-year US Treasury yield ( $i^{US}$ ).

All our econometric exercises employ the logarithm of the Colombian stock market index ( $ligbc$ ), the logarithm of the foreign exchange rate ( $lcop$ ) and its

<sup>7</sup> Ten-year CDS spreads data are not available for our full sample period. Nevertheless, five-year CDS contracts are highly liquid and effectively represent the country risk.

<sup>8</sup>  $Cope$  is calculated as  $cope = cop \times \left( \left[ \frac{rer_t^{trend}}{rer_t} \right] \times \left[ \frac{1+bei_t^{Col}}{1+bei_t^{US}} \right] \right)$

where  $rer_t$  is the real exchange rate;  $rer_t^{trend}$  is its trend; and  $bei_t^{Col}$  and  $bei_t^{US}$  are, respectively, Colombian and US break-even inflation to 10 years. The construction of  $cope$  assumes that agents expect a correction of real exchange rate misalignments. The variables  $rer_t$  and  $rer_t^{trend}$  correspond to the Colombia-US bilateral trade weighted real exchange rate. The real exchange rate trend is computed with the Hodrick-Prescott filter. The data for  $rer$  are from the Bank of the Republic, and those for the remaining variables are from Bloomberg.

expected value for a horizon of 10 years (*lcope*). CDS spreads are expressed as a percentage.

The variable  $i^{Col}$  is constructed with the Nelson-Siegel methodology with data from the Bank of the Republic. Data for variables other than  $i^{Col}$  and *rer* are taken from Bloomberg.

## 4.2 Moving window linear regression (MWLR) analysis

We perform two sets of MWLR exercises to understand how the relationship between US and Colombian long-term bond yields and other local asset prices have changed over time.<sup>9</sup> In particular, these exercises provide evidence on the pattern of this link during the global financial crisis and the US implementation of the QE program.

The MWLR exercises are run with a rolling sample of 435 business days (approximately two years). The first sample period is from 29 January 2003 to 4 January 2005 and the final one is from 28 December 2011 to 7 November 2013.<sup>10</sup> In our estimates, we use a GJR-GARCH(1,1) process to model the variance of errors and thereby take into account the changing volatility commonly found in high-frequency financial data (see Appendix A).

### First exercise

The first MWLR exercise examines the relationship between changes in the US long-term Treasury rate and changes in a Colombian asset price. The latter could be the local long-term sovereign bond yield, the foreign exchange rate, CDS spreads or the stock market index. In particular, we estimate the model stated by

$$\Delta y_{t_k} = \beta_0(k) + \beta_1(k)\Delta i_{t_k}^{US} + a_{t_k} \quad (1)$$

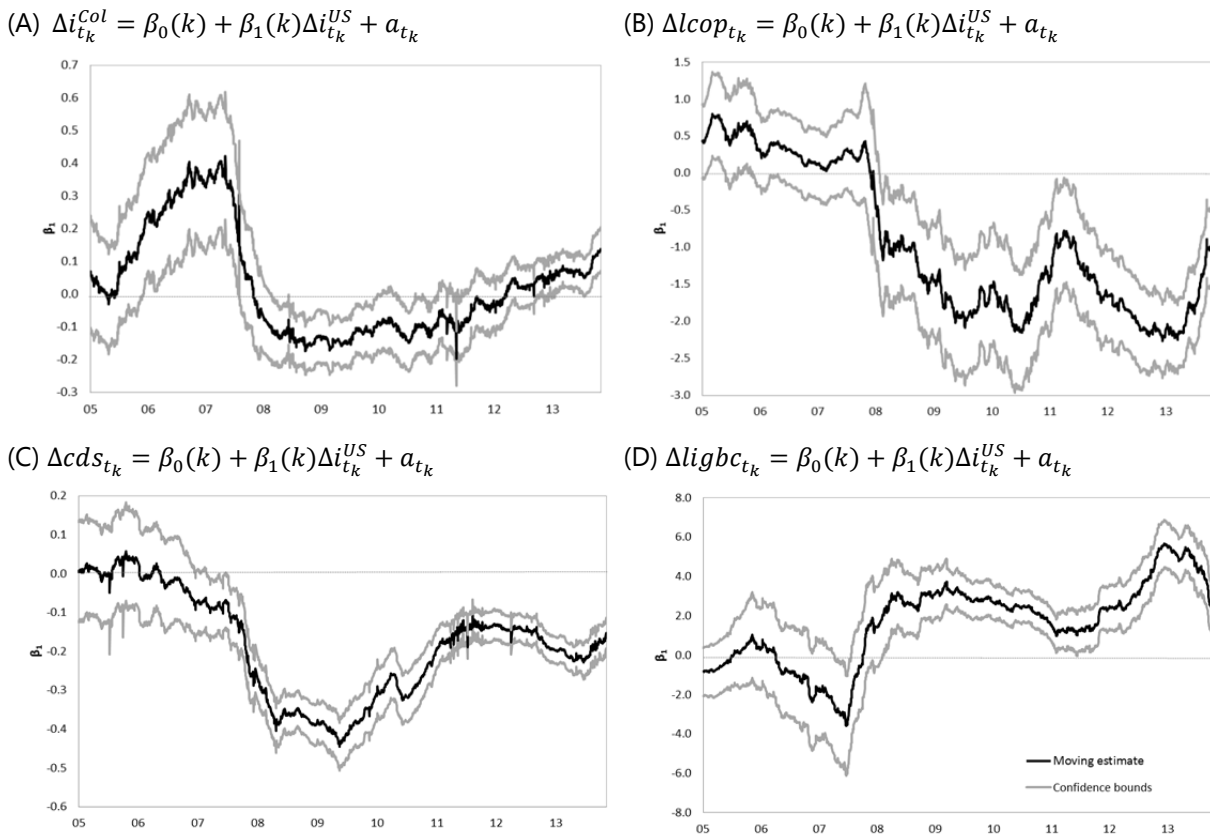
where  $k$  indexes the rolling sample. For each window  $k$ ,  $\beta(\mathbf{k}) = [\beta_0(k), \beta_1(k)]$  is the estimated coefficient vector,  $y_{t_k}$  denotes the dependent variable,  $i_{t_k}^{US}$  is the US long-term Treasury rate, and  $a_{t_k}$  are the errors. The latter are assumed to be heteroscedastic.

We carry out four MWLR calculations of equation (1), and Figure 4 plots the estimated coefficient and its confidence interval for each (the dates on the horizontal axes correspond to the end-day of each rolling regression). In each calculation,  $y_{t_k}$  takes the value of one of the following variables for Colombia: the long-term interest rate  $i_{t_k}^{Col}$  (panel A), the logarithm of the foreign exchange rate *lcope* <sub>$t_k$</sub>  (panel B), the value of CDS spreads *cds* <sub>$t_k$</sub>  (panel C) or the logarithm of the stock value index *ligbc* <sub>$t_k$</sub>  (panel D).<sup>11</sup>

<sup>9</sup> Cronin (2014) highlights that the analysis over time of the relationship between financial variables provides more information than a static assessment with the full sample.

<sup>10</sup> For the MWLR exercises, we have extended our original sample period back to 2003 to have at least two years of observations in the first window of our rolling regression.

<sup>11</sup> This econometric exercise could suffer from omitted-variable bias. However, this MWLR meets the aim of providing evidence on the changing relationship between  $i^{US}$  and the four Colombian financial variables. In addition, our results are robust to including the VIX in the MWLR.



Source: Authors' calculations

The MWLR estimate of  $\beta_1$  for the values of the Colombian long-term interest rate  $i^{Col}$  illustrates a remarkable shift through time in the relationship between the long-term interest rates (Figure 4, panel A). The positive correlation from 2006 to 2007 turns negative between 2008 and 2011 and then turns positive again after 2012.

Regarding the remaining three calculations for the relationship between local asset prices and the US long-term Treasury rate, our estimates show that before the second half of 2007 those relationships were not statistically significant. Nevertheless, during and after the crisis (since the end of 2007) the link between the US Treasury rate and  $l_{cop}$  (negative),  $c_{ds}$  (negative) and  $l_{igbc}$  (positive) became significant (Figure 4, panels B, C and D, respectively). These sharp variations suggest important changes in the nature of shocks hitting these variables throughout the sample period, and particularly, during the crisis.

### Second exercise

In our second MWLR exercise, we again analyse the relationship between changes in the Colombian and US long-term interest rates, controlling for other relevant financial variables. In this case we estimate the equation stated by

$$\Delta i_{t_k}^{Col} = \beta_0(k) + \beta_1(k)\Delta i_{t_k}^{US} + \beta_2(k)\Delta l_{t_{k-1}}^{cop} + \beta_3(k)\Delta c_{t_{k-1}}^{ds} + a_{t_k} \quad (2)$$

where  $k$  indexes the rolling sample. For each window  $k$ ,  $\beta(\mathbf{k}) = [\beta_0(k), \beta_1(k), \beta_2(k), \beta_3(k)]$  is the new estimated coefficient vector, and  $a_{t_k}$

denotes the regression errors. The latter are again assumed to be heteroscedastic. In this new specification we include both  $cds_{t_k-1}$  and  $lcop_{t_k-1}$  as explanatory variables to control for the effects that changes in the sovereign risk premium and the expected depreciation could have on  $i_{t_k}^{col}$ . Note that we consider the lagged values of  $cds_{t_k}$  and  $lcop_{t_k}$  to minimise problems of endogeneity.

The moving window estimate of  $\beta_1$  for the explanatory variable  $i^{US}$ , shows that the link between US and Colombian bond yields is positive before the second half of 2007 and then turns negative up to the end of 2012 (Figure 5, panel A). Moreover, the dynamics of our estimated  $\beta_1$  coefficient are similar to those found in Figure 4, panel A.

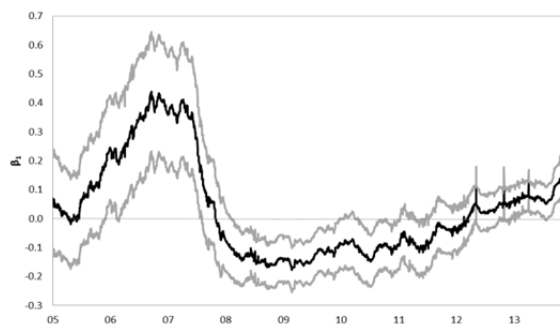
On the other hand, the relationships between  $i^{col}$  and the variables  $lcop$  and  $cds$  have also changed through time. These two links are positive and significant before the second half of 2007 (Figure 5, panels B and C). Even more, in the first part of that year, the impact of changes in CDS spreads on the Colombian sovereign bond yield is stronger than in the previous period. Nevertheless, beginning in 2008, these relationships are not statistically significant.<sup>12</sup>

<sup>12</sup> We also carry out the MWLR of Equation (2) using the difference between  $lcope_t$  and  $lcop_t$  as a proxy for the expected devaluation  $dlcope_t^*$ . Figure 10 in Appendix B presents the moving window estimate of coefficients  $\beta_1$ ,  $\beta_2$  and  $\beta_3$  of this exercise. The dynamics of the estimated coefficients  $\beta_1$  and  $\beta_3$  in Figure 10 are very similar to those exhibited in Figure 5. These findings show that there are no relevant differences in the estimated coefficient  $\beta_1$  when either  $lcop_{t_k-1}$  or  $dlcope_{t_k-1}^*$  is used as a proxy for expected depreciation.

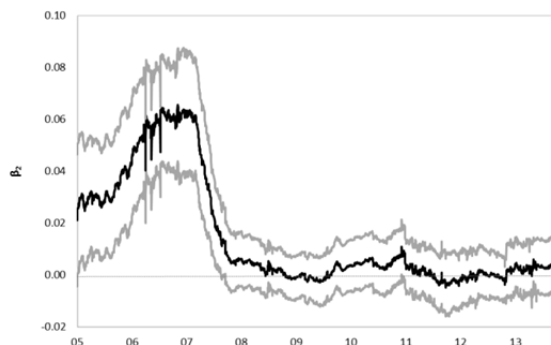
$$\text{MWLR: } \Delta i_{t_k}^{\text{Col}} = \beta_0(k) + \beta_1(k)\Delta i_{t_k}^{\text{US}} + \beta_2(k)\Delta \text{lcop}_{t_k-1} + \beta_3(k)\Delta \text{cds}_{t_k-1} + a_{t_k}$$

Figure 5

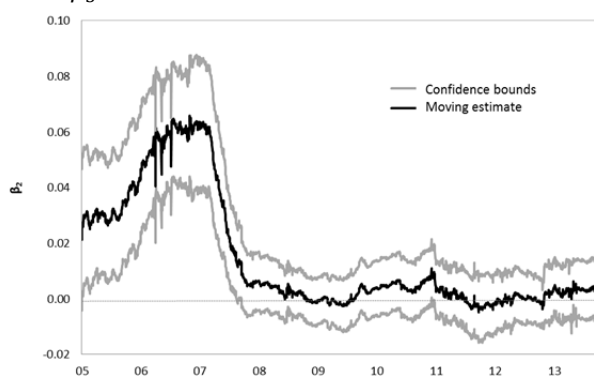
(A) Moving window estimate of  $\beta_1$



(B) Moving window estimate of  $\beta_2$



(C) Moving window estimate of  $\beta_3$



Source: Authors' calculations

### 4.3 The link between US and Colombian long-term interest rates from the perspective of shocks

US and Colombian long-term interest rates bear a positive relationship before the second half of 2007 and after the end of 2012 (Panel A, Figures 4 and 5). This link is negative during the crisis, and it is also negative afterwards, when the Federal Reserve was implementing unconventional stimulus policies.

The relationship between long-term bond yields can be understood from the source of the shock affecting the US Treasury rate. For example, a positive link between both interest rates could be explained by a shock whose origin is the tightening or expected tightening of monetary policy in the US. This shock induces the sale of local bonds, capital outflows from Colombia and, consequently, an increase of long-term interest rates and the depreciation of the local currency.

This response would be greater if the change in the US Treasury rate is perceived as permanent or highly persistent. In this case, domestic factors determining the short-term local interest rate or its expected future path would also be affected by the shock (eg increases in the "natural interest rate" of the small open economy, inflationary pressures derived from the depreciation of the currency, or the reaction of the local central bank to these effects).

Similarly, a shock to the US Treasury rate stemming from a rise in the term premium induces a positive link between US and Colombian long-term bond yields.

This shock reflects an increase in the uncertainty on the future path of the US short-term interest rate. The latter effect could also be associated to increases in the risk and the uncertainty derived from US economic conditions. In this case, the shock would also affect the country risk.

On the other hand, a negative relationship between US and Colombian long-term bond yields could be explained by the role of US Treasuries as a safe haven asset during the global financial crisis. Under this context, shocks buffeting the US long-term interest rate are linked to movements toward or away from safe assets. Therefore, a reduction in the US Treasury rate is associated to a larger appetite for safe assets, capital outflows from emerging market economies and the decline in the prices of local assets, including sovereign bonds.

In this case, shocks to the US long-term Treasury rate would not only include surprises associated with expectations about US monetary policy or the Treasury term premium, but also a safe haven premium during the crisis period. Moreover, shifts in the appetite for safe assets would not only be reflected in the prices of US Treasuries, but also in the price of emerging market economy assets. This hypothesis highlights the usefulness of including a measure of global risk and economic uncertainty in our analysis. The VIX index is the natural candidate for this purpose.

In the next section we explore the short-term responses of some Colombian asset prices to external financial shocks.

## 5. The short-term responses of Colombian asset prices to external financial shocks

US Treasury interest rates are endogenous variables subject to various shocks that may likewise affect emerging market economy asset prices, including long-term interest rates, the foreign exchange rate, CDS spreads and the stock market index. Hence, the “transmission” of changes in US long-term bond yields to local asset prices implies the response of all these variables to shocks from different sources. Moreover, the frequency and predominance of these shocks change over time.

To capture this transmission idea, we estimate the response of the Colombian long-term interest rate and other asset prices to three shocks, namely (i) global volatility and economic uncertainty, (ii) the term premium and (iii) the stance of monetary policy. These shocks can impact asset prices directly and through the US Treasury rate channel.

Consider the following VARX( $p, q$ ) model:

$$\Delta \mathbf{Y}_t = \boldsymbol{\mu} + \sum_{i=1}^p A_i \Delta \mathbf{Y}_{t-i} + \sum_{i=0}^q B_i \Delta \mathbf{X}_{t-i} + \boldsymbol{\epsilon}_t \quad (3)$$

where  $\boldsymbol{\mu}$  is a vector of means,  $A_i$  and  $B_i$  stand for the coefficient matrices associated with the endogenous and exogenous variables, respectively, and  $\boldsymbol{\epsilon}_t \sim WN(0, \boldsymbol{\Sigma}_t)$  is a vector of errors. We assume a Baba-Engle-Kraft-Kroner (BEKK) multivariate GARCH model as defined in Engle and Kroner (1995). The latter is used to model the high volatility of financial time series with daily frequency in the sample.

Vectors  $\mathbf{Y}_t = (i_t^{col}, ligbc_t, lcop_t, lcope_t, cds_t, i_{MP,t}^{col})$  and  $\mathbf{X}_t = (VIX_t, i_t^{US}, MOVE_t)$  stand for the sets of endogenous and exogenous variables, respectively.<sup>13</sup> These vectors are included in first differences in the estimation.<sup>14</sup>

Variables  $i_t^{col}, ligbc_t, lcop_t, lcope_t, cds_t$  and  $i_t^{US}$  were already defined in Section 4.1. The variable  $i_{MP,t}^{col}$  denotes the monetary policy rate for Colombia. We use the Colombian interbank rate as a proxy for  $i_{MP,t}^{col}$ . The VIX<sup>15</sup> and MOVE<sup>16</sup> indexes are proxies for US market volatility and the US Treasury term premium, respectively. The VIX picks the effects of global uncertainty shocks, while the MOVE captures the uncertainty on the future path of short-term interest rates in the US market. Tobias et al (2013) and Cieslak and Povala (2013) point out that the *MOVE* is highly correlated with the 10-year US Treasury term premium.

*VARX* equations consider contemporaneous and lagged values of our exogenous variables ( $i_t^{US}, VIX_t, MOVE_t$ ). Hence, the responses of local asset prices to an  $i_t^{US}$  shock capture the impact of changes in the US long-term Treasury rate that are not explained by movements of the VIX or MOVE. Therefore, the responses to  $i_t^{US}$  shocks must reflect changes in the stance of monetary policy in the US and "other effects".

The VARX-MGARCH model is estimated for our three sample periods (before, during and after the financial crisis), as defined in Section 3.2. The estimates of the VARX-MGARCH for each sample period are used to perform the impulse-response analysis to shocks to exogenous variables (multiplier analysis). In particular, we study the responses of Colombian asset prices<sup>17</sup> to shocks to *VIX*, *MOVE* and  $i_t^{US}$ . Appendix C presents the technical details of the estimation method, compiles the main results and briefly summarises the specification test.

## 5.1 Pre-crisis period

Figure 6 shows the multiplier analysis for the pre-crisis period (June 2004 to February 2007). For this sample, shocks to *VIX* lead to positive responses in *cds*, *lcop*, *lcope*,  $i^{col}$  and a negative reaction in *ligbc*. Hence, an increase in US market volatility induces a rise in Colombian long-term interest rates (ie a fall in the value of the long-term bond portfolio) and a decline in stock prices. Investors reallocate their

<sup>13</sup> To control for specific events like FOMC meetings and the publication of its minutes (as suggested by Wright (2012) and Londoño and Sapriza (2014)), we included a dummy variable that covers the dates of those events.

<sup>14</sup> As a robustness check, we estimated the VARX-MGARCH model without the variable *lcope*. The latter is not observable, and hence it may introduce noise in the estimation. The results do not alter the findings reported in this section.

<sup>15</sup> The VIX is the Chicago Board Options Exchange Market Volatility Index. It is a measure of the implied volatility of S&P 500 Index options over the next 30 days.

<sup>16</sup> The MOVE is the Merrill Lynch Option Volatility Estimate Index. It is a weighted average of the normalised implied yield volatility for 1-month Treasury options on the two-year (20%), five-year (20%), 10-year (40%) and 30-year (20%) maturities. The weights are based on option trading volumes in each maturity.

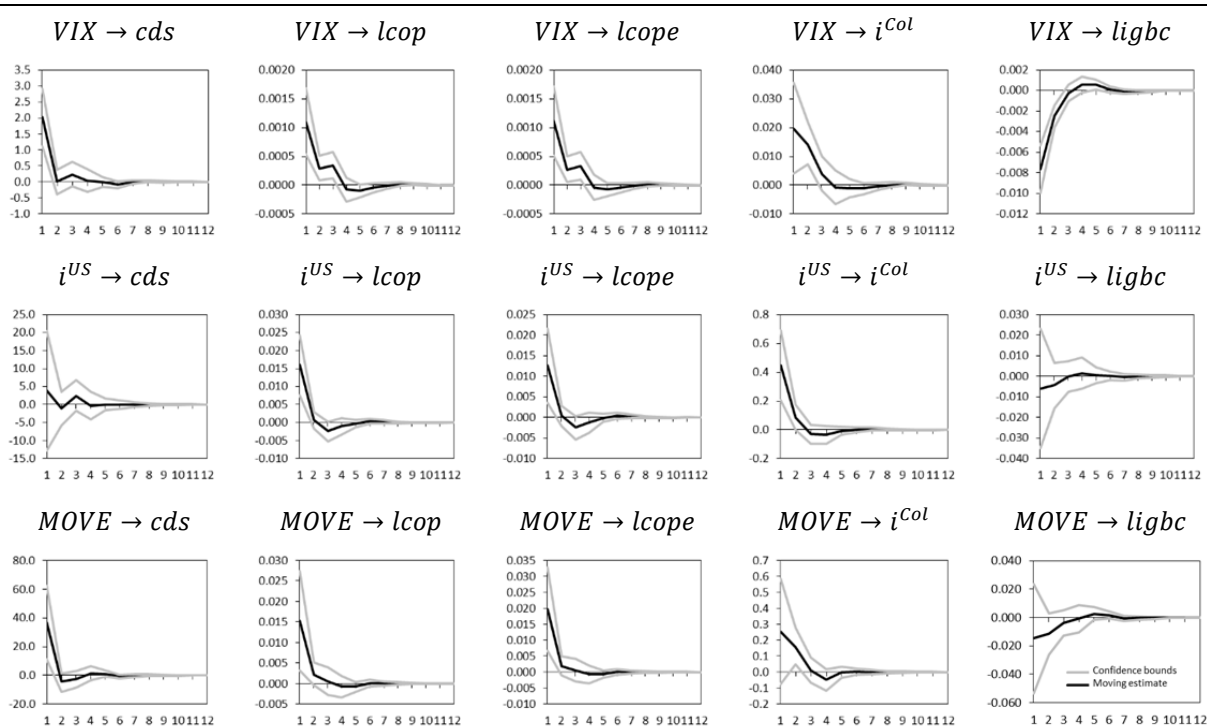
<sup>17</sup> For all sample periods, the responses of  $i_{MP}^{col}$  to external financial shocks in this analysis are not statistically significant.

resources away from local markets, which causes a depreciation of the currency and an increase in the perception of country risk.

For the same period, a positive shock to either the stance of US monetary policy ( $i_t^{US}$ ) or the term premium ( $MOVE$ ) produces a depreciation of the currency ( $lcop$ ) and an increase in the Colombian long-term bond yield ( $i^{Col}$ ). Moreover, the shock to  $MOVE$  also leads to a positive response in the country risk perception ( $cds$ ) and higher expectations of future devaluation. On the other hand, none of these two shocks has a significant effect on the stock market index ( $ligbc$ ).

Pre-crisis period: multiplier analysis (*Impulse* → *Response*)

Figure 6



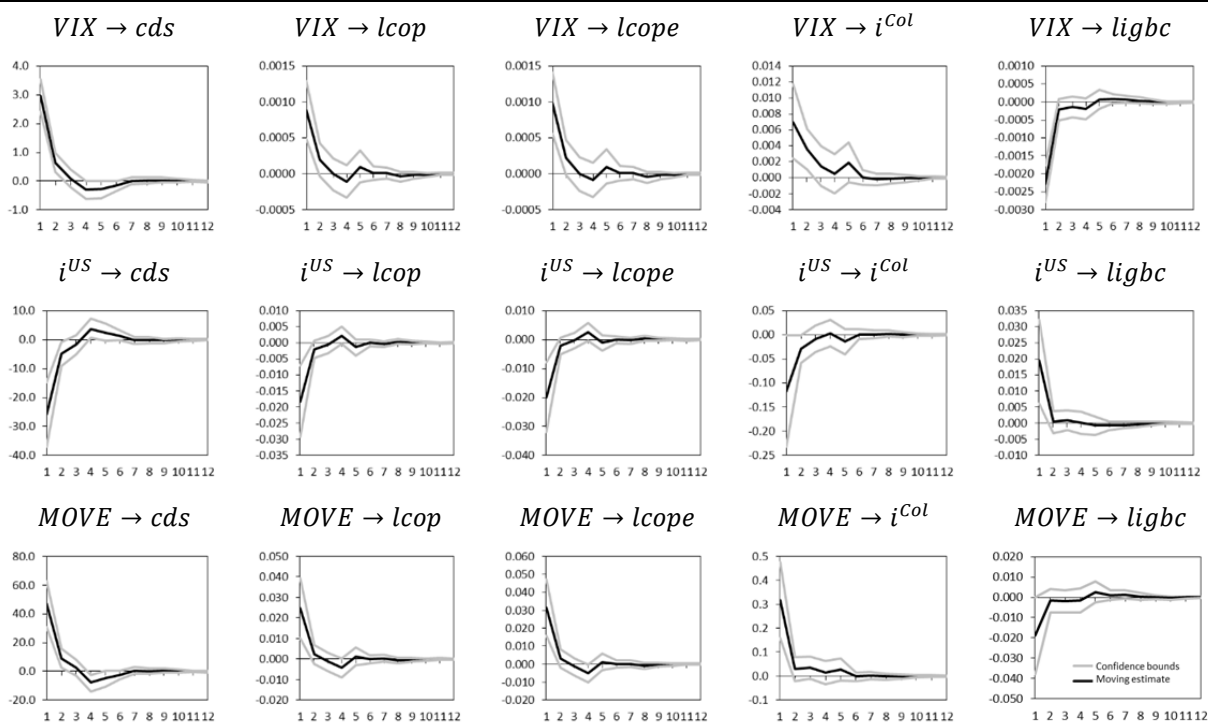
Source: Authors' calculations.

These results suggest that in this period of relative stability in the market, bond investment decisions are characterised mainly by the search for high returns. Positive shocks to global volatility, the term premium and the stance of US monetary policy increase the Treasury yield and lead to sales of sovereign long-term bonds and local currency. Further, in this period only risk shocks produce significant shifts in stock prices.

## 5.2 The crisis period

Figure 7 illustrates the multiplier analysis for the crisis period (February 2007 to October 2009). As in the results for the pre-crisis period, positive shocks to either US risk ( $VIX$ ) or the Treasury term premium ( $MOVE$ ) lead to a decline in the prices of sovereign bonds (ie a rise in the long-term bond yield ( $i_t^{Col}$ )), a depreciation of the local currency ( $lcop_t$ ), a decline in stock prices ( $ligbc_t$ ) and a rise in the perception of sovereign risk ( $cds_t$ ).





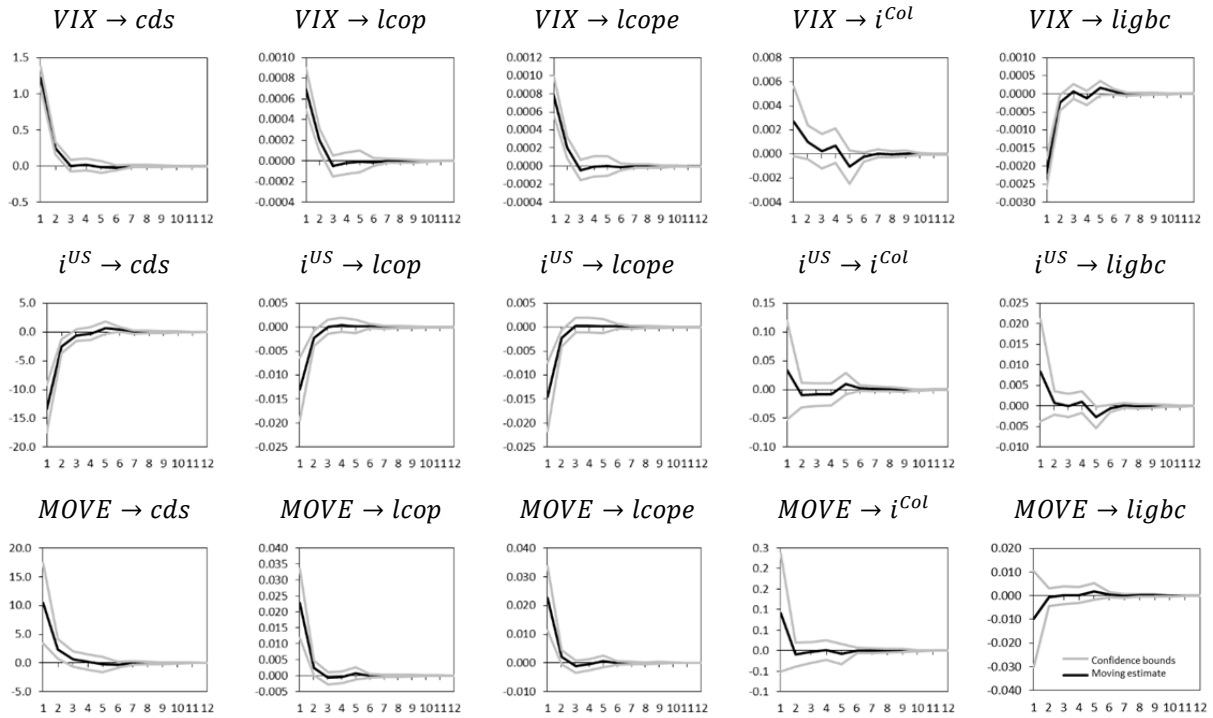
Source: Authors' calculations.

However, a positive shock to  $i_t^{US}$  provides a qualitatively different story. The response to this shock is opposite to that observed in the pre-crisis period: a reduction in local long-term bond yields, an appreciation of local currency, an increase in stock prices, and a decline in country risk perception. Moreover, the impacts on *cds* and *ligbc* become statistically significant.

In contrast, if the  $i_t^{US}$  shock is negative (a reduction in the US long-term Treasury rate and hence a higher market value of these securities), the response is an increase in the Colombian long-term bond yield (a decline in the value of the local bond portfolio), a devaluation of the currency and a decline in stock prices.

In the crisis period, the response of local asset prices to a shock to  $i_t^{US}$  suggests that US Treasuries became a safe haven in the midst of economic uncertainty and high levels of risk. Under these circumstances, Colombia and other emerging market economies are observed as a potential source of losses in an episode of crisis. In this scenario, a negative shock to  $i_t^{US}$  reduces the US long-term Treasury rate and leads to capital outflows from the Colombian financial market and into the US Treasury market.

As already mentioned, the shock to  $i_t^{US}$  encompasses surprises in the stance of monetary policy and "other effects". We suggest that in the climate of high uncertainty, the "other effects" component captures the desire of investors to hedge exposures to emerging market economies by using safe haven assets. These results also suggest that the *VIX* does not completely capture changes in global risk aversion or in the fear of a generalised economic collapse. In the crisis period, shocks to  $i_t^{US}$  captures mostly movements toward or away from safe assets.



Source: Authors' calculations.

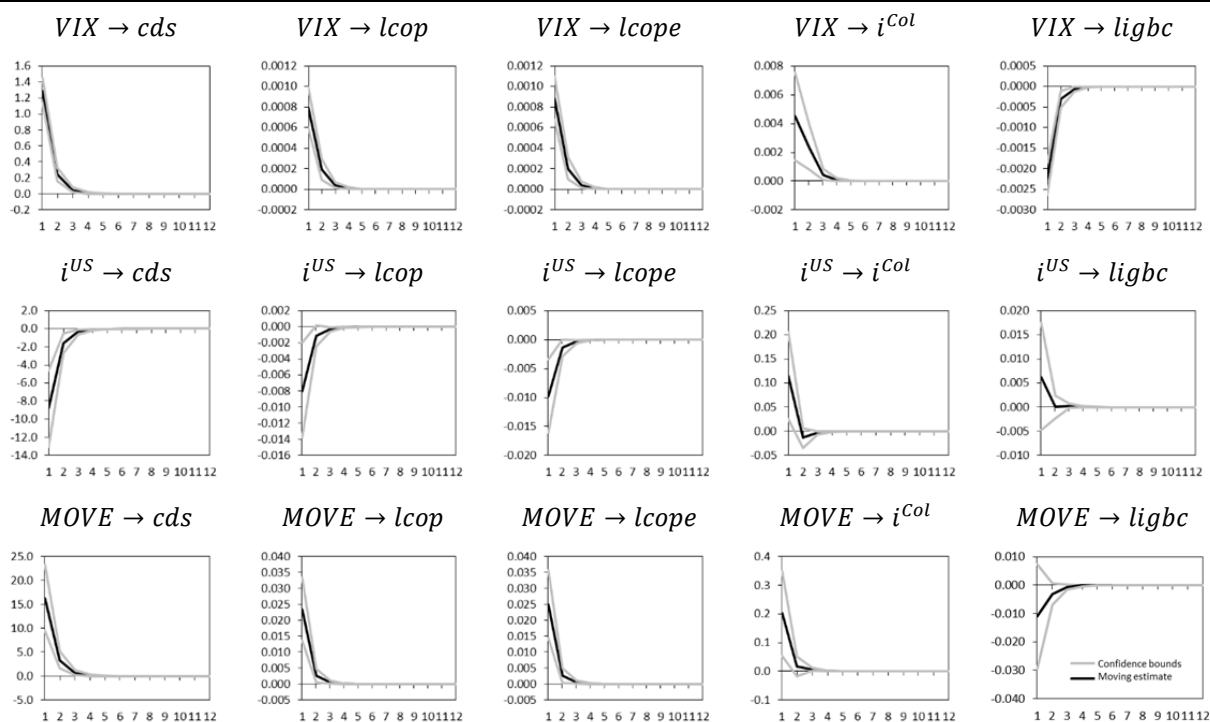
### 5.3 Post-crisis period

For the post-crisis period we examine a subsample as well as the overall period. The subsample covers data between November 2009 and April 2013 (ie before the tapering-related turmoil that began in May); the overall interval runs beyond, to November 2013.

#### Post-crisis period before tapering-related statements

In the November 2009–April 2013 subsample, the qualitative responses of the country risk spread ( $cds$ ), the foreign exchange rate ( $lcop$ ) and the expectations of future depreciation ( $lcope$ ) to  $VIX$ ,  $i_t^{US}$  and  $MOVE$  shocks do not change with respect to the results presented for the crisis period (Figure 8).

These results suggest that US Treasuries partially kept their safe-haven status. Possible explanations for this condition are the uncertainty associated with the slow recovery of the US economy, with the Greek debt crisis and with the unconventional policies adopted by advanced economies.



Source: Authors' calculations.

However, for this period, not all local assets behave in the same fashion. In particular, unlike in the crisis period,  $MOVE$  and  $i_t^{US}$  shocks do not produce statistically significant responses of the long-term interest rates. This result provides evidence of market differentiation that distinguishes long-term sovereign bonds from other Colombian assets. Accordingly, the sensitivity of local long-term bonds to external financial shocks may have been reduced.

### Overall post-crisis period

Figure 9 exhibits the multiplier analysis for the overall post-crisis period (November 2009 to November 2013). The qualitative responses of local asset prices other than the local long-term interest rate are similar to those observed in the post-crisis subsample, which precedes the tapering-related statements in May.

The responses of the long-term bond yield to shocks to  $VIX$ ,  $i_t^{US}$  and  $MOVE$  are positive and statistically significant. These results are consistent with the gradual retrenchment of the unconventional monetary policy adopted by the Federal Reserve through its QE3 program. These findings suggest that the local long-term interest rate was more sensitive to the tapering-related turmoil than other local asset prices.

## 6. Conclusions

Understanding the relationship between the long-term interest rates of advanced and emerging market economies requires the identification of specific shocks that

affect their dynamics. Our findings suggest that changes in the nature and importance of these shocks are behind the time-varying link between the US Treasury rate and Colombian asset prices, including local long-term bond yields.

In particular, our results show that the short-run response of the local long-term interest rate, CDS spreads, the foreign exchange rate and the stock market index to shocks to the US Treasury rate have been qualitatively different depending on both the sample period (ie before, during and after the global financial crisis) and the source of the shock.

Our findings suggest that in the pre-crisis period, investment decisions were characterised mainly by the search for high returns. Positive shocks to global volatility, the term premium and the stance of US monetary policy increase the Treasury yield and lead to a rise in local long-term interest rates, a decline in stock prices, a depreciation of the currency and a perception of higher country risk.

During the crisis, shocks to the US Treasury rate caused by changes in global volatility or the Treasury term premium show the same qualitative responses observed in the pre-crisis period. However, the responses to an  $i^{US}$  shock provide a different story, leading to a reduction in local long-term bond yields, the appreciation of the local currency, an increase in stock prices, and a fall in the country risk perception.

We suggest that in the atmosphere of economic uncertainty and high levels of risk that characterised the crisis period, a shock to  $i_t^{US}$  captures the desire of investors to hedge against such stress by using safe haven assets. The latter effect was dominant during the crisis. These results also suggest that the *VIX* does not completely capture changes in global risk aversion or in the fear of a generalised economic collapse.

In the post-crisis period, the responses of the Colombian long-term bond yield and other asset prices are similar to those observed during the crisis. Our findings indicate that this period is also characterised by a safe haven role for US Treasuries. Nevertheless, there are signals of a possible differentiation between local asset types.

## Appendix

### A Moving window linear regression (MWLR) with GJR-GARCH variance

We discuss here the main details of the econometric strategy used to estimate the rolling coefficients of the regression exercises presented in Section 4.

To provide an estimation of time-varying model parameters, we carry out an analysis based on a MWLR (see Zivot and Wang (2006) and Stock and Watson (2011)). Moreover, we capture the changing volatility of financial time series used in these exercises by assuming that the volatility follows a conditional heteroscedastic model.<sup>18</sup>

In particular, we consider a MWLR model with fixed windows of length  $n$ . The model is defined as

$$Y_{t_k} = \beta_0(k) + \beta_1(k)X_{t_k} + a_{t_k}, \quad \text{for } k = 1, \dots, T - n + 1, \quad \text{and } t = k, \dots, n + k - 1, \quad (4)$$

where  $k$  indexes the rolling window,  $t$  indexes the time in the regression and  $T$  is the total number of observations.

For each window  $k$ ,  $Y_{t_k}$  denotes an  $(n \times 1)$  vector of observations on the dependent variable;  $X_{t_k}$  is an  $(n \times 1)$  vector of values on the explanatory variable;  $[\beta_0(k), \beta_1(k)]$  are scalars that stand for the intercept and slope of the regression, respectively; and  $a_{t_k}$  is an  $(n \times 1)$  vector of error terms. For the window  $k$ , the  $n$  observations in  $Y_{t_k}$  and  $X_{t_k}$  correspond to the  $n$  most recent values of the sample for time  $t = k: n + k - 1$  (see also Zivot and Wang (2006)).

We also consider that heteroscedastic errors  $a_{t_k}$  are given by

$$a_{t_k} = \sigma_{t_k} \epsilon_{t_k} \quad (5)$$

and that the conditional variance  $\sigma_{t_k}^2$  evolves over time following a GJR-GARCH(1,1) process

$$\sigma_{t_k}^2 = \alpha_0(k) + (\alpha_1(k) + \gamma_1(k)N_{t_{k-1}})a_{t_{k-1}}^2 + \delta_1(k)\sigma_{t_{k-1}}^2 \quad (6)$$

where  $N_{t_{k-1}}$  is an indicator for negative values of  $a_{t_{k-1}}$ , that is,

$$N_{t_{k-1}} = \begin{cases} 1 & \text{if } a_{t_{k-1}} < 0, \\ 0 & \text{if } a_{t_{k-1}} \geq 0, \end{cases}$$

with parameters  $\alpha_0(k) > 0$ ,  $\alpha_1(k) \geq 0$ ,  $\delta_1(k) \geq 0$ ,  $\gamma_1(k) \geq 0$  and  $\alpha_1(k) + 0.5\gamma_1(k) + \delta_1(k) < 1$  (for more details see Tsay (2010)). The  $\alpha_1(k)$ ,  $\delta_1(k)$  and  $\gamma_1(k)$  are referred as the *ARCH*, *GARCH* and *Leverage* parameters, respectively. The GJR-GARCH is commonly used to model asymmetry in the ARCH process. We also assume that  $\epsilon_{t_k}$  is a sequence of Student's t errors.

<sup>18</sup> The volatility modelling can improve the efficiency in parameter estimation and the accuracy in confidence intervals (Tsay (2010)).

For each window  $k$ , the estimation is performed by maximum likelihood. All regression exercises are performed using the Matlab econometric toolbox. Each figure in Section 4 shows the moving window estimate of the coefficient  $\hat{\beta}_{(c)}$  and its 95% confidence interval.

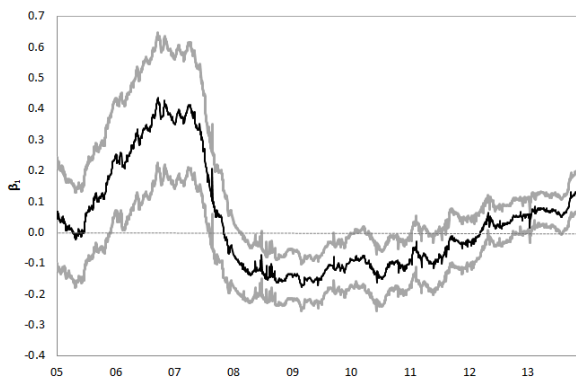
## B MWLR: multivariate model

Figure 10 shows the evolution of the moving window estimate of  $\beta_1$ ,  $\beta_2$  and  $\beta_3$  for the explanatory variables  $i^{US}$ ,  $dlcpe$  and  $cds$  (panels A, B and C, respectively). This exercise also provides evidence on the changing relationship between long-term sovereign interest rates. The relationships between  $i^{Col}$  and variables  $dlcpe$  and  $cds$  have also changed through time.

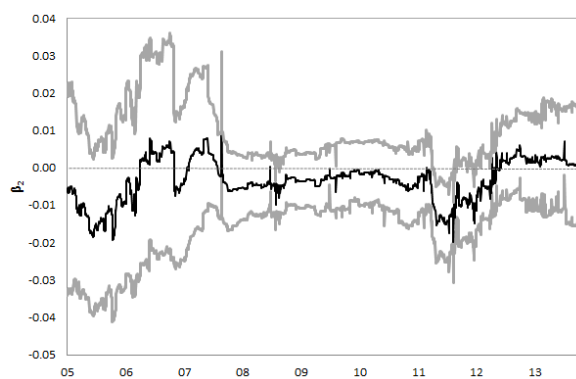
$$\text{MWLR: } \Delta i_{t_k}^{Col} = \beta_0(k) + \beta_1(k)\Delta i_{t_k}^{US} + \beta_2(k)\Delta dlcpe_{t_k-1} + \beta_3(k)\Delta cds_{t_k-1} + a_{t_k}$$

Figure 10

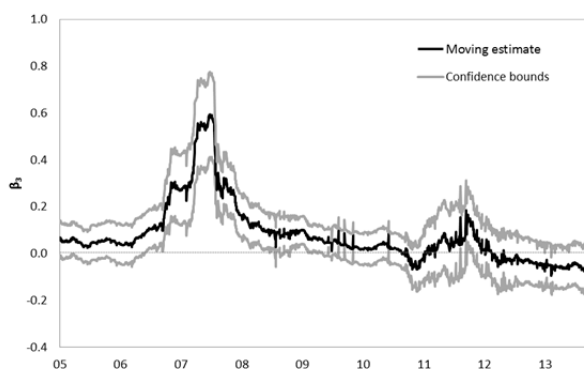
(A) Moving window estimate of  $\beta_1$



(B) Moving window estimate of  $\beta_2$



(C) Moving window estimate of  $\beta_3$



Source: Authors' calculations.

## C VARX-MGARCH model

We discuss here the main details of the econometric methodology used to estimate the VARX( $p,q$ )-MGARCH( $l,m$ ) model and the impulse-response to shocks to exogenous variables (ie multiplier analysis) considered in Section 5.

We consider the VARX( $p,q$ )-MGARCH( $l,m$ ) model,

$$\Delta \mathbf{Y}_t = \mu + \sum_{i=1}^p A_i \Delta \mathbf{Y}_{t-i} + \sum_{i=0}^q B_i \Delta \mathbf{X}_{t-i} + \boldsymbol{\varepsilon}_t \quad (7)$$

$$\boldsymbol{\Sigma}_t = C_0' C_0 + \sum_{i=1}^l F_i' \boldsymbol{\varepsilon}_{t-i} \boldsymbol{\varepsilon}_{t-i}' F_i + \sum_{j=1}^m G_j' \boldsymbol{\Sigma}_{t-j} G_j \quad (8)$$

where  $\mu$  is a vector of means;  $A_i$  and  $B_i$  stand for the coefficient matrix associated with the endogenous and exogenous variables, respectively; and  $\boldsymbol{\varepsilon}_t \sim WN(0, \boldsymbol{\Sigma}_t)$  is a vector of errors. We assume a Baba-Engle-Kraft-Kroner (BEKK) multivariate GARCH model as defined in Engle and Kroner (1995). The vectors

$$\mathbf{Y}_t = (i_t^{col}, ligbc_t, lcop_t, lcope_t, cds_t, i_{MP,t}^{col})$$

and

$$\mathbf{X}_t = (VIX_t, i_t^{US}, MOVE_t)$$

denote the sets of endogenous and exogenous variables. These vectors are order-one integrated  $I(1)$ .

The estimation of the model is carried out in two steps.<sup>19</sup> First, we estimate the VARX model defined in equation (7). Second, we use residuals obtained from the previous step to estimate the MGARCH model stated by equation (8). We then perform the multiplier analysis.<sup>20</sup>

We carry out the estimation of the VARX – MGARCH model in Section 5 for four specific periods. The first one corresponds to dates from June 2004 to February 2007 (ie pre-crisis period). The second time span goes from February 2007 to October 2009 (ie crisis period). The third period includes dates from November 2009 to April 2013 (ie post-crisis period before the “Tapering” announcement). Our fourth period considers the post-crisis period until November 2013.

Table 1 reports the unit-root tests performed in our analysis. The order of integration is determined using the augmented *Dickey-Fuller*, *Phillips & Perron (PP)*, *Elliott, Rothenberg & Stock (ERS)* and *KPSS* tests. These results indicate that variables are order-one integrated.<sup>21</sup> We assume that variables are not cointegrated.

Tables 2, 3 and Figures 11, 12, 13 and 14 show the specification test for each model. These tests were carried out on MGARCH standardised residuals. There is no evidence of misspecification.

<sup>19</sup> The lags of the endogenous and exogenous variables  $p$  and  $q$  in equation 7 are determined using standard information criteria. The lags in equation 8 are determined from the specification tests of MGARCH models.

<sup>20</sup> Two points are clarified. First, problems on simultaneity and identification are precluded because the shock occurs on an exogenous variable. Second, as endogenous and exogenous variables are assumed  $I(1)$ , the resulting multipliers do not need to be integrated to obtain the responses of endogenous variables in levels.

<sup>21</sup> Unit-root tests were also carried out on the first difference of the variables to confirm the order of integration.

The multiplier analysis presented in Section 5 shows the response of the level of endogenous variables to a one-unit shock on the level of exogenous variables *VIX*, *i<sup>US</sup>* and *MOVE*. Confidence bounds for our multiplier analysis are estimated by bootstrapping techniques after controlling for *GARCH* effects. Our results are based on 5,000 replications.

Unit-root test					Table 1				
Test	Variable	Stat	Critical value	Evidence	Test	Variable	Stat	Critical value	Evidence
ADF trend	<i>cds</i>	-3.93	-3.96	Unit root	ERS	<i>cds</i>	-0.94	-3.48	Unit root
	<i>i<sup>col</sup></i>	-2.72	-3.96	Unit root		<i>i<sup>col</sup></i>	-2.39	-3.48	Unit root
	<i>lcop</i>	-2.88	-3.96	Unit root		<i>lcop</i>	-2.19	-3.48	Unit root
	<i>lcope</i>	-2.75	-3.96	Unit root		<i>lcope</i>	-2.01	-3.48	Unit root
	<i>ligbc</i>	-2.21	-3.96	Unit root		<i>ligbc</i>	-0.58	-3.48	Unit root
	<i>VIX</i>	-3.34	-3.96	Unit root		<i>VIX</i>	-2.38	-3.48	Unit root
	<i>i<sup>US</sup></i>	-2.99	-3.96	Unit root		<i>i<sup>US</sup></i>	-2.29	-3.48	Unit root
	<i>MOVE</i>	-3.03	-3.96	Unit root		<i>MOVE</i>	-2.96	-3.48	Unit root
PP	<i>cds</i>	-3.67	-3.97	Unit root	KPSS	<i>cds</i>	3.04	0.22	Non-stationary
	<i>i<sup>col</sup></i>	-2.33	-3.97	Unit root		<i>i<sup>col</sup></i>	1.85	0.22	Non-stationary
	<i>lcop</i>	-2.65	-3.97	Unit root		<i>lcop</i>	1.93	0.22	Non-stationary
	<i>lcope</i>	-2.65	-3.97	Unit root		<i>lcope</i>	2.06	0.22	Non-stationary
	<i>ligbc</i>	-1.88	-3.97	Unit root		<i>ligbc</i>	4.50	0.22	Non-stationary
	<i>VIX</i>	-4.24	-3.97	No unit root		<i>VIX</i>	2.01	0.22	Non-stationary
	<i>i<sup>US</sup></i>	-3.04	-3.97	Unit root		<i>i<sup>US</sup></i>	2.72	0.22	Non-stationary
	<i>MOVE</i>	-3.34	-3.97	Unit root		<i>MOVE</i>	2.25	0.22	Non-stationary

Q-test		Table 2			
Sample period	Standardised residuals		Standardised square residuals		
	Statistic	P-value	Statistic	P-value	
1 June 2004–26 February 2007	3617.43	0.4157	2775.87	0.9162	
26 February 2007–5 November 2009	3729.26	0.0651	2941.91	0.2064	
5 November 2009–30 April 2013	3626.32	0.3755	2930.61	0.2509	
5 November 2009–7 November 2013	3502.18	0.8761	2835.85	0.7177	



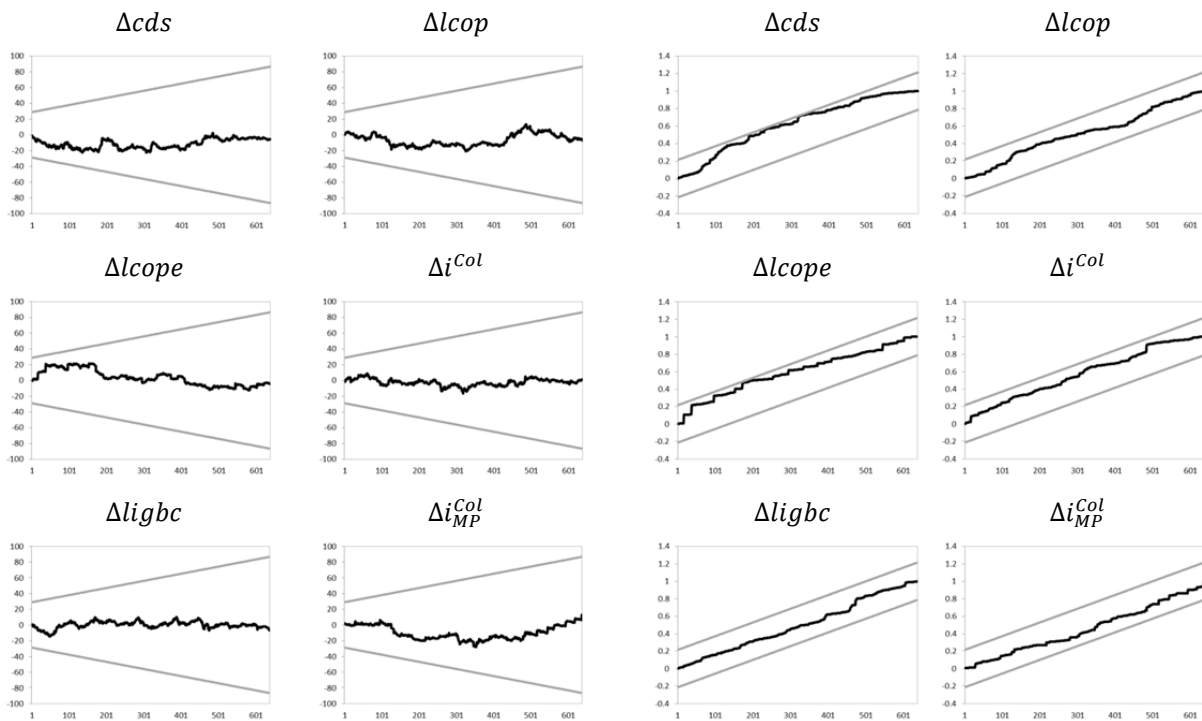
Maximum eigenvalue

Table 3

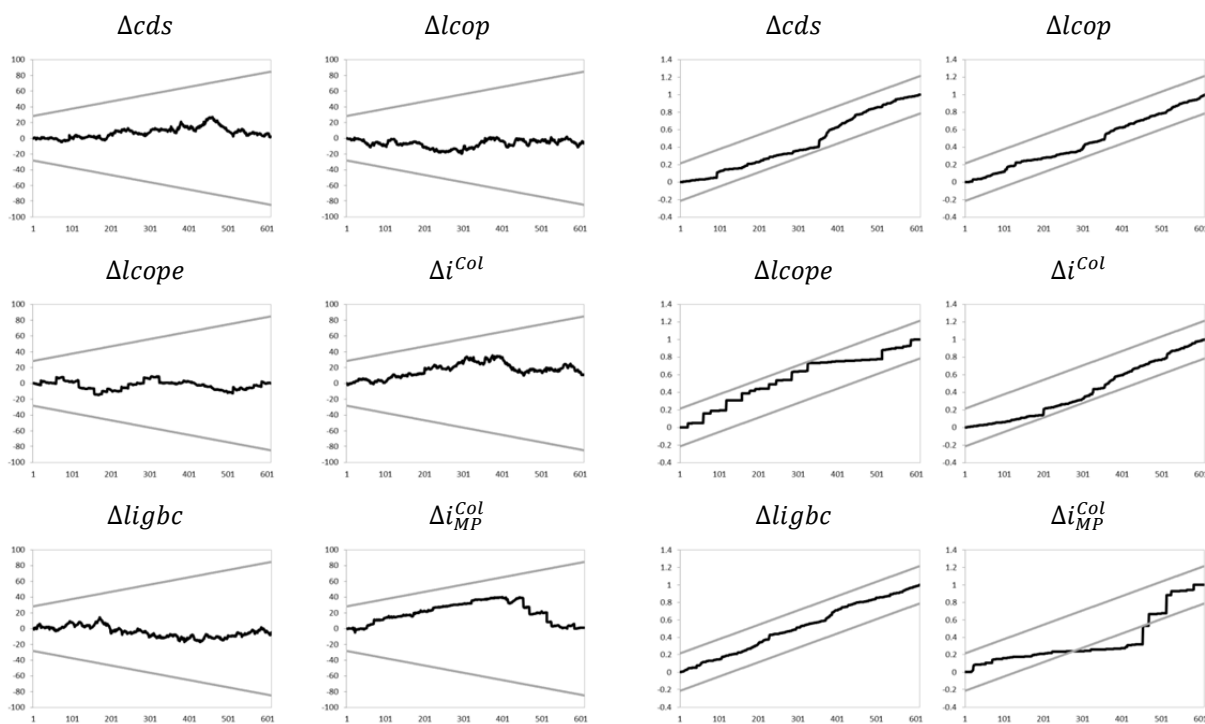
Sample period	Maximum eigenvalue	
	<i>VAR</i>	<i>MGARCH</i>
1 June 2004–26 February 2007	0.5203	0.9433
26 February 2007–5 November 2009	0.6473	0.9262
5 November 2009–30 April 2013	0.2006	0.9378
5 November 2009–7 November 2013	0.5549	0.9467

CUSUM and CUSUM-squared tests: pre-crisis period

Figure 11



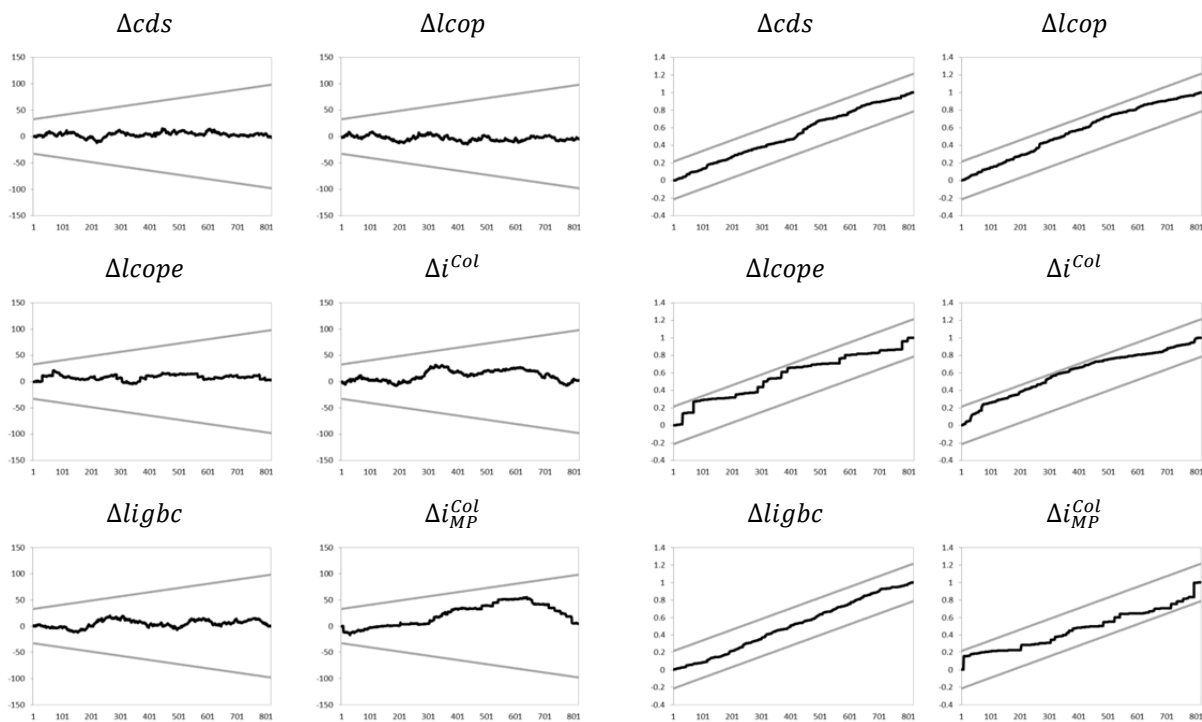
Source: Authors' calculations.



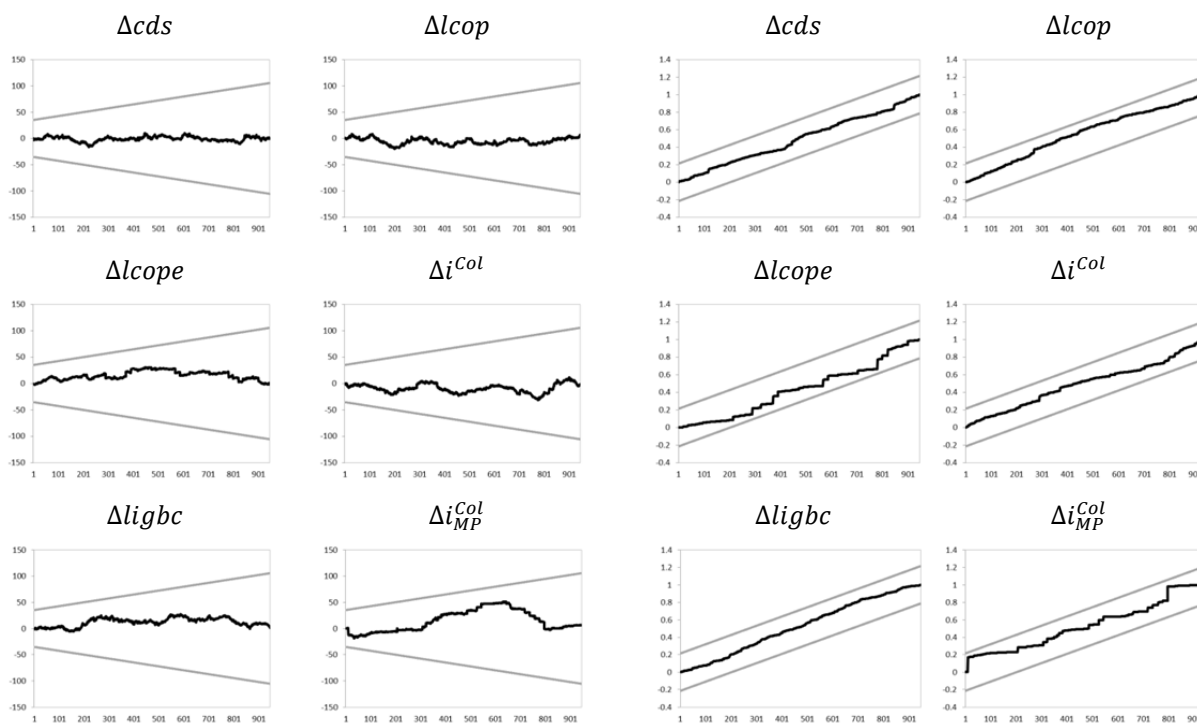
Source: Authors' calculations.

CUSUM and CUSUM-squared tests: post-crisis period before the "Tapering" announcement

Figure 13



Source: Authors' calculations.



Source: Authors' calculations.

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