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# The effects of global excess liquidity on emerging stock market returns: Evidence from a panel threshold model☆

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ABSTRACT

The expansion of global liquidity, exacerbated by the unconventional monetary policies implemented by the major central banks over the past several years, has contributed to the debate on the cross-border impact of those measures. This paper examines the impact of global excess liquidity on asset prices for a set of seventeen emerging market countries taking into account nonlinearity by using a panel threshold model. We find that in a period of global investors' high risk appetites, global excess liquidity is a positive determinant of asset prices in emerging market countries. However, the link between the two variables changes when global risk aversion strengthens.



## 1. Introduction

The 2007–2009 global financial crisis sparked a renewed interest in the topic of global liquidity by those involved in the policy debate. In the context of severe financial market disruptions and impaired financial intermediation following the Lehman Brothers collapse, central banks aggressively lowered their rates to near zero and ultimately used unconventional measures to address issues related to shortages in many financial market segments. These non-standard monetary policy measures, which ranged from forward guidance to credit and quantitative easing, have contributed to the boom in global liquidity. In particular, quantitative easing measures have consisted of a massive expansion of central bank balance sheets via several asset purchase programs.

Monetary base growth has indeed exploded in most advanced economies. The monetary base created by the Bank of England tripled between mid-2007 and mid-2010 to reach U.K. £336 billion by the end of 2012 (from £64 billion in mid-2007), [1] whereas in the Euro zone, the monetary base surged by almost €2900 billion between mid-2007 and the end of 2012. [2] During the same period, the monetary base of the U.S. Federal Reserve Bank more than tripled; [3] in Japan, it increased by two-thirds. [4] However, not only advanced economies but also emerging markets participated in the global monetary expansion. Central banks in many emerging countries intervened in foreign exchange markets to prevent their currencies from excessive appreciation and to prevent a deterioration in competitiveness resulting from strong capital inflows from international investors searching for higher yields (Brana et al., 2012). Foreign exchange reserves have thus risen strongly, particularly in Asian countries, oil exporting countries and Brazil. The partial sterilization of reserve inflows contributed to the expansion of domestic monetary bases and ultimately, to the global monetary expansion (Filardo and Yetman, 2012). According to Borio (2013), the buildup of foreign currency reserves is not precautionary but instead is a by-product of zero lower bound and unconventional monetary policies in advanced economies.

Global excess liquidity provides international investors with relatively cheap liquidity, inducing them to increase their portfolio returns by investing in assets that earn a higher rate. The major channel for global spillovers in emerging countries is capital flows, along with the

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[1] Source: Bank of England.
[2] Source: European Central Bank via Eurostat.
[3] Source: Federal Reserve Bank.
[4] Source: Bank of Japan.

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impact on exchange rates and other asset prices (Chinn, 2013). To what extent is global liquidity responsible for upward pressure on asset prices, especially in emerging countries? Few studies have tackled this question.

Due to substantial increases in international capital flows, the concept of "global liquidity" and the analysis of spillover effects attracted growing attention at the beginning of the 2000s. Most studies consider the impact of global monetary growth on interest rates, GDP and inflation rates only in industrialized countries (Rüffer and Stracca, 2006; Sousa and Zaghini, 2008). They identify significant consumer price reactions to global liquidity shocks, but the link to asset prices is mixed. Baks and Kramer (1999) find that for the G7 countries, global monetary growth has a positive impact on equity prices; however, Belke et al. (2010), who study the interaction between global liquidity and prices levels for goods and assets in 11 OECD countries, find that equity prices do not react to liquidity shocks. Their results are consistent with those of Giese and Tuxen (2007) and Darius and Radde (2010), who show that global liquidity has an impact on housing prices but not on stock prices. All of these authors use VAR models and impulse functions. Other authors find significant impacts of global liquidity on commodity prices (Beckmann et al., 2014; Ratti and Vespignani, 2013).

Studies of the impact of global liquidity on emerging countries are scarcer. Chudik and Fratzscher (2011) compare the role of the tightening of monetary conditions (estimated by the change in the 3-month money market interest rate) and the collapse in risk appetite (evidenced by a shock on the VIX index or the TED spread) in the global transmission of financial crises, as measured by changes in the stock market index. They show that liquidity shocks are felt more in leading countries, whereas changes in risk appetite are felt more in emerging economies. The IMF (2010) analyzes the link between global liquidity and equity prices in emerging countries and presents evidence of a positive impact between 2003 and 2009. Matsumoto (2011) finds the same result in some Latin American countries. Finally, Brana et al. (2012), using a PVAR model, confirm the positive impact of surplus global liquidity on asset prices for a sample group of 16 emerging economies in Latin America and Asia.

Following the financial crisis of 2008, a growing body of literature has studied the effects of unconventional monetary policies on international financial markets[5] but empirical works about cross-border effects on emerging markets remain scarce. Fratzscher et al. (2013) analyze the effects of the Federal Reserve's unconventional policies on 65 foreign markets. They highlight the opposite effects of QE1 and QE2 on emerging asset prices via substantial rebalancing in global portfolios. Investors seem to have shifted out of emerging markets into U.S. equity and bond funds during the QE1 program, whereas the QE2 program prompted portfolio rebalancing in the opposite direction with strong capital flows into emerging markets. Using event study methodology and a GVECM model, Chen et al. (2011) provide empirical evidence on the short-run, cross-border effects of unconventional policies on asset prices in emerging economies, especially in Asia and Latin America. Over the long term, the expansionary impact seems to be stronger for some emerging economies than it is for the U.S. These results are in line with those of the IMF (2013) and Chinn (2013), although to this author, the impact seems to be mitigated by the exchange rates of some emerging economies. Morgan (2011) analyzes the impact of Federal Reserve LSAPs on Asian economies and financial markets and concludes that the LSAPs do not have a significant impact. In the same way, Moore et al. (2013) conduct an empirical analysis of the impact of LSAP announcements on ten emerging government bond market yields. They found that unconventional policies have contributed to U.S. outflows into emerging economies and explain marginal reductions in long-term government bond yields.

Our paper is part of the recent literature on the impact of unconventional monetary policies on asset prices especially in emerging markets. However, our approach, which extends recent research, differs from previous studies on different aspects, and our main contribution is threefold.

First, previous research has focused exclusively on linear models and neglected the possibility of nonlinearities in the relationship between monetary policy and asset prices. However, as noted by Beckmann et al. (2014) there are several reasons for nonlinearities in the context of a global monetary policy analysis. As previously mentioned, several authors have noted different—even opposing—effects of different programs of quantitative easing on asset prices in advanced and emerging economies according to different phases of the cycle (see Chen et al., 2011; Darius and Radde, 2010; Fratzscher et al., 2013; Glick and Leduc, 2012). Moreover, the usual channels of monetary policy transmission may have been impaired following the global financial crisis and pre-crisis relationships may have become obsolete (Chen et al., 2011). Our study introduces non-linearity into empirical methodology. To consider the non-linear response of emerging asset prices to quantitative easing measures of monetary policy, we use a panel threshold model developed by Hansen (1999). To our knowledge, such an empirical specification has not yet been used to account for the non-linear process between monetary policy and asset prices.

Second, along with the impact of global liquidity, our paper considers the literature on the impact of variations in global investor sentiments on financial stability (Bruno and Shin, 2012; Forbes and Warnock, 2012; González-Hermosillo, 2008). These empirical works focus on investors' risk appetites as a key determinant of capital flows and financial contagion. Jaramillo and Weber (2012) estimate the impact of a large drop in investor sentiment on bond yield for a set of emerging countries. We extend this literature to study the non-linear impact of unconventional measures on asset prices in emerging economies after controlling for the shift of international investor sentiment. In this paper, we present empirical evidence of a non-linear impact of global excess liquidity on equity prices using a panel threshold model for a set of 17 emerging market economies. More specifically, we use an index of global investor sentiment as a transition variable that separates "tranquil periods" from periods of financial stress. We find that global excess liquidity has a positive impact on asset prices during "tranquil" periods. However, when investors' risk aversion increases suddenly—i.e., when financial markets are under stress—the impact on asset prices changes.

Third, we calculate an original exhaustive global excess liquidity index for each country in our sample. In previous studies, global liquidity has referred only to monetary expansion in the U.S. or in some advanced economies. Cerutti et al. (2014) confirm the explanatory power of U.S. financial conditions on cross-border bank flows, but show that similar variables for other countries, like the U.K. and the Euro zone, are also important, sometimes even more so. Thus, the global liquidity variable cannot be restricted to U.S. monetary variables but must include other developed countries, including emerging countries that contribute to the growth of the global monetary base through the accumulation of foreign exchange reserves. [6] Our global excess liquidity indicator, based on monetary bases, takes into account 49 countries, including developed, newly industrialized and emerging countries. Moreover, in our empirical study, this global excess liquidity indicator is exogenous for each country, which allows us to consider spillover effects between countries.

The remainder of the paper is organized as follows. Section 2 describes the data and the methodology employed to calculate our global excess liquidity index. Section 3 presents the panel threshold

[5] See for example Neely (2013), Glick and Leduc (2013), or Bauer and Neely (2014).

[6] For example, Ratti and Vespignani (2014) show that increases in the BRIC countries' liquidity is associated with significant increases in commodity prices that are much larger than the effect of increases in G3 liquidity.

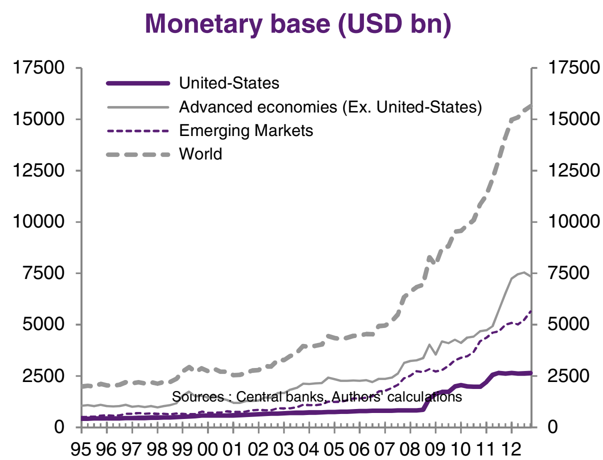


**Fig. 1.** Monetary base (USD bn).

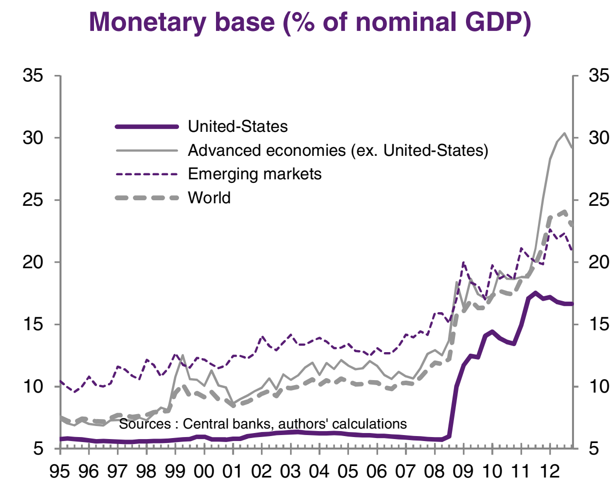


**Fig. 2.** Monetary base (% of nominal GDP).

model. Section 4 provides the empirical results and section 5 is the conclusion.

## 2. Data description

To investigate the impact of global excess liquidity on emerging stock markets, we use asset returns[7] as the dependent variable approximated by nominal equity returns (in U.S. dollars) for 17 emerging and newly industrialized economies: Argentina, Brazil, Chile, Colombia, Mexico, Peru, India, Indonesia, Malaysia, the Philippines, Hong Kong, South Korea, Taiwan, Thailand, Singapore, Turkey and South Africa. Those countries were considered because of their high degree of capital account openness and also because economic and financial data were available for a relative long period. Equity indices are obtained from Bloomberg.

We use a quarterly dataset ranging from 1995Q3 to 2011Q4, with the time series sample and country coverage determined for a most part by data availability.

Next, we build a new indicator to measure excess global liquidity. Because unconventional measures have directly impacted the balance sheets of central banks around the world, we consider the monetary base (i.e., the money supply aggregate M0) as the indicator of liquidity rather than broader monetary or credit aggregates, which have not been affected by quantitative easing policies (Mohan, 2013).

Then, we compute the ratio of global base money to global nominal GDP, which is commonly used as an indicator of global excess liquidity. However, as opposed to previous empirical research on the impact of global liquidity on asset prices (see, e.g., Baks and Kramer, 1999; Giese and Tuxen, 2008; Becker, 2007, 2009; Psalida and Sun, 2011), we calculate an indicator of excess liquidity for each country in our sample that considers money creation not only from some advanced economies but also from emerging markets.

In a preliminary step, we collect data on domestic monetary bases (M0) and nominal GDPs for a large sample of 49 advanced and emerging market countries.[8] Data are drawn from national central banks and the International Financial Statistics database, respectively. Because monetary bases and GDP are expressed in local currencies, we convert all domestic time series at each period in the same unit (U.S. dollar) by using current nominal exchange rates against the US dollar observed at the end of each quarter. Next, we create a series called the "world" monetary base by adding up the monetary bases of all those countries for each period. We also build a series called "world" GDP by adding nominal GDPs for this broad set of countries expressed in dollar terms for each period.[9]

The monetary bases of different areas (in billions of dollars and in % of GDP) are plotted in Figs. 1 and 2.

As Figs. 1 and 2 clearly illustrate, global liquidity increased sharply from the mid-1990s. It began to grow on the back of several interest rate cuts led by the Japanese monetary authorities, who were prompted by that economy's banking and financial crisis. Japan's low interest rate environment was coupled with the introduction of the Euro in 1999, which was accompanied by an increase in the monetary supply above the target established by the ECB (+4.5%);[10] the introduction of the Euro could also have played a major role in the increase in global liquidity. The excess liquidity indicator started to climb again on the back of large expansionist monetary policies pursued by central banks in advanced countries (the Federal Reserve, ECB, etc.) after the 2000 dot-com crisis. Not only advanced economies but also emerging markets and the OPEC countries contributed to the strong increase in global liquidity, essentially through the accumulation of foreign exchange reserves during the 2000s (see Figs. 3 and 4). Finally, the recent global financial crisis and the measures adopted by monetary authorities have again boosted liquidity. The impact of the recent financial crisis on global liquidity appears to be extremely large.

In the second step, we compute for each of the 17 countries of the sample our global excess liquidity indicator ($M0Y_{it}$). This indicator is calculated as the ratio of "world" monetary base from which we subtract the domestic money supply of each country to "world" nominal GDP, from which we also subtract domestic nominal GDP to avoid endogeneity. [11] The ratio is expressed as a percentage. Thus, $M0Y_{it}$ may be interpreted as the global excess liquidity the country $i$ faces at each period $t$, or as the excess liquidity that is likely to be invested in each country $i$ at each time.

In addition to our global excess liquidity indicator, the final dataset covers a set of various macroeconomic indicators largely used in empirical research as explanatory variables of asset prices (see references

[7] We focus on stock returns rather than stock prices to avoid the issue of non-stationary series (see below).

[8] The sample includes the largest advanced economies (in Europe: the Eurozone countries, The United Kingdom, Denmark and Sweden; in North America: the United States and Canada; In Asia/Pacific: Japan, Australia and New-Zealand), the traditional set of emerging countries (in Latin America: Argentina, Brazil, Chile, Colombia, Mexico, Peru and Venezuela; in Asia including newly industrialized countries: China, India, Indonesia, Malaysia, Philippines, Thailand, South Korea, Taiwan, Hong-Kong, Singapore), along with the largest CEEC (Poland, Hungary, Czech Republic, Bulgaria and Romania), including Russia and Turkey. We add South Africa and the sample also includes three major oil-exporting countries for which data was available: Qatar, Kuwait and Saudi Arabia. For these three last countries, the huge oil revenues lead to a sharp rise in foreign exchange reserves, and so, on monetary bases.

[9] "World" monetary base and nominal GDP are both expressed in billions of dollars.

[10] This argument must be viewed carefully in the context of exogenous factors linked to institutional and statistics changing.

[11] For example for Argentina, (%).

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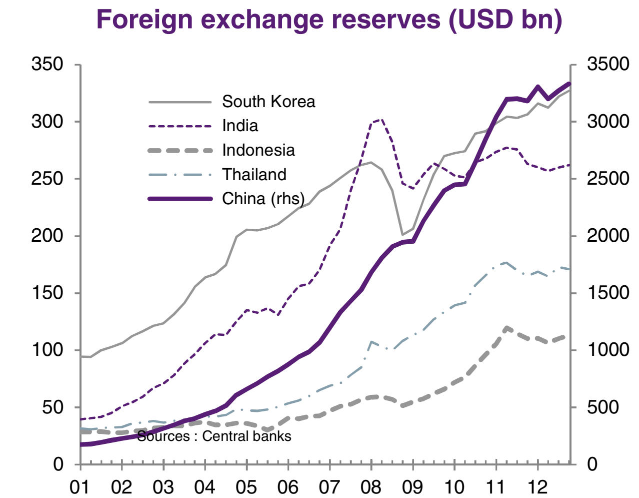


**Fig. 3.** Foreign exchange reserves (USD bn).

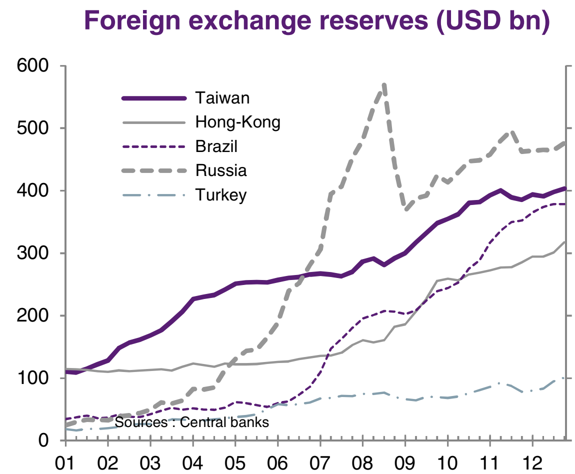


**Fig. 4.** Foreign exchange reserves (USD bn).

listed above). Following the IMF(2010), the explanatory factors can be divided into two groups:

- Domestic or fundamental factors include real GDP growth (GDP), the inflation rate based on the CPI (CPI), the three-month interbank rate (3Mrate), the domestic monetary base (M0) and the domestic broad money M2 (M2) expressed both in growth rates;
- A global factor defined as our global excess liquidity indicator (M0Y).

Expansionary monetary policy in normal times is supposed to have a positive impact on asset prices, as it changes the cost of capital (Bernanke and Kuttner, 2005) which in turn will increase the firms expected present value of cash flows. Monetary policy shocks also affect financial and housing wealth and leads to a shift of financial stress and of risk premium (Sousa, 2010, 2012)[12].

Finally, the regime-switching variable should reflect global investors' sentiment. As discussed by, inter alia, Kumar and Persaud (2002), McGuire and Schrijvers (2003), Cowan et al. (2006), Gonzales-Rozada and Levy-Yeyati (2008), Coudert and Gex (2008), Baldacci and Kumar (2010) and Jaramillo and Weber (2012), all concerning emerging market countries studies, the implied volatility of the *S&P500* stock index option prices (the Chicago Board of Options Exchange Market Volatility Index, hereafter the VIX index) has been traditionally used as a measure of risk aversion. Following this strand of literature, we proxy global investor sentiment by the VIX index. [13]

## 3. The panel threshold model methodology

To investigate the impact of excess global liquidity on asset prices in emerging market countries, we employ a panel threshold model. The evidence suggests that global liquidity has significant effects on the stock market. This result is consistent with other studies, like Ehrmann and Fratzscher (2009), Hausman and Wongswan (2011), or Kim and Nguyen (2009) who find, using a fixed coefficient approach, that foreign equity returns respond positively to an unexpected Fed rate cut. We extend this result by using global liquidity, instead of the sole US monetary policy. Moreover, we allow for a time varying response of foreign stock markets to global monetary surprise supposing a threshold effect of global liquidity on emerging asset prices depending on the level of global risk aversion in international financial markets. Kadilli (2014) investigates the predictability of stock returns in the financial market for a panel of developed countries and find evidence that a regime-switching model containing a normal and a crisis regime fits better the data than a linear specification. In the same way, Gang and Li (2014) find that the relationship between the expected S&P 500 index return and the volatility index (VIX) shows non-linearity and asymmetries. The risk-return behavior depends on the signs as well as the magnitudes of the perceived risk. We use a global investor sentiment indicator as a regime-switching indicator that separates periods of financial 'tranquility' from periods of 'financial stress'. According to recent empirical research, a shift in global risk aversion could affect the traditionally positive relationship between global excess liquidity and emerging asset prices (see *supra*).

Considering that the transition from a 'tranquil period' to a period of 'financial stress' is brutal, our empirical approach is based on Hansen's (1999) estimation and inference theory for non-dynamic panel data models. The panel threshold regression (hereafter PTR) model with individual specific effects is given by the following equation:

$$y_{it} = \mu_i + \beta_1' x_{it} \mathbb{I}(q_{it} \le \gamma) + \beta_2' x_{it} \mathbb{I}(q_{it} > \gamma) + \varepsilon_{it} \tag{1}$$

where $\mathbb{I}(\cdot)$ is the indicator function, $q_{it}$ is the threshold variable and $\gamma$ is the optimal threshold value. The subscripts $i$ and $t$ stand for the cross-section and time dimensions, respectively. The error term $\varepsilon_{it}$ is assumed independent and identically distributed (*iid*) with zero mean and a finite variance $\sigma^2$.

The dependent variable $y_{it}$ and the threshold variable $q_{it}$ are scalar matrices. The regressor $x_{it}$ is a $k \times 1$ vector of explanatory variables. All variables are assumed stationary to avoid a spurious regression model.

The observations are divided into two regimes depending on whether the threshold variable $q_{it}$ is smaller or larger than the threshold value $\gamma$. The individual effects $\mu_i$ are assumed the same in both regimes. Thus, the two regimes are distinguished by differing regression slopes $\beta_1$ and $\beta_2$.

Eq. (1) can also be written in a compact form:

$$\begin{aligned} y_{it} &= \mu_i + \beta' x_{it}(\gamma) + \varepsilon_{it} \\ x_{it}(\gamma) &= \begin{pmatrix} x_{it}\mathbb{I}(q_{it} \le \gamma) \\ x_{it}\mathbb{I}(q_{it} > \gamma) \end{pmatrix} \end{aligned} \tag{2}$$

where $\beta = (\beta_1' \, \beta_2')'$.

Following Hansen (1999), taking the averages of Eq. (2) over the time index $t$ produces the following equation:

$$\bar{y}_i = \mu_i + \beta' \bar{x}_i(\gamma) + \bar{\varepsilon}_i \tag{3}$$

[12] For monetary policy transmission in emerging market economies, see Mallick and Sousa (2012).
[13] The VIX index has been obtained from Bloomberg.

where $\bar{y}_i = \frac{1}{T}\sum_{t=1}^{T} y_{it}$, $\bar{\varepsilon}_i = \frac{1}{T}\sum_{t=1}^{T}\varepsilon_{it}$ and

$$\bar{x}_i(\gamma) = \frac{1}{T}\sum_{t=1}^{T} x_{it}(\gamma) = \begin{pmatrix} \frac{1}{T}\sum_{t=1}^{T} x_{it}\mathbb{I}(q_{it} \le \gamma) \\ \frac{1}{T}\sum_{t=1}^{T} x_{it}\mathbb{I}(q_{it} > \gamma) \end{pmatrix}.$$

The difference between Eqs. (2) and (3) yields the following:

$$y^*_{it} = \beta' x^*_{it} + \varepsilon^*_{it} \tag{4}$$

where $y^*_{it} = y_{it} - \bar{y}_i$, $x^*_{it}(\gamma) = x_{it}(\gamma) - \bar{x}_i(\gamma)$, and $\varepsilon^*_{it} = \varepsilon_{it} - \bar{\varepsilon}_i$.

Let

$$y^*_i = \begin{bmatrix} y^*_{i2} \\ \vdots \\ y^*_{iT} \end{bmatrix}; \quad x^*_i = \begin{bmatrix} x^*_{i2}(\gamma') \\ \vdots \\ x^*_{iT}(\gamma)' \end{bmatrix} \quad \text{and} \quad \varepsilon^*_i = \begin{bmatrix} \varepsilon^*_{i2} \\ \vdots \\ \varepsilon^*_{iT} \end{bmatrix}.$$

Then, let $Y^*$, $X^*(\gamma)$ and $\varepsilon^*$ denote the data stacked over all individuals. Using this notation, Eq. (4) is equivalent to

$$Y^* = X^*(\gamma)\beta + \varepsilon^*_{it}. \tag{5}$$

Then, for any given value of the threshold parameter $\gamma$, the slope coefficients $\beta_1$ and $\beta_2$ can be estimated by OLS. That is,

$$\hat{\beta}(\gamma) = \left(X^*(\gamma)'X^*(\gamma)\right)^{-1} X^*(\gamma)'Y^*. \tag{6}$$

Furthermore, the sum of squared errors (SSEs) dependent on any given value of $\gamma$ given by

$$SSE_1(\gamma) = \hat{\varepsilon}^*(\gamma)'\hat{\varepsilon}^*(\gamma). \tag{7}$$

To estimate endogenously the threshold parameter $\gamma$, Chan (1993) and Hansen (1999) recommend estimating the threshold value by least squares. This can be achieved by minimizing the sum of squared errors $SSE_1(\gamma)$. Therefore, the least square estimator *of* $\gamma$ is

$$\hat{\gamma} = \operatorname{Arg}\min_{\gamma} SSE_1(\gamma) \tag{8}$$

According to Hansen (1999), to avoid the issue of estimating a threshold value that sorts too few observations into one or the other regime, it would be convenient to restrict the set of values of $\gamma$ by excluding the smallest and largest $\eta$ % values of the threshold variable $q_{it}$ to assure that a minimal percentage of the observations is situated in each regime. In this paper, the lowest and highest 5% values are excluded. Then, given the estimated values of $\hat{\gamma}$, coefficients for each regime are given by $\hat{\beta}_1(\hat{\gamma})$ and $\hat{\beta}_2(\hat{\gamma})$.

The following step determines whether the threshold effect is statistically significant, thus validating the non-linearity of our model. This can be achieved by testing the null hypothesis $H_0 : \beta_1 = \beta_2$, for which there is no threshold effect.

However, under $H_0$, the threshold value $\gamma$ is not identified and the asymptotic distribution of $F_1$ is not standard. To overcome this issue, generally known as the 'Davies Problem' (Davies, 1977, 1987), Hansen (1996) suggests using a bootstrap procedure to attain the first-order asymptotic distribution of the likelihood ratio test of $H_0$.

$$F_1 = \frac{SSE_0 - SSE_1(\hat{\gamma})}{\hat{\sigma}^2}$$

where $\hat{\sigma}^2$ is the residual variance of the PTR, and $SSE_0$ is the sum of squared errors obtained from the linear model. Therefore, the *p*-values constructed from the bootstrap are asymptotically valid. The null hypothesis of no threshold effect is rejected if the *p*-value is smaller than the desired critical value.

Empirically, our threshold regression model can be specified in a compact form as follows:

$$EqReturns_{it} = \mu_i + \xi_1' X_{it}\mathbb{I}(q_t \le \gamma) + \xi_2' X_{it}\mathbb{I}(q_t > \gamma) + \varepsilon_{it} \tag{9}$$

where $i = 1, \ldots, N$ denotes the country-index and $t = 1, \ldots, T$ the time index. $\mu_i$ are the country specific effects and $\varepsilon_{it}$ is the i.i.d error term with zero mean and finite variance $\sigma^2$.

$\xi_j = (\delta_j\ \lambda_j\ \theta_j\ \varphi_j\ \alpha_j\ \nu_j\ \mu_j)'$ for $j = 1, 2$ and $X_{it} = (MOY_{it}\ GDP_{it}\ CPI_{it}\ 3Mrate_{it}\ M0_{it})'$ represent the vector of explanatory regime-dependant variable, and $\mathbb{I}(\cdot)$ is the indicator function.

The dependent variable $EqReturns_{it}$ represent nominal equity returns (in USD) for each country *i* at time *t*.[14]

As previously mentioned, the regime-switching variable $q_t$ is represented by the VIX index. Actually, the model allows the exogenous variables to have differing regression slopes depending on whether the threshold variable, the VIX, is above or below threshold $\gamma$.

## 4. Estimation results

The PTR model requires the variables to be stationary to avoid spurious regressions. Before proceeding to the estimation, we make sure that all variables are stationary. The results are shown in Table 3 in the Appendix. Overall, the tests strongly reject the null hypothesis of a unit root.

The first step of the estimation is to examine the threshold effect. Repeating the bootstrap procedures 300 times, we obtain the approximation of the *F*-statistic and associated *p*-value. The reported *F*- statistic assessing the null hypothesis of no threshold is 163.8533 with a bootstrap *p*-value of 0.0000 allowing us to clearly reject the linear structure of the model.

The estimated threshold value of the VIX index ($\hat{\gamma}$) is 25.61 with a 95% confidence interval = [25.61 26.85] (Fig. 5). [15] This level appears to be consistent with the estimated threshold value of the VIX in Jaramillo and Weber (2012) and also in Baldacci and Kumar (2010) where the VIX threshold was chosen exogenously.[16]

Fig. 5 allows us to clearly characterize periods of financial stress over the last twenty years, during which several financial crises affected not only emerging markets but also advanced economies. In our model, periods of financial stress are defined as occurring when the VIX index goes beyond the threshold level of 25.61.

The next step consists of estimating the slope coefficients of the PTR model with two regimes. As suggested by Hansen (1999), we specify that each regime must include at least 5% of all observations. The estimation includes our global excess liquidity variable (*M0Y*), a domestic liquidity variables (*M0 expressed in growht rate*)[17] and three main control variables (*GDP*, *CPI*, *3Mrate*).

Final results are reported in Table 1.

As previously mentioned, the threshold variable allows the sample to be divided into two regimes: a period of financial stress and a 'tranquil' period. During a 'tranquil' period, that is, when global risk aversion, proxied by the VIX index, is below the threshold value, emerging asset returns are positively influenced by global excess liquidity as expected. The estimated coefficient is found to be 0.85 and appear to be highly significant. This finding supports empirical evidence of

[14] Reverse causality may be present in the model, given that stock returns are also likely to influence the dynamics of the regressors. For the linkages between monetary policy and asset prices, see for example Castro and Sousa (2012) or Sousa (2014).

[15] The Matlab codes have been provided by Candelon et al. (2011) and Hurlin (2012).

[16] The authors find a threshold value of 25.56, which is statistically significant.

[17] The broad money proxied by *M2* did not appear to be statistically significant in all panel threshold regressions, so we do not report it.

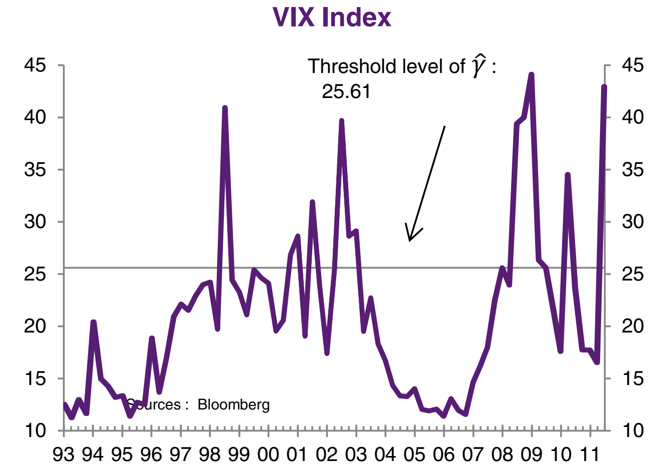


**Fig. 5.** VIX index.

**Table 1**
Threshold regression estimations results.

| | Regime 1: VIX ≤ 25.61 | | Regime 2: VIX > 25.61 | |
|---|---|---|---|---|
| | Coefficient estimate | t-Stats | Coefficient estimate | t-Stats |
| Global excess liquidity (M0Y) | 0.8520*** | 2.5730 | −0.6602** | −2.0203 |
| Real GDP growth (GDP) | 2.6230*** | 6.6595 | 1.5919*** | 5.0475 |
| Inflation rate (CPI) | 1.4455*** | 4.6847 | 0.2559 | 0.8628 |
| Three-month interbank rate (3Mrate) | −1.0332*** | −3.7238 | −0.7817*** | −2.7719 |
| M0 growth | 0.3577*** | 3.8845 | 0.1972* | 1.7380 |
| Estimated threshold parameter $\gamma = 25.61$ | | | | |
| Confidence interval (95%) on $\gamma$ = [25.61 28.62] | | | | |
| F-stat = 163.5101 | | | | |
| p-Value = 0.0000 | | | | |
| Number of simulations (for F statistic p-value) = 300 | | | | |
| Cross-unit dimension N = 17 | | | | |
| Time dimension T = 66 (from 1995Q3–2011Q4) | | | | |

Notes: ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

international spillovers of central bank balance sheet policies, especially with respect to emerging market countries. However it seems that when global risk aversion rises – the VIX index goes beyond the threshold value – the impact of global excess liquidity on equity returns in emerging markets turns negative. The estimated coefficient is found to be −0.6602 and statistically significant at the 5% level. This result appears to be consistent with several studies that suggest that the impact of monetary policy varies over time. Indeed empirical evidence supports the view that monetary policy exerts asymmetric effects on output and prices depending on the economic conditions (Weise, 1999, Peersman and Smets, 2005, Lo and Piger, 2005; Santoro et al., 2014).[18] Some authors study more precisely the impact of monetary policy on stock returns (Basistha and Kurov, 2008; Hsu and Chiang, 2011; Kontonikas et al., 2013; Kurov, 2010). They present evidence that the stock market response to monetary surprises is highly asymmetric and is stronger when the economy is in a recession. All these studies focus on the response of the US stock returns to US monetary policy, while Guo et al. (2013) find the same result for China. Kishor and Marfatia (2013) find a significant time-variation in the stock market response to US monetary policy surprises for all 35 countries they study, especially for the emerging markets.

In accordance with previous studies, we find that the underlying relationship between monetary policy and stock returns is non-linear. An important strand of the literature emphasizes financial-market stress as a crucial source of nonlinearity (Afonso et al., 2011, Mittnik and Semmler, 2013). The feature of investors' behavior changes in time of crisis. They decrease their risky investment stronger during recession than they increase them during booms (Apergis, 2014). This nonlinearity is amplified by bank's behavior. During the downside phase of the cycle, banks become more cautious and switched to more conservative lending rules. They are less willing to provide loans since the value of assets used as collateral started to decline.

However, while many studies show that the impact of monetary policy is reinforced in times of crisis, we find a different result: while global liquidity has a positive impact on asset prices in normal times, this impact becomes negative in times of financial stress.

This result is similar to that of Kontonikas et al. (2013). These authors examine the response of US stock returns to federal funds rate (FFR) surprises between 1989 and 2012 and conclude that an important structural shift occurred during the 2007–2009 financial crisis, changing the stock response to FFR shocks and the nature of state dependence. They find that during the crisis period, stocks did not react positively to FFR cuts. On the contrary, their estimates reveal a negative stock market response to unexpected FFR cuts within a sharply deteriorating economic and financial environment. They explain this result by a flight to safety behavior during the crisis, associated with a rebalancing of global investment portfolios, away from equities and towards less risky assets. In the same way, Kishor and Marfatia (2013) find a strong positive impact of the U.S. monetary policy shocks on emerging countries' stock price indexes during the normal time, but the relation gets weakened if the recent financial crisis period is included. The cut in the US interest rate during the 2008 financial crisis led to a fall in the stock returns across Europe and the North American markets. In a context of high financial stress, high money creation is not sufficient to halt the decline in stock prices. It may moreover reveal information about deteriorating economic outlook.

The other explanatory variables, the domestic macro-fundamental ones, have a more stable impact on the equity returns. The real GDP growth, the inflation rate and the domestic money supply have a positive impact on asset returns, whereas the three-month interest rate negatively influences equity returns, as expected. All those variables appear to be highly statistically significant, at the 1% level. Except for the inflation rate that becomes insignificant, all those variables still influence equity returns during period of low, as high, financial stress.

## 5. Robustness analysis

We focus on the high yield spread to assess the robustness of our previous estimates of the impact of global excess liquidity over the period of analysis. In particular, the US Corporate High-Yield spread is calculated as the difference between the yield of the USD-denominated, non-investment grade, fixed rate, taxable corporate bonds and the yield of the 10-year Treasury bonds expressed in basis point. To be included, securities must be rated high yield, that is Ba1/BB+/BB+ (the middle rating of Moody's, Fitch and S&P respectively) or below, excluding emerging market debt. In most studies, the high yield spread is also widely used as a proxy of global risk aversion, so as the VIX index (Kumar and Persaud, 2002, Cowan et al., 2006, Gonzalez Rozada and Levy Yeyati, 2008, Hesse et al., 2014). As expected, during period of financial stress where investors' appetite for risk shrink, the high yield spread is found to increase (Fig. 6).

The base sample spans the period 1995Q3–2011Q4. We perform the panel threshold methodology to test the impact of global excess liquidity on emerging asset prices taking into account the high yield spread as the transition variable. As for the VIX index, we expect a nonlinear relationship between those two variables, where it should be positive when

[18] Contributions focusing on the state dependence effects of the fiscal stance lead to the same result: fiscal multipliers are more pronounced during recessions than during booms (see for example Proaño et al., 2014).

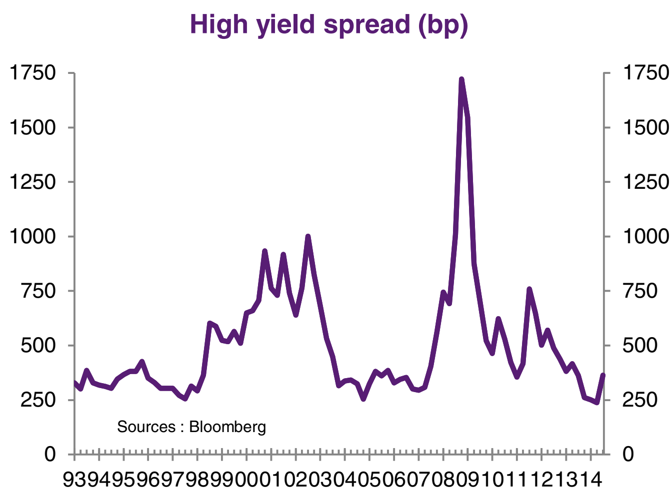


**Fig. 6.** High yield spread (bp).

global risk aversion is low and becomes non-significant or negative when global financial tensions increase.

In a first step, the model determines the existence of a threshold for the high yield spread variable. Repeating the bootstrap procedure 300 times, the *F*-stat and the associated *p*-value assessing the null hypothesis of no threshold is 164.3667 with a bootstrap *p*-value equal to 0.0000. This result confirms the nonlinear structure of the model. The second step is to estimate the threshold value for the high yield spread ($\hat{\gamma}$). It is found to be 565 (basis point) with a 95% confidence interval = [524.00 759.00]. All observations are split into two regimes depending on whether the threshold variable – the high yield spread – is smaller or larger than the estimated threshold value $\hat{\gamma}$. Accordingly, we define the two regimes to be low global risk aversion and high global risk aversion depending if the high yield spread is below 565 basis point or exceeds 565 basis point.

Table 2 reports the regression slope coefficients and the associated t-stats.

The results confirm the nonlinear relationship between the global excess liquidity variable and stock returns in emerging markets depending on the relative global financial stress level in financial markets. In the first regime, that is, when investors' appetite for risk is quite high, an increase in global liquidity leads to a rise in stock returns as expected. However, when global risk aversion builds up, international investors tend to sell emerging assets, triggering sudden-stop episodes and in worst case sudden withdrawal, with potential harmful consequences for financial stability. According to Mallick and Sousa (2013), the nexus between monetary stability and financial stability is strong and financial stress conditions can have an important impact on the likelihood of "boom-bust" episodes.

Concerning other macro-fundamental variables, the estimated coefficients of real GDP growth and inflation are found to be positive as expected, and are statistically significant. The three-month interbank interest rate is also significant with a negative sign whatever the level of global risk aversion. When domestic short rates rise, implying tightened monetary conditions, asset returns tend to fall.

These findings confirm that global excess liquidity has a nonlinear impact on stock returns in emerging countries depending on the global investors' risk aversion. These results therefore highlight that increasing global liquidity can be blamed to have fuelled asset price bubble in normal period in emerging markets, but also the ineffectiveness of monetary policy in a period of intense financial stress.

**Table 2**
Threshold regression estimation results with high yield spread.

| | Regime 1: HY Spread ≤ 565 | | Regime 2: HY Spread > 565 | |
|---|---|---|---|---|
| | Coefficient estimate | t-Stats | Coefficient estimate | t-Stats |
| Global excess liquidity (M0Y) | 1.0529*** | 3.3616 | −0.8679*** | −2.9810 |
| Real GDP growth (GDP) | 2.1527*** | 3.8815 | 1.8927*** | 5.0505 |
| Inflation rate (CPI) | 1.7007*** | 4.7040 | 1.0323*** | 2.3651 |
| Three-month interbank rate (3Mrate) | −1.4026*** | −5.0477 | −0.8413*** | −3.7214 |
| M0 growth | 0.3885*** | 3.8661 | 0.1292 | 1.1259 |
| Estimated threshold parameter $\gamma = 565.000$ | | | | |
| Confidence interval (95%) on $\gamma$ = [524.000 759.000] | | | | |
| F-stat = 158.9951 | | | | |
| p-Value = 0.0000 | | | | |
| Number of simulations (for F1 statistic p-value) = 300 | | | | |
| Cross-unit dimension N = 17 | | | | |
| Time dimension T = 66 (from 1995Q3-2011Q4) | | | | |

Notes: ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

## 6. Concluding remarks

Global excess liquidity, proxied by the ratio of world monetary base to world nominal GDP, increased steadily from the mid-1990s before accelerating sharply because of unconventional monetary policies established by central banks in advanced economies. Only a few studies have focused on the international spillovers of those policies, particularly with respect to their impact on emerging market countries. The primary conclusions that can be drawn from those works are mixed with respect to a positive relationship between liquidity and asset prices, although they do not consider the probable empirical nonlinearities in the relationship.

This paper explores the impact of global excess liquidity on asset returns for a large sample of emerging market economies over the period 1995–2012 using the panel threshold methodology proposed by Hansen (1999). Our main findings are twofold. First, the relationship between global liquidity and emerging equity returns appears to be non-linear. Second, we find evidence that global excess liquidity contributes to the rise in equity returns in emerging markets only during tranquil periods, i.e., when global risk aversion across financial markets is low. On the contrary, during periods of financial stress, the impact of global excess liquidity becomes negative, while domestic liquidity variables remain important determinants of asset returns in these countries.

Our results contribute to the recent debate around the cross-border impact of monetary policies defined by major central banks, in a context in which the U.S. Federal Reserve plans to slow money printing (FED QE tapering) before an expected normalization of monetary conditions. Moreover, if the massive money creation has proven to be efficient in countering liquidity shortages that have affected some market segments, it seems that money creation is insufficient to counter the collapse of asset markets, in particular when global risk aversion is high. Finally, our empirical study seems to confirm the cross-border effects of monetary policies. Our results could justify the need for better coordination at global level and could explain the return of capital controls in several emerging markets as a tool against financial instability.

## Appendix. Unit root tests

The null hypothesis of non-stationary versus the alternative, in which the variable is stationary, has been tested. First, we run the Pesaran (2007) second-generation panel unit root test to address the problem of cross-sectional dependencies.[19] The Pesaran test assumes that cross-section dependence takes the form of a single, unobserved factor. It has satisfactory size and power even for relatively small values of N and T. As the individual CADF (Cross Augmented Dickey Fuller) and the panel statistic (CIPS) have non-normal distributions, their critical values (for different N and T) are obtained by Monte Carlo simulations.

**Table 3**
Unit root tests.

| | Pesaran (2007) panel unit root test (CIPS[a]) | | | Pesaran panel unit root test CADF[b] | | |
|---|---|---|---|---|---|---|
| Lag length | 1 | 2 | 3 | 1 | 2 | 3 |
| Real GDP growth | −8.952 | −8.496 | −8.737 | −3.831 | −3.588 | −3.528 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Inflation rate | −9.695 | −7.941 | −6.604 | −3.535 | −3.192 | −3.078 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Equity returns | −9.154 | −9.822 | −12.749 | −3.153 | −3.126 | −3.673 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| M2 growth | −5.682 | −6.589 | −6.568 | −3.447 | −3.473 | −3.555 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| M0 growth (domestic) | −5.344 | −5.185 | −7.078 | −2.972 | −2.937 | −3.359 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Three-month interbank rate | −5.702 | −3.686 | −2.563 | −3.052 | −2.602 | −2.352 |
| | (0.000) | (0.000) | (0.005) | (0.000) | (0.000) | (0.005) |
| Global excess liquidity | −4.26 | −1.252 | 0.291 | −2.731 | −2.059 | −1.715 |
| | (0.000) | (0.105) | (0.614) | (0.000) | (0.105) | (0.614) |
| | Maddala and Wu (1999) panel unit root test (MW) | | | Fisher test for panel unit root using the Phillips Perron test (3 lags) | | |
| VIX | 148.446 | 106.735 | 78.789 | $Chi^2(32) = 272.49$ | | |
| | (0.000) | (0.000) | (0.000) | (0.000) | | |

Null for MW, CADF and CIPS tests: series is I(1).
As shown in the table, the nulls of the unit root are all rejected.
[a] Cross-sectionally augmented IPS (Im, Pesaran and Shin).
[b] Cross-sectionally augmented DF.

Second, for the VIX (which is constant across countries), we use the Maddala and Wu (1999) first-generation test. The results are presented in Table 3 below.

[19] The Breusch–Pagan LM test for independence clearly reject H0 at 1% level [$Chi^2(136) =$ 2424,471]. This result is confirmed by the Pesaran test of cross sectional independence (46,646).

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