Journal of Banking and Financial Economics

No 2(6)2016

University of Warsaw Faculty of Management

ISSN 2353-6845

Editorial Board

Editorial Office: Faculty of Management, University of Warsaw, Szturmowa Str. 1/3, Postal Code 02-678 Warsaw, Email: jbfe@wz.uw.edu.pl

Publisher: Faculty of Management Publishing House, University of Warsaw, Szturmowa Str. 1/3, Postal Code 02-678 Warsaw
Telephone: +48 22 55 34 164; Fax: +48 22 55 34 001; jbfe@wz.uw.edu.pl

Webpage of Journal of Banking and Financial Economics:

ttp://www.wz.uw.edu.pl/portale/journal-of-banking-and-financial-economics

Overview

The ***Journal of Banking and Financial Economics*** **(JBFE)** is an open access journal. The submission of manuscripts in free of fee payment. This journal follows a double-blind reviewing procedure.

Aims and Scope

JBFE publishes high quality empirical and theoretical papers spanning all the major research fields in banking and financial economics. The aim of the journal is to provide an outlet for the increasing flow of scholarly research concerning banking, financial institutions and the money and capital markets within which they function. The journal also focuses on interrelations of financial variables, such as prices, interest rates and shares and concentrates on influences of real economic variables on financial ones and vice versa. Macro-financial policy issues, including comparative financial systems, the globalization of financial services, and the impact of these phenomena on economic growth and financial stability, are also within the *JBFE*'s scope of interest. The Journal seeks to promote research that enriches the profession's understanding of the above mentioned as well as to promote the formulation of sound public policies.

Main subjects covered include, e.g.: [1] ***Valuation of assets***: Accounting and financial reporting; Asset pricing; Stochastic models for asset and instrument prices; [2] ***Financial markets and instruments:*** Alternative investments; Commodity and energy markets; Derivatives, stocks and bonds markets; Money markets and instruments; Currency markets; [3] ***Financial institutions, services and regulation***: Banking efficiency; Banking regulation; Bank solvency and capital structure; Credit rating and scoring; Regulation of financial markets and institutions; Systemic risk; [4] ***Corporate finance and governance****:* Behavioral finance; Empirical finance; Financial applications of decision theory or game theory; Financial applications of simulation or numerical methods; Financial forecasting; Financial risk management and analysis; Portfolio optimization and trading.

Special Issues

JBFE welcomes publication of Special Issues, whose aim is to bring together and integrate work on a specific theme; open up a previously under-researched area; or bridge the gap between formerly rather separate research communities, who have been focusing on similar or related topics. Thematic issues are strongly preferred to a group of loosely connected papers.

Proposals of Special Issues should be submitted to at jbfe@wz.uw.edu.pl. All proposals are being reviewed by the Editorial Team on the basis of certain criteria that include e.g.: the novelty, importance and topicality of the theme; whether the papers will form an integrated whole; and the overall 'added value' of a Special Issue.

CONTENTS

Comparing the Performance of Logit and Probit Early Warning Systems for Currency Crises in Emerging Market Economies

Fabio Comelli[1]
International Monetary Fund, USA
fcomelli@imf.org

Received: 23 January 2016 / Revised: 13 April 2016 / Accepted: 13 April 2016 / Published online: 16 June 2016

ABSTRACT

We compare how logit (fixed effects) and probit early warning systems (EWS) predict in-sample and out-of-sample currency crises in emerging markets (EMs). We look at episodes of currency crises that took place in 29 EMs between January 1995 and December 2012. Stronger real GDP growth rates and higher net foreign assets significantly reduce the probability of experiencing a currency crisis, while high levels of credit to the private sector increase it. We find that the logit and probit EWS out-of-sample performances are broadly similar, and that the EWS performance can be very sensitive both to the size of the estimation sample, and to the crisis definition employed. For macroeconomic policy purposes, we conclude that a currency crisis definition identifying more rather than less crisis episodes should be used, even if this may lead to the risk of issuing false alarms.

JEL Classification: F30, F32, F37

Keywords: Early warning systems, currency crises, out-of-sample performance.

I. INTRODUCTION

The global financial crisis of 2008–2009 has revived the interest of professional economists in designing and assessing the performance of early warning systems (EWS), a class of models employed to quantify the likelihood of observing financial crisis episodes in the short term. In this context, the goal of this study is to compare how two competing parametric limited dependent variable (fixed effects logit and random effects probit) EWS predict in-sample and out-of-sample currency crises in emerging market economies (EMs).[2] What makes our empirical analysis interesting is that we use a rich panel dataset which includes macroeconomic and external vulnerability indicators for 29 EMs, with monthly data between January 1995 and December 2012.

[1] Corresponding author: Fabio Comelli, International Monetary Fund, 700 19th Street N.W., Washington, D.C. 20431, USA. Telephone: +1/202/623-5403, Fax: +1/202/589-5403.

[2] Ideally one would compare the performance of a fixed effects logit estimator with that of a fixed effects probit estimator. However, as in Wooldridge (2002), the estimation of unobserved country-specific effects along with the estimation of the explanatory variables coefficients leads to obtain inconsistent estimates of the latter, particularly if the length of the panel is small.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

In the EWS literature, currency crises are usually defined as large depreciations of the nominal exchange rate and/or extensive losses of foreign exchange reserves over a 24–month forecast horizon. Specifically, a currency crisis is said to occur when the exchange market pressure index – a weighted average of one-month changes in the exchange rate and foreign exchange reserves – is two or three (country-specific) standard deviations above its (country-specific) mean.[3] In this context, a relevant question is: should a currency crisis be defined as a situation when the exchange rate pressure index is two or three standard deviations above its mean? Since there is no clear consensus in the EWS literature about which crisis definition should be used, we attempt to fill this gap by taking an agnostic approach. We use two definitions of currency crisis. According to one definition, a currency crisis occurs when the exchange rate pressure index is *two* standard deviations above its mean, while according to the other a crisis occurs when the index is *three* standard deviations above its mean. For each EWS, we are interested to establish how the performance changes if we use one crisis definition or the other.

In addition, this study contributes to the EWS literature as follows. We conduct a cross-country empirical analysis to compare how two competing limited dependent variable EWS perform in predicting out-of-sample currency crises episodes. As in Candelon and others (2012) and Comelli (2013), we assess in-sample and out-of sample EWS performance by calculating optimal cut-off values for the estimated crisis probability, while in most of the EWS literature those cut-off values are selected arbitrarily. This matters because the cut-off value for the crisis probability determines the total misclassification error of a EWS.[4] Selecting cut-off values arbitrarily implies that the quantification of the total misclassification error is also arbitrary.

We find that stronger real GDP growth rates and higher net foreign assets significantly reduce the probability of experiencing a currency crisis, while high credit to the private sector increases it. By contrast, the current account balance and the measure of real exchange rate misalignment are not always statistically significant. Overall, the logit and probit EWS out-of-sample performances are broadly similar. The logit EWS is able to classify correctly between 42% and 66% of the total out-of-sample observations (e.g. crisis and tranquil periods), while the probit EWS is able to classify correctly between 41% and 64% of the total out-of-sample observations. We also find that the EWS performance is sensitive to the size of the estimation sample, and to the crisis definition used. In particular, both EWS perform better when a crisis episode is defined as a situation when the country–specific exchange market pressure index is two standard deviations above its mean.

The results offer two macroeconomic policy conclusions. First, as is common in the EWS literature, the EWS out-of-sample performance can be very sensitive to the size of the estimation sample. Specifically, the EWS total misclassification error and the probability of experiencing a currency given a crisis alarm can vary considerably if a particular year with many outlying observations (this is the case for the year 2008) is included in the estimation sample. This suggests that the total misclassification error and the probability of observing a currency crisis may crucially depend on new economic and financial data. Second, the results imply that selecting a crisis definition as a situation when the exchange rate pressure index is two standard deviations above its average value reduces the EWS total misclassification error. Therefore, for macroeconomic policy purposes, a currency crisis definition identifying more rather than less crisis episodes should be employed in a EWS, even if this may lead to the risk of issuing false alarms.

This paper is organized as follows. Section 2 reviews the relevant literature, while section 3 discusses the methodology used in this study. Section 4 presents the results obtained with the logit (fixed effects) and probit EWS, while in section 5 we compare the out-of-sample performance of the logit and probit EWS. Concluding remarks are presented in section 6.

---

[3] See IMF (2002) and Bussiere and Fratzscher (2006).

[4] The total misclassification error of an early warning system (EWS) is the sum between the percentages of missed crisis episodes and of false alarms issued by the EWS.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

2. LITERATURE REVIEW

Following the episodes of severe financial distress in Mexico (1994–95) and Asia (1997–98), economists became interested in thinking about frameworks that could help policymakers anticipating episodes of financial crises, whose economic costs are well documented (Cerra and Saxena, 2008).

We divide the EWS literature contributions relevant for this study in two groups. The first group includes those studies that propose parametric (i.e. regression-based) and non-parametric (i.e. crisis signal extraction) EWS and assess in-sample and out-of-sample performances of different EWS. Kaminsky, Lizondo and Reinhart (KLR) look at the evolution of those indicators which exhibit an unusual behavior in periods preceding financial crises. When the indicator exceeds a given threshold then that indicator is issuing a signal that a crisis could take place within the next 24 months. They find that exports, measures of real exchange rate overvaluation, GDP growth, the ratio between the money stock and foreign exchange reserves and equity prices have the best track record in terms of issuing reliable crisis signals. Berg and Pattillo (1999) test the KLR model out-of-sample and show that their regression-based approach tends to produce better forecasts compared to the KLR model.

Bussiere and Fratzscher (2006) develop a multinomial logit regression-based EWS, which allows distinguishing between tranquil periods, crisis periods and post-crisis periods. They show that the multinomial logit model tends to predict better than a binomial logit model episodes of financial crisis in emerging market economies. Beckmann and others (2007) compare parametric and non-parametric EWS using a sample of 20 countries during the period included between January 1970 and April 1995. They find that the parametric EWS tends to perform better than non-parametric EWS in correctly calling financial crisis episodes. However, as noted by Candelon and others (2012), in these studies the choice of the crisis probability cut-off value is arbitrarily made and not optimally derived. Comelli (2013) compares the performance of parametric and non-parametric EWS for currency crises in 28 emerging market economies and finds that the parametric EWS achieves superior out-of-sample results compared to the non-parametric EWS.

The second group of relevant EWS literature contributions for this study includes studies that discuss the significance of the various macroeconomic indicators to explain crisis incidence. Berkmen and others (2012) looked at the change in growth forecasts by professional economists before and after the global financial crisis. They found that countries with more leveraged domestic financial systems and rapid credit growth tended to suffer larger downward revisions to their growth forecasts, while international reserves did not play a significant role. Similarly, Blanchard and others (2010) do not find a significant role played by reserves in explaining unexpected growth, which is defined as the forecast error for output growth in the semester from October 2008 until March 2009. Rose and Spiegel (2012) find that the only robust predictor of crisis incidence in the 2008 global financial crisis is the size of the equity market prior to the crisis. They are unable to link most of the other commonly cited causes of the global financial crisis to its incidence across countries. By contrast, Gourinchas and Obstfeld (2011) look at financial crisis episodes in advanced and emerging economies from 1973 until 2010. They find that for both advanced and emerging market economies, the two most robust predictors are domestic credit growth and real currency appreciation. In addition, they find that in emerging market economies the country's level of foreign exchange reserves is a significant factor in determining the probability of future crises. Borio and Drehmann (2008) build indicators to quantify financial imbalances (based on credit, equity prices and property prices). These indicators are informative about financial strains in a given country but, since they do not take into account cross-border exposures of banking systems, are unable to anticipate crisis episodes associated with losses on foreign portfolios, even when the domestic economy does not exhibit credit or asset price booms.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Llaudes and others (2010) find that foreign exchange reserve holdings helped to mitigate the growth collapse in EMs provoked by the global financial crisis.

Frankel and Saravelos (2012) estimate the crisis incidence of the 2008–2009 global financial crisis. They surveyed the existing literature on early warning indicators to see which leading indicators were the most reliable in explaining the crisis incidence. They find that foreign exchange reserves, the real exchange rate, credit growth, real GDP growth and the current account balance as a percentage of GDP are the most reliable indicators to explain crisis incidence and conclude that the large accumulation of foreign exchange reserves has played an important role in reducing countries' vulnerability during the global financial crisis. The results obtained in this study are in line with the notion that the stock of foreign exchange reserves is significantly negatively related with our measure of crisis incidence. Finally, Goldman Sachs (2013) estimates the probability of a sudden stop in capital inflows across emerging market economies. They find that an increase of 25 basis points in the ratio between foreign exchange reserves and short-term external debt has the same impact as a one percentage point improvement in the current account balance (as a percentage of GDP).

3. METHODOLOGY

We build two competing early warning systems (EWS) and compare their ability to correctly predict in-sample and out-of-sample episodes of currency crises in EMs in the period between January 1995 and December 2012.[5] We proceed as follows. We build an exchange rate pressure index from which we derive a crisis variable (or crisis incidence) that identifies episodes of currency crisis in EMs. The crisis variable is binary, as it assumes the value of one if a currency crisis takes place within the next 24 months, and 0 otherwise.

Once defined the crisis variable, we proceed to construct the logit and probit EWS, where the crisis variable is regressed on a set of selected external vulnerability indicators of EMs. A crisis probability is then calculated with the coefficient estimates obtained from the regression. Afterwards, we describe how we select the optimal cut-off value for the estimated crisis probability.

Formally, we follow Bussiere and Fratzscher (2006) and assume that that there are N countries, $i = 1, 2, ..., N$, that we observe during T periods $t = 1, 2, ..., T$. For each country and month, we observe a forward-looking crisis variable Y_{it} that can assume as values only 0 (non-crisis) or 1 (crisis). To derive the crisis binary variable, we follow Kaminsky and others (1998) and build an exchange rate pressure index.[6] The exchange rate pressure index for country i at time t ($ERPI_{it}$) is defined as a weighted average between the monthly change in the nominal exchange rate and that in the stock of foreign exchange reserves:

$$ERPI_{it} = \frac{e_{it} - e_{it-1}}{e_{it-1}} - \left(\frac{s_{e_i}}{s_{fxr_i}}\right)\frac{fxr_{it} - fxr_{it-1}}{fxr_{it-1}} \tag{1}$$

where e_{it} is the nominal exchange rate of country i's currency against the U.S. dollar at time t, and fxr_{it} is the stock of foreign exchange reserves held by country i at time t. Finally, s_{e_i} and s_{fxr_i} are the standard deviations of the nominal exchange rate and the stock of foreign exchange reserves, respectively.

[5] The emerging market economies that we consider are: Argentina, Brazil, Bulgaria, Chile, China, Colombia, Croatia, Czech Republic, Egypt, Hungary, India, Indonesia, Kazakhstan, Korea, Malaysia, Mexico, Pakistan, Peru, Philippines, Poland, Romania, Russia, South Africa, Taiwan, Thailand, Turkey, Ukraine, Uruguay and Vietnam.

[6] For a discussion on exchange rate pressure indices, see Eichengreen and others (1995).

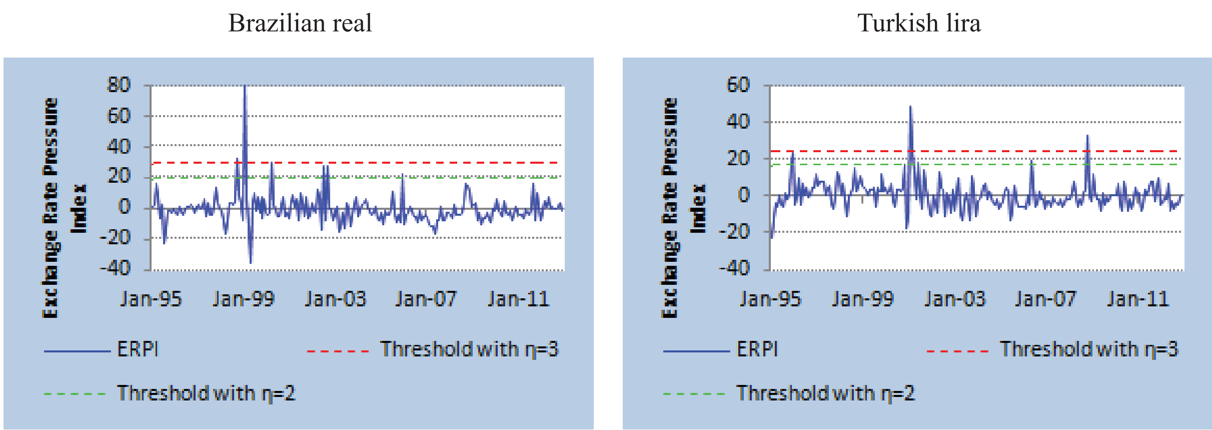
DOI: 10.7172/2353-6845.jbfe.2016.2.1

As a next step, we define a currency crisis hitting country *i* at time *t*, CC_{it}, as a binary variable that can assume either 1 (when the *ERPI* is above its mean by a number η of standard deviations) or 0 (otherwise):

$$CC_{it} = \begin{cases} 1 & \text{if } ERPI_{it} \geq \overline{ERPI_i} + \eta\sigma_{ERPI_i} \\ 0 & \text{otherwise} \end{cases} \tag{2}$$

where $\overline{ERPI_i}$ denotes the (country–specific) mean of the exchange rate pressure index, σ_{ERPI_i} denotes its standard deviation, which is multiplied by the weight η. Condition (2) states that a currency crisis is observed if the exchange rate pressures index of country i at time t is equal or larger than a country–specific threshold. The threshold is calculated as the sum between the mean of the (country–specific) exchange rate pressure index and the product between a coefficient η and the standard deviation of the (country–specific) exchange rate pressure index. In the EWS literature, η typically assumes the values of three.[7] The choice of values to assign to η determines the position of the exchange rate pressure index threshold: if η=3, condition (2) implies that the threshold of the exchange rate pressure index is higher than when η=2. Because of the higher threshold, the index identifies less crisis episodes compared to when η=2. The choice of η involves a trade-off. With a low crisis threshold, an early warning system may miss few crisis episodes but, at the same time, issue several false alarms. By contrast, with a high threshold, an early warning system may issue few false alarms, but may miss several crisis episodes. As an illustration, figure 1 plots the exchange rate pressure index for selected EMs, and the thresholds for the index when η=2 and when η=3.

Figure 1
Exchange Rate Pressure Index and Thresholds: January 1995–December 2012.



Source: International Financial Statistics and Author's calculations.

Unlike most of the EWS empirical literature, we let η assuming both values (two and three), and then compare how the EWS perform when η assumes one value or the other. Put differently, we are interested to see how the choice of the value to assign to η affects the EWS in-sample and out-of-sample performance. When η=2, the exchange rate pressure index identifies 191 crisis episodes across the set of emerging market economies, between January 1995 and December 2012. When η=3, the exchange rate pressure index threshold is higher and the index identifies only 77 crisis episodes in the panel.

Next, the variable CC_{it} is converted into the forward-looking crisis variable Y_{it} which is defined as follows

[7] See IMF (2002).

DOI: 10.7172/2353-6845.jbfe.2016.2.1

$$Y_{it} = \begin{cases} 1 \text{ if } \exists k = 1, 2, ..., 24 \text{ s.t. } CC_{it} = 1 \\ 0 \text{ otherwise} \end{cases} \tag{3}$$

The forward-looking crisis variable Y_{it} is equal to 1 if within the next 24 months a currency crisis is observed in country i, and to 0 otherwise. As in Bussiere and Fratzscher (2006), the crisis definition adopted in this study allows capturing both successful and non-successful speculative attacks to a given currency. Finally, conditions (2) and (3) imply that the crisis variable Y_{it} will also depend on the choice of η. Since we allow η to assume the value of two or three, as a result we will have two crisis variables, one defined when η=2, and one when η=3.

We define Pr(Y_{it}=1) as the probability of country i to experience a currency crisis at time *t*. We estimate the probability of a currency crisis following two approaches. In the first approach, we estimate the probability of currency crisis using a fixed effects logit model. In the second approach we estimate the probability of currency crisis with a probit model. More formally, in each model the probability of a currency crisis is expressed as a non-linear function of a given set of explanatory variables *X*:

$$(\Pr Y_{it} = 1) = \Lambda(X'\beta) = \frac{e^{X'\beta}}{1 + e^{X'\beta}} \tag{4}$$

$$(\Pr Y_{it} = 1) = \Phi(X'\beta) = \int_{-\infty}^{X'\beta} j(z)dz \tag{5}$$

where $\Lambda(X'\beta)$ denotes the cumulative distribution function (cdf) of the logistic distribution, while $\Phi(X'\beta)$ denotes the cdf of the normal distribution. Conditions (4) and (5) express the conditional probabilities that country *i* experiences a currency crisis at time *t* as a function of selected external vulnerability indicators, denoted by *X*. The crisis binary variable Y_{it} is regressed on the external vulnerability indicators *X* in the period January 1995–December 2012, using logit (fixed effects) and probit estimation techniques.

The explanatory variables that we use in (4) and (5) are external vulnerability indicators. Following Goldman Sachs (2013), we select the following external vulnerability indicators: the ratio between the stocks of foreign exchange reserves and short-term external debt, the current account balance as a percentage of nominal GDP, the real GDP growth rate and a measure of the real effective exchange rate misalignment. In addition, we use the stocks of net foreign assets and credit to the private sector, both expressed as percentages of nominal GDP.

Once obtained the logit (fixed effects) and probit coefficient estimates, we derive the estimated probability of experiencing a currency crisis from conditions (4) and (5). Then, we choose a cut-off value for the estimated crisis probability in order to assess the performance of the logit and probit EWS. The cut-off value is chosen such that the total misclassification error (TME) is minimized. The TME is calculated as the sum between the percentage of missed crisis episodes (expressed as the ratio between missed crisis calls over the total number of crisis called) and the percentage of false alarms (expressed as the ratio between false alarms over the total number of tranquil period called). Formally:

$$\text{TME} = \text{Type 1 error} + \text{Type 2 error} \tag{6}$$

Where

$$\text{Type 1 error} = \frac{\text{Total missed crisis episodes}}{\text{Total crisis episodes}} \tag{7}$$

$$\text{Type 2 error} = \frac{\text{Total false alarms}}{\text{Total non-crisis episodes}} \tag{8}$$

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Finally, note that in (6), type 1 and type 2 errors are equally weighted. In Comelli (2013), type 1 and type 2 errors are allowed to assume different weights in the TME. Changing the weights of type 1 and type 2 errors affects the EWS in-sample and out-of-sample performances.

4. RESULTS

We use a panel containing monthly observations of external vulnerability indicators for 29 emerging market economies, for the period included between January 1995 and December 2012.[8] The indicators are the ratio between the stocks of foreign exchange reserves and short-term external debt, the current account balance as a percentage of nominal GDP, the real GDP growth rate, a measure of real effective exchange rate misalignment, the stocks of net foreign assets and credit to the private sector, both expressed as a percentages of nominal GDP. The dependent variable is the crisis binary variable, which is regressed on the lagged indicators using logit (fixed effects) and probit estimation techniques. Since we allow the coefficient η in the exchange rate pressure index (section III) to assume either the value of two or three, we have two different dependent variables: one where the crisis variable is defined with $\eta=2$, and another where the crisis variable is defined with $\eta=3$. Tables 1 and 2 report the coefficient estimates obtained with logit and probit panel regressions. In each table, panel A reports the estimates obtained when $\eta=2$, while panel B reports the estimates when $\eta=3$.

For each estimation technique, seven different EWS specifications have been estimated.[9] The coefficient estimates across the seven EWS specifications tend to have the correct sign. Stronger real GDP growth rates and higher net foreign assets as a percentage of nominal GDP tend to reduce significantly the probability of experiencing a currency crisis episode. By contrast, in most of the specifications a higher stock of credit to the private sector as percentage of nominal GDP is significantly associated to a higher currency crisis incidence. The ratio between foreign exchange reserves and short–term external debt is correctly signed but is statistically significant only when $\eta=2$, in which case there are 191 crisis episodes identified in the panel. Conversely, the ratio loses significance when $\eta=3$, in which case there are only 77 crisis episodes identified in the panel. Put differently, the statistical significance of the ratio between foreign exchange reserves and short–term external debt depends on the choice of η, hence on the number of crisis episodes identified. The current account balance as a percentage of nominal GDP has the correct sign – e.g. a higher current account balance is associated with a decline in crisis incidence – but the estimates are not significant. Similarly, the measure of real exchange rate misalignment (deviations from the 3-year moving average) has the correct sign – a systematically overvalued currency raises the likelihood of experiencing a currency crisis – but the estimates are not always significant.[10]

8 See annex for a description of the variables.

9 We also run regressions for the seven EWS specifications using a logit random effects estimator. For some (but not all) of the EWS specifications, the Hausman Specification Test leads to reject the null hypothesis according to which the logit random effects estimator is efficient.

10 As an alternative measure of real exchange rate misalignment, in the logit and probit regressions we tried to include among the explanatory variables the real exchange rate deviations from a deterministic time trend. The coefficient estimates turned out not to be significant across the logit and probit specifications.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Table 1
Logit Fixed Effects Regression: Coefficient Estimates

				Panel A: η=2			
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
FXR/STED	–0.006*** (0.002)	–	–0.006*** (0.002)	–0.006*** (0.002)	–0.003*** (0.002)	–0.007*** (0.002)	–0.006*** (0.001)
CAB/Y	–0.029 (0.037)	–0.057 (0.036)	–	–0.011 (0.034)	–0.022 (0.036)	–0.043 (0.035)	–0.057 (0.035)
ΔY	–0.107** (0.040)	–0.098** (0.040)	–0.098** (0.038)	–	–0.118*** (0.039)	–0.136*** (0.036)	–0.136*** (0.037)
REERM	0.002 (0.002)	0.003** (0.001)	0.002 (0.001)	0.003** (0.001)	–	0.002 (0.001)	0.001 (0.001)
NFA/Y	–0.028 (0.020)	–0.057*** (0.018)	–0.028 (0.020)	–0.051*** (0.018)	–0.029 (0.020)	–	–0.027 (0.019)
PRCR/Y	0.022** (0.009)	0.026*** (0.009)	0.023** (0.009)	0.031*** (0.009)	0.018* (0.009)	0.014* (0.008)	–
Observations	4029	4029	4029	4029	4029	4255	4110
ROC Statistics	0.668	0.624	0.638	0.638	0.677	0.665	0.691
				Panel B: η=3			
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
FXR/STED	–0.003 (0.003)	–	–0.004 (0.003)	–0.003 (0.003)	–0.004 (0.003)	–0.008** (0.003)	–0.005 (0.003)
CAB/Y	–0.077 (0.061)	–0.089* (0.060)	–	–0.036 (0.053)	–0.067 (0.058)	–0.096* (0.057)	–0.118** (0.058)
ΔY	–0.124* (0.068)	–0.121* (0.067)	–0.093 (0.062)	–	–0.141** (0.067)	–0.188*** (0.061)	–0.192*** (0.062)
REERM	0.004* (0.002)	0.004* (0.002)	0.003 (0.002)	0.005** (0.002)	–	0.002 (0.002)	0.001 (0.002)
NFA/Y	–0.067* (0.036)	–0.082** (0.033)	–0.061* (0.035)	–0.094*** (0.032)	–0.066* (0.004)	–	–0.048 (0.003)
PRCR/Y	0.046** (0.020)	0.049** (0.019)	0.051*** (0.019)	0.059*** (0.019)	0.035** (0.017)	0.012 (0.013)	–
Observations	2961	2961	2961	2961	2961	3176	3042
ROC Statistics	0.649	0.616	0.615	0.596	0.699	0.723	0.753

***: Significant at 1%, **: Significant at 5%, *: Significant at 10%.

Sources: Joint External Debt Hub (www.jedh.org), Author's calculations based on International Financial Statistics.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Table 2
Probit Regression: Coefficient Estimates

Panel A: η=2							
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
FXR/STED	–0.001** (0.000)	–	–0.001*** (0.000)	–0.001*** (0.000)	–0.001** (0.000)	–0.001*** (0.000)	–0.001** (0.000)
CAB/Y	–0.010 (0.012)	–0.019* (0.011)	–	–0.005 (0.011)	–0.007 (0.012)	–0.012 (0.011)	–0.012 (0.011)
ΔY	–0.035** (0.015)	–0.040*** (0.014)	–0.033** (0.014)	–	–0.040*** (0.014)	–0.045*** (0.013)	–0.043*** (0.014)
REERM	0.002** (0.001)	0.002** (0.000)	0.002* (0.000)	0.002*** (0.000)	–	0.002* (0.001)	0.001 (0.000)
NFA/Y	–0.012* (0.006)	–0.017*** (0.006)	–0.013** (0.006)	–0.017*** (0.006)	–0.012* (0.006)	–	–0.007 (0.006)
PRCR/Y	0.008*** (0.002)	0.007*** (0.002)	0.008*** (0.002)	0.009*** (0.002)	0.007*** (0.002)	0.006*** (0.002)	–
C	–1.835	–1.985	–1.798	–1.911	–1.750	–1.849	–1.483
Observations	4161	4161	4161	4161	4161	4387	4242
ROC Statistics	0.659	0.634	0.655	0.654	0.664	0.640	0.684
Panel B: η=3							
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
FXR/STED	–0.001 (0.001)	–	–0.001* (0.000)	–0.001* (0.000)	–0.001 (0.001)	–0.002** (0.001)	–0.001 (0.001)
CAB/Y	–0.024 (0.015)	–0.031** (0.014)	–	–0.017 (0.015)	–0.023 (0.015)	–0.022 (0.014)	–0.028* (0.016)
ΔY	–0.037** (0.019)	–0.044** (0.017)	–0.031* (0.018)	–	–0.044** (0.008)	–0.044** (0.017)	–0.043** (0.019)
REERM	0.002* (0.001)	0.002* (0.001)	0.002 (0.001)	0.002** (0.001)	–	0.001 (0.001)	0.001 (0.001)
NFA/Y	–0.017** (0.008)	–0.020*** (0.008)	–0.017** (0.008)	–0.021*** (0.007)	–0.015* (0.008)	–	–0.010 (0.007)
PRCR/Y	0.008*** (0.002)	0.007** (0.003)	0.007*** (0.002)	0.008*** (0.002)	0.007** (0.003)	0.005 (0.002)	–
C	–2.316	–2.421	–2.209	–2.341	–2.247	–2.231	–2.021
Observations	4161	4161	4161	4161	4161	4387	4242
ROC Statistics	0.716	0.694	0.713	0.713	0.719	0.698	0.748

***: Significant at 1%, **: Significant at 5%, *: Significant at 10%.

Sources: Joint External Debt Hub (www.jedh.org), Author's calculations based on International Financial Statistics.

At this stage, following Minoiu and others (2013), for each estimation technique employed, we select that EWS specification whose Receiver Operating Characteristic (ROC) statistics is the largest. ROC analysis provides a quantitative measure of the accuracy of diagnostic tests to discriminate between two states or conditions (e.g. crisis and non-crisis).[11] For our analysis we use the ROC curve, which depicts the relationship between the fractions of positive cases correctly

[11] See STATA Press (2013).

DOI: 10.7172/2353-6845.jbfe.2016.2.1

classified (true positives) and that of positive cases incorrectly classified (false positives), for a range of probability thresholds (see figure 1). The fraction of positive cases that are correctly identified (the true-positive rate) is also called *sensitivity*, while *specificity* is the true-negative rate. The measure *1–specificity* is also called the false-positive rate. Therefore, for every cut-off value the ROC curve measures the trade-off between the true positive rate (sensitivity) and the false-positive rate (1–specificity).

The ROC curve is interpreted as follows. If the curve lies above the 45-degree line, for every cut-off value the true-positive rate is higher than the false-positive rate and the model generates crisis predictions that are superior to random guessing. By contrast, along the 45-degree line, sensitivity is equal to (1–specificity), meaning that for every cut-off value, the true-positive rate is exactly equal to the false-positive rate. The larger the ROC statistics, the most accurate is the diagnostic test to discriminate between crises and non-crises. For each estimation technique, the ROC curves of the logit and probit EWS specifications having the lowest and highest ROC statistics are plotted in figure 1. The ROC curves in panel A have been obtained when η=2, while those in panel B have been generated when η=3.

Figure 2
Areas Under Receiver Operating Characteristic (ROC) Curves

Panel A: η=2

Chart 1. Logit EWS

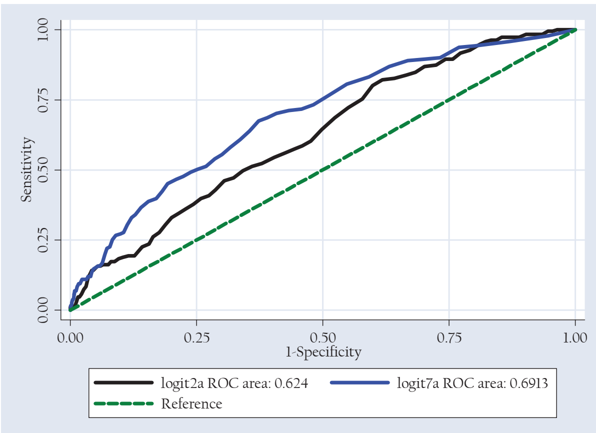
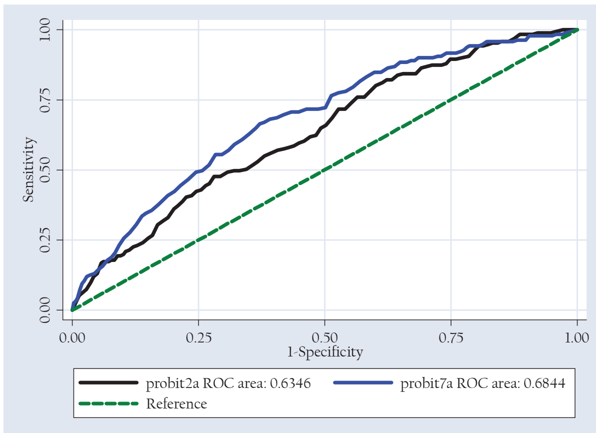


Chart 2. Probit EWS



Panel B: η=3

Chart 3. Logit EWS

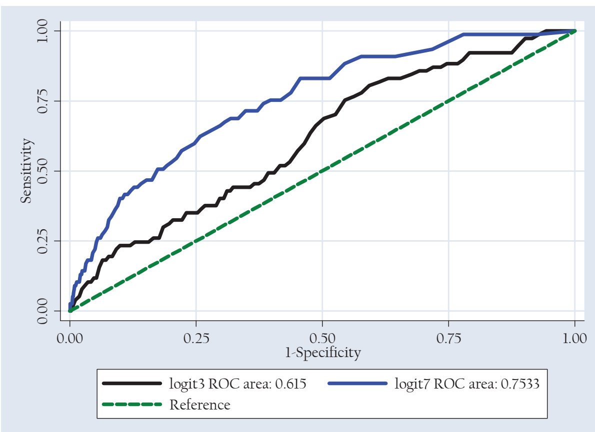
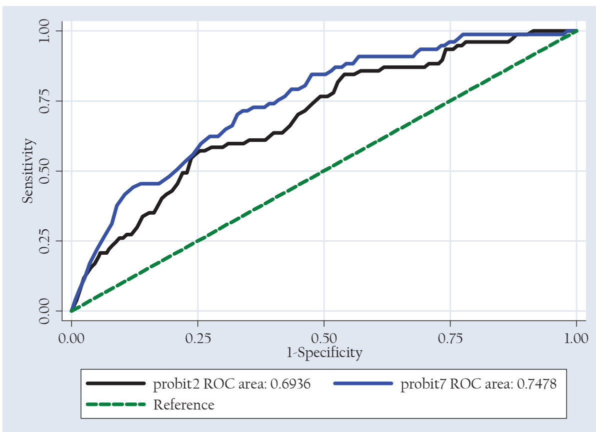


Chart 4. Probit EWS



Sources: Author's calculations based on International Financial Statistics.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

We select those logit and probit specifications having the largest ROC statistics. For each estimation technique and for each definition of the crisis variable used, specification (7) is the one with the largest ROC statistics. On this basis, we select the logit and probit EWS with specification (7) and assess their ability in correctly predicting in-sample and out-of-sample crisis and non-crisis episodes.[12]

5. THE PERFORMANCE OF THE SELECTED EARLY WARNING SYSTEM

We measure the performance of the two competing logit and probit EWS by looking primarily at their out-of-sample total misspecification error (TME). For each EWS, tables 3 and 4 report the following measures: the TME, the percentages of crisis episodes and tranquil periods correctly called, the percentages of missed crisis episodes and false alarms, the probability of observing a crisis within the next 24 months given a crisis alarm, the probability of observing a crisis if no prior alarm has been issued and the estimated probability cut-off values. As usual, in each table there are two panels. Panel A reports the EWS in-sample and out-of-sample performance results obtained when η=2, while panel B reports the results when η=3.

To assess the EWS in-sample performance, we proceed as follows. We restrict the original sample period (January 1995– December 2012) to the period January 1995–December 2006, which is the new estimation sample. For each of the two competing EWS, we obtain coefficient estimates, derive the probability of experiencing a currency crisis and calculate the TME. We begin by considering the estimation sample corresponding to the period January 1995–December 2006. Then, we gradually extend the estimation sample by one year at a time, for the following four years. We stop the procedure when the estimation sample corresponds to the period January 1995–December 2010. Then, we look at the one-year-ahead out-of-sample performance of each of the two competing EWS, as we are primarily interested to assess how the EWS perform in the first year outside the estimation sample.

The results in tables 3 and 4 show that the logit and probit EWS out-of-sample performances deteriorate compared to their in-sample performances, as reflected by the higher out-of-sample TME scores and the higher percentages of missed crisis episodes.[13] Both EWS tend to perform out-of-sample reasonably well until 2008 as indicated by the TME scores. From 2009 onwards, however, the TME scores of both EWS rise considerably, as well as the probability of experiencing a currency crisis given a crisis alarm. Summing up, including 2008 in the estimation sample – the year with most disruptive episodes of market turbulence observed during the global financial crisis – considerably deteriorates the EWS out-of-sample performance.

The data in table 3 suggest that when η=2 the logit EWS TME out-of-sample scores vary between 68 and 111, meaning that the logit EWS is able to classify correctly between 44% and 66% of the total out-of-sample observations. When η=3 – hence when less crisis episodes (77) are identified in the panel – the logit EWS out-of-sample performance deteriorates, as the TME out-of-sample scores vary between 83 and 116, meaning that the EWS correctly classifies only between 42% and 58% of the total out-of-sample observations.

[12] We estimated the EWS specification (7) with a logit random effects estimator. The Hausman Specification Test rejects the null hypothesis of random effects efficiency for the EWS specification (7). Therefore, we limit ourselves to compare the EWS specification (7) estimated with logit fixed effects and probit estimating techniques.

[13] See also charts in the Annex.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Table 3
Logit EWS: In-sample and Out-of-sample Performances

Sample Windows	Prob. cut-off value	Crisis Episodes Correctly Called (%)	Non-Crisis Episodes Correctly Called (%)	Missed Crisis Episodes (%)	False Alarms (%)	Crisis Prob. Given Alarm	Crisis Prob. Given No Alarm	TME
				Panel A: η=2				
				In-sample				
1995–06	0.06	83.3	53.5	16.7	46.5	0.26	0.06	63
1995–07	0.08	73.4	61.0	26.6	39.0	0.25	0.07	66
1995–08	0.07	63.4	61.4	36.6	38.6	0.29	0.13	75
1995–09	0.11	60.9	62.7	39.1	37.3	0.28	0.13	76
1995–10	0.09	60.8	59.5	33.2	40.5	0.27	0.11	74
				Out-of-sample				
2007	0.06	44.8	86.8	55.2	13.2	0.74	0.35	68
2008	0.08	52.1	78.1	47.9	21.9	0.63	0.30	70
2009	0.07	30.0	58.9	70.0	41.1	0.04	0.07	111
2010	0.11	51.6	77.4	48.4	22.6	0.19	0.06	71
2011	0.09	47.8	76.0	52.2	24.0	0.13	0.05	76
				Panel B: η=3				
				In-sample				
1995–06	0.06	66.5	65.9	33.5	34.1	0.28	0.09	68
1995–07	0.05	66.3	68.3	33.7	31.7	0.27	0.08	65
1995–08	0.08	46.7	73.2	53.3	26.8	0.30	0.15	80
1995–09	0.18	36.6	85.7	63.4	14.3	0.38	0.15	78
1995–10	0.24	35.1	86.9	64.9	13.1	0.37	0.14	78
				Out-of-sample				
2007	0.06	23.4	86.8	76.6	13.2	0.60	0.43	90
2008	0.05	35.0	81.6	65.0	18.4	0.58	0.36	83
2009	0.08	30.0	63.6	70.0	36.4	0.05	0.07	106
2010	0.18	0.0	83.6	100.0	16.4	0.00	0.11	116
2011	0.24	0.0	95.5	100.0	4.5	0.00	0.07	105

Sources: Joint External Debt Hub (www.jedh.org), International Financial Statistics and Author's calculations.

Table 4 reports the in-sample and out-of-sample performance results of the probit EWS. When η=2, the TME out-of-sample scores vary between 72 and 97, meaning that the EWS correctly classifies between 51% and 64% of the total out-of-sample observations. By contrast, when η=3, the probit EWS out-of-sample performance deteriorates, as the out-sample TME scores vary between 87 and 117. This implies that the EWS correctly classifies only between 41% and 53% of the total out-of-sample observations.

Overall, the logit and probit EWS out-of-sample performances are broadly similar. When η=2, no clear hierarchy emerges among the two EWS: the logit EWS out-of-sample TME scores vary between 68 and 111, while the probit out-of-sample TME scores vary between 72 and 97. When η=3, the logit EWS performs slightly better than the probit EWS, as the logit EWS out-of-sample TME scores are included between 83 and 116, while the probit EWS out-of-sample TME scores are included between 87 and 117. In addition, the results show that the EWS performance is sensitive to the crisis variable definition. Specifically, both EWS perform better (e.g. the TME scores are lower)

DOI: 10.7172/2353-6845.jbfe.2016.2.1

when the ERPI is defined by setting η=2 – when more currency crisis episodes are identified in the panel – then when setting η=3 – when less currency crisis episodes are identified in the panel.

Finally, the results show that crisis alarms issued out-of-sample are not always reliable, as the conditional probability of observing a crisis given an alarm declines considerably once that the estimation sample includes the year 2008. From a macroeconomic policy perspective, these results offer two implications. First, these results underscore the importance that similar early warning exercises having the goal to estimate the likelihood of experiencing financial crises should be run at least once every year. This is motivated by observing that the probability of experiencing a currency crisis may crucially depend on new incoming economic and financial data. Second, the results confirm that while running an EWS today can help identifying past crisis episodes with some accuracy (in-sample performance), predicting crisis episodes outside the estimation sample is much more challenging because of the presence of uncertainty, as the available information set outside the estimation sample is much more limited than within the estimation sample.

Table 4
Probit EWS: In-sample and Out-of-sample Performances

Sample Window	Prob. cut-off value	Crisis Episodes Correctly Called (%)	Non-Crisis Episodes Correctly Called (%)	Missed Crisis Episodes (%)	False Alarms (%)	Crisis Prob. Given Alarm	Crisis Prob. Given No Alarm	TME
				Panel A: η=2				
				In-sample				
1995–06	0.51	92.1	51.5	7.9	48.5	0.45	0.06	56
1995–07	0.48	86.9	53.2	13.1	46.8	0.43	0.09	60
1995–08	0.62	53.4	68.7	46.6	31.3	0.49	0.27	78
1995–09	0.52	67.8	54.9	32.2	45.1	0.45	0.24	77
1995–10	0.63	57.3	67.3	42.7	32.7	0.48	0.25	75
				Out-of-sample				
2007	0.51	31.5	81.8	68.5	18.2	0.83	0.70	87
2008	0.48	49.8	72.9	50.2	27.1	0.80	0.60	77
2009	0.62	47.6	64.3	52.4	35.7	0.31	0.21	88
2010	0.52	45.8	56.5	54.2	43.5	0.37	0.35	98
2011	0.63	39.6	88.5	60.4	11.5	0.60	0.23	72
				Panel B: η=3				
				In-sample				
1995–06	0.75	67.9	63.7	32.1	36.3	0.27	0.09	68
1995–07	0.63	69.3	65.7	30.7	34.3	0.27	0.08	65
1995–08	0.63	55.0	66.2	45.0	33.8	0.29	0.14	79
1995–09	0.74	48.8	74.1	51.2	25.9	0.31	0.14	77
1995–10	0.80	39.1	83.0	60.9	17.0	0.34	0.14	78
				Out-of-sample				
2007	0.75	9.7	96.7	90.3	3.3	0.71	0.44	94
2008	0.63	30.7	82.1	69.3	17.9	0.55	0.38	87
2009	0.63	30.0	58.2	70.0	41.8	0.04	0.07	112
2010	0.74	3.2	79.3	96.8	20.7	0.02	0.11	117
2011	0.80	0.0	95.8	100.0	4.2	0.00	0.07	104

Sources: Joint External Debt Hub (www.jedh.org), International Financial Statistics and Author's calculations.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

6. CONCLUSIONS

In this study we compared the performance of logit (fixed effects) and probit EWS in correctly predicting in-sample and out-of sample currency crisis in selected emerging market economies.

We found that stronger real GDP growth rates and higher net foreign assets significantly reduce the probability of experiencing a currency crisis, while high credit to the private sector increases it. By contrast, the current account balance and the measure of real exchange rate misalignment are not always statistically significant. The ratio between foreign exchange reserves and short–term external debt has the correct sign but it is significant only when the currency crisis is defined as a situation where the (country–specific) exchange rate pressure index is two standard deviations above its mean. The ratio is no longer significant if the exchange rate pressure index is three standard deviations above its mean.

The logit and probit EWS out-of-sample performances are broadly similar. The logit EWS is able to classify correctly between 42% and 66% of the total out-of-sample observations (e.g. crisis and tranquil periods), while the probit EWS s able to classify correctly between 41% and 64% of the total out-of-sample observations. We also find that the EWS performance is sensitive to the size of the estimation sample, and to the crisis definition used. In particular, both EWS perform better when a crisis episode is defined as a situation when the (country–specific) exchange market pressure index is two standard deviations above its mean. All in all, the results show that the model's ability to produce reliable out-of-forecasts currency crises prediction is limited. However, the model can be a useful framework to identify which vulnerabilities may significantly affect the probability of experiencing a currency crisis.

The results offer two macroeconomic policy conclusions. First, the EWS out-of-sample performance can be very sensitive to the size of the estimation sample. Specifically, the EWS total misclassification error and the probability of experiencing a currency given a crisis alarm can vary considerably if a particular year with many outlying observations is included in the estimation sample. Second, the results imply that selecting a crisis definition as a situation when the exchange rate pressure index is two standard deviations above its average value reduces the EWS total misclassification error. Therefore, the results obtained in this study suggest that a crisis definition identifying more rather than less currency crisis episodes should be employed when setting up an EWS model, even if this may lead to the risk of issuing several false alarms.

Finally, the analysis in this study can be extended in a number of ways. First, the two EWS employed in this study rely mainly on the information conveyed by standard macroeconomic and external vulnerability indicators. It would be interesting to assess if and how the EWS out-of-sample performance changes if indicators quantifying cross–country contagion, spillover effects or cross-border financial linkages were included in the EWS. Second, it would be interesting to modify the analysis in this study to estimate the probability of sudden stops in capital inflows in emerging market economies, and to check the model out–of–sample performance. Thirdly, as regards the definition of currency crisis adopted in this paper, η (the coefficient that multiplies the standard deviation of the exchange rate pressure index in the definition of currency crisis) was chosen arbitrarily. As an extension, it would be interesting to find out what is the value that η should assume in order to minimize the total misclassification error of the EWS.

References

Abiad, A. (2003), Early Warning Systems: A Survey and a Regime-Switching Approach. IMF Working Paper, No. WP/03/32, (Washington: International Monetary Fund).

Beckmann, D., Menkhoff, L., Sawischlewski, K. (2006), Robust Lessons About Practical Early Warning Systems. *Journal of Policy Modeling*, Vol. 28, pp. 163–193.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Berg, A, Pattillo, C. (1999), Predicting Currency Crises: The Indicators Approach and an Alternative. *Journal of International Money and Finance*, Vol. 18, pp. 561–586.

Berkmen, S.P., Gelos, G., Rennhack, R., Walsh, J. (2012), The global financial crisis: Explaining cross-country differences in the output impact. *Journal of International Money and Finance*, Vol. 31, pp. 42–59.

Blanchard, O., Das, M., Faruqee, H. (2010), The Initial Impact of the Crisis on Emerging Market Countries. *Brooking Papers on Economic Activity*, Spring 2010, pp. 263–323.

Borio, C., Drehmann, M. (2008), Assessing the Risk of Banking Crises – Revised. *BIS Quarterly Review*, March 2009, pp. 29–46 (Basel: Bank for International Settlements).

Bussiere, M., Fratszcher, M. (2006), Towards a New Early Warning System of Financial Crises. *Journal of International Money and Finance*, Vol. 25, pp. 953–973.

Candelon, B., Dumitrescu, E., Hurlin, C. (2012), How to Evaluate an Early-Warning System: Toward a Unified Statistical Framework for Assessing Financial Crises Forecasting Methods. *IMF Economic Review*, Vol. 60, No.1 (Washington: International Monetary Fund).

Cerra, V., Saxena, S.C. (2008), Growth Dynamics: The Myth of Economic Recovery. *American Economic Review*, Vol. 98, No.1 (March 2008), pp. 439–457.

Comelli, F. (2013), Comparing Parametric and Non-parametric Early Warning Systems For Currency Crises in Emerging Market Economies. IMF Working Paper, No. WP/13/134, (Washington: International Monetary Fund).

Eichengreen, B., Rose, A.K., Wyplosz, C. (1995), Exchange Market Mayhem. *Economic Policy*, Vol. 10, No. 21, pp. 249–312.

Frank, N., Hesse, H. (2009) Financial Spillovers to Emerging Markets during the Global Financial Crisis. IMF Working Paper, No. WP/09/104 (Washington: International Monetary Fund).

Frankel, J., Saravelos, G. (2012), Can Leading Indicators Assess Country Vulnerability? Evidence From the 2008-09 Global Financial Crisis. *Journal of International Economics*, Vol. 87, pp. 216–231.

Goldman Sachs (2013), 'Sudden–stops' in Capital Inflows and The Role of FX Reserves. *CEEMEA Economics Analyst*, Issue No: 13/25.

Gourinchas, P.-O., Obstfeld, M. (2011), Stories of The Twentieth Century For The Twenty-first. NBER Working Paper No. 17252, (Cambridge, Massachusetts: National Bureau of Economic Research).

International Monetary Fund (2002), Global Financial Stability Report. March 2002, Chapter 4 (Washington: International Monetary Fund).

Kaminsky, G. (1998), Currency and Banking Crises: The Early Warning of Distress. International Finance Discussion Paper No. 629, (Washington: Board of Governors of the Federal Reserve System).

Kaminsky, G., Lizondo, S., Reinhart, C. (1998), Leading Indicators for Currency Crisis. *IMF Staff Papers*, Palgrave Macmillan Journals, 45(1).

Kaminsky, G., Reinhart, C. (1999), The Twin Crises: The Causes of Banking and Balance-of-Payments Problems. *American Economic Review*, Vol. 89, No. 3, pp. 473–500.

Llaudes, R., Salman, F., Chivakul, M. (2010), The Impact of the Great Recession on Emerging Markets. *International Monetary Fund Working Paper*, No. WP/10/237 (Washington: International Monetary Fund).

Milesi-Ferretti, G., Tille, C. (2011), The great retrenchment: international capital flows during the global financial crisis. *Economic Policy*, Vol. 26, No. 66, pp. 289–346.

Minoiu, C., Kang, C., Subrahmanian, V.S., Berea, A. (2013), Does Financial Connectedness Predict Crises? *IMF Working Paper*, No. WP/13/267, (Washington: International Monetary Fund).

Rose, A., Spiegel, M. (2012), Cross-country Causes and Consequences of the 2008 Crisis: Early Warning. *Japan and the World Economy*, Vol. 24, pp. 1–16.

Stata Press (2013), Stata Base Reference Manual, College Station, TX: StataCorp LP.

Wooldridge, J.M. (2002), Econometric Analysis of Cross Section and Panel Data, Cambdrige, MA: MIT Press.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

ANNEX

A1. COUNTRY LIST

Argentina, Brazil, Bulgaria, Chile, China, Colombia, Croatia, Czech Republic, Egypt, Hungary, India, Indonesia, Kazakhstan, Korea, Malaysia, Mexico, Pakistan, Peru, Philippines, Poland, Romania, Russia, South Africa, Taiwan, Thailand, Turkey, Ukraine, Uruguay and Vietnam.

A2. DESCRIPTION OF THE VARIABLES

(1) Ratio between foreign exchange reserves and short term external debt, FXR/STED: This is calculated as the ratio between the stocks of foreign exchange reserves and short-term external debt (i.e. maturing within one year). Both numerator and denominator are expressed in U.S. dollars. It is a reserve adequacy ratio which is often used in early warning exercises. Quarterly data have been interpolated in order to have monthly time series. Source: Joint External Debt Hub (www.jedh.org).

(2) Current account balance as a percentage of GDP, CAB/Y. Ratio between the current account balance and nominal GDP. Both numerator and denominator are expressed in U.S. dollars. Annual data have been interpolated in order to have monthly time series. Source: World Economic Outlook Database, International Monetary Fund.

(3) Real GDP growth, ΔY: Annual percentage change in real GDP. Annual data have been interpolated in order to have monthly time series for real GDP growth. Source: World Economic Outlook Database, International Monetary Fund.

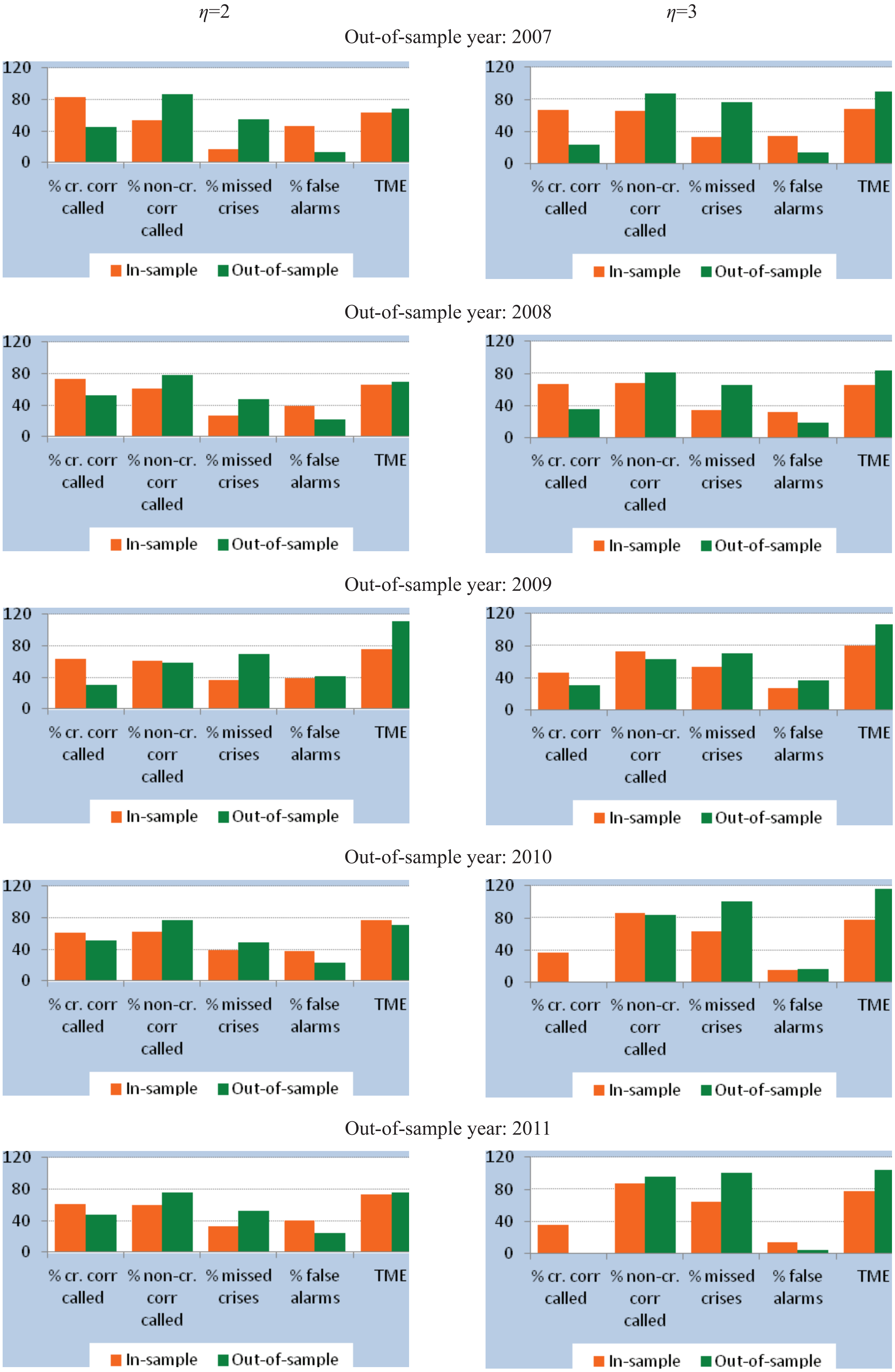
(4) Real effective exchange rate misalignment, REERM. The series has been obtained by taking the real effective exchange rate (REER) deviation from the three-year moving average. Monthly data. Source: International Financial Statistics, International Monetary Fund.

(5) Ratio between the stock of net foreign assets and nominal GDP, NFA/Y. Both numerator and denominator are expressed in U.S. dollars. Source: International Financial Statistics, International Monetary Fund.

(6) Ratio between private credit and nominal GDP, PRCR/Y. Available for most emerging economies only from January 2001 onwards. Source: International Financial Statistics, International Monetary Fund.

DOI: 10.7172/2353-6845.jbfe.2016.2.1

Figure 3
Logit In-sample and Out-of-sample Performances *(Percentages)*

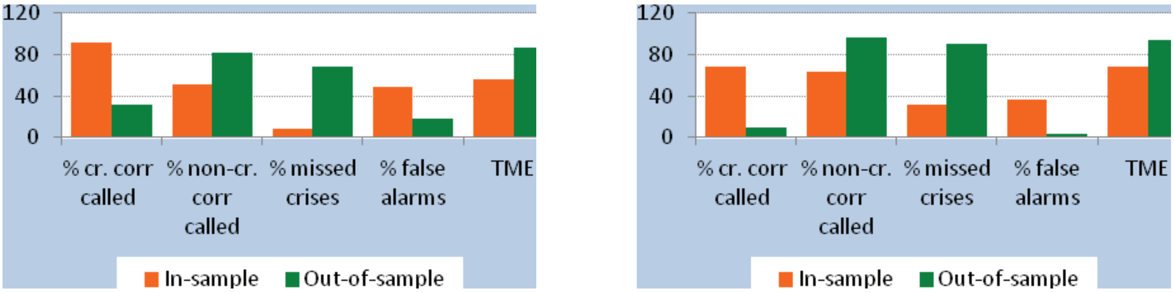


DOI: 10.7172/2353-6845.jbfe.2016.2.1

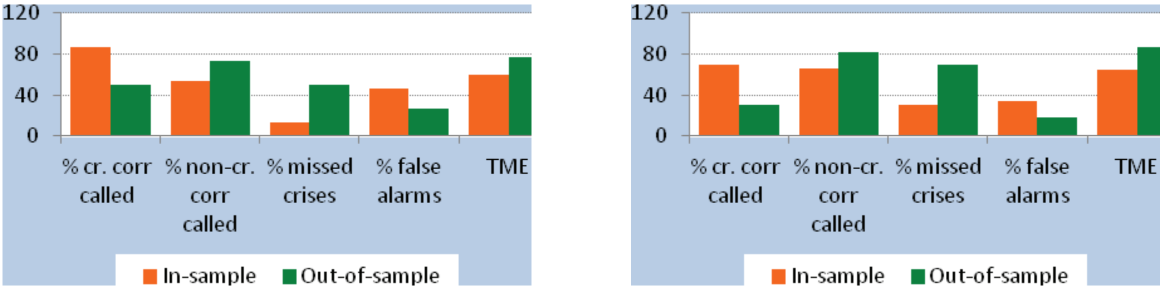
Figure 4
Probit In-sample and Out-of-sample Performances *(Percentages)*

$\eta=2$ $\eta=3$

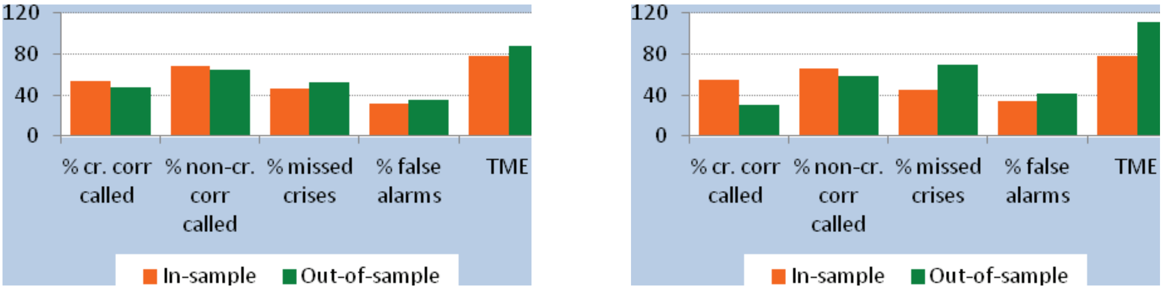
Out-of-sample year: 2007



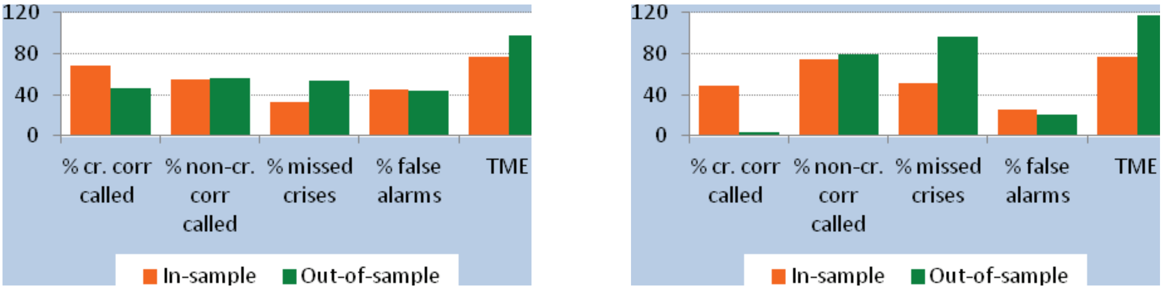
Out-of-sample year: 2008



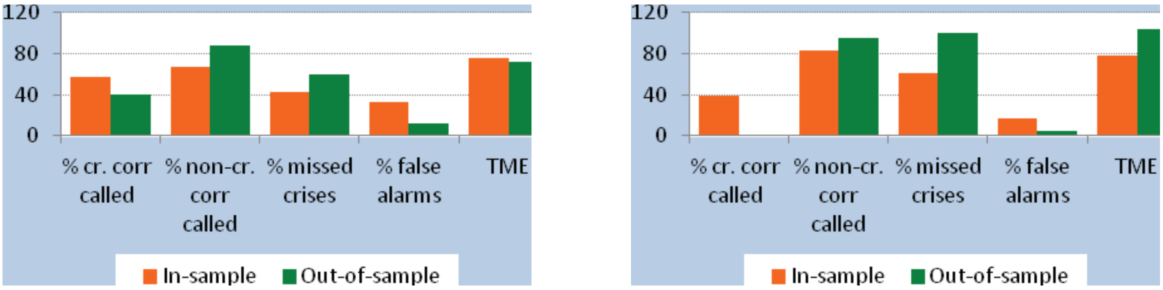
Out-of-sample year: 2009



Out-of-sample year: 2010



Out-of-sample year: 2011



DOI: 10.7172/2353-6845.jbfe.2016.2.1

Investigating Impact of US, Europe, Frontier and BRIC Stock Markets on Indian Financial Stress Index

Amanjot Singh[1]
University School of Applied Management, Punjabi University, Patiala, India 147002
Email: amanjot_42@yahoo.com

Manjit Singh
University School of Applied Management, Punjabi University, Patiala, India 147002
Email: smanjitt@gmail.com

Received: 25 December 2015 / Revised: 3 April 2016 / Accepted: 20 April 2016 / Published online: 16 June 2016

ABSTRACT

The emerging markets are slowly opening up their respective financial markets to foreign investments, thereby making the latter markets more sensitive to cross-market information transmissions. There are different transmission mechanisms ranging from trade related to financial linkages. However, statistically, both price discovery and conditional volatility act as transmission mechanisms, whereby information in one stock market has an impact on another. In this regard, the present study attempts to empirically analyse the impact of global information transmissions, i.e., stock market returns and conditional volatility on overall Indian financial stress and its various sub-components by employing different econometric models comprising Johanson Cointegration, Vector Autoregression and its various counterparts, Component GARCH (1,1) model and multivariate OLS regression models ranging from October 2003 to October 2014. The study firstly constructed Indian financial stress index owing to non-existence of a standardised index. The results reported that the one month lagged returns in the BRIC stock markets have an impact on the financial stress index of India. The stress in the Indian financial system responds statistically significantly to the Brazilian and Chinese market returns, with a greater degree of integration after two months. A statistically significant impact of the short-run volatility has also been observed running from the European markets to the Indian financial system contemporaneously. Furthermore, unexpected volatility in the BRIC markets also has an impact on the Indian financial stress contemporaneously as well as dynamically. The present study provides an insight to the international investors regarding the response of Indian financial system and its sub-components toward global information transmissions.

JEL classification: F36; G10; G15

Keywords: BRIC, financial stress, frontier markets, transmission

[1] Corresponding author; Postal address #303-C, Rajguru Nagar, Ludhiana, Punjab, India 141012.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

1. INTRODUCTION

The markets all over the world and especially the emerging ones are slowly opening up their equity markets to the foreign direct as well as portfolio investors. The increased globalisation and the development of the trading platforms have made the countries prone to an international crisis and the country specific news and events gets transferred from one country to another impacting the embedded trading nations (Angkinand et al., 2010). A simple example to comprehend the integration of the markets can be that of the United States (US). The developments in the US and the likely decision of the Monetary Authorities to end the Quantitative Easing cycle started as a result of the 2008 crisis make the Indian equity markets or in a broader sense the emerging markets witness a downward rally as the quest to transfer the 'Hot Money' from the emerging nations to the safer ones increases. Frank and Hesse (2009) reported spillover of the crisis from the developed markets to the emerging markets highlighting the safety concerns of the international investors in their act of transferring money from the emerging nations.

Price discovery can be denoted as the speed at which an asset's price reacts to new information (Booth et al., 1999). Apart from price discovery, conditional volatility also acts as another information transmission mechanism, wherein information in one stock market has an impact on another country's stock market (Gagnon & Karolyi, 2006; Rittler, 2012). So, due to increased integration and international portfolio allocations, the stock market returns and volatility in one market get transferred to the stock markets of other countries. Numerous studies have captured the contagion impact of the country specific events on the other integrated economies (see Worthington & Higgs, 2004; Mukherjee & Mishra, 2010 and Kharchenko & Tzvetkov, 2013; etc.). Now a question arises whether these spillovers have an impact on the financial stress of the recipient country owing to international flow of funds. The present study attempts to answer the question posited by employing Vector Autoregression model (VAR model) and Component Generalised Autoregressive Conditional Heteroskedastic model [CGARCH (1,1) model]. The impact of the first moment as well as the second moment of the stock market has been covered in the study; the impact of the global stock market returns and volatility on the Indian financial stress respectively. We have considered the stock markets of the US, Europe, frontier markets and the BRIC (Brazil, Russia, India and China) markets, thereby taking into account the case of two developed economies (the US and Europe) and two developing economies (BRIC countries and frontier countries). The frontier markets like Kuwait, Tunisia, Pakistan, etc., which are economically lesser developed even compared to the emerging markets, are also included in the study as an endogenous variable making the study first of its kind. A priori one would expect a higher degree impact of the US and BRIC equity markets on the Indian financial system and its sub-components because of the increasing integration in the sense of real as well as financial linkages among the countries concerned and the US being the dominant economy worldwide. Moreover, the steps taken in the direction of incorporation of the BRICS bank and increasing flow of foreign funds act as a base for considering the higher impact of the BRIC markets on the Indian financial system. Besides this, we expect that falling returns in the respective equity markets will have an increasing impact on the financial stress in the Indian economy.

A financial stress in a general sense implies commotion in the asset prices and the failure of financial institutions (Manamperi, 2015). There is no specific definition available for the financial stress, but in a layman terms, it is a stress or uncertainty in the financial sector of an economy which further has an impact on the macroeconomic conditions. A stress in the financial system which not only comprises equity market but also debt market, money market, commodity market as well as currency market has an impact on the fundamental health of an economy. Any disruption in a financial system makes the economies feel the heat of the lower output, higher bank rates, increased unemployment, lower GDP growth, higher inflation, etc. Hakkio and Keeton (2009) explained the various features of the financial stress, ranging from uncertainty about the fundamental value of

DOI: 10.7172/2353-6845.jbfe.2016.2.2

assets, uncertainty about the behaviour of investors, increased asymmetry of information, flight to quality investment avenues to flight to liquidity. The review highlights the fact that the global macroeconomic conditions or a country specific crisis does have an impact on the financial stress of a country like what happened during the subprime episode in the US, as studied by Bianconi et al. (2013). During the period of a financial stress, this increased integration across different markets proves to be a bane as the financial stress gets spillovered from one market to another through the channels of trade as well as the financial markets. Therefore, a study to account for the impact of the global market returns and volatility on the Indian financial stress has been undertaken.

Over a period of time, many researchers have tried to capture the financial stress in an economy. Particularly the work relates to the developed markets as compared to the emerging markets. A reason that could be attributed to this can be the development level of financial markets in the emerging nations and the availability of the data therein. Kliesen et al. (2012) explored various financial stress indices across different countries and found out the co-movement among them. The authors gave a very comprehensive review of the stress indices, like STLFSI (St. Louis Fed Financial Stress Index), KCFSI (Kansas City Financial Stress Index) and CFSI (Cleveland Financial Stress Index) are some of the indices to account for the US financial stress. Apart from the US, the indices have been designed keeping in view the financial aspects of other countries as well, like Canada (Illing & Liu, 2006), Sweden (Sandhal et al., 2011), Colombia (Morales & Estrada, 2010), Hong Kong (Yiu, Ho, & Jin, 2010), etc. But there are some studies which have also concentrated on the emerging markets and have developed the financial stress indices. Balakrishnan et al. (2011) and Park and Mercado (2013) investigated the determinants of the financial stress in the emerging markets as well as the transmission of the global financial crisis. The studies relating to the financial stress are not only limited to the designing of the indices yet an effort has been made by every author to either explore the impact of a global or inter-regional or intra-regional shock on the financial stress index. Furthermore, the impact of the financial stress on the economic activities has also been studied by the scholars (Davig & Hakkio, 2010). Sum (2013) examined the impulse response functions of the Federal Reserve Bank of St. Louis financial stress index and excess returns on the CRSP (Centre for Research in Security Prices) value-weighted index. The author observed that the financial stress Granger-causes market risk premiums to drop significantly. Moreover, there is no reverse causation. The studies, like Goldstein and Xie (2009), Roye (2014) and Wallace (2013), have captured the impact of the financial stress on economic activities across different countries with the findings that the financial stress does have an impact on different economic activities of an economy like GDP, inflation numbers, etc. There are some studies which have gone one step ahead and tried to account for the impact of the financial stress on the stock markets as well (see for detail Christopoulos et al., 2011 and Rachdi, 2013). Almost all of the studies have analysed the impact of the financial stress on some type of economic activity, but till now not much work has been done to see what factors have an impact on the financial stress, particularly the impact of global stock markets on the domestic financial stress. The present study attempts to fill this research gap.

A study relating to the impact of the global stock markets on the domestic overall financial stress is an imperative task to be performed by the policy makers as well as the investors. The results reported by the models employed signify the existence of an impact from the BRIC and the European nations to the Indian financial system, making a case for the financial market investors and the policy makers to discount this type of information well in advance. Moreover, the results provide an insight to the international investors regarding the response of Indian financial system and its sub-components toward global information transmissions.

The paper has been divided into five sections. Section 2 explains the construction of the financial stress index in the Indian economy context. Section 3 highlights the methodology used to account for the impact of the markets on the financial stress. Sections 4 and 5 introduce the readers to the empirical findings and the concluding remarks thereof.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

2. CONSTRUCTION OF FINANCIAL STRESS INDEX

A financial system comprises various segments like the equity markets, debt markets, foreign exchange markets and money markets. In order to construct the financial stress index, we have taken all of these four segments into account. The index has been constructed following the work of Balakrishnan et al. (2011), Yiu, Ho, and Jin (2010) and Park and Mercado (2013). The latter studies have considered these four segments to account for the possible macroeconomic channels having an impact on the financial stress of an economy. We have collected monthly data for a period of 11 years with effect from October 2003 up till October 2014 from the Bloomberg and Yahoo Finance Database as per its availability. Our reason for taking monthly values instead of daily values is that monthly values of the indicators would reduce the sensitivity and enhance the reliability of the data in comparison to the daily data.

To capture the impact of global stock markets on the Indian financial stress index, we have taken Morgan Stanley Capital International (MSCI) indices: MSCI US, MSCI Europe, MSCI frontier markets and MSCI BRIC markets. Likewise, we have taken monthly closing values from the website of MSCI ranging from October 2003 to October 2014. The period assumed for the study has been decided keeping in view the availability of the data. All of the MSCI indices are designed by taking large cap and mid cap scripts from the respective nations. For instance, MSCI BRIC index comprises good quality stocks from Brazilian, Russian, Indian and Chinese markets. The monthly continuously compounded returns are calculated for the MSCI indices.

2.1. FSI Variables

To capture the stress in the equity market, we have taken the NIFTY index returns and volatility. The NIFTY index monthly continuously compounded returns have been calculated as:

$$R_t = Ln(P_t/P_{t-1}) * 100 \tag{1}$$

where R_t is the monthly return, P_t is the current month close price and P_{t-1} is the previous month close price. We have taken the NIFTY index returns *as it is* without really considering the impact of only negative returns on the financial stress index because it has been observed that increased conditional volatility is also concerned with the negative market returns (Christie, 1982). So, the study has considered overall stock market returns. To model the volatility in the NIFTY index, we have employed GARCH (1,1) model. The standardised values[2] are taken for the equity market return and volatility series. To calculate the standardised values, the mean value is firstly deducted from the series and then divided by the series' standard deviation.

The stress in the banking sector has been captured by the spread between MIBOR 3 monthly rate and Treasury Bill 3 monthly yield. The spread exhibits stress in the liquidity and the interbank lending risks. An increased spread indicates increased interbank lending rates, thereby depicting the liquidity funding risk in an economy. Again, we have taken the standardised values. The banking sector BETA is generally taken as a parameter for measuring the banking stress; however, due to non-availability of the data during the period concerned, we have taken only MIBOR 3-month rate – Treasury bill 3-month yield spread as the stress measurer.

$$\text{FSI} = \textit{Equity Market Return} + \textit{Equity Market Volatility} + \textit{Debt Market (Spread)} \\ + \textit{Baking Sector (Spread)} + \textit{Exchange Rate Volatility} \tag{2}$$

[2] $Y_t = \dfrac{(X_t - \textit{Mean Value})}{\textit{Standard Deviation}}$

DOI: 10.7172/2353-6845.jbfe.2016.2.2

For an emerging market like India, the major worry with regard to the foreign exchange market is the volatility in the exchange rate due to very high twin deficits (fiscal deficit and current account deficit) and interest rates. The GARCH (1,1) model has been used to account for the volatility in the Dollar/Rupee exchange rate. Again, we have taken the standardised values. To capture the stress in the sovereign debt, we have taken standardised spread between the India Government securities 10-year yield and the US Government securities 10-year yield. An increased value of the spread means an increased sovereign debt servicing risk. These five components of the financial markets represent overall financial stress in an economy. A reason for taking only these four segments of the financial system into account is that these segments are the core segments to measure a financial stress in an economy. Other indicators which can also be taken as a part of measuring the overall financial condition can be the GDP growth rate, unemployment levels, inflation index, etc. However, in the present study we are relying only on the core financial stress indicators for measuring overall financial stress instead of financial condition. Another important part of the construction of the index is the weighting scheme of different components.

2.2. Weighting Scheme of the Index

The literature describes various weighting schemes used by the researchers over a period of time. For instance, unweighted index, equal weighted index, market weighted index, principal component analysis, etc., are some of the methods that have been used to construct the financial stress index. In the present study, we have employed unweighted index, equal weighted index and principal component analysis so as to capture the financial stress in the Indian economy. Under the unweighted index approach, an aggregative value of respective standardised variables is taken up as an index value at time *t* and so on. The aggregate values are further re-based in the range of 0 to 100, by employing the following formula, as inspired from Lall et al. (2008):

$$New\ Scale\ Value = \frac{Old\ Scale\ Value - Lowest\ Value}{Highest\ Value - Lowest\ Value} * 100 \tag{3}$$

Another method of creating an index is principal component analysis, which identifies the best possible combination of the variables that explains the total variance in the five variables. Firstly, the standardised values are factored into the model and then the normalised component loadings/coefficients (component loadings divided by the square root of respective communalities) are calculated. The coefficients identify the impact of one standard deviation change in the variable on the financial stress index. In the present study, the second component explained around 66 percent of the total variation in the five variables. So, the normalised component loadings of the second component pave the way for the computation of financial stress index values by taking an aggregative value of respective variables over a period of time. In the last step, the index values are re-based in the range of 0 to 100.

Lastly, to calculate equal weighted financial stress index, we have computed average values of the sub-components of the financial stress. Before taking an average, respective standardised series of each sub-component are re-based within the range of 0 to 100. So, by taking the average values of different components of the financial stress, we allocated equal weights to each component in the financial stress index.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

3. RESEARCH METHODOLOGY

In order to model the impact of the returns on the financial stress, we have employed VAR model (Vector Autoregression model) and to model the impact of the global stock market volatility on the Indian financial stress index, CGARCH (1,1) model has been used. Before applying the VAR model, an effort has also been made to check a long-run stochastic trend among the variables concerned by using Johansen Cointegration Approach.

3.1 Vector Autoregression Model

The efficiency of the VAR model in capturing dynamic relationships among the underlying variables is well documented. Under the VAR model, a dependent variable is a function of its own lagged values as well as the lagged values of some other variable. The model has been popularised by Sims (1980). Due to its dynamic nature, the VAR model is observed to be an optimal candidate accounting for first moment linkages among the markets. Say there are two variables Y_1 and Y_2, the VAR model equation shall be defined as follows:

$$Y_{1,t} = c_1 + A_{1,1}Y_{1,t-1} + A_{1,2}Y_{2,t-1} + e_{1,t} \tag{4}$$

$$Y_{2,t} = c_2 + A_{2,1}Y_{1,t-1} + A_{2,2}Y_{2,t-1} + e_{2,t} \tag{5}$$

where c_1 and c_2 is a $k \times 1$ vector of constants, A_i is a $k \times k$ matrix (for every $i = 0, ..., p$) and e_t is a $k \times 1$ vector of error terms known as impulses or shocks. The lagged values of the dependent as well as the independent variables help in analysing the dynamic impact of global stock market returns on the Indian financial stress index. To analyse the results of the VAR model, we have further employed Granger causality test, impulse response functions and variance decomposition analysis.

3.2 Component GARCH (1,1) Model

To analyse the impact of the second moment, i.e. market volatility, on the financial stress in India, we have employed CGARCH (1,1) model. The model demarcates conditional variance into two components: Transitory/Short-run component and the Permanent/Long-run component. The CGARCH model was introduced by Ding et al. (1993) as an advancement to the plain vanilla GARCH (1,1) model introduced by Bollerslev (1986). Volatility is not directly observable in the market; however, it can be gathered from the past behaviour of the respective market prices. Subsequently, the GARCH based models are found to be effective in capturing the conditional variances or time-varying volatility. Under the plain vanilla GARCH model, a dependent variable is a function of its own past squared error terms and the past volatility. For the sake of simplicity, the GARCH models exhibit the impact of recent news/shock as well as the past volatility on the current conditional variance. The CGARCH model is an extension of the plain vanilla GARCH model wherein a conditional variance comprises two components: permanent and transitory. The conditional mean equation [eq (6)] and the variance equation under the CGARCH model framework are defined as follows:

$$R_t = c + \varepsilon_t \tag{6}$$

where R_t is monthly continuously compounding return of the respective nations. The monthly returns are a function of only the constant term and ε_t is the residual part. It may be noted that overall GARCH diagnostic tests support the inclusion of only the constant term in the mean

DOI: 10.7172/2353-6845.jbfe.2016.2.2

equation with respect to all of the markets, except for the frontier equity markets. So, one and two months' lagged values are included in the mean equation of the latter markets, evidenced from significant autocorrelation coefficients. The residuals are further checked for the existence of the ARCH effects because there should be a volatility clustering phenomenon in the residuals derived from the mean equation so as to employ a GARCH model.

Variance Equation

$$q_t = \gamma_0 + \gamma_1(q_{t-1} - \gamma_0) + \gamma_2(e^2_{t-1} - h_{t-1}) \tag{7}$$

$$h_t = q_t + \gamma_3(e^2_{t-1} - q_{t-1}) + \gamma_4(h_{t-1} - q_{t-1}) + \varepsilon_t \tag{8}$$

where q_t is the long-run component, $(e^2_{t-1} - h_{t-1})$ highlights the time-varying movement of permanent/long-run component and γ_1 represents persistency in the long-run component. $(h_{t-1} - q_{t-1})$ is transitory/short-run component of the conditional variance. The model has been estimated by assuming a normal distribution of the error terms. We have used the short-run and long-run component of conditional volatility to account for the impact of the volatility on the Indian financial stress. The short-run component of the conditional volatility is computed by deducting long-run component from the total conditional variance. Equations (7) and (8) bring to light the computational aspect of long- and short-run component of the conditional volatility in the context of the respective countries and regions considered. After the computation of respective conditional volatilities, we have tried to analyse the impact of those on the Indian financial stress index by using multivariate Ordinary Least Squares (OLS) regression model.

Equation (9) explains the impact of monthly transitory component on the monthly financial stress. On a similar note, we have captured the impact of monthly long-run component. The equation highlights the 'Contemporaneous Impact', i.e. same month impact of the volatility on the Indian financial stress index.

$$\Delta FSI = \text{Ø}_1 + \text{Ø}_2 Tran.Vol_1 + \text{Ø}_3 Tran.Vol_2 + \text{Ø}_4 Tran.Vol_3 + \text{Ø}_5 Tran.Vol_4 + \varepsilon_t \tag{9}$$

where Ø_1 is the constant term and Ø_2, Ø_3, Ø_4 and Ø_5 are the coefficients of the short-run components of the conditional variance relating to the BRIC, US, Europe and frontier nations respectively. However, one-month lagged impact of the conditional long-run as well as short-run volatility has also been estimated to account for the 'Dynamic Impact'. Lastly, to capture the impact of the unexpected market volatility on the financial stress index, the standardised residuals $(e_{k,t} = \frac{\varepsilon_{k,t}}{\sqrt{\sigma_{k,t}}})$ are derived from the respective variance equations (8).

$$\Delta FSI = \infty_1 + \infty_2 Uexp.Vol_1 + \infty_3 Uexp.Vol_2 + \infty_4 Uexp.Vol_3 + \infty_5 Uexp.Vol_4 + \varepsilon_t \tag{10}$$

where ∞_1 is the constant term and ∞_2, ∞_3, ∞_4 and ∞_5 are the coefficients of the unexpected volatility relating to the BRIC, US, Europe and frontier nations respectively capturing the monthly contemporaneous impact. On a similar note, dynamic impact of one-month lagged unexpected volatility on the financial stress has been captured. For instance, equation (11) depicts one-month lagged impact of the unexpected component on the monthly financial stress:

$$\begin{aligned}\Delta FSI = \infty_1 + \infty_2 Uexp.Vol_1(-1) + \infty_3 Uexp.Vol_2(-1) + \infty_4 Uexp.Vol_3(-1) + \\ + \infty_5 Uexp.Vol_4(-1) + \varepsilon_t\end{aligned} \tag{11}$$

DOI: 10.7172/2353-6845.jbfe.2016.2.2

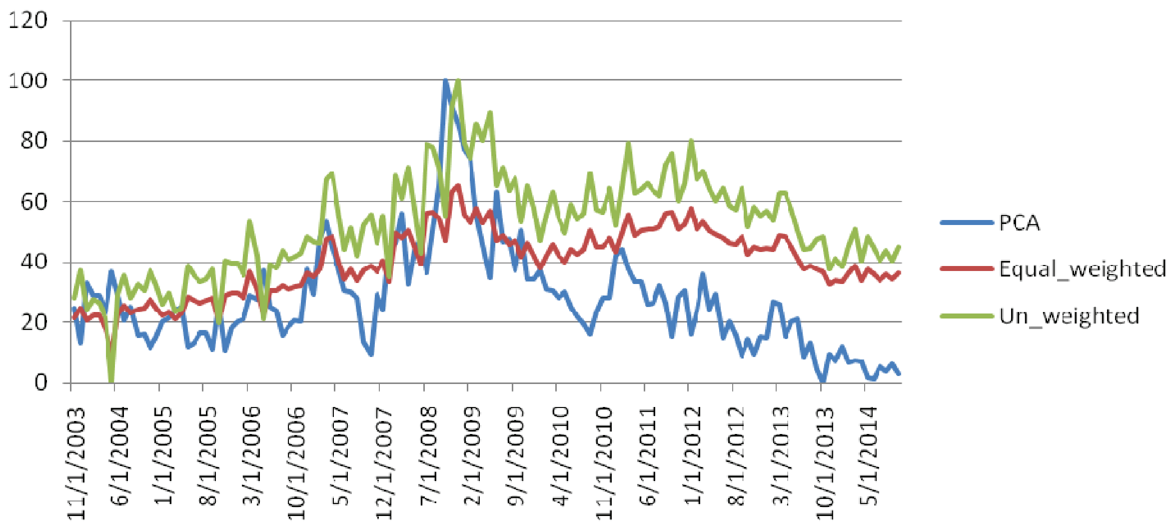
where ∞_1 is the constant term and ∞_2, ∞_3, ∞_4 and ∞_5 are the coefficients of the unexpected volatility capturing the dynamic impact relating to the BRIC, US, Europe and frontier nations respectively. The whole analysis has been done by using MS Excel and EVIEWS software.

4. EMPIRICAL FINDINGS AND DISCUSSION

Exhibit 1 reports graphical image of the Indian financial stress index by employing three different techniques. The index has captured stress in four different sub-sectors of the Indian financial system. The financial stress is very high in the last quarter of year 2008, thereby highlighting the presence of the US subprime crisis. It clearly shows spillover of the crisis from the US financial system to the Indian financial system.

As discussed earlier, we have used three different techniques to construct the financial stress index. The financial stress index measured through principal component analysis[3] is quite volatile in comparison to the other techniques during the period 2003 to 2014. Interestingly, the time-varying movements of unweighted as well as equally weighted series are somewhat similar. At the same time, all of the financial stress indices have captured the existence of a high level of stress in the Indian financial system during the US financial crisis period, thereby confirming the adequacy of the Indian financial stress indices.

Exhibit 1
Indian Financial Stress Index



Source: Computed by the Authors

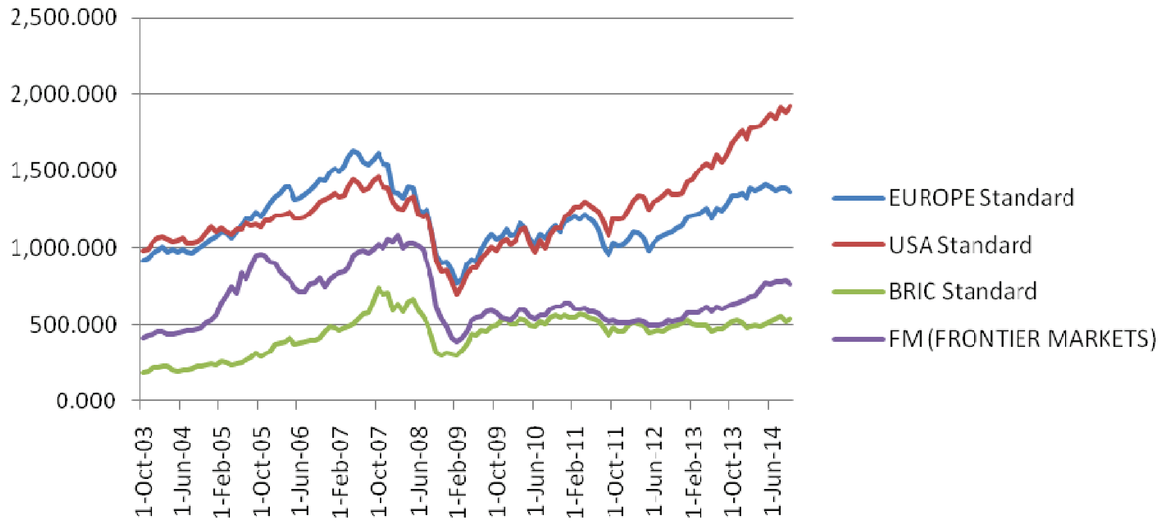
Considering the above facts, we have adopted a conservative approach to measuring the financial stress through equally weighted approach; relatively lesser as well as stable. So, the rest of the discussion and analysis has been done by taking equally weighted financial stress index. On a monthly average basis, the financial stress for years 2003 to 2014 is about 39.69 coupled with a very high standard deviation of 10.95. The financial stress series is normally distributed at level because we failed to reject the null hypothesis of normal distribution of Jarque-Bera test (2.768, $p>0.05$). Exhibit 2 is the graphical presentation of respective MSCI indices for years 2003 to 2014. A downward rally can easily be seen during the period 2008–2009, thereby signifying

[3] The respective standardised weights are 0.97, 0.23, –0.29, 0.94 and –0.65 for bank spread, exchange rate volatility, debt spread, equity market volatility and equity market returns. A negative weight highlights an increasing impact of the respective variable during negative phases, for instance, negative equity market returns enhance financial stress.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

the presence of the US subprime crisis during that period worldwide. The crisis that started in the US spillovered to other countries owing to strongly integrated financial system and increasing international trade relations (Dooley & Hutchison, 2009).

Exhibit 2
Graphical Presentation of MSCI Indices



Source: Computed by the Authors

Exhibit 3 reports descriptive statistics of MSCI index returns with respect to all of the countries. The average monthly returns are observed to be highest for the BRIC countries during the sample period. The highest average returns for the BRIC countries signify positive behaviour of the investors toward opportunities available in the emerging markets. The emerging markets like the BRIC nations with increasing middle class population, technological up-gradation and infrastructure development provide immense opportunities to the domestic as well as international investors to reap out the investment benefits. The average returns in the frontier markets are also higher coupled with a higher level of standard deviation compared to Europe. The average values signify the fact that the emerging and frontier markets act as an investment opportunity for the international investors and more to those who are ready to digest an increased risk level. Notwithstanding, the BRIC countries witness higher average monthly returns yet the volatility is the highest comparing to other countries selected for the study justifying the adage: the higher the risk, the higher the returns. The skewness values are negative in nature with respect to each nation making a case that probability of a negative return is higher comparing to positive return. The probability values of the Jarque-Bera test indicate non-normal distribution of the respective return series and further higher kurtosis values (greater than three) imply clustering nature of the returns.

The financial time series data is required to be stationary as non-stationary series would entail spurious regression results (Gujarati et al., 2013). We have used Augmented Dickey Fuller test to check stationarity of the data. The Augmented Dickey Fuller test is a function of lagged values of the dependent variable. The alternate hypothesis signifies the stationarity of the data. The MSCI data is found to be non-stationary at level but stationary after taking the first difference at the 5 percent significance level. We have also checked the stationarity of the financial stress index. The index is also found to be non-stationary at the level but stationary after taking the first difference at the 5 percent significance level.

After checking the stationarity of the data, we performed Johansen Cointegration test to check the presence of any stochastic trend or a long-run co-movement between the financial stress index and MSCI indices. One of the main conditions for the Johansen Cointegration test is that the data should be integrated of the same order I (1). The test comprises two alternate test statistics:

DOI: 10.7172/2353-6845.jbfe.2016.2.2

Trace test and the Maximum Eigenvalue test. Under the Trace test, the alternative hypothesis of cointegration is that the cointegrating vectors are greater than 0 (h_1: r > 0), whereas the alternative hypothesis tests the number of cointegrating vectors as r+1 in the case of the Max Eigenvalue test. Under the Johansen methodology, the lag lengths should be appropriate and to fulfil our requirements, we employed VAR model first and determined the maximum number of lags on the basis of Akaike's Information Criteria (AIC) values. The results of the Johansen test are reported in Exhibit 4. The trace test indicates the presence of one cointegrating vector, but the Max Eigenvalue test indicates no cointegrating vectors.

Exhibit 3
Descriptive Statistics (MSCI Index Returns)

	Return_BRIC	Return_Europe	Return_Frontier	Return_USA
Mean	0.789966	0.296069	0.461415	0.506650
Median	1.558207	1.249729	1.292051	1.142933
Maximum	15.71472	11.31091	16.31162	10.28517
Minimum	–30.52406	–14.58630	–26.86788	–18.93110
Std. Dev.	6.591723	4.144947	5.619984	4.180697
Skewness	–1.006184	–0.974301	–0.947880	–1.140631
Kurtosis	6.380101	4.858294	7.360727	6.219436
Jarque-Bera	85.11089	39.87670	124.3542	85.62908
Probability	0.0000	0.0000	0.0000	0.0000

Source: Computed by the Authors.

Exhibit 4
Johansen Cointegration Results

Null Hypothesis	Alternate Hypothesis		95% Critical Value
Trace test		Trace value	
r = 0	r > 0	75.73	69.82
r <= 1	r >1	44.43	47.86*
Max test		Max value	
r = 0	r = 1	31.29	33.87

* Trace test indicates 1 cointegrating equation at the 5 percent significance level.

Source: Computed by the Authors.

As both results are conflicting, we will go by the results reported by the Max Eigenvalue test as the results of the Trace test are indicative in nature (Pentecost & Moore, 2006). On the other hand, it may be noted that the Johansen approach supports the results reported by the Trace test. So, there is no long-run co-movement or a stochastic trend among the financial stress index and the other MSCI indices taken at logged level. Though there is no long-run co-movement yet there can be short-run relationships among the underlying variables, which we have checked by employing a VAR model. To understand the results of the VAR model, three branches of the model, Granger causality test, impulse responses and variance decomposition analysis are reported.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

4.1. Stock Market Returns and Financial Stress Index

The Granger causality test states that if the lagged values of variable Y_2 help in predicting the values of the dependent variable Y_1 then Y_2 Granger-causes Y_1. Exhibit 5 reports the results of the Granger causality test. The Akaike Information Criteria (AIC) values support the usage of one-month lagged values. One-month lagged returns in the BRIC equity markets (–0.2439, p-value < 0.05) Granger-cause the financial stress index of India. The p-values are not found to be significant in the case of other countries examined. As a result of growing integration among the BRIC nations (Bhar & Nikolova, 2008; Dasgupta, 2014), the returns in the BRIC nations have an impact on the Indian financial stress. The coefficient is found to be negative in nature, which makes the case that if the returns in the BRIC countries are positive or in other words a buying spree is present in the BRIC equity markets, then the financial stress in India gets reduced. In the same way, when the returns are negative in the BRIC equity markets as a whole, then that downtrend does add to the financial stress of India. One-month lagged value of the financial stress index has a significant negative (-0.1718, p-value < 0.05) impact on the current financial stress in the Indian economy. On an overall basis, all the markets Granger-cause the financial stress index.

Exhibit 5
Granger Causality Test Results

Dependent variable: D (FSI)			
Excluded	**Chi-square**	**Degree Of Freedom**	**Probability**
Return_BRIC	7.455534	1	0.0063*
Return_EUROPE	0.039158	1	0.8431
Return_FRONTIER	0.318088	1	0.5728
Return_USA	0.747094	1	0.3874
All	33.03215	4	0.0000*

* Reject the null hypothesis of no significant relationship at the 5 percent significance level.

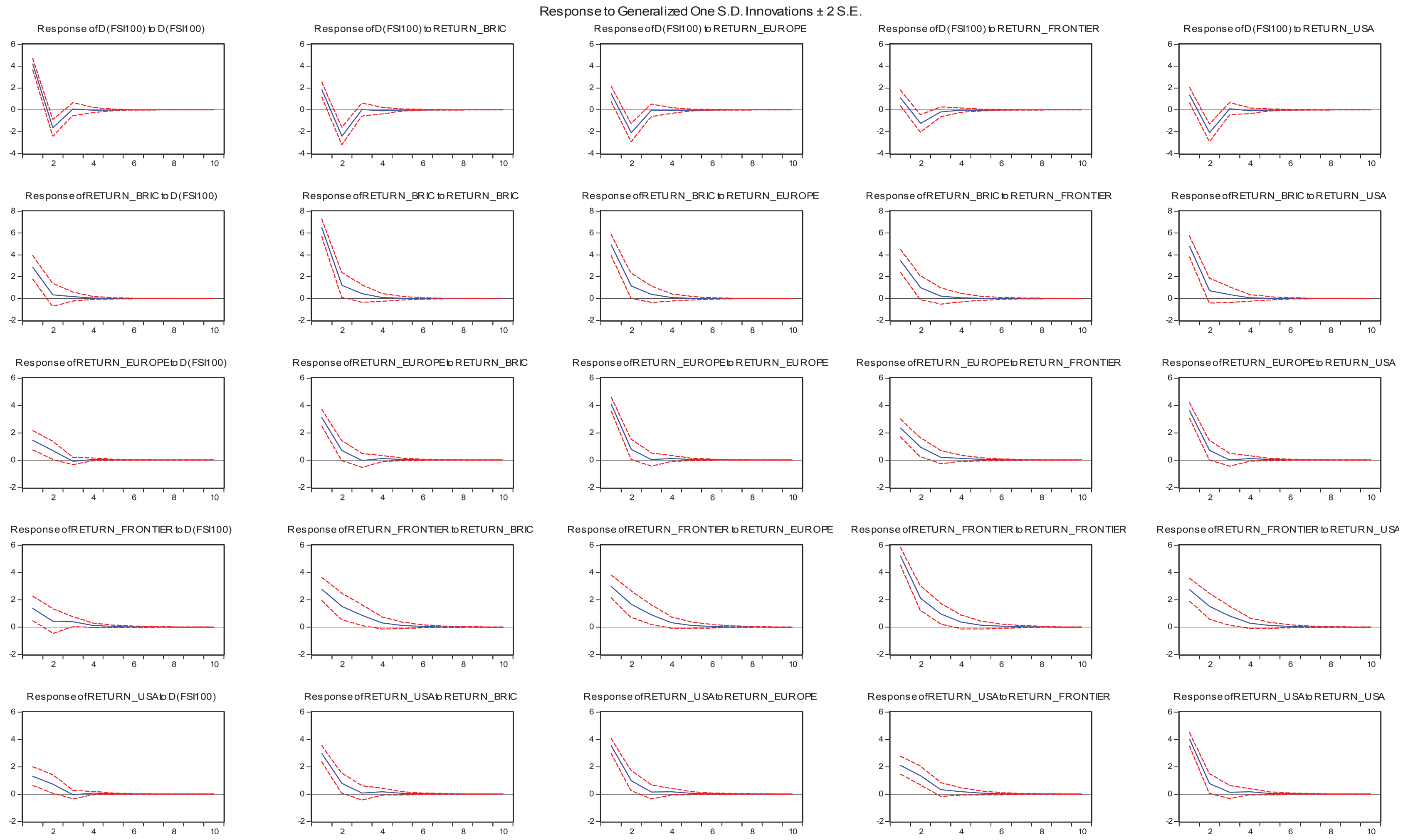
Source: Computed by the Authors.

The impulse responses highlight response of respective variables when a shock is subjected to an error term of an endogenous variable. As each variable enters into VAR equation in its lagged form, so it does have an impact on the other variable as well. Exhibit 6 reports the results of generalised impulse responses of the financial stress index as well as the other variables. We are concerned only with the results of the financial stress index which are in the first row. Initially, the response is positive when a shock is given to the US, Europe, frontier and BRIC stock markets. A positive response means that initially the return(s) shocks in these respective markets enhance the financial stress in India. The response becomes negative only after two months in each case and completely dies out after three to four months. In a nutshell, the shock in the equity index returns of the US, Europe, frontier and BRIC markets initially increases the Indian financial stress and the effect dies out after a few months. The investors in the Indian financial markets should consider the impact of a return(s) shock on the Indian financial stress.

After impulse response functions, another branch of the VAR model is the variance decomposition analysis. The analyses highlight contribution of one variable when a shock is subjected to the error terms of the latter in explaining variations in another variable at the time of forecasting. The time-varying contribution of the shocks shows spillover impact of the returns on the financial stress in the Indian economy. Under variance decomposition analysis, the ordering of the variables is very important as per the Cholesky Decomposition framework. We have done the ordering by assuming and placing the US market at the first place, considering its development

DOI: 10.7172/2353-6845.jbfe.2016.2.2

Exhibit 6
Generalised Impulse Responses



Source: Computed by the Authors.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

level and European and the BRIC markets at the second and third place respectively. The frontier markets and the financial stress index take the fourth and fifth place respectively. We have drawn out the results for twelve months.

Exhibit 7 reports the results of variance decomposition analysis. In the first month, when a shock is given to the financial stress index, then that contributes around 80 percent of the variation in the financial stress index itself. The US market contributes around 11 percent, whereas the contribution of the BRIC market is low. The contribution of the frontier markets remains negligible throughout the twelve months. During the fifth month, the US market accounts for around 26 percent of the variations, whereas the contribution of the financial stress reduces to approximately 60 percent. The contribution remains more or less similar with respect to all of the endogenous variables after five months.

The results reported by the Vector Autoregression model have further prompted us to study the relationship among the Brazilian, Russian, Chinese equity market returns and Indian financial stress index (excluding Indian market) under the VAR framework. To analyse the impact of the countries, the study uses monthly returns series of the benchmark indices of the respective countries, i.e. BOVESPA (Brazil), Russian Trading System RTS (Russia) and Shanghai Composite index SSE (China). The monthly returns are calculated in the similar fashion as mentioned in eq (1). The Augmented Dickey Fuller test reports stationarity in the data. The Akaike Information Criteria (AIC) values support the usage of two months' lagged values in the model.

Exhibit 7
Variance Decomposition of D (FSI)

Period	S.E.	D (FSI)	R_BRIC	R_Europe	R_Frontier	R_USA
1	4.173322	80.19323	7.171757	2.019459	0.050563	10.56499
2	4.884772	60.26348	10.86562	2.437928	0.172829	26.26014
3	4.899858	59.92068	10.80023	2.753622	0.386530	26.13894
4	4.900903	59.89559	10.80164	2.754271	0.386736	26.16177
5	4.901038	59.89233	10.80151	2.756472	0.387825	26.16187
6	4.901043	59.89221	10.80150	2.756482	0.387875	26.16193
7	4.901044	59.89218	10.80150	2.756489	0.387884	26.16195
8	4.901045	59.89217	10.80150	2.756490	0.387885	26.16195
9	4.901045	59.89217	10.80150	2.756490	0.387886	26.16195
10	4.901045	59.89217	10.80150	2.756490	0.387886	26.16195
11	4.901045	59.89217	10.80150	2.756490	0.387886	26.16195
12	4.901045	59.89217	10.80150	2.756490	0.387886	26.16195

Source: Computed by the Authors.

One-month and two-month lagged values of the financial stress index have a negative and statistically significant impact on the current financial stress at the 5 percent significance level. The impact of one-month lagged returns in the Brazilian market is positive (0.2059, $p<0.05$) and statistically significant. A positive sentiment in the Brazilian market increases the financial stress in the Indian financial market and vice versa for the negative returns. On the other hand, the impact of two months' lagged returns in the Brazilian market becomes negative (–0.2097, $p<0.05$) and significant with a similar magnitude. It exhibits that after two months the interaction between the countries concerned becomes more integrated, wherein the positive returns in the Brazilian market also reduce the stress in the Indian financial system and vice versa. Similarly, the

DOI: 10.7172/2353-6845.jbfe.2016.2.2

impact of two months' lagged returns is negative and statistically significant (–0.0781, p<0.10) at the 10 percent significance level in the context of Chinese market returns. The Granger causality results report the impact of only the Brazilian market on the financial stress index at the 5 percent significance level owing to increasing Brazil-India trade relations.

The generalised impulse responses came out with a finding that the response of the Indian financial stress towards the return(s) shocks in the Brazilian, Russian and Chinese markets becomes negative only after two months. As per the variance decomposition analysis, when a shock is given to the Brazilian market then that contributes around 12 percent variation in the financial stress in the second month but the contribution increases to 27 percent in the third month. The contribution of the Russian and Chinese markets remains 1 percent and 2 percent, respectively throughout the 12 months ahead variances. The ordering of the variables has been done considering the VAR model results. Overall, the VAR model is found to be stationary as all of the inverse roots lie inside the unit circle.

Furthermore, we have tried to capture the contemporaneous and dynamic impact of the Brazilian, Russian and Chinese equity market returns (excluding India) on the sub-components of the financial stress index by employing multivariate OLS regression model, wherein we have taken different sub-components as dependent variables and equity market returns in the respective markets as the independent variables. Before employing the model, the stationarity of the sub-components has been assured first. Exhibit 8 reports the contemporaneous impact of the markets on the sub-components.

Exhibit 8
Contemporaneous Impact on Sub-Components of the Financial Stress Index

	Bank Spread	Exchange Volatility	Debt Market	NIFTY Returns	NIFTY Volatility
Brazil	0.1631	–0.2320**	0.2846*	–0.2364	0.2360
Russia	–0.4619*	0.0683	–0.0775	0.3343*	–0.1142
China	0.4750*	0.0476	–0.0627	0.0807	–0.0805

Reject null hypothesis of no significant relationship at 5* and 10** percent significance level.

Source: Computed by the Authors.

There is a negative and statistically significant impact of the Russian market returns on the spread between the MIBOR and Treasury bills at the 5 percent significance level depicting increased integration among the markets. The falling returns in the Russian market increase the banking stress in the Indian economy. However, the impact of the Chinese market returns is positive on the banking spread signifying an increased level of banking stress with the positive returns in the Chinese markets. The Brazilian market has a negative (–0.2320, p<0.10) impact on the exchange rate volatility, wherein the falling returns increase the volatility in the Indian exchange rate. However, the Brazilian market has a positive impact on the spread between the 10-year Indian government securities and 10-year US government securities, thereby adding to the financial stress. The returns in the Russian market have a positive and statistically significant impact on the Indian market returns at the 5 percent significance level. There is no evidence of impact of the Brazilian, Russian and Chinese markets on the Indian equity market volatility. Exhibit 9 reports the dynamic impact of one-month lagged returns in the Brazilian, Russian and Indian markets on the sub-components.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

Exhibit 9
Dynamic Impact on Sub-Components of the Financial Stress Index

	Bank Spread	Exchange Volatility	Debt Market	NIFTY Returns	NIFTY Volatility
Brazil	–0.0423	0.0428	–0.1109	1.0048*	0.0554
Russia	–0.4689*	0.0525	0.0680	0.2659*	–0.1670**
China	0.3710	–0.0097	0.0155	0.1340	0.0978

Reject null hypothesis of no significant relationship at 5* and 10** percent significance level

Source: Computed by the Authors

The one-month lagged returns in the Russian market have a statistically significant impact on the bank spread, NIFTY returns and NIFTY volatility. The impact is negative with respect to the bank spread and NIFTY volatility. A negative return in the Russian market has a positive impact on the financial stress as it would entail an increase in the banking stress and equity volatility. Besides this, there is a strong positive impact of the one-month lagged return in the Brazilian market on the NIFTY returns (1.0048, $p<0.05$) at even 1 percent significance level. All of this empirical evidence testifies the importance of liberalising international financial flows, whereby global information transmissions are having an impact on different sub-components of the financial stress index in the form of stress in the money market as well as debt market.

4.2. Impact of US, Europe and Frontier Equity Market Returns on the Sub-Components

We further extended our analysis to study the impact of the US, Europe and frontier markets (excluding BRIC markets) on the sub-components of the financial stress index through multivariate OLS regression model. The returns in the frontier markets have a negative impact on the banking spread at the 5 percent significance level both contemporaneously and dynamically. A positive return in the frontier markets reduces the spread in the Indian banking sector, thereby exhibiting an increased level of integration among the markets. Besides this, a positive return in the European markets reduces spread in the Indian debt market segment (–0.8627, $p<0.05$) and increases the NIFTY returns (1.6499, $p<0.05$) contemporaneously and with a larger magnitude. There is a dynamic negative impact of one-month lagged returns in the European markets on the NIFTY volatility (–0.7584, $p<0.05$). A falling return increases volatility in the Indian equity market. Similarly a positive return in the US (0.4998, $p<0.10$) and frontier markets (0.2083, $p<0.10$) increases the spread in the debt market contemporaneously as the coefficients are found to be significant at the 10 percent level. This means that positive flows to the US and frontier markets increase long-term sovereign debt market risks in the Indian economy.

4.3. Conditional Volatility and Financial Stress Index

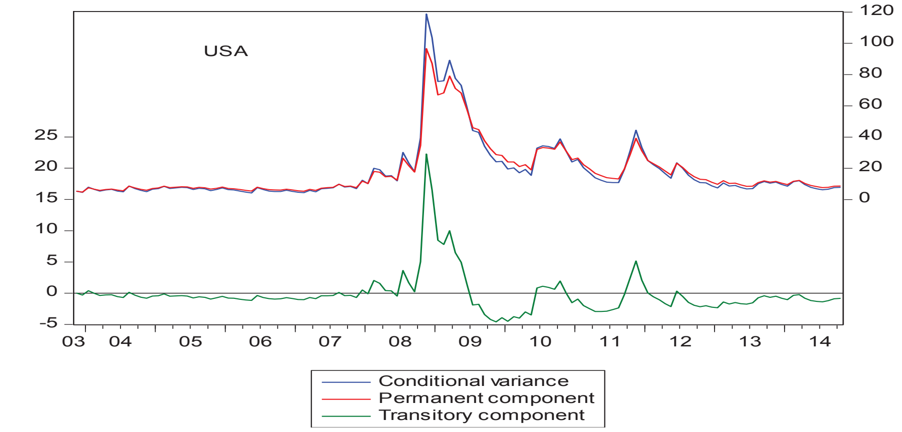
Lastly, we have studied the impact of the stock market volatility on the financial stress. Exhibit 10 presents the graphical images of the permanent, transitory and the overall conditional variances. In the case of the European markets, the transitory component of the volatility is very volatile in comparison to the other countries examined. A spike can easily be witnessed during the fourth quarter of year 2008 with respect to all of the markets, highlighting the existence of the US subprime loan crisis during that period. We have used the conditional variances derived from the CGARCH (1,1) model to account for the impact of the market volatility in the respective nations on the Indian financial stress and its sub-components. As mentioned earlier, one and two months' lagged values are included in the frontier markets' mean equation model. Both of the lagged values are observed to be having a significant impact on the current frontier markets' returns to

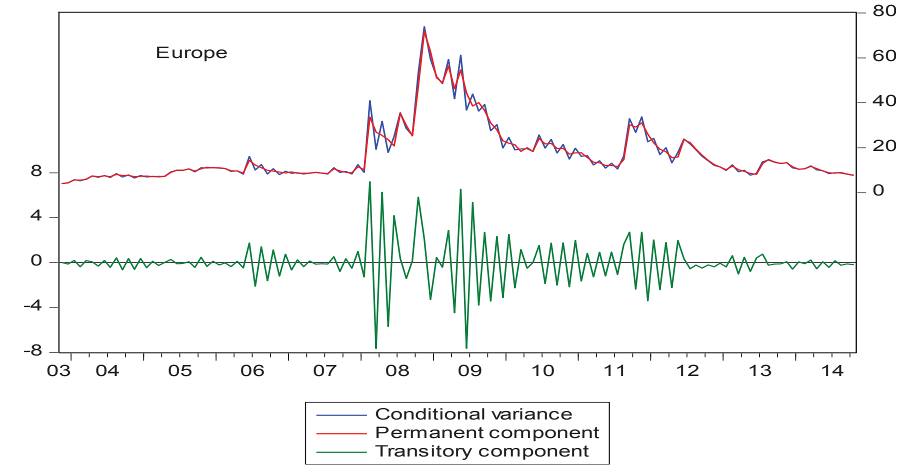
DOI: 10.7172/2353-6845.jbfe.2016.2.2

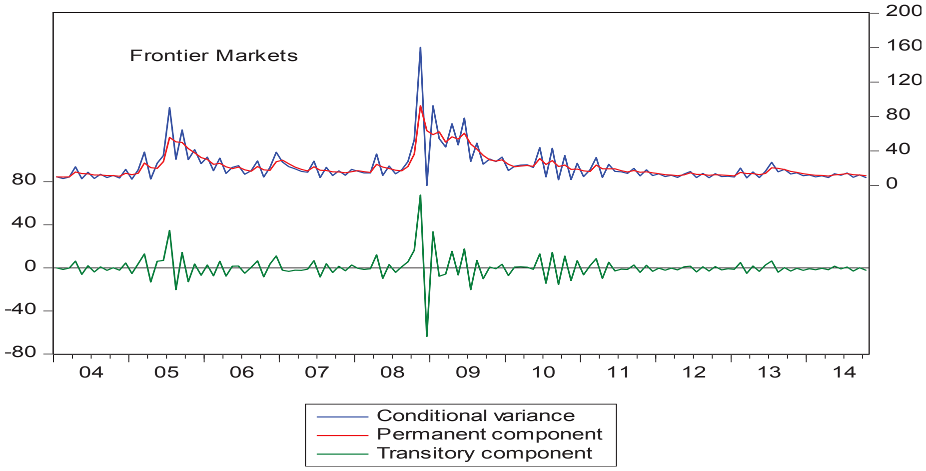
the tune of around 0.20 percent. This further supports inefficiency in the latter markets owing to significant impact of past month values on the current ones. It is pertinent to mention that all the GARCH based models are found to be adequate in the context of non-existence of serial autocorrelation and heteroskedasticity in the standardised error terms.

Exhibit 10

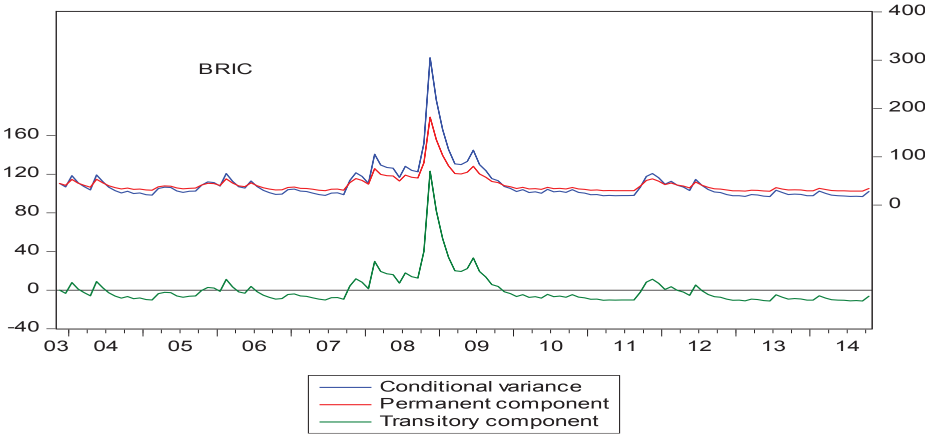
Graphical Presentation of the Conditional Variances







DOI: 10.7172/2353-6845.jbfe.2016.2.2



Source: Computed by the Authors.

Exhibit 11 reports the results of impact of the transitory component of volatility on the financial stress index of India. The results are derived by employing multivariate OLS regression model. A statistically significant contemporaneous impact of the short-run volatility running from the European stock market to the Indian financial system has been observed. The coefficient is positive in nature signifying a positive impact on the financial stress whenever there is an increase in the short-run component of the volatility. For the rest of the nations, the impact is not found to be statistically significant at the 5 percent level.

Similarly, when we tried to capture the impact of the long-run component of the conditional variance on the financial stress, then the results are not found to be significant at the 5 percent level. The results of the long-run component of the volatility on the financial stress have not been reported but can be provided on request. Even the impact of one-month lagged volatility (dynamic volatility) on the financial stress is not found to be statistically significant at the 5 percent level in the context of both the short-run as well as long-run component of the conditional volatility.

Exhibit 11
Transitory Component of Volatility and Financial Stress Index

	Coefficient	**T-statistic**	**P-value**
$Ø_2$	–0.0019	–0.0416	0.9668
$Ø_3$	0.2429	1.0123	0.3133
$Ø_4$	0.6024	3.1213	0.0022*
$Ø_5$	–0.0055	–0.1498	0.8811

*Reject null hypothesis of no significant relationship at the 5 percent significance level.

Source: Computed by the Authors.

Exhibit 12 reports the results of impact of unexpected volatility in the markets of the US, Europe, BRIC and frontier markets on the Indian financial stress. The unexpected volatility in the BRIC stock markets has a positive and statistically significant impact on the Indian financial stress contemporaneously with a larger magnitude. The dynamic impact (one-month lagged) of the unexpected volatility in the BRIC stock markets statistically reduces the financial stress in the Indian economy.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

Exhibit 12
Impact of Unexpected Component of Volatility on Financial Stress Index

	Contemporaneous Impact			Dynamic Impact		
	Coefficient	**T-stat**	**P-value**	**Coefficient**	**T-stat**	**P-value**
∞_2	1.2370	2.0818	0.0394*	–1.1450	–2.0877	0.0389*
∞_3	–0.3982	–0.5077	0.6125	–0.7249	–0.9903	0.3239
∞_4	0.6695	0.8107	0.4190	–0.3814	–0.4951	0.6214
∞_5	–0.2414	–0.4981	0.6193	–0.3764	–0.8370	0.4042

* Reject null hypothesis of no significant relationship at the 5 percent significance level.

Source: Computed by the Authors.

In other words, the same month unexpected volatility in the BRIC markets adds to the financial stress, whereas the one-month lagged unexpected volatility in the BRIC markets reduces the financial stress exhibiting a reducing impact of the past unexpected variations. The values are not found to be significant for the other countries. Similar types of results were reported by the regression model when we took transitory component and unexpected volatility component simultaneously in the regression equation in terms of both the contemporaneous as well as dynamic impact at 5 and 10 percent level. The unexpected variations in the market make the investors expect a risk premium for holding riskier emerging market assets (Kumar & Dhankar, 2009). The expectations of high risk premiums in the BRIC equity markets further reduce the stock market returns coupled with increased volatility adding to the financial stress.

To understand the behaviour of the Indian financial stress in a much more calibrated manner, we have tried to analyse the impact of the individual Brazilian, Russian and Chinese markets' conditional volatility (excluding India) on the Indian financial stress index by employing multivariate OLS regression model, wherein the conditional variances are taken as independent variables and financial stress index as dependent variable. Surprisingly, we could not find any ARCH effects in the Brazilian market. So we have studied the impact of only the Russian and Chinese markets' conditional volatility on the Indian financial stress. The results reported that there is no statistically significant impact of the unexpected volatility in the Russian and Chinese markets on the Indian financial stress index either contemporaneously or dynamically. This means that the impact of the BRIC markets' unexpected volatility is largely due to the impact of the Indian market itself on the financial stress. On a similar note, the stress in the Indian financial system does not get affected by the permanent volatility in the Russian and Chinese markets. However, the short-run volatility in the Chinese market has a reducing impact on the Indian financial stress (–0.0642, $p<0.05$) contemporaneously, whereas the impact becomes positive after one month (0.0462, $p<0.05$) with a lesser magnitude. Furthermore, the impact of the unexpected volatility in the Russian and Chinese markets on the sub-components of the financial stress index is not found to be statistically significant at the 5 percent significance level, either contemporaneously or dynamically barring the positive and stronger dynamic impact of one-month lagged unexpected volatility in the Russian (5.5164, $p<0.05$) and Chinese market (2.3239, $p<0.05$) on the NIFTY returns. The permanent volatility in the Russian market has a significant positive and negative impact on the banking spread and exchange rate volatility respectively in a longer period (dynamic impact). Even the transitory component has a positive impact on the banking spread. The debt market spread and NIFTY volatility also gets positively influenced by the transitory volatility in the Chinese market (0.0729, $p<0.05$) and the Russian market (0.0676, $p<0.05$) respectively, after one month. There is a contemporaneous negative impact of transitory volatility in the Chinese market (–0.1391, $p<0.05$) and the Russian market (–0.0446, $p<0.05$) on the NIFTY returns. If the short-run volatility in the Chinese market increases by 1 percent then the returns in the Indian market reduces by 0.1391 percent.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

4.4. Impact of US, Europe and Frontier Markets Volatility on the Sub-Components

We extended our analysis to study the impact of the US, European and frontier markets' volatility (excluding BRIC markets) on the sub-components of the financial stress index by employing multivariate OLS regression model. The permanent component of the volatility in the European market increases banking spread in the Indian economy both contemporaneously and dynamically. However, permanent volatility in the US market dynamically reduces NIFTY volatility (–0.3389, $p<0.05$) and has an increasing impact on the NIFTY returns (0.4127, $p<0.05$ and 0.4507, $p<0.05$) both contemporaneously and dynamically. The permanent volatility in the European markets has a reducing impact on the NIFTY returns (–0.6184, $p<0.05$ and –0.6826, $p<0.05$) both contemporaneously and dynamically with a stronger magnitude. Similarly, one-month lagged permanent volatility in the European market has a reducing dynamic impact on the exchange rate volatility in the Indian economy (–0.2571, $p<0.10$), whereas it dynamically increases the volatility in the Indian equity market (0.4515, $p<0.05$).

The transitory component of the volatility in the US market increases banking spread in the Indian market, thereby adding to the banking stress in the short run (2.8587, $p<0.01$ and 2.6183, $p<0.01$) both contemporaneously and dynamically, whereas, on the other hand, the transitory volatility in the US (–0.3106, $p<0.10$) and frontier markets (–0.1002, $p<0.05$) contemporaneously reduces exchange rate volatility in the Indian market to the tune of around 0.31 and 0.10 percent respectively. On a similar note, transitory volatility in the US (0.6814,$p<0.01$) and frontier markets (0.2774, $p<0.01$) contemporaneously increases volatility in the Indian equity market, and the impact of the US market is observed to be higher. One-month lagged transitory volatility in the European markets has a reducing dynamic impact on the exchange rate volatility (-0.5163,$p<0.05$) at the 5 percent significance level, whereas the impact is strongly positive on the Indian equity market volatility (1.0148, $p<0.01$) dynamically. Similarly, one-month lagged transitory volatility in the frontier markets has a reducing dynamic impact on the debt spread (–0.0894, $p<0.10$).

The unexpected component of the volatility in the European market has a strong reducing impact on the debt spread in the Indian market (–2.4148, $p<0.05$) contemporaneously. There is a contemporaneous negative impact of the unexpected volatility in the European market on Indian exchange rate volatility at the 10 percent significance level (–1.8518, $p<0.10$). The Indian equity market returns are greatly influenced by the unexpected volatility in the US, European and frontier markets, as all the coefficients are found to be significant at 5 and 10 percent significance level contemporaneously. Moreover, one-month lagged unexpected volatility in the frontier markets has a dynamic reducing impact on the banking spread at the 10 percent significance level. Interestingly, unexpected volatility in the respective markets does not have a statistically significant impact on the Indian equity market volatility either dynamically or contemporaneously.

5. CONCLUDING REMARKS

Studies relating to information transmission across different countries are of paramount interest to the international investors. The cross-market impact highlights sensitivity of the domestic economies to the foreign information elements. However, the present study adds to the literature by capturing the impact of the said global information transmissions on Indian financial stress index and its various sub-components. To study the impact of the US, Europe, frontier and BRIC stock markets on the Indian financial stress index, we have primarily employed Vector Autoregression and Component GARCH (1,1) models with the monthly data ranging from year 2003 to 2014. In order to construct the financial stress index, four major segments of the financial system: Equity market, Debt market, Foreign Exchange market and the Money market are taken into consideration. The Johansen Cointegration test indicates that there is no long-run

DOI: 10.7172/2353-6845.jbfe.2016.2.2

co-movement between the financial stress index and the MSCI indices of the respective nations. So, there are short-run dynamic interactions among the respective equity markets and financial stress index that channelize the impact of equity markets on the latter. Overall, the results of the VAR model report that only the BRIC market returns contribute to the financial stress index. The short-term dynamic relationship is negative in nature, wherein positive returns in the BRIC nations reduce stress in the Indian financial system highlighting increasing integration among the markets. The impulse responses of the financial stress index dies out after three to four months, whereas the contribution of the US market to the variations in the financial stress index increases over a period of time. The stress in the Indian financial system responds statistically significantly to the Brazilian and Chinese market returns. The banking spread, NIFTY returns and NIFTY volatility component of the stress index gets strongly influenced by the one-month lagged Russian market returns but there is no significant impact of the Russian market returns on the overall financial stress. Moreover, a positive return in the European market reduces the spread in the Indian debt market segment and increases the NIFTY returns contemporaneously and with a larger magnitude.

A statistically significant impact of the short-run component of the volatility in the European market on the Indian financial stress has also been found. Further, the unexpected volatility in the BRIC markets has a significant impact on the Indian financial stress. But the impact of the BRIC markets' unexpected volatility is largely due to the impact of the Indian market itself on the financial stress. There is a positive and a stronger dynamic impact of one-month lagged unexpected volatility in the Russian and Chinese markets on the NIFTY returns. Both the permanent and transitory components of the volatility in the European market have a strong and significant positive impact on the NIFTY volatility dynamically. On the other hand, the permanent component of the conditional volatility in the US market helps in reducing overall financial stress because it increases the NIFTY returns both contemporaneously and dynamically and also reduces the NIFTY volatility with a lagged impact. The results show that, notwithstanding, most of the countries do not have a statistically significant impact on the overall financial stress but they do have an impact on the sub-components in a much more calibrated manner owing to response of respective sub-components towards global information transmissions. The impact of the stock markets on the sub-components of the financial stress also spotlights the transmission channels through which these equity market spillovers have an impact on the financial stress.

A stock market discounts every type of information in advance and positive behaviour in the market indicates wellness of the economy as a whole. The investments in the BRIC stock markets by the international investors exhibit overall positive behaviour of the investors toward the emerging nations, which further entails increased inflows of foreign funds in the latter markets. This positive behaviour also helps in reducing stress in the overall financial system because these foreign financial flows have a spillover impact on the other sub-components of the financial system as well, like favourable banking spreads, reduced debt market spreads and lower exchange rate volatility. A general conclusion that can be drawn from the study is that the BRIC markets in the first as well as second moment and the European markets in the second moment have a direct and statistically significant impact on the Indian financial stress index in the short run. A possible reason that could be attributed to the Indian financial stress being sensitive to the European short-run volatility in the markets can be monetary stimulus policies adopted by the European nations, hence increasing financial flows. The policy makers and especially the investors in the financial markets comprising equity, debt and currency should discount the information coming from the European stock markets and the BRIC stock markets because these markets have an impact not only on the overall financial stress in the Indian economy but also on the core sub-components of the financial system with a greater magnitude as compared to others. Moreover, different monetary policy initiatives are also undertaken considering the co-movement and dynamic interactions among the underlying markets due to the international transmission of shocks through equity

DOI: 10.7172/2353-6845.jbfe.2016.2.2

markets and confidence levels (Berben & Jansen, 2005). As a future scope of research, the other components or participants of the global financial markets may have an impact on overall Indian financial stress and its sub-components.

Acknowledgements

We are grateful to the anonymous referees for providing many insightful suggestions. We would also like to express our gratitude to Dr. Navneet Kaur for assisting in data collection.

References

Angkinand, A.P., Barth, J.R., Kim, H. (2010), Spillover effects from the U.S. financial crisis: Some time-series evidence from national stock returns. In G. Benton (Ed.), *The Financial and Economic Crises: An International Perspective* (pp. 24–52). Cheltenham, UK: Edward Elgar.

Balakrishnan, R., Danninger, S., Elekdag, S., Tytell, I. (2011), The Transmission of Financial Stress from Advanced to Emerging Economies. *Emerging Markets Finance and Trade, 47*(2), pp. 40–68.

Berben, R., Jansen, W.J. (2005), Comovement in international equity markets: A sectoral view. *Journal of International Money and Finance*, *24*(5), pp. 832–857.

Bhar, R., Nikolova, B. (2007), Analysis of Mean and Volatility Spillovers using BRIC Countries, Regional and World Equity Index Returns. *Journal of Economic Integration, 22*(2), pp. 369–381.

Bianconi, M., Yoshino, J.A., Machado, M.O. (2013), BRIC and the U.S. financial crisis: An empirical investigation of stock and bond markets. *Emerging Markets Review, 14*, pp. 76–109.

Bollerslev, T. (1986), Generalized Autoregressive Conditional Heteroskedasticity. *Journal of Econometrics, 31*, pp. 307–327.

Booth, G.G., So, W.R., Tse, Y. (1999), Price discovery in the German equity index derivatives markets. *The Journal of Futures Markets*, *9*(6), pp. 619–643.

Christie, A.A. (1982), The stochastic behavior of common stock variances: Value, leverage and interest rate effects. *Journal of Financial Economics*, *10*, pp. 407–432.

Christopoulos, A.G., Mylonakis, J., Koromilas, C. (2011), Measuring the Impact of Financial crisis on International Markets: An Application of the Financial Stress Index. *Review of European Studies, 3*(1), pp. 22–34.

Dasgupta, R. (2014), Integration and dynamic linkages of the Indian stock market with BRIC: An empirical study. *Asian Economic and Financial Review*, *4*(6), pp. 715–731.

Davig, T., Hakkio, C. (2010), *What Is the Effect of Financial Stress on Economic Activity?* Retrieved from https://www.kansascityfed.org/publicat/econrev/pdf/10q2Davig_Hakkio.pdf

Ding, Z., Granger, C.W.J., Engle, R.F. (1993), A long memory property of stock market returns and a new model. *Journal of Empirical Finance, 1*(1), pp. 83–106. doi: 10.1016/0927-5398(93)90006-D.

Dooley, M.P., Hutchison, M.M. (2009), *Transmission of the U.S. subprime crisis to emerging markets: Evidence on the decoupling-recoupling hypothesis.* Retrieved from http://sciie.ucsc.edu/workingpaper/2009/DooleyHutchison%20%20Emerging%20Mrkt%20linkages%20March%2031.pdf

Enders, W. (2014), *Applied Econometric Time Series.* New Jersey: John Wiley & Sons.

Frank, N., Hesse, H. (2009), *Financial Spillovers to Emerging Markets During the Global Financial Crisis.* Retrieved from http://core.ac.uk/download/pdf/6605059.pdf

Gagnon, L., Karolyi, A.G. (2006), Price and volatility transmission across borders. *Financial Markets, Institutions and Instruments, 17*(3), pp. 107–158.

Goldstein, M., Xie, D. (2009), *The Impact of the Financial Crisis on Emerging Asia.* Retrieved from http://www.frbsf.org/economic-research/files/09_Goldstein.pdf

Gujarati, D., Porter, D., Gunasekar, S. (2013), *Basic Econometrics.* New Delhi: McGraw Hill Education.

Hakkio, C.S., Keeton, W.R. (2009), Financial stress: What is it, How can it be measured, and Why does it matter?. *Economic Review, Q*(II), pp. 5–50.

Illing, M., Liu, Y. (2006), Measuring financial stress in a developed country: An application to Canada. *Journal of Financial Stability*, *2*(3), pp. 243–265.

Kharchenko, I., Tzvetkov, P. (2013), *Estimation of Volatilities and Spillover Effects Between Developed and Emerging Market Economies*. Retrieved from http://lup.lub.lu.se/luur/download?func=downloadFile&recordOId=3801343&fileOId=3813321

Kliesen, K.L., Owyang, M.T., Vermann, E.K. (2012), Disentangling Diverse Measures: A Survey of Financial Stress Indexes. *Federal Reserve Bank of St. Louis Review, 94*(5), pp. 369–397.

DOI: 10.7172/2353-6845.jbfe.2016.2.2

Kumar, R., Dhankar, R.S. (2009), Asymmetric Volatility and Cross Correlations in Stock Returns under Risk and Uncertainty. *Vikalpa: The Journal for Decision Makers, 34*(4), pp. 25–36.

Lall, S., Cardarelli, R., Elekdag, S. (2008), *Financial Stress and Economic Downturns.* Retrieved from http://www.imf.org/external/pubs/ft/weo/2008/02/pdf/c4.pdf

Manamperi, N.P. (2015), A Comparative Analysis on US Financial Stress Indicators. *International journal of Economics and Financial issues, 5*(2), pp. 613–623.

Morales, M.A., Estrada, D. (2010), A financial stability index for Colombia. *Annals of Finance*, *6*, pp. 555–581.

Mukherjee, K.N., Mishra, R. (2010), Stock Market Integration and Volatility Spillover: India and its Major Asian Counterparts. *Research in International Business and Finance, 24*(2), pp. 235–251.

Park, C., Mercado, R.V. (2013), *Determinants of financial stress in Emerging Market Economies*. Retrieved from http://www10.iadb.org/intal/intalcdi/PE/2013/12502.pdf

Pentecost, E.J., Moore, T. (2006), Financial Liberalisation in India and a New Test of the Complementarity Hypothesis. *Economic Development and Cultural change, 54*(2), pp. 487–502.

Rachdi, H. (2013) The Impact of the International Financial Crisis on the Stock Market Return: The Case of Tunisian Stock Exchange. *Journal of Empirical Economics, 1*(2), pp. 67–74.

Rittler, D. (2012) Price discovery and volatility spillovers in the European Union emissions trading scheme: A high-frequency analysis. *Journal of Banking & Finance, 36*(3), pp. 774–785.

Roye, B.V. (2014) Financial Stress and economic activity in Germany and the Euro Area. *Empirica, 41*(1), pp. 101–126. doi: 10.1007/s10663-013-9224-0

Sandahl, F., J., Holmfeldt, M., Rydén, A., Strömqvist, M. (2011), An index of financial stress for Sweden*. Sveriges riksbank economic review, 2,* pp. 49–67.

Sims, C.A. (1980), Macroeconomics and Reality. *Econometrica, 48*(1), pp. 1–48.

Sum, V. (2013), Impulse Response Functions and Causality Test of Financial Stress and Stock Market Risk Premiums. *International Journal of Financial Research, 4*(1), pp. 1–4.

Wallace, C. (2013), *Financial Stress and its Impact on Economic Activity: Evidence from Jamaica.* Retrieved from http://www.ccmf-uwi.org/files/publications/conference/2013/1_2-Wallace-p.pdf

Worthington, A., Higgs, H. (2004), Transmission of equity returns and volatility in Asian developed and emerging markets: A multivariate GARCH analysis. *International journal of finance and economics, 9*(1), pp. 71–80.

Yiu, M.S., Ho, W.Y.A., Jin, L. (2010), *A Measure of Financial Stress in Hong Kong Financial Market — the Financial Stress Index.* Retrieved from http://www.hkma.gov.hk/media/eng/publication-and-research/research/research-notes/RN-02-2010.pdf

DOI: 10.7172/2353-6845.jbfe.2016.2.2

Foreign Investor Flows and Sovereign Bond Yields in Advanced Economies

Serkan Arslanalp[1]
International Monetary Fund, USA
sarslanalp@imf.org

Tigran Poghosyan
International Monetary Fund, USA
tpoghosyan@imf.org

Received: 11 January 2016 / Revised: 28 April 2016 / Accepted: 25 May 2016 / Published online: 16 June 2016

ABSTRACT

Asset allocation decisions of international investors are at the core of capital flows. This paper explores the impact of these decisions on long-term government bond yields, using a quarterly investor base dataset for 22 advanced economies over 2004–2012. We find that a one percentage point increase in the share of government debt held by foreign investors can explain a 6–10 basis point reduction in long-term sovereign bond yields over the sample period. Accordingly, international flows to core advanced economy bond markets over 2008–12 are estimated to have reduced 10-year government bond yields by 40–65 basis points in Germany, 20–30 basis points in the U.K., and 35–60 basis points in the U.S. In contrast, foreign outflows are estimated to have raised 10-year government bond yields by 40–70 basis points in Italy and 110–180 basis points in Spain during the same period. These results suggest that changes in the foreign investor base for sovereign debt can have economically and statistically significant effects on sovereign bond yields, independent of other standard macroeconomic determinants of bond yields.

JEL classification: E4, E6, G1.

Keywords: Government bond yields, investor base, interest rate determinants.

1. INTRODUCTION

The divergence in long-term sovereign bond yields experienced by advanced economies in the aftermath of the global financial crisis has featured prominently in policy discussions over the last several years (IMF 2011, IMF 2012a, IMF 2012b). On the one hand, long-term yields in core advanced economies are perceived to be below their fundamental value, which is commonly attributed to the quantitative easing policies pursued by the central banks. On the other hand, sovereign bond yields in a number of euro area countries are perceived to be above their fundamental level due to capital outflows and elevated perceptions of tail risks.

The purpose of this paper is to analyze the factors driving the divergence in sovereign bond yields among advanced economies (AEs) from the perspective of foreign investor decisions.

[1] Corresponding author: International Monetary Fund, 700 19th Street, N.W., Washington, D.C. 20431, USA. Tel: +1 (202) 623-5420, SArslanalp@imf.org
The views expressed in this paper are those of the authors and do not necessarily represent the views of the IMF, its Executive Board, or IMF management.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

While we acknowledge the importance of fundamental *macroeconomic, monetary, and fiscal policy determinants* for explaining the dynamics of long-term sovereign bond yields in AEs, we focus our analysis on the impact of the *foreign investor base* (FIB) of sovereign debt, which has received less attention in cross-country studies and policy discussions.[2]

In particular, we argue that shifts in the foreign investor base observed in AEs in the aftermath of the global financial crisis have contributed to the diverging movements in long-term sovereign bond yields observed through end-2012. More specifically, foreign inflows to the United States, the United Kingdom, and core euro area countries have put downward pressure on sovereign bond yields, while foreign outflows have resulted in upward pressure in the periphery euro area countries.[3]

In order to analyze the impact of the FIB on long-term sovereign bond yields, we utilize a comprehensive dataset on the holders of government debt compiled by Arslanalp and Tsuda (2014). This dataset contains quarterly information on the composition of government debt over 2004–2012 for 22 advanced economies. We focus our attention on the FIB, including whether the investors represent foreign official or foreign private investors.

Our analysis suggests that:

- A rising foreign share of government debt holdings is associated with a statistically and economically significant decline in long-term sovereign bond yields. On average, a one percentage point increase in the share of general government debt held by non-residents can explain a 6–10 basis point decrease in 10-year sovereign bond yields.
- The type of foreign investors also matters. The impact of official foreign investors on long-term sovereign bond yields is slightly smaller (7 basis points) compared to that of foreign private investors (8.5 basis points). However, the difference in coefficients is not statistically significant.
- The increase in the FIB in the aftermath of the global financial crisis contributed to a decline in the long-term sovereign bond yields in the United States, the United Kingdom, and Germany of 35–60, 20–30, and 40–65 basis points, respectively. By contrast, the decline in the FIB in the aftermath of the global financial crisis contributed to an increase in the long-term sovereign bond yields in Spain and Italy of 110–180 and 40–70 basis points, respectively.

Our results have several implications. They suggest that "normalization" of macroeconomic, monetary, and fiscal policy determinants of bond yields may be insufficient to bring long-term rates back to their pre-crisis level unless this is accompanied by a similar "normalization" of the FIB. To the extent that the current benchmark portfolio allocations by foreign investors are distorted by "safe-haven" considerations and that these portfolio allocations can be persistent, our analysis suggests that the currently observed divergence in long-term sovereign bond yields in advanced economies may continue in the foreseeable future.

The remainder of the paper is structured as follows. Section II reviews the existing literature, with a focus on the foreign investor base as a determinant of sovereign bond yields. Section III describes the methodology employed in the analysis. Section IV describes the data and provides descriptive statistics. Section V presents the empirical findings, and the last section concludes with some policy implications.

2. LITERATURE REVIEW

On February 16, 2005, Federal Reserve Chairman Alan Greenspan described in a congressional testimony the recent behavior of long-term U.S. Treasury yields as a "conundrum." Specifically, he was referring to the atypical situation in the middle of 2004 when, despite monetary tightening,

[2] The importance of the investor base for the sustainability of public debt in market access countries has been emphasized in the recently revised IMF staff guidance note on debt sustainability (IMF, 2013b).

[3] Following the 2013 Article IV report for the euro area, periphery countries include Greece, Ireland, Italy, Portugal, and Spain.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

the Treasury yields continued their decline and were fluctuating at levels well below what one would expect on the basis of economic fundamentals (inflation, growth, fiscal and monetary policy stances). A number of follow-up academic studies have shown that a possible cause of the conundrum was the structural change in the investor base of U.S. Treasury securities, with the expansion in foreign demand for U.S. bonds (especially from Asia) depressing the long rates by tens of basis points (e.g., Bernanke, 2005). Moreover, it seems that the conundrum was not limited to the United States only, but applied to euro area countries as well around the same period (e.g., Hördahl, Tristani, and Vestin, 2006).

Foreign purchases of government debt could lead to a decline in bond yields for a number of reasons. First, a foreign purchase of an asset is, by definition, a capital inflow. And capital inflows, all else equal, can reduce the cost of funding for domestic borrowers to the extent that they expand the domestic savings pool available for domestic borrowers (i.e., as long as capital inflows are not offset by capital outflows one-for-one). Second, financial integration, driven by a desire for portfolio diversification by global investors and an expanding pool of world savings, increase cross-border flows; as a result, foreign inflows can lead to a convergence in real interest rates. Finally, foreign investors may have a higher demand for liquidity and safety than domestic investors (e.g., domestic banks), which can lead to market segmentation and deviations from the standard term structure of interest rates due to supply-demand imbalances.[4]

2.1. Single-country studies

Single-country studies are mostly limited to the case of the United States. Bernanke (2005) has attributed some of the decline in U.S. long-term bond yields since 2000 to a "global savings glut." Starting from the early 2000s, a significant share of global foreign exchange reserves were invested in U.S. Treasury securities (36 percent in 2010), and foreign official holdings of U.S. Treasury securities increased from $400 billion in 1994 to $3 trillion in 2010, suppressing long-term sovereign bond yields in the U.S. (Beltran and others, 2012).

Motivated by the above arguments, Warnock and Warnock (2009) conduct a quantitative assessment of the change in the FIB, driven by the rising foreign demand on long-term bond yields in the United States, using monthly data spanning January 1984–May 2005. They find that foreign purchases of U.S. Treasury securities have an economically large and statistically significant impact on long-term interest rates. Their simulations suggest that if foreign holdings of U.S. debt had not accumulated over the 12 months ending May 2005, then 10-year Treasury yields would have been around 80 basis points higher. They also consider other financial instruments, including 2-year Treasury yields, high- and low-quality corporate debt (Aaa and Baa), and long-term fixed and short-term adjustable mortgage rates. The impact of foreign inflows differs across these instruments, but it is always statistically significant and often economically large. In particular, they find that the 2-year bond yields are less affected by foreign flows, explaining this by the fact that they are more closely linked to short-term monetary policy rates rather than macroeconomic fundamentals.

Other studies confirming the negative association between foreign purchases of U.S. Treasuries and long-term rates are Bernanke, Reinhart, and Sack (2004); Beltran and others (2012); and Kaminska, Vayanos, and Zinna (2011). By contrast, Rudebush, Swanson, and Wu (2006) find that changes in the FIB were not so important in explaining the conundrum and the decline in long-term sovereign bond yield volatility observed during the same period.

[4] For example, Bernanke (2013) argues that the global economic and financial stresses of recent years—triggered by the financial crisis and then by the problems in the euro area—may have elevated the safe-haven demand for Treasury securities, pushing down Treasury yields and implying a lower, or even negative, term premium. Krishnamurthy and Vissing-Jorgensen (2012) estimate that U.S. Treasury bond yields may have been reduced by 73 basis points, on average, from 1926 to 2008 given their extreme safety and liquidity. Hördahl, Tristani, and Vestin (2006) show that safe-haven demand has decreased yields also in the core euro area countries. Kaminska, Vayanos, and Zinna (2011) provide a structural model of the term structure of interest rates that is consistent with no arbitrage, but allows for market segmentation between arbitrageurs and preferred-habitat investors with preferences for specific maturities.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

2.2. Cross-country studies

Few studies have analyzed the impact of the FIB on long-term sovereign bond yields in a cross-country context. Andritzky (2012) put together a database on the composition of FIB for government securities in the G-20 AEs (six countries) and the euro area (seven countries). Using quarterly data for 2000-2010, he finds that a 10 percentage point increase in the share of debt in AEs held by non-residents leads to a reduction in long-term sovereign bond yields of between 32 and 43 basis points. The impact is stronger at around 60 basis points for the sample of euro area countries.

Hauner and Kumar (2006) explicitly focus on the impact of capital flows in their attempt to resolve the "conundrum" of low government bond yields and high fiscal imbalances observed in G-7 advanced economies before the crisis. Their results suggest that the upward pressures on government bond yields due to chronic weakening of budgetary positions was more than offset by foreign inflows triggered by "safe-haven" considerations. However, they warn about the temporary nature of these effects and predict that upward correction in bond yields is inevitable in the long run.

Unlike the above studies, Lam (2013) finds no significant relationship between the FIB and 5-year forward contracts on 5-year sovereign bond yields in 12 AEs over 1990–2012. However, the dependent variable in his analysis captures only the 5-year maturity and may not fully capture the long-term borrowing costs of the sovereign. In addition, liquidity and default risk characteristics of forward contracts may be different from those of long-term sovereign bond yields, which could affect the estimations. Nevertheless, Lam (2013) finds a significant negative association between forward rates and central bank holdings of sovereign bonds, suggesting that changes in the domestic investor base may have implications for sovereign bond yields. Furthermore, in a similar study for Japan, Ichiue and Shimizu (2012) find that when an increase in government debt is financed entirely by borrowing from external sources, which leads to identical increases in government and foreign debt, the increase in the forward rate is approximately twice that when financed domestically.

As opposed to the previous literature analyzing bond yields, Dell'Erba, Hausmann, and Panizza (2013) focus their attention on the impact of debt composition on bond spreads in 15 AEs. Their analysis suggests that there is no significant association between bond spreads and the share of external debt. However, the coefficient on the debt composition variable turns significant and negative when an interaction term of the debt composition and debt level is added to the regression.

The relationship between the FIB and bond yields was also studied in the context of emerging markets (EMs). Peiris (2010) analyzes the impact of foreign participation on local-currency government bond yields in a panel of 10 emerging markets (EMs) for the period 2000:Q1–2009:Q1. The estimation results suggest a slightly stronger impact in EMs, with a 10 percentage point increase in the share of foreign debt leading to a 60 basis points decline in domestic bond yields. Using a panel of 13 EMs and 30 AEs between 2000 and 2012, IMF (2013c) and Jaramillo and Zhang (2013) show that "buy and hold" investors, including national and foreign central banks, are able to provide a more stable source of demand for government debt, contributing to the reduction of sovereign bond yields and their volatility. Dell'Erba, Hausmann, and Panizza (2013) analyze the impact of debt composition on bond spreads in 26 EMs. They find that a larger share of foreign debt is associated with higher spreads in EMs. However, this relationship turns insignificant when an interaction term of the debt composition and debt level is added to the regression.

Recognizing the importance of the investor base for sovereign borrowing costs, a separate stream of literature analyzes macroeconomic and institutional determinants of the investor base in a cross-country setting. One of the first studies on the topic is Burger and Warnock (2006), who

DOI: 10.7172/2353-6845.jbfe.2016.2.3

use BIS cross-sectional data on domestic securities in 49 countries (27 EMs and 22 AEs) and find that low inflation, rule of law, and country size positively affect the development of the domestic government bond market, while fiscal balance and GDP growth are negatively correlated with the size of the government bond market. Eichengreen and Luengnaruemitchai (2004) and Claessens, Klingebiel, and Schmukler (2007) extend the BIS data analysis to a panel setting. Consistent with Burger and Warnock (2006), they find that country size, size of the banking system (measured as total deposits/GDP), good institutions, low inflation, and fiscal burden are positively correlated with the size of the domestic bond market. Borensztein and others (2008) distinguish between developments in government, corporate, and financial sector bonds, rather than considering them as one aggregate. They confirm that country size is significantly correlated with the size of bond markets, but this relationship is non-linear and, in the case of government bonds, their point estimates imply that the level off GDP that maximizes the size of the government bond market relative to GDP is US$6 trillion. They also show that other factors positively affecting domestic bond market activity include trade openness, public debt, institutional quality, lack of capital controls, and privatization of the pension system.

Contrary to the previous studies, Forslund, Lima, and Panizza (2011) find a much weaker association between macroeconomic and institutional factors and the share of domestic government debt in total debt for a wider sample of 95 countries, 33 of which are low income countries. The most puzzling finding is the insignificant impact of inflation history. The authors explain this result by the presence of capital controls, as this relationship turns negative and significant when a subsample of countries with moderate capital controls is considered.

Finally, in addition to studies on the FIB, a number of studies have examined the impact of purchases of government bonds by domestic central banks— including through quantitative easing policies—on long-term interest rates. These studies have focused on the United States, the United Kingdom, and Japan where such purchases have mainly taken place. The literature has focused on event studies and single country regressions. A recent IMF study (IMF 2013a) reviews this literature and finds that the cumulative effects of bond purchase programs were between 90 and 200 basis points in the U.S., between 45 and 160 basis points in the U.K. and between 10 and 30 basis points in Japan.

3. METHODOLOGY

We employ panel data methods to analyze the relationship between the FIB and long-term sovereign bond yields in AEs. Our empirical specification includes the standard macroeconomic determinants of long-term sovereign bond yields used in previous studies. We also control for the domestic central bank purchases of government debt. In addition, we introduce the FIB variable as an additional determinant of long-term sovereign bond yields. The empirical specification takes the following form:

$$y_{i,t}^{10Y} = \underbrace{\alpha_i + \beta_1 y_{i,t}^{2Y} + \beta_2 g_{i,t} + \beta_3 \pi_{i,t} + \beta_4 D_{i,t}}_{standard\ determinants\ of\ gov.\ bond\ yields} + \rho CB_{i,t} + \gamma FIB_{i,t} + \lambda_t + \varepsilon_{i,t} \quad (1)$$

where i and t indices denote country and time, respectively, y^{10Y} and y^{2Y} are the 10- and 2-year nominal government bond yields, respectively, g is the output growth (y-o-y), p is the CPI inflation (y-o-y), D is the debt-to-GDP ratio, CB is the share of domestic official (central bank) holdings of government debt in total, FIB is the foreign investor base variable, and e is the random error. Estimations are performed using the fixed effects estimator, with a_i capturing unobserved country-specific time-invariant determinants of long-term sovereign bond yields (e.g., institutional characteristics, political stability) and l_t capturing unobserved time-specific

DOI: 10.7172/2353-6845.jbfe.2016.2.3

common effects influencing all countries simultaneously (e.g., movements of capital between riskier equity and safer fixed income security markets in periods of financial stress).

The first four determinants (y^{2Y}, g, p, and D) are the standard determinants that were used for assessing the "fair value" of long-term sovereign bond yields in AEs (see Poghosyan, 2012, and references therein):

- **Short-term bond yields** (y^{2Y}) summarize the impact of monetary policy stance on long-term sovereign bond yields. The pass-through from the short-term rate is expected to be positive but less than 1, leaving room for other determinants of long-term sovereign bond yields to have an impact as well.
- **Output growth** (g) can have a positive or negative impact on long-term sovereign bond yields. On the one hand, an increase in output growth can be driven by a positive shift in potential output growth, which theoretically should have a positive effect on long-term sovereign bond yields as envisaged by the intertemporal utility maximization problem of a representative household (Laubach, 2009; Poghosyan, 2012). On the other hand, the increase may be cyclical and temporary, improving the tax capacity of the country, lowering the sovereign risk, and having a negative effect on bond yields (Cottarelli and Jaramillo, 2012). Which of these opposite effects prevails is an empirical question.
- **CPI inflation** (p) is expected to have a positive impact on long-term sovereign bond yields. According to the Fisher equation, an increase in expected inflation by one percentage point will lead to a commensurate increase in nominal long-term sovereign bond yields, all else equal, implying that the pass-through effect from expected inflation should be 1. However, in practice, it is difficult to come up with precise measures of inflation expectations and investors may not be totally rational. This implies that in practice the pass-through effect may be less than 1 (Caporale and Williams, 2002).
- **Debt-to-GDP ratio** (D) is expected to have a positive impact on long-term sovereign bond yields through two main channels. First, higher government debt crowds out private investments (assuming Ricardian equivalence does not hold) resulting in a lower stock and higher marginal product of capital in the steady state. Second, higher government debt may boost sovereign bond yields through a higher risk premium requested by investors (Engen and Hubbard, 2004).

In addition to standard determinants, we control for the impact of domestic central bank purchases of government debt by introducing a variable that measures the share of domestic official (central bank) holdings of government debt in total. The importance of central bank purchases of government securities on long-term government bond yields was evidenced in a series of recent papers studying the impact of quantitative easing policies launched by the central banks of major AEs in the aftermath of the crisis (Joyce et al., 2011; Krishnamurthy and Vissing-Jorgensen, 2012; and Ueda, 2012).

We augment these standard determinants by an additional variable to assess the impact of the FIB.[5] We use several FIB measures. First, we use the share of foreign bond holdings in total government debt as an overall measure of the FIB. Motivated by the "conundrum" argument put forward by Alan Greenspan, our prior is that an increase in the share of debt held by foreign investors would lead to a reduction of long-term sovereign bond yields. Second, we use subcomponents of foreign bond holdings to get a more refined view of channels through which the FIB affects long-term sovereign bond yields. For this reason, we use the following three non-overlapping and exhaustive subcomponents of the share of foreign bond holdings: (i) the share

[5] Changes in the FIB would occur as a result of *net* purchases of government bonds by non-resident investors (that is total purchases minus total sales and redemptions). As such, these transactions would be recorded as part of *gross capital inflows* in the balance of payments. It is important to note that this is different from *gross capital outflows* (i.e. residents buying foreign assets), or *net capital inflows* (i.e. gross capital inflows minus gross capital outflows). Movements in the exchange rate are likely to be sensitive to both capital inflows and outflows, but for our analysis of sovereign bond yields, we can restrict our attention to only transactions in the domestic bond market between non-resident and resident investors, rather than all portfolio decisions by residents and non-resident investors.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

of official foreign debt in total, (ii) the share of foreign bank debt in total, and (iii) the share of foreign non-bank debt in total.

The impact of foreign official bond holdings on long-term sovereign bond yields is ambiguous. On the one hand, the increase in foreign demand for long-term sovereign bonds will push the yields down. On the other hand, private sector bond holders may require an additional premium following large foreign official bond purchases, in recognition of the fact that official lenders (especially international financial organizations) have a more senior status than private bond holders.

The impact of foreign private bond holdings on long-term sovereign bond yields is expected to be negative. Overall, the negative impact of foreign private bond holdings may more than offset the ambiguous impact of official bond holdings, leading to an overall negative effect of total foreign debt holdings on long-term sovereign bond yields.

We also conduct several robustness checks. First, we use actual data (including smoothed) on macro variables rather than their expectations. Second, we use lagged values of independent variables to alleviate simultaneity issues. Finally, we replace time fixed effects with a crisis dummy and two measures of global uncertainty (VIX and the policy uncertainty index). The impact of the FIB on long-term sovereign bond yields remains qualitatively unchanged in all specifications.

4. DATA AND DESCRIPTIVE STATISTICS

Our sample covers 22 AEs that make up 98 percent of the general government debt of all AEs and can, therefore, provide a comprehensive view of the global demand for advanced economy sovereign debt. All data are compiled on a quarterly basis and cover the period from 2004 to 2012. Table 1 describes all variables, their sources, and the country coverage. Descriptive statistics are shown in Table 2.

The dependent variable is the nominal 10-year benchmark bond yield extracted from Bloomberg (daily average), which measures long-term government borrowing costs. The standard drivers of bond yields include the macroeconomic, monetary, and fiscal policy determinants used in previous studies. We capture the impact of monetary policy decisions on long-term bond yields through: (i) 2-year government bond yields (Bloomberg, daily averages), which are closely linked to the monetary policy stance,[6] and (ii) the share of domestic official holdings of government debt (Arslanalp and Tsuda, 2014), which capture the scale of quantitative easing policies in some AEs where the monetary policy rate hit the zero lower bound. The impact of macroeconomic factors is captured through y-o-y changes in the CPI index (International Financial Statistics) and real output (Haver Analytics), while the impact of fiscal policy is captured by the government debt ratio (Haver Analytics; Arslanalp and Tsuda, 2014).

For the outstanding amount of debt, we use a common definition of government debt—consolidated general government gross debt—to facilitate international comparability. General government debt covers the debt of central, local, and state governments, and social security funds. When it is consolidated, all intra-governmental holdings, such as central government debt held by social security funds, are netted out. All debt figures are expressed in face value and on a gross basis. For European Union (EU) countries, this definition matches the definition of "Maastricht debt," for which data are readily available from Eurostat's Quarterly Government Finance Statistics.[7] For others, they are constructed by Arslanalp and Tsuda (2014) from national flow of funds data using the same definition of government debt. The debt-to-GDP ratio is

[6] We use the 2-year rates as an indicator of not just current, but also expected, policy rates in the near term.

[7] Switzerland also provides government debt figures consistent with the definition of Maastricht debt. The data for Switzerland are on an annual basis, so quarterly figures are interpolated.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

calculated as the stock of debt in the referenced quarter divided by the 4-quarter moving sum of GDP. Quarterly GDP data are available from Eurostat for EU countries and Haver Analytics for other advanced economies.

Table 1
Description of Variables and Their Sources

Variable	Description	Expected sign	Source
Dependent variable			
Nominal long-term interest rate	Nominal 10 year benchmark bond yield (daily average)		Bloomberg*
Standard determinants			
Nominal short-term interest rate	Nominal 2 year bond yield (daily average)	(+)	Bloomberg
Real GDP growth (actual and expected)	Percentage change in y-o-y quarterly real output	(?)	Haver Analytics, Consensus Forecast
CPI inflation (actual and expected)	Percentage change in y-o-y quarterly CPI index (end of period)	(+)	IFS, Consensus Forecast
Debt ratio (actual and expected)	Ratio of general government debt to four-quarter moving sum of GDP (in percent)	(+)	Arslanalp and Tsuda (2012), Haver Analytics, IFS
Domestic official debt share	Ratio of domestic official debt to total government debt	(–)	Arslanalp and Tsuda (2012)
Foreign investor base			
Foreign debt share	Ratio of foreign debt to total government debt	(–)	Arslanalp and Tsuda (2012)
Foreign official debt share	Ratio of foreign official debt to total government debt	(+)	Arslanalp and Tsuda (2012)
Foreign bank debt share	Ratio of foreign bank debt to total government debt	(–)	Arslanalp and Tsuda (2012)
Foreign non-bank debt share	Ratio of foreign non-bank debt to total government debt	(–)	Arslanalp and Tsuda (2012)

Note: The sample covers the following advanced economies: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Korea, the Netherlands, New Zealand, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

* Data for Ireland comes from ECB's harmonized long-term interest rates warehouse.

For the foreign investor base of debt, we use the dataset constructed for advanced economies by Arslanalp and Tsuda (2014). This dataset has several advantages for the purpose of our analysis and provides a major improvement relative to the datasets used in previous cross-country studies. First, it is based on a common definition of government debt, as explained above (general government gross debt on a consolidated basis). Second, a common estimation methodology is used based on harmonized international data sources, such as the BIS International Banking Statistics (IBS), IMF International Financial Statistics (IFS), and IMF/World Bank Quarterly External Debt Statistics (QEDS). This ensures that all data are based on the same residency principle of the investor, include comparable definitions of general government, and use similar definitions of debt instruments. Third, all data are compiled either in face value or adjusted for valuation changes, where appropriate, to track investor transactions as well as holdings. This is important to eliminate any spurious correlation between long-term sovereign bond yields and

DOI: 10.7172/2353-6845.jbfe.2016.2.3

investor holdings. Finally, foreign investor holdings are estimated separately for the foreign official sector, foreign banks, and foreign nonbanks, in contrast to national data sources that usually classify them under one category ("rest of the world"). A more detailed description of the dataset, including stylized facts about recent trends (composition of foreign holdings, differences across countries, etc.) is provided in Arslanalp and Tsuda (2014).

Table 2
Descriptive Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
Nominal long-term interest rate	792	4.1	2.4	0.5	30.9
Nominal short-term interest rate	792	4.1	14.1	0.2	191.9
Real GDP growth	792	1.5	3.0	–10.1	8.7
CPI inflation	792	2.1	1.4	–6.1	7.4
Debt ratio	792	65.1	39.2	15.4	223.0
Doestic official debt share	792	3.5	4.4	0.0	27.6
Foreign debt share	792	41.1	21.1	3.2	83.3
Foreign official debt share	792	13.5	11.5	1.3	76.1
Foreign bank debt share	792	10.4	7.2	0.3	39.6
Foreign non-bank debt share	792	17.2	13.3	0.0	56.2
Foreign non-official (bank and non-bank) debt share	792	27.6	18.5	0.8	70.8

We use a number of variables for the robustness test of our main results. To capture the forward looking nature of markets we use projections of inflation and real GDP growth from Consensus Forecasts.[8] The projected ratios are compiled from different vintages of the IMF World Economic Outlook (WEO) database. For each vintage, we calculate the maximum debt-to-GDP ratio over the 5-year projection horizon, as a way to capture market concerns about the debt trajectory. The WEO database is updated semi-annually, so quarterly figures are interpolated. The interpolation assumes that debt-to-GDP projections change steadily over a six-month period for most of the countries in the sample.

Finally, the descriptive statistics suggest that data is quite heterogeneous in some cases, such as real growth rates. This can lead to misspecification of the fixed-effects model. Hence, we checked the estimates on a winsorized sample as well. The results remained qualitatively unchanged and are available from the authors upon request.

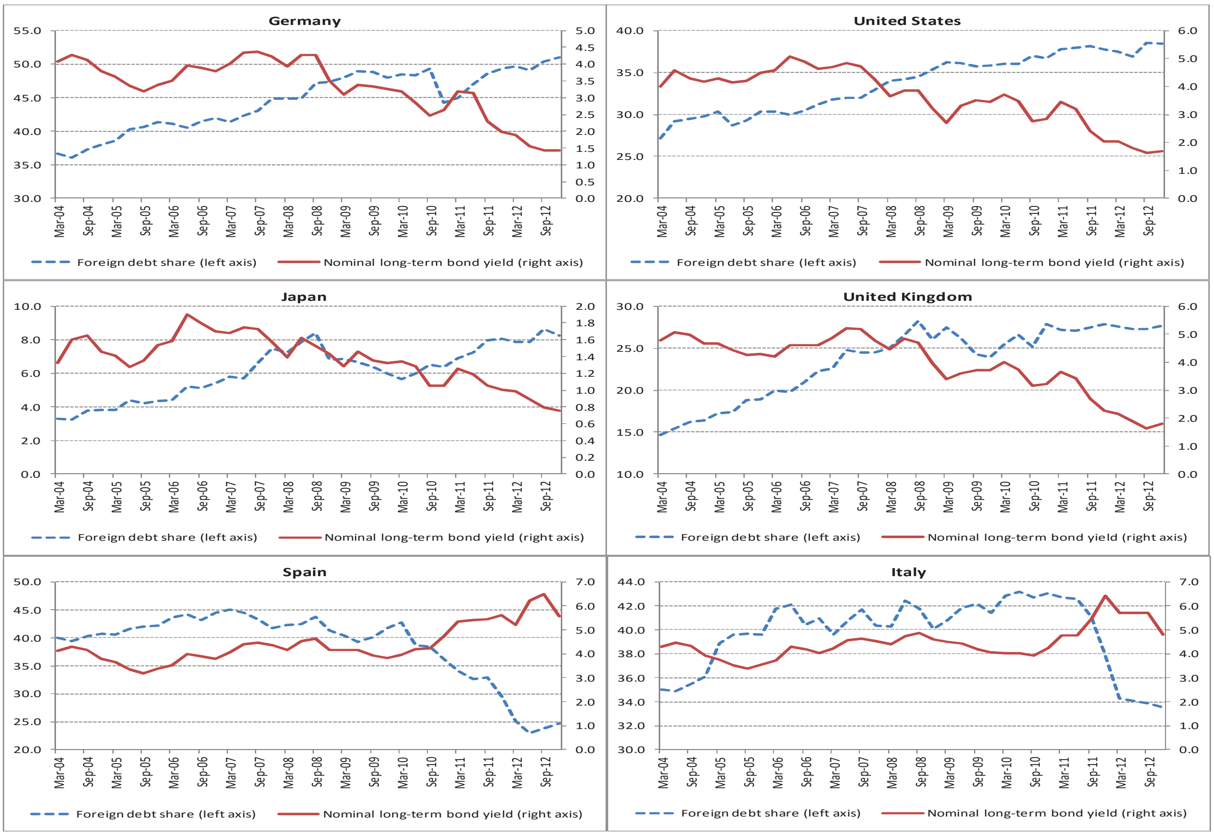
To illustrate our main hypothesis, we present the relationship between long-term government bond yields and the FIB in selected AEs (Figure 1). As shown in the figure, bond yields and the FIB move in the opposite direction, and this divergence became particularly pronounced following the crisis. The share of the FIB has been growing in "safe haven" recipient AEs, such as the United States, the United Kingdom, Germany, and to a lesser extent Japan, and this increase has coincided with a decline in bond yields in these countries. By contrast, the share of the FIB has dipped in periphery euro area countries, such as Spain and Italy, and this decline has coincided with a rapid increase in bond yields in these countries. It is also interesting to observe the high persistency of the FIB variable, suggesting that foreign investors adjust their holdings of sovereign securities gradually.[9]

8 Consensus Forecast provides projections for the current year and the following year. To construct a one-year ahead projection for each quarter, we took the weighted average of these two projections, where the weights were determined as follows: ¾ and ¼, respectively, for the first quarter; ½ and ½ for the second quarter, and so forth.

9 The autocorrelation coefficient of the foreign investor base variable is 0.995 for the first lag and gradually declines to 0.915 for the tenth lag.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Figure 1
Sovereign bond yields and foreign investor base in selected advanced economies



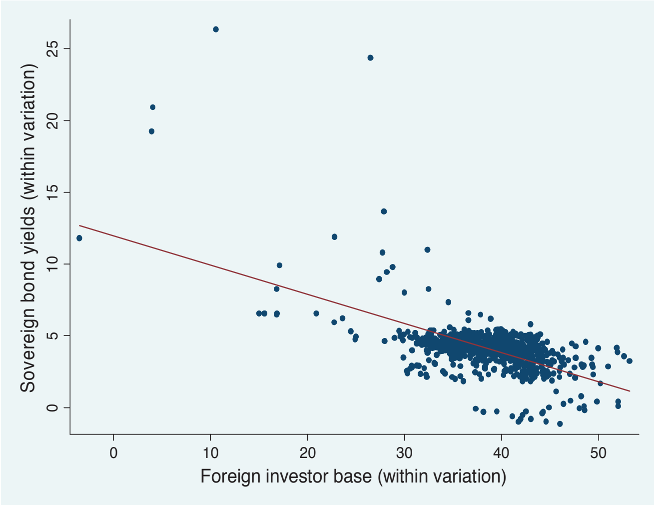
Note: The figure shows 10-year government bond yields and the share of government bond market held by foreigners.

Another way to illustrate the relationship between the FIB and sovereign bond yields is to juxtapose the within variation of the two variables in a scatter plot (Figure 2). The within transformation measures the difference between each data point of a variable from its country-specific mean and eliminates country-specific unobserved heterogeneity (fixed effects).[10] As shown in the figure, the within transformations of sovereign bond yields and the FIB are inversely related and this relationship is not driven by a few outliers. The inverse association suggests that an increase in a country's FIB over time was on average associated with a reduction in sovereign bond yields. Of course, the dynamics of bond yields also reflect the impact of other fundamental determinants (e.g., monetary policy rate, output growth, inflation) as well as central bank purchases of government debt, which we will take into account in our empirical analysis to separate the impact of the FIB from other factors.

10 More formally, the within transformation of variable X_{it} can be written as $(X_{it} - X_{i.})$, where $X_{i.}$ is the average for country i. In Figure 2, we have also added sample means $(X_{..})$ of the FIB and sovereign bond yields to their respective within transformations in order to move the scatter plot away from the axis origin.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Figure 2
Sovereign bond yields and foreign investor base (within variation)



4. ESTIMATION RESULTS

In this section, we discuss the results from the baseline specification. Then we present robustness checks. Finally, we assess the impact of changes in the FIB on sovereign bond yields in the aftermath of the crisis.

4.1. Baseline specification

Table 3 reports estimation results from the baseline specification.[11] Column 1 shows results from the specification with the foreign share of total government debt securities issued as a measure of the FIB. Column 2 shows results from the specification with the two subcomponents of the foreign debt share: official and non-official, respectively.

The standard determinants of sovereign bond yields have the right sign and are significant in both specifications. A 100 basis points increase in the nominal short-term bond yield leads to an increase of 8 basis points in long-term sovereign bond yields. This result is comparable to estimates found in studies on AEs that do not (e.g., Poghosyan, 2012) and do (e.g., Andritzky, 2012) consider the impact of the FIB. Given that the short-term sovereign bond yields are closely related to the monetary policy rate, this result suggests that monetary policy has a less than one-to-one pass-through effect on long-term sovereign bond yields, with the rest of the impact linked to changes in macroeconomic and fiscal factors.

In terms of macroeconomic determinants, bond yields are negatively affected by the expected real output growth rate and positively by the expected inflation rate. A 1 percent increase in the expected real growth rate leads to a 48–49 basis points decline in bond yields, while a 1 percent increase in expected inflation leads to an increase of 18–23 basis points in bond yields. The negative impact of growth is consistent with the findings of Baldacci and Kumar (2010) and can be explained by the stronger influence of cyclical (relative to structural) factors on growth in the relatively high frequency (quarterly) data used in the analysis. The smaller impact of inflation could be driven by the fact that, in AEs, inflation expectations have been firmly anchored at a low level, diminishing their importance for long-term investors.

[11] We have checked variables for stationarity. Although individual/country-specific ADF tests suggest that some variables are I(1), co-integration tests using Johansen procedure confirm co-integration for those variables.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Table 3
Baseline Specification: Total Sample

	1	2
Short-term bond yield (2 year)	0.0790***	0.0772***
	(25.666)	(24.212)
Real GDP growth (y-o-y) [Consensus Forecast]	–0.4960***	–0.4807***
	(–10.481)	(–10.058)
CPI inflation (y-o-y) [Consensus Forecast]	0.2318***	0.1819**
	(2.603)	(1.977)
Debt/GDP*100 [WEO projection, 5 year max]	0.0268***	0.0282***
	(8.120)	(8.392)
Domestic official debt share	–0.0537***	–0.0522***
	(–3.656)	(–3.553)
Foreign debt share	–0.0811***	
	(–10.625)	
Foreign official debt share		–0.0691***
		(–7.243)
Foreign non-official (bank and non-bank) debt share		–0.0851***
		(–10.837)
Constant	6.3002***	6.3030***
	(13.931)	(13.969)
Observations	792	792
R-squared	0.804	0.805
Number of countries	22	22
R-square adjusted	0.788	0.789
R-square overall	0.184	0.167

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator controlling for time effects (not reported).
Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

As for fiscal determinants, government debt has a positive and significant impact on bond yields. A 1 percentage point increase in the expected debt-to-GDP ratio leads to an increase of 3 basis points in bond yields. This result is comparable to the 2–7 basis points range found in studies that do not consider the impact of the FIB (see Poghosyan, 2012, and references therein). Domestic official debt purchases also have a significant impact on bond yields. A 1 percentage point increase in the share of debt held by the central banks contributes to a reduction of 5 basis points in bond yields. This suggests that central bank purchases of government bonds may have reduced bond yields by about 15 basis points in Japan, 120 basis points in the United Kingdom, and 30 basis points in the United States. With the exception of the United States, these are within the range of estimates found in other studies, including event studies.[12]

Finally, the FIB variable has a significant impact on bond yields in most specifications. This emphasizes the importance of the FIB as an additional determinant of bond yields in AEs. As expected, the coefficient has a negative sign: a 1 percentage point higher share of foreign debt in total leads to a decline in bond yields of 8 basis points. This result supports the "conundrum" effect, according to which a substantial increase in foreign demand for AE securities led to a compression of their long-term bond yields in the mid-2000s.

[12] The U.S. may be an exception in the results because it also benefits from significant *foreign* central bank purchases of its government bonds, which may have amplified the downward pressure on U.S. Treasury bond yields.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

However, the impact of the foreign debt share slightly differs across its subcomponents. A 1 percentage point rise in the foreign official share in total debt leads to a 7 basis points decrease in bond yields, while a 1 percentage point increase in the foreign private share in total debt leads to a decline in bond yields of 8.5 basis points. This supports the view that capital flows from foreign official investors are relatively more stable, with a long-term horizon and less commercial orientation, making them less sensitive to changes in market sentiment. However, these results should be interpreted with caution since the difference between the coefficients of FIB subcomponents is not statistically significant.

We have also estimated the baseline model for 11 euro area countries in the sample (Table 4). The main difference between these countries and the other AEs is that they share a single currency. As a result, risks stemming from fiscal variables are likely to be more pronounced in these countries compared to the rest of the sample. Indeed, the sensitivity of bond yields to the debt ratio is slightly higher in euro area countries (5 basis points). In addition, the impact of the domestic official investor base is stronger in these countries (19 basis points), which is consistent with the fact that ECB interventions in the aftermath of the crisis had a strong impact on sovereign bond yields in the euro area. However, the impact of the FIB is comparable to that of the total sample, providing evidence in support of the "conundrum" effect in euro area countries.[13]

Table 4
Baseline Specification: Euro Area Countries

	1	2
Short-term bond yield (2 year)	0.0737***	0.0732***
	(17.884)	(16.859)
Real GDP growth (y-o-y) [Consensus Forecast]	–0.5061***	–0.4938***
	(–5.649)	(–5.242)
CPI inflation (y-o-y) [Consensus Forecast]	0.5662***	0.5269**
	(2.868)	(2.419)
Debt/GDP*100 [WEO projection, 5 year max]	0.0506***	0.0511***
	(5.853)	(5.844)
Domestic official debt share	–0.1895*	–0.1946*
	(–1.829)	(–1.864)
Foreign debt share	–0.0741***	
	(–5.329)	
Foreign official debt share		–0.0696***
		(–3.997)
Foreign non-official (bank and non-bank) debt share		–0.0748***
		(–5.339)
Constant	4.6796***	4.6938***
	(3.562)	(3.567)
Observations	396	396
R-squared	0.849	0.850
Number of countries	11	11
R-square adjusted	0.827	0.827
R-square overall	0.681	0.681

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator controlling for time effects (not reported).
Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

13 Moreover, several studies, such as Acharya and Steffen (2013), find that increasing "home bias"—greater exposure of domestic banks to sovereign bonds—after the European banking crisis may have played a role in pushing down bond yields in periphery countries. All else equal, that would suggest that the impact of foreign outflows from euro area periphery countries on bond yields may have been even higher than suggested in Table 4.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

4.2. Robustness Checks

We conduct several robustness checks to assess the sensitivity of our key findings to various assumptions. First, we replace expectations of macro and fiscal variables with actual data (Table 5). The magnitude of coefficients on macro and fiscal variables has been affected by this replacement, providing support to the argument advanced in Engen and Hubbard (2004) and Laubach (2009). According to this argument, in the presence of forward-looking market participants, sovereign borrowing costs depend on expected rather than actual macro and fiscal determinants and using expectations of determinants helps to disentangle the effect of fiscal policy from other factors influenced by the business cycle. Nevertheless, the impact of the FIB is not affected when using actual data, suggesting that our results on the importance of the FIB are not sensitive to this assumption.

Table 5
Robustness Check: Using Actual Data, Instead of Expectations

	1	2
Short-term bond yield (2 year)	0.0886***	0.0852***
	(15.997)	(11.967)
Real GDP growth (y-o-y)	–0.1663**	–0.1617**
	(–2.186)	(–2.272)
CPI inflation (y-o-y)	0.0828	0.0291
	(1.561)	(0.613)
Debt/GDP*100	0.0552***	0.0591***
	(4.108)	(3.914)
Domestic official debt share	–0.0685*	–0.0668**
	(–1.899)	(–2.092)
Foreign debt share	–0.0771***	
	(–4.050)	
Foreign official debt share		–0.0548**
		(–2.282)
Foreign non-official (bank and non-bank) debt share		–0.0849***
		(–4.326)
Constant	4.0049***	3.9255***
	(3.770)	(3.480)
Observations	792	792
R-squared	0.816	0.820
Number of countries	22	22
R-square adjusted	0.806	0.810
R-square overall	0.135	0.114

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator controlling for time effects (not reported).
Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

Second, we follow the approach by Correia-Nunes and Stemitsiotis (1995) and use smoothed values of actual macro and fiscal variables, instead of market analyst expectations and WEO projections (Table 6).[14] Using this approach leads to slightly different coefficient estimates compared to the baseline, but the impact of the FIB and its components remains unchanged.

[14] The smoothing is performed using MA (4,1,4) representation: $1/9*(x_{t-4}+x_{t-3}+x_{t-2}+x_{t-1}+x_t+x_{t+1}+x_{t+2}+x_{t+3}+x_{t+4})$, where x_t is the macro and fiscal variable of interest (real GDP growth, inflation, and debt-to-GDP ratio).

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Table 6
Robustness Check: Using Smoothed Values of Macro Variables, Instead of Expectations

	1	2
Short-term bond yield (2 year)	0.0827***	0.0802***
	(12.630)	(10.026)
Real GDP growth (y-o-y) [smoothed]	–0.3418**	–0.3319**
	(–2.457)	(–2.474)
CPI inflation (y-o-y) [smoothed]	0.2066	0.1416
	(1.634)	(1.083)
Debt/GDP*100 [smoothed]	0.0551***	0.0577***
	(4.382)	(4.102)
Domestic official debt share	–0.0762**	–0.0734**
	(–2.164)	(–2.338)
Foreign debt share	–0.0688***	
	(–3.492)	
Foreign official debt share		–0.0540**
		(–2.190)
Foreign non-official (bank and non-bank) debt share		–0.0750***
		(–3.482)
Constant	4.0042***	4.0304***
	(3.652)	(3.569)
Observations	792	792
R-squared	0.825	0.826
Number of countries	22	22
R-square adjusted	0.815	0.817
R-square overall	0.161	0.141

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator controlling for time effects (not reported).
Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

Third, we use lagged dependent variables for standard determinants of bond yields (Table 7). The main motivation is to alleviate the possible simultaneity between sovereign borrowing costs and the macroeconomic environment. Lagging standard determinants of bond yields slightly altered their magnitude and significance compared to the baseline. However, the sign and the significance of the FIB variables have remained unchanged. The sensitivity of sovereign bond yields to the foreign debt share is slightly higher in this specification (10 basis points), which is mostly due to the larger impact of the non-official foreign debt (11 basis points).

Fourth, we assess the sensitivity of the results to the inclusion of the crisis dummy and observable global factors, instead of time-fixed effects capturing unobserved heterogeneity. The crisis dummy or global risk aversion indicators could capture the impact of changing preferences or risk-appetite of foreign (and domestic) investors during the crisis period. Table 8 shows estimation results using the VIX index as a measure of global risk aversion. The impact of the FIB and standard determinants remain unchanged in this specification. As expected, bond yields tend to increase with rising global risk aversion (2 basis points per unit increase in the VIX). In addition, sovereign bond yields in all AEs have declined by 117–125 basis points in the aftermath of the crisis.[15] This decline could be driven by the shift of capital from riskier equity to safer fixed income securities markets following the crisis. Table 9 shows estimation results using the

[15] The results are robust to the inclusion of the crisis dummy and VIX variables one at a time. We also included an interaction term between the crisis dummy and domestic official debt share variable to address the issue that central bank bond purchases may have been more powerful during the crisis. The results for the FIB remain robust to these changes in the model specification.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

news-based economic policy uncertainty index by Bloom (2009) and Baker, Bloom, and Davis (2013).[16] Once again, the impact of the FIB and standard determinants remains unaffected. The crisis dummy coefficient is slightly lower (54–59 basis points), while the policy uncertainty index is negative and significant. The latter suggests that policy uncertainty triggers capital outflows from riskier equity to safer fixed income markets.

Table 7
Robustness Check: Using Lagged Independent Variables

	1	2
Short-term bond yield (2 year) [lagged]	0.0640***	0.0584***
	(4.654)	(3.657)
Real GDP growth (y-o-y) [Consensus Forecast, lagged]	–0.5308***	–0.4978***
	(–2.923)	(–3.647)
CPI inflation (y-o-y) [Consensus Forecast, lagged]	0.0259	–0.1123
	(0.175)	(–0.535)
Debt/GDP*100 [WEO projection, 5 year max, lagged]	0.0270**	0.0316**
	(2.364)	(2.432)
Domestic official debt share [lagged]	–0.0577*	–0.0526*
	(–1.820)	(–1.854)
Foreign debt share	–0.1043***	
	(–5.447)	
Foreign official debt share		–0.0666**
		(–2.610)
Foreign non-official (bank and non-bank) debt share		–0.1148***
		(–4.978)
Constant	7.8774***	7.8068***
	(7.791)	(7.708)
Observations	770	770
R-squared	0.713	0.723
Number of countries	22	22
R-square adjusted	0.697	0.708
R-square overall	0.0696	0.0473

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator controlling for time effects (not reported).
Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

Fifth, we examine whether the FIB may be driven by the same macro variables driving bond yields. For this, we look for potential signs of multicollinearity among independent variables. The correlation matrix for independent variables shows that bilateral correlations are not very high, including between the FIB and macro variables (Table 10). The highest correlation (0.86) is between the foreign share and the foreign non-official share of government debt, but we include these variables separately in the regressions. Small correlation coefficients do not support the hypothesis that the FIB may be driven by the same macro variables driving bond yields, rather than having its own impact on bond yields.

16 The index of economic policy uncertainty refers to uncertainty surrounding economic policies in the United States and euro area and is a weighted average of three indicators: the frequency with which terms like "economic policy" and "uncertainty" appear together in the media; the number of tax provisions that will expire in coming years; and the dispersion of forecasts of future government outlays and inflation.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Table 8
Robustness Check: Replacing TE with VIX and Crisis Dummy

	1	2
Short-term bond yield (2 year)	0.0852***	0.0835***
	(9.305)	(7.693)
Real GDP growth (y-o-y) [Consensus Forecast]	–0.2838**	–0.2770**
	(–2.166)	(–2.291)
CPI inflation (y-o-y) [Consensus Forecast]	0.4013*	0.3596**
	(1.995)	(2.284)
Debt/GDP*100 [WEO projection, 5 year max]	0.0294**	0.0298**
	(2.383)	(2.367)
Domestic official debt share	–0.0786*	–0.0767**
	(–2.018)	(–2.121)
Foreign debt share	–0.0682***	
	(–4.041)	
Foreign official debt share		–0.0587**
		(–2.611)
Foreign non-official (bank and non-bank) debt share		–0.0724***
		(–4.049)
Crisis dummy (=1 for 2008Q3 onwards)	–1.1738***	–1.2488***
	(–3.732)	(–3.333)
VIX	0.0201***	0.0207***
	(3.120)	(3.213)
Constant	4.4929***	4.5502***
	(3.803)	(3.946)
Observations	792	792
R-squared	0.721	0.722
Number of countries	22	22
R-square adjusted	0.719	0.719
R-square overall	0.200	0.185

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator. Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

Finally, we run country-specific Granger causality tests to see which way the causality between the FIB and bond yields is likely to run. The tests suggest that for the vast majority of countries, the causality runs from the FIB to bond yields, and not vice versa (Table 11). Exceptions are Greece, Ireland, Italy, and Portugal, i.e. only some of the euro area periphery countries (or 4 countries out of 22 in the sample).[17] For these countries, Granger causality tests suggest that the FIB may have reacted to rising bond yields (e.g., presumably, foreign investors cut exposure to these countries after taking large losses due to sharp rises in bond yields).

[17] This causality test is rejected consistently for these countries using lag periods up to 3 quarters. The test is rejected for Spain (3 lags), Switzerland (3 lags), and Sweden (1 and 2 lags) in only some specifications.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Table 9
Robustness Check: Replacing TE with Policy Uncertainty Index and Crisis Dummy

	1	2
Short-term bond yield (2 year)	0.0849***	0.0832***
	(9.334)	(7.770)
Real GDP growth (y-o-y) [Consensus Forecast]	–0.3552**	–0.3505***
	(–2.749)	(–2.868)
CPI inflation (y-o-y) [Consensus Forecast]	0.4378**	0.3972**
	(2.273)	(2.625)
Debt/GDP*100 [WEO projection, 5 year max]	0.0267**	0.0271**
	(2.390)	(2.388)
Domestic official debt share [lagged]	–0.0829**	–0.0811**
	(–2.273)	(–2.398)
Foreign debt share	–0.0647***	
	(–3.657)	
Foreign official debt share		–0.0551**
		(–2.308)
Foreign non-official (bank and non-bank) debt share		–0.0688***
		(–3.700)
Crisis dummy (=1 for 2008Q3 onwards)	–0.5419*	–0.5965*
	(–1.859)	(–1.804)
Policy uncertainty index (Bloom, 2009; Baker et al, 2013)	–0.0061***	–0.0063***
	(–3.521)	(–3.472)
Constant	5.4983***	5.5873***
	(5.400)	(5.778)
Observations	792	792
R-squared	0.721	0.722
Number of countries	22	22
R-square adjusted	0.718	0.719
R-square overall	0.230	0.215

Note: The dependent variable is the nominal long-term bond yield (10 year). Estimations are performed using the fixed effects estimator. Robust t-statistics are in parentheses. ***, **, and * denote significance at 1, 5, and 10 percent level, respectively.

Table 10
Correlation Matrix of Independent Variables

	Long-term bond yield (2 years)	Real GDP growth	CPI inflation	Debt to GDP ratio	Domestic official debt share	Foreign debt share	Foreign official debt share	Foreign non-official debt share
Long-term bond yield (2 years)	1.00							
Real GDP growth	–0.27	1.00						
CPI inflation	–0.01	0.48	1.00					
Debt to GDP ratio	0.15	–0.43	–0.55	1.00				
Domestic official debt share	0.03	0.09	0.15	0.20	1.00			
Foreign official debt share	–0.04	–0.20	0.08	0.00	–0.28	1.00		
Foreign debt share	0.01	–0.11	0.03	–0.08	–0.14	0.44	1.00	
Foreign non-official debt share	–0.04	–0.17	0.08	0.04	–0.23	0.86	–0.09	1.00

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Table 11
Granger Causality Tests for Individual Countries

	Foreign investor base does not cause interest rates?	Interest rates do not cause foreign investor base?
United States	0.00	0.12
United Kingdom	0.12	0.67
Austria	0.07	0.61
Belgium	0.03	0.53
Denmark	0.01	0.35
France	0.47	0.89
Germany	0.44	0.74
Italy	0.03	0.01
Netherlands	0.77	0.07
Sweden	0.86	0.04
Switzerland	0.14	0.10
Canada	0.04	0.26
Japan	0.02	0.78
Finland	0.00	0.62
Greece	0.03	0.01
Ireland	0.96	0.00
Portugal	0.03	0.00
Spain	0.26	0.15
Australia	0.00	0.74
New Zealand	0.00	0.35
Korea	0.35	0.76
Czech Republic	0.04	0.78

Note: Country-specific Granger causality tests were run using 2 lags. Reported are p-values. P-values below 0.05 (highlighted in red) indicate rejection of the hypothesis stated on top of the column at 5 percent confidence level.

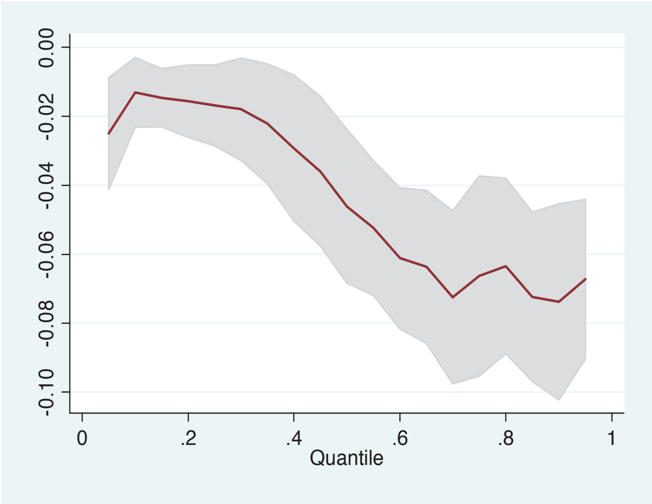
4.3. Quantile Regression

The panel fixed effects regressions discussed above assume that the impact of the FIB variable on bond yields is the same across all countries. To relax this assumption, we also estimate specification (1) using quantile regression. The quantile regression explicitly allows the impact of the FIB variable to vary across different quantiles of bond yields. As shown in Figure 3, the impact of the FIB on bond yields remains negative and significantly different from zero across all quantiles of bond yields. However, the negative impact tends to be larger in the upper quantile (up to –0.07), suggesting that the FIB has a stronger negative impact on bond yields when the level of interest rates is high. By contrast, the impact tends to be smaller in lower quantiles (up to –0.02), suggesting a smaller impact of FIB on bond yields when the level of interest rates is low.

These results are in line with models of creditor discrimination (Broner et al., 2013). In these models, creditor discrimination arises because, in turbulent times, sovereign debt offers a higher expected return to domestic creditors than to foreign ones. This provides incentives for domestic purchases of debt. In the context of recent developments, these results suggest that the recent outflow of foreign investment from the periphery countries may have had a larger impact on bond yields than the "safe haven" inflow of foreign funds to core AEs.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Figure 3
Quantile regression: coefficient of the foreign investor base variable



4.4. Assessing the Impact of Changes in the Foreign Investor Base

In this section, we assess the contribution of the FIB on bond yields, drawing on the results of the regression analysis.

Our regression results suggest that a 1 percentage point increase in the share of government debt held by foreigners is associated with a reduction in 10-year bond yields of about 8 basis points for a panel of 22 advanced economies. The impact remains qualitatively unchanged (within a range of 6.5 to 10.4 basis points) and statistically significant in all specifications. These estimates are in line with Warnock and Warnock (2009), whose corresponding estimate would be around 7 basis points for the United States.[18] At the same time, they are somewhat higher than Andritzky (2012), whose estimates are around 6 basis points for euro area countries and 3–4 basis points for non-euro G-20 advanced economies.

Table 12 translates these estimates into contributions of the FIB to the changes in the long-term bond yields of major core and periphery countries. It shows that foreign inflows to bond markets may have reduced long-term rates by 35–60 basis points in the United States, 20–30 basis points in the United Kingdom, and 40–65 basis points in Germany since 2007. In contrast, foreign outflows from Italy and Spanish government bonds may have raised long-term yields by 40–70 and 110–180 basis points in these countries, respectively.

Table 12
Impact of Foreign Investor Base (FIB) on Government Bond Yields, 2008–12

Country	Change in foreign ownership of government debt 1/	Coefficient of foreign Investor base		Contribution to change in yields (bps) 1/	
	(in percentage points)	Low	High	Low	High
Germany	6.1	–0.065	–0.104	–40	–64
Italy	–6.8	–0.065	–0.104	44	71
Spain	–17.1	–0.065	–0.104	110	178
United Kingdom	3.1	–0.065	–0.104	–20	–33
United States	5.4	–0.065	–0.104	–35	–57

Note: * change in ownership from end-2007 to end-2012.

18 Warnock and Warnock (2009) find that foreign flows into the U.S. Treasury market in the amount of 1 percent of GDP are associated with a 19 basis point reduction in long-term rates. This would correspond to a 2.7 percentage point increase in foreign ownership of U.S. Treasuries, based on figures from 2005 when their study ended.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

The implication of these estimates is that "flight to safety" flows in the form of foreign purchases of U.S., U.K., and German bonds after the global financial crisis may have made a substantial contribution to the decline in long-term interest rates in these countries. At the same time, foreign outflows from periphery countries can explain a substantial amount of the rise in their bond yields, in addition to the deterioration in their macroeconomic fundamentals immediately after the crisis.

Put differently, under a scenario in which the FIB normalizes to its pre-crisis level at end-2007, long-term bond yields may rise by 20–65 basis points in major core economies and decline by 40–180 basis points in major periphery countries. These are substantial amounts and highlight the important role of foreign investors in determining long-term bond yields even as the macroeconomic environment normalizes.

5. CONCLUSIONS

This paper analyzes the impact of the foreign investor flows on sovereign bond yields of 22 AEs. Our analysis suggests that an increase in foreign ownership is associated with a statistically and economically significant decline in long-term bond yields. In particular, we find that a 1 percentage point increase in the share of government debt held by non-residents can account for a 6–10 basis point decrease in 10-year government bond yields across advanced economies, controlling for other determinants for interest rates. This result is consistent with the "conundrum" phenomenon highlighted by U.S. policymakers in the early 2000s.

Moreover, country-specific Granger causality tests suggest that, for the vast majority of countries, the causality runs from foreign investor base to bond yields, and not vice versa. Exceptions are Greece, Ireland, Italy, and Portugal, i.e. only some of the periphery countries (or 4 out of 22 countries in the sample). For these countries, Granger causality tests suggest that foreign investors may have reacted to rising bond yields instead.

Overall, our results suggest that changes in the foreign investor base for sovereign debt can have economically and statistically significant effects on sovereign bond yields, independent of other standard macroeconomic determinants of bond yields.

Acknowledgements

The authors would like to thank to Luc Everaert, Abdelhak Senhadji, Ali Abbas, Ruchir Agarwal, Jochen Andritzky, Sergei Antoshin, Andrij Blokhin, Francesco Columba, Heiko Hesse, Bradley Jones, Luc Laeven, Prachi Mishra, Sergejs Saksonovs, Tahsin Saadi Sedik, Anita Tuladhar, and an anonymous referee for helpful comments and discussion.

References

Acharya, V. and Steffen, S. (2013), "The Greatest Carry Trade Ever? Understanding Eurozone Bank Risks" NBER Working Paper No. 19039, May 2013.

Andritzky, J. (2012). "Government Bonds and Their Investors: What Are the Facts and Do They Matter?", IMF Working Paper WP/12/158.

Arslanalp, S. and Tsuda, T. (2014). "Tracking Global Demand for Advanced Economy Sovereign Debt." *IMF Economic Review*, Vol. 62, No. 3, International Monetary Fund, Washington, DC.

Baker, S., Bloom, N., and Davis, S. (2013). "Measuring Economic Policy Uncertainty", Stanford University Working Paper, available at: http://www.policyuncertainty.com/papers.html.

Bernanke, B. (2005). "The Global Saving Glut and the U.S. Current Account Deficit", The Homer Jones Lecture St. Louis, Missouri.

Bernanke, B. (2013). "Long-Term Interest Rates", Remarks at the Annual Monetary and Macroeconomics Conference: The Past and Future of Monetary Policy, sponsored by the Federal Reserve Bank of San Francisco, San Francisco, California, March 1, 2013
Bernanke, B., Reinhart, V., Sack, B. (2004). "Monetary Policy Alternatives at the Zero Bound: An Empirical Assessment", *Brookings Papers on Economic Activity*, 35: pp. 1–100.
Beltran, D., Kretchmer, M., Marquez, J., and Thomas, C. (2012). "Foreign Holdings of U.S. Treasuries and U.S. Treasury Yields", Federal Reserve Board, Washington D.C.
Bloom, N. (2009). "The Impact of Uncertainty Shocks", *Econometrica*, 77 (3): pp. 623–85.
Borensztein, E., Cowan, K., Eichengreen, B., Panizza, U. (2008). "Prospects for Latin American Bond Markets: A Cross-Country View", in: Borensztein, Cowan, Eichengreen, Panizza (Eds.), Bond Markets in Latin America: On the Verge of a Big Bang? MIT Press, Cambridge, MA.
Broner, F., A. Erce, A. Martin, and Ventura, J. (2013). "Sovereign Debt Markets in Turbulent Times: Creditor Discrimination and Crowding-Out Effects," NBER Working Paper Series No. 19676.
Burger, J., and Warnock, F. (2006). "Local Currency Bond Markets", *IMF Staff Papers*, 53: pp. 133–146.
Caporale, G. and Williams, G. (2002). "Long-Term Nominal Interest Rates and Domestic Fundamentals", *Review of Financial Economics,* 15: pp. 370–413.
Claessens, S., Klingebiel, D., and Schmuckler, S. (2007). "Government Bonds in Domestic and Foreign Currency: The Role of Macroeconomic and Institutional Factors", *Review of International Economics*, 15: pp. 370–413.
Correia-Nunes, J. and Stemitsiotis, L. (1995). "Budget Deficit and Interest Rates: Is There a Link? International Evidence", *Oxford Bulletin of Economics and Statistics,* 57(4): pp. 425–49.
Cottarelli, C. and Jaramillo, L. (2012). "Walking Hand in Hand: Fiscal Policy and Growth in Advanced Economies", IMF Working Paper WP/12/137.
Dell'Erba, S., Hausmann, R., and Panizza, U. (2013). "Debt Levels, Debt Composition, and Sovereign Spreads in Emerging and Advanced Economies", *Oxford Review of Economic Policy (forthcoming)*.
Eichengreen, B., and Luengnaruemitchai, P. (2004). "Why Doesn't Asia Have Bigger Bond Markets?", NBER Working Paper 10576. National Bureau of Economic Research, Cambridge, MA.
Engen, E. and Hubbard, G. (2004). "Federal Government Debt and Interest Rates", *NBER Macroeconomic Annual*, 19: pp. 83–138.
Forslund, K., Lima, L., and Panizza, U. (2011). "The Determinants of the Composition of Public Debt in Developing and Emerging Market Countries", *Review of Development Finance*, 1: pp. 207–222.
Hauner, D. and Kumar, M. (2006), "Fiscal Policy and Interest Rates--How Sustainable Is the New Economy?" IMF Working Paper No. 06/112, May 2006.
Hördahl, P., Tristani, O. and Vestin, D. (2006). "A Joint Econometric Model of Macroeconomic and Term-Structure Dynamics", *Journal of Econometrics*, 131 (1–2): pp. 405–444.
Ichiue, H. and Shimizu, Y. (2012). "Determinants of Long-term Yields: A Panel Data Analysis of Major Countries and Decomposition of Yields of Japan and the U.S.", Bank of Japan Working Paper, No. 12-E 7.
IMF (2011). Global Financial Stability Report, October 2011, IMF, Washington D.C.
IMF (2012a). Fiscal Monitor, October 2012, IMF, Washington D.C.
IMF (2012b). Global Financial Stability Report, October 2012, IMF, Washington D.C.
IMF (2013a). "Unconventional Monetary Policies—Recent Experience and Prospects," April 2013, IMF, Washington D.C.
IMF (2013b). "Staff Guidance Note for Public Debt Sustainability Analysis in Market-Access Countries," May 2013, IMF, Washington D.C. (available at: http://www.imf.org/external/pubs/ft/dsa/mac.htm).
IMF (2013c). Fiscal Monitor, April 2013, IMF, Washington D.C.
Joyce, M., Lasaosa, A., Stevens, I., and Tong, M. (2011). "The Financial Market Impact of Quantitative Easing in the United Kingdom," *International Journal of Central Banking*, 7(3): pp. 113–161.
Jaramillo, L., and Zhang, Y. (2013). "Real Money Investors and Sovereign Bond Yields", IMF Working Paper No. 13/254, December 2013
Kaminska, I., Vayanos D., and Zinna, G. (2011) "Preferred-habitat investors and the US term structure of real rates", Bank of England Working Paper No. 435.
Krishnamurthy, A. and Vissing-Jorgensen, A. (2011). "The Effects of Quantitative Easing on Interest Rates," *Brooking Papers on Economic Activity*, 43(2): pp. 215–287.
Krishnamurthy, A., and Vissing-Jorgensen, A. (2012). "The Aggregate Demand for Treasury Debt," *Journal of Political Economy*, 2012, vol. 120, no. 2.
Kumar, M., and Baldacci E. (2010), "Fiscal Deficits, Public Debt, and Sovereign Bond Yields" IMF Working Paper No. 10/184, August 2010.
Lam, R. (2013). "What Would Happen to Long-Term Interest Rates in Japan Under "Abenomics"?" paper presented at the IMF Surveillance Meeting, April 30, Washington, D.C.
Laubach, T. (2009). "New Evidence on the Interest Rate Effects of Budget Deficits and Debt", *Journal of European Economic Association*, 7: pp. 858–85.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Mehl, A. and Reynaud, J. (2010). "Risky Public Domestic Debt Composition in Emerging Economies", *Journal of International Money and Finance*, 29: pp. 1–18.

Panizza, U. (2008). "Domestic and External Public Debt in Developing Countries", UNCTAD Discussion Paper. United Nations Conference on Trade and Development, Geneva.

Peiris, S. (2010). "Foreign Participation in Emerging Markets' Local Currency Bond Markets", IMF Working paper WP/10/88.

Poghosyan, T. (2012). "Long-Run and Short-Run Determinants of Sovereign Bond Yields in Advanced Economies" IMF Working Paper No. 12/271.

Rudebusch, G., Swanson, E., and Wu, T. (2006). "The Bond Yield "Conundrum« from a Macro-Finance Perspective«, *Monetary and Economic Studies*, 24(S1): pp. 83–109.

Ueda, K. (2012). "The Effectiveness of Non-traditional Monetary Policy Measures: the Case of the Bank of Japan," *The Japanese Economic Review*, 63(1): pp. 1–22.

Warnock F., and Warnock, V. (2009). "International Capital Flows and U.S. Interest Rates," *Journal of International Money and Finance*, 28(6): pp. 903–919.

DOI: 10.7172/2353-6845.jbfe.2016.2.3

Financial Inclusion, Growth and Inequality: A Model Application to Colombia

Izabela Karpowicz[1]
International Monetary Fund, United States of America
ikarpowicz@imf.org

Received: 13 January 2016 / Revised: 29 April 2016 / Accepted: 25 May 2016 / Published online: 16 June 2016

ABSTRACT

Financial inclusion has been one of the key pillars of Colombia's development strategy for a number of years. Financial inclusion policies have aimed at channeling microcredit to poor, spreading formal banking system usage, fostering electronic payment acceptance, and making financial services more affordable. Using simulations from a general equilibrium model it is possible to identify the most binding financial sector frictions that preclude financial inclusion of enterprises, and study the effects on growth and inequality of efforts to remove these frictions. The study finds that lowering contraints on collateral promises higher growth while inequality is better tackled through measures that lower the financial participation cost.

JEL classification: G2, G21, G28, O16

Keywords: Financial deepening, financial inclusion, access to finance, inequality

1. INTRODUCTION

While delivering strong economic growth is most policymakers' concern, inequality and financial inclusion have been Colombia's foremost preoccupations over the past several years. The government has invested efforts and resources into eliminating constraints to access to financial services and increasing efficiency, depth and breath of financial instruments. On the supply side there have been substantive improvements in physical infrastructure, regulatory framework and costs, while demand constraints were adressed by targeting financial literacy. Frictions were identified from the perspective of households, firms and banks, addressed, measured, and reported, making government's initiatives focused and transparent, and progress measurable.

The potential effect of financial inclusion efforts on growth in Colombia have not been studied, and neither has their implication for income inequality. This paper attempts to fill this literature gap by analyzing the state of financial inclusion in Colombia and the link between reforms implemented mainly on the micro side and their longer-term macroeconomic consequences. The model used is borrowed from Dabla-Norris et al. (2014).

The findings suggest that relaxing collateral requirements precluding greater financial sector inclusion promises higher growth while inequality is better tackled through measures that lower the financial participation cost. This result is important inasmuch as efforts to address inequality

[1] Corresponding author Izabela Karpowicz. Email:ikarpowicz@imf.org. Tel. (+1) 202-623-7007. International Monetary Fund, Western Hemisphere Department, 700 19th street NW, Washington D.C., USA
The views expressed in this paper are those of the author(s) and do not necessarily represent the views of the IMF, its Executive Board, or IMF management.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

through financial sector policy are called to complement those aimed at eliminating distortions in Colombia's fiscal policy framework, that have hindered a wider redistribution of economic gains.[2]

The paper is organized as follows. Section 2 presents a literature review; section 3 describes the state of financial inclusion in Colombia, section 4 identifies obstacles precluding greater financial inclusion and takes stock of authorities' efforts to eliminate them; section 5 presents the financial deepening model applied to Colombia and discusses model outcomes; and section 6 concludes with policy recommendations.

2. LITERATURE REVIEW

In line with governments' and private sector's efforts to embrace a larger share of population into the financial sector, by increasing access as well as effective usage of formal financial services, the literature measuring financial inclusion has bloomed in recent years. Hohonan (2007) and Sarma (2008), for instance, have used the access to financial services as a measure of financial inclusion. Roja-Suarez and Amado (2014), on the other hand, used the percent of people who have an account at a formal financial institution as a proxy to measure financial inclusion. However, as the World Bank Global Financial Inclusion dataset (Findex) become available in 2012, recording in great detail how people in 148 countries save, borrow, and make payments, the literature on usage of financial services has also expanded. Camara and Tuesta (2014), for instance, complemented the work of Roja-Suarez and Amado (2014) by constructing a composite index that included both the access and the usage of financial services. Dabla-Norris et al (2015) contribute to this stream of literature by constructing a composite financial inclusion index, which comprises households' as well as SME's access, that they use to gauge determinants of financial inclusion and financial inclusion gaps in Latin America following Suraez (2014).

The existing literature focuses, however, mainly on financial inclusion of households, while studies on financial access of small and medium size enterprises (SMEs) remain scarce. Moreover, whereas informal finance is prevalent, as is the case of many Latin America countries, studies on informal finance have mainly focused on its contribution to firm growth (Aiyagari et al., 2010 on China) and its relevance for households consumption smoothing (Nigeria, 2015 on Nigeria; and Townsend and Alem, 2014 on Thailand), while determinants of use are less frequently found, with a notable exception of Klapper and Singer (2015) Findex-based study on Africa.

The literature on the link between financial development and growth and the relationship between financial development and poverty alleviation predates the studies on households access to finance. King and Levine (1993) and Levine (2005) showed in an empirical framework that financial deepening spurs growth. Aggregate financial depth has also been linked to poverty reduction and income inequality in Beck et al. (2007) and Clarke et al. (2006). In the specific case of firms, access to finance has been positively associated with innovation, job creation, and growth (Beck et al., 2005 and Aiyagari et al., 2008). However, establishing causality and evaluating policies in a regression framework has proven challenging. Recent papers, such as Moll et al. (2014) and Blaum (2013) have used quantitative models whose structural framework allows for a normative policy analysis. The model used in this study is borrowed from Dabla-Norris et al. (2014) who develop a micro-founded general equilibrium framework with heterogeneous agents to identify constraints to financial inclusion and evaluate policy effects of relaxing these constraints on GDP and inequality.

[2] Colombia is reported to have had the weakest track record on equality compared to major Latin American countries, and the highest Gini coefficient, with inequality levels comparable to Haiti and Angola. This result appears at odds with the country's relatively strong and stable growth profile over the last two decades (IMF, 2013).

DOI: 10.7172/2353-6845.jbfe.2016.2.4

3. THE STATE OF FINANCIAL INCLUSION IN COLOMBIA

Over the past decade Colombia has witnessed substantial financial deepening. Supported by political stability, sound macroeconomic policies, and favorable external developments domestic private credit grew strongly in Colombia, at 14 percent in real terms on average since 2003, outpacing credit growth in regional comparators. At end-2012, the stock of credit-to-GDP amounted to 37 percent, still somewhat below the regional average (Figure 1, Appendix I).

The record on financial inclusion has not, however, kept pace with credit growth. Large amounts of credit do not always correspond to broad use of financial services as credit may be concentrated among the largest firms and highest income individuals. As in other middle-income countries in Latin America, this has also been the case in Colombia, where in 2011 only 15 percent of people belonging to the bottom 40 percent income share held an account at a formal financial institution against 45 percent in the top 60 income share. Young adults and the poor were much less likely to hold an account in a formal institution. The former were also much less likely to hold a formal loan (Figure 2, Appendix I).[3] Only 41 percent of small companies, with less than 20 employees, held a bank loan or a line of credit in 2010, against 72 percent of large firms (Figure 1, Appendix I). Disparities in financial access are one potential explanation for persistent income inequality. In fact, the Gini coefficient improved only marginally since 2000, from 58.7 to 55.9 percent in 2010, when the lowest quintile held only 3 percent of the income share.

Colombia scored below the upper-middle-income average and the average for LACs on financial inclusion indicators related to households. Fewer people in 2011 held debit and credit cards (23 and 10 percent of the population respectively), less than 5 percent of the population received government payments through bank accounts, and less than 10 percent held savings in a formal financial institution (Figure 3, Appendix I). Statistics on frequency of use of accounts for savings and payments were equally grim.[4] In contrast, informal finance was widespread, with a relatively larger share of adults declaring having received a loan from, or having saved through, informal channels. Among closest comparators, Colombia's usage of formal finance was slightly below average, while use of informal finance was on the higher end (Figure 4, Appendix I).[5]

Financial deepening was also not fully "shared" across enterprises. While from the perspective of firms progress on inclusion was recorded in a number of variables reported in the World Bank Enterprise Survey in 2010 compared to 2006, a greater share of enterprises claims to have been affected by insufficient financing more recently. Particularly affected were the firms in the food industry. Among all companies, over 50 percent of smaller ones (with less than 20 employees) have identified access to finance as a major constraint for their operations in 2010 (Figure 5, Appendix I).[6]

4. DETERMINANTS OF FINANCIAL INCLUSION

Obstacles precluding greater financial inclusion may vary widely, and may be micro- or macro-focused in nature. At the macro level, price volatility dissuades savers whose real wealth tends to erode with inflation while trust in institutions may be recouped with great difficulty following

[3] These data are from the Global Financial Inclusion Database, which provides 506 country-level indicators of financial inclusion summarized for all adults and disaggregated by key demographic characteristics—gender, age, education, income, and rural or urban residence. It covers 148 economies.

[4] Results from the 2012 national survey of financial capabilities suggest that 45 percent of the population does not have any financial products, and 72 percent has no savings products. Informal borrowing (mainly from family and friends) was commonly reported as a coping strategy for easing financial strain for 56 percent of the population. Meanwhile, 65 percent of the population reported having been short of money to cover basic needs. (Reddy, et al., 2013)

[5] A useful description of the coverage of different data sources on financial products usage in Colombia is available in Reddy et al. (2013).

[6] Non-financial corporations rely mainly on retained earnings as a source of funding and have low levels of leverage. Loans with - mainly domestic - banks represent less than half of their liabilities. In 2012, 7 percent of largest corporate borrowers accounted for 90 percent of loans (IMF, 2012).

DOI: 10.7172/2353-6845.jbfe.2016.2.4

a banking system failure. A variety of obstacles to greater access to and use of financial services exist also at a micro-institutional level. High cost of services, aside from lack of savings, is the most often quoted reason for avoiding formal finance around the world.[7] This finding appears robust across regions as well as country income types (Demirguc-Kunt and Klapper, 2012).

In practice, obstacles to financial inclusion can be broadly grouped into three distinct categories: access, depth, and efficiency.

- Obstacles to access typically reflect distortions related to scarcity of physical infrastructure, high documentation requirements by banks for opening, maintaining, and closing accounts and for applying to loans, as well as various forms of immeasurable rationing, including red tape and the need for informal guarantors as connections to access finance. These obstacles increase the cost of participation in the financial system.
- Depth is generally determined by collateral requirements that can be high when the rule of law and, more generally, institutions are weak. These can include the state of creditors' rights, information disclosure requirements, and contract enforcement procedures, among others. In fostering greater transparency on practices, credit information, revealed through public credit registries and private credit bureaus, makes assessing risk easier (thereby lowering collateral requirements) and supports trust in the financial system.
- Intermediation efficiency is generally associated with the state of competition and the degree of asymmetric information facing financial institutions, and is reflected in interest spreads and banks' overhead costs.

Some of these obstacles may be particularly binding for poor households, especially those living in distant rural areas, and with lower financial literacy. Whatever the cost of access, it absorbs a higher share of the income of the poor and is likely to weigh more heavily on the choice of how to save and borrow. Therefore, distance to facilities, burdensome paperwork requirements, and other such inclusion barriers are likely to discourage both individuals and enterprises from using formal finance.

4.1. Access

Colombia has implemented a number of improvements to address constraints affecting cost of access.

- Physical infrastructure, the number of access points for financial services, such as commercial bank branches, points of sale, and ATM machines, has increased, although it is still below the average for upper-middle-income countries.[8]
- Banks have been allowed to provide financial services (such as payment, withdrawal, and deposit) through correspondents for social transfers programs (such as *Familias en Accion, Banca de las Opportunidades, and others*) since 2006 and over 38,000 correspondents were registered as of 2013.

The government has subsidized the opening of accounts for most *Familias en Accion* transfers recipients and lowered the financial transaction tax (the "4*1,000") on low account balances.[9] The program of interest subsidies on new mortgages granted to over 5,000 low income families since 2009 has been extended into 2014 and will cover up to 5 percentage points of the agreed interest rate for a 7-year period.

[7] In a survey reported by Maldonado and Tejerina (2010), about 70 percent of respondents claimed not to have savings.

[8] Over the past year only, 320 branches and over 1,500 ATMs were added to the network. Financial services were absent in only 3 out of over 1,100 municipalities as of June 2013 as opposed to 28 percent of total in 2006.

[9] This debt tax had been initially introduced temporarily in 1998, during the banking crisis, but was maintained and increased twice since then, from 0.2 percent to 0.4 percent. It covers all financial transactions, including banknotes, promissory notes, wire transfers, internet banking, bank drafts checks, money and term deposit, overdrafts, installment loans, letters of credit, guarantees, performance bonds, securities underwriting commitments, safekeeping of documents, currency exchange, unit trusts and similar financial products. Its current phasing out is planned to start in 2015 and be concluded in 2018.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

- An electronic money decree was issued to regulate financial transactions between individuals who are not necessarily linked to a formal financial intermediary.
- The National Treasury makes payments exclusively through commercial banks and uses the banks to collect taxes.

Moreover, new consumer credit products are being offered by the banks and are penetrating the market, while internet and mobile banking are becoming increasingly more popular.

Physical Access of Financial Services

	2006	2008	2012
Commercial bank branches			
per 1,000 km^2	3.7	4.0	4.6
100,000 adults	13.3	14.0	14.9
ATMs			
per 1,000 km^2	...	7.7	11.1
per 100,000 adults	...	27.0	35.8

Sources: IMF Financial Access Survey.

4.2. Depth

Colombia's score on the strength of legal rights according to "Doing Business" (2014) is average but depth of credit information is considerably strong. Colombia does not have public registries; however, the two private credit bureaus' coverage has increased substantially over the past years. At 72.5 percent of adults, coverage is more in line with advanced OECD countries. At present, operations of over 750,000 firms and over 20 million individuals are covered by private credit bureaus whose legislation was strengthened in 2010. Both positive and negative information is shared. Nevertheless, some deficiencies with handling of historical data exist inasmuch as "negative" information is kept in the system only for a maximum of 4 years. Moreover, the very lengthy judicial enforcement procedures, and the absence of special treatment for secured creditors in insolvency procedures, have induced financial institutions to seek collateralization of loans, thereby increasing costs faced by borrowers.

4.3. Efficiency

Banks concentration in Colombia may be a phenomenon correlated with depth as well as efficiency. Asset concentration is believed to discourage banks from extending loans to smaller firms. When banks make high profits by lending to a narrow base of customers, they lose incentives to assess riskier customers and diversify their portfolio. In this case, low coverage of small firms is typically coupled with high collateral requirements and high spreads that compensate banks for the risk of failure but also act as gate-keeping expedient.[10]

Colombia's link between asset concentration and the record of financial inclusion are not at odds with developments in its peer economies. Colombia's high bank concentration, with over 70 percent of bank assets held by the five largest institutions, still scores relatively well in terms of regional peers, with Peru and Uruguay displaying much greater concentration (Figure 6). Brazil and Mexico have, however done better in terms of financial inclusion of households. Inclusion

[10] Average interest rate spread was 7.2 percent in 2012.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

of enterprises in Colombia has also lagged behind Brazil and Peru in 2010, and was considerably worse than Chile's. Indeed, in recent years, credit growth in Colombia has mainly derived from an increase in the average size of loans, rather than an increase in the number of debtors (IMF, 2012).

4.4. Progress and challenges

Recent years have witnessed steady progress in fostering financial inclusion in Colombia. The authorities have been closely tracking access to financial services through semi-annual reports (Asobancaria, 2013) documenting the evolution in the number of users of different products based on banks' data. According to data on individual users, since 2011:

- The number of adults owning at least one financial product, the so-called "bancarization", has increased from 63 to over 69 percent, supported by a substantial increase in the use of electronic deposit accounts, which more than trebled over this period;
- Credit and debit cards are becoming increasingly popular although their coverage is still low;
- The growth in the number of people with housing loans was also pronounced although the number of those holding consumer credit is still six times greater;
- On the side of enterprises, the strong increase in the number of checking and savings accounts has far outpaced the increase in access to commercial credit.

Yet, actual usage of financial services is still low and costs are considerable. It is important to distinguish between financial *access* and financial *usage*. Less than 13 percent of account holders made three or more deposits in a month in 2011, against only 5 percent in rural areas. While most recent statistics by Asobancaria suggest a steady increase in the total number of financial transactions, it is less clear if frequency of use has been spread out to a large share of individuals. At $5.50 for entry-level savings, monthly charges on accounts are prohibitively expensive for a large share of the poor population and may be indicative of low market competition.[11]

5. MODEL APPLICATION

The model is borrowed from Dabla-Norris et al. (2014) and focuses on the financial inclusion of enterprises.[12] This micro-funded, general equilibrium, overlapping generation model features heterogeneous agents who are distinguished from each other by wealth and talent and who can chose their occupations between workers and entrepreneurs. In equilibrium, only talented agents with some wealth choose to be entrepreneurs while untalented and those talented but with no wealth choose to be workers. There are two states of world, or "regimes," one with credit and one with savings only. Individuals in the savings regime can save but cannot borrow. Participation in the savings regime is free, but to borrow, i.e. to move into the finance regime, individuals must pay a participation cost whose size is one of the determinants of financial inclusion. Once in the finance regime individuals may obtain credit but its size is constrained due to limited commitment (i.e. poor contract enforceability) which leads to the need to post collateral. Thus collateral is another determinant of financial inclusion affecting financial sector depth. Finally, because of asymmetric information between banks and borrowers, interest rates charged on borrowing account for costly monitoring of highly leveraged firms.[13] Because more productive and poorer agents are more likely to be highly leveraged the higher intermediation cost would be another source of inefficiency and financial exclusion but also inequality.

[11] Basic ATM operations cost US$0.60 per transaction at the bank's own ATM. (IMF, 2012)

[12] The authors actually refer to financial "deepening". However, while financial deepening often denotes and increase in the stock of credit in the economy—which can occur even if the number of borrowers remains unchanged—the model allows for crowding in of enterprises that were initially excluded from the financial sector. Hence, we are using the term "financial inclusion".

[13] Since only highly leveraged firms are monitored, firms face different costs of capital and may chose not to borrow even when credit is available.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

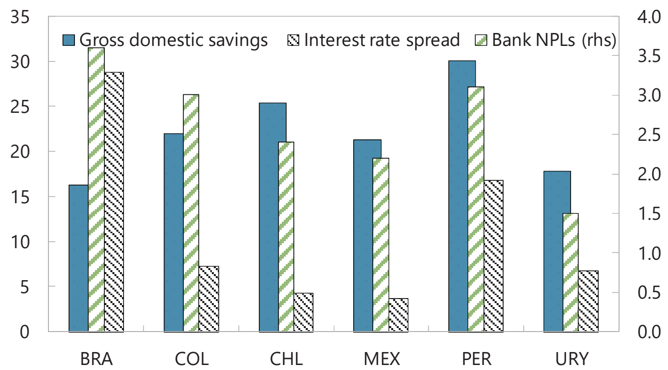
In the model, financial inclusion affects growth and inequality through three channels. First, more developed financial markets channel more funds to entrepreneurs, thereby increasing their output; second, more efficient contracts limit waste from frictions leading to higher growth; and third, more efficient allocation of funds in the financial system brings about an increase in TFP.[14] This occurs as financial deepening speeds up the process in which initially wealth constrained but talented workers become constrained entrepreneurs, while wealth constrained entrepreneurs become unconstrained entrepreneurs.[15]

Two data sets are used: the 2010 World Bank enterprise survey provides firm-level cross-sectional data (from 942 firms) and the development data platform includes data on gross savings, non-performing loans, and the interest rate spread.

In terms of variables used in the model Colombia does not appear to be an outlier compared to regional peers and other developing countries. The savings rate, representing the overall funds available for financial intermediaries in a closed economy, is below that of Chile and Peru, and interest rates spreads are higher Chile's and Mexico's. Yet, NPLs are low and have declined further below 3 percent more recently. Although not excessive by regional standards, Colombia's collateral requirements, at 169 percent, are rather high, with some upper middle-income developing countries, namely Brazil, Malaysia, and Egypt, requiring between 60 and 90 percent collateral. At 57 percent of total registered firms, the number of firms with credit compares favorably. However, as identified above, small firms continue to face severe financial constraints.

Main Model Variables, 2012

(Percent, unless otherwise indicated)[a)]



a) Gross domestic savings are expressed in percent of GDP; interest rate spread equals average lending minus deposit rate; Bank NPLs are expressed in percent of total gross loans.

Source: Author's calculations.

The model was calibrated with Colombian data using standard measures from the literature for some of the parameters as in the original paper. The other parameters are estimated by matching the simulated moments to actual data. The gross savings rate is matched to estimate the bequest rate, ω; the average value of collateral is used to calibrate the degree of financial friction stemming from limited commitment, λ; while the financial participation cost, ψ, intermediation cost, χ, recovery rate, η, probability of failure, p, and the parameter governing the talent distribution, ρ, are jointly estimated to match the moments of the percentage of firms with credit, NPLs as a percent of total loans, interest rate spread, and the employment share distribution. In the model, the share of firms with credit is endogenous and is affected by ψ, λ, and χ. We conduct three isolated policy experiments that can help identify key constraints to financial sector inclusion and

14 However, financial inclusion can also crowd in relatively untalented agents, decreasing TFP.

15 GDP is calculated as the sum of all individuals' income; TFP is the average entrepreneur's talent weighted by their respective output.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

study the macro effects of their removal. The first experiment consists of reducing the financial participation cost, ψ. The second experiment consists of relaxing borrowing constraints in the form of collateral requirements, λ. The third experiment assumes an increase in intermediation efficiency, χ.

Calibration: Data, Model, and Estimated Parameters

<table>
<tr><th>Target Moments</th><th>Data</th><th>Model</th><th>Parameter</th></tr>
<tr><td>Savings (%GDP)</td><td>20.0</td><td>20.0</td><td>ω=0.2</td></tr>
<tr><td>Collateral (% loan)</td><td>169.00</td><td>169.0</td><td>λ=1.59</td></tr>
<tr><td>Firms with credit (%)</td><td>57.2</td><td>57.4</td><td rowspan="7">ψ=0.06
χ=0.3
η=0.37
p=0.17
ρ=3.8</td></tr>
<tr><td>NPLs (%)</td><td>4.0</td><td>4.2</td></tr>
<tr><td>Top 5% emp. share</td><td>52.1</td><td>54.9</td></tr>
<tr><td>Top 10% emp. share</td><td>65.7</td><td>67.3</td></tr>
<tr><td>Top 20% emp. share</td><td>80.3</td><td>79.1</td></tr>
<tr><td>Top 40% emp. share</td><td>92.8</td><td>89.3</td></tr>
<tr><td>Interest rate spread (%)</td><td>6.2</td><td>5.0</td></tr>
</table>

5.1. Reducing the participation cost

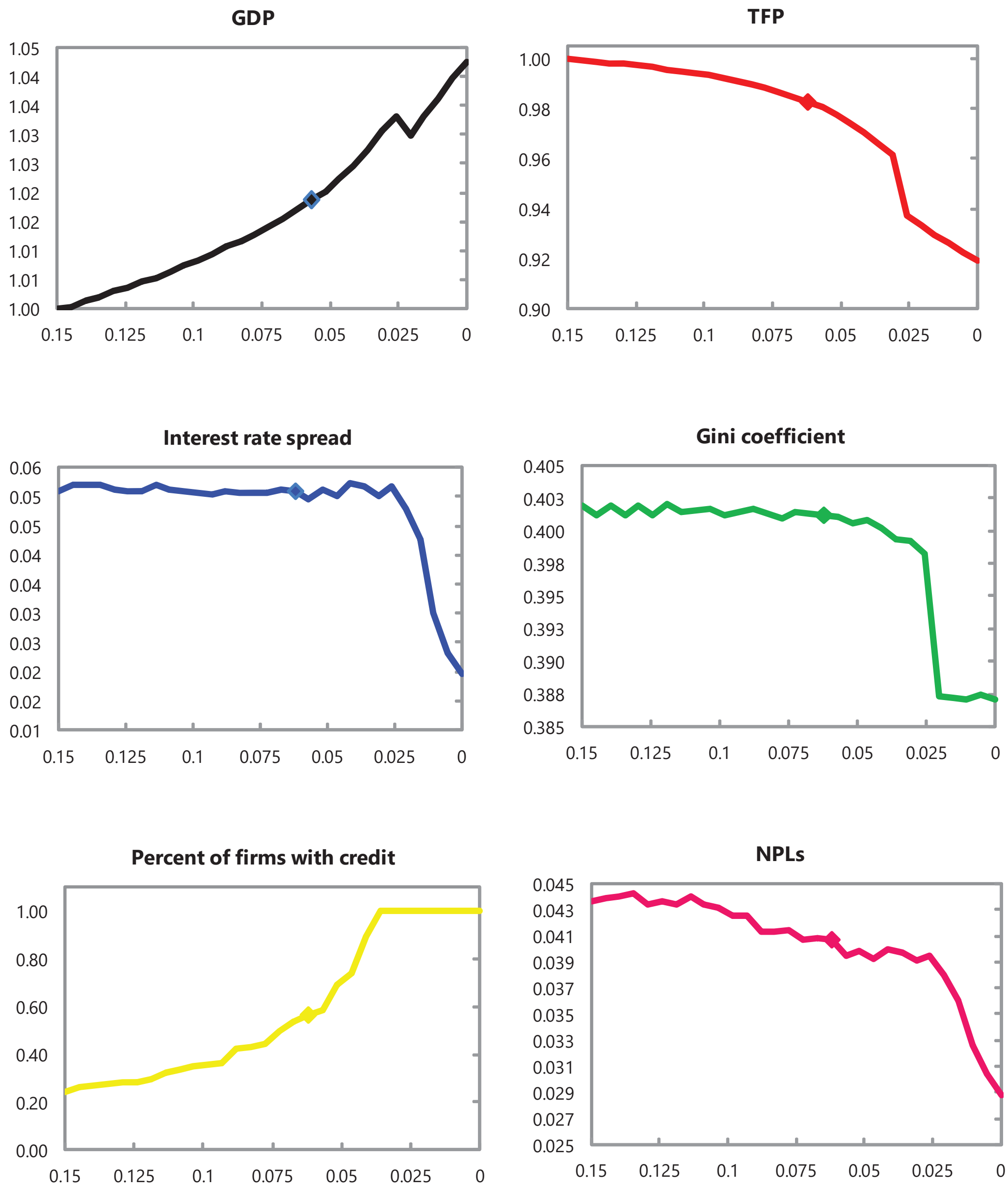
The impact of a decline in the financial participation cost, ψ, from 0.15 to 0 on GDP reported in Figure 1 is favorable. A decrease in the participation cost pushes up GDP through its positive effect on investment for two reasons: (i) a lower ψ enables more firms to have access to credit, and (ii) fewer funds are wasted in unproductive contract negotiation freeing up more capital for investment. However, aggregate TFP declines, implying efficiency losses in the allocation of capital. This occurs because the participation cost, which is fixed, has a higher weight in small firms' income. As the previously excluded firms enter the financial sector they push down TFP of the economy.

The interest rate spread is very stable when financial participation cost is high, but decreases as ψ approaches zero. This is because a decrease in ψ has two countervailing effects on interest rates in the model. First, the wealth effect—entrepreneurs become "richer", and tend to deleverage, which results in a lower average interest rate spread. Second, a smaller ψ enables some severely wealth constrained workers to become entrepreneurs. These entrepreneurs choose a very high leverage ratio, driving the average interest rate spread up. The first effect dominates the second effect when borrowing constraints are very tight, thus discouraging constrained workers' access.

As the financial market develops, income inequality decreases. A decrease in ψ is disproportionally more beneficial for constrained workers and entrepreneurs without credit. It allows them to invest capital into production driving down the Gini coefficient. The share of firms with credit increases until all firms have access to finance as ψ approaches 0, while the share of non-performing loans (NPLs) declines. The decline in inequality reaches a plateau the process hits other binding constraints to inclusion.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

Figure 1
Comparative Statics: Reducing Participation Cost-ψ



5.2. Relaxing borrowing constraints

Relaxing borrowing constraints by varying λ from 1 to 3 in Figure 2 has a positive effect on GDP and TFP. The increase in aggregate GDP is greater than in the experiment related to financial participation costs. The relatively high savings rate implies that the decline in the collateral requirement unlocks financial resources, leading to a significant increase in GDP. As λ declines, TFP increases, implying a more efficient resource allocation across firms.[16] The effect on GDP is very large suggesting that credit constraints are one of the major obstacles to financial development in Colombia.

The interest rate spread increases in this scenario. The spread is zero when λ is low, because firms leverage is low and no default happens even when production fails. As λ increases above

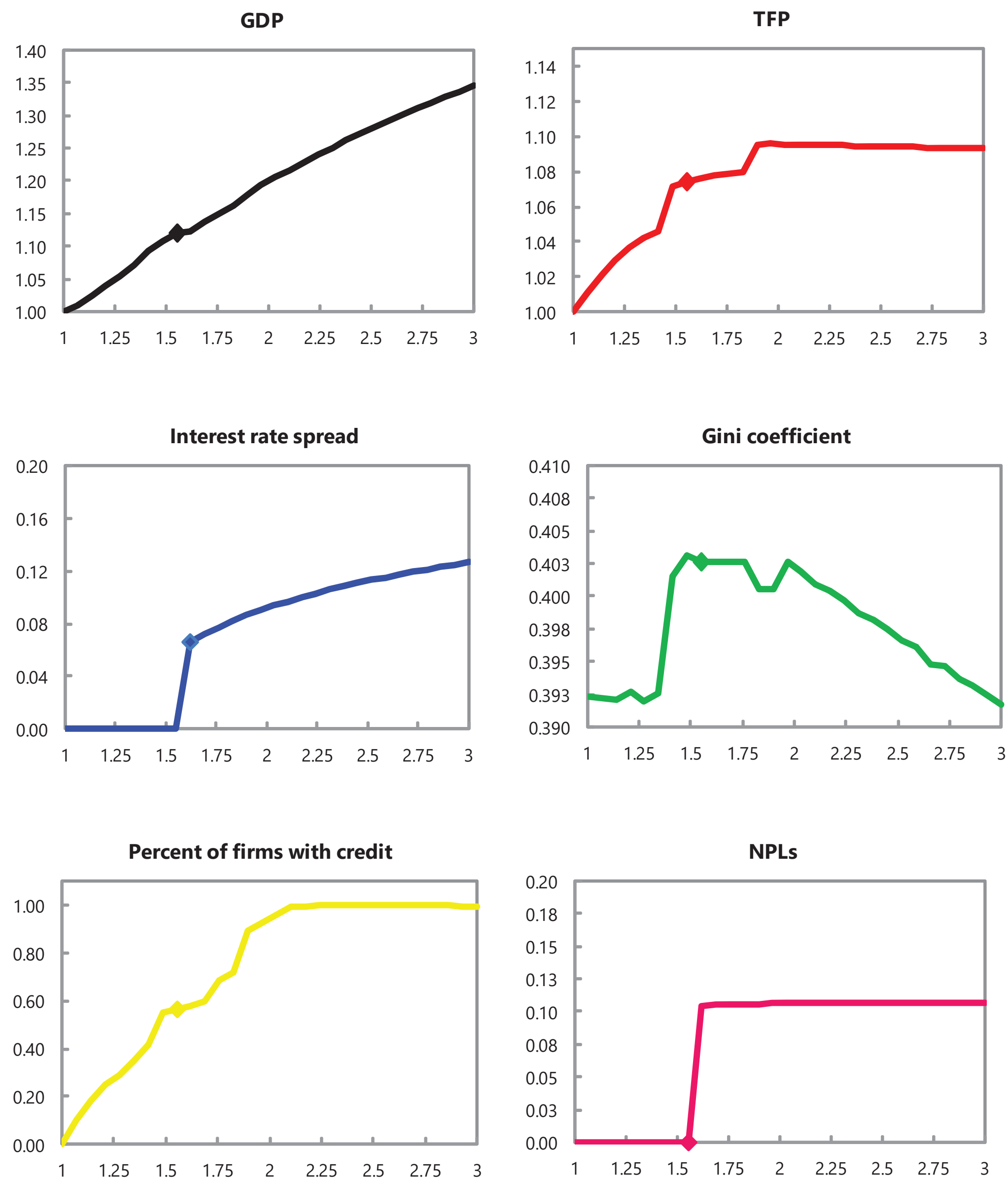
[16] Dabla-Norris et al. (2014) describe this process in the following way: a relaxation of the borrowing constraint benefits talented entrepreneurs more as they often desire to operate firms at a larger scale than untalented entrepreneurs. Relaxing the borrowing constraint allows all entrepreneurs to borrow more, but, on average, untalented ones do not borrow as much because their small maximum business scale may have already been achieved. As a result, more talented entrepreneurs expand business scales, driving up TFP in the "finance regime".

DOI: 10.7172/2353-6845.jbfe.2016.2.4

a threshold, agents leverage more, the share of non-performing loans increases, and the interest rate spread starts increasing. Also, in line with Kuznets theory, when λ increases from low levels, talented entrepreneurs leverage more and increase their profits, driving up the Gini coefficient. However, as λ becomes larger, the sharp increase in the interest rate shrinks entrepreneurs' profits, leading to a lower Gini coefficient. The stage in which Columbia is now (i.e. its current value of λ) suggests that inequality should be declining.

A relaxation of borrowing constraints pushes up the share of firms with credit but also increases NPLs. Relaxing the borrowing constraint provides more external credit to entrepreneurs once they pay the participation cost. This induces more entrepreneurs to join the financial regime. However, NPLs increase. This occurs as a relaxation of collateral constraints opens up the doors for small new entrants who tend to be more leveraged. This phenomenon underlines a trade-off between growth and stability that needs to be carefully managed.[17]

Figure 2
Comparative Statics: Relaxing Borrowing Constraints-λ



[17] Note that caution should be made in interpreting the magnitude of the changes in the variables of interest across experiments in the figures. The scales on the y-axis of the figures are intentionally different to allow appreciating the various turning points of the variables.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

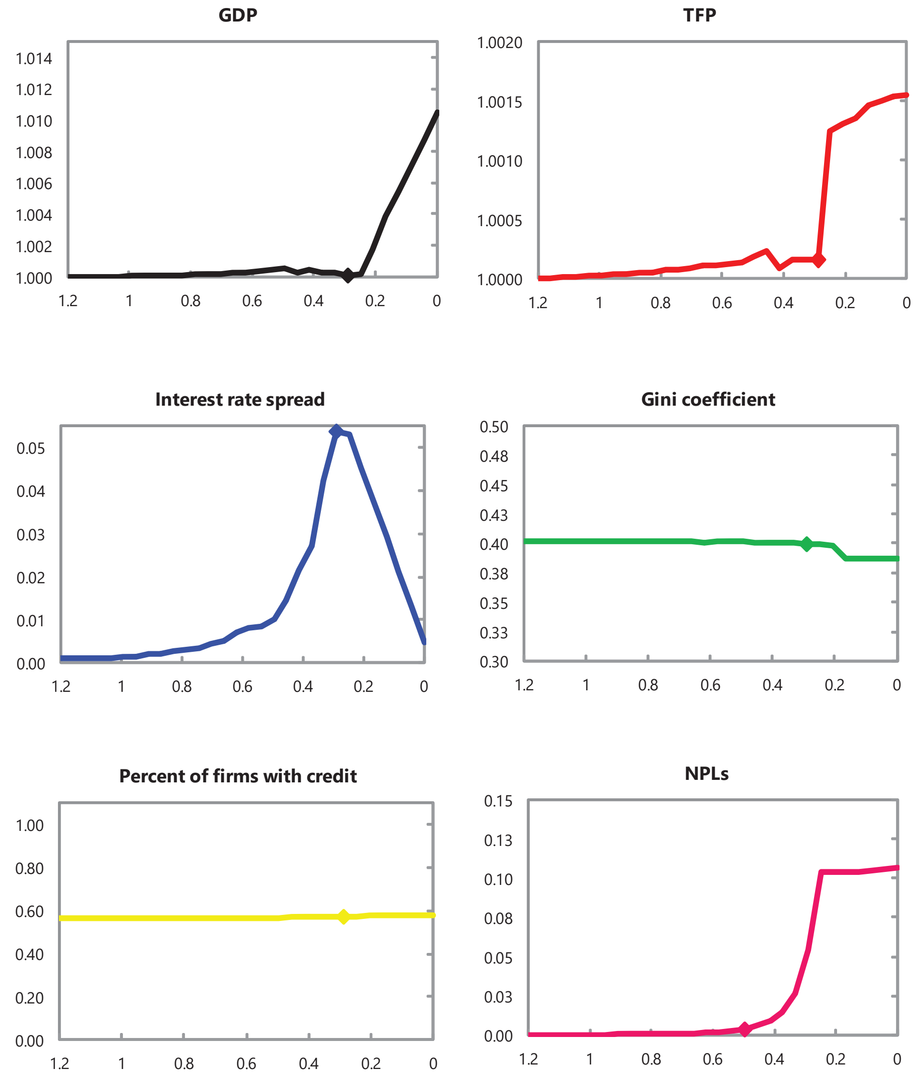
5.3. Increasing intermediation efficiency

Varying the financial intermediation cost, χ, from 1.2 to 0 in Figure 3 pushes up growth and TFP. GDP and TFP are responsive to a decrease in χ although less so compared to the case where λ is lowered. At higher levels of χ, better intermediation efficiency only benefits the highly leveraged firms which are few (due to the low financial inclusion ratio and tight borrowing constraints). As χ decreases further TFP increases because the lower intermediation cost facilitates the allocation of capital to more efficient entrepreneurs.

The interest rate spread can be expected to decrease. The spread increases initially for lower levels of χ and decreases sharply as χ approaches zero, displaying an inverted V shape. There are two opposing forces affecting the spread stemming from a decline in χ: first, the decline in the cost of borrowing induces risky firms to leverage more, pushing up NPLs and increasing the endogenous interest rate spread; second, the decline in χ decreases the interest spread directly. Whether the interest rate spread increases or decreases depends on which effect dominates.

However, the percent of firms with credit remains unchanged. Efficient intermediation appears to be disproportionately benefiting a small number of highly leveraged firms, while the general equilibrium effects on wages and the interest rate may be preventing smaller firms from entering the financial system. The Gini coefficient declines only marginally at very low parameter levels.

Figure 3
Comparative Statics: Increasing Intermediation Efficiency-χ



DOI: 10.7172/2353-6845.jbfe.2016.2.4

5.4. Discussion of results

5.4.1. Comparative statics

Comparison of results across measures shows that different financial inclusion strategies have differential effects on the variables of interest. First, relaxing constraints on collateral appears to offer the greatest benefits in terms of growth, TFP and inclusion of firms. Yet, the effect on inequality is much lower compared to the case when the cost of access decreases, and the increase in the share of firms with credit is strong, at 76 percent. In fact, entrepreneurs who are already included in the financial system benefit more from the reduction in collateral and less so from a reduction in participation cost which is a fixed cost and a relatively lower share of their income. The latter, however, benefits new entrepreneurs more decreasing inequality. Nevertheless, the "poor" may still be better off overall under the lower collateral scenario, albeit not relative to the "rich".[18]

Different financial inclusion strategies may imply trade-offs and present undesired side effects that need to be closely monitored. A side effect of a decrease in collateral constraints is increasing spreads and NPLs. Low NPLs are not necessary welcome as they may precisely be a reflection of limited lending, possibly circumscribed to low-leveraged, rich entrepreneurs. Entry of new entrepreneurs would however still point to the need for close monitoring of NPLs and possibly mitigating macro prudential measures.

Some financial inclusion measures may not have the result policymakers are hoping for. Increasing intermediation efficiency does not appear to bear a particularly strong effect on any variable. This most likely occurs because collateral constraints and participation costs are more binding financial sector frictions. Greater intermediation efficiency would be enjoyed only (or disproportionally more) by entrepreneurs that are already included in the financial system and would not affect inequality.

The Impact of Financial Inclusion (percent)

	GDP	TFP	Interest rate spread	Gini coefficient	Percent of firms with credit	NPLs
↓ ψ to 0	4.3	–8.1	–3.1	–1.5	75.9	–1.5
↑ λ to 3	34.6	9.3	12.6	–0.1	99.6	10.7
↓ χ to 0	1.1	0.2	0.4	0.0	1.1	10.7

These examples are illustrative, as the calibration for the financial inclusion process is chosen arbitrarily. It may well be possible to increase λ beyond 3 in a shorter period of time compared to that necessary to achieve other changes, with greater positive effects on the Gini coefficient. Moreover, as many reforms are implemented on various fronts contemporaneously they are likely to affect the frictions in unison with additive effects. The results of the calibration to Colombia are similar to the emerging economies' experiments in the original paper by Dabla-Norris et al. (2014), in particular to the results of Philippines, suggesting that there may be similarities in the process of financial inclusion for countries with similar economies and similar level of development.

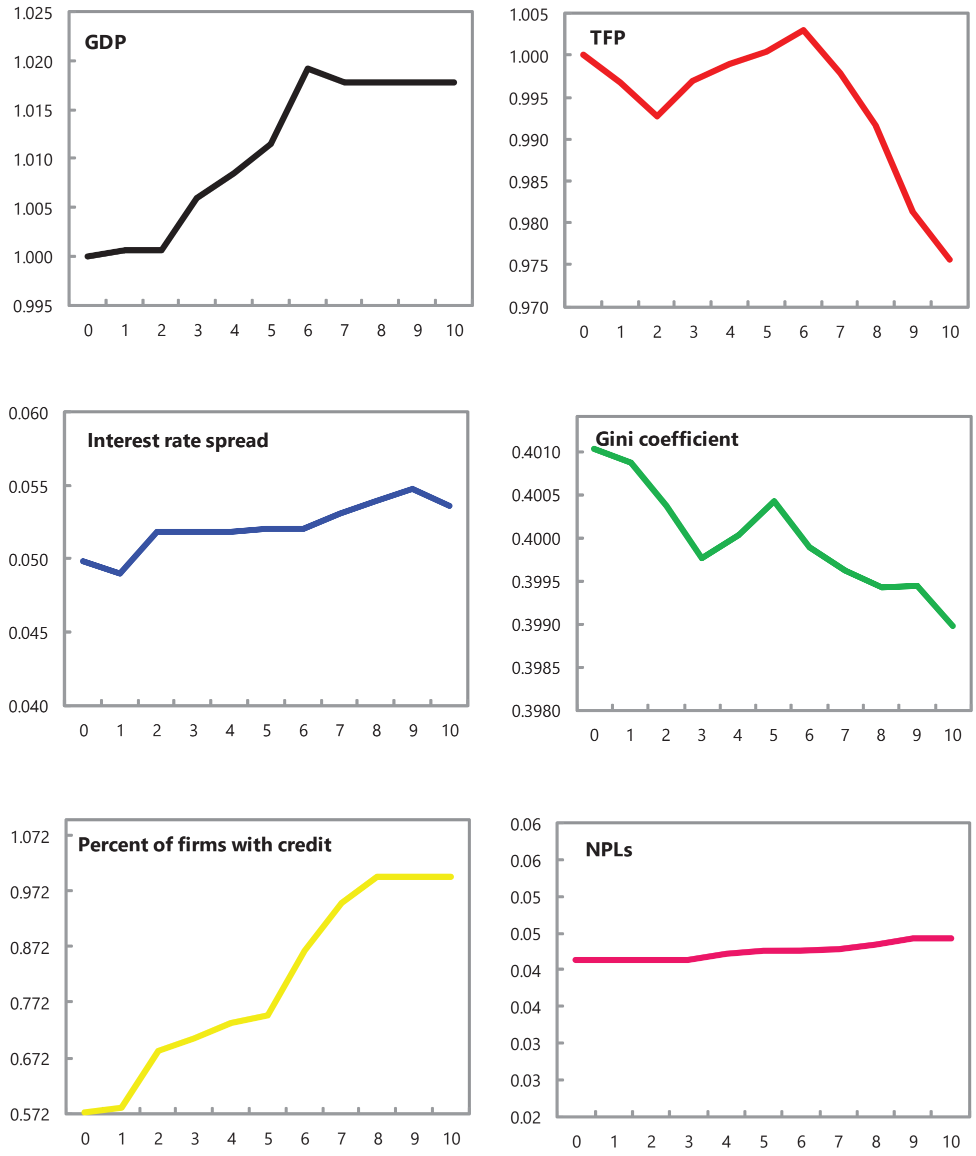
[18] Inequality does not decrease substantially with lower λ because "rich" entrepreneurs (possibly also more talented and more productive) can borrow much more when collateral constraints are released increasing firm size and profits, thus becoming richer. The optimal production scale of new entrants is lower and, even if they can borrow, they are not likely to achieve the same profits.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

5.4.2. Transitional dynamics

Figures 4-6 included below show the transitional dynamics of various measures. Starting at year 0, the figures show the dynamics reflecting a linear decrease in φ and χ by 50 percent, and an increase in λ by 30 percent over 10 years. The interpretation of results remains the same with the addition of the time dimension of financial inclusion. Nevertheless, the transitional dynamics is important inasmuch as it points to possible temporary trade-offs of various measures. For instance, lowering cost of access in Figure 4 implies a temporary increase in the Gini coefficient in the transition period before it declines to a lower level.

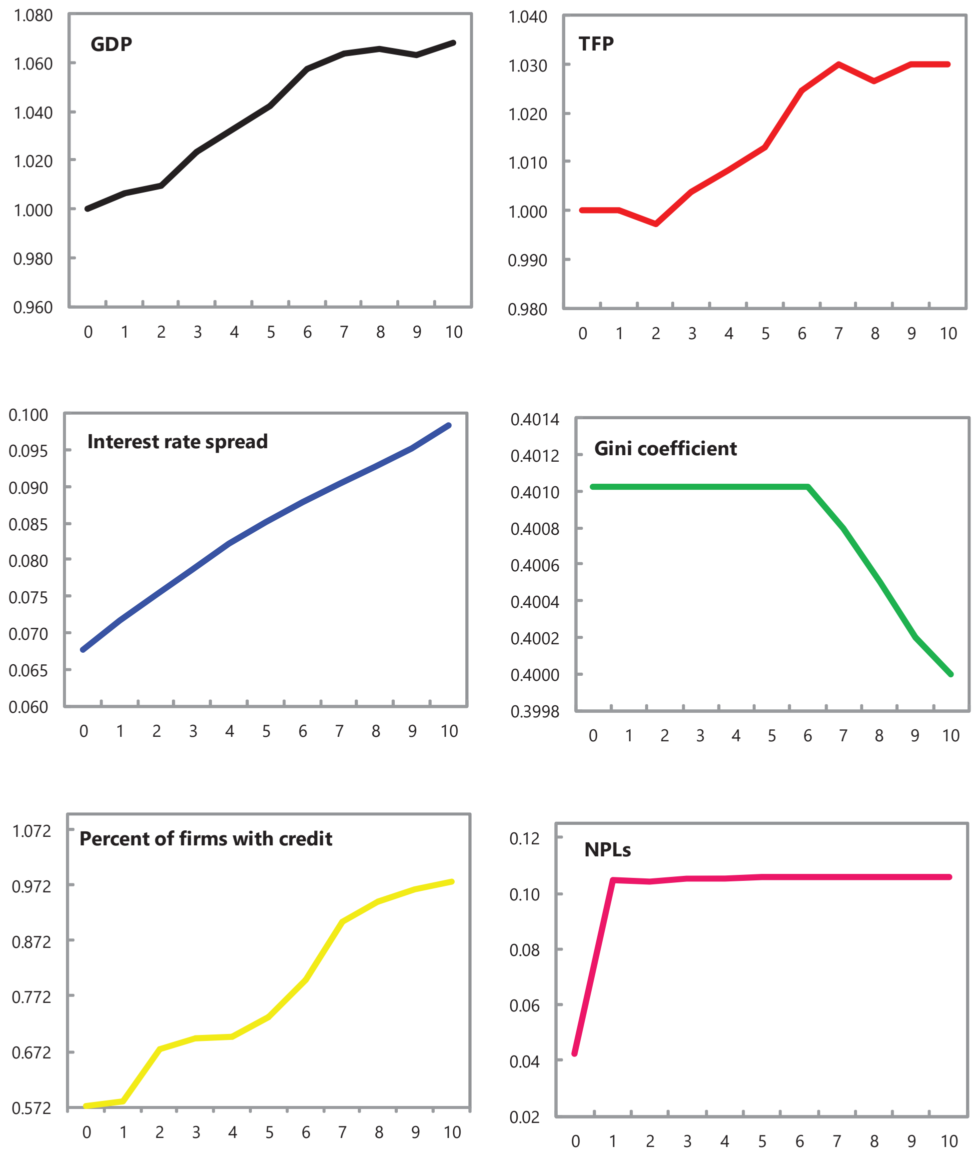
Figure 4
Transitional Dynamics: Relaxing Constraints to Access[a)]



a) Effect of a 50 percent decrease in the participation cost.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

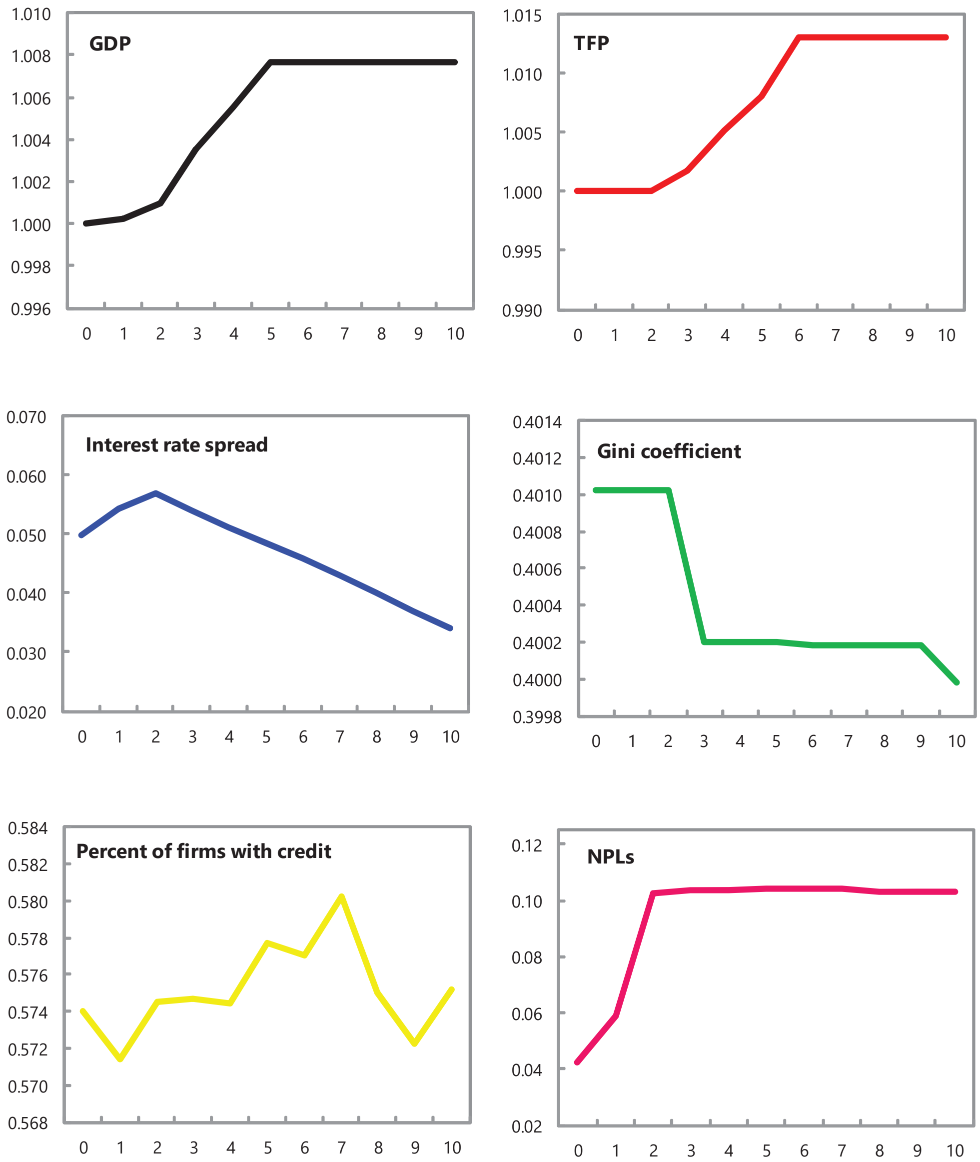
Figure 5
Transitional Dynamics: Relaxing Borrowing Constraints[a)]



a) Effect of a relaxation of borrowing constraints by 30 percent.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

Figure 6
Transitional Dynamics: Increasing Intermediation Efficiency[a)]



[a)] Effect of a 50 percent decrease in the intermediation cost.

6. CONCLUSIONS

Boosted by government support in various areas and financial sector innovation, financial inclusion is progressing in Colombia. Microcredit is growing, "bancarization" is spreading, and electronic payments are increasingly being accepted for economic transactions. The financial inclusion agenda continues to gain momentum, supported by domestic policy interest as well as global focus on financial inclusion. Authorities' efforts in this area can only be expected to intensify going forward.

The effects of governments' financial inclusion actions on growth and inequality will depend upon the pace and choice of measures implemented. Grouping the various micro initiatives and the remaining challenges into three broad areas of financial frictions—participation costs (access), borrowing constraints (depth), and intermediation efficiency—it is possible to assess the effects the removal of constraints has on main macroeconomic variables in a general equilibrium model. Simulations suggest that relaxing various financial sector frictions may affect growth and

DOI: 10.7172/2353-6845.jbfe.2016.2.4

inequality differently in the transition and in the steady state. Lowering contraints on collateral precluding greater financial sector inclusion promises higher growth while inequality is better tackled through measures that lower the financial participation cost. However, some measures may imply tradeoffs that need to be monitored closely.

Some ideas already in the implementation phase are promising and new areas of intervention could also be explored. The financial inclusion model is theoretical by nature and does not allow for identifying country-specific micro-level measures that may be most successful in removing financial sector friction. However, the authorities are already acting on several different fronts. The recent proposal to license electronic-money issuers, that would be entitled to collect deposits and offer electronic payment services, goes in the right direction towards creating more competition in the financial sector. This can in turn have positive effects on collateral requirements but also on the other two financial inclusion barriers, participation costs and intermediation efficiency. The recently passed Law on movable property should also relax borrowing constraints by increasing transparency and improving access to information. Moreover, supporting policies to improve the regulatory flexibility—by, for instance, simplifying account opening (as discussed in the recent FSAP)—and policies to enhance consumer protection, could also contribute to lowering the participation cost in a more substantial way. Going forward, some areas for identifying remaining frictions may include possible regulatory obstacles to bank entry, market practices on the use of collateral, and options for further improving access to and adequacy of credit information.

Acknowledgements

I would like to thank Era Dabla-Norris, Valerie Cerra, Marina Tavares, and Filiz Unsal (all IMF) and Professor Robert Townsend, Stacy Carlson, and Yu Shi (all MIT) for their helpful suggestions and comments, and Yan Ji (MIT) for the calibration of the model to Colombia.

References

Aiyagari, M., Demirguc-Kunt, A., Maksimovic, V. (2008) How well do institutional theories explain firms' Perceptions of Property Rights?, *Review of Financial Studies*, No. 21 (4), pp. 1833–1871.

Aiyagari, M., Demirguc-Kunt, A., Maksimovic, V. (2010) Formal versus informal finance: evidence from China, *Review of Financial Studies, Society for Financial Studies*, vol. 23(8), pp. 3048–3097.

Asobancaria, *Informe Semestral de Inclusion Financiera*, June 2013.

Bebczuk, R.N. (2008), Financial inclusion in Latin America and the Caribbean: Review and Lessons, *Centro de Estudios Distributivos, Laborales y Social*.

Beck, T., Demirguc-Kunt, A., Levine, R. (2007) Finance, inequality and the poor, *Journal of Economic Growth*, No. 12 (1), pp. 27–49.

Beck, T., Demirguc-Kunt, A., Maksimovic, V. (2005) Financial and legal constraints to growth: does firm size matter? *Journal of Finance*, No. 60 (1), pp. 137–177.

Blaum, J. (2013) Wealth inequality and the losses from financial frictions, Mimeo.

Bruhn, M., Love, I. (2009) The economic impact of banking the unbanked—evidence from Mexico, *WB Policy Research Working Paper* No. 4981 (Washington: World Bank).

Camara, N., Tuesta, D. (2014) Measuring financial inclusion: a multidimensional index, *BBVA Working Paper* (14/26), BBVA Research.

Clarke, G. R. G., Xu, L. C., Fu Zou, H. (2006) Finance and income inequality: what do the data tell us?, *Southern Economic Journal*, No. 72 (3), 578–596.

IMF (2013) Colombia: Staff Report for the 2013 Article IV Consultation, International Monetary Fund

Doing Business (2013) (Washington: The World Bank).

Dabla-Norris, E., Ji, Y., Townsend, R., Unsal, F. (2014), Distinguishing constraints on financial inclusion and their impact on GDP, TFP and inequality, *NBER Working Paper*, No. 20821, (Cambridge, MA: National Bureau of Economic Research).

DOI: 10.7172/2353-6845.jbfe.2016.2.4

Dabla-Norris, E., Deng, Y., Ivanova, A., Karpowicz, I., Unsal, F., VanLeemput, E., Wong, J. (2015) Financial inclusion: zooming in on Latin America, *IMF Working Paper 15/206*, International Monetary Fund.

Demirguc-Kunt, A., Klapper, L. (2012) Measuring financial inclusion: The global Findex database, *World Bank Policy Research Paper* 6025 (Washington: World Bank).

Demirguc-Kunt, A., Klapper, L., Singer, D., Van Oudheusden, P. (2014) The global Findex database 2014, *Policy Research Working Paper* 7255, The World Bank.

IMF,(2012) Financial Stability Report, International Monetary Fund.

Global Financial Development Report (2014) (Washington: World Bank).

Hohonan, P. (2007) Cross country variation in households access to financial services, Conference manuscript.

Karpowicz, I., 2014, Financial inclusion, growth and inequality: a model application to Colombia, *IMF Working Paper*, WP/14/166, International Monetary Fund.

King, R.G. Levine, R. (1993), Finance and growth: Schumpeter might be right, *The Quarterly Journal of Economics*, No. 108 (3), 717–37.

Klapper, L., Singer, D. (2015) The role of informal financial services in Africa, *Journal of African Economies*, Vol. 24.

La Gran Encuesta PYME, ANIF Centro de Estudios Económicos, 2nd Semester 2013.

Levine, R. (2005) Finance and growth: theory and evidence, *Handbook of Economic Growth*, Vol. 1, edited by Aghion P. and Durlauf, S., Elsevier, pp. 865–934.

Moll, B., Townsend, R., Zhorin, V. (2014) Economic development and the equilibrium interaction of financial frictions, mimeo.

Nigeria: Selected Issues Paper, 2015, International Monetary Fund.

Reddy, R., Bruhn, M., Congyan, T. (2013) *Financial capability in Colombia: results from a national survey on financial behaviors, attitudes, and knowledge* (Washington: World Bank).

Rojas-Suarez, L., Amado, M.A. (2014) Understanding Latin America's financial inclusion gap, *CGD Working Paper* 367, Center for Global Development.

Sarma, M. (2008) Index of financial inclusion, *ICRIER Working Paper* 215, Indian Council for Research on International Economic Relations.

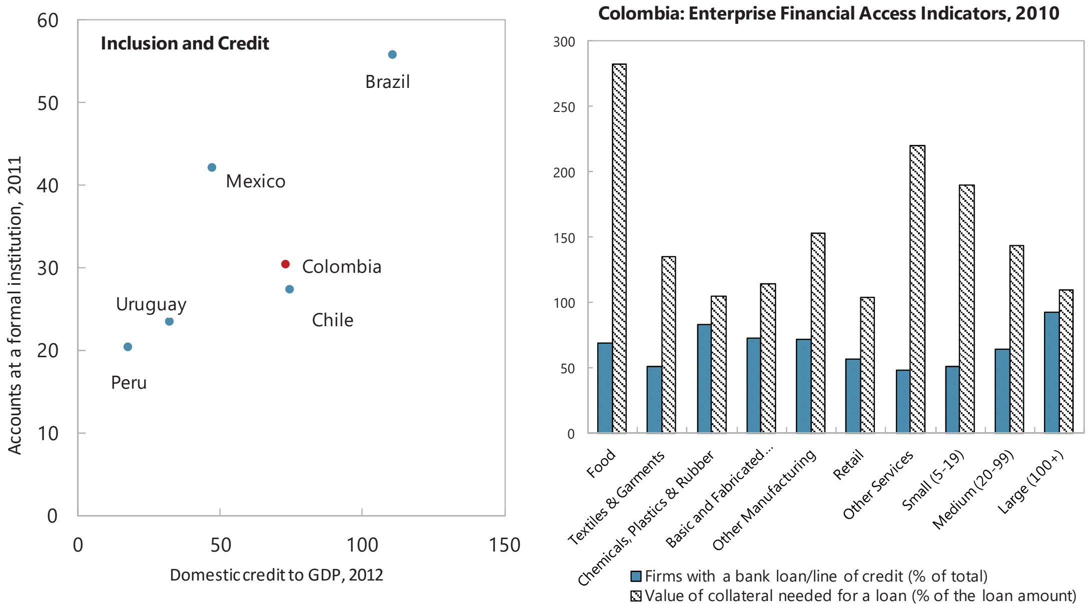
Townsend, R., Alem, M. (2014) An evaluation of financial institutions: impact on consumption and investment using panel data and the theory of risk-bearing, *Journal of Econometrics*, 183(1): 91–103.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

APPENDIX I.

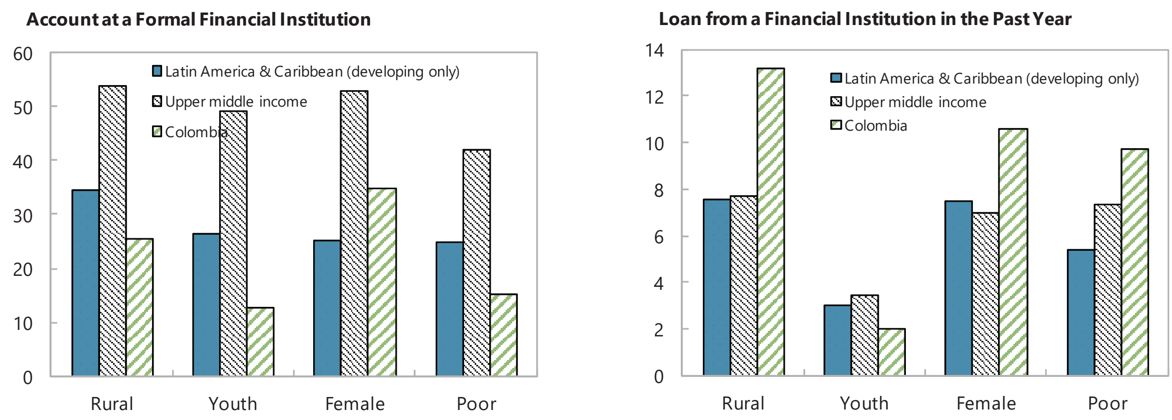
Taking Stock of Financial Inclusion

Figure 1
Colombia: Inclusion – Households and Enterprises, 2010–12 (percent)



Source: Findex database and Enterpise Survey, The World Bank

Figure 2
Colombia: Formal Finance, 2011
(Percent of population age 15 and above, unless otherwise indicated)[a)]

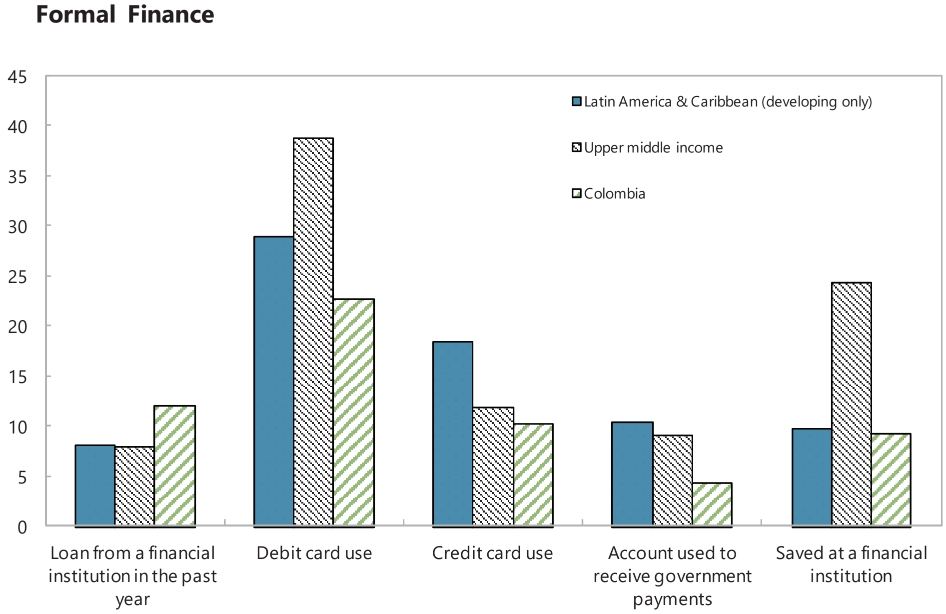


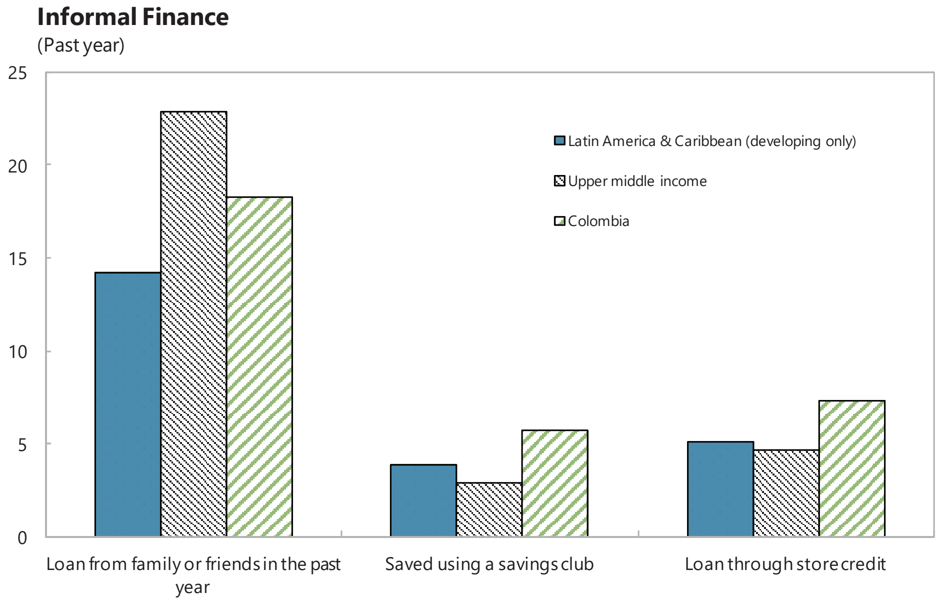
[a)] Youth – percent of population aged 15–24; Poor – percent of population aged 15 and above whose income is in the bottom 40 percent.

Sources: Findex database, World Bank

DOI: 10.7172/2353-6845.jbfe.2016.2.4

Figure 3
Colombia: Financial Inclusion Indices, 2011
(percent of population aged 15 and above)

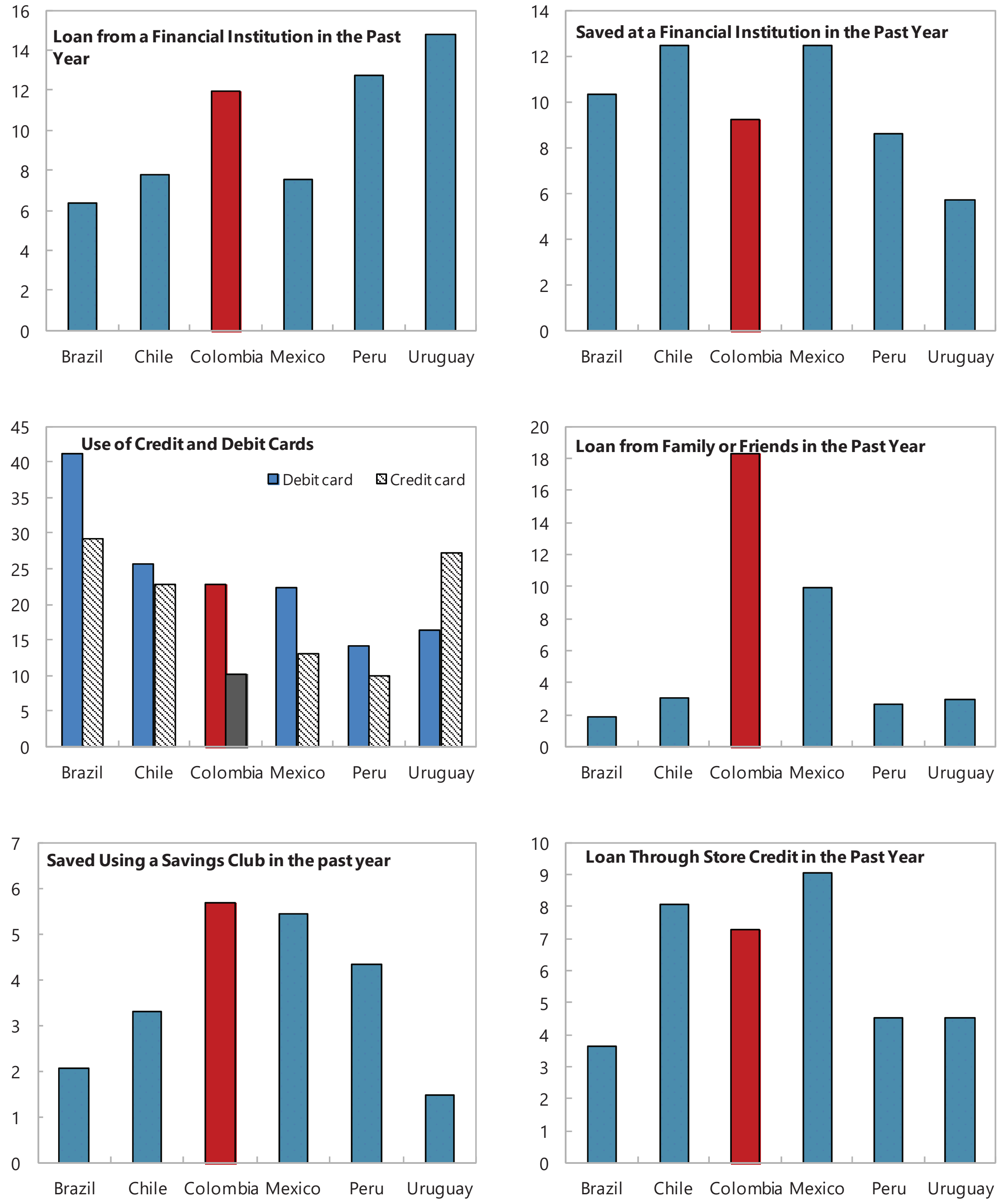




Source: Findex Database, World Bank.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

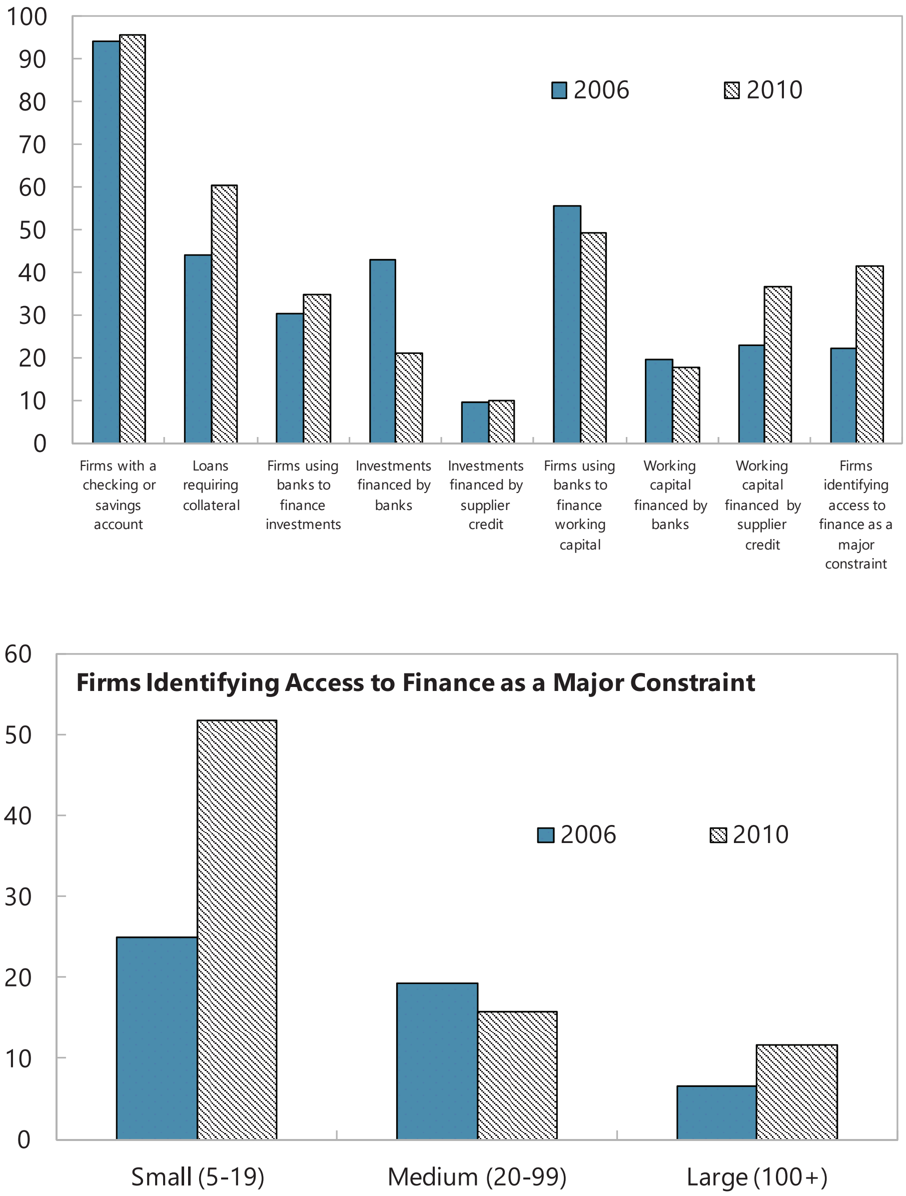
Figure 4
Colombia: Formal and Informal Finance, 2011 (percent)



Source: Findex database; The World Bank.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

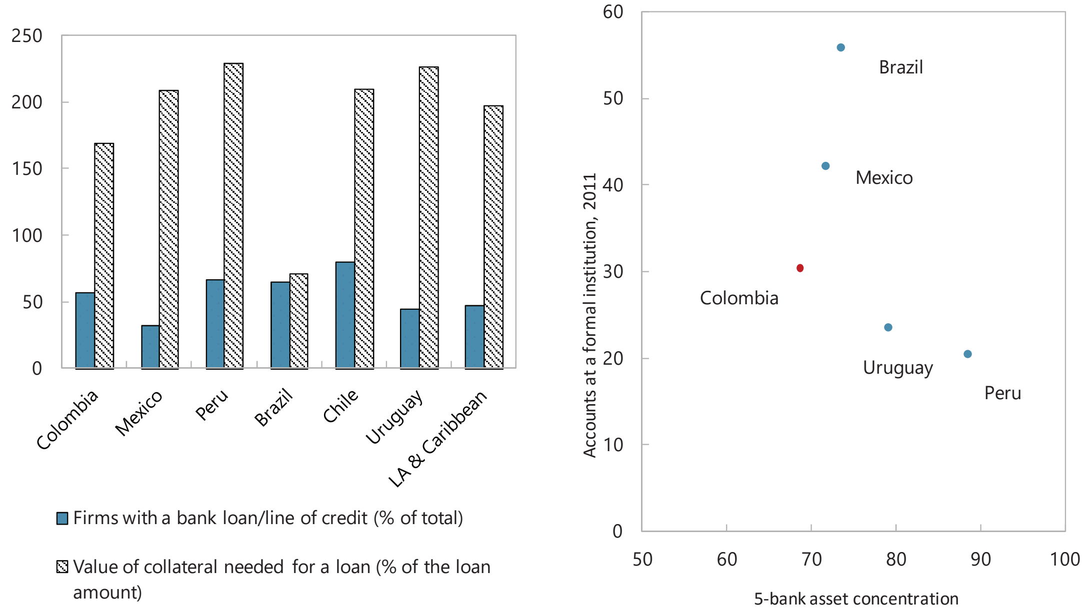
Figure 5.
Colombia: Enterprise Survey Indicators, 2006–10 (percent)



Source: Enterpise Survey, The World Bank.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

Figure 6
Colombia: Banks Concentration, Households and Entreprise Inclusion, 2010–12 (percent)



Source: WDI; Findex and Enterprise Survey Banking Data, The World Bank.

DOI: 10.7172/2353-6845.jbfe.2016.2.4

Spotting Bubbles: A Two-Pillar Framework for Policy Makers

Bradley A. Jones[1]
International Monetary Fund, United States of America
bjones@imf.org

Received: 5 March 2016 / Revised: 12 July 2016 / Accepted: 15 July 2016 / Published online: 29 July 2016

ABSTRACT

In the aftermath of the global financial crisis, the issue of how best to identify speculative bubbles remains in flux. This owes to the difficulty of disentangling irrational investor exuberance from the rational response to lower risk, based on price behavior alone. In response, I introduce a two-pillar (price and quantity) approach for financial market surveillance. While asset pricing models comprise a valuable component of the surveillance toolkit, risk taking behavior, and financial vulnerabilities more generally, can also be reflected in subtler, non-price terms. Though policy makers will always encounter uncertainty when attempting to measure imbalances in financial markets, 'perfect should not be the enemy of the good.' In this spirit, the framework in this paper seems to capture some of the stylized facts of asset booms and busts, and thus could offer policy makers a practical guide as to when to consider leaning against the wind.

JEL classification: E44, F37, G12, G15, G18

Keywords: asset bubbles, asset pricing, market efficiency, macroprudential policy

1. INTRODUCTION

Financial history reads in many respects as a history of booms, bubbles and busts. The Dutch Tulip Mania (1634–1637), the French Mississippi Bubble (1719–20), the South Sea Bubble in the United Kingdom (1720), the first Latin American debt boom (1820s), and railway manias in the United Kingdom (1840s) and United States (1870s) are all notable early examples.[2] In the past century, no busts have been more devastating than the Great Depression ushered in by the collapse of world stock markets in 1929. Over the past few decades, the Japanese Heisei bubble in the late 1980s, the various emerging market booms and busts in the 1980s and 1990s, and the equity mania in the late 1990s, offer examples of speculative frenzies gone awry. The threat to financial stability posed by large asset price movements has come into sharper focus over the past decade as the boom during the Great Moderation gave way to the collapse in global credit, real estate, and equity markets. Most recently, questions have been raised as to whether the prolonged use of unusually accommodative monetary policies may be fermenting another asset price bubble.

[1] Corresponding author Bradley A. Jones. Email: BJones@imf.org. Tel +1 202-623-9914. International Monetary Fund, Monetary and Capital Markets Department, 700 19th street NW, Washington D.C., USA.

[2] As with many aspects of the speculative bubble literature, it should be noted that there remains considerable disagreement among scholars as to which of these episodes actually represent bona-fide bubbles.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Much of the debate over the existence and implications of speculative bubbles stems in the first instance from disagreement as to their measurable properties: the term "bubble" has been widely used to mean very different things.[3] Bubble models typically emphasize the self-fulfilling nature of expected future price changes based on the concept of 'resale optionality.' This enables asset prices to be decomposed into a rational intrinsic yield component (based on discounted future cash flows), and an irrational bubble component (based on expectations of future capital gains independent of fundamentals). Other researchers have focused upon the broader social dimensions of bubbles, with their tendency to engulf members of society who typically have little interest in financial matters (Mackay, 1841; Keynes, 1936; Kindleberger, 1978; Chancellor, 2000; Shiller, 2000a; Bonner and Rajiva, 2007; Reinhart and Rogoff, 2009; and Akerlof and Shiller, 2009).[4] Yet definitional ambiguity and inference problems have long plagued formal studies of speculative bubbles. Distinguishing irrational investor exuberance from the rational response to lower perceived risk is made difficult in real time by numerous issues, not least that it can only be known with absolute certainty *ex-post* whether the optimistic *ex-ante* projections embedded in asset prices were in fact justified.

Though much of the post-crisis literature has focused, understandably, on the role of credit growth in fermenting asset bubbles, the next major threat to financial stability may well take a different form. Moreover, as capital markets continue to expand in scale and scope, there are reasons to expect their dynamics to increasingly capture the attention of policy makers (Haldane, 2014; Stein, 2014; Feroli and others, 2014; Jones, 2015). In response, I introduce a simple two-pillar approach to bubble surveillance, based on both price and quantity data in the capital markets.[5] Though by no means a conclusive solution to the age-old difficulties of crisis prediction, the framework appears to offer promise in capturing some of the stylized facts of asset booms and busts: some of the largest in history have been associated not just with below average risk premia (captured by the 'pricing pillar'), but also with unusually elevated patterns of issuance, trading volumes, fund flows, and survey-based return projections (i.e., the 'quantities pillar'). The ability to cross-reference signals from both pillars may give policy makers a richer understanding of the dynamics of asset price cycles and the threats they pose (if at all) to economic stability.

The analysis proceeds as follows. A synthesis of measurement and inference issues that arise in the identification of speculative bubbles is presented in Section 2. Section 3 outlines the contours of the two-pillar approach, drawing upon past asset boom and bust episodes to demonstrate the concept. Concluding remarks and suggestions for future research are presented in Section 4.

2. LITERATURE REVIEW – BUBBLE MEASUREMENT AND INFERENCE ISSUES

Speculative bubbles are intuitively recognized to represent situations where market prices significantly exceed the level dictated by fundamentals.[6] Yet broad agreement as to the properties of speculative bubbles has remained elusive virtually ever since the concept of speculation has been invoked. These have not been debates over semantics, but rather quite fundamental issues with important policy implications. For instance how large must be the deviation of prices from those suggested by a fundamental-based model in order for it to be considered 'speculative' or

[3] References to "bubbles" throughout this paper are made in the general sense: (i) of an asset price so high that no reasonable (probability-weighted) future scenario for fundamentals could justify it, and (ii) where the expectation of future short-term price gains drives explosive self-fulfilling increases in prices (and possibly transaction volumes). Section 2 discusses measurement issues in more detail.

[4] It is not uncommon for policy makers to monitor anecdotal information along these lines. For instance, to complement their formal quantitative analysis of home price dynamics, the Reserve Bank of Australia has been known to monitor the number of property seminars held by finance companies targeting retail investors.

[5] This two-pillar approach is analogous to the European Central Bank's (ECB) approach to maintaining price stability by cross-checking both real and monetary developments.

[6] Under this general concept, bubbles could include episodes where prices do not rise at all—for instance where (expectations of) fundamental values collapse but prices are unchanged or decline only modestly. Alternatively, a 'negative bubble' represents a situation where asset prices far undershoot the level implied by fundamentals. For the purposes of this paper and consistent with much of the related literature, our discussion of asset booms and bubbles refers to episodes where prices are rising rapidly (in absolute terms and relative to fundamentals).

DOI: 10.7172/2353-6845.jbfe.2016.2.5

'irrational'?[7] And for how long must the discrepancy between model-predicted and observed prices persist? Reflecting the so-called 'joint hypothesis problem,' how do we know a model of fair value, upon which the determination of a bubble is made, is in fact correctly specified in the first instance? Offering a related defense of his seminal efficient markets asset pricing paradigm, Eugene Fama asserted,

> *"I don't even know what a bubble means. These words have become popular. I don't think they have any meaning … They have to be predictable phenomena … It's easy to say prices went down, it must have been a bubble, after the fact. I think most bubbles are twenty-twenty hindsight. Now after the fact you always find people who said before the fact that prices are too high. People are always saying that prices are too high. When they turn out to be right, we anoint them. When they turn out to be wrong, we ignore them. They are typically right and wrong about half the time … I didn't renew my subscription to the The Economist because they use the word bubble three times on every page. People have become entirely sloppy."*[8]

Theoretical studies have often focused upon the extrapolation of recent capital gains into the expectation of future capital gains, based on the concepts of 'resale optionality' and the self-fulfilling nature of expected future price changes. Analytically, this allows for a clean delineation between the rational and irrational component of asset prices.[9] For instance, Hirshleifer (1977) suggests speculation refers to the purchase (sale) of a good for later re-sale (re-purchase), rather than for use, in the hope of profiting from an intervening price change. Harrison and Kreps (1978) suggest investors exhibit speculative behavior if the right to resell a stock makes them willing to pay more for it than they would pay if obliged to hold it forever. On this basis, an asset bubble exists where investors make a purchase only if they have the ability to subsequently sell the asset at some future date. Kindleberger (1987) defines a speculative bubble as a sharp rise in price of an asset or a range of assets in a continuous process, with the initial rise generating expectations of further rises and attracting new buyers—generally speculators, interested in profits from trading in the asset rather than its use or earning capacity. Stiglitz (1990) defines a bubble where the reason that the price is high today is only because investors believe that the selling price will be high tomorrow—when 'fundamental' factors do not seem to justify such a price. Flood and Garber (1994) categorize a bubble where the positive relationship between price and its expected rate of change implies a similar relationship between price and its actual rate of change. In such conditions, the arbitrary, self-fulfilling expectation of price changes may drive actual price changes independently of market fundamentals. Shiller (2003, pp. 35, 38) describes a bubble in behavioral terms where irrational investors are attracted to an investment because "rising prices encourage them to expect, at some level of consciousness, more price increases. A feedback develops—as people become more and more attracted, there are more and more price increases … the amplification mechanisms that make a bubble grow strong are just that price increases beget more price increases, through human psychology."

From a different perspective, Siegel (2003, p. 14) states formulaically that "a period of rising (or falling) prices in an asset market can be described as a bubble (or negative bubble) at time t if it can be shown that the realized return of the asset over a given future time period, that time period defined by the duration of the asset, can be shown to be inconsistent, i.e. more than two standard deviations from the expected return, given the historical risk and return characteristics of that asset at time t." This is an *ex-post* measure where the real time identification of irrational optimism (or pessimism) is impossible—the presence of a bubble can only be established once

[7] Black (1986) notes that a market could still be considered efficient even if prices deviated in a range of plus 200 percent and minus 50 percent of fundamental value.

[8] Cassidy (2010).

[9] The fundamentally-derived cash flow yield is a stationary process, in contrast to the irrational bubble component which is non-stationary.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

fundamental data have been realized over the maturity of the asset. In acknowledging there will almost always be *ex-ante* and *ex-post* disagreement about the objective measurement of bubbles, Asness (2014) concludes more generally that to have content, the term should indicate a price that no reasonable future outcome can justify.

Other sweeping historical analyses have emphasized the predominant feature of speculative asset bubbles as their tendency to draw in members of the general public who are typically aloof in monetary matters. In other words, bubbles can be distinguished from other episodes through their broader impact on society.[10] Yet as colorful as these socio-behavioral descriptions might seem, an obvious limitation is they are not amenable to formal testing.

Empirical tests of speculative bubbles, including those assessing early warning indicators in the context of financial crises, are forced to contend with other difficult measurement and inference issues. In setting threshold levels for asset price misalignments, policy makers have to balance the tradeoff between false negatives and false positives (see Kaminsky and others, 1998; Alessi and Detken, 2009; and Gerdesmeier and others, 2009). If thresholds are set too high, this will increase the likelihood of failing to predict subsequent busts (Type I errors), while setting them too low can come at the cost of incurring frequent warnings of impending busts that do not materialize (Type II errors). Other complications arising from speculative bubble testing include the problems of small sample sizes (dealing with relatively rare events), the stability of estimated coefficients (in vs. out of sample), and quantitatively accounting for the pervasive irrational behavioral/social phenomema that are emphasized in descriptive accounts of speculative manias.[11]

Broadly speaking, formal tests of speculative asset price bubbles are plagued by estimation and measurement limitations to such a degree that they achieve little of substance in advancing the policy debate over the existence of bubbles in real time—the domain in which policy makers operate. As Gurkaynak, (2005, p. 27) concludes, "Bubble tests do not do a good job of differentiating between misspecified fundamentals and bubbles. This is not only a theoretical concern: For every test that 'finds' a bubble, there is another paper that disputes it … The bubble tests teach us little about whether bubbles really exist or not."

3. METHODOLOGY AND RESULTS – A TWO PILLAR FRAMEWORK TO OPERATIONALIZE BUBBLE SURVEILLANCE

The difficulties associated with identifying asset bubbles in real time suggests a need for policy makers to survey and cross-validate information from a variety of approaches and metrics. Unusually stretched asset prices might reflect rationally lower compensation for risk, the limits to arbitrage, or irrational behavioral errors. It is unlikely any single variable or model specification will ever be able to offer irrefutable evidence, in real time, that an irrational bubble is in progress—by definition, bubbles can only be identified with complete certainty *ex-post*. Without the benefit of hindsight to inform their real time assessments, policy makers need to rely on a diverse and timely suite of measures to highlight the accumulation of financial vulnerabilities and concomitant threats to financial stability. It is in this context that supplemental non-price data might help provide authorities with a richer, more nuanced understanding of risk taking

[10] In an early example, Mackay (1841) states of the seventeenth century Dutch Tulip Mania, "Nobles, citizens, farmers, mechanics, seamen, footmen, maid-servants, and even chimney-sweeps and old clotheswomen dabbled in tulips." The editorial of the *New York Herald* wrote in June 1857 that the U.S. railway boom appeared to "infect all classes of society - the country merchant is becoming a city stockjobber, and the honest country farmer has gone off among the gamblers in western land" (Sobel, 1968, p. 96). As Kindleberger (1978, p. 13) points out in his classic study of speculative manias: "There is nothing so disturbing to one's well being and judgment as to see a friend get rich. When the number of firms and households indulging in these practices grows large, bringing in segments of the population that are normally aloof from such ventures, speculation for profit leads away from normal, rational behavior to what has been described as a mania." Shiller (2000a, p. 2) depicts a society-wide mania as the spread of "psychological contagion from person to person, in the process amplifying stories that might justify the price increase and bringing in a larger and larger class of investors, who, despite doubts about the real value of the investment, are drawn to it partly through envy of others' successes and partly through a gambler's excitement."

[11] Useful summaries of issues encountered in econometric tests of market efficiency and speculative bubbles are presented in Campbell and others (1997) and Gurkaynak (2005).

DOI: 10.7172/2353-6845.jbfe.2016.2.5

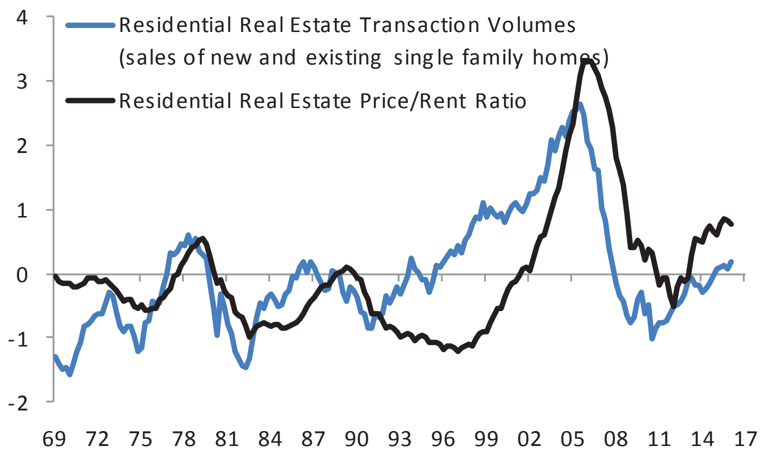
behavior—potentially enhancing the quality of decision making under uncertainty—even if (as is almost certainly the case) this approach falls short of the proverbial magic bullet solution to the difficulties of crisis prediction. In a pragmatic sense, 'perfect should not be the enemy of good.'

In recognition of these issues, this paper proposes a conceptually straightforward surveillance approach based on two distinct though complementary pillars: one that is *price*-based (capturing swings in risk premia or required returns), and another that is *quantity*-based (tracking issuance, transaction volumes, investor fund flows, and investor surveys; see Table 1). Periods where (i) risk premiums have been compressed to abnormally low levels, and (ii) issuance, trading activity, fund flow data, and survey-based return expectations are unusually elevated, are likely to warrant particular attention from policy makers.[12] Though most of the following analysis is focused on capital markets, this framework can be equally applied to real estate markets. To briefly illustrate, Figure 1 depicts the relationship between the rent/price ratio (i.e., the rental yield, a conventional valuation metric) and transaction volumes in the U.S. housing market over the 1969-2016 period. The U.S. housing bubble of the mid 2000s was notable in that it constituted a three standard deviation event not just in valuation terms, but also in (non-price) quantity terms. As outlined below, asset booms characterized by extremely unusual valuations *and* turnover present a significant challenge to benign rational-based explanations. More forcefully, one of the paper's main contentions is that they should put policy makers on high alert.

Table 1
A Two-Pillar Asset Bubble Surveillance Framework

Pricing Pillar (top down)	**Quantities Pillar (bottom up)**
Risk Premia (asset valuations)	Quality and quantity of new issuance
	Trading volumes
	Investor Fund Flows
	Surveys of Expected Returns

Figure 1
The Rent/Price Ratio and Transaction Volumes in the U.S. Housing Market (shown as a z-score)



Source: Author's estimates, Census Bureau

Notes: The z-score depicts the number of standard deviations from the mean. The correlation between the two series is 0.54. Quarterly data from March 1969 – March 2016.

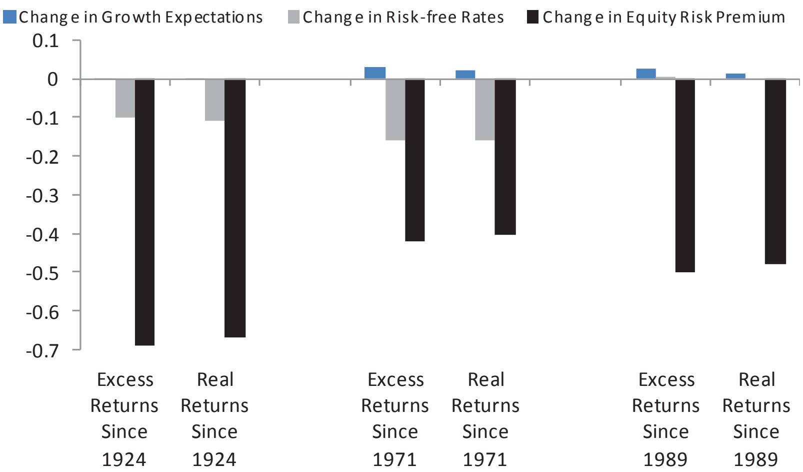
12 This framework is designed to address the 'identification problem.' As to the 'implication problem,' policy makers will need to be guided by other metrics, including but not limited to those assessing the degree of interconnectedness across institutions, credit growth dynamics, potential wealth effects, etc.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

3.1. The Pricing Pillar

Over time, the search for explanations of asset price movements has shifted in focus. Early asset pricing theories emphasized the role of changes in expected cash flows as the key driver of variability in asset prices, with discount rates (comprising risk free rates and a risk premium) assumed constant (Fama, 1970). Unusually high valuation ratios could be justified only by expectations of unusually strong cash flow (dividend or rental) growth in the future. However, subsequent research has showed large asset price increases to have been poor predictors of future cash flow growth (see most recently, Cochrane, 2011; and Williams, 2013), with time variation in discount rates now established as the key source of variation in asset returns. As Figure 2 illustrates, while U.S. stock market returns are strongly negatively correlated with contemporaneous changes in risk premia, the link between stock returns and changes in either long-term bond yields or long-term growth expectations is more muted.

Figure 2
Empirical Determinants of Stock Returns



Source: Author’s estimates, Consensus Economics.

Notes: The chart depicts the correlation of excess and real S&P 500 returns to contemporaneous changes in growth expectations, risk-free interest rates, and the equity risk premium. All data are expressed in monthly terms. The sample periods in the above chart reflect the beginning of the data set (1924), the end of the Bretton Woods system of fixed exchange rates (1971), and the beginning of the dataset on survey-based expectations of real growth and inflation compiled by Consensus Economics (1989). Prior to 1989, long-term nominal growth expectations are proxied by a ten year moving average of nominal GDP growth. The risk free rate is the 10 year Treasury yield. The equity risk premium is an average of three separate measures (see Annex).

For surveillance purposes, valuation metrics (ideally based on real-time information) should demonstrate at least some degree of predictive power over subsequent returns. More importantly, they should display unusual behavior preceding large busts. Figure 3 plots the explanatory power of valuation measures over subsequent returns (since 1953) for the major U.S. asset classes (see the R^2 from regressions of asset returns over different holding periods on initial levels of valuation). The implied equity discount rate is used to forecast stock returns; the rent/price ratio forecasts returns to housing; the credit spread is used to forecast excess credit returns; and the gap between long term Treasury yields and long-term GDP growth forecasts is used to forecast excess Treasury returns (see Annex for details).

In each asset class, valuations appear to have only modest explanatory power over one year returns, but an increasing degree of explanatory power as the investment holding period extends out to a multi-year basis (i.e., asset returns are noisy in the short-run).[13] Of greater consequence

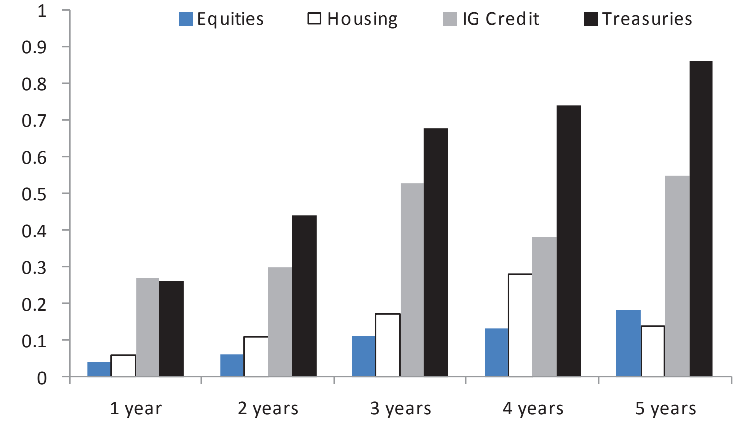
[13] For reasons of data availability, only the results for U.S. asset classes are reported here to demonstrate the concept. Based on the shorter (post-1989) cross-country sample using survey based expectations of long-term growth and inflation, the average R^2 (from regressions of realized returns on model-based expected returns) over a one-year holding period for stocks was 0.11, and 0.18 for bonds. Note this exercise attempts to examine the empirical features of standard return forecasting models – it does not seek to find an ‘optimal’ model *per-se*.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

is that in the years preceding the three largest crashes in history for each of the major asset classes,[14] risk premiums (required returns) declined to unusually low levels—around 1 to 2 standard deviations below the long-term average (Figure 4 and 5). Following a bust, they rapidly reverted back to more normal levels over the next two years. Notably, each bust was followed by recession—on average, six months after a peak in the case of equities, eight months for housing, and around two years for credit and Treasuries (Figure 6).

Figure 3

Valuation-Model based Return Predictability (R2 of valuation models) Across Holding Periods

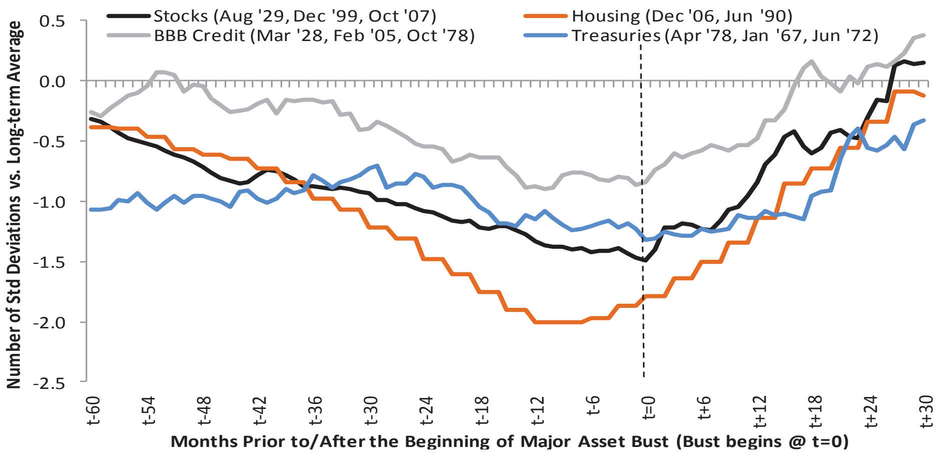


Source: Author's estimates, Haver.

Notes: Chart depicts the R^2 from regressions of real (equities and housing) or excess (credit and treasury) returns (measured across different holding periods) on the starting level of valuations for each asset class. Regressions are based on non-overlapping holding periods over the 1953-2013 sample. In the case of stocks, real returns are regressed on the market-implied real equity discount rate; in the case of housing, real returns are regressed on the rent/price ratio; for investment grade credit, excess returns on BBB credit are regressed on BBB spreads over duration-matched Treasuries; and for Treasuries, 5 year bond returns 5 years forward are regressed on the bond risk premium, measured as the 5y5y rate less the 5y5y forward consensus estimate of growth and inflation. See Annex for more details on the valuation measures.

Figure 4

Risk Premiums (as a z-score) in the Years Before and After Large Busts



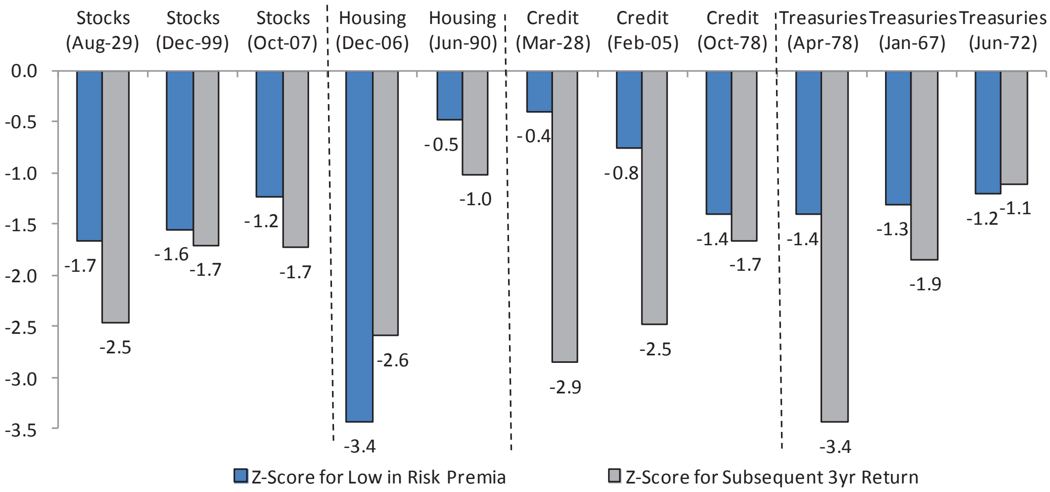
Source: Author's estimates, Haver, Bloomberg.

Notes: Based on the average z-score (depicting the number of standard deviations from the mean) of required returns for the three largest crashes, for each asset class, from five years prior to the commencement of a bust to two and half years after. Dates in parenthesis for each asset class refer to the month in which the bust commenced. Sample period commences in 1924 for stocks and BBB credit, and 1953 for housing and Treasuries, ending in 2013.

[14] For the purposes of this exercise, crashes were defined as the largest decline in real (equity and housing) or excess (credit and Treasury) total return terms, measured over a three-year observation window (rolling monthly). In order to distinguish separate crash episodes, a new regime is signified whenever the three-year change crosses zero (the series is stationary). Figures are based on the largest three crash episodes, except for housing, which experienced only two episodes of negative three-year real returns between 1953 and 2013.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

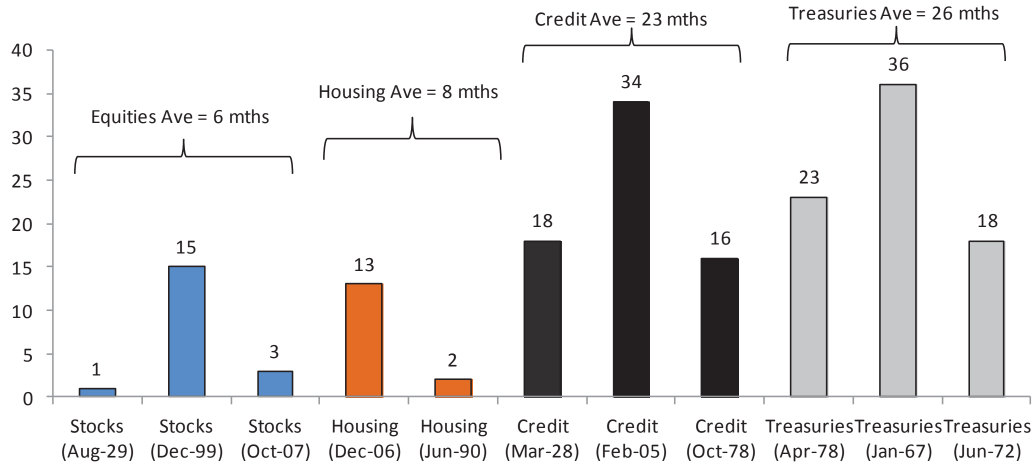
Figure 5
Risk Premiums (as a z-score) Preceding Large Busts, and Subsequent Asset Returns



Source: Author's estimates, Haver, Bloomberg.

Notes: The z-score depicts the number of standard deviations from the mean. Dates in parenthesis for each asset class refer to the month in which the bust commenced. Sample period commences in 1924 for stocks and BBB credit, and 1953 for housing and Treasuries, ending in 2013. Asset returns are measured over a three-year observation window, rolled monthly.

Figure 6
Lead Time between Start of Asset Bust and Onset of Recession (in months)



Source: Author's estimates, Haver, Bloomberg.

Notes: Dates in parenthesis for each asset class refer to the month in which the bust commenced. Sample period commences in 1924 for stocks and BBB credit, and 1953 for housing and Treasuries, ending in 2013. Recession dating is based on the NBER classification.

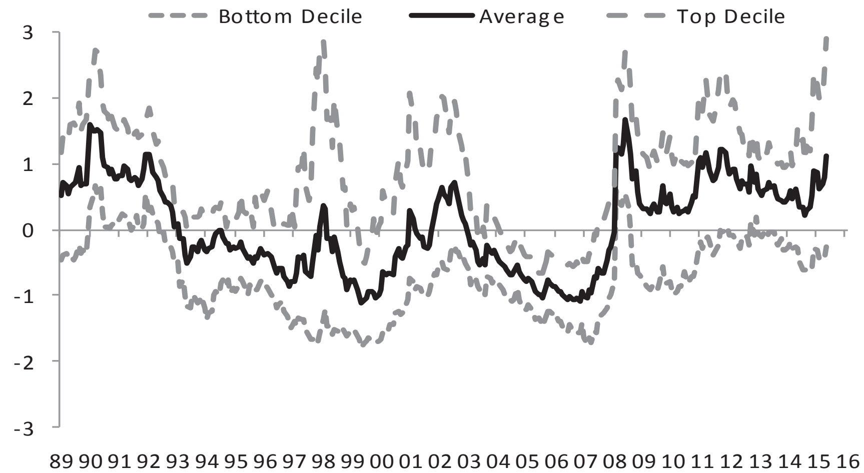
From a broader perspective, policy makers might also glean information as to the generalized environment for risk taking by examining equity and bond valuations across global markets. In this regard, it is interesting to note that based on a sample of 15 developed and 10 emerging markets since 1989, the market-implied real cost of equity (i.e. the required return to hold stocks) for the average country index was more than 1 standard deviation below its historical average prior to the worldwide equity crashes commencing in 2000 and 2008 (Figure 7).[15] In both instances, even the cheapest decile of country equity indices were more expensive than historical averages. World bond markets have not suffered the same type of generalized crash over the post-1989 period,

[15] Proxies for the equity risk premium provide a similar message. I report results here for the market-implied cost of equity rather than the equity risk premium for two reasons: (i) the cost of equity is a superior forecaster of future stock returns than the equity risk premium; and (ii) stocks are a perpetuity with a duration that is constantly in flux, meaning that deflating the cost of equity by a bond yield at one point in the curve could be akin to comparing apples with oranges.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

though it should be noted that the average cross-country measure of risk premia in sovereign bonds (the difference between the 5 year rate, 5 years forward, and consensus estimates for growth and inflation over the same period) is currently close to record lows (Figure 8).[16] Moreover the most expensive decile of countries on this measure are now almost 3 standard deviations richer than their historical norm.

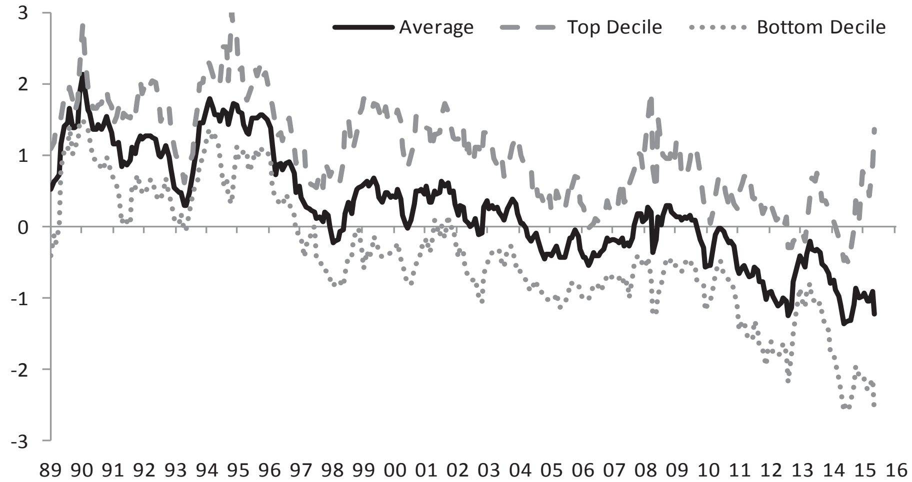
Figure 7
Market-Implied Real Cost of Equity across 25 Countries (as a z-score)



Source: Author's estimates, Consensus Economics, Datastream, Haver, Bloomberg.

Notes: Chart depicts the average, top decile and bottom decline z-score (depicting the number of standard deviations from the mean) for the market-implied real cost of equity, displayed as an average derived from three models: a cyclically adjusted earnings yield; a single stage perpetuity model; and a multi-stage dividend discount model. See Annex for details. The sample consists of 15 developed and 10 emerging markets, comprising 95 percent of the market capitalization of the MSCI All Country benchmark index. Data from September 1989 – January 2016.

Figure 8
'Wicksellian' Sovereign Bond Risk Premia across 25 Countries (as a z-score)



Source: Author's estimates, Consensus Economics, Datastream, Haver, Bloomberg.

Notes: Chart depicts the average, top decile and bottom decline z-score (depicting the number of standard deviations from the mean) for the 'Wicksellian' sovereign bond risk premia, measured as the 5 year rate, 5 years forward, minus consensus estimates for growth and inflation over the same period. See Annex for details. The sample includes 15 developed markets and 10 emerging markets, comprising 94 percent of the market capitalization of the Citigroup World Government bond benchmark index. Data from September 1989 – January 2016.

16 A negative 'Wicksellian' bond risk premium suggests bond yields are too low relative to the neutral rate for the economy, as proxied by long-run estimates of growth and inflation. By examining the relationship in long-term forward space, the impact of near-term monetary policy settings is ameliorated.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

3.2. The Quantities Pillar

The aforementioned results suggest asset pricing models should comprise a key component of the surveillance toolkit. Nevertheless, risk taking behavior, and financial vulnerabilities more generally, might also be reflected in subtler, non-price terms—beyond what asset valuations alone can signify. Additionally, 'top down' asset pricing models are subject to considerable estimation error: forward looking inputs are unobservable and small changes in discount rates for long duration assets can exert a very large impact on estimates of fair value (recall Figure 1). As such, (non-price) quantity data may have a useful cross-referencing role for the purposes of financial surveillance.[17] Against this backdrop, the analysis that follows is based on the 'bottom up' information content in: (i) the quantity and quality of capital market issuance; (ii) trading volumes; (iii) investor fund flows; and (iv) investor surveys. This may be the first proposal for a multi-faceted cross-asset 'quantities pillar' to be employed as a supplement to more traditional valuation-based surveillance analysis.[18] It should be noted however that the absence of lengthy time series and comparable cross country data necessarily makes the following more interpretative than the earlier analysis.

(i) Quantity and Composition of Capital Market Issuance

The financial crisis of 2008 yielded important insights regarding the information content embedded in the increasing quantity and declining quality of capital market issuance. Credit markets are particularly amenable to scrutiny in non-price terms as implicit leverage, subordination, illiquidity, and a paucity of investor protection features allow investors to take on risk in ways that might transcend conventional measurement through spreads alone. Issuance patterns in the structured credit, asset-backed and riskier bond and loan markets underwent profound shifts in the years leading up to global financial crisis, notably in that securitization issuance volumes surged in the most complex, risky, and opaque market segments that had previously played only a peripheral role in the industry (Segoviano and others, 2013).

For instance, between 2000 and 2005, annual U.S. subprime mortgage issuance rose from $100 billion to more than $600 billion, lifting the subprime share of total U.S. mortgage origination from a low of 6.9 percent to a peak of 20.1 percent. These loans featured heavily in the explosive growth in collateralized debt obligations (CDOs). Over the same period, the rapid emergence of new and federally-unregulated players as a force in U.S. mortgage markets saw private-label residential mortgage backed security (MBS) issuance increase from $150 billion to $1.2 trillion, increasing their share of total MBS issuance from 18 percent to 56 percent.[19] Between 2000 and 2007, global issuance of CDOs increased more than six times to $1 trillion, while issuance of CDO-squared product increased eleven-fold to $300 billion. Far from reflecting just a U.S. phenomena, between 2000 and the onset of the crisis, annual European securitization issuance increased from €80 billion to just over €700 billion. More broadly, other forms of high-risk debt issuance also increased markedly, notably high yield bonds, leveraged loans, covenant-lite loans, and payment-in-kind (PIK) notes (Figure 9).[20]

[17] Traditionally, to the extent that non-price data has been examined in the context of asset bubbles, the focus has been almost exclusively on credit growth, dating back to at least Kindleberger (1978) and Minsky (1986, 1992). Yet credit growth can occur in ways that escape conventional measurement (especially in shadow banking), and moreover, bubbles might still present a major threat to financial stability even in the absence of leverage, particularly where their bursting drives risk premiums out to unusually wide levels that prohibit capital formation.

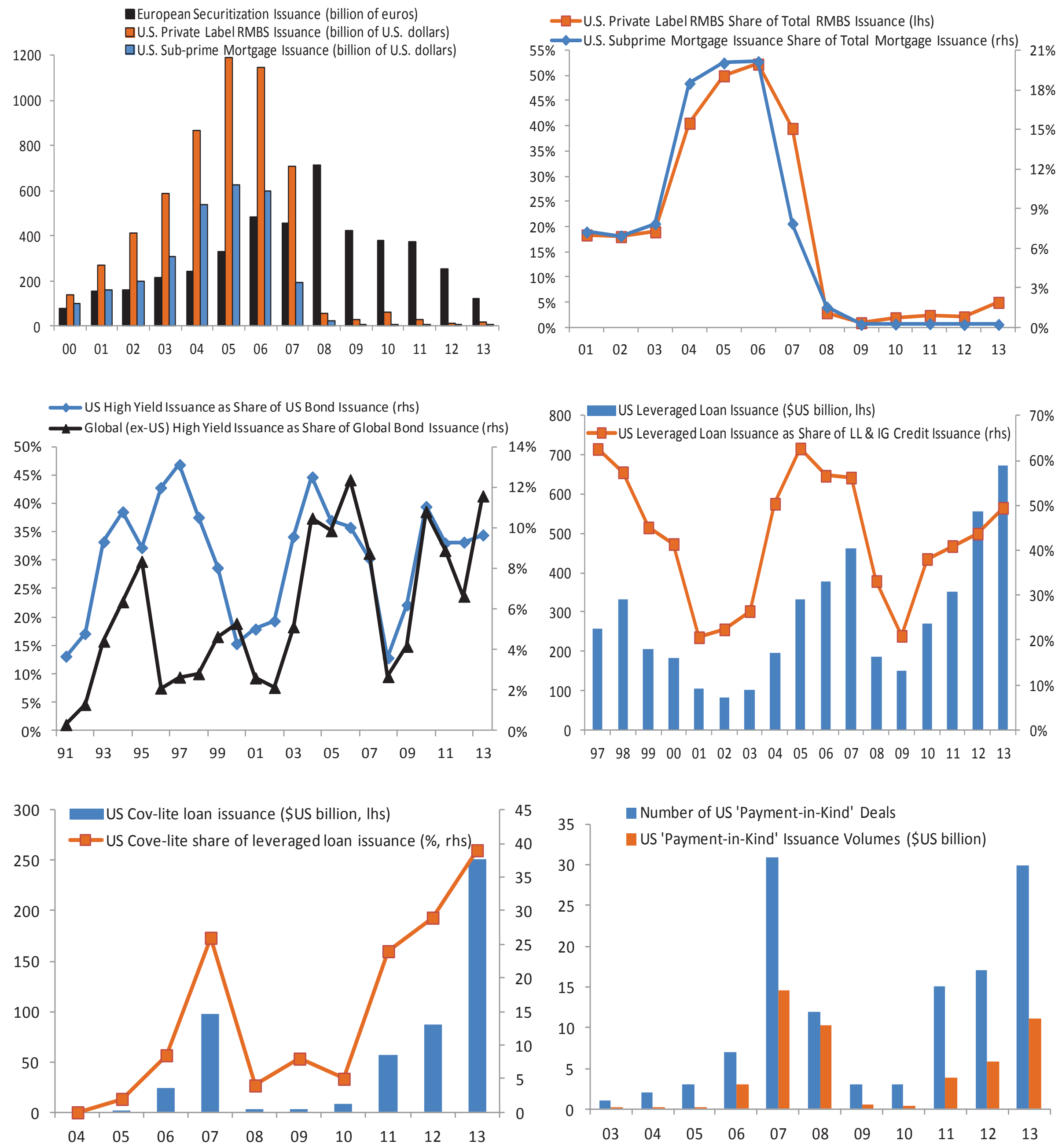
[18] Related analysis of both price and quantity-based data appear in BIS (2012) and IMF (2014). In a more limited context, Stein (2013) examines the financial stability implications derived from patterns in credit issuance, while Feroli and others (2014) focus on investor fund flows.

[19] More than 45 percent of high cost first-time mortgages were originated by independent mortgage companies that were not regulated by federal agencies (Bernanke, 2008).

[20] PIK securities are a financial instrument that pay interest to investors in the form of additional debt or equity instead of a cash coupon. They are attractive to companies who need (or prefer) to avoid making cash outlays to investors. Cove-lite issuance refers to debt obligations which do not contain the usual protective covenants for the benefit of the lending party.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Figure 9
The Pre-Crisis Boom in Risky Credit Issuance



Source: Author's estimates, Inside Mortgage Finance, Dealogic, AFME, SIFMA ,

Based on data back to the 1920s, Greenwood and Hanson (2013) establish that compositional shifts in the pattern of debt issuance between high and low quality firms, including the high yield share of total nonfinancial debt issuance, have had strong predictive power over future corporate bond returns (exceeding the explanatory power of spreads alone). A high and/or rising share of debt market issuance from lower quality firms can portend sharp credit spread widening episodes (Figure 10). As lower quality firms face binding financing constraints (unlike higher quality peers, they do not enjoy an indefinitely open window to raise capital), a relative flurry of lower quality debt issuance can signify unusually easy financing conditions and thus should be interpreted as an inherently procyclical development (Korajczyk and Levy, 2003).[21] Similarly, in a study of the first high yield credit boom and bust cycle of the 1980s, Kaplan and Stein (1993) describe how the

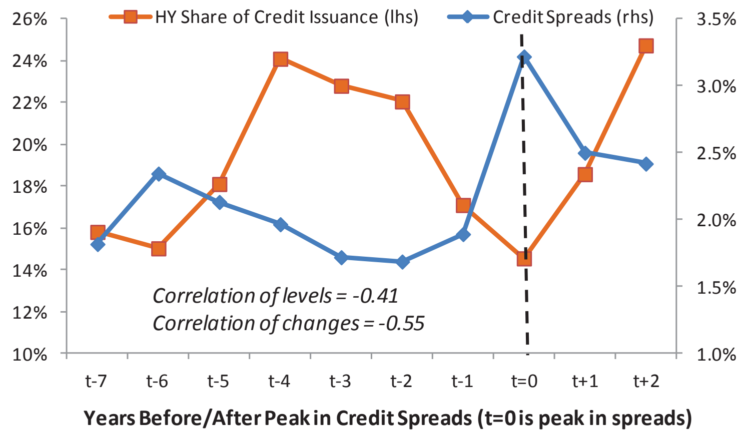
[21] A similar result holds for the debt and equity raisings of relatively small firms (Covas and den Hann, 2006).

DOI: 10.7172/2353-6845.jbfe.2016.2.5

value of leveraged buyouts (LBOs) grew dramatically from just under $1 billion in 1980 to over $60 billion in 1988 (before collapsing below $20 billion in the credit market bust of the following year). This growth was fuelled in part by the introduction of deferred or payment-in-kind interest features in high-yield bonds, which further reduced the protection of already junior creditors and was analogous to investors willingly accepting low or even negative *ex-ante* risk premia.

The implication from these findings is that policy makers need to go beyond simply monitoring the level of credit spreads or the growth in aggregate credit—the composition and quality of credit issuance may be even more important.

Figure 10
High Yield Share of Credit Issuance Before/After Credit Blow-ups



Source: Author's estimates, Greenwood and Hanson (2013).

Notes: Figures are based on an average of the ten largest annual BBB spread widening episodes between 1924 and 2016.

Drawing on the pecking order theory of Myers and Majluf (1984) where corporate managers ('insiders') exploit their informational advantages (in that they know more about the value of their firms than outsiders), patterns of equity issuance that accelerate into a large run up in stock valuations can also be of interest for financial surveillance purposes. Bernstein and Arnott (2003) show that the two largest booms (and busts) in U.S. stock market history—the late 1920s and late 1990s—coincided with the fastest ever rates of aggregate U.S. net equity issuance.[22] Nelson (1999) finds each percentage point of U.S. net share issuance is associated with a market that is 5 percent overvalued relative to historical averages. At the single stock level, inflated valuations have been found to be a key element of IPOs which helps to explain their subsequent long-run underperformance (Ritter, 1991; Schultz, 2003, Baker and Xuan, 2009).[23] Since stock prices are positively correlated, and all firms are incentivized to issue equity when valuations are high, aggregate equity issues tend to cluster around market peaks (Korajczyk and others, 1990; Baker and Wurgler, 2000). Gilchrist and others (2004) show that as stock prices move well above measures of fundamental value, managers rationally respond by issuing new equity which has the effect of reducing the cost of capital and increasing the desire to invest.[24] The corporate

[22] The trend pace of net dilution was 5 percent in the late 1920s, and 3 percent of market capitalization in the late 1990s. The cyclically-adjusted price-earnings multiple peaked at 33 times and 47 times respectively.

[23] Earnings manipulation has also been found to contribute to procyclical issuance at the single-stock level: companies issuing new stock tend to enjoy particularly good earnings and stock price performance prior to an offering, as earnings are managed upwards by incorporating all possible accruals into income. However the accelerating recognition of income simply borrows from the future. Investors fall into the frequent trap of extrapolating the good times to last indefinitely, only to be greeted with a sharp subsequent drop off in stock price performance (Loughran and Ritter, 1995, 1997; Teoh and others, 1998).

[24] Neoclassical investment theories such as 'Tobin's Q' posit a direct, simple link between market valuation and investment decisions: firms invest when the increase in market value due to investment exceeds the costs.

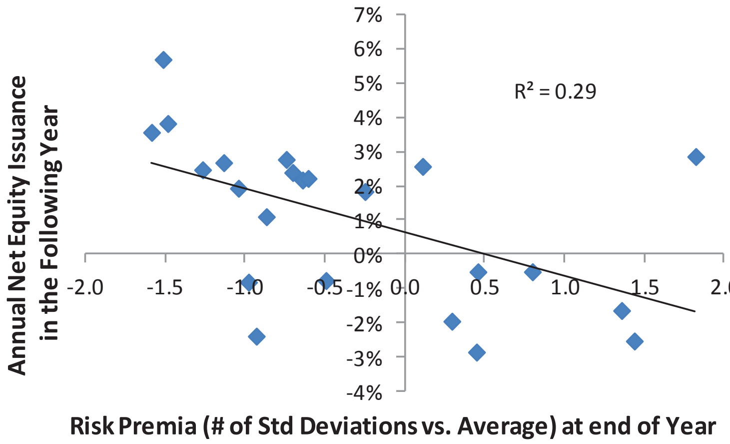
DOI: 10.7172/2353-6845.jbfe.2016.2.5

investment boom in the U.S. accompanying the late 1990s equity market bubble is a notable recent example.[25]

Based on U.S. data since 1965, Figure 11 illustrates the empirical link between an unusually low cost of capital, and elevated net equity issuance (gross issuance less gross buybacks) the following year.[26] Drawing on the experience of the largest stock booms in the U.S., Germany, China and India, Figure 12 similarly shows that initial public offering (IPO) volumes tend to surge when the cost of raising capital (i.e. the required return for investors to hold stock) is unusually low. At the global level, the surge in the number (and value) of global IPOs was especially notably in the late 1990s. These observations reflect the notion that corporate managers are incentivized to issue securities when prices are unusually high relative to fundamentals.[27] Collectively, the procyclical findings on debt and equity issuance suggest a role for policy makers to closely monitor changes in the quantity and composition of capital market financing.

Figure 11

U.S. Aggregate Net Equity Issuance vs. Equity Valuations



Source: Author's estimates, Haver.

Notes: Annual data based on the S&P 500 from 1965- 2013. The regression measures the 'voluntary' equity issuance response (i.e. when real returns have been positive over the past five years) of corporate managers to different levels of risk premia (required return) in the stock market. The required return on stocks (an average of three model outputs, as defined in the Annex) is a statistically significant predictor (at the 1 percent level) of net equity issuance the following year.

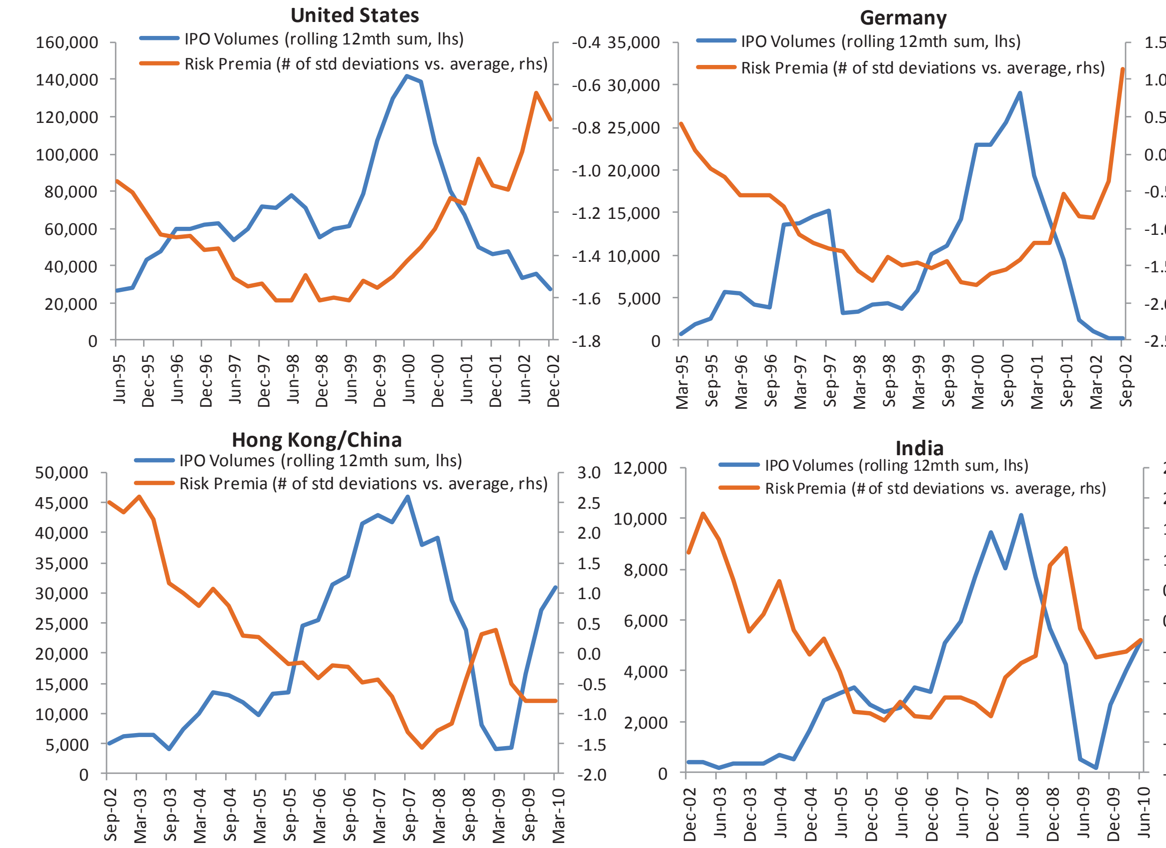
25 From 1992 to 2000, the non-residential investment share of GDP rose from 8.5 to 13.2 percent (a record).

26 Capital raisings occurring around recessions, when valuations are depressed, are not of interest here.

27 In another illustration of the 'insider selling' concept, private equity management teams have also been found to invest a substantially smaller fraction of their net worth in post-buyout equity vis-à-vis pre-buyout equity: managers tend to "cash out" a large fraction of their pre-buyout equity investment at the time of the exiting buyout, and may therefore have an incentive to participate in overpriced exits (Kaplan and Stein, 1993). Kaplan and Stein (1993) also find that ahead of a downturn in company fortunes, financiers with the most intimate knowledge of deals such as relatively risk averse bank lenders and private subordinated debt financiers ('insiders') begin exiting while public subordinated lenders ('outsiders') increase their exposure.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Figure 12
IPO Volumes vs. Stock Valuations in Equity Booms



Source: Author's estimates, Dealogic.

Notes: IPO volumes in millions of $US, shown 5 years prior to and two years after the peak in IPO volumes.

(ii) Trading/Transaction Volumes

Periods of unusually elevated trading activity may also have a role to play in surveillance work. Classical asset pricing theory suggests that where investors share the same beliefs and information, and perceive one another to be rational, the incentive to transact (at the expense of one another) would evaporate. As a result, trading activity would collapse to reflect only unanticipated liquidity and portfolio rebalancing needs of investors. However, these motives seem far too small to account for the enormous trading volumes observed in reality, most notably during periods of rapidly rising asset prices. Though trading volume is frequently relegated to a separate and effectively unconnected area of inquiry from studies of market efficiency, it is hard to imagine a fully satisfying asset pricing model—in either the rational or behavioral genres—that does not give a front-and-center role to trading volume (Hong and Stein, 2007). As Cochrane (2011, p. 1079) observes in a sweeping survey, "every asset price bubble—defined by popular use of the label—has coincided with a trading frenzy, from Dutch tulips in 1620 to Miami condos in 2006."[28]

Figure 13 displays the relationship between the growth in real stock prices and growth in trading volumes across a variety of equity markets during some the largest bull markets in history: the 'Roaring 1920s' episode, Japan's Heisei bubble in the late 1980s, the U.S. technology bubble in the late 1990s,

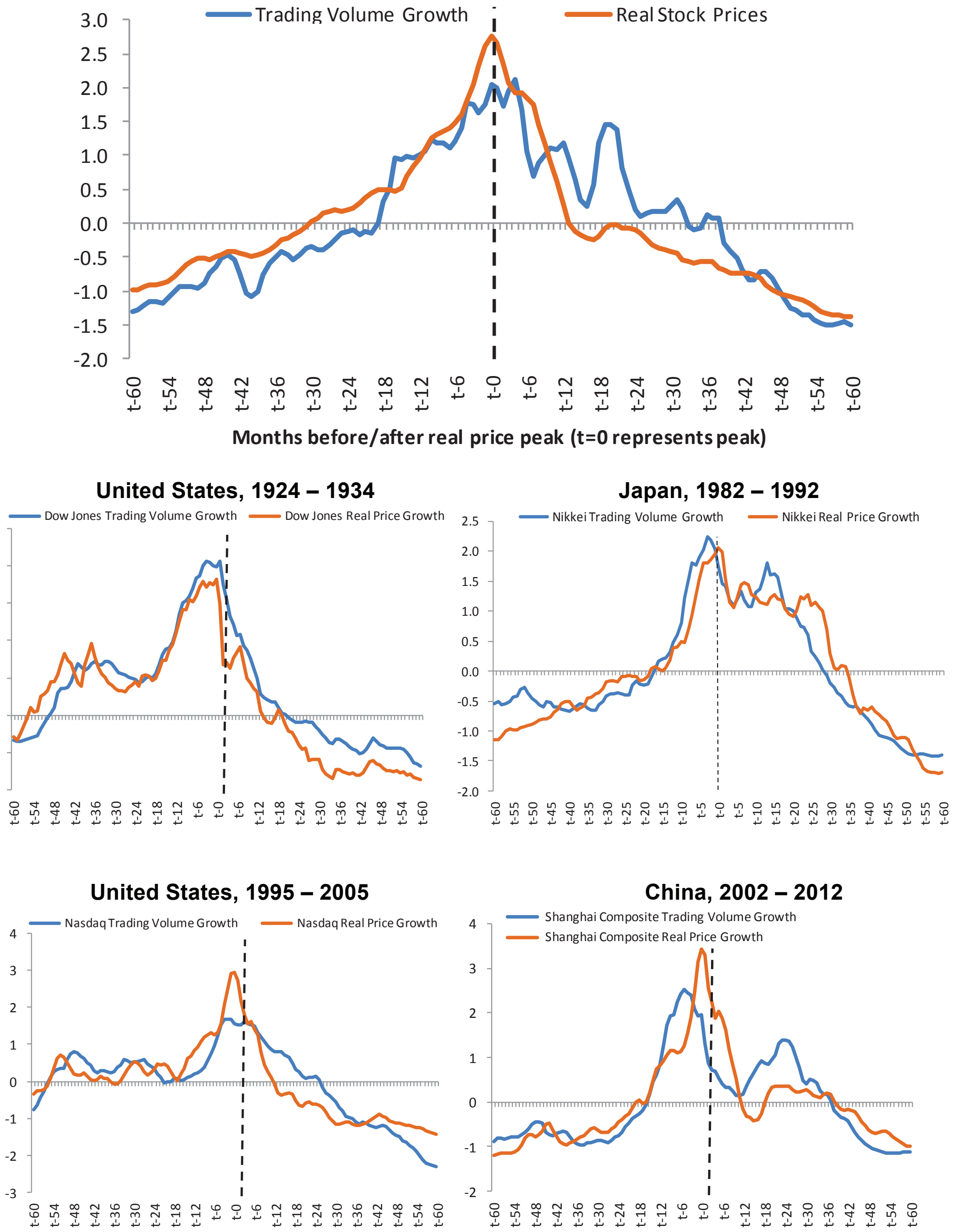
[28] Kindleberger (2011, p. 15) likens speculative manias to a "frenzied pattern of purchases" reflected in rapidly rising trading volumes, a phenomena synonymous with the 'greater fool theory' whereby more and more euphoric investors purchase securities solely in anticipation of future short-term capital gains. For a related discussion of transaction volumes in the Dutch Tulip Mania, see Mackay (1841) and Garber (2001); for the South Sea stock bubble of 1720, see Carlos, Neal, and Wandschneider (2006); for the U.S. stock bull market of the 1920s, see Thomas and Morgan-Witts (1979) and Hong and Stein (2007); for a similar phenomena during the 1990s technology bubble, see Cochrane (2003) Scheinkman and Xiong (2003); Ofek and Richardson (2003); and Hong and others (2006).

DOI: 10.7172/2353-6845.jbfe.2016.2.5

and China's equity boom in the mid-2000s. A similar pattern for transaction volumes unfolded in the case of the recent U.S. housing market bubble (see Figure 1). Average daily trading volumes in the U.S. mortgage backed security market increased five-fold in absolute terms between 2000 and 2008 (and doubled relative to lower risk Treasury and corporate bond markets), while high yield trading activity also rose sharply relative to investment grade credit trading activity prior to the crisis.[29]

Figure 13

Equity Trading Volumes Around Equity Market Booms (as a z-score)



Source: Author's estimates, Bloomberg.

Notes: The top panel depicts the simple average z-score (depicting the number of standard deviations from the mean) for trading volume and real price growth (measured over five years) during the largest real price boom in the following 24 countries: Australia, Canada, Chile, China, France, Germany, Hong Kong, India, Indonesia, Italy, Japan, Korea, Mexico, Netherlands, Norway, Poland, Russia, Singapore, Spain, Sweden, Switzerland, Turkey, UK, and the U.S.

[29] Hong and Sraer (2013) document "quiet bubbles" as a phenomena unique to high grade debt markets, as bonds have a smaller embedded resale option value than infinite-lived assets like stocks and housing, and hence have less disagreement, volatility, and turnover.

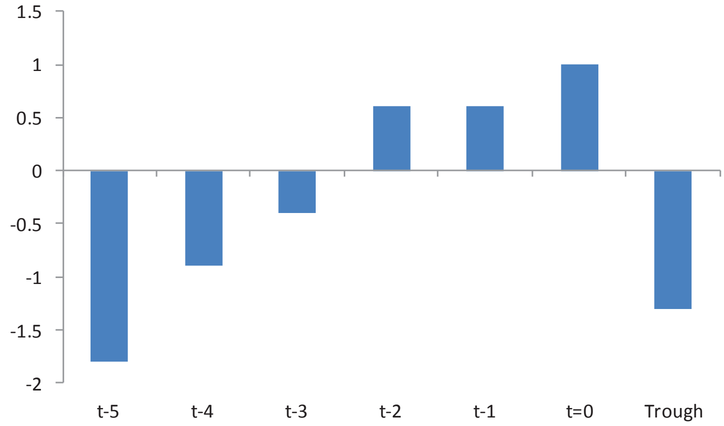
DOI: 10.7172/2353-6845.jbfe.2016.2.5

(iii) Investor Fund Flows

Though cross-border capital flows have long featured in the literature on balance of payments crises, only in recent times have researchers begun to examine the financial stability implications arising from herding and redemption patterns in institutional investment funds. Elevated fund flows can put upward pressure on prices (particularly in small or illiquid asset classes), which in turn attract more flows from underinvested and/or underperforming investors (Chevalier and Ellison, 1997, 1999; Vayanos and Woolley, 2013; Jones, 2015). There are few natural circuit breakers to this feedback loop. The fund flows of unlevered investment managers operating in the 'relative performance derby' vis-à-vis their peer group can be a locus of financial instability to the extent they are motivated by herding considerations (Woolley, 2010; Cai and others, 2012; Feroli and others, 2014; Jones, 2015).[30]

The pattern of investor fund flows into and out of asset classes suggests they should be monitored by authorities. Investor fund flows (based on EPFR data) appear to gradually rise over a period of years and peak at around one standard deviation above average just prior to a large decline in asset prices (Figure 14). In the subsequent bust, cumulative fund flows fall to more than one standard deviation below average before markets trough.

Figure 14
Fund Flows in the Years Preceding a Bust (average z-score across asset classes)



Source: Author's estimates, EPFR, Haver.

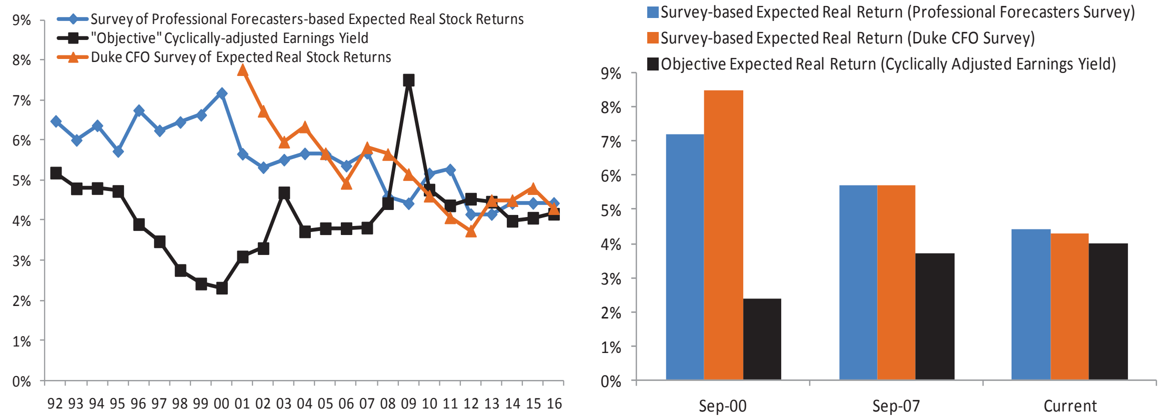
Notes: Average z-score (depicting the number of standard deviations from the mean) of cumulative three-year flows (normalized as a share of total fund assets) across developed market equity, emerging market equity, emerging market hard currency fixed income, emerging market local currency fixed income, global high yield credit, and U.S. Treasuries. A 'bust' is defined as the largest peak-to-trough decline in real returns for each asset class between 1996–2014 (or earliest available).

[30] Positive feedback trading can result from no manager wanting to be last in or last out, with the effect most pronounced in relatively risky and illiquid markets. Chen and others (2010) show that redemptions from mutual funds holding illiquid assets create incentives like those facing depositors in a bank run (see the classic model of Diamond and Dybvig, 1983). Money market funds can face a particularly acute form of vulnerability to runs.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

(iv) Surveys of Return Expectations

Figure 15
Survey-based vs. Objective Measures of Future U.S. Stock Returns



Source: Author's estimates, Duke Quarterly CFO Survey, Philadelphia Fed Survey of Professional Forecasters.

Notes: Latest estimates are as at Q1-2016. 'Objective' measure of future stock returns is the cyclically adjusted earnings yield calculated as the ratio of the ten-year moving average of earnings to stock prices.

To the extent they highlight the irrational extrapolation of past returns into investor estimates of future returns, survey data may also serve as another tool in the asset bubble surveillance toolkit. Shiller (2000b) developed a survey-based indicator to assess whether stock market investors were buying stocks purely on the basis of expectations of a short term increase in the market. Muellbauer (2012) argues that central banks should regularly survey home buyers regarding their expectations for capital appreciation for this express purpose, especially where there might be a large degree of uncertainty over model-based estimates of 'fair value' (as is typically the case for housing). Survey data reveal that after a run up in asset prices, subjective expectations of returns tend to be high while objective expected returns tend to be low (Ilmanen, 2011; Greenwood and Shleifer, 2013; Williams, 2013).

Unlike past cyclical highs in the U.S. stock market, survey-based estimates of expected returns appear relatively benign at the present time (as a relatively new field of research, lengthy time series of investor expectations for risky asset returns have been compiled only for the U.S. stock market). At the peak of the 1990s equity bubble, investor return expectations (as measured by the Survey of Professional Forecasters, and the Duke University CFO Survey) were around three times higher than the (objective) cyclically-adjusted earnings yield (Figure 15). Not only were return expectations high in the Duke CFO survey, they were also rising at the same time the earnings yield collapsed in response to booming stock prices. Just prior to the late-2007 market peak, survey-based return expectations were again rising and well above the earnings yield measure of future returns.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

4. CONCLUDING REMARKS AND FUTURE RESEARCH

The cost of bursting bubbles, and the inability of authorities to identify the accumulation of excesses in asset markets prior to the global financial crisis, suggests the need for an augmented surveillance framework. While acknowledging that the identification of asset bubbles will always remain a difficult task and require some element of subjectivity—a task that cannot be reduced to a single equation—the analysis presented in this paper suggests that a framework anchored in both price and non-price terms offers an encouraging starting point. Asset pricing models, no matter how elaborate, should be the place to begin, not finish, surveillance work. By measuring risk taking behavior, and financial vulnerabilities more generally, in ways that escape conventional asset valuation-based analysis, this paper argues quantity data offer a promising, complementary way to enrich our understanding of asset market dynamics.

The broader issue of what policy makers should do about asset bubbles, while beyond the scope of the present analysis, constitutes a critical related objective of future research. In the aftermath of the global financial crisis, much emphasis has been placed on the role of monetary and macroprudential policy in circumventing conventional leverage-driven asset booms. However, the dimensions of the next crisis will not necessarily follow those of the most recent cycle. Characteristics of the rapidly growing asset management industry—including incentives for asset managers to knowingly 'ride bubbles'—present policy makers with challenges that will likely require examination of a new suite of policy tools if authorities are to mitigate future threats to financial instability. In conjunction with ongoing analysis of early warning surveillance techniques, these areas should be fertile ground for future research.

Acknowledgements

The author is grateful for feedback from colleagues at the International Monetary Fund, in particular Tamim Bayoumi, Olivier Blanchard, Luc Everaert, Gaston Gelos, Shuntaro Hara, Tommaso Mancini-Griffoli, and Miguel Savastano; and for comments from seminar participants at the University of California's Center for Analytical Finance (Santa Cruz). The views expressed herein are those of the author alone and should not be attributed to the IMF, its Executive Board, or its management. The author has no conflicts to declare in support of the publication of this research.

References

Alessi, L., and Detken, C. (2009). Real time early warning indicators for costly asset price boom/bust cycles: A role for global liquidity. Frankfurt, Germany: *ECB Working Paper Series*, *No. 1039.*

Akerlof, G., and Shiller, R.J. (2009). *Animal spirits: How human psychology drives the economy, and why it matters for global capitalism.* Princeton, NJ: Princeton University Press.

Asness, C.S. (2014). My top 10 peeves. *Financial Analysts Journal*, *70*(1), 22–30.

Baker, M., and Wurgler, J. (2000). The equity share in new issues and aggregate stock returns, *The Journal of Finance, 55*(5), 2219-57.

Baker, M., and Xuan, Y. (2009). Under new management: equity issuance and the attribution of past returns. Cambridge, MA: Harvard Business School Working Paper.

Bank for International Settlements. (2012). Operationalising the selection and application of macroprudential instruments. Basel, Switzerland: *CFGS Papers*, 48.

Bernanke, B. (2008). Fostering sustainable homeownership. *National Community Reinvestment Coalition Annual Meetings*, Washington DC.

Bernstein, W.J., and Arnott, R.D. (2003). Earnings growth: The two percent dilution. *Financial Analysts Journal*, *59*(5), 47–55.

Black, F. (1986). Noise. *The Journal of Finance*, *41*(3), 529–543.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Bonner, W., and Rajiva, L. (2007). *Mobs, messiahs and markets: Surviving the public spectacle in finance and politics.* Hoboken, NJ: John Wiley & Sons Inc.
Cai, F., Han, S., and Li, D. (2012). Institutional herding in the corporate bond market. Washington DC: Board of Governors of the Federal Reserve System, *International Finance Discussion Papers, Number 1071*.
Campbell, J.Y., and Thompson, S.B. (2008). Predicting excess stock returns out of sample: Can anything beat the historical average? *Review of Financial Studies*, *21*(4), 1509–1531.
Campbell, J.Y., Lo, A.W., and MacKinlay, A.C. (1997). *The econometrics of financial markets*, Princeton, NJ: Princeton University Press.
Carlos, A., Neal, L., and Wandschneider, K. (2006). Dissecting the anatomy of exchange alley: The dealings of stockjobbers during and after the south sea bubble. Champaign, IL: Working Paper, University of Illinois.
Cassidy, J. (2010, January 13). Interview with Eugene Fama. *New Yorker Magazine*. Retrieved from: http://www.newyorker.com/online/blogs/johncassidy/2010/01/interview-with-eugene-fama.html.
Chancellor, E. (2000). *Devil take the hindmost: A history of financial speculation*. New York, NY: Farrar, Straus and Giroux.
Chen, Q., Goldstein, I., and Jiang, W. (2010). Payoff complementarities and financial fragility: Evidence from mutual fund outflows. *Journal of Financial Economics*, *97*(2), 239–262.
Chevalier, J., and Ellison, G. (1997). Risk taking by mutual funds as a response to incentives. *The Journal of Political Economy*, *105*(6), 1167–1200.
Chevalier, J., and Ellison, G. (1999). Career concerns of mutual fund managers. *Quarterly Journal of Economics*, *114*(2), 389–432.
Cochrane, J.H. (2003). Stock as money: Convenience yield and the tech-stock bubble. In W.C. Hunter, G.G. Kaufman, and M. Pomerleano (Eds.), *Asset Price Bubbles: The Implications for Monetary, Regulatory and International Policies (pp. 174-201)* Cambridge, MA: MIT Press.
Cochrane, J.H. (2011). Presidential address: discount rates. *The Journal of Finance*, *66*(4), 1047–1108.
Covas, F., and den Haan, W.J. (2006). The role of debt and equity finance over the business cycle. Ottawa, Canada: *Bank of Canada Working Paper 2006-45*.
Diamond, D. and Dybvig, P. (1983). Bank runs, deposit insurance, and liquidity. *Journal of Political Economy*, *91*(3), 401–19.
Fama, E.F. (1970). Efficient capital markets: A review of theory and empirical work. *The Journal of Finance, 25*(2), 383–417.
Feroli, M., Kashyap, A.K., Schoenholtz, K., and Shin, H.S. (2014). Market tantrums and monetary policy. Working Paper presented at the 2014 U.S. Monetary Policy Forum, New York, NY. Retrieved from: http://research.chicagobooth.edu/igm/usmpf/2014.aspx?source=igm-em-usmpf14-20140221-initial.
Flood, R.P., and Garber, P.M. (1994). *Speculative bubbles, speculative attacks, and policy switching*. Cambridge, MA: MIT Press.
Fuller, R.J., and Hsia, C.C. (1984). A simplified common stock valuation model. *Financial Analysts Journal*, *40*(5), 49–56.
Gallin, J. (2008). The long-run relationship between house prices and rents. *Real Estate Economics*, *36*(4), 635–658.
Garber, P. M. (2001). "Famous first bubbles: The fundamentals of early manias," *The Journal of Political Economy, 109*(5), 1150–1154.
Gerdesmeier, D., Reimers, H., and Roffia, B. (2009). Asset price misalignments and the role of money and credit. Frankfurt, Germany: *ECB Working Paper Series, No. 1068*.
Gilchrist, S., Himmelberg, C.P., and Huberman, G. (2004). Do stock price bubbles influence corporate investment. New York, NY: *Federal Reserve Bank of New York Staff Reports, No. 177*.
Greenwood, R., and Hanson, S.G. (2013). Issuer quality and corporate bond returns. *Review of Financial Studies, 26*(6), 1483–1525.
Greenwood, R., and Shleifer, A. (2013). Expectations of returns and expected returns. Cambridge, MA: *NBER Working Paper No. 18686*.
Gurkaynak, R.S. (2005). Econometric tests of asset price bubbles: Taking stock. Washington, DC: *Finance and Economics Discussion Series, The Federal Reserve Board Working Paper 2005-04*.
Haldane, A.G. (2014). *The age of asset management*? Speech at the London Business School, London. Retrieved from: http://www.bankofengland.co.uk/publications/Documents/speeches/2014/speech723.pdf.
Hamilton, J.D., and Whiteman, C.H. (1985). The observable implications of self-fulfilling expectations. *Journal of Monetary Economics*, *16*(3), 353–373.
Harrison, M.J., and Kreps, D. (1978). Speculative investor behavior in a stock market with heterogeneous expectations. *Quarterly Journal of Economics, 92*(2), 323–336.
Hirshleifer, J. (1977). The theory of speculation under alternative regimes of markets. *The Journal of Finance, 32*(4), 975–999.
Hong, H., Scheinkman, S., and Xiong, W. (2006). Asset float and speculative bubbles. *The Journal of Finance, 61*(3), 1073–1117.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Hong, H., and Stein, J.C. (2007). Disagreement and the stock market. *Journal of Economic Perspectives, 21(*(2), 109–128.

Hong, H. and Sraer, D. (2013). Quiet bubbles. *Journal of Financial Economics, 110*(3), 596–606.

Ilmanen, A. (2011). *Expected returns: An investor's guide to harvesting market rewards*. Chichester, UK: John Wiley & Sons Ltd.

International Monetary Fund. (2014). Global liquidity—issues for surveillance. Washington, DC: *IMF Policy Paper*.

Jones, B. (2015). Asset bubbles: Rethinking policy for the age of asset management. Washington, DC: *IMF Working Paper 15/27*.

Kaminsky, G., Lizondo, S. and Reinhart, C. (1998). Leading indicators of currency crises. *IMF Staff Papers, 45*(1), 1–48.

Kaplan, S., and Stein, J.C. (1993). The evolution of buyout pricing and financial structure in the 1980s. *Quarterly Journal of Economics*, *108*(2), 313–357.

Keynes, J.M. (1936). *The general theory of employment, interest, and money*. New York, NY: Harcourt Brace.

Kindleberger, C.P. 1978 (2011). Manias, panics and crashes. New York, NY: Macmillan.

Kindleberger, C. (1987). Bubbles. In J. Eatwell, M. Milgate and P. Newman (Eds.), *The new palgrave*: *A dictionary of economics*. London, UK: MacMillan.

Korajczyk, R.A., and Levy, A. (2003). Capital structure choice: Macroeconomic conditions and financial constraints. *Journal of Financial Economics, 68*(1), 75–109.

Korajczyk, R.A., Lucas, D., and McDonald, R.L. (1990). Understanding stock price behavior around the time of equity issues. In R.G. Hubbard (Ed.), *Asymmetric information, corporate finance and investment (pp. 257-277)*. Chicago, IL: University of Chicago Press.

Loughran, T. and Ritter, J.R. (1995). The new issues puzzle. *The Journal of Finance, 50*(1), 23–51.

Loughran, T. and Ritter, J.R. (1997). The operating performance of firms conducting seasoned equity offerings. *The Journal of Finance, 52*(5), 1823–1850.

Mackay, C. 1841 (1852). *Memoirs of extraordinary popular delusions and the madness of crowds*. London, UK: Office of the National Illustrated Library.

Minsky, H. (1986). Stabilizing an unstable economy. New York, NY: Columbia University Press.

Minsky, H. (1992). The financial instability hypothesis. New York, NY: *The Jerome Levy Economics Institute of Bard College Working Paper 74*.

Muellbauer, J. (2012). When is a housing market overheated enough to threaten stability? In A. Heath, F. Packer and C. Windsor (Eds.), *Property markets and financial stability (pp. 73–105)*. Sydney, Australia: Reserve Bank of Australia.

Myers, S.C., and Majluf, N.S. (1984). Corporate financing decisions when firms have information investors do not have. *Journal of Financial Economics, 13*(2), 187–221.

Nelson, W.R. (1999). The aggregate change in shares and the level of stock prices. Washington, DC: Board of Governors of the Federal Reserve System, *Finance and Economics Working Paper Discussion Series 1999-08*.

Ofek, E., and Richardson, M. (2003). Dotcom mania: The rise and fall of internet stock prices. *The Journal of Finance, 58*(3), 1113–1137.

Reinhart, C.M., and Rogoff, K.S. (2009). *This time is different: Eight centuries of financial folly*. Princeton, NJ: Princeton University Press.

Ritter, J.(1991). The long-run performance of initial public offerings. *The Journal of Finance, 156*(1), 3–27.

Scheinkman, J. and Xiong, W. (2003). Overconfidence and speculative bubbles. *Journal of Political Economy, 111*(6), 1183–1219.

Schultz, P. (2003). Pseudo market timing and the long-run underperformance of IPOs. *The Journal of Finance*, 68(2), 483–518.

Segoviano, M., Jones, B., Lindner, P., and Blankenheim, J. (2013). Securitization: lessons learned and the road ahead. Washington DC: *IMF Working Paper 13/255*.

Siegel, J.J.(2003). What is an asset price bubble? An operational definition. *European Financial Management, 9*(1), 11-24.

Shiller, R.J. (2000a). *Irrational exuberance*. Princeton, NJ: Princeton University Press.

Shiller, R.J. (2000b). Measuring bubble expectations and investor confidence. *The Journal of Psychology and Financial Markets, 1*(1), 49–60.

Shiller, R. J.(2003). Diverse views on asset bubbles. In W. C. Hunter, G. G. Kaufman, and M. Pomerleano (Eds.), *Asset Price Bubbles: The Implications for Monetary, Regulatory and International Policies (pp. 35–40).* Cambridge, MA: MIT Press.

Sobel, R. (1968). *Panic on wall street*. New York, NY: Macmillan.

Stein, J.C. (2013). Overheating in credit markets: origins, measurement, and policy responses. *Restoring household financial stability after the great recession: Why household balance sheets matter*. St. Louis, MO: Federal Reserve Bank of St. Louis.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Stein, J.C. (2014). *Incorporating financial stability considerations into a monetary policy framework*. Washington, DC: Board of Governors of the Federal Reserve System, Remarks at the International Research Forum on Monetary Policy.

Stiglitz, J.E. (1990). Symposium on bubbles. *Journal of Economic Perspectives, 4*(2), 13–18.

Teoh, S.H., Welch I., and Wong, T.J. (1998). Earnings management and the underperformance of seasoned equity offerings. *Journal of Financial Economics, 50*(1), 63–99.

Thomas, G., and Morgan-Witts, M. (1979). *The day the bubble burst*. New York, NY: Doubleday & Company.

Vayanos, D. and Woolley, P. (2013). An institutional theory of momentum and reversal. *Review of Financial Studies*, *26*(5), 1087–1145.

Welch, I., and Goyal, A. (2008). A comprehensive look at the empirical performance of equity premium prediction. *Review of Financial Studies, 21*(4), 1455–1508.

Williams, J. (2013). *Bubbles tomorrow and bubbles yesterday, but never bubbles today?* Speech to the National Association for Business Economics,San Francisco, CA. Retrieved at: http://www.frbsf.org/our-district/press/presidents-speeches/williams-speeches/2013/september/asset-price-bubbles-tomorrow-yesterday-never-today/.

Woodford, M. (2003). *Interest and prices*. Princeton, NJ: Princeton University Press.

Working, H. (1962). New concepts concerning futures markets and prices. *American Economic Review, 52*, 431–459.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

ANNEX
Estimates of Required/Expected Returns

Throughout the paper, the terminology 'expected' or 'required' returns is applied to the valuation of housing and stock markets (both of which are real assets with an undefined maturity that can be modeled as a perpetual running yield), while 'excess returns' refer to corporate bond spreads and risk premia in the government bond markets (i.e. nominal assets with a fixed maturity). Valuation estimates are derived as follows:

- *Housing*—constructed as a quarterly index of the ratio of rents to home prices.[31] The rental series is the 'rent of primary residence' published by the Bureau of Labor Statistics. Nominal home price data are based on the series published in Shiller (2000a), with updates available in Haver.
- *Stocks*—expected real returns (into perpetuity) are defined as an equally weighted average of three models for the real discount rate backed out of current prices:[32]
  - (i) The cyclically-adjusted earnings yield, defined as the reciprocal of the ratio of prices to the seven year moving average of annual earnings per share for the MSCI country indices (aside from the longer U.S. series, which is based on S&P500 data back to 1914 from the Robert Shiller website);
  - (ii) The forward-looking single stage Gordon growth dividend model, where: Price = dividend per share / (long-term bond yield + equity risk premium – long term GDP growth). Prices, (cyclically adjusted) dividends, and the long-term risk free rate can be directly observed. Long-term inflation and real GDP growth expectations are based on survey data reported by Consensus Economics beginning in 1989. Prior to this, expectations for nominal growth are proxied by the 10-year moving average of inflation and real GDP growth.
  - (iii) The forward-looking multi-stage 'H-model' of Fuller and Hsia (1984), where the growth rate of earnings per share in the first seven years is based on matching year GDP growth expectations, with an upward or downward adjustment based on the current level of profit margins so as to stabilize the profit to GDP ratio in the steady state. A constant (60 percent) payout ratio is applied to (cyclically adjusted) earnings to ameliorate cross-country differences in dividend taxation policies. The yield curve out to seven years is used to discount the initial set of cash flows. Cash flows beyond seven years are modeled as a constant growth-rate perpetuity based on long-term growth and inflation expectations, and long term bond yields. Using observable spot prices, the pricing equation is solved iteratively such that observed market prices are consistent with future cash flows and discount rates. Long-term inflation and real GDP growth expectations are based on survey data reported by Consensus Economics from 1989. Prior to this, expectations for nominal growth are proxied by the 10-year average of inflation and real GDP growth.

- *Credit*—for U.S. investment grade bonds, duration matched spreads are based on the difference between Moody's seasoned Baa and Aaa bond yield series available via the St Louis Federal Reserve FRED service. For high yield bonds, duration-matched spreads are derived from the Bank of America Merrill Lynch BB+ corporate bond yield series available from Global Financial Data, and US Treasury bond yields.
- *Government Bonds*—motivated in part on Woodford (2003), the 'Wicksellian' bond risk premium is defined as the spread between the 5 year 5 year forward bond yield, and the equilibrium (or neutral) rate, which is proxied by long term consensus expectations of real

[31] Gallin (2008) finds the rent/price ratio to be useful for forecasting future U.S. home prices.

[32] Welch and Goyal (2008) and Campbell and Thompson (2008) run a series of horse races for competing models of equity valuation. Although not the primary purposes of this paper, like these studies, it was difficult to find any specification which could rival the forecasting power of the simple cyclically-adjusted earnings yield, despite it containing no forward looking inputs.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

growth and inflation over the same period (a negative Wicksellian risk premium suggests bond yields are too low relative to the neutral rate, as proxied by long-run estimates of growth and inflation). This is essentially a special case of the Taylor rule for the bond market, where the neutral rate is captured by long term consensus expectations of real growth and inflation, and the inflation and output gap are closed. Wherever possible, zero-coupon bond yields are used (in their absence, the yield to maturity).

Finally, real GDP growth and inflation expectations are measured out to ten years via a survey of professional forecasters compiled by Consensus Economics. These data, though not entirely free of limitations, are interesting from a number of perspectives: they are not subject to revisions or serious time lags as is the case for many macrofinancial time series routinely employed in studies of asset price predictability; they are available on a consistent cross-country basis; and they can capture market expectations of structural breaks or changes in macro variables, upon which estimates of fundamentals are based. To the best of the author's knowledge, this is the first time such data have been used in a comparable study.

DOI: 10.7172/2353-6845.jbfe.2016.2.5

Does it pay to be good? An analysis of vice and virtue stock performance in the Eurozone

Toni Vide
University of Warsaw, Faculty of Management, Poland[1]
videtoni60@gmail.com

Received: 7 March 2016 / Revised: 10 July 2016 / Accepted: 7 August 2016 / Published online: 5 September 2016

ABSTRACT

This paper provides a performance analysis of vice and virtue stocks in the Eurozone for the period between January 2005 and December 2014. In order to do so, a vice index consisting of listed Eurozone companies operating in selected vice industries is created and subsequently matched with a corresponding virtue index, which for the purpose of this analysis is represented by the DJSI Eurozone. The tools used to conduct the performance evaluation are the Sharpe ratio, the capital asset pricing model and the Carhart four-factor model. The analysis indicates no consistent outperformance or underperformance of one or the other index, yet the realised performance over the whole period favours the vice index. Consequently, it can be concluded that from a statistical point of view, there is no substantial advantage or disadvantage in being "good" when investing into stocks, as such it is a matter of investor preference, with the note that historical returns do favour vice stocks.

JEL classification: G11; G15; G17

Keywords: CAPM, Eurozone, four-factor model, Sharpe ratio, virtue stocks, vice stocks.

1. INTRODUCTION

The drive for return maximisation and risk minimisation is a central issue for investors and plays an important role in their actorness on the market. This assumption can be trailed back to the work of Bernoulli (1968) and von Neumann and Morgenstern (1944) and is known as the *expected utility hypothesis*. Yet conducting a more pragmatic examination, we are soon to discover a number of cognitive biases as important determinants in the actions of financial investors. These were prominently described by Kahneman and Tversky (1979) and termed *prospect theory*.

One of the biases intrinsic to the financial investment world are social norms. This type of discrimination in decision-making in the area of economics was first elaborated in Becker's 'The Economics of Discrimination' (1957). He describes how agents bear the economic costs of not interacting with certain people for reasons of social norms imposed onto them by society.

[1] Corresponding author is Toni Vide. Phone +352 661 389 050. Address: Route d'Arlon 226a, L-8010 Strassen, Luxembourg. E-mail address: videtoni60@gmail.com.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Research by Hong and Kacperczyk (2009) suggests that financial investors are indeed willing to pay a price in order to oblige to social norms. In other words, they are willing to forgo higher returns or lower risk by not investing into stocks of industries deemed as a vice, even though this is not a rational choice from an economic point of view.

Continuing from that, we enter two distinct investment strategies: *vice investing* and *virtue investing*.[2] From a semantic point of view, virtue implies moral excellence, goodness and righteousness. More specifically, it is to be understood as investing that generates profit and at the same time produces positive externalities (e.g. environmentalism). A vice, on the other hand implies an immoral or evil habit or practice. In the investment world, it refers to investing into industries with a real or perceived negative externality, with the sole objective of the investor being economic benefits. Finally, it has to be noted that the definition of what exactly constitutes a vice or virtue investment varies and is therefore endogenous. A very clear example of this can be provided when contrasting Christian and Muslim societies. Muslim societies would define interest (*riba*) as a vice (*haram*) (Ahmad, 2015), while Christian societies would not, even though historically, there was some impetus to do so, perhaps most prominently by Thomas Aquinas (for further reading on this topic research the topics *theory of just price* and *concept of usury*).

This paper analyses the performance of vice and virtue investments within the Eurozone in the time period of January 2005 to December 2014 in order to assess which of the two strategies is more efficient. The question examined is whether there is a significant difference in performance between vice and virtue investments. The hypothesis is that there are no significant differences between vice and virtue investments, under the assumption that these assets are correctly valued, that the Eurozone financial market is efficient and finally that the return characteristics of the stocks are derived exclusively through their risk characteristics. While these assumptions do appear to be logical, they do not hold in the case of vice assets, as will be examined in more detail in the literature review. Vice assets are assumed to be a victim of social stigma producing higher returns for their risk profile. The performance of these assets is determined through the utilisation of the Sharpe ratio, the capital asset pricing model (CAPM) and the four-factor model.

The paper is structured in the following manner: Section 2 provides a literature review of the existing research on the topic, Section 3 describes the collection, organisation and preparation of the data for analysis, Section 4 explains the methods used in the performance examination, Section 5 provides the results and their interpretation, Section 6 provides the conclusions and Section 7 provides a list of references.

2. LITERATURE REVIEW

The examination of vice and virtue asset performance is not unprecedented in the academic literature. Examinations of the factors that are indicative for the performance of virtue investments are especially frequent. In this section, a review of the papers deemed most relevant for this study is provided.

We begin with the paper presented by Jo, Saha, et al. (2010), as using only the Sharpe ratio, it employs the most straightforward methodology of performance measurement. The paper compares the DS400 (a virtue index) to the S&P500, and then also compares two funds, the DSEFX (a virtue fund) and the VICEX (a vice fund now known as the Barrier fund) with each other. The findings were that on the long term, the DS400 outperforms the S&P500, while the VICEX outperforms the DSFEX in almost every period (Jo, Saha, Sharma, & Wright, 2010, p. 8).

[2] It should be noted that the terms *vice investing* and *virtue investing* were taken up on a partially arbitrary and partially semantic basis. Therefore, alternative terms are equally valid to describe these types of investing. Alternative terms for virtue investing found in academic literature also include: socially responsible investing (SRI), investing on the basis of environmental, social and corporate governance (ESG) criteria, investing on the basis of corporate social responsibility (CSR) criteria, green investing, sustainable investing, faith-based investing and faith-compliant investing. The only significant alternative term for vice investing found in the academic literature is sin investing.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Lobe and Walkshäusl (2011) conduct a more extensive analysis of vice and virtue asset performance, as they cover the period between 1995 and 2007. They constructed a number of region and industry specific vice and virtue indices. In order to examine their performance, the authors use the Sharpe ratio, the CAPM and the four-factor model. The authors find no evidence for a statistically significant difference between vice indices and their comparables, even after employing the four-factor model analysis. They conclude that choosing vice or virtue bears no significant advantage and is up to the investors' non-financial preferences.

A fairly significant paper, in terms of being cited by the literature reviewing the performance of vice and virtue equities, is that of Hong and Kacperczyk (2009). The question addressed within their study is whether in order to adhere to social norms, investors forgo bigger potential returns by not investing into stocks that can be classified as vice investments. The paper also addresses many contextual questions that give an overall insight into the behaviour and perception of vice stocks, namely that vice stocks receive less analyst coverage and are held in smaller quantities by institutions in comparison to other kinds of stocks. The performance measurement applied by the authors is the four-factor model, which they used to analyse the period of 1965–2006, focusing mostly on Western Europe and North America. They conclude that the vice stocks outperform their comparables and that consequently, social norms have a significant effect on the decision-making of the investors, given that vice stocks are undervalued by up to 20%. The authors suggest that vice firms should rather rely on debt for financing their operations, because of the presumably lower discrimination on that market.

The research conducted by Fabozzi, Ma and Oliphant (2008) is a fairly extensive analysis of vice stock returns, as it covers the timeframe of January 1970 to June 2007 and examines Asia, North America and Europe. After collecting the sample of stocks that were to form the baseline of the research, the authors first computed the simple returns and then the excess market returns as well as the risk-adjusted excess returns. They confirmed that vice stocks do produce abnormal returns, even though these vary significantly from country to country. The reasons given for the outperformance were high potential costs of adhering to social norms for non-vice firms and high barriers of entry into vice industries, which as a consequence comes to facilitate positive monopolistic returns. Further, the evidence provided is consistent with the notion of the undervaluation of vice stocks because of the negative perception of these assets by the average investor.

A fairly interesting approach to the issue can be found in the paper by Salaber (2007), as it focuses mainly on Europe and examines whether religion, litigation risk and the level of excise taxation impact the performance of vice stocks. The author sets up three hypotheses for testing. The first hypothesis states that vice stocks exhibit higher risk-adjusted returns only in Protestant countries. The reason for this is that Protestant countries tend to have stricter regulations regarding alcohol and gambling, going by the assumption that Protestants are less willing to promote vice. As such, they will be more vice-averse or alternatively will require higher returns to justify their investment into vice. The second hypothesis is the assumption that for reasons of negative externalities caused by tobacco, alcohol and gambling, these industries have an increased litigation risk. This should have the effect that these industries yield higher risk-adjusted returns, which has a depressing effect on the stock prices. The third hypothesis is similar to the second and assumes that vice stocks with a high excise taxation have higher risk-adjusted returns. All three hypotheses the author made were confirmed. Salaber states that vice stocks perform best in protestant countries (because of "sin aversion"), countries with a high litigation risk (because of "high external costs") and countries with higher excise taxation. This is a significant finding as previous research based the excess returns of the vice assets solely on them being neglected by investors. Yet because of the limitation of the data, the paper does not provide to what extent these factors are significant.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Finally, research on virtue asset performance indicators is available in far greater quantities than the one examining vice assets. Therefore, two articles that are themselves literature reviews were reviewed, namely the work of RBC Global Asset Management (2012) and Sjöström (2011), both reviewing about 20 papers. The conclusions of these are that there is no disadvantage having a virtue investing strategy in comparison to general investing strategies, yet at the same time there is no indication that a virtue investing strategy would produce abnormal returns, unlike in some of the papers described above regarding the vice investing strategy.

3. DATA

This paper covers a timeframe of ten years, from January 2005 to December 2014, where the data is collected in monthly intervals. These intervals represent the closing price of the stock on the first trading day of a particular month (the reason why the data was not derived for the last trading day is a matter of convenience, as Yahoo! Finance provides only monthly data for the first trading day of a particular month). As a result, 120 data points are accumulated for each variable of the analysis, which is perceived as sufficient to produce conclusive results. The market frame (investable stock universe) of this paper is limited to the Eurozone. The Eurozone is defined as the territory of all member states of the EU that have adopted the Euro as their currency. In order to consider a stock to be a part of this stock universe, the company that has issued the stock needs to be incorporated in one of these countries. If the information for a stock included in the analysis is already available from the starting date of the analysis (January 2005) but the country in question joined the Eurozone only later in time (e.g. Malta joining in January 2008), the stock is only included into the analysis from the date when the country in question joined the Eurozone.

To execute the analysis, the following data had to be collected: information on the historical market development in the Eurozone in the form of a benchmark index (EURO STOXX index); information on the historical development of virtue assets in the form of a virtue index (DJSI Eurozone index); historical stock returns of stocks of industry branches defined as a vice, from which the vice index would be constructed (from this point on referred to as Vicex, which is not to be confused with the former name of the Barrier fund, which also carried that name); the historical risk-free rate; and the factors for the four-factor model.

What regards the question which industries to include as vice industries, the initial intention was to include the alcohol, defence, gambling, nuclear, tobacco and sex industries, as these are most frequently classified as vice industries in consideration of exclusion practices of virtue indices (one of the reasons why the DJSI Eurozone was chosen as the virtue index is its exclusion of these industry categories). In order to represent a vice industry in a substantive manner, the aim was to collect at least ten stocks of any particular vice industry. In some cases, as described below, this was not possible due to the limited presence of certain industries on the stock market.

The tobacco industry could not be included in the analysis as it was found that no public tobacco companies exist in the Eurozone. Overall, four tobacco companies on stocks exist in the European Economic Area: British American Tobacco (UK), Imperial Tobacco (UK), Swedish Match (Sweden) and Japan Tobacco International (Switzerland). The sex industry was found to have a negligible presence on the stock markets, which is in tandem with other literature consulted (e.g. Hong and Kacperczyk 2009, 20). Only two companies from this industry were found, namely Beate Uhse AG (Germany) and Private Media (Spain). Similarly, when searching for gambling companies, only seven such companies were collected. By consulting the Stockholm International Peace Research Institute (SIPRI), nine companies belonging to the defence industry could be found. Therefore, even though the threshold of ten companies was not reached, confidence is high that these stocks represent most if not all of the defence industry's market capitalisation in the Eurozone. Ten alcohol and eleven nuclear industry stocks were collected. The reason for the

DOI: 10.7172/2353-6845.jbfe.2016.2.6

number of nuclear energy stocks is that Siemens AG announced that it would exit the nuclear industry in 2011 as such it had to be removed from the index and was replaced by Areva SA. In total, 39 vice stocks were collected, representing the alcohol, defence, gambling, nuclear and sex industries (for the full list of companies whose stocks were used in this paper, please refer to the Annex, Table A1).

The historical stock information was taken from Yahoo! Finance. Because the market capitalisation was also an information of significance for the analysis and given that in some cases, Yahoo! did not provide this information, Google Finance was also used as a complementary tool of data collection. The factors for the four-factor model were attained from Kenneth French's website, which provides this information for different geographical regions, one of these being Europe. The methodology for the calculation of the rate and the factor loadings can be found on the webpage (French, 2015). The risk-free rate that was taken to be representative for the Eurozone was the yield of the 10-year German government bond (FRED, 2015). The reason for this choice was that it is the biggest and arguably the most stable Eurozone economy, thus being least risky. The reason why the 10-year bond was chosen and not the 30-year bond is that the timeframe of the analysis is 10 years.

Having attained the necessary information, the next step was to construct the individual vice indices as well as the collective Vicex. The specific type of index constructed was the total return index (also called gross return index by STOXX), given that the EURO STOXX and the DJSI Eurozone are also total return indices. This type of index, other than only taking into consideration the price change of the stock, also assumes that the dividends yielded by the stock are reinvested. Because of that, the stock prices used within the analysis are the adjusted stock prices provided by Yahoo! Finance. In practice, this means that they are already adjusted for stock splits and the payments of dividends. The equation used for the computation of the industry indices is:

$$Index_{(t)} = \sum_{i=1}^{n} \frac{p_{it} * w_{it}}{p_{i1}} \tag{1}$$

where $Index_{(t)}$ represents the value of the index at time t, n is the number of companies in the index, p_{it} is the price of company i at time t, p_{i1} is the price of company i at time $t = 1$ and w_{it} is the weight of company i at time t and was computed using the equation below:

$$W_{(i,t)} = \frac{market\ capitalization\ of\ stock\ i}{total\ industry\ capitalization\ t} \tag{2}$$

where the market capitalisation of stock i is the market capitalisation of the company in August 2015 (an explanation for this is given below) and the total market capitalisation t is the total industry capitalisation at time t. Because the capitalisations of individual stocks in specific industries could have excessively disproportional weights, there was a need to cap the weights to a certain maximum. The maximum weight a single stock could take in an index at any point was calculated using the equation:

$$w_{\max (t)} = \frac{200\%}{n_t} \tag{3}$$

where $w_{max\ (t)}$ is the maximum weight the stock could take and n_t is the number of stocks in the index at time t. The value produced from this equation is the maximal percentage of the weight a stock could take. The numerator of 200% is arbitrary, yet there are practical reasons for this

DOI: 10.7172/2353-6845.jbfe.2016.2.6

number. Recalling that the maximal number of stocks an industry portfolio would include at any point is ten or less, if the numerator was set to 100% it is clear that once the portfolio would reach the ten stock benchmark, the maximal weight any stock in the portfolio could take would be 10%, which would in practice mean that the portfolio would become equally weighted. This would consequently eliminate the purpose of weighting.

There are two reasons that necessitated this weighting. First, if we assume that there is no weighting, it means that all companies contribute with an equal effect to the index, which is deemed to not be a realistic representation of the performance of a specific industry, especially if we consider that certain companies have capitalisations that differ from one another 100 times or more. Secondly, if we apply an indiscriminate weighting, it may happen that a single company would influence nearly all movement in the index, which practically eliminates the need to create an index. If we consider that almost all of the variance within an index is derived from one stock alone, it would make more sense to simply examine the single stock. A good example for the need to apply this measure, yet not representative for all industries examined, is the alcohol industry, where the stocks of Anheuser-Busch InBev and LVMH SA together represent more than 80% of the total capitalisation.

In the case where assigning a maximal weight to a stock would have the result that another stock would surpass $w_{max\,(t)}$, that stock was also assigned the maximal weight. In the indices where $w_{max\,(t)}$ was applied, an adjusted $w_{(i,t)}$ was calculated for the rest of the stocks in the portfolio by using the equation:

$$W_{(i,t)} = \frac{market\ capitalization\ of\ stock\ i}{total\ industry\ capitalization\ t - excluded\ capitalization\ t} \tag{4}$$

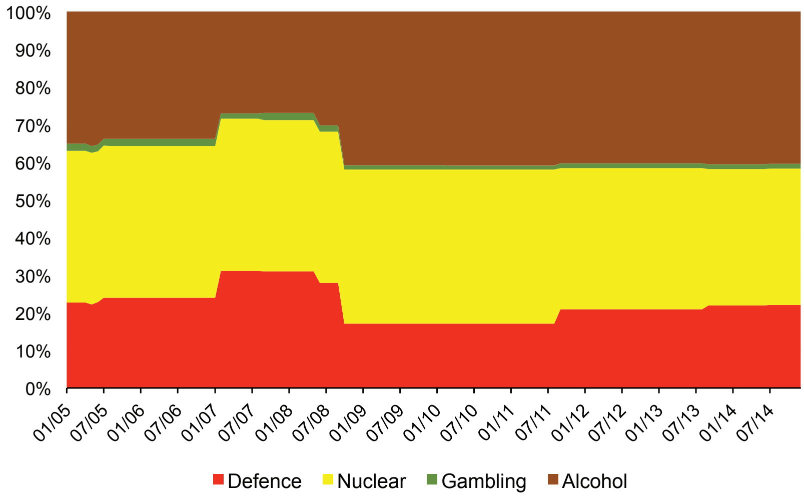
where the excluded capitalisation is the capitalisation of the stocks that surpassed the maximum weight and had to have their weights capped.

While choosing the market capitalisation of August 2015 might not seem outright logical, as the assumption that the proportional capitalisation of individual companies and industries remains constant through time does not hold in practice, this had to be done, as there is a general lack of information regarding the market capitalisation of a particular company in the past. Therefore, this solution was conceived as a compromise, considering that not all the stocks in an industry index have existed for the whole lifetime of the index. It is to be noted that the weights of individual stocks within an index are rebalanced through time as stocks are added to and removed from the index.

It was decided that the divisor for the index would be the price of each individual stock at time $t = 1$. As a result, all indices begin with the value of one. This decision was made as it was felt that starting with a value of one would provide for the cleanest examination of data, yet in the end, the decision was fully arbitrary in nature and in practice, the starting value of the indices could have been any number. The Vicex itself was calculated in the same way as the other vice industry indices and essentially represents a weighted combination of the defence, alcohol, nuclear and gambling indices, as presented in Figure 1.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Figure 1
Industry weights in the Vicex through time



Description: This figure shows how the proportionality of weights in the Vicex developed over time. Time is represented on the x-axis of the index. It should further be noted that the gambling industry constitutes a very small amount of the total capitalisation of the index. Therefore, at no point the gambling industry makes up more than 2% of the index.

The first step of the analysis itself was the calculation of the monthly return of an index. This was achieved using the following equation:

$$r_{(t)} = \frac{P_t - P_{t-1}}{P_{t-1}} \tag{5}$$

where $r_{(t)}$ is the return at time t, P_t is the price of the stock at time t and P_{t-1} is the price of the stock at time t–1. The average returns were calculated for the ten-year period, using the arithmetic mean equation:

$$\bar{x} = \frac{\sum_{n=1}^{k} x_n}{n} \tag{6}$$

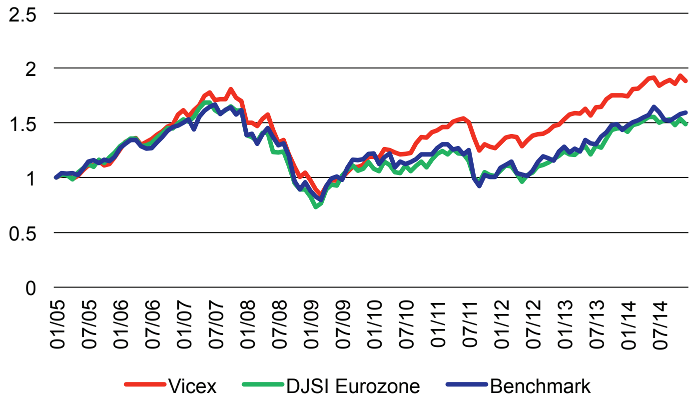
where the average is the sum of all values of returns in the sample, divided by n, which represents the total number of terms in the sample. In practice, the averages were calculated using the Excel command =AVERAGE. The standard deviation represents how much the members of a group in a certain dataset differ from the mean value of the group and as such represents a measure of risk. For the calculation of the standard deviation, the following equation was used:

$$\sigma = \sqrt{\frac{\sum_{i=1}^{n}(x_i - \bar{x})^2}{n-1}} \tag{7}$$

where n is the number of data points, x_i is each value of the dataset and $\bar{x}$ is the mean of all values in the dataset. In practice, the standard deviation was calculated through the usage of the Excel function =STDEV.S.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Figure 2
Vice vs. virtue indices



Description: The figure is a visualisation of the performance of the vice, virtue and benchmarks indices.

4. METHODOLOGY

In order to examine the performance of the indices compared, the Sharpe ratio, CAPM and four-factor model were used. The Sharpe ratio was chosen in order to provide an upfront and easy comparison in the form of a ratio between the selected indices. The CAPM and the four-factor model were selected because they provide for the possibility of a regression analysis through which one can determine if the results have statistical significance and because one is built on the other. Therefore, it is possible to gain insight into the factors that are significant in determining the outperformance of one index over the other. Another reason for choosing these three measurement criteria is that these are most frequently employed in performance measurement in the main articles examined the literature review and as such the comparability of results is enhanced.

The Sharpe ratio was introduced by William F. Sharpe (1966). It is a means of measuring the performance of an asset by adjusting it by its total risk. More specifically, it measures excess return per unit of deviation in an investment. The risk in this case is the total risk of a company, while for example, the Treynor ratio only includes systematic risk. The Sharpe ratio is calculated using the following equation:

$$Sharpe\ ratio_i = \frac{\overline{r_i} - \overline{r_f}}{\sigma_i} \tag{8}$$

where $\bar{r}_l$ is the average return of asset i over the selected time period, $\bar{r}_f$ is the average risk-free rate over the same period and σ is the standard deviation of the selected asset over the given time period.

The CAPM was introduced by a number of authors independently (e.g. Sharpe (1964)). It was elaborated as a tool through which one could determine a theoretically appropriate return on an asset in regard to the systematic risk of the asset. It represents the idea that the investors need to be compensated for the time value of money and risk. The time value of money in this case is represented by r_f, while the (systematic) risk is represented by the so called beta (β). The risk provides us with the assets' sensitivity to non-diversifiable risk. The CAPM is calculated using the following equation:

DOI: 10.7172/2353-6845.jbfe.2016.2.6

$$E(r_i) = r_f + \beta_i(r_m - r_f) \tag{9}$$

where $E(r_i)$ is the expected return of an asset, r_f is the risk-free rate, β_i is the β of the asset in question in regard to the market, and r_m represents the market returns. The term $(r_m - r_f)$ represents the market premium, which is calculated by subtracting the risk-free rate from the market returns.

The main point of interest within the analysis of the CAPM is to find out whether there is a statistically significant positive alpha (α) (also known as the Jensen's alpha in the case of CAPM). If so, this would imply that the index in question outperformed the market (has produced returns that are abnormal). α is calculated through the usage of linear regression in SPSS and received as a constant (intercept).

Because of the well-known and empirically proven fallacies inherent in the CAPM (e.g. different predicted and realised returns, as well as other risk factors for which the CAPM does not account for), there is a need to complement the model with the four-factor model. The four-factor model includes the three factors of the French and Fama three-factor model, which by itself constitutes an improvement over the one factor model (referred to as the CAPM, but also known as the market model) and also includes the momentum factor (MOM) introduced by Carhart (1997).

French and Fama (1993) determined that there are two other asset classes tending to perform better than the market. On the one hand, these are stocks with a small market capitalisation, represented as SMB (which means Small (market capitalisation) Minus Big and describes the size premium one would expect to earn on small caps, which tend to be riskier). On the other hand, these are stocks with a high book-to-market ratio, represented as HML, which stands for High (book-to-market ratio, more popularly termed as the price/book ratio) Minus Low, and describes the circumstance in which companies with a high book-to-market ratio (value stocks) outperform those with low ones (growth stocks). Because these factors could not be explained by the CAPM, the so-called three-factor model was created, which is represented by the following equation:

$$r_t - r_{t,f} = \alpha + \beta_{t,m}(r_{t,m} - r_{t,f}) + \beta_{SMB}(SMB_t) + \beta_{HML}(HML_t) + \varepsilon_t \tag{10}$$

For the means of brevity, only terms not covered when explaining the CAPM are explained: β_{SMB} represents the size loading factor; β_{HML} represents the value loading factor; SMB_t represents the size premium at time t; HML_t represents the value premium at time t; and ε_t represents an error term which can interpreted as the firm-specific risk and as such cannot be explained by the model. The three-factor model has a higher explanatory power than the CAPM, which is not surprising from a statistical point of view, considering that it includes more factors.

To further explain the SMB and HML factor loadings, if $\beta_{SMB} > 0$, it implies that the index in question consists of stocks with a small market capitalisation (small caps) or at the very least that it behaves as if it was made up of such stocks (this holds true for any of the factor loadings described). If $\beta_{SMB} < 0$, then it would suggest that this index is made up of stocks with a high market capitalisation (big caps). If $\beta_{HML} > 0$ the index in question is made up out of value stocks, while a value of $\beta_{HML} < 0$ would indicate that the index in question is mostly made up of growth stocks (Bernstein, 2001).

Building on the three-factor model, the Carhart four-factor model includes a so-called momentum factor (in the model it is represented by MOM, which stands for MOnthly Momentum) into the model as the fourth factor. The momentum factor describes the tendency of stocks that are rising to continue rising and for stocks that are falling to continue to fall further. The four-factor model is defined through the following equation:

$$r_t - r_{t,f} = \alpha + \beta_{t,m}(r_{t,m} - r_{t,f}) + \beta_{SMB}(SMB_t) + \beta_{HML}(HML_t) + \beta_{MOM}(MOM_t) + \varepsilon_t \tag{11}$$

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Continuing from the explanation of the three-factor model, β_{MOM} represents the momentum loading factor and MOM_t represents the momentum premium at time *t*. If $\beta_{MOM} > 0$ it means that the returns of the asset in question were significantly influenced by the momentum factor (alternatively interpreted also as seasonality), while $\beta_{MOM} < 0$ would suggest the absence of such an effect (Carhart, 1997). As in the CAPM, the actual interest of the model is the so-called four-factor α. Achieving such a statistically significant value implies that the index in question produced abnormal returns.

In regard to the statistical analysis of the data, a p-value that is equal or less than 0.1 will be interpreted as statistically significant. Statistical significance will be reported at 10%, 5% and 1% levels. The R^2 measure is also reported (it explains how much of the variance of the assets' risk premium can be accounted by the factors of the model).

The existence of heteroscedasticity (a circumstance where there is a sub-population of variables that have different values from other variables within a population or sample) and autocorrelation (a mathematical representation of the degree of similarity between a given time series and a lagged version of itself over successive time intervals), is also tested for. Heteroscedasticity is tested for by looking at the residual statistics table in order to see whether the mean of the residual is a value significantly higher than zero (if not then there is no heteroscedasticity) and autocorrelation is tested by applying the Durbin-Watson measure. It is deemed that autocorrelation exists if the value of the measure is either lower than 1.5 or higher than 2.5 (University of Minnesota, 2015). Because these two statistical circumstances did not manifest in the frame of the analysis, they are not reported.

5. RESULTS

This section describes the results of the analysis. The numbers ^{1}, ^{5} and ^{10} noted in the upper right corner of the values produced through linear regressions signify the p-value of a certain variable (e.g. 1 means that the output is significant within the 1% level, thus the p-value was $p \geq 0.01$). If no such number is reported, it means that the value in question is statistically not significant (thus the p-value is $p > 0.1$). The coefficients of the analysis are interpreted only within the description of the ten-year period, as these factor loadings are deemed to be the most representative. The information provided on the performance of the benchmark is there only for reference reasons.

Looking at the ten-year period and observing the Sharpe ratio, we can see that several vice assets have substantially beaten the benchmark. The DJSI just barely did not manage to beat the benchmark. The clear outperformer of the period is the gambling industry, indeed so much so that its returns are statistically significant in both the CAPM and the four-factor model. Yet the model does not very well explain the returns of this industry with R^2 being only 7%. At the same time, we cannot find any statistically significant alphas for other indices analysed (observing the four-factor model). The conclusion that we can draw from this is that there is no general tendency for vice or virtue industries to outperform the market in a consistent manner.

Looking at the statistically significant coefficients in the four-factor model table, we can see that the virtue index consists largely of big companies (unsurprising given the nature of the DJSI) unaffected by seasonality and is made up mostly of value stocks. This by itself is interesting as previous research suggests that virtue assets should mostly behave like growth stocks. At the same time, the only other conclusion that one can make for the Vicex is that it is made up of growth stocks, even though substantially less than the virtue index. Looking at the vice industries, we can see that in general they are unaffected by cyclical business (low or insignificant MOM factor loading), they mostly behave as value stocks (high HML factor loading, except for the alcohol and sex industries), they significantly differ from industry to industry what regards the size of their constituents (high factor loading for the sex industry and low factor loading for the nuclear industry) and finally that they are far less volatile than the market.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

Table 1
Sharpe ratio and CAPM January 2005–December 2014

	Avg. Ret.	Std. Dev.	Sharpe	CAPM (α)	CAPM (β)	CAPM (R^2)
Defence	0.99%	0.07	0.13	.601	.648[1]	.273
Nuclear	0.41%	0.05	0.06	.067	.542[1]	.371
Gambling	1.18%	0.06	0.18	.948[10]	.284[1]	.067
Sex	-1.52%	0.10	-0.16	-1.766[10]	.279	.021
Alcohol	0.74%	0.05	0.13	.411	.507[1]	.313
Vicex	0.61%	0.04	0.12	.277	.522[1]	.488
DJSI	0.46%	0.05	0.07	.074	.636[1]	.473
Benchmark	0.54%	0.05	0.08	/	/	/

Notes: "Avg. Ret." stands for average return; "Std. Dev." stands for the standard deviation; "Sharpe" stands for the Sharpe ratio; "CAPM (α)" stands for the alpha of the CAPM as given by the intercept; "CAPM (β)" stands for the beta of the CAPM; "CAPM (R^2)" is the *r squared* measure as provided by the CAPM.

Table 2
Four-factor model January 2005–December 2014

	α	β	SMB	HML	MOM	R^2
Defence	.405	.584[1]	.666[5]	.509[10]	.265[10]	.328
Nuclear	.037	.445[1]	-.331[10]	.901[1]	.130	.528
Gambling	1.052[10]	.247[5]	.099	-.016	-.114	.073
Sex	- .986	-.220	1.159[5]	1.199[5]	-.704[1]	.196
Alcohol	.390	.572[1]	-.236	-.282	-.008	.330
Vicex	.237	.507[1]	-.176	.276[10]	.073	.514
DJSI	.268	.525[1]	-.663[1]	.699[1]	-.131[10]	.671

Notes: α stands for the alpha of the four-factor model as given by the intercept; β, SMB, HML and MOM are the factor loadings of the four-factor model and explain the return behaviour of the index in question. Finally R^2 is the *r squared* measure as provided by the four-factor model.

6. CONCLUSIONS

Looking at the realised returns and primarily at the Sharpe ratio, we can conclude that vice investments did perform better than virtue investments, yet observing the expected returns derived from the alpha value, we can see that no statistically significant outperformance of vice or virtue assets can be detected. As such it is taken that the hypothesis which was set out is confirmed—by default one should not expect that vice assets would outperform virtue assets or vice versa. One exception of this was the alpha of the gambling index, yet the returns of this industry are not well explained by the four-factor model and CAPM, so it remains unclear why the industry performed so well.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

The findings of this study are similar to the findings of Lobe and Walkshäusl (2011) and conclude that there is no significant advantage or disadvantage when applying either investment strategy, or at least that one should not expect one strategy to outdo the other by default. One question that would be interesting to approach is whether the results would be different when continuing the research of Salaber (2007), which suggests that vice assets would normally excel in performance in Protestant countries. By her definition, most of the countries within this study are Catholic, where vice assets tend to perform worse given the reduced neglect effect, thus extending the scope of this study to the United Kingdom and Sweden could substantially enhance the conclusions.

While the results of this study were significant, the research could be further improved by including delisted companies into the list of companies and thus avoiding the effect of the survivor bias (explained through the notion that a company that survived for a significant amount of time also had to be successful to a certain extent) and by using the real (historical) market capitalisation weights instead of the generalised weights used in this paper. Both of these considerations could have been very easily addressed by accessing databases such as Bloomberg, Capital IQ or DataStream, yet considerable financial resources would have been required in order to do so.

REFERENCES

Yousri Ahmad, A.-R. (2015). *Riba, Its Economic Rationale and Implications*. Retrieved from http://www.islamic-banking.com/iarticles_8.aspx

Becker, G. S. (1957). *The Economics of Discrimination*. Chicago: University of Chicago Press.

Bernoulli, Daniel (1968). *Specimen theoriae novae de mensura sortis*. Gregg, Farnborough

Bernstein, W. J. (2001). Rolling Your Own: Three-Factor Analysis. Retrieved from http://www.efficientfrontier.com/ef/101/roll101.htm

Carhart, M. M. (1997). On Persistence in Mutual Fund Performance. *The Journal of Finance*, *52*(1), 57–82. doi:10.1111/j.1540-6261.1997.tb03808.x

Fama, E. F., & French, K. R. (1993). Common risk factors in the returns on stocks and bonds. *Journal of Financial Economics*, *33*(1), 3–56. doi:10.1016/0304-405X(93)90023-5

FRED. (2015, August 11). *Long-Term Government Bond Yields: 10-year: Main (Including Benchmark) for Germany©*. Retrieved August 11, 2015, from https://research.stlouisfed.org/fred2/series/IRLTLT01DEM156N

French, K. R. (2015). *Description of Fama & French European Factors and Portfolios*. Retrieved August 15, 2015, from http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/Data_Library/details_european.html

Hong, H., & Kacperczyk, M. (2009). The price of sin: The effects of social norms on markets. *Journal of Financial Economics*, *93*(1), 15–36. doi:10.1016/j.jfineco.2008.09.001

Jo, H., Saha, T., Sharma, R., & Wright, S. (2010). Socially Responsible Investing vs. Vice Investing. In AABRI (Ed.), *Academic and Business Research Institute Conference – Las Vegas 2010*.

Kahneman, D., & Tversky, A. (1979). Prospect theory: An analysis of decision under risk. *Econometrica*, *47*(2), 263. doi:10.2307/1914185

Lobe, S., & Walkshäusl, C. (2014). Vice versus virtue investing around the world. *Review of Managerial Science*, *10*(2), 303–344. doi:10.1007/s11846-014-0147-3

RBC Global Asset Management. (2012). *Does Socially Responsible Investing Hurt Investment Returns?* Retrieved from http://funds.rbcgam.com/_assets-custom/pdf/RBC-GAM-does-SRI-hurt-investment-returns.pdf

Salaber, J. (2007). *The determinants of sin stock returns: Evidence on the European market*. Retrieved from BIRD, https://basepub.dauphine.fr/handle/123456789/3028

Sharpe, W. F. (1964). Capital asset prices: A theory of market equilibrium under conditions of risk. *The Journal of Finance*, *19*(3), 425–442. doi:10.2307/2977928

Sharpe, W. F. (1966). Mutual fund performance. *The Journal of Business*, *39*(1), 119–138. doi:10.2307/2351741

Sjöström, E. (2011). The performance of socially responsible investment – A review of scholarly studies published 2008-2010 by Emma Sjöström: *SSRN*. doi:10.2139/ssrn.1948169

University of Minnesota. (2015). *Durbin Watson Test*. Retrieved August 30, 2015, from University of Minnesota: http://www.tc.umn.edu/~ryoox001/images/DurbinWatson_test.pdf

Von Neumann, J., & Morgenstern, O. (1944). *Theory of games and economic behavior*. United States: Princeton University Press.

DOI: 10.7172/2353-6845.jbfe.2016.2.6

ANNEX: LIST OF COMPANIES

Table A1
List of companies used to construct the vice index

#	Name	Country	Industry	Date	Market Cap.
1	Anheuser-Busch InBev	Belgium	Alcohol	10/2008	176.82 bn
2	Brauerei Ottakringer	Austria	Alcohol	1/2005	0.23 bn
3	C&C Group plc	Ireland	Alcohol	1/2005	1.18 bn
4	Davide Campari	Italy	Alcohol	5/2005	4.2 bn
5	Groupe Laurent-Perrier	France	Alcohol	1/2005	0.49 bn
6	Heineken	Netherlands	Alcohol	1/2005	19.34 bn
7	Lanson-BCC	France	Alcohol	1/2005	0.25 bn
8	LVMH SA	France	Alcohol	1/2005	88.19 bn
9	Pernod Ricard	France	Alcohol	6/2008	29.28 bn
10	Rémy Cointreau SA	France	Alcohol	1/2005	3.11 bn
11	Airbus Group SE	France	Defence	2/2007	52.37 bn
12	CNH Industrial N.V.	Netherlands	Defence	9/2013	11.34 bn
13	Dassault Aviation SA	France	Defence	1/2005	17.45 bn
14	Fincantieri S.p.A.	Italy	Defence	7/2014	1.24 bn
15	Finmeccanica SpA	Italy	Defence	7/2005	7.62 bn
16	Rheinmetall AG	Germany	Defence	6/2005	2.18 bn
17	Safran SA	France	Defence	1/2005	29.19 bn
18	Thales SA	France	Defence	1/2005	12.91 bn
19	ThyssenKrupp AG	Germany	Defence	1/2005	13.49 bn
20	Bet-At-Home.com	Germany	Gambling	8/2005	0.291 bn
21	GOFP S.A.	Greece	Gambling	1/2005	2.33 bn
22	Groupe Partouche SA	France	Gambling	1/2005	0.2 bn
23	Mybet Holding	Germany	Gambling	1/2006	0.024 bn
24	Paddy Power plc	Ireland	Gambling	1/2005	3.62 bn
25	Snai S.p.A.	Italy	Gambling	1/2005	0.15 bn
26	Unibet SDR	Malta	Gambling	9/2007	1.95 bn
27	Ansaldo STS S.p.A.	Italy	Nuclear	3/2006	1.88 bn
28	Areva SA	France	Nuclear	9/2011	3.25 bn
29	Bouygues	France	Nuclear	1/2005	11.61 bn
30	E.ON	Germany	Nuclear	12/2007	24.18 bn
31	Electricite de France SA	France	Nuclear	11/2005	40.75 bn
32	Endesa SA	Spain	Nuclear	1/2005	20.62 bn
33	Enel SpA	Spain	Nuclear	1/2005	41.15 bn
34	Engie SA	France	Nuclear	1/2005	43.62 bn
35	Fortum Oyj	Finland	Nuclear	1/2005	14.33 bn
36	Iberdrola	Spain	Nuclear	5/2005	40.5 bn
37	Siemens AG	Germany	Nuclear	1/2005	87.39 bn
38	Beate Uhse AG	Germany	Sex	1/2005	0.036 bn
39	Private Media Group	Spain	Sex	3/2008	0.00041 bn

Description: The companies are sorted by the alphabetical name of their contextual industry and afterwards by their name. The date is the date on which the company stock data became available, and was thus included into the vice index. The market capitalisation is indicated in Euros.

DOI: 10.7172/2353-6845.jbfe.2016.2.6