Is There a Global Financial Cycle in the Stock Market? The Case for Developed Economies Since the Interwar Years.

*Germán Forero-Laverde*

Abstract

This paper explores the global financial cycle (GFCy) hypothesis by extracting a common component from indices that represent the behaviour of the stock markets of twelve advanced economies from 1922 until 2013. To further test the hypothesis, this common component is compared to the behaviour of the US stock market. Finally, we tested whether this GFCy is consistent with the usual story of international financial market disintegration that occurred worldwide during the Bretton Woods period. We find evidence of a common factor that explains over 50% of the variance-covariance matrix of the stock markets of advanced economies and is, consistently, a significant driver of stock valuations to different time horizons. Additionally, we find that this factor is strongly correlated, contemporaneously, with the behaviour of the US stock market, and that this correlation is stable throughout their joint empirical distributions. Finally, we find evidence that the GFCy is present regardless of the policy choices on exchange rates and on capital controls, which contradicts the story of financial integration/disintegration. We suggest three possible mechanisms for the presence of the GFCy. First, under open capital accounts, global capital flows suggest the presence of a “communicating vessels” phenomenon. Second, in the more recent period, companies that trade in several stock markets at the same time may explain part of the commonality. Finally, in scenarios of financial repression, we expect that information flows, and a cycle of global risk aversion may explain common movements in the valuation of equity.

Keywords

Global financial cycle; Stock markets; Market co-movement; Economic history

JEL Codes

C14, E60, F33, G15, N20

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Introduction

The literature on the history of international capital markets, spearheaded by Obstfeld & Taylor (2004), describes the evolution of global financial integration as a “U” shaped process. It first peaked during the classical gold standard (1870-1914), troughed during the Bretton Woods years (1944-71), and peaked again in the present day, particularly after the wave of deregulation of the 1980s.\(^1\) This compelling story has been built using the framework of the macroeconomic trilemma, which posits that policymakers must choose two out of three competing objectives: free capital flows, stable exchange rates, and an autonomous monetary policy (Obstfeld & Taylor, 1997; Obstfeld, Shambaugh & Taylor, 2005; Obstfeld & Taylor, 2017). The underlying assumption is that under closed capital accounts, such as occurred in the post-war arrangements of Bretton Woods, international financial markets disintegrate.

However, in recent work Hau & Rey (2006), Rey (2015, 2016), Miranda-Agrippino & Rey (2015) and Passari & Rey (2015) have argued that the macroeconomic trilemma is indeed a dilemma since there exists a global financial cycle (GFCy), proxied by the VIX index, that explains over 25% of the volatility of risky assets. In their view, this global financial cycle, for which they find evidence since the 1990s, operates as a central mechanism in the transmission of investor expectations, their pricing and perception of risk and, consequently, it renders the possibility of an autonomous monetary policy moot in the long run. Additionally, they show that the GFCy exists regardless of the exchange rate regime in place, raising further challenges to the idea of the impossible trinity.

To address these conflicting pieces of evidence, we propose two separate research questions that will drive the rest of the paper. First, is there evidence for the GFCy that explains the evolution of the prices of risky assets in the long-run? To tend to this issue, we use principal component analysis to identify a common cycle in the stock markets of twelve advanced economies between 1922 and 2013.\(^2\) Additionally, we confirm through a standard asset pricing model that this cycle has significant explanatory power in a panel data setting. Furthermore, using a time series specification, we contrast whether this common component can be proxied by the behaviour of the US stock market as measured by the Standard & Poor’s 500 index. Second,

\(^1\) Bordo (2017) reflects on whether the current level of financial integration has surpassed that of the classical gold standard.

\(^2\) Most historical studies that use the framework of the trilemma begin in the last third of the nineteenth century, with the start of the classical gold standard. However, we argue that the rigours of the straitjacket imposed by the macroeconomic trilemma were not entirely evident during this period. According to Eichengreen (2008), it was not until the interwar period, when the increasing demands for a domestic policy that fostered full employment and growth made themselves evident through social movements and the rise of labour parties in many European countries. During the classical gold standard (1870-1914) countries with open capital accounts and fixed exchange rates, while evidencing the trade-off of losing monetary policy autonomy, did not care much for the political cost of forfeiting its use. Consequently, we do not expect to find a role for the trilemma regime prior to the 1920s.
is the behaviour of the GFCy consistent with the story of financial integration/disintegration posited by Obstfield & Taylor (2004)? To tend to this issue, we perform a structural break test of the coefficients in the pricing model under different configurations of the macroeconomic trilemma. Additionally, we test whether the break exists only for the GFCy coefficient.

Regarding the first question, we find evidence of a common factor that explains over 50% of the variance-covariance matrix of an indicator that measures expansions and contractions in the stock markets of the countries in our database. This common cycle bears a positive and statistically significant coefficient to different time horizons, even after controlling for a variety of factors. Moreover, we find that this factor is strongly correlated, contemporaneously, with the behaviour of the US stock market, and that this correlation is not driven by extreme cases but instead is stable across their joint empirical distributions. Consequently, the main contribution to the global cycle debate is to find evidence of a common factor that drives the stock market of advanced economies and providing an economic interpretation for it.

Concerning the second question, we find evidence of a joint break in all the coefficients of the valuation model by trilemma regime in the short-run. However, the coefficients for the GFCy to different time horizons and under different trilemma configurations are statistically equal. This shows that the common factor across stock markets is present regardless of the policy choices on exchange rates, as in Rey (2015), and on capital controls, which contradicts the story of a U-shaped financial integration. We contribute to the debate on the history of international financial integration by suggesting three possible mechanisms for the presence of the GFCy. First, under open capital accounts, global capital flows suggest the presence of a “communicating vessels” phenomenon. Second, in the more recent period, companies that trade in several stock markets at the same time may explain part of the commonality. Finally, in scenarios of financial repression, we expect that information flows, and a cycle of global risk aversion may explain common movements in the valuation of equity.

The rest of the paper is structured as follows. Section I presents the data and discusses the construction of trilemma regimes. Section II summarises the methodological approaches employed to construct the variables that measure the evolution of stock prices and to extract the common component that will proxy for the GFCy. Section III performs a time series analysis of the common factor in several directions. First, it argues that the new variables presented in Section II mirror the behaviour of the VIX index employed by Rey (2015). Second, it contrasts the common component extracted in Section II to the S&P 500 index to test Rey’s proposition further. Finally, it verifies whether the relationship is strictly contemporaneous and if it is driven by extreme tail events. Section IV presents the stock market valuation model and the panel data evidence on the behaviour of the GFCy under different trilemma configurations. Section V summarises the results and offers avenues for further research.
When discussing stock market integration and the existence of a global financial cycle, the natural impulse is to follow the literature on the issue and build a large panel of countries. Doing so will increase time series and cross-sectional variation and improve the chances for identification. Such is the case in the papers by Rey and coauthors discussed in the introduction. An alternative route is proposed by Forero-Laverde (forthcoming) where he studies the co-movement between the UK and US stock markets since 1922. In this paper, we choose to follow a middle route and focus exclusively on advanced economies: Australia, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Sweden, Switzerland, and the United Kingdom. Additionally, we include the US not as a dependent but as an independent variable to contrast the findings by Passari & Rey (2015).

The choice of restricting the database to a handful of countries is aimed at minimising the bad-pooling effect that occurs when aggregating stock markets that may be driven by different fundamental factors. For example, the determinants of financial instability and crises in rich and developing countries may be substantially different (Borio & Lowe, 2002). Additionally, Bordo, Meissner & Stucker (2010) argue that capital flow reversals and sudden stops, which bear an effect on the stock market, are likelier if levels of debt denominated in foreign currency are high and international reserves are low. Some of these issues, are specific to emerging market economies (EMEs) and are rarely present in advanced economies (AEs). While the natural consequence is to focus on one of those two groups, the natural extension of this work is to verify that the findings are robust to a change in the database.

The different stock market indices have been taken from Jordà, Schularick & Taylor (2017) Macrohistory database (JST). In this database, they offer market-wide, capitalisation-weighted nominal annual stock market indices for the whole selection of countries. The indices have been re-expressed in real terms using the CPI series from the same database. The authors present a full description of the sources in the online documentation to their database.

The choice of countries, however, raises the question of whether referring to a global financial cycle is pertinent. To tend to this caveat from the onset, Figure 1 shows the percentage of worldwide GDP represented by the countries in the database. As the graph shows, the share consistently represents over 50% of worldwide GDP.

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3 The database is available online at http://www.macrohistory.net/data/.
Figure 1: Percentage of World GDP represented by the countries in the database.

Note: Data on real GDP in 2011 US dollars has been taken from the Maddison Project rebasing presented in Bolt et al., (2018).

Additionally, Kuvshinov & Zimmerman (2018), using the full JST database, have calculated stock market capitalisation in advanced economies (Figure 2). They have shown that, while the ratio to GDP exploded in the 1990s, it has remained consistently around 25% since the 1920s. Given that the database is composed by the countries with the most advanced financial systems, we believe it is safe to assume that they represent the lion’s share of stock market capitalisations world-wide at least until recently.

Figure 2: Stock market capitalisation in advanced economies

Note: Taken from Kuvshinov & Zimmerman (2018, p. 6). The note to the figure reads: “Stock market capitalisation to GDP ratio, 17 countries. The solid line and the shaded area are, respectively, the median and interquartile range of the individual country capitalisation ratios in each year”.

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4 We thank Dmitri Kuvshinov for authorizing the use of this figure in this and earlier versions of this paper.
Trilemma Regimes
The idea of the macroeconomic trilemma, formalised by Obstfeld & Taylor (1997), states that fixed exchange rates, open capital accounts and autonomous monetary policies are incompatible, and thus policymakers are forced to choose two out of the three desirable goals (Obstfeld, Shambaugh & Taylor, 2005). Bordo & James (2015) show that rarely a country will choose one of the corners (the pure trilemma solutions), and instead will try to locate somewhere between the corner positions. Moreover, Aizenman, Chinn & Ito (2010, 2013) find that this framework is a tie that binds both developed and developing economies since the end of the Second World War. Obstfeld, Shambaugh & Taylor (2004) find similar results for the interwar period.

The usual portrayal in the literature, defines six broad trilemma regimes: classic gold standard (1870-1914), gold exchange standard during the interwar years, pre-convertible Bretton Woods (1944-58), convertible Bretton Woods (1959-71), inflationary post-Bretton Woods (1972-82), and the Great Moderation (1983-2015). However, this periodisation is too blunt to reflect the nuances of each particular country. Consequently, we follow the approach by Klein & Shambaugh (2015) who define four different states of the world, our trilemma regimes, as the interaction between capital control and exchange rate regime dummies: closed pegs, open pegs, closed floats, and open floats.

On the one hand, we construct an annual dummy series of capital controls using several sources that present data for the countries in the database. On the other hand, to define the exchange rate regime, we follow the methodology in Shambaugh (2004). The author studies the exchange rate between a given currency and a base currency, defined as that to which the country is pegged or more likely to peg its exchange rate. He allows for hard pegs (a ±2% band) and soft pegs (a ±5% band). All remaining periods are treated as floating exchange rates. To transform this series into a binary sequence, we can either treat all pegs as fixed exchange rates, a lax definition of the peg, or apply a strict definition where only hard pegs register as fixed exchange rates. We compare both the lax and strict exchange rate classifications with other available in the literature and find that the strict classification coincides with the updated version of Shambaugh (2004) in over 90% of the country-year observations.

Consequently, to keep comparability with other studies, we will define the four different trilemma regimes like the ones resulting from the interaction of capital control and strict exchange rate dummies. Figure 3 presents the time series evolution of the four trilemma regimes by country. Both time series and cross-sectional variability are evident in the figure and suggest that the choice of countries, while restricted in number seems to be enough.

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5 In this periodization I follow the works of Bordo & Schwartz (1999), Obstfeld & Taylor (2004), and Fatas et al. (2009).
**Figure 3: Trilemma regimes by country**

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In their work on the global financial cycle, Passari & Rey (2015) and Miranda-Agrippino & Rey (2015) use the single-period growth rates (returns) of the prices of risk assets. While this measure is standard in the literature, we take issue with the fact that returns do not account for risk and restrict the time horizon the researcher can observe. To tend to these issues, we follow the Local Bull Bear Indicator (LBBI) methodology presented in Forero-Laverde (Forthcoming). The goal of this methodology is to present three different indicators, short, medium and long-run, that characterise the behaviour of stock market indices as a risk-adjusted measure of above or below trend growth rates measured in standard deviations. Different time horizons allow for the distinction between persistent and non-persistent shocks to the series. A schematic description of the construction of LBBIs is presented below.

Let \( R \) be a matrix of dimensions \( t \times n \) where each position \( r_{t,n} \) corresponds to \( (P_t/P_{t-n}) - 1 \), and \( P_t \) corresponds to the value of the index at time \( t \). For annual returns, \( n \) takes consecutive integer values from one to five years. The short-run indicator (LBBIS) covers stock market index annual growth rates (returns). The medium-run indicator (LBBIM) covers returns from 2 to 3 years. The long-run indicator (LBBIL) covers returns from 4 to 5 years. We then perform a rolling standardization of each vector \( r_n \) using

\[
d_{t,n} = \frac{(r_{t,n} - \mu_{t,n})}{\sigma_{t,n}}
\]

where \( \mu_{t,n} \) is obtained as an exponentially weighted moving average for the last 5 observations and \( \sigma_{t,n} \) is the contemporaneous standard deviation obtained from fitting a GARCH (1,1) model. In this case, each observation \( d_{t,n} \) is measured in standard deviations. This rolling standardization serves the purpose of re-expressing returns considering the volatility context at each point in time. After all, a 10% annual return may seem like a strong boom when annual volatility is 1% but may seem as a quiet year when volatility is 20%. Additionally, (1) can be interpreted as the risk-adjusted above or below trend return. This is an added benefit of the methodology as it allows to integrate, in a single measure, characteristics of profitability and risk.

The different LBBIs are obtained from

\[
\begin{align*}
\text{LBBIS} &= D_1 \\
\text{LBBIM} &= \frac{1}{2} (D_2 + D_3) \\
\text{LBBIL} &= \frac{1}{2} (D_4 + D_5)
\end{align*}
\]
The following figures present the resulting indicators by country in the database. We exclude the United States as we will present it in section IV. Short-run LBBIs are presented in Figure 4, medium-run LBBIs in Figure 5 and long-run LBBIS in Figure 6.

**Figure 4: Short-run stock market Local Bull Bear Indicator by country**

![Short-run LBBIs](image)

**Figure 5: Medium-run stock market Local Bull Bear Indicator by country**

![Medium-run LBBIs](image)
After obtaining LBBIs, we follow the literature on dimensionality reduction to extract the first principal component from each matrix of twelve LBBIs. Following Tsay (2002), given a collection of random variables $y_n$, with variance-covariance matrix $\Sigma_y$, PCA will find a small set of vectors $c_m$, defined as linear combinations of $y_n$ that explain the underlying structure of $\Sigma_y$. Dimensionality reduction occurs since $m$ is smaller than $n$. This process requires any two components $c_j$ and $c_k$ to be orthogonal. Additionally, if the variance-covariance matrix $\Sigma_y$ is positive definite, then it has a spectral decomposition, and principal components $c_m$ of $\Sigma_y$ are its eigenvectors.\textsuperscript{6} The associated eigenvalues are directly related to the proportion of variance within the variance-covariance matrix that component $c_m$ explains.\textsuperscript{7} The intuition and benefits of employing PCA to reduce the dimension of a dataset are discussed in detail in Henning et. al. (2011). Figure 7 presents the first principal components for the short, medium and long run. The percentage value in the south-west corner of each panel represents the explanatory power of each vector over the corresponding variance covariance matrix of country LBBIs. From now on we will refer to this variables as the global financial cycle variables to the short, medium and long-run.

\textsuperscript{6} For matrix $\Sigma_y$ to be positive definite, it has to fulfill that the scalar $v = \theta^T \Sigma_y \theta$ is always positive for every non-zero column vector $\theta$. From the standard financial literature, we know that $v$ can be thought of as the variance of a portfolio with weights $\theta$ which is positive by definition.

\textsuperscript{7} To obtain the proportion of the variance of $\Sigma_y$ explained by component $c_1$, one needs to calculate the ratio between the associated eigenvalue $e_1$ and the sum of all eigenvalues $E$. 

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Figure 6: Long-run stock market Local Bull Bear Indicator by country

Figure 7: First principal components for short, medium and long run.
III

Passari & Rey (2015), in a study covering 1990-2012, indicate that ‘risky asset prices (equities, corporate bonds) around the world are largely driven by one global factor. This global factor is tightly negatively related to the [Chicago Board Options Exchange Volatility Index] VIX’ (p. 681). This result is not unique to this paper as Miranda-Agrippino & Rey (2015), Adrian & Shin (2014), and Rey (2015, 2016) also find an inverse relationship between measures of market fear and the evolution of asset prices. For example, Passari & Rey (2015), argue that the link can be established using the VIX which is extracted from the S&P 500 index, the VSTOXX which is the European equivalent, the VFTSE extracted from the London Stock Exchange FTSE index, and the VNKY which reflects market fear in the Japanese stock market. In choosing to restrict the analysis to the VIX to the detriment of all other market-fear indices, Rey (2016) argues that:

“Given the prevalence of dollar funding and of dollar assets in world balance sheets and the reliance on some type of collateral constraints or value-at-risk constraints in many parts of the financial system, assessing the effect of U.S. monetary policy on the dynamics of this global component is of interest to test for the existence of an international credit or risk-taking channel” (p.14).

As we Shown by Eichengreen & Sussman (2000), Eichengreen (2008), and Neal (2015), the role of the US as a creditor to the rest of the world was well underway after the First World War and, willingly or unwillingly, they have remained as such ever since. Consequently, it is a
possibility that, if there exists a global financial cycle, it is well represented by the evolution of the VIX index.

A final argument for using US stock market information to measure the global cycle is that, in terms of market capitalization, the US has held over 40% of global equity value since the 1930s, with only a short-lived drop at the end of the 1980s and early 1990s, as shown in Figure 8 (Kuvshinov & Zimmermann, 2018, p. 12). The mechanism at work for the transmission of financial conditions across markets runs through international investors participating in the US market and through the international flow of information as we will discuss ahead. For example, French & Poterba (1991) show that, even though there is a significant home bias in their construction, there is evidence of international diversification in stock market portfolios.

**Figure 8: World market capitalisation shares (Figure 6)**

![Figure 8: World market capitalisation shares](image)

**Note:** “Shares of individual countries’ capitalisation in world total. Capitalisation shares are computed by transforming domestic stock market capitalisation into US dollars using historical exchange rates and dividing it by the sum of capitalisations of all 17 countries. Shares of the United States, the United Kingdom, France, Germany and Japan are shown separately. All other countries are combined together into one joint item” Kuvshinov & Zimmermann (2018, p.12).

In the ideal experiment to test for the existence of the GFCy since 1922, we would need to obtain data for the VIX index since the interwar years. However, the VIX series is only available from the Chicago Board Options Exchange (CBOE) since 1990 because it originates from the implied volatility of plain vanilla call and put options on the S&P500 index to different time horizons. These financial products have only been traded for a few decades, which makes constructing the series for earlier periods impossible (CBOE, 2014).

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8 We thank Dmitri Kuvshinov for authorizing the use of this figure.
As we show in the statistical annex to this paper, the VIX index is a function of, among other things, the expected dividend growth rate and the volatility of the stock market index, where it covaries negatively with the former and positively with the latter. The relationship between the expected growth rate in dividends, volatility and LBBIs, mirrors the one with the VIX. On the one hand, LBBIs increase with the expected dividend growth rate, as it affects the numerator in (1) and increases the current level of the index. On the other hand, LBBIs decrease with increases in volatility as they affect the numerator in the construction of vectors $d_n$. Consequently, we expect the correlation between the VIX index and LBBIs to be negative and the correlation between the global cycle and LBBIs to be positive.

The natural next step is to study whether the first principal component bears any relation to the LBBIs for the US stock market. The stock market index for the US was also obtained from the JST database. The authors indicate that the series corresponds to the S&P 500 index. FIGURE plots the LBBIs for the US stock market and the GFCy variable obtained in Section II.

Figure 9: GFCy and US LBBIs to different time horizons

While simple inspection reveals striking similarities in the three pairs of series, further statistical testing is relevant. We will run regressions of the US LBBIs (independent) onto the
GFCy (dependent) variable to each time horizon. Since the GFCy series has a strong autoregressive component, we will also include its first lag in the correlates and run regressions with Newey-West (1987) corrected standard errors. Table 1 presents the results.

### Table 1: Regressions of US LBBIs onto GFCy variable

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<th>Medium-run GFCy</th>
<th>Long-run GFCy</th>
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<tr>
<td>GFCy (first lag)</td>
<td>-0.0069</td>
<td>0.2924***</td>
<td>0.3627***</td>
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<td>US stock market LBBI</td>
<td>0.0986***</td>
<td>0.1055***</td>
<td>0.1032***</td>
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<tr>
<td>Constant</td>
<td>0.0015</td>
<td>-0.0006</td>
<td>-0.0002</td>
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<td>Number of lags (Newey West)</td>
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<td>N</td>
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<tr>
<td>R-squared (OLS)</td>
<td>0.4583</td>
<td>0.5483</td>
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**Note:** All regressions run using Newey West standard errors. Choice of lags is based on the behaviour of the error term as white noise. Significance levels * 10%, ** 5%, *** 1%.

There are two main takeaways from Table 1. First, the value of the R-squared measure is high and indicates that the explanatory power of the US LBBIs over the global financial cycle variables is notable. Second, the regression coefficient for the US LBBIs is positive and statistically significant beyond 99% to every time horizon. This is an indication that the linear relationship between both variables is strong. Three questions remain to be answered. Is the relationship strongest when measured contemporaneously? Is the relation stable across the whole distribution or do tail events drive it? Can we argue for causality in the relationship? We will address the first two questions in what follows while the question about causality is left for further research.

### Leads & Lags analysis

To confirm whether the linear relationship between US LBBIs and the GFCy is strongest when measured contemporaneously, we ran five regressions as the one presented in Table 1 where we changed the US LBBI variables to include one and two lags and one and two leads with respect to the GFCy variable. Results for the coefficient of the variable of interest is presented in Figure 10.

#### Figure 10: Leads and lags analysis of the US LBBI coefficient

**Note:** Note: X-axis indicates the number of leads (positive) or lags (negative) of US LBBI in a regression following the one presented in Table 1. The continuous line represents the estimator, dotted lines represent the 95% confidence interval, and the broken line represents the BIC criterion for each specification.
The main takeaway from Figure 10 is that both significance and goodness of fit in the model peak contemporaneously. We take this as a first piece of evidence in favour of the global cycle hypothesis presented by Rey (2016) and quoted above.

**Quantile Regressions**

Results from the lead and lag analysis do not imply that the relationship between GFCy and US LBBIs is stable across the whole joint empirical distribution. An alternative is that the link may be driven by extreme events which cause an increase or decrease synchronisation between both variables. Consequently, we will run quantile regressions by decile of the same form as the ones presented in Table 1 and check whether there are statistically significant differences in the coefficients by quantile and the full sample contemporaneous coefficient. Results are presented in Figure 11.

**Figure 11: Coefficient for the US LBBIs under quantile regressions by decile.**

Note: Grey continuous line shows the full sample coefficient for US LBBIs as obtained from Table 1. Black continuous line shows the coefficient for the US LBBIs by decile. Dotted lines show the 95% confidence interval around the coefficient by decile.

As results in Figure 11 show, the linear relationship between US LBBIs and GFCy remains stable, regardless of the portion of the joint empirical distribution that is being observed. This serves as further indication that, as suggested by the authors of the global financial cycle hypothesis, the US stock market, as proxied by the VIX seems to be a sufficiently robust representative of the cycle. This finding is innovative as it allows extending the stylized fact back until the interwar years.

**IV**

To understand the possible sources of variation in stock prices and their links to the macroeconomic trilemma, we resort to Gordon’s dividend discount model as in Stuart (2017), which we present in the following equation:

\[
P_0 = \sum_{t=0}^{T} \frac{D_t}{(1+r_t)^t}
\]  

(3)

According to the model, the price of a stock at time 0 \(P_0\) is the present value of the expected payout \(D_t\) stockholders will receive in the future. The discount factor is associated with an interest rate \(r_t\) which accounts for the time-value of money and the riskiness of the investment.
It can be thought of as the hurdle rate that makes the net present value of a correctly priced investment 0 in equilibrium (Damodaran, 2011).

To control for the different policy choices regarding the trilemma, we will include variables of interest related to capital flows and exchange rates. First, we include the overall current balance to GDP as increases in this variable indicate a more robust economic activity and more income for export-oriented firms.\(^9\) To account for international capital flows, we include the evolution of the net capital account as a proportion of GDP.\(^10\) To account for the effect of exchange rates we resort to two distinct equilibrium conditions: the purchasing power parity (PPP) and the covered interest rate parity (CIRP). PPP establishes that in the absence of transaction costs and trade barriers, two identical goods in two distinct markets should cost the same amount of money when expressed in the same currency. In that sense, the critical determinant of the exchange rate in this model is the differential of inflation rates between the domestic and foreign economies.

On the other hand, CIRP establishes that under free capital flows and no transaction costs (0 bid-offer spread), the nominal exchange rate, the domestic and foreign interest rates are jointly set to eliminate arbitrage opportunities. The nominal exchange rate is such that no risk-free profit can be obtained by borrowing abroad (domestically), selling (buying) the foreign currency domestically, investing at the local (foreign) interest rate and using the proceeds to pay the loan in foreign (domestic) currency. Under this model, the critical determinant of the nominal exchange rate is the differential of interest rates. We summarise these relationships in the following equation:\(^{11}\)

\[
FX_{Nominal} = f[(r_d - r_f), (\pi_d - \pi_f)]
\]  

(4)

Where \(r_d\) is the domestic interest rate, \(r_f\) is the foreign interest rate, \(\pi_d\) is the domestic inflation rate, \(\pi_f\) is the foreign inflation rate.

A first issue has to do with the underlying assumptions in (4). Both PPP and CIRP require the law of one price to function correctly, which optimally occurs when there are no transaction costs and in the presence of free capital flows. This does not mean, however, that permanent deviations from these two equilibria are sustainable in time if capital controls are established. If the exchange rate is fixed, and the capital account is closed, as was the case during Bretton Woods, a large inflation differential would make imports cheap and exporters loose competitiveness.

---

\(^9\) Data for the net current account to GDP comes from Mitchel’s (2013) International Historic Statistics 1750-2010.

\(^10\) Data for the net capital account to GDP comes from Mitchel’s (2013) International Historic Statistics 1750-2010. We obtain the net capital account from the identity: net capital account + net current account + net changes in reserves = 0. We thank Barry Eichengreen for this suggestion.

\(^11\) Data for the interest rate and inflation differentials is obtained from Jordà, Schularick & Taylor Macrohistory database (2017). We always define the foreign rate to be the one from the United States which is the largest developed economy not included in our dependent variables.
When this happened, persistent deficits in the current would motivate a one-time change in the value of the peg, as was the case for France, or the UK.

By the same token, a sustained increase in the interest rate differential, could hinder domestic economic activity, reduce the number of viable investment projects for domestic companies, restrict access to credit and aggregate demand, and drive down the competitiveness of the export sector. Persistent long-run deviations in the differentials would bring forth a permanent payment imbalance via the current account which, under the Articles of the Agreement, would have been grounds for a change in the level of the peg (one-time devaluation). In that sense, we expect the determinants of the nominal exchange rate to affect the stock price under all trilemma regimes.

Regarding additional control variables, we follow Jorion (1990) who indicates that companies with sizeable international presence will benefit from openness to the international trade system while domestic companies may suffer from competition from abroad. We include changes in openness to trade, defined as the first difference of the sum of imports and exports to GDP, to control for this situation.

Furthermore, the level of financial development will be critical in determining the liquidity in the stock market as well as its overall size. To measure this, we include the changes in M2, a measure of broad money, to GDP. This is a measure of the overall size of the financial system which, according to the finance and growth literature bears a strong relationship with the general economy and consequently should have a strong positive correlation with the stock market (Levine, 2005). This variable will affect both the numerator and denominator in (3) as it relates both to economic activity and to the discount rate employed in valuation.

Finally, economic growth is critical as it alters expectations about future cash flows and alters the investment opportunity set for companies. Consequently, we include the percentage change in real GDP per capita among our covariates. We can summarise the variables of interest and their relationship to the model in the following two equations:

\[
D_t = f\left[\frac{NKA}{GDP}, (r_d - r_f), (\pi_d - \pi_f), \frac{OCB}{GDP}, (M + X), \frac{M2}{GDP}, \left(\frac{RGDP_{pc}}{RGDP_{pc_t-1}} - 1\right)\right]
\]

\[
r_t = f\left[\frac{NKA}{GDP}, (r_d - r_f), \frac{M2}{GDP}\right]
\]

Where NKA is the net capital account, OCB is the overall current balance, M is the total value of imports, X is the total value of exports, M2 is a measure of broad money, and RGDP_{pc} represents real GDP per capita.
Fixed effects VS pooled regressions

We run a first panel regression of LBBIs to each time horizon as the dependent variable, a linear time trend, and the different correlates in (5) and (6) as explanatory variables. We perform the F test for the significance of country fixed effects and time fixed effects where the null is that the effects are not significant. We also run the Breusch & Pagan (1979) test for the significance of random effects. Results are presented in Table 2.

Table 2: Panel specification tests for the use of random or fixed effects

<table>
<thead>
<tr>
<th>Horizon</th>
<th>Short-run</th>
<th>Medium-run</th>
<th>Long-run</th>
</tr>
</thead>
<tbody>
<tr>
<td>Statistic (F)</td>
<td>0.22</td>
<td>1.00</td>
<td>1.21</td>
</tr>
<tr>
<td>Prob &gt; F</td>
<td>1.00</td>
<td>0.44</td>
<td>0.28</td>
</tr>
<tr>
<td>Conclusion</td>
<td>Pool</td>
<td>Pool</td>
<td>Pool</td>
</tr>
</tbody>
</table>

Note: Panel A tests for the joint significance of country fixed effects, under the null that they are jointly insignificant. Panel B tests for the joint significance of time fixed effects, under the null that they are jointly insignificant. Panel C performs the Breusch & Pagan (1979) test for a random effects model under the null that random effects are unnecessary.

According to Table 2, there is insufficient heterogeneity across countries to warrant the use of country fixed effects (Panel A). Similarly, in Panel C the results for the Breusch & Pagan (1979) tests suggest that the use of a random effects model is unnecessary and that pooling the data is a sufficient alternative. However, Panel B shows that there seems to be a role for time fixed effects, which indicates that there are shocks, over time, that are common to all countries in the database. This is precisely the idea behind the “global financial cycle hypothesis” argued by Rey (2015) and Passari & Rey (2015) among many others.

Consequently, we run the pooled OLS regression of LBBIs on the dependent variables in (5) and (6) under two alternative specifications. In the first one, we include time fixed effects, while in the second one we only include the global financial cycle variable to the corresponding time horizon. The underlying idea is that both these specifications offer the inclusion of common shocks to all stock markets while the second specification has the added benefit of offering an underlying economic and historical interpretation. In their respective specifications, time fixed effects are jointly significant, and the coefficient for the GFCy is statistically significant with 99% confidence. For the sake of brevity, in Table 3 we only present the BIC measure for goodness-of-fit in each regression.
Table 3: Goodness-of-fit for time fixed-effects model VS GFCy model

<table>
<thead>
<tr>
<th></th>
<th>Short-run</th>
<th>Medium-run</th>
<th>Long-run</th>
</tr>
</thead>
<tbody>
<tr>
<td>BIC</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Time FE</td>
<td>1908</td>
<td>1712</td>
<td>1653</td>
</tr>
<tr>
<td>PC</td>
<td>1353</td>
<td>1173</td>
<td>1116</td>
</tr>
</tbody>
</table>

Note: Goodness-of-fit measured by the Bayesian Information Criterion for the two regressions discussed in the text. Lower values of the statistic indicate a better fit of the data.

Panels specification tests

As we have discussed, the optimal regression model, according to the preliminary tests in Table 3, requires running OLS regressions without country fixed effects and merits the inclusion of time fixed effects that account for common contemporaneous shocks across all countries. As we argued extensively, the role of a common shock to all countries will be proxied by the inclusion GFCy variable to the different time horizons among the independent variables. However, from an economic perspective, even if the twelve countries in the database can be characterised as advanced economies, there are country-specific criteria that remain unobserved and that we may wish to control for. For example, the difference between common or civil law judicial systems, the existence of parliamentary or presidential executive branches, as well as variations in religious preferences, language, government institutions, and several other characteristics that may render, for example, Canada and Japan to be incomparable.

Including country fixed effects, however, would lead to an over-specified model according to the contrasts we have performed. We know that the inclusion of these variables has no effects on the unbiasedness of the estimators for the other variables although it may increase the standard errors. In that sense, including fixed effects acts as a litmus test against our testable hypothesis and suggests that statistically significant coefficients would be even more significant if the “true model” were used (Wooldridge, 2002).

Since we run a panel regression with country fixed effects, we need to perform tests to confirm the OLS estimations are appropriate. The three primary assumptions underlying the ordinary least squares estimation for the panel fixed effects model is that there is group-wise homoskedasticity of the errors, that each error series is not autocorrelated, and that there is no contemporaneous correlation in the error terms for the different countries. We perform regressions where the dependent variables are the LBBIs for the twelve stock markets to the different time horizons, and the regressors are the dependent variables in (5) and (6), and the LBBI for the US stock market as a proxy for the global cycle. In information that is available upon request, we perform panel unit root tests for all the dependent and independent variables and find they are all panel-stationary. We then perform an additional battery of tests for the panel’s specification. First, the modified Wald test for group-wise heteroskedasticity (Greene, 2017). Secondly, Wooldridge’s (2002) test for first-order autocorrelation in panel data. Finally, regarding
the contemporaneous cross-correlation of error terms, we include Pesaran’s (2004) and Frees’ (1995, 2004) tests. We present the results in Table 4.

Table 4: Panel specification tests for structure in the residuals

<table>
<thead>
<tr>
<th>Horizon</th>
<th>Short-run</th>
<th>Medium-run</th>
<th>Long-run</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Panel A</strong>: Modified Wald statistic for group-wise heteroskedasticity in fixed effect model (Greene, 2017)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Statistic (chi 12 DoF)</td>
<td>41.48</td>
<td>32.67</td>
<td>25.02</td>
</tr>
<tr>
<td>Prob &gt; CHI2</td>
<td>0.00</td>
<td>0.00</td>
<td>0.01</td>
</tr>
<tr>
<td>Conclusion</td>
<td>Heteroskedastic</td>
<td>Heteroskedastic</td>
<td>Heteroskedastic</td>
</tr>
<tr>
<td><strong>Panel B</strong>: Wooldridge (2002) test for autocorrelation in panel data</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Statistic (F 1,11)</td>
<td>10.59</td>
<td>182.19</td>
<td>175.64</td>
</tr>
<tr>
<td>Prob &gt; F</td>
<td>0.01</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>Conclusion</td>
<td>Autocorrelation</td>
<td>Autocorrelation</td>
<td>Autocorrelation</td>
</tr>
<tr>
<td><strong>Panel C1</strong>: Test for cross-correlation of errors Pesaran (2004)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Statistic N(0, sigma)</td>
<td>-5.38</td>
<td>-4.88</td>
<td>-4.87</td>
</tr>
<tr>
<td>Prob &gt; N</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Conclusion</td>
<td>No correlation</td>
<td>No correlation</td>
<td>No correlation</td>
</tr>
<tr>
<td>Statistic Chi2</td>
<td>12.33</td>
<td>14.14</td>
<td>8.53</td>
</tr>
<tr>
<td>Prob &gt; N</td>
<td>0.34</td>
<td>0.23</td>
<td>0.67</td>
</tr>
<tr>
<td>Conclusion</td>
<td>No correlation</td>
<td>No correlation</td>
<td>No correlation</td>
</tr>
</tbody>
</table>

*Note:* Panel A shows the modified Wald test statistic for group-wise heteroskedasticity in the error under the null of homoskedasticity (Greene, 2017). Panel B shows Wooldridge’s (2002) test for autocorrelations in the errors in a panel data framework under the null that there is no first-order correlation. Panel C1 contains the test for cross-correlation of errors as in Pesaran (2004) under the null of no cross-correlation between groups. Panel C2 contains Frees’ Q distribution test under the null of no cross-correlation between groups.

Panel A of Table 4 shows that there is evidence of heteroskedasticity for the short and long-run regressions. Panel B in the table indicates that there is evidence of first-order autocorrelation in all specifications. Finally, Panels C1 and C2 indicate that we cannot reject the null hypothesis of no contemporaneous cross-correlation of the error terms for the different countries.

Consequently, in the panel regressions that follow we will present panel corrected standard errors (PCSE) estimates obtained through Prais-Winsten (1954) regressions. This estimation methodology, when in a panel data framework, corrects for the two situations discussed above: heteroskedasticity and autocorrelation in the individual error terms. While this regression methodology belongs to the family of feasible generalised least squares (FGLS) estimations, it is preferred over standard GLS estimation procedures as it allows for unbalanced panels, as is our case.12

---

12 Our panel is unbalanced due to some gaps in the independent variables, particularly during the Second World War or during the interwar years. Gaps are not common to all countries.
The unconditional model

In Table 5 we present the unconditional regressions of LBBIs the twelve countries’ stock markets on the independent variables. 13

### Table 5: Unconditional regressions for the stock market model

<table>
<thead>
<tr>
<th></th>
<th>Short-run</th>
<th>Medium-run</th>
<th>Long-run</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.0284</td>
<td>0.073</td>
<td>0.0807</td>
</tr>
<tr>
<td>Trend</td>
<td>-0.00083</td>
<td>-0.0014</td>
<td>-0.0013</td>
</tr>
<tr>
<td>Change in real GDP per capita</td>
<td>0.1901</td>
<td>1.127**</td>
<td>0.5116</td>
</tr>
<tr>
<td>Change in trade openness</td>
<td>0.5198</td>
<td>0.3639</td>
<td>0.1752</td>
</tr>
<tr>
<td>Change in financial development</td>
<td>.8647*</td>
<td>.7013*</td>
<td>0.5379</td>
</tr>
<tr>
<td>Overall current balance to GDP</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Capital account to GDP</td>
<td>1.216*</td>
<td>1.987***</td>
<td>2.161***</td>
</tr>
<tr>
<td>Short term rate differential</td>
<td>0.5795</td>
<td>0.5335</td>
<td>0.401</td>
</tr>
<tr>
<td>(domestic-foreign)</td>
<td>-0.0052</td>
<td>-.0208***</td>
<td>-0.0091</td>
</tr>
<tr>
<td>Inflation differential</td>
<td>-.457***</td>
<td>-.2443*</td>
<td>-.3457***</td>
</tr>
<tr>
<td>(domestic-foreign)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Global cycle (First Principal Component)</td>
<td>4.786***</td>
<td>4.163***</td>
<td>4.33***</td>
</tr>
<tr>
<td>N</td>
<td>938</td>
<td>938</td>
<td>938</td>
</tr>
<tr>
<td>R squared</td>
<td>0.5507</td>
<td>0.5496</td>
<td>0.5196</td>
</tr>
</tbody>
</table>

**Note:** Dependent variables are short-run, medium-run and long-run stock market LBBIs. The regressions were run using Panel Corrected Standard Errors (PCSE). Statistically significant coefficients in bold. Variables in the top panel refer to domestic controls. Variables in the bottom panel refer to controls associated with the macroeconomic trilemma and the GFCy. Significance levels * 10%, ** 5%, *** 1%.

A first element that is worth highlighting from Table 5 is the little relevance that domestic variables, in the top half of the panel, have in explaining the evolution of the stock market. This may be because all common movement across economic conditions is already accounted for within the principal component that proxies for the GFCy. Since we are dealing with advanced economies, commonalities among them can be high throughout the period.

Regarding the results for the variables in the bottom half of the table, we find that the coefficients are statistically significant to every time horizon for the current account, the inflation differential, and the GFCy component. In all cases, the sign is consistent with what has been discussed throughout the text.

---

13 To tend to the issue of collinearity, following Greene (2017) we calculated the condition numbers for the three sets of variables used in the regressions. A condition number is the square root of the ratio between the largest and smaller eigenvalues of the correlation matrix of variables. We obtained values below 5. As a rule of thumb values above 20 should be worrisome as they indicate the presence of a single factor with large explanatory power over all explanatory variables which may be driving collinear relationships between them.
A final noteworthy finding in Table 5 is the high value of the R2 statistic. Each of the unconditional regressions is able to explain over 50% of the variance in the stock market of the twelve countries, which is substantially better than what is usually achieved by standard models such as the CAPM model (Markowitz, 1952; Lintner, 1965; Sharpe, 1966) or the Fama & French (1993) three-factor model and its extensions to 5 factors (Fama & French, 2008). The goodness-of-fit we achieve is also higher than the one obtained by Passari & Rey (2015) which is close to 35%.

A possible criticism of this result arises from the “bottles of ketchup” analogy by Summers (1985) in that we may be just observing no-arbitrage conditions that allow for the synchronisation of markets. However, we have built a stock pricing model from Gordon’s dividend discount model that includes the interest rate differential in the denominator and several fundamental variables that can affect the dividend stream among our correlates. The aim of including so many variables is to allow us to shift from a no-arbitrage scenario to a more fundamental valuation. Additionally, it is worthwhile recalling that we are not asking whether the changes in stock prices are fundamental or bubbly. In the following sub-section, we verify if these relationships are stable across trilemma regimes, expansions and contractions.

The conditional model: Structural breaks by trilemma regime

To test whether the findings in Table 5 are contingent on the trilemma regime in place we perform a Chow (1960) test for structural breaks in all coefficients by trilemma regime. Additionally, we perform a standard F-test test for the presence of a break only in the GFCy coefficients by trilemma regime. Table 6 contains the results. In all cases, the null hypothesis is the equality of coefficients by group.

Table 6: Statistical test for structural breaks in coefficients

| Hypothesis I: There is no break by trilemma regime in all coefficients |
|---------------------------------|---------|---------|---------|
| Horizon                        | Short   | Medium  | Long    |
| Statistic                      | 1.53    | 0.90    | 1.13    |
| P-value                        | 0.01    | 0.70    | 0.24    |

| Hypothesis II: There is no break by trilemma regime in global cycle coefficients |
|---------------------------------|---------|---------|---------|
| Horizon                        | Short   | Medium  | Long    |
| Statistic                      | 0.98    | 1.06    | 0.32    |
| P-value                        | 0.40    | 0.36    | 0.81    |

Note: The top panel presents the results of a Chow test for structural breaks in the coefficients of the regressions in Table 5 by trilemma regime. The bottom panel presents the results of a standard F-test to check whether there is a break in the GFCy coefficients by trilemma regime. The null hypothesis of both tests is that there is no break.

Results in Table 6 show that the only evidence of a joint break in all coefficients by trilemma regime occurs in the short-run specification. In the medium and long-run specifications coefficients seem to be stable across regimes. This is puzzling and seems to suggest a new avenue
for further research. In Table 7 we present the regression results for the full sample and by trilemma regime.

Table 7: Short-run regression results for the full sample and by trilemma regime

<table>
<thead>
<tr>
<th>Short run LBBI</th>
<th>Full sample</th>
<th>Closed peg</th>
<th>Open peg</th>
<th>Closed float</th>
<th>Open float</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.0284</td>
<td>0.2343</td>
<td>-0.1978***</td>
<td>-0.0245</td>
<td>0.2185**</td>
</tr>
<tr>
<td>Trend</td>
<td>-0.0008</td>
<td>-0.0073***</td>
<td>-0.0020</td>
<td>0.0006</td>
<td>-0.003***</td>
</tr>
<tr>
<td>Change in real GDP per capita</td>
<td>0.1901</td>
<td>1.729**</td>
<td>-1.0850</td>
<td>-0.4504</td>
<td>-2.422**</td>
</tr>
<tr>
<td>Change in trade openness</td>
<td>0.5198</td>
<td>0.0057</td>
<td>1.073*</td>
<td>-0.0820</td>
<td>1.391*</td>
</tr>
<tr>
<td>Change in financial development</td>
<td>.8647*</td>
<td>1.404*</td>
<td>0.4525</td>
<td>-0.5644</td>
<td>1.299*</td>
</tr>
<tr>
<td>Overall current balance to GDP</td>
<td>1.216*</td>
<td>-0.0584</td>
<td>2.7020</td>
<td>3.9130</td>
<td>2.178***</td>
</tr>
<tr>
<td>Capital account to GDP</td>
<td>0.5795</td>
<td>0.2439</td>
<td>1.3210</td>
<td>3.2360</td>
<td>0.6638</td>
</tr>
<tr>
<td>Short term rate differential (domestic-foreign)</td>
<td>-0.0052</td>
<td>-0.0120</td>
<td>-0.0002</td>
<td>-0.0331*</td>
<td>0.0145</td>
</tr>
<tr>
<td>Inflation differential (domestic-foreign)</td>
<td>-.457***</td>
<td>-.3339*</td>
<td>-1.4410</td>
<td>-0.1472</td>
<td>-1.1580</td>
</tr>
<tr>
<td>Global cycle (First Principal Component)</td>
<td>4.786***</td>
<td>4.421***</td>
<td>5.226***</td>
<td>4.702***</td>
<td>4.794***</td>
</tr>
<tr>
<td>N</td>
<td>938</td>
<td>260</td>
<td>206</td>
<td>160</td>
<td>312</td>
</tr>
<tr>
<td>R squared</td>
<td>0.5507</td>
<td>0.4458</td>
<td>0.7306</td>
<td>0.4309</td>
<td>0.7072</td>
</tr>
</tbody>
</table>

Note: Dependent variables are the short-run stock market LBBI. The regressions were run using Panel Corrected Standard Errors (PCSE). Statistically significant coefficients in bold. Variables in the top panel refer to domestic controls. Variables in the bottom panel refer to controls associated with the macroeconomic trilemma and the GFCy. Significance levels * 10%, ** 5%, *** 1%.

The main takeaway from Table 7 is that the principal component that proxies for the GFCy is statistically significant with 99% confidence to every time horizon. Additionally, in tests that are available upon request, all coefficients are statistically equal so we can conclude that the relationship does not change contingent on the institutional setup. This means that the story of disintegrated capital markets under Bretton Woods may need some revision. In theory, strong capital market integration under open capital accounts is expected. On the one hand, it may be due to a “communicating vessels” phenomenon in which investors flood all advanced economies. On the other hand, it may be because the same company can trade its equity in two or three exchanges simultaneously, giving way to no-arbitrage arguments for co-movement. The question of what happens when the capital account is closed remains open to debate. We suggest that it is information flows and the transmission of risk perspectives and valuations across agents what explains the co-movement. It is not necessarily a story of effective stock market transactions but the transmission of investor expectations across markets.
Additionally, we find that several variables that according to the theory should be statistically significant, do not appear to determine the behaviour of stock prices: GDP growth, capital flows, short-term interest rate differentials. This, which is also true for the unconditional model presented above, suggests that the GFCy variable may be capturing not only the joint movement of stock markets but also the joint movement across economies. The counterintuitive nature of this result suggests that further research is necessary.

V

The goal of this paper was twofold. On the one hand, we wished to find evidence of a global financial cycle, as posited by Passari & Rey (2015), that would explain part of the stock market valuations since the interwar years for twelve advanced economies. On the other hand, we wanted to confirm whether this global financial cycle would confirm the usual story of financial integration/disintegration posited by Obstfeld & Taylor (2004). To do so, measure the evolution of stock market indices we used the three LBBIs proposed by Forero-Laverde (Forthcoming).

Regarding the first question, we find evidence of a common factor that explains over 50% of the variance-covariance matrix of LBBIs for the stock market indices of the countries in our database. This common cycle bears a positive and statistically significant coefficient to different time horizons, even after controlling for a variety of factors. Additionally, as suggested by the global financial cycle literature, we find that this factor is strongly correlated, contemporaneously, with the behaviour of the US stock market LBBI, and that this correlation is not driven by extreme cases but instead is stable across their joint empirical distributions. Consequently, the main contribution to the global cycle debate is to find evidence of a common factor that drives the stock market of advanced economies and providing an economic interpretation for it.

Concerning the second question, we find evidence of a joint break in all the coefficients of the valuation model by trilemma regime in the short-run. However, the coefficients for the GFCy to different time horizons and under different trilemma configurations are statistically equal. This shows that the common factor across stock markets is present regardless of the policy choices on exchange rates, as in Rey (2015), and on capital controls, which contradicts the story of a U-shaped financial integration. We contribute to the debate on the history of international financial integration by suggesting three possible mechanisms for the presence of the GFCy. First, under open capital accounts, global capital flows suggest the presence of a “communicating vessels” phenomenon. Second, in the more recent period, companies that trade in several stock markets at the same time may explain part of the commonality. Finally, in scenarios of financial repression, we expect that information flows, and a cycle of global risk aversion may explain common movements in the valuation of equity.
We offer three avenues for further research. First, our results indicate the relevance of showing whether the relationship between US LBBIs and GFCy is causal. It is possible that it is the US which is driving the relationship, as suggested by Rey (2016). However, an alternative possibility is that there is bi-directional causality which would imply that stock market movements are all jointly and contemporaneously determined. The implications this has for financial stability, stock valuation, and macroprudential regulation are significant.

A second avenue for further research has to do with changing our sample towards emerging market economies and check whether the same regularities are present. This idea is challenging on several accounts. First, stock markets in EMEs do not have necessarily a long history, nor are they as liquid and efficient as their counterparts in advanced economies. Second, data is not difficult to compare across countries but also difficult to come by. Finally, EMEs are usually bank-based systems and rather than trading equities tend to trade either public debt instruments or foreign currencies. It might be interesting to check whether there is a common cycle driving the prices of these different asset classes.\(^\text{14}\)

The final line of further inquiry has to do with the puzzling results of TABLES (chow and by trilemma). It is surprising and counterintuitive that changes in GDP, the level of free capital flows to GDP, or the interest rate differential do not affect stock valuations to any time horizon. It is possible that the part of the synchronous movement across economies is being captured by our GFCy variable. Consequently, given that our main contribution is to show the existence of a GFCy in the long run, a natural next step would be to study its determinants and how they change depending on the institutional setup.

**References**


\(^{14}\) We are thankful to Jose Antonio Ocampo for bringing this to our attention.


Annex 1. A Comparison between VIX and LBBIs: Are they related?

Following Wilmott (2006), financial derivatives are contracts signed at a time \( t_0 \), which mature at time \( t_k \). The two best-known derivatives contracts are forwards and European options. The forward contract is one in which the issuer of the contract commits him or herself to buy or sell a given asset (the underlying), at a future date (expiration), at a set price that is known at time \( t_0 \) (strike price). The beneficiary, the counterparty to the issuer, also commits him or herself to sell or buy the underlying asset under the agreed upon conditions. The strike price of a forward contract is set in such a way that there is no need to exchange cash flows on the date of issuance.

A European option contract is similarly defined, but in this case, while the issuer commits to buying or selling the underlying, the beneficiary of the contract can choose to sell or buy depending on current market conditions.\(^{15}\) In a call option, the beneficiary of the contract can buy the underlying at the strike price on the maturity date. He or she will only do so if the market price is above the strike price (the option is “in the money”). In a put option, the beneficiary of the contract can sell the underlying at the strike price on the maturity date. He or she will only do so if the market price is below the strike price. On the date of issuance, options are quoted and negotiated out of the money, meaning that for a call (put) option, the strike price must be above (below) the market price (spot price). Since the bearer of the option has no downside in his or her future cash flows, just as in an insurance contract, options have a positive price at the time of issuance (premium) which the beneficiary pays to the issuer. The premium for call and put options on a stock market index are calculated as follows (Wilmott, 2016, V1 pp 116-118):

\[
c = S e^{-D(T-t)} N(d_1) - K e^{-r(T-t)} N(d_2) \tag{7}
\]

\[
p = -S e^{-D(T-t)} N(-d_1) + K e^{-r(T-t)} N(-d_2) \tag{8}
\]

\[
d_1 = \frac{\log \left(\frac{S}{K}\right) + \left(r - D + \frac{1}{2} \sigma^2\right)(T - t)}{\sigma \sqrt{T - t}} \tag{9}
\]

\[
d_2 = \frac{\log \left(\frac{S}{K}\right) + \left(r - D - \frac{1}{2} \sigma^2\right)(T - t)}{\sigma \sqrt{T - t}} = d_1 - \sigma \sqrt{T - t} \tag{10}
\]

Where \( S \) is the spot value of the index, \( D \) is the continuous expected dividend growth rate, \( (T-t) \) represents time to maturity, \( N \) is the standard normal distribution function, \( K \) is the agreed upon strike price, \( r \) is the continuous risk-free interest rate between time \( t \) and time \( T \), and \( \sigma \) is the volatility of the underlying asset. Several important relationships from equations (7) and (8) are relevant:

\(^{15}\) A relevant characteristic of European options is that they can only be exercised at the maturity date. Conversely, American options can be exercised at any point in time between issuance and maturity.
All else equal, if the strike price increases call (put) options become more in the (out of the) money and their price \( c \) (\( p \)) increases (decreases). With regards to the pay out, higher expected dividend growth rates \( D \) imply a lower value for both put and call options. On the other hand, higher \( \sigma \) increases the value of both types of contracts. Note that the relationship between expected dividend growth, volatility, and LBBIs mirror their relations to the price of call and put options. When expected dividends increase, the stock price increases and so do LBBIs. When volatility increases the denominator in the LBBI function increases, and thus the LBBI decreases.

According to CBOE (2014, p. 4), the generalised formula for calculating the VIX index is:

\[
\begin{align*}
\frac{VIX}{100} = \sigma^2 &= \left( \frac{2}{(T-t)} \sum_{i} \frac{\Delta K_i}{2T} e^{r(T-t)} Q(K_i) \right) - \\
&\left( \frac{1}{(T-t)} \left( \frac{K + e^{r(T-t)}(c - p)}{K_0} - 1 \right) \right) 
\end{align*}
\]  

(33)

Where \( c \) and \( p \) correspond to the price of out-of-the-money call and put options centred around an at-the-money strike price \( K \). \( K_0 \) is the first traded strike price below strike price \( K \), \( K_i \) is the strike price of the \( i \)-th out-of-the-money option such that for call options \( K_i > K_0 \) and for put options \( K_i < K_0 \), and \( Q(K_i) \) is the mid-point of the bid-ask spread for each option with strike \( K_i \). According to Wilmott (2006), buying an out-of-the-money call option and selling an out-of-the-money put option is an investment structure called an at-the-money straddle. The payoff structure of that investment strategy at time \( T \) is depicted in Figure 12.

**Figure 12: Payoff profile of an at-the-money straddle**

![Payoff profile of an at-the-money straddle](image)

**Note:** In the payoff structure depicted above the current stock market price is $100. The out-of-the-money call option (green line) has a strike price of $110. The out-of-the-money put option (orange line) has a strike price of $90. To exemplify, the premium paid for each option is $5. The payoff for the portfolio of options is depicted in the blue line.

If the current stock price is $100, such that both options are out-of-the-money, higher expected volatility will increase the probability that the structure has a positive payoff. Consequently, higher volatility will increase the value of the structure and affect the VIX index.
positively. In the payoff structure shown in Figure 36 we assume there are no dividend payments. If a dividend payment occurs during the life of the contracts, that is a payoff that the beneficiary is not entitled to and thus would cause a parallel downward shift of the payoff structure. Consequently, as the expected dividend growth rate increases the payoff in the structure and the value of the VIX is reduced.