How does sovereign bond market integration relate to fundamentals and CDS spreads?

We have revised the title to better reflect the new content and contributions of the paper. However, we are submitting under the old title so as to avoid any confusion.

New Title: Measuring sovereign bond market integration

Abstract

There is significant heterogeneity in the degree and dynamics of sovereign bond market integration across 21 developed and 18 emerging countries. We show that better spanning can significantly enhance market integration through local risk premia dissipation. Integration of the sovereign bond markets increases on average by about 10%, when a country moves from the 25th percentile to the 75th percentile as a result of higher political stability and credit quality, lower inflation and inflation risk, and lower illiquidity. The 10% increase in integration leads to, on average, a decrease in the sovereign cost of funding of about 1% per annum.

Market integration is central to the study of international finance but has so far received less attention in the sovereign bond market literature. Yield spreads, bond returns and the funding cost faced by sovereigns clearly depend on the degree of market integration. Hence, for sovereign bonds, we examine the pricing of global and local risks, the level and dynamics of integration, and its key drivers. We show that better spanning affects sovereign bond prices through local risk premia dissipation and leads to higher integration and lower cost of funding.

Scholars have primarily measured sovereign bond market integration based on the law-of-one price under the no-arbitrage condition that bonds with the same risk should command the same expected return in the same currency denomination regardless of the nationality of the issuer or the location of its listing. Indeed, in addition to their exposure to the common global market risk, differences in country specific risks as well as differential explicit and implicit barriers that asymmetrically impact countries should be taken into account for estimating the degree of market integration.[1] Hence, we estimate market integration based on the theoretical international asset pricing model of Errunza and Losq (henceforth EL, 1985). It is a two factor (global and unspanned local factor) model that incorporates barriers to capital flows and is applicable to all types of financial assets. Nonetheless, since it has not been used in the context of sovereign bond pricing, we test the model and compare our results with different alternative models including full integration that do not depend on EL assumptions.

Under the model, integration is the square of the correlation between the returns on the *Ith* country diversification portfolio and the *Ith* sovereign bond. The diversification portfolio (DP) is the portfolio of substitute assets that is most highly correlated with the *Ith* sovereign bond. Substitute assets are all

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[1] Explicit barriers include legal restrictions on ownership, foreign exchange controls that are imposed by the governments of borrowing and creditor countries as well as those related to institutional constraints/mandates, for example, investors may have limited funding capacity. Implicit barriers include risks related to political uncertainty, incomplete, inaccurate or asymmetric information, quality of governance, market size, illiquidity, and market regulation.

assets that are freely available to investors regardless of their nationality.[2] Increasing integration implies increase (decrease) in the importance of the global (local) factors, and declining expected return on the sovereign bond. An increasing degree of integration should, ceteris paribus, lead to declining sovereign funding cost.

Previous studies on sovereign bond market integration have primarily focused on major developed markets, European markets and on the impact of the Euro's introduction on regional and global integration.[3] Other studies measure market integration using correlations.[4] Several studies also examine the international bond market co-movements and determinants of the yield spreads and CDS spreads.[5]

Our paper makes three important contributions. First, we test whether local risk is priced and estimate the level and time variation in the degree of market integration for a large sample of sovereign bonds from both the developed markets (DMs) and the emerging markets (EMs). We would expect the tremendous increase in the availability of substitute assets over the years, ceteris paribus, to affect foreign bond prices through local risk premia dissipation and to further integrate the international bond markets. The time varying integration index (*II*) based on the EL model accounts for the role of such assets and spans the entire range from full integration (*II=1*) to complete segmentation (*II=0*).

Second, we would expect countries with a higher level of integration to exhibit a lower borrowing cost (that is a lower expected rate of return on their sovereign bonds). Hence, we examine the impact that a 10 % increase in market integration would have on the sovereign cost of funding of both DM and EM countries.

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[2] The list of substitute assets includes a total of 255 international (global, regional, and country-specific) sovereign bond funds, 46 corporate bond (country-specific and regional) funds and 847 international (country-specific and regional) equity funds.

[3] See, for example, Barr and Priestley (2004), Abad, Chuliá and Gómez-Puig (2010, 2014), and Christiansen (2014).

[4] See, for example, Kumar and Okimoto (2011). Note that the literature on stock market integration shows that correlation may not be an appropriate measure of market integration, see for example, Carrieri, Errunza and Hogan (2007) and Pukthuatong and Roll (2009).

[5] See for example, Mauro, Sussman, and Yafeh (2002), Pagano and von Thadden (2004), Pan and Singelton (2008), Longstaff et al. (2011), Ehrmann et al. (2011), and Du and Schreger (2016).

Third, we investigate local characteristics that drive sovereign bond market integration. Indeed, a deeper understanding of these characteristics is critical for devising monetary policy in an increasingly global environment. For that purpose, we draw on the large body of literature on bond pricing, term structure models and the determinants of yield spreads to make a number of testable predictions regarding the main local characteristics that could explain the level and dynamics of integration. We identify four primarily local characteristics, namely: (1) political stability, (2) credit quality, (3) macroeconomic conditions, including, inflation, and real economic activity, and (4) liquidity. In particular, under imperfect spanning of sovereign political and credit risks, we would expect countries with lower political stability and lower credit quality to experience a lower level of integration compared to higher political stability and higher credit quality countries. We would also expect distressed countries with more illiquid bonds to experience lower levels of integration. Likewise, we would expect countries with low and stable inflation to exhibit higher levels of co-movement with each other, even after controlling for their economic globalization. Furthermore, we examine whether a lower global investment sentiment and reduced confidence in the global equity markets as proxied by Baker and Wurgler (2006) sentiment index leads to higher integration of high credit quality sovereign bond markets reflecting a "flight-to-safety" phenomenon (see, for example, Connolly, Stivers, and Sun 2005). We propose a set of hypotheses and use panel regressions to test these conjectured relationships.

We first validate the model and compare our results with different alternative models including full integration that do not depend on EL assumptions. We estimate six models for 21 DMs from 1993 or later to 2017 and 18 EMs from 2001 or later to 2017 using weekly returns. The first model is the world Capital Asset Pricing Model (WCAPM) where only the world market risk is priced. The second model is the world CAPM augmented with world sovereign and world liquidity risk factors (SOV-LIQ WCAPM). The third model is the EL model where both the world market and local unspanned market risks are priced. The fourth model is the EL model augmented with world sovereign and world liquidity risk factors (SOV-LIQ EL). The fifth model is the mixed model where the world market risk and the total variance of the bond index return are priced. The sixth model is the generalized model where the

variance of the bond index return and the covariance between the bond return and its DP are priced separately. This specification permits us to isolate the pricing effect of the integration measure over and above the local variance and world market covariance risks.

In view of the importance of domestic-currency bonds, which are extensive for advanced economies and growing rapidly for emerging economies, we conduct the analysis for local-currency denominated sovereign bonds. This is in sharp contrast with most existing studies that focus on foreign-currency denominated external debt for EMs. Note that the literature uses the terminology sovereign and government bonds interchangeably to define local and foreign currency denominated bonds. We primarily use sovereign bonds to define bonds issued by a sovereign (government) for our sample of local currency bonds.

Based on the test of zero intercept, we find stronger support for models that allow some form of segmentation such as EL, SOV-LIQ EL, mixed and generalized models than for either of the two full integration models (WCAPM and SOV-LIQ WCAPM) not only among EMs but also for DMs. The price of local unspanned risk estimated from the EL model is significant for most of our sample markets. In addition, the generalized model shows that the price of local covariance risk (between the bond index return and its DP) matters over and above local variance risk. Hence, the estimated price of local risk from the mixed model that omits this covariance risk is biased.

We next estimate the integration index for the 39 countries. There are interesting differences across countries in the dynamics of integration. In general, the Eurozone countries are more integrated compared to European Union non-euro countries. However, there are clear differences between core and periphery countries in the Eurozone specifically after the European sovereign debt crisis. The Euro sovereign debt crisis has negatively affected bond market integration of the Eurozone Periphery countries. The integration of EM sovereign bond markets is lagging behind DMs. During the sample period, the average integration for the EM pool is 0.48 compared to 0.65 for DMs.

We measure the effect of an increase in integration of sovereign bonds by 10% on the average cost of funding. Overall, a 10% increase in integration leads to a large average drop in the cost of funding of 0.86% per annum for DMs and 1.02% per annum for EMs. We obtain similar evidence regarding the effect of an increase in market integration on the cost of sovereign funding from a cross-sectional regression of the expected excess returns on the average integration measure. Thus, higher market integration can significantly mitigate sovereign funding costs.

Next, we test a number of hypotheses to study the relationship between integration and countries' political stability, credit quality, inflation rates and their volatility, industrial production growth rates and their volatility, and illiquidity. We run panel regressions and find that the better the spanning of sovereign risk, the lower the effect of political stability and credit quality on integration. Conditional on the level of economic globalization, countries with higher inflation uncertainty are less integrated but there is no evidence of significant positive association between industrial production growth rate and integration. We also find that more illiquid countries are less integrated and the effect of illiquidity is larger among distressed countries. The contribution of political stability, credit quality and to a lesser extent inflation risk decreases monotonically with increased spanning, while no such pattern is detected for illiquidity. Furthermore, the integration of the sovereign bond markets increases on average by about 10%, when a country moves from the 25th percentile to the 75th percentile as a result of higher political stability and credit quality, lower inflation and inflation risk, and lower illiquidity. We also find that reduced confidence in the global equity market leads to higher integration for high credit quality sovereign bonds but has no effect on the integration of low credit quality sovereign bonds. Additionally, we examine the effect of currency risk in terms of its impact on the integration measure and its drivers. We control for changes in foreign exchange (FX) and FX volatility in the panel regressions and find insignificant coefficients on both FX changes and FX volatility. Furthermore, we study the link between foreign holdings and bond market integration. The link is positive as expected but insignificant. Finally, we examine the effect of recent Central Bank bond purchasing programs on market integration. The public sector purchase programs reduced the sovereign bonds' free float (see Coeuré (2018)). To assess

the impact of reduced free float, we add to the panel regressions a measure of free float. We also interact free float with illiquidity to examine the effect of illiquidity on integration conditional on free float. Our results suggest that the reduced free float due to Central Bank bond purchasing programs substantially amplified the negative effect of illiquidity on market integration.

The rest of the paper is organized as follows. Section 1 presents the asset pricing model, the integration measure, and the empirical implementation. Section 2 discusses the data. Section 3 reports the asset pricing model tests and the integration estimates for DMs and EMs. It also discusses the evolution over time of the integration measures and their effect on the cost of funding. Section 4 presents and tests the main hypotheses regarding characteristics that could affect the degree of integration over time and discusses their economic significance. Conclusion follows. We report additional estimation results in the online appendix.

1. Asset pricing model and bond market integration

In this section, we lay out the asset pricing model that underlies the integration measure. We then discuss empirical implementation.

1.1 The asset pricing model and the integration measure

Measurement of sovereign bond market integration is primarily based on the law-of-one price. Empirically, the simplest price based measure would compare yields on identical maturity bonds issued by different sovereigns in the same currency denomination. If the bond market is fully integrated, the law-of-one price should hold, i.e. bonds with the same risk and currency denomination should command the same expected return regardless of the nationality of the issuer or the location of listing. Further, in an integrated market, for bonds with same risk characteristics, the priced common factor (for example the global factor) should be much more important than local factor suggesting the proportion of variance explained by the global factor as an alternate measure of integration. This measure is sometimes referred to as a news-based measure of integration (see Baele et al., 2004). In this paper, we wish to measure the

degree of integration of a large sample of sovereign bonds from both developed and emerging markets. These markets are not similar in their risk characteristics with differences in terms of local economic risk factors, explicit and implicit barriers to investments etc. Thus, deviations from law-of-one price could result from priced systematic (for example country specific) risk differences. Further, since the law-of-one-price is difficult to test empirically for our sample, we operationalize the price, news, and quantity based measures in a systematic manner that takes into account these differences.[6] More precisely, we use the integration measure as defined in EL model to estimate a time-varying comprehensive integration index. EL is a two factor (global and unspanned local factors) model that incorporates barriers to capital flows and is applicable to all types of financial assets. It uses a two-country set-up and two sets of securities. All securities traded in the U.S. (domestic) market are eligible for investment by all investors whereas securities traded in the foreign (national) market are ineligible in the sense that they can only be held by national investors. This is a reasonable description of the world market structure since in most financial markets, cross-border capital flows encounter explicit and implicit barriers. The nature, extent and severity of these barriers vary widely among markets. Generally, they are not onerous among major developed markets during tranquil times but they may be prohibitive for markets that are not well developed, undergoing a financial/currency/political crisis or have defaulted in the recent past.[7] Together, these barriers determine international investors' ability to access and willingness to invest in foreign securities either directly or indirectly through substitute assets such as different types of bond funds. Thus, capital controls, derivatives like ETFs as well as other substitute

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[6] Adam et.al. (2002) and Adjaoute and Danthine (2003) propose price-based measure such as interest rate convergence in the Euro area, and quantity based measures which take into account the differential explicit and implicit barriers that asymmetrically impact countries. They report increasing levels of integration for the Eurozone prior to the financial crisis. Baele et al. (2004) extends their framework and investigates the impact of common versus country specific information on integration. The paper shows that the proportion of local variance explained by benchmark (German) bond market increased significantly in the post-euro period for government bonds.

[7] Their bonds may be more prone to fire sale risk, and therefore investors could abstain from investing in public bonds of such markets. The reluctance of foreign investors to buy sovereign bonds from Greece and Argentina is well documented.

assets such as mutual/country funds should jointly determine the degree of integration.[8] Using a time-varying specification of the EL model, we can write,

$$E_t(r_{i,t+1}) = r_{f,t} + \theta_{W,t} cov_t(r_{i,t+1}, r_{W,t+1}) + \lambda_{I,t} cov_t(r_{i,t+1}, r_{I,t+1} | \boldsymbol{r}_{S,t+1}), \quad (1)$$

where $E_t(.)$ and $cov_t(.)$ represent moments conditional on the information available to investors at the end of time t, $r_{i,t+1}$ is the return on the *ith* bond in the *Ith* market between time *t* and *t+1*, $r_{f,t}$ is the risk free rate, $r_{W,t+1}$ is the return on the world portfolio which is a weighted sum of the world market equity return $r_{E,t+1}$, and the world sovereign bond market return, $r_{B,t+1}$, i.e., $r_{W,t+1} = M_{E,t} r_{E,t+1} + M_{B,t} r_{B,t+1}$ where $M_{B,t}$ $(M_{E,t})$ is the market value of the world sovereign bond (equity) portfolio, $r_{I,t+1}$ is the return on the *I*th sovereign bond market, and $\boldsymbol{r}_{S,t+1}$ is the vector of returns on all substitute securities that can be bought by all investors irrespective of their nationality, the covariance between the ith sovereign bond return and the world market return, $cov_t(r_{i,t+1}, r_{W,t+1})$, is the world market covariance risk, $\theta_{W,t}$ is the price of world market covariance risk, $cov_t(r_{i,t+1}, r_{I,t+1} | \boldsymbol{r}_{S,t+1})$ is the covariance between the ith sovereign bond return and the Ith market return conditional on the set of substitute asset returns, $\lambda_{I,t}$ is the price of local unspanned market risk. Thus, the expected return on the *ith* bond commands a global risk premium and a local risk premium that is proportional to the unspanned risk. At the sovereign bond market level, we can aggregate Equation (1) across bonds in each *Ith* market to obtain,

$$E_t(r_{I,t+1}) = r_{f,t} + \theta_{W,t} cov_t(r_{I,t+1}, r_{W,t+1}) + \lambda_{I,t} var_t(r_{I,t+1} | \boldsymbol{r}_{S,t+1}). \quad (2)$$

Let the diversification portfolio, DP, be the return on the portfolio (of $\boldsymbol{r}_S$) that is most highly correlated with the *Ith* sovereign bond. We can write $var_t(r_{I,t+1} | \boldsymbol{r}_{S,t+1}) = var_t(r_{I,t+1})(1 - II_{I,t})$, where $II_{I,t}$ is the time-varying integration index given by,

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[8] See for example, Bekaert and Harvey (1995), Bekaert, Harvey, Lundblad and Siegel (2011) and Carrieri, Chaieb and Errunza (2013).

$$II_{I,t} = \frac{cov_t(r_{I,t+1}, r_{DP,t+1})^2}{var_t(r_{I,t+1})var_t(r_{DP,t+1})} = \rho^2_{I,DP,t} \, , \tag{3}$$

where $\rho_{I,DP,t}$ is the correlation coefficient between the *Ith* sovereign bond and its DP at time t. Equation (2) can then be written as,

$$E_t(r_{I,t+1}) = r_{f,t} + \theta_{W,t} \, cov_t(r_{I,t+1}, r_{W,t+1}) + \lambda_{I,t} var_t(r_{I,t+1})(1 - II_{I,t}). \tag{4}$$

If the DP return fully spans the *Ith* bond return, the index $II_{I,t}$ takes on the value of one and the market is effectively integrated. In this case, only global factor shocks will determine excess returns on the *Ith* bond market. The *Ith* bond market will be completely segmented if none of the variation can be explained by the returns on substitute assets. In this case, the index $II_{I,t}$ and $cov_t(r_{I,t+1}, r_{W,t+1})$ take on the value of zero and the bond market only commands the local risk premium determined by the local price of risk and total variance risk.

Note that under the null, $r_{I,t} = r_{DP,t} + u_{I,t}$. We can then write Equation (4) as,

$$E_t(r_{I,t+1}) = r_{f,t} + \theta_{W,t} \, cov_t(r_{DP,t+1}, r_{W,t+1}) + \lambda_{I,t} \left(var_t(r_{I,t+1}) - cov_t(r_{I,t+1}, r_{DP,t+1}) \right) \tag{5}$$

$$= r_{f,t} + \theta_{W,t} \, cov_t(r_{DP,t+1}, r_{W,t+1}) + \lambda_{I,t} var_t(u_{I,t+1}). \tag{6}$$

Equation (6) implies that exposure to the global risk is the same for the Ith sovereign bond return and its DP. Thus a sudden increase in global risk, or its price, will result in a joint drop in sovereign bond and the DP prices with no effect on the level of the integration index. However, an increase in local risk, or its price, will affect sovereign bond prices with no similar effect on DP. Thus, to the extent that DP spans the globally priced risk, we should only capture the impact of the local factor on a given sovereign bond market integration measure if the local risk factor is priced.

We construct DP from the projection of the *Ith* sovereign bond return on the space of substitute asset returns. The return on DP, $r_{DP,t}$, is the fitted value $\widehat{\boldsymbol{\beta}}'_t \boldsymbol{r}_{S,t}$ from the regression,

$$r_{I,t} = \boldsymbol{\beta}'_t \boldsymbol{r}_{S,t} + u_{I,t}, \tag{7}$$

where $\boldsymbol{\beta}'_t = \Sigma^{-1}_{S,S,t} \boldsymbol{\Sigma}_{I,S,t}$ is the vector of time-varying weights of DP, $\Sigma_{S,S,t}$ is the conditional covariance matrix of substitute assets, and $\boldsymbol{\Sigma}_{I,S,t}$ is the vector of conditional covariance between the return on the sovereign bond market I and the vector of substitute assets $\boldsymbol{r}_{S,t}$. We follow the approach proposed by Chaieb, Errunza and Langlois (2018) and obtain the conditional covariance matrices from an asymmetric GARCH model. This approach has three main advantages. First, it allows the weights to vary over time because the covariance matrix of substitute assets and the covariance between the sovereign bond index and the substitute assets are also time-varying. Second, it accounts for the variation over time in the set of substitute assets as funds are listed or delisted. Third, it handles the curse of dimensionality with the extended set of substitute assets. We impose a well-conditioned correlation matrix of substitute assets to remove some of the highly correlated (redundant) substitute assets over some periods.[9]

The EL model assumes purchasing power parity over the holding period. This is not likely for the markets we consider. Given the importance of the currency factor in sovereign bond returns, we examine the effect of currency risk on the integration measures and their determinants by controlling for FX changes and FX volatility in the panel regressions as detailed in Section 4.4.

One could also argue that the EL model does not consider several factors, such as, liquidity risk, sovereign (default and political) risk. A potential concern is that the integration measure will decrease in case sovereign or liquidity risk suddenly emerges, even if these risk factors are priced *globally*. An increase in the exposure to these risk factors or their global price will result in a drop in sovereign bond prices. This should not affect the integration measure if substitute assets span the globally priced risk factors. To address these issues we consider four experiments.

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[9] We impose a condition number of the correlation matrix below 15. The condition number is the square root of the ratio of the maximum eigenvalue to the minimum eigenvalue of the correlation matrix.

First, since for most countries, the set of substitute assets does not include bond funds focusing on that country's sovereign debt, we include country equity and corporate bond funds. If sovereign risk goes up, returns on local currency sovereign bonds and local stocks go down. Therefore, equity funds could help span the variance of local-currency sovereign bond returns. As such, according to EL model, they help integrate sovereign bond markets. Similarly, because European banks and European governments' balance sheets are intertwined (Acharya, Drechsler, and Schnabl 2014), European corporate bond funds could help span the variance of European sovereign bond returns.

Second, we include the difference of the natural logarithm of the sovereign CDS spreads in the substitute assets to help span the increase in the country's probability of sovereign default.

Third, we augment the EL model with world liquidity and world sovereign risks to examine their effect on the estimated price of local unspanned risk. This also allows us to test the impact of world sovereign risk and world liquidity risk on the level and dynamics of integration. The augmented EL model (SOV-LIQ EL) is,

$$E_t(r_{I,t+1}) = r_{f,t} + \theta_{W,t}\, cov_t(r_{I,t+1}, r_{W,t+1}) + \theta_{SOV,t}\, cov_t(r_{I,t+1}, r_{SOV,t+1}),$$

$$+\theta_{L,t}\, cov_t(r_{I,t+1}, r_{L,t+1}) + \lambda_{I,t} var_t(r_{I,t+1})(1 - II_{I,t}), \qquad (8)$$

where $r_{SOV,t+1}$ is the return on the proxy for world sovereign risk factor between time t and $t+1$, $r_{L,t+1}$ is the return on the proxy for world bond illiquidity between time t and $t+1$, $cov_t(r_{I,t+1}, r_{SOV,t+1})$ is the world sovereign risk, $\theta_{SOV,t}$ is the price of world sovereign risk, $cov_t(r_{I,t+1}, r_{L,t+1})$ is the world liquidity risk, and $\theta_{L,t}$ is the price of world liquidity risk.

Fourth, in the panel regressions of Section 4, we relate the estimated integration measures to political stability and credit quality proxies, nominal and real macro variables and local bond market liquidity characteristics and examine the effect of better spanning on these relationships. We expect these country

or bond characteristics to have less effect on market integration as we gradually expand the set of substitute assets to the extent that they help better span these characteristics.

We estimate a conditional version of the model allowing the first and second moments to vary over time. Note however that introducing dynamics in the first and second moments would imply additional intertemporal state variables a la Merton (1973) and is thus internally inconsistent as argued by Dumas and Solnik (1995). As with most conditional asset pricing tests, these are not considered in the estimation of the model and the integration indices, however state variables that usually proxy for changes in investment opportunities are used as conditioning variables (see Section 2.3).

1.2. Empirical implementation

For each country, we estimate the following system of equations to test the asset-pricing model given by Equation (4),

$$er_{W,t+1} = \alpha_W + \theta_{W,t} var_t(r_{W,t+1}) + \varepsilon_{W,t+1},$$

$$er_{DP,t+1} = \alpha_{DP} + \theta_{W,t} cov_t(r_{DP,t+1}, r_{W,t+1}) + \varepsilon_{DP,t+1},$$

$$er_{I,t+1} = \alpha_I + \theta_{W,t} cov_t(r_{I,t+1}, r_{W,t+1}) + \lambda_{I,t} var_t(r_{I,t+1})(1 - II_{I,t}) + \varepsilon_{I,t+1}, \quad (9)$$

where $er_{W,t+1}$, $er_{DP,t+1}$, and $er_{I,t+1}$ denote the continuously compounded (log) excess return on the world market, *DP*, and the sovereign bond index of country *I* at *t+1*, respectively. Excess returns are defined as $er_{t+1} = r_{t+1} - r_{f,t}$. $\alpha_W, \alpha_{DP}, and \ \alpha_I$ are asset-specific intercepts. We obtain the conditional covariances of returns from the $(m \times m)$ conditional covariance matrix H_{t+1} of the $(m \times 1)$ vector of residuals $\boldsymbol{\varepsilon}_{t+1} = (\varepsilon_{W,t+1}, \varepsilon_{DP,t+1}, \varepsilon_{I,t+1})'$, where *m* is the number of assets included in the system of equations.

To test the augmented model, SOV-LIQ EL, i.e. Equation (8), we estimate the following system of equations for each country,

$$er_{SOV,t+1} = \alpha_{SOV} + \theta_{W,t}cov_t(r_{SOV,t+1}, r_{W,t+1}) + \theta_{SOV,t}var_t(r_{SOV,t+1}) + \theta_{L,t}cov_t(r_{SOV,t+1}, r_{L,t+1}) + \varepsilon_{SOV,t+1},$$

$$r_{L,t+1} = \alpha_L + \theta_{W,t}cov_t(r_{L,t+1}, r_{W,t+1}) + \theta_{SOV,t}cov_t(r_{L,t+1}, r_{SOV,t+1}) + \theta_{L,t}var_t(r_{L,t+1}) + \varepsilon_{L,t+1},$$

$$er_{DP,t+1} = \alpha_{DP} + \theta_{W,t}cov_t(r_{DP,t+1}, r_{W,t+1}) + \theta_{SOV,t}cov_t(r_{DP,t+1}, r_{SOV,t+1}),$$
$$+\theta_{L,t}cov_t(r_{DP,t+1}, r_{L,t+1}) + \varepsilon_{DP,t+1},$$

$$er_{I,t+1} = \alpha_I + \theta_{W,t}cov_t(r_{I,t+1}, r_{W,t+1}) + \theta_{SOV,t}cov_t(r_{I,t+1}, r_{SOV,t+1}) + \theta_{L,t}cov_t(r_{I,t+1}, r_{L,t+1}),$$
$$+\lambda_{I,t}var_t(r_{I,t+1})(1 - II_{I,t}) + \varepsilon_{I,t+1}. \tag{10}$$

We measure world sovereign risk of a security i, $cov_t(r_{i,t+1}, r_{SOV,t+1})$, as the covariance of the return on security i with the difference of the natural logarithm of the equally-weighted 5-year sovereign CDS spreads aggregated across all countries in our sample. We measure world liquidity risk of a security i, $cov_t(r_{i,t+1}, r_{L,t+1})$, as the covariance of the return on security i with the equally-weighted sum of the illiquidity of individual local-currency denominated sovereign bonds aggregated across all countries in our sample. We measure individual local-currency sovereign bond illiquidity with the quoted bid ask spread. There is high persistence in the global sovereign bond market liquidity of 0.86 at the weekly frequency. We, then use an AR(p) specification to compute market illiquidity innovations similar to Pastor and Stambaugh (2003) and Acharya and Pedersen (2005). Specifically, an AR(6) produces innovations in illiquidity with low autocorrelation (-0.001) and no rejection of the Ljung-Box test at the 5% level.

We model the covariance matrix H_{t+1} as a diagonal asymmetric GARCH process in which the variances depend only on past squared residuals and an autoregressive component while the covariances depend on the past cross-product of residuals and an autoregressive component,

$$H_{t+1} = H_0 \circ (ii' - aa' - bb') - \overline{H}_0 \circ cc' + aa' \circ \varepsilon_t\varepsilon_t' + bb' \circ H_t + cc' \circ \overline{\varepsilon}_t, \tag{11}$$

where i is a ($m \times 1$) vector of ones, a, b, and c are ($m \times 1$) vectors of unknown parameters, $m=3$ for EL and $m=6$ for SOV-LIQ EL, and $\circ$ denotes the Hadamard (element by element) matrix product. The shocks $\overline{\varepsilon}_{i,t} = -I_{\{\varepsilon_{i,t}<0\}}\varepsilon_{i,t}$ capture the asymmetric response of covariances to lagged shocks (see Bekaert and

Wu, 2000; Cappiello, Engle, and Sheppard, 2006). The matrices H_0 and $\bar{H}_0$ are set using the sample covariance matrix of $\boldsymbol{\varepsilon}$ and $\bar{\boldsymbol{\varepsilon}}$, respectively. The advantage of this multivariate GARCH parameterization is that it ensures positive definiteness of the covariance matrix while reducing the number of parameters to be estimated. Because the weekly returns display high levels of non-normality, we use a Student t distribution for the shocks.

Since the theory predicts that the prices of world risk factors should be the same for each country, we follow Bekaert and Harvey (1995) and use a two-stage estimation procedure. In the first stage of EL estimation, we estimate the world return equation to obtain estimates of the time-varying world market price of risk and the coefficients of the time-varying world variance. For the test of the augmented EL model, we also use estimates of the time-varying prices of world liquidity and world sovereign risks and the coefficients of their time-varying variances. In the second stage, we impose these risk prices estimates in the country estimations. This procedure results in sampling errors from the first stage, but it is more in line with the theory and produces more powerful tests. As robustness, for each country, we also use a one-step approach where we estimate the world price of market risk from the world market portfolio, DP and the sovereign bond index. This approach allows us to test whether world market covariance risk is priced for sovereign bond returns. See Online Appendix Table A.6, which shows that the world market risk premium matters for many sovereign bond markets.

We allow world and local prices of risk to change through time as suggested in the literature (see among others, Harvey, 1991 and De Santis and Gerard, 1997). Given that the model implies the prices of global and local market risks to be positive, we use a square function to model their dynamics as follows,

$$\theta_{W,t} = \left(k_{W,0} + k_W' Z_{W,t} \right)^2, \tag{12}$$

$$\lambda_{I,t} = \left(k_{I,0} + k_I' Z_{I,t} \right)^2, \tag{13}$$

where $Z_{W,t}$ and $Z_{I,t}$ are the vectors of time-varying global and local information variables. If world market risk is priced, we should reject the hypothesis that $k_{W,j} = 0$, for $j \geq 0$. If world market risk is time-

varying, we should reject the hypothesis that $k_{W,j} = 0$, for $j > 0$. If local unspanned risk is priced, we should reject the hypothesis that $k_{I,j} = 0$, for $j \geq 0$. If local unspanned risk is time-varying, we should reject the hypothesis that $k_{I,j} = 0$, $for\ j > 0$.

We use linear specifications to model the time variation of the prices of world sovereign and world liquidity risks, respectively, $\theta_{SOV,t}$ and $\theta_{L,t}$. Specifically,

$$\theta_{SOV,t} = k_{SOV,0} + k'_{SOV} Z_{W,t}, \tag{14}$$

$$\theta_{L,t} = k_{L,0} + k'_L Z_{W,t}. \tag{15}$$

In the augmented model, SOV-LIQ EL (see Equation (8)), we expect the covariance between bond return and bond illiquidity to affect required returns negatively because investors are willing to accept a lower return on an asset with a high return in times of market illiquidity (see, e.g. Pastor and Stambaugh (2003) and Acharya and Pedersen (2005)). By the same reasoning, an investor is willing to accept a discounted return on assets with high return in states of high sovereign risk. We then expect negative prices of world sovereign risk and world liquidity risk.

To test the full integration models WCAPM and SOV-LIQ WCAPM, we estimate for each country the systems of equations (9) and (10) imposing a zero price of local market risk.

We provide further validation of the EL model by estimating two other specifications. First, we estimate a restricted version of EL where we assume $(1 - II_{I,t})$ is constant equal to 1. This restricted version implies that total variance of the bond index rather than the unspanned variance is priced. We call it the mixed model. Second, we express the unspanned variance risk as the difference between the total variance of the bond index and the covariance between the bond index and its DP as in Equation (5) and allow $var_t(r_{I,t+1})$ and $cov_t(r_{I,t+1}, r_{DP,t+1})$ to have different prices. That is, we consider the relation below, termed, generalized model,

$$E_t(r_{I,t+1}) = r_{f,t} + \theta_{W,t}\, cov_t(r_{I,t+1}, r_{W,t+1}) + \lambda_{I,t} var_t(r_{I,t+1}) - \gamma_{I,t} cov_t(r_{I,t+1}, r_{DP,t+1}). \tag{16}$$

We parametrize the price of local covariance risk as a square function of local instruments,

$$\gamma_{I,t} = \left(\gamma_{I,0} + \gamma_I' Z_{I,t}\right)^2. \tag{17}$$

If local covariance risk is priced, we should reject the hypothesis that $\gamma_{I,j} = 0$, for $j \geq 0$. If local covariance risk is time-varying, we should reject the hypothesis that $\gamma_{I,j} = 0$, for $j > 0$. This specification permits us to isolate the pricing effect of the integration measure over total variance risk and world market covariance risk.

2. Data

The estimation of the asset pricing models and of the integration indices requires three groups of data. First, returns data on the sovereign bond indices. Second, data on the substitute assets used to construct the diversification portfolios. Third, the global and local conditioning variables to parametrize the global and local prices of risk (Equations (12), (13), (14), (15) and (17)). The data used for the panel regressions is detailed in Appendix B and discussed in Section 4.

2.1 Sovereign bond indices

We use local currency-denominated sovereign bond indices with all maturities. Local currency debt is significant in DMs and is increasingly important in EMs. As of 2013, EM local currency sovereign bond market represented 50% of the total EM bond market. Brandão-Marques et al. (2015) document that local-currency bond funds have expanded more rapidly than hard-currency bond funds. For a sample of 14 emerging markets and over the past decade, Du and Schreger (2016) document that the average fraction of external sovereign debt in local-currency increased from around 15% to almost 60%.

To be included in the sample, we require that sovereign bond return data begin at the latest by December 2007. This criterion results in a total of 39 countries, with 21 developed (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, and UK) and 18 emerging markets (Brazil, Chile, China, Colombia, Czech Rep., Hungary, India, Indonesia, Malaysia, Mexico, Poland,

Peru, Russia, South Korea, South Africa, Taiwan, Thailand, Turkey). Our sample of EMs encompass the largest issuers of local currency sovereign bonds. For each country, we use the bond index with the longest historical span. Thus, we get the bond indices of 12 DMs from Citigroup (CITI/SSS). For Belgium, Germany, Greece, Japan, Norway, Portugal, and U.K., we use the Bank of America Merrill Lynch (BOAML). For Canada, we use S&P, and for Singapore, we use JP Morgan. For the emerging market bonds, we use JP Morgan GBI except for Taiwan, we use Bank of America Merrill Lynch. All bond indices are market cap-weighted and rebalanced monthly.

The return data are available through Datastream. Since the substitute assets used to construct the DPs are in USD, the local currency bond returns are also expressed in USD and hence are unhedged dollar returns. The sample has different starting dates for each country, depending on when the data become available. We choose to study returns at a weekly frequency to have enough observations to capture the time-dynamics in covariances and to alleviate the problems caused by nonsynchronous trading at higher frequencies.

Since a substantial percentage of stale prices would prevent proper estimation of the diversification portfolio, and hence proper measurement of the integration index, we use filters discussed in Appendix A to minimize the occurrence of stale prices. For example, if we detect three consecutive observations with zero returns we eliminate these observations. At the index level, this is due to data discontinuity. Appendix A reports details on data filters and Online Appendix Table A.1 shows summary statistics for sovereign bond returns across the sample of DMs and EMs. In the DM sample, the average bond returns are highest for Australia and New Zealand. In the EM sample, the average bond returns range from 3.6% in Malaysia to 14.4% in Brazil. Non-normality is strongly present in the weekly bond returns of DMs and EMs. Among DMs, only Greece, New Zealand, and Sweden show evidence of significant autocorrelation in their returns at 5% level as indicated by the Ljung-Box test statistic $Q(z)_{12}$ for 12th-order serial correlations in the returns. Among EMs, $Q(z)_{12}$ is significant at 5% level in 12 countries. Furthermore, the Ljung-Box test statistic for 12th-order serial correlations in the squared returns,

$Q(z^2)_{12}$ strongly suggests the presence of time-varying volatility for all DMs except Switzerland and for all EMs. The Engle–Ng test statistic indicates the presence of significant negative asymmetry at 5% level in 11 DMs and 16 EMs and significant positive asymmetry at 5% level in 12 DMs and 6 EMs.

2.2. Substitute assets and the diversification portfolios

The EL model suggests that we should use as large a set of substitute assets as possible to hedge out local market risks. We thus include the following financial assets in our largest set: the world sovereign bond index proxied by the Citigroup World Government Bond Index, and the world market equity index proxied by the Datastream TOTMK World index. We add the sovereign bond open-end funds (OEFs), closed-end funds (CEFs) and exchange traded funds (ETFs). From the universe, we select funds with at least 50% exposure to government bonds as classified by Bloomberg. We include global, regional and country-specific sovereign bond funds. For each country, we add the regional funds if the country is part of that region. We also add the global sovereign bond funds to all countries.[10] For EMs, we also add the JP Morgan EMBI dollar-denominated sovereign bonds, which are widely accessible to US institutional investors.[11]

Because sovereign bond funds that focus on a specific country are not available for many countries, we further add the equity and corporate bond funds. Since, the bond and equity asset classes are integrated to some extent, equity and corporate bonds could help better span the sovereign bond market. Also, equity and corporate bond funds could help better span country-specific risks such as sovereign risk. From the universe, we select corporate bond funds with at least 50% exposure to corporate bonds and equity funds with at least 95% exposure to equity as classified by Bloomberg. We include country and regional equity and corporate bond funds.

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[10] Funds are listed on any of the following primary exchanges: New York, NYSE Arca, NASDAQ CM, NASDAQ GM, NASDAQ GS, OTC US, OTC Markets, OTC BB, LSE European Quotes Service, London, Euronext Amsterdam, Euronext Brussels, Euronext Lisbon, Euronext Paris, Xetra, Xetra ETF, Xetra Intl Mkt, Luxembourg, Singapore, Hong Kong.
[11] JP Morgan EMBI includes U.S. dollar denominated Brady bonds, loans, and Eurobonds issued or guaranteed by emerging market governments.

Lastly, we include the change in 5-year sovereign CDS spread natural logarithm to help span the changes in the country's probability of default.[12] We label SET 4 this largest set of substitute assets and use it to build the diversification portfolio DP_4. We use this portfolio for all subsequent main asset pricing tests as well as in the main panel regressions results. We also construct the diversification portfolios using reduced sets to examine the effect of better spanning on the relationship between market integration and its drivers as discussed in Section 4. We define SET 3 as SET 4 excluding the 5-year CDS, SET 2 is SET 3 excluding equity funds and corporate bond funds, and SET 1 is SET 2 excluding sovereign bond funds.

[Insert Table 1]

Table 1 reports the distribution of the funds by asset class (sovereign bond, corporate bond, equity), by type (CEF, ETF, OEF), and geographical focus (country, region, global). In total and after screening for data quality and time series length, we have more than a thousand funds.[13] Only Australia, Canada, Chile, China, France, Germany, Italy, Japan, Spain, and UK have country-specific sovereign bond funds (CEF, ETF, or OEF). Most of the country-specific sovereign bond funds are ETFs. Corporate bond funds with country-focus are only available for Australia, China, Germany, Russia, and UK. But, equity funds with country-focus are available for most countries. See Online Appendix Figure A.1 for the time evolution of the number of funds by asset class (Panel A), and by type for each asset class (Panels B-D).

Notwithstanding the large number of substitute securities initially used in the estimation of the DPs for each country, the final number of substitute securities retained in the DPs for each country does not exceed 40 after we impose the constraint on the condition number of the correlation matrix to remove

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[12] We use Markit sovereign 5-year CDS spreads quoted in US dollars on foreign currency sovereign debt. For each country, we select the most common restructuring clauses that ensures the highest liquidity and longest time span. We then select Cumulative Restructuring clause for Europe, North America, Asia, and Emerging Markets. For Australia and New Zealand, we select the Modified Restructuring. Because of occasionally missing observations in the time series, we use linear interpolation techniques to obtain a complete set of weekly estimates of credit quality for all countries.

[13] We filter out funds with time series returns data shorter than 100 weeks and funds, which had less than 95% data available for the period between their inception and end of sample period. See Appendix A for further details.

redundant substitute assets (see footnote 9 for further details). The number of retained substitute securities vary across countries and over time but is not trending up. Indeed, in many countries the trend is downward because of new substitute assets that better span the sovereign bond indices. Online Appendix Figure A.2 plots for representative countries, the average number of substitute assets retained each year in DP_4.

Online Appendix Table A.2 shows the pairwise unconditional correlations between the sovereign bond indices and their diversification portfolios obtained from the four different substitute asset sets. The unconditional correlations between the bond indices and their respective DP_4 are larger or equal to unconditional correlations with the other DPs in 18 countries. Country equity and corporate bond funds help better span the variance of local-currency sovereign bond returns. The spanning effect of equity and corporate bond funds is larger for EMs than for DMs. The spanning effect of 5-year sovereign CDS is rather marginal.[14]

2.3. Global and local instruments

We follow the extant literature in specifying the instruments.[15] The global instruments include: the US term spread measured by the yield difference between the 10-year T-bond and the 3-month T-bill, and the world dividend yield in excess of the US T-bill. The local instruments include: the local dividend yield in excess of the US T-bill and the local bond excess return.[16] All instruments are one-period lagged.

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[14] The time-varying weights of DP maximize the conditional correlation between the country sovereign bond index and its DP. However, the unconditional correlation is not necessarily higher when the set of substitute assets is enlarged.

[15] There is strong evidence for predictability of sovereign bond returns. Excess bond returns are predictable by the dividend yield (see, Fama and French, 1989), and the yield spread (see, for instance, Fama and Bliss (1987) and Campbell and Shiller (1991). See, also, Cochrane and Piazzesi (2005), and Ludvigson and Ng (2009) for US evidence and Solnik (1993), Ilmanen (1995), and Dahlquist and Hasseltoft (2013) for international evidence.

[16] As robustness, we use instead the local stock market index return proxied by TOTMK equity index provided by Datastream and the change in exchange rate as local instruments to parametrize the local price of risk. Untabulated results show that this experiment did not affect our main findings on asset pricing specification tests and market integration. However, the magnitude and dynamics of the local prices of risk for some countries are sensitive to the choice of instruments. For Canada, China, Columbia, Singapore, Switzerland, and UK, we report results based on this set of instruments.

3. Empirical results

In this section, we discuss estimation results of the world market equation, validate the EL model, and test whether world and local unspanned risks are priced. We also examine the level and dynamics of the integration indices as well as the impact of better spanning and higher integration on the country's cost of funding.

3.1 Prices of world risk factors

In the first stage, we estimate the world market equation based on WCPAM over January 06th, 1993 to August 30th, 2017. The world market portfolio is the market capitalization weighted world sovereign bond index proxied with Citigroup World Government Bond Index (WGBI) and world equity index proxied with Datastream TOTMK World market index (TOTMKWD). We also estimate the world market equation based on WCAPM augmented with world sovereign and liquidity risk factors (SOV-LIQ WCAPM) over January 10th, 2001 to August 30th, 2017 when CDS data are available.

[Insert Table 2 and Figure 1]

First, we run the zero intercept test where the intercept of the world market portfolio equation is obtained from WCAPM or SOV-LIQ WCAPM. The intercept estimated from WCAPM is significant at 5% level and amounts to 0.13% per week (6.76% per annum). The intercept estimated from SOV-LIQ WCAPM remains significant at 5% level and amounts to 0.2% per week (10.4% per annum). Second, we run specification tests on the prices of world risk factors. Panel A of Table 2 reports robust Wald test statistics for the significance and time variation in the price of world risk factors and their average prices over time. Columns (i)-(iii) report these results for the price of world market risk estimated from WCAPM. The price of world market risk is significant at 1% level and averages 2.62. We do not reject the null hypothesis that the price of world market risk is constant. However, Figure 1 shows large fluctuations over time in the price of world market risk and a peak during the global financial crisis and the Euro sovereign crisis. Shaded areas in Figure 1 indicate NBER recession periods and major financial

crisis. Columns (iv)-(xii) report the test statistics, and average prices of the three risk factors from SOV-LIQ WCAPM. The price of world market risk is significant but not time varying and averages 4.33. The price of world sovereign risk is positive but insignificant, and small. The price of world liquidity risk is not statistically significant but is negative as expected with an average of -4.4.[17]

The estimation results of the WCAPM and SOV-LIQ WCAPM are robust to other specifications of the set of global instruments. Specifically, adding the US Fed Funds rate and the US default spread measured by the yield difference between Moody's Baa and Aaa rated bonds do not change inferences on the pricing of the global risk factors. These additional robustness checks are untabulated and available from the authors.

3.2 Validation of the EL model

Although the EL framework is general in terms of its application to all types of financial assets, to-date, it has not been used in the context of sovereign bond pricing. Hence, we validate the model and examine whether the price of local unspanned risk is significant as predicted by the model. We also compare our results with alternative full integration models that do not depend on EL assumptions. In addition, we test other specifications that enable us to isolate the pricing effect of the integration measure over total variance risk and world market covariance risk.

Panel B of Table 2 shows the average country estimation results across DMs and across EMs. For each country, we first estimate four models; (1) full integration WCAPM, (2) full integration SOV-LIQ WCAPM, (3) mild segmentation EL, (4) mild segmentation EL augmented with world sovereign and

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[17] Results on SOV-LIQ WCAPM are overall robust to whether aggregating CDS and bid ask spreads across only DM countries or across all countries when we construct the proxies for world sovereign and liquidity risk factors. In one experiment, we construct the sovereign risk factor excluding, every week, the countries with the highest CDS spreads that fall above the 90th percentile. Online Appendix Table A.5 shows that inferences about the intercept (α (%)) are robust to the alternative proxies of the sovereign risk factor. Nevertheless, we acknowledge that test results of the augmented models could be partially driven by measurement errors in the sovereign risk or liquidity risk proxies.

liquidity risk factors (SOV-LIQ EL).[18] Based on the significance and magnitude of the intercept, we find stronger support for the mild segmentation models than for full integration models among not only EMs but also DMs. For example, the intercept of WCAPM is significant at 5% level for 13 DMs and 15 EMs, while the intercept of the EL model is significant only for 4 DMs and 4 EMs. The magnitude of the intercept across WCAPM and EL models is comparable though lower in absolute terms for EL in most DMs and EMs. The per week cross-sectional average intercept $|\alpha|$ across DMs are 0.07% and 0.05% for WCAPM and EL. That is a difference in average $|\alpha|$ of about 1.04% per annum for DMs. The per week average intercept $|\alpha|$ across EMs are 0.15% and 0.11% for WCAPM and EL. That is a difference in average $|\alpha|$ of about 2.08% per annum for EMs. We next augment WCAPM and EL with sovereign and liquidity world risk factors. We reject the null of zero intercept in 14 EMs for SOV-LIQ WCAPM but only for one EM for SOV-LIQ EL. The null of zero intercept is rejected in 17 DMs for SOV-LIQ WCAPM but only for 8 DMs for SOV-LIQ EL. The per week average intercept $|\alpha|$ across DMs are 0.12% and 0.07% for SOV-LIQ WCAPM and SOV-LIQ EL. That is a difference in average $|\alpha|$ of about 2.6% per annum for DMs. The per week average intercept $|\alpha|$ across EMs are 0.19% and 0.09% for SOV-LIQ WCAPM and SOV-LIQ EL. That is a difference in average $|\alpha|$ of about 5.2% per annum for EMs. Overall, based on the significance and magnitude of the intercept, the EL model performs better than the full integration models WCAPM or SOV-LIQ WCAPM. Including exposures to world sovereign and liquidity risk factors to the EL model provides further support to the EL model.

The price of local unspanned market risk estimated from the EL model is significant for 13 DMs and 14 EMs and is time varying for 6 DMs and 5 EMs. There is large variation in the average prices of local unspanned risk across countries. The cross-sectional average is 5.78 among DMs and 6.89 among EMs. Based on the SOV-LIQ EL model, Online Appendix Table A.4 shows that local unspanned market risk is significantly priced in the presence of world sovereign and liquidity risk factors in 16 DMs and 14

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[18] The model estimations for SOV-LIQ WCAPM and SOV-LIQ EL start in 2001 or later depending on CDS and sovereign bond return data availability.

EMs and is time varying in 11 DMs and 7 EMs. Specifically, we find strong evidence for pricing of the local unspanned market risk in 7 of the 11 Eurozone countries over the more recent period 2001-2017.[19] The evidence regarding the significant price of local unspanned market risk based on EL and SOV-LIQ EL suggests that local market risk is an important driver of sovereign bond expected returns not only in EMs but also in DMs.

We next consider two additional alternative models that allow both global and local risk sources, (1) mixed model that shuts down the integration measure, i.e. we impose a zero integration level for all countries and thus price world market covariance risk and total rather than unspanned variance risk, and (2) generalized model that allows for different prices of risk for the total variance of the sovereign bond return and for its covariance with DP. Based on the tests of zero intercept, the mixed model shows similar performance to the EL model (average $|\alpha|$ of 0.05) for DMs and very similar (average $|\alpha|$ of 0.10 for mixed and 0.11 for EL) for EMs. Both models outperform the full integration models WCAPM and SOV-LIQ WCAPM for DMs and EMs. The generalized model shows on average a lower intercept compared to full integration models (average $|\alpha|$ of 0.06 versus 0.07 and 0.12) but a slightly larger intercept (0.06 versus 0.05) compared to mixed and EL models for DMs. It also has a higher intercept compared to mixed, SOV-LIQ EL and the EL model for EMs.

Specification tests (untabulated) of the mixed model shows that the price of total variance risk is significant for 14 DMs and 13 EMs. Similarly, as reported in Panel B of Table 2, specification tests of the generalized model show that the price of total variance, $\lambda_{I,t}$, is significant in 13 DMs and 15 EMs. Also the price, $\gamma_{I,t}$, of the covariance between the bond return index and its DP is significant for 16 DMs and 16 EMs. Indeed, $\text{cov}_t(r_{I,t+1}, r_{DP,t+1})$ matters over and above total variance risk and world market

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[19] Unfortunately, we could not conduct a clinical analysis of Euro adoption as per the Euro experiment in the equity market of Karolyi and Wu (2018) because weekly data on bond returns before Euro adoption is not available for a long enough period to conduct GARCH in mean analysis for the Eurozone countries.

covariance risk. The mixed model that omits this risk factor is underspecified and the estimates of the local price of risk from this model are biased.

We further consider a comparison across models by evaluating the cross-sectional correlation between the average excess returns and the predicted excess returns generated by the various models; WCAPM, SOV-LIQ WCAPM, SOV-LIQ EL, mixed, generalized, and EL. Panel B of Table 2 shows that the EL and the SOV-LIQ EL models are superior to the mixed and the generalized models because they display a higher cross-sectional correlation between average excess returns and predicted excess returns for both DMs and EMs. Although, the EL and the SOV-LIQ EL models outperform the standard World CAPM or the SOV-LIQ WCAPM for EMs they do not perform as well in the case of DMs. This is consistent with the financial markets evidence that DMs are more integrated than EMs. However, based on this metric, the mixed and generalized models underperform all the other models for both DMs and EMs. Online Appendix Table A.3 reports all above tests by country.

In summary, models that allow some degree of segmentation outperform fully integrated models based on the zero intercept test for both DMs and EMs. The asset pricing estimations demonstrate the existence of the local unspanned market risk premium predicted by the EL model in the case of both DMs and EMs. The best performing models for EMs are the ones with an important role for unspanned (rather than total) local risk as they also generate the highest cross-sectional correlation between average returns over the sample with the expected returns predicted by these models. But for DMs, the evidence is mixed since fully integrated models generate the highest cross-sectional correlation between average returns and the expected returns. Taken together, the evidence suggests that the EL model is superior for pricing local EM sovereign bonds. However, the evidence for DMs is weaker. We caution the readers about the fragility of the evidence in favor of the EL model especially for DMs. Nevertheless, the integration index based on the conditional second moments is robust to misspecification in estimation of the expected returns as further detailed in Section 3.3.

Online Appendix Table A.3, Panel B, provides some diagnostics for the residuals of the EL model. For most countries, the Bera-Jarque (B-J) test statistic for the standardized residuals is lower than the corresponding test statistic for the excess returns (see Online Appendix Table A.1). Nevertheless, the hypothesis of normality is rejected in all cases. This suggests that, the GARCH parameterization and fat-tailed t distribution can accommodate some but not all of the kurtosis in the weekly returns data. At the 5% level and based on the Ljung-Box statistic of serial autocorrelation up to order 12 ($Q(z^2)_{12}$), there is no more significant serial correlation in the squared standardized residuals except for 2 DMs and 4 EMs. The Engle–Ng tests indicate that there is no evidence of positive asymmetry in the residuals for DMs and most EMs. But, negative asymmetry remains in three DMs and 10 EMs. Overall, the diagnostic results support our use of the multivariate asymmetric GARCH process. We next examine the integration measure, its trends and the economic impact of higher integration on the sovereign cost of funding.

3.3 Integration measure estimates

Table 3 contains our main results for the integration measure estimated from the EL model. We report statistics for the three most and three least integrated markets among DMs and among EMs. Among DMs, the estimated integration measures are highest for Germany, Belgium, and France, and lowest for New Zealand, Switzerland, and Sweden. Among EMs, Poland, Turkey and Hungary are the most integrated, while Taiwan, Chile, and China are the least integrated.[20] The sovereign bond markets integration of EMs is lagging behind DMs. This is not surprising in view of the recent history of the EM local sovereign bond market and the youth of its yield curve. The average integration for EMs is 0.48

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[20] Since the average integration measure could suffer from noise in the estimation process, we run the stationary block-bootstrap of Politis and Romano (1994) on the most integrated countries of our sample, namely, Belgium, France and Germany. Based on 1000 bootstrap replications, we obtain an integration mean of 0.85, 0.84, and 0.86 compared to the EL model estimated mean of 0.84, 0.80 and 0.85 for Belgium, France and Germany, respectively. At 95% level, the block-bootstrap confidence intervals for the averages integration measures are [0.82; 0.88], [0.80; 0.87], [0.83; 0.88] for Belgium, France, and Germany respectively. Therefore, within the confidence intervals, Belgium, France and German bond markets are highly integrated with global bond markets.

compared to 0.65 for DMs. Emerging Europe stands as the most integrated among EMs, while Emerging Asia is the least integrated. [21]

[Insert Table 3 and Figure 2]

We also test for the presence of time trends in the integration measures using Bunzel and Vogelsang (2005) linear time trend test, which is robust to strong serial correlation and unit root in the data and has good size and power properties both asymptotically and in finite samples. [22] We estimate the error variance non-parametrically using "Daniell kernel" since tests based on this kernel maximize power among a wide range of kernels. Table 3 shows the trend coefficient in percent per annum and its t-stat for the three most and three least integrated countries. Online Appendix Table A.8 reports these tests by country (Panel A) and for all spanning sets (Panel B). [23] The trend coefficient is positive and highly significant for only 8 DMs and 7 EMs and is significantly negative for Switzerland. Also, we run panel regressions of the integration measures by region, on a trend and with country fixed effects. The standard errors are two-way clustered by country and time. Table 3 reveals a statistically significant upward trend at 1% confidence level but the magnitude of the trend of 1.03% and 0.59% per annum for DMs and EMs, respectively, is small. Among DMs, the largest trend is for Eurozone core countries and among EMs the largest trend is for Asian countries.

Panel A of Figure 2 plots the cross-sectional averages (equally-weighted) of the integration indices across DMs and across EMs. The shaded areas correspond to March to November 2001 NBER recession and five key crises periods. They are the European exchange rate mechanism crisis (September 1992 to August 1993), the Tequila crisis (December 1994 to January 1995), the East Asia crisis (June to December 1997), the Russian default and Long-Term Capital Management crisis (August to December 1998), the October 2007-May 2009 global financial crisis (GFC), and the January 2010 to December

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[21] Inferences on global risk premia (see Section 3.1) and local risk premia (see Section 3.2) as well as the estimated integration indices are robust to the use of a pure world bond market instead of a combination of bonds and stocks as proxy for the world market portfolio. Untabulated results are available from the authors.

[22] We use the Bunzel and Vogelsang (2005) test based on the *J* unit root test statistic of Park (1990) and Park and Choi (1988) to scale the linear trend test statistic.

[23] We convert the coefficient to annual numbers by multiplying the coefficient by 52 (number of weeks in 1 year).

2012 Euro-sovereign debt crisis (ESC). The figure shows that the estimated level of integration varies over time. There is a general upward trend but reversals occur during crises. Interestingly, the addition of world sovereign and liquidity risk factors to the EL model does not affect the level and dynamics of market integration as shown in Panels B and C of Figure 2 for DMs and EMs. Indeed, the integration index computed from the conditional second moments is rather robust to misspecification in the estimation of the expected returns. Similarly, Chaieb and Errunza (2014) find that, though currency risk is conditionally priced, it does not affect the level and the dynamics of the integration measure for international equity markets.

We also examine the effect of USD denominated EM bonds on the level of integration of local-currency denominated EM bonds. On average, across EMs, the effect of USD denominated EM bonds on the level of integration is marginal when using the largest set of substitute assets, SET 4. However, with the smallest set SET 1, the availability of USD denominated EM bonds helps integrate the local currency sovereign bonds. The average differential in integration across EMs and over time, under SET 1, is about 0.2. (See Online Appendix Figure A.4).[24]

3.4 Global financial and Euro sovereign debt crises

Next, we estimate the effect of the global financial crisis (GFC) and the euro sovereign debt crisis (ESC) on the integration indices. For each country, we run regressions of the integration measures on a constant, a trend, and the dummies for GFC and ESC. Table 3 shows the coefficients of the dummies and their t-stats for the selective countries. (See Online Appendix Table A.8 for results on all countries of our sample.) The standard errors are heteroskedasticity and autocorrelation consistent obtained from Newey-West (1987) correction with six lags. We find a statistically significant but small drop in 6 DMs and 9 EMs during GFC. We also find a significant drop in integration during ESC of some but not all

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[24] In our main specifications, we do not include the U.S. Treasury bond indices in the sets of substitute assets. Their inclusion would bias upward the trend coefficient of market integration because of the correlation among international yield curves that trended up over time (see, for example, Kaminska, Meldrum, and Smith (2013), and Dahlquist and Hasseltoft (2013). As robustness, we include the U.S. Treasury bond indices in SET 4. We find no significant change on the mean integration measures but the trend coefficients and their t-stats increase for some countries (see Online Appendix Table A.7).

periphery countries. Greece, Ireland and Portugal experienced a drop of about 0.2 in their level of integration, with no drop for Italy and a small one in Spain. The inclusion of substitute assets such as ETFs did not help integrate Greece, Ireland and Portugal during the ESC crisis. This could be driven by a deterioration of ETF's tracking performance during the crisis period (see, for example, Drenovak, Urosevic and Jelic, 2014). For EMs, with few exceptions, the impact of the euro crisis on integration measure is positive and small.

Table 3 also shows panel regressions of the integration measures on crises dummies, a trend and country fixed effects for DM and EM regions. GFC has a small positive impact on DM regions and a small negative impact on EM regions. ESC has no effect on the integration of the core DM countries while periphery DM countries witnessed a drop of about 0.12 in their integration. The reversals for the periphery countries could be due to the "wake-up call" contagion (see Goldstein 1998, and Beirne and Fratzscher, 2013). The negative impact of the ESC crisis is consistent with the increased importance of the local factors.

3.5 Cost of funding

We now examine the impact of an increase in market integration on the cost of sovereign funding. Specifically, we measure the effect of an increase in integration of sovereign bonds by 10% on the average cost of funding (COF) given by,

$$\Delta E\left(r_{I,t+1}\right) = -10\% \times E\left(\lambda_{I,t} var_t\left(r_{I,t+1}\right)\right). \tag{18}$$

The last columns of Table 3 show the results for the selective DMs and EMs. We also show the averages across regions. The magnitude varies widely across countries because of the variation in the estimated prices of risk and of the level of volatilities of the sovereign bond markets. For DMs, the average decrease in the cost of funding ranges from as low as 0.11% per annum in Switzerland to as high as 5.8% per annum in Greece. For EMs, it ranges from 0.12% per annum in Malaysia to 4.3% per annum in Hungary. Overall, a 10% increase in integration results in a large drop in average cost of

funding by 0.86% per annum for DMs and 1.02% per annum for EMs. (See Online Appendix Table A.8 for results on all countries of our sample.)

We obtain similar evidence on the effect of an increase in market integration on the cost of sovereign funding from cross-sectional regression of the expected excess returns (per annum) on the average integration measure. The slope is negative and its absolute value is seven. That is a country with 10% higher integration, on average, shows a lower average cost of funding of about 0.7% per annum. The negative relation between expected returns and integration is the crucial prediction associated with the EL model. Thus, our integration measure for the sovereign bonds is useful because it is associated with the risk pricing effects predicted by the model.

4. Characteristics influencing sovereign bond markets' integration

The results of Section 3 suggest that most countries' sovereign bonds are not fully integrated and that their excess returns command local risk premia. In this context, we further examine whether, local bond market characteristics and countries macroeconomic fundamentals, that have been shown to play a key role as determinants of the term structure of interest rates and yield spreads, also influence the integration measures of sovereign bond markets. More precisely, we focus on four main local characteristics: political stability, credit quality, macro-economic conditions, and sovereign bond market liquidity. Indeed, as spanning increases, these local characteristics should gradually lose significance in explaining our integration measure. In addition, we examine if global market confidence – as proxied by the Baker and Wurgler (2005) sentiment index- also influences sovereign bond markets' integration. Finally, we examine the effect of currency risk, foreign holdings, and the reduced free float resulting from the Central Banks' bond purchasing programs on market integration.

4.1 Hypothesis Development

<u>Political stability</u>:

The role of the legal system and of political institutions on financial development and economic growth is well established in the literature (see among others La Porta et al. 1997, 1998; Rajan and Zingales, 2003; Stulz, 2005; Karolyi, 2015). Duffie, Pedersen and Singleton (2003) show that Russian yield spreads respond to political events. Foreign investors are attracted to safe countries with strong institutions. We conjecture that political stability should affect the integration of sovereign bond markets under imperfect spanning of sovereign risk. We hypothesize that,

H1: Under imperfect spanning of sovereign risk, greater political stability should lead to higher integration.

We expect that better spanning should decrease the influence of political stability proxies on the integration measure. As a corollary, moving from the set of substitute assets SET 1 to SET 4, we should observe a gradual reduction in the impact of country specific political stability on the integration measure if the enlarged set better spans sovereign political risk. We use the political risk index (*POL*) computed by the Political Risk Services' International Country Risk Guide (ICRG) that combines several components, such as quality of institutions, conflict, democratic tendencies, and government actions. The range of the rating index goes from 0 to 1. A higher number indicates lower political risk.

<u>Credit Quality:</u>

To the extent that unspanned country specific sovereign risk is locally priced, a lower credit quality should be associated with lower level of integration.[25] In particular, a change in the perceived probability

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[25] Duffie, Pedersen and Singleton (2003) model both the systematic and sovereign-specific components of sovereign credit risk. Geyer, Kossmeier and Pichler (2004), Pan and Singleton (2008), and Longstaff et al. (2011), among others, show that a large part of sovereign credit risk is related to common global factors. Ang and Longstaff (2013) find that systemic credit risk constitutes about 31% of the total credit risk of the Eurozone sovereigns. Remolona, Scatigna, and Wu (2008) find that country-specific fundamentals drive sovereign risk

of default could affect sovereign bond prices and returns. As long as the set of substitute assets does not allow perfect spanning of local credit risk, a period of volatile probability of default should be associated with lower integration. Our second hypothesis is,

H2: Under imperfect spanning of sovereign risk, higher country-specific sovereign credit quality should be associated with higher level of market integration.

Here as well, we primarily focus on SET 4 but also examine integration indices estimated with reduced DPs to test H2. Specifically, we examine whether improved spanning of sovereign credit risk reduces the impact of a change in the probability of default on the integration measure. To proxy for country-specific credit quality, we use the S&P credit rating (*CREDIT*) linearly transformed into a numerical format ranging from 1 (Default) to 21 (AAA).[26]

Macroeconomic conditions:

Uncertainty about future inflation is often cited by financial market participants as an important source of risk in nominal bond returns. For the US, Ang and Piazzesi (2003) stress the role of macro characteristics (inflation and real economic activity) in explaining the yield curve dynamics specifically at the short end and middle of the yield curve. Ludvigson and Ng (2009) show that real macroeconomic and inflation variables have important forecasting power for future excess returns on U.S. government bonds. Wright (2011) emphasizes the role of inflation uncertainty on term premia. Hilscher and Nosbusch (2010) find local macroeconomic fundamental's levels and volatility to be the dominant characteristics for emerging markets sovereign yield spreads. Baele, Bekaert and Inghelbrecht (2010) show that uncertainty about inflation and output are important in fitting bond return volatility. Burger et

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while global risk aversion is the dominant determinant of time-variation in sovereign risk premia. Augustin (2015) shows that the relative importance of global vs. local risk factors depends on the slope of the CDS term structure.

[26] Country credit rating is highly correlated to political risk proxies. Country credit rating could then partially capture political risk, see, for example, Bekaert et al. (2014). Furthermore, political and credit ratings could also capture macroeconomic and, in particular, inflation conditions.

al. (2015) examine US investments in global bonds and find a significant role of macroeconomic fundamentals, especially inflation volatility. To the extent that real and nominal national macro variables affect the local risk premia, their level and volatility should negatively affect the level of integration. However, real and nominal macro factors could affect local sovereign bond prices not only through risk premia but also through economic globalization. We then conjecture,

H3a: Conditional on the level of economic globalization, higher country-specific inflation and inflation volatility should lead to lower level of integration.

H3b: Conditional on the level of economic globalization, weaker country-specific real macroeconomic growth and higher real macroeconomic growth volatility should lead to lower level of integration.

To capture the nominal and real macroeconomic conditions of the sovereign bond issuing country, we use inflation (π), inflation risk ($\sigma(\pi)$), change in industrial production (ΔIP), industrial production risk ($\sigma(IP)$). The inflation measure is based on the Consumer Price Index (CPI). All growth rates, including inflation, are measured as the difference in logs of the index at time t and t-12, t in months. We measure inflation risk with inflation volatility and industrial production risk with industrial production growth volatility. For each country, we estimate the volatility dynamics of inflation and of industrial production growth rate by using similar approach to Schwert (1989) and fitting an ARMA(p,q)-GARCH(1,1). We use the Bayesian Information Criterion (BIC) to select the best ARMA(p,q) specification.

In the panel regressions, we control for the level of economic globalization through covariation in nominal and real macroeconomic conditions using the three-year rolling correlation between the US and the country's inflation rates to test H3a and the three-year rolling correlation between the US and the country's industrial production growth rates to test H3b.

<u>Illiquidity</u>

Illiquidity level and risk affect the pricing of bonds (Alquist, 2010) and are important determinants of yield spreads especially in times of distress. Beber, Brandt, and Kavajecz (2009) show that liquidity plays a non-trivial role in explaining sovereign yield spreads for low credit risk countries and in times of heightened market uncertainty. We thus conjecture,

H4a: Sovereign bond market illiquidity should be negatively related to the level of integration.

H4b: The impact of illiquidity on sovereign bond market integration should be more pronounced for distressed countries

As illiquidity measure, *ILIQ*, we use the quoted bid-ask spread. We construct the measure for each country from her individual local-currency denominated sovereign bonds (from Bloomberg) and build equally weighted monthly averages in order to test H4a.

To proxy for distressed countries, we use a dummy, $D_{CDS10-CDS1}$, equal to one when the slope of the term structure of CDS spreads is negative and 0 otherwise. The slope of the term structure is the difference between the 10 and 1-year CDS spreads. Lando and Mortensen (2005) show that the term structure of CDS spreads is closely linked with conditional default probabilities and this link suggests a downward sloping term structure of credit spreads for highly risky issuers. To test H4b, we interact $D_{CDS10-CDS1}$ with *ILIQ*. H4b holds if the coefficient on $D_{CDS10-CDS1} \times ILIQ$ is negative.

We finally examine if global investor sentiment could also affect bond market integration.

<u>Global investor sentiment</u>

Past studies show that investor sentiment is an important driver of emerging market bond spreads (see, for example, Eichengreen and Mody, 1998, Baek, Bandopadhyaya and Du, 2005, Diaz-Weigel and Gemmill, 2006). As investors lose their general appetite for risk or face higher volatility in the equity

markets, they may reallocate to high credit quality sovereign bonds reflecting a "flight-to-safety" phenomenon (see, for example, Connolly, Stivers, and Sun 2005). We thus conjecture that,

H5: Under a pure flight-to quality phenomenon, reduced confidence in the global equity market should lead to higher integration for high credit quality sovereign bonds.

We use Baker and Wurgler (2006) sentiment index (*SENT*), which is based on first principal component of five sentiment proxies. We use a dummy, D_{Invt_Grade}, equal to one for investment grade countries i.e. when *CREDIT* exceeds 20 and 0 otherwise. To test H5, we interact D_{Invt_Grade} with *SENT*. H5 holds if the coefficient on *SENT* is insignificant and $D_{Invt\_Grade} \times SENT$ is negative and large. To the extent that there is a flight not only from risky equities but also from lower rated bonds to top rated bonds, the integration of low rated bonds could decrease with reduced market sentiment. A significant positive coefficient on *SENT* might partially capture such phenomenon.

Appendix B provides a detailed explanation of all the variables and their sources.

4.2 Descriptive statistics of the explanatory variables

Table 4 reports averages of the time series and cross-sectional variables used in the panel regressions for DM, DM excluding Eurozone, Eurozone, and EM.[27] Given the monthly frequency of most of the explanatory variables, we time aggregate the weekly integration measures for each country. The averages of political risk rating are, respectively, 0.83 and 0.68 for DM and EM confirming that DM are politically more stable and safer than EM. On average, DM countries are rated AA over the period, while EM countries are rated BBB+. The average inflation rate and inflation volatility are higher in EM compared to DM. The average rolling correlation between the inflation rates in the US and in the other countries is 0.26. The volatility of real macro conditions is also higher in EM compared to DM. The average rolling correlation between the industrial production growth rate in the US and in DM or EM is rather low but varies a lot over time and ranges between -0.56 (Singapore, Jun 1996-1999) and 0.78

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[27] Online Appendix Table A.9 shows the averages of these variables by country and their cross-correlations.

(Canada, Oct 2008-2011).[28] EM bonds are more illiquid than DM bonds. The high illiquidity for Eurozone is due to peripheral countries. Trade/GDP is of similar average magnitude in DM and EM. Except for the higher illiquidity measure in the Eurozone, the averages of the other variables are similar among Eurozone and the other DMs.

[Insert Table 4]

4.3 Panel regression results

Next, we test the null hypotheses developed above based on various specifications of the following panel regression,

$$II_{i,t} = \beta_1 (POL)_{i,t-1} + \beta_2 (CREDIT)_{i,t-1} + \beta_{3,1} (Macro\ Growth)_{i,t-1} + \beta_{3,2} (Macro\ Volatility)_{i,t-1}$$
$$+ \beta_4 (ILIQ)_{i,t-1} + \beta_5 (SENT)_{t-1} + X'_{i,t-1}\gamma + Z'_{t-1}\theta + c_i + \delta_t + \varepsilon_{i,t}, \tag{19}$$

where X_{it} is the set of local control variables, Z_t is the set of global control variables with only time series and c_i and δ_t are, respectively, country (C) and time (T) fixed effects. Obviously, the correlation patterns are subject to endogeneity and omitted variables critique. However, relying on lagged variables alleviates the former issue. Country fixed effects account for unobserved country characteristics that are constant over the sample period. We use double-clustered robust standard errors by country and time to account for serial and cross-country correlations (see Petersen, 2009). The use of the estimated integration indices as dependent variables in the panel yields consistent estimates of the coefficients. However, the reported standard errors ignore the sampling error and hence likely understate the true standard errors.

X_{it} includes country characteristic variables used as determinants of yield spreads and CDS spreads in past studies. For example, Aizenman, Hutchison, and Jinjarak (2016) find evidence for macroeconomic fundamentals, specifically trade openness, in explaining the CDS spreads of emerging

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[28] Note that the unconditional correlation between US and Canada industrial production growth rate is 0.8.

markets. We control for the level of trade openness proxied with the sum of monthly exports and imports of goods and services measured as a share of GDP (*Trade/GDP*). We also add the local stock market return as a proxy for the state of the local economic conditions. Z_t includes proxies of the state of the global economy captured with the US stock market return (*US_R_EQUITY*), the investment-grade corporate bond spread (*US_Invt_Grade*) and the high-yield corporate bond spread (*US_High_Yield*). The investment-grade spread is the spread between five-year BBB and A rated bonds. The high-yield spread is the spread between five-year BB and BBB rated bonds.

We report the estimated coefficients and their p-values from the various specifications of Equation (19) in Table 5. In all specifications, we include X_{it} variables. In all specifications except (5), we include country and time fixed effects. To examine the role of global sentiment in model (5), we remove the time fixed effects but include country fixed effects, a trend and Z_t variables. The country or time fixed effects estimates are not reported to save space. We run model (6) to evaluate the joint impact of political stability, credit quality, inflation and illiquidity on integration.

[Insert Table 5]

In Column (1) of Table 5, we report the test results of H1. The coefficient on *POL* is insignificant for SET 4. But, it is positive and highly significant for SET 1 and positive and marginally significant for SET 2 (see Online Appendix Table A.10 for results using SET 1-3) supporting H1 and suggests that greater political stability is associated with higher level of bond market integration under imperfect spanning of sovereign risk. We measure economic significance by multiplying the coefficient estimate with the standard deviation of the explanatory variable (see Panel C of Online Appendix Table A.9). A one standard deviation increase in *POL* which, for example, corresponds to moving from the political rating of Brazil to that of Italy increases the bond market integration on average by 3.5% (=σ(*POL*) of 0.1 times the coefficient on *POL* of 0.35). The effect is much larger at 9.8% (=0.1× 0.98) with less spanning (SET 1) of sovereign risk.

In Column (2), we report test results of H2. We find that credit rating is only marginally significant for SET 4, while it is significantly positively associated with the level of integration for SET 1-3 (see Online Appendix Table A.10). Also, the magnitude of the coefficient is smaller for SET 4. A one standard deviation increase in *CREDIT* which corresponds for example to a move from speculative grade BB+ to investment grade A- is associated with an increase in bond market integration on average by 6.1% for SET 4 but 11.1% increase under less spanning of credit risk (SET1).

In column (3), we test H3a and H3b. The coefficient on inflation is insignificant, while the coefficient on inflation risk is negative and marginally significant. Controlling for the correlation of country inflation with US inflation, we find that countries with higher inflation uncertainty are less integrated. A one standard deviation increase in annual inflation rate risk of 0.61% is associated with a decrease of market integration by an average 3.5%. Conditional on the level of economic globalization, we find no evidence of a significant positive association between industrial production growth rate or its volatility and integration.

In column (4), we report the tests of H4a and H4b. More illiquid countries are less integrated. A one standard deviation (of 36.83 basis points) increase in illiquidity is associated with a decrease of market integration on average by 1.8%. The interaction term between the dummy for CDS negative slope and illiquidity is significantly negative and large. The negative relationship between sovereign bond market illiquidity and integration is thus larger for distressed countries as conjectured in H4b.

In column (5) of Table 5, we find a significant negative coefficient on the interaction term $D_{Invt\_Grade} \times SENT$ suggesting that the lower the global investor sentiment, the higher the integration of the investment grade sovereign bond market. However, we find no evidence of a significant effect of *SENT* on the lower credit quality speculative grade sovereign bonds.

In all specifications, the coefficients on *Trade/GDP* that measures economic openness and on *R_EQUITY* that captures the state of the local economy are insignificant. In model (5), the US control variables are also insignificant.

In Online Appendix Table A.10, we present the panel regression coefficients from the model specifications that independently evaluates each main variable of interest (*POL*, *CREDIT*, $\sigma(\pi)$, or *ILIQ*) in the presence of the control variables for all the four spanning sets. Notwithstanding the loss of economic significance from increased spanning, the economic effect of political stability, credit quality, and inflation risk remains important. However, the statistical and economic significance of illiquidity hardly decreases in absolute terms as we move from SET 1 to SET 4. For all four sets, a one standard deviation increase in illiquidity is associated with a drop in integration of about 2%. Although the economic effect of illiquidity on integration is smaller than that of political, credit and inflation risk, it is rather constant as we gradually expand the set of substitute assets. This finding implies that there is only marginal spanning of bond market illiquidity within our substitute assets.

We report the full multivariate specification that jointly evaluates the impact of political stability, credit quality, inflation and illiquidity on integration in column 6 of Table 5. Online Appendix Table A.10 reports the regression results of this specification with the reduced sets SET 1-3. These additional tests confirm our hypotheses. The better the spanning of sovereign risk, the lower the effect of political stability and credit quality on integration. Political stability and credit rating are positively and significantly associated with the level of integration only for SET 1. The coefficients *POL* and *CREDIT* decrease monotonically as we move from SET 1 to SET 4. The coefficients on inflation and inflation risk are negative but insignificant. Illiquidity remains significant and negatively related to integration for all sets.

We use model (6) with the different levels of spanning for further analysis of the economic significance of our results. We combine the estimated coefficients with the corresponding cross-sectional distribution of the explanatory variables and assume a joint move from the 25th percentile to

the 75th percentile in the variables proxying for political stability, credit quality, inflation (level and risk) and illiquidity. Figure 3 confirms the monotonic decreasing pattern of the contribution of political stability, credit quality, and inflation risk. It also shows that the sovereign bond market integration increases by about 26%, 16%, 13%, and 10% for SET 1, SET 2, SET 3, and SET 4, respectively when a country moves from the 25th percentile to the 75th percentile as a result of higher political stability and credit quality, lower inflation and inflation risk, and lower illiquidity.

[Insert Figure 3]

4.4. Currency effects

To examine the effect of currency risk on our results, we control for the monthly change in foreign exchange (FX) rate expressed as US dollar per foreign currency (ΔFX) and volatility of changes in exchange rates ($\sigma(FX)$) in model 6. $\sigma(FX)$ is measured by cumulating daily squared changes in foreign exchange rate (see Andersen et al., 2003). We then take a 12-month moving average of the monthly FX volatility measures. Both coefficients on ΔFX and $\sigma(FX)$ are insignificant for all four sets (see Online Appendix Table A.11). Further, the inclusion of FX changes and FX volatility has no impact on our results. The association between integration and political stability, credit quality, inflation and illiquidity are not subsumed by FX changes and FX volatility.[29]

4.5. Foreign holdings and market integration

Next, we examine the link between the integration measure and the foreign holdings of local debt, which include local-currency and dollar denominated since the global holdings dataset does not differentiate by currency denomination. Data are available quarterly over 2004:Q1-2017:Q4 from Arslanalp and Tsuda (2014) and the data extension on their website. We find a positive but insignificant

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[29] Alternatively, we could repeat the analysis with currency hedged bond returns. However, of the 1148 funds, only 42 manage currency risk. Also, the currency composition of regional and global funds is not available to us. Given these limitations, we could not run this experiment.

association between foreign holdings and bond integration (see Online Appendix Table A.12).[30] Inclusion of foreign holdings does not change any of our key findings. Specifically, we still obtain a monotonic decrease in the effect of political stability and credit quality as we gradually increase spanning. Illiquidity is significant for all four sets.

4.6 Bond purchases, free float and bond market integration

During 2009-2017, there were large bond purchases by some Central Banks especially in the US and in Europe. Coeuré (2018) documents that the free float of German Bund decreased from 50% to 10% after 2015 as the ECB initiated its sovereign Bond purchase program. To assess the impact of reduced free float, we add to the panel regressions a measure of free float. Further, in view of the evidence of Christensen and Gillan (2014), Kandrac and Schlusche (2013), Kandrac (2018) and Steeley (2015) regarding the effect of large scale bond purchases on bond market liquidity, we also interact free float with illiquidity. That is we estimate the following model,

$$II_{i,t} = \alpha + \beta_1(POL)_{i,t-1} + \beta_2(CREDIT)_{i,t-1} + \beta_{3,1}\pi_{i,t-1} + \beta_{3,2}(\sigma(\pi))_{i,t-1} + \beta_4(ILIQ)_{i,t-1},$$

$$+\beta_5(Ffloat)_{i,t-1} + \beta_6(ILIQ \times Ffloat)_{i,t-1} + X'_{i,t-1}\gamma + c_i + \delta_t + \varepsilon_{i,t}, \qquad (20)$$

where *Ffloat* is the free float estimated as percentage share of outstanding central government bonds. Our methodology follows Coeuré (2018). We construct *Ffloat* by subtracting from outstanding central government bonds the bond holdings of domestic and foreign Central Banks available quarterly over 2004:Q1-2017:Q3 from Arslanalp and Tsuda (2014) and the data extension on their website. However, for Eurozone countries, over the period 2015:Q2-2017:Q3, we use bond holdings of the Eurosystem under the public sector purchase program. We then transform the quarterly to monthly free float using piecewise cubic interpolation. Online Appendix Table A.13 reports the estimated coefficients. We show that while free float has no direct effect on the integration measure, illiquidity has a much stronger effect

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[30] Few recent papers examine foreign holdings and the extent of sovereign bond home bias, see, for example, Burger et al. (2015) and Burger, Warnock and Warnock (2018).

on market integration when free float decreases. We measure the marginal effect of illiquidity on integration conditional on the level of free float from,

$$\frac{\partial II}{\partial ILIQ} = \hat{\beta}_4 + \hat{\beta}_6 \times Ffloat. \tag{21}$$

The coefficients $\hat{\beta}_4$ and $\hat{\beta}_6$ are estimated from Equation (20) – Results as reported in Online Appendix Table A.13. Online Appendix Figure A.5 reports the marginal effect across possible ranges of the free float and the 95% confidence interval for the four spanning sets. There is a statistically significant effect whenever the upper and lower bounds are both below the zero line. For all four sets, $\hat{\beta}_4 + \hat{\beta}_6 \times Ffloat$ is significant when *Ffloat* is below 0.9. These results suggest that the reduced free float resulting from the Central Banks' purchase programs significantly amplified the negative effect of illiquidity on sovereign bond market integration.

5. Conclusion

We estimate time-varying integration for 21 developed and 18 emerging sovereign bond markets based on the EL model. Our integration measure accounts for the role of substitute assets such as open-end funds, closed-end funds, and ETFs across the three asset classes: sovereign bonds, corporate bonds, and equity. The substitute assets play a major role in integrating bond markets. We find that not only sovereign bond funds but also corporate bond funds and equity funds help span the variance of the sovereign bond returns. We also examine the economic importance of four important country characteristics, namely political stability, credit quality, macroeconomic conditions, and illiquidity of the sovereign bonds that may explain the differences in the level and dynamics of integration.

We find that local risk is significantly priced and matters not only for EMs but also for DMs. Based on the zero intercept test, the EL model that allows both global and local unspanned risk to be priced performs better than fully integrated models. Local unspanned risk is priced for DMs and EMs even after accounting for exposure to world sovereign and liquidity risks.

We uncover substantial heterogeneity in the level and dynamics of integration across countries. The integration of EM sovereign bond markets is lagging behind DMs. The average integration for the EM pool is 0.48 compared to 0.65 for DMs. Although there is no significant upward trend in the integration measure for many individual countries, we do observe a small statistically significant upward trend for DMs and EMs. Among DMs, the largest statistically significant upward trend is for Eurozone core countries and among EMs the largest statistically significant upward trend is for Asian countries. We do observe reversals during the global financial crisis. The euro sovereign debt crisis has no effect on the integration of the core countries but a significant drop in integration of some periphery countries. Further, allowing for exposure to world sovereign and liquidity risks does not affect the dynamics and the average level of the integration indices.

We show that the integration of the sovereign bond markets increases on average by about 10% when a country moves from the 25th percentile to the 75th percentile as a result of higher political stability and credit quality, lower inflation and inflation risk, and lower illiquidity. A 10% increase in integration leads to a large decrease in the cost of funding of about 1% per annum on average across countries. Further, reduced confidence in the global equity market leads to higher integration for high credit quality sovereign bonds. Finally, we show that the reduced free float resulting from the recent Central Bank bonds' purchasing programs substantially amplified the negative effect of illiquidity on market integration.

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Appendix A- Details on data filters

We use several filters for the sovereign bond index returns and for the returns on funds (CEF, ETF, OEF) used as substitute assets. Specifically,

1. Filter for zeros: stale prices result in zero returns and could be due to lack of liquidity. We remove the weekly zero returns observations if we find more than three consecutive zeros in the weekly returns.

2. Filter for outliers: we remove return observations with absolute value greater than 150%.

3. Filter for time-series length: we remove funds with time series returns data shorter than 104 weeks.

4. Filter for holes in time series: we remove funds with less than 95% data available for the period between their start date and end date.

5. We do not include funds without geographical allocation.

6. We include both active and inactive funds to avoid survivorship bias.

Appendix B - Definition of the variables used in the panel regressions (Section 4 of the paper)

Variable	Name	Description
Sovereign variables		
Political Stability	*POL*	Political risk ratings based on the sum of 12 weighted variables covering both political and social attributes. The index has 100 points. It is scaled to range from 0 (high risk) to 1 (low risk). Frequency: Monthly. Source: International Country Risk Guide (ICRG).
Credit Quality	*CREDIT*	S&P sovereign ratings of long term foreign bond transformed linearly into a numerical format ranging from 1 (Default) to 21 (AAA). Frequency: Monthly. Source: Bloomberg and Standard&Poor's.
CDS Dummy	$D_{CDS10-CDS1}$	A dummy equal to 1 if slope of sovereign CDS term structure computed from 10-year CDS spread minus 1-year CDS spread is negative
Investment Grade Dummy	D_{Invt_Grade}	A dummy that takes the value of 1 for investment grade countries with RATING ≥ 20
Inflation		
Inflation	π	Inflation rate measured as difference in logs of the Consumer Price Index (CPI) at time t and t-12, t in months. Frequency: Monthly. Source: IFS
Inflation Volatility	$\sigma(\pi)$	Inflation volatility measured by fitting a GARCH(1,1) to the shocks to monthly inflation rates. Inflation rate shocks are estimate from the ARIMA(p,q). We use the BIC to select the best ARMA(p,q) specification.
Inflation Correlation with US	$\rho(\pi_j, \pi_{US})$	Three-year rolling correlation between a country's inflation rate and that of the US.
Real Macroeconomic Variables		
Industrial Production growth rate	ΔIP	Growth rate of industrical production (IP). Frequency: Monthly. Source: OECD and IFS through Datastream.
IP volatility	$\sigma(IP)$	IP volatility measured by fitting a GARCH(1,1) to the shocks to monthly IP growth rates. IP shocks are estimated from an ARIMA(p,q).
IP growth Correlation with US	$\rho(IP_j, IP_{US})$	Three-year rolling correlation between a country's industrial production growth rate and that of the US.
Iliquidity		
Bid-ask spread	*ILIQ*	Equally-weighted quoted bid-ask spread expressed relative to mid price in basis points. The measure is constructed from individual sovereign local currency-denominated bonds with at least one year to maturity and no special features (that is bonds with options or floating rates or inflation-indexed bonds are eliminated). The measure is winsorized at the first and 99th percentile to limit the influence of outliers. Frequency: Monthly. Source: Bloomberg
Other Local variables		
Trade to GDP	*Trade/GDP*	Sum of monthly exports and imports of goods and services measured as a share of annual GDP. Frequency: Monthly. Source: International Financial Statistics (IFS) of IMF and WDI.
Local stock market return	*R_EQUITY*	Local stock market total return denominated in local currency. Frequency: Monthly. Source: Datastream Indexes
Change in FX	ΔFX	Percentage changes in the exchange rate, expressed as US dollar units per local currency. Frequency: Monthly. Source: Datastream
FX volatility	$\sigma(FX)$	Realized volatility measured by cumulating daily squared changes in foreign exchange rate. We then take a 12-month moving average of the monthly FX volatility measures. Frequency: Monthly. Source: Datastream and authors calculation.

Appendix B (continued)

Variable	*Name*	Description
Free float	*Ffloat*	Free float is the fraction of outstanding central government bonds not held by domestic and foreign central banks. Frequency: Monthly from Quarterly. Source: Quarterly holdings data from Arslanalp and Tsuda (2014) and the data extension on their website over 2004:Q1-2017:Q3, and bond holdings of the Eurosystem under the PSPP over 2015Q2-2017Q3 from ECB.
Foreign Holdings	*FHoldings*	Foreign holdings is the fraction of outstanding central government bonds held by foreign investors excluding foreign official sector. Frequency: Monthly from Quarterly. Source: Quarterly holdings data from Arslanalp and Tsuda (2014) and the data extension on their website over 2004:Q1-2017:Q3

Global investor sentiment

SENT	*SENT*	Sentiment index in Baker and Wurgler (2006); updated version of Eq. (2) in that paper; based on first principal component of FIVE (standardized) sentiment proxies.

Global variables

US market return	*US_R_EQUITY*	Total return US equity index. Frequency: Monthly. Source: Datastream Indexes
Corporate yield spread-Investment grade	*US_Invest_Grade*	Change in basis point yield spread between BBB and A industrial bond indexes. The indexes represent the average yields of non-callable A and BBB rated bonds with maturities about five years. Frequency: Monthly. Source: Bloomberg (fair market curves).
Corporate yield spread- High yield	*US_High_Yield*	Change in basis point yield spread between BB and BBB industrial bond indexes. The indexes represent the average yields of non-callable BBB- and BB- rated bonds with maturities about five years. Frequency: Monthly. Source: Bloomberg (fair market curves).
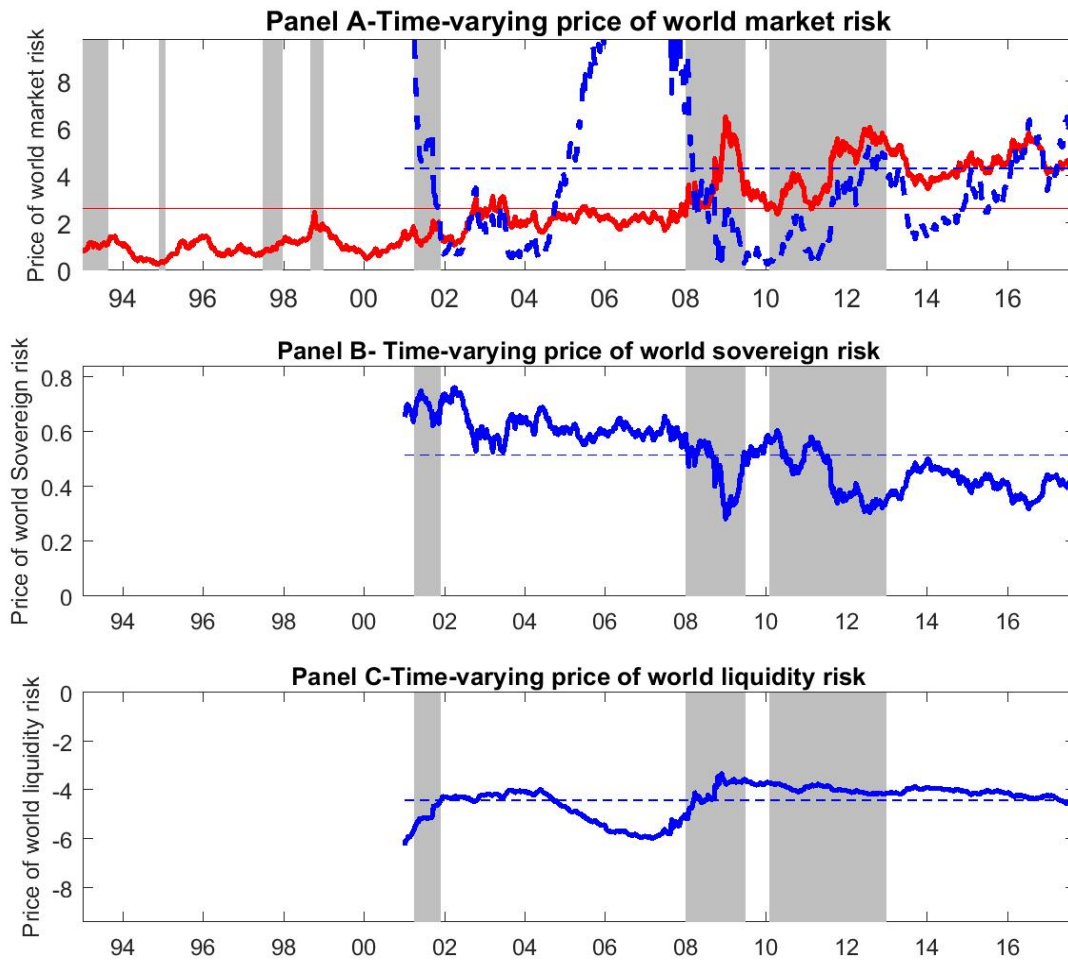
Figure 1

Time-varying prices of world covariance risk factors

The figure plots the time-varying prices of world covariance risk factors from first-stage estimation. The horizontal lines report the time-series averages of the prices of world risk factors. Panel A shows the prices of world market risk estimated from WCAPM (solid line) and SOV-LIQ WCAPM (dashed line). Panels B and C show the prices for the world sovereign risk and world illiquidity risk estimated from SOV-LIQ WCAPM. The world market portfolio is the value-weighted world equity market index and world government bond index. World sovereign risk factor is the difference of the logarithm of the equally-weighted 5-year sovereign CDS spread aggregated across all countries. World bond illiquidity is the equally weighted sum of the quoted bid-ask spread of individual local-currency sovereign bonds from the 39 countries of our sample. The first equation in the system of equations (9) shows the WCPAM for the world market portfolio. The first three equations of the system of equations (10) show the SOV-LIQ WCAPM for the world market portfolio, the aggregate 5-year sovereign CDS, and the aggregate bid-ask spread. Gray areas indicate March to November 2001 NBER recession, the European exchange rate mechanism crisis (September 1992 to August 1993), the Tequila crisis (December 1994 to January 1995), the East Asia crisis (June to December 1997), the Russian default and Long-Term Capital Management crisis (August to December 1998), the October 2007-May 2009 global financial crisis, and the Euro-sovereign debt crisis (January 2010 to December 2012).
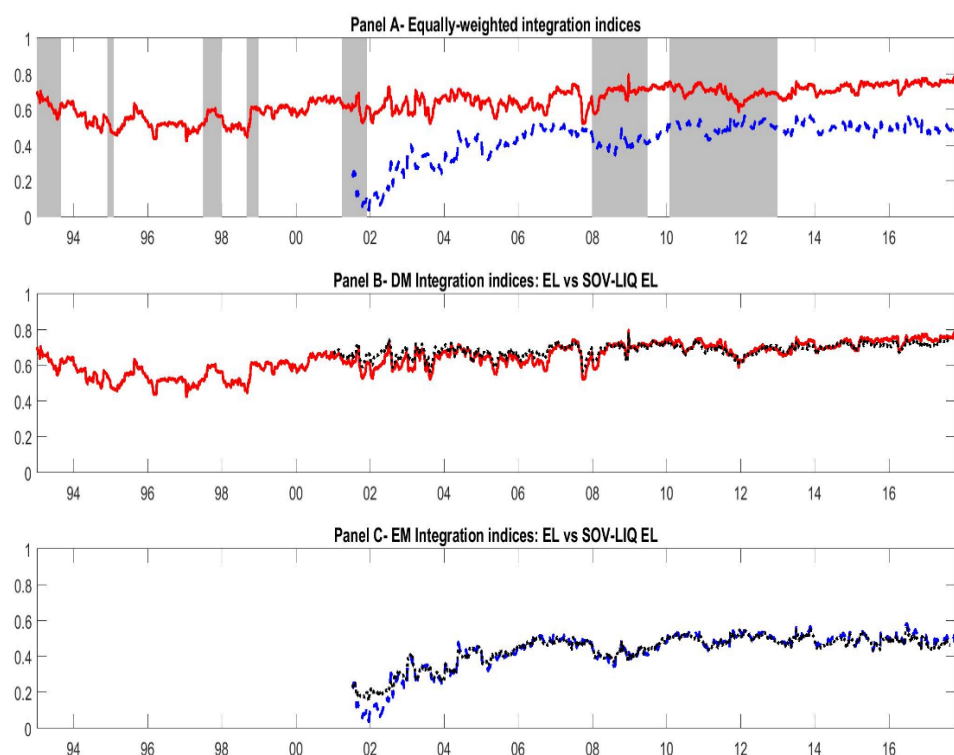
Figure 2

Cross-section average integration measures

Panel A plots the equally-weighted averages across 21 developed markets (solid line) and 18 emerging markets (dashed line) at each point in time of the integration measures of the sovereign bond indices estimated from the EL model. Panel B plots the equally-weighted averages across 21 developed markets of the integration measures estimated from the EL model (solid line) and the augmented SOV-LIQ EL model (dotted line). Panel C plots the equally-weighted averages across 18 emerging markets of the integration measures estimated from the EL model (dashed line) and the augmented SOV-LIQ EL model (dotted line). Gray areas indicate March to November 2001 NBER recession, the European exchange rate mechanism crisis (September 1992 to August 1993), the Tequila crisis (December 1994 to January 1995), the East Asia crisis (June to December 1997), the Russian default and Long-Term Capital Management crisis (August to December 1998), the October 2007-May 2009 global financial crisis, and the Euro-sovereign debt crisis (January 2010 to December 2012).
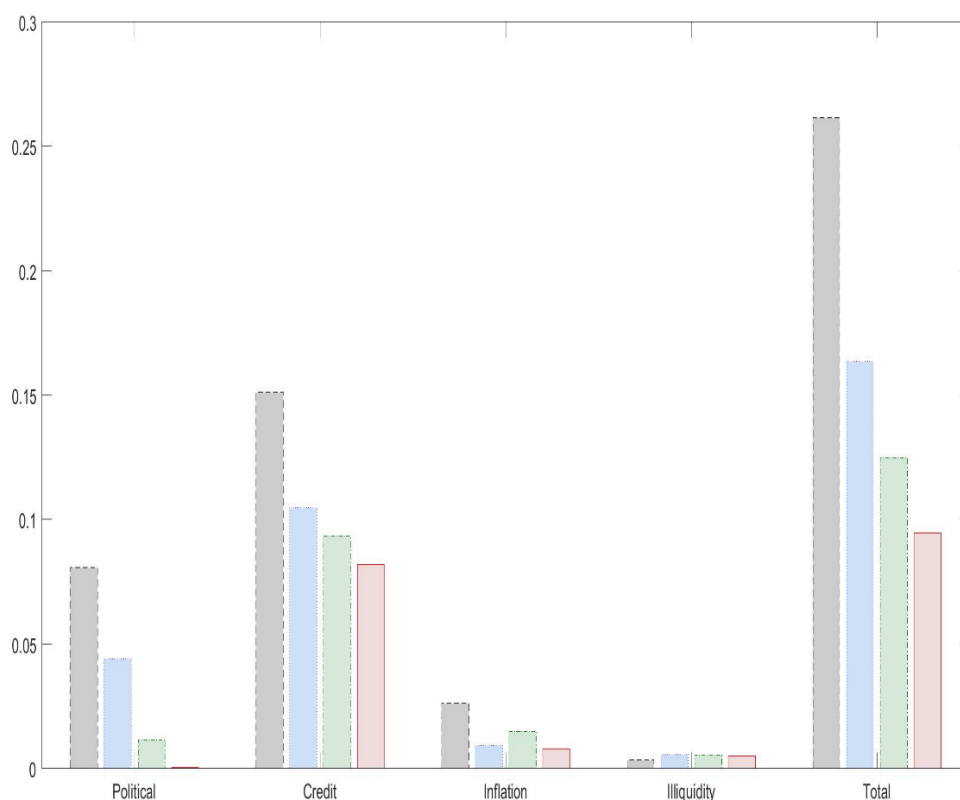
Figure 3
Economic Significance

Economic impact is estimated as countries move from 25th percentile to 75th percentile. We plot the economic impact on sovereign bond market integration of political ratings (*POL*), credit ratings (*CREDIT*), inflation, and Illiquidity (*ILIQ*). Inflation is the sum of inflation level (π). and risk ($\sigma(\pi)$). Definition of the variables and data source are in Appendix B. We plot for the integration measures estimated from the EL model using the four spanning sets SET 1 (dashed), SET 2 (dotted), SET 3 (dash-dotted) and SET 4 (solid). SET 4 includes the world bond index, world equity index, sovereign bond funds, equity funds, corporate bond funds, 5-year CDS, and USD denominated JP Morgan EMBI bond indices for EMs. SET 3 is SET 4 excluding 5-year CDS. SET 2 is SET 3 excluding the equity and corporate bond funds. SET 1 is SET 2 excluding sovereign bond funds.

Table 1
Distribution of the funds by funds type, asset class, and geographical allocation

	Closed-End Funds				ETFs				Open-End Funds				All Funds			
	SOV	*CORP*	*EQUI*	*TOTAL*	*SOV*	*CORP*	*EQUI*	*TOTAL*	*SOV*	*CORP*	*EQUI*	*TOTAL*	*SOV*	*CORP*	*EQUI*	*TOTAL*
Australia	0	0	2	**2**	1	1	7	**9**	0	0	0	**0**	1	1	9	**11**
Austria	0	0	0	**0**	0	0	3	**3**	0	0	0	**0**	0	0	3	**3**
Belgium	0	0	0	**0**	0	0	1	**1**	0	0	0	**0**	0	0	1	**1**
Canada	0	0	3	**3**	1	0	4	**5**	0	0	3	**3**	1	0	10	**11**
Denmark	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**
Finland	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**
France	0	0	1	**1**	2	0	10	**12**	0	0	0	**0**	2	0	11	**13**
Germany	0	0	1	**1**	17	1	33	**51**	0	0	0	**0**	17	1	34	**52**
Greece	0	0	0	**0**	0	0	1	**1**	0	0	0	**0**	0	0	1	**1**
Ireland	0	0	1	**1**	0	0	1	**1**	0	0	0	**0**	0	0	2	**2**
Italy	0	0	0	**0**	8	0	5	**13**	0	0	0	**0**	8	0	5	**13**
Japan	0	0	7	**7**	1	0	43	**44**	0	0	10	**10**	1	0	60	**61**
Netherlands	0	0	0	**0**	0	0	5	**5**	0	0	3	**3**	0	0	8	**8**
New Zealand	0	0	0	**0**	0	0	1	**1**	0	0	0	**0**	0	0	1	**1**
Norway	0	0	0	**0**	0	0	1	**1**	0	0	0	**0**	0	0	1	**1**
Portugal	0	0	1	**1**	0	0	2	**2**	0	0	0	**0**	0	0	3	**3**
Singapore	0	0	2	**2**	0	0	4	**4**	0	0	0	**0**	0	0	6	**6**
Spain	0	0	1	**1**	2	0	3	**5**	0	0	0	**0**	2	0	4	**6**
Sweden	0	0	0	**0**	0	0	1	**1**	0	0	0	**0**	0	0	1	**1**
Switzerland	0	0	1	**1**	0	0	6	**6**	0	0	0	**0**	0	0	7	**7**
UK	0	0	29	**29**	9	5	30	**44**	0	0	2	**2**	9	5	61	**75**
Emerging Markets																
Brazil	0	0	3	**3**	0	0	9	**9**	0	0	0	**0**	0	0	12	**12**
Chile	2	0	3	**5**	0	0	3	**3**	0	0	0	**0**	2	0	6	**8**
China	0	0	6	**6**	3	1	58	**62**	0	0	9	**9**	3	1	73	**77**
Colombia	0	0	0	**0**	0	0	2	**2**	0	0	0	**0**	0	0	2	**2**
Czech Republic	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**
Hungary	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**	0	0	0	**0**
India	0	0	7	**7**	0	0	13	**13**	0	0	3	**3**	0	0	23	**23**
Indonesia	0	0	1	**1**	0	0	5	**5**	0	0	0	**0**	0	0	6	**6**
Malaysia	0	0	1	**1**	0	0	3	**3**	0	0	0	**0**	0	0	4	**4**
Mexico	0	0	3	**3**	0	0	3	**3**	0	0	0	**0**	0	0	6	**6**

Table 1 (continued)

	Closed-End Funds				ETFs				Open-End Funds				All Funds			
	SOV	CORP	EQUI	TOTAL	SOV	CORP	EQUI	TOTAL	SOV	CORP	EQUI	TOTAL	SOV	CORP	EQUI	TOTAL
Peru	0	0	0	0	0	0	1	1	0	0	0	0	0	0	1	1
Poland	0	0	1	1	0	0	3	3	0	0	0	0	0	0	4	4
Russia	0	0	2	2	0	1	9	10	0	0	1	1	0	1	12	13
South Africa	0	0	1	1	0	0	5	5	0	0	0	0	0	0	6	6
South Korea	0	0	1	1	0	0	7	7	0	0	1	1	0	0	9	9
Taiwan	0	0	2	2	0	0	6	6	0	0	0	0	0	0	8	8
Thailand	0	0	1	1	0	0	3	3	0	0	0	0	0	0	4	4
Turkey	0	0	0	0	0	0	5	5	0	0	0	0	0	0	5	5
Regional and Global Funds																
African Region	0	0	0	0	0	0	3	3	0	0	2	2	0	0	5	5
Asian Pacific Region	0	0	4	4	1	0	16	17	0	3	12	15	1	3	32	36
Asian Pacific ex Japan	0	0	10	10	1	1	25	27	0	0	11	11	1	1	46	48
BRIC	0	0	0	0	0	0	4	4	0	0	0	0	0	0	4	4
Eastern Europe Region	0	0	1	1	0	0	5	5	0	0	1	1	0	0	7	7
European Region	0	1	3	4	8	13	179	200	2	0	33	35	10	14	215	239
European Reg. ex UK	0	0	8	8	0	0	8	8	0	0	1	1	0	0	17	17
European Union	1	0	1	2	1	1	2	4	0	0	0	0	2	1	3	6
Eurozone	0	0	0	0	58	15	65	138	1	1	0	2	59	16	65	140
Greater China	0	0	2	2	0	0	2	2	0	0	7	7	0	0	11	11
Indian Sub-Continent	0	0	0	0	0	0	0	0	0	0	1	1	0	0	1	1
Latin American Region	0	0	2	2	0	0	6	6	0	0	5	5	0	0	13	13
Middle East Region	0	0	0	0	0	0	1	1	0	0	0	0	0	0	1	1
Nordic Region	0	0	0	0	0	0	3	3	0	0	1	1	0	0	4	4
North American Region	0	1	0	1	0	0	5	5	0	0	7	7	0	1	12	13
OECD Countries	0	0	0	0	0	0	0	0	0	1	0	1	0	1	0	1
South East Asia Region	0	0	0	0	0	0	0	0	0	0	1	1	0	0	1	1
Tiger Region	0	0	0	0	0	0	0	0	0	0	1	1	0	0	1	1
Western Europe	0	0	0	0	0	0	0	0	1	0	0	1	1	0	0	1
Emerging Market	6	0	6	12	20	4	53	77	52	15	134	201	78	19	193	290
Global/International	2	19	42	63	14	22	260	296	41	63	677	781	57	104	979	1140

For each developed or emerging country and for region and global funds, we report the number of sovereign (SOV), corporate (CORP) and equity (EQUI) funds by fund type (Closed-end, ETF, Open-end) as well as the total number of funds by asset class (in bold).

Table 2

Asset Pricing Models Estimation and Validation

Panel A- Prices of world risk factors from first-satge estimation

	WCAPM. Sample: 06-Jan-1993 to 30-Aug-2017			SOV-LIQ WCAPM Sample: 10-Jan-2001 to 30-Aug-2017								
	Price of world market risk			Price of world market risk			Price of world sovereign risk			Price of world liquidity risk		
	H_0: zero	H_0: constant	Mean	H_0: zero	H_0: constant	Mean	H_0: zero	H_0: constant	Mean	H_0: zero	H_0: constant	Mean
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)	(ix)	(x)	(xi)	(xii)
World market	**12.20**	0.57	2.62	**24.40**	4.18	4.33	4.49	0.35	0.52	0.23	0.01	-4.40
	(0.01)	(0.75)		(0.00)	(0.12)		(0.21)	(0.84)		(0.97)	(1.00)	

Panel B- Summary of the country test specifications from second-satge estimation

	WCAPM	SOV-LIQ WCAPM	SOV-LIQ EL	Mixed	Genralized	EL	EL		Generalized			
	Intercepts						Price of local unspanned risk		Price of local cov risk		Price of local var risk	
	average $|\alpha|$ in % per week						average price	price range	average price	price range	average price	price range
DM	0.07	0.12	0.07	0.05	0.06	0.05	5.78	[0.73, 11.90]	5.96	[0.61, 19.34]	4.89	[0.85, 16.91]
EM	0.15	0.19	0.09	0.10	0.18	0.11	6.89	[2.22, 26.46]	12.79	[2.27, 35.81]	6.94	[1.68 23.78]
	# (pvalue (H_0: α=0) $\leq$5%)						# (pvalue (H_0: zero) $\leq$5%)	# (pvalue (H_0: constant) $\leq$5%)	# (pvalue (H_0: zero) $\leq$5%)	# (pvalue (H_0: constant) $\leq$5%)	# (pvalue (H_0: zero) $\leq$5%)	# (pvalue (H_0: constant) $\leq$5%)
DM	13	17	8	1	7	4	13	6	16	13	13	8
EM	15	14	1	6	7	4	14	5	16	8	15	10
	Cross-sectional correlation between average excess returns and predicted excess returns											
DM	0.57	0.56	0.29	0.09	0.07	0.30						
EM	0.22	0.22	0.42	0.18	0.20	0.31						

This table presents evidence on the global and local risk pricing for sovereign bonds. The prices of world risk factors are pre-estimated. Panel A reports first-stage estimation results of the world market model from WCAPM and SOV-LIQ WCAPM models. For WCAPM, we report the test statistic for a significant price of world market risk, the test statistic for a time-varying price of risk, and the time-series averages of the prices of risk in columns (i)-(iii). For SOV-LIQ WCAPM, we report similar statistics for the prices of world market risk, world sovereign risk and world liquidity risk in columns (iv)-(xii). p-values in parentheses appear below their corresponding statistics. Numbers in bold represent significance at 5% level or lower. For each country, we estimate in a second-satge six models: WCAPM, SOV-LIQ WCAPM, SOV-LIQ EL, Mixed, Generalized, and EL. EL is our main model estimated from system of equations (9). We consider alternative models. WCAPM and SOV-LIQ WCAPM are fully integrated models. SOV-LIQ EL augmente EL with world sovereign and world liquidity risk factors (see system of Equations (10)). The mixed model is a restricted version of EL where we impose zero value on the integration index. The generalized model allows different prices of risk for the total variance of the bond index and the covariance between the bond index and its DP (see Equation (16)). In all these estimations, we take the pre-estimated prices of global risk factors from first-stage as given. Panel B contains evidence on the EL and the alternative models. We show across DMs and EMs, the cross-section average absolute intercepts and number of countries with significant interecept for the six models. We also show the number of countries with significant and time-varying local unspanned prices estimated from EL, their time-series and cross-section averages as well as the range of the time series averages prices across countries. We also report the specification tests, averages and ranges for the price of covariance between bond index return and its DP as well as the price of total variance estimated from the generalized model. Finally, we report a comparison across the six models based on the cross-sectional correlation between realized average excess returns and predicted expected excess returns estimated from the various models. We count the number of countries where we reject the null hypothesis at 5% level. Estimations from weekly (wednesday-to-wednesday) returns. For SOV-LIQ WCAPM and SOV-LIQ EL models the start date is January 10th, 2001 or later depending on data availability. All time series end on August 30th, 2017. See Online Appendix Table A.3 for the detailed estimation results by country.

Table 3

Integration index estimated from the EL model

	start date	Mean	β_{trend} (% per annum)	D_{GFC}	D_{ESC}	COF (% per annum)		start date	Mean	β_{trend} (% per annum)	D_{GFC}	D_{ESC}	COF (% per annum)
Developed markets							**Emerging Markets**						
three most integrated markets							**three most integrated markets**						
Germany	6-Jan-93	0.85	**0.43**	**-0.06**	-0.01	0.64	Poland	8-Sep-04	0.69	0.43	**-0.04**	**0.04**	2.40
			(2.91)	(-4.52)	(-1.11)					(1.35)	(-2.32)	(4.04)	
Belgium	12-Oct-94	0.84	0.32	**-0.03**	**-0.06**	0.66	Turkey	13-Apr-05	0.69	-0.06	**-0.05**	**0.04**	1.25
			(1.52)	(-2.17)	(-4.43)					(-0.21)	(-2.53)	(2.28)	
France	3-Apr-96	0.80	1.22	0.03	-0.03	1.31	Hungary	2-Nov-05	0.67	0.39	**-0.04**	-0.01	4.32
			(0.68)	(1.40)	(-1.66)					(0.92)	(-2.40)	(-0.25)	
three least integrated markets							**three least integrated markets**						
Sweden	5-Jan-94	0.53	0.76	0.02	**0.07**	0.94	Taiwan	11-Jul-01	0.26	**0.70**	**-0.04**	0.00	0.30
			(1.54)	(1.29)	(8.52)					(2.43)	(-2.14)	(-0.10)	
Switzerland	5-Jan-94	0.52	**-0.41**	**0.07**	-0.01	0.11	Chile	12-Nov-03	0.25	0.23	-0.03	0.03	0.53
			(-2.17)	(2.70)	(-0.22)					(0.36)	(-1.23)	(1.43)	
New Zealand	5-Jan-94	0.41	**1.03**	**-0.03**	**0.04**	0.56	China	12-Jan-05	0.08	**0.59**	0.01	-0.01	0.28
			(4.49)	(-1.86)	(3.66)					(1.76)	(0.48)	(-1.52)	
Averages and trend tests across regions													
DM	21	0.65	**1.03**	0.01	0.00	0.86	EM	18	0.48	**0.59**	**-0.04**	**0.03**	1.02
			(4.90)	(0.35)	(-0.08)					(6.34)	(-3.59)	(3.45)	
DM ex Eurozone	10	0.58	**0.84**	0.00	**0.03**	0.62	Emerging Europe	4	0.64	**0.29**	**-0.09**	0.01	2.01
			(3.47)	(0.08)	(2.66)					(5.71)	(-3.34)	(0.73)	
Eurozone	11	0.71	**1.21**	0.01	-0.03	1.15	Emerging Asia	8	0.41	**0.73**	**-0.02**	**0.04**	0.68
			(3.61)	(0.38)	(-0.97)					(6.89)	(-1.29)	(2.95)	
Eurozone Core	6	0.74	**1.68**	0.00	0.04	0.67	Latin America	5	0.44	**0.30**	**-0.05**	0.02	0.88
			(3.38)	(0.18)	(1.26)					(4.06)	(-2.64)	(1.41)	
Eurozone Periphery	5	0.67	0.66	0.02	**-0.12**	1.74	DM & EM	39	0.57	**0.96**	-0.01	0.01	0.94
			(1.57)	(0.36)	(-3.17)					(5.19)	(-1.09)	(1.16)	

The table reports for the estimated integration measures, the start date, mean, trend coefficient in % per annum (β_{trend}) , the global financial crisis (2007/10-2009/05) dummy coefficient (D_{GFC}) and the euro sovereign debt crisis (2010/01-2012/12) dummy coefficient (D_{ESC}) for the three most integrated and the three least integrated among developed markets and among emerging markets. The t-stat of the trend coefficient is based on Bunzel and Vogelsang (2005) using Daniell kernel variance estimator. The 5% critical value (two-sided) for the Bunzel and Vogelang (2005) t-stat is 1.71. The last columns show the decrease in cost of funding in % per annum (COF) for a 10% increase in market integration (see Equation 18). The integration measures are estimated from the EL model (see Equations (3) and (9)). We also report mean, trend, crises dummies coefficients and their t-stats by region based on panel regressions with country fixed effects and a trend. The number of countries in the different pools is reported in the first column. T-stats appear below their corresponding coefficients in parentheses. The standard errors for the dummy crises tests of the country regressions are heteroskedasticity and autocorrelation consistent obtained from Newey-West (1987) correction with six lags. The standard errors for the trend and dummy crises coefficients of the panel regressions are clustered by country and time. Numbers in bold represent significance at 5% level or lower. Estimations from weekly (wednesday-to-wednesday) returns. All time series end on August 30th, 2017. See Online Appendix Table A.8 for the detailed estimation results by country.

Table 4

Panel variables averages

Panel A-Variables with time series and cross-section

	DM &EM	DM	DM ex. Eurozone	Eurozone	EM
II	0.57	0.65	0.58	0.71	0.48
POL	0.76	0.83	0.85	0.82	0.68
CREDIT	17.41	19.73	20.79	18.76	14.70
Dummy D $_{CDS10-CDS1}$	0.02	0.03	0.00	0.06	0.01
π *(% p.a)*	2.98	1.84	1.58	2.09	4.31
$\sigma(\pi)$ *(% p.a)*	1.46	1.17	1.27	1.08	1.79
$\rho(\pi_j, \pi_{US})$	0.26	0.31	0.29	0.33	0.20
ΔIP *(% p.a)*	2.08	1.62	1.65	1.60	2.62
$\sigma(IP)$ *(% p.a)*	8.20	7.43	7.16	7.67	9.10
$\rho(IP_j, IP_{US})$	0.08	0.08	0.09	0.07	0.08
ILIQ (bps)	30.38	26.96	14.05	38.69	34.60
Trade/GDP	0.68	0.68	0.68	0.67	0.69
R_EQUITY	9.60	8.04	8.91	7.25	11.41
ΔFX *(% p.a)*	-0.39	0.00	0.32	-0.29	-0.85
$\sigma(FX)$ *(% p.a)*	8.55	8.76	8.71	8.81	8.30

Panel B-Variables with time series only

SENT	0.13				
US_R_EQUITY (% p.a.)	8.63				
US_Invt_Grade (BBB-A) (bps)	0.61				
US_High_Yield (BB-BBB) (bps)	1.52				

The table lists average values of the regressors and regressands for the group of DM & EM (39), DM (21), DM ex. Eurozone (10), Eurozone (11), and EM (18) economies. Panel A reports on the variables with time series and cross-section and Panel B reports on the variables with time series only. The regressands are the integration measures estimated from the EL model (see Equations (3) and (9)). Appendix B details the set of regressors and their sources. The values by country are reported on Online Appendix Table A.9.

Table 5

Sovereign bond market integration and country characteristics

	(1) Political Stability	(2) Credit quality	(3) Nominal and Real Macro		(4) Illiquidity		(5) Sentiment	(6) Full Model
Hypothesis	H1	H2	H3a	H3b	H4a	H4b	H5	
POL	0.35							0.00
	(0.28)							(1.00)
CREDIT		**0.02**						0.01
		(0.10)						(0.13)
π $(\times 10^2)$			0.31					-0.12
			(0.57)					(0.73)
$\sigma(\pi)$ $(\times 10^2)$			**-5.64**					-1.19
			(0.06)					(0.69)
$\rho(\pi_j, \pi_{US})$			-0.01					
			(0.80)					
ΔIP $(\times 10^2)$				0.04				
				(0.50)				
$\sigma(IP)$ $(\times 10^2)$				0.08				
				(0.60)				
$\rho(IP_j, IP_{US})$				-0.02				
				(0.55)				
ILIQ $(\times 10^4)$					**-4.78**	-2.20		**-3.30**
					(0.05)	(0.17)		(0.04)
$D_{CDS10-CDS1}$						**-0.14**		
						(0.00)		
$D_{CDS10-CDS1} \times ILIQ$						**-2.57**		
						(0.01)		
SENT							-0.01	
							(0.73)	
D_{Invt_Grade}							-0.01	
							(0.86)	
$D_{Invt\_Grade} \times SENT$							**-0.06**	
							(0.01)	

Table 5 (continued)

	(1) Political Stability	(2) Credit quality	(3) Nominal and Real Macro		(4) Illiquidity		(5) Sentiment	(6) Full Model
Hypothesis	H1	H2	H3a	H3b	H4a	H4b	H5	
Trade/GDP	0.03	0.03	0.02	0.03	0.04	-0.07	0.00	0.04
	(0.66)	(0.51)	(0.68)	(0.64)	(0.55)	(0.12)	(0.93)	(0.47)
R_EQUITY	0.04	0.05	0.04	0.04	0.03	0.02	0.01	0.04
	(0.23)	(0.11)	(0.24)	(0.27)	(0.36)	(0.75)	(0.84)	(0.18)
US_R_EQUITY							0.00	
							(0.99)	
US_Invt_Grade							-0.02	
							(0.11)	
US_High_Yield							0.00	
							(0.61)	
Trend (×100)							**0.07**	
							(0.01)	
FE	T, C	T, C	T, C	T, C	T, C	T, C	C	T, C
# obser.	7731	7647	7731	7494	7494	4972	5253	7411
Adjusted R^2	66.7%	68.4%	66.6%	67.1%	67.1%	75.4%	72.2%	68.4%

The table reports the estimated coefficients from panel regressions of the sovereign bond integration measures estimated from the EL model on proxies for political stability (*POL*), credit quality (*CREDIT*), nominal macro (inflation level π, volatility $\sigma(\pi)$ and rolling corrleation between country *j* inflation with US inflation ($\rho(\pi_j,\pi_{US})$) and real macro (industrial production growth rate ΔIP, volatility $\sigma(IP)$ and rolling corrleation between country *j* industrial production growth rate with US ($\rho(IP_j,IP_{US})$)), illiquidity (*ILIQ*), and global sentiment (*SENT*). The estimated models are based on the general equation below,

$$II_{i,t} = \beta_1(POL)_{i,t-1} + \beta_2(CREDIT)_{i,t-1} + \beta_{3,1}(Macro\ Growth)_{i,t-1} + \beta_{3,2}(Macro\ Volatility)_{i,t-1} + \beta_4(ILIQ)_{i,t-1} + \beta_5(SENT)_{t-1} + X'_{i,t-1}\gamma + Z'_{t-1}\theta + c_i + \delta_t + \varepsilon_{i,t}$$

where c_i are country fixed effects (C) and δ_t are time fixed effects (T). In model (5) we remove the time fixed effects and add a time trend, *Trend*. In all specifications, we add a set of country variables controls (*X*), which include Trade to GDP (*Trade/GDP*) and the local stock market return (*R_EQUITY*). $D_{CDS10\text{-}CDS1}$ is a dummy that takes the value of 1 if the sovereign CDS slope measured by the difference between 10-year CDS and 1-year CDS is negative, D_{Invt_Grade} is a dummy that takes the value of 1 for investment grade countries with *RATING*≥20. In model (5), we also control for global market conditions, Z, proxied with the US equity return (*US_R_EQUITY*), the spread between 5-year BBB and A rated US coporate bonds (*US_Invt_Grade*), the spread between 5-year BB and BBB rated US coporate bonds (*US_High_Yield*). We run unbalanced regressions as not all the explanatory variables are available for all countries. All explanatory variables are lagged. p-values appear below their corresponding coefficients in parentheses and are obtained from standard errors that are clustered by country and time. The sample period is monthly from 01/1993 to 08/2017. Definition of variables and data source are in Appendix B. Numbers in bold represent significance at 10% level or lower.