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Kei-Ichiro Inaba*

Abstract

This article analyzes the developments and determinants of the country-specific dependence of sovereign bond returns on global factors for 41 advanced and emerging countries over the last decade. The dependence was cyclical and substantial: the average for the sample countries and period is around 56 percent. This is consistent with a global financial cycle hypothesis stressing the dominant role played by global factors in the synchronization of asset price changes across countries. The dependence was smaller for emerging countries than for advanced ones. Differences in the dependence among countries and over time were attributable to country-fixed effects and time-varying factors. These factors include the size and openness of domestic bond market, the variability of foreign exchange rates, the impact of macro-economic policies, and the indebtedness of the national finance. One policy implication of the hypothesis is examined, namely, the dilemma between international capital mobility and monetary policy effect.

Keywords: Sovereign bonds; Market integration; Global financial cycle; Monetary policy; Capital control.

JEL classification: F3, G1, O1

* Director and Senior Economist, Institute for Monetary and Economic Studies, Bank of Japan (E-mail: keiichirou.inaba@boj.or.jp)

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1. Introduction

The synchronization of asset price changes across countries is "a key topic in finance studies, as it has important implications for asset allocation, risk management, and international diversification" (Chuluun, 2017, p. 53). The more the market for a specific asset in a country is integrated with foreign markets, the more likely its price will respond to global factors (GFs). Such a passive response of domestic financial asset prices is referred to hereafter as the dependence on global factors (DGF).

I investigate the developments and determinants of country-specific DGFs for ten-year sovereign bonds for 41 advanced and emerging countries over the period 2007–16. I measure their DGFs using a method proposed by Pukthuanthong and Roll (P&R, 2009). They define a DGF of the national market for a specific financial asset as the percentage of the asset return's total variation accounted for by a number of latent GFs and regard this percentage as "a sensible intuitive quantitative measure of financial market integration" (P&R, 2009, p. 214). Identifying the GFs is beyond the scope of their method. They used it for countries' stocks while I use that for countries' sovereign bonds.

This investigation can add on the literature simply because previous finance studies have focussed mainly upon the DGFs of countries' stocks.[1] Exceptional studies on the DGFs of countries' sovereign bonds include Driessen *et al*. (2003), Barr and Priestley (2004), Diebold *et al.* (2008), Kumar and Okimoto (2011), Jotikasthira *et al.* (2015), and Byrne *et al.* (2019). However, their range of sample countries has been limited to some or all of the Group of Seven (G7) countries, despite the fact that, over the last decade, sovereign bonds denominated in local currencies have been rapidly increasing in emerging countries, and a major proportion of these bonds are held by global investors (Agur *et al.*, 2018). Although the previous studies present a broad consensus that, while differing in size across countries, G7 sovereign-bond DGFs have been substantial, the driving forces behind these DGFs still merit investigation.

By analyzing the relevance of monetary authorities' policy tools to these driving forces, I address a potential dilemma between international capital mobility and monetary policy effect for countries' sovereign bond price changes. The costs of high DGFs for specific financial assets became apparent with the reduction of domestic monetary policy effect due to monetary policy spillovers from the U.S. after the global financial crisis in 2008 (Passari and Rey, 2015). Now, "(where experts) need to make more progress is the links among monetary policy, international capital flows, and domestic financial fragility" (Rajan, 2018, p. 22, terms in parentheses added by the author). Rey (2013; 2016), Passari and Rey (2015), and Coeurdacier *et al.* (2019) claim that a global financial cycle sets the tone of financial conditions globally beyond domestic circumstances. They regard two factors as GFs driving that cycle – global investors' risk preference and global uncertainty, and argue that these two factors are affected by U.S. monetary policy. This global financial cycle hypothesis puts the Mundellian trilemma into

[1] Using different methods to measure national stock DGFs, experts have found mixed evidence for how the DGFs have changed in advanced countries, including a seminal study by Bekaert *et al.* (2009); see, for a survey, Inaba (2019, appendix A). By adding emerging countries to their own samples, P&R (2009) and Inaba (2019) find upward trends in many sample countries' DGFs and in the average for all sample countries. Beine and Candelon (2011), Chuluun (2017), and Inaba (2019) go ahead and investigate the determinants of national stock DGFs.

question by arguing, "letting the exchange rate float may not be enough to insulate the domestic economy, even if it is a large economy, from global factors and permit monetary policy independence" (Rey, 2016, p. 7). A dilemma referred to hereafter as the Rey-type dilemma is, "independent monetary policies are possible if and only if the capital account is managed, directly and indirectly" (Rey, 2013, p. 287).

This article proceeds as follows. Section 2 measures national DGFs. Section 3 investigates the driving forces behind DGFs. Section 4 concludes with policy discussions.

2. Global Factors and Countries' Sovereign Bond Price Changes

2.1. National Sovereign Bond Prices

I start by assuming a global sovereign-bond investor who rolls over a one-week U.S. dollar debt and manages a GDP-weighted sum of countries' zero-coupon ten-year sovereign bonds without hedging foreign exchange (FX) fluctuation risks in local currencies. She invests in 41 countries, including both advanced and emerging countries.[2] The sum of these members' GDPs accounted for around 90 percent of world GDP in 2015. Allocation rates to the members will change every year in response to changes in their GDP percentage shares.[3]

The investor's dataset is on a weekly basis over the period 2007–16. She calculates the prices of the countries' ten-year sovereign bonds by using their zero-coupon yields calculated by Bloomberg. The sovereign bond prices are converted into U.S. dollars with reference to FX rates in the market. As detailed in Appendix A, the weekly excess returns accruing from investing in individual countries' ten-year sovereign bonds are the weekly changes in dollar-based prices of those bonds minus U.S. dollar one-week interest costs. The weekly excess return accruing from the global portfolio is the GDP-weighted average of the country-specific excess returns.

2.2. Estimating National DGFs

Suppose that the hypothetical global investor attempts to gauge countries' DGFs at the end of a sample year by using historical weekly data for that year. To help her, I follow P&R (2009) and define a country's DGF as a R_{adj}^2 of the model below estimated every sample year:

[2] They include in alphabetical order: Australia (AUS), Austria (AUT), Belgium (BEL), Brazil (BRA), Canada (CAN), Chile (CHL), China (CHN), the Czech Republic (CZE), Denmark (DNK), Finland (FIN), France (FRA), Germany (DEU), Greece (GRC), Hong Kong (HKG), Hungary (HUN), India (IND), Indonesia (IDN), Ireland (IRL), Israel (ISR), Italy (ITA), Japan (JPN), Malaysia (MYS), Mexico (MEX), the Netherlands (NLD), New Zealand (NZL), Norway (NOR), the Philippines (PHL), Poland (POL), Portugal (PRT), Russia (RUS), Saudi Arabia (SAU), Singapore (SGP), South Africa (ZAF), South Korea (KOR), Spain (ESP), Sweden (SWE), Switzerland (CHE), Thailand (THA), Turkey (TUR), the United Kingdom (GBR), and the United States (USA). This sample includes 13 emerging countries belonging to the Group of Twenty and/or the Executives' Meeting of East Asia and Pacific Central Banks (EMEAP). This forum's members are JPN, AUS, NZL, KOR, HKG, SGP, CHN, IDN, MYS, THA, and PHL. The investor believes that her world sovereign bond portfolio reflects well information incorporated into sovereign bond prices in all parts of the globe because it covers Asia, Africa, and Latin America.

[3] She uses those shares as weights, believing that GDP is a good proxy for the size of a national economy. An alternative method is to use country-specific market capitalisations as weights. She does not employ this method because it has the risk of disproportionally weighting countries whose governments are greatly indebted, such as Japan, Italy, and Greece.

$$ER_{i,t} = \beta_{i0} + \beta_{i1}GF1_t + \beta_{i2}GF2_t + \beta_{i3}GF3_t + \beta_{i4}GF4_t + \beta_{i5}GF5_t + e_{i,t}, \quad (1)$$

where t is a weekly point in time, ER_i is country i's government bond excess return mentioned above, $GF1$–$GF5$ are the GFs (global factors) considered, β_is are coefficients, and e is assumed to be independent and identically normally distributed. Eq. (1) relies on the arbitrage pricing theory (APT) proposed by Ross (1976): ER_i is represented as a set of latent risk factors common to all ERs and a component (e) for a specific sample year. The assumption that there are no omitted variables correlating with GFs makes β_is free of omitted-variable biases. I will discuss how to obtain the GFs and why there are five later. It should be noted here that identified risk factors can differ by sample year, depending on the global market conditions in each sample year.

The R_{adj}^2 of Eq. (1) is written as:

$$R_{adj}{}^2{}_i = 1 - \{\textstyle\sum_{t=1}^{n} \acute{e}_{i,t}{}^2 / \sum_{t=1}^{n}(ER_{i,t} - \overline{ER_i})\,(ER_{i,t} - \overline{ER_i})\} \times \{(n-1)/(n-5)\}$$

$$= R\_DGF_{i,t}, \quad (2)$$

where $\acute{e}$ is estimated residuals, $\overline{ER}$ is the mean, n is the number of observations, and 5 is the number of GFs. This formulation of R_DGF has an advantage of avoiding technical difficulties with inter-country correlation coefficients. First, a correlation coefficient between two countries' ERs tends to decrease due to the non-proportional differences in β_is for their counterpart GFs. Second, this tendency becomes prominent as the number of GFs increases. Last, to interpret an upward trend in the correlation coefficient as an increasing comovement, it is necessary to assume that the volatility of e is zero. In this regard, R_DGF is expected to increase "over time even if factor exposures (β_is) or factor volatilities decrease rather than increase, as long as country-specific residual volatility is not zero" (P&R, 2009, p. 218, terms in parentheses added by the author).

The principal component (PC) analysis is often used in APT based models to identify GFs (Chamberlain and Rothschild, 1983; Connor and Korajczyk, 1988). That is, the five GFs are the first five PCs extracted from individual countries' ERs. $GF1$, or the first PC, is a latent factor accounting for as much of the variability in those ERs as possible. $GF1$ is a liner combination of products of country-specific ERs (risk premium) for the most influential risk factor and their exposures to that. Good proxies for these exposures are eigenvectors obtained by the PC analyses.

To be more precise, I obtain five GFs by conducting the PC analyses every sample year using weekly data of all individual sample countries' GDP-weighted ERs. I use the countries' GDP percentage shares as weights because they are allocation rates to the member countries of the hypothetical portfolio under study. When constituents are not equally weighted in a portfolio, treating them equally runs the risk of creating biased PCs (Brown, 1989). Since P&R (2009) treat sample countries equally in obtaining their GFs by making PC analyses, they seem to implicitly assume a portfolio consisting of equally-weighted member countries. This assumption does not match that hypothetical portfolio.[4]

[4] Another difference from P&R's (2009) computation is that they use out-of-sample PCs while I use in-sample PCs. There are two reasons why I use them: one is that, as mentioned above, the hypothetical global investor can use historical weekly data at the end of a sample year when estimating countries' DGFs for the year. The other reason is that in-sample

When P&R (2009) gauge national DGFs for stock returns, they regard as GFs the first ten PCs whose cumulative eigenvalues are around 90 percent. Fig. 1 shows that, in my case, using the first five PCs meets this criterion.[5]

[Fig. 1 near here]

I obtain *GF1–GF5* on a weekly basis every year. A precise explanation of *GF1–GF5* is that they are unidentifiable GFs which are uncorrelated with each other. Such complete absence of correlation is indispensable for meeting the above-mentioned assumption made to avoid potential omitted-variable biases in estimating Eq. (1).[6] Fig. 2 plots their annual averages by year. The upper panel of Appendix B shows their descriptive statistics across the sample years, although there is no guarantee that a specific *GF* keeps representing an identical economic or financial factor for different years; for example, *GF1* could reflect changes in U.S. monetary policy in a particular sample year, but changes in energy prices in a different sample year.

[Fig. 2 near here]

Finally, I make ordinary least squares (OLS) estimations of Eq. (1) with around 52 weekly observations every sample year for all individual sample countries. Based upon Eq. (2), I gain one *R_DGF* for one sample country every sample year.[7]

2.3. Individual and Grouped National DGFs

Fig. 3 plots *R_DGF*s by country-group and shows three observations: first, all kinds of group DGFs changed stably around their own levels; second, national DGFs were larger in advanced countries than in emerging ones; and last, the European DGF was apt to be larger than other groups.

[Fig. 3 near here]

Fig. 4 plots *R_DGF*s by country and shows two observations: first, there are no clear trends in any of the national DGFs, except for SGP, ITA, and GRC; and last, the USA's DGF looks almost constant and one in all sample years. The last observation tells a subtle characteristic of *GF*s. *GF*s are the regressors for a country's *ER* in Eq. (1) while being obtained by the PC analysis using *ER*s of all countries, including the country. This suggests that there should be the endogeneity between countries'

PCs are exactly orthogonal. P&R (2009) obtain ten PCs in a year for multiplying countries' stock returns in the year by eigenvectors (factor loadings) gained for the returns in the previous year. The resulting PCs are not exactly orthogonal; that is, there is a risk of multicollinearity. In this regard, P&R (2009) argue that they separately find the correlations among their ten PCs too mild to make the risk serious.

[5] Driessen *et al.* (2003) apply APT-based models to U.S., German, and Japanese government bond returns of several maturities and find that five common factors explain almost all the variation in the returns.

[6] The absence cannot be fully secured by any data-based and meaningful indicators for GFs because they are more-or-less correlated with each other.

[7] Given space constraints, I present only two observations on the results of 410 plain OLS estimations of Eq. (1), as follows. First, on the above-assumed normality of *e*, the Jarque-Bera test does not reject null hypotheses that *e*s have the normalities in 309 regressions, but the tests do in 101 regressions, applying the 10% significance level. The rejections take place more frequently in emerging countries than in advanced ones, excluding GRC. Although the rejection ratio, 24.6%, appears to insufficiently low, I do not think that the ratio prevents me from using Eq. (1) to gauge national DGFs. This is because the normality assumption does not directly affect their size (although its collapse affects the statistical significances of estimated βs). Last, very small negative R_{adj}^2s are gained in five regressions. These R_{adj}^2s appear irregular because a R_{adj}^2 is interpreted here as the percentage of non-diversifiable systematic risks in total risks of *ER*. Therefore, I regard the negative R_{adj}^2s as zero.

*ER*s and five *GF*s. Because the PC analyses are based upon GDP-weighted *ER*s, *GFs* are exposed to the greater risk of endogeneity with respect to sample countries of larger economy. As discussed in Appendix C, I do not take that risk to be a concern and consider the U.S. ten-year sovereign bond price changes to be a good proxy for GFs for which I control with five *GF*s. I will drop U.S. from the sample in the next section.

[Fig. 4 near here]

Finally, individual countries and country-groups' DGFs were substantial: the average for the sample countries and period is 56 percent.

3. Determinants of National DGFs

3.1. A Panel-Data Regression: Model

This article does not fall into a strand of studies applying term-structure models to yield curve fluctuations across some or all G7 countries, including Diebold *et al.* (2008), Jotikasthira *et al*. (2015), and Byrne *et al.* (2019). These studies find that the G7 countries' long-term sovereign-bond yields are dominated by factors which are common to all of the countries – a finding in line with that in Section 2. Nevertheless, they do not set great store by differences over time and among G7 countries in the dependence on G7 common factors.[8] This section looks into such differences at a global level.

Time-series and cross-country differences in national DGFs gauged using P&R's (2009) method are attributable to country-specific factors: omitted variables carried by e (residuals) in Eq. (1). To test the global financial cycle hypothesis, these factors include the effects of monetary authorities' policies: FX management, capital control, and short-term interest rates. My baseline regression equation is:

$$L\_DGF_{i,\tau} = h_0 + h_1 Size_{i,\tau} + h_2 FXV_{i,\tau} + h_3 CBM_{i,\tau-1} + h_4\, MP_{i,\tau} + h_{\mathrm{IT}}[CBM_{i,\tau-1} \times MP_{i,\tau}] + h_5 FP_{i,\tau} + h_6 SR1_{i,\tau} + h_7 SR2_{i,\tau} + h_8 TO_{i,\tau} + IE_i + \varepsilon_{i,\tau}, \quad (3)$$

where L_DGF is a national DGF which is slightly different from R_DGF as will be explained later, i stands for individual sample countries, τ stands for a yearly-point in time, hs are coefficients, IE stands for i's heterogeneities carried by omitted variables and unobservable factors, and ε is residuals. Before explaining the nine regressors, I mention here two points as follows. First, L_DGF is the generalised logit-transformation of the square root of R_DGF gained in Section 2 by estimating Eq. (1). The logit-transformation is applied to i's R_DGF in order to transform its range [0, 1] to [0, +∞]:

$$L\_DGF_{i,\tau} = \ln\{(1 + \sqrt{R\_DGF_{i,\tau}})/(1 - \sqrt{R\_DGF_{i,\tau}})\}. \quad (4)$$

The other point is that time effects common to all sample countries (is) in individual sample years (τs) are unnecessary because *GF*s in Eq. (1) include such common effects.

Based on the global financial cycle hypothesis, I suppose that foreign investors access a country's sovereign bond market in response to GFs. They would affect more the price of a country's sovereign

[8] On the other hand, they make some clarifications of what the G7 common factors are, including G7 level/slope factors, G7 fundamental/non-fundamental factors, and G7 future policy-rate expectations/term premia.

bonds when the country's bond market is smaller in comparison with their investment flows.[9] *Size* controls for the impact of flow size on sovereign bond prices. It is the ratio of the absolute value of foreign investors' portfolio-bond investment flows to the market capitalisation of all domestic bonds. An estimate ($\hat{h}_1$) for *Size* should be positive.

Because *L_DGF* refers to sovereign bond prices in U.S. dollars, both a country's sovereign bond and FX markets are areas where GFs can affect the price of its sovereign bonds. *FXV* controls for the exposure of a country's FX market to GFs. It is the coefficient of variation for the country's FX rates, or the rates' variability, based on the assumption that these rates tend to more volatile as its FX market is open more to GFs. This assumption will be verified if an estimate ($\hat{h}_2$) for that is positive.

CBM represents the closedness of *i*'s bond market, taken from Fernández *et al.* (2016). *CBM* is equal to one if there are some control measures on both foreign investors' "purchasing" and "selling" of local bonds, 0.5 if there are measures controlling either of them, and otherwise zero.[10] Since a *CBM* of zero indicates a bond market that is fully open to foreign investors, I expect its estimate ($\hat{h}_3$) to be negative. *CBM* refers to a previous point in time (τ–1), in order to avoid potential statistical problems.[11]

To verify the Rey-type dilemma, I examine the direct impact of short-term interest-rate management on a national DGF with *MP* which stands for monetary policy effect on the domestic economy and prices. *MP* is the absolute value of changes in "real short-term interest-rate gaps" – a theoretically suitable indicator for that effect (Woodford, 2003). Real short-term interest rates minus natural interest rates – hypothetical interest rates that are neutral to business climate – leaves these gaps.[12] A larger *MP* means a greater effect. A negative estimate ($\hat{h}_4$) for that, if gained, will suggest that *MP* is likely to have been an independent local factor in sovereign bond prices: a disconfirmation of the dilemma.

I also look into the dependence of monetary policy effect on *CBM* by adding an interaction term: $CBM_{-1} \times MP$. I will interpret a negative $\hat{h}_{\mathrm{IT}}$, if estimated, as suggesting that the negative impact of *MP* on a national DGF is likely to have been strengthened as foreign investors were regulated.

The four regressors in the second line of Eq. (3) deal with factors apart from the global financial cycle hypothesis. In analogy to monetary policy, fiscal policy (*FP*) may also reduce a country's DGF by affecting its effective demand. *FP* is the absolute value of changes in fiscal surplus/deficit over nominal GDP. Because a larger *FP* means a greater policy effect, its estimates ($\hat{h}_5$) should be negative.

Investors can add a credit-risk premium on yields on a bond issued by a government whose fiscal sustainability is weak. *SR1* and *SR2* control for sovereign risks. *SR1* is a common indicator for the indebtedness of the national finance, or outstanding government debt over GDP. *SR2* is a World Bank

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[9] Transaction costs can give rise to illiquidity discounts on asset prices (Amihud and Mendelson 1991; Lo *et al.* 2004). A more liquid financial asset can be bought and sold in the market with a relatively small impact on its market price. The size of a financial market is one of the conventionally-used indicators for market liquidity.

[10] Due to the nature of data availability, I ignore an unfortunate gap: *Size* and *CBM* refer to both private and sovereign bonds, while *L_DGF* refers to sovereign bonds of a specific ten-year maturity.

[11] For example, a country's larger DGF could encourage its authorities to regulate foreign investors to maintain domestic financial stability. Another example is a potential correlation between *CBM* and $|BI|$ (the numerator of *Size*).

[12] As explained in Appendix A, these two rates are simply formulated due to the nature of computability and data-availability for all sample countries.

indicator for countries' political stability. Because I interpret a larger *SR1* and a smaller *SR2* as implying larger sovereign risks, their estimates ($\hat{h}_6$ and $\hat{h}_7$) should be negative and positive, respectively.

TO stands for the exposure of a national economy to the world business climate and global inflation. Such GFs can affect a country's inflation.[13] *TO* is Index of Trade Freedom which The Heritage Foundation calculates for individual countries by considering tariffs, taxes, and bans. A larger *TO* means fewer restrictions. Byrne *et al.* (2019) find that a G7-common inflation affects G7 sovereign bond yields, but two of their components – future policy-rate expectations and term premia – offset one another. If this is the case on a global average basis, *TO*'s estimate ($\hat{h}_8$) will be tiny or insignificant.

3.2. Estimation Results

I estimate Eq. (3) by using an unbalanced panel dataset covering 33 countries over the period 2007–15. Seven countries are excluded from the sample due to the nature of data availability.[14] U.S. is excluded too because its DGF is stably close to one, as shown in Section 2. The character of *IE*s makes Eq. (3) take one of three potential forms: first, a pooling model represented by dropping *IE*s from Eq. (3); second, a fixed-effect model, or Eq. (3) in which *IE*s are country-specific constants; and last, a random-effect model, or Eq. (3) in which *IE*s are country-specific stochastic variables.[15]

As detailed in Appendix D, I follow the conventional procedure and select the fixed-effect model. The result of estimating Eq. (3) using a generalised least squares (GLS) method is:

$$L\_DGF = 5.29^{***} + 0.01^{*}Size + 0.04^{***}FXV - 1.19^{**}CBM_{-1} - 0.02^{**}MP - 0.01\,[CBM_{-1} \times MP] \\ - 0.07^{**}FP_{i,\tau} - 0.48^{**}SR1 - 0.00SR2 - 0.02TO \qquad (\#\text{ of observations} = 287,\ R_{adj}^2 = 0.79)$$

where the superscripts ***, **, and * stand for one percent, five percent, and ten percent statistical significances, respectively, and the p-values used are the averages of two cases in which ε's four potentially irregular aspects are separately adjusted for. The fixed-effect estimates which are statistically significant are *Size* (+), *FXV* (+), CBM_{-1} (–), *MP* (–), *FP* (–), and *SR1* (–). The signs in parentheses are those of their counterpart $\hat{h}$s, and they are the same as expected above.

Thus, the driving forces behind national DGFs are individual effects (*IE*) and time-varying factors. National DGFs tended to be larger in countries with (i) smaller and (ii) less-closed bond markets, where (ii) FX rates were more variable, where (iii) macro-economic policies were more neutral, and where (iv) the national finance was more sound. Specifically, (ii) above is line with my assumption that a country with larger *FXV* has an FX market which is more sensitive to GFs. More interestingly, the (iii) for *MP* and the insignificance of the interaction term disconfirm the Rey-type dilemma.

[13] See, for example, Ciccarelli and Mojon (2010), Borio and Filardo (2007), and Borio (2017).

[14] CHL, CZE, HUN, IND, ISR, POL, and SAU are excluded.

[15] When either a fixed-effect model or a random effect model is selected, four potential irregular aspects of residuals (ε) need to be addressed to gain asymptotically consistent estimates ($\hat{h}$s): first, cross-section heteroskedasticity; second, period heteroskedasticity; third, contemporaneously correlation; and last, serial correlation. These can reduce the reliability of the results of t-tests on the estimates. Among these irregular aspects, cross-section and period heteroskedasticity could be acute for *L_DGF*s because *L_DGF*s are logit-transformed variables (Kataoka, 2005).

Using the estimation result above, I calculate the average marginal effects of one-unit increases in the effective regressors on *R_DGF*s. The average of 33 sample countries' *L_DGF*s over the period 2007–15 is 2.55, which translates into *R_DGF* of 0.73.[16] Increasing *CBM* from zero to one, or regulating foreigners' ability to both purchase and sell local bonds, has a considerable impact; that is, doing so contributed toward reducing *R_DGF*s by 0.39.

By contrast, the impacts of other regressors on *R_DGF*s are estimated to be rather insubstantial. For example, a one-unit increase in *FXV* contributed toward enhancing *R_DGF*s by 0.01. The sample-period average of *FXV* is 4.30 for JPN, CAN, and European countries, all of which employ floating systems. The average for HKG with a fixed-rate system is 0.1. The transition from a floating system to a fixed-rate system could result in a reduction in *R_DGF*s of only 0.04. One-unit increases in *MP* and *FP* contributed toward reducing *R_DGF*s by 0.004 and 0.016, respectively, suggesting that monetary and fiscal policies need to be extremely drastic so as to have any impact on national DGFs.[17]

3.3. Two Extensions

I make two extensions to check the robustness of the independent monetary policy effect. First, I analyze changes in sovereign bond prices quoted in domestic currencies because monetary authorities would be much more interested in these prices than those in U.S. dollars.[18] Following the same procedure as in Section 2 for ER^{LC}s, I estimate local currency based DGFs, or $\mathrm{DGF}^{\mathrm{LC}}$s. The lower panel of Appendix B shows the descriptive statistics of five *GF*s on which ER^{LC}s are regressed. Fig. 5 plots the simple averages of all countries' *R_DGF* and R_DGF^{LC}. This average is larger for *R_DGF* than for R_DGF^{LC} in all sample years.[19] The gap between the two is stable over time, except for the aftermath of the 2008 crisis.

[Fig. 5 near here]

The lower panel of Appendix D shows the result of a GLS estimation of Eq. (3) for L_DGF^{LC} (instead of *L_DGF*). I select the fixed-effect model. *MP* gains a statistically significant and negative estimate, as in the baseline estimation for *L_DGF*. Notably, the interaction term also gains a statistically

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[16] Because of the non-linearity of the logit-transformation, this value is different from 0.59 – the value gained by straightforwardly averaging 33 sample countries' *R_DGF*s over that period.

[17] The average marginal effects of *Size* and *SR1* are as follows. A one-unit (one percent point) increase of *Size* contributed toward enhancing *R_DGF*s by 0.002. *Size*'s sample-country and sample-period average and standard deviation are 6.1 and 5.1, respectively. Even a one-standard-deviation increase in *Size* can result in a reduction in *R_DGF*s of only 0.01. A unit-increase of *SR1* contributed toward reducing *R_DGF*s by 0.13. It is necessary for public finances to deteriorate massively to achieve this impact because *SR1* is a natural logarithm of debt-to-GDP ratios whose sample-country and sample-period average is 62.5 percent. *SR1*'s unit-increase from its sample-country and sample-period average of 4.13 is equivalent to a 107 percent point increase in that average ratio.

[18] The other reason is that analyzing changes in sovereign bond prices in U.S. dollars assumes that the hypothetical global investor is willing to take FX fluctuation risks or is obliged to do so due to the lack of tools to hedge the risks. This assumption reflects reality, but only partially, because cheap hedging-tools are available in some FX markets. Here, I suppose another extreme case where the global investor holds their sovereign bond portfolio while fully hedging those risks; that is, local currency based *ER*s do matter. The reality is somewhere between these two hypothetical cases.

[19] This excess is the impact of GFs through the FX evaluation channel: the FX-rate change directly alters the U.S. dollar value of the sovereign bond prices. The sample-period averages of these two global DGFs are 56 percent and 48 percent for *R_DGF* and R_DGF^{LC}, respectively. On average, the impact of GFs through the FX evaluation channel accounted for 15 percent of the *R_DGF*-type global DGF: (1 – 48 percent / 56 percent) $\times$ 100 = 15 percent. Both *R_DGF* and R_DGF^{LC} carry the impact of GFs through the FX bond pricing channel: the FX-rate change has implications for domestic prices, such as import-inflation/deflation, thereby affecting the sovereign bond prices quoted in local currency.

significant and negative estimate. This suggests that regulating foreign investors is likely to have helped increase the impact of *MP*. Other significant regressors are the same as in the baseline estimation.

The second extension is to split the sample period into a crisis period (2008–11) and a post-crisis period (2012–15), in the former of which national monetary authorities used their policy tools aggressively in response to sudden changes in domestic asset prices due to GFs. I drop *SR2* and *TO*, which are insignificant in the baseline estimation, from Eq. (3) so as to increase the degree of freedom, and estimate such a regression equation for each of the two periods for both *L_DGF* and L_DGF^{LC}.

[Table 1 near here]

Table 1 summarises the results of four cases, for all of which I select the fixed-effect models for the comparability. The significant regressors in the two crisis cases are *FXV*, *CBM*, and *SR1*. In the post-crisis cases, *CBM*, *MP*, *FP*, and *SR1* are significant regressors for *L_DGF*, while so are *FXV* and the interaction term for L_DGF^{LC}. All of these regressors are effective also in the baseline estimation.

Additional findings are the following. First, the independent monetary policy effect was not persistent across the cases. The effect is found in the crisis and post-crisis periods for L_DGF^{LC} and *L_DGF*, respectively. It depended on the presence of regulations on foreign bond investors for L_DGF^{LC} not in the crisis period but in the post-crisis period. These disconfirm the Rey-type dilemma for L_DGF^{LC} in the crisis period. This could be a kind of relief to national monetary authorities because "L_DGF^{LC} in the crisis period" was of great interest to them in the context of domestic financial stability. Second, the positive association of national DGFs with *FXV* was more stable as it is found in all cases, except for the post-crisis period for L_DGF^{LC}. Last, either *CBM* itself or its interaction with *MP* is a significant regressor in both of the two periods for both *L_DGF* and L_DGF^{LC}.

4. Concluding Remarks

By analyzing national DGFs for 41 countries' 10-year sovereign bond returns over the period 2007–16, this article finds as follows. First, countries' sovereign bond markets have been well integrated. Second, such integration has not necessarily proceeded rapidly over the last decade. This is because the developments of different country-group DGFs, including emerging markets, were not on upward trends. Third, the similar developments of the group DGFs over time provide credible evidence of a global financial cycle. Fourth, the progress of sovereign bond market integration has differed by country-group. The markets in emerging countries were less integrated with the rest of the world than those in advanced countries. The markets in European countries were more integrated with the rest of the world than those in other countries. These appear to be consistent with investors' broad view of international sovereign bond market integration. Finally, U.S. sovereign bond market is special in the sense that its fluctuations are dominant factors of GFs. U.S. ten-year T-note futures thus would be a useful tool for hedging risks of price changes in global sovereign-bond portfolios.

By investigating the driving forces behind national DGFs in relation to the global financial cycle hypothesis, this article, on the one hand, shows that its policy implication, or the Rey-type dilemma, cannot be taken as read. This is because controlling FX rates, conducting non-neutral interest-rate policy, and regulating foreign bond investors all had the capacity to affect national DGFs. The monetary policy effect was negatively associated with a national DGF, independently of the presence of regulations on the investors.

This article, on the other hand, finds that the policy effectiveness differed by tool. The effect of regulating foreign bond investors was greater and more stable than the other two. In addition, an implication of the dilemma that such regulations help the monetary policy effect improve was confirmed with respect to national DGFs estimated separately for changes in sovereign bond prices quoted in local currencies. These findings could tempt the authorities to make those regulations, despite the IMF's (2012) proposal that capital control measures should be temporary and non-discriminatory between residents and non-residents.

This article closes with three caveats that do not allow the findings to be fully generalised. The first is that the findings are gained through the lens of a realistic but very specific global investor investing in 41 countries' 10-year sovereign bonds using GDP percentage shares as allocation rates. Expanding the coverage of member countries and changing allocation rates to them could affect the measurements of both GFs and national DGFs. The second caveat is that the GFs used for measuring countries' DGFs are blurred because proxies for them are not specific economic and financial variables but the PCs extracted from changes in countries' sovereign bond prices. I accept this limitation in order to enjoy the tractability and simplicity of P&R's (2009) method. The last caveat is that the implications for monetary policy effect could be sensitive to the simple formulation of real short-term interest-rate gaps and to the coarse content of the indicator for regulations on them. The latter indicator tells us merely whether or not national authorities impose these regulations. To my knowledge, internationally comparable indicators are not available for the strictness of such regulations. Beyond these caveats, this study will hopefully serve as a good initial step for further research on international sovereign bond market integration and the global financial cycle hypothesis.

Appendix A

[Appendix Table A here]

Appendix B

[Appendix Table B here]

Appendix C

The Risk of Endogeneity between *GF*s and Sample Countries' *ER*s

My *GFs* are exposed to the greater risk of endogeneity with respect to sample countries of larger economy. P&R (2009) exclude a sample country in making the PC analysis to estimate GFs for the country. Although this method would be beneficial in hedging that risk, I do not employ it for three reasons. First, their GFs are not common to their sample countries. I believe that GFs should be common to all sample countries. Second, when sample countries are of larger GDP, the method is likely to estimate GFs making sense less to the hypothetical global sovereign-bond investor investing in a GDP-weighted portfolio. Last, the method is more likely to grasp too small GFs for countries in which domestic factors for their asset prices are more directly linked with GFs, such as oil prices for oil-exporting countries and interest rates on the U.S. dollar, or the sole key currency, for the U.S.

I regard the endogeneity risk as negligible for the following reasons. First, an observation that a number of small-economy countries, especially European ones, have large DGFs may serve as counter-evidence for that risk.

Second, if the risk had been systematic and substantial over that sample period, there should have been downward trends and upward trends in DGFs for advanced and emerging countries, respectively, because GDP percentage shares have broadly decreased and increased for the two groups, respectively.

The last reason is that, by calculating the Wu-Hauman statistics to investigate the endogeneity of the U.S. *ER* with respect to *GF1–GF5* every year over the period 2007–16, I find that the risk can be rejected in 45 out of 50 cases. To be specific, I assume that instrument variables for *GF*s at t are *ER* at $t–1$ and their own *GF*s at $t–1$; for example, $GF2_{t-1}$ and U.S. ER_{t-1} for $GF2_t$. As mentioned in the text, *GF1–GF5* are exactly orthogonal with each other. I regress individual *GF*s on their instruments every year over the 10-year sample period and gain counterpart 50 residuals. By repeatedly adding as a regressor one of the residuals in Eq. (1), I gain 50 augmented equations. As a result of their OLS estimations, a null hypothesis that one of the added regressors is exogenous in its augmented equation cannot be rejected in only five regressions: U.S. *ER* has the potential to explain *GF2* and *GF4* in 2010, *GF1* and *GF4* in 2014, and *GF3* in 2015.

Consequently, the fact that U.S. DGF is almost full, I consider, means that the U.S. ten-year sovereign bond price changes is a good proxy for GFs for which I control with five *GF*s.

Appendix D

[Appendix Table D here]

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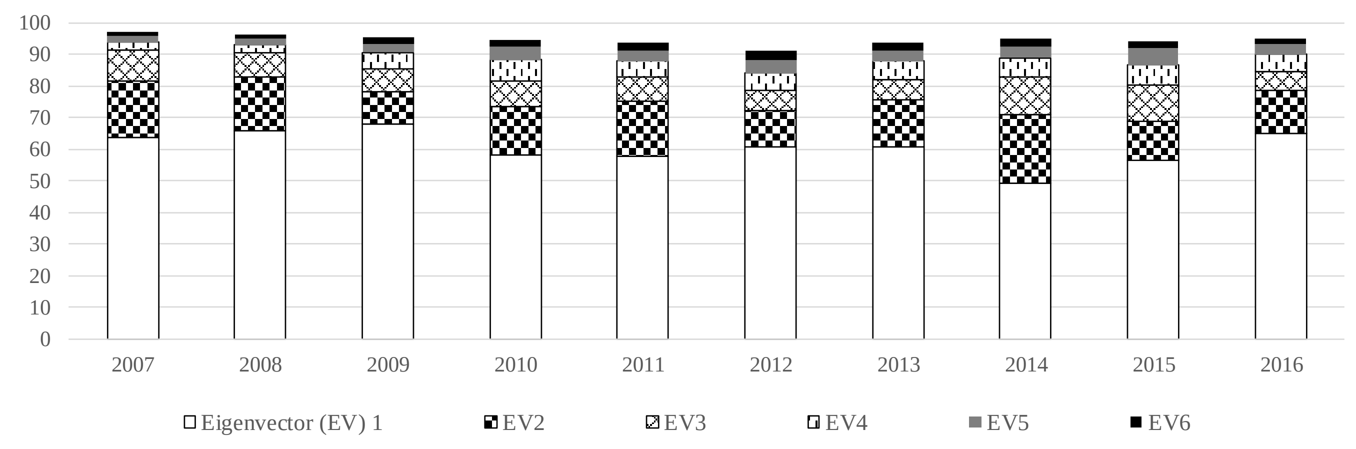
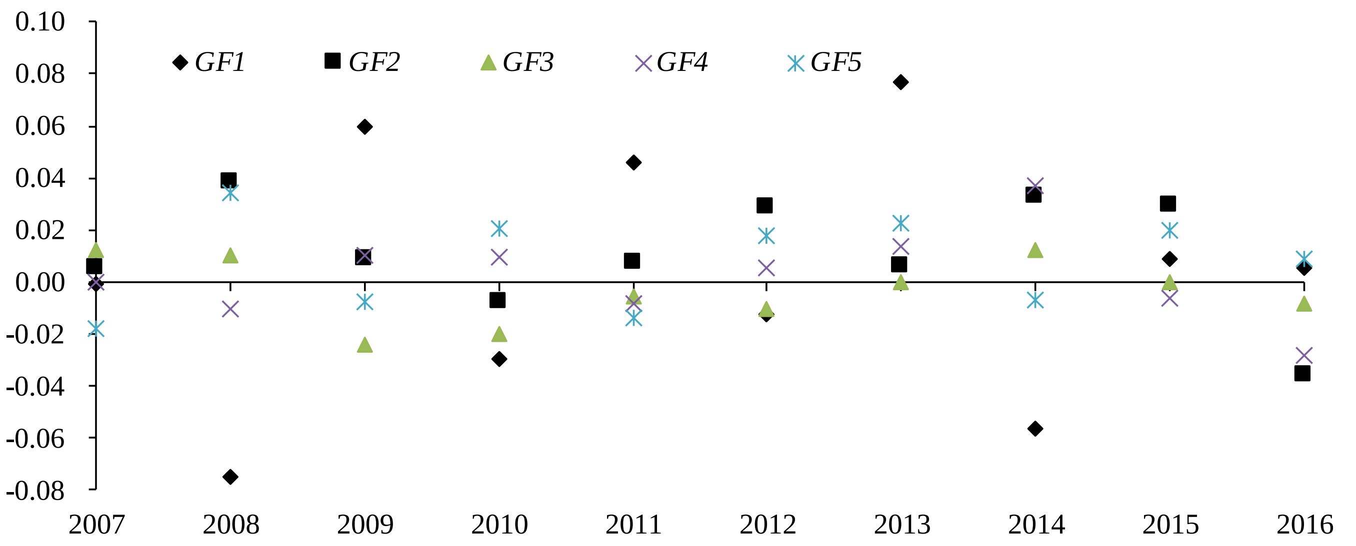
Figures

Fig. 1. Percent cumulative eigenvalues of the Principal Components (PCs)



Note: The PC analyses are made every sample year for all sample countries using weekly data of individual sample countries' GDP-weighted *ER*s (excess returns of 10-year sovereign bonds).

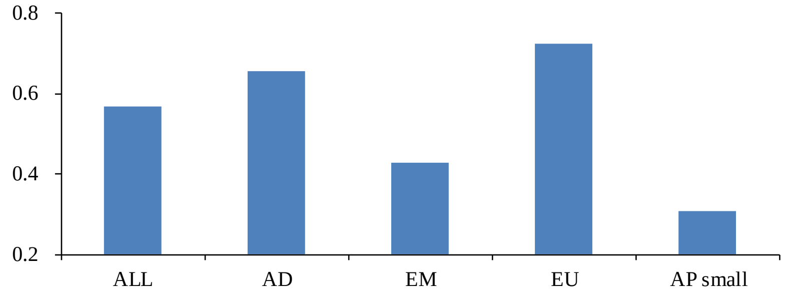
Fig. 2. Developments in *GF1–GF5*



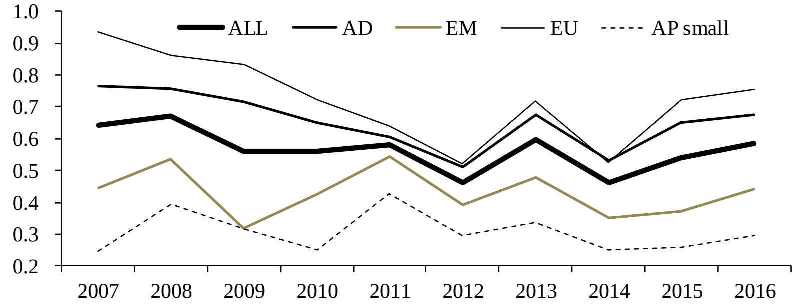
Note: Annually-averaged values of the first five PCs on a weekly basis.

Fig. 3. *R_DGF*s by Country-Group

Sample-period averages of *R_DGFs* by country-group



Simple averages of country-group members

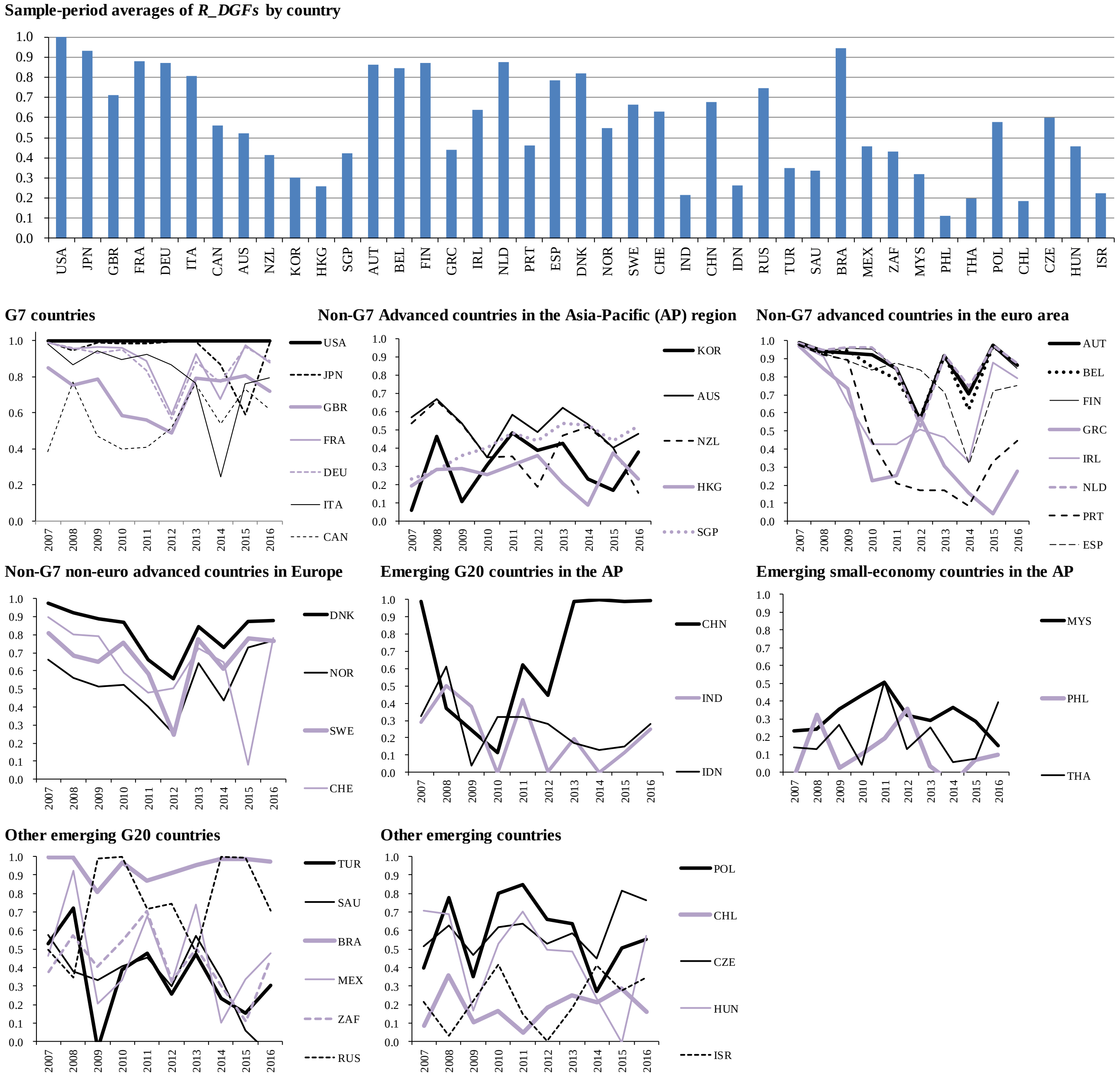


Notes 1: A national DGF at τ (a yearly point in time) is a *R_DGF* defined in Eq. (2), or a R_{adj}^2 gained by estimating Eq. (1) for individual sample countries with around 52 weekly observations.

2: ALL stands for all sample countries, AD for advanced countries, EM for emerging countries, EU for European countries, and AP small for EMEAP countries, excluding JPN and CHN.

3: The distinction between advanced and emerging countries is based on the International Monetary Fund's *World Economic Outlook*.

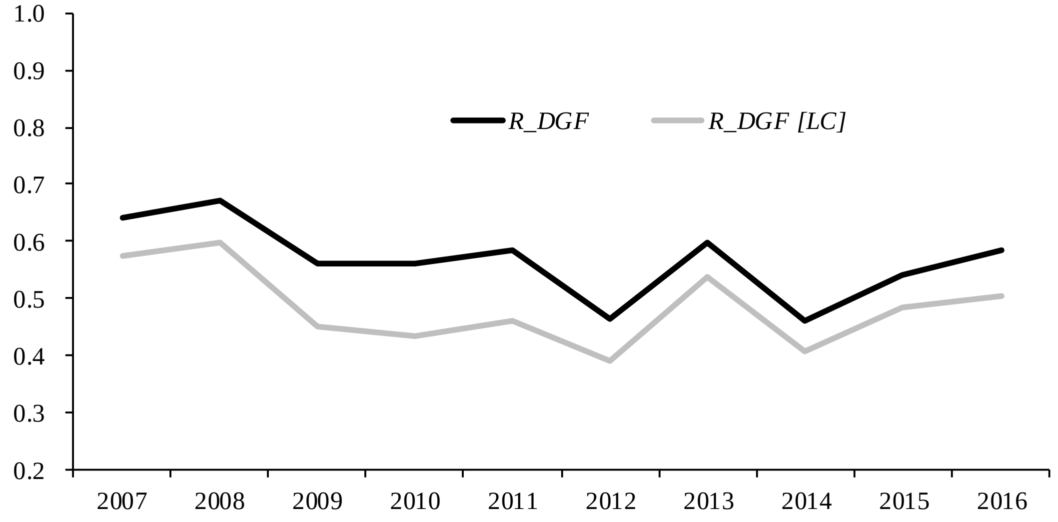
Fig. 4. *R_DGF*s by Country



Notes 1: A national DGF at τ (a yearly point in time) is a *R_DGF* defined in Eq. (2), or a R_{adj}^2 gained by estimating Eq. (1) for individual sample countries with around 52 weekly observations.

2: The distinction between advanced and emerging countries is based upon the International Monetary Fund's *World Economic Outlook*.

Fig. 5. Two Kinds of *R_DGF*s



Notes 1: The simple averages of the sample countries' *R_DGF*s in U.S. dollar value and in local currency value, the latter of which is denoted by *[LC]* here.

2: These two kinds of *R_DGF*s are defined in Eq. (2), or R_{adj}^2s gained by estimating Eq. (1) for individual sample countries with around 52 weekly observations.

Tables

Table 1

Summary of the Results of Estimating Modified Eq. (3) in Two Sub Sample Periods

	For *L_DGF* in 2008–2011 [129 observations]	**For L_DGF^{LC} in 2008–2011** [129 observations]	**For *L_DGF* in 2012–2015** [128 observations]	**For L_DGF^{LC} in 2012–2015** [128 observations]
C	6.56 ***	6.85 ***	3.07 ***	2.37 **:
Size	0.35	0.53 **	1.23	-1.04
FXV	0.04 **	0.04 ***	0.02	0.05 *
CBM_{-1}	-5.76 **	-3.89 **	-1.16 ***	-0.82
MP	-0.01	-0.03 **	-0.03 *	0.01
$CBM_{-1} \times MP$	0.14	-0.12	-0.09	-0.23 ***
FP	-0.04 **	-0.01	-0.14 ***	-0.06
SR1	-0.88 ***	-0.94 ***	-0.17 **	0.02
IE	Fixed effect (FE)	FE	FE	FE
R_{adj}^2	0.96	0.88	0.97	0.67

Notes 1: *SR2* and *TO* are dropped from Eq. (3).
2: Weighted GLS estimates are shown.
3: ***, **, and * stand for one percent, five percent, and ten percent statistical significances. The p-values used refer to the case where residuals' (ε's) three potentially irregular aspects (cross-section heteroscedasticity, period heteroscedasticity, and serial correlation) are adjusted for.

Appendix Table A

Definitions and Sources of the Data Used

Indicators	Notations	Definitions	Sources	Notes
Excess returns of sovereign bonds (value in U.S. dollars).	*ER*	{(10-year zero-coupon sovereign bond prices at t / those prices at t–1) – one week interest costs of U.S. dollar at t–1}.	Bloomberg	% points. Sovereign bond prices are converted into U.S. dollars using weekly foreign exchange rates. The one-week interest costs are linearly interpolated with FF effective rates and one-year Treasury bill yields.
Excess returns of sovereign bonds (value in local currencies).	ER^{LC}	{(10-year zero-coupon sovereign bond prices at t / those prices at t–1) – one week interest costs of U.S. dollar at t–1}.	See the above.	% points. Sovereign bond prices are quoated in local currencies. The one-week interest costs are linearly interpolated with FF effective rates and one-year Treasury bill yields. Hedging costs are abstracted from.
Flow-size impact	*Size*	(|BI| / outstanding amounts of domestic private and public debt securities)*100. Both the numerator and denominator are % of nominal GDP. BI = Foreigners' purchasing of portfolio-bonds minus their selling of the bonds.	IMF, *Balance of Payments Statistics* (BOPS). World Bank, *Global Financial Development Database* (GFDD).	%. Nominal GDPs are taken from IMF, *World Economic Outlook* (WEO). This is applicable to all indicators divided by nominal GDPs.
Foreign exchange variability	*FXV*	Coefficients of variation of foreign exchange rates: annual standard deviations / annual averages × 100.	Bloomberg.	Weekly data is used.
Closedness of a national bond market	*CBM*	*CBM* = (*CFP* + *CFS*)/2	Fernández et al. (2016)	0, 0.5, or 1.0. 0 means no regulations on foreign bond-investors.
	CFP	The presence of controls on non-residents' purchasing of local bonds.	See the above.	1 when some regulations exist, otherwise 0.
	CFS	The presence of controls on non-residents' selling of local bonds.	See the above.	1 when some regulations exist, otherwise 0.
Monetary policy effect	*MP*	|Real short-term interest-rate gaps at t – real short-term interest-rate gaps at t–1|. Real short-term interest-rate gaps = Real short-term interest-rates – natural rates. The real short-term interest rates = one-year sovereign bond yields – annual inflation rates. The natural interest rates = Potential growth rates based on local-currency real GDP smoothed by applying the Hodrick-Prescott filter with a multiplier of 100.	Bloomberg; IMF *International Financial Statistics* (IFS); World Bank *World Development Indicators* (WDI).	% points.
Fiscal policy effect	*FP*	|Cash surplus/deficit (% of nominal GDP) at t – cash surplus/deficit (% of nominal GDP) at t–1 | .	World Bank, WDI.	% points.
Sovereign risk 1: Indebtedness	*SR1*	Log (The outstanding debts of the general government <% of nominal GDP>)	World Bank, WDI.	
Sovereign risk 2: Political stability	*SR2*	An index reflecting measured perception of the likelihood of political instability and/or politically-motivated violence, including terrorism.	World Bank, *Worldwide Governance Indicators.*	Percentile rank among all countries ranges from 0 (least stable) to 100 (most stable) rank.
Institutional openness of trade.	*TO*	*Index of Trade Freedom*. A composite measure of the absence of tariff and non-tariff barriers that affect imports and exports of goods and services. Country i's score is written as: $(((T_{max}–T_i) / (T_{max}–T_{min})) \times 100) – NTB_i$, where T_{max} and T_{min} represent the upper and lower bounds for tariff rates (%), T_i represents the weighted average tariff rate (%), and NTB stands for a non-tariff barrier. NTB is 0, 5, 10, 15, or 20: a value dermined using both qualitative and quantitative information on the extensiveness of using non-tariff barriers.	The Heritage Foundation	A larger *TO* means a freer trade.

Appendix Table B

Descriptive Statistics of *GF1– GF5*

(1) The first five PCs on which countries' *ER*s are regressed.

	GF1	*GF2*	*GF3*	*GF4*	*GF5*
Mean	0.0024	0.0114	-0.0038	0.0021	0.0016
Standard Deviation	0.3534	0.1747	0.1296	0.0973	0.0618
Median	-0.0068	0.0050	0.0033	-0.0040	0.0032
Max	1.2296	1.3148	0.4288	0.2906	0.2187
Min	-2.4759	-0.6808	-0.5388	-0.4003	-0.2138

(2) The first five PCs on which countries' ER^{LC}s are regressed.

	GF1	*GF2*	*GF3*	*GF4*	*GF5*
Mean	0.0044	-0.0037	-0.0005	-0.0033	-0.0025
Standard Deviation	0.3407	0.1178	0.0843	0.0645	0.0516
Median	-0.0060	-0.0061	0.0011	-0.0041	-0.0028
Max	2.4637	0.5530	0.3779	0.2170	0.2479
Min	-1.1551	-0.8154	-0.3245	-0.3590	-0.2477

Notes 1: See Appendix A for details of *ER* and ER^{LC}.
2: Weekly data are used over the period April 2007–December 2016. The number of observations is 509.

Appendix Table D

Determinants of National DGFs

(1) Dependent variable: *L_DGF*

Model		A: Pooling	B: Fixed effect			C: Random effect	
Specification of *IE*		No	Yes: Constant			Yes: Stochastic	
Estimation method		OLS	LSDV	Weighted GLS		GLS	
Adjustments on residuals (ε)		-	-	White period	White cross-section	White period	White cross-section
				CSH, PH, & SC are adjusted for.	CSH & CCE are adjusted for.	PH & SC are adjusted for.	CSH & CCE are adjusted for.
Regressors	Estimators	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s
Constant term	*C*	0.185	5.188	5.286	5.286	5.055	5.055
		[0.868]	[0.026]	[0.000]	[0.003]	[0.020]	[0.075]
Flow-size impact	*Size*	-0.007	0.012	0.009	0.009	0.009	0.009
		[0.510]	[0.155]	[0.078]	[0.049]	[0.252]	[0.117]
FX rate variability	*FXV*	0.079	0.048	0.037	0.037	0.059	0.059
		[0.016]	[0.064]	[0.010]	[0.001]	[0.013]	[0.098]
Controls on foreign bond investors	CBM_{-1}	0.569	-2.688	-1.191	-1.191	-1.898	-1.898
		[0.270]	[0.000]	[0.064]	[0.001]	[0.142]	[0.035]
Monetary policy effect	*MP*	-0.031	-0.026	-0.024	-0.024	-0.029	-0.029
		[0.239]	[0.162]	[0.027]	[0.036]	[0.000]	[0.001]
Interaction term	$MP \times CBM_{-1}$	-0.073	-0.028	-0.010	-0.010	-0.052	-0.052
		[0.673]	[0.814]	[0.883]	[0.870]	[0.379]	[0.711]
Fiscal policy effect	*FP*	-0.045	-0.052	-0.066	-0.066	-0.053	-0.053
		[0.283]	[0.108]	[0.000]	[0.003]	[0.086]	[0.103]
Sovereign risk 1: Indebtedness	*SR1*	0.141	-0.632	-0.477	-0.477	-0.192	-0.192
		[0.112]	[0.008]	[0.085]	[0.001]	[0.421]	[0.287]
Sovereign risk 2: Political stability	*SR2*	0.004	-0.002	0.003	0.003	-0.006	-0.006
		[0.358]	[0.850]	[0.724]	[0.470]	[0.618]	[0.343]
Trade openness	*TO*	0.012	0.000	-0.015	-0.015	-0.018	-0.018
		[0.336]	[0.987]	[0.215]	[0.426]	[0.271]	[0.541]
R_{adj}^2		0.022	0.596	0.787		0.067	
F-test on H_0: Pooling model > Fixed-effect model		13.321 (p-value: 0.000)					
Hausman test on H_0: Random-effect model > Fixed-effect model			17.880 (p-value: 0.037)				

(2) Dependent variable: *L_DGF* ^{LC}

Model		A: Pooling	B: Fixed effect			C: Random effect	
Specification of *IE*		No	Yes: Constant			Yes: Stochastic	
Estimation method		OLS	LSDV	Weighted GLS		GLS	
Adjustments on residuals (ε)		-	-	White period	White cross-section	White period	White cross-section
				CSH, PH, & SC are adjusted for.	CSH & CCE are adjusted for.	PH & SC are adjusted for.	CSH & CCE are adjusted for.
Regressors	Estimators	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s	$\hat{h}$s
Constant term	*C*	1.702	7.560	7.920	7.920	5.620	5.620
		[0.133]	[0.002]	[0.000]	[0.003]	[0.000]	[0.032]
Flow-size impact	*Size*	-0.015	0.015	0.011	0.011	0.009	0.009
		[0.203]	[0.092]	[0.085]	[0.001]	[0.212]	[0.186]
FX rate variability	*FXV*	0.116	0.087	0.041	0.041	0.104	0.104
		[0.000]	[0.002]	[0.013]	[0.096]	[0.000]	[0.002]
Controls on foreign bond investors	CBM_{-1}	-0.159	-2.108	-1.290	-1.290	-1.609	-1.609
		[0.761]	[0.001]	[0.000]	[0.001]	[0.069]	[0.012]
Monetary policy effect	*MP*	-0.040	-0.017	-0.024	-0.024	-0.021	-0.021
		[0.132]	[0.387]	[0.000]	[0.056]	[0.000]	[0.059]
Interaction term	$MP \times CBM_{-1}$	-0.219	-0.302	-0.107	-0.107	-0.321	-0.321
		[0.212]	[0.018]	[0.099]	[0.074]	[0.033]	[0.005]
Fiscal policy effect	*FP*	-0.014	-0.042	-0.052	-0.052	-0.036	-0.036
		[0.744]	[0.212]	[0.025]	[0.019]	[0.123]	[0.447]
Sovereign risk 1: Indebtedness	*SR1*	0.272	-0.816	-0.631	-0.631	-0.139	-0.139
		[0.003]	[0.001]	[0.061]	[0.001]	[0.542]	[0.300]
Sovereign risk 2: Political stability	*SR2*	0.007	-0.018	-0.016	-0.016	-0.005	-0.005
		[0.167]	[0.185]	[0.250]	[0.047]	[0.690]	[0.610]
Trade openness	*TO*	-0.010	-0.007	-0.020	-0.020	-0.026	-0.026
		[0.450]	[0.811]	[0.248]	[0.520]	[0.207]	[0.434]
R_{adj}^2		0.105	0.599	0.683		0.111	
F-test on H_0: Pooling model > Fixed-effect model		11.663 [0.000]					
Hausman test on H_0: Random-effect model > Fixed-effect model			20.572 [0.015]				

Notes 1: The result of estimating Eq. (3): $L\_DGF_{i,\tau}$ (or $\_DGF^{LC}_{i,\tau}$) $= h_0C + h_1Size_{i,\tau} + h_2FXF_{i,\tau} + h_3CBM_{i,\tau-1} + h_4 MP_{i,\tau} + h_{IT}[CBM_{i,\tau-1} \times MP_{i,\tau}] + h_5FP_{i,\tau} + h_6SR1_{i,\tau} + h_7SR2_{i,\tau} + h_8TO_{i,\tau} + IE_i + \varepsilon_{i,\tau}$.

2: The number of observations is 287.

3: I follow the conventional procedure to specify the type of *IE*. I estimate the pooling model using the OLS method, and I estimate the fixed-effect model with the Least Squares Dummy Variables (LSDV) method. I justify the addition of constant *IE*s by checking with the F-test by how many and how significantly that addition reduces residual squared sums. If the fixed-effect model is selected, then, to compare it with the random-effect model, I test a null hypothesis with the Hausman test that *IE*s are uncorrelated with explanatory variables.

4: Shading indicates regressors with statistically significant estimators and a specification of *IE* with statistical adequacy. Finally, I select the fixed-effect model.

5: Values in [ ] are p-values.

6: Potential irregular aspects of residuals (ε) are cross-section heteroskedasticity (CSH), period heteroskedasticity (PH), serial correlation (SC), and contemporaneously correlated errors (CCE). I address these aspects by using the *EViews 10* statistical software package when I estimate the fixed-effect and random-effect models. *EViews 10*'s option for a panel-data regression, *White period*, is used to gain standard errors adjusted for the risks of PH and SC, with *White cross-section* used to gain those adjusted for CSH and CCE. In estimating the fixed-effect model by GLS, I additionally use the *Cross-section weights* option, which also makes it possible to control for the risk of CSH. Reed and Ye (2011) demonstrate that estimators gained using the weighted-GLS method together with each of the two options for adjusted standard errors are excellent in terms of the estimates' asymptotical efficiency and the accuracy of confidence intervals across them. The random effect estimators depend on the Swamy-Arora method which uses residuals gained in the within (fixed-effect) and between-means regressions.