On the Global Financial Market Integration "Swoosh" and the Trilemma

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Geert Bekaert[†] and Arnaud Mehl[*]

Abstract

We propose a measure of financial market integration based on a factor model of equity returns computed back to the first era of financial globalization for 17 countries. Global financial integration follows a "swoosh" shape – high pre-1913, higher post-1990, low in the interwar period – rather than other shapes hypothesized in earlier literature. We find no evidence of financial globalization reversing since the Great Recession, as claimed in other recent studies. We use our measure to revisit the debate on whether the classic monetary policy trilemma has recently morphed into a dilemma and find no evidence for such change.

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[†]Columbia Business School, NBER and CEPR; [*]European Central Bank. Email addresses: gb241@columbia.edu and arnaud.mehl@ecb.int. We would like to thank Mark Aguiar, Bernard Dumas (our discussant), Menzie Chinn (our editor), Georgios Georgiadis, Luca Dedola, Frank Moss, Livio Stracca, Jay Shambaugh, participants in the IMF 18th Jacques Polak Annual Research Conference on 2-3 November 2017 and in the JIMF-EU Commission conference on international financial integration on 1-2 March 2018 for helpful comments and suggestions. We are also grateful to Dennis Quinn for sharing his data on de jure financial openness pre-World War II. The views expressed in this paper are solely our own and do not necessarily reflect those of the European Central Bank or the Eurosystem.

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1. Introduction

The process of international financial integration is not a gentle climb towards ever higher peaks. This is true from both a short- and long-run perspective. Bekaert and Harvey (1995), focusing on the post-1990 financial globalization wave in emerging equity markets, find that de facto integration may exhibit reversals and does not always increase over time. Recent papers, focusing on banks, or bonds, but not on equity markets, show that financial globalization partly reversed in the wake of the Great Recession (see e.g. Rose and Wieladek (2014), van Rijckeghem and Weder (2014), Giannetti and Laeven (2012, 2016)). There is also a thriving literature documenting the recent surge in capital controls in emerging markets, along with their economic effects (see e.g. Jeanne and Korinek (2010), Ostry et al. (2012), Forbes, Fratzscher and Straub (2013), Pasricha et al. (2015), Ito and McCauley (2018)).

From a historical perspective, there is an old-standing debate among macroeconomists and economic historians as to whether international financial integration was, in fact, "stronger" pre-1913, a period also known as the first era of financial globalization, compared to the modern era of financial globalization, which started with capital account liberalizations in advanced economies in the 1980s and in emerging markets in the 1990s. Bordo and Flandreau (2003), Bordo and Murshid (2006) and Quinn (2003) deem the early period more globalized; Bordo, Eichengreen and Irwin (1999), Mauro et al. (2002), and Quinn and Voth (2008) claim the opposite is true. Bordo and Flandreau (2003), Obstfeld and Taylor (2003, 2004), and Goetzmann et al. (2005) argue that global financial integration follows a U-shape pattern with equal degrees of integration before 1914 and after 1970. Volosovych (2011), focusing on sovereign bond markets, claims that global financial integration is rather characterised by a J-shape pattern, with a trough in the 1920s. Rangvid et al. (2016) look at equity market integration over 1875-2012 and find that financial integration in the later part of their sample is "very high" relative to earlier periods.

The interest from macroeconomists in measuring international financial market integration over long time periods has been spurred by recent policy debates on the trilemma, the trade-offs between the exchange rate regime, financial openness and monetary policy autonomy.[1] In particular, several articles stress the critical role played by the US dollar and US monetary policy in setting global liquidity and credit conditions.[2] They suggest that non-US central banks have lost their ability to influence domestic long-term interest rates, even in the presence of flexible exchange rates, due to the existence of "US-driven" global financial cycles in liquidity and credit. As a result, the trilemma may have morphed into a dilemma between financial openness and monetary policy autonomy.

In this paper, we focus on equity market integration and propose a simple measure that can be computed back to the first era of financial globalization for 17 countries accounting for two-thirds of world GDP throughout the sample. The key strengths of our measure are that it describes integration at relatively high monthly frequencies; captures de facto, and not simply de jure, integration; and provides a framework to test formally for the various shapes of the temporal pattern of integration hypothesized in earlier literature. We can also use our measure to distinguish global from regional patterns of integration and to uncover the economic sources of financial integration, both at the global and regional level.

The measure employs conditional betas of a country's stock return with respect to global and regional equity market returns. While betas may be affected by both cash flow comovements and discount rates, they provide an economically meaningful measurement of the sensitivity of a country's equity market to global and regional shocks. Moreover, they do not suffer from the volatility bias plaguing simple correlation statistics, which arises because much of the time-

[1] See e.g. Shambaugh (2004), Obstfeld, Shambaugh, and Taylor (2005), Miniane and Rogers (2007), Bluedorn and Bowdler (2010), Klein and Shambaugh (2015), Aizenmann, Chinn and Ito (2014), Pasricha et al. (2015).
[2] See e.g. Rey (2013), Miranda-Agrippino and Rey (2014), Bruno and Shin (2015a, 2015b), Passari and Rey (2015) as well as Obstfeld (2015) for a discussion.

variation in correlations is accounted for by changes in factor volatilities.[3] While market integration has clear implications for first moments (see Bekaert, Harvey, 1995, for instance), most models admitting time-variation in discount rates and heteroskedasticity should imply that betas with respect to global factors increase when going from a segmented to an integrated equilibrium. This would be true e.g. in the dynamic reduced form pricing model described in Bekaert, Harvey, Lundblad and Siegel (2011), as long as the discount rate under segmentation of a particular country is not perfectly correlated with the global discount rate. Empirical studies focusing on liberalizations in emerging markets, such as Bekaert and Harvey (1997), European equity markets, such as Baele (2005) and American Depository Receipt introductions (a firm-specific liberalization), such as Lewis (2015), show jumps in betas around these events.

We use this measure to test several hypotheses. First, we assess which factors explain the time series and cross-country variation in de facto financial market integration over the long run. We find that de jure financial openness is a statistically significant determinant of de facto integration, while trade openness and financial development are not, which confirms the results of Bekaert et al. (2011) for the modern era of financial globalization. In terms of explained variation, however, we find that global growth uncertainty explains an equally important share of global financial market integration, while a third significant determinant, namely a variable measuring the incidence of high volatility across markets, explains only a minuscule share.

Second, we formally test whether the long-run temporal pattern of de facto financial market integration follows a flat line, a U shape, a J shape or even a “swoosh” shape (i.e. the trademark logo of a famous athletic shoe and clothing manufacturer). The exact shape of global financial integration is part of important discussions in economic history, but it also has potential implications for the recent policy debate regarding the classic Mundell-Flemming trilemma, as we show in greater detail below. For this debate, whether integration has continued to increase

[3] See e.g. Forbes and Rigobon (2002), Dungey et al. (2004), Bekaert, Harvey and Ng (2005), Bekaert et al. (2014).

since the global financial crisis (as would be the case under the “swoosh”) is a highly relevant input. In so doing, we distinguish explicitly between global and regional financial market integration patterns. We fail to reject the presence of a swoosh pattern for de facto global financial market integration, i.e. high pre-1913, still higher post-1990, low in the interwar period, but statistically reject the other shapes previously hypothesized. We do not find a clear regional financial market integration pattern.

Third, we use the measure to test whether the Great Recession has been associated with a reversal in the process of de facto financial globalization, as claimed by recent studies, and do not find evidence in support of this claim.

Fourth, we use our measure of de facto global financial market integration to revisit the debate on the existence of a monetary policy trilemma in history. We find evidence that pass-through from base country to domestic interest rates – at both short and long maturities – depends on whether an economy is open to global finance or closed, and on whether it has pegged or flexible exchange rates, in line with the trilemma hypothesis. For the recent period, the evidence also points on balance more toward the trilemma than the dilemma, even though it is difficult to conduct inference in an increasingly globalized world.

The paper is organized as follows. Section 2 presents the empirical framework which we use to measure global financial market integration over the long run and discusses how de facto integration evolves over time. While our integration measure uses equity market data, we argue later that the degree of market integration across different asset classes is highly correlated. Section 3 presents our formal test of the long-run temporal pattern of financial market integration and provides evidence consistent with a swoosh shape. Section 4 employs our measure to revisit the debate on the trilemma versus dilemma hypotheses. Section 5 concludes.

2. Measuring Global Financial Market Integration over the Long Run

This section outlines the model we estimate, elaborates on the concept of time-varying de facto financial market integration, and discusses how integration evolves over time.

2.1. The Factor Model

2.1.1. General Specification

We formulate an international factor model with two factors – a global market factor, and a regional market factor, $\mathbf{F_t}' = [F_t^{Glob}, F_t^{Reg}]$. The two factors are value-weighted market indices, so that the model potentially embeds different conditional CAPMs as special cases: a regional (world) CAPM when the beta on the first (regional) factor is zero. As in any factor model, the correlation between portfolios is increasing in the factor exposures of the portfolios and the magnitude of the factor volatilities. The use of these two factors ensures that the model satisfactorily fits comovements across countries.[4]

The full model is:

$$R_{i,t} = \alpha_i + \lambda_t + \beta_{i,t}^{glo} F_t^{glo,\backslash i} + \beta_{i,t}^{reg} F_t^{reg,\backslash i} + \boldsymbol{\gamma}^{\mathbf{glo}\prime} \mathbf{X}_{\mathbf{i,t-k}}^{\mathbf{glo}} + \boldsymbol{\gamma}^{\mathbf{reg}\prime} \mathbf{X}_{\mathbf{i,t-k}}^{\mathbf{reg}} + \varepsilon_{i,t} \tag{1}$$

$$\beta_{i,t}^{glo} = b_0^{glo} + \mathbf{b_1^{glo}}' \mathbf{X_{i,t-k}^{glo}} \tag{2}$$

$$\beta_{i,t}^{reg} = b_0^{reg} + \mathbf{b_1}^{reg}{}' \mathbf{X_{i,t-k}^{reg}} \ , \tag{3}$$

where $R_{i,t}$ is the excess return on the local equity index in country i during month t, expressed in dollars (i.e., the dollar equity return minus the 10-year U.S. Treasury yield in monthly units), α is a country fixed effect, λ is a year effect, F_t^{glo} is the global market factor, F_t^{reg} is the regional market factor and $\mathbf{X}$ is a vector of control variables designed to capture time and cross-sectional

[4] The analysis in Bekaert, Hodrick, and Zhang (2009), Bodnar, Dumas, and Marston (2003), and Brooks and Del Negro (2006) motivates the use of both global/international and domestic factors from a statistical perspective, even for developed markets. Rangvid et al. (2016) use the cross-country dispersion of stock returns as their main measure of global financial market integration but they also calculate a measure based on a world-CAPM in robustness checks.

variation in factor exposures.[5] These variables are country-specific, typically lagged by one year, and also enter the conditional mean. If the dimension of **X** is k, the vectors $\mathbf{b_1}^{\mathbf{l}}$, γ^{l}, where ***l*=glo, reg**, are $k \times 1$. The sample period is January 1885 to June 2014. It contains up to 1,554 monthly observations for each of 17 country-equity portfolios, which are split into three regions: Europe (Austria, Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Spain, Sweden, Switzerland, and the U.K.); Northern America (the U.S. and Canada); and Asia-Pacific (Australia and Japan). While the sample is dominated by advanced, typically European countries, they account on average for about two-thirds of global GDP over the sample period, but their dominance has decreased over time.

To avoid adding-up constraints and spurious correlations, the factors are value-weighted across countries, but exclude returns of country i itself. To facilitate the interpretation of the factor loadings, representing global and regional patterns of market integration, we orthogonalize the two factors as in, for example, Bekaert et al. (2014). The regional market factor is the residual of a regression of regional market returns on global market returns over the full sample period; the regression is estimated for each country individually as country i itself is excluded from the market factors.

While financial integration should increase stock return correlations and betas, one should ideally control for economic fundamentals, including cash flows, as stressed by Dumas, Harvey and Ruiz (2003). However, in our sample it is impossible to correct for these fundamentals as cash flow data going back to the early 1900s are not available. If it were true that most of the time variation in equity market comovements arises from discount rate variation, this assumption would be relatively innocuous. Empirically, Xu (2017) claims this to be true for a recent post-1995 sample on G7 countries, showing that the time-variation in comovements across G7 equity market returns is primarily driven by global risk aversion shocks. We implicitly assume that this

[5] Expressing equity returns in dollars is common and provides the perspective of a US investor but when returns are expressed in local currency the results remain unchanged. This is not surprising because the variance of equity returns dominates the variance of exchange rate returns (by a factor of about 4).

is true for our sample period too. Finally, the country and time fixed effects in the model help absorb variation not explained by the factor model, but empirically contribute little to explain equity return variation.

In a related vein, our monthly equity data from Global Financial Data do not systematically adjust for dividends. This data issue justifies our focus on second moments, as it prevents us from investigating expected returns. Dividend yields are an important part of the expected return and would be particularly important in measuring integration jumps (see the intuition described in Bekaert and Harvey, 2000). However, the online appendix shows that our price returns correlate highly at the annual frequency with the annual dividend-adjusted data from Jordà, Schularick, and Taylor (2017).

Table 1 contains an overview of the data and selected descriptive statistics, with further details on data sources and construction relegated to the online appendix.

2.1.2. Financial vs. Equity Market Integration

The use of equity market data to measure financial integration raises a number of issues. First, a common view is that financial globalization rested overwhelmingly on debt and foreign direct investment flows prior to 1914. However, recent studies show that cross-border equity flows were not negligible in this earlier era.[6]

Second, the degree of integration may differ across different asset classes. However, in many cases capital account restrictions tend to apply to a broad range of asset classes simultaneously. For the modern era, there is concrete evidence that capital flow restrictions in bond and equity markets tend to go hand in hand. In the online appendix, we show strong comovements between the average values of an index of restrictions on cross-border flows in equities and on bonds, respectively, across the 16 countries of our sample and the U.S., drawing

[6] For instance, Esteves (2006, 2011) estimates that about 30% (15%) of the U.K.'s (Germany's) capital exports were in the form of shares between 1883 and 1913 and that more foreign equity shares than foreign bonds were issued in Paris between 1880 and 1913. In addition, Van Hombeeck (2017) shows that 18% of the foreign securities traded on British exchanges were foreign equity shares.

from the data set of Fernández et al. (2015) which extend from 1995 to 2015. The panel correlation coefficient between the index values for the two asset classes is 0.75. This suggests that countries tend to increase or reduce openness to bond and equity flows simultaneously. Finally, Klein (2012) shows that an economy must fully close its financial account (i.e. erect "walls" as opposed to "gates") to be insulated from global financial influences. In all, the degree of equity integration is likely informative about financial market integration in a broader sense.

Third, the measure is price- not flow-based. Whereas flow-based measures of integration are common in international economics, shocks can be transmitted across borders without capital flows occurring (see also Bekaert et al. (2014)).

2.1.3. Instruments to Model Cross-Sectional and Time Variation in Exposures

Equations (1) to (3) contain a set of lagged instruments, $\mathbf{X}_{\mathbf{i,t\text{-}k}}$, which are used to model the cross-sectional and time variation in the factor loadings $\beta_{i,t}^{glo}$ and $\beta_{i,t}^{reg}$. This practice has a long tradition in finance; see, for example, Ferson and Harvey (1991) and Dumas and Solnik (1995).[7] We entertain seven potential instruments, which are listed in Table 1, to distinguish between different channels and hypotheses regarding the sources of the variation in global and regional factor exposures. When observations are annual they are kept constant over the ful year at their annual values.

Several studies have suggested that equity return comovements increase with financial and economic integration (see e.g. Mendoza and Quadrini (2010), Brière, Chapelle, and Szafarz (2012), Fratzscher (2012)). The trade channel in particular has often been associated with international spillovers and, in some cases, contagion (see e.g. Kaminsky and Reinhart (2000), Forbes (2004), Caramazza, Ricci and Salgano (2004), and Baele and Inghelbrecht (2009)). Hence, we use trade openness, measured as exports plus imports scaled by GDP in country i and

[7] These "instruments" are not "exogenous" in the strict sense of econometric identification but indicate variables that are not returns, are pre-determined (in a temporal sense) and are used to model time-variation in factor exposures or prices of risk. Also, the instruments are too slow-moving to reflect public information that may instantaneously change prices and potentially cause contagion (see Connolly and Wang (2003)).

year *t*, as a first potential determinant of the cross-sectional and time variation in factor exposures. Another potential determinant, specific to regional financial integration, is regional trade openness, which is defined as the sum of country *i*'s exports and imports of goods to/from the countries belonging to the country *i*'s region, scaled by total trade in year *t*.

A third potential determinant is de jure capital account openness, a natural determinant of de facto financial integration (see Kose et al. (2006), Bekaert et al. (2011)). We use the indices of capital account openness assembled by Quinn and Voth (2008) and Quinn and Toyoda (2008), which measure the extent of restrictions to capital outflows and inflows by residents and nonresidents in country *i* and year *t*.

The fourth instrument is domestic financial development. Several researchers have stressed that poorly developed financial systems may impair financial integration (see Bekaert and Harvey, (1995); Bekaert et al. (2011)). Equity market illiquidity may prevent foreign institutional investors from investing in emerging markets according to some surveys (see e.g. Chuhan (1994)). Poor liquidity as a priced local factor may also lead to valuation differentials and different betas relative to global benchmarks (see Acharya and Pedersen (2005) or Bekaert, Harvey and Lundblad (2007) for models incorporating liquidity risks). The metric of financial development we use is the ratio of equity market capitalization to output, inferred from Rajan and Zingales (2003) and Beck, Demirgüç-Kunt, and Levine (2010).

Factor exposures may also vary over time with global shocks, such as oil and other commodity price shocks or shifts in global risk aversion. Given our data limitations, we consider only two variables. The first one is a measure of global oil price spikes, defined as the deviation (in logarithms) between the current oil price and its five-year moving average. Hamilton (2005) shows that 9 out of 10 U.S. recessions since World War II were preceded by a sudden increase in oil prices. A global recessionary shock induced by changes in oil prices may increase global factor exposures. Because the likely channel is through increased cash flow comovement, this potential time variation should not be interpreted as higher financial market integration. If recessions increase (global) risk aversion, the opposite effect may result as investors flee from foreign equity markets considered as risky towards domestic equity markets or other financial assets considered as safe, leading to a divergence in valuations and increased segmentation (see the discussion in Bekaert et al. (2011), for instance). In most models, high risk aversion increases the volatility of stock returns (see e.g. Bekaert, Engstrom and Xing (2009)) so we measure risk aversion indirectly through volatility. Specifically, we first estimate the conditional volatility of stock returns for each country of our sample using a GARCH(1,1) model. We then normalise the conditional volatilities of each country's stock returns and define a high market volatility variable as the proportion of the 17 country-specific volatilities in excess of 1.65 in a given month. This yields a global "volatility spike" time series with monthly observations over January 1885-June 2014. Note that high return volatility itself may lead to higher return correlations not associated with financial integration, which is captured in our model through the factor volatilities and does not affect our integration measure (see below for further discussion).

The last potential determinant of the time variation in factor loadings is uncertainty in earnings growth, which is another possible source of financial market segmentation. For instance, in a pricing model with stochastic growth opportunities and discount rates, Bekaert et al. (2011) show that under a strong notion of integration, encompassing both financial and economic integration, the time-varying components of industry price-to-earnings ratios are identical across countries, and are determined entirely by variation in the world discount rate and world growth opportunities. However, even under the null hypothesis of full financial and economic integration, industry earnings yield differentials between a country and the world market can still arise because of differences in earnings growth volatility. Because harmonised and consistent data on earnings growth are not directly available for our century-long panel, we measure the conditional volatility of real GDP growth instead, using three different measures. The first two employ the logarithm of the standard deviation of real GDP growth in each country over periods of 5 years, which yields 17 country-specific times series of annual observations on local growth uncertainty.[8] In the first measure, we use non-overlapping periods keeping the measure constant over 5-year intervals; in the second we use overlapping windows of 5 years centred around the current observation (with one year increments). An alternative metric obviates the need for time series observations by employing the natural logarithm of the cross-sectional dispersion of real GDP growth for the 17 countries of our sample in a given year. This yields a global time series with annual observations over 1885-2014. The cross-sectional variance can be decomposed into an estimate of the country-specific variance (the average country-specific volatility minus the "world" variance) and an estimate of the variance of the country averages (see Bekaert, Harvey, Kiguel and Wang (2016)). The first component is correlated with the times series uncertainty of growth opportunities worldwide; the second component with the divergence of growth opportunities across countries at a given point in time. Increases in both components of this global growth uncertainty measure would tend to decrease de facto integration.

2.2. Model Estimation and Measuring Time-Varying Financial Market Integration

2.2.1. Model Estimation and Parameter Estimates

The panel model is estimated using pooled OLS, with standard errors accounting for heteroskedasticity and clustered by country. We lag the instruments, $\mathbf{X}_{i,t-k}$, by one year with the exception of the trend deviation of oil prices and the high volatility variable, available at the monthly frequency, which are lagged by one month instead. All the estimates control for country fixed effects, year effects and for the direct effects of the instruments included in vector $\mathbf{X}$ (whose coefficients are not reported to save space).

[8] The rationale for using logs rather than levels is that the distribution of real GDP growth is heavily fat-tailed because of two observations in 1945 and 1946, when output collapsed (or jumped from an extremely low base) in several countries in the wake of the end of World War II and the move to a postwar economy.

Using the parameter estimates we define a benchmark, $\bar{\beta}^{glo}$, for global market integration as the (weighted) average across countries and time of the $\beta_{i,t}^{glo}$ estimates, i.e.

$$\bar{\beta}^{glo} = \frac{1}{N}\frac{1}{T}\sum_{i=1}^{N}\sum_{t=1}^{T} w_{i,t}\beta_{i,t}^{glo} \ , \tag{4}$$

where N = 17, T = 1,554; $w_{i,t}$ denotes the market capitalisation of country i at time t. The relative global market integration of country i at time t then is defined as $\beta_{i,t}^{glo} / \bar{\beta}^{glo}$. The benchmark $\bar{\beta}^{reg}$ for regional market integration is defined analogously, with $\beta_{i,t}^{reg} / \bar{\beta}^{reg}$ indicating the relative regional market integration of country i at time t.

Table 2 reports the conditional global and regional beta estimates, with each instrument included individually in the estimates reported in columns 2 to 7, while all instruments are included in column 8. The regression only includes 7 instruments, using only global growth uncertainty among the three uncertainty measures. The local growth uncertainty measures are neither significant in univariate nor multivariate specifications. The three uncertainty measures are also the only ones displaying pairwise correlations larger than 40% in absolute value (see the online appendix).

We obtain a parsimonious model in column 9 using model selection techniques (see, for instance, Hendry and Krolzig (2005)) to pare down the regression to a more manageable number of independent variables. Starting from the full model including all instruments, we reduce the model step-by-step by excluding the interaction variables with insignificant parameters. If all interaction effects are insignificant, the variable is dropped from the regression. Convergence was reached in two steps.

In column (1), we report a specification without instruments; the global factor beta is 0.68 and the regional factor beta is 0.29, both significantly different from zero. De jure capital account openness exerts a positive and statistically significant effect on global betas, an effect that is preserved in the multivariate specifications. Trade openness and financial development are statistically significant determinants of global betas individually, but not in the multivariate specification, which confirms earlier results in Bekaert et al. (2011) for the modern era of financial globalization. The global oil price variable is statistically significant individually, but it is not in the multivariate specification. Higher uncertainty in real GDP growth reduces global betas significantly, in line with the model predictions of Bekaert at al. (2011). Global betas increase significantly in periods of heightened market volatility, but the economic magnitude of the effect is economically very small (more on this below). Finally, while there are some significant univariate results, only de jure capital account openness exerts a statistically significant – and positive – effect on regional betas, both in the univariate and multivariate specifications. Therefore, the final specification reported in column (9) contains capital openness (for both regional and global betas), and growth uncertainty and the market volatility variable (only for global betas).

2.2.2. Understanding the Variation in Financial Market Integration

To examine the relative economic importance of the determinants of global financial market integration, we conduct a variance ratio analysis. For each of the three statistically significant instruments *j* (i.e. de capital account openness, global growth uncertainty and high market volatility periods) of the parsimonious specification, we calculate the variance ratio for the conditional global beta estimates as

$$VR^{j} = \frac{\operatorname{cov}[\hat{\mathbf{b}}_{\mathbf{1}}^{\mathbf{glo}\prime}\mathbf{X}_{\mathbf{i,t-k}}^{\mathbf{glo}}, \hat{b}_{1,j}^{glo} X_{i,t-k}^{glo}]}{\operatorname{var}[\hat{\mathbf{b}}_{\mathbf{1}}^{\mathbf{glo}\prime}\mathbf{X}_{\mathbf{i,t-k}}^{\mathbf{glo}}]} \tag{5}$$

By definition, these variance ratios sum to one. De jure capital account openness explains 53% of global financial market integration, against 47% for global growth uncertainty. The proportions are statistically significantly different from zero but we cannot reject that they are equal (with the standard errors computed using a bootstrap with 1,000 replications). In contrast, high market volatility periods explain a negligible part of global financial market integration, which is statistically insignificant. As for regional equity returns, recall that their predictable variation is fully explained by de jure capital account openness.

The temporal pattern in financial integration is therefore completely driven by the de jure integration and global growth uncertainty. Figure 1 (Panel A) shows the evolution between 1885 and 2014 of the unweighted (thick grey lines) and value-weighted (light grey lines) cross-country averages of the measures of financial market integration along with the corresponding conditional beta estimates.

The temporal pattern of global financial market integration clearly follows a swoosh shape. During the first era of financial globalization, de facto global financial market integration was close to its century-long average. It then decreased significantly in the wake of World War I, but recovered temporarily until the early 1930s. A nadir was reached immediately after World War II, when de facto global financial market integration stood at roughly 90% below its century-long average. Since the 1950s, de facto global financial market integration has increased steadily. However, it exceeded pre-1913 levels only after 1990. De facto global financial market integration has remained at historically high levels since the global financial crisis broke out in 2007, at about 30% above its century-long average in 2014, notwithstanding the capital controls and other financial protectionist measures recently taken in some countries. The online appendix shows that the temporal pattern of global financial market integration is nicely swoosh-shaped in all countries. The temporal pattern of regional financial market integration, relegated to the online appendix, is less clear, being in between a swoosh and a U-shaped pattern.

2.2.3. Model Validation

As stressed e.g. by Cochrane (2001), a challenge to our conditional factor model is that it requires the econometrician to know the ‘‘true’’ state variables. Lewellen and Nagel (2006) propose to estimate the factor regressions over a short window – using no conditioning variables – providing direct estimates of conditional betas. To implement their approach, we split the sample in 5-year non-overlapping periods and compute the betas over these 5 years. For each starting point of a 5 year period, the beta is set equal to that rolling beta; for periods in-between the beta is a linearly interpolated number between the current and next beta. The choice of a forward window is consistent with the idea that our factor model produces conditional betas. A well specified factor model should then produce beta estimates that are insignificantly different from the rolling beta estimates.

Figure 1 (Panel B) shows the evolution between 1885 and 2014 of the conditional global betas (shown as thick grey lines) together with 90% confidence bands obtained from the corresponding pooled rolling beta estimates (shown as light grey lines) and the point estimates (shown as black dashed lines). The simple rolling global beta estimates also follow a swoosh shape. Moreover, the conditional betas fall mostly well within the confidence bands of the simple rolling beta estimates. The conditional global betas fall within the bands 81% of the time. When conditional betas are outside the bands, they tend to be quite close to them. The conditional betas overestimate the extent of global financial market integration relative to what rolling betas would predict during World War I a bit, which might suggest that the conflict led to a reversal in financial globalization that was partially unexpected, but they do a good job during World War II. From the early 2000s, the conditional betas systematically underestimate the extent of global financial market integration relative to what rolling betas would predict, an issue we revisit below. The online appendix shows that country-specific conditional betas mostly fall well within the confidence bands of the simple rolling beta estimates. Similar patterns emerge for regional betas (see online appendix).

A simpler measure to quantify de facto integration is the average correlation between equity markets (see Quinn and Voth (2008)). However, correlations suffer from the volatility bias described in the seminal work of Forbes and Rigobon (2002). As volatilities tend to dramatically increase during crises, increased correlations are not necessarily indicative of higher interdependence between equity markets. Under the null of our model, the comovement between equity markets is determined by the factor exposures (the betas) and the variance-covariance matrix of the factors. Such a model can potentially fit the observed increase in correlations during a crisis through an increase in factor volatilities, while betas – the true measure of interdependence – remain stable. Assuming uncorrelated factors, this is true because the correlation between a particular equity market and a factor is then the beta with respect to that factor, times the ratio of factor to equity market volatility, which can be shown to be increasing in the factor's volatility (see also the discussion in Bekaert et al. (2014) for further details). This is of particular importance during the global financial crisis of 2007-2008 when volatility reached exceptionally high levels, which could have biased correlations in international equity markets upwards. We relegate a full analysis of how our beta measures and a correlation measure compare to the online appendix, but note that over the recent globalization period (2001-2014) the comovement between the beta and correlation measures is close to zero.

3. The Swoosh in Financial Market Integration

This section sets out to formally test the swoosh pattern in de facto financial market integration apparent in our full model estimates.

3.1. Testing for a Swoosh Pattern

We estimate the following simple variant of our two-factor model in Equations (1) – (3):

$$R_{i,t} = \alpha_i + \lambda_t + (\boldsymbol{\beta}_{\mathbf{j}}^{\mathbf{glo}}{}'\mathbf{D}_{\mathbf{j}})\, F_t^{glo,\backslash i} + (\boldsymbol{\beta}_{\mathbf{j}}^{\mathbf{reg}}{}'\mathbf{D}_{\mathbf{j}})\, F_t^{reg,\backslash i} + \varepsilon_{i,t} \tag{6}$$

where j =1, 2, 3 and D_j denotes a dummy variable which equals one over time period j and zero otherwise. The first period is 1885-1913, which is often referred to as the first era of financial globalization; see e.g. Bordo, Cavallo, and Meissner (2010); the second period is 1914-1990, which includes the interwar period (when several countries adopted protectionist and capital control measures in the run-up to World War II), and the Bretton Woods period (when capital controls, albeit possibly leaky, were still prevalent), and its immediate aftermath. The third subperiod is 1990-2014, which is often referred to as the second era of financial globalization, despite the alleged reversal since the Great Recession.

With this model, we can formally test whether the temporal pattern of de facto financial integration follows a U shape, as hypothesised by e.g. Obstfeld and Taylor (2003, 2004), among others; a J (or L-inverted) shape, as argued by Volosovych (2011); or a swoosh shape, as we posit. All tests can formally distinguish between global and regional financial integration patterns. Considering again the three aforementioned subperiods (i.e. pre-1913, 1914-1990 and 1990-2014) and three dummy variables D_j, j=1, 2, 3, we use Wald tests for the (in)equality restrictions: $H_0 : \beta_1^f = \beta_3^f, \beta_1^f > \beta_2^f, \beta_3^f > \beta_2^f$ for the U shape hypothesis,

$H_0 : \beta_1^f = \beta_2^f, \beta_3^f > \beta_1^f, \beta_3^f > \beta_2^f$ for the J (or L-inverted) shape hypothesis, and

$H_0 : \beta_1^f > \beta_2^f, \beta_3^f > \beta_1^f, \beta_3^f > \beta_2^f$ for the swoosh shape hypothesis, with $f = glo, reg$.

3.2. Empirical Results

Table 3 reports the estimates for the model in equation (6), obtained by pooled OLS, and the corresponding Wald tests. Panel A reports the parameter estimates. The *R*-squared of the regression is on the order of 20%-30%, which is in the same ballpark as the goodness of fit of the analogous models estimated by e.g. Bekaert et al. (2014) on a post-1995 sample including emerging markets. The time variation in betas confirm the swoosh pattern we detected with our full model. The estimate for the unconditional global beta prior to 1913 is 0.67, which is close to the full sample estimate. For the period 1914 and 1990, the global beta drops to 0.54, which is consistent with the adoption by several countries of protectionist and capital control measures in between the two World Wars and the slow road toward integration afterwards. It is only after 1990 that global financial market integration exceeds pre-1913 levels, with the global beta estimate reaching 0.94 for the 1990-2014 period. Regional betas decrease from 0.36 in the pre-war period to 0.21 in the 1914-1990 period, but then increase again to 0.58 in the post-1990 period. Parenthetically, there is not only substantial heterogeneity in betas over time, but also across countries, which is further illustrated in the online appendix. For example, global betas are as low as 0.4 in Austria and Japan, and as high as over 0.8 in the Netherlands, Germany and Canada.

The online appendix also shows the results to be very robust. First, the results are robust to the presence of fixed effects and outliers.[10] Second, the results remain robust when the factors are GDP – rather than value-weighted. Third, the estimates are essentially unchanged when we use sterling returns and Britain's long-term interest rate as the risk-free rate prior to 1914, i.e. when sterling was the leading international currency and the Bank of England "conducting the international orchestra" (see e.g. Eichengreen (1987)). Finally, using equity returns in local currency and long-term interest rates as risk free rates over the full sample delivers estimates essentially unchanged relative to the baseline specification.

In Panels B and C of Table 3, we test the various patterns of time variation in the degree of financial market integration. First, a Wald test overwhelmingly rejects the null of equality of the global and regional beta coefficients over the three subperiods. Second, we reject the null hypotheses that de facto global financial market integration over the last century is characterised by a U-shape process, or by a J-shape process. However, we cannot reject the hypothesis that the temporal pattern follows a swoosh shape. That is, de facto global financial market integration was high in the first era of financial globalization before World War I, but not as high as during the second era after 1990. Still, de facto global financial market integration in both eras was substantially stronger than between 1914 and 1990.

The results for de facto regional financial market integration are different, however. While the coefficient pattern is numerically consistent with a "swoosh," we fail to reject either the U, J or swoosh shapes.

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[10] The country and time effects are used to capture deviations from our simple multi-factor pricing model and to soak up variation not explained by the model. As Table D2 in the online appendix shows these fixed effects almost contribute nothing (about 1-2%) to the return variation we are trying to explain, and the key results remain the same when dropping them. This also shows that the dependence on the world and regional market returns soaks up most of the common variation in country stock returns.

In the face of the capital controls and other financial protectionist measures taken by advanced and emerging market economies (see, e.g., Ostry et al. (2012), and many others) including public interventions in the financial sector of advanced economies (as stressed e.g. by Rose and Wieladek (2014)), effective financial market integration may have partly reversed since the onset of the global financial crisis in 2007. In Table 4 we present estimates of model equation (6) with two period dummies (1990-2007 and 2007-2014).[11] Wald tests overwhelmingly reject the hypothesis that the global betas in the two subperiods are equal. They also reject the hypothesis that the pre-crisis beta is higher than the post-crisis beta. The converse hypothesis is not rejected. This evidence suggests that the process of de facto global financial market integration has not reversed since the Great Recession, despite claims made in recent studies. Because the subperiod 2007-2009 coincided with the acute phase of the global financial crisis, it may be contaminated by contagion effects (see Bekaert et al. (2014)). We therefore obtained estimates of model equation (6) with three period dummies (1990-2006, 2007-2009 and 2010-2014). The hypothesis that the global betas in subperiods 2007-2009 and 2010-2014 are equal is not rejected, however, while the hypothesis that the global beta in subperiod 1990-2006 is larger than the global beta in subperiod 2010-2014 is rejected (see the estimates reported in the online appendix).

4. Revisiting the Monetary Policy Trilemma in History

We now use our benchmark measure of de facto financial market integration over 130 years to revisit the debate on the monetary policy trilemma in history. Standard macroeconomic theory posits that an economy can have at most two out of an open capital account, a fixed exchange rate and an independent monetary policy. Specifically, if capital is allowed to move freely across borders, domestic interest rates can deviate from interest rates abroad only if the exchange rate is flexible. Alternatively, if policy-makers seek to stabilise the exchange rate under

[11] The specification also includes pre-1914 and 1914-1990 period dummies, which are not shown in the table to save space.

free capital mobility, domestic interest rates have to shadow foreign interest rates. This is the classic Mundell-Flemming's "trilemma" or "impossible trinity".

Early empirical tests of the trilemma suggest that it describes reasonably well the trade-offs between international capital mobility, the choice of the exchange rate regime and monetary policy autonomy over the last century or so (e.g. Obstfeld, Shambaugh and Taylor (2005)). More recently, however, Rey (2013) and co-authors argue that the classic trilemma has morphed into a "dilemma" and the impossible trinity into an "irreconcilable duo". Central banks outside the U.S., the world's foremost financial centre, have lost their ability to influence domestic long-term interest rates, even in the presence of flexible exchange rates, due to the existence of global financial cycles that are set in motion by US monetary policy shocks.[12] There is related evidence that bank leverage cycles are key determinants of the global transmission of US financial conditions across borders through banking sector capital flows (Bruno and Shin (2015a)) and that spillovers between US monetary policy, cross-border capital flows, and the US dollar exchange rate through the banking sector are substantial (Bruno and Shin (2015b)). This debate rekindled empirical work on the trilemma as we indicated earlier.

4.1. Testing the Trilemma Hypothesis

We start from the regression in Obstfeld, Shambaugh and Taylor

$$\Delta R_{i,t} = a_0 + b\Delta R^{base}_{i,t} + u_{i,t} \ , \tag{7}$$

where $R_{i,t}$ is the domestic interest rate at time t, $R^{base}{}_{i,t}$ is the base interest rate at time t in the anchor country; and Δ is the difference operator. Under full capital mobility and a credible peg, it is expected that $b = 1$, i.e. domestic and base-country interest rates move one for one, which implies that monetary policy in the pegging country is fully dependent on monetary policy in the base country. In contrast, $b = 0$ implies full independence from monetary policy in the base

[12] See also Miranda-Agrippino and Rey (2014), and Passari and Rey (2015)); Farhi and Werning (2014) study a small open economy model in which, in contrast with the Mundellian view, capital controls are desirable even when the exchange rate is flexible as they help to lean against the wind and smooth out capital flows.

country, which is to be expected if the exchange rate is floating, or if capital does not move freely across borders.

First, we seek to replicate Obstfeld, Shambaugh and Taylor's results by estimating equation (7) using similar yearly averages of monthly data, similar time periods and similar country groups as they have. Specifically, the time periods and country groups we consider for the replications are: 1885-1914 (gold standard), 1959-1970 (Bretton Woods), and 1973-2000 (post-Bretton Woods), nonfloaters vs. floaters, countries highly integrated into global finance vs. countries highly segmented from global finance. Floaters are defined as countries having floating or freely falling exchange rates according to the updated classification of Ilzetzki, Reinhart, and Rogoff (2004). Nonfloaters are defined as the remaining countries, including those which were on the classical gold standard or gold exchange standard prior to World War II according to the classification of Reinhart and Rogoff (2011). Countries with high (low, respectively) global financial market integration in period t are defined as those for which $\overline{\beta}_{i,t}^{glo} > 1$ ($\overline{\beta}_{i,t}^{glo} \leq 1$, respectively). Because the trilemma hypothesis concerns general capital mobility, it is important to repeat that equity and bond market integration are highly correlated so that our integration measure remains relevant. We take the U.K. as the base country prior to 1914 (classical gold standard); the average of the U.K. and the U.S. as the base for the 1920s (gold exchange standard), in line with the sterling-dollar duopoly of this earlier era (see e.g. Eichengreen and Flandreau (2009)); the U.K., U.S. and Germany as base countries for the sterling bloc, U.S. dollar bloc and Reichsmark bloc, respectively, for the 1930s; the U.S. as the base country for the Bretton Woods period (1959-1970); Germany as the base country for European countries (in the European Monetary System) and the U.S. for the remaining countries, respectively, for the period 1973-1999; and the U.S. for all countries for the period 1999-2014.[13] As measures of interest

[13] Our base countries are the same as those of Obstfeld, Shambaugh and Taylor for the pre-1914 period, Bretton Woods period and 1973-1999 period (they did not consider the interwar period and the period 1999-2014). That Germany is considered as the base country for European countries for the period 1973-1999 is motivated by the "German dominance hypothesis" (see e.g. Giavazzi et al. (1986); Giavazzi and Pagano (1988); von Hagen and Frattiani (1990)). Germany's monetary policy was so central

rates we use both short-term policy rates and long-term government bond yields. Because of the introduction of the euro and the European Central Bank in 1999, we use Germany as the representative country for the euro area post-1999 in the short-term interest rate regressions. See the online appendix for details on the data.

Next we modify equation model (7) to the form

$$\begin{aligned}\Delta R_{i,t} &= a_0 + b_1 \Delta R_{i,t}^{base} + b_2 nonfloater_{i,t} + b_3 \overline{\beta}_{i,t}^{glo} \\ &+ c_1(\Delta R_{i,t}^{base} \times nonfloater_{i,t}) + c_2(\Delta R_{i,t}^{base} \times \overline{\beta}_{i,t}^{glo}) + c_3(nonfloater_{i,t} \times \overline{\beta}_{i,t}^{glo}) \;, \\ &+ d_1(\Delta R_{i,t}^{base} \times nonfloater_{i,t} \times \overline{\beta}_{i,t}^{glo}) + u_{i,t}\end{aligned} \tag{8}$$

to test explicitly for the monetary policy trilemma hypothesis conditionally on the exchange rate regime and the degree of de facto global financial market integration, where *nonfloater* is a binary dummy variable which equals one if country *i* in period *t* is a nonfloater and zero otherwise. In so doing, we extend Obstfeld, Shambaugh and Taylor's study in three ways. First, we investigate both short-term and long-term interest rates whereas Obstfeld, Shambaugh and Taylor focused on short-term interest rates. Second, we extend the time dimension of their sample, insofar as we consider almost half a century of additional data by looking at the periods 1914-1945 and 2001-2014 as well. Third, in terms of data measurement, we employ a measure of de facto global financial market integration over the full sample period.[14] From equation (8) one can derive the elasticity of the domestic interest rate with respect to the base interest rate for both nonfloaters and floaters, respectively, conditional on the degree of global financial market integration, namely:

in the E.M.S. that many European countries simply shadowed the Bundesbank's interest rates. In fact, one reason why some countries pushed for the creation of the euro was to end the dominance of Germany's monetary policy by sharing Germany's influence and credibility through the single currency.

[14] Obstfeld, Shambaugh and Taylor (2005) rely on more limited data regarding capital market openness and assume that all countries had open capital markets during the gold standard era, that none did during Bretton Woods, and that the official I.M.F. coding from the *Exchange Rate Arrangements* yearbooks is a reasonable approximation for measuring the use of capital controls during the post-Bretton Woods era. Our metric of global financial market integration provides a direct measure for 17 countries over the full sample period. The cross-sectional dimension of Obstfeld, Shambaugh and Taylor (2005)'s panel is larger than ours for the Bretton Woods and post-Bretton Woods periods, however. As a result our sample has more closed non-floaters in general because of the presence of the 1914-1959 period.

$$\frac{\partial \Delta R_{i,t}}{\partial \Delta R_{i,t}^{base}} = b_1 + (c_2 \times \overline{\beta}_{i,t}^{glo}) \text{ if } nonfloater_{i,t} = 0 \text{, and}$$

$$\frac{\partial \Delta R_{i,t}}{\partial \Delta R_{i,t}^{base}} = (b_1 + c_1) + \left[(c_2 + d_1) \times \overline{\beta}_{i,t}^{glo}\right] \text{ if } nonfloater_{i,t} = 1 .$$

When the exchange rate is fully floating, capital controls are not a necessary condition for monetary policy independence. That is, $b_1 = 0$ no matter what the degree of financial market integration is. However, it is conceivable that increased capital market integration increases international interest rate dependence (that is $c_2 > 0$). A nonfloating exchange rate should only constrain monetary policy independence when markets are integrated, so the sign of c_1 is not necessarily clear ex-ante. If $\overline{\beta}_{i,t}^{glo} = 0$ represents fully binding capital controls, then c_1 may in fact be zero and pass-through only increases when $\overline{\beta}_{i,t}^{glo}$ increases and capital is more mobile. In any case, pass-through should increase with more openness, that is we expect $d_1 > 0$.

4.2. Empirical Results

4.2.1. Unconditional estimates

As a starting point, Table 5 reports the estimates of the parameters of the unconditional model in equation (7) using short-term policy interest rates, in the spirit of Obstfeld, Shambaugh and Taylor (2005). Column 1 reports pooled estimates, Columns 2 to 4 report estimates over three subperiods, namely: the classical gold standard era (pre-1914); the Bretton Woods regime (1959-1970); and the post-Bretton Woods era (1973-2000) (these subperiods match those of Obstfeld, Shambaugh and Taylor). Columns 5 to 8 report estimates by country groups (nonfloaters vs. floaters; high vs. low global financial market integration). Our estimate of b, the degree of pass-through from base country to domestic policy interest rates, for the full sample is about 0.30.[15] Our estimates by subperiod are qualitatively consistent with those of Obstfeld, Shambaugh and Taylor. Interest rate pass-through is found to be higher during the classical gold

[15] This is remarkably close to Hofmann and Takáts (2015)'s estimate of 0.34 for a panel of 30 emerging market and advanced economies over the period 2000-2014.

standard and post-Bretton Woods era, with estimates for the coefficient *b* of 0.19 and 0.56, respectively, than during the Bretton Woods regime, with a *b*-estimate of 0.10. Obstfeld, Shambaugh and Taylor's estimates are 0.42 (gold standard), -0.20 (Bretton Woods) and 0.36 (post-Bretton Woods), respectively. The country group estimates suggest that there are no discernible differences in the extent of interest rate pass-through between nonfloaters and floaters, insofar as the *b*-estimate, at about 0.30, is virtually identical for both groups of countries. This is unlike Obstfeld, Shambaugh and Taylor's estimates, which point to differences between nonfloaters (0.43) and floaters (0.26). Of course, this result may reflect differences in the extent of de facto global financial market integration between and within groups. Indeed, when the sample is restricted to countries highly integrated into global finance, the *b*-coefficient is close to 0.40, but when it is restricted to countries segmented from global finance, it drops to 0.14 only, a finding consistent with that of Obstfeld, Shambaugh and Taylor whose estimates are 0.56 (no capital controls) and 0.26 (capital controls).

In the online appendix, we show that the results for long-term interest rates are in line with those for short-term policy rates.

4.2.2. Conditional estimates

How does interest rate pass-through change if we condition on global financial market integration and the exchange rate regime? We now turn to the estimation of the parameters of the conditional model in equation (8). The estimates for short-term policy interest rates are reported in columns (9) to (10) of Table 5 (where the latter column considers a specification excluding the world war periods).

The full sample estimate for the direct pass-through effect of base-country policy interest rates to domestic policy interest rates, b_1, is about half the economic magnitude of the unconditional estimate, i.e. 0.16 (0.14 excluding World War I and II) versus 0.30 (see column 9 (10) of Table 5). It is only significantly different from zero at the 20% level. Moreover, interaction effects between interest rates, global financial market integration, and the exchange rate regime are statistically significant and strong in economic magnitude. This suggests that pass-through from base to domestic policy interest rates depends on whether an economy is open to global finance or closed, and on whether it has pegged or flexible exchange rates, potentially in line with the trilemma hypothesis. The c_1 coefficient is negative suggesting that a pegged exchange rate system can decrease pass-through. However, we also find that d_1 is positive and statistically significant at the 11% level, suggesting that the dependence on openness is more pronounced for countries with non-floating exchange rate systems, but insignificantly positive for floaters.

How large are these effects economically? Figure 2 plots the estimated pass-through from base short-term policy interest rates to domestic policy interest rates against the extent of de facto global financial market integration for both nonfloaters and floaters, as implied by the full sample estimates reported in column (10) of Table 5. First, the extent of financial market integration has little effect on pass-through coefficients for floaters; pass-through is relatively low, increasing from about 0.10 for fully segmented countries to about 0.30 for fully integrated countries. This is largely consistent with the trilemma hypothesis: floating exchange rates should suffice to guarantee monetary policy independence. Second, for nonfloaters financial market integration dependence is much more pronounced. Specifically, pass-through of nonfloaters well integrated into global finance is high, at about 0.60. Segmentation from global finance should protect domestic policy interest rates from movements in base-country policy interest rates. Indeed, pass-through is nil or even negative for nonfloaters with integration levels lower than 0.5, which suggests that they can decouple from movements in base-country policy interest rates, or even lean against them. These findings again support the trilemma hypothesis, and are statistically significant. The positive estimate for the triple interaction coefficient d_1, which is significant at the 11% level, suggests that financial integration dependence is stronger for nonfloaters than floaters in a statistically significant manner. Third, pass-through of floaters well integrated into global finance is only half as large as that of nonfloaters. This again suggests that a flexible exchange rate acts as a shock absorber of movements to base-country policy interest rates, in line with the trilemma hypothesis. In the online appendix, we show that the results for long-term interest rates are in line with those for short-term policy rates.

How has interest rate pass-through evolved over the last century? Figure 3 shows the evolution between 1885 and 2014 of the average pass-through estimates from base to short-term interest rates for nonfloaters and floaters, respectively, as predicted by the full sample estimates reported in column (10) of Table 5.[16] Pass-through on the short end of the yield curve for nonfloaters follows a nice swoosh shape, which largely reflects the temporal pattern of de facto global financial market integration over the last century. Pass-through for floaters is more stable over time, in contrast. Interestingly, short-term interest rate pass-through remains appreciably higher for nonfloaters than for floaters in the modern era of financial globalization, at about 0.50 and 0.30, respectively, on average in the 2000s. This suggests that central banks outside the U.S. still exert more control on their domestic short-term interest rates in the presence of flexible exchange rates, which can act as a shock absorber, despite the potential existence of global financial cycles set in motion by US monetary policy impulses. This finding is consistent with the trilemma hypothesis, but not with the dilemma hypothesis. Results for long-term rates are broadly similar, as we show in the online appendix.

4.2.3. Trilemma vs. dilemma hypotheses

While our full sample results do not point to evidence in favour of a "dilemma," this hypothesis has surfaced only recently emphasizing the increasing importance of U.S. financial cycles in the world economy.

[16] There are no estimates for nonfloaters during World War I because only the U.S. had stuck to the gold standard in this period. The 14 countries shown in the figure were all floaters as they had either suspended gold convertibility or, in the case of Spain, were previously not on the gold standard.

It is important to qualify what our results indicate about the recent dilemma/trilemma debate. The results in the extant studies (e.g. Rey (2013), Miranda-Agrippino and Rey (2014), Passari and Rey (2015)) concern the worldwide transmission of global financial cycle shocks which seem to be correlated with the VIX, an indicator of option implied volatility on the S&P500 stock index. The period considered is post-1990. Here we narrowly focus on the transmission of short-term and long-term interest shocks over a very long historical period. Yet, it remains interesting to translate the dilemma/trilemma debate more precisely to our setting. Essentially, the dilemma hypothesis would suggest that pass-through is now large, irrespective of the exchange rate regime. That is, nonfloaters can no longer as easily escape the global financial or interest rate cycle by introducing capital controls. Also, presumably, even countries with a floating exchange rate should experience pass-through, as the exchange rate no longer plays the role of a shock absorber in increasingly globalized markets.

With this translation in hand, what do our results really contribute to the debate?

First, our results are overall certainly inconsistent with the dilemma hypothesis: nonfloaters can achieve a high degree of monetary policy autonomy when global financial integration is low and are less exposed to foreign interest rate shocks; and countries with floating exchange rate regimes are subject to much less-pass through than pegged currency countries when capital markets are financially integrated globally, and significantly so when the degree of financial integration is very high (see again Figure 2).

However, in interpreting these findings, it is important to recall our results on the "integration swoosh." These results strongly suggest that the degree of global financial integration is very high for the countries in our sample post-1990, and higher than in the earlier globalization wave. As Figure 2 further shows, we therefore do observe a positive non-negligible pass-through for floating-exchange-rate countries, consistent with the dilemma hypothesis. However, for the most integrated countries and the largest part of our sample period, our model predicts significantly more pass-through for nonfloaters than for floaters, which remains inconsistent with the dilemma hypothesis.

Second, our results may be erroneous if the model parameters are unstable. Perhaps the model parameters have changed recently and become more consistent with the dilemma hypothesis. Upon reflection, examining this is fraught with difficulty, exactly because of the previous point we made. The identification of our conditional model relies on substantial time and cross-country variation in the degree of financial integration and exchange rate regimes. However, post-1990 this heterogeneity has diminished, which challenges the identification of the model.[17] This also makes it conceivable that the significance results at high levels of integration is based on full sample observations rather than on recent data. To examine this a bit more formally, we re-estimated the model allowing all parameters to break in 1990 (see the results in the online appendix). We do find that a likelihood ratio test typically rejects the null of no break at the 5% level for both short and long-term rates. However, the parameter changes do not support the dilemma hypothesis and very much confirm the identification problems discussed above. For example, for short-term interest rates one parameter change, which is significant at the 10% level, indicates that for countries with pegged exchange rates financial integration decreases pass-through, which makes little sense. We also find that the b_1 coefficient increases (overall pass-through), but that global financial integration reduces pass-through, and more considerably so for countries with pegged exchange rate systems. This model would imply that pass-through is lower than before for reasonably integrated countries. Although this is not very plausible, it is surely inconsistent with the dilemma hypothesis.

Finally, it is conceivable that the recent dilemma results are heavily influenced by the Great Recession, where economic conditions in the US spilled over into other countries. Of course, trying to estimate the conditional model over such a short period with even more

[17] In particular, 38% of the observations are nonfloaters post-1990, against 74% pre-1990; for instance in Europe nonfloaters comprise EU currencies managed within the Exchange Rate Mechanism in the 1990s, versus only Denmark and Switzerland in the 2000s.

homogenous integration and currency regimes is likely to be even less advisable. We therefore propose a simpler methodology to provide an alternative test of the implications of our conditional model for the dilemma/trilemma hypotheses. We divide the data post-1990 in four compartments, analogously to the aggregate results in Table 5, namely nonfloaters versus floaters and high versus low financial integration.[18] This immediately reveals the problem with the analysis. We have only four observations that qualify as "low financial integration" and thus cannot provide a meaningful statistical analysis for that regime (see the online appendix). While the empirical estimate of pass-through for the low financial integration regime is indeed low, statistically this has little meaning. For the high financial integration regime, we find that there is significant pass-through for both nonfloaters and floaters and for both short-term and long-term interest rates. For short-term interest rates, the estimate is 0.437 for nonfloaters and 0.318 for floaters. For long-term interest rates, the corresponding numbers are 0.653 and 0.510. This confirms our discussion above. In a world where capital controls are no longer in place, pegging a currency exposes the country to shock transmission. But this is simply the trilemma at work. Floating-currency countries also experience significant interest rate spillovers from the base countries, but the coefficient is lower than it is for countries attempting to peg their currencies – which is in line with the trilemma, too – albeit not significantly so, which might reflect the relatively low number of observations from which we can draw inference. In any case, the dilemma findings are hard to interpret in a world of largely globalized capital flows.

5. Conclusions

We propose a simple measure of de facto global and regional equity market integration using the beta exposure of the stock market returns of 17 markets to either the global or regional equity market portfolio. The beta exposure depends significantly on de jure market integration and global growth uncertainty, both accounting for about 50% of the total variation.

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[18] As aforementioned, countries with high (low, respectively) global financial market integration in period t are defined as those for which $\overline{\beta}_{i,t}^{glo} > 1$ ($\overline{\beta}_{i,t}^{glo} \leq 1$, respectively).

When viewed over time from 1885 to 2014, we uncover a “swoosh pattern” in de facto global financial market integration. That is, global financial market integration was high pre-1913, still higher post-1990, and low in the interwar period. In fact, we statistically reject the presence of other shapes hypothesized in earlier literature, such as a flat line, a U shape, a J shape, but cannot reject this distinct “swoosh” pattern. For regional integration, we do not find a clear statistically significant pattern. Also, we do not find integration to have reversed after the recent global crisis, contrary to claims in a number of recent papers.

Our results have implications for the recent debate on the trilemma hypothesis, which posits that a country can only run two of the three following policies: open capital markets, an independent monetary policy and a pegged exchange rate. We investigate the role of de facto financial market integration and the exchange rate regime on monetary policy interdependence measured by the sensitivity of local interest rate changes to international base rate changes, using both short and long-term interest rates.

Our evidence is consistent with the trilemma hypothesis. First, for countries with flexible exchange rates, interest rate pass-through is rather limited and is not affected by the extent of de facto financial market integration. However, for nonfloaters, a higher degree of financial integration increases interest rate pass-through, undermining monetary policy independence. For segmented markets, interest rate pass-through is close to zero or even negative, hence enabling these countries to decouple from base interest rates. For integrated markets, in contrast, pass-through can be as high as 0.60 for short and 0.75 for long-term interest rates. For the recent period, we find that the trilemma is alive and well and has not morphed into a dilemma as recent papers claim, although it is natural to witness larger pass-through in more globalized markets.

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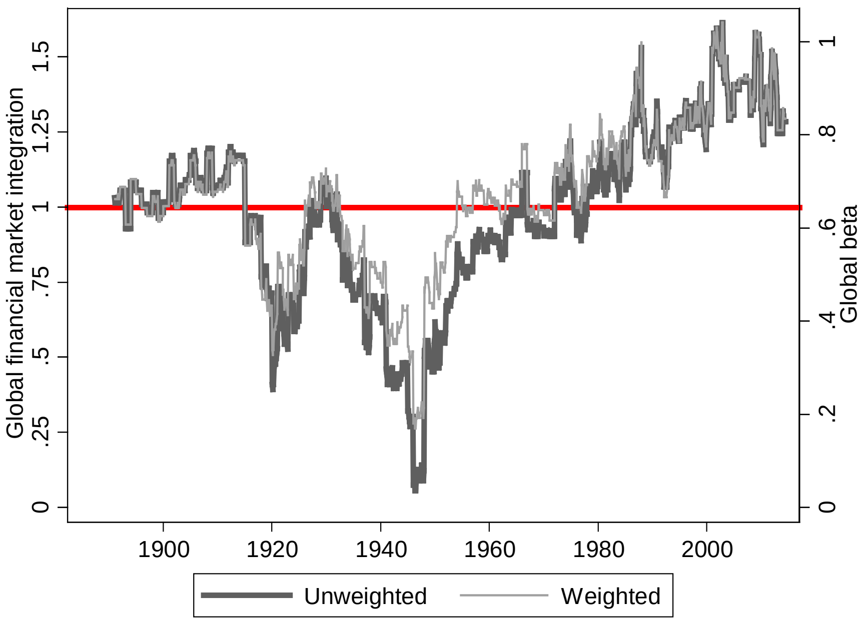
Panel A. Global Financial Market Integration and Conditional Global Betas



Panel B. Global Betas – Conditional vs. Rolling

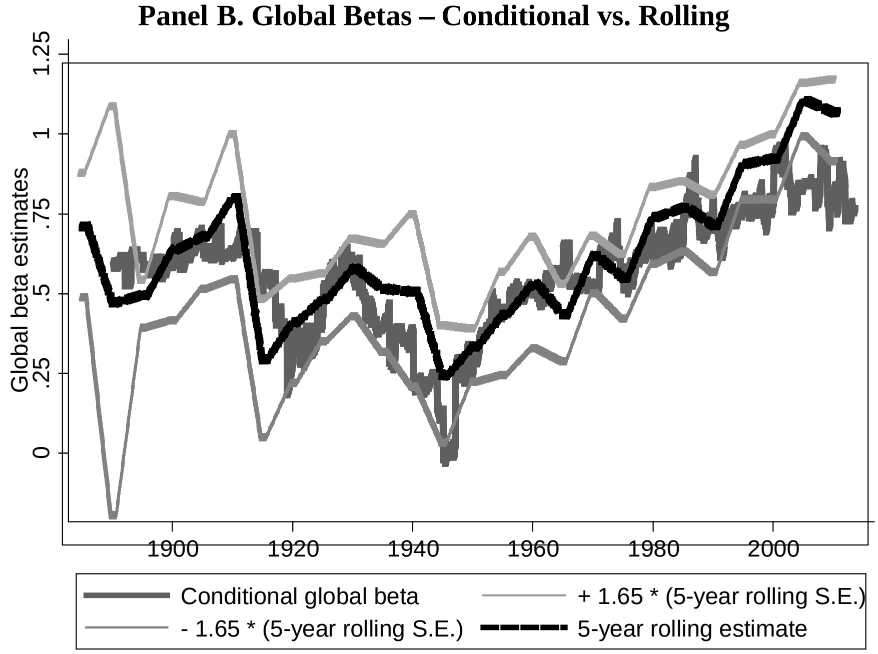


Figure 1. Global betas. Panel A of the figure shows the evolution between 1885 and 2014 of the unweighted (thick grey lines) and value-weighted (light grey lines) averages (17 countries) of our measures of global financial market integration (defined in equations (4) and (5)) and corresponding conditional beta estimates obtained from the full model equations (1) to (3). Panel B of the figure shows the evolution between 1885 and 2014 of the unweighted averages (17 countries) of the estimated cross-sectionally heterogeneous and time-varying parameters $\beta_{i,t}^{glo}$ (thick grey lines) obtained from the full model equations (1) to (3) together with 90% confidence bands (light grey lines) obtained from pooled estimates with a 5-year rolling forward window, with non-overlapping observations, of β_t^{glo}, and the point estimates (shown as black dashed lines).

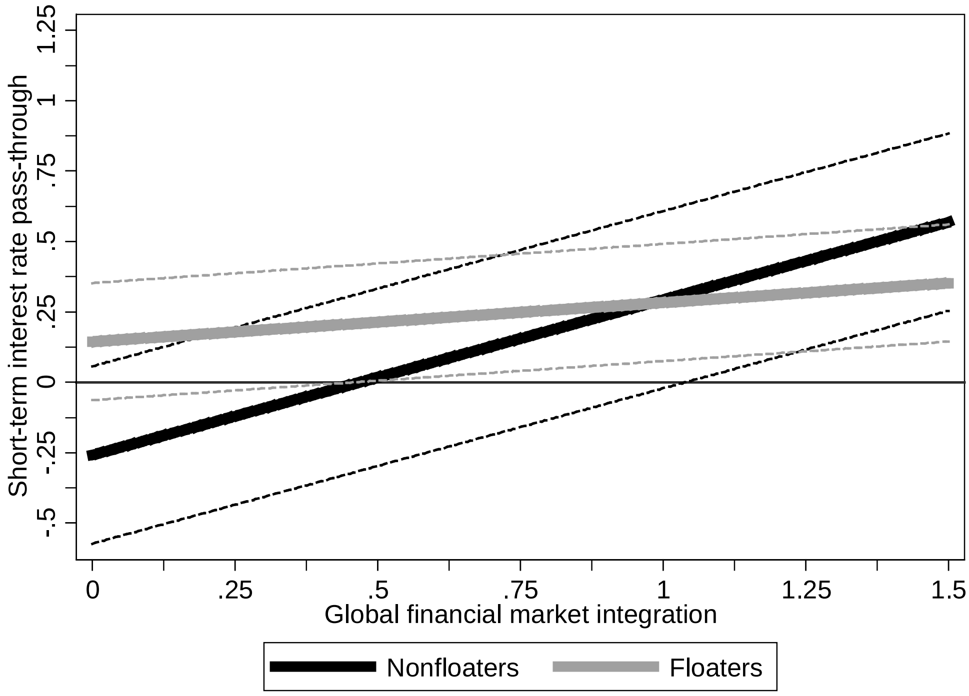


Figure 2. Estimated interest rate pass-through vs. global financial market integration. The figure plots the estimated pass-through from base (i.e. US, UK or German) short-term interest rates to domestic interest rates against the extent of global market integration for both nonfloaters and floaters as predicted by the full sample estimates reported in column (10) of Table 5. 90% confidence bands are shown as dotted lines.

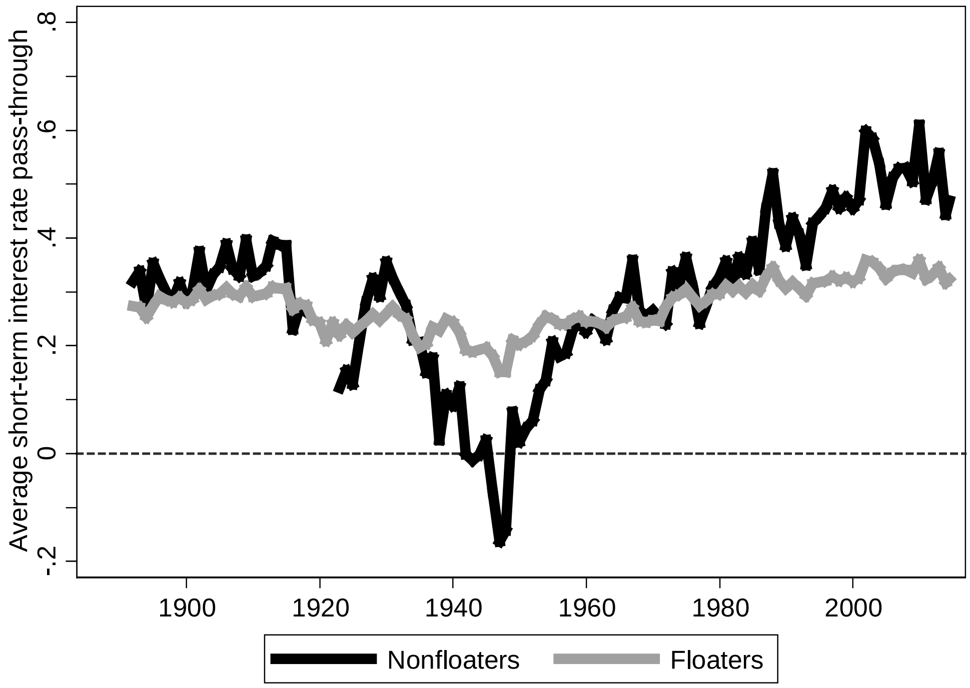


Figure 3. Estimated interest rate pass-through: nonfloaters vs. floaters – 1885-2014. The figure shows the evolution between 1885 and 2014 of the average (14 countries) pass-through estimates from base (i.e. US, UK or German) short-term interest rates to domestic interest rates for nonfloaters and floaters as predicted by the full sample estimates reported in column (10) of Table 5.

Table 1
Data Overview

The table reports summary statistics for the various variables used in the model. All statistics shown in the table are calculated for the sample's 17 economies over the period January 1885-June 2014.

Variables	Units	Frequency	Definition	Unit of observation	Source	mean	median	s.d.	min.	max.
Returns										
Equity returns	in % per month	Monthly	Exact return of the local equity market index in dollar terms	Country	Global Financial Data	0.48	0.44	6.57	-92.45	179.64
Risk free rate	in % per month	Monthly	10-year US Treasury yield in domestic currency terms	Country	Global Financial Data	0.37	0.30	0.19	0.12	0.12
Instruments										
Trade openness	% of GDP	Annual	Sum of total exports and imports of goods relative to output	Country	Mitchell (1998a, 1998b and 1998c) and IMF Direction of Trade Statistics	47.23	40.10	34.40	2.40	352.80
Regional trade openness	% of total trade	Annual	Sum of a country's exports and imports of goods to/from its neighbours relative to total trade	Country	Mitchell (1998a, 1998b and 1998c) and IMF Direction of Trade Statistics	51.28	59.00	24.03	0.00	100.00
Capital account openness	index from 0 to 100	Annual	Extent of the restrictions to capital outflows and inflows from residents and nonresidents	Country	Quinn and Voth (2008) and Quinn and Toyoda (2008)	73.59	80.00	30.20	0.00	100.00
Financial development	in %	Annual	Equity market capitalization relative to output	Country	Rajan and Zingales (2003) and Beck, Demirgüç-Kunt, and Levine (2010)	60.40	50.00	46.00	3.00	323.00
Oil prices	in %	Monthly	Log. deviation of the dollar price of an oil barrel from a 5-year moving average	Global	Global Financial Data	6.92	4.48	27.21	-107.38	123.10
Global growth uncertainty	in % points per year	Annual	Logarithm of the standard deviation of real GDP growth across countries in the sample	Global	Maddison (2010) and IMF World Economic Outlook	1.02	0.99	0.68	-0.36	3.31
Local growth uncertainty	in % points per year	Annual	Logarithm of the standard deviation of real GDP growth in each country over non-overlapping windows of 5 years	Country	Maddison (2010) and IMF World Economic Outlook	1.22	1.13	0.56	0.21	3.78
Local growth uncertainty	in % points per year	Annual	Logarithm of the standard deviation of real GDP growth in each country over overlapping windows of 5 years	Country	Maddison (2010) and IMF World Economic Outlook	1.23	1.13	0.56	0.05	3.78
High market volatility periods	%	Monthly	Share of the countries with normalised log conditional volatility of stock returns from GARCH(1,1) models above 1.65 in a given month	Global	Authors' calculations	15.14	11.76	17.24	0.00	100.00
Other data										
Equity market capitalization	in %	Annual	Equity market capitalization relative to total sample capitalization	Country	Rajan and Zingales (2003) and Beck, Demirgüç-Kunt, and Levine (2010)	5.88	1.90	10.58	0.00	56.40
Central bank policy rates	in % per year	Monthly	Main policy interest rate in domestic currency terms	Country	Global Financial Data	4.93	4.50	2.95	0.00	90.00
Nonfloaters	Dummy variable (0/1)	Annual	Dummy variable which equals zero for floaters (floats, managed floats or freely falling exchange rates) and one for nonfloaters (other countries, including those on the gold standard)	Country	Ilzetzki, Reinhart, and Rogoff (2004) and Reinhart and Rogoff (2011)	0.67	1.00	0.47	0.00	1.00

Table 2
Full Model Estimates

The table reports the estimates of the parameters $\beta_{i,t}^{glo}$ and $\beta_{i,t}^{reg}$ from the full model equations (1) to (3). Each instrument is included individually in the estimates reported in columns 2 to 7, while all seven instruments are included in column 8. We then obtain a parsimonious model in column 9 by excluding the variables with insignificant parameters. All the estimates control for country fixed effects, year effects and for the direct effects of the instruments included in vector **X** (whose coefficients are not reported to save space).The standard errors reported in parentheses are robust to heteroskedasticity and are clustered by country. ***, **, *, and $^+$ indicate statistical significance at the 1%, 5%, 10%, and 15% levels, respectively.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Global factor	0.679***	0.515***	0.198**	0.643***	0.678***	0.892***	0.640***	0.430***	0.443***
	(0.033)	(0.083)	(0.087)	(0.047)	(0.032)	(0.036)	(0.036)	(0.139)	(0.106)
Regional factor	0.286***	0.117	-0.145**	0.346***	0.283***	0.425***	0.290***	-0.087	-0.129*
	(0.083)	(0.117)	(0.062)	(0.094)	(0.086)	(0.087)	(0.077)	(0.104)	(0.062)
Global factor × trade openness		0.003*						0.000	
		(0.002)						(0.001)	
Regional factor × regional trade		0.004**						0.002	
		(0.002)						(0.001)	
Global factor × capital openness			0.006***					0.004***	0.004***
			(0.001)					(0.001)	(0.001)
Regional factor × capital openness			0.007***					0.006***	0.006***
			(0.001)					(0.002)	(0.001)
Global factor × financial development				0.100+				0.011	
				(0.063)				(0.053)	
Regional factor × financial development				-0.087				-0.087	
				(0.070)				(0.083)	
Global factor × oil prices					0.001*			0.001	
					(0.001)			(0.001)	
Regional factor × oil prices					-0.001*			-0.001**	
					(0.000)			(0.000)	
Global factor × growth uncertainty						-0.254***		-0.167***	-0.175***
						(0.031)		(0.036)	(0.029)
Regional factor × growth uncertainty						-0.133***		-0.026	
						(0.037)		(0.042)	
Global factor × high market volatility							0.001***	0.002***	0.002***
							(0.000)	(0.000)	(0.000)
Regional factor × high market volatility							-0.000	-0.001	
							(0.001)	(0.001)	
Constant	0.315**	0.700+	-0.538	-0.921*	-0.283	-0.047	0.228*	-0.653	0.401
	(0.132)	(0.430)	(0.650)	(0.444)	(0.287)	(1.167)	(0.128)	(0.491)	(1.627)
Observations	20,800	20,546	20,587	18,142	20,798	20,778	20,800	17,850	20,587
R^2	0.209	0.270	0.228	0.221	0.210	0.223	0.210	0.311	0.233

Table 3
Testing for the Shape of Global Financial Market Integration over the Last Century

The table reports in Panel A the estimates of the following model:

$$R_{i,t} = \alpha_i + \lambda_t + (\boldsymbol{\beta}_{\mathbf{j}}^{\mathbf{glo}}{}'\mathbf{D}_{\mathbf{j}})\, F_t^{glo,\backslash i} + (\boldsymbol{\beta}_{\mathbf{j}}^{\mathbf{reg}}{}'\mathbf{D}_{\mathbf{j}})\, F_t^{reg,\backslash i} + \varepsilon_{i,t} \tag{6}$$

where j =1, 2, 3; D_1 denotes a dummy variable equal to one between 1885 and 1913 and zero otherwise; D_2 a dummy variable equal to one between 1914 and 1990 and zero otherwise; and D_3 a dummy variable equal to one between 1990 and 2014 and zero otherwise. The estimates, obtained by OLS, control for country fixed effects and for year effects. The standard errors reported in parentheses are robust to heteroskedasticity and are clustered by country. ***, **, *, and $^+$ indicate statistical significance at the 1%, 5%, 10%, 15% levels, respectively. In addition, Panel B reports four hypothesized shapes that may characterise global financial market integration over the last century, while Panel C reports the *p*-value of Wald restriction tests on the estimated coefficients of the betas interacted with D_1, D_2 and D_3 corresponding to each of the four hypothesized shapes.

Panel A. Full sample estimates (1885-2014)	
Global factor × D_1	0.669***
	(0.051)
Global factor × D_2	0.536***
	(0.046)
Global factor × D_3	0.943***
	(0.041)
Regional factor × D_1	0.355**
	(0.160)
Regional factor × D_2	0.212**
	(0.076)
Regional factor × D_3	0.585***
	(0.146)
Constant	0.338**
	(0.118)
Observations	20,800
R^2	0.221

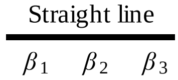
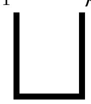
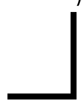
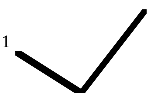
Panel B. Hypothesized shape	**Panel C. *p*-value of Wald test**
Straight line  β_1 β_2 β_3	$\beta_1^{glo} = \beta_2^{glo} = \beta_3^{glo}$: 0.000 $\beta_1^{reg} = \beta_2^{reg} = \beta_3^{reg}$: 0.009
U-shape  β_1 β_3 β_2	$\beta_1^{glo} = \beta_3^{glo}$: 0.000 , $\beta_1^{reg} = \beta_3^{reg}$: 0.276 $\beta_1^{glo} > \beta_2^{glo}$: 0.965 , $\beta_1^{reg} > \beta_2^{reg}$: 0.798 $\beta_3^{glo} > \beta_2^{glo}$: 0.999 , $\beta_3^{reg} > \beta_2^{reg}$: 0.998
J (or inverted L)-shape  β_3 β_1 β_2	$\beta_1^{glo} = \beta_2^{glo}$: 0.068 , $\beta_1^{reg} = \beta_2^{reg}$: 0.404 $\beta_3^{glo} > \beta_1^{glo}$: 0.999 , $\beta_3^{reg} > \beta_1^{reg}$: 0.861 $\beta_3^{glo} > \beta_2^{glo}$: 0.999 , $\beta_3^{reg} > \beta_2^{reg}$: 0.998
Swoosh-shape  β_3 β_1 β_2	$\beta_1^{glo} > \beta_2^{glo}$: 0.965 , $\beta_1^{reg} > \beta_2^{reg}$: 0.798 $\beta_3^{glo} > \beta_1^{glo}$: 0.999 , $\beta_3^{reg} > \beta_1^{reg}$: 0.861 $\beta_3^{glo} > \beta_2^{glo}$: 0.999 , $\beta_3^{reg} > \beta_2^{reg}$: 0.998

Table 4
Testing for a Reversal in Global Financial Market Integration since the Great Recession

The table reports in Panel A the estimates of the following model:

$$R_{i,t} = \alpha_i + \lambda_t + (\boldsymbol{\beta}_{\mathbf{j}}^{\mathbf{glo}}{}'\mathbf{D}_{\mathbf{j}})\, F_t^{glo,\setminus i} + (\boldsymbol{\beta}_{\mathbf{j}}^{\mathbf{reg}}{}'\mathbf{D}_{\mathbf{j}})\, F_t^{reg,\setminus i} + \varepsilon_{i,t} \tag{6}$$

where j =1, 2; D_1 denotes a dummy variable equal to one between 1990 and 2006 and zero otherwise; D_2 a dummy variable equal to one between 2007 and 2014 and zero otherwise. The specification also includes pre-1914 and 1914-1990 period dummies, which are not shown in the table to save space. The estimates, obtained by OLS, control for country fixed effects and for year effects. The standard errors reported in parentheses are robust to heteroskedasticity and are clustered by country. ***, **, *, and $^+$ indicate statistical significance at the 1%, 5%, 10%, 15% levels, respectively. In addition, Panel B reports the p-value of Wald restriction tests on the estimated coefficients of the betas interacted with D_1 and D_2.

Panel A. Full sample estimates (1885-2014)		**Panel B. p-value of Wald test H_0**	
Global factor × D_1	0.860***		
	(0.049)		
Global factor × D_2	1.091***		
	(0.063)	$\beta_1^{glo} = \beta_2^{glo}$:	0.010
Regional factor × D_1	0.558***	$\beta_1^{glo} > \beta_2^{glo}$:	0.005
	(0.146)	$\beta_1^{glo} < \beta_2^{glo}$:	0.995
Regional factor × D_2	0.500***		
	(0.162)	$\beta_1^{reg} = \beta_2^{reg}$:	0.557
Constant	0.333**	$\beta_1^{reg} > \beta_2^{reg}$:	0.721
	(0.119)	$\beta_1^{reg} < \beta_2^{reg}$:	0.278
Observations	20,800		
R^2	0.222		

Table 5

Testing for the Monetary Policy Trilemma in History – Short-Term Interest Rates

The table reports the estimates of the parameters of the unconditional model equation (7) using short-term policy interest rates, in the spirit of Obstfeld, Shambaugh and Taylor (2005). Column 1 reports pooled estimates, Columns 2 to 4 report estimates over three periods. Columns 5 to 8 report estimates by country groups. The estimates of the parameters from the conditional model equation (8) are reported in columns 9 and 10 over the full sample. The standard errors reported in parentheses are robust to heteroskedasticity and are clustered by country. ***, **, *, and + indicate statistical significance at the 1%, 5%, 10%, and 15% levels, respectively.

	Unconditional estimates à la Obstfeld, Shambaugh and Taylor (2005)								Conditional estimates	
		By time period			By country group				1885-2014	
	Pooled	Gold standard (Pre-1914)	Bretton Woods (1959-1970)	Post-Bretton Woods (1973-2000)	Floaters	Nonfloaters	High global financial market integration	Low global financial market integration	Full sample	Ex. World War I & II
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Interest rate (b_1)	0.303***	0.187***	0.096**	0.561***	0.306***	0.293***	0.404***	0.136*	0.158	0.145
	(0.044)	(0.052)	(0.035)	(0.120)	(0.049)	(0.051)	(0.038)	(0.077)	(0.109)	(0.110)
Global market integration (b_3)									-0.022**	-0.028**
									(0.009)	(0.010)
Nonfloater (b_2)									-0.012**	-0.015**
									(0.005)	(0.007)
Interest rate × global market integration (c_2)									0.132	0.139
									(0.126)	(0.126)
Interest rate × nonfloater (c_1)									-0.415*	-0.402*
									(0.212)	(0.213)
Global market integration × nonfloater (c_3)									0.014	0.017
									(0.010)	(0.011)
Interest rate × global market integration × nonfloater (d_1)									0.419+	0.413+
									(0.238)	(0.239)
Constant	-0.001**	0.002*	0.010***	-0.001	-0.002	0.000	-0.006***	0.005***	0.019***	0.026***
	(0.000)	(0.001)	(0.003)	(0.002)	(0.001)	(0.004)	(0.001)	(0.002)	(0.005)	(0.006)
Observations	1,596	326	168	378	1,149	440	887	709	1,439	1,305
Adjusted R^2	0.0815	0.120	0.0658	0.0950	0.0842	0.0704	0.167	0.0123	0.0917	0.0921
Log likelihood	1189	454.3	235.9	69.30	816	369.3	716.6	492.8	1027	874.2