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ON THE GLOBAL FINANCIAL MARKET INTEGRATION "SWOOSH" AND THE TRILEMMA

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ABSTRACT

We propose a simple measure of de facto financial market integration based on a factor model of monthly equity returns, which can be computed back to the first era of financial globalization for 17 countries. Global financial market integration follows a “swoosh” shape – i.e. high pre-1913, still higher post-1990, low in the interwar period – rather than the other shapes hypothesized in earlier literature. We find no evidence of financial globalization reversing since the Great Recession as claimed in other recent studies. De jure capital account openness and global growth uncertainty are the two main determinants of long-run global financial market integration. We use our measure to revisit the debate on the trilemma between financial openness, the exchange rate regime, and monetary policy autonomy, and on whether the trilemma has recently morphed into a dilemma due to global financial cycles. We find evidence consistent with the trilemma and inconsistent with the dilemma hypothesis, both throughout history and for the recent decades; non-US central banks still exert more control over domestic interest rates when exchange rates are flexible in economies open to global finance.

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The process of international financial integration is not a gentle climb towards ever higher peaks. This is true both from a short-run and from a long-run perspective. Bekaert and Harvey (1995), focusing on the post-1990 financial globalization wave in emerging equity markets, suggested that de facto integration may exhibit reversals and does not become necessarily stronger over time. Evidence that financial globalization might have partly reversed has resurfaced in the wake of the Great Recession (see e.g. Rose and Wieladek (2014), van Rijckeghem and Weder (2014), Giannetti and Laeven (2012, 2016)). These recent papers have focussed on banks, or bonds, but not on equity markets. 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), among many others).

From a long-run 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 globalization wave which started with capital account liberalizations in advanced economies in the 1980s and in emerging markets in the 1990s, a period also referred to as the modern era of financial globalization. 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 (see e.g. Shambaugh (2004), Obstfeld, Shambaugh, and Taylor (2005), Miniane and Rogers (2007), Bluedorn and Bowdler (2010), Klein and Shambaugh (2013), Aizenmann, Chinn and Ito (2014), Pasricha et al. (2015)). In particular, several articles (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) stress the critical role played by the US dollar and US monetary policy in setting global liquidity and credit conditions. 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 propose a simple measure of equity market integration which can be computed back to the first era of financial globalization for 17 countries. 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-variation in correlations is accounted for by changes in factor volatilities (see e.g. Forbes and Rigobon (2002), Dungey et al. (2004), Bekaert, Harvey and Ng (2005), Bekaert et al. (2014)). In contrast, the fundamental change in risk occurring upon capital market liberalization naturally increases dynamic betas (see Chari and Henry (2004) for a simple model). 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). 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 I 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. Section II presents our formal test of the long-run temporal pattern of financial market integration and provides evidence consistent with a swoosh shape. Section III employs our measure to revisit the debate on the trilemma versus dilemma hypothesis.

I. 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.

A. The Factor Model

A.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. When the beta on the first factor is zero, the model becomes a regional CAPM; when the beta of the regional factor is set to zero, the model can act as a world CAPM. 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.[1]

The full model is:

$$R_{i,t} = \alpha_i + \lambda_t + \beta_{i,t}^{glo} F_t^{glo,\setminus i} + \beta_{i,t}^{reg} F_t^{reg,\setminus i} + \boldsymbol{\gamma}_{\mathbf{i,t}}^{\mathbf{glo}}{}' \mathbf{X}_{\mathbf{i,t-k}}^{\mathbf{glo}} + \boldsymbol{\gamma}_{\mathbf{i,t}}^{\mathbf{reg}}{}' \mathbf{X}_{\mathbf{i,t-k}}^{\mathbf{reg}} + \varepsilon_{i,t} \tag{1}$$

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

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

[1] 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.

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 **X** is a vector of control variables designed to capture time and cross-sectional variation in factor exposures. These variables are country-specific, and are typically lagged by one year. If the dimension of **X** is k, the vectors $\mathbf{b_1}^{\mathbf{glo}}$ and $\mathbf{b_1}^{\mathbf{reg}}$ are $k \times 1$. When the model includes control variables **X**, the conditional mean also depends automatically on lagged **X**s. 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, Northern America and Asia-Pacific).[2]

To avoid adding-up constraints and spurious correlations, the factors are value-weighted across countries, but exclude returns of country i itself. To obtain an intuitive interpretation of the estimates of the factor loadings, we orthogonalize the two factors as in Bekaert, Harvey and Ng (2005), Bekaert, Hodrick, and Zhang (2009) and Bekaert et al. (2014). The regional market factor is orthogonalized by regressing regional market returns on global market returns over the full sample period and then using the residuals of this regression as the regional market factor. The orthogonalized factors are estimated for each country individually as country i itself is excluded from the market factors. This enables us to distinguish global from regional patterns of financial market integration. Table I contains an overview of the data and selected descriptive statistics. Further details on the sources of the data and on the methodology used to assemble them are provided in Appendix A.

Table I

A.2. Instruments to Model Cross-Sectional and Time Variation in Exposures

Equations (1) to (3) contain a set of lagged instruments, $\mathbf{X_{i,t-k}}$, 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

[2] The three regions include 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).

tradition in finance; see, for example, Ferson and Harvey (1991) and Dumas and Solnik (1995).[3] We entertain seven potential instruments, which are listed in Table I, to distinguish between different channels and hypotheses regarding the sources of financial market integration, both at the global and regional level.[4]

The first three channels include measures of external exposures through trade and financial openness. 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 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 its neighbours, that is 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). These indices measure the extent of restrictions to capital outflows and inflows by residents and nonresidents in country i and year t.

Domestic financial development is the fourth potential determinant of the cross-sectional and time variation in factor exposures we consider. 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 is one reason preventing foreign institutional investors

[3] Note that we do not mean to suggest that these "instruments" are "exogenous" in the strict sense of econometric identification. In the asset pricing literature, as discussed in Ferson and Harvey (1991) for instance, this term is simply used for 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)).

[4] We use annual observations to fill in for monthly observations.

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, which we obtained 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. Data limitations prevent us from casting a wide net in terms of variables. We consider just two specific 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 is likely to increase global factor exposures. Increases in (global) risk aversion may generate the opposite effect. Higher risk aversion may lead investors to retreat away 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 almost any model, high risk aversion should increase the volatility of stock returns (see e.g. Bekaert, Engstrom and Xing (2009)) so we measure risk aversion indirectly through volatility. Specifically, we measure the share of countries in the sample with high equity market volatility in a given month. We estimate the conditional volatility of stock returns for each country of our sample using a GARCH(1,1) model. We normalise the conditional volatilities of each country's stock returns and define the 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 section B.5 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 use real GDP growth instead. Measuring a conditional volatility is challenging and we use three different measures. The first is the logarithm of the standard deviation of real GDP growth in each country over non-overlapping windows of 5 years, which yields 17 country-specific times series of annual observations on local growth uncertainty which are kept constant over 5-year intervals.[5] Analogously, we use the logarithm of the standard deviation of real GDP growth in each country over overlapping windows of 5 years centred around the current observation (with one year increments), which yields for each country local growth uncertainty measures varying in each year. These two measures are imperfect indictors of country-specific real uncertainty both requiring a number of time series observations to obtain a proxy for an estimate at time *t*. An alternative metric we employ is 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)). It is therefore correlated with the times series uncertainty of growth opportunities worldwide. Because of its second component, it also measures the divergence of growth opportunities

[5] 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.

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.

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

B.1. Model Estimation

We estimate our model for all countries jointly by means of pooled OLS. Standard errors account for heteroskedasticity and are clustered by country. Note that the instruments $\mathbf{X}_{\mathbf{i,t\text{-}k}}$ are lagged by one year to prevent an unobserved factor from simultaneously influencing both returns and the fundamental **X** in a given period and thereby generating a spurious relationship between both.[6]

When estimating the full model in equations (1) to (3), we consider two different model specifications. In a first step, we include each of the instruments individually. In a second step, we build on the work of David Hendry (see, for instance, Hendry and Krolzig (2005)) to pare down the regression to a more manageable number of independent variables. We start out with the full model including all instruments simultaneously, and then step-by-step reduce the model by excluding the interaction variables with insignificant parameters. If all interaction effects are insignificant, the variable is dropped from the regression. This approach aims to reduce the dimension of the model and to arrive at a model that can be interpreted in an economically meaningful way. Convergence was reached in two steps.

B.2. Measuring Time-Varying Global and Regional Financial Market Integration

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

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

[6] The trend deviation of oil prices and the high volatility variable, which are available at the monthly frequency, are lagged by one month instead.

where here $N = 17$ and $T = 1{,}554$. The relative global market integration of country i at time t then is defined as $\beta_{i,t}^{glo} \big/ \overline{\beta}^{glo}$. Similarly, the benchmark $\overline{\beta}^{reg}$ for regional market integration is the (weighted) average across countries and time of the $\beta_{i,t}^{reg}$ estimates, i.e.

$$\overline{\beta}^{reg} = \frac{1}{N}\frac{1}{T}\sum_{i=1}^{N}\sum_{t=1}^{T} w_{i,t}\beta_{i,t}^{reg} \ , \tag{5}$$

and the relative regional market integration of country i at time t is defined as $\beta_{i,t}^{reg} \big/ \overline{\beta}^{reg}$

B.3. The Temporal Pattern of Financial Market Integration

Table II reports pairwise correlations of the instruments in vector **X**. The correlations, albeit statistically significant, are generally low in terms of economic magnitude (at around 15-20% or less in absolute value in most cases, and close to 40% for only a few pairs of variables). This suggests that each instrument has the potential to contribute specific information regarding the underlying determinants of the cross-sectional and temporal variation in global and regional financial market integration. The only exception, unsurprisingly, is the correlation between global and local growth uncertainty measures, which reaches almost 60%.

Table II

Next, Table III reports the estimates of the conditional global and regional beta estimates 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. Note that we do not report estimates including the local growth uncertainty measures, which are never significant in multivariate specifications. Moreover, in univariate specifications only the interactions with the global factor were significant – albeit with the expected (negative) sign – which suggests that our results are driven primarily by global growth uncertainty.

We 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). 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. The model estimates suggest that 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 earnings growth reduces global betas significantly, in line with the model predictions of Bekaert at al. (2011). Global betas tend to increase significantly in periods of heightened market volatility, although the economic magnitude of the effect is economically very small (more on this below). Finally, while there are some significant univariate results, among all the instruments 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), growth uncertainty and the market volatility variable (the latter two only for global betas). We now further analyse the implications of this model for the time-variation in financial market integration.

Table III

Figure 1 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. Global measures are shown in Panel A, while regional measures are in Panel B.

Figure 1

That the temporal pattern of global financial market integration follows a swoosh shape is apparent from the figure. 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 taken in some countries, recently. The temporal pattern of regional financial market integration seems less clear, being in-between a swoosh and a U-shaped pattern.

A complementary perspective is provided by Figures 2 and 3 which show the conditional global and regional beta estimates for each country and for selected years, namely: 1913, 1928, 1945, 1973, 1990, and 2008. The temporal pattern of global financial market integration is nicely swoosh-shaped in all countries, which suggests that it is not only a broad overall trend but also a country-level phenomenon. The temporal pattern of regional financial market integration is less clear.

Figures 2, 3

B.4. Determinants of Variation in Financial Market Integration

What is the relative economic importance of the determinants of global financial market integration? Figure 4 gets at the issue by showing the evolution between January 1885 and June 2014 of the value-weighted averages of the conditional global beta estimates of our 17 countries when only one of each of the three significant instruments of vector **X** remains active, in turn. This is achieved by setting the loadings on the other significant instruments at their respective means. The figure makes clear that de jure capital account openness and effective global financial market integration go hand in hand. It also makes clear that heightened global growth uncertainty pulls in the opposite direction, reducing global financial market integration in an economically meaningful way. In contrast, the economic importance of high market volatility periods on financial market integration is comparably much smaller, as is evident from the figure.

Figure 4

That de jure capital account openness and global growth uncertainty opportunities explain the lion's share of the predictable variation in global equity returns is confirmed more directly by the variance ratio analysis of Table IV. 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

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

By definition, these variance ratios sum to one. The analysis confirms our earlier observations. 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. 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.

Table IV

B.5. Model Validation

As stressed e.g. by Cochrane (2001), Lewellen and Nagel (2006) and Brusa, Ramadorai and Verdelhan (2014), a challenge to our conditional factor model is that it requires the econometrician to know the ‘‘true’’ state variables. Lewellen and Nagel (2006) propose a methodology to circumvent this problem which does not require specifying the set of conditioning information. As long as betas are relatively stable within a certain period, simple factor regressions estimated over a short window – using no conditioning variables – provide direct estimates of assets' conditional betas. Using rolling 5-year windows of observations, it is possible to obtain time series of time-varying betas. In particular, the windows are "forward" and non-overlapping. That is, we split the sample in 5-year 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 previous 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 5 shows the evolution between 1885 and 2014 of the conditional global (Panel A) and regional (Panel B) betas (both 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). Our factor model does, in fact, pretty well. The simple rolling global beta estimates also follow a swoosh shape. The conditional betas fall mostly well within the confidence bands of the simple rolling beta estimates. The conditional global (regional) betas fall within the bands 81% (82%) 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. Excluding the two world wars, the conditional global and regional betas fall within the bands 82% and 84% of the time, respectively, which is a marginal improvement relative to the full sample. It is only starting from the early 2000s that the conditional betas underestimate systematically the extent of global financial market integration relative to what rolling betas would predict. Because this is indirect evidence that financial globalization did not reverse since the Great Recession in 2007-2009, as claimed in other studies, we come back to this finding below.

Figure 5

Figures B1 and B2 in Appendix B show the corresponding betas and confidence bands broken down by country. The figures confirm that the country-specific conditional betas mostly fall well within the confidence bands of the simple rolling beta estimates.

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 upwards correlations in international equity markets. As a result, if the conditional betas of our reduced-form factor model did not increase during the global financial crisis, while unconditional correlations did, this is additional evidence that the model is well specified.

We examine the difference between our beta measures and a correlation measure in Figure 6. It shows the evolution between 1885 and 2014 of the unweighted averages of the conditional global and regional beta estimates (shown as thick grey lines) together with 1-year forward rolling (non-overlapping) pooled correlations between the equity excess returns $R_{i,t}$ and the global and regional market factors, respectively. The correlation between the two global measures is relatively high, at 0.80, but it is a paltry 0.15 for the two regional measures, suggesting that betas and correlations may produce different inferences. The differences are indeed pronounced during the recent global crisis. Consider Panel B of Figure 6 first. Bilateral correlations between country returns and the regional market factor peak at the time of the collapse of Lehman Brothers, which is suggestive of a possible volatility bias. In contrast, the conditional regional betas remain more stable. For global market integration, the correlation between the beta and correlation measures varies substantially over time. For the recent globalization period (2001-2014) it is in fact zero, but even that number hides very different sub-sample behaviour. Over the 2001-2007 pre-crisis period, the correlation is -0.67, whereas over the recent crisis period (2008-2014) it is 0.44. Thus, while a correlation measure may reproduce some of the long-run patterns of long-run financial integration as Figure 6 demonstrates, it cannot really be used to make precise inferences.

Figure 6

II. The Swoosh in Financial Market Integration

While the swoosh pattern in de facto financial market integration is apparent in our full model estimates, we here attempt to formally test for it. This section sets out a framework to do so and reports the results.

A. Testing for a Swoosh Pattern

To test for the swoosh pattern, we start by estimating the following simple variant of our two-factor model:

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

where j =1, 2, … k; and D_k denotes a dummy variable which equals one over time period k and zero otherwise. All variables are defined similarly as before, and we include the same two factors – a global market factor, and a regional market factor. In practice, we set k = 3 and focus on three subperiods of interest, namely: 1885-1913 (which is often referred to as the first era of financial globalization; see e.g. Bordo, Cavallo, and Meissner (2010)); 1914-1990, which includes the interwar period (when several countries adopted protectionist and capital control measures in the run-up to World War II), 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. This model embeds a simple constant-beta model, which can be straightforwardly tested through a simple Wald test.

However, we can also formally test whether the temporal pattern of de facto international financial integration follows a U shape, as hypothesised by e.g. Bordo and Flandreau (2003), Obstfeld and Taylor (2003, 2004), and Goetzmann et al. (2005); 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_k, j =1, 2, 3, the corresponding Wald restriction tests are, respectively

$$H_0 : \beta_1^f = \beta_3^f , \beta_1^f > \beta_2^f , \beta_3^f > \beta_2^f$$

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 $f = glo, reg$.

B. Empirical Results

B.1. Simple Constant-Beta Model Estimates

Table V reports the unconditional beta estimates from the constant-beta version of the model in equation (6) obtained by OLS in columns 1 to 3. Estimates obtained by excluding outliers from the sample (i.e. excess equity returns larger than 30% within a month in absolute value) are reported in column 4. Those obtained with a random-effect estimator are reported in column 5. The estimates of columns 3 to 5 control for both country and year effects. The unconditional global beta is about 0.7, while the regional market factor beta is estimated to be about 0.3. The economic magnitude of these estimates is not too far off those obtained over much shorter samples. For instance, Bekaert *et al.* (2014) obtain unconditional beta estimates for a sample of 415 country-sector portfolios over the period January 1995-March 2009 of about 0.4 for both their global financial and US equity market factors, as well as about 0.5 for their domestic market factor. However, their portfolios included many emerging market portfolios. 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 models estimated by e.g. Bekaert *et al.* (2014). What is also striking is that the estimates remain remarkably robust across estimation methods. The rho statistic indicates that the country fixed effects are in fact not that important, which is also obvious comparing the results in columns (1) and (2) with the rest.

Table V

In Table C1 in Appendix C. we report estimates of the constant-beta model in which factors are GDP-weighted rather than value-weighted. The estimated betas remain largely unaltered in terms of sign, statistical significance and economic magnitude relative to our base estimates, suggesting that our results do not hinge upon a particular weighting scheme.

Table VI reports estimates of the constant-beta model obtained by OLS on the full sample (in column 1) together with estimates obtained on alternative subperiods, namely: pre-1913, 1914-1990, 1990-2014, 1990-2006 and 2007-2014 (in columns 2 to 6). The time variation in betas confirm the swoosh pattern we detected with our dynamic model. The estimate for the unconditional global beta prior to 1913 is close to the full sample estimate, at about 0.7. It is 25% larger than the estimate for the period between 1914 and 1945 (i.e. roughly 0.5), which is consistent with the decline in global financial market integration in the interwar period noted by previous scholars, and the adoption by several countries of protectionist and capital control measures then. Global financial market integration picks up between 1945 and 1990, with the estimated global beta increasing to 0.6, albeit still remaining lower than prior to World War I. It is only after 1990 that global financial market integration exceeds pre-1913 levels. Our global beta estimate reaches indeed about 0.9 for the 1990-2006 period, and 1.1 for the 2007-2014 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 exceed 0.50 in the post-1990 period.

Table VI

A formal test of the constant-parameter model is in Table VII, which presents estimates of model equation (6) with three period dummies (pre-1913, 1914-1990 and 1990-2014). A Wald test overwhelmingly rejects the null of equality of the global and regional beta coefficients over the three subperiods (see the first row of Panel C in Table VII). Parenthetically, there is not only substantial heterogeneity in betas over time, but also across countries. This is suggested in Figure B3 in Appendix B which shows country-by-country estimates over the full sample of the global and regional market factor betas obtained from the simple constant-beta model. 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. Regional betas are as low as almost zero for Australia and as high as 0.6 for the US or Germany.

B.2. Testing Temporal Patterns of Financial Market Integration

We now test formally whether financial market integration follows a U shape, a J (or L-inverted) shape, or a swoosh shape, along the lines described in Section II.A, which requires the testing of equality and inequality restrictions.

The results of the corresponding Wald restriction tests, based on the estimates of model equation (6) on the full sample, are in Table VII (see the second to fourth row of Panel C). We can reject the null hypotheses that de facto global financial market integration over the last century is characterised by a U-shape process, as hypothesised by e.g. Bordo and Flandreau (2003), Obstfeld and Taylor (2003, 2004), and Goetzmann et al. (2005), or by a J-shape process, as posited by Volosovych (2011). However, we cannot reject the hypothesis that the temporal pattern follows a swoosh shape. In other words, 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 during most of the twentieth century, namely between 1914 and 1990.

Interestingly, the results for de facto regional financial market integration are different. Although we can clearly reject the hypothesis that its temporal pattern follows a flat line, we fail to reject either the U, J or swoosh shapes. Nevertheless, the coefficient pattern is numerically consistent with a “swoosh”. It is also possible that the temporal pattern followed by regional financial market integration in the last century is different from the global pattern, but we lack statistical precision and power to distinguish different shapes.

Table VII

We also formally test whether effective financial market integration has partly reversed since the onset of the global financial crisis in 2007, as some observers have argued recently in the face of the capital controls and other financial protectionist measures taken by advanced and emerging market economies (in line with evidence in e.g. by Ostry et al. (2012), Forbes, Fratzscher and Straub (2013), Pasricha et al. (2015), and many others) and lingering public interventions in the financial sector of advanced economies, such as bank nationalizations, aimed at influencing the quantity and/or price of loans that banks from one country make to borrowers resident in another country (as stressed e.g. by Rose and Wieladek (2014)). This question is addressed in Table VIII which presents estimates of model equation (6) with two period dummies (1990-2007 and 2007-2014). Wald tests overwhelmingly reject the hypothesis that the global betas in the two subperiods are equal. They also reject the hypothesis that the precrisis beta is higher than the postcrisis 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) reported in Table C2 in Appendix C. The hypothesis that the global betas in subperiods 2007-2009 and 2010-2014 are equal is not rejected, however.

Table VIII

III. 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 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, it has been argued that the classic trilemma had morphed into a "dilemma" and the impossible trinity into an "irreconcilable duo". Central banks outside the U.S., the world's foremost financial centre, would 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 impulses (see Rey (2013), Miranda-Agrippino and Rey (2014), and Passari and Rey (2015)).[7] 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 has triggered strong interest in testing for the trilemma empirically (see e.g. Miniane and Rogers (2007), Bluedorn and Bowdler (2010), Klein and Shambaugh (2013), Aizenmann, Chinn and Ito (2014), Pasricha et al. (2015), Obstfeld (2015)).

Note that the trilemma hypothesis regards general capital mobility, which includes bond and equity flows, whereas our measure of financial market integration is based on equity market data only. While equity and bond flows may not be perfectly correlated, it is typically the case that international capital restrictions are comprehensive and apply uniformly to bond and equity markets (see Fernández, Klein, Rebucci, Schindler and Uribe (2015) for an analysis of recent data). In fact, many countries were either largely open or closed in terms of de jure capital market openness over several periods. We use this information in computing our de facto integration measure. It would be interesting to investigate bond market integration as well, but government bond markets were in

[7] 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.

several cases less well developed and liquid in the early part of our sample than were equity markets (although corporate debt markets were in some cases quite developed).

A. Testing the Trilemma Hypothesis

As a starting point, we rely on Obstfeld, Shambaugh and Taylor's benchmark regression model (which is also in the spirit of the specification used by Shambaugh (2004), Klein and Shambaugh (2013), and Obstfeld (2015))

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

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 (see Obstfeld, Shambaugh and Taylor (2005), p. 427). 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 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.[8] We take the U.K. as the base country prior to 1914 (classical gold standard); the mean of the U.K. and the U.S. as the base for the 1920s (gold exchange standard); 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

[8] 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), peggers vs. nonpeggers, countries highly integrated into global finance vs. countries highly segmented from global finance. Nonpeggers are defined as countries having floating or freely falling exchange rates according to the updated classification of Ilzetzki, Reinhart, and Rogoff (2004). Peggers 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).

countries, respectively, for the period 1973-1999; and the U.S. for all countries for the period 1999-2014.[9] As measures of interest rates we use both short-term policy rates and long-term government bond yields. See Appendix A for details on the data.

Next we modify equation model (7) to the form

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

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 *peg* is a binary dummy variable which equals one if country *i* in period *t* is a pegger 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.[10] From equation (8) one can derive the elasticity of the domestic interest rate with respect to the base interest rate for both peggers and nonpeggers, respectively, conditional on the degree of global financial market integration, namely:

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

[9] 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; more on this below). 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 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.

[10] Obstfeld, Shambaugh and Taylor (2005) relied on more limited data on capital market openness and assumed 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 was 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.

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

If a country maintains a peg, capital controls are necessary to maintain monetary policy independence. That is, we expect $d_1 > 0$. 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 pegged 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.

B. Empirical Results

We first review unconditional estimates of our test of the trilemma hypothesis in Section III.B.1., then condition the estimates on our measure of de facto global financial market integration in Section III.B.2., and finally test the trilemma hypothesis against the dilemma hypothesis in Section III.B.3.

B.1. Unconditional estimates

As a starting point, Table IX 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 (peggers vs. nonpeggers; 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.[11] 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 standard and post-Bretton Woods era, with estimates for the coefficient *b* of 0.19 and 0.57, 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 peggers and nonpeggers, insofar as the *b*-estimate, at about 0.30, is virtually identical for both groups of countries.[12] 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).

Table IX

The results for long-term interest rates are in line with those for short-term policy rates. The estimates are somewhat larger in economic magnitude, which points to a stronger degree of transmission of base interest rate movements to domestic interest rates at the long end of the yield curve relative to the short end, a finding consistent with Obstfeld (2015). Our estimate of *b* for the full sample is about 0.43.[13] We again find that interest rate pass-through is higher during the classical gold standard and post-Bretton Woods era, with *b*-estimates of 0.25 and 0.53, respectively, than during the Bretton Woods regime, with a *b*-estimate of about 0.12. The magnitude of the differences in interest rate pass-through between peggers and nonpeggers (0.46 vs. 0.40) is economically small for long-

[11] 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.

[12] This is unlike Obstfeld, Shambaugh and Taylor's estimates, which point to differences between peggers (0.43) and nonpeggers (0.26).

[13] This is somewhat lower than estimates for the recent period. For instance Hofmann and Takáts (2015) find an estimate of 0.59 in a panel of 30 emerging market and advanced economies over the period 2000-2014.

term interest rates, too. Again, a higher degree of global financial market integration results in higher pass-through estimates, i.e. 0.50 for integrated countries relative to 0.29 for segmented countries.

B.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 IX (where the latter column considers a specification excluding the world war periods). The estimates for long-term interest rates are reported in the corresponding columns of Table X.

Table X

Consider short-term interest rates first. 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. 0.18 (0.16 excluding World War I and II) versus 0.30 (see column 9 (10) of Table IX). It is still significantly different from zero at the 10% level, however. 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 c_2 is positive (only significant at the 15% level), that is, financially open countries experience more pass-through, and d_1 is positive and statistically significant at the 5% level, suggesting that dependence on openness is more pronounced for countries with pegged exchange rate systems.

How large are these effects economically? Figure 7 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 peggers and nonpeggers, as implied by the full sample estimates reported in column (10) of Table IX (see Panel A of Figure 7). First, the extent of financial market integration has little effect on pass-through coefficients for nonpeggers; 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, as floating exchange rates should suffice to guarantee monetary policy independence. Second, for peggers financial market integration dependence is much more pronounced. Specifically, pass-through of peggers 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 peggers 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 1% level, shows that financial integration dependence is stronger for peggers than nonpeggers in a statistically significant manner. Third, pass-through of nonpeggers well integrated into global finance is only half as large as that of peggers. 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. The difference is positive and statistically significant for values of global financial market integration above 1.17 (i.e. 17% above the average across countries and time).[14] This comprises roughly 30% of the observations in our sample, i.e. those mainly immediately before World War I and after 1985 (basically the two waves of global financial integration highlighted in the literature).

Figure 7

Consider now long-term interest rates. The full sample estimates for the direct pass-through effect of base-country long-term interest rates to domestic interest rates, b_1, is similar in economic magnitude to the unconditional estimate, i.e. about 0.45 (see columns 9 and 10 of Table X). Pass-through from base to domestic long-term interest rates again depends on whether an economy is open

[14] The difference between peggers and nonpeggers is significant at the 10% level for values of $\overline{\beta}_{i,t}^{glo}$ such that $(\hat{c}_1 + \overline{\beta}_{i,t}^{glo}\hat{d}_1)/[(Var(\hat{c}_1) + \overline{\beta}_{i,t}^{glo^2}Var(\hat{d}_1) + 2\overline{\beta}_{i,t}^{glo}Cov(\hat{c}_1, \hat{d}_1)]^{1/2} > 1.65$ i.e. $\overline{\beta}_{i,t}^{glo} > 1.17$ given our sample estimates of 0.0272, 0.0316 and -0.0288 for $Var(\hat{c}_1), Var(\hat{d}_1)$ and $Cov(\hat{c}_1, \hat{d}_1)$.

to global finance or closed, and on whether it has pegged or flexible exchange rates, in line with the trilemma hypothesis, as evidenced by the statistically significant interaction effects. The only exception is the interaction between interest rates and de facto global financial market integration. Panel B of Figure 7 sheds light on the economic magnitude of these effects. First, the extent of financial market integration has no effect on pass-through coefficients for nonpeggers; pass-through remains constant, at 0.45, and does not depend on the extent of de facto global financial market integration. This is again consistent with the trilemma hypothesis, as floating exchange rates should suffice to guarantee monetary policy independence although the degree of pass-through remains relatively high. Second, for peggers financial market integration dependence is again much more pronounced in a statistically significant manner as evidenced by the estimate for the triple interaction coefficient d_1. Pass-through to domestic long-term interest rates of peggers well integrated into global finance is even higher than for short-term policy rates, at about 0.75. Segmentation from global finance again dampens pass-through: for peggers which are largely or fully closed to global finance, pass-through is nil or even negative, as predicted by the trilemma hypothesis. The difference is positive and statistically significant for values of global financial market integration above 1.26 (i.e. 26% above the average across countries and time). This comprises roughly 20% of the observations in our sample. The observations featuring such high degrees of freedom correspond mostly to the second wave of globalization, post-1985.

How has interest rate pass-through evolved over the last century? Figure 8 shows the evolution between 1885 and 2014 of the average pass-through estimates from base to short-term (Panel A) and long-term (Panel B) interest rates for peggers and nonpeggers, respectively, as predicted by the full sample estimates reported in column (10) of Table IX and X.[15] Pass-through on the short end of the yield curve for peggers 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 nonpeggers is more stable over time, in contrast. Interestingly, short-term interest rate pass-through

[15] There are no estimates for peggers 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 nonpeggers as they had either suspended gold convertibility or, in the case of Spain, were previously not on gold.

remains appreciably higher for peggers than for nonpeggers in the modern era of financial globalization, at about 0.50 and 0.30, respectively, on average in the 2000s (see Panel A of Figure 8). 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. Long-term interest rate pass-through for nonpeggers follows a nice swoosh shape, too, while pass-through for nonpeggers is stable over time. In addition, long-term interest rate pass-through is also much higher for peggers than for nonpeggers in the modern era of financial globalization, at about 0.70 versus 0.45, respectively, on average in the 2000s (see Panel B of Figure 8).

Figure 8

B.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.

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, peggers 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: peggers can achieve a high degree of monetary 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 Figure 7).

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 7 further shows, we therefore do observe a positive non-negligible pass-through for floating-exchange-rate countries for both short and long-term interest rates, consistent with the dilemma hypothesis.[16] However, for the most integrated countries and the largest part of our sample period, our model predicts significantly more pass-through for peggers than for non-peggers, 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 Table XI). We do find that a likelihood ratio test 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

[16] For short-term interest rates, this is purely as an artefact of the higher degree of financial globalization.
[17] In particular, 38% of the observations are peggers post-1990, against 74% pre-1990; for instance in Europe peggers comprise EU currencies managed within the Exchange Rate Mechanism in the 1990s, versus only Denmark and Switzerland in the 2000s.

dilemma hypothesis and very much confirm the identification problems discussed above. For example, for short-term interest rates, only one parameter change is significant at the 10% level. But the change indicates that for countries with pegged exchange rates, financial integration decreases pass-through, which makes little sense. For both short and long term interest rates, we find that the b_1 coefficient increases (overall pass-through), which is true in a statistically significant manner for long-term interest rates, 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.

Table XI

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 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 Tables IX-X, namely peggers versus non-peggers 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 Table XII). 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 peggers and for non-peggers and for both short-term and long-term interest rates. For short-term interest rates, the estimate is 0.437 for peggers and 0.323 for non-peggers. 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

[18] As aforementioned in footnote 9, 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).

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.

Table XII

IV. 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.

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 peggers, 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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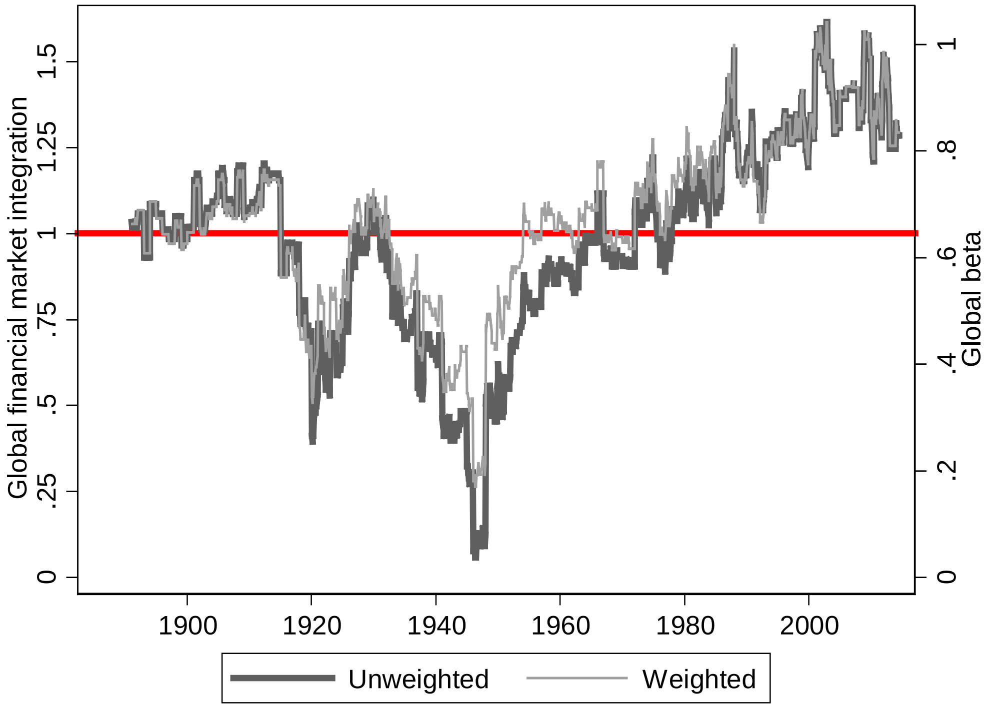
Panel A. Global Financial Market Integration and Conditional Global Betas



Panel B. Regional Financial Market Integration and Conditional Regional Betas

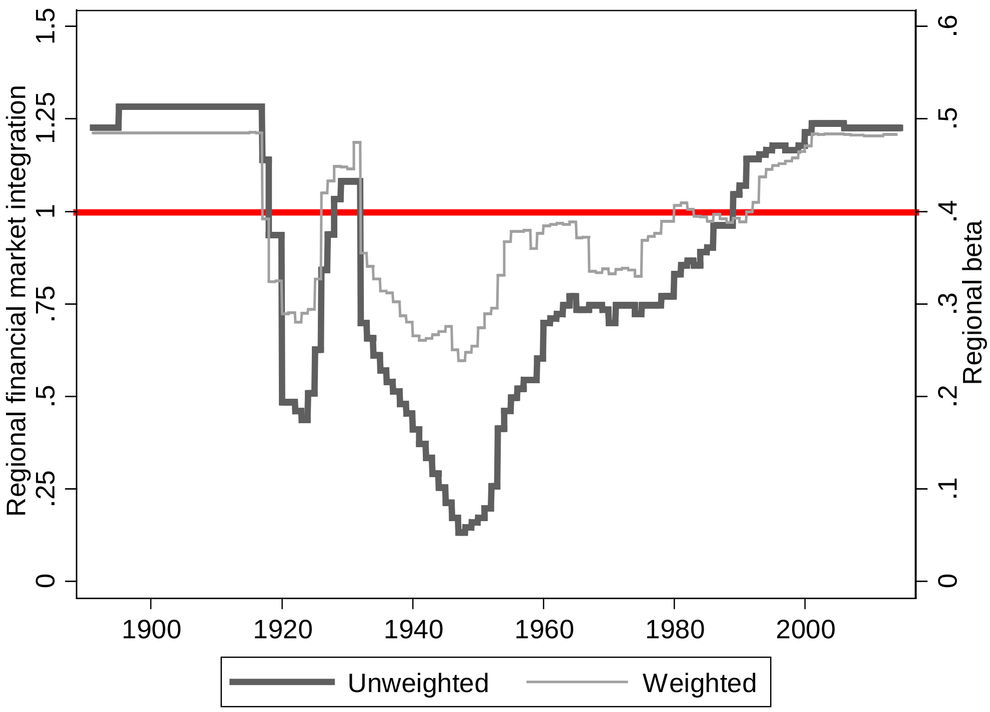


Figure 1. Average conditional betas – 1885-2014. The figure shows the evolution between 1885 and 2014 of the unweighted (thick grey lines) and value-weighted (light grey lines) averages (17 countries) of our measures of financial market integration (defined in equations (4) and (5)) and corresponding conditional beta estimates obtained from the full model equations (1) to (3). Global estimates are shown in Panel A and regional estimates in Panel B.

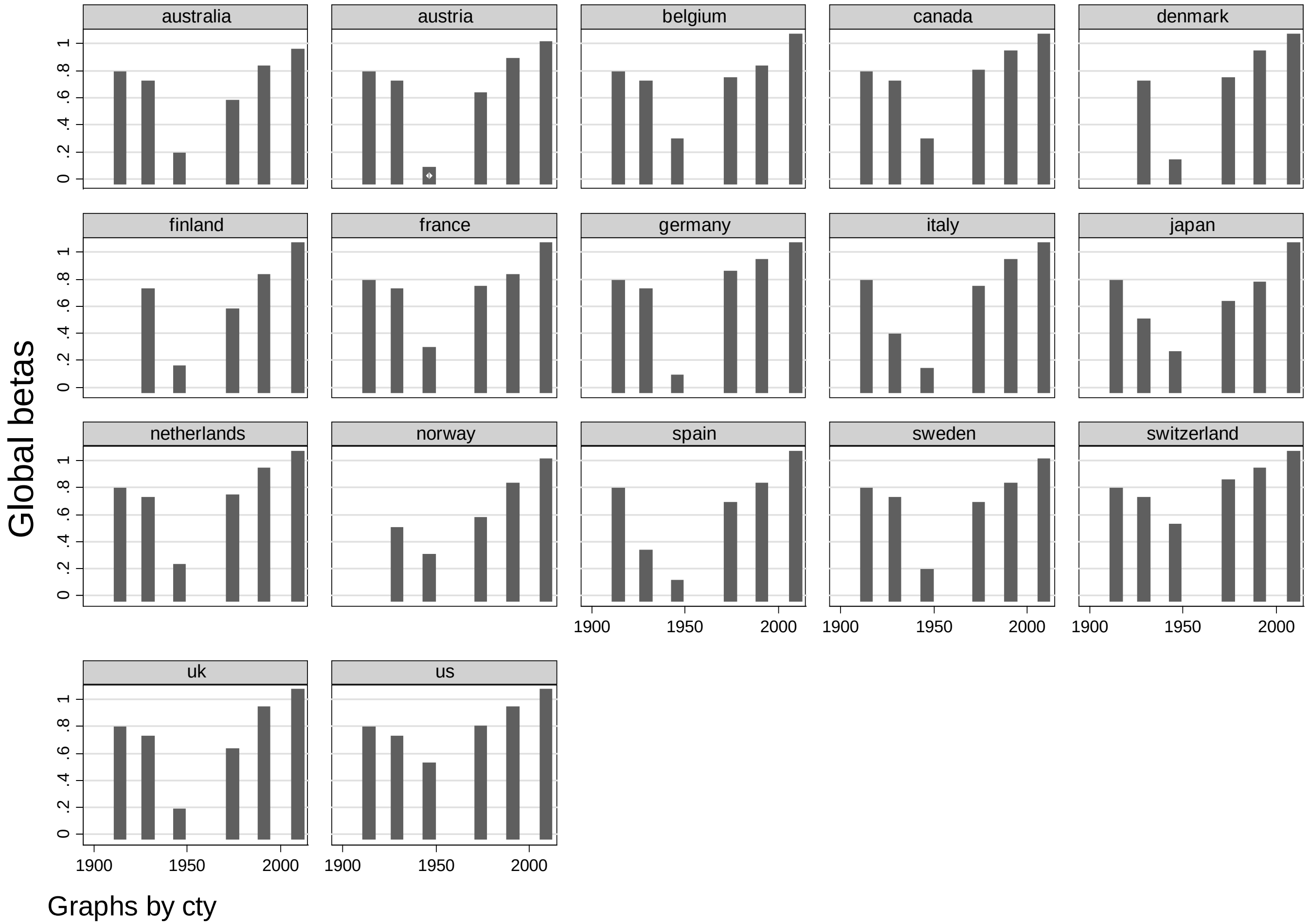


Figure 2. Conditional global betas – Breakdown by country and selected years. The figure shows the conditional global beta estimates obtained from the full model equations (1) to (3) for each country of our sample in selected years, namely: 1913, 1928, 1945, 1973, 1990, and 2008.

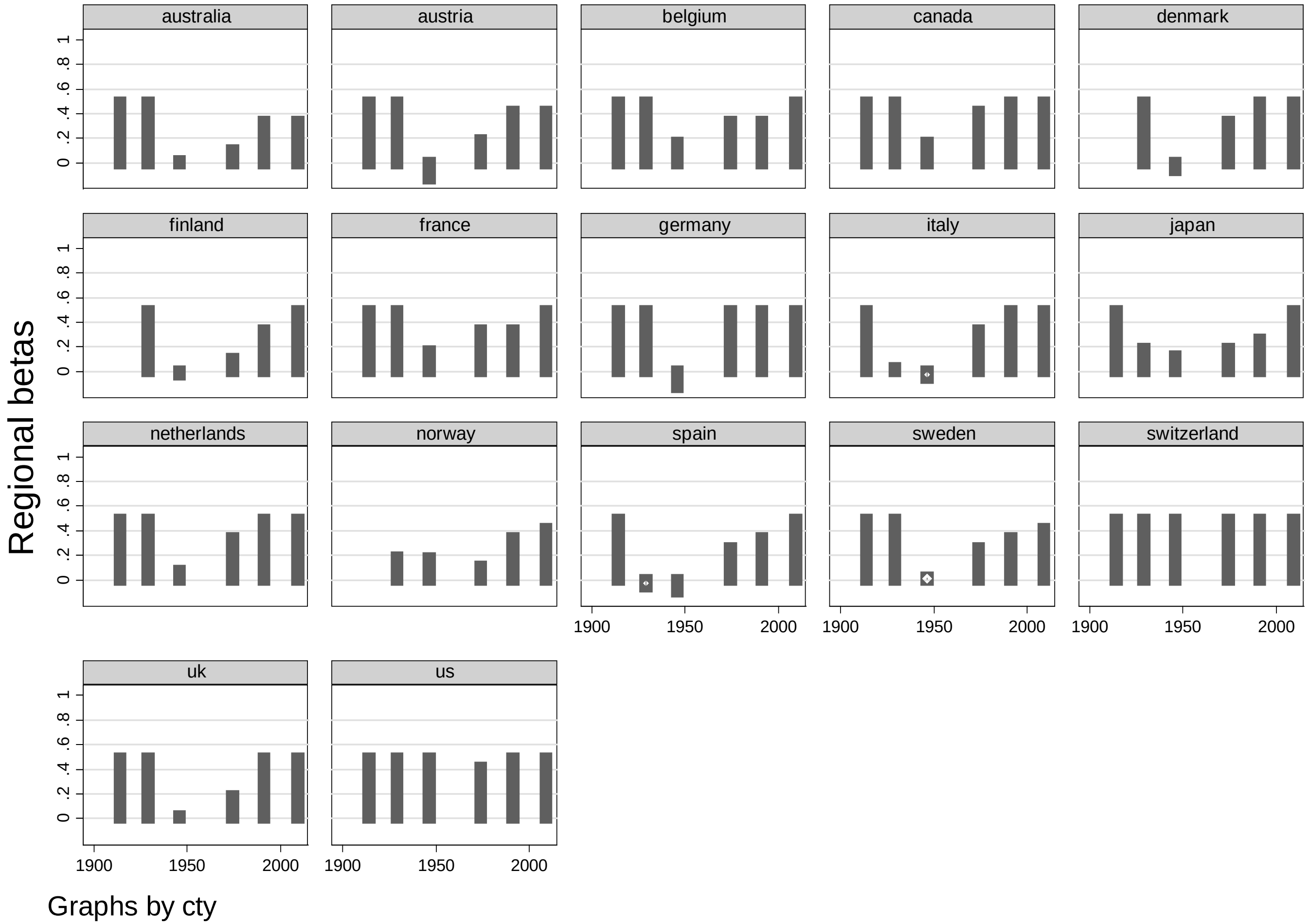


Figure 3. Conditional regional betas – Breakdown by country and selected years. The figure shows the conditional regional beta estimates obtained from the full model equations (1) to (3) for each country of our sample in selected years, namely: 1913, 1928, 1945, 1973, 1990, and 2008.

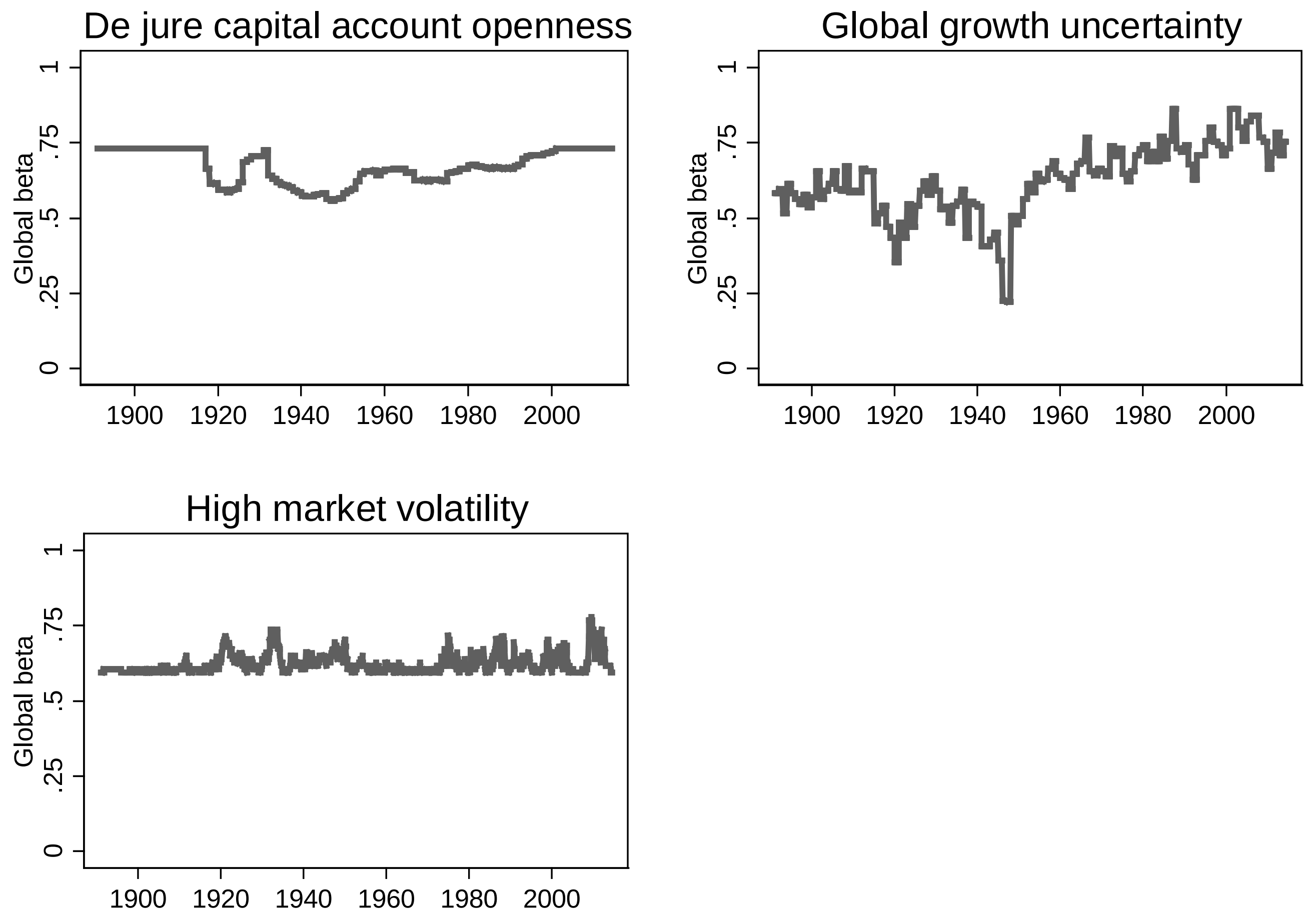
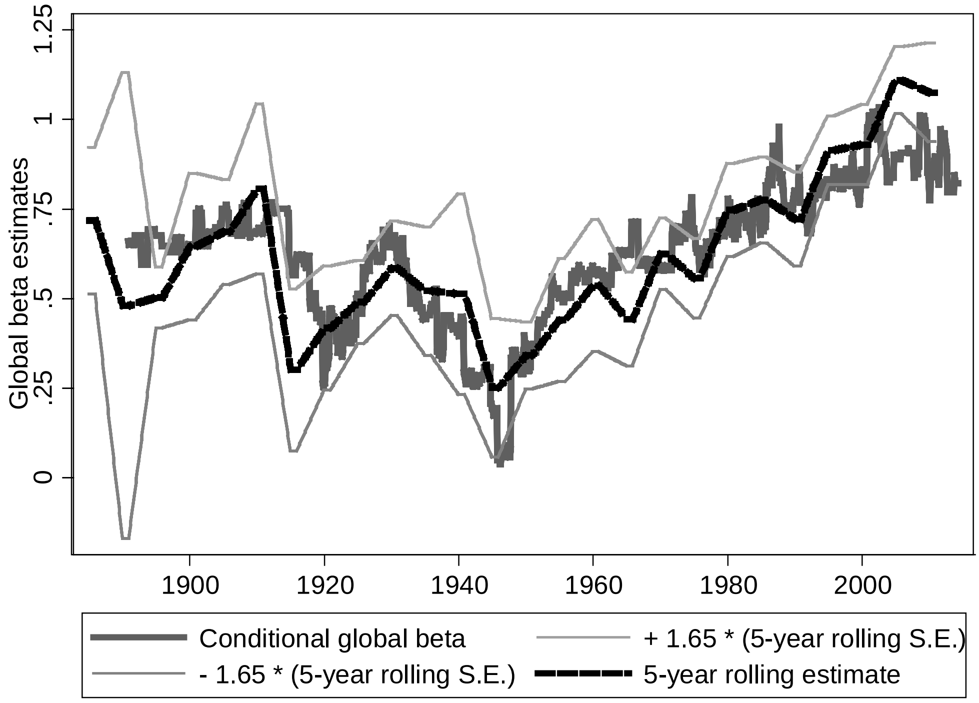


Figure 4. Economic importance of the instruments for global financial market integration – 1885-2014. The figure shows the evolution between 1885 and 2014 of the value-weighted averages (17 countries) of the estimates of the conditional global betas obtained from the full model equations (1) to (3) when only one of each instrument in vector **X** remains active, in turn. This is achieved by setting the loadings on the other instruments at their respective means.

Panel A. Global Betas – Conditional vs. Rolling



Panel B. Regional Betas – Conditional vs. Rolling

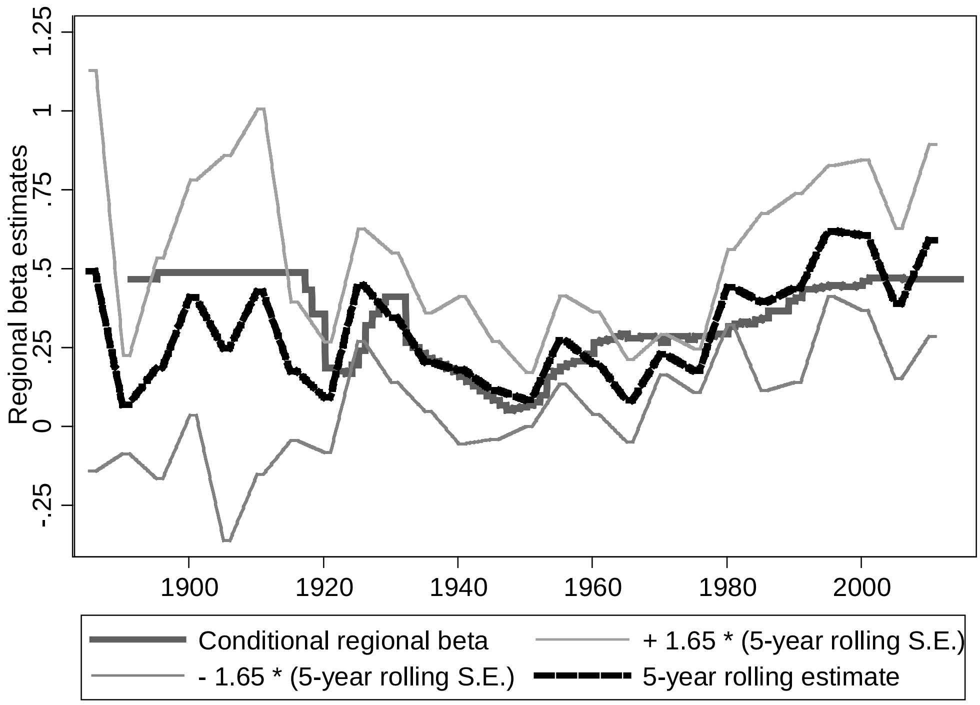
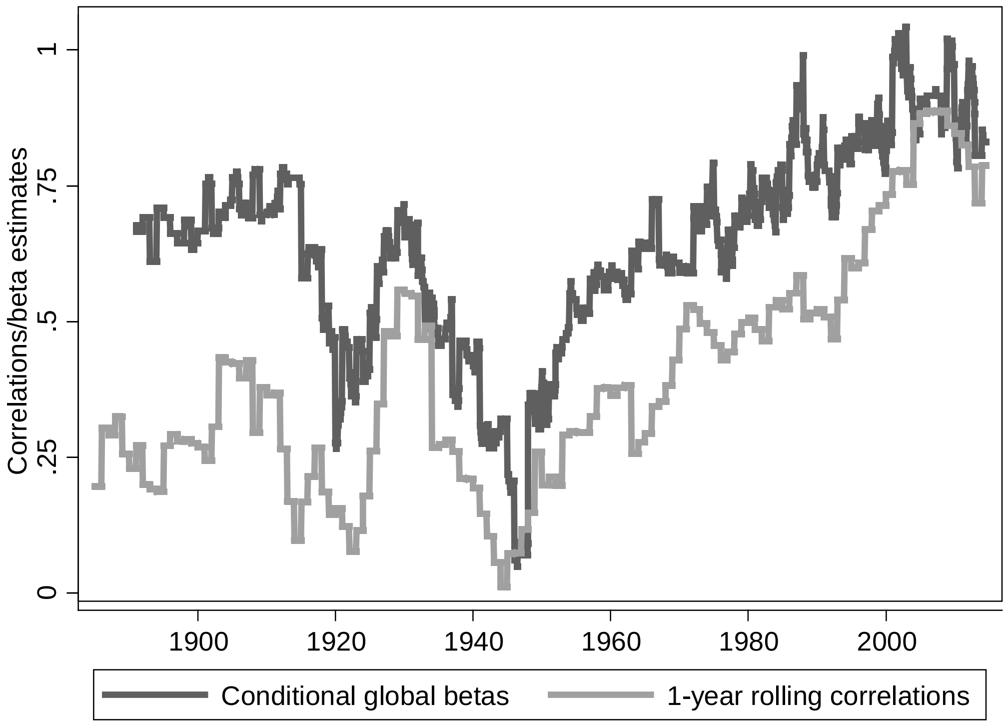


Figure 5. Average conditional betas vs. 5-year rolling betas – 1885-2014. The figure shows the evolution between 1885 and 2014 of the unweighted averages (17 countries) of the estimated cross-sectionally heterogeneous and time-varying parameters $\beta_{i,t}^{glo}$ and $\beta_{i,t}^{reg}$ (thick grey 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 β_t^{reg}, and the point estimates (shown as black dashed lines).

Panel A. Global Betas vs. Correlations



Panel B. Regional Betas vs. Correlations

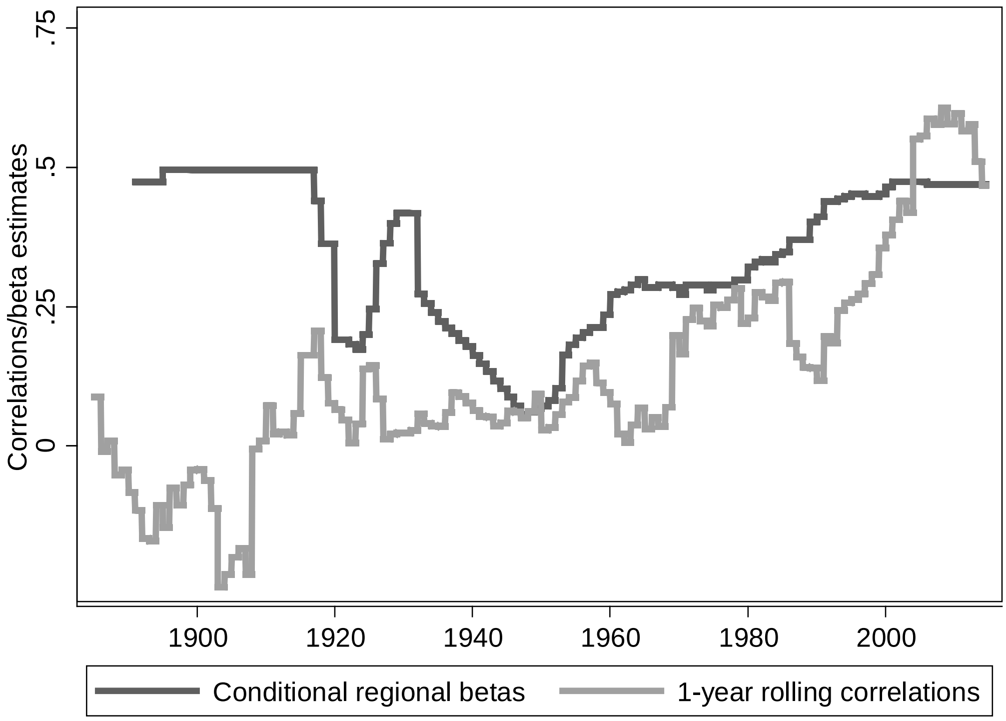
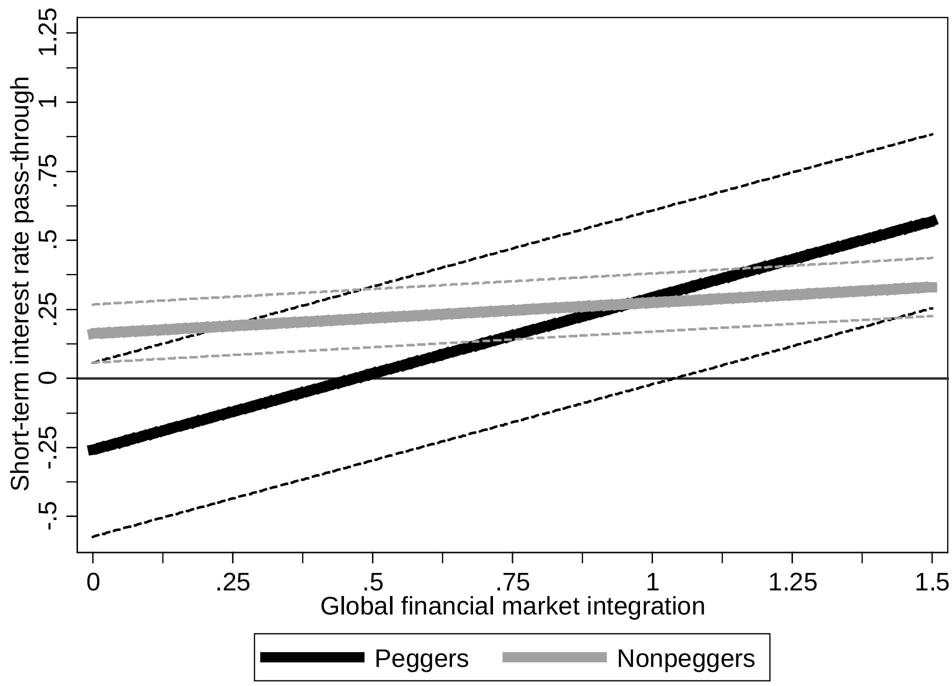


Figure 6. Average conditional betas vs. 1-year rolling correlations – 1885-2014. The figure shows the evolution between 1885 and 2014 of the unweighted averages (17 countries) of the conditional global and regional beta estimates (thick grey lines) together with 1-year rolling (non-overlapping) bilateral pooled correlations between equity excess returns $R_{i,t}$ and the global and regional market factors, respectively.

Panel A. Short-Term Interest Rates



Panel B. Long-Term Interest Rates

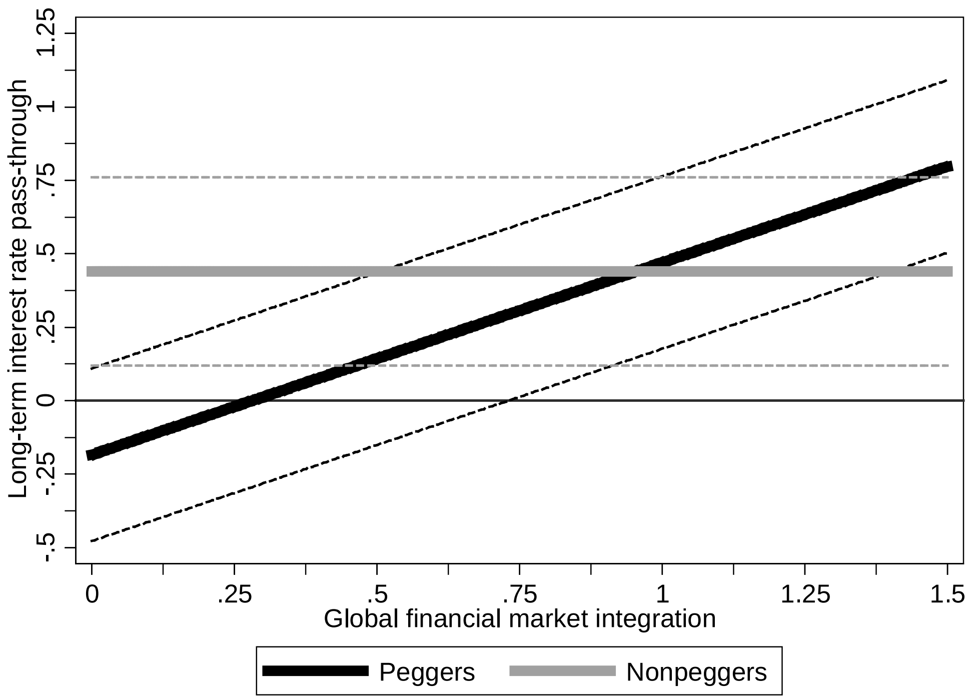
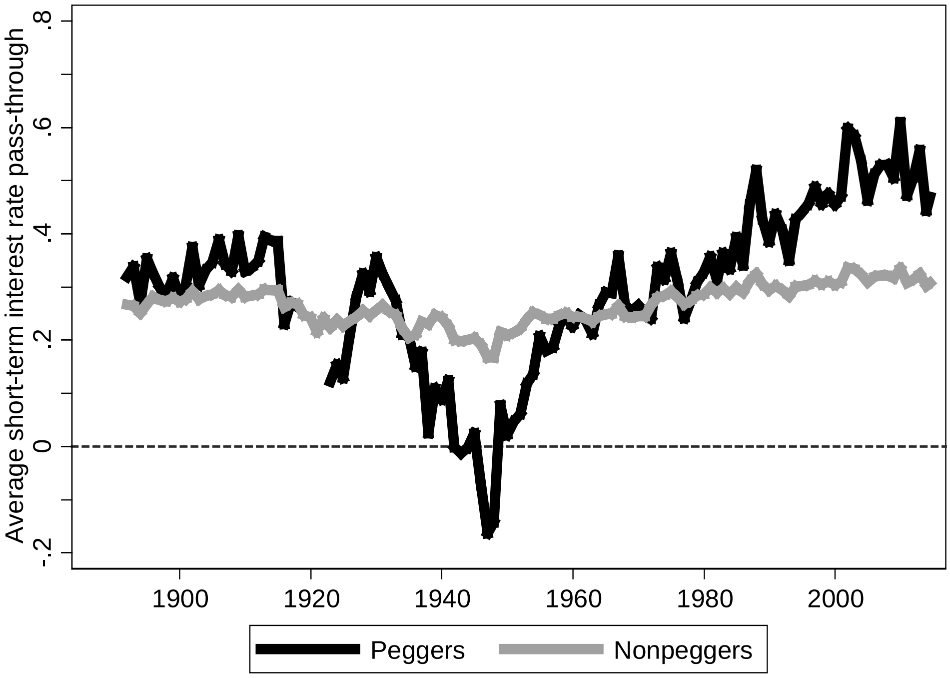


Figure 7. Estimated interest rate pass-through vs. global financial market integration. The figure plots the estimated pass-through from base (i.e. US, UK or German) short-term (Panel A) and long-term (Panel B) interest rates to domestic interest rates against the extent of global market integration for both peggers and nonpeggers as predicted by the full sample estimates reported in column (10) of Table IX and X, respectively. 90% confidence bands are shown as dotted lines.

Panel A. Short-Term Interest Rates



Panel B. Long-Term Interest Rates

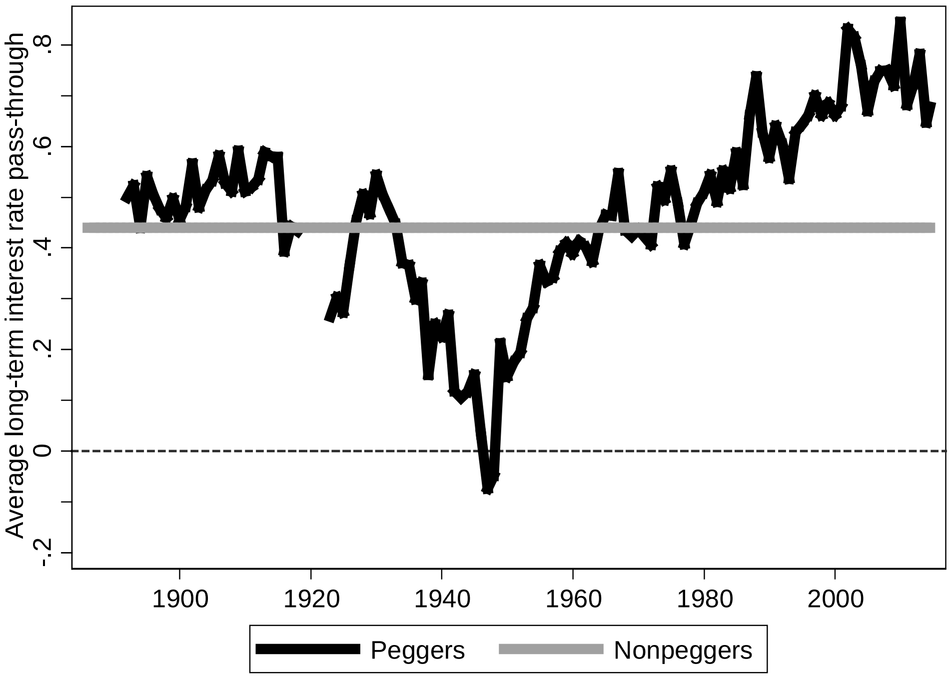


Figure 8. Estimated interest rate pass-through: peggers vs. nonpeggers – 1885-2014. The figure shows the evolution between 1885 and 2014 of the average (14 countries) pass-through estimates from base (i.e. US, UK or German) short-term (Panel A) and long-term (Panel B) interest rates to domestic interest rates for peggers and nonpeggers as predicted by the full sample estimates reported in column (10) of Table IX and X, respectively.

Table I
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	1.13	1.22	0.56	0.21	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
Peggers	Dummy variable (0/1)	Annual	Dummy variable which equals zero for nonpeggers (floats, managed floats or freely falling exchange rates) and one for peggers (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 II
Pairwise correlations of the instruments

The table reports the pairwise correlations of the variables contained in vector **X**. The local growth uncertainty variable reported below is the logarithm of the standard deviation of real GDP growth in each country over non-overlapping windows of 5 years. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

	Trade openness	Regional trade	Capital account openness	Financial depth	Oil prices	Global growth uncertainty	Local growth uncertainty	High market volatility periods
Trade openness	1.00							
Regional trade	0.43 ***	1.00						
Capital account openness	0.30 ***	0.02 ***	1.00					
Financial depth	0.15 ***	-0.19 ***	0.30 ***	1.00				
Oil prices	0.11 ***	0.01 **	0.06 ***	-0.01	1.00			
Global growth uncertainty	-0.23 ***	-0.09 ***	-0.37 ***	-0.10 ***	-0.06 ***	1.00		
Local growth uncertainty	-0.20 ***	-0.20 ***	-0.26 ***	-0.10 ***	-0.03 ***	0.59 ***	1.00	
High market volatility periods	0.06 ***	-0.02 ***	-0.13 ***	0.02 ***	-0.16 ***	0.13 ***	0.13 ***	1.00

Table III
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 clustered heterogeneity. ***, **, *, 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)
Country fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES
Year effects	YES	YES	YES	YES	YES	YES	YES	YES	YES
Instrument main effects	YES	YES	YES	YES	YES	YES	YES	YES	YES
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 IV
Variance Ratio Analysis

The table reports the estimates of a variance ratio analysis. For each statistically significant instrument j (i.e. de jure capital account openness, global growth uncertainty and high market volatility periods) of our parsimonious specification, we calculate the variance ratio for the conditional global beta estimates as

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

and where the standard errors are estimated by bootstrap with 1,000 replications. Each replication in the bootstrap consists in scrambling the residuals from the parsimonious model (shown in column 9 of Table III), recreating the dependant variable using the data's independent variables and the estimated parameters, and re-estimating the model. We then re-compute the variance ratios as indicated above. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

	Capital account openness	Global growth uncertainty	High market volatility periods
Variance contribution	0.528 ***	0.471 ***	0.001
	(0.060)	(0.059)	(0.007)

Table V
Constant-Beta Model Estimates

The table reports the estimates of the following model:

$$R_{i,t} = \alpha_i + \lambda_t + \beta_0^{glo} F_t^{glo,\backslash i} + \beta_0^{reg} F_t^{reg,\backslash i} + \varepsilon_{i,t}$$

Estimates obtained by OLS are reported in columns 1 to 3; estimates obtained by excluding outliers from the sample (i.e. returns larger than 30% within a month in absolute value) are reported in column 4; estimates obtained with a random effects estimator are reported in columns 5. The estimates of columns 2 to 5 control for country fixed (or random) effects. The estimates of columns 3 to 5 control for year effects. The ρ-statistic in column (5) is the percent contribution to the total variance of the panel-level variance component. The standard errors reported in parentheses are robust to heteroskedasticity and clustered heterogeneity. ***, **, *, and + indicate statistical significance at the 1%, 5%, 10% and 15% levels, respectively.

	(1) OLS	(2) OLS	(3) OLS	(4) Outliers	(5) Random effects
Global factor	0.707***	0.707***	0.679***	0.665***	0.679***
	(0.031)	(0.032)	(0.033)	(0.032)	(0.033)
Regional factor	0.300***	0.300***	0.286***	0.287***	0.286***
	(0.081)	(0.081)	(0.083)	(0.081)	(0.084)
Constant	0.018	-0.007	0.315**	-0.336***	0.471*
	(0.033)	(0.017)	(0.132)	(0.089)	(0.254)
Country fixed/random effects	NO	YES	YES	YES	YES
Year effects	NO	NO	YES	YES	YES
Observations	20,800	20,800	20,800	20,719	20,800
R^2	0.199	0.199	0.209	0.285	0.209
# panel units					17
ρ					0.000

Table VI
Estimates for Alternative Sample Periods

The table reports the estimates of the following model:

$$R_{i,t} = \alpha_i + \lambda_t + \beta_0^{glo} F_t^{glo,\setminus i} + \beta_0^{reg} F_t^{reg,\setminus i} + \varepsilon_{i,t}$$

Estimates obtained by OLS on the full sample are reported in column 1; estimates obtained on alternative subperiods (pre-1913, 1914-1990, 1990-2014, 1990-2006 and 2007-2014) are reported in columns 2 to 6. The estimates control for country fixed effects and for year effects. The standard errors reported in parentheses are robust to heteroskedasticity and clustered heterogeneity. ***, **, *, and + indicate statistical significance at the 1%, 5%, 10%, 15% levels, respectively.

	(1) Full sample	(2) Pre-1913	(3) 1914-1990	(4) 1990-2014	(5) 1990-2006	(6) 2007-2014
Global factor	0.679***	0.669***	0.536***	0.943***	0.860***	1.091***
	(0.033)	(0.055)	(0.046)	(0.041)	(0.049)	(0.064)
Regional factor	0.286***	0.357+	0.212**	0.586***	0.558***	0.502***
	(0.083)	(0.176)	(0.076)	(0.145)	(0.146)	(0.163)
Constant	0.315**	-0.184	0.210	-0.283	0.937***	0.523*
	(0.132)	(0.119)	(0.944)	(0.418)	(0.222)	(0.248)
Country fixed effects	YES	YES	YES	YES	YES	YES
Year effects	YES	YES	YES	YES	YES	YES
Observations	20,800	1,071	14,748	4,981	3,468	1,513
R^2	0.204	0.177	0.119	0.598	0.473	0.783

Table VII
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 clustered heterogeneity. ***, **, *, 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.

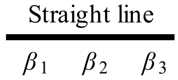
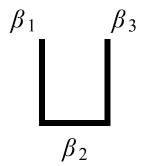
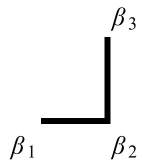
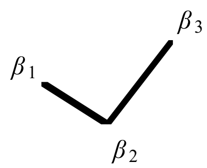
Panel A. **Full sample estimates (1885-2014)**		**Panel B.** **Hypothesized shape**	**Panel C.** ***p*-value of Wald test**
Global factor × D_1	0.669***	Straight line 	$\beta_1^{glo} = \beta_2^{glo} = \beta_3^{glo}$: 0.000 $\beta_1^{reg} = \beta_2^{reg} = \beta_3^{reg}$: 0.009
	(0.051)		
Global factor × D_2	0.536***		
	(0.046)	U-shape 	$\beta_1^{glo} = \beta_3^{glo}$: 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
Global factor × D_3	0.943***		
	(0.041)		
Regional factor × D_1	0.355**		
	(0.160)		
Regional factor × D_2	0.212**	J (or inverted L)-shape 	$\beta_1^{glo} = \beta_2^{glo}$: 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
	(0.076)		
Regional factor × D_3	0.585***		
	(0.146)		
Constant	0.338**		
	(0.118)		
Year effects	YES	Swoosh-shape 	$\beta_1^{glo} > \beta_2^{glo}$: 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
Country fixed effects	YES		
Observations	20,800		
R^2	0.221		

Table VIII
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 estimates, obtained by OLS, control for country fixed effects and for year effects. The standard errors reported in parentheses are robust to heteroskedasticity and clustered heterogeneity. ***, **, *, 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)	$\beta_1^{reg} = \beta_2^{reg}$:	0.554
Constant	0.155	$\beta_1^{reg} > \beta_2^{reg}$:	0.723
	(0.119)	$\beta_1^{reg} < \beta_2^{reg}$:	0.277
Year effects	YES		
Country fixed effects	YES		
Observations	20,800		
R^2	0.155		

Table IX
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 clustered heterogeneity. ***, **, *, 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)	Peggers	Nonpeggers	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	0.305***	0.187***	0.096**	0.567***	0.306***	0.304***	0.393***	0.136*	0.176*	0.163*
	(0.040)	(0.052)	(0.035)	(0.118)	(0.049)	(0.035)	(0.032)	(0.077)	(0.082)	(0.084)
Global market integration									-0.021***	-0.026***
									(0.004)	(0.005)
Peg									-0.012***	-0.015**
									(0.004)	(0.006)
Interest rate × global market integration									0.105+	0.112+
									(0.063)	(0.064)
Interest rate × peg									-0.433**	-0.420**
									(0.165)	(0.166)
Global market integration × peg									0.013**	0.015**
									(0.006)	(0.006)
Interest rate × global market integration × peg									0.447**	0.439**
									(0.178)	(0.178)
Constant	-0.002***	0.002*	0.010***	-0.001	-0.002	-0.001	-0.007***	0.005***	0.019***	0.025***
	(0.000)	(0.001)	(0.003)	(0.002)	(0.001)	(0.003)	(0.001)	(0.002)	(0.003)	(0.004)
Observations	1,715	326	168	392	1,149	552	1,006	709	1,551	1,417
Adjusted R^2	0.0880	0.120	0.0658	0.0976	0.0842	0.0986	0.172	0.0123	0.0990	0.0995

Table X
Testing for the Monetary Policy Trilemma in History – Long-Term Interest Rates

The table reports the estimates of the parameters of the unconditional model equation (7) using long-term government bond yields. Column 1 reports pooled estimates, Columns 2 to 4 report estimates over three subperiods. 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 clustered heterogeneity. ***, **, *, and + indicate statistical significance at the 1%, 5%, 10%, and 15% levels, respectively.

	Unconditional estimates								Conditional estimates	
		By time period			By country group				1885-2014	
	Pooled	Gold standard (Pre-1914)	Bretton Woods (1959-1970)	Post-Bretton Woods (1973-2000)	Peggers	Nonpeggers	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	0.432***	0.250**	0.124+	0.529***	0.459***	0.404***	0.502***	0.286***	0.455**	0.440**
	(0.074)	(0.092)	(0.076)	(0.113)	(0.114)	(0.064)	(0.086)	(0.062)	(0.188)	(0.195)
Global market integration									-0.023***	-0.026***
									(0.006)	(0.007)
Peg									-0.025	-0.027+
									(0.017)	(0.018)
Interest rate × global market integration									-0.053	-0.043
									(0.145)	(0.151)
Interest rate × peg									-0.588**	-0.624***
									(0.219)	(0.189)
Global market integration × peg									0.030*	0.031*
									(0.016)	(0.016)
Interest rate × global market integration × peg									0.626**	0.654***
									(0.229)	(0.206)
Constant	0.000	0.006***	0.016***	0.003	0.001	-0.001	-0.000	0.002	0.021**	0.024**
	(0.001)	(0.001)	(0.004)	(0.004)	(0.002)	(0.002)	(0.001)	(0.002)	(0.007)	(0.008)
Observations	1,232	174	167	132	779	453	624	608	1,199	1,097
Adjusted R^2	0.139	0.0241	0.0101	0.369	0.111	0.191	0.298	0.0332	0.151	0.188

Table XI
Testing for the Monetary Policy Trilemma vs. Dilemma Hypotheses – Stability Tests

The table reports stability test results for a break in the coefficients of model equation (8) in 1990 and 2007, respectively, using two interaction dummy variables: the first dummy is denoted *prebreak* and equals one before 1990 (respectively 2007) and zero afterwards; the second dummy is denoted *postbreak* and equals zero before 1990 (respectively 2007) and zero afterwards. The results are obtained using the full sample and short-term policy rates (in columns 1 to 3) and long-term government bond yields (in columns 4 to 6). Columns 1 and 4 report the estimates without breaks, columns 2 and 5 the estimates with a break in 1990, and columns 3 and 6 the estimates with a break in 2007. The standard errors reported in parentheses are robust to heteroskedasticity and clustered heterogeneity. ***, **, *, and + indicate statistical significance at the 1%, 5%, 10%, and 15% levels, respectively.

	Short-term interest rates			Long-term interest rates		
	Full sample	Break in 1990	Break in 2007	Full sample	Break in 1990	Break in 2007
	(1)	(2)	(3)	(4)	(5)	(6)
Interest rate	0.176*			0.455**		
	(0.082)			(0.188)		
Global market integration	-0.021***			-0.023***		
	(0.004)			(0.006)		
Peg	-0.012***			-0.025		
	(0.004)			(0.017)		
Interest rate × global market integration	0.105+			-0.053		
	(0.063)			(0.145)		
Interest rate × peg	-0.433**			-0.588**		
	(0.165)			(0.219)		
Global market integration × peg	0.013**			0.030*		
	(0.006)			(0.016)		
Interest rate × global market integration × peg	0.447**			0.626**		
	(0.178)			(0.229)		
Interest rate × prebreak		0.100	0.190**		0.821***	0.283
		(0.192)	(0.083)		(0.259)	(0.188)
Global market integration × prebreak		-0.001	-0.010*		-0.015	-0.014*
		(0.010)	(0.005)		(0.013)	(0.007)
Peg × prebreak		-0.012**	-0.007*		-0.026	-0.024
		(0.005)	(0.004)		(0.021)	(0.017)
Interest rate × global market integration × prebreak		0.217	0.097+		-0.683**	0.122
		(0.286)	(0.061)		(0.254)	(0.152)
Interest rate × peg × prebreak		-0.520+	-0.469**		-1.037***	-0.493*
		(0.297)	(0.158)		(0.308)	(0.233)
Global market integration × peg × prebreak		0.008	0.003		0.031	0.026+
		(0.010)	(0.006)		(0.023)	(0.017)
Interest rate × global market integration × peg × prebreak		0.498	0.479**		1.333***	0.531*
		(0.380)	(0.171)		(0.367)	(0.265)
Interest rate × postbreak		0.289	-4.359+		2.813***	2.080**
		(0.956)	(2.816)		(0.841)	(0.906)
Global market integration × postbreak		-0.018***	-0.030***		-0.019**	-0.027***
		(0.004)	(0.005)		(0.007)	(0.007)
Peg × postbreak		-0.174***	0.242		-0.098*	-0.127*
		(0.052)	(0.235)		(0.050)	(0.067)
Interest rate × global market integration × postbreak		0.010	3.343+		-1.730**	-1.265*
		(0.644)	(2.017)		(0.577)	(0.599)
Interest rate × peg × postbreak		1.238	-6.453*		0.608	1.300
		(1.301)	(3.423)		(0.973)	(1.006)
Global market integration × peg × postbreak		0.119***	-0.160		0.072*	0.097*
		(0.038)	(0.179)		(0.035)	(0.047)
Interest rate × global market integration × peg × postbreak		-0.792	4.660*		-0.289	-0.731
		(0.892)	(2.524)		(0.654)	(0.698)
Constant	0.019***	0.009+	0.013***	0.021**	0.015+	0.016*
	(0.003)	(0.005)	(0.003)	(0.007)	(0.010)	(0.008)
Observations	1,551	1,551	1,551	1,199	1,199	1,199
Adjusted R^2	0.0990	0.107	0.104	0.151	0.182	0.158
log likelihood	1145	1155	1153	1533	1559	1542
p-value of likelihood ratio test		0.005	0.036		0.000	0.021
(H_0: Full sample model nested in alternative model with break)						

Table XII
Additional Tests of the Dilemma Hypothesis

The table reports the estimates of the parameters of the unconditional model equation (7) using short-term interest rates (Panel A) and long-term government bond yields (Panel B) using a sample restricted to observations post-1990 and for various country groups. The standard errors reported in parentheses are robust to heteroskedasticity and clustered heterogeneity. ***, **, *, and + indicate statistical significance at the 1%, 5%, 10%, and 15% levels, respectively.

	(1) Peggers	(2) Nonpeggers	(3) High financial integration	(4) Low financial integration	(5) High financial integration & peggers	(6) High financial integration & nonpeggers	(7) Low financial integration & nonpeggers
				A. *Short-term interest rates*			
Interest rate	0.442***	0.306***	0.364***	0.111	0.437***	0.323***	1.667
	(0.108)	(0.050)	(0.036)	(1.963)	(0.107)	(0.042)	(0.943)
Constant	-0.027***	-0.017**	-0.021***	-0.097	-0.027***	-0.015**	-0.292
	(0.006)	(0.006)	(0.003)	(0.245)	(0.007)	(0.005)	(0.118)
Observations	151	199	360	4	150	196	3
Adjusted R2	0.184	0.108	0.143	-0.499	0.181	0.121	0.351
Log likelihood	127.5	159.7	301	2.717	126.6	159.6	5.886
				B. *Long-term interest rates*			
Interest rate	0.653***	0.511***	0.547***	0.316	0.653***	0.510***	0.316
	(0.098)	(0.104)	(0.083)	(0.218)	(0.098)	(0.105)	(0.218)
Constant	-0.008**	-0.011***	-0.010***	-0.055	-0.008**	-0.011***	-0.055
	(0.002)	(0.001)	(0.001)	(0.025)	(0.002)	(0.001)	(0.025)
Observations	67	193	257	3	67	190	3
Adjusted R2	0.650	0.326	0.393	0.166	0.650	0.326	0.166
Log likelihood	117.3	253.2	356.3	8.488	117.3	248.4	8.488

Appendix A. Data definitions, coverage, and sources

This appendix outlines the data definitions, coverage, and sources. The sample includes 17 economies, namely: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Norway, Spain, Sweden, Switzerland, the U.K., and the U.S. The data are sampled at the monthly frequency. The baseline estimation period is January 1885 to June 2014. We have an unbalanced panel due to missing observations, with up to 1,554 observations per economy.

Equity prices: We take data on local equity market indices – initially gathered from national sources – from *Global Financial Data* (G.F.D.). G.F.D. assembled long time series using benchmark national indices (market capitalisation-weighted) at given points in time that are chain-linked with one another. The resulting long time series are expressed in nominal local currency terms. They are not systematically adjusted for dividend payments. The data are available since January 1885 for Australia, Germany, the U.K. and the U.S.; January 1897 for Belgium; January 1898 for France; September 1905 for Italy; January 1906 for Sweden; September 1912 for Finland; January 1914 for Norway; July 1914 for Japan; January 1915 for Spain; January 1916 for Switzerland; January 1918 for Canada; January 1919 for the Netherlands; January 1921 for Denmark; and January 1922 for Austria.

Risk free rate: We take the 10-year US government yield as a proxy for the risk free rate, also from G.F.D., which is available since January 1885 (T-bill rates are not available for the pre-World War I period, in contrast). G.F.D. initially gathered these data from a range of scholarly and official sources.[19] Yields on Treasury securities at constant, fixed maturity were constructed by the U.S. Treasury department based on the most actively traded marketable Treasury securities. Yields on

[19] G.F.D. used the 4% U.S. government bond of 1907 from January 1885 to January 1895; from February 1895 until September 1918, the 4% U.S. government bond of 1925 (when no trades were recorded during a given month, the previous month's yield was used). The source for this data was William B. Dana Co., *The Financial Review*, New York: William B. Dana Co. (1872-1921) which reprinted data published by *The Commercial and Financial Chronicle*. Beginning in 1919, the Federal Reserve Board's 10-15 year Treasury bond index was used and 10-year bonds were used starting in 1941. The data dating back to 1919 were taken from the Federal Reserve, *National Monetary Statistics*, New York: Federal Reserve Board, which was published in 1941, 1970 and annually since then. To obtain constant maturity yields, Treasury staff constructed a yield curve each business day and yield values were obtained from the curve at fixed maturities.

these issues are based on composite quotes reported by U.S. government securities dealers to the Federal Reserve Bank of New York.

Dollar exchange rates: We take the exchange rate against the U.S. dollar of the currency of the 16 economies of our sample (i.e. all economies minus the U.S.) from G.F.D. to convert local currency returns to U.S. dollar returns. The exchange rate is defined as the number of local currency units per U.S. dollar (i.e. an upward movement indicates a depreciation of the local currency relative to the U.S. dollar). We adjusted the data for re-denominations. They were available since January 1885 for Austria, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Sweden, Switzerland, the U.K. and the U.S; April 1901 for Australia; January 1910 for Canada; January 1914 for Spain; and January 1920 for Belgium and Finland.

Trade openness: Our measure of trade openness is defined as the sum of total exports and imports of goods scaled by output. We took the data from Mitchell (1998a, 1998b and 1998c) for the period 1885-1947, and from the I.M.F.'s Direction of Trade Statistics (D.O.T.S.) for the period 1948-2013. Missing observations were replaced with their last known values. The data were available since 1885 for Australia, Canada, Denmark, Finland, France, Germany, Italy, Japan, Norway, Sweden, the U.K. and the U.S, against 1900 for the Netherlands; 1901 for Spain; and 1913 for Austria, Belgium and Switzerland.

Regional trade openness: We split the sample into three regions, namely: 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). Regional trade is defined as the sum of a country's exports and imports of goods to the countries in the same region, scaled by total trade.[20] We took the data from Mitchell (1998a, 1998b and 1998c) for the period 1885-1947, and from the D.O.T.S. for the period 1948-2013. Missing observations were replaced with their last known values.[21] We obtained long time series that start in

[20] Note that the U.S. is Canada's sole regional trading partner (and vice-versa). Likewise, Australia is Japan's sole regional trading partner (and vice-versa).

[21] Insofar as bilateral trade data were available for a subset of the 13 European countries prior to 1948, we used the share of intra-European trade obtained from this subset as a proxy, and rebased it to the actual share of intra-European trade (i.e. the share obtained with all 13 European countries) in 1948.

1885 for Australia, Austria, Denmark, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the U.K. and the U.S; 1897 for Belgium; 1910 for Canada; and 1920 for Finland.

Capital account openness: We use the de jure index of financial openness of Quinn and Voth (2008) for the pre-World War II period as well as the index of Quinn and Toyoda (2008) for the post-World War II period. This index measures the extent of restrictions to capital outflows and inflows by residents and nonresidents in a given country. It runs from 0 (financial autarky) to 100 (complete financial openness). Annual data for all our countries were available for the period 1890-1931 and for the period 1950-2007. In other words, these indices leave gaps for parts of our sample. To fill in the missing observations between 1932 and 1949, we assumed that each country index evolved commensurately with the global index calculated by Quinn and Voth (see Quinn and Voth (2008), Table 1, p. 536). When there were large discrepancies between our own estimates and those of Quinn and Toyoda (2008) for 1950, however, i.e. in cases when a country's capital account openness had, in fact, evolved quite differently from the global average, we used linearly interpolated observations between 1931 and 1950 instead.[22] We used 2008 observations to fill in missing values up to 2014. This is consistent with the observation that none of the economies in our sample changed capital account regulations in the global financial crisis, an observation supported by the data base of Chinn and Ito (2006) which compile information reported to the I.M.F. up to 2012 and confirm that, indeed, none of our economies took such measures.[23]

Financial development: We took the ratio of equity market capitalization to output as our metric of financial depth from Rajan and Zingales (2003) for the period 1913-1999 and from Beck, Demirgüç-Kunt, and Levine (2010) for the period 2000-2011.[24] We obtained updates for 2012 and

[22] This was the case of Finland, Japan, Netherlands, Norway, Sweden, Switzerland and the U.S. for which the discrepancy between our own estimates and those of Quinn and Toyoda (2008) for 1950 exceeded 20 percentage points.

[23] This is also consistent with the observation that capital controls were introduced in recent years by emerging market economies, but not by advanced economies (with the exception of Greece, Cyprus and Iceland, which, as readers will remember, are not in our sample).

[24] The Rajan and Zingales data were available for selected years only and were linearly interpolated to annual data. Rajan and Zingales gathered these data from an array of sources, including the official publication of the stock exchanges; those of the Fédération Internationale des Bourses Valeurs (FIBV); private guides to stock exchanges; estimates based on a stock exchange handbook in 1913 (or the closest year before 1913) to identify the number of domestic companies listed, the

2013 observations using data from Bloomberg (for equity market capitalization) and the I.M.F.'s World Economic Outlook database (for nominal GDP in U.S. dollars).

Oil prices: We first took data on the dollar price of an oil barrel (West Texas Intermediate), as assembled from G.F.D. from an array of primary and secondary sources.[25] Then we calculated the deviation (in logarithms) between oil prices and a five-year moving average to capture periods of "high" and "low" global oil prices.

Global growth uncertainty: 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. Real GDP growth is calculated as the sum of real GDP per capita growth and population growth using data from Maddison (2010). Observations for 2010-2014 are taken from the I.M.F.'s World Economic Outlook database.

High market volatility periods: Share of the 17 countries in our sample with high equity market volatility in a given month. We estimate the conditional volatility of stock returns for each country of our sample using GARCH(1,1) models. We normalize the conditional volatilities of each country's stock returns and define high market volatility periods as the proportion of the 17 country-returns in excess of 1.65 in a given month. This yields a global time series with monthly observations over January 1885-June 2014.

Market capitalization: Market capitalization data are typically not available for our century-long time period. For instance, MSCI value weights extend only to the 1970s (see e.g. Brusa, Ramodorai, and Verdelhan (2014)). To obtain rough estimates of equity market capitalization relative to world output, we multiplied each country's equity market capitalization-to-output ratio with its share in world output (itself taken from Maddison (2008)).[26] We then derived each country's equity

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number of shares of each company, and the price per share; and various issues of the *Bulletin of the International Institute of Statistics*.

[25] Data for the period 1885-1987 are taken by G.F.D. from *The Derrick's Hand-Book of Petroleum*, Oil City, PA; those from 1898 to 1912 from the NBER; those from 1912 to 1941 refer to the price of an oil barrel in Oklahoma City; G.F.D. used data collected by Platt's for the period 1941-1968, and by the Bureau of Labour Statistics for the period 1969-1982; G.F.D. also used the price for West Texas Intermediate Crude Oil for the period 1983-2014.

[26] The Maddison data on GDP weights end in 2008. We updated them with IMF World Economic Outlook data on country shares in world GDP at purchasing power parity exchange rates.

market capitalization relative to the total sample capitalization from the latter metric.[27] To replace missing observations prior to World War I, we used the observations in 1913 for all countries, except Finland and Spain (for which data were not available prior to 1989 and 1980, respectively).[28] We used the observations in 2011 to replace missing observations between 2012 and 2014.

Central bank policy rates: We took data on the main central bank policy interest rate from G.F.D., which assembled long time series using data from national sources on the main policy interest rate used at a given point in time.[29] The data are expressed in annual percent and in local currency terms. They were available since January 1885 for Austria, Belgium, France, Germany, Denmark, Finland, Italy, Japan, Norway, Netherlands, Spain, Sweden and the U.K.; against June 1907 for Switzerland; November 1914 for the U.S.; July 1920 for Australia; and March 1935 for Canada.

Domestic bond yields: We took data on government bond yields from G.F.D. We took the 10-year maturity for all countries, with the exception of Finland (5-year bond yield) and the U.K. (5-year note yield) due to the absence of long term series for longer maturities. G.F.D. initially gathered these data from national sources and assembled long time series using yields on the benchmark bonds closest to the stated maturity, albeit not exceeding it, at given points in time. The resulting series are expressed in annual percent and in nominal local currency terms. They were available since January 1885 for Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Spain, Sweden, and the U.S.; November 1899 for Switzerland; January 1902 for the U.K. and August 1910 for Finland.

Peggers vs. nonpeggers: dummy variable which equals zero for nonpeggers and one for peggers. Nonpeggers are defined as countries having floating, managed floating or freely falling

[27] Global GDP at market prices was not calculated prior to World War II. Existing historical series are later reconstructions. As Schularick (2006) notes, Maddison's estimates for real GDP in constant 1990 'international', i.e. purchasing power adjusted, dollars were 'deflated to historical market value' by the U.S. GDP deflator, a method which hinges on a purchasing power parity assumption, but which also remains the best –albeit crude – available approximation. Although this method may lead to overestimates of the GDP of developing economies, readers should note that the bias should be smaller and similar across our sample of advanced economies.

[28] Our estimates are broadly consistent with estimates obtained from other sources for the modern period. For instance, we find that the U.S. accounts for 52% of total sample market capitalization at the end of our baseline period, against e.g. 8% for Japan and 4% for Canada. This compares with 51%, 10% and 6%, respectively, according to the estimates of the World Federation of Exchanges for end-2012, and with 55%, 8% and 4% according to the MSCI World index (in May 2014).

[29] For instance, for the U.S. G.F.D. used the Fed's discount rate from 1914 to 1950 (there are no data prior to 1914 insofar as the Federal Reserve was established in December 1913 only); the Fed fund market rate from 1951 to 1979 and the Fed fund target rate from 1980 to 2014.

exchange rates according to the classification of Ilzetzki, Reinhart, and Rogoff (2004). Peggers 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).

Appendix B. Supplementary Figures

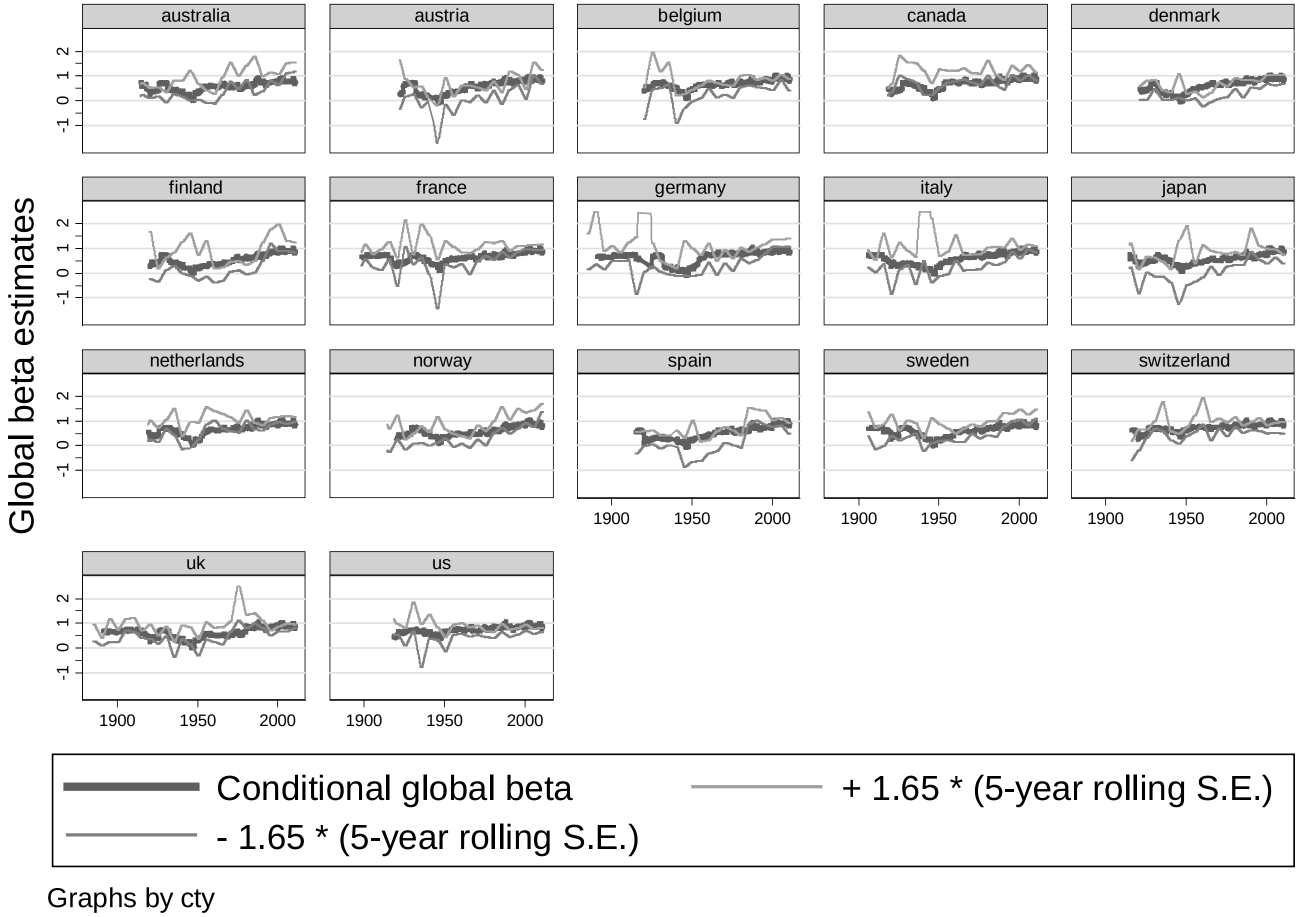


Figure B1. Conditional global betas vs. 5-year rolling betas – 1885-2014. The figure shows the evolution between 1885 and 2014 of the conditional global beta estimates (thick grey lines) obtained from the full model equations (1) to (3) for each country of our sample together with 90% confidence bands (light grey lines) obtained from 5-year rolling (non-overlapping) global estimates.

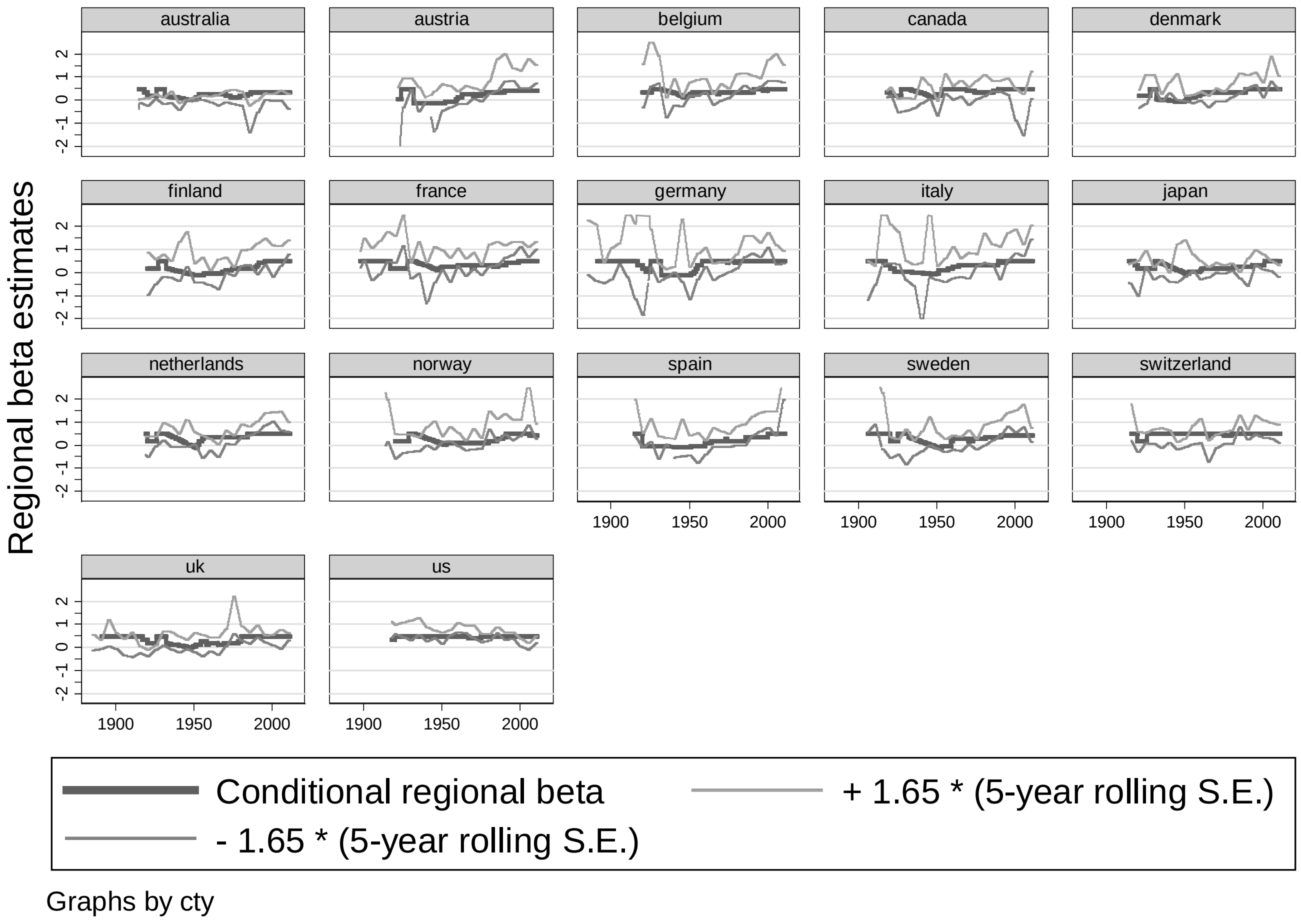
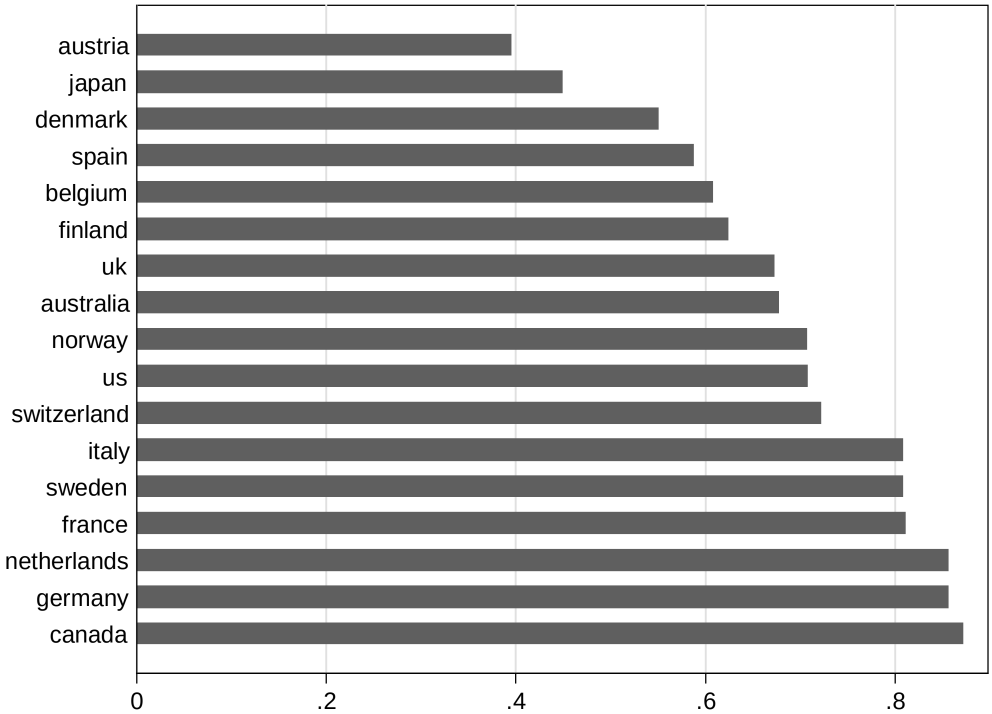


Figure B2. Conditional regional betas vs. 5-year rolling betas – 1885-2014. The figure shows the evolution between 1885 and 2014 of the conditional regional beta estimates (thick grey lines) obtained from the full model equations (1) to (3) for each country of our sample together with 90% confidence bands (light grey lines) obtained from 5-year rolling (non-overlapping) regional beta estimates.

Panel A. Unconditional Global Betas by Country



Panel B. Unconditional Regional Betas by Country

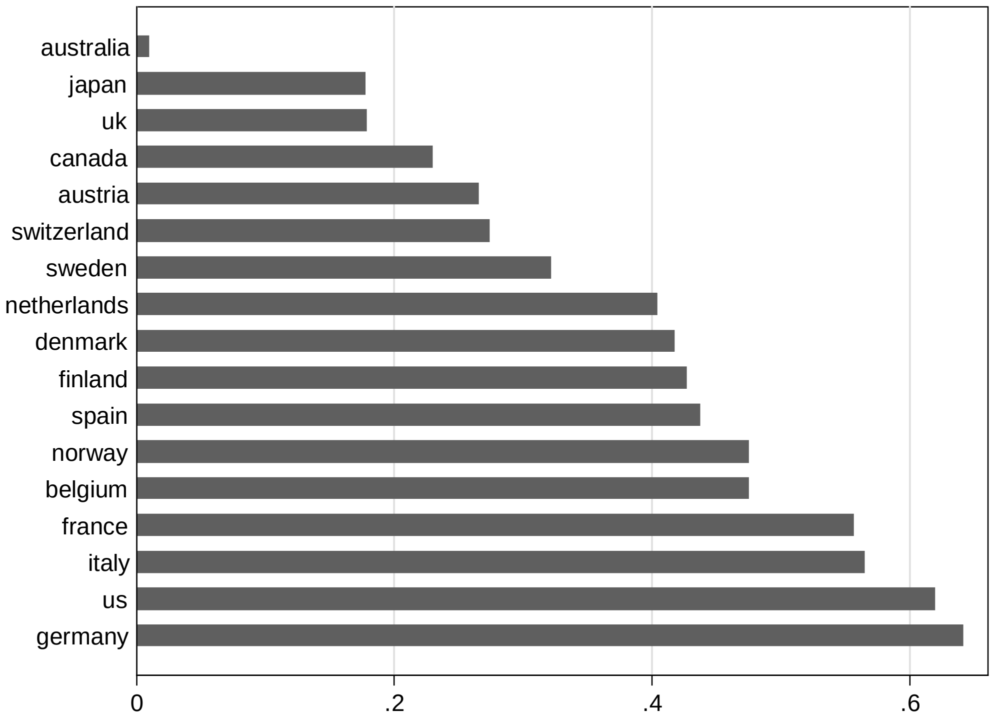


Figure B3. Unconditional betas – 1885-2014. The figure shows the full sample country-by-country estimates of the β_0^{glo} and β_0^{reg} parameters obtained from the model equation (6)

Appendix C. Supplementary Table

Table C1
Constant-Beta Model Estimates with GDP-Weighted Factors

The table reports the estimates of the following model:

$$R_{i,t} = \alpha_i + \lambda_t + \beta_0^{glo} F_t^{glo,\setminus i} + \beta_0^{reg} F_t^{reg,\setminus i} + \varepsilon_{i,t}$$

in which both the global factor and the regional factor are GDP-weighted (rather than value-weighted as in the estimates of Table V). Estimates obtained by OLS are reported in columns 1 to 3; estimates obtained by excluding outliers from the sample (i.e. returns larger than 30% within a month in absolute value) are reported in column 4; estimates obtained with a random effects estimator are reported in column 5. The estimates of columns 2 to 5 control for country fixed (or random) effects. The estimates of columns 3 to 5 control for year effects. The ρ-statistic in column (5) is the percent contribution to the total variance of the panel-level variance component. The standard errors reported in parentheses are robust to heteroskedasticity and to clustered heterogeneity. ***, **, *, and $^+$ indicate statistical significance at the 1%, 5%, 10%, and 15% levels, respectively.

	(1) OLS	(2) OLS	(3) OLS	(4) Outliers	(5) Random effects
Global factor	0.697***	0.698***	0.664***	0.655***	0.664***
	(0.032)	(0.032)	(0.032)	(0.030)	(0.032)
Regional factor	0.262***	0.262***	0.249***	0.253***	0.249***
	(0.076)	(0.076)	(0.076)	(0.073)	(0.076)
Constant	0.019	-0.007	0.331**	-0.375***	0.402+
	(0.034)	(0.017)	(0.143)	(0.068)	(0.269)
Country fixed effects	NO	YES	YES	YES	YES
Year effects	NO	NO	YES	YES	YES
Observations	20,800	20,800	20,800	20,719	20,800
R^2	0.196	0.196	0.205	0.281	0.205
# panel units					17
ρ					0

Table C2
Testing for a Reversal in Global Financial Market Integration since the Great Recession - Robustness

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

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

where j =1, 2, 3; 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 2009 and zero otherwise; and D_3 a dummy variable equal to one between 2010 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 clustered heterogeneity. ***, **, *, 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, D_2 and D_3.

Panel A. Full sample estimates (1885-2014)		**Panel B. *p*-value of Wald test H_0**	
Global factor × D_1	0.860***	$\beta_1^{glo} = \beta_2^{glo}$:	0.003
	(0.049)	$\beta_1^{glo} = \beta_3^{glo}$:	0.039
Global factor × D_2	1.115***	$\beta_2^{glo} = \beta_3^{glo}$:	0.492
	(0.052)	$\beta_1^{glo} > \beta_2^{glo}$:	0.001
Global factor × D_3	1.076***	$\beta_1^{glo} > \beta_3^{glo}$:	0.020
	(0.083)	$\beta_2^{glo} > \beta_3^{glo}$:	0.754
Regional factor × D_1	0.558***	$\beta_1^{glo} < \beta_2^{glo}$:	0.999
	(0.146)	$\beta_1^{glo} < \beta_3^{glo}$:	0.980
Regional factor × D_2	0.352**	$\beta_2^{glo} < \beta_3^{glo}$:	0.246
	(0.151)	$\beta_1^{reg} = \beta_2^{reg}$:	0.102
Regional factor × D_3	0.588***	$\beta_1^{reg} = \beta_3^{reg}$:	0.781
	(0.182)	$\beta_2^{reg} = \beta_3^{reg}$:	0.091
Constant	0.155	$\beta_1^{reg} > \beta_2^{reg}$:	0.949
	(0.119)	$\beta_1^{reg} > \beta_3^{reg}$:	0.390
Year effects	YES	$\beta_2^{reg} > \beta_3^{reg}$:	0.045
Country fixed effects	YES	$\beta_1^{reg} < \beta_2^{reg}$:	0.051
Observations	20,800	$\beta_1^{reg} < \beta_3^{reg}$:	0.610
R^2	0.161	$\beta_2^{reg} < \beta_3^{reg}$:	0.955