# FOREIGN DIRECT INVESTMENT AS A DETERMINANT OF CROSS-COUNTRY STOCK MARKET COMOVEMENT*

ALEXIOS ANAGNOSTOPOULOS† ORHAN EREM ATESAGAOGLU‡ ELISA FARAGLIA§
CHRYSSI GIANNITSAROU¶

February 15, 2019

*Preliminary and incomplete*

ABSTRACT. We develop a theoretical framework in order to investigate the link between two recent trends: (i) the rise in cross-country stock market correlations over the past three decades, and (ii) the increase in global foreign direct investment (FDI) positions over the same period. Our objective is twofold: first, we investigate empirically the channel through which the rise in global stock market correlations is associated with the observed increase in global FDI. Second, we develop a two-country stochastic asset pricing model with multinational firms that allows us to quantify the extent to which the recent rise in global FDI can account for the observed increase in cross-country stock market comovement. Calibrating three versions of the model (financial autarky, incomplete markets and complete markets) to the US and the rest-of-the-world, we find that a permanent inrcease in FDI positions, as observed in the late 1990s to mid 2000s, leads to substantial increase in cross-country stock market comovements, in line with that seen in the data.

KEYWORDS: stock market comovements, foreign direct investment, business cycle synchronisation, multinational firms, asset pricing

JEL CLASSIFICATION: G15, F21, F23, G12, F44

## 1. INTRODUCTION

Historically, the correlation between international stock markets had been fairly low, providing significant diversification benefits. However, this pattern started to change in the mid-1990s. Over the past two decades, international stock markets have steadily become more correlated. During the period mid-1990s to late 2000s, we observe some very sharp increases in correlations of major stock markets, often at the scale of almost threefold increases, from around 30% to 80% or more.

What drives the observed rise in cross-country stock market correlations? We posit that the bilateral foreign direct investment (FDI) positions between the big world economies act as an important determinant of stock market comovements. The link between the comovement of stock markets and global FDI positions is multinational firms. Generally, the equity price of a multinational firm is determined by their earnings generated all over the world. Since the earnings of a foreign subsidiary are directly affected by the economic conditions of the country that it operates in, as multinationals increase their overseas investment, their earnings and thus equity value, will be affected more from

*We thank the *British Academy* for generously funding this project via the Newton Advanced Fellowship scheme, and the *Cambridge Endowment for Research in Finance (CERF)* for financial support. We are grateful to internal seminar participants at Department of Economics of Stony Brook University.

†Department of Economics at Stony Brook University. *E-mail*: alexis.anagnostopoulos@stonybrook.edu

‡Istanbul Bilgi University and Newton Advanced Fellow at University of Cambridge. *E-mail*: eremates@gmail.com.

§Faculty of Economics at the University of Cambridge, and CEPR. *E-mail*: ef307@cam.ac.uk.

¶Faculty of Economics at the University of Cambridge, and CEPR. *E-mail*: cg349@cam.ac.uk.

cross-border business cycles. Thus, the correlation of cross-country stock market indices will increase, as the global FDI openness increases. Indeed, actual FDI positions were roughly on a stable level until the early 1990s, but there was a permanent increase in them since then and up until the mid 2000s. This level change seems to be mostly driven by increased FDI activity between developed economies, mainly within and between two regions, the US and the European Union and coincides with the period of rising stock market correlations.

In this paper, we have two main objectives. First, we carefully document evidence that uncovers the relationship between the rise in global stock market correlations and the rise in global FDI. Second, we develop a two-country stochastic asset pricing model which incorporates multinational firms, in order to quantify the extent to which the recent rise in global FDI can account for the observed increase in cross-country stock market correlations.

To understand the mechanisms at work, we first work with a stripped down model with exogenous production, in order to show that stock markets may comove even when business cycles are not synchronised. In this setting, we can generate simple increases in stock price correlations that are only due to increased openness between the two economies, and over and above any increases resulting from an increase in GDP correlations. Next, we move to a full production model of two countries with multinationals calibrated to the US and the rest of the world, and consider three versions of this setting, namely economies in financial autarky, incomplete markets and complete markets. The case of financial autarky serves as useful benchmark for quantifying the importance of FDI in a world where, apart from the presence of multinationals, the two economies do not have any other interaction. In this setting, we show that a permanent increase in FDI positions in the order of magnitude of what we observe in the data may well generate up to threefold increases in the stock market correlations. We also explain how this happens, by a heuristic decomposition of the comovements in stock prices into the comovement of the two multinational firm equity payouts and the comovement of the stochastic discount factors in the two countries. Our numerical results are broadly in line with the correlations and other statistics we observe in the data.

Our work relates to various distinct literatures. The empirical observation that stock market correlations have been increasing has been documented in some papers, most recently in Jorda, Schularik, Taylor and Ward (2018), who document a general synchronisation in financial cycles across developed countries. Our model extends the framework of McGrattan and Prescott (2009), Kapicka (2012) and Atesagaoglu and Anagnostopoulos (2019), in that we model explicitly the stock markets in the two countries. Finally our work is broadly placed in the large literature that studies international business and financial cycles comovement (both empirical and theoretical). This literature is mature (seminal papers include Backus, Kehoe and Kydland, 1995; Baxter and Kouparitsas, 2004) but also has important more recent contributions, including di Giovanni, and Levchenko (2010), Kalemli-Ozcan, Papaioannou and Peydro (2013), Jorda, Schularik, Taylor and Ward (2018) and Cesa-Bianchi, Imbs and Saleheen (2019).

The paper is organised as follows: the next section gives some empirical motivation for the paper and explains some basic facts about stock market and FDI comovements. Section 3 goes over the simple illustrative example that guides our intuition, and section 4 describes the full model with production and multinational firms. Section 5 gives the numerical results and section 6 provides some more detailed empirical support to the link between stock market correlations and FDI positions. The last section summarises and concludes.

## 2. Empirical motivation

Historically, the correlation between international stock markets had been fairly low (less than 50% on average), providing significant diversification benefits. However, this pattern started to change in the mid-1990s. Over the past two decades, global stock markets have steadily become more correlated; this fact has been identified in different contexts in several papers, e.g. Bekaert and Harvey (2000), Quinn and Voth (2008) and recently Jorda, Schularik, Taylor and Ward (2018). Here we provide evidence showing this increase in correlations that is consistent with those studies and, at the same time, evidence on the increase in FDI linkages. The evidence provided suggests a potential link between the two phenomena.

Starting with stock markets, our data consists of MSCI indices for a number of developed economies with large and well functioning stock markets, specifically United States (US), Canada (CA), Japan (JN), United Kingdom (UK), France (FR) and Germany (DE).[1] We also make use of the MSCI index of 22 of 23 developed markets, i.e. excluding US, (ROW), to look at correlations of the US stock markets versus those in the 'rest of the world'.[2] We use the MSCI indices in weekly frequency to first calculate weekly returns for a country $i$ at the end of week $t$ as

$$r_{it} = \frac{MSCI_{it} - MSCI_{i,t-1}}{MSCI_{i,t-1}}$$

and then calculate a measure of correlation $SMC^h_{i,t}$ between the stock markets of two countries $i$ and $h$ at time $t$, using the following definition

$$SMC^h_{i,t} = corr\left(\mathbf{r}^i_t, \mathbf{r}^h_t\right)$$

where $\mathbf{r}^i = \left(r^i_{t-w/2}, ..., r^i_{t+w/2}\right)$, with $w$ a pre-specified time 'window'. Using a rolling window of $w = 208$ weeks, Figures 1 and 2 show correlations between US and ROW, and all bilateral correlations between the countries listed above, respectively. Each point reported in these graphs is therefore dated in the middle of the rolling window used for calculating the reported correlation, e.g. $SMC^h_{i,(1/1/1990)}$ is the correlation of weekly returns for the range 1/1/1988 to 31/12/1991. Looking at the figures, it becomes apparent that there has been a substantial upshift in the correlations from some time in late 1990s to late 2000s (more or less pronounced, depending on the country pair).

What drives the observed rise in global stock market correlations? Here we explore the relationship between the rise in global stock market correlations and FDI activity, and as a first pass we look at how FDI between the same pairs of countries has changed in the last 30 years. The equity price of a multinational firm is determined by its earnings generated all over the world. Since the earnings of a foreign subsidiary are directly affected by the business cycles of the countries that it operates in, as multinationals increase their overseas investment, they will be affected more from cross-border business cycles. Therefore changes in international stock market correlations may be related to changes in FDI activity between countries. With this intuition in mind, we next construct a measure of correlations of FDI positions between the six countries of interest.

[1] Our choice of countries is restricted mostly by availability of data on FDI positions. Importantly, the six countries we look at have very large stock markets, that account for the majority of stock market activity across the world.

[2] The countries included in MSCI world index excl. US are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Hong Kong, Ireland, Israel, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland and the UK.

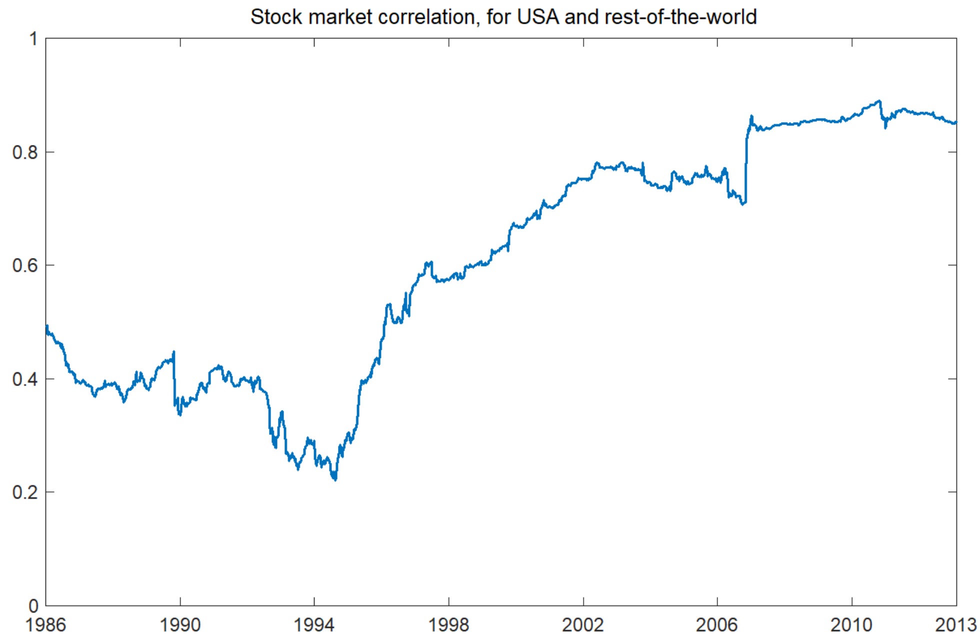


Figure 1: Stock market correlation, SMC(US, ROW). Calculated over four year rolling windows, using weekly MSCI US and MSCI World Exc. US. Data range 1/1/1984 - 31/12/2015.

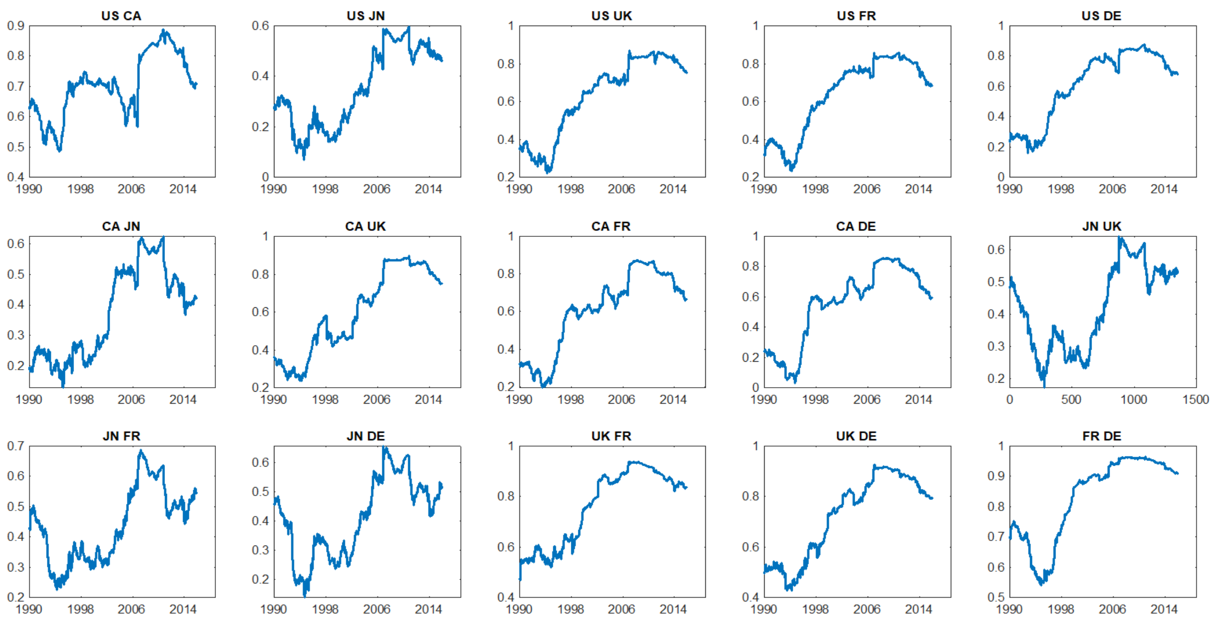


Figure 2: Stock market correlations, bilateral, between United States (US), Canada (CA), Japan (JN), United Kingdom (UK), France (FR), Germany (DE).

Our measure of FDI activity between two countries $i$ and $h$ at the start of a given year $t$, is defined relative to the size of the two economies:[3]

$$RF_t^{ih} = \frac{FDI_{i,t}^h + FDI_{h,t}^i}{GDP_{i,t} + GDP_{h,t}}$$

where $FDI_{i,t}^h$ is the nominal FDI position of country $h$ in country $i$, $FDI_{h,t}^i$ is the nominal FDI position of country $i$ in country $h$, and $GDP_{i,t}$ and $GDP_{h,t}$ denote the nominal GDPs of the two countries, all reported in US dollars.

Figure 3 shows the measure of relative FDI, $RF$, between US and the rest of the world, plotted together with the corresponding $SMC$ measure. In the figure, the left hand vertical axis shows the scale for $SMC$ and the right hand axis shows the scale for $RF$, expressed in percent. We see that during the 1980s and well into the 1990s, the FDI of US relative to world GDP was stable and around 5%, and then increases steadily for a few years, until it doubles to a permanent higher level of approximately 10%. During the period from mid 1990s to early 2000s, we observe an increase in relative FDI, and at the same time a dramatic increase in the stock market correlation measure from an average of 0.3 to almost 0.8.

Turning to specific country pairs, good quality data for country bilateral FDI positions is not broadly available. We use bilateral FDI positions from *OECD Foreign Direct Investment Database* at a yearly frequency. OECD has recently revised the definition of FDI and provides a new series from 2013 onwards, which unfortunately causes a break in the series after 2012.[4] For this reason we restrict our data and analysis up to and including 2012. Theoretically, the outward FDI position of country $i$ in country $h$ should be equal to the inward FDI position of country $h$ in country $i$. However, because countries may have different ways of reporting inward and outward FDIs, these two statistics are mostly different. Moreover, for Japan, there are some missing data values, and for France, FDI inward and outward position data only starts in 1998. For these reasons, we use the best and longest available measured version of FDI position between the two countries. We have done the analysis for a variety of alternative combinations and find the same qualitative features of FDI positions in the last 2.5 decades. With a few exceptions, the pattern we observe for the bilateral relative FDIs is the same as that for the US and the rest of the world. They appear to be increasing at around the same period as the corresponding shifts in stock market correlations. For many of these pairs, the shift up occurs around late 1990s to early 2000s. These increases seem to take place at around the same time that the corresponding stock market correlations increase, as shown in Figure 4.

## 3. An illustrative example

Before presenting a fully calibrated model that formalizes the link between FDI linkages and cross-country stock market correlations that quantifies the proposed mechanism, we first illustrate the main idea in the simplest possible environment. The objective here is to show how our model has the potential to generate increases in stock price correlations over and above any increases resulting from an increase in GDP correlations.[5]

[3] We follow Kalemli-Ozcan et al. (2012) and Menno (2017) who define FDI linkages similarly.

[4] See OECD Benchmark Definition of Foreign Direct Investment, 4th Edition (BMD4).

[5] Jorda et al. (2018) find empirically that recent increases in the synchronization of financial variables (and specifically equity prices) are larger than the relative increases in synchronization of real variables. In other words, the recently observed increased correlations of international stock markets cannot be attributed solely to the increased synchronisation of international business cycles.

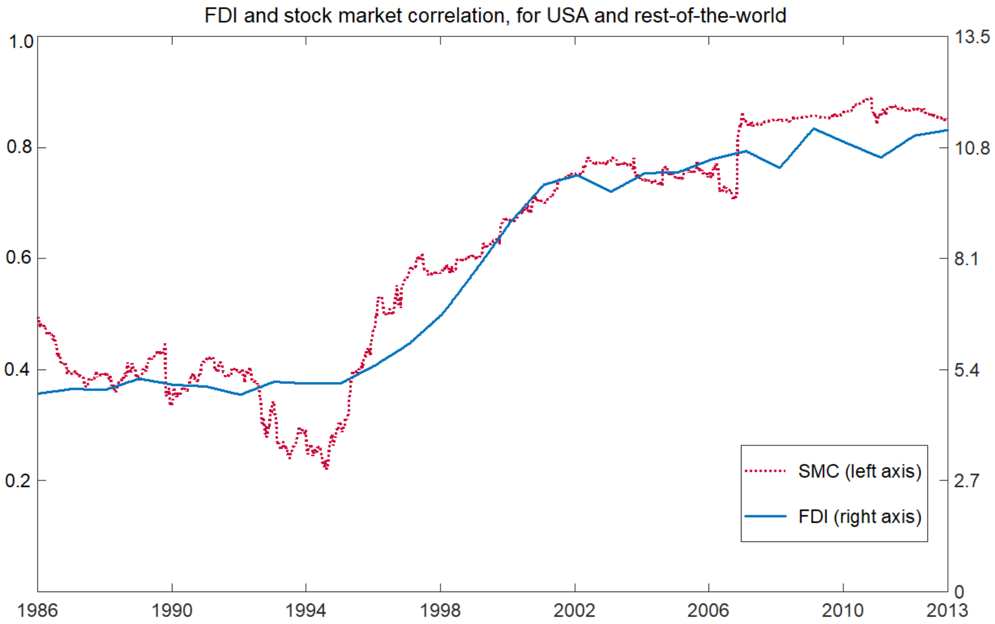


Figure 3: Stock market correlation and relative FDI, for the US and the rest of the world. FDI data source: *FRED*. The blue line, scale on left axis, is $SMC_{US,ROW}$. The red line, scale on right axis, is $RF_{US,ROW}$.

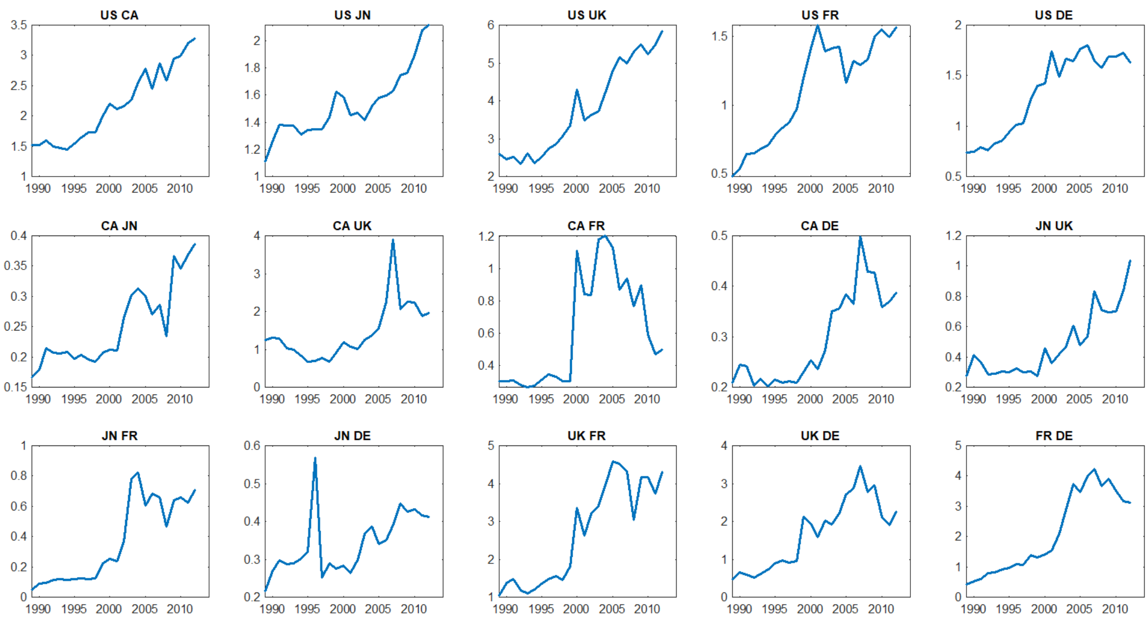


Figure 4: Bilateral relative FDIs, between United States (US), Canada (CA), Japan (JN), United Kingdom (UK), France (FR), Germany (DE). Data in annual frequency, source: *OECD*.

We consider the following setup. Time is discrete, infinite and indexed by $t = 0, 1, 2, ...$ There are two multinational firms and two countries, each one populated by a representative household. Each multinational firm operates two productive units (plants), one located within the country where the multinational is incorporated and one located in the other country (abroad). Thus, there are four plants in total. In what follows, superscripts $h = 1, 2$ are used to denote the multinational that *owns* the plant and subscripts $i = 1, 2$ are used to denote the country in which the plant is *located*.

In this first simple version of the model, all production is exogenously determined: each plant produces $Y_{it}^h$ goods according to

$$Y_{it}^h = \sigma_i^h Z_{it}, i = 1, 2, h = 1, 2$$

where $Z_{1t}$, $Z_{2t}$ are exogenous shocks and $\sigma_i^h$ are exogenous parameters with $\sigma_1^1 = \sigma_2^2 = 1$ and $\sigma_1^2, \sigma_2^1 \in [0, 1]$. The parameters $\sigma_i^h$, $i \neq h$, capture the amount of production by foreign firms $h$ in a given country $i$ relative to the production by the domestic firm of that country. This production function suggests that production of a given firm $h$ depends on the specific productivity shocks hitting the local economy (country $i$), scaled by the parameter $\sigma_i^h$. Gross domestic product (GDP) in each country can be expressed as

$$\begin{aligned} GDP_{1t} &= Y_{1t}^1 + Y_{1t}^2 = \left(1 + \sigma_1^2\right) Z_{1t}, \\ GDP_{2t} &= Y_{2t}^2 + Y_{2t}^1 = \left(1 + \sigma_2^1\right) Z_{2t}. \end{aligned}$$

Therefore, the cross country correlation of GDP is exogenously given by the correlation of the shocks $Z_{1t}, Z_{2t}$ and is independent of the linkages captured by the parameters $\sigma_i^h$, since correlation is scale invariant.

The goods produced are perishable and the firm (tree) simply pays its output to its shareholders, so that the equity payouts $D_t^1$ and $D_t^2$ of the two firms are given by

$$\begin{aligned} D_t^1 &= Z_{1t} + \sigma_2^1 Z_{2t}, \\ D_t^2 &= Z_{2t} + \sigma_1^2 Z_{1t}. \end{aligned}$$

Contrary to GDPs, the equity payouts are linear combinations of both shocks that affect the two economies, and therefore their correlation now generally depends on the scale parameters $\sigma_1^2$ and $\sigma_2^1$. For example, if $\sigma_1^2 = \sigma_2^1 = 0$ then the correlation of dividends is exactly the same as the correlation of GDP. In an alternative extreme, if $\sigma_1^2 = \sigma_2^1 = 1$, then the dividends are perfectly correlated regardless of the correlation of GDP. The implication is that, as the level of linkages $\sigma_\iota^h$ increases, the correlation of equity payouts increases even if business cycles are not necessarily more synchronized.

Since stock prices are functions of the expected sums of discounted equity payouts, this mechanism that separates the business cycle synchronization and equity payout correlations will matter for stock market correlations. If we want to be more precise regarding this effect, we will need to be explicit about the stochastic discount factor. This is done in the next section where we build a general equilibrium model with endogenous production, in which the stochastic discount factor is linked to the intertemporal marginal rate of substitution (MRS) of stockholders. Note that assumptions regarding the level of international financial market integration will matter for the behavior of MRS.

Qualitatively, we can obtain some intuition regarding the effect of different risk-sharing assumptions, in this simple exogenous production framework. First, we analyse the economies in financial autarky, assuming a representative household in each country that holds all the stocks of the home multinational and has no income other than from assets. If there are no assets traded with foreign households, then consumption equals the equity payout of the home multinational. Assuming $Z_{1t}$ and $Z_{2t}$ are independent and follow the same AR(1) process with persistence $\rho$, it is straightforward to show that, up to a log-linear approximation, the correlation of stock prices coincides with the correlation of equity payouts

$$Corr\left(\hat{p}_{1t}, \hat{p}_{2t}\right) = Corr\left(\hat{d}_{1t}, \hat{d}_{2t}\right) = \frac{1}{\sqrt{1 + \frac{\left(1-\sigma_1^2\sigma_2^1\right)^2 Var(\hat{z}_{1t})Var(\hat{z}_{2t})}{\sigma_1^2 Var(\hat{z}_{1t}) + \sigma_2^1 Var(\hat{z}_{2t})}}}$$

so that the correlation is one if $\sigma_1^2 = \sigma_2^1 = 1$ and zero if $\sigma_1^2 = \sigma_2^1 = 0$ as expected.

At the other extreme, with complete international financial markets, the MRS of the stock owners in the two countries will be perfectly correlated. Under log-linearization, stock prices will reflect a weighted average of MRS and payout so that even if payouts are uncorrelated ($\sigma_1^2 = \sigma_2^1 = 0$) stock price correlations will be positive. It will still be the case, however, that as the $\sigma_i^h$ increase stock price correlations increase too.

## 4. The model

The model developed in McGrattan and Prescott (2009, 2010), and used by Kapicka (2012), McGrattan (2012) and Anagnostopoulos and Atesagaoglu (2019), is augmented with country-specific productivity shocks and capital adjustment costs.

Time is discrete and indexed by $t = 0, 1, 2, ...$ There are two countries, each one populated by a representative household, and two multinational firms. Each multinational firm operates two productive units (plants), one located within the country where the multinational is incorporated and one located abroad. Thus, there are four plants overall. In what follows, superscripts $h = 1, 2$ are used to denote the multinational that *owns* the plant and subscripts $i = 1, 2$ are used to denote the country in which the plant is *located*.

**4.1. Firms.** Consider the plant located in country $i$ and owned by multinational firm $h$. At any time $t$, the plant's output is denoted by $Y_{it}^h$. The physical capital stock and labour used for production are denoted by $K_{it}^h$ and $N_{it}^h$ respectively. Each multinational also has technology capital $M_t^h$ which is used as an additional input to production in both of its plants (hence no $i$ subscript). The production function for firm $h$ in country $i$ at time $t$ is

$$Y_{it}^h = Z_{it}\sigma_i^h F(\upsilon_i M_t^h, K_{it}^h, N_{it}^h), \quad i, h = 1, 2. \tag{1}$$

where $Z_{it}$ denotes a country-specific TFP shock and the $\sigma_i^h$ are parameters governing the degree of openness of each country $i$. More precisely, we assume $\sigma_1^1 = \sigma_2^2 = 1$ so that $\sigma_1^2, \sigma_2^1 \in [0, 1]$ can be used to control the amount of production by the foreign multinational relative to the home multinational in country $i$. The parameter $\upsilon_i$ captures population size in country $i$.[6]

[6] In the micro-founded version of the production function by McGrattan and Prescott (2009), $\upsilon_i$ captures the number of plant locations in country $i$, assumed to be proportional to population in $i$.

Aggregate domestic production in country $i$ is

$$Y_{it} = \sum_{h=1}^{2} Y_{it}^{h},\ i = 1, 2. \tag{2}$$

Physical capital and technology capital accumulation are described by

$$K_{it+1}^{h} = (1-\delta_K)K_{it}^{h} + X_{K,it}^{h} - \Phi(K_{it+1}^{h}, K_{it}^{h}),\ \ i, h = 1, 2, \tag{3}$$

$$M_{t+1}^{h} = (1-\delta_M)M_{t}^{h} + X_{M,t}^{h} - \Phi(M_{t+1}^{h}, M_{t}^{h}),\ \ h = 1, 2, \tag{4}$$

where $X_{K,it}^{h}$ and $X_{M,t}^{h}$ are investment in technology and physical capital respectively, $\delta_K,\ \delta_M$ are depreciation rates and $\Phi$ represents the capital adjustment cost function

The multinational incorporated in country $h$ maximizes the discounted value of worldwide dividends $D_t^h$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \Psi_{0,t}^{h} D_t^h,$$

where $\Psi_{0,t}^{h}$ is the stochastic discount factor (SDF) used by the firm. Since we will assume perfect home bias below, the SDF then simply corresponds to the intertemporal marginal rate of substitution of the representative household in country $h$. Dividends $D_t^h$ are given by

$$D_t^h = Y_{1t}^h + Y_{2t}^h - W_{1t}N_{1t}^h - W_{2t}N_{2t}^h - X_{K,1t}^h - X_{K,2t}^h - X_{M,t}^h, \tag{5}$$

**4.2. Households.** The representative household in each country $i$ maximizes lifetime expected utility

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u\left(\frac{C_{it}}{\upsilon_i}, \frac{H_{it}}{\upsilon_i}, \frac{N_{it}}{\upsilon_i}\right)\upsilon_i$$

where $C_{it}$ represents aggregate consumption, $H_{it} = bC_{it-1}$ captures an external habit term, $N_{it}$ is aggregate labour supply and the instantaneous utility satisfies the usual assumptions. Households supply labour $N_{it}$ to the plants that operate domestically and earn a wage $W_{it}$. They can buy and sell shares of the domestically incorporated firm only (perfect home bias), the number of shares bought at time $t$ are denoted by $\Theta_{it}$ and the price at which they are bought denoted by $P_t^i$.

We consider three distinct financial market structures, namely financial autarky (FA), incomplete markets (IM) and complete markets (CM). In one extreme, we do not allow any cross-country trade in financial assets by households (FA)

$$C_{it} + P_t^i\Theta_{it+1} = W_{it}N_{it} + \left(D_t^i + P_t^i\right)\Theta_{it} \tag{6}$$

Under incomplete markets, we allow households to also trade a non-contingent bond $B_{it}$ across countries. The budget constraint then looks like

$$C_{it} + P_t^i\Theta_{it}^i + B_{it+1}\left(s^t\right) + \frac{\chi}{2}\left(B_{it+1}\left(s^t\right)\right)^2 = W_{it}N_{it} + \left(D_t^i + P_t^i\right)\Theta_{it-1}^i + R_t\left(s^{t-1}\right)B_{it}\left(s^{t-1}\right) \tag{7}$$

Finally, at the other end, under complete markets (CM), households can trade a full set of state

| | | | | | |
|---|---|---|---|---|---|
| Discounting | $\beta =$ | 0.9900 | Capital share | $\alpha_K =$ | 0.2400 |
| IES parameter | $\theta =$ | 1.0000 | Labour share | $\alpha_N =$ | 0.6400 |
| Shock persistence | $\rho_1 =$ | 0.9500 | Tech capital share | $\alpha_M =$ | 0.1200 |
| Spillover parameter | $\rho_2 =$ | 0.0000 | Depreciation of physical capital | $\delta_K =$ | 0.0280 |
| Shock st. deviation 1 | $\sigma_{\varepsilon,1} =$ | 0.0163 | Depreciation of tech capital | $\delta_M =$ | 0.0280 |
| Shock st. deviation 2 | $\sigma_{\varepsilon,2} =$ | 0.0163 | Openness parameter (low) | $\sigma_{low} =$ | 0.7673 |
| Investment adj. costs | $\phi =$ | 5.0000 | Openness parameter (high) | $\sigma_{high} =$ | 0.8959 |
| Habit persistence | $b =$ | 0.0000 | Relative FDI (low) | $RF_{low} =$ | 0.6846 |
| Adj. cost for bond (IM) | $\chi =$ | 0.0010 | Relative FDI (high) | $RF_{high} =$ | 1.9738 |

Table 1: Baseline calibration.

contingent claims

$$C_{it} + P_t^i \Theta_{it}^i + \int q_t \left(s^t, \bar{s}\right) B_{it+1} \left(s^t, \bar{s}\right) d\bar{s} = W_{it} N_{it} + \left(D_t^i + P_t^i\right) \Theta_{it-1}^i + B_{it} \left(s^{t-1}, s_t\right) \tag{8}$$

where $s^t$ denotes the history of shocks $\left(Z_1^t, Z_2^t\right)$, $B_{it}\left(s^{t-1}, s_t\right)$ is the number of contingent claim bought in the previous period at state $s^{t-1}$ and promising to pay at state $s^t = \left(s^{t-1}, s_t\right)$ today and $q_{t-1}\left(s^{t-1}, s_t\right)$ is the corresponding price.

**4.3. Equilibrium.** Labor markets clear in each country

$$\upsilon_i N_{it} = N_{it}^1 + N_{it}^2,\ i = 1, 2$$

and the stock market clears in each country

$$\Theta_{it+1}^i = 1,\ i = 1, 2$$

In case of cross-country asset trade, the bond markets clear

$$\text{IM: } B_{1t+1} + B_{2t+1} = 0$$

or the contingent claims markets clear

$$\text{CM: } B_{1t+1}\left(s^t, \bar{s}\right) + B_{2t+1}\left(s^t, \bar{s}\right) = 0 \text{ for all } \bar{s} \text{ and all } s^t$$

Finally, in all cases, the aggregate resource constraint implies that total consumption plus total investment equals total GDP across both countries, i.e.:

$$\begin{aligned} & C_{1t} + C_{2t} + X_{K,1t}^1 + X_{K,2t}^1 + X_{K,1t}^2 + X_{K,2t}^2 + X_{M,t}^1 + X_{M,t}^2 \\ = \;& Y_{1t}^1 + Y_{1t}^2 + Y_{2t}^1 + Y_{2t}^2 \end{aligned}$$

## 5. Numerical results

**5.1. Calibration.** We assume that the two countries are symmetric. The model is calibrated on a quarterly basis for the US ($i = 1$) versus the rest of the world ROW ($i = 2$). The full set of parameters is provided in Table 1. We assume inelastic labor supply. The momentary utility function of households has the standard CRRA form with habit persistence

$$u\left(\frac{C_{it}}{v_i}, \frac{H_{it}}{v_i}, 1\right) = \frac{\left(\frac{C_{it}}{v_i} - \frac{H_{it}}{v_i}\right)^{1-\theta}}{1-\theta},$$

where the coefficient of risk aversion $\theta$ is set to 1. The habit persistence term is given by $H_{it} = bC_{i,t-1}$. In the baseline we consider no habit persistence ($b = 0$). The discount factor $\beta$ is set to 0.99, so that the quarterly real interest rate is equal to 1%.

The production of multinational $h$ in country $i$ is represented by the following Cobb-Douglas technology

$$Y_{it}^h = Z_{it}\sigma_i^h(v_i M_t^h)^{\alpha_M}(K_{it}^h)^{\alpha_K}(N_{it}^h)^{\alpha_N},$$

where $\alpha_K$, $\alpha_M$ and $\alpha_L$ denote, respectively, the income shares of tangible capital, technology capital and labor, $0 < \alpha_K, \alpha_M, \alpha_L < 1$ and $\alpha_K + \alpha_M + \alpha_L = 1$. We set $\alpha_K = 0.24$, $\alpha_M = 0.12$ , $\alpha_N = 0.64$ . The population sizes are normalized to $v_1 = 1$. Without loss of generality, the TFP level of the economies are normalized to $Z_1 = Z_2 = 1$. Capital adjustment costs are commonly used in international macro models to avoid excessive investment volatility. Accordingly, in our model, tangible and technology capital are subject to adjustment costs and following Fogli and Perri (2015) the cost functions takes the following form:

$$\begin{aligned}\Phi\left(K_{it}^h, K_{it+1}^h\right) &= \frac{\phi}{2}\left(\frac{K_{it+1}^h}{K_{it}^h} - 1\right)^2 K_{it}^h \\ \Phi\left(M_t^h, M_{t+1}^h\right) &= \frac{\phi}{2}\left(\frac{M_{t+1}^h}{M_t^h} - 1\right)^2 M_t^h,\end{aligned}$$

where the adjustment cost parameter is calibrated and set to $\phi = 5$, to match the observed relative volatility of tangible capital investment with respect to output for the US economy, in the range of $[2.39, 2.65]$.[7] The depreciation rate for tangible capital stock is set to $\delta_K = 0.028$ to match the ratio of US corporate tangible capital to US corporate GDP which is approximately 6.9.[8] Following McGrattan and Prescott (2010) and Kapicka (2012), the depreciation rate for technology capital stock is set to $\delta_M = 0.028$. In terms of the productivity shocks, Following Backus, Kehoe and Kydland (1992), Baxter and Crucini (1995) and Kehoe and Perri (2002), we assume that the shocks $(Z_{1t}, Z_{2t})$ follow a vector autoregressive (VAR) process of the form

$$\begin{bmatrix}\log(Z_{1t+1}) \\ \log(Z_{2t+1})\end{bmatrix} = \begin{bmatrix}\rho_1 & \rho_2 \\ \rho_2 & \rho_1\end{bmatrix}\begin{bmatrix}\log(Z_{1t}) \\ \log(Z_{2t})\end{bmatrix} + \begin{bmatrix}\varepsilon_{1t+1} \\ \varepsilon_{2t+1}\end{bmatrix}$$

The innovations $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t})$ are serially independent, multivariate normal random variables with contemporaneous correlation between innovations in US and ROW economy. Following the estimates of Baxter and Crucini (1995) and Kehoe and Perri (2002), we set the persistence parameter $\rho_1$ to 0.95 and the spillover parameter $\rho_2$ to 0. Based on the same set of papers, we set the correlation of innovation shocks to two alternative values. In our first parameterization, we set $corr(\varepsilon_{1t}, \varepsilon_{2t})$ to 0,

[7] FILL IN SOURCES HERE.

[8] Corporate tangible capital stock is calculated as the sum of non-residential equipment and structures plus residential assets and inventories, but excluding intellectual property products.

| **Uncorrelated Shocks** | Financial Autarky | | Incomplete Markets | | Complete Markets | |
|---|---|---|---|---|---|---|
| | $RF_{low}$ | $RF_{high}$ | $RF_{low}$ | $RF_{high}$ | $RF_{low}$ | $RF_{high}$ |
| $corr(p_1, p_2)$ | 0.0557 | 0.1873 | 0.1015 | 0.3193 | 0.2535 | 0.6273 |
| $corr(\Delta p_1, \Delta p_2)$ | 0.0562 | 0.1985 | 0.0810 | 0.3247 | 0.2344 | 0.6183 |
| $corr(d^1, d^2)$ | -0.1554 | -0.3270 | -0.0827 | -0.1417 | 0.2437 | 0.4980 |
| $corr(sdf_1, sdf_2)$ | 0.4401 | 0.7709 | 0.9704 | 0.9620 | 1.0000 | 1.0000 |
| $corr(c_1, c_2)$ | 0.1086 | 0.2939 | 0.2067 | 0.3913 | 1.0000 | 1.0000 |
| $corr(gdp_1, gdp_2)$ | -0.0139 | -0.0450 | -0.0260 | -0.0701 | -0.0607 | -0.1231 |
| **Correlated Shocks** | Financial Autarky | | Incomplete Markets | | Complete Markets | |
| | $RF_{low}$ | $RF_{high}$ | $RF_{low}$ | $RF_{high}$ | $RF_{low}$ | $RF_{high}$ |
| $corr(p_1, p_2)$ | 0.3000 | 0.4140 | 0.3416 | 0.5241 | 0.4731 | 0.7581 |
| $corr(\Delta p_1, \Delta p_2)$ | 0.3025 | 0.4278 | 0.3250 | 0.5319 | 0.4581 | 0.7523 |
| $corr(d^1, d^2)$ | 0.0941 | -0.1021 | 0.1685 | 0.1021 | 0.4640 | 0.6647 |
| $corr(sdf_1, sdf_2)$ | 0.6217 | 0.8559 | 0.9821 | 0.9770 | 1.0000 | 1.0000 |
| $corr(c_1, c_2)$ | 0.3485 | 0.5063 | 0.4338 | 0.5838 | 1.0000 | 1.0000 |
| $corr(gdp_1, gdp_2)$ | 0.2361 | 0.2066 | 0.2246 | 0.1823 | 0.1913 | 0.1300 |

Table 2: Numerical results with uncorrelated and correlated shocks.

and in our second parameterization, we set $corr(\varepsilon_{1t}, \varepsilon_{2t})$ to 0.25.

Finally, we choose $\sigma_2^1 = \sigma_1^2 = \sigma_{low} = 0.7673$ to match the US FDI position abroad as a share of tangible capital stock in the US owned by US firms which was approximately 11% between 1982 and 1993, based on the BEA's International Investment Position (IIP) data.[9] Moreover, we choose the higher value for openness $\sigma_{high} = 0.8959$. Recall that we have also assumed $\sigma_1^1 = \sigma_2^2 = 1$. The implied steady state values for the relative FDI measure defined as

$$RF = \frac{K_1^2 + K_2^1}{GDP_1 + GDP_2}$$

is then $RF_{low} = 0.6846$ and $RF_{high} = 1.9738$, for low and high openness parameter respectively.[10]

### 5.2. Results and discussion.

For each economy, we run experiments for low and high relative FDI, and examine how the correlation of stock prices and stock price growth changes. Apart from stock market correlations, Table 2 presents correlations for key variables of the two countries, for our baseline calibration (uncorrelated shocks) and for shocks that are correlated with $corr(\varepsilon_{1t}, \varepsilon_{2t}) = 0.25$ respectively.. Moments are generated by a simulation of 100,000 periods of the first order approximation of the model, from which we drop the first 1,000 periods and then take averages. All series are HP filtered with $\lambda = 1600$.

A first look at the results from the baseline calibration show a clear pattern: higher relative FDI position implies stronger correlations of equity prices and equity price growth rates. At the extreme case when country shocks are completely uncorrelated and the economies are operating under financial autarky, the increase in the openness parameter from $\sigma_{low}$ to $\sigma_{high}$ generates *threefold* increases in stock market correlations, from approximately 0.06 to about 0.19. The case of financial autarky is

[9] This is proxied by total US corporate sector capital stock minus the FDI position of foreign countries in the US.

[10] When we calibrate the model to asymmetric countries, we will use the figures $RF_{low} = 5$ and $RF_{high} = 10$ to match the steady state values for US vs ROW, as shown in Figure 3.

clearly not realistic, but it is a very useful benchmark because here we shut down as much as possible all other channels via which the two economies may affect each other, apart from the production parameters $\sigma_1^2 = \sigma_2^1 \equiv \tilde{\sigma}$ (and thus the relative FDI position). As we add model features which make the economies more financially integrated, i.e. as we move to complete markets and eventually to correlated shocks, the comovement of stock markets becomes stronger, although the relative increases in the correlations are of course not as large.

Next, we discuss the mechanism that generates increases in the stock market correlations, whenever the relative FDI positions increase (i.e. when the relative production of the foreign multinational to that of the home multinational increases from $\sigma_{low}$ to $\sigma_{high}$). To understand why increases in $\tilde{\sigma}$ lead to higher correlations of the stock price indices of the two countries, we start with a heuristic decomposition the covariance of the two stock prices and explain what determines the components of the covariance. Each stock price is determined by the infinite stream of expected discounted equity payouts, where the stochastic discount factor is determined by the marginal rate of intertemporal substitution. The comovement of the two stock prices will, as a result, depend on (a) the comovement of the two stochastic discount factors and (b) the comovement of the two equity payouts (as well as cross-terms).

The correlations of the two stochastic discount factors are increasing as $\tilde{\sigma}$ increases, and they also become higher as we move towards complete markets. In fact, in the context of complete markets and since the model is for now symmetric, consumptions in the two countries are identical and therefore the stochastic discount factors are always perfectly correlated. Generally speaking our results indicate that the comovement of the SDFs is strong and increasing when $\tilde{\sigma}$ increases.

Turning to the comovement of equity payouts, we note that the correlations of total payouts between the two countries can be negative or positive, as well as decreasing or increasing in $\tilde{\sigma}$. The reason why this happens can be explained by looking at the decomposition of a firm's payouts as the sum of payouts originating from each location $D_t^h = D_{1t}^h + D_{2t}^h$, where we define

$$\begin{aligned} D_{it}^i &= Y_{it}^i - W_{it} N_{it}^i - X_{K,it}^i - X_{M,t}^i \\ D_{it}^j &= Y_{it}^j - W_{it} N_{it}^j - X_{K,it}^j \,, \qquad j \neq i \end{aligned}$$

Using these definitions, the covariance of $D_t^1$ and $D_t^2$ can be decomposed as follows

$$Cov\left(D_t^1, D_t^2\right) = Cov\left(D_{1t}^1, D_{2t}^2\right) + Cov\left(D_{1t}^1, D_{1t}^2\right) + Cov\left(D_{2t}^1, D_{1t}^2\right) + Cov\left(D_{2t}^1, D_{2t}^2\right).$$

We then want to trace how each of these components changes as the parameter $\tilde{\sigma}$ increases. Generally, we anticipate that as $\tilde{\sigma}$ increases, the relative importance of the last three terms versus the first term should be bigger, since higher $\tilde{\sigma}$ means more openness between the two economies.

To understand how these components operate, it is useful to do an impulse response experiment, whereby we shock the productivity $Z_1$ in country 1 and then trace the effects of the shock to key variables of the model for low and high $\tilde{\sigma}$. Figures 5 and 6 show the impulse response functions of total payouts and their components, as well as investments and their components for financial autarky and complete markets respectively. The blue and red lines show responses for $\sigma_{low}$ and $\sigma_{high}$ respectively. By comparing the two economies under CM and FA, we observe that the reaction of the different components of the dividends between the two economies are very similar apart from the responses of

$D_2^2$.

Tracing the effects of the shock, we observe that due to the increase in productivity, firm 1 in country 1 finds it more profitable to invest and therefore both $X_M^1$ and $X_{K,1}^1$ increase. Moreover, the TFP shock also increases wages $W_1$, which is why we always have a negative response of $D_1^1$, as can be seen from

$$D_{1t}^1 = Y_{1t}^1 - W_{1t}N_{1t}^1 - X_{K,1t}^1 - X_{M,t}^1$$

At the same time, in country 2 there are two effects that determine how equity payouts respond, which are due to how $X_{K,2t}^2$ responds. Investment of each firm is determined by how much the firm discounts the future and the incentive to reallocate capital across countries in order to benefit from higher returns. When we have a positive shock to productivity, consumption in country 1 increases under both financial autarky and complete markets. With complete markets, consumption in country 2 is identical to that in country 1, and therefore responds in the same way as in country 1, increasing by the same amount as TFP in country 1 increases. At the other end, in financial autarky consumption in country 2 increases by a lot less. Therefore, the effects on the increased SDFs for the two economies are qualitatively the same across countries but much less pronounced in financial autarky. A higher SDF encourages the firms to invest in each plant, but on the other hand firms have an incentive to reallocate physical capital from country 2 to country 1. The higher investment returns due to the higher productivity in country 1 therefore make firm 2 willing to reallocate capital across countries and disinvest in country 2, in order to invest in country 1. These two incentives work against each other, and which of the two dominates can be different in the two economies. The SDF effect dominates under complete markets, so that investment increases everywhere and in particular $X_{K,2t}^2$ increases and $D_{2t}^2$ behaves similarly to $D_{1t}^1$ in country 1. At the other end, in financial autarky, and given that there is a much smaller effect in the SDF in country 2, the reallocation effect prevails and $X_{K,2t}^2$ decreases allowing firm 2 to distribute more dividends so $D_{2t}^2$ increases. In conclusion, even when the payouts of the two firms move in opposite directions, the effect that comes from the stochastic discount factor dominates and results in increases of the stock market correlations, as $\tilde{\sigma}$ and relative FDI increase.

## 6. Some additional empirical evidence

We aim to uncover a relationship between the correlations of international stock markets and foreign direct investment (FDI). In particular we examine the impact of FDI position between two countries on the correlation of the stock market returns of the two countries (i.e. on the correlation of the growth rate of stock market prices), controlling for a set of relevant macroeconomic variables. Our data set consists of six countries, as listed in Section 2 (US, Canada, Japan, UK, France and Germany), i.e. we have $N = 15$ country pairs over $T = 24$ years, up to 2012, due to the restrictions that relate to how FDI positions are calculated. In our specifications, the dependent variable is stock market correlations, as calculated in Section 2 based on the MSCI indices and the explanatory variable is the relative FDI position measure $RF$. We also allow for a variety of other controls, namely a measure of bilateral trade, as well correlations of industrial production, interest rates and inflation rates, defined in the same way as the correlations of the MSCI indices.[11]

Macro time series panels such as the one we work with are plagued by a variety of problems

[11] For our trade measure we use the database of Lane and Milesi-Ferretti (2007), which is extended up to 2012. The source for the remaining controls is OECD.

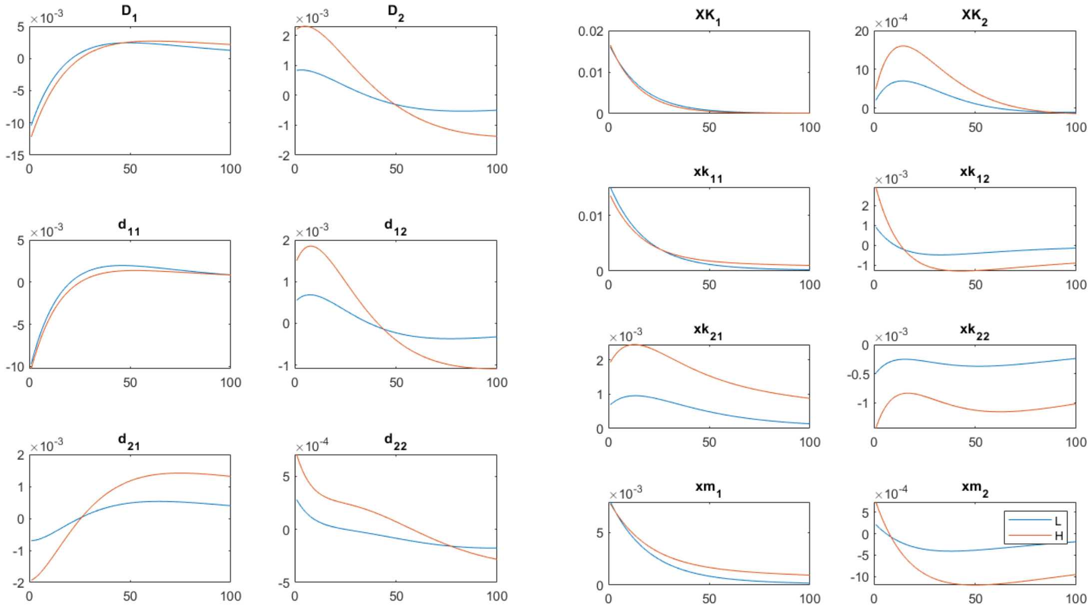


Figure 5: Impulse response functions, uncorrelated shocks, in financial autarky (FA). Blue line is for $\sigma_{low}$ and red line is for $\sigma_{high}$.

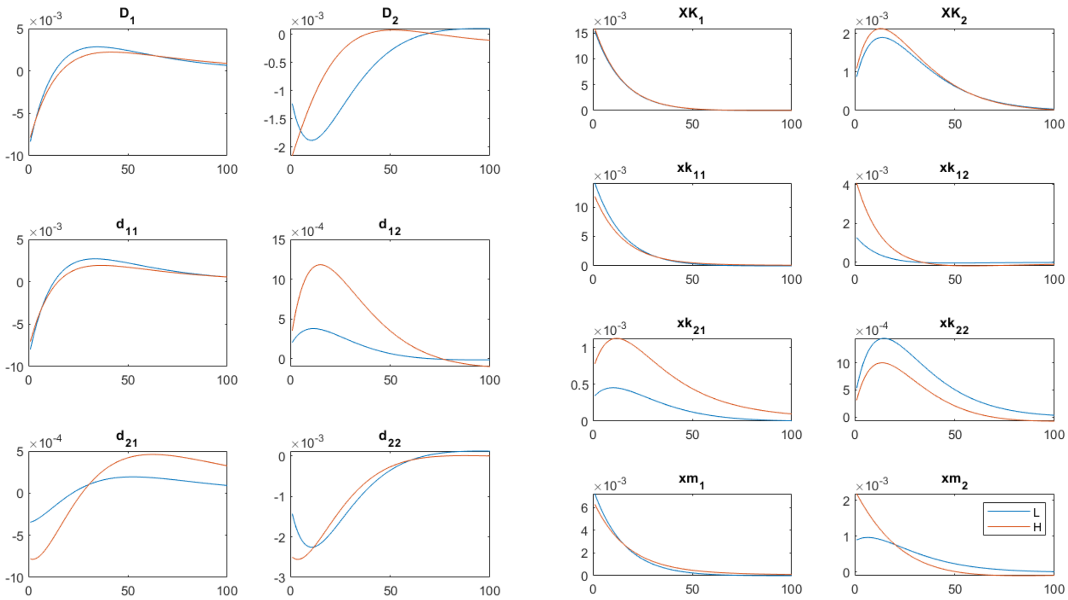


Figure 6: Impulse response functions, uncorrelated shocks, under complete markets (CM). Blue line is for $\sigma_{low}$ and red line is for $\sigma_{high}$.

that are typically absent in standard micro, longitudinal panels, such as non-stationarities, pervasive endogeneity, as well as cross-sectional dependence. Moreover, as such panels are typically of small sizes, i.e. the number of groups $N$ is relatively small and often of the same order or magnitude as the number of time periods $T$, estimation is potentially subject to small-sample bias. All these are issues that we account for carefully here. On the positive side, because the number of time periods for time series panels is longer than micro panels, it is possible to account for possible parameter heterogeneity. Indeed, we tested our data for the presence of (i) panel unit roots (using the tests of Levin, Li and Chu, 2002; Im, Pesaran and Shin, 2007), (ii) cross-sectional dependence (using the tests of Pesaran, 2004; Friedman, 1937; Frees, 1995) and (iii) serial correlation in panel data (using the test of Woolridge, 2002). We find strong evidence of the presence of both cross sectional dependence and serial correlation. Also, we cannot confidently reject the hypothesis of the presence of unit roots in some of the variables, such as our FDI measure and the trade measure, even when a deterministic linear trend is removed from those.

Given these results and together with the fact that we are essentially after estimating a long run relationship between stock market correlations and FDI, we resort to estimation methods that are more suitable for panel time series, which can account for some these issues and provide reliable estimates. Nevertheless, we also report results from standard micro panel estimation methods for comparison, being however aware of the fact that under the presence of cross-sectional dependence, endogeneity or non-stationarity these will be inconsistent and/or biased. We do two sets of estimations, by assuming that the panel is first static and second it is dynamic. The tables in the Appendix show the estimation results for (a) static panels, using fixed effects estimation (assuming homogeneous slopes), and mean group estimates and common correlated effects (CCE) estimates (assuming heterogeneous slopes) and (b) dynamic panels, based on the autoregressive distributed lag model, using dynamic fixed effects (DFE), pooled mean group (PMG) and mean group (MG) estimates. For details on the ARDL and CS-ARDL models and proposed estimators, see Chudik et. al. (2016). The results from the tables that follow generally support the importance of FDI positions as a determinant for cross-country stock market comovements, and the effects are especially strong under the ARDL specification, which we believe is more appropriate for the panel we work with.

## 7. Conclusion

This paper aims to establish the relationship between the rise in cross-country stock market correlations and the increase in FDI investment positions in the last thirty years. Following the literature, we have calculated measures of stock market correlations and of cross country relative FDI positions first between US and the rest of the world and subsequently among several OECD countries. We establish that at a first inspection most countries have experienced an increase in the stock market correlations and cross country FDI positions in the first part of the nineties. We then turn to a theoretical framework to establish the transmission channel between FDI investment increases and variations in stock market correlations in a two-country stochastic asset pricing model with multinational firms extending McGrattan and Prescott (2010). We show that the model generates an increase in the correlations when relative FDI positions increase and the result is robust to several financial market structures. We establish that the increase in the correlation follows different channels: a dividend channel in complete markets and a SDF channel in financial autarky. Given the theoretical results we carry out further empirical analysis and examine the impact of FDI position between two countries on the correlation of the stock market returns of the two countries controlling for a set of relevant macroeconomic variables using panel data analysis. The results generally support the importance of FDI positions as a determinant for cross-country stock market comovements, and the effects are especially strong under the ARDL specification.

## References

[21] Anagnostopoulos, A., Atesagaoglu, O. E., 2019. 'Technology Capital and the Taxation of Multinational Corporations', mimeo.

[21] Backus, D. K., Kehoe, P. J. and Kydland, F. E., 1995. 'International Business Cycles: Theory vs. Evidence', in T. F. Cooley (Ed.), Frontiers of Business Cycle Research.

[21] Baxter, M. and M. Kouparitsas, 2005. 'Determinants of Business Cycle Comovement: A Robust Analysis', Journal of Monetary Economics, 52(1): 113–157.

[21] Bekaert, G. and C. R. Harvey, 2000. 'Foreign Speculators and Emerging Equity Markets', Journal of Finance, 55(2): 565–613

[21] Cesa-Bianchi, A., J. Imbs and J. Saleheen (2019). 'Finance and Synchronization', Journal of International Economics, forthcoming.

[21] Chudik, A. K. Mohaddes, H. Pesaran and M. Raissi, 2016. 'Long-Run Effects in Large Heterogeneous Panel Data Models with Cross-Sectionally Correlated Errors', in the book *Advances in Econometrics* (Volume 36): Essays in honor of Aman Ullah, Emerald Group Publishing Limit.

[21] di Giovanni, J. and A. A. Levchenko. 'Putting the Parts Together: Trade, Vertical Linkages, and Business Cycle Comovement, with Andrei A. Levchenko.' American Economic Journal: Macroeconomics, 2:3 (April 2010), 95–124.

[21] Frees, E.W. (1995) 'Assessing cross-sectional correlations in panel data', Journal of Econometrics, 64, 393-414.

[21] Friedman, M. (1937) 'The use of ranks to avoid the assumption of normality implicit in the analysis of variance', Journal of the American Statistical Association, 32, 675-701.

[21] Fogli, A., Perri, F., 2015. 'Macroeconomic Volatility and External Imbalances', Journal of Monetary Economics, 69, pp. 1-15.

[21] Im, K. S., M. H. Pesaran, and Y. Shin. 2003. 'Testing for unit roots in heterogeneous panels', Journal of Econometrics 115: 53–74.

[21] Jordà, O., M. Schularick, A. M. Taylor and F. Ward, 2018. 'Global Financial Cycles and Rish Premiums', NBER WP 24677.

[21] Kalemli-Ozcan, S., E. Papaioannou, and J-L. Peydro, 2013. 'Financial regulation, financial globalization, and the synchronization of economic activity', Journal of Finance, 68 (3), 1179-1228.

[21] Kapicka, M., 2012. 'How Important is Technology Capital for the United States?', American Economic Journal: Macroeconomics, 4 (2), pp. 218-248.

[21] Levin, A., C.-F. Lin, and C.-S. J. Chu. 2002. 'Unit root tests in panel data: Asymptotic and finite-sample properties', Journal of Econometrics 108: 1–24.

[21] McGrattan, E.R., 2012. 'Tranisition to FDI Openness: Reconciling Theory and Evidence', Review of Economic Dynamics, 15, pp. 437-458.

[21] McGrattan, E.R. and E. C. Prescott, E.C., 2010. 'Technology Capital and the US Current Account', American Economic Review, 100 (4), pp. 1493-1522.

[21] McGrattan, E.R. and E. C. Prescott, E.C., 2009. 'Openness, Technology Capital, and Development', Journal of Economic Theory, 144, pp. 2454-2476.

[21] Pesaran, M.H. (2004) 'General diagnostic tests for cross section dependence in panels', Cambridge Working Papers in Economics, 0435, University of Cambridge.

[21] Quinn, D. P., and H. J. Voth, 2008. 'A Century of Global Equity Market Correlations', American Economic Review: Papers & Proceedings 2008, 98:2, 535–540.

[21] Wooldridge, J. M. 2002. 'Econometric Analysis of Cross Section and Panel Data', Cambridge, MA, MIT Press.

**STATIC PANEL MODELS WITH HOMOGENEOUS SLOPES**

| | (a) | | | (b) | | |
|---|---|---|---|---|---|---|
| | FE | FE | FE | FE | FE | FE |
| $\hat{\beta}_{fdi}$ | 0.2704*** | 0.0142 | -0.0103 | 0.2466*** | -0.0089 | -0.0266 |
| | (0.0503) | (0.0486) | (0.0526) | (0.0469) | (0.0399) | (0.0446) |
| $\hat{\beta}_{trade}$ | | | | 0.1674 | 0.0210*** | 0.1951*** |
| | | | | (0.1523) | (0.0662) | (0.0533) |
| $trend$ | | 0.0209*** | | | 0.0210*** | |
| | | (0.0034) | | | (0.0029) | |
| $\hat{\beta}_{ip}$ | | | | 0.0220 | 0.0268 | -0.0142 |
| $\hat{\beta}_{ir}$ | | | | 0.0878 | 0.0615 | 0.0188 |
| $\hat{\beta}_{\pi}$ | | | | 0.0511 | 0.0311** | 0.0426*** |
| time dummies | no | no | yes | no | no | yes |
| $N \times T$ | 360 | | | 360 | | |

Legend: The econometric specification is given by $y_{it}= c_i+\beta' x_{it}+u_{it}$, where $y_{it}$ is the correlation of the difference of logs of stock market prices for country pair $i$ at $t$, and $x_{i,t}$ contains the explanatory variables. The symbols *, **, *** denote significant estimates at 1%, 5% and 10% level. Numbers in parentheses are robust standard errors. The explanatory variables include only FDI in specification (a) and FDI and other controls in specification (b).

**STATIC PANEL MODELS WITH HETEROGENEOUS SLOPES**

| | (a) | | | | | (b) | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | MG | MG | MG | CCE | CCE | MG | MG | MG | CCE | CCE |
| $\hat{\beta}_{fdi}$ | 0.4057*** | 0.1414*** | 0.0451 | 0.0696 | 0.1095 | 0.3097*** | 0.1203** | 0.0846 | 0.1361 | 0.1693* |
| | (0.0561) | (0.0497) | (0.0666) | (0.0753) | (0.0733) | (0.0432) | (0.0509) | (0.0768) | (0.1069) | (0.1024) |
| $\hat{\beta}_{trade}$ | | | | | | 0.1580 | 0.1106 | 0.2031* | -0.0559 | -0.1525 |
| | | | | | | (0.1093) | (0.0879) | (0.1038) | (0.0958) | (0.1000) |
| $trend$ | | 0.0153*** | | | -0.0025 | | 0.0150*** | | | -0.0033 |
| | | (0.0031) | | | (0.0036) | | (0.0035) | | | (0.0035) |
| $\hat{\beta}_{ip}$ | | | | | | 0.0326 | 0.0398** | 0.0107 | 0.0010 | -0.0128 |
| $\hat{\beta}_{ir}$ | | | | | | 0.0232 | 0.0224 | 0.0265*** | 0.0473*** | -0.0369** |
| $\hat{\beta}_{\pi}$ | | | | | | 0.1132*** | 0.0978*** | 0.0329 | 0.0592** | -0.1005 |
| time dummies | no | no | yes | n/a | n/a | no | no | yes | n/a | n/a |
| $N \times T$ | 360 | | | | | 360 | | | | |

Legend: The econometric specification is given by $y_{it}= c_i+\beta_i' x_{it}+u_{it}$, where $y_{it}$ is the correlation of the difference of logs of stock market prices for country pair $i$ at $t$, and $x_{i,t}$ contains the explanatory variables. For CCE estimation it is assumed that $u_{it} = \gamma_i' f_t + \varepsilon_{it}$, where $f_t$ is the unobserved common correlated effect. The symbols *, **, *** denote significant estimates at 1%, 5% and 10% level. Numbers in parentheses are robust standard errors. The explanatory variables include only FDI in specification (a) and FDI and other controls in specification (b).

**ARDL(p,q) APPROACH - DFE, MG and PMG Estimates**

| | ARDL(1,1) | | | | | | ARDL(2,2) | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (a) | | | (b) | | | (a) | | | (b) | | |
| | DFE | PMG | MG | DFE | PMG | MG | DFE | PMG | MG | DFE | PMG | MG |
| $\hat{\theta}_{fdi}$ | 0.2916*** | 0.2743*** | 0.4346*** | 0.2869*** | 0.2407*** | 0.3644*** | 0.3223*** | 0.2764*** | 0.4625*** | 0.3133*** | 0.2716*** | 0..3116*** |
| | (0.0363) | (0.0204) | (0.0573) | (0.0349) | (0.0213) | (0.0588) | (0.0399) | (0.0134) | (0.0631) | (0.0398) | (0.0136) | (0.0738) |
| $\hat{\theta}_{trade}$ | | | | 0.2732** | 0.0959 | 0.2667 | | | | 0.1200 | -0.1285** | 0.2515 |
| | | | | (0.1147) | (0.0895) | (0.1637) | | | | (0.1369) | (0.0723) | (0.1946) |
| $\hat{\lambda}$ | -0.4925*** | -0.5712*** | -0.7065*** | -0.5093*** | -0.5681*** | -0.7753*** | -0.4632*** | -0.6079*** | -0.8276*** | -0.4658*** | -0.5720*** | -0.9149*** |
| | (0.0473) | (0.0585) | (0.0625) | (0.0476) | (0.0671) | (0.0602) | (0.0556) | (0.0925) | (0.0910) | (0.0566) | (0.1020) | (0.1125) |
| Haus $\chi^2$ | | 7.03 | | | 2.16 | | | 7.02 | | | 2.85 | |
| $p$ | | 0.008 | | | 0.3401 | | | 0.008 | | | 0.2403 | |
| $N \times T$ | | 345 | | | 345 | | | 330 | | | 330 | |

Legend: The ARDL$(p,q)$ specification is given by $y_{it}= c_i+\Sigma_{l=1}^{p}\varphi_{il}y_{i,t-l}+\Sigma_{l=0}^{q}\beta'_{il}x_{i,t-l}+u_{it}$ or in EC form $\Delta y_{it}= \lambda_i\left(y_{it-1}-\theta'_i x_{it}\right) + \Sigma_{l=1}^{p-1}\varphi^*_{il}\Delta y_{i,t-l}+\Sigma_{l=0}^{q-1}\beta^{*\prime}_{il}\Delta x_{i,t-l}+\tilde{u}_{it}$, where $y_{it}$ is the correlation of the difference of logs of stock market prices for country pair $i$ at $t$, and $x_{i,t}$ contains the explanatory variables (measures of FDI and trade, and the correlations of difference logs of industrial production, inflation rate and the correlation of nominal interest rates). The symbols *, **, *** denote significant estimates at 1%, 5% and 10% level. Numbers in parenthesis are standard errors. The Hausman test reports the $\chi^2$ for testing that "$H_0$: difference in coefficients MG - PMG is not systematic", with corresponding $p$ values below. Whenever $p$ value $< 0.05$, the MG estimator is prefered to the PMG. Standard errors not reported for industrial production, $y$, interest rates, $i$ and inflation rate $\pi$. The explanatory variables include only FDI in specification (a) and FDI and other controls in specification (b).

**CS-ARDL(p,q) APPROACH - DFE, MG and PMG Estimates**

| | **CS-ARDL(1,1)** | | | | | | **CS-ARDL(2,2)** | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (a) | | | (b) | | | (a) | | | (b) | | |
| | DFE | PMG | MG | DFE | PMG | MG | DFE | PMG | MG | DFE | PMG | MG |
| $\hat{\theta}_{fdi}$ | 0.0011 | 0.2965*** | 0.1838*** | -0.0107 | 0.2706*** | 0.1600 | -0.0087 | 0.3389*** | 0.1206** | -0.0196 | 0.3542*** | 0.1500 |
| | (0.0371) | (0.0320) | 0.0710 | (0.2037) | (0.289) | (0.1336) | (0.0448) | (0.0267) | (0.0623) | (0.0419) | (0.0325) | 0.2876 |
| $\hat{\theta}_{trade}$ | | | | 0.2037*** | -0.0037 | 0.0336 | | | | 0.1950** | -0.1927** | -0.0005 |
| | | | | 0.0659 | (0.0746) | (0.2104) | | | | (0.0820) | (0.0190) | (0.2775) |
| $\hat{\lambda}$ | -0.6349*** | -0.7518*** | -0.8893*** | -0.6715 | -0.8211*** | -1.0355*** | -0.5241*** | -0.7341*** | -0.9958*** | -0.5626*** | -0.7326*** | -0.9396*** |
| | (0.494) | (0.0998) | (0.0888) | (0.0506) | (0.1267) | (0.1299) | (0.0572) | (0.1197) | (0.0911) | 0.0601 | (0.1582) | 0.2129 |
| Haus $\chi^2$ | | 2.45 | | | 0.47 | | | 12.41 | | | 1.44 | |
| $p$ | | 0.117 | | | 0.7917 | | | 0.0004 | | | 0.4856 | |
| $N \times T$ | | 330 | | | 330 | | | 330 | | | 330 | |

Legend: The CS-ARDL(p,q) specification is given by $y_{it}= c_i+\Sigma_{l=1}^{p}\varphi_{il}y_{i,t-l}+\Sigma_{l=0}^{q}\beta'_{il}x_{i,t-l}+\Sigma_{l=1}^{2}\psi'_{il}\bar{z}_{t-l}+u_{it}$, where $y_{it}$ is the correlation of the difference of logs of stock market prices for country pair $i$ at $t$, and $x_{i,t}$ contains the explanatory variables (measures of FDI and trade). The vector $\bar{z}_t$ contains the cross-sectional means of all variables. The symbols *, **, *** denote significant estimates at 1%, 5% and 10% level. Numbers in parenthesis are standard errors. The Hausman test reports the $\chi^2$ for testing that "$H_0$: difference in coefficients MG - PMG is not systematic", with corresponding $p$ values below. Whenever $p$ value $< 0.05$, the MG estimator is prefered to the PMG. The explanatory variables include only FDI in specification (a) and FDI and other controls in specification (b).