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RETURN COMOVEMENT

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Abstract

We examine intra-market return comovement within each of 33 economies' stock exchanges from 1995 through 2013 using a model-free comovement gauge. We find that the stability of international macroeconomic trilemma policies, the number of crises, and the extent of turnover overshadow the empirical relevance of many variables previously thought to be important for intra-market comovement, including country risk, corruption, and investor protections.

We also use a much longer historical sample of US firms to examine compositional explanations of the well-known US comovement decline and to decompose the comovement into trend and cycle. Our findings challenge the compositional explanations of the decline and suggest that the most recent uptick reflects short-term conditions, rather than a trend reversal.

Keywords: Asset pricing; Information and market efficiency; International financial markets; Financial economics; Trilema

JEL classification: G12; G14; G15; G38; N20

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1. Introduction

A firm's return reflects both the individual vagaries of its underlying business and the conditions shared by other businesses in the economy as a whole. Within each market, the relative importance of the two parts – the individual part and the economy-wide part – varies over time and across countries. Using a market model to separate the individual and economy-wide parts for the market as a whole, Roll (1988) noted that his measure of the individual part had accounted for most of the variation in US returns. The market-wide portion was relatively small. Later, Morck, Yeung, and Yu (2000) showed that measures of the market-wide portion, which they called synchronicity, had fallen over time as US markets evolved in the 20th century. They also showed that their synchronicity measures were larger within individual emerging markets and seemed to decline in the individual markets that increased their financial openness. They interpreted declines in synchronicity as evidence of increasing informational efficiency in some of the individual markets.[1] More deeply, they suggested declining intra-market synchronicity may evince a market's improved financial institutions: investor protections that ultimately support greater economic dynamism.

However, in more recent work re-examining the US, the same authors (Morck, Yeung, and Yu, 2013) and others now note that the decline in US synchronicity has partly reversed itself since the late 1990s. That is, US stock returns have become more synchronised in the current century. If the earlier decline in synchronicity was evidence of rising efficiency due to better financial institutions, then one might be tempted to interpret the subsequent reversal in synchronisation as the opposite: falling efficiency and deteriorating financial institutions.

This temptation is partly forestalled by the sense that US institutional financial arrangements as a whole seldom change abruptly. To the extent that institutions evolve only relatively slowly, sudden changes in synchronisation are likely to reflect other considerations. Likewise, as noted by Alves, Peasnell and Taylor (2010) and by Morck, Yeung, and Yu (2013), the intra-market synchronicity rankings of countries change substantially between periods. Major changes in international rankings

[1] Their interpretation reflects the private informational side of Roll's (1988) assessment that the degree of individual variation might represent (p. 566) "the existence of either private information or else occasional frenzy unrelated to concrete information". On the other side of Roll's assessment are models that allow rational and irrational agents to coexist in equilibrium. See, for example, the emblematic models of Russell and Thaler (1987) in a consumer economy, De Long, Shleifer, Summers, and Waldmann (1990) in a purely financial model, and Barberis and Shleifer (2003) for a model with cross-sectional implications.

make it hard to convincingly interpret a market's ranking as an indicator of its relative financial institutional development.[2]

This paper provides evidence for a new (though complementary) explanation, namely that international financial policy instability and crises underlie the striking repositioning of countries' relative comovement rankings noted by Alves, Peasnell, and Taylor (2010). That is, while intra-market return comovement indeed may reflect the kinds of investor protections that have been suggested, it also may reflect a different aspect of institutions: macroeconomic policy instability, along with other changes that are more transitory in nature. Macroeconomic policy stability is itself a time-varying institutional backdrop to markets' informational efficiency.

We use an international panel that includes observations of policy instability (measured in terms of the international financial policy trilemma, discussed in section 4.2.1) and crises within each country; and we compare the empirical importance of such variables to the relevance of standard measures of countries' institutional quality. The panel includes 33 economies that have had stock markets in place since 1995. Our key results show that international macroeconomic policy instability and crises are important in explaining intra-market comovement behaviour. In international panel regressions, such considerations are better able to account for return comovement within a market than, for example, country risk, corruption and investor protections.

Our work relies on a model-free gauge of comovement.[3] The model-free gauge is closely related to the conventional comovement measure, the R^2 from a market model. Since using the R^2 measure requires imposing a particular empirical version of the market model, conclusions about comovement may inadvertently arise from limitations in the particular model's relevancy or from model instability. So, we use the related, but atheoretic gauge of comovement instead. Specifically, we decompose the total market variation within an economy into two parts: the sum of individual variances and the sum of covariances. Then we gauge comovement within each market using the market's average

[2] Besides changes in financial institutional development, some explanations have focused on specific market features, including, e.g., index inclusion (Barberis, Shleifer, and Wurgler (2005); Chen, Singal, and Whitelaw (2016)), analyst following (Chan and Hameed (2006); Claessens and Yafeh (2013); Hameed, Morck, Shen, and Yeung (2015)), and correlated trading (Kumar and Lee (2006); Kumar, Page, and Spalt (2016)).

[3] The appendix describes how using the conventional R^2 comovement measure to make inferences about informational efficiency requires that the empirical relevance of the underlying model is unchanging. This requirement motivates the use of a model-free gauge.

covariance of returns relative to its average individual variance of returns.[4] Using the model-free gauge, we examine the behaviour of return comovement within each of the markets in the international panel, and we also reexamine the comovement of US returns over a longer period.

Extending the US series to the earlier period (not available for the rest of the international panel) allows us to do two additional things. First, it allows us to address a notable, compositional explanation of the nearly century long decline in the return comovement within the US. Brown and Kapadia (2007) argue that changes in the market's mix of firms explain the downward pattern. To explore this possibility, we construct a sample that is invulnerable to the compositional criticism: a sample of "old" firms that have been listed for nearly all of the sample. We find that return comovement within the US exhibits similar behaviour in both the full US sample and the nearly constant-composition US subsample. Thus, compositional changes do not explain the US trend.

Second, the longer US series allows us to apply well-known time series filters to the model-free measure of comovement to explore the meaning of a recent upturn in US comovement. Such filters are used widely to obtain smoothed-curve representations of economic time series. These representations are designed to separate long-term trend from short-term cyclical fluctuations. The trend and cycle decomposition provides an explicit framework for assessing whether the recent US comovement upturn is a cyclical phenomenon or the reversal in the long trend. We find that the recent upturn is at least partly due to cyclical factors.

Section 2 describes the model-free gauge. The long, US sample is explored in section 3. Section 4 presents the international data and the panel results. Section 5 concludes.

2. A Model-Free Comovement Measure

The model-free, variance-based decomposition begins with the market return, r_m, to a portfolio of N

[4] While we focus on the behaviour of intra-market return comovement, which requires the use of a time series to estimate variances and covariances within each market, we note that the approach builds directly on the work of Allen and Bali (2007), Goyal and Santa-Clara (2003), and Garcia, Mantilla-Garca, and Martellini (2014). Those authors focus on returns' idiosyncratic portions, which they estimate with the cross-sectional variation of returns. The model-free measure also builds on the related work of Goetzmann, Li, and Rouwenhorst (2005), who constructs such a measure with market indices to study the comovement across markets.

firms: $r_m = \sum_{i=1}^{N} \omega_i r_i$ where r_i is the return to the i^{th} firm, and ω_i is the weight of each firm in the portfolio, and we omit the time subscripts. The weights may be chosen arbitrarily, and we use equal weights in the empirical implementation below. Constructing the portfolio over a particular period, it is straightforward to decompose the portfolio's return variance, σ_m^2, into the sum of two parts: a part made up of the underlying individual returns' variances, and a part made up of their covariances.

$$\sigma_m^2 = \frac{1}{N^2} \sum_{i=1}^{N} \sigma_i^2 + \frac{1}{N^2} \sum_{i=1}^{N} \sum_{\substack{j=1 \\ i \neq j}}^{N} \sigma_{i,j}^2. \tag{1}$$

In essence, the pure comovement portion of returns is captured by the second of the two terms in equation 1.

Next, we look at the sample counterparts of each of these two terms. We denote the sample value of each individual return variance by s_i^2, the sample value of each covariance by $c_{i,j}^2$. We denote their corresponding average values by s^2 and c^2, where $s^2 = \frac{1}{N} \sum_{i=1}^{N} s_i^2$, and, $c^2 = \frac{1}{N^2 - N} \sum_{i=1}^{N} \sum_{\substack{j=1 \\ i \neq j}}^{N} c_{i,j}^2$. In terms of these two averages, the sample portfolio variance is:

$$\hat{\sigma}_m^2 = \frac{1}{N^2}\, s^2 + \frac{N-1}{N}\, c^2. \tag{2}$$

This compact expression allows us to readily see that we can construct the average covariance without having to calculate all of the covariances between every pair of firms. That is, the average covariance can be expressed simply in terms of the market variance and the average of the individual variances:[5]

[5] The value weighted average covariance and individual variance portions are given by:

$c_{vw}^2 = \frac{N}{N-1}\left(\hat{\sigma}_m^2 - \frac{1}{N} s_{vw}^2\right)$ and $s_{vw}^2 = \sum_{i=1}^{N} \omega_i \hat{\sigma}_i^2$.

$$c^2 = \frac{N}{N-1}(\hat{\sigma}_m^2 - \frac{1}{N}\, s^2). \tag{3}$$

Finally, we construct a gauge of comovement using c^2 as a share of the average individual variance, s^2:

$$\frac{c^2}{s^2} = \frac{N}{N-1}(\frac{\hat{\sigma}_m^2}{s^2} - \frac{1}{N}). \tag{4}$$

This simple, variance-based comovement gauge, c^2/s^2, comes from the decomposition of the total market variance into its average variance and average covariance pieces.[6] So, it is independent of any model. It also avoids the piecemeal estimation procedure that is needed in order to cobble together a lengthy time series for the index-model-based comovement gauges.[7] This covariance share is simple to calculate, is derived from a straightforward variance decomposition, and is easy to compare with the well-known R^2 comovement measure.[8]

We next calculate this comovement gauge for the sample of all US returns for each month from 1926-2013, then we compare the model-free measure with the traditional, index-model based R^2 measure of comovement that is common in the literature. In section 4, we construct the model-free gauge for the international panel of returns, and we examine the relative importance of institutional influences – such as investor protections and corruption – and macroeconomic influences, such as crises and international policy instability.

[6] As mentioned in section 1, this measure builds on the work of Allen and Bali (2007), Goyal and Santa-Clara (2003), Garcia, Mantilla-Garca, and Martellini (2014), and Goetzmann, Li, and Rouwenhorst (2005).

[7] An index-model based R^2 measure requires periodically re-estimating the underlying model since each firm's coefficients evolve over time. See footnote 29 for an explanation of how the evolution of the estimates of an index model's coefficients can contaminate the behaviour and interpretation of the model's R^2 measure.

[8] Additionally, this measure, c^2/s^2 will equal the average correlation coefficient when $\sigma_i^2 = \sigma_j^2 \forall ij$; and, like the correlation coefficient, $c^2/s^2 \le 1$.

3. US Comovement

The US has been the starting point for most comovement studies, and US returns data are readily available for a relatively long time period. So, we begin by constructing the c^2/s^2 measure for the US. We use data from the Center for Research in Securities Prices (CRSP) from 1926 through 2013. The data include ordinary common stocks of companies headquartered in the US.[9] We construct the comovement gauge annually for the US from 1926 through 2013 using the returns of a random sample of 400 firms.[10]

Figure 1 shows the resulting measures. The yearly average covariance shares, c^2/s^2, are given by the dots, and the red line fits a simple, linear trend. As shown, the covariance share exhibits a downward trend for the sample period as a whole.

As discussed above, the long decline has been interpreted as evidence of improved informational efficiency. However, a number of authors have proffered additional explanations for the observed secular decline in comovement, including several hypotheses about changes over time in the nature of listed firms themselves. For example, in a careful study, Brown and Kapadia (2007) account for earlier reported decline in terms of the increasing number of new, risker firms over the sample period. Here, we address this concern by recalculating the model-free comovement measure for a constant sample: a sample that includes only "old" firms.[11;12] If the long decline is explained by such firm dynamics, then we should expect that the trend would be absent from our sample of old firms.

Figure 2 illustrates the results. As shown, the trend persists in the sample with a constant set of firms. While the comovement share is slightly larger, and its decline is slightly more modest for the old firms,

[9] Our dataset excludes those with fewer than 11 monthly observations, those classified as utilities or as bank and financial firms, and those with a price below $3.00 in any year. See Waszczuk (2014) for a discussion of the exclusion of penny stocks.

[10] Our results are unaffected by various other sample choices as discussed below.

[11] Specifically, our "old" firms sample includes 62 firms, all those in CRSP with at least 84 years of data. That is, we exclude firms with more than three missing years. We also repeat the exercise for only those firms that existed for the complete period. (By allowing for three years of missing data, we are able to increase the sample size from 35 to 62 firms.) The results are similar for both samples.

[12] Dasgupta, Gan, and Gao (2010) also discuss firm demographics, but the issue they raise is somewhat different. They point out that comovement should be higher, not lower, for old firms since investors know more about those firm's time-invariant characteristics. So, keeping the composition constant, as we do here, might by itself suggest that the comovement for such a sample would rise over time as investors get to know the firms. However, the empirical relevance of this point diminishes for a fixed set of firms once the firms are established enough that agents have already learned most of what they can about the firms' stable characteristics.

the differences are minor, and they are not statistically significant at any standard confidence levels. Firm demographics cannot explain the long decline observed in this sample.

We also construct separate, smaller portfolios of 50 random firms, the 50 smallest firms, and the 50 largest firms. The patterns are again the same, and we cannot reject the hypotheses that any of the portfolios differ in terms of either their levels or their trends. In addition to the one-year windows used here, we also construct portfolios with variances and covariances computed over two-year windows. Again, the patterns are the same.

While we have calculated a simple, long-term linear trend, the potential roles of shorter-term conditions and policies suggests a richer exploration of the trend. The behaviour of return comovement potentially encompasses short-term, cyclical changes and slow-moving, secular changes. Return comovement may move systematically with business cycle and policy instability in ways that are relatively temporary. For example, comovement may vary over the business cycle because information production rises with aggregate economic activity, as in Veldkamp (2005) and Veldkamp (2006). Or, it may be cyclical because risk premia increase more in bad times than in good, as described by Mele (2007).[13]

At the same time, the institutional framework of financial markets may continue to change. The arguably slower and more persistent institutional changes may generate a non-constant trend around which the cyclical portion moves. The potential importance of both components – substantial secular changes over relatively long time periods and meaningful fluctuations over the business cycle – suggests that trend and cycle decompositions might be useful. Such decompositions can help distinguish the slow-moving, more persistent institutional influences from the business cycle influences that are more likely to be reversed.

This decomposition is captured by modeling comovement in each period, t, as having two components, a deterministic or stochastic trend, $(\frac{c^2}{s^2})_t^T$ and stationary cyclical part, $(\frac{c^2}{s^2})_t^C$. That is, in each period:

[13] The trend and cycle decomposition is closely related to the empirical work of Brockman, Liebenberg, and Schutte (2010). Building on the theoretical work of Veldkamp (2005), they find an empirical link between return comovement and economic activity.

$$\left(\frac{c^2}{s^2}\right)_t = \left(\frac{c^2}{s^2}\right)_t^T + \left(\frac{c^2}{s^2}\right)_t^C$$

There are many different techniques available to decompose the comovement measure into its trend and cycle components. Here, we use two well-known methods: the Butterworth filter and the Hodrick-Prescott filter, which both treat the trend and cycle as uncorrelated with each other as they allow the trend to change.[14] As before, the decomposition is performed for an annually randomised 400-firm sample.

Figure 3 depicts the estimated trend components using the two techniques. Like the simple, linear trend depicted in Figures 1 and 2, both estimated trend components decline over most of the 20th century. Using either decomposition technique, the trend comovement reaches its nadir in the late 1990s and then begins to rise. However, the reversal seems to be short-lived: the smoothed comovement share renews its decline at the end of the sample. Thus, results from the standard time-series filtering technique support the idea that the recent reversal is at least in part due to cyclical factors.

4. International Data

In this section, we explore the c^2/s^2 comovement measure in an international panel of economies. The use of international data is important for its own sake – to illustrate the patterns of return comovement in other countries. In addition, the inclusion of additional variables in a panel can help expose the country characteristics that may determine comovement patterns within the US and elsewhere. In particular, it can give us the statistical power needed to explore the role of such things as economy-wide investor protections and corruption, or, of macroeconomic conditions, crises, and policy stability. Many characteristics of the US economy that may be related to its observed decline in return comovement have been changing in concert over the period we have examined. We attempt to parse the empirical roles of some of these potentially important characteristics by looking across

[14] These approaches are described in StataCorp (2013), which itself references Butterworth (1930), Hodrick and Prescott (1997), and Baum (2006).

economies where the characteristics have changed to different extents and at different points in time. So, in this section, we first construct and describe the covariance gauges for the economies in the panel; then we explore how they are linked to some of the key characteristics that have been suggested by the literature and for which we were able to obtain data.

The international panel includes 33 economies that have had stock exchanges operating over the period 1995 to 2013.[15] Twenty are OECD countries; and, among those, 12 are European. The return data are taken from Datastream.

4.1 Individual Country Return Comovement

We construct each economy's comovement gauge using equally-weighted portfolios of 50 randomly chosen stocks each period.[16] Figures 4 and 5 display the comovement gauges for the OECD countries and for the non-OECD economies. The most striking feature of these graphs is that the return comovement patterns – their trends, levels, and variability – differ across countries.

While most countries exhibit a decline in return comovement, about one-fifth of them exhibit an overall comovement rise over the sample period. All but one of those with increasing comovement are outside of Europe. As shown in figure 4, the non-European OECD countries that exhibit an increase in return comovement include Australia and Canada, which have quite low levels of comovement to begin with; Chile, for which the rise is negligible; and the US, which was discussed in detail in section 3. Within Europe, only the Netherlands exhibits an overall rise over the sample period. Among the non-OECD countries, only Sri Lanka exhibits an upward trend in return comovement. For the most part, where we do observe increases in return comovement, they appear to be relatively minor.

Perhaps surprisingly, the overall level of return comovement is not necessarily tightly tied to a prominent marker of affluence and development, namely, whether or not a country is part of the OECD. The US and Japan (figure 4) have relatively high comovements, but so does China (figure 5). Likewise, Canada (figure 4) has consistently low values, but so does South Africa (figure 5).

[15] We use 1995 as a starting point in the regressions to allow for a representative balanced panel. When data are available, we use somewhat longer series when looking at the countries individually.

[16] While we have not constructed value-weighted international portfolios, as mentioned in the discussion of US. data in section 5, we found no statistically significant differences across the equally-weighted and value-weighted portfolios, or between portfolios made of different sizes, ranging from 50 to 400.

On the surface, the comovement patterns are in some ways at odds with the detailed and influential study of Morck, Yeung, and Yu (2000), who concluded that richer countries exhibit lower comovement than poorer ones. Their measures of comovement put the US, at the time, at the opposite end of the spectrum from where it is here. The figures here show that their early result, based on a 1995 cross-section, was obtained just as the US neared its comovement lows. The patterns we document here suggest that countries' return comovements change continually and dramatically. In the following section, we use a panel regression to explore the observed comovement patterns and their interpretation more fully.

4.2 International Panel Regressions

The observed changes over time in countries' comovement measures suggest that international rankings of return comovement are likely to be sensitive to the year in which the measures are constructed. In an explicit look at how countries stack up in this regard, Alves, Peasnell, and Taylor (2010) provide year-by-year R^2 rankings over a 20-year period for the same 40 countries originally used by Morck, Yeung, and Yu (2000). Indeed, they document considerable changes in countries' R^2 rankings.[17] In later work, Morck, Yeung and Yu (2013) provide a bar graph with a series of snapshots of countries' R^2 rankings by year over much of the same period. In contrast with the interpretation of Alves, Peasnell, and Taylor (2010), Morck, Yeung, and Yu (2013) interpret their own graph as being consistent with the patterns that they emphasised in their original 2000 paper.[18]

In this section, we use panel regressions to explore the interpretations of the international patterns of comovement. In particular, we are interested in how the observed national comovement patterns are related to countries' underlying characteristics.[19] We begin with the earlier suggestion that stock returns seem to comove more within poor (low per capita GDP) countries than within rich (high per capita GDP) ones. We use multivariate regressions to see if per capita GDP can explain return

[17] Focusing on idiosyncratic volatility in the US, Brandt, Brav, Graham and Kumar (2010) document similar US instability over time.

[18] Morck, Yeung, and Yu (2013) emphasise the relationship between comovement and income, noting that there still appears to be a bivariate, negative relationship between R^2 and per capita GDP. We explore that relationship in more detail below.

[19] Note that, as in most of the extant literature, the problem of endogenous regressors prevents us interpreting regression results causally. However, the multivariate regressions do allow us to compare the explanatory power of policy and macroeconomic variables relative to most of the institutional variables examined in the past.

comovement once we account for crises and the stability of international macroeconomic policy, and for other characteristics that have been suggested by the recent literature. We find that we cannot resuscitate the role of per capita GDP: it remains only tenuously linked to comovement. That is, a country's level of economic development – whether it is rich or poor – does not seem to matter for return comovement.

Prior literature also suggests that various forms of investor protections may drive comovement behaviour. Notably, Morck, Yeung, and Yu (2000) argue that property rights explain the largest differences in comovement that they observe across countries. Jin and Myers (2006) then provide a model to show that, in the presence of limited information, poor investor protections lead to a higher R^2. So, we explore whether measures that previously have been described as being related to investor property rights – such as indicators of investor protections, disclosure, and corruption – can help explain the patterns of comovement. We confirm that many of these measures are significant in bivariate regressions. However, we find that none of these variables appear to matter in multivariate regressions. Instead, comovement appears to be linked to various, shorter-lived aspects of countries' conditions.

As described in section 3, theoretical advances have suggested that return comovement may be affected by short-term macroeconomic conditions. We indeed find that comovement is linked over time to detrended GDP growth, which (in contrast to the level of per capita GDP) represents the stage of the business cycle, not whether a country is rich or poor. While to our knowledge, the extant theoretical literature does not, in addition, explicitly tie comovement to economic crises or to international macroeconomic policy stability, some of the same theoretical forces are also suggestive there. For example, changes in risk premia, as in Mele (2007), and disruptions in information provision, along the lines of Veldkamp (2005) and Veldkamp (2006) may be at work. The early model of Berk, Green, and Naik (1999) may also be relevant: there, a firm's asset composition changes when production is reorganized or productivity changes. The model links economic activity to comovement in the presence of macroeconomic disturbances. One might interpret those disturbances to include crises or to be affected by policy instability. In the empirical work below, we explore this possibility, and we find that comovement indeed is: high during GDP downturns, high during economic crises, and high when international macroeconomic policies are unstable.

We also find that stock market turnover helps explain comovement: greater turnover goes hand in hand with greater comovement.[20] Overall, these variables appear to be more important than many of the earlier variables in explaining changes in observed comovement behaviour.

4.2.1 Regressors

To examine the potential role of investor protections, we begin by including the Djankov, La Porta, de Silanes, and Shleifer (2008) anti-self-dealing index, the Spamann (2010) corrected antidirector rights index, and the World Bank's Business Extent of Disclosure Index. In addition, we include an indicator of corruption: Transparency International's corruption perceptions index. Because of its prominence in the literature, we also include the International Country Risk Guide's Composite Political, Financial, and Economic Risk Rating, which is an amalgam of 22 country characteristics.

To examine the role of shorter-term conditions, we include detrended GDP growth (as distinct from the per capita GDP level), inflation, and stock market turnover. These data are taken from the IMF's International Financial Statistics. In addition, we include a crisis variable constructed from the indicators defined in Reinhart and Rogoff (2011).[21]

To these standard measures, we add an indicator of international macroeconomic policy stability. Specifically, we use a gauge of the stability of a country's triad of: exchange rate arrangements, financial market openness, and monetary sovereignty. The classic, open-economy trilemma constrains a country's choices among these arrangements. (For example, with open capital markets, monetary actions spill over into exchange rate markets. So, a country with open capital markets and an exchange rate peg cannot use its monetary policy to manage inflation.) To capture the stability – or instability – of these arrangements, we use the trilemma stability indicator of Popper, Mandilaras, and Bird (2013). The trilemma indicator gauges the stability of policy, not of the underlying variables. For example, the measure accurately captures the relative stability of US trilemma policies since the mid-70s: the US has largely maintained open financial markets, a dollar float, and monetary sovereignty.

[20] The role of turnover is in keeping with the finding of Karolyi, Lee, and Van Dijk (2012) that comovement is greater in countries with high turnover. It is also related – though more loosely – to the findings of Gassen, Skaife, and Veenman (2016) and Chan, Hameed and Kang (2013) who connect the R^2 measure of comovement to liquidity.

[21] Specifically, we create a summary variable for banking, currency, default, inflation, or stock market crises as defined by Reinhart and Rogoff (2011), and that equals zero in the absence of any of these crises.

This policy stability contrasts with the fluctuations in the underlying variables, such as the foreign exchange value of the US dollar. That is, a constant policy of floating exchange rates means that the value of the exchange rate fluctuates; and a constant policy of monetary sovereignty allows for periods of monetary tightening and monetary ease. The policy can be constant while the affected variables are allowed to fluctuate.

Because differences in comovement might arise as an artifact of size, or of industry structure, we also include the IMF's standard measures of market capitalisation and the country's share of world output; and we include UNCTAD's concentration index.

Separately, we also include an alternative set of institutional variables. These include the World Bank's governance indicators developed by Kaufmann et al. (2010), World Bank data on the extent of business disclosure and the strength of legal rights, and an indicator of press freedom provided by Freedom House. We use these variables to explore the robustness of our results and to provide a more detailed and disaggregated alternative to the ICRG composite.

We transform several of the variables to make linear regression more appropriate and to make the coefficient estimates easier to interpret. Specifically, we use a logistic transformation of the covariance share, $ln\left(\frac{\frac{c^2}{s^2}}{1-\frac{c^2}{s^2}}\right)$;[22] we standardise the subjective indices; and we take logs of countries' per capita GDP, share of world output, market capitalisation, and turnover.[23]

4.2.2 Regression Results

Tables 1, 2, and 3 provide the details of the baseline panel regressions. The regressions in the three tables differ only in their treatment of fixed effects. The first table provides the estimates of regressions that include both country fixed effects and time fixed effects. The second table provides the estimates from regressions that include fixed country effects, but no time fixed effects. By leaving

[22] The transformation follows Morck, Yeung, and Yu (2000), and it is necessary because the covariance share, c^2/s^2, is bounded between zero and one.

[23] We also re-estimate the panel regressions without the log of per capita GDP; instead, we use an indicator for whether or not a country is a member of the OECD in each year. The estimates change little. The purpose of the re-estimation is to ensure that the results are not an artifact of a possible unit root in log per capita GDP, which could not conclusively be rejected and which would make the standard errors unreliable.

out the time fixed effects, we can unpack the role of variables that change over time. Finally, the third table provides estimates from the regressions that include time fixed effects, but no country fixed effects. This allows us to explore the cross-country variation that is hidden in – subsumed by – the country fixed effects of the regressions in the first two tables. That is, table 3 allows us to focus on the variables that change little over time, but vary across countries. Each of the three tables first lists the estimates from regressions on individual variables; then, each table lists the estimates from the full, multivariate regression.

As shown in the first table, the coefficients on both standard measures of investor protections – the antidirector rights index and the anti-self-dealing index – along with the coefficient on the country risk measure are negative, as expected, and statistically significant at the 5 percent or 1 percent confidence levels when included individually. (The country risk indicator taken from the PRS Group, which defines the variable such that a larger value represents a less risk.) However, none of them remains statistically significant in the full regression. Nor is the coefficient on the corruption perceptions index sizable or statistically significant in the univariate regression, or in the full regression. We also see that the coefficient on per capita GDP is statistically insignificant individually, and in the full regression (where it changes sign).

Because multicollinearity could potentially mask the significance of some of these variables in the full regression, we also provide two sets of joint hypotheses. The first tests whether the coefficients on both country risk and corruption perceptions equal zero; and the second tests whether the coefficients on the indices of both antidirector-rights and anti-self-dealing equal zero. As shown at the bottom of the table, neither hypothesis can be rejected at standard confidence levels.

Instead, some of the shorter-term conditions appear to be important. The coefficients on trilemma stability, turnover, and inflation are all statistically significant, and the coefficients on trilemma stability and turnover are large enough to be economically meaningful.[24] A few examples may give some insight into their sign and quantitative relevance.

The Asian crisis of the 1990s provides a telling trilemma stability example. At the time of the crisis, Indonesia's international macroeconomic policies changed considerably and then stabilised. The

[24] While the coefficient on inflation is small, its sign is puzzling and will be seen to persist in all of the regressions.

policy instability is captured by a trilemma measure of about -0.6, which is substantial relative to the rest of the sample. As the crisis began to abate, Indonesia's policies steadied and the trilemma measure shrank to about -0.2. The estimated trilemma coefficient of -0.06 given in table 1 implies that Indonesia's return to stability corresponds to a decline in return comovement of about 0.03.[25] While smaller than the actual covariance decline of about 0.05 at the time, it accounts for a large portion of the decline.[26]

The quantitative relevance of turnover can be seen clearly in a US example. Turnover changes of as much as 50 to 100 percent are large, but not rare in the sample. The US experienced such a change between 2007 and 2008, when US stock market turnover doubled at the time of the global financial crisis. As shown in table 1, the coefficient on the natural log of turnover is about 0.13. The implied increase in US comovement would be about 0.01, which is about a fifth of the actual US comovement change of about 0.05. Like the coefficient on trilemma policy stability, the coefficient on turnover is both statistically significant and of an economically meaningful magnitude.

The next table gives the results from regressions that include country fixed effects, but no time fixed effects. In the individual regressions, we again find that the coefficients on country risk, antidirector rights, and anti-self-dealing are negative and statistically significant. The coefficient on per capita GDP is now statistically significant as well, and negative, as suggested by Morck, Yeung, and Yu (2000, 2013). However, as before, none of these variables is significant in the full regression.

The bottom of the table gives the test statistic for the joint hypotheses that the political risk and corruption coefficients are zero, and that the antidirector rights and anti-self-dealing coefficients are zero. Neither hypothesis can be rejected at any standard confidence level: the full regression again provides no support for the idea that corruption and country risk, or traditional measures of investor protections drive comovement.

In the full regression, the estimated coefficients on several of the shorter-term variables are again statistically significant, including the coefficients on turnover and inflation. (The trilemma coefficient

[25] To assess the implied quantitative impact, we divide the estimated coefficient by the logistic transformation's derivative, which equals $\frac{1}{\frac{c^2}{s^2}\left(1-\frac{c^2}{s^2}\right)}$ and varies with the covariance share. Here, we evaluate the derivative at the value of $\frac{c^2}{s^2}$ equal to Indonesia's actual covariance share, 0.18. We also account for the standardisation of the trilemma measure.

[26] Indonesia's average comovement is 0.11.

point estimate is little changed, but its larger standard error renders it statistically indistinguishable from zero.) In addition, this table allows us to see the significance of variables whose variation occurs largely in the time dimension. With country fixed effects, but no time fixed effects, table 2 highlights the roles of GDP growth and crises: their estimated coefficients are now statistically significant and economically meaningful. At the average covariance share of .11, a one standard deviation decline in GDP implies a rise in the covariance share of about 9 percent; and a crisis implies a rise of about seven percent. Without time dummies, we can now see that low growth and crisis periods are accompanied by higher comovement.

The next table provides results from regressions that include time fixed effects, but no country fixed effects. Most of the main results of the first table are also evident here. The estimated anti-self-dealing, antidirector rights, country risk and corruption coefficients are again negative and statistically significant at standard confidence levels in the bivariate regressions, but not in the full regression; and the two traditional measures of investor protections again appear jointly unimportant. The joint insignificance of corruption and country risk (the amalgam of 22 variables) is rejected at the five percent significance level; however, that significance disappears when time dummies are also omitted. As in tables 1 and 2, it is again other variables that appear to be important in the multivariate regression.

The coefficients on trilemma stability, turnover, and inflation are again all statistically significant, and the coefficients on trilemma stability and turnover remain large enough to be economically meaningful.

The final table provides similar regressions, but it substitutes a broader set of indicators for the anti-self-dealing and antidirector rights indices. The measures used in these regressions are intended to capture the traditional idea of legal and institutional protections somewhat more broadly. Specifically, the regressions use the World Bank’s indicators described above: the business index of the extent of disclosure of ownership and financial information; the index of strength of legal rights, which gauges protections in lending; and the Kaufmann, Kraal, and Mastruzzi measures of government effectiveness, regulatory quality, political stability, press freedom, voice and accountability, and rule of law.

Here we show only the full multivariate regressions, but – as before – many of the individual variables are significant in individual regressions. The table's first column gives the results from the regression that includes both time and country fixed effects, as in Table 1. The middle column gives the results with country fixed effects, but no time fixed effects, as in Table 2. The final column gives the results with only time fixed effects, but not country fixed effects, as in Table 3.

As before, the estimated coefficients on trilemma policy stability and turnover are again negative and statistically significant, and their point estimates are of similar magnitude to their earlier values. Likewise, removing the time fixed effects again reveals the coefficients on demeaned GDP growth and crises to be statistically significant; that is, comovement again appears to be high when GDP growth is low and in periods of crisis. Additionally, as was the case in the original full regressions, none of the variables that might be considered indicative of traditional legal and institutional protections or country risk are significant in any of these full regressions. Finally, neither of the joint hypotheses given at the bottom of the table can be rejected. None of these multivariate regressions provide support for the idea that traditional measures of legal and institutional protections and country risk are important in explaining variation in comovement; instead other aspects economic and institutional environments seem to be important, namely: trilemma policy stability, turnover, and crises.

5. Conclusions

In this paper, we use a simple gauge of within-market return comovement that relies only on a variance decomposition of the market return, and not on any particular model of returns to explore the behaviour of stock return comovement.

Applying the simple gauge to a recent international panel, we observe that return comovement is not, as had once been thought, tied to whether a country is rich or poor. Nor is it tightly tied to traditional measures of investor protection and country risk. Instead, it is more closely related to variables that may reflect different, shorter-term aspects of institutions, including international macroeconomic policy stability. These variables help explain the dramatic changes in the behaviour of countries' return comovement over the last several decades.

We also challenge findings that attribute the long-term downward trend in US comovement to compositional factors. We find that the trend appears even in a sample that has a nearly constant composition. Finally, we attribute much of the recent US comovement uptick to cyclical factors, rather than to a reversal in the long trend observed over much of the 20th century.

Many existing cross-country comparisons of return comovement highlight differences in the quality of the markets' financial institutions differences that correspond to variation in informational efficiency. Our findings suggest that institutional quality and its link to informational efficiency should be considered in a broad context, one that includes the stability of macroeconomic institutions as well as financial ones. Future firm-level research linking return comovement to other variables, such as analyst following, foreign ownership, corporate structure, and corporate culture, would benefit from also considering interactions with broader and more transitory economic conditions, such as international macroeconomic policy instability.

Appendix: The R^2 Comovement Measure

In this appendix, we discuss some aspects of the construction of the R^2 measure that may confound its interpretation. In particular, we note that using R^2 to make inferences about informational efficiency requires the empirical relevance of the underlying model to be unchanging. We also replicate and extend the US results of Morck, Yeung, and Yu (2000); and, using the R^2 measure, we confirm the model-free result that the co-movement of "old" firms follows the same pattern as the larger sample, and we decompose the R^2 into trend and cycle.

The R^2 measure, used by Roll (1988) and many others since then, is motivated by the Capital Asset Pricing Model (CAPM) and the model's simplest empirical incarnation, the single index model. While Roll (1988) begins with the single index model for the US, he also includes a five-factor model; and many other researchers incorporate specific additional factors, such as the Fama and French (1993) factors (size and book-to-market ratio), momentum (as in Carhart (1997)), or industry, regional, and global returns. Here, we focus on the benchmark US, single-index model of Morck, Yeung, and Yu (2000).

Following others, we denote the average R^2-based gauge of this synchronicity by f_R. So, in each sub-sample period, t, with equally-weighted returns, we have:

$$f_{R,t} = \frac{1}{N}\sum_{i=1}^{N} R_{i,t}^2 = \frac{1}{N}\sum_{i=1}^{N} \frac{\hat{\beta}_{i,t}^2\hat{\sigma}_{m,t}^2}{\hat{\beta}_{i,t}^2\hat{\sigma}_{m,t}^2 + \hat{\sigma}_{e_i,t}^2}$$

where $R_{i,t}^2$ is the R^2 from the regression of the i^{th} return on the market return in the subsample; $\hat{\beta}_{i,t}$ is the i^{th} stock's estimated coefficient on the market return in the subsample; $\hat{\sigma}_{m,t}^2$ is the sub-sample variance of the market return; and $\hat{\sigma}_{e_i,t}^2$ is the residual, sub-sample variance. That is, the market model's average ability to explain returns in each subperiod provides a comovement gauge in that subperiod.

Replicating and Extending f_R for US Data

Following Roll (1988), Morck et. al. (2000, 2013), and others, we use five-year estimates of the βs and monthly returns. However, unlike previous studies, we construct the R^2 measures using annual subsamples so that we can begin to consider the role of the business cycles.[27] We report the US results for value-weighted returns, but the results change little when equally-weighted returns are used.

This benchmark R^2 measure of comovement is shown in the blue line in figure 6, which uses the same randomised sample of 400 firms. The overall pattern replicates Morck, Yeung, and Yu's (2000) finding that the R^2 declines over much of the 20th century, and it replicates and extends the later finding (Morck, Yeung, and Yu, 2013) that the measure seems to reach its nadir and begin to reverse itself after the mid-1990s.[28] The vertical line indicates the end of the 2003 to 2007 boom and the onset of the financial and economic crisis. Since the onset of the financial crisis, the measure has returned to its levels of the 1960s and 1970s.

As we did with the model-free measure, we must again consider the compositional arguments of Brown and Kapadia (2007) and others who suggest that the decline may represent changing firm demographics. So, we recalculate the average R^2 for the smaller sample that includes only the 62 "old" firms. The result of this exercise is shown by the red line in figure 6. While the long decline is no longer quite as compelling, we can see that, like the overall R^2 measure, the average old-firms' R^2s also reaches its nadir in the mid-1990s and rises afterward.

A Generic R^2

While many researchers have used the R^2 as a comovement gauge, returning to an R^2 statistic's traditional meaning provides additional insight. Typically, (and as characterised by Roll's 1988 address on the subject), an R^2 statistic is used to describe the empirical fit of a model. That traditional meaning reminds us that the numerical value of the R^2 is influenced largely by two things: the

[27] We also separately calculate f_R using the full five-year sub-periods as in Morck, Yeung, and Yu (2000); and, while we report the results from annual returns constructed from monthly data, we also separately calculate the R^2 using daily and weekly observations to construct the annual returns. The results differ little. Most importantly, the measures exhibit the same long decline and recent reversal.

[28] Morck, Yeung and Yu's (2013) data ends at 2010. Here, we extend the sample through 2013.

variability of the explanatory variables, and the empirical validity of the model.[29] An R^2 increases with variability in the explanatory variables; and it increases with the model validity.

In this case, the explanatory variable is the market return. So, even without any knowledge of such things as informational efficiency, property rights, or compositional changes, we should expect the average R^2 to rise in periods of greater market volatility and fall in periods of tranquility. Figure 7 depicts these patterns for the same US returns used above. The figure's red line gives the standard deviation of the US market return, and the blue line gives the average US R^2. While the two measures do not move in lock step, periods of substantial declines in R^2 are accompanied by periods of decline in market volatility; and the recent rise in R^2 has been accompanied by a rise, albeit an uneven one, in the standard deviation of the market return. This observation is in keeping with the findings of Bekaert, Hodrick, and Zhang (2012), who carefully attribute time variation in related variables to overall market volatility.

Since an R^2 also depends on the model's validity, an interpretation of comovement as being indicative of informational efficiency requires a joint hypothesis. Specifically, in addition to the hypothesis of interest, it requires the auxiliary hypothesis that the relevance of the market model is unchanging. The relevance of a particular market model will change if, for example, the importance of other factors changes over time or across countries. If the model's validity is changing, the R^2 will change as well, and the corresponding changes would be conflated with the interpretation of the R^2. Ignoring the change in validity would lead to incorrect inferences. In general, when a model's R^2 falls, one may reasonably be concerned that the model's relevance is declining. So, model relevance complicates inferences about informativeness.[30]

[29] In addition, for each firm, $R_i^2 = \frac{\hat{\beta}_i^2 \hat{\sigma}_m^2}{\hat{\sigma}_i^2}$. So, its value also changes as $\hat{\beta}_i^2$ changes when it is re-estimated in every sub-period, whether every year or every five years. While β_i must sum to one across all firms, the sum across firms of its squared value can change over time. Thus, the periodic reestimation of the$\beta_i s$ also affects the average R^2. We find the $\hat{\beta}_i^2$ term itself does indeed fall over much of the sample period, and then rise. However, the quantitative significance of this in accounting for the change in the average R^2 is modest.

[30] Morck, Yeung, and Yu (2000), and others have used the proportion of firms with returns moving in the same direction as a model-free gauge. Specifically, the measure counts the number of firm returns of the same sign as a fraction of the total number of firms. While this fraction captures a sense of directionality without recourse to a model, it exhibits two important drawbacks. First, it lumps small and large movements together. That is, it does not take into account the extent of the comovement; instead it is all or nothing. Second, the measure rises and falls with the magnitude of the mean return. This latter problem is particularly important when comparing time periods or countries, where the average nominal returns can differ greatly. Time periods with high nominal returns, for example, will appear to have more comovement.

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Figure 1: U.S. Covariance Share, $\frac{c^2}{s^2}$

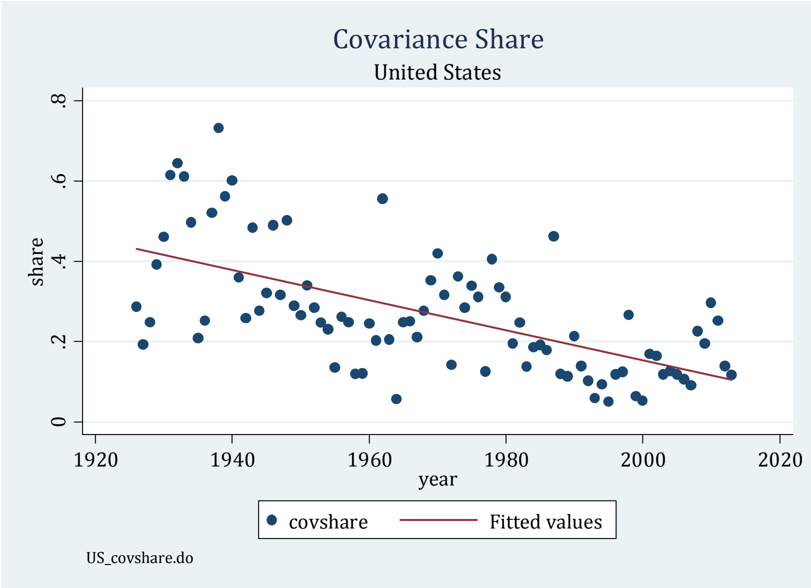

Figure 2: U.S. Old Firms Covariance Share, $\frac{c^2}{s^2}$

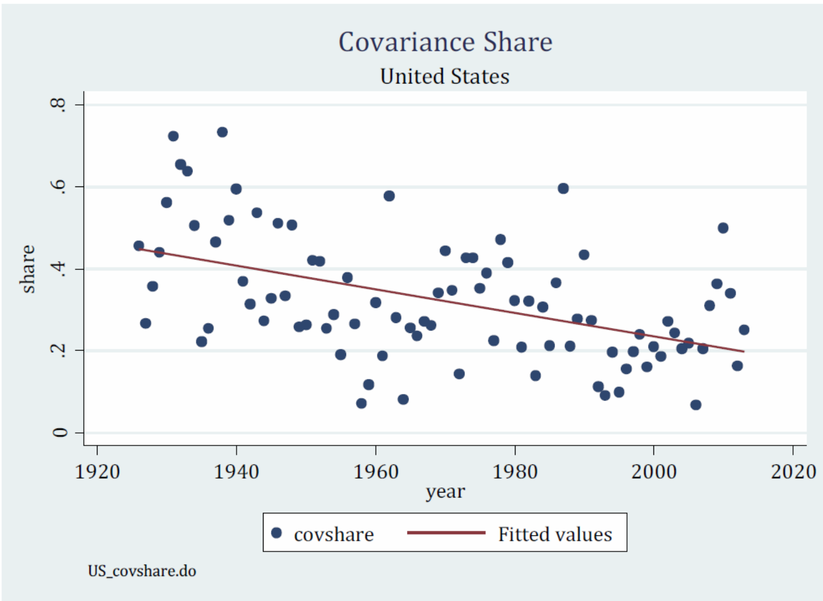

Figure 3: Trend component of Covariance Share, $\frac{c^2}{s^2}$

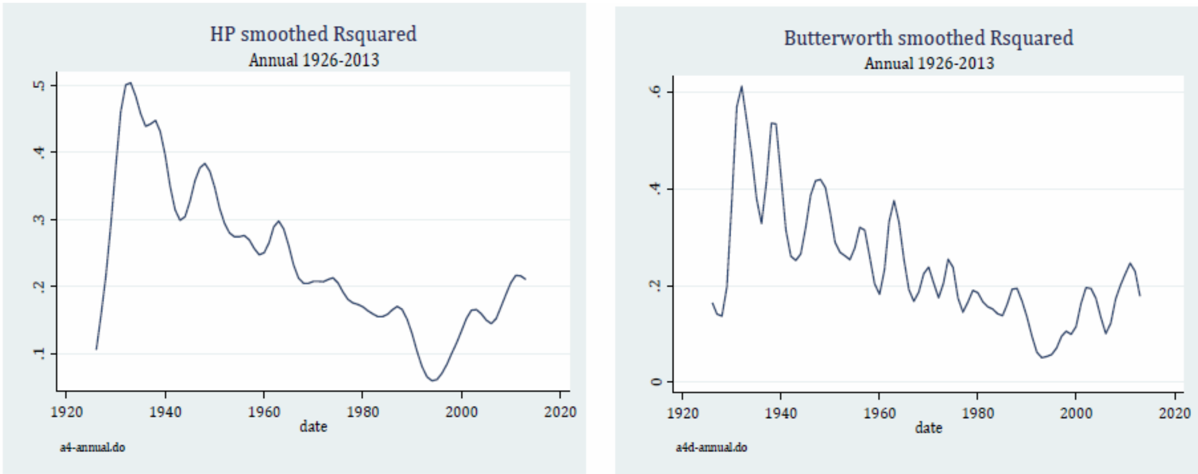

Figure 4: Covaraiance Share in OECD Countries

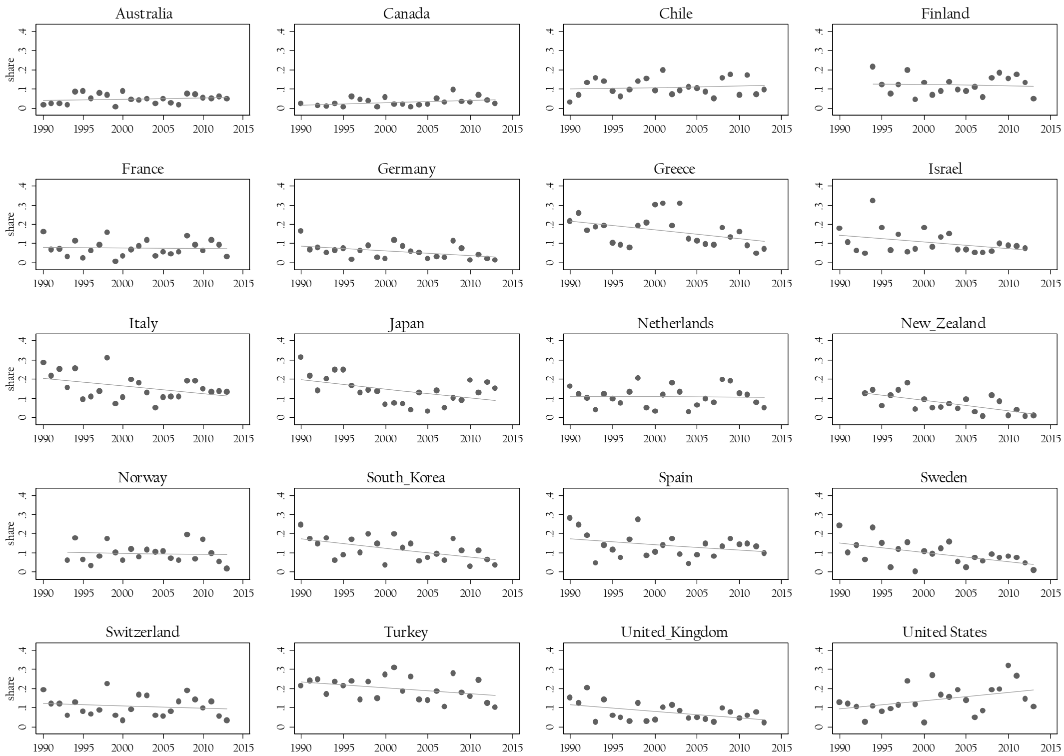

Figure 5: Covaraiance Share in non-OECD Countries

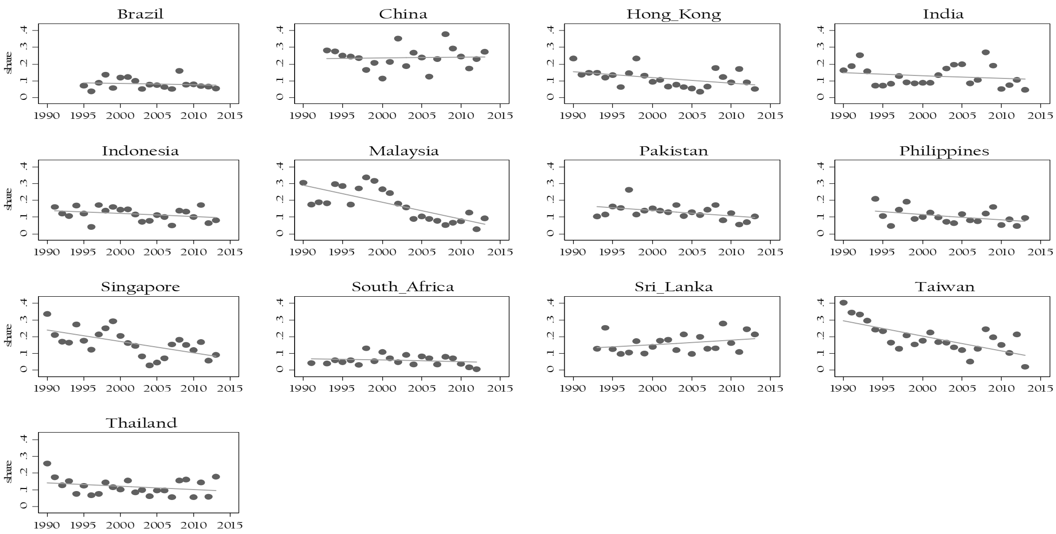

Figure 6: R^2 of the single index model: old-firm sample

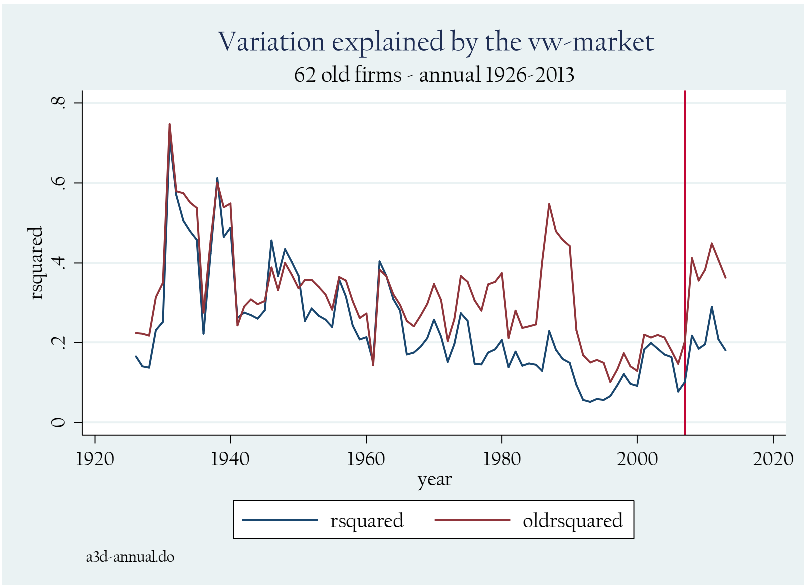

Figure 7: Variability of explanatory variable and R^2

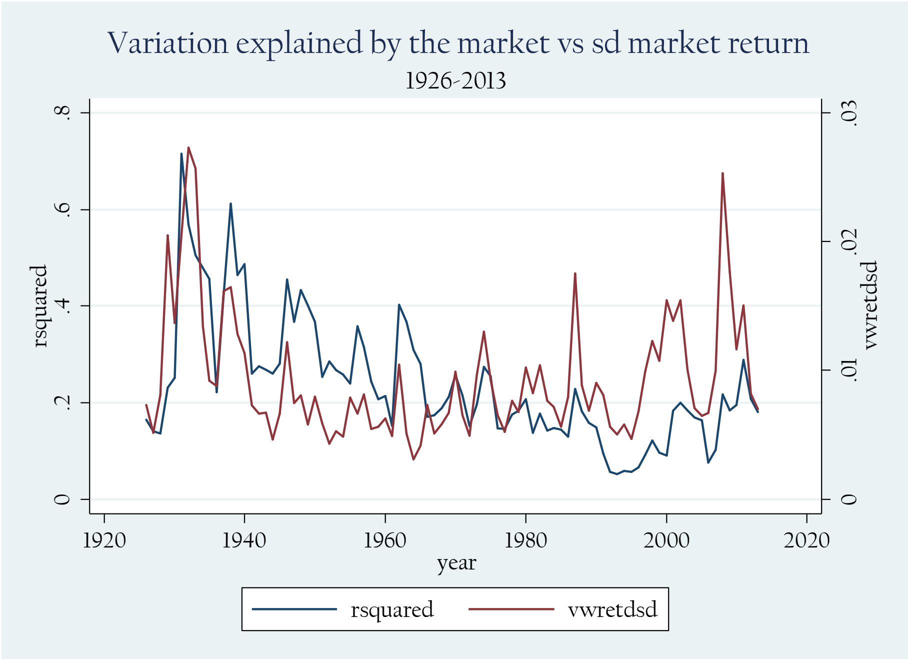

Table 1: Comovement panel regression with time and country fixed effects

Covariance Share	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Country risk	-0.1950**							-0.1924
	(0.084)							(1.619)
Corruption		0.0039						-0.0245
		(0.050)						(0.055)
Antidirector rights			-0.3496***					-2.6650
			(0.000)					(2.933)
Anti-self-dealing				-1.8520***				5.6993
				(0.000)				(6.620)
Concentration					0.0040			0.0529
					(0.097)			(0.125)
ln(Output share)						-0.3831		-1.1284
						(0.383)		(1.490)
ln(per capita RGDP)							-0.1958	0.0917
							(0.378)	(1.619)
Demeaned growth								0.0158
								(0.015)
Trilemma stability								-0.0551**
								(0.021)
ln(Market cap)								0.1487
								(0.127)
ln(Turnover)								0.1380**
								(0.060)
Crisis								0.0688
								(0.053)
Inflation								-0.0080*
								(0.005)
Ho_1: Country Risk = Corruption = 0								1.81
Probability > χ^2								(0.4054)
Ho_2: Country Risk = Corruption = 0								3.16
Probability > χ^2								(0.2064)
Observations	626	626	569	626	626	626	626	417
Number of countries	33	33	30	33	33	33	33	26
R^2	0.564	0.557	0.543	0.557	0.557	0.560	0.558	0.611
Country dummies	yes	yes	yes	yes	yes	yes	yes	yes
Time dummies	yes	yes	yes	yes	yes	yes	yes	yes

Notes: The dependent variable is the logistic transformation of Covariance share, c^2/s^2, defined as the average equity return covariance scaled by the average equity return variance. Data for the 33 countries and 19 years (1995-2013) are included. The explanatory variables are: Country risk, defined as the (standardised) composite country risk indicator produced by PRS Group; the Corruption perceptions risk indicator from transparency international; the (standardised) corrected Antidirector rights indicator from Spamann (2010); the (standardised) Anti-self-dealing indicator taken from Djankov, et al (2008); a (standardised) measure of trade product Concentration taken from the United Nations Conference on Trade and Development statistical system; countries' (logged) Output share of world GDP, (logged) PPP-based per capita Real GDP, and Demeaned GDP growth expressed as a deviation from its 1995-2013 country mean, all taken from the World Economic Outlook Database, October 2014; a measure of macroeconomic policy stability (Trilemma stability) taken from Popper, Mandilaras and Bird (2013); Market capitalisation Market cap and Turnover taken from the World Development Indicators; a Crisis indicator taken from Reinhart and Rogoff (2011); and Inflation, taken from the World Economic Outlook Database, October 2014. Robust standard errors are reported in parentheses and are clustered at the country level, and asterisks indicate statistical significance at the one (***), five (**) and 10 percent (*) levels. Ho_1 and Ho_2 are the hypotheses: (1) country risk and corruption coefficients jointly equal zero, and (2) both corporate governance indicators (antidirector rights and anti-self-dealing) jointly equal zero.

Table 2: Comovement panel regression with country fixed effects

Covariance Share	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Country risk	-0.1591**							-0.0340
	(0.080)							(0.122)
Corruption		0.0149						0.0091
		(0.063)						(0.060)
Antidirector rights			-0.3496***					-1.3571
			(0.000)					(1.348)
Anti-self-dealing				-1.8520***				2.7988
				(0.000)				(2.855)
Concentration					-0.0847			-0.0327
					(0.103)			(0.125)
ln(Output share)						-0.1559		-0.4245
						(0.312)		(0.669)
ln(per capita RGDP)							-0.3472**	-0.1610
							(0.148)	(0.340)
Demeaned growth								-0.0305**
								(0.014)
Trilemma stability								-0.0599
								(0.037)
ln(Market cap)								-0.0886
								(0.124)
ln(Turnover)								0.1217*
								(0.068)
Crisis								0.1439***
								(0.045)
Inflation								-0.0064*
								(0.004)
Ho_1: Country Risk = Corruption = 0								1.81
Probability > χ^2								(0.4054)
Ho_2: Country Risk = Corruption = 0								3.16
Probability > χ^2								(0.2064)
Observations	626	626	569	626	626	626	626	417
Number of countries	33	33	30	33	33	33	33	26
R^2	0.406	0.401	0.361	0.401	0.402	0.401	0.405	0.475
Country dummies	yes	yes	yes	yes	yes	yes	yes	yes
Time dummies	yes	yes	yes	yes	yes	yes	yes	yes

Notes: See table 1.

Table 3: Comovement panel regression with time fixed effects

Covariance Share	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Country risk	-0.1741***							-0.2190
	(0.056)							(0.148)
Corruption		-0.0632**						-0.0636
		(0.029)						(0.044)
Antidirector rights			-0.0208					-0.0806
			(0.060)					(0.061)
Anti-self-dealing				-0.0491				-0.0660
				(0.078)				(0.066)
Concentration					0.0201			0.0655
					(0.065)			(0.069)
ln(Output share)						-0.0667		-0.2308***
						(0.084)		(0.080)
ln(per capita RGDP)							-0.1848**	0.0880
							(0.074)	(0.115)
Demeaned growth								0.0117
								(0.013)
Trilemma stability								-0.0568**
								(0.022)
ln(Market cap)								0.1035
								(0.120)
ln(Turnover)								0.2057***
								(0.066)
Crisis								0.0814
								(0.051)
Inflation								-0.0086*
								(0.005)
Ho_1: Country Risk = Corruption = 0								6.19**
Probability > χ^2								(0.04054)
Ho_2: Country Risk = Corruption = 0								4.05
Probability > χ^2								(0.1322)
Observations	626	626	569	626	626	626	626	626
Number of countries	33	33	30	33	33	33	33	33
R^2	0.189	0.226	0.180	0.158	0.157	0.153	0.195	0.355
Country dummies	yes	yes	yes	yes	yes	yes	yes	yes
Time dummies	yes	yes	yes	yes	yes	yes	yes	yes

Notes: See table 1.

Table 4: Comovement panel regression with alternate indicators

Covariance Share	(1)	(2)	(3)
Government effectiveness	-0.2757	-0.2588	-0.2677
	(0.218)	(0.245)	(0.236)
Regulatory quality	0.2529	0.2595	0.2468
	(0.213)	(0.259)	(0.188)
Political stability	-0.1135	0.0151	-0.0285
	(0.101)	(0.097)	(0.093)
Press freedom	-0.0063	-0.0045	0.0054
	(0.007)	(0.008)	(0.005)
Voice and accountability	0.0700	-0.0750	-0.0421
	(0.185)	(0.246)	(0.186)
Rule of law	-0.2269	-0.2902	-0.1872
	(0.353)	(0.371)	(0.335)
Business disclosure	0.4414	0.0089	-0.0179
	(0.896)	(0.280)	(0.025)
Strength of legal rights	-30.0592	-3.1698	-0.0242
	(51.288)	(17.962)	(0.037)
Concentration	0.0629	-0.0058	0.0169
	(0.125)	(0.129)	(0.067)
Ln(Output share)	-0.7274	0.1371	-0.1482*
	(1.570)	(0.529)	(0.080)
Ln(per capita Real GDP)	0.3915	0.0145	0.0549
	(1.597)	(0.265)	(0.104)
Demeaned growth	0.0062	-0.0296***	0.0066
	(0.011)	(0.010)	(0.009)
Trilemma stability	-0.0559***	-0.0611*	-0.0451**
	(0.019)	(0.034)	(0.019)
Ln(Market capitalization)	0.0964	-0.1125	0.0413
	(0.118)	(0.106)	(0.099)
Ln(Turnover)	0.1282**	0.1001*	0.1831***
	(0.063)	(0.059)	(0.059)
Crisis	0.0771	0.1463***	0.0935*
	(0.054)	(0.039)	(0.051)
Inflation	-0.0065*	-0.0077**	-0.0091***
	(0.004)	(0.004)	(0.003)
Ho_1:	7.57	3.49	8.70
Probability > χ^2	0.2713	0.7455	0.1912
Ho_2:	2.38	0.76	1.10
Probability > χ^2	0.3035	0.6854	0.5757
Observations	449	449	449
Number of countries	28	28	28
R^2	0.625	0.509	0.380
Country dummies	yes	yes	no
Time dummies	yes	no	yes

Notes: The explanatory variables are: *Government effectiveness, Regulatory quality, Political stability, Voice and accountability*, and *Rule of law*, all taken from the Worldwide Governance Indicators, 2014 update; the extent of *Business disclosure* and the *Strength of legal rights* taken from the World Bank; *Press freedom* taken from Freedom House; and additional variables (*Concentration, Output share, per capita Real GDP, Demeaned growth Trilemma stability, Market capitalisation, Turnover, Crisis, and Inflation*), which are defined in Table 1, which also defines the dependent variable and gives the sample size. Robust standard errors are clustered at the country level and reported in parentheses; and asterisks indicate statistical significance at the one (***), five (**) and 10 percent (*) levels. The table also reports tests of two hypotheses, Ho_1: the coefficients on the first six variables jointly equal zero; and Ho_2: the coefficients on the extent of *Business disclosure* and the *Strength of legal rights* jointly equal zero.