Are Common Swings in International Stock Returns Justified by Subsequent Changes in National Outputs?

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Abstract

In an integrated world capital market, the same pricing kernel is applicable to all securities. If the kernel is excessively volatile, as has been found in some studies, it should translate into an excessive degree of correlation in the returns of different equities. In this paper, we apply this idea to the stock returns of different countries. First, we determine, for a given, measured degree of commonality of country outputs, what should be the degree of correlation of national stock returns. We then match the correlations of the combined model containing the statistical model for output and the financial model for stock returns. We find that actual correlations are not excessive and, in fact, are about what they should be in a unified market but that they are much larger than they should be in fully segmented financial markets.

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Shiller (1981) shows that in the United States “stock prices are too volatile to be justified by subsequent changes in dividends”. This is confirmed by Campbell (1996) for a number of other countries. The “excess volatility” puzzle is commonly ascribed to an excessive degree of volatility of the pricing kernel.\(^1\) In integrated world capital markets, the same pricing kernel is applicable to all securities. If the kernel is excessively volatile, this should translate into an equally excessive degree of correlation of world equity returns.

The goal of our paper is to explore the notion of correlation, much like Shiller (1981) did for volatility. We ask the question, for a measured commonality in country outputs, what should the correlation among equity returns be? To answer this question, we combine a statistical model of the business cycle with an asset pricing framework.

Our approach offers an opportunity to understand the interplay between the real economy and stock returns. Traditionally, four routes have been taken. First, a number of papers (see, for example, Fama, 1990) show that there is a relation between expected output and stock returns. These are generally statistical exercises and it is not possible to use these results to understand correlations of different countries or sectors. Second, Hamao, Masulis and Ng (1990) and the many papers that followed this work study the spillover of information from one economy to another. While these studies are important in tracing the type of information that causes common movement in expected returns and volatility, they do not give us a starting point. That is, they do not answer the main question in our study: What should the level of correlation be?

Asset pricing tests offer a third possible route of examination. These tests specify common factors which each country has sensitivity to (see Ferson and Harvey, 1993). We can deduce from the estimated sensitivities to the common factors what the correlation of equity returns should be. For example, in a one factor world, if one country as a positive sensitivity and the other has a negative sensitivity, the correlation of the two countries’ returns is negative. Correlation is determined by a statistical model that determines the relative movement of each country’s return versus some global benchmarks.

Recent asset pricing literature (see Fama and French 1992, Daniel and Titman 1997) has pursued yet another approach. Instead of measuring risk exposure has a sensitivity to a common factor, the stock returns are directly determined by company characteristics. This involves the measuring of certain fundamental data. Differences in fundamental data determine differences in price behavior. Unfortunately, there is little theoretical guidance on how to select the fundamental characteristics.

In a way, the fundamental approach is most closely related to our contribution. Instead of working with accounting and financial data, the fundamentals we focus on are the economic fundamentals. We use a fairly standard approach that is often applied in international business cycle research. We specify a model

\(^1\)“Excessive” is understood relative to the observed degree of volatility of consumption.
that identifies a common world business cycle. The deviation of any country’s business cycle from the world business cycle is defined to be the country-specific business cycle.

The intuition is that, if most of the variation in economic activity in two countries is associated with the world business cycle, then the two countries should have high equity correlations. We formalize this intuition using a log-linear asset pricing model. After making some choices on the form of the utility function and some distributional assumptions, we are able to determine the model’s implied level of correlation for two countries’ returns.

In the end, international asset pricing poses a number of challenges. For example, when it is based on an asset pricing model which assumes that markets are completely integrated, one could interpret a failure of our approach (model correlations different from observed correlations) as either evidence against our specification or against the hypothesis of market integration. But our idea is more general. We are able, for purposes of comparison, to apply the technique under the hypothesis of market segmentation.

Our paper is organized as follows. Section 1 explores the data and the phenomena that we are trying to explain. In Section 2, we develop the dynamic single-index model of Stock and Watson (1993) which we will use to define each country’s business cycle. The log-linear pricing kernel of Restoy and Weil (1993) is explained in Section 3. Section 4 applies the log-linear pricing kernel to the dynamic single-index business cycle model to derive equilibrium security returns. We then examine, in Section 5, the correlations implied by the model and the actual correlations observed in the data. Section 6 develops a statistical test of the hypothesis of financial market integration. Some concluding remarks are offered in the final section.

1  A first look at the data

In the empirical analysis below, we focus on the behavior of industrial production and stock returns in twelve OECD countries on a monthly basis from January 1970 to June 1996. Figure 1 shows a positive relation between the correlation of a country’s output with OECD output and the correlation of that country’s stock market returns with OECD stock returns. In order to abstract from the effect of a country’s size, we consider a country’s correlation with the other countries, itself excluded. The index of the other countries’ output is calculated with weights that reflect each country’s output size. To account for the possibility of lags, the correlations of a country’s output with the OECD output is, in fact, the square root of the $R^2$ in a multiple regression of the country’s industrial production on aggregate OECD industrial production (the country itself excluded), contemporaneous and with eleven monthly lags.
Most of the work in our paper is calibrated to industrial production growth. There are two important questions: (i) how good of a proxy for the business cycle is industrial production and (ii) how closely related are the industrial production growth correlations with measures of consumption growth correlations? The second question is important because the model we use assumes that consumption is equal to output.

The main reason that we chose industrial production was that it was available on a monthly basis for all 12 countries in our sample. We also collected real GDP data which was available on a quarterly basis for nine of 12 countries and on an annual basis in the other three countries. We then calculated for each country their GDP correlations with the rest of the world and compared these to the industrial production correlations (these results are available on request). We found very similar patterns between the GDP and industrial production correlations. Both GDP and industrial production correlations are positively related to equity correlations. There were two countries, Germany and Japan, who had GDP growth correlations with the rest of the world that were far smaller than the industrial production correlations. Nevertheless, there seemed to be a reasonable correspondence between GDP and industrial production.

We also conducted sensitivity analysis on the number of lags and the frequency of measurement of industrial production. We found that the multiple correlation measures produced considerably higher correlations than using two lags or no lags. We also compared quarterly correlations with the monthly correlations and found broad similarity.

Perhaps predictably, the international real consumption growth correlations appeared to be different from the other measures of output. Indeed, the correlation between the consumption growth correlations and the industrial production growth correlations is positive but not statistically significant. As with the GDP data, consumption was available on a quarterly basis for nine of the 12 countries and on an annual basis for the other countries. The problems with consumption data are well known. While our model requires the assumption that consumption equals production, we are much more comfortable using production data.

Before we proceed, one caveat is in order. It is related to exchange rate movements. The data we use for each country are expressed in U.S. dollars. Randomly fluctuating exchange rates can cause a disconnection of realized returns expressed in local currency, since, in theory, they ought to be linked by an equilibrium pricing relationship applicable to returns expressed in a common currency.\(^3\) Although this could be investigated further, one actually finds that stock returns expressed in dollars exhibit approximately the same measured correlations as do stock returns expressed in the respective local currencies.

Consistent with the observations in Ammer and Mei (1996), the correlations

\(^3\)On that count, see the empirical results of Dumas and Solnik (1995). Similar results were reached by Dumas (1994) who relates the international asset pricing relationship to business conditions.
of stock returns are always higher than the correlations of industrial productions. Ammer and Mei and Campbell and Mei (1993) attack the correlation issue by decomposing the innovations in stock returns into three components: news about future dividends, interest rates, and equity premiums. Our approach will be different. We look beyond financial data and tackle the real economy. We attempt to answer the question: Are the equity correlations higher than can be justified by a dynamic model of the world financial market equilibrium? And, if they are not found to be higher, we attempt to answer the alternative question: Are they about what they should be under financial market integration or are they closer to what they should be under segmentation?

Our research has a simple goal. We aim to understand the empirical observations contained in the above diagram. Specifically, for a given degree of commonality in real activity growth, our model will predict a level of correlation of market returns.

Our work is related to, but unfortunately does not encompass, those studying

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4A similar question had been raised earlier by Shiller (1989) and Beltratti and Shiller (1993) but their asset pricing model was a “present-value” model with unspecified, potentially stochastic discount rates. In the present paper, the valuation equation is derived from optimal portfolio choices.
time-varying correlation. Longin and Solnik (1995) show by means of a statistical model, how correlations change through time. Both Longin and Solnik (1995) and Erb, Harvey and Viskanta (1994) try empirically to explain, on the basis of economic variables, how correlations vary over time. By contrast, our approach offers an economic framework to understand why correlations are different across countries and industries, but the correlations are assumed to be constant over time.

2 The “dynamic single-index” model

We now describe the model that captures the evolution of the vector of national outputs. This will be the first component of our overall model. It is a purely statistical model of international business cycles. It represents a short-cut for Real Business Cycle (RBC) models, such as Backus, Kehoe and Kydland (1992), which contain (i) a statistical model for productivity shocks and (ii) an explicit representation of the households’ consumptions and work decisions and the firms’ investment and production decisions. In our framework, we postulate a pure-exchange economy in which the dynamics of output is exogenous and people consume the entire output. It is hoped that not much will be lost by this short cut, since it is generally agreed that most of the dynamics in RBC models comes from the exogenous dynamics of productivity shocks and very little comes from the endogenous capital accumulation process.

The statistical model decomposes each country’s industrial output growth into two unobserved components: the “world” business cycle which is common to all and the “country-specific” business cycle. Each of the cycles, whether common or specific, is assumed to follow an autoregressive process of order two (parsimony). We assume that the volatilities of the innovations for each cycle are constant (homoskedasticity) and that the innovations are independent of each other across cycle types. The assumption of homoskedasticity is not the best assumption since volatilities of stock returns are known to vary over the business cycle.

Throughout, \( d_t \) denotes a vector of logarithms of outputs of a number of countries. We postulate a dynamic single index model. The comovements at all leads and lags among the output variables are modeled as arising from a single common source \( c_t \), a scalar unobserved variable that portrays the world business cycle. The idiosyncratic component, \( u_t \), which is the part not arising

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5 Hodrick (1989) derives the multivariate GARCH process followed by stock prices when dividends themselves follow a multivariate GARCH process.

6 For instance, non US stock returns tend to have a higher correlation with US stock returns while the US is in a recession than while it is in an expansion. Volatility of returns is also larger while the US is in a recession. See also Perez-Quiros and Timmermann (1996) and Ang and Bekaert (1999).

from leads and lags of $c_t$, is assumed to be stationary and uncorrelated across countries. Otherwise, it follows a general autoregressive process. The statistical representation of the system is:

$$\Delta d_t = \xi \times c_t + u_t$$
$$\chi(L)c_t = \nu + \eta_t$$
$$D(L)u_t = \varepsilon_t$$

where $L$ is the lag operator, $(\varepsilon_t, \eta_t)$ are serially uncorrelated with a diagonal covariance matrix and $D(L)$ is diagonal.

We chose to model the log-growth rate instead of the level of the industrial production for one economic and one econometric reason. First, we use the log-linear pricing kernel which directly applies to log-growth rates, as will be presented in Section 3. Second, we know that macroeconomic time-series are often integrated (see Nelson and Plosser (1982)). Indeed, we could not reject the hypothesis that our industrial production series are integrated (at the 10 percent level, with Dickey and Fuller’s (1979) test). The single-index trend model would then imply that the series are cointegrated, but we could not reject the hypothesis that the 12 series are not cointegrated (at the 10 percent level, with Stock and Watson’s (1988) test). The assumptions of the model are verified, however, for the growth rate series, because they are all stationary, at the one percent level.

The above statistical model is estimated by means of a linear Kalman filter. We use for the purpose the SCOREM algorithm of Raynauld, Simonato and Sigouin (1993). The results for the countries in our sample are presented in Table 1.

Observe that the autoregressive behavior of the world business cycle is very different from the autoregressive behavior of the country-specific cycles. The world component is driven by positive coefficients which sum to 0.697 which implies no deterministic cycle but random shocks with a persistent behavior, whereas the country cycles mostly have negative coefficients implying a much more transient (strongly mean reverting) character. The persistent world cycle will play a driving role in the determination of stock returns.

3 The log-linear pricing kernel

Restoy and Weil (1996) take the lead of Campbell (1993) in log-linearizing the budget constraint of a household. They obtain an approximate pricing kernel for multiperiod securities which is based solely on consumption behavior. Their

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8We are very grateful to Jacques Raynauld who generously provided us with the GAUSS code to run the algorithm.
<table>
<thead>
<tr>
<th>Country</th>
<th>Loadings on world</th>
<th>Country-specific business cycles</th>
<th>Variance of innovations</th>
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<td></td>
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<td>Second AR coeff</td>
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</tr>
<tr>
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<td>(2.866)</td>
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Table 1: Coefficients of the single-index statistical model applied to international data. t statistics are reported in parenthesis below coefficients.
economy consists of many identical, infinitely-lived consumers who are endowed with an intertemporal, recursive utility of the Epstein-Zin (1989) or Kreps-Porteus (1978) isoelastic form. This type of utility function allows a distinction, which we find useful (see below), between two behavioral parameters: (i) a person’s relative risk aversion on the one hand and (ii) a person’s elasticity of intertemporal substitution (e.i.s.) on the other. The latter is a measure of the person’s willingness to shift her consumption over time.

Let relative risk aversion be denoted $\gamma$, elasticity of intertemporal substitution be denoted $1/\rho$ and $\beta$ denote the discount factor of utilities, all of which are assumed to be constant and equal for all individuals. Let $\Delta x_{t+1}$ denote the increment in the logarithm of the households’ consumption. Epstein and Zin (1989, 1991) and Weil (1990) have shown that the increment, $\Delta m_{t+1}$, of the logarithm of the pricing kernel between time $t$ and time $t+1$ is given by:

$$\Delta m_{t+1} = \theta \ln \beta - \rho \theta \Delta x_{t+1} + (\theta - 1) r_{W,t+1}$$

where: $\theta = \frac{1-\gamma}{1-\rho}$ and $r_{W,t+1}$ denotes the logarithmic rate of return on aggregate wealth between times $t$ and $t+1$. This pricing kernel corresponds to an asset pricing model containing two risk premia: one based on the covariance with consumption, the other based on the covariance with wealth. 9 But recall that, in our pure-exchange economy, consumption is equal to output.

Campbell (1993) and Restoy and Weil (1996) point out, however, that in this expression $\Delta x_{t+1}$ and $r_{W,t+1}$ are not independent quantities since wealth equals the present value of consumption: $\Delta x$ represents changes in output and $r_{W}$ captures changes in “discounted” future output. It is possible to derive an approximate relationship between these two quantities. Assuming that the households’ consumption is one-step-ahead lognormal and conditionally homoskedastic, Restoy and Weil present an expression for the financial market pricing kernel which does not involve the rate of return on wealth and, in fact, allows returns to be endogenous. 10

Restoy and Weil’s (1996) work can be interpreted as meaning that the increment, $\Delta m_{t+1}$, in the logarithm of the pricing kernel is given by:

$$\Delta m_{t+1} = \ln \beta - (\rho - \gamma) \frac{1-\gamma}{2} \text{var}_t [\Delta x_{t+1} + h_{t+1}] - \rho \text{Ex}_t [\Delta x_{t+1}]$$

$$- \gamma S_{t+1} [\Delta x_{t+1}] + (\rho - \gamma) S_{t+1} [h_{t+1}]$$

where, because of homoskedasticity, $\text{var}_t(x_{t+1} + h_{t+1})$ is a constant, to be determined on the basis of the stochastic process for consumption, and $\delta$ is a linearization constant (equal to one minus the exponential of the unconditional expected value of the log-ratio of consumption over wealth) arising in the log-linear approximation to the budget constraint. In addition:

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10Whereas Campbell (1993), by the same reasoning, derives an expression for the pricing kernel which does not involve consumption.
\[ h_{t+1} = E_{t+1} \left[ \sum_{j=1}^{\infty} \delta^j \Delta x_{t+j+1} \right] \]  \hspace{1cm} (4)

\[ S_{t+1} [\Delta x_{t+1}] = \Delta x_{t+1} - E_t [\Delta x_{t+1}] \]  \hspace{1cm} (5)

\[ S_{t+1} [h_{t+1}] = h_{t+1} - E_t [h_{t+1}] . \]  \hspace{1cm} (6)

\( S_{t+1} \) is the “surprise” operator.

The above pricing kernel may be used to price any security in an exchange economy (see Lucas (1978)) in which production and consumption are equal. For instance, the conditional expected value of the pricing kernel provides the one-period riskless rate of interest:

\[ r_{f,t} = -\ln \beta + (\rho - \gamma) \frac{1}{2} \text{var}_t [\Delta x_{t+1} + h_{t+1}] \]

\[ + \rho E_t [\Delta x_{t+1}] - \frac{1}{2} \text{var}_t [-\gamma \Delta x_{t+1} + (\rho - \gamma) h_{t+1}] . \]

Because of homoskedasticity, both \( \text{var}_t \) terms are time invariant. Applying the kernel to an asset which pays aggregate consumption provides a value for the aggregate stock market return. This last task has been also undertaken by Restoy and Weil (1996) who show that:

\[ r_{W,t+1} = \mu + \rho \Delta x_{t+1} + (1 - \rho) S_{t+1} [\Delta x_{t+1} + h_{t+1}] , \]  \hspace{1cm} (7)

where:

\[ \mu = -\ln \beta - \frac{(1 - \gamma)(1 - \rho)}{2} \text{var}_t [\Delta x_{t+1} + h_{t+1}] . \]  \hspace{1cm} (8)

Similarly, applying the pricing kernel to an asset which pays a dividend \( d_{i,t} \) at time \( t \), Restoy and Weil get the equilibrium rate of return on individual assets.

Stock market returns in country \( i \) are:

\[ r_{i,t+1} = \pi_i + \rho \Delta x_{t+1} + S_{t+1} [\Delta d_{i,t+1} + f_{i,t+1}] - \rho S_{t+1} [\Delta x_{t+1} + h_{i,t+1}] , \]  \hspace{1cm} (9)

where:\(^{11}\)

\[ \pi_i = -\ln \beta + (\rho - \gamma) \frac{1}{2} \text{var}_t [\Delta x_{t+1} + h_{t+1}] \]

\[ -\frac{1}{2} \text{var}_t [(\rho - \gamma)(\Delta x_{t+1} + h_{t+1}) \]

\[ + \Delta d_{i,t+1} + f_{i,t+1} - \rho (\Delta x_{t+1} + h_{i,t+1})] . \]  \hspace{1cm} (10)

\[ f_{i,t+1} = E_{t+1} \left[ \sum_{j=1}^{\infty} \delta^j \Delta d_{i,t+j+1} \right] \]  \hspace{1cm} (11)

\[ h_{i,t+1} = E_{t+1} \left[ \sum_{j=1}^{\infty} \delta^j \Delta x_{t+j+1} \right] , \]  \hspace{1cm} (12)

\(^{11}\)Recall that the \( \text{var}_t \) terms are assumed time invariant.
and $\delta_i$ is a Taylor-expansion coefficient arising from the log-linearization of the definition of a rate of return.\footnote{The following quote from Restoy and Weil (1998) explains this approximation. “Let $p_{i,t}$ denote the log of the cum-dividend price-dividend ratio of asset $i$ at date $t$, and $d_{i,t}$ the log rate of growth of the dividends paid off by asset $i$ between $t$ and $t+1$. Then, by definition, the log return on asset $i$ satisfies the identity:

$$r_{i,t+1} = d_{i,t+1} + p_{i,t+1} - \ln (e^{p_{i,t}} - 1).$$

Following Campbell and Shiller (1988), we assume that the log dividend growth process is stationary and use a Taylor expansion similar to the one applied [by Campbell (1993)] to the budget constraint to find that:

$$r_{i,t+1} \approx d_{i,t+1} + p_{i,t+1} - \frac{1}{\delta_i} p_{i,t} - k_i$$

where $k_i$ and $\delta_i$ ($0 < \delta_i < 1$) are two linearization constants.”} This constant is related to the unconditional expected value of the dividend yield of each security. In equation (9), observe the respective roles of country vs. world outputs. Out of the three random terms, two (the first and last one) stand for the current and future behavior of world output whereas only the center term refers to the future behavior of asset $i$’s specific output stream. The terms related to world output reflect the movement in the pricing kernel applicable to all assets worldwide.

A fascinating result falls out of equation (9). Whereas the financial asset conditionally expected return depends on both risk aversion $\gamma$, the e.i.s. $1/\rho$, and, of course, on the impatience parameter $\beta$, the second moments (volatilities and correlations) of the asset return depend on only one utility parameter: the elasticity of intertemporal substitution. This result would hold exactly in the case of Epstein-Zin utilities with constant risk aversion and e.i.s., and identically, independently distributed returns (see Epstein (1988)). In the case of our output model, the result holds under the log-linear approximation made by Restoy and Weil. This remarkable property is the reason why we have chosen to adopt this type of utility function. It will prove most convenient in what follows.

It may seem surprising that the second moments of rates of return only depend on the e.i.s. and not on the risk aversion. In fact, the “level” of rates of returns does change when people become more risk averse, but, in a homoskedastic world, it changes by a constant. As a consequence, the time-series “volatility” of each individual return is not affected by changes in risk aversion. The effect of risk aversion is time invariant in a homoskedastic world.\footnote{We are grateful to Philippe Weil for a helpful discussion on this point.} The elasticity of intertemporal substitution, on the other hand, governs the price response to shocks. Suppose output undergoes a positive shock; if consumers are willing to absorb this shock into consumption without further ado, there is no need to adjust asset prices. But, if their e.i.s. is low, they will have to be induced to consume the increased current output by the device of higher market prices of assets relative to current consumption.

It was asserted in the introduction that, were the kernel excessively volatile, this should translate into an equally excessive degree of correlation of world
equity returns. No doubt, this is true in the context of our model since an increase of the parameter $\rho$ simultaneously increases volatilities and correlations.

4 The log-linear pricing kernel combined with the dynamic single-index model of output

Rodriguez et al. (1996) specialize (7) to the case in which the growth rate in aggregate consumption is AR(2):

$$\left(1 - \phi_1 L - \phi_2 L^2 \right) \Delta x_{t+1} = \varepsilon_{t+1}$$

(13)

This particular autoregressive process implies that:

$$S_{t+1} (\Delta x_{t+1} + h_{t+1}) = \frac{1}{1 - \phi_1 \delta - \phi_2 \delta^2} \varepsilon_{t+1}$$

(14)

and, therefore:

$$\text{var}_t (\Delta x_{t+1} + h_{t+1}) = \left[ \frac{1}{1 - \phi_1 \delta - \phi_2 \delta^2} \right]^2 \text{var} (\varepsilon)$$

(15)

Our approach is similarly to apply the pricing kernel to the dynamic single-index model in Equation (1) and obtain the behavior of individual stock returns where stocks are defined as claims on individual output series. Since we have made the assumption of an exchange economy, aggregate consumption growth is equal to the weighted sum of output growth rates of individual countries:

$$\Delta x_t = \sum_j w_j \Delta d_{j,t}.$$  

For the dynamic single-index model, the terms of (9) can be particularized as follows:

$$S_{t+1} [\Delta d_{i,t+1} + f_{i,t+1}] = A_{i,i} \eta_{t+1} + B_{i,i} \varepsilon_{i,t+1}$$

(16)

$$S_{t+1} [\Delta x_{t+1} + h_{t+1}] = \sum_j [A_{j,i} \eta_{t+1} + B_{j,i} \varepsilon_{j,t+1}]$$

(17)

$$\text{var}_t [\Delta x_{t+1} + h_{t+1}] = \text{var} \left\{ \sum_j [A_{j,0} \eta + B_{j,0} \varepsilon_j] \right\}$$

(18)

$$A_{j,i} = \frac{1}{\sum_{s=0}^{\infty} \chi_s \delta_s^i}; \quad A_{j,0} = \frac{1}{\sum_{s=0}^{\infty} \chi_s \delta_s^0}$$

(19)

$$B_{j,i} = \frac{1}{\sum_{s=0}^{\infty} D_{j,s} \delta_s^i}; \quad B_{j,0} = \frac{1}{\sum_{s=0}^{\infty} D_{j,s} \delta_s^0}$$

(20)

---

14 The result can be extended trivially to an autoregressive process of any order.
15 This, of course, assumes constant weights for all countries or, at least, that the risk of weight fluctuations is not priced, which is probably a minor approximation.
\[ \text{var}_t \left[ (\rho - \gamma) (\Delta x_{t+1} + h_{t+1}) + \Delta d_{i,t+1} + f_{i,t+1} - \rho (\Delta x_{t+1} + h_{i,t+1}) \right] = \]

\[
\text{var} \left\{ (\rho - \gamma) \sum_{j} [A_{j,0}\eta + B_{j,0}\varepsilon_{j}] + A_{i,\eta}\eta + B_{i,\varepsilon_i} - \rho \sum_{j} [A_{j,\eta} + B_{j,\varepsilon_j}] \right\} = \]

(21)

Our next goal is to determine whether the second moments of observed stock returns can be matched with those of the theoretical model above.

5 Calibration of the model

The system of equations (1) coupled with equations (9-12, 16-21) provide a strong set of restrictions on the output and stock returns series. The unknown parameters are: those of the dynamic single-index model \( \zeta, \xi, D, \chi, \nu \), the initial value for the common cycle \( c_0 \), the variances of the residual terms \( \varepsilon \) and \( \eta \), the utility parameters \( \beta, \rho \) and \( \gamma \) (and the linearization constants \( \delta, \delta_i \), which, however, are not in principle independent entities).

While it is possible conceivably to estimate the full system of equations, the amount of computing required would be enormous and, above all, it is not clear how useful such an exercise would be. We already know that a model of an exchange economy, even if coupled with a generalized utility formulation such as the Epstein-Zin utility function, cannot account for observed mean rates of return on equity. This is the “equity premium” puzzle of Mehra and Prescott (1985).\(^{16}\) Hence, we already know that, if most countries are like the United States (but see Goetzmann and Jorion (1999)), the model would be strongly rejected because of the first moment of returns alone.

We also know from Shiller (1981) that stock return volatilities tend to be larger than can be explained by a pricing model. Furthermore, output is not dividend. In other words, leverage (both operational and financial) magnifies volatilities in a way that we cannot control for, in the absence of an explicit measurement of leverage.\(^ {17}\)

In what follows, therefore, we imagine counterfactually that the equity-premium and excess-volatility puzzles have been solved and put aside (perhaps, by the device of a very large level of risk aversion; but see Weil (1989) to for an exposition of why this is unattractive). We ignore the first and own second moments to focus exclusively on correlations. Under the simplifying assumption that dividends and output are exactly linearly related, leverage magnifies

\(^{16}\)See also Kocherlakota (1996) who considers Epstein-Zin utility functions.

\(^{17}\)We are grateful to Huntley Schaller for helping us to articulate this distinction.
volatilities but leaves correlations unchanged. These correlations, viewed as moment conditions will provide us with a calibrated version of the model. In a later section, we use the Generalized Method of Moments to test the validity of these moment conditions.

5.1 Calibration under integration

The calibration is carried out in a simple way. We have already observed that, according to equation (9), under the assumption of homoskedasticity, the second moments are all dictated by the elasticity of intertemporal substitution of the market participants. Based on the dynamic single-index statistical model, which we estimate in a first stage, we select at a second stage the degree of e.i.s. that will best match the levels of a number of correlations between stock returns and output. Once that is done, we have pinned down all the parameters of the model.\textsuperscript{18} We then calculate the equilibrium stock returns for the history of shocks which we have identified statistically and compute their correlations.

The range of reasonable values for the coefficient $\rho$ is dictated by measurements of e.i.s. ($= 1/\rho$) that have been conducted in the past. Regressing the rate of growth of aggregate consumption on changes in the rate of interest, Hall (1988) found a coefficient equal to 0.1 which is lower than most previous estimates which ranged as high as e.i.s. $= 1$.\textsuperscript{19} Epstein and Zin (1991), in their test of the CAPM implied by their preferences and applied to stock returns, found a values ranging from 0.2 to 0.8. In short, values for $\rho$ ranging from 1 to 10 seem reasonable.

Figure 2 illustrates the calibration trade-offs that we are facing. It shows three kinds of “moments” that will ultimately have to be matched. They are: the correlations of a country’s stock return with the country’s own output, the correlations of a country’s stock returns with the rest of the world output and the correlations of a country’s stock returns with the rest of the world stock returns. The figure displays the straight arithmetic average across countries of the theoretical correlations and the average level of the corresponding observed correlations that we try to match. As mentioned earlier, given that we may not observe in the real-world the exact synchronocity between stock returns and output shocks that is postulated in the model, the correlations in question are not simple correlations. They are, in fact, square roots of the $R^2$ of multiple regressions of stock returns on contemporaneous, plus eleven lagged, output growth rates.

The correlations between stock markets are really what we try to explain.\textsuperscript{18}\textsuperscript{19}

\textsuperscript{18}The parameters $\beta$ and $\gamma$ are ignored since they play no role in determining correlations. The same is true for the linearization constant $\delta$. For $\delta_i$, we use the mean dividend yields calculated over the entire sample.

\textsuperscript{19}In order to carry out this type of analysis, it would be best to examine disaggregated consumption according to social categories (borrowers vs. lenders, old vs. young, employed vs. unemployed, etc.). See Deaton (1992).
While these will ultimately be taken into account when we run a test of the model (see Section 6), let us try, at first, to calibrate the model on the basis of the relationship between stock returns and output. The figure shows that, when trying to do that, we are confronted with a dilemma. The value of $\rho$ that would best match the average within-country correlation with output is about equal to 6.9 whereas the value of $\rho$ that would best match the average correlation with world output is about 2.1. It is easy to show (and the picture illustrates) that a higher value of $\rho$ implies a higher value of the cross-country correlations of stock returns. Hence we shall obtain a lower bound of the estimate of cross-country correlations under the integration hypothesis if we choose to calibrate the model on the basis of correlations with the rest of the world output rather than on the basis of within-country correlations. Using an objective function which downweights the countries with more volatile correlation estimates and searching for the best fitting value of $\rho$, we find an optimal value equal to 2.1.

For each country, Figure 3 shows the comparison between actual and model stock market correlations resulting from the value $\rho = 2.1$. There are two factors that play a role in the derivation of model correlations. First, the world pricing kernel, which applies to all securities by construction, has been set in such a way as to match the observed correlations of output with stock returns. Secondly, in our dataset, the world business cycle (see Table 1) has been found to be fairly persistent. A component of stock returns fluctuations comes from the anticipation of discounted future dividends (see Equation (9)). If a time
series is persistent, any movement occurring today is the harbinger of a lasting movement in future realizations and produces a large immediate effect on returns. In our statistical model, this large component is common to all countries since it originates in the world business cycle.

The model correlations turn out to be of a magnitude similar to the observed ones.\textsuperscript{20} As mentioned, if the pricing kernel were excessively volatile, it would be the case, if the same kernel applies worldwide, that international stock returns are excessively correlated.\textsuperscript{21} However, no such excess correlation appears in our results.

Belgium, Canada and especially the Netherlands are the countries for which actual correlations fall well above their full-integration levels. But the Netherlands is a problematic case as the Amsterdam stock market covers an industrial base which, in fact, is a world-wide one.

To confirm this, we measured the percent of foreign sales for the companies in each of the 12 countries in 1997. Using the Worldscope universe, we

\begin{figure}[h]
\centering
\includegraphics[width=\textwidth]{figure3.png}
\caption{Stock returns correlations with rest of the world under integration.}
\end{figure}

\textsuperscript{20}Canova and de Nicolo (1995), in the context of a full-blown, calibrated model of international business cycles find much larger model correlations. The difference is due to the fact that their choice of parameters was dictated by first and own second moments.

\textsuperscript{21}In fact, if we had matched volatilities – which we do not want to do because of leverage, – the resulting value of $\rho$ would have generated values of the cross-country correlations of the order of 0.95 or more.
construct country aggregates by value-weighting these ratios by the total revenues of each firm. Belgium, Canada, and the Netherlands have the highest proportions (64.7%, 64.2% and 65.2% respectively). The same ratio averages only 40.1% for the other countries in our sample.

It is tempting to make a scale adjustment to each country’s correlation based on the amount of foreign activity. This adjustment would have to change through time. For example, in 1991 the weighted proportion of foreign sales in Canada was 47.2% and it has increased to 64.2% in 1997. But the proportion of sales is an imperfect measure because it only measures one part of earnings – the revenues. We have no information as to the extranational costs of the firms.

One alternative is to look at aggregate exports in proportion to GDP. This is a natural measure of the openness of an economy and could be transformed into a scaling factor for the correlations. This ratio ranges from 0.07 for the U.S. to 0.49 in the Netherlands (in December 1997). Unfortunately, it is not obvious how to map these ratios into a scaling factor. In the end, we choose not to apply an arbitrary scaling factor to the correlations.

The other countries that stand out are Japan, Germany and Austria because the full integration correlation is much higher than the observed correlation. But we know from previous research that Japan was segmented from the world market by regulation until 1981 at least (see Gultekin, Gultekin and Penati (1989)) and it seems reasonable that Austria may have been segmented because of its status as a quasi-Eastern block country for a good part of our sample period. We have no explanation for the German deviation.

5.2 Calibrating local-currency returns, deflated returns and excess returns

We need to check whether our results depend on the way we measure the stock returns. Table 2 compares the results of the calibration, conducted as above, depending on whether the returns are measured in dollars (as has been done so far), in local currency, in local currency but deflated by the local Consumer Price Index or in local currency in excess of the local rate of interest.

Whether returns are measured in dollars, in local currency or in local currency deflated makes very little difference to the results. When examining returns measured in local currency in excess of the local rate of interest, we have had to reduce the number of countries and reduce the length of the sample period because some data on one-month Euro-rates of interest were not available.22 The results in terms of excess returns are markedly at variance with the results in terms of other units. The fit of stock return correlations is poorer. Some of the difference is accounted for by the change in sample. An additional discrepancy arises because the output process was not refitted to the shorter sample. But most of the difference is the result of interest rate behavior. It is

22The reduced sample starts in October 1978.
Table 2: The calibration of the integration hypothesis, using various units

clear that the model does not explain actual one-month interest rates very well. While interest rates make little difference to the variance of returns, they have a clear impact on correlations.

5.3 Calibration under segmentation

With the same estimated “dynamic single-index” business cycle model as in Section 2, we now modify the log-linear pricing kernel, taking each national stock market as a stand-alone financial market. The required change in the pricing kernel is straightforward: equation (7), where \(x\) now stands for each country’s output, instead of equation (9), is used to obtain individual country stock returns. In this formulation, each country lives in autarky. The correlation in output behavior which happens to exist statistically, is the only source of common behavior in stock returns. The pricing kernel is a different one in each country although the pricing kernels of different countries do exhibit some degree of cross-correlation since outputs are cross-correlated.

The calibration trade-offs that we face in this case are displayed in Figure 4. It is immediately apparent from this picture that no value of \(\rho\) will allow us to match the actual between-country stock return correlation of 0.595; the model values for these correlations barely reach the value 4.7 when \(\rho\) is as high as 10. As far as the correlations with output are concerned, the correlations of a country’s stock returns with its own output is very large in this model. Even with a value of \(\rho = 10\), the correlation falls to 0.67 which is still far greater than
the observed correlation of 0.207.

Focusing on the only correlation that can reasonably be matched, we choose the value of the e.i.s. to get the best possible match of correlations of stock returns with world output, weighted by the reliabilities of correlation estimates. We find: \( \rho = 1.42 \). The combined result of the calibration exercise for each country is shown in Figure 5.

With the single exception of the U.S., we find that theoretical stock market correlations now fall far below actual ones.\(^{23}\) This suggest that the observed levels of international stock returns correlations are inconsistent with the hypothesis of market segmentation.

Several researchers have examined the correlations of stock returns internationally.\(^{24}\) It has been tempting to interpret the measured correlations as indications of the degree of integration of financial markets. For instance, if one finds that correlations have been rising, one is tempted to conclude that financial markets are in the process of gradual integration. This conclusion was

---

\(^{23}\) Even if the value of \( \rho \) selected under integration had been maintained, the correlations under segmentation would have been similar to what they are in Figure 5.

\(^{24}\) Among the more recent investigations, see Longin and Solnik (1995), Erb, Harvey and Viskanta (1994), and Ang and Bekaert (1999). See also Bansal and Lundblad (1999).
premature, however, for as long as one did not control for the degree of correlation of economic fundamentals. Here, we control for the correlation in the fundamentals, where “fundamentals” is taken to mean “output”, and we find that correlations are about equal to their full-integration levels and markedly above their complete-segmentation levels, given the common behavior of outputs. This gives us some reason to try and see whether we can actually reject either one of the two hypotheses.

6 A statistical test of the integration hypothesis

It is evident from the international calibration exercise of Section 5 that reality is very much at variance with the full segmentation hypothesis. Two of the three correlation categories that we have chosen to look at, present no prospect of coming reasonably close to their measured counterparts, no matter what value of the crucial parameter $\rho$ we choose. Hence, we focus in this section exclusively on the design and implementation of a test of the full integration hypothesis.\textsuperscript{26}

\textsuperscript{25}Bekaert and Harvey (1995) link correlation with the degree of market integration. Freimann (1998) offers an alternative, entirely statistical procedure based on randomization of industrial sector returns, to compare country correlations to what they would have been under integration.

\textsuperscript{26}Under the integration hypothesis, it is not easy to allow for a different value of the parameter $\rho$ for each country’s sub-population of investors. This is because the aggregate of
6.1 Test design

Define \( y_t \) as a variable which has been regressed on a set of explanatory variables and call \( \varepsilon_t \) the residuals of that regression. We construct a variable \( u_t \) in the following way. Let,

\[
\xi_t = \frac{(\varepsilon_t)^2}{\text{var}(y)}.
\]  (22)

Notice that \( \sum \xi_t \) is equal to one minus the \( R^2 \) of the regression. In our application, \( y_t \) is each country’s monthly rate of stock return and the regressors are the contemporaneous and eleven lagged values of either the country’s own output growth rate, or the rest of the world output growth rate or the rest of the world stock return, as the case may be. We calculate this variable in two versions; one denoted \( \xi_t \) is based on realized observations; the other \( \xi_t \) is calculated from model outputs and is, therefore, a function of the unknown parameter \( \rho \).

The deviations between model and reality are calculated as:

\[
u_t = \xi_t - \hat{\xi}_t.
\]  (23)

Since we have twelve countries and three categories of residuals that we try to match, we have thirty six such deviations at each point in time. We stack them in a 36-element vector which we then use to construct a weighted objective function in the manner of the Generalized Method of Moments. The moments form a vector \( g = \sum_t u_t \) and the weighting matrix is the inverse of the variance-covariance matrix of \( u \). This objective function can be minimized to obtain an estimate of the single unknown parameter \( \rho \). Asymptotically, the minimized objective function is \( \chi^2 \) distributed with 35 degrees of freedom.

We make two adjustments to that procedure. The first one aims to take into account of the possible serial dependence of the vector \( u \). The adjustment involves an optimal number of lags of the vector \( u \). It follows the method proposed by DenHaan (1996). We allow a maximum lag of fifteen months. We only perform a univariate correction: a series’ own past values only are considered in this correction for serial dependence. The lag length is chosen on the basis of the Schwarz Bayesian Information Criterion. Call \( w \) the inverse variance-covariance matrix of \( u \) after adjustment for serial correlation.

Our second adjustment takes parameter uncertainty into account. Recall that, in a first stage of our procedure, the dynamic single-index model has been estimated to model output behavior while, in a second stage of the estimation, we now estimate the preference parameter \( \rho \). This is acceptable because the output model is independent of the financial model.\(^{27}\) While the structure of the

\(^{27}\)Some improvement in efficiency could still be achieved if the two model components were estimated jointly but that is not feasible.
procedure is sound, the parameter uncertainty of the first stage must be taken into account at the second stage. This is easily achieved by first computing the Jacobian matrix $\partial g/\partial \theta$ where $\theta$ stands for all the first-stage parameters, and then adjusting the weighting matrix $w$ as follows:

$$w_{1}^{-1} = w^{-1} + [\partial g/\partial \theta] \Omega [\partial g/\partial \theta]^T,$$

where $\Omega$ is the $50 \times 50$ variance-covariance matrix of the first-stage parameter estimates.

The GMM iterates over the choice of the parameter $\rho$ and over the choice of the weighting matrix $w$. Once that is done, the matrix $w$ is replaced by the matrix $w_1$ and one more iteration series is performed over the choice of $\rho$.

### 6.2 Test results

A test conducted over the thirty six moment conditions of the twelve countries of our sample leads to the following results:

<table>
<thead>
<tr>
<th>Sample period</th>
<th>$\rho$</th>
<th>Std Err</th>
<th>$\chi^2$</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>January 1970-June 1996</td>
<td>2.8432</td>
<td>0.0709</td>
<td>102.82</td>
<td>&lt;.001</td>
</tr>
</tbody>
</table>

Hence, we reject the hypothesis of full financial market integration for all twelve countries against an unspecified alternative. It is notable that our study of correlations has produced a test powerful enough to reject the integration hypothesis whereas extant tests based on the first moment and a partial equilibrium model such as an international Capital Asset Pricing Model, for the most part, have had too little power to reject.\footnote{For an exception, see Jorion and Schwartz (1986). Note that partial-equilibrium models would not be rejected because of the equity-premium puzzle, except if they are based on consumption behavior.}

This result is subject to one caveat. It is possible that our rejection is a result of imposing the assumption that $\rho$ is the same for all countries. An extension of the test to accommodate the possibility of country-specific $\rho$s will await future research.

### 6.3 Robustness: breaking the sample and other variations

One should be careful about the meaning of this conclusion. It is very much dependent on the fact that we tried to match three types of correlations: correlations of stock returns with each country’s output growth, with the rest of the world output and with the rest of the world stock returns. The rejection of the integration hypothesis is a rejection of the adequate match of these three moments. In that respect, the rejection of integration could arguably be viewed as a misnomer. We have pointed out that the two moments involving output did
not quite agree with each other. Given the values of $\rho$ found, we could equally say that we are rejecting the ability of the model to explain the correlations of stock returns with a country’s own output. To underscore that alternative interpretation, we present below the result of a test of the integration model in which only the two moment categories, involving the rest of the world only,\(^{29}\) are matched:

<table>
<thead>
<tr>
<th></th>
<th>$\rho$</th>
<th>Std Err</th>
<th>$\chi^2$</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Two moment groups only</td>
<td>2.256</td>
<td>0.075</td>
<td>36.65</td>
<td>0.0354</td>
</tr>
</tbody>
</table>

One should not be surprised to observe that we no longer reject the hypothesis at the usual 1% level (even though we still reject it at the 5% level).

We are able to use our test to determine whether the world financial market has evolved over time towards more integration. To that aim, we break the sample into two halves. It is out of the question to re-estimate the output model over each half. But, taking output behavior as given and independent of the workings of the financial market, one can nonetheless estimate the second-stage financial component of the model over two subsamples. The results are the following:

<table>
<thead>
<tr>
<th>Subsample</th>
<th>$\rho$</th>
<th>Std Err</th>
<th>$\chi^2$</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>January 1970-Mars 1983</td>
<td>2.54</td>
<td>0.082</td>
<td>78.06</td>
<td>&lt;.001</td>
</tr>
<tr>
<td>April 1983-June 1996</td>
<td>3.84</td>
<td>0.141</td>
<td>101.15</td>
<td>&lt;.001</td>
</tr>
</tbody>
</table>

Hence, we reject the integration hypothesis for all twelve countries over both subsamples.

It is also interesting to see which country, if any, causes the full-integration hypothesis to fail. While commenting the calibration results, we had second thoughts about including the Netherlands in the sample and we recognized that Japan and Austria may have been segmented away from the world financial market. However, excluding one country at a time, we still reject the integration hypothesis at the 1% level (the complete results are available on request). One thus finds no evidence in favor of the idea that one country, being perhaps segmented financially from the rest of the world, would have caused the overall integration test to fail. The calibration exercise that we performed provides us with a reason to try and exclude one pair of countries that may both have been at some point segmented away from the world financial market, namely Japan and Austria. But excluding these two countries together also does not allow us to accept the hypothesis at the 1% level.

We checked in the calibration section that our results were not very sensitive to the way we measure stock returns. In order to confirm this intuition, we conducted a test for each convention we considered. We get the following results:

\(^{29}\)Namely, the correlation of stock returns with the rest-of-the-world stock returns and the correlation of output with the rest-of-the-world output.
<table>
<thead>
<tr>
<th>Stock Returns...</th>
<th>$\rho$</th>
<th>Std Err</th>
<th>$\chi^2$</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>...in local currency</td>
<td>2.94</td>
<td>0.081</td>
<td>94.32</td>
<td>&lt; .001</td>
</tr>
<tr>
<td>...in local currency and deflated</td>
<td>2.89</td>
<td>0.080</td>
<td>94.04</td>
<td>&lt; .001</td>
</tr>
<tr>
<td>...in excess of the risk free rate</td>
<td>3.35</td>
<td>0.092</td>
<td>94.66</td>
<td>&lt; .001</td>
</tr>
</tbody>
</table>

The conclusion remains the same: we reject the integration hypothesis at the 1% level.

7 Conclusion

In this paper, we have linked the correlations of stock returns to their fundamental determinants. These determinants were taken to be the behavior of output in the various countries. We have represented the behavior of output by means of a “dynamic single-index” statistical model, designed to capture the “covariation” of outputs in a dynamic framework, over the business cycle. The coefficients of the statistical model seem reasonable, and produce a common world cycle which is fairly persistent.

The theory of integrated stock markets which we have applied to the estimated behavior of output, has yielded levels of theoretical correlations of rates of return about equal to the measured correlations and, above all, the alternative hypothesis of financial-market segmentation hypothesis has produced correlations markedly lower than the actual ones. The likely interpretation is that the stock markets of the world are reasonably integrated.

One type of correlation, however, has not been explained satisfactorily by our model. It is the correlation of each country’s stock return with the own-country industrial production. The theoretical value is quite a bit higher than the observed one, at the value of the unknown parameter (elasticity of intertemporal substitution) that matches the other correlation moments. This is the single reason for which the full-integration model was rejected by the data.

One often hears the assertion that increased global integration implies higher global stock market correlations. This assertion is problematic because it does not control for the economic fundamentals of each country. This is exactly the motivation of our paper. Our framework allows us to give international stock market correlations an interpretation in terms of degree of integration vs. segmentation.
Appendix Table A1

Summary Statistics for 12 OECD Countries

A. Equity Market Capitalization to GDP Ratios

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>0.040</td>
<td>0.370</td>
<td>0.154</td>
<td>0.102</td>
</tr>
<tr>
<td>Belgium</td>
<td>0.060</td>
<td>0.204</td>
<td>0.159</td>
<td>0.067</td>
</tr>
<tr>
<td>Canada</td>
<td>0.130</td>
<td>0.313</td>
<td>0.206</td>
<td>0.106</td>
</tr>
<tr>
<td>France</td>
<td>0.070</td>
<td>0.174</td>
<td>0.031</td>
<td>0.041</td>
</tr>
<tr>
<td>Germany</td>
<td>0.09</td>
<td>0.051</td>
<td>0.028</td>
<td>0.051</td>
</tr>
<tr>
<td>Italy</td>
<td>0.058</td>
<td>0.138</td>
<td>0.030</td>
<td>0.041</td>
</tr>
<tr>
<td>Japan</td>
<td>0.08</td>
<td>0.047</td>
<td>0.012</td>
<td>0.026</td>
</tr>
<tr>
<td>Netherlands</td>
<td>0.038</td>
<td>0.023</td>
<td>0.012</td>
<td>0.021</td>
</tr>
<tr>
<td>Sweden</td>
<td>0.102</td>
<td>0.073</td>
<td>0.038</td>
<td>0.019</td>
</tr>
<tr>
<td>U.K.</td>
<td>0.058</td>
<td>0.053</td>
<td>0.019</td>
<td>0.013</td>
</tr>
<tr>
<td>U.S.</td>
<td>0.610</td>
<td>0.500</td>
<td>0.305</td>
<td>0.235</td>
</tr>
</tbody>
</table>

Average

| 0.263 |

B. Equity Market Capitalizations as a Ratio of MSCI Universe

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>0.001</td>
<td>0.005</td>
<td>0.011</td>
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Average

| 0.041 |

C. GDP as a Proportion of OECD GDP

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Average

| 0.041 |

Figure 6: Summary statistics for sample of countries.

8 Data appendix

The data used in this article are monthly time series covering the industrial production and the stock returns of a subset of OECD countries. The data are from two different sources, both available on DATASTREAM: OECD for industrial production series and Morgan Stanley Capital International (MSCI) for stock return data.

We selected the following twelve countries based on data availability and on the joint sizes of their economy and their stock market during the last twenty years: Austria, Belgium, Canada, France, Germany, Italy, Japan, Netherlands, Spain, Sweden, the United Kingdom and the United States. There exist no monthly output series of any kind for Switzerland and Australia.

Some summary statistics are presented in Appendix Table A1. Panel A shows the equity capitalization to GDP ratios for the twelve countries in 1971, 1980, 1990 and 1995. The proportional size of the equity market increases in every country in our sample except for Canada and Spain. Panel B shows the equity capitalization as a ratio of the MSCI world. The twelve countries in our sample encompass 81% of world market capitalization in 1995. Finally, we examine the GDP as a proportion of OECD GDP. The twelve countries we choose have 91% of OECD GDP (reported in Panel C) in 1995.
8.0.1 Industrial Production

We have used the monthly time series of real industrial production with a 1990 basis year, deseasonalized, as published for each of the twelve countries by the OECD.

The series codes of the series in DATASTREAM are: OEOCIPRDG, BGO-CIPRDG, CNOCIPRDG, FROCIPRDG, BDOCIPRDG, ITOCIPRDG, JPOCIPRDG, NLOCIPRDG, ESOCIPRDG, SDOCIPRDG, UKOCIPRDG and USOCIPRDG.

For the weighting of each country in the world aggregate economy, we have used the yearly values of Gross Domestic Product (GDP), with a 1990 basis year for the prices and for the exchange rates, as published by the OECD.

The series codes of the series in DATASTREAM are: OEGDP90, BG-GDP90, CNGDP90, FRGDP90, BDGDP90, ITGDP90, JPGDP90, NLGDP90, ESGDP90, SDGDP90, UKGDP90 and USGDP90.

8.0.2 Stock returns

We have used the monthly time series of MSCI indices, measured in U.S. dollars, with dividends re-invested, with a 1970 basis year, for the twelve countries.

For the weighting of each country in the world stock market, we have used the yearly values of Gross Domestic Product (GDP).

The average dividend yields of each country come also from MSCI. We should caution that in some countries, the dividend yield of the index is not available in the early years of the sample period. We have assumed the yield to be constant over the period with missing data.

The Consumer Price Indices (CPI) used to deflate the stock returns and the risk free rates for the excess returns also come from DATASTREAM. The interest rates are the one month Eurodollar deposit rate.
References


Indicators as Instrumental Variables,” in J. A. Frankel, ed., The Internationalization of Securities Markets, University of Chicago Press for the NBER.


