Are Correlations of Stock Returns Justi¯ed 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. We apply this idea to the stock returns of different countries. We investigate the underlying determinants of cross-country stock return correlations. First, we determine, for a given, measured degree of commonality of country outputs, what should be the degree of correlation of national stock returns. To that end, we develop a model containing a statistical model for output and an intertemporal financial market model for stock returns. We then match the correlations generated by the model with measured correlations. We find that actual correlations can be matched to 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 as 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[2] 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.[3]

[2] Austria, Belgium, Canada, France, Germany, Italy, Japan, The Netherlands, Spain, Sweden, United Kingdom and the United States of America.

[3] 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, that 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.

Another possibility that we did not pursue is to directly use dividends or earnings for the countries that we study, as in Bansal and Lundblad (2000). While there are issues with the macroeconomic data (consumption and industrial production), one encounters a different set of problems using dividends and earnings. For example, both the dividends and earnings available from Morgan Stanley Capital International are smoothed with a 12-month moving average. In addition, managers tend to smooth both dividends and earnings. While the macroeconomic data is far from ideal, we elected not to use the smoothed financial data. More fundamentally, we are interested in the relationship between the macroeconomies and the financial sector. As such, it seems logical to focus on the macroeconomic data.[4]

[4] Aside from the smoothing issue, there are two additional differences between our research and Bansal and Lundblad. First, they allow for heteroskesdasticity. However, second, they

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.[5] In subsection 5.2 below, we find 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 of stock returns, in Figure 1, 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.[6] 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 Figure 1. Specifically, for a given degree of commonality in real activity growth, our model will predict a level of correlation of market returns.

Several researchers have examined the correlations of stock returns internationally.[7] 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 premature, however, for as long as one did not control for the degree of correlation of economic fundamentals.[8] Here, we control for the correlation in the fundamentals, where "fundamentals" is taken to mean "output".

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use a static CAPM whereas we use a more appropriate intertemporal framework.

[5]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.

[6]A 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.

[7]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 (2000).

[8]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.

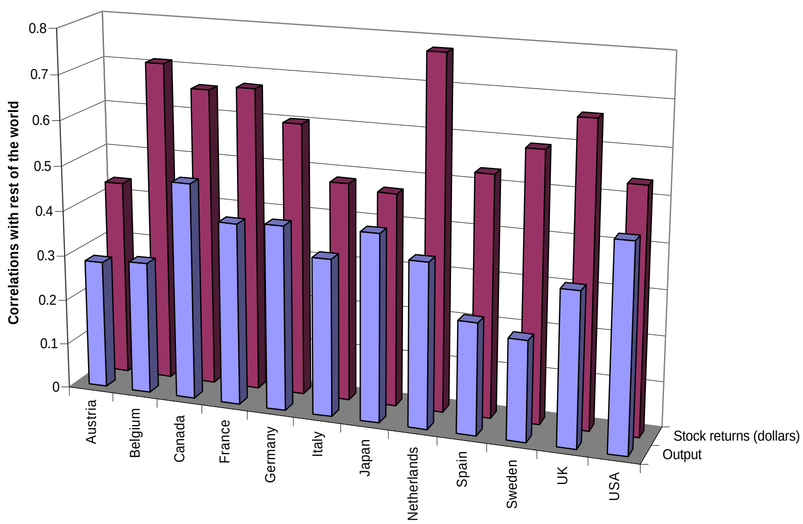

Figure 1: **Industrial production and stock returns correlations**. Correlations of each country's output growth with the rest of the world and correlation of each country's equity return with the rest of the world. Monthly data from January 1970 to June 1996.

Our work is related to, but unfortunately does not encompass, those studying time-varying correlation. Longin and Solnik (1995) show by means of a statistical model, how correlations change through time.[9] 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.[10] 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.

Indeed, the analysis in Erb *et al.* shows that while there is some time-variation in the correlations of the G7 countries' equity returns through time, the ranking of the correlations rarely changes. That is, while there is variation in both the U.S.-U.K. and U.S.-Japan correlations through time, the U.S.-U.K. correlation is always higher than the U.S.-Japan correlation. Similarly, the U.S.-Canada correlation is always higher than the U.S.-U.K. correlation. While it is clear that correlations are not constant, our assumption should not interfere with the main point of our paper, to explain why correlations are different across different countries.

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. For reasons of parsimony, each of the cycles, whether common or specific, is assumed to follow an autoregressive process of order two. A model with lag order three was also estimated but the additional coefficients were not significantly different from zero. We assume that

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

[10] 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).

the volatilities of the innovations for each cycle are constant (homoskedasticity) and that the innovations are independent of each other across cycle types.

Throughout, $\mathbf{d}_t$ denotes a vector of logarithms of outputs of a number of countries. We postulate a dynamic single index model.[11] 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, $\mathbf{u}_t$, which is the part not arising 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:

$$\begin{aligned} \Delta\mathbf{d}_t &= \boldsymbol{\xi}\times c_t + \mathbf{u}_t \\ \chi(L)c_t &= \nu + \eta_t \\ \mathbf{D}(L)\mathbf{u}_t &= \boldsymbol{\varepsilon}_t \end{aligned} \tag{1}$$

where L is the lag operator, $(\boldsymbol{\varepsilon}_t, \eta_t)$ are serially uncorrelated with a diagonal covariance matrix and $\mathbf{D}(L)$ is diagonal.

The model is formulated in terms of log-growth rates. In terms of levels, the output series of all twelve countries would have been integrated (this is confirmed at the 10 percent level by a Dickey-Fuller (1979) test). Had it been written in terms of levels, the model would say that all twelve series are cointegrated with one common trend. This is a testable proposition. The null hypothesis in such a test (see Stock and Watson (1988)), is that the twelve series are not cointegrated. We performed the test, with ambiguous results; at the 10 percent level, the hypothesis could not be rejected.[12] Had it been rejected, the estimation of the common trend could have been done in level form by cointegration methods; but such was not the case.

The log-growth rates, on the other hand, are all stationary at the one percent level, a feature which validates the estimation method we use. One added advantage of this formulation is that the log-linear pricing kernel to be used below (Section 3) directly applies to log-growth rates.

The statistical model (1), formulated in terms of log-growth rates, is estimated by means of a linear Kalman filter. We use for the purpose the SCOREM algorithm of Raynauld, Simonato and Sigouin (1993).[13] The results for the countries in our sample are presented in Table 1.

Two observations are in order. First, practically all parameter estimates are significantly different from zero. Second, the autoregressive behavior of the

[11] See Sargent and Sims (1977), Geweke (1977) and Singleton (1980). This model is discussed at length in Stock and Watson (1989, 1991, 1993).

[12] Kasa (1992) presents evidence of a single stochastic trend in GNP in a sample of five our the markets that we study.

[13] We are very grateful to Jacques Raynauld who generously provided us with the GAUSS code to run the algorithm.

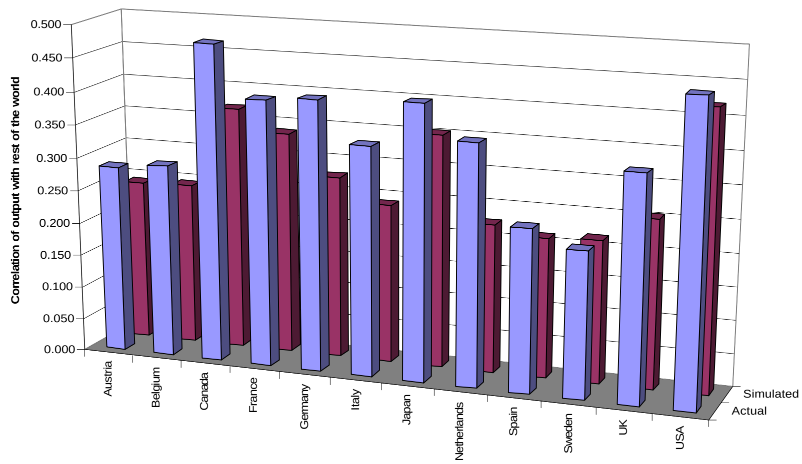

Figure 2: **Actual and model simulated correlations of country output with the rest of the world**.

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 0.705 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.

As a measure of the success, or descriptive quality, of the output model, we present Figure 2 which compares the correlations of each country's output with the rest of the world to correlations obtained by a simulation performed under the assumptions of the model. The model does a good job of matching these correlations. For a number of countries, the simulated correlations are a bit lower than the actual ones. But the reader should keep in mind that, in subsequent analysis, we use the actual innovations rather than the simulated ones. This provides an even closer match because, even after fitting the model optimally, the actual innovations contain some residual correlations across countries.[14]

[14]In fact, the model is more important as a tool to represent the persistence of the various cycles, especially the common one which will play a crucial role in asset pricing, than it is as a tool to capture contemporaneous correlations.

	Loadings	Country-specific business cycles		
Country	**on world**	**First**	**Second**	**Variance of**
	business cycle	**AR coeff**	**AR coeff**	**innovations**
Austria	0.158	-0.582	-0.225	0.696
	(5.303)	(-10.309)	(-4.020)	(12.232)
Belgium	0.163	-0.554	-0.289	0.712
	(5.466)	(-9.968)	(-5.240)	(12.209)
Canada	0.343	-0.122	0.069	0.790
	(6.654)	(-1.993)	(1.150)	(11.459)
France	0.298	-0.494	-0.233	0.657
	(8.072)	(-8.017)	(-3.843)	(11.321)
Germany	0.217	-0.553	-0.312	0.678
	(6.917)	(-9.838)	(-5.574)	(11.907)
Italy	0.151	-0.586	-0.273	0.701
	(5.202)	(-10.488)	(-4.932)	(12.256)
Japan	0.339	-0.463	-0.139	0.687
	(8.260)	(-7.296)	(-2.188)	(11.190)
Netherlands	0.127	-0.675	-0.309	0.643
	(4.911)	(-12.161)	(-5.625)	(12.300)
Spain	0.098	-0.804	-0.396	0.543
	(4.640)	(-15.237)	(-7.541)	(12.347)
Sweden	0.101	-0.410	-0.147	0.841
	(2.934)	(-7.343)	(-2.625)	(12.457)
UK	0.176	-0.221	-0.103	0.894
	(4.178)	(-3.850)	(-1.805)	(12.303)
USA	0.407	0.378	-0.017	0.618
	(6.629)	(5.167)	(-0.247)	(9.421)
		World business cycle		
World		0.454	0.251	1.000
		(5.000)	(2.866)	

Table 1: **Coefficients of the single-index statistical model applied to international data**. t statistics are reported in parenthesis below coefficients.

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 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 γ, elasticity of intertemporal substitution be denoted $1/\rho$ and β denote the discount factor of utilities, all of which are assumed to be constant and equal for all individuals. Let Δx_{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, Δm_{t+1}, of the logarithm of the pricing kernel between time t and time $t+1$ is given by:

$$\Delta m_{t+1} = \theta \ln \beta - \rho\theta \Delta x_{t+1} + (\theta - 1) r_{W,t+1} \tag{2}$$

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.[15] 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 Δx_{t+1} and $r_{W,t+1}$ are not independent quantities since wealth equals the present value of consumption: Δx 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.[16]

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

$$\begin{aligned} \Delta m_{t+1} = \ln \beta - (\rho - \gamma)\frac{1-\gamma}{2} var_t\left[\Delta x_{t+1} + h_{t+1}\right] - \rho E_t\left[\Delta x_{t+1}\right] \\ - \gamma S_{t+1}\left[\Delta x_{t+1}\right] + (\rho - \gamma) S_{t+1}\left[h_{t+1}\right] \end{aligned} \tag{3}$$

[15]See Epstein and Zin (1989, 1991) and Giovannini and Weil (1989).

[16]Whereas Campbell (1993), by the same reasoning, derives an expression for the pricing kernel which does not involve consumption.

where, because of homoskedasticity, the conditional variance $var_t(x_{t+1} + h_{t+1})$ is a constant, to be determined on the basis of the stochastic process for consumption, and δ 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:

$$h_{t+1} = E_{t+1}\left[\sum\nolimits_{j=1}^{\infty} \delta^j \Delta x_{t+j+1}\right] \tag{4}$$

$$S_{t+1}\left[\Delta x_{t+1}\right] = \Delta x_{t+1} - E_t\left[\Delta x_{t+1}\right] \tag{5}$$

$$S_{t+1}\left[h_{t+1}\right] = h_{t+1} - E_t\left[h_{t+1}\right]. \tag{6}$$

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:

$$\begin{aligned} r_{f,t} = &-\ln\beta + (\rho - \gamma)\frac{1-\gamma}{2} var_t\left[\Delta x_{t+1} + h_{t+1}\right] \\ &+ \rho E_t\left[\Delta x_{t+1}\right] - \frac{1}{2} var_t\left[-\gamma\Delta x_{t+1} + (\rho - \gamma)\, h_{t+1}\right]. \end{aligned}$$

Because of homoskedasticity, both 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:

$$r_{W,t+1} = \mu + \rho\Delta x_{t+1} + (1-\rho)\, S_{t+1}\left[\Delta x_{t+1} + h_{t+1}\right], \tag{7}$$

where:

$$\mu = -\ln\beta - \frac{(1-\gamma)(1-\rho)}{2} var_t\left[\Delta x_{t+1} + h_{t+1}\right]. \tag{8}$$

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:

$$r_{i,t+1} = \pi_i + \rho\Delta x_{t+1} + S_{t+1}\left[\Delta d_{i,t+1} + f_{i,t+1}\right] - \rho S_{t+1}\left[\Delta x_{t+1} + h_{i,t+1}\right], \tag{9}$$

where:[17]

$$\pi_i = -\ln\beta + (\rho-\gamma)\frac{1-\gamma}{2}var_t\left[\Delta x_{t+1} + h_{t+1}\right]$$
$$-\frac{1}{2}var_t\left[(\rho-\gamma)\left(\Delta x_{t+1} + h_{t+1}\right)\right.$$
$$\left. + \Delta d_{i,t+1} + f_{i,t+1} - \rho\left(\Delta x_{t+1} + h_{i,t+1}\right)\right] \tag{10}$$
$$f_{i,t+1} = E_{t+1}\left[\sum\nolimits_{j=1}^{\infty}\delta_i^j \Delta d_{i,t+j+1}\right] \tag{11}$$
$$h_{i,t+1} = E_{t+1}\left[\sum\nolimits_{j=1}^{\infty}\delta_i^j \Delta x_{t+j+1}\right], \tag{12}$$

and δ_i is a Taylor-expansion coefficient arising from the log-linearization of the definition of a rate of return.[18] 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 γ, the e.i.s. $1/\rho$, and, of course, on the impatience parameter β, 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

[17] Recall that the var_t terms are assumed time invariant.

[18] 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:

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

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 δ_i $(0 < \delta_i < 1)$ are two linearization constants."

"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.[19] 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 ρ 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):

$$\left(1 - \phi_1 L - \phi_2 L^2\right) \Delta x_{t+1} = \varepsilon_{t+1} \tag{13}$$

This particular autoregressive process implies that:[20]

$$S_{t+1}\left(\Delta x_{t+1} + h_{t+1}\right) = \frac{1}{1 - \phi_1 \delta - \phi_2 \delta^2} \varepsilon_{t+1} \tag{14}$$

and, therefore:

$$var_t(\Delta x_{t+1} + h_{t+1}) = \left[\frac{1}{1 - \phi_1 \delta - \phi_2 \delta^2}\right]^2 var\left(\varepsilon\right) \tag{15}$$

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:[21] $\Delta x_t = \sum_j w_j \Delta d_{j,t}$.

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

[19]We are grateful to Philippe Weil for a helpful discussion on this point.

[20]The result can be extended trivially to an autoregressive process of any order.

[21]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.

as follows:

$$S_{t+1}\left[\Delta d_{i,t+1}+f_{i,t+1}\right]=A_{i,i}\eta_{t+1}+B_{i,i}\varepsilon_{i,t+1} \tag{16}$$

$$S_{t+1}\left[\Delta x_{t+1}+h_{i,t+1}\right]=\sum_{j}\left[A_{j,i}\eta_{t+1}+B_{j,i}\varepsilon_{j,t+1}\right] \tag{17}$$

$$var_t\left[\Delta x_{t+1}+h_{t+1}\right]=var\left\{\sum_{j}\left[A_{j,0}\eta+B_{j,0}\varepsilon_j\right]\right\} \tag{18}$$

$$A_{j,i}=\xi_j\frac{1}{\sum_{s=0}^{\infty}\chi_s\delta_i^s};\quad A_{j,0}=\xi_j\frac{1}{\sum_{s=0}^{\infty}\chi_s\delta^s} \tag{19}$$

$$B_{j,i}=\frac{1}{\sum_{s=0}^{\infty}D_{j,s}\delta_i^s};\quad B_{j,0}=\frac{1}{\sum_{s=0}^{\infty}D_{j,s}\delta^s} \tag{20}$$

$$var_t\left[(\rho-\gamma)\left(\Delta x_{t+1}+h_{t+1}\right)+\Delta d_{i,t+1}+f_{i,t+1}-\rho\left(\Delta x_{t+1}+h_{i,t+1}\right)\right]=$$

$$var\left\{(\rho-\gamma)\sum_{j}\left[A_{j,0}\eta+B_{j,0}\varepsilon_j\right]+A_{i,i}\eta+B_{i,i}\varepsilon_i\ -\rho\sum_{j}\left[A_{j,i}\eta+B_{j,i}\varepsilon_j\right]\right\} \tag{21}$$

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), provides a strong set of restrictions on the output and stock returns series. The unknown parameters are: those of the dynamic single-index model $\boldsymbol{\zeta}$, $\boldsymbol{\xi}$, $\mathbf{D}$, χ, ν, the initial value for the common cycle c_0, the variances of the residual terms $\boldsymbol{\varepsilon}$ and η, the utility parameters β, ρ and γ (and the linearization constants δ, δ_i, which, however, are not in principle independent entities).

While it is possible (but extremely difficult) to estimate the full system of equations, 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, has difficulties accounting for observed mean rates of return on equity in the United States. This is the "equity premium" puzzle of Mehra and Prescott (1985). Goetzmann and Jorion (1997) have pointed out that most countries are not like the United States and have argued that U.S. equity rates of return presumably represent a repeated sequence of surprises, not to be confused with high expected returns.

Like Kocherlakota (1996) and Campbell and Koo (1997), we use Epstein-Zin utility functions. Previous research shows that the freedom to separately choose the risk aversion and the elasticity of intertemporal substitution of the representative individual allows a somewhat better fit of first moment (equity returns and interest rate) than standard time additive utilities. Nonetheless, it is clear that, at generally accepted levels of risk aversions, our model would not account for the U.S. equity premium.

In what follows, therefore, we employ the result obtained above (see Equation (9)) that, to an approximation, the elasticity of intertemporal substitution alone determines second moments while, once second moments have been determined, the two taste parameters jointly determine the first moments. We rely on this argument to separate out the debate on the equity premium and we focus exclusively on second moments.

Further, we know from Shiller (1981) that stock return volatilities tend to be larger in reality than can be explained by a pricing model based on dividends. In our model, however, dividends, or even earnings, are not the basis for the determination of stock returns. Output is. In the real world, two layers of leverage (operational and financial) should normally tend to magnify the volatility of dividend growth rates relative to that of output growth rates. In our dataset, the average taken across countries of the volatilities of industrial production growth rates is equal to 2.02%/month whereas the average volatility of dividend growth rates is equal to 4.7%/month and the average volatility of earnings is equal to 8.4%/month.[22] Under the simplifying assumption that dividends and output are exactly linearly related, leverage magnifies volatilities but leaves correlations unchanged.[23,24]

Below, we report the correlations and the volatilities produced by our model, knowing, however, that actual volatilities are magnified by leverage, which is unobserved. For this reason, we focus mostly on correlations. The correlations, viewed as moment conditions, provide us with a calibrated version of the model. In a later section (Section 6), we use the Generalized Method of Moments to test the validity of the 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

[22] The average of earnings growth rates given here is based on eight countries only. In the US, these numbers are 0.8%, 2.4% and 3.6% respectively. In all cases, earnings and dividends are measured as twelve-month moving averages.

[23] We are grateful to Huntley Schaller for helping us to articulate this distinction.

[24] We admit readily that heteroskedasticity of output would produce time varying expected returns which would contribute, to a small extent, to an explanation of the high volatility of stock returns.

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.[25] We then calculate the equilibrium stock returns for the history of shocks which we have identified statistically and compute their correlations.

To be clear, we are executing calibrations (and, later, statistical tests) using unconditional moment conditions. This choice is unrelated to the assumptions of the model, such as the homoskedasticity assumption, since any model, whether homoskedastic or heteroskedastic can be tested on the basis of its unconditional predictions. The choice of moment conditions is also not a limitation placed on the rational-choice paradigm. In our model, the representative investor dynamically optimizes his/her portfolio on the basis of all the information at his/her disposal. The first-order condition of portfolio choice being implemented by the investor would imply a full conditional, intertemporal CAPM. But that is not the restriction which is tested here.

The range of reasonable values for the coefficient ρ 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 an elasticity equal to 0.1 which is lower than most previous estimates which ranged as high as e.i.s. = 1. Epstein and Zin (1991), in their test of the CAPM implied by their preferences and applied to stock returns, found values for the elasticity ranging from 0.2 to 0.8. In short, values for ρ ranging from 1 to 10 seem reasonable.[26]

Figure 3 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 synchroneity between stock returns and output shocks that is postulated in the model, the correlations in question are not simple correlations. They are actually square roots of the R^2 of multiple regressions of stock returns on contemporaneous, plus eleven lagged, output growth rates.

[25] The parameters β and γ are ignored since they play no role in determining correlations. The same is true for the linearization constant δ. For δ_i, we use the mean dividend yields calculated over the entire sample.

[26] 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). Attanasio and Weber (1989), placing themselves explicitly in the Epstein-Zin paradigm, and using a single cohort of household found a value $\rho = 0.514$ or an e.i.s. approximately equal to 2.

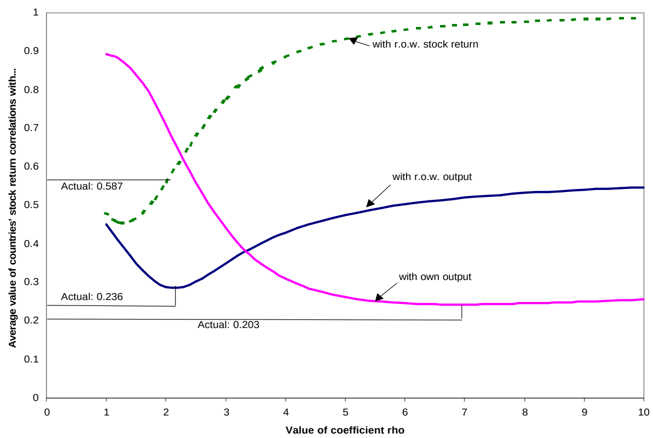

Figure 3: **The calibration trade-off under integration.** For each of the correlations, the level marked "actual" is the target (observed) correlation which is to be matched.

Let us try to calibrate the model on the basis of the relationship between stock returns and output and then see whether the calibrated model would account for what we are really trying to explain, viz. the correlations between stock markets. The figure shows that, when trying to do that, we are confronted with a dilemma. The value of ρ that would best match the average within-country correlation with output is about equal to 6.9 whereas the value of ρ 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 ρ implies a higher value of the cross-country correlations of stock returns. Hence we shall obtain a lower bound of the theoretical value 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 downweighs the countries with more volatile correlation estimates and searching for the best fitting value of ρ, we find an optimal value equal to 2.1.

For each country, Figure 4 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 as we just explained. Secondly, in our dataset, the world business cycle (see Table 1) has been found to be fairly persistent. A component of stock returns fluctuations

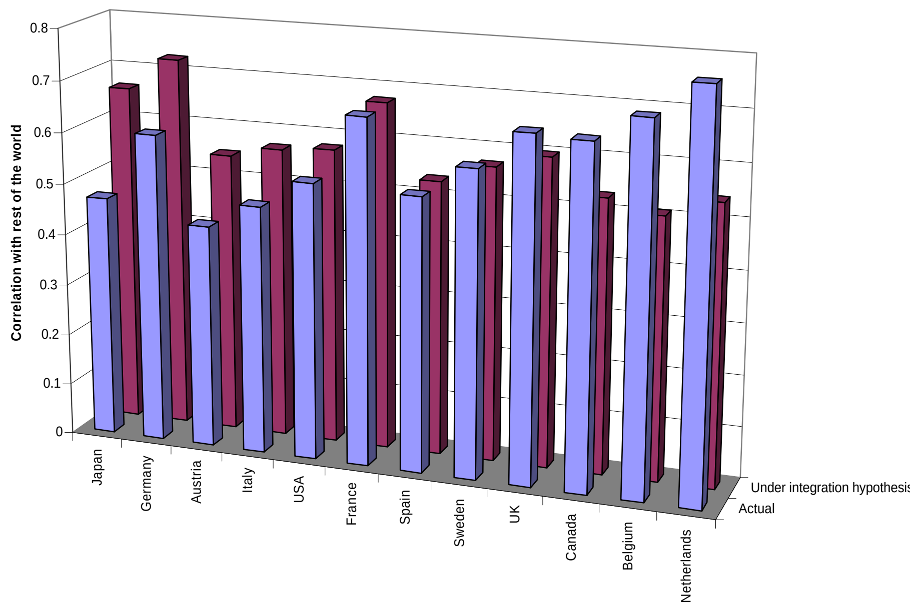

Figure 4: **Stock returns correlations with rest of the world under integration**.

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.[27] 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. However, no such excess correlation appears in our results.

5.2 Discussion

If the above calibration errs, the direction in which it errs is perfectly clear; the correlations observed in the data can only be viewed as being lower than (or equal to) those of the model, for two reasons. Remember, first, that the model correlations that we have calculated are lower bounds. By choosing a relatively low value of ρ, we can only have *understated the model correlations*. If, instead

[27]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.

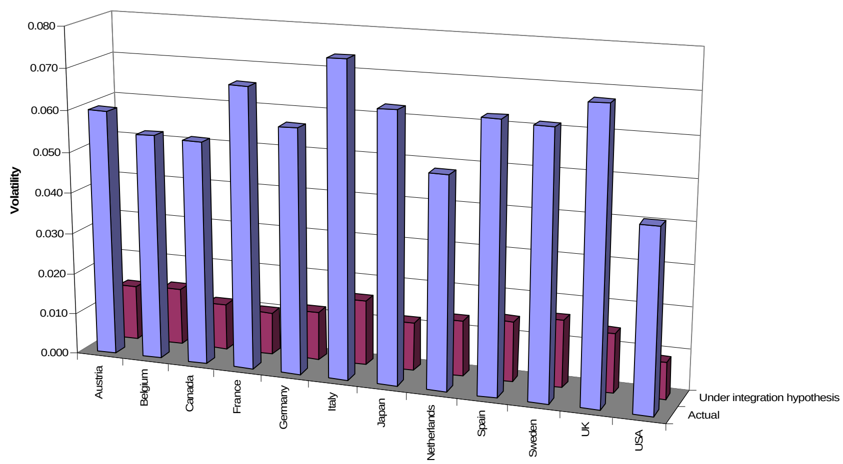

Figure 5: **Actual and model stock return volatilities**.

of what we did, we had attempted to match a country's stock return correlation with its own output, we would have picked a value of ρ equal approximately to 7 (see Figure 4). For that value of ρ, the model would have given us an average cross-country correlation of stock returns equal to: 0.97.

What if we had tried instead to match volatilities? It is evident from Figure 5 that the model, at the chosen value $\rho = 2.1$, explains only about 25 percent of the actual volatilities (which goes to show again that excess volatility does not in actuality imply excess correlation). Based on the leverage argument that we have put forward, we do not intend to take the measured volatilities literally. The figures we gave for the relative volatilities of output *vs.* dividend or earnings growth rates are sufficient to explain away the discrepancy in stock return volatilities. Had we tried to match the observed volatilities, however, the needed value of ρ would have been slightly above 9, a value at which the model would indicate an average cross-country correlation of stock returns equal to 0.98.

The second reason why our calibration can only err in the direction we have stated is that the *correlations observed in the data are, if anything overstated.* This is because, in several countries, many of the companies listed in the stock exchange typically have levels of foreign activities markedly larger than the share of exports in the corresponding output series. When an overstated share of profits originates abroad, one would surmise, the correlation of the market value of these profits relative to the rest of the world is also overstated. This conjecture is buttressed by the fact that Netherlands is a clear outlier in Figure 4; the Amsterdam stock market covers an industrial base which, in fact, is a world-wide one. Besides the Netherlands, Belgium and Canada are the countries for which actual correlations fall above their full-integration levels. To confirm this interpretation, we measured the percent of foreign sales for the companies in each of the 12 countries in 1997. Using the Worldscope universe, we constructed 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.[28]

Other countries that stand out are Japan, Germany and Austria because their 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.

Except for the cases of Japan, Germany and Austria, it seems reasonable to conclude that there is no evidence of excess correlation in the data. If anything, the cross-correlations may be lower than what they should be under full integration,[29] this being especially true if we insist on trying to match the correlation between a country's stock return and its own output (see Section 6 below).

5.3 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

[28] It is tempting to make a scale adjustment to each country's correlation based on the level of foreign activity, to bring that level down to equal the share of exports in GDP (in December 1997, this ratio ranges from 7% for the U.S. to 49% in the Netherlands). 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. We choose not to apply a scaling factor of that type. First, we felt that imposing the scaling factors based on these measures would be arbitrary as we do not observe for each country the composition of foreign trade by destination. Second, we were worried about introducing another level of estimation error.

[29] Within an integrated country such as France, the correlation of stock returns across industrial sectors is equal to 0.818 whereas the cross-country correlation, as we saw, is equal to 0.587. Freimann (1998), as mentioned, 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. He finds that cross-correlations are lower than they should be under full integration.

	US$		US$ excess		Local		Local		Local excess	
	return		**return**		**return**		**real return**		**return**	
ρ	2.10		1.67		2.10		2.08		1.77	
Country	Actual	Fit	Actual	Fit	Actual	Fit	Actual	Fit	Actual	Fit
Austria	0.435	0.547	0.470	0.197	0.358	0.545	0.364	0.537		
Belgium	0.710	0.508	0.723	0.183	0.668	0.506	0.672	0.499	0.683	0.192
Canada	0.660	0.531	0.649	0.263	0.702	0.529	0.696	0.524	0.705	0.321
France	0.671	0.675	0.696	0.253	0.645	0.674	0.646	0.670	0.685	0.258
Germany	0.603	0.726	0.615	0.388	0.575	0.725	0.577	0.721	0.602	0.362
Italy	0.483	0.568	0.506	0.451	0.452	0.567	0.458	0.565	0.470	0.433
Japan	0.470	0.663	0.491	0.349	0.456	0.661	0.462	0.657	0.474	0.333
Netherlands	0.778	0.542	0.789	0.307	0.743	0.541	0.741	0.534	0.772	0.328
Spain	0.531	0.534	0.532	0.366	0.529	0.533	0.535	0.530		
Sweden	0.593	0.571	0.600	0.256	0.553	0.568	0.557	0.559		
UK	0.666	0.598	0.684	0.227	0.662	0.596	0.662	0.589	0.739	0.166
USA	0.537	0.576	0.551	0.657	0.619	0.576	0.627	0.574	0.649	0.705

Table 2: **The calibration of the integration hypothesis, using various units**

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.

The first set of returns in Table 2 shows that, whether returns are measured in dollars, in local currency or in local currency deflated makes very little difference to the results (columns 1, 3 and 4). 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.[30] The results in terms of excess returns (columns 2 and 5) are markedly in contrast 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 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.

5.4 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

[30] The reduced sample starts in October 1978.

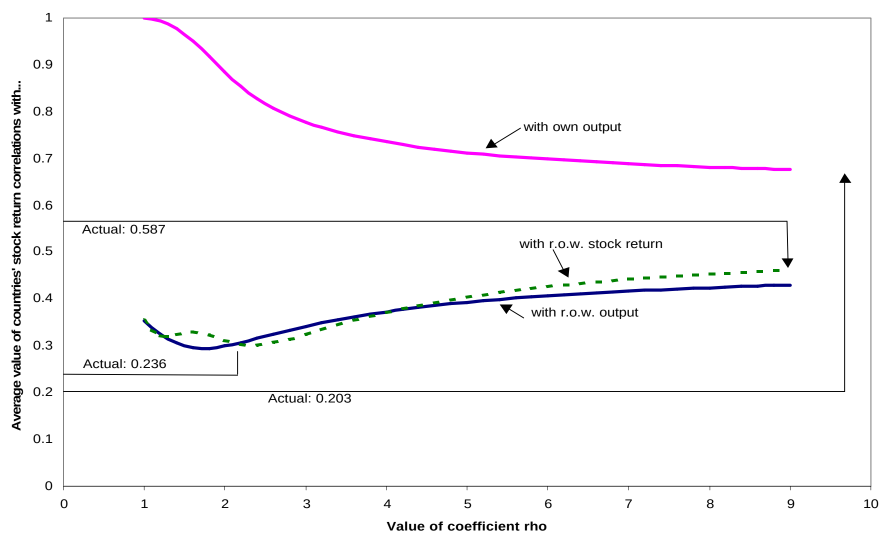

Figure 6: **The calibration trade-off under segmentation.** For each of the correlations, the level marked "actual" is the observed correlation which is to be matched.

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 6. It is immediately apparent from this picture that no value of ρ 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 ρ 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

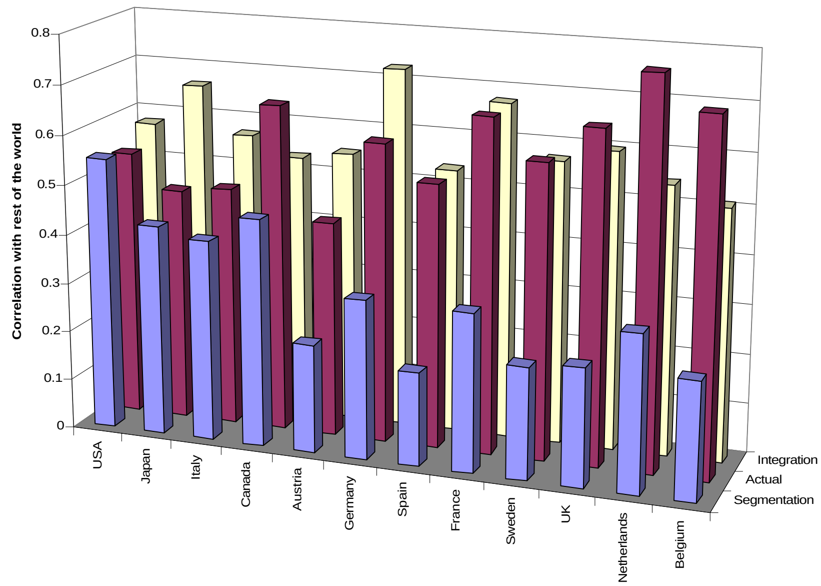

Figure 7: **Stock market correlations with the rest of the world under two hypotheses: integration and segmentation**.

is shown in Figure 7.

With the single exception of the U.S., we find that theoretical stock market correlations now fall far below actual ones.[31] This suggests that the observed levels of international stock returns correlations are inconsistent with the hypothesis of market segmentation.

In short, we find that correlations are about equal to (or lower than) 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 odds 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 ρ we choose. Hence, we focus in this section exclusively

[31]Even if the value of ρ selected under integration had been maintained, the correlations under segmentation would have been similar to what they are in Figure 7.

on the design and implementation of a test of the full-integration hypothesis.[32]

6.1 Test design

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

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

Notice that $\sum_t \xi_t$ 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 ξ_t is based on realized observations; the other $\widehat{\xi_t}$ is calculated from model outputs and is, therefore, a function of the unknown parameter ρ.

The deviations between model and reality are calculated as:

$$u_t = \xi_t - \widehat{\xi_t}. \tag{23}$$

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 ρ. Asymptotically, the minimized objective function is χ^2 distributed with 35 degrees of freedom.

We make two adjustments to that procedure. The first one aims to take 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

[32]Under the integration hypothesis, it is not easy to allow for a different value of the parameter ρ for each country's sub-population of investors. This is because the aggregate of a world population of investors with recursive utility is not a representative investor with recursive utility (see Dumas, Uppal and Wang (1999)). Under the segmentation hypothesis, it would have been possible, of course, to allow such a difference from country to country.

estimated to model output behavior while, in a second stage of the estimation, we now estimate the preference parameter ρ. This is acceptable because the output model is independent of the financial model.[33] While the structure of the 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 θ stands for all the first-stage parameters, and then adjusting the weighting matrix w as follows:

$$w_1^{-1} = w^{-1} + [\partial g/\partial\theta]\text{-}\ [\partial g/\partial\theta]', \tag{24}$$

where - is the 50×50 variance-covariance matrix of the first-stage parameter estimates.

The GMM iterates over the choice of the parameter ρ 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 ρ.

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:

Sample period	ρ	Std Err	χ^2	P-value
January 1970-June 1996	2.8432	0.0709	102.82	$< .001$

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.[34]

This result is subject to one caveat. It is possible that our rejection is a result of imposing the assumption that ρ is the same for all countries. An extension of the test to accommodate the possibility of country-specific ρ_is 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

[33]Some improvement in efficiency could still be achieved if the two model components were estimated jointly but that is not feasible.

[34]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.

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 ρ 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,[35] are matched:

	ρ	Std Err	χ^2	P-value
Two moment groups only	2.256	0.075	36.65	0.0354

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:

Subsample	ρ	Std Err	χ^2	P-value
January 1970-Mars 1983	2.54	0.082	78.06	$< .001$
April 1983-June 1996	3.84	0.141	101.15	$< .001$

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.

[35]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.

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:

Stock Returns...	ρ	Std Err	χ^2	P-value
...in local currency	2.94	0.081	94.32	$< .001$
...in local currency and deflated	2.89	0.080	94.04	$< .001$
...in excess of the risk free rate	3.35	0.092	94.66	$< .001$

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.

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Appendix Table 1

Summary Statistics for 12 OECD Countries

	Austria	Belgium	Canada	France	Germany	Italy	Japan	Netherlands	Spain	Sweden	U.K.	U.S.	Average
A. Equity Market Capitalization to GDP Ratios													
1971	NA	NA	0.367	0.071	NA	0.079	0.057	NA	0.257	NA	0.610	0.398	0.263
1980	0.013	0.060	0.238	0.051	0.050	0.039	0.119	0.104	0.047	0.056	0.273	0.225	0.106
1990	0.096	0.188	0.240	0.139	0.129	0.067	0.513	0.298	0.108	0.204	0.534	0.305	0.235
1995	0.110	0.255	0.292	0.233	0.176	0.078	0.631	0.583	0.137	0.367	0.717	0.520	0.342
B. Equity Market Capitalizations as a Ratio of MSCI Universe													
													Sum
1971	0.001	0.006	0.045	0.020	0.028	0.009	0.055	0.014	0.013	0.006	0.107	0.664	0.969
1980	0.001	0.005	0.046	0.023	0.037	0.011	0.166	0.015	0.006	0.004	0.092	0.500	0.907
1990	0.003	0.007	0.026	0.032	0.039	0.014	0.295	0.016	0.010	0.009	0.102	0.331	0.887
1995	0.002	0.006	0.020	0.032	0.035	0.011	0.211	0.021	0.009	0.010	0.087	0.369	0.814
C. GDP as a Proportion of OECD GDP													
													Sum
1970	NA	NA	0.034	0.092	NA	0.081	0.198	NA	0.035	NA	0.085	0.416	0.940
1980	0.009	0.012	0.030	0.071	0.089	0.064	0.154	0.017	0.027	0.014	0.056	0.310	0.853
1990	0.009	0.012	0.034	0.071	0.097	0.065	0.175	0.017	0.029	0.014	0.058	0.328	0.907
1997	0.009	0.011	0.033	0.070	0.095	0.061	0.172	0.017	0.029	0.013	0.057	0.338	0.906

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, BGOCIPRDG, 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, BGGDP90, 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.