International Financial Integration

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INTERNATIONAL FINANCIAL INTEGRATION

by

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INTERNATIONAL FINANCIAL INTEGRATION

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During the last two decades, the degree of international financial integration (IFI) has increased substantially. This increased level of IFI has a number of benefits. First, it can lead to more efficient allocation of saving and investment across countries and, therefore, facilitate consumption smoothing. Second, it can enable domestic investors to achieve a higher level of diversification. Third, the industrial sector can benefit from having better access to the world's capital supply and, eventually an increase in the level of IFI will have a positive impact on countries' output growth. In all, higher levels of IFI can lead to a more efficient economy and ultimately to a higher level of economic well-being.

This dissertation consists of three essays, each presenting different approaches for measuring the degree of IFI across countries. Different from the vast empirical literature we focus mainly on analyzing the behavior across international bond market returns rather than the behavior across international stock market returns. In the first essay, we study the dependence structure among international bond returns by focusing on two common approaches: (1) rolling correlations, and (2) cointegration analysis. How the observed increase in the dependence structure across the international bond market is reflected in business cycles is also analyzed. In the second essay, the level of IFI is investigated by testing for the presence of a common volatility process across

international bond market returns. In the third essay, a dynamic measure of IFI is obtained. The saving-investment relationship and cross-country correlations also are used in this essay to determine the degree of capital mobility across countries. In this way, we are able to compare the benefits of using the new time-varying measure of IFI to these two traditional measures.

In conclusion, we find an increase in the level of IFI especially across major EU country members beginning in the mid-1990s. However, during the last three to four years there is some evidence that this integration may be trending downward. The results with regard to the U.K suggest that the U.K. is not highly integrated with the European countries, or with the rest of the world.

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CHAPTER 1

INTRODUCTION

1.1 International Financial Integration and its Traditional Measures

During the last two decades the degree of international financial integration (IFI) has increased substantially. This increased level of integration between countries has a number of benefits. First, more integrated financial markets can lead to more efficient allocation of saving and investment across countries and, therefore, facilitate consumption smoothing. Second, higher degrees of international capital mobility will enable domestic investors to achieve a higher level of diversification in their investments. Third, the industrial sector will benefit from having better access to the world's capital supply and, eventually this increase in the level of financial integration will have a positive impact on countries' output growth. In all, higher levels of financial integration can lead to a more efficient economy and ultimately to a higher level of economic well being.

That the level of international financial integration between countries is increasing over time is a common belief. Capital controls have been reduced and the share of foreign capital holdings by domestic investors has increased. This phenomenon is not limited to developed countries, as there has been an increase in the level of IFI even among developing countries (see Bekaert et al., 2003).

The removal of capital controls in the 1970s and 1980s in developed countries, and during the 1990s in developing countries (Neely, 1999), is believed to have increased the degree of international capital mobility and financial integration. However, the quantification of the degree of international capital mobility is problematic as there is no widely accepted measure of international financial integration (IFI). While it is sometimes stated that IFI has taken place when the law of one price holds, in reality many different measures have been proposed and used in the literature. These can be broadly classified as follows: IFI is said to have taken place if (a) we observe interest rate parity; (b) there is a lack of correlation between saving and investment ratios; (c) we observe high cross-country consumption growth, and (d) international asset markets are integrated.

In this dissertation the focus is mainly on the behavior of the international bond markets to examine the degree of international financial integration. The work will consist of three essays, each using a different approach to track the degree of IFI. The current research uses information from international asset markets and focuses mainly on stock market behavior to measure the degree of IFI. In contrast, we focus on the international bond markets instead of equity markets, and also we apply a new approach for measuring time variation in IFI.

A brief outline of the three essays is presented in the next section and then a discussion of the main measures used to measure international financial integration in the literature is presented. The main approaches to measuring IFI that have been used in the literature thus far are discussed below. This provides the context for this dissertation, which focuses on alternative approaches to measure IFI. An overview of the contributions that are made in the three essays is then presented.

Interest Rate Parity Conditions

Interest rate parity conditions are one of the tools used to determine the degree of international capital mobility. Some of the early applications of these parity conditions can be found in Akhtar and Weiller (1987) and Reinhart and Weiller (1987). Table 1.1 summarizes the family of parity conditions known as covered interest rate parity (CIP), uncovered interest rate parity (UIP) and real interest rate parity (RIP).

Table 1.1: Parity Conditions

CIP	$i = i^{*} + (f - s)$
UIP	$i = i^{*} + (\mathrm{E}s - s)$
RIP	$\mathrm{E}r = \mathrm{E}r^{*}$

Source: Eijffinger and Lemmen (2003)

Domestic and foreign nominal interest rates are represented by i and i^*, f and s and Es are the forward, spot and the expected spot exchange rates respectively. Er and Er^* are the expected values of domestic and foreign real interest rates respectively.

Under perfect capital mobility, CIP should equalize the return on any two assets issued in two different countries that are identical in terms of maturity, liquidity and default risk. That is, there would be an absence of country premia with: $i - i^{*} - (f - s) = 0$. Frankel and MacArthur (1998) argue that CIP is the appropriate measure of capital mobility. In contrast, factors such as transaction costs, capital controls, and political risk will lead to deviations from CIP conditions. Of interest, therefore is to empirically examine whether we observe deviations from CIP in order to determine where the conditions for capital mobility seem to exit. However, there may be some limitations in using this approach. Montiel (1994) for example, argues that examination of CPI, especially for developing countries, is of limited empirical relevance due to a lack of the appropriate data. Montiel emphasized UIP as the most relevant measure of IFI.

UIP requires, in addition to perfect capital mobility, that perfect capital substitutability apply. Under such conditions, capital flows will equalize the expected rate of return on countries' bonds without covering for exchange rate risk. Hence, for the UIP condition to hold, we require a zero exchange risk premium. It is also necessary to assume that agents' expectations are formed rationally implying that agent's expectations of future exchange rates differ from the ex-post realizations of exchange rates by only a random expectation error. This assumption, however, is often rejected in the literature (see Cumby and Obstfeld, 1984 and Davidson, 1995). Therefore the rejections of UIP conditions can be attributed to either the existence of a time-varying risk premium or to the way that agents form their expectations. MacDonald and Torrance (1990) empirically check for the validity of UIP and conclude that its rejection is attributed to both the existence of a time-varying risk premium and to an expectation factor.

The third parity condition often used when testing for IFI is RIP, and it entails both the perfect mobility of financial and non-financial capital. Mobility in non-financial capital refers to the mobility of goods and services and to the mobility of factors of production—labor and physical capital (Eijffinger and Lemmen, 1993). RIP requires that the expected real exchange rate change equals the ex-ante real interest differential. For

RIP to hold, in addition to UIP it is necessary that the expected real exchange rate change equal zero.

Montiel (1994) argues that this measure “confounds” financial with goods market integration, making it impossible to properly determine the level of financial integration. Frankel and MacArthur (1988) decompose real interest rate parity into two parts: the CIP part and the currency premium (the differentials between local currency and foreign currency interest rates) part. In their analysis, the empirical examination of RIP per se indicates that there is a low level of capital mobility. They therefore examine which part of RIP best explains the deviations. Deviations from CIP can explain only a small portion of deviations from RIP. The currency premium explains most of the real interest rate differentials observed.

Correlation between Savings and Investments

Feldstein and Harioka (1980) (FH hereafter) proposed an alternative empirical method for measuring international capital mobility. In an open economy, under perfect capital mobility, there should be no relation between domestic savings and domestic investment. Under these conditions the level of domestic investment would rely totally on the world’s capital supply and not be limited by domestic saving. Using data from 16 OECD countries covering the period from 1960 to 1974, they estimate the following equation

$$\left({}^{I}\!/_{Y}\right)_i = \alpha + \beta\left({}^{S}\!/_{Y}\right)_i + \varepsilon_i .$$

Where I/Y denotes the ratio of gross domestic investment to gross domestic product (GDP) in country i, S/Y denotes the ratio of gross domestic savings to GDP, α is a constant, and β is a coefficient that measures the degree of financial integration. For a small open economy, perfect capital mobility would imply a value of β equal to zero, meaning that investment decisions in the domestic country do not depend on domestic savings, and a value of β equal to 1 would suggests that domestic investments rely totally on domestic savings. For large open economies, under perfect capital mobility the β coefficient should approximate the country's share of the world's capital stock. Changes in savings in large countries would affect the world interest rate, therefore leading to higher levels of correlation between saving and investments. Feldstein and Harioka (1980) obtain a value of β equal to 0.89, which is not significantly different from 1, suggesting a very low level of capital mobility between the 16 OECD countries for the period 1960–1974.

Although the FH approach did not provide an explicit theoretical framework for measuring capital mobility, it has been extensively applied. Using various techniques, different researchers have estimated FH saving–investment relationships to empirically assess the degree of capital mobility. The results have been robust. Based on the FH approach there seems to be a low degree of capital mobility. This contradicts the common belief of an increase in capital mobility, especially during the last two decades following the removal of capital controls across the globe (see Coakley, 1998).

The FH approach suggests that there is a high saving and investment correlation and also that real interest rate parity holds. Frankel (1992) and Eijffinger and Lemmen (1995), among others, argue that if the domestic real interest rate is not tied to the foreign interest rate, then we cannot expect a zero correlation between savings and investment. Thus, a failure of real interest rate parity can explain the high correlations observed between savings and investment in the various empirical studies. Alternative hypotheses have been suggested to explain the high saving investment correlations observed. Summers (1988) and Bayoumi (1990), argue that current account targeting by the government can result in higher degrees of saving (*S*) and investment (*I*) correlation, regardless of capital mobility. Other factors such as population growth, productivity shocks (Summer, 1988 and Obstfeld, 1986) or the existence of non-traded goods and factor immobility (see Murphy, 1986 and Wong, 1990) will also induce higher correlations between savings and investment.

The FH approach of measuring capital mobility can be undertaken using a cross sectional or a time series approach. Cross sectional analysis seems to give relatively robust results, generally finding the degree of capital mobility to be low (see Feldstein, 1983; Obstfeld, 1986; and Obstfeld, 1995). However, an important limitation of using cross sectional analysis (Sinn, 1992; Jansen, 1994, 1996a) is that cross-sectorial analysis ignores the dynamics of saving-investment correlations and does not take into consideration differences in the economic structure between countries. The cross sectional approach also does not account for nonstationarity in the variables. Consequently more attention has recently been placed on using time-series analysis.

While time series studies are not as plentiful as cross sectional studies, these studies provide us with another chance to sort out evidence for and against IFI. Jansen (1996b) used an error correction model (ECM) to analyze the dynamic relationship between saving and investment

$$\Delta I_t = \alpha + \beta \Delta S_t + \gamma \left(S_{t-1} - I_{t-1} \right) + \delta S_{t-1} .$$

He argues that this model provides us with more than one way to detect capital mobility. In the absence of cointegration between *S* and *I*, this would imply the presence of capital mobility ($\gamma = 0$). Nevertheless, even if cointegration is found between *S* and *I*, this does not necessarily imply capital immobility. There are two possible cases for capital mobility even when there is cointegration between *S* and *I*. First, if δ is different from zero, the current account is not converging to a constant in the long-run, and therefore there is capital mobility. Second, if δ is zero and γ is equal to zero, this would imply that there is no short-run correlation between saving and investment. Applying this ECM to 23 OECD countries, he finds evidence of capital mobility. His results support the idea that *S* and *I* correlations are positively related to country size. Rensselaer and Copeland (2000) use the same ECM to investigate the S-I correlation between 15 Latin America countries and find evidence of capital mobility.

Cross-Country Consumption Growth Correlations

An alternative method used to measure the degree of international capital mobility is the analysis of consumption smoothing and risk diversification. This measure, proposed by Obstfeld (1986), is based on Euler's equation. The main idea is that, if individuals across countries have access to the same set of financial instruments, then under perfect capital mobility, there should be perfect co-movement of a country's consumption growth with world consumption growth. Bayoumi and MacDonald (1995) argue that the analysis of consumption growth correlations based on Euler equation restrictions has the attractive feature that the underlying theory is stronger than that underlying the FH approach. Comparison of consumption growth correlations is different from other methods also because it does not require comparisons of dissimilar assets.

Most of the empirical literature finds low levels of risk sharing between countries, since consumption growth rates are not as highly correlated as one would expect. Eijjfinger and Lemmen (1995) argue that there are numerous factors that might affect the level of correlation. The basic assumptions of this approach are: complete markets, economic integration, identical time preference rates, and constant relative risk aversion. Any violation of these assumptions would result in lower levels of correlation that would complicate the interpretation of the tests for integration. Olivei (2000) also notes that the lack of high consumption risk sharing among the G7 countries may be related to the large levels of non-tradable consumption that is not taken into account when analyzing consumption risk sharing.

Equity Market Integration

In analyzing the level of international financial integration, most recent literature has focused on the behavior of international asset markets. An increase in the level of financial integration will promote faster adjustment of equity prices to information flows, leading to more efficient markets. Therefore an increase in the degree of market integration in general should be associated with an increase in equity market correlations. During the late 1980s and up to the mid 1990s, the analysis of correlations between international equity markets was commonly used to measure IFI. As asset markets become more integrated they are more sensitive to common global or regional shocks, and therefore we should expect a higher level of cross-market correlations. Although this positive relationship between financial integration and equity market correlations has been criticized, this approach allows us to explore IFI as an ongoing process and not as a static event.[1]

One of the earliest studies analyzing stock market correlations is Kaplanis' (1988) study. She divided equity series into four equal sub-samples and investigated whether international correlations between monthly stock markets differed over these sub-periods. No evidence of changes in correlations was found. Ragunathan and Mitchell (1997) tested for the existence of a time trend in the estimated time varying conditional correlations among different national equity returns. Using the diagonal *vech* approach of Bollerslev, Engle and Wooldridge (1988), they did not find sufficient evidence of time varying conditional correlation between equity market returns for the period 1970 to 1990. They also concluded that the 1987 crash had no effect on the time varying conditional correlations. Fratszcher (2002), using daily data on European equity markets, found that there has been an increase in correlations between these market indexes since 1998 when the Euro members were announced. Kearney and Poti (2003) found evidence of an increase in the conditional correlation among euro-zone market indexes. This is interpreted as an evidence of financial integration between European Union countries.

A considerable number of empirical studies have employed the capital asset pricing model (CAPM) to test for international financial integration. Under this model, "asset(s) within a particular country are rewarded in terms of their contributions to the

[1] In earlier literature, typically we would see the analysis of equity market correlations for the period before the 1987 crash, the correlations during the 1987 crash, or the correlations during the 1990s. These sub–periods correspond with particular periods of extreme negative returns or extreme positive returns. Therefore the way that these sub-periods are determined would imply a change in the conditional correlation and not necessarily because of changes in level of financial integration or other fundamental factors. These changes in the correlation level between international equity markets might be simply a result of changes in volatility.

well-diversified world portfolio" (Harvey, 2000, p. 3). Therefore what is important in this model is the covariance between local assets and the world portfolio (known as the beta coefficient). Higher values of the beta coefficient would imply higher levels of IFI. An important limitation of this model is that IFI is considered fixed over time, ignoring time variation in the financial integration process. Bekaert and Harvey (1995) and Bekaert et al. (2003) provide a parameterization of the beta coefficient that allows for time variation of financial integration. But the proposed measure is highly parametric and the results are very sensitive to the choice of instrumental variables used in the estimation.

It is interesting that the focus of most empirical research on IFI appears to be the analysis of international equity markets, while the analysis of the international bond markets has largely been ignored. However, examination of the relative sizes of these markets points to the indisputable fact that bond markets are nearly as large, if not larger, than equity markets in many cases.

Tables 1.2 through 1.5 present the structure of financial markets for the U.S., Euro area, the U.K. and Japan. The size of equity versus bond market differs from one region to another as indicated in these tables. For the U.S. and Japan, the market size for equity holdings appears to be larger than the bond market size for the period 1994–2002, while for the E.U. and the U.K. the bond market is larger than the equity market. Differences observed in the financial market structure can be attributed to different factors such as the stage of development and regulatory restrictions.[2] Nonetheless, the information reported in these tables indicates that bonds are an important component of the investor's

[2] See Rajan and Zingales (2003) for a discussion of the differences between U.S. capital markets and European capital markets and Hartmann, Maddaloni, and Manganelli (2003) for a discussion of the differences between the U.S. and Japan financial markets.

Table 1.2: United States: Financial Market Size[a] (in billions of U.S. dollars)

	1994	1995	1996	1997	1998	1999	2000	2001	2002
Equity	6,318	8,475	10,276	13,293	15,547	19,523	17,627	15,311	11,871
Bonds	7,927	8,562	9,233	9,805	10,768	11,713	12,219	13,491	14,831
Of which:									
Treasuries	3,466	3,609	3,755	3,778	3,724	3,653	3,358	3,353	3,610
Agencies	2,199	2,405	2,635	2,848	3,321	3,912	4,345	4,971	5,525
Financial corporate debt	1,009	1,205	1,383	1,569	1,878	2,080	2,286	2,588	2,985
Nonfinancial corporate debt	1,253	1,344	1,460	1,611	1,846	2,068	2,230	2,579	2,711
Bank loans of nonfinancial corporations	681	766	836	930	1,031	1,122	1,214	1,149	1,066
Total equity, bonds, and loans	14,926	17,804	20,345	24,028	27,346	32,357	31,060	29,950	27,768

Source: Global Financial Stability Report: Market Developments and Issues, April 2004.
[a]Claims on residents.

Table 1.3: Euro Area: Financial Market Size[a] (in billions of U.S. dollars)

	1998	1999	2000	2001	2002	2003:Q1
Equity	4,591	5,498	5,054	4,104	3,279	4,066
Bonds	6,834	6,382	6,278	6,406	7,977	9,260
Of which:						
Government Bonds	3,862	3,460	3,292	3,310	4,122	4,795
Financial corporate debt	2,625	2,570	2,583	2,634	3,293	3,801
Nonfinancial corporate debt	347	353	403	462	562	664
Loans of nonfinancial corporations (NFC)	2,690	2,440	2,499	2,559	3,117	3,448
Total equity, bonds, and loans to NFC	14,114	14,321	11,248	13,068	14,373	16,744
Asset-backed securities (issuance)[b]	…	68	80	80	134	…
Collateralized debt obligations (issuance)[c]	…	42	71	71	114	…

Source: Global Financial Stability Report: Market Developments and Issues, April 2004.
[a]Claims on residents.
[b]For 2002, data shown are as year-to-data as of September 30.
[c]For 2002, data refer to first half of 2002.

Table 1.4: Japan: Financial Market Size[a] (in billions of U.S. dollars)

	1994	1995	1996	1997	1998	1999	2000	2001	2002
Equity	3,232	3,940	3,005	2,443	2,966	4,850	3,173	2,420	2,027
Bonds	5,478	5,755	5,526	5,147	5,919	7,096	6,770	6,257	7,484
Of which:									
Government Bonds	3,490	3,771	3,704	3,585	4,262	5,225	5,121	4,896	6,028
Financial corporate debt	768	710	607	471	482	535	416	315	298
Nonfinancial corporate debt	1,220	1,275	1,215	1,091	1,175	1,336	1,233	1,045	1,159
Loans of nonfinancial corporations	11,918	11,712	10,318	9,234	10,336	11,464	9,983	8,454	8,824
Total equity, bonds, and loans to NFC	20,629	21,407	18,848	16,824	19,222	23,410	19,926	17,131	18,335

Source: Global Financial Stability Report: Market Developments and Issues, April 2004.

[a]Claims on residents.

Table 1.5: United Kingdom: Financial Market Size[a] (in billions of U.S. dollars)

	1994	1995	1996	1997	1998	1999	2000	2001	2002
Equity	1,583	1,860	2,247	2,751	3,107	3,838	3,643	3,156	2,856
Bonds issued[b]	878	1,033	1,252	1,386	1,536	1,637	1,686	1,714	2,059
Of which:[c]									
Government	282	348	398	443	483	449	398	367	441
Financial corporate debt	54	59	74	78	89	109	111	113	130
Nonfinancial corporate debt	105	130	149	170	209	257	300	306	370
Bank loans to nonfinancial corporations	337	368	427	449	471	516	543	583	692
Total equity, bonds, and loans to NFC	8,538	9,617	11,379	12,998	14,112	15,638	15,740	15,564	16,894

Source: Global Financial Stability Report: Market Developments and Issues, April 2004.
[a]Claims on resident.
[b]Includes bond issued by nonresidents.
[c]Following are selected components of the above aggregate.

portfolio, suggesting that international bond markets deserve the same attention as the equity markets for deriving conclusions about IFI.

But the importance of the international bond markets does not rely only on the size of these markets. The behavior across international bond markets, in addition to the financial and economic integration across these countries, reflects also the political efforts to increase the degree of international financial integration. As pointed out by Barassi et al. (2001) the degree of integration across international financial markets can be viewed also as an increase in policy convergence across countries. Hence by analyzing the behavior of international bond market returns, we will be able to look at a different dimension, and a broader picture of IFI. In the next section an overview of this dissertation, which focuses on analyzing IFI via international bond markets, is presented.

1.2 Dissertation Summary

The goal of this dissertation is to measure the degree of IFI by focusing mainly on the relationship across international bond market returns. Three different essays develop different approaches to measuring the degree of IFI across countries. These essays are presented in Chapters 2 through 4. Two important aspects of the work distinguish this dissertation from the vast empirical literature on detecting the degree of international financial integration. First, the focus is mainly on the behavior of the international bond markets, which appears to have been neglected in the literature of IFI. Second, a new approach of detecting the dynamics of IFI across international bond market returns is applied.

The data used in the three essays of this dissertation will consist of daily returns for 13 international bond markets (Australia, Austria, Canada, Denmark, France, Germany, Ireland, Japan, Netherlands, Sweden, Switzerland, the U.K. and the U.S.). These data cover the period 22 June 1989 up to 22 June 2004. The financial markets considered in these essays do not have the same trading hours. The trading times for the financial markets of the U.S. and European countries overlap only partially and some other markets, such as the U.S. and Japanese financial markets, do not have common trading hours. This lack of common trading hours between these international financial markets leads to different problems when these asynchronous data are used in estimation. Martens and Poon (2001) have shown that the use of such asynchronous data will result in a significant downward bias of the correlations. In order to avoid the problem of asynchronous data, the daily indexes are converted into weekly indexes. The rest of this section presents a brief outline of the three essays presented in Chapters 2 through 4.

An Analysis of the Dependence Structure between International Bond Markets

An extensive literature exists that analyzes international equity markets: their distribution, correlations, co-movement and whether they have changed over time. In contrast, the literature on international bond markets lacks this detailed analysis. In the first essay, this gap in the literature is filled. The dependence structure among international bond returns is documented by analyzing the correlations between these returns. Is the same asymmetry in the correlations across bond markets observed as has been observed in the equity markets? Are periods of extreme negative returns associated with higher correlations than periods of extreme positive returns? Have these correlations changed over time? What is the dependence between these bond returns? Is the level of co-movement among bond markets related to the level of synchronization of business cycles across different countries?

The analysis of the correlations among international equity returns has been widely analyzed. An extensive literature has documented asymmetry in the conditional correlations among international stock indexes. The correlation appears to be higher for values in the negative tail than for values in the positive tails and this asymmetry cannot be explained by the normal distribution. Different characteristics across the asset markets, such as the risk of these assets and their recent history (see Ang and Chen, 2002) or contagion (see Dornbusch et al., 2000) are two possible explanations of this observed asymmetry. In general one can argue that higher asymmetry correlations can be an indicator of heterogeneity across international asset markets. On the other hand, lack of this asymmetry in correlations or lower values can be an indicator of homogeneity across asset market returns.

Whether the same behavior is also observed across the international bond market returns is tested. We do not find strong evidence of asymmetric correlations. This differs from previous analyses. Since IFI can be measured by comparing returns between similar assets issued in different countries, we can use the results of tests regarding symmetry or asymmetry to provide better inferences with regard to the degree of IFI.

Significant evidence of asymmetry in the correlations across European bond market returns is not found, indicating that the European bond market returns share similar characteristics, or there is a lack of contagion across these markets. But this is not true for other international bond market returns. These results are not surprising. The sample period considered in our analysis reflects a period of significant changes within European countries. This is reflected in the observation that the bond markets of these countries have become more similar. However, these results are not robust to the definition of extreme bond returns.

Rolling correlation of bond returns over time is analyzed and an overall increase in the co-dependence between international bond returns is found, especially during the latter part of the sample period, from the mid-1990s onward. The results indicate full financial integration starting from the mid-1990s, especially for the bond market returns of Austria, France, Germany and the Netherlands. On the other hand, evidence is found that the U.K. bond market appears to be more financially integrated with the bond markets across the Atlantic, particularly the U.S. bond markets. The U.K. bond market does not appear to be financially integrated with the European bond markets.

But when analyzing the business cycles across these countries strong evidence of business cycle synchronization is not found. This suggests a lower level of integration than the rolling correlations and cointegration analysis indicated. To further understand the inconsistencies provided by the analysis of business cycle synchronizations relative to the analysis of the international bond market and with respect to rolling correlations and cointegration, different approaches of measuring the degree of integration across international bond market returns are offered in the next two essays. These different approaches will enable a better understanding of these inconsistencies.

Do International Bond Market Returns Share a Common Volatility Process?

In the third chapter of this dissertation a different approach is taken to investigate the degree of IFI across international bond market returns. The existence of common features for the volatility of international bond markets is tested. Investigating the common feature in bond markets will add to the literature of international financial integration and in particular to the relatively small literature that analyzes the relationship among international bond markets. The common feature approach is closely related to the concept of integration, and as Engle and Kozincki (1993) state, it is a generalization of cointegration in the first moments. Using this approach, the date when IFI became pronounced cannot be pinpointed, nor can how it has evolved over time be explained. But, as in the cointegration methodology, a conclusion can be made whether there are common factors determining co-dependence in the second moments across the international bond markets.

The results presented in this essay indicate that the bond market returns of Germany, France, Austria and Netherlands share the same volatility process. This indicates the presence of regional integration among these markets. With respect to the U.K. bond market returns, strong evidence that the bond returns of this country share the same volatility process with other European bond returns is not found. The volatility process for this market is not closely related with those of other European countries indicating a weak integration of this market with other European markets. There is evidence of cointegration in the second moments between the U.S. and U.K. bond markets. In addition, the common feature results presented in this essay suggest that U.S.

bond market impacts bond markets globally. Evidence of common volatility process between U.S. bond market returns and European bond market returns is found.

A Time Varying Measure of International Financial Integration

The IFI measures of international bond market returns presented in the previous two essays have consisted mainly of three methodologies: rolling correlations, cointegration in the first moment (Johansen, 1988 methodology), and cointegration in the second moments (common feature methodology). Each of these three approaches seems to suggest that, by and large, IFI is indeed present and strong. There is nonetheless doubt that the IFI process is complete for several reasons. First, the results with respect to the output synchronization do not point to full IFI. Second, while the cointegration analyses (with respect to the first and second moments) indicates that there are two cointegrated regions, details are deficient because of the lack of dynamics in these approaches. IFI could be rising over time or falling over time and these methodologies do not allow observation of these trends, which would have important implications for deriving conclusions about the actual degree of IFI across countries. Third, while the rolling correlations allow observation of the dynamics in IFI and how they have changed over time, this methodology is suspect. Trends in the degree of correlations over time and the changes in the volatility are confounded. Therefore, our results are compared with yet another dynamic approach to IFI.

In this essay a new method for measuring time variation of international financial integration is applied. This measure differs from the rolling correlations approach that is analyzed in Chapter 2, and corrects for the changes in the dynamic dependence across international bond market returns that are due to higher volatilities and does not rely on highly parametric models. Using the recently developed dynamic conditional correlations multivariate—generalized autoregressive conditional heteroskedasticity (DCC MV-GARCH) model of Engle (2002) time-varying beta coefficients are obtained in the capital asset pricing model (CAPM). In addition, the IFI results obtained from using this new method for measuring the time variation of international financial integration are compared with the results obtained from two traditional approaches of measuring IFI: saving investment correlations and cross-country consumption correlations.

The results indicate a significant increase in the level of financial integration between Austria, France, Germany and Netherlands. However, these countries are far from being fully integrated. Their level of financial integration has increased, in particular beginning in the mid1990s. However, during the last 3–4 years there is some evidence that this integration may be trending downward. The U.K. does not show any significant increase in the level of integration with these European countries. The results obtained in the previous chapters indicate a higher level of integration between U.K. and U.S. than the integration between U.K. and other European countries. In this chapter, although there is some evidence of a higher level of integration between U.K. and U.S., the results are not as strong as the previous findings, suggesting that U.K. is not highly integrated with the European countries, or with the rest of the world.

In conclusion, the results presented in this dissertation indicate that among the nine European bond market returns analyzed only the bond market returns of Austria, France, Germany and the Netherlands appear to have reached a stronger level of financial integration. However, even for these markets, the last 2–3 years indicate the possibility of a decline, in their level of integration. Difficulties in fulfilling the Maastricht Treaty criteria and the global recession are possible explanations for the observed decline in the level of IFI.[3] Consistent with previous work, a low level of integration between the U.K. and these European countries is found. However, different from the common belief, the results indicate that the U.K. has a low level of integration even with the U.S.

[3] The five Maastricht convergence criteria are: (1) each country's inflation rate should not be more than the average of the lowest three inflation rates in the European Monetary System; (2) each country's long-term interest rates should be within 2% of the average long-term interest rates of the three countries with the lowest inflation rates; (3) each country must have been a member of the narrow band of fluctuation of the exchange rate mechanism for at least two years without realignment; (4) each country's budget deficit should not be greater than 3% of GDP; (5) each country's national debt should not be more than 60% of its GDP.

CHAPTER 2

AN ANALYSIS OF THE DEPENDENCE STRUCTURE ACROSS INTERNATIONAL BOND MARKETS

2.1 Introduction

In this chapter, an attempt is made to measure the degree of international financial integration (IFI) by analyzing the behavior of international bond market returns. We parallel international stock market analysis of IFI using the international bond market returns focusing mainly on two common approaches to IFI: (1) rolling correlations, and (2) cointegration analysis.

The dynamics of rolling correlations across international bond market returns are analyzed in order to determine whether the relationship across these markets has increased or decreased over time. The dynamics of these rolling correlations would indicate whether there has been an increase or decrease in IFI over time. Next, vector autoregression cointegration methodology is used to analyze the degree of cointegration across international bond market returns. This allows discussion regarding whether the common trends across these market returns exist. That is, whether or not international market returns tends to move together over time.

One common empirical finding on the international stock market is that there is asymmetry in the correlations of these market returns. The period of extremely negative returns in these stock markets is associated with higher correlations than periods of extreme positive returns. While no theoretical explanation exists for the observed asymmetry, several hypotheses have been offered. Ang and Chen (2002) suggest that the asymmetry in correlation is related to differences in stock characteristics, such as risk level, and most recent trends of the returns of these assets. An alternative explanation of these asymmetric correlations is contagion (rational or irrational). Dornbush et al., (2000) argue that contagion, which is the spread of market disturbances mostly during downside periods, is reflected in asymmetric correlations across asset markets.

Heterogeneity across the asset markets can be a possible explanation for asymmetry observed in the correlations across asset markets. Therefore the more similar, or homogeneous, these asset markets are we would expect to find no (or little) evidence of asymmetry in correlations across these asset market returns. In contrast, the more dissimilar these asset markets are, the higher the level of asymmetry is expected to be.

In this chapter whether the same asymmetric behavior is observed across the international bond market returns as has been observed in the U.S. and international equity markets is tested. Lack of this asymmetry would indicate similar characteristics across international bond market returns. Therefore, by analyzing the international bond market returns better inferences with regards to the degree of international financial integration would be received.

In addition, an attempt is made to find how the relationships across international bond market returns are reflected in output correlations. Is the higher degree of financial integration observed across international bond market returns also reflected in higher output synchronization across these countries? Before presenting results on the degree of IFI across international bond market returns, the main findings are reviewed with regard to the degree of IFI by analyzing the behavior across international asset markets.

The analysis of correlations across international asset market returns, especially across international stock market returns, has become a common measure for detecting the degree of IFI across these markets. An increase in the level of financial integration will promote faster adjustment of equity prices to information flows, leading to more efficient markets. Therefore, an increase in the degree of market integration in general should be associated with an increase in asset market correlations. As asset markets become more integrated, they are more sensitive to common global or regional shocks and therefore a higher level of cross-market correlations should be expected.

One of the earliest studies analyzing stock market correlations is by Kaplanis (1988). She divided equity series into four equal sub-periods and investigated whether international correlations between monthly stock markets differed over these sub-periods. No evidence of changes in correlations was found. Ragunathan and Mitchell (1997) tested for the existence of a time trend in the estimated time varying conditional correlations among different national equity returns. Using the diagonal *vech* approach of Bollerslev, Engle and Wooldridge (1988), they did not find sufficient evidence of time varying conditional correlation between equity market returns for the period 1970 to 1990.[4] They also concluded that the 1987 crash had no effect on the time varying conditional correlations. Fratszcher (2002), using daily data on European equity markets, found that there has been an increase in correlations between these market indexes since 1998 when the Euro members were announced. Kearney and Poti (2003) found evidence of an increase in the conditional correlation among euro-zone market indexes. This was interpreted as evidence of financial integration between European Union countries.

[4] The term *vech* comes from the column-stacking operator, VECH(.), applied to the upper triangle of a symmetric matrix.

One important finding emerges from different studies analyzing the degree of integration across international asset market returns. There is no evidence of an increase in the degree of integration between the U.K. and other European asset market returns. Moreover, U.K. asset markets appear to have become more financially integrated with the U.S. asset markets. Alexander (1995b), using Granger (1986) causality tests and Engle and Kozicki's (1993) common feature methodology analyzed the dependence across international bond market returns. Her results indicated that the causal influence of European bond markets on the U.K. bond market returns has decreased, especially after 1992. In addition, from this period onward the U.K. bond market returns appear to be influenced more by the U.S. bond market returns. Fraser and Oyefeso (2002) showed that U.K. stock market returns are becoming more sensitive to the shocks originating in the U.S. market relative to those coming form EU countries.

In contrast to the equity literature, there are relatively few studies that examine the co-dependence between international bond markets and, in particular, their correlations. The few empirical studies that analyze international bond markets focus mainly on four countries: the U.S., the U.K., Germany and Japan.

Clare and Lekkos (2000) investigated the relationship among government bonds issued by Germany, the U.K. and the U.S. Using a vector autoregression (VAR) approach for weekly data covering the period 1990–1999, they found that the yield curves for each of these countries is influenced by international factors, especially during the financial crises periods (although this increase in dependence is not as significant as in international equity markets).

Hunter and Simon (2003), using a bivariate conditional correlation GARCH model, investigated the lead-lag relationship between 10-year government bond returns between the U.S., the U.K., Germany and Japan. Their results suggested that, different from equity markets, the correlations between international bond markets do not increase during turbulent events. Therefore, the benefits of portfolio diversification in bond markets do not diminish during extreme negative events. They suggested that the observed increase in the correlation between these countries' bond returns can be explained with the fact that business cycles are becoming more synchronized.

A description of the data used in the analysis and summary statistics are presented in the next section. Section 3 presents the analysis of asymmetric correlations across international bond returns. In Section 4 rolling correlations are obtained to examine the dynamic structure across these markets. In order to be able to better understand the dependence among these international bond returns, a cointegration analysis of the bond returns is conducted. Section 5 presents the methodology that is used to determine the level of cointegration across international bond markets. The cointegration results of the bond markets are presented in Section 6. In addition, whether the level of co-movements among bond markets is related to the level of synchronization in business cycles across these countries is analyzed. These results are presented in Section 7 and are followed by a short conclusion.

2.2 Data and Summary Statistics

Data used in this essay consists of 10-year DataStream Benchmark Bond indexes, measured in U.S. dollars, for 13 international bond markets: Australia, Austria, Canada, Denmark, France, Germany, Ireland, Japan, Netherlands, Sweden, Switzerland, the U.K. and the U.S.[5] These bond indexes are available for daily returns. The data cover the period June 22, 1989 to June 22, 2004. In order to avoid the problem of asynchronous data, these daily indexes are converted into weekly frequencies.[6] Thus, the bond data in levels will consist of the natural logarithm of these weekly indexes.

Before analyzing the dependence structure of the international bond markets, the stationary of the bond data in log levels is tested. Two commonly used tests, the augmented Dickey-Fuller (ADF, 1979) and Kwiatkowski, Phillips, Schmidt, and Shin (KPSS, 1992) tests are conducted to test for the presence of a unit root in the bond data expressed in log levels. The results are presented in Table 2.1. The ADF test takes the unit root as the null hypothesis. The test regression used to test for the presence of a unit root where both an intercept and a trend are included is as follows:

$$\Delta y_t = a_0 + \gamma y_{t-1} + a_1 t + \sum_{i=2}^{p} \beta_i \Delta y_{t-i+1} + \varepsilon_t$$

Where p is the lag level used using the Akaike information criteria (AIC) and Schwarz information criteria (SIC). The null hypothesis of a unit root implies $\gamma = 0$. The test

[5] It is common practice to analyze the dependence structure across international asset market returns using returns converted into a common currency. Often these returns are converted into U.S. dollars (see Bekaert et al., 2003, and Engle et al., 2003). By using a common currency the underlying assumption is that the investors are not able to hedge any of the foreign exchange risk. In contrast, the use of returns dominated in local currency implies that the investors are able to hedge the currency risk. However, this approach does not take into consideration the transaction costs that would incur through currency hedging. We use bond indexes measured in U.S. dollars; in this way we are able to maintain comparability of our results. Fratzcher (2001) analyzed the dependence structure across international asset market returns using both common and local currency indexes. He obtained similar results for both cases.

[6] The financial markets studied in this dissertation do not have the same trading hours. The trading times between these financial markets overlap only partially or they do not have common trading hours. The lack of common trading hours will lead to problems when these asynchronous data are used in estimation. Martens and Poon (2001) have shown that the use of such asynchronous data will result in a significant downward bias of the correlations. In order to avoid problems raised from the use of such asynchronous data, the daily indexes are converted into weekly indexes.

Table 2.1: Unit Root Test for Bond Indexes in Log Levels

		ADF							KPSS
		Lag	No Const.	Const.	Const &Trend	φ_1	φ_2	φ_3	H_0: Level Stationarity
Australia		1	3.22	-1.73	-2.33	1.96	2.46	1.72	10.50*
Austria		2	2.77	-0.86	-1.82	0.53	1.22	1.3	9.84*
Canada		1	3.18	-0.51	-2.52	0.25	2.19	3.04	11.27*
Denmark		2	3.21	-1.33	-2.21	1.81	2.45	1.85	10.33*
France	(AIC)	2	2.99	-1.15	-1.94	0.91	1.48	1.31	9.92*
	(SIC)	1	3.18	-1.36	-2.06	1.22	1.73	1.38	
Germany		2	2.43	-0.9	-1.9	0.54	1.3	1.4	9.59*
Ireland		2	3.1	-0.93	-2.13	0.63	1.67	1.87	10.49*
Japan		1	0.16	-1.63	-1.56	3.67	2.75	0.46	9.94*
Netherlands		1	2.79	-1.11	-1.87	0.82	1.35	1.2	9.61*
Sweden		1	2.69	-0.99	-2.05	0.82	1.66	1.66	10.51*
Switzerland		1	2.47	-1.21	-2.03	0.94	1.56	1.4	9.69*
UK		2	3.19	0.95	-2.71	0.68	2.66	3.31	11.45*
US		1	3.87	-1.15	-2.48	0.95	2.45	2.72	11.60*

Note: For ADF unit root test, the null hypothesis is the existence of unit root. The lag length for this test was selected using AIC and SIC criteria. For all, but French returns, these two criteria suggests the same optimal lag length. For the French bond returns the AIC suggests an optimal lag of 2 and SIC suggests an optimal lag of 1. At both these lag lengths we cannot reject the null of unit root. The 5% critical values for the test statistics φ_1, φ_2, and φ_3 are 4.71, 4.88 and 6.49 respectively. For the KPSS test under the null the series are assumed to be local stationary. The lag length for KPSS test is equal to 5 and was selected using the level selection procedure proposed by Newey and West (1994). * indicates the significance level at the 1% level.

results are sensitive to the inclusion or exclusion of an intercept, an intercept and a trend or neither in the test regression.[7]

Dickey-Fuller (1981) provides three additional tests, φ_1, φ_2, and φ_3, to test for the significance of the inclusion of an intercept and a trend in the above regression. The φ_1 statistic tests the null hypothesis of $\gamma = a_0 = 0$. The φ_2 statistic tests the null of $\gamma = a_0 = a_1 = 0$, and the φ_3 statistic tests the null of $\gamma = a_1 = 0$.[8] In Table 2.1 the ADF test for all three cases are reported: intercept or trend, intercept, intercept and trend. The ADF test cannot reject the null hypothesis of a unit root when the bond data in log levels is first differenced. In addition, the φ_1, φ_2, and φ_3 test statistics indicate that no intercept and no trend are needed in the above regression.

The last column of Table 2.1 reports the KPSS test for a unit root. This test differs from the ADF test in that the series are assumed to be stationary under the null. This test is conducted under the null of level stationarity. Based on these results the hypothesis of stationarity for bond series in log levels is rejected while the null of stationarity for bond returns is not.

Table 2.2 reports the summary statistics for the weekly bond returns. The skewness and the standardized kurtosis coefficients are reported in order to determine whether there is a departure from normality for these bond returns. For a normally

[7] AIC and SIC are two common information criteria used to determine the lag length. However, these criteria often lead to different optimal lag lengths. The AIC tends to overfit the optimal lag length, while the SIC tends to underfit the optimal lag length. The underfit leads to biased estimated coefficients. In cases where AIC and SIC optimal lags lead to different conclusions with regards to the presence of a unit root, we pick the results suggested by AIC. However, even in cases where these two criteria suggest different optimal lag lengths, we cannot reject the null of unit root at both these lags.

[8] See Enders (2004) pp. 181–183 for a more detailed description of these tests.

Table 2.2: Summary Statistics of Bond Returns

	Mean	Skewness	Kurtosis	Observations
Australia	0.20*	-0.19*	0.74*	782
Austria	0.17*	0.05	0.11	782
Canada	0.15*	-0.09	0.49*	782
Denmark	0.20*	0.00	0.13	782
France	0.18*	0.03	-0.10	782
Germany	0.16*	0.06	0.16	782
Ireland	0.18*	-0.24*	0.39*	782
Japan	0.13**	0.82*	4.29*	782
Netherlands	0.16*	0.06	0.08	782
Sweden	0.18*	-0.28*	0.95*	782
Switzerland	0.15*	0.08	0.29	782
U.K.	0.19*	-0.19*	1.62*	782
U.S.	0.13*	-0.51*	0.73*	782

Note: The skewness and standardized kurtosis are tested against the null of zero; * and ** show the significant levels at 1% and 5% respectively.

distributed series the skewness coefficient, which measures the asymmetry of the distribution, should equal zero. A positive skewness coefficient implies that the series has a long positive tail and a negative skewness coefficient implies that the series has a long negative tail. Skewness is an important measure in evaluating the riskness of an asset. In general a positive skewness coefficient indicates that the asset is favored by the market and it is priced at a premium, while a negative skewness coefficient indicates that the asset is not favored by the market, and it is priced at a discount.

The standardized kurtosis coefficient measures the peakness or flatness of the distribution of the series relative to normal distribution. Negative standardized kurtosis coefficient implies that the distribution of the series is flat relative to the normal distribution and a positive value indicate that the distribution of the series has a sharp peak and fatter tails relative to the normal distribution. A positive standardized kurtosis implies that extreme events, positive or negative, are more likely to happen than in the normal distribution case. Therefore, a risky bond index will be characterized by a negative skewness coefficient and a positive standardized kurtosis coefficient.

The results presented in Table 2.2 indicate that there appears to be a departure from normality for the bond returns of Australia, Ireland, Japan, Sweden, the U.K. and the U.S. Among these bond returns only Japanese bond returns display positive skewness. The rest of these returns display negative skewness. The presence of negative skewness is an indicator that these markets give higher probability to the decreases than to the increases of these bond returns. The standardized values for the kurtosis coefficients of the bond returns of Australia, Canada, Ireland, Japan, Sweden, the U.K. and the U.S. displayed in Table 2.2 suggest that these bond returns have thicker tails than in the case of the normal distribution. This implies that extreme observations are more likely than in the normal distribution case. Meanwhile the rest of the bond returns, which represent most of the European countries, do not indicate any departure from the normal distribution. Therefore, these returns are relatively less risky than the bond returns of Australia, Canada, Ireland, Japan, Sweden, the U.K. and the U.S.

Table 2.3 presents the variance, covariance and unconditional correlations for these returns. The elements of the main diagonal give the variance for each return. The covariances are shown on the lower triangle of the table while unconditional correlations are displayed in the upper triangle.[9] The unconditional variances for the international bond returns appear to vary within a small range. For most of the bond returns, the variance is between 2 and 3, except for the U.S. bond returns with an unconditional variance less than 2, and Japanese and Sweden's bond returns with unconditional

[9] Except for the correlation between U.S. and Japanese bond returns, the correlation coefficients are significant at the 1% level. The U.S.–Japanese bond returns correlation is significant at the 10% level.

Table 2.3: Covariance – Correlations across International Bond Market Returns

	Australia	Austria	Canada	Denmark	France	Germany	Ireland	Japan	Netherlands	Sweden	Switzerland	U.K.	U.S.
Australia	**3.00**	0.39	0.49	0.43	0.39	0.41	0.43	0.20	0.39	0.39	0.30	0.39	0.33
Austria	*1.07*	**2.54**	0.27	0.92	0.91	0.98	0.83	0.36	0.95	0.60	0.85	0.65	0.34
Canada	*1.15*	*0.57*	**1.80**	0.33	0.32	0.29	0.38	0.11	0.34	0.35	0.20	0.40	0.58
Denmark	*1.19*	*2.36*	*0.71*	**2.60**	0.92	0.93	0.88	0.33	0.92	0.70	0.82	0.70	0.37
France	*1.11*	*2.36*	*0.69*	*2.42*	**2.64**	0.93	0.87	0.32	0.94	0.66	0.82	0.70	0.39
Germany	*1.14*	*2.51*	*0.63*	*2.43*	*2.43*	**2.61**	0.85	0.37	0.97	0.62	0.86	0.67	0.38
Ireland	*1.24*	*2.21*	*0.85*	*2.37*	*2.38*	*2.29*	**2.81**	0.25	0.88	0.66	0.75	0.78	0.42
Japan	*0.64*	*1.04*	*0.27*	*0.95*	*0.92*	*1.06*	*0.75*	**3.21**	0.35	0.16	0.37	0.17	0.07
Netherlands	*1.09*	*2.47*	*0.73*	*2.43*	*2.48*	*2.55*	*2.40*	*1.03*	**2.65**	0.64	0.86	0.70	0.42
Sweden	*1.31*	*1.84*	*0.92*	*2.17*	*2.06*	*1.92*	*2.14*	*0.54*	*2.02*	**3.71**	0.51	0.55	0.31
Switzerland	*0.88*	*2.34*	*0.47*	*2.27*	*2.29*	*2.40*	*2.17*	*1.14*	*2.41*	*1.69*	**2.95**	0.58	0.30
U.K	*1.00*	*1.51*	*0.78*	*1.65*	*1.66*	*1.59*	*1.91*	*0.44*	*1.67*	*1.56*	*1.47*	**2.14**	0.50
U.S.	*0.59*	*0.56*	*0.80*	*0.61*	*0.66*	*0.63*	*0.73*	*0.13*	*0.71*	*0.62*	*0.52*	*0.75*	**1.06**

Note: The lower triangle gives the covariances of international bond market returns and the main diagonal gives their variances. The upper triangle gives the correlations across these international bond market returns. Except for the correlations between the U.S. and Japanese bond returns, all correlations are significant at 1% level. The correlation coefficient between the U.S. and Japanese bond returns is significant at the 10% level.

variances greater than 3.0. The covariances between these bond markets are positive, indicating a positive covariation across these markets, although a relatively wide range of dependence can be observed. The analysis of the correlation coefficients enables us to compare the dependence across markets. The range of correlations is very wide. It ranges from 0.07, indicating relatively low correlation between the U.S. and Japanese bond returns, to 0.98, indicating relatively high correlation between German and Austrian's bond returns.

Across European countries the lowest level of correlation is that between Sweden and Switzerland (0.51). In most cases, the correlations are above 0.80. In contrast, Japanese bond returns seem to be the least correlated return with other international bond returns. Japanese bond market returns have the highest level of correlation with German bond returns, 0.37. U.S. bond returns are more correlated with Canadian and U.K. bond returns with correlations of 0.58 and 0.50 respectively. U.S. bond return correlations with other international bond returns range between 0.30 and 0.42, except for Japan. These unconditional correlations suggest a grouping of these international bond returns. The bond returns of the European countries appear to be highly correlated.

The sample period covered in this chapter includes the three stages of European Monetary Union (EMU) that were laid down by the Dolores Report of June 1989. In the first stage, covering the period January 1990 to December 1993, all restrictions on the movement of capital between member states were abolished. In the second stage, January 1994 to December 1998, the focus was on strengthening central bank cooperation and monetary policy coordination and the preparation for the establishment of the European System of Central Banks. The purpose of the European System of Central Banks was the conduct of a single monetary policy and the creation of a single currency. The third stage of EMU, began in January 1999, with participating states in the Monetary Union conducting a single monetary policy under the responsibility of the European Central Bank. These events may explain the high level of correlations of bond returns across most of the European countries.

Among few papers analyzing the international bond returns, Smith (2002) analyzes the seasonality across monthly bond returns for the U.S., Canada, the U.K., Germany, France and Japan. The presence of a January effect is an indicator of market inefficiency. This means that there is information available to the investors that can help them earn abnormal returns. Smith (2002) tested whether there is January effect in government bonds across these countries—a specific type of seasonality where excess returns are observed for the series each January. His results did not show any clear evidence of seasonality. The January effect in the equity market was first documented by Rozeff and Kinney (1976) and later on by Bhardwaj and Brooks (1992) and Eleswarapu and Reinganum (1993).

There is no clear theoretical explanation for why one should observe a January effect. Various explanations have been offered. Some believe that the January effect is caused by capital gains taxes (Ritter, 1988) or anomalies related with the business cycle (Kohers and Kohli, 1991) or to higher trade volumes and lower real interest rates in January (Ligon, 1997). Other researches (DeBondt and Thaler, 1987; Rubinstein, 2000) argue that the January effect is simply due to investor irrationality. But no matter what the source of this January effect is, the fact is that it is often observed in the asset markets. Bhadra, Dhillon and Ramirez (1999) found that the January effect has become stronger since 1986.

The January effect is mainly evidenced in the equity market. Previous work on the seasonality of bond returns has given mixed results with regard to the presence of a January effect in bond markets. Clayton, Delozier and Ehrhardt (1989) found the January effect among U.S. bond returns, while Schneeweis and Wooldridge (1979) do not find evidence of seasonality in the U.S. bond returns.

For the purpose of analyzing the dependence structure between international bond market returns, an appropriate stationary transformation of these data is needed. Given the above literature on the possibility of a January effect in asset markets, and in particular in international bond markets, in this section the presence of a January effect across international bond returns is tested. The following regression is used

$$y_{it} = a + b_0 \times dummy1 + \varepsilon_{it}$$

where y_{it} is the bond return of country *i*, *dummy*1 takes value 1 the first week of each January. The results are presented in the first part of Table 2.4. The numbers in parentheses give White's (1980) robust standard errors.[10] The results indicate that there is no January effect in the weekly international bond returns considered here.

The major part of the bond returns considered in this analysis are from European countries. Therefore, we test whether January 1994 and January 1999, two important dates with respect to European Union, had an impact on international bond returns. The following regression is used

[10] The use of White's robust standard errors allows us to make appropriate inferences based on the least squares without actually specifying the type of heteroscedasticity, if it exists.

Table 2.4: January Effect Tests across International Bond Market Returns

	Australia	Austria	Canada	Denmark	France	Germany	Ireland	Japan	Netherlands	Sweden	Switzerland	UK	US
						$y_{it} = a + b_0 \times dummy1 + \varepsilon_{it}$							
b_0	-0.22	-0.02	0.04	0.04	-0.14	-0.12	0.13	-0.66	-0.24	0.04	0.01	-0.31	-0.16
	(0.43)	(0.42)	(0.34)	(0.44)	(0.43)	(0.44)	(0.45)	(0.47)	(0.44)	(0.50)	(0.45)	(0.43)	(0.52)
						$y_{it} = a + b_1 \times dummy2 + b_2 \times dummy3 + \varepsilon_{it}$							
b_1	0.83^{*}	0.20^{*}	0.66^{*}	0.61^{*}	-0.46^{*}	-0.14^{*}	2.05^{*}	0.25^{*}	-0.18^{*}	2.52^{*}	0.89^{*}	0.02	0.95^{*}
	(0.06)	(0.06)	(0.05)	(0.06)	(0.06)	(0.06)	(0.06)	(0.13)	(0.06)	(0.07)	(0.06)	(0.06)	(0.04)
b_2	0.18^{*}	1.10^{*}	1.97^{*}	0.59^{*}	0.42^{*}	1.17^{*}	1.03^{*}	2.86^{*}	0.99^{*}	-0.38^{*}	0.15^{*}	-0.14^{*}	1.58^{*}
	(0.06)	(0.06)	(0.05)	(0.06)	(0.06)	(0.06)	(0.06)	(0.13)	(0.06)	(0.07)	(0.06)	(0.06)	(0.04)
Q(4)	2.54	3.36	10.89^{**}	4.47	3.91	2.72	6.15	4.12	5.03	2.87	1.16	4.52	9.85^{**}
Q(8)	11.02	8.67	14.00^{***}	7.43	7.31	6.06	7.98	7.93	7.94	5.16	2.76	16.28^{***}	19.86^{**}
Q(12)	16.31	17.98	17.39	13.73	15.02	12.77	14.02	12.16	15.91	5.67	4.31	19.58^{**}	22.27^{**}
$Q^2(4)$	70.93^{*}	13.82^{*}	23.65^{*}	17.25^{*}	26.46^{*}	21.52^{*}	44.83^{*}	12.62^{*}	29.94^{*}	11.33^{*}	4.41	12.27^{*}	20.56^{*}
$Q^2(8)$	87.10^{*}	19.73^{*}	32.62^{*}	19.27^{*}	41.12^{*}	23.58^{*}	49.39^{*}	35.45^{*}	35.13^{*}	51.06^{*}	9.72	13.87^{***}	27.86^{*}
$Q^2(12)$	101.46^{*}	26.40^{*}	37.43^{*}	41.35^{*}	58.65^{*}	30.96^{*}	54.82^{*}	54.71^{*}	45.86^{*}	67.92^{*}	15.98	19.42^{***}	29.55^{*}

Note: The first part of this table presents the results when we test for a January effect in international bond returns. The variable *dummy1* takes value one for each first week of January each year. The second part of this table presents the results when we test for the impact of January 1, 1994 (*dummy2*) and January 1, 1999 (*dummy3*) on international bond returns for period June 1989 to June 2004. The numbers in parentheses are robust standard errors. The superscripts *, ** and *** indicate significance at the 1, 5 and 10 percent levels.

$$y_{it} = a + b_1 \times dummy2 + b_2 \times dummy3 + \varepsilon_{it} \qquad (1)$$

where y_{it} is the bond return of country *i*, *dummy2* takes the value 1 on the first week of January 1994 and *dummy3* takes on the value 1 on the first week of January 1999. The results are presented in the second part of Table 2.4. The numbers in parentheses are White's (1980) robust standard errors. Except for the U.K. on January 1, 1994, these two dates appear to have a significant impact on the international returns. On January 1, 1994, there is a negative effect on the bond returns for France, Germany and Netherlands. This result may be related to the great skepticism observed during this time on whether the 3rd stage of European integration would go through. The January 1, 1999 effect is negative only for Sweden. Note that in January 1999, Sweden did not adopt the euro as its official currency. In May 1998, the EU Council of Minister stated that Sweden did not fulfill the criteria for joining.

The last part of the Table 2.4 reports the Q-statistics for the residuals and squared residuals from equation (1) at lags 4, 8 and 12. The Q-statistics for these residuals are insignificant for most of the international bond returns, indicating no autocorrelation. For Canada, the U.K. and the U.S., the Q-statistics indicate the presence of autocorrelation. For these particular residuals, the Box-Jenkins (1976) method is used to remove the autocorrelation observed. For the U.S. and U.K. residuals an MA(4) process is fitted and for the Canadian residuals an MA(3) process is fitted. The Q-statistics for the squared residuals are significant, indicating the presence of ARCH (autoregressive conditional heteroscedasticity) in these series.

Thus, the results presented in Table 2.4 indicate that there is no January effect in weekly international bond returns, but the start of the second and the third stages of EMU have had a significant effect in most of these markets. We correct for the January 1994 and January 1999 effects and the autocorrelation observed in the bond returns of Canada, the U.K. and the U.S.[11] We then use these corrected bond in the analysis that follows.

2.3 Correlation Profiles

In this section, the asymmetry in the correlations between international bond returns is analyzed. In contrast to the vast work in analyzing the asymmetry between international equity returns, international bond markets lack this thorough analysis. Ang and Chen (2002) suggest that the asymmetry in correlation is related to differences in stock characteristics, such as riskness level, and most recent trends of the returns of these assets. An alternative explanation of these asymmetric correlations is contagion (rational or irrational). Dornbush et al. (2000) argue that contagion, which is often referred to as the spread of market disturbances mostly during downside periods, is reflected in asymmetric correlations across asset markets.

Heterogeneity across the asset markets can be a possible explanation of asymmetry observed in the correlations across asset markets. Therefore, it would be expected that the more similar, or homogeneous, these asset markets are, the less likely it is that evidence would be found of asymmetry in correlations across these asset market returns. In contrast, the more dissimilar these asset markets are, the higher the level of asymmetry is expected to be.

In this section, we try determine if the same asymmetry observed across international stock market returns is also observed across international bond market

[11] The correction implies that the residuals obtained from regression (1) are used in the following analysis. From this point on we refer to these residuals as the bond returns.

returns. First, the correlations of each of these countries' bond returns with U.S. bond returns are calculated. If there is no asymmetry across the international bond markets then there is no significant difference in the correlations during bear and bull markets. As Boyer, Gibson and Loretan (1999) showed, the conditional correlation between two random variables x and y from a normal distribution increases as x is in the tail of the distribution. Let the event A $\subset$ R be such that 0 < Pr(A) <1. Then the correlation between x and y conditional on the event $x \in$ A derived by Boyer, Gibson, and Loretan (1999) is

$$\rho_A = \rho\left(\rho^2 + \left(1-\rho^2\right)\frac{Var(x)}{Var(x \mid x \in A)}\right)^{-1/2} \qquad (2)$$

where ρ and ρ_A are unconditional and conditional correlation coefficients between x and y. As x takes values further in the tails, the variance ratio in (2) becomes smaller, and therefore, the conditional correlation between x and y increases resulting in symmetric U-shape correlations.

We are interested in testing if the correlations during bear and bull markets are significantly different. Following Forbes and Rigobon (2002), a *t*-test is used to evaluate if there is a significant increase in the correlations during the bear market relative to the correlations during the bull market. Let ρ^{Bear} be the correlation during the bear market and let ρ^{Bull} be the correlations during the bull market then the test hypotheses are:

$$H_0: \quad \rho^{Bear} > \rho^{Bull}$$

$$H_1: \quad \rho^{Bear} \leq \rho^{Bull}.$$

In the literature, different ways of defining extreme events have been used. Choosing a very high threshold value will lead to fewer observations falling in the area of extreme returns. When a very small threshold value is chosen there would be too many observations satisfying the condition of being considered as an extreme return. So Longin and Solnik (2001) consider as thresholds levels ± 3%, ± 5%, ± 8% and ± 10% away from the empirical mean. Hunter and Simon (2003) consider the mean plus two standard deviations as a threshold values. Butler and Joaquin (2002) suggested that by grouping the returns in three equal subsets we can avoid the problem raised by choosing a low or high threshold value.

To be able to analyze whether the correlation changes when U.S. returns take on values in the negative tail or the positive tail, Butler and Joaquin's (2002) approach is followed. Twelve pairs of bond returns are created where each pair contains the U.S. bond return and one of 12 international bond market returns. Each pair of returns is sorted by the U.S. returns in ascending order, and then is grouped into three subsets with 261, 260 and 261 observations each corresponding to the bear, calm and bull markets respectively. The bear group contains the lowest bond return observations while the bull group contains the highest bond returns observations. These correlation profiles between the U.S. and other bond market returns are presented in the first part of Table 2.5.

There are no negative correlations between these bond returns. The test statistic of whether the correlations during bear market are significantly higher than those during bull markets are given in the fourth column of the first part of Table 2.5. The critical value for this test statistic at the 5 percent level is 1.65, therefore any value greater than 1.65 indicates the presence of asymmetry. The results show clear evidence of asymmetry in the correlations of these international bond returns with the U.S. returns. The correlations are higher during extreme negative returns relative to the correlations in the

Table 2.5: Correlation Profiles

	Correlations of international bond returns with the U.S. bond returns				**Correlations of international bond returns with German bond returns**			
	Bear	Calm	Bull	Test Stat	Bear	Calm	Bull	Test Stat
Australia	0.11	0.18	0.13	-0.33	0.12	0.06	0.13	-0.16
	(0.08)	(0.00)	(0.04)		(0.06)	(0.34)	(0.03)	
Austria	0.28	0.21	0.16	2.03	0.86	0.75	0.88	-1.32
	(0.00)	(0.00)	(0.01)		(0.00)	(0.00)	(0.00)	
Canada	0.30	0.11	0.23	1.21	0.15	0.06	0.00	2.43
	(0.00)	(0.07)	(0.00)		(0.02)	(0.36)	(0.97)	
Denmark	0.31	0.24	0.16	5.15	0.76	0.57	0.78	-0.79
	(0.00)	(0.00)	(0.01)		(0.00)	(0.00)	(0.00)	
France	0.32	0.21	0.20	2.07	0.78	0.50	0.78	0.00
	(0.00)	(0.00)	(0.00)		(0.00)	(0.00)	(0.00)	
Germany	0.30	0.22	0.17	2.21	1.00	1.00	1.00	0.00
	(0.00)	(0.00)	(0.01)					
Ireland	0.35	0.20	0.17	3.11	0.67	0.45	0.63	1.11
	(0.00)	(0.00)	(0.01)		(0.00)	(0.00)	(0.00)	
Japan	0.00	0.12	0.09	NA	0.09	0.09	0.30	-3.52
	(0.94)	(0.05)	(0.15)		(0.16)	(0.13)	(0.00)	
Netherlands	0.33	0.21	0.23	1.75	0.90	0.75	0.92	-1.88
	(0.00)	(0.00)	(0.00)		(0.00)	(0.00)	(0.00)	
Sweden	0.23	0.18	0.16	1.17	0.37	0.29	0.29	1.44
	(0.00)	(0.00)	(0.01)		(0.00)	(0.00)	(0.00)	
Switzerland	0.23	0.24	0.18	0.84	0.63	0.30	0.68	-1.41
	(0.00)	(0.00)	(0.00)		(0.00)	(0.00)	(0.00)	
UK	0.38	0.16	0.16	3.83	0.38	0.19	0.35	0.56
	(0.00)	(0.01)	(0.01)		(0.00)	(0.00)	(0.00)	
US	1.00	1.00	1.00		0.15	0.06	0.15	0.00
	-	-	-		(0.02)	(0.34)	(0.02)	

Note: The numbers in parentheses are the p-values. The test statistic tests the null hypothesis that there is a significant increase in the correlations during the bear markets relative to the correlations during bull markets, i.e. $\boldsymbol{H_0}:\ \rho^{Bear} > \rho^{Bull}$. The critical value for the test statistic at the 5% significant level is 1.65.

right tails, which correspond to bull markets. It is important to point out that the correlations between the U.S. and Japan's bond returns are not significantly different from zero. The movements in Japan's bond market do not appear to be affected by the changes in the U.S. bond market.

The bond returns of Austria, Denmark, France, Germany, Ireland, Netherlands and the U.K. appear to be more correlated during the bear market with the U.S. bond returns. These results confirm those found by Cappiello, Engle and Sheppard (2003) but not those found by Hunter and Simon (2003). There are two reasons why the results presented here differ from those of Hunter and Simon (2003). First, the way that the extreme returns are defined in this paper differs from the methodology used by Hunter and Simon (2003). Second, in determining the extreme events we condition only on the U.S. returns as the main interest is in determining how the international bond returns react to extreme negative or positive returns in the U.S. bond market.

A different perspective is taken in the second set of results displayed in Table 2.5. The correlation profiles between German bond returns and other international bond returns are analyzed. Typically, in the literature we would find that the analysis of the correlations among international financial markets is mainly focused on the correlations between the U.S. and other financial markets, with the U.S. market considered as a proxy for the world market. However, there is considerable evidence (see Fratzcher, 2002; Bekhaert and Harvey, 2002) that the level of *global* financial integration is different from the level of *regional* financial integration. In particular, for the European countries, the level of regional financial integration is believed to have increased, especially during the last decade. In this case, one would expect a different dependence structure among the European bond markets as compared to other international bond markets. Germany has often been considered the reference country when analyzing the European economies.

Following this approach, thus we consider the correlations of German bond returns with other international bond returns.

The results presented in Table 2.5 do not present any significant asymmetry in the correlations of German bond market with other European bond markets, except for the correlations of Netherlands' bond returns with Germany's bond returns. There is evidence of the asymmetry between the correlations of German's bond returns with other non-European countries bond returns, in particular Canada and Japan.[12]

In summary, the results presented in Table 2.5 indicate that the European bond market returns share similar characteristics such as riskness level. But this is not true for other international bond market returns. These results are not surprising. The sample period consider in our analysis reflects a period of significant changes within European countries. This is reflected in the fact that the bond markets of these countries have become more similar.

2.4 Moving Correlations

In order to analyze the dynamic structure of correlations between the international bond indexes, in this section we look at 52-week moving correlations between bond returns. Figure 2.1 presents the moving correlations between the U.S. and other international bond returns. There are several patterns that one can point out when

[12] We test the robustness of results presented in Table 2.5 by changing the definition of bear, calm and bull market. In Table 2.5, bear and bull markets are defined as each containing 33% of the extreme observations. When we change the definition of the bear and bull markets, each containing 25% of the extreme observations, the results are similar to those presented in Table 2.5. However, these results change when bear and bull markets are defined as containing 20%, or less, of the extreme observations. It is also interesting that in this case we do not always get the U-shaped correlation profiles. For some bond market return pairs, the correlations during calm markets appear to be greater than the correlations during the bear and/or bull markets. These results suggest that the asymmetric correlation results presented in Table 2.5 are not robust to the definition of the outliers.

Figure 2.1: 52-Week Rolling Correlations of International Bond Returns with the U.S. Bond Returns

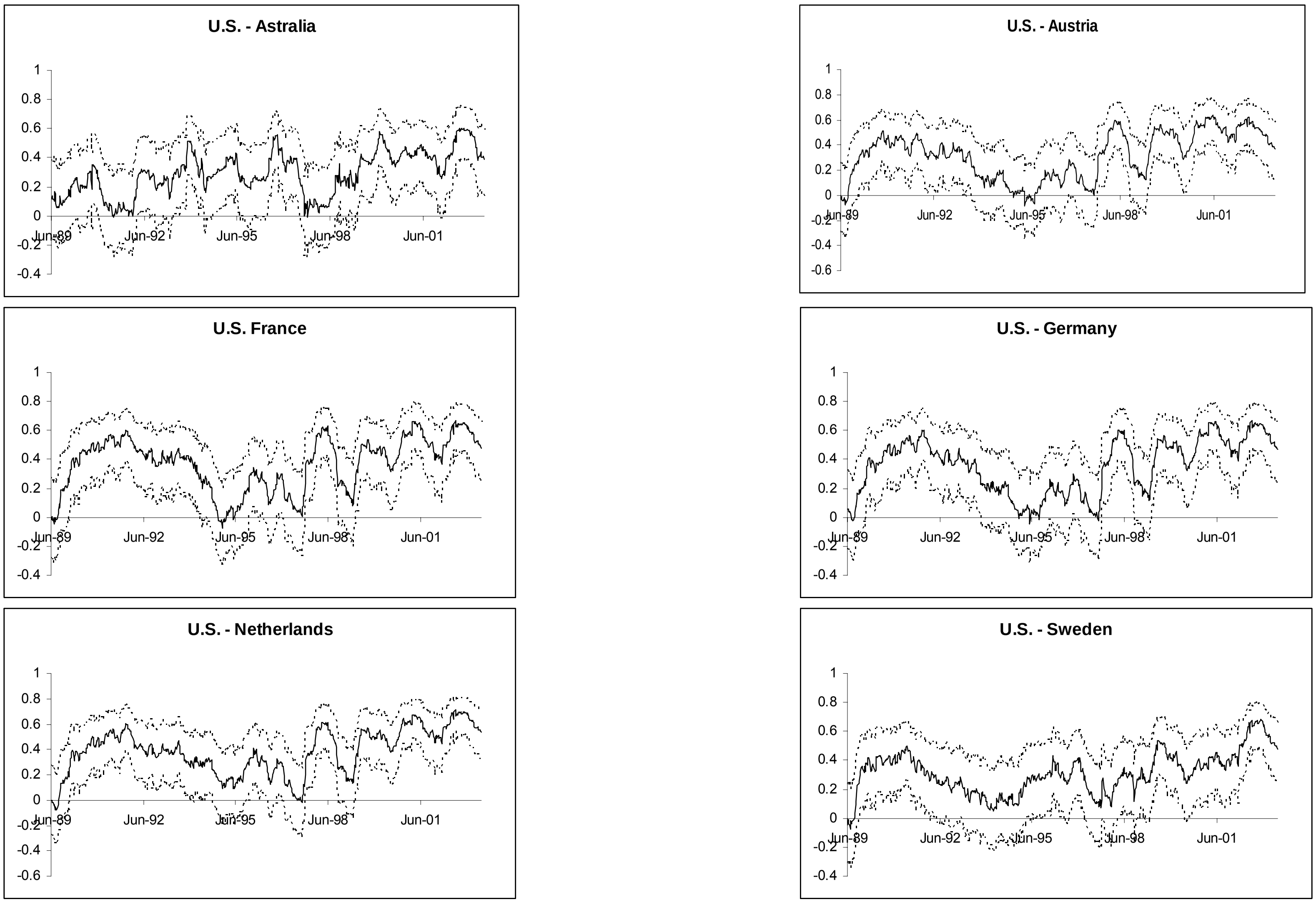
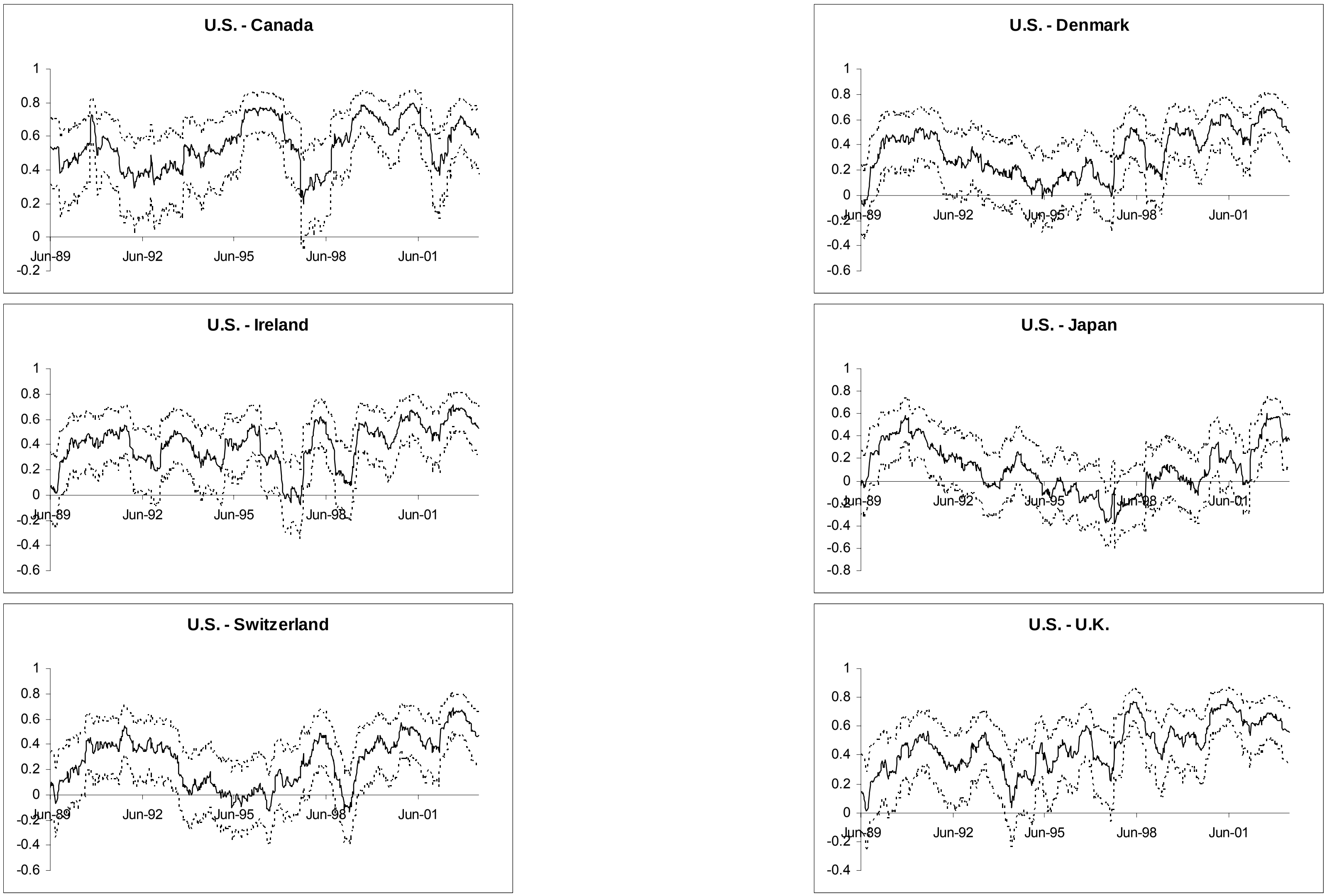


Figure 2.1 (continued)



Note: The solid lines represent the correlations, while the dotted lines represent 95% confidence bounds.

observing these rolling correlations. First, the U.S. correlation with European countries shows a decline during the period 1991 to 1996, except for the correlations with Ireland and the U.K. With respect to Ireland, the correlations show clear evidence of an upward trend only during the last 3 to 4 years. For the U.K., there appears to be a sharp decline in the correlations with the U.S. only during 1994, and afterward these correlations seem to trend upward.

The U.S. correlations with other non-European countries follow different patterns. With respect to Canada, they remain stable almost the entire period, with a correlation of around 0.6. Only in 1997, there appear to be a temporary drop in the correlations between U.S. and Canadian bond market returns. The correlations with Australia show a positive trend, although not a very significant one. The correlations of the U.S. bond returns with the Japanese bond market returns are negative for a considerable period, from 1994 to 2000. This period corresponds with the asset bubble burst and the sluggish economic performance of Japan. This period is often known as the lost decade of the Japanese economy (Fukao, 2003). However, we should point out that the correlations of the U.S. bond market returns with the Japanese bond market returns are significantly different from zero only during the 1990–1992 period, in 1997, and from 2002 onward.

Figure 2.2 presents the moving correlations between German and other international bond returns. The graphs display different patterns of correlations between German bond market returns and other European bond market returns versus the correlations between German bond market returns with other non-European bond market returns. Even between European countries the pattern of the rolling correlations is different. It is important to mention here that not all these countries are members of the

Figure 2.2: 52-Week Rolling Correlations of International Bond Returns with German Bond Returns

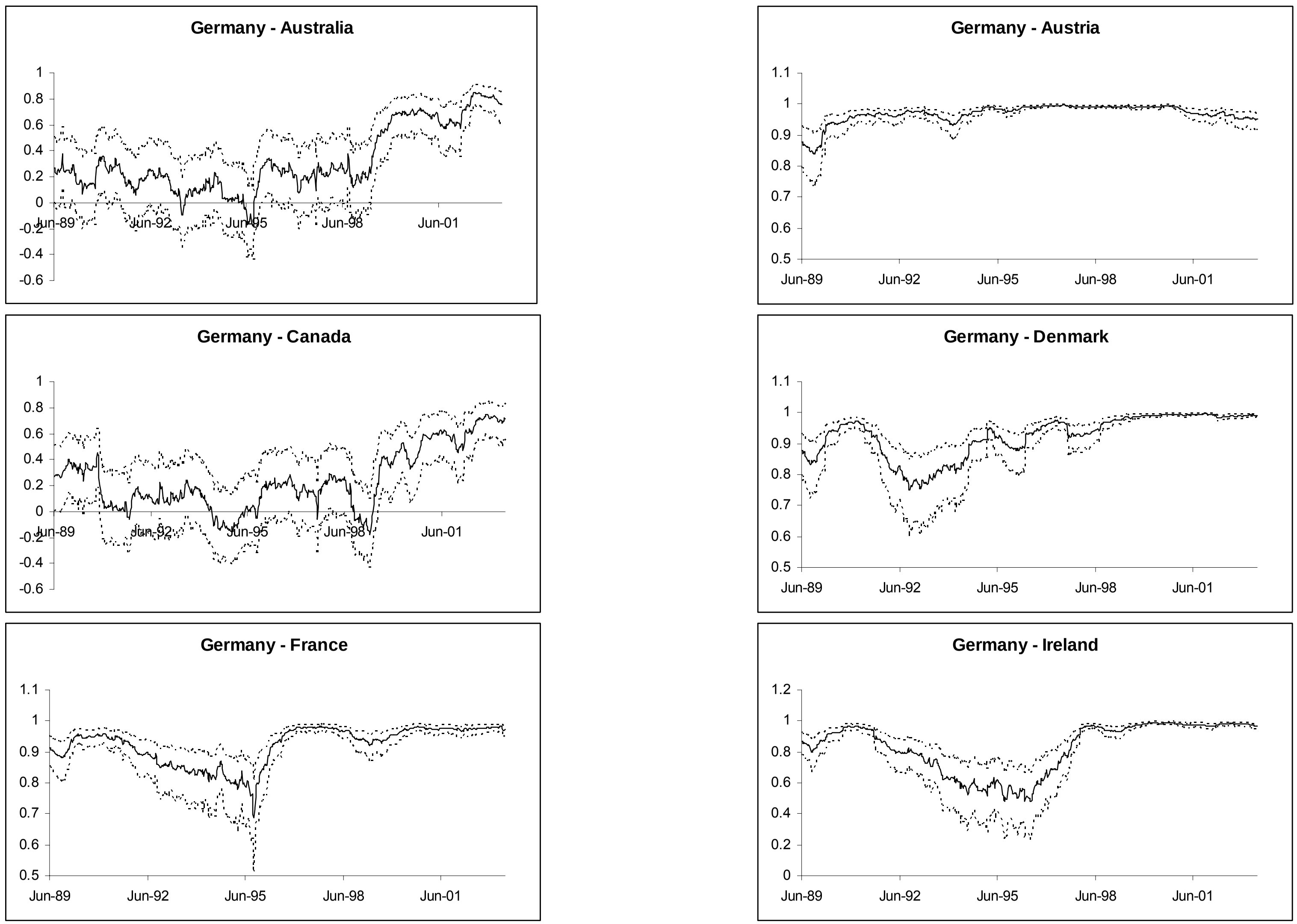
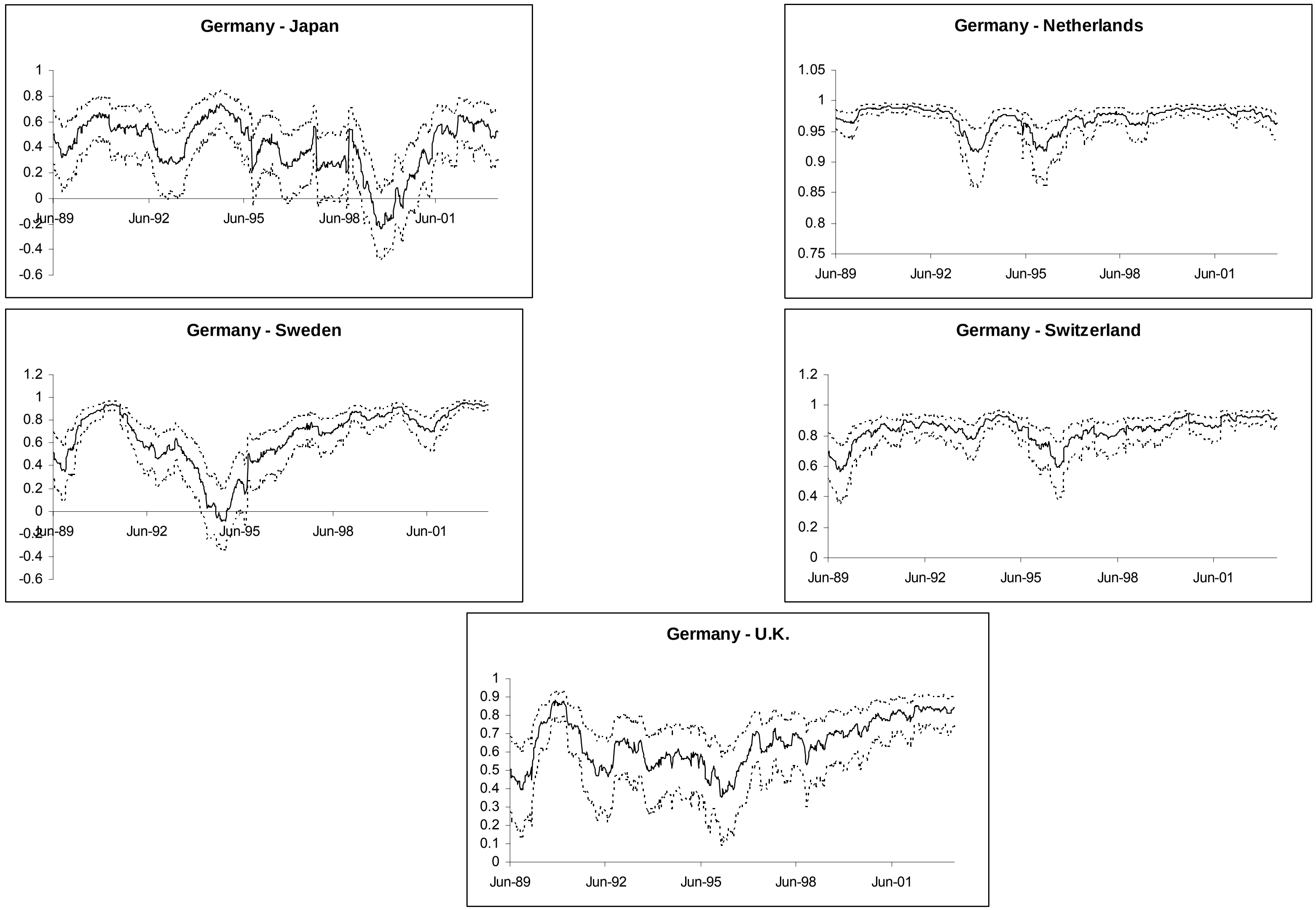


Figure 2.2 (continued)



Note: The solid lines represent the correlations, while the dotted lines represent 95% confidence bounds.

Euro area. Among European bond market returns we consider in this analysis, Austria, France, Germany, Ireland and Netherlands were part of the Euro area with the start of the third stage of the EMU on January 1, 1999. Denmark and the U.K. negotiated an "opt-out" protocol to the EU Treaty that gave them an option of joining or not joining the euro area. Presently these countries are not part of the Euro area. Sweden will join when the necessary conditions imposed by the EMU are met.[13]

Within members of the Euro area, from the mid-1990s onward, there is a clear increase in the correlations with German bond returns that remain close to one for the rest of the period under consideration. The correlations of German bond returns with other European countries' bond returns show a clear positive trend starting in the mid to the late 1990s. The correlations between German bond returns and other non-European countries' bond returns exhibit a positive trend, especially after 1999, indicating an increasing codependence within these international bond markets.

In conclusion, the results presented in the Figures 2.1 and 2.2 indicate an overall increase in the co-dependence among international bond returns especially during the last part of our sample period. As expected, this increase is greater among members of the Euro area. For these countries the rolling correlations indicate that, in the last period of our analysis (from the mid 1990s onward), they have become almost fully financially integrated.

[13] These convergence criteria laid down in the Maastricht Treaty involve restrictions on inflation, interest rates, exchange rates and budget deficits.

2.5 Cointegration Methodology

In order to be able to better understand the dependence across international bond market returns, we conduct a cointegration analysis. Before analyzing the cointegration between international bond returns, in this section we present the cointegration methodology. Granger (1981) first introduced the concept of cointegration into the literature. It is a statistical implication of the existence of a long-run equilibrium between economic variables. A set of variables, each integrated by order one, $I(1)$, is said to be cointegrated if a linear combination is $I(d)$ where d is any number less than one. Most of the economic variables are found be $I(1)$, therefore in the conventional cointegration analysis, a linear combination of $I(1)$ variables is required to be $I(0)$.

Early empirical work applied the Engle and Granger (1987) two-step procedure for modeling the relationship between cointegrated variables. This procedure is as follows: First, using the data in levels or log levels, the long-run relationship (i.e. the linear combination of a set of $I(1)$ variables) is estimated by OLS. This is called the cointegrating regression. Then the stationarity of the residuals from this regression is tested. If these residuals are found to be stationary, then the non-stationary variables are said to be cointegrated.[14]

A shortcoming of this procedure is that it assumes at most one cointegrating vector. Also, the results in Engle and Granger method are potentially dependent on the choice of the dependent variable.

Johansen (1988) and Johansen and Juselius (1990) introduced a maximum likelihood test procedure for cointegration that allows multiple cointegrating vectors in a

[14] McKinnon (1991) critical values are used.

multivariate framework and assumes all the variables to be endogenous. This approach starts with a pth-order VAR model for *Xt*,

$$X_t = A_1 X_{t-1} + \ldots A_p \mathrm{X}_{\mathrm{t-p}} + \mathrm{BZ}_{\mathrm{t}} + \varepsilon_{\mathrm{t}}$$

where *Xt* is a (nx1) vector of non-stationary *I*(1) variables, *Zt* is a (kx1) vector of deterministic variables, and ε_{t} is a (nx1) vector of innovations. Following Johansen (1988), the above equation can be rewritten in an error-correction, or differenced, form

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots \Gamma_{p-1} \Delta X_{t-p} + \Pi X_{t-1} + \Psi Z_t + \varepsilon_t$$

where

$$\Pi = \sum_{i=1}^{p} A_i - I \text{, and} \qquad \Gamma_i = \sum_{j=i+1}^{p} A_j$$

The Π matrix contains information on the long-run relationships. If the rank of Π is r ≤ n-1, then there exist *n*x*r* matrices α and β such that Π = αβ’ and β’Xt is stationary, *I*(0). From the number of cointegrating vectors (r) and the numbers of variables in the system (n) we can infer the number of common stochastic trends driving the system (equal to n-r). When r = 0 there is no cointegration, implying that there are no linear combinations of the X_t that are *I(0)*. When Π has full rank, r = n, all the variables in X_t are stationary in levels, $X_t \sim I(0)$.

In determining the existence of a long-run relationship between a set of n variables it is important to distinguish between the cases when r = n - 1 and 0 < r < n - 1. If there exists r = n – 1 cointegrating relationships then it is said that the n variables are “perfectly” cointegrated. Otherwise if 0 < r < n – 1 then the n variables are said to be “partially” cointegrated. We will comment more on this point as it will be important in our analysis of determining the degree of integration among international bond market returns.

The cointegration results are sensitive to the assumption made with respect to the deterministic components. There are five possible models in the Johansen procedure (Johansen, 1994):

(1) Series X have no deterministic trends and the cointegrating equations (CE) do not have intercepts;

(2) Series X have no deterministic trends and the CE have intercepts;

(3) Series X have a deterministic trends and the CE have intercepts;

(4) Both series X and CE have deterministic trends; and

(5) Series X have quadratic trends and the CE have trends.

It should be noted here that the first model is the most restrictive one and the last model is considered the least restrictive one.

In determining which of the above models better represent the data the following test statistics is proposed by Johansen (1992)

$$-T\sum_{i=r+1}^{n}\left[\ln\left(1-\hat{\lambda}_i^{*}\right)-\ln\left(1-\hat{\lambda}_i\right)\right] \qquad (3)$$

where T is the sample size, $\hat{\lambda}_i^{*}$ and $\hat{\lambda}_i$ are the characteristic roots of the restricted and unrestricted model respectively, and *r* is the number of the nonzero characteristic roots of the unrestricted model. This test statistics has a χ^2 distribution with (n - r) degrees of freedom.[15] Rarely do economic series exhibit a quadratic trend. Therefore, in the empirical applications of the cointegration analysis, only models *i* through *iv* are

[15] See Enders (2004) pp 354–357 for a detailed description of this test.

considered. In the cointegration analysis of the bond returns reported in the next section we use the above test statistics to determine which of the models 1 through 4 better represents the data.

2.6 Cointegration across International Bond Markets

In this section, we look at the level of cointegration among international bond markets. First, the cointegration of the U.S. bond market with other international bond markets is analyzed. The results are presented in Table 2.6. The Johansen (1988) method was used to determine the rank order of the cointegration. The lag length is selected using Sims' (1980) likelihood ratio (LR) corrected for small samples

$$LR = (T-k) * \left(\log|\Omega_r| - \log|\Omega_u|\right)$$

where T is the sample size, k is the number of coefficients in each equation, $\log|\Omega_r|$ is the log determinant of the residual covariance matrix when the model is restricted and $\log|\Omega_u|$ is the log determinant of the residual covariance matrix when the model is unrestricted.[16] During the sample period considered in this analysis, several important events have occurred. In the previous section, we found that January 1, 1994 and January 1, 1999 are two important dates that have had a significant impact on the international

[16]The optimal lag length was selected using data in first differences, with a lagged level term in the model. The maximum lag length tested for each model is 14. In addition to this LR test, we get the optimal lag length suggested by AIC and SIC. For the bivariate systems the AIC suggests the same optimal lag length as Sims' (1980) LR test, while the SIC tends to suggest a lower optimal lag length. The cointegration results are robust, and do not depend on the lag length chosen. When we test for the presence of cointegration in a system with more than two returns both AIC and SIC suggest a lower lag length than Sims' (1980) LR test. But the cointegration results are robust to the lag length chosen except for the case when the cointegration between Canadian, the U.K. and the U.S. bond returns is analyzed. Both AIC and SIC suggest an optimal lag length equal to 1, with residuals being white noise. At this lag length we found one cointegration vector, suggesting that these markets are partially cointegrated. When lag length is greater than 1 the results indicate that these markets are not cointegrated.

Table 2.6: Cointegration of U.S. Bond Market Returns with Other International Bond Markets Returns

US cointegration with	Lag	No Time Trend in VAR: No Intercept in CE*	No Time Trend in VAR: Intercept in CE	p-value
Australia	1	r=1	r=1	0.25
Austria	1	r=1	r=0	0.65
Canada	1	r=1	r=0	0.19
Denmark	1	r=1	r=0	0.11
France	1	r=1	r=0	0.60
Germany	1	r=1	r=0	0.89
Ireland	1	r=1	r=0	0.43
Japan	5	r=1	r=0	0.12
Netherlands	1	r=1	r=0	0.76
Sweden	1	r=1	r=0	0.10
Switzerland	1	r=1	r=0	0.59
UK	2	r=1	r=1	0.04
Canada & UK	4	r=0	r=0	0.06
France & Germany	5	r=1	r=1	0.00

Note: This table reports cointegration tests when no trend is included in the VAR. CE stands for cointegrating equation, and r gives the cointegration rank. The lag length is selected using Sims (1980) likelihood ratio (LR) test corrected for small samples. The last column gives the p-value of the LR test statistic for the null hypothesis of H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. The results of the model suggested by this LR test are underlined.

bond market returns we analyze. To account for the impact of these two important dates, we include two dummies in the cointegration analysis.

As was mentioned in the previous section, the cointegration results are sensitive to the assumptions about the constant and/or a drift in the data and the cointegration equation. For this reason the cointegration test is performed under different assumptions. We allow for a time trend, or not, in the data and for an intercept, or not, in the cointegrating equation. For all the systems for which we tested the presence of cointegration the test statistics in (3) rejects the existence of a deterministic trend in the

VAR. Therefore, the cointegration results when a trend is assumed in the VAR are not reported.

Table 2.6 presents the cointegration results between the U.S. bond markets and other international bond markets. The lag length reported is the optimal lag length suggested by Sims' (1980) LR test. At the optimal lag length the residuals are white noise. The last column in Table 2.6 reports the p-value of the test statistic (3) for the hypothesis H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. The results of the model suggested by this test are underlined. Overall the results presented in Table 2.6 show evidence of cointegration of international bond market returns with the U.S. bond market returns, except for the case between the U.S. and Sweden's bond market returns.

It is often believed that Canada, the U.K. and the U.S. are becoming more economically integrated. Therefore, we analyze whether there is evidence of cointegration across these three markets. No evidence of cointegration between these international bond markets is found. In addition, we analyze the cointegration order between the U.S., Germany and France. There is evidence of cointegration between these three bond markets but the cointegration order is equal to one, therefore we cannot conclude that these markets are fully cointegrated.

Table 2.7 presents the cointegration results of German bond market returns with other international bond market returns. Similar to Table 2.6, we present the results only when no trend is assumed in the VAR and the p-value for the test statistic (3) for the hypothesis H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. There results indicate no evidence of cointegration between the German bond

Table 2.7: Cointegration of German Bond Market Returns with Other International Bond Markets Returns

German cointegration with	Lag	No Time Trend in VAR No Intercept in CE*	 Intercept in CE	p-value
Australia	2	r=1	<u>r=1</u>	0.06
Austria	2	r=1	<u>r=2</u>	0.01
Canada	2	<u>r=0</u>	r=0	0.26
Denmark	4	r=1	<u>r=1</u>	0.01
France	1	r=1	<u>r=1</u>	0.05
Ireland	2	<u>r=1</u>	r=0	0.19
Japan	2	<u>r=0</u>	r=0	0.17
Netherlands	2	r=1	<u>r=1</u>	0.03
Sweden	1	r=0	<u>r=0</u>	0.09
Switzerland	2	r=0	<u>r=0</u>	0.03
UK	2	<u>r=1</u>	r=0	0.22
US	1	<u>r=1</u>	r=0	0.89
Austria & France & Netherlands	7	r=3	<u>r=3</u>	0.00
Austria & France & Netherlands & Ireland	8	r=3	<u>r=3</u>	0.00
France & UK	4	<u>r=1</u>	r=2	0.33
France & Ireland & Denmark & Netherlands	8	r=2	<u>r=1</u>	0.00
France & Ireland & Denmark & Netherlands & UK	8	r=2	<u>r=2</u>	0.00
France & Ireland & Denmark & Netherlands & UK & Sweden	8	r=1	<u>r=1</u>	0.00

Note: This table reports cointegration tests when no trend is included in the VAR. CE stands for cointegrating equation, and r gives the cointegration rank. The lag length is selected using Sims (1980) likelihood ratio (LR) test corrected for small samples. The last column gives the p-value of the LR test statistic for the null hypothesis of H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. The results of the model suggested by this LR test are underlined.

market and bond markets of Canada, Japan, Sweden and Switzerland. Evidence of cointegration is found between the German bond market and bond markets of Austria, France and Netherlands. Also when these four countries are considered together in a system of 4, we find 3 cointegrating vectors. This indicates that these international markets are fully cointegrated.

We test whether Germany, France and U.K. bond markets are fully cointegrated. The results, presented in Table 2.7, show some evidence of the cointegration between these countries. If a constant in the cointegrating equation is assumed, the Johansen (1988) cointegration test indicates that in a system of 3 bond returns there are 2 cointegrating vectors, indicating that these markets are fully cointegrated. However, the test statistic indicated in equation (3) suggests no constant in the CE; therefore, we conclude that these markets are not fully cointegrated. In addition, we analyze whether the bond markets across European countries are fully cointegrated during the period covered in this paper. Only one cointegrating vector is found in a system of 5 countries' bond returns (Germany, France, Ireland, Denmark and Netherlands). This indicates partial cointegration across these markets. The results do not change when the U.K. and Sweden's bond markets are added to the system.

In conclusion, over the time period considered in this analysis, the results presented in Tables 2.6 and 2.7 indicate partial integration across international bond markets. There is evidence of the cointegration between U.S. and all other international bond markets. Different from what may have become a common belief, when we considered the U.S., Canada and the U.K. bond markets as a single system, the results do not indicate full cointegration across these markets. German, French, Austrian and Dutch bond markets are fully integrated. On the other hand, when German, French and the U.K. bond returns are analyzed together we fail to find evidence of full cointegration.

Recall that the moving correlations across international bond markets, analyzed in Section 2.4, indicated a stronger correlation, especially starting from the mid-1990s. Therefore analyzing the cointegration level between these markets from this point on may be of interest in order to determine the actual degree of integration across these markets. However, the cointegration analysis looks for the existence, or not, of a long-run equilibrium between these markets. What is important is not the frequency of the data but the time span considered. Analyzing a short span of bond returns would not be sufficient to determine whether the long-run relation between international bond markets has changed significantly from what we found during for the period 1989 to 2004 (see Hakkio and Rush, 1991). It will be the focus of future research to determine how the cointegration relationship across international bond markets has changed, especially after January 1, 1999.

2.7 Business Cycles Synchronization and Bond Returns

The analysis of rolling correlations across international bond markets indicated that the dependence structure among them has increased over time, especially within the Euro area countries. This dependence has increased considerably, especially during the last 6–7 years of our sample. As a check of our results, we analyze how this increase in dependence structure across international bond markets is reflected in business cycles. In particular, we analyze whether the business cycles across countries have become more synchronized and how that is related with the dynamics observed among international bond markets.

In general, an increase in the capital mobility across countries is likely to lead to a higher interdependence of business cycles across these countries, i.e. more synchronized business cycles. However, an increase in the integration level across these countries could alternatively, lead to more specialized production, and thus less output synchronization

(Krugman, 1993). Different methodologies have been used to determine the degree of synchronization in business cycles across countries.[17] One is the concordance correlation coefficient, proposed by Harding and Pagan (2002). It measures the number of periods during which national cycles are in the same phase. Analysis of the common factors explaining the business cycle of a particular country is another measure used to analyze synchronization. When using this methodology, one is trying to determine the importance of global versus local shocks on the output production of a particular country with shocks originating in the U.S. economy considered as a good proxy for world shocks. This methodology often involves the analysis of the importance of common factors in explaining the output volatility of a country. A third measure used in analyzing business cycle synchronization, and also the most commonly used one, involves output correlations across different countries. The higher the correlations the more synchronized are the business cycles across countries.

Following this literature, correlations across growth rates of output (that is the first difference of log levels) are used to determine the degree of business cycle synchronization across countries. In addition, we conduct a cointegration analysis across international industrial production to examine the degree of business cycle synchronization (see Table 2.8). The data used consists of seasonally adjusted monthly industrial production indexes taken from the IMF's International Financial Statistics covering the period June 1989 to April 2004.[18] For each series, an ARMA process is fitted in order to remove any possible autocorrelation.

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[17] See Bordo and Helbling (2003) for a more detailed explanation of different methodologies used to measure business cycles synchronization.

[18] The industrial production series for Australia and Switzerland are not available at monthly frequencies for this period; therefore, these two countries are excluded from the analysis.

Table 2.8: Unit Root Tests for Industrial Production in Log Levels

		ADF							KPSS
		Lag	No Const.	Const.	Const &Trend	φ_1	φ_2	φ_3	H_0: Level Stationarity
Austria		2	4.05	0.30	-1.41	0.01	0.81	1.21	4.43*
Canada		3	2.15	0.02	-2.87	0.00	2.95	4.43	4.47*
Denmark		5	0.70	-0.85	-3.19[c]	0.40	3.43	4.72	4.41*
France	(AIC)	4	1.80	-0.27	-1.76	0.05	1.12	1.64	4.12*
	(SIC)	1	1.97	-0.23	-1.71	0.03	1.08	1.58	
Germany		1	1.26	-0.70	-1.49	0.26	0.81	0.96	2.93*
Ireland		2	4.53	-0.03	-2.49	0.19	2.24	3.16	4.61*
Japan	(AIC)	3	0.39	-2.17	-2.17	2.36	1.62	0.08	0.20*
	(SIC)	1	0.56	-1.65	-1.65	1.36	0.95	0.06	
Netherlands		3	1.51	-1.42	-2.57	1.04	2.36	2.48	4.31*
Sweden		1	0.93	-1.13	-2.83	0.66	2.78	3.49	3.58*
UK		2	1.21	-0.84	-0.77	0.36	0.29	0.08	3.72*
US		3	2.60	-0.71	-1.31	0.33	0.68	0.69	4.48*

Note: For ADF unit root test, the null hypothesis is the existence of unit root. The lag length for this test was selected using AIC and SIC criteria. For IP of France and Japan, these two criteria suggest different optimal lag length. We report the results for both these lag lengths and we cannot reject the null hypothesis of the existence of unit root at both these lag lengths. The critical values at 5% significance level for the test statistics φ_1, φ_2, and φ_3 are 4.71, 4.88 and 6.49 respectively. For the KPSS test under the null the series are assumed to be local stationary. The lag length for KPSS test is equal to 5 and was selected using the level selection procedure proposed by Newey and West (1994). * shows significancy at the 1% level.

We first analyze the correlation across industrial production growth. Figure 2.3 presents the correlations between U.S. industrial production (IP) and the IP of other countries. Using rolling correlation techniques we found an increase in bond market dependence, especially beginning in the mid 1990s. Therefore, we would expect higher synchronization of business cycles between these countries. Positive correlations between output growth series will indicate synchronization of business cycles and as these positive correlations increase, we would conclude that the synchronization has increased.

While we observe that correlations are sometimes positive, the results presented in Figure 2.3 do not indicate that there is any clear positive trend in the correlations for industrial production growth with respect to the U.S. What is more important, for almost all the sample period considered these correlations are not significantly different from zero at 5% level of significance. These results indicate the lack of business cycle synchronization.

Figure 2.4 presents the correlations of German output growth with output growth of other countries. Similar to the results presented in Figure 2.3, for almost all our sample period the correlations are not significantly different from zero. Only for a short period, 1995–1996, the correlations of Germany's IP with the IP of France, Sweden and U.K. appear to be significantly positive. Sweden join the EU in 1995, therefore this temporary increase in the correlation of its IP with Germany's might be a reflection of this important event. This period also corresponds with important events in the integration process of the EU that laid down the transition to a single currency (for example the European Council in Cannes). In summary, we conclude that the output correlations do not give evidence of the business cycle synchronization across countries.

Figure 2.3: Correlations of U.S. Industrial Production (IP) With Other Countries' IP

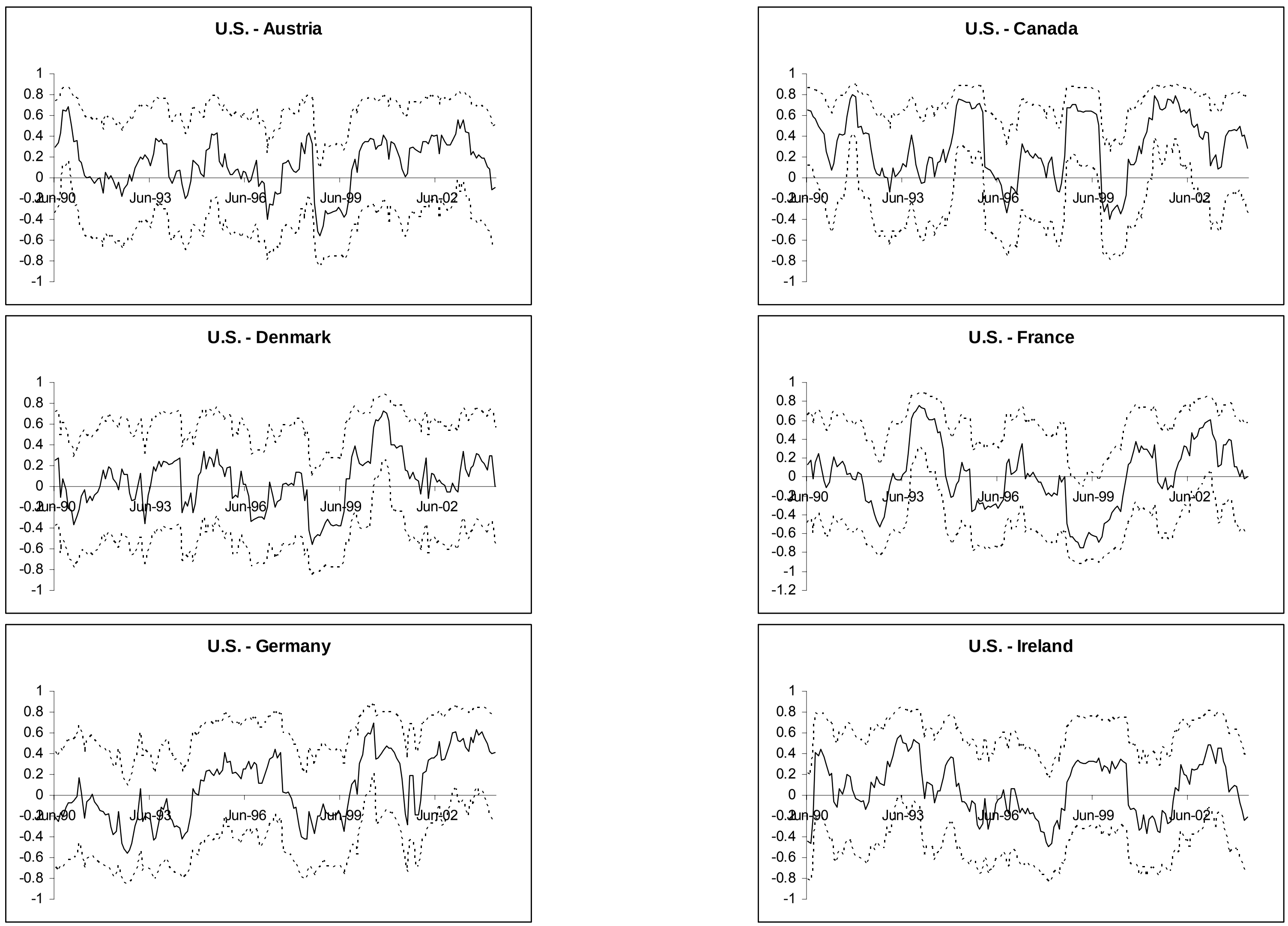
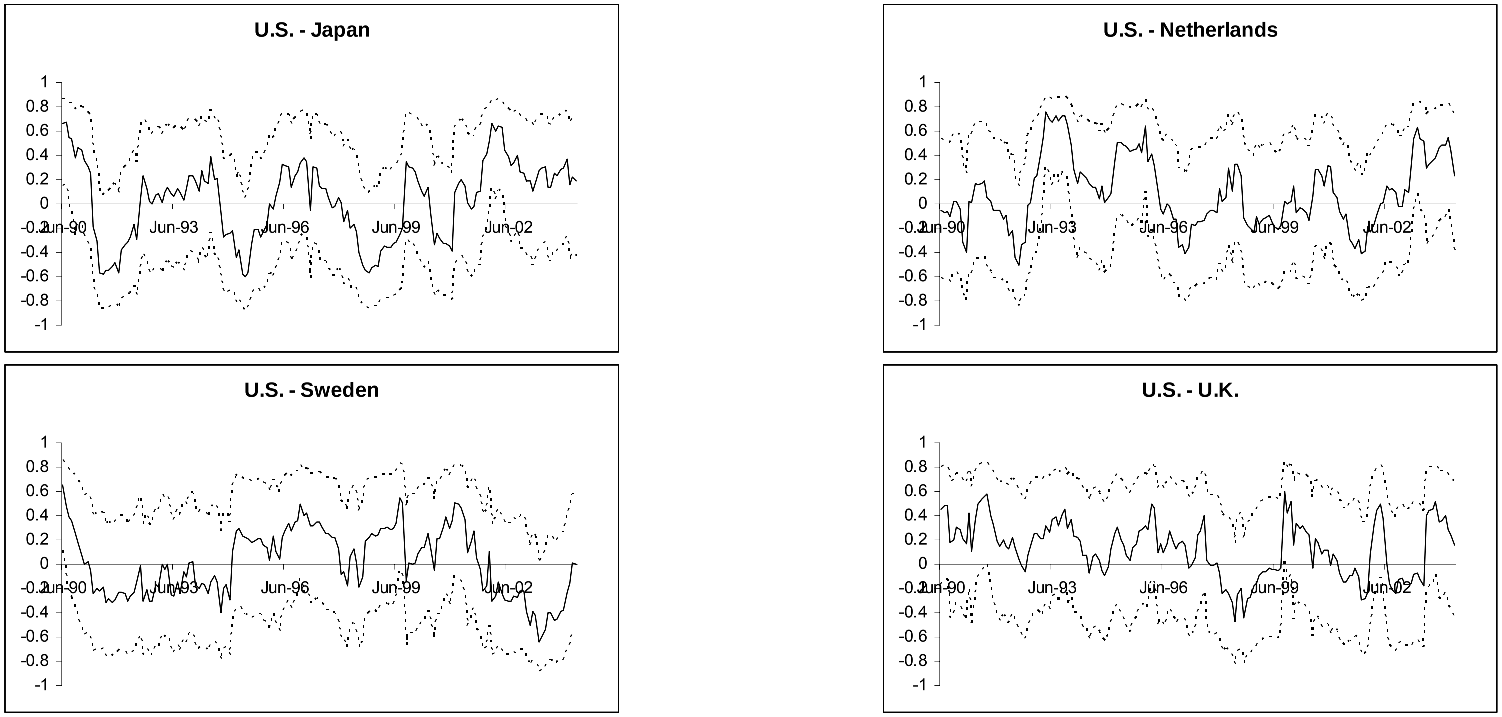


Figure 2.3: (continued)



Note: The solid lines represent the correlations, while the dotted lines represent 95% confidence bounds.

Figure 2.4: Correlations of the German Industrial Production (IP) With Other Countries' IP

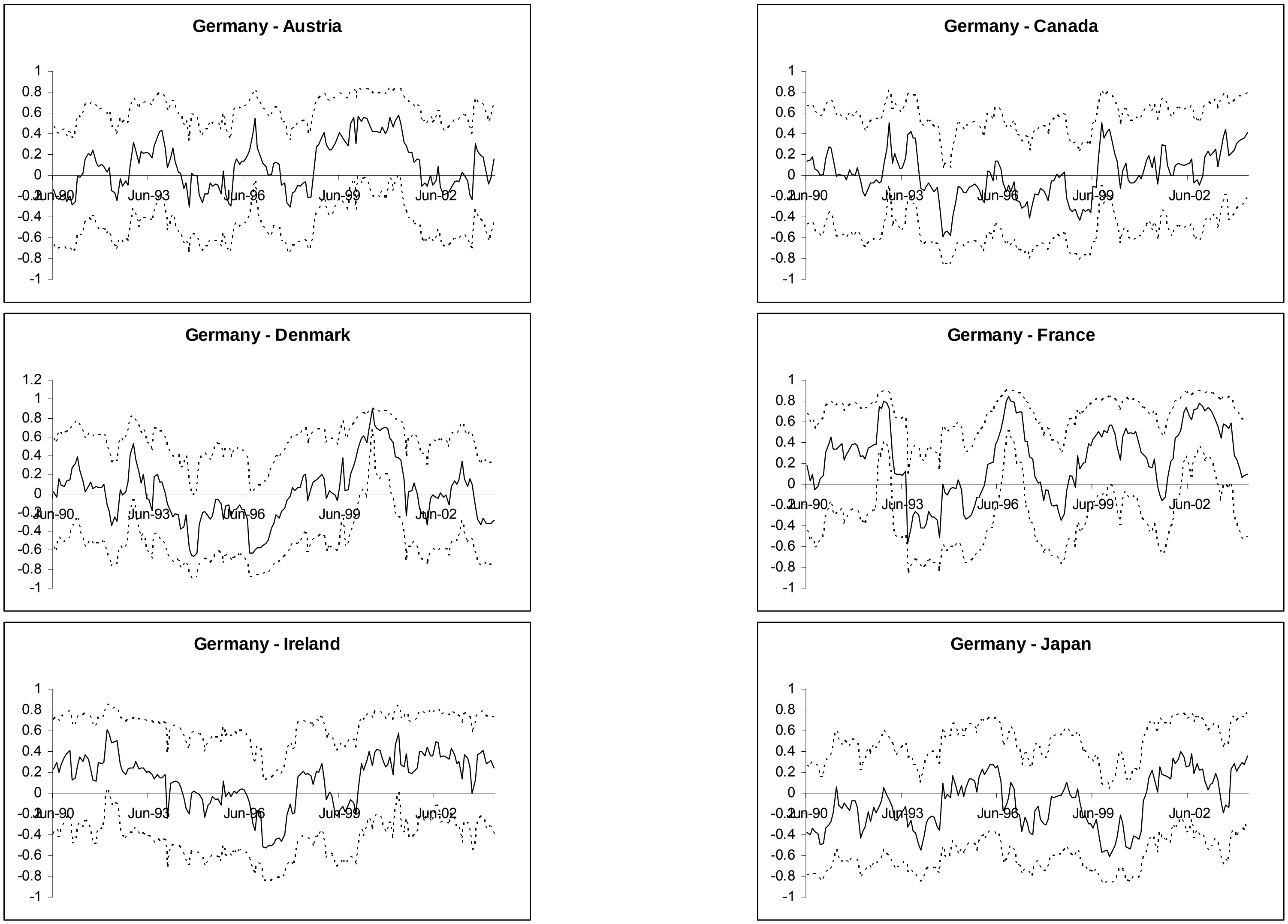
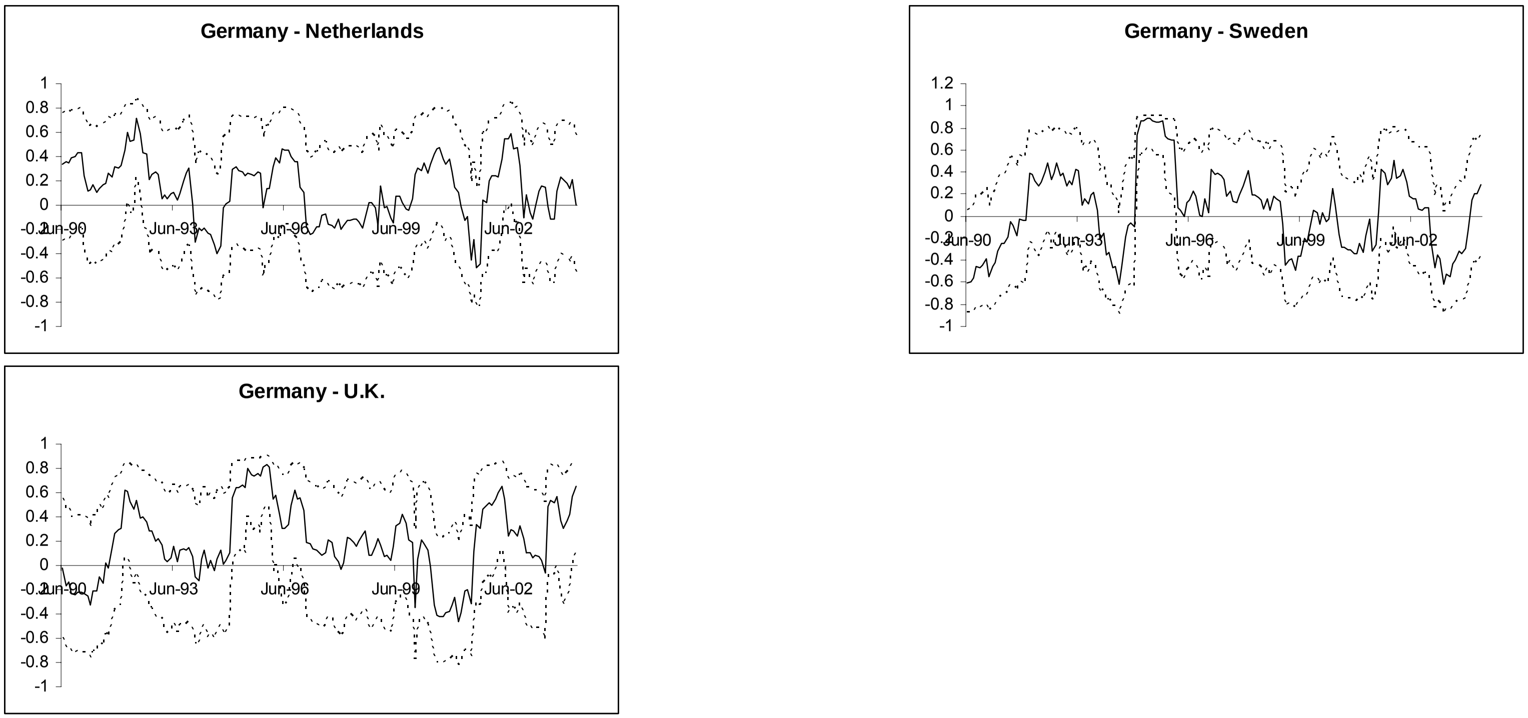


Figure 2.4: (continued)



Note: The solid lines represent the correlations, while the dotted lines represent 95% confidence bounds.

The final analysis of this section consists of the examination of the cointegration of industrial production series across countries.[19] These results are reported in Table 2.9. The first part of this table presents the cointegration results between Germany's IP and the IP of other countries. There is a lack of cointegration between the IP of Germany, Austria, Denmark, France, Netherlands and the U.K. This indicates that there is no evidence of business cycles synchronization across these countries. The German business cycles appear to be synchronized only with the business cycles of Canada, Ireland, Sweden and U.S.

The second part of Table 2.9 presents the cointegration results between the U.S.'s IP and the IP of other countries. The cointegration results between the U.S and Canadian industrial production are interesting. With an increase in trade between these two countries, we would expect higher output synchronization. However, the results presented in Table 2.9 indicate lack of cointegration between the IP series of these two countries. We also find no evidence of cointegration between the IP series of the U.S. and the U.K.

Overall, the cointegration results presented in Table 2.9 are consistent with the results obtained from the examination of the correlations across international industrial production growth. There is lack of evidence of business cycles synchronization between the countries considered in our analysis. Note that the economic theory suggests that an improvement in capital mobility would increase financial integration among countries, which then would lead to risk diversification and consumption smoothing. This would

[19] We follow the same procedure as the one we used to test for cointegration across international bond market returns. The optimal lag length was determined using Sims' (1980) LR test corrected for small samples. At each optimal lag length the VAR residuals are white noise.

Table 2.9: Cointegration of Industrial Production (IP)

		No Time Trend in VAR		
German IP cointegration with	Lag	No Intercept in CE	Intercept in CE	p-value
Austria	2	r=1	$\underline{r=0}$	0.06
Canada	2	$\underline{r=1}$	r=1	0.90
Denmark	4	$\underline{r=0}$	r=0	0.23
France	3	r=0	$\underline{r=0}$	0.08
Ireland	2	$\underline{r=1}$	r=1	0.18
Netherlands	3	r=0	$\underline{r=0}$	0.09
Sweden	1	$\underline{r=1}$	r=0	0.19
U.K.	2	r=1	$\underline{r=0}$	0.01
U.S.	2	$\underline{r=1}$	r=1	0.84

		No Time Trend in VAR		
U.S. IP cointegration with	Lag	No Intercept in CE	Intercept in CE	p-value
Austria	2	r=1	$\underline{r=1}$	0.10
Canada	2	$\underline{r=2}$	r=1	0.26
Denmark	2	$\underline{r=2}$	r=1	0.20
France	2	$\underline{r=1}$	r=1	0.60
Germany	2	$\underline{r=1}$	r=1	0.84
Ireland	2	$\underline{r=1}$	r=1	0.17
Netherlands	3	r=0	$\underline{r=1}$	0.02
Sweden	2	$\underline{r=1}$	r=1	0.31
U.K.	2	r=1	$\underline{r=0}$	0.11

Note: The above table reports cointegration tests when no trend is included in the VAR. CE stands for cointegrating equation, and r gives the cointegration rank. The lag length is selected using Sims (1980) likelihood ratio corrected for small samples. The last column gives the p-value of the LR test statistic for the null hypothesis of H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. The results of the model suggested by this LR test are underlined.

improve specialization in production and capital allocation, and therefore, leads to more economic growth (see Obstefeld, 1994 and Acemoglu and Zilibotti, 1997). If the process of financial integration were associated with intra-industry specialization across countries, and therefore a larger trade volume of intermediate inputs, then we would expect more synchronization of the business cycles across countries. However, if the increase in the financial integration is associated with a higher level of inter-industry specialization, then the production structure may actually become more vulnerable to idiosyncratic shocks (see Krugman, 1993; Bekaert, Harvey and Lundblad, 2001). This might lead to lower business synchronization across countries.

Although output specialization might be a plausible explanation for the lack of evidence of business synchronization, there is skepticism that this might be an explanation of the results. Note that the output of the world's largest economies such as U.S, U.K., Germany, and France are analyzed. Therefore, another explanation is needed for the inconsistencies found in analyzing correlations and cointegration across international bond market returns versus the business cycle synchronizations results. Perhaps a deeper look is needed to find a more appropriate or a stronger measure of IFI. In the next chapters, alternatives are provided to measuring IFI across international bond market returns.

2.8 Conclusions

This chapter constitutes the first of three essays where the dependence structure among international bond markets is examined. Whether the same asymmetric correlations observed across international stock market returns are also observed across international bond market returns is tested. The analysis indicates that the U.S. bond market correlations with other bond market returns appear to be higher during extreme negative returns relative to the correlations during extreme positive returns. However, there appear to be no asymmetry in the correlations of German bond market returns with other European bond market returns, while there is still evidence of the asymmetry between the correlations of German bond market returns with other non-European bond market returns. This appears to indicate a different dependence structure between European bond markets relative to the other international bond markets.

By looking at rolling correlation over time an overall increase is found in the co-dependence between international bond returns especially during the latter part of the sample period, that is, beginning in the mid-1990s. As expected, this increase is more significant between countries that are members of the Euro area. The cointegration analysis across these markets indicates weak evidence of cointegration between U.S. and other international bond markets. When considered in a system, the U.S., Canadian and U.K. bond markets do not show evidence of cointegration. Across European bond markets the cointegration results are stronger. German, French, Austrian and Dutch bond markets are fully integrated. Similar results are found when the bond markets of Germany, France and the U.K. were analyzed as a system.

In addition, the business cycles across these countries are analyzed. Strong evidence of business cycles synchronization is not observed between these countries. Hence, while some measures show evidence of the existence of IFI, others do not. To better analyze the inconsistencies provided by the analysis of business cycle synchronizations relative to the analysis of the international bond market rolling correlations and cointegration, in the next chapters alternate approaches are provided to measuring the degree of integration across international bond market returns. These alternatives will enable us to better understand these inconsistencies.

CHAPTER 3

DO INTERNATIONAL BOND MARKETS SHARE A COMMON VOLATILITY PROCESS?

3.1 Introduction

In the previous chapter the rolling correlations and cointegration across international bond market returns was analyzed and strong evidence of financial integration across many of these markets was found. On the other hand, when business cycles were examined they did not appear to be synchronized, indicating lack of financial integration across these countries. In an attempt to lend credibility to one result versus another, in this chapter another approach is employed to measure the degree of financial integration across international bond market returns. In particular, in this chapter, the level of integration across the international bond markets is investigated by testing for the presence of a common volatility process across these markets.

In order to determine whether international bond market returns share the same volatility process, the common feature methodology developed by Engle and Kozicki (1993) is used. This methodology tests whether a feature that is present in a variable is also present in a group of variables. For example, when two series individually exhibit an ARCH effect and a linear combination of the series does not exhibit an ARCH effect, then it is said that these two series share the same volatility process. This approach is similar to the cointegration analysis across international bond market returns that we conducted in the previous chapter. Johansen's (1988) cointegration methodology, presented in the previous chapter, tests for the existence of a common trend in the first moments of the international bond returns, while in this chapter Engle and Kozicki's (1993) common feature methodology is used to test whether a common variance is observed across these bond markets.

Testing for a common feature is of interest, as it will provide evidence on whether the bond markets of different countries share the same volatility process. The presence of a common volatility process would indicate that there is integration across the international bond markets. Alexander (1995b) is the only paper that tests for the common feature in the second moments among international bond markets. Using one-year and five-year weekly bond indices for the period June 19, 1987 to April 4, 1993, she found weak evidence of the existence of a common volatility process among the international bond indices of Canada, France, Germany, Japan, Holland, the U.K. and the U.S. Her results from the U.K. asset market are particularly interesting. She found that the U.K. appears to have become less dependent on European asset markets beginning in 1992. Instead, U.K. asset market movements seem to follow U.S. asset market movements.

Analyzing the existence of common volatility among financial markets is not new in the empirical literature. There is a considerable amount of research that tests for volatility across financial markets. This research focuses mainly on stock and exchange rate markets. In the case of stock markets, Black (1976) and Engle and Susmel (1993) have looked for evidence of a common volatility process across stocks within a market as well as across international equity markets, while Bollerslev and Melvin (1994), and

King, Sentana and Wadhwani (1994) have examined common volatility across foreign exchange rate markets.

Engle and Susmel (1993) analyzed the relationship among 18 major international stock markets during the 1980s. Using the common feature model developed by Engle and Kozicki (1993), they test whether these equity markets share the same volatility process. The existence of a common feature across the stock market volatility is an indication of the integration of these markets. Their results indicated that there is regional integration across some international equity markets.

King, Sentana and Wadhwani (1994) used an arbitrage pricing theory (APT) model to test for the impact of observable and unobservable common factors on the covariance of 16 international equity markets during the period 1970:01 to 1988:10. Their results indicated that idiosyncratic shocks have a significant impact on each of these markets suggesting a relatively low level of financial integration across these markets. Alexander (1995a) used the common feature methodology in the foreign exchange market to test whether intra-currency variability is dominated by regional factors, global factors or speculative investments for the period 1982–1992. She did not find strong evidence of the existence of a common global factor.

Knif and Pynnonen (1998) test for the existence of common volatility among the stock market returns of Asian-Pacific, European and North America countries for the period September 1991 to November 1997. In contrast to Engle and Susmel (1993), they did not find evidence of regional common factors. In fact, their results indicate that small markets are sensitive to world factors, represented by the U.S. market. Farrell (2001), considering various Southern African exchange rates, used the Engle and Kozincki

(1993) methodology to test for common volatility among these exchange rates. He found no evidence of common volatility.

3.2 Common Features Methodology

Engle and Kozicki (1993) developed the common feature model. More specifically, a feature is said to be common if each individual series has this feature but a linear combination of these series does not. In practice, serial correlation, trends, heteroskedasticity, skewness, kurtosis, ARCH or seasonality are some of the possible features displayed in a series. Testing for common features involves two steps. First, we test each series to determine whether the feature is present. Second, if the feature is detected in each individual series, we test whether a linear combination of these series exhibits the feature. The null hypothesis is that the feature is common and the alternative is that it is not.

In this chapter, we test whether two international bond market indexes share the same volatility process using the common volatility approach. Thus, the common feature of interest is the autoregressive conditional heteroskedasticity (ARCH) effect. Consider two stationary time series, which in this case would be two bond returns, x and y (Engle and Susmel, 1993; Alexander, 1995). Each series individually exhibits an ARCH effect

$$x_t = v_t + e_{xt} \quad \text{and} \quad y_t = w_t + e_{yt}$$

where,

$$v_t \mid I_t \sim D\left(0, h_t^2\right) \quad \text{and} \quad w_t \mid I_t \sim D\left(0, k_t^2\right)$$

h_t^2 and k_t^2 are time varying and follow an ARCH process. I_t is the information set on which agents condition their decision at time t, and e_x and e_y are mutually independent homoscedastic error terms. The common feature effect asks whether there is a linear combination of *x* and *y* that exhibits no ARCH effect. Therefore

$$Var_t(x_t + \rho y_t) = h_t^2 + \rho^2 k_t^2 + 2\rho \operatorname{cov}_t(v_t, w_t) + \text{constant}$$

The variance of the linear combination of *x* and $y, Var_t(x_t + \rho y_t)$, will be constant (independent of time) if and only if

$$v_t = -\rho w_t + \text{constant} \qquad (1)$$

If (1) is true then $h_t^2 = \rho^2 k_t^2$, and $\operatorname{cov}_t(v_t, w_t) = -\rho k_t^2$. Then the series can be written with the common ARCH factor w_t as $x_t = -\rho w_t + e_{xt}$ and $y_t = w_t + e_{yt}$ which implies that $Var_t(x_t + \rho y_t)$ is constant. In this case the ARCH effect is a common feature for both series, *x* and *y* share the same volatility process.

The first step in testing for a common feature is to test each series, *x* and *y*, for the presence of an ARCH effect (Engle and Kozicki, 1993). Engle's (1982) Lagrange multiplier (LM) test for ARCH effects in the residuals is used. To test the null hypothesis of no ARCH effect, the squared residuals are regressed on a constant and lagged squared residuals up to order *q*.[20] The ARCH LM test statistic is asymptotically distributed as $\chi^2(q)$.

If an ARCH effect is found in both series, then we test whether there is a linear combination $u_t(\rho) = x_t + \rho y_t$ that does not exhibit an ARCH effect. The test calls for finding the parameter ρ that minimizes the TR^2 from the regression of u_t^2 on lagged values of y_t^2, x_t^2 and cross products of y_t and x_t,

[20] In order to test for the presence of ARCH(1) the lag order of the LM ARCH test will be equal to 1, for ARCH(2) lag order would be equal to 2, and so on. To be able to capture the GARCH behavior more lags should be included in the LM ARCH test.

$$u_t^2 = \sum_{j=1}^{\gamma}\left(\alpha_{1,j} y_{t-j}^2 + \alpha_{2,j} x_{t-j}^2 + \alpha_{3,j} y_{t-j} x_{t-j}\right) \qquad (2)$$

where γ is the lag order and T is the sample size.[21]

Engle and Kozicki (1993) showed that this test statistic, TR^2, is asymptotically distributed as a $\chi^2_{(d)}$, with the degrees of freedom, d, equal to the number of overidentifying restrictions. Note that the null hypothesis is H_0: the ARCH effect is a common feature, i.e. the linear combination of *x* and *y* does not exhibit an ARCH effect, versus the alternative H_1: no common feature. Engle and Susmel (1993) used a grid search to find the value of ρ that minimizes TR^2.

3.3 Data Analysis

The data consists of 10-year DataStream Benchmark Bond indexes measured in U.S. dollars for 13 international bond markets. The bond markets are located in Australia, Austria, Canada, Denmark, France, Germany, Ireland, Japan, the Netherlands, Sweden, Switzerland, the U.K. and the U.S. These bond indexes are available on daily returns and covers the period June 22, 1989 to June 22, 2004. In order to avoid the problem of asynchronous data, daily indexes are converted into weekly frequencies.[22] The weekly bond returns are calculated as log differences using Friday-to-Friday closing prices.

[21] Engle and Susmel (1993) and Alexander (1995a, 1995b) do not give a specific test to determine the lag order γ. In their empirical applications they both consider γ = 4. For the univariate ARCH LM test the lag order of the LM test depends on the order of ARCH process. The higher the order of ARCH the larger the lag order. In their empirical applications, Engle and Susmel (1993) and Alexander (1995a, 1995b), consider γ = 4. In practice a lag order equal to 4 would be sufficient to detect any (G)ARCH effect on the u_t^2. In this chapter this lag order is adopted to detect the presence of an ARCH effect on u_t^2.

[22] The financial markets taken in consideration here do not have the same trading hours. The trading times between these financial markets overlap only partially or they do not have common trading hours. This lack of common trading hours will lead to different problems when these asynchronous data are used in estimation. Martens and Poon (2001) have shown that the use of asynchronous data will result in a

Table 3.1 reports the summary statistics for weekly international bond returns for each country. A more detailed analysis of the results presented in this table is given in the second chapter of this dissertation. The results reported in Table 3.1 indicate evidence of departure from normality for the bond returns of Australia, Ireland, Japan, Sweden, the U.K. and the U.S. Note that the bond returns show evidence of negative skewness, except for the Japanese bond return. The standardized values of the kurtosis coefficient indicate that these returns have thicker tails than in the case of a normal distribution.

Table 3.1: Summary Statistics of Bond Returns

	Mean	Skewness	Kurtosis	Observations
Australia	0.20*	-0.19*	0.74*	782
Austria	0.17*	0.05	0.11	782
Canada	0.15*	-0.09	0.49*	782
Denmark	0.20*	0.00	0.13	782
France	0.18*	0.03	-0.10	782
Germany	0.16*	0.06	0.16	782
Ireland	0.18*	-0.24*	0.39*	782
Japan	0.13**	0.82*	4.29*	782
Netherlands	0.16*	0.06	0.08	782
Sweden	0.18*	-0.28*	0.95*	782
Switzerland	0.15*	0.08	0.29	782
U.K.	0.19*	-0.19*	1.62*	782
U.S.	0.13*	-0.51*	0.73*	782

Note: The mean, skewness and standardized kurtosis are tested against the null of zero; * and ** show the significant levels at 1% and 5% respectively.

The correlation coefficient and the unconditional volatilities of these international bond returns are presented in Table 3.2. The elements of the main diagonal give the variance for each return. The covariances are shown on the lower triangle of the table and the upper triangle gives the unconditional correlations. The unconditional variances for

significant downward bias of the correlations. In order to avoid any problem raised from the use of asynchronous data, the daily indexes are converted into weekly indexes.

Table 3.2: Covariance – Correlations across International Bond Market Returns

	Australia	Austria	Canada	Denmark	France	Germany	Ireland	Japan	Netherlands	Sweden	Switzerland	U.K.	U.S.
Australia	**3.00**	0.39	0.49	0.43	0.39	0.41	0.43	0.20	0.39	0.39	0.30	0.39	0.33
Austria	*1.07*	**2.54**	0.27	0.92	0.91	0.98	0.83	0.36	0.95	0.60	0.85	0.65	0.34
Canada	*1.15*	*0.57*	**1.80**	0.33	0.32	0.29	0.38	0.11	0.34	0.35	0.20	0.40	0.58
Denmark	*1.19*	*2.36*	*0.71*	**2.60**	0.92	0.93	0.88	0.33	0.92	0.70	0.82	0.70	0.37
France	*1.11*	*2.36*	*0.69*	*2.42*	**2.64**	0.93	0.87	0.32	0.94	0.66	0.82	0.70	0.39
Germany	*1.14*	*2.51*	*0.63*	*2.43*	*2.43*	**2.61**	0.85	0.37	0.97	0.62	0.86	0.67	0.38
Ireland	*1.24*	*2.21*	*0.85*	*2.37*	*2.38*	*2.29*	**2.81**	0.25	0.88	0.66	0.75	0.78	0.42
Japan	*0.64*	*1.04*	*0.27*	*0.95*	*0.92*	*1.06*	*0.75*	**3.21**	0.35	0.16	0.37	0.17	0.07
Netherlands	*1.09*	*2.47*	*0.73*	*2.43*	*2.48*	*2.55*	*2.40*	*1.03*	**2.65**	0.64	0.86	0.70	0.42
Sweden	*1.31*	*1.84*	*0.92*	*2.17*	*2.06*	*1.92*	*2.14*	*0.54*	*2.02*	**3.71**	0.51	0.55	0.31
Switzerland	*0.88*	*2.34*	*0.47*	*2.27*	*2.29*	*2.40*	*2.17*	*1.14*	*2.41*	*1.69*	**2.95**	0.58	0.30
U.K.	*1.00*	*1.51*	*0.78*	*1.65*	*1.66*	*1.59*	*1.91*	*0.44*	*1.67*	*1.56*	*1.47*	**2.14**	0.50
U.S.	*0.59*	*0.56*	*0.80*	*0.61*	*0.66*	*0.63*	*0.73*	*0.13*	*0.71*	*0.62*	*0.52*	*0.75*	**1.06**

Note: The lower triangle gives the covariances of international bond market returns and the main diagonal gives their variances. The upper triangle provides correlations across these international bond market returns. Except for the correlation between U.S. and Japanese bond returns, all correlations are significant at 1% level. The correlation between U.S. and Japanese bond returns is significant at 10% level.

the international bond returns appear to vary within a small range. For most of the bond returns, the variance is between 2 and 3 except for the U.S., Japan and Sweden's bond returns. The U.S. bond returns have an unconditional variance smaller than 2 and Japan and Sweden bond returns each have an unconditional variance greater than 3. The correlations between the international bond market returns analyzed here are positive and significantly different from zero, indicating a positive correlation across these markets although a relatively wide range of dependence can be observed.[23]

In order to analyze the existence of a common conditional volatility process across international bond returns, first the conditional volatility needs to be examined for each individual bond series. Table 3.3 presents the results of different univariate ARCH tests. Following Engle and Susmle (1993), the ARCH tests have been calculated using two different information sets. The first two columns of Table 3.3 present the results of the traditional univariate Langrage multiplier (LM) ARCH test. The square of bond returns, r_t^2 are regressed on a constant and its own lagged values.

$$r_t^2 = c + \phi_1 r_{t-1}^2 + \phi_2 r_{t-2}^2 + \ldots + \phi_p r_{t-q}^2 + \varepsilon_t \qquad (3)$$

where q is the lag order used in the univariate ARCH test. The LM test statistic is obtained by multiplying the R^2 with the sample size and has an approximate $\chi^2_{(q)}$ distribution under the assumption that the residuals of the above regressions are white noise. The LM test statistics for the univariate ARCH test presented in the first two columns of Table 3.3 where the lag order is 2 and 4 respectively.[24] The LM test statistics

[23] Except for the correlation between U.S. and Japanese bond returns, the correlation coefficients are significant at the 1% level. The U.S. – Japanese bond returns correlation is significant at the 10% level.

[24] These lag orders will be sufficient to capture the presence of any (G)ARCH behavior in the series. The same lag orders were considered by Engle and Susmel (1993) and Alexander (1995).

Table 3.3: ARCH Test Results

	ARCH(2)	ARCH(4)	MARCH(1) -US	MARCH(2) -US	MARCH(1) -GM	MARCH(2) -GM	MARCH(1) - US-GM	MARCH(2) -US-GM
Australia	56.94^{*}	59.02^{*}	44.82^{*}	58.94^{*}	44.45^{*}	60.26^{*}	45.44^{*}	61.31^{*}
Austria	7.29^{**}	12.26^{*}	2.83	8.98^{***}	12.16^{**}	14.32^{*}	17.31^{*}	14.80^{**}
Canada	…	…	18.02^{*}	18.10^{*}	17.99^{*}	23.96^{*}	18.06^{*}	24.44^{*}
Denmark	12.52^{*}	16.31^{*}	9.69^{*}	11.85^{**}	9.86^{**}	19.68^{*}	13.47^{*}	19.73^{*}
France	14.73^{*}	21.97^{*}	8.70^{**}	15.58^{*}	8.86^{***}	15.99^{*}	11.35^{*}	16.16^{**}
Germany	13.15^{*}	17.93^{*}	6.35^{**}	13.48^{*}	…	…	…	…
Ireland	26.32^{*}	34.43^{*}	19.30^{*}	28.31^{*}	21.23^{*}	32.80^{*}	21.73^{*}	33.55^{*}
Japan	5.02^{***}	13.14^{*}	7.92^{**}	8.12^{***}	5.98^{**}	6.59^{***}	10.09^{**}	9.13
Netherlands	15.55^{*}	24.06^{*}	10.36^{*}	16.82^{*}	18.66^{*}	22.60^{*}	18.94^{*}	23.84^{*}
Sweden	3.51	9.87^{*}	3.11	5.97	4.58	6.52	4.59	8.91
Switzerland	3.79	4.80	0.64	6.07	3.69	7.59	5.43	9.15
UK	11.52^{*}	12.65^{*}	7.33^{**}	11.82^{**}	7.27^{***}	10.92^{**}	9.66^{**}	12.02^{**}
US	5.13^{***}	17.70^{*}	…	…	5.34	9.12	…	…

Note: The univariate ARCH LM tests (Engle, 1982) are reported under the columns ARCH(i). Multivariate ARCH tests with only the U.S. bond returns included as exogenous variables are reported under the columns MARCH(i) – US. Multivariate ARCH tests with only the German bond returns included as exogenous variables are reported under the columns MARCH(i) – GM. Multivariate ARCH tests with both the U.S. and German bond returns included as exogenous variables are reported under the columns MARCH(i) – US – GM. *, ** and *** denote 1%, 5% and 10% significance level respectively. i denotes the lag values $i = 1,2$.

indicate the presence of an ARCH process in each of international bond returns with the exception of Sweden and Switzerland.

The ARCH test results presented in the third through sixth columns of Table 3.3 make use of the multivariate information set. The squared of bond returns ($r_{i,t}^2$) are regressed on a constant, its own lagged values of $r_{i,t}^2$ and also on the squared bond returns of other countries ($r_{j,t}^2$)

$$r_{i,t}^2 = c + \phi_1^i r_{i,t-1}^2 + \phi_2^i r_{i,t-2}^2 + ... + \phi_p^i r_{i,t-q}^2 + \sum_{j=1}^{k} \left(\phi_1^j r_{j,t-1}^2 + \phi_2^j r_{j,t-2}^2 + ... + \phi_p^j r_{j,t-q}^2 \right) + \varepsilon_t \qquad (3)$$

where k is the number of squared bond returns of other countries. The U.S. is used to represent the global bond market. To account for the impact of the U.S. market, the squared bond returns of each country are regressed on a constant, its own lagged values and lagged values of the U.S. squared returns. The results for lagged values equal to 1 and 2 are presented under the columns named MARCH(1)-US and MARCH(2)-US respectively.

The majority of the international bond returns are from European markets, and Germany has often been considered the reference country when analyzing the European economies. The above exercise is repeated substituting the lagged values of the German squared bond returns for lagged values of the U.S. squared returns. The results for lagged values 1 and 2 with German squared returns substituted into equation (3) are presented under the columns MARCH(1)-GM and MARCH(2)-GM respectively. The last two columns of Table 3.3, MARCH(1)-US-GM and MARCH(2)-US-GM, presents the ARCH test with one lag and two lags respectively when both the values for the U.S. and German squared returns are included in Equation (3). Wald test for the German and the U.S. terms in the MARCH tests are provided in Table 3.4.

Using the multivariate information sets, the MARCH test results confirm the test results of the univariate ARCH test. There is strong evidence of the presence of the ARCH disturbance in the international bond returns with the exception of Sweden and Switzerland. As the ARCH test results indicate no evidence of the ARCH disturbance for the bond returns of these two countries, they are not included in the analysis that follows. In order for the ARCH effect to be considered as a common feature, each individual series should exhibit ARCH effects, while their linear combination should not exhibit any ARCH effect. That is, only when the portfolio, which is a linear combination of bond returns, does not follow an ARCH process can we conclude that these series share a common volatility process.

Before testing for the existence of a common ARCH effect, time varying conditional volatilities of individual countries' returns is tested. Table 3.5 reports the estimated coefficients of the univariate GARCH(1,1) models for each bond return. Note that conditional volatilities for Sweden and Switzerland are not displayed since ARCH effects were not detected for either of these series. Different specifications of (G)ARCH models were examined. If the (G)ARCH model is correctly specified, the standardized residuals should be independent, and identically distributed random variables with mean zero and variance one. Based on this analysis GARCH(1,1) best fits the data. This model has the following form,

$$x_t = c + \xi_t$$

Table 3.4: Wald Test of the Exogenous Variables in the Multivariate ARCH Test

	MARCH(1) -GM	MARCH(2) -GM	MARCH(1) -US	MARCH(2) -US	MARCH(1) - US-GM	MARCH(2) -US-GM
Australia	0.069	1.746	0.358	0.717	1.250	0.707
	(0.792)	(0.186)	(0.549)	(0.378)	(0.263)	(0.400)
Austria	23.830	12.724	0.778	0.000	23.609	4.545
	(0.000)	(0.000)	(0.377)	(0.986)	(0.000)	(0.033)
Canada	0.156	4.783	0.195	0.010	0.033	0.322
	(0.692)	(0.029)	(0.658)	(0.920)	(0.855)	(0.576)
Denmark	0.317	9.761	0.021	0.003	5.487	2.199
	(0.573)	(0.001)	(0.884)	(0.953)	(0.019)	(0.138)
France	0.501	1.117	0.092	0.126	5.987	0.128
	(0.479)	(0.290)	(0.761)	(0.721)	(0.014)	(0.720)
Germany	…	…	0.124	0.000	…	…
	…	…	(0.724)	(0.999)	…	…
Ireland	2.682	0.190	0.433	0.018	2.167	0.002
	(0.100)	(0.662)	(0.510)	(0.890)	(0.141)	(0.957)
Japan	1.202	0.962	3.144	1.186	5.638	1.573
	(0.273)	(0.326)	(0.076)	(0.276)	(0.017)	(0.209)
Netherlands	17.833	8.055	0.075	0.014	2.632	4.897
	(0.000)	(0.004)	(0.431)	(0.903)	(0.105)	(0.027)
U.K.	0.013	0.025	0.066	0.697	2.406	0.577
	(0.906)	(0.873)	(0.797)	(0.404)	(0.121)	(0.447)
U.S.	2.464	5.493	…	…	…	…
	(0.116)	(0.019)	…	…	…	…

Note: This table presents the F-statistics for the significance of the exogenous variables in the multivariate ARCH test. The MARCH-GM(i) column gives the F-statistics when only German bond returns are included as exogenous variables in the multivariate ARCH test at the lag level equal to *i*. The MARCH-US(i) column gives the F- statistics when only the U.S. bond returns are included as exogenous variables in the multivariate ARCH test at the lag level equal to *i*. The MARCH-US-GM(i) gives the F- statistics when both the German and U.S. bond returns are included as exogenous variables in the multivariate ARCH test at the lag level equal to *i*. *i* takes value (1,2). The numbers in parenthesis give the p-values.

Table 3.5: Univariate GARCH (1,1)

	α_0	α_1	β_1	Wald Test $H0$: $\alpha_1 + \beta_1 = 1$
Australia	0.37**	0.10*	0.76*	5.64**
	(0.16)	(0.03)	(0.08)	
Austria	0.13***	0.06*	0.89*	2.86***
	(0.08)	(0.02)	(0.05)	
Canada	0.15**	0.07*	0.84*	5.12***
	(0.07)	(0.03)	(0.06)	
Denmark	0.06	0.06*	0.92*	2.53***
	(0.04)	(0.02)	(0.03)	
France	0.09***	0.07*	0.90*	2.81***
	(0.05)	(0.02)	(0.03)	
Germany	0.23***	0.07*	0.85*	3.21***
	(0.13)	(0.03)	(0.07)	
Ireland	0.18***	0.10*	0.84*	2.87***
	(0.11)	(0.04)	(0.07)	
Japan	0.14**	0.08*	0.88*	2.73***
	(0.07)	(0.03)	(0.04)	
Netherlands	0.16***	0.07*	0.87*	3.02***
	(0.09)	(0.03)	(0.05)	
U.K.	0.38	0.09***	0.77*	2.77***
	(0.24)	(0.05)	(0.13)	
U.S.	0.04***	0.04*	0.92*	2.62***
	(0.02)	(0.01)	(0.03)	

Note: *, ** and *** indicate significance at 1%, 5% and 10% level respectively. The numbers in parentheses give the Bollerslev and Wooldrige (1992) standard errors.

where

$$\xi_t \mid I_t \sim N\left(0, h_t^2\right),$$

$$h_t^2 = a_0 + a_1 \xi_{t-1}^2 + \beta_1 h_{t-1}^2, \qquad \text{and} \qquad a_0 > 0, \qquad \mathrm{a}_1, \beta_1 \geq 0.$$

The last column of Table 3.5 gives the Wald test H_0: $\alpha_1 + \beta_1 = 1$ against the alternative of H_1: $\alpha_1 + \beta_1 < 1$. Under the null hypothesis the series x_t is not weakly stationary since it does not have finite variance. The results reported in Table 3.5 indicate that we can reject the null hypothesis of $\alpha_1 + \beta_1 = 1$ indicating that all bond return series considered have finite variances, and therefore are weakly stationary.

The two most important characteristics of the financial time series are the relatively high kurtosis coefficient and volatility clustering, or persistence. The standardized kurtosis coefficients presented in Table 3.1 indicate that the international bond returns exhibit fat tails (i.e. high kurtosis coefficients), while the estimated GARCH(1,1) parameters presented in Table 3.5 are typical parameters obtained when analyzing financial time series. The sum of the estimated GARCH(1,1) parameters α_1 and β_1 are close to 1 for all the international bond returns. The closer the sum of these parameters is to 1, the higher the persistence of these returns. This implies that after a shock the volatility reverts slowly to its long-term mean.

Figure 3.1 displays the estimated time varying conditional volatilities, GARCH(1,1) for each bond return (x_t). There appears to be considerable variation in the volatility of these bond returns over time. The bond returns for Austria, Denmark, France, Germany, the Netherlands and the U.S. appear to be particularly volatile during the last 4–5 years of our sample period. Except for the U.K., there appears to be a sharp increase in the conditional volatility of the bond returns around 1998. The observed increase in the volatility in the international bond market returns may correspond to the Russian bond default in 1998. Different studies have documented a significant increase of

Figure 3.1: Conditional Volatilities of Bond Returns Obtained from the GARCH (1,1) Model

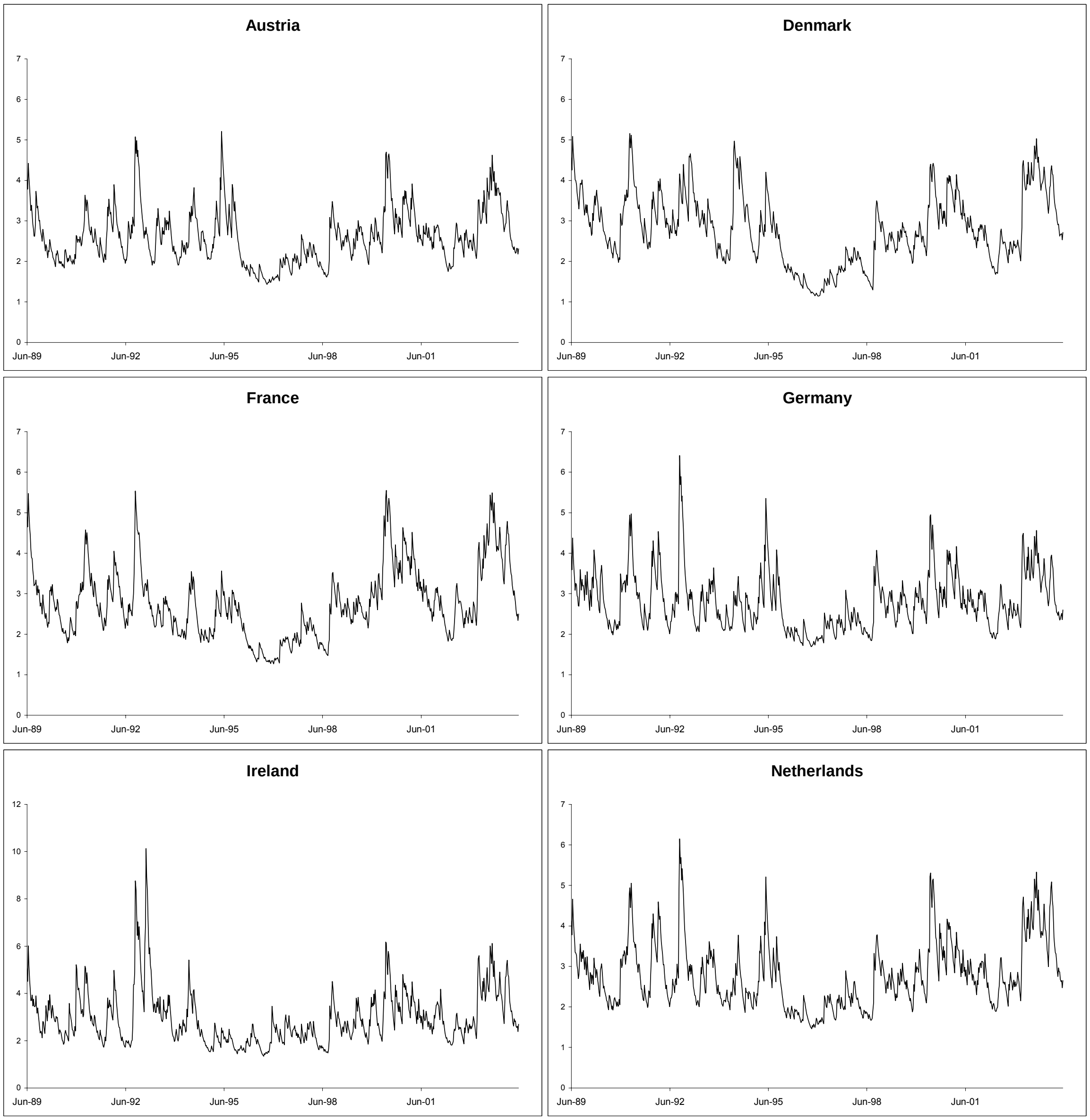
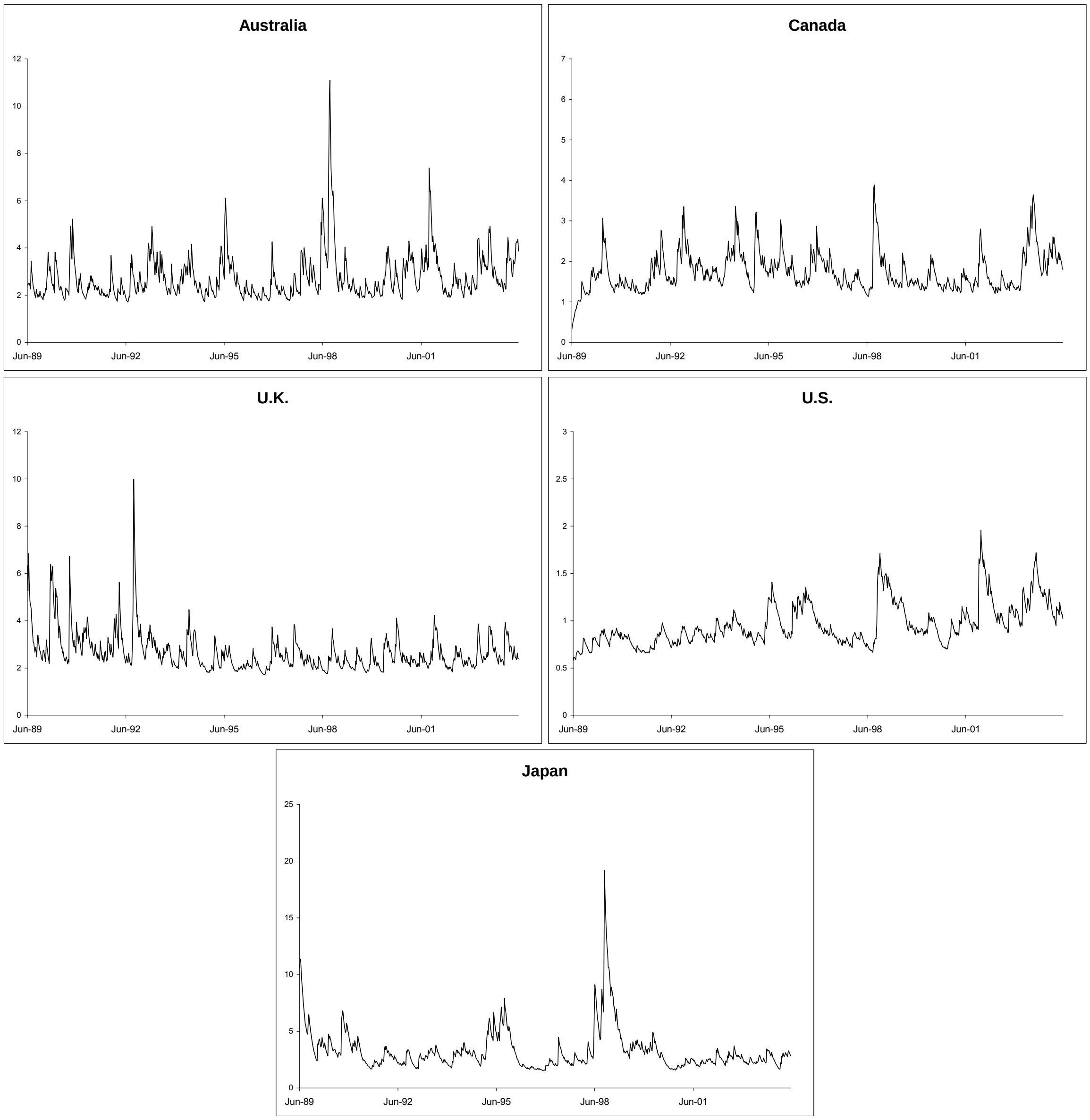


Figure 3.1: (Continued)



international bond market volatility following the Russian bond market crisis.[25] This increase is more evident for the Australian and Japanese markets. The spike observed for

[25] See the survey of the Bank of International Settlements, Committee of the Global Financial System (1999), for an analysis of the impact of the Russian crisis on the international bond markets.

the conditional volatility for Japanese bond market returns has the highest magnitude of any of the bond returns.

The conditional volatility processes presented in Figure 3.1 suggest that the bond returns of Austria, Denmark, France, Germany, Netherlands and the U.S. might share a common volatility process. Given the similarities observed in conditional volatilities, the pattern that emerges is that there is a cluster of high volatility in the first 4–5 years of these series. This is followed by a dip in the volatility in 1997–1998 and a spike in 2003. This pattern appears to be somewhat repeated for the Netherlands and Canada, suggesting that the path of volatility might be common for this group of countries. From the analysis of these graphs, we would expect these countries to exhibit a common feature.

There does not appear to be any similarity between the volatility processes of the U.K. bond market with other European markets. This suggests that the U.K. might not share the same volatility process as these countries. There appears to be some similarity between the volatility process of the U.K., Canada and Australia, although not a very strong one. Note that on October 1997 the U.K. decided not to adopt the single currency on 1 January 1999.[26]

Other empirical papers have also found that while the U.K. financial market is becoming less integrated with European financial markets; it is becoming more integrated with the non-European financial markets. Alexander (1995b) found that, starting in 1992,

[26] The U.K. government's decision to opt for a single currency was made based on five economic tests: (1) the business cycles in the U.K. must be compatible with the Euro area; (2) participation of the U.K. in the single currency area should have a positive effect on its employment and growth; (3) this participation should increase the competitiveness of the U.K. financial services industry with EU country members; (4) participation in the EU must promote investment in the U.K. in the long term; and (5) if problems emerge from the single currency area, the U.K. economy should have sufficient flexibility to deal with them. The analysis of the five economic tests by the U.K. government, reported on 9 June 2003, indicated that the benefits of joining the Euro area are not very clear for the U.K. economy. This decision was left to be reconsidered in the near future.

the U.K. bond market seems to have departed from other European bond markets. Fraser and Oyefeso (2002) showed that the U.K. stock market returns are becoming more sensitive to the shocks originating in the U.S. market relative to those coming from EU countries.

Batavia et al. (2004), using quarterly data for the period 1985–2002, measured the degree of financial integration among the member countries of the European Union and the U.S. Using various measure of financial integration: (a) covered interest rate parity, (b) uncovered interest rate parity, (c) real rate interest rate parity, and (d) the Feldstein–Horioka index, they found that U.K. financial integration has increased more significantly with the U.S. than with the rest of the Europe. These results, and the estimated conditional volatilities presented in Figure 3.1, would suggest that we may not observe evidence of the existence of a common volatility process between the U.K. and other European countries.

3.4 Common ARCH Test Results

The univariate analysis of weekly bond returns indicated the presence of ARCH effects in the bond returns of Australia, Austria, Canada, Denmark, France, Germany, Ireland, Japan, Netherlands, the U.K. and the U.S. In this section, using Engle and Susmel's (1993) common feature methodology, we look at each pair of these bond returns and test to determine which of these portfolios display an ARCH effect. If the linear combination between two bond returns, x and y, $u_t(\rho) = x_t + \rho y_t$, does not display ARCH effect, then we can say that they share a common volatility process. Common feature results are reported in Tables 3.6 and 3.7. Table 3.6 reports the results for those pairs that show no ARCH effect. Therefore, they share a common volatility process. The test results for pairs not having common volatility process are reported in Table 3.7. The column under ρ reports the parameter that minimizes the TR^2 from the regression of u_t^2 on lagged values of y_t^2, x_t^2 cross products of y_t and x_t,

$$u_t^2 = \sum_{j=1}^{\gamma}\left(\alpha_{1,j}y_{t-j}^2 + \alpha_{2,j}x_{t-j}^2 + \alpha_{3,j}y_{t-j}x_{t-j}\right) \qquad (4)$$

where the lag value γ is equal to four.[27] The column under min-TR^2 reports the minimum value of TR^2, which asymptotically has a $\chi^2_{(d)}$ distribution with degrees of freedom $d = 12$ because four lags of each series and their cross products are used. The sign of the parameter ρ indicates the direction of second moments for each pair. A negative value for this parameter indicates that the second moments move in the same direction.[28]

The results presented in Table 3.6 indicate that Germany, France, Austria and the Netherlands form a group of bond markets that share a common volatility process, at least in a bivariate setting. The only pair in this group that shows ARCH effects is Austria–Netherlands. Note that the cointegration analysis in the second chapter of this dissertation indicated that the bond markets of Germany, France, Austria and the Netherlands appear to be integrated. The cointegration analysis indicated that there appears to be a long-run relationship in the first moments of these bond returns.

The common feature results reported in Table 3.6 look at the second moments of these returns and indicate that these bond market returns are also “cointegrated” in the second moments (Engle and Kozicki, 1993). This analysis indicates that Austria, France,

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[27] Following Engle and Susmel (1993) and Alexander (1995), we choose the lag order equal to 4.

[28] The programming is done in MATLAB. The FMINSEARCH function is used to find the parameter ρ that minimizes TR^2.

Table 3.6: Common Feature Results

Market	ρ	min-TR^2	Market	ρ	min- TR^2
North America and European countries					
U.S./Austria	-0.19	16.41	Germany/Austria	-0.66	14.44
U.S./Canada	-0.85	19.13	Germany/Denmark	-0.87	16.27
U.S./Denmark	-0.56	16.63	Germany/France	-1.22	7.39
U.S./France	-0.83	6.66	Germany/Ireland	-1.98	12.48
U.S./Germany	-1.59	13.22	Germany/Netherlands	-0.98	17.21
U.S./Netherlands	-1.33	8.00	France/Austria	-0.73	10.28
U.S./U.K.	0.44	13.59	France/Netherlands	-0.79	9.69
U.K./Denmark	2.50	16.90	Netherlands/Ireland	-1.89	12.18
			Austria/Denmark	-1.42	13.67
			Austria/Ireland	-2.28	15.48
Japan and Australia					
Japan/U.K.	0.51	7.43	Australia/Austria	-2.81	15.35
Japan/Denmark	0.31	16.00	Australia/France	-1.17	15.36
Australia/Denmark	-0.94	13.61	Australia/U.K.	0.27	12.24

Note: The results in this table present the pairs for which the TR^2 does not satisfy the 5% criteria, indicating the no ARCH effect in the residuals of these portfolios. Therefore, a common feature is detected.

Table 3.7: Common Feature Results

Market	ρ	min-TR^2	Market	ρ	min- TR^2
U.S./Australia	-3.14	25.62*	Canada/Australia	-1.34	21.87*
U.S./Ireland	-0.39	18.76***	Canada/Denmark	-0.67	24.94*
U.S./Japan	-2.35	24.23*	Canada/Ireland	-2.75	32.23*
Germany/Australia	-1.98	20.78**	Canada/Japan	-3.38	18.75***
Germany/Canada	-0.50	24.33*	Denmark/Netherlands	-0.80	23.15**
Germany/Japan	-0.15	28.55*	Ireland/Australia	-0.02	35.13*
France/Canada	-0.99	19.66**	Australia/Japan	1.53	18.75***
France/Denmark	-0.76	20.88**	Denmark/Ireland	0.11	25.38*
France/Ireland	-0.43c	18.39***	Ireland/Japan	0.66	19.82**
France/Japan	0.16	25.34*	U.K./Germany	0.51	19.58**
Netherlands/Australia	-1.34	27.42**	U.K./Austria	1.31	19.52**
Netherlands/Austria	-0.62	21.87**	U.K./Canada	0.23	25.66**
Netherlands/Canada	-0.55	27.42*	U.K./France	2.36	21.91**
Netherlands/Japan	4.21	34.38*	U.K./Ireland	10.32	25.38*
Austria/Canada	0.55	27.80**	U.K./Netherlands	1.74	25.53*
Austria/Japan	9.59	19.16**	U.S./Canada	-0.85	19.13**

Note: *, **, and *** indicates the pairs for which the TR^2 satisfies 1%, 5% and 10% criteria respectively. This indicates the presence of an ARCH effect in the residuals of these portfolios. Therefore, no common feature is detected.

Germany and Netherlands create a regional group of bond markets that are fully integrated. The parameter ρ is negative for all these pairs indicating that the movements in the conditional volatility are in the same direction. For example, in the case of Germany–France, the estimated parameter is -1.22, which suggests that the movement in the conditional volatility of German bond returns are 22% larger than those in conditional volatility of the French bond returns and in addition, these movements are in the same direction.

The U.S. bond returns show no ARCH effects when regressed against any of the bond returns of this group. In fact, the U.S. bond returns show no ARCH effect when regressed against any of the European bond markets. Similar results are obtained when the German bond returns are regressed against other European countries' bond returns, except for the U.K. These results confirm the regional importance of the German financial market within the European area and the global impact that the U.S. financial markets exert on the world financial market and, in particular, within the European area. Except for the U.K., the estimated ρ parameter is negative for all pairs exhibiting a common feature. This indicates that the movements in the conditional volatility in these bond returns move in the same direction.

The common feature results for U.K. bond returns with other international bond returns do not indicate the existence of common volatility. With regard to the European markets, we fail to find evidence of a common feature in volatility except for the case when the U.K. bond returns are regressed against Danish bond returns. For this pair the estimated ρ parameter is positive, indicating that the conditional volatilities for these two particular bond returns move in opposite directions. When the U.K. bond returns are regressed against other non-European bond market returns, no ARCH effect is found except with respect to Canada. These results indicate lack of a fully integrated process between the U.K. and other European bond markets.

The international bond market returns that are analyzed in this chapter include only two bond returns from the Far East countries: Australia and Japan. When Japanese bond returns are regressed against other international bond returns, evidence of ARCH effects are found, except for the case when Japanese bond returns are regressed against Danish and U.K. bond returns. The Australian bond returns appear to share a common volatility process with only Austria, France and the U.K. In conclusion, the common feature test results for the Australian and Japanese bond markets suggest that there is no strong evidence for the integration of these two bond markets with other international bond markets considered in this study.

Table 3.7 presents the results for all the pairs for which no common feature is found. For these pairs the TR^2 test statistics satisfies the 1%, 5% and 10% significance criteria. There are only four portfolios for which the TR^2 satisfies the 10% criteria: the portfolio consisting of U.S. and Irish bonds, the portfolio consisting of French and Irish bonds, the portfolio of Canadian and Japanese bonds, and the portfolio of Australian and Japanese bonds. The presence of an ARCH effect for these portfolios is not very strong.[29]

[29] We rerun the results by using a lag level of 3 and 5 in Equation (4). For both these lags we still cannot reject the presence of an ARCH for these four portfolios. For all other portfolios presented in Table 3.7 the TR^2 satisfies the 1% and 5% criteria.

3.5 Conclusions

The main focus of this chapter was to examine whether the conditional volatility process of the international bond returns share the same common volatility process. The results presented in this chapter parallel the results obtained earlier with respect to cointegration in the first moments across international bond market returns. The approach in this chapter differs from the previous chapter in that, instead of looking at the cointegration in first moments, we analyze cointegration in the second moments. Using the common feature methodology introduced by Engle and Kozicki (1993), we tested whether two international bond market indexes share the same volatility process. The presence of common volatility process between international bond markets is interpreted as evidence of integration between these markets. In addition, we examined whether the countries that share the same volatility process are within a region or not. If the common feature in the volatility process is more characteristic of countries within a particular region that would indicate the existence of a regional integration rather than a global one.

The majority of bond market returns examined in this chapter are from the European area. These countries have gone through a significant process of formal integration that was finalized with the introduction of a common currency, the EURO, in January 2002. Therefore, we expected to find relatively strong evidence of a common volatility process among these international bond market returns. The results presented in this chapter confirm this assumption. The results for Germany, France, Austria and the Netherlands suggest that they share the same volatility process. This indicates the presence of regional integration among these markets.

The common feature results with respect to the U.K. bond market do not provide us with strong evidence that the bond returns of this country share the same volatility process with other international bond returns. The volatility process for this market is not closely related with those of other European countries, indicating weak integration of this market with other European markets. These results seem to confirm those found by Alexander (1995b), Fraser and Oyefeso (2002), and Batavia et al. (2004) and suggest that the U.K. appears to have become more financially integrated with non-European countries than with European countries. With respect to Australian and Japanese bond market returns, the results do not indicate strong evidence for the integration of these two bond markets with other international bond markets considered.

The common feature results presented in this chapter are similar to the results obtained from rolling correlations and cointegration analysis on the first moment, which give evidence of the presence of the integration. These three approaches seem to suggest that, by and large, the IFI is indeed present and strong between Germany, France, Austria and the Netherlands on one hand and between U.S. and U.K. on the other.

Nonetheless, we are still suspicious of these results for a variety of reasons. One, the results with respect to the output synchronization does not point to full IFI. Two, while the cointegration analyses (with respect to the first and second moments) indicate that there are two cointegrated regions; we lack details because of the lack of dynamics in these approaches. IFI could be rising over time or falling over time, and these methodologies do not allow us to observe these trends that would have important implications for the main conclusions about the actual degree of IFI. Third, while the rolling correlations allow us to observe the dynamics in IFI and how it has change over time, this methodology is suspect. Trends in the degree of correlations over time and changes in the volatility are confounded. Therefore, we need to contrast our results with yet another dynamic approach to better discern IFI. In the next chapter a dynamic measure of IFI is presented, which corrects for the changes in dynamic dependence across international bond markets due to higher volatilities.

CHAPTER 4

A TIME-VARYING MEASURE OF INTERNATIONAL FINANCIAL INTEGRATION

4.1 Introduction

In the second chapter of this dissertation, the correlations among international bond markets were looked at. The 52-week rolling correlations among the bond returns from these markets indicated an overall increase in the co-dependence between international bond returns during the last part of the sample period. The correlations of German bond returns with other European countries' returns clearly increased starting in the mid to the late 1990s, while the correlations between German and other non-European countries exhibit an increase, especially after 1999. The increased correlations indicate increasing codependence among these international bond markets. The correlations between the U.S. and other international bond returns also show evidence of an increase in the co-dependence between these bond market returns.

The use of correlations as a measure of international financial integration, in place of more traditional measures such as saving-investment relationships or cross–country correlations has its own advantage. Instead of looking at the integration process as a state, as the traditional measures do, the correlations across international asset markets allow us to view international financial integration as an ongoing process. Nevertheless, the use of the correlations as a measure of financial market integration has recently been criticized.

Boyer, Gibson and Loretan (1999) argued that higher correlations across markets may not be due to an increase in the covariance, but rather to higher volatilities in asset markets.

In this chapter, a new measure is used to determine the dynamics of international financial integration. This measure makes use of the Capital Asset Pricing Model (CAPM), which relates domestic asset returns to world asset returns. This measure differs from the correlation method as it corrects for the changes in dynamic dependence across international bond markets due to higher volatilities. In addition, in this chapter, the saving-investment relationship and cross-country correlations are used to determine the degree of capital mobility across countries. In this way we are able to compare the benefits of using the new time-varying measure of IFI to these two traditional measures.

While the capital asset pricing model (CAPM) has become popular to determine the level of integration across international asset markets, it has been used mostly with respect to the stock market (see, among others, Bekaert and Harvey, 1995 and Bekaert et al., 2003). CAPM relates the domestic market return to the world market return, where the market beta indicates the sensitivity of the market returns to changes in the world market. In the international CAPM, the appropriate measure of international financial integration (IFI) is the market beta. Higher values of beta would imply higher levels of IFI. However, the traditional applications of this model have considered the beta coefficient to be fixed over time, therefore ignoring the dynamics of international financial integration.

To account for the dynamics of IFI, Bekaert and Harvey (1995) presented a parameterization of beta that allows for time variation of financial integration. One drawback of this parameterization is that it is highly parametric, making the results very sensitive to the choice of instrumental variables used in estimation. In this chapter, the dynamic conditional correlations multivariate-generalized autoregressive conditional heteroskedasticity (DCC MV-GARCH) model of Engle (2002) is used to obtain a time-varying beta coefficient in the CAPM without relying on highly parametric CAPM. The beta coefficient gives the covariance between a country's asset return and the global asset returns divided by the variance of global asset returns.

In previous chapters, rolling correlations were used to determine the linear relationship across international bond market returns. The Johansen (1988) cointegration procedure was looked at to find whether there is a long-run equilibrium across these returns, and Engle and Kozicki's (1993) common feature methodology was also used to determine whether these returns share the same volatility process. Either each of these measures analyzes the mean or variance of the international bond market returns in an attempt to determine the degree of IFI. By using a time-varying beta coefficient to determine the level of IFI across international bond market returns we are able to account for both mean and variance of these returns as well as their offsetting values. Therefore the time-varying measure of IFI used in this chapter will give a more general and comprehensive measure of financial integration. This time varying measure of IFI encompasses the previous measures of IFI used: rolling correlations, cointegration in the first moment and common volatility. In addition, it also allows us to detect the dynamics of the financial integration process.

The next section describes the data and countries analyzed. Before obtaining the time varying measure of IFI, in Section 4.3 two traditional measures are reported that are often used in the literature to determine the degree of capital mobility: the saving-investment relationship and cross-country correlations. Section 4.4 presents the model used to estimate a time varying measure of IFI, and section 4.5 presents the new time-varying IFI estimation results. Section 4.6 presents a short conclusion.

4.2 Data Descriptions

In this chapter, the DCC MV-GARCH model will be used to obtain a time varying measure of international financial integration. This model has the flexibility of the univariate GARCH model and at the same time allows correlations to change over time. What is more important, the DCC MV-GARCH model does not restrict the number of series considered in the system. However, the DCC MV-GARCH model assumes the same structure of dynamic conditional correlation across countries. The results presented in the previous chapters indicate different degrees of IFI across these countries. Therefore including all international bond returns in the DC MV-GARCH system would not give a correct time varying measure of the international financial integration.

The analysis of international bond returns in the previous chapters of this dissertation indicated the existence of two integrated groups. Austria, France, Germany and the Netherlands appear to create one group with integrated financial markets. On the other hand, evidence was found of financial integration between the U.S. and the U.K. bond market returns. Therefore, the DCC MV-GARCH analysis will be conducted on two groups of international bond returns. The first group will include Austria, France, Germany and the Netherlands bond returns, and the second group the U.S. and U.K. bond returns.

The data consists of 10-year DataStream Benchmark Bond Indexes measured in U.S. dollars for Austria, France, Germany, the Netherlands, the U.S. and U.K. These data are available on a daily frequency covering the period June 22, 1989 to June 22, 2004. To avoid any problem raised by the asynchronomy of these daily data, they are converted into weekly frequency.[30] Weekly bond returns are calculated as log differences using Friday-to-Friday closing prices.

In the second chapter of this dissertation, we found that January 1994 and January 1999 were two dates that had significant impacts on the returns of these international bond markets. Therefore, for each bond returns series, the following regression is used

$$y_{it} = a + b_1 \times dummy2 + b_2 \times dummy3 + \varepsilon_{it} ,$$

where y_{it} is the bond return of country *i*, *dummy2* takes value 1 the first week of January 1994 and *dummy3* takes value 1 the first week of January 1999. The residuals from the above regression are used in the following analysis. In addition, following the same procedure for the U.S. and U.K. bond returns, an MA(4) process is fitted, and for the Canadian bond returns an MA(3) process is fitted. The residuals from these MA models are used in the following analysis.

For the same group of countries, the saving-investment relationship and cross-country correlations were also estimated to determine the degree of capital mobility across countries. To obtain these traditional measures of capital mobility, quarterly data

[30] The financial markets analyzed here do not have the same trading hours. The trading times between these financial markets overlap only partially or they do not have common trading hours. This lack of common trading hours will lead to a series of problems including a significant downward bias of correlations if used in estimation (Martens and Poon, 2001). In order to avoid this problem of asynchronous data, the daily indexes are converted into weekly indexes.

covering the period 1970:Q1 to 2004:Q3 was used. The data are taken from IMF, International Financial Statistics (IFS).[31]

In obtaining these traditional measures of capital mobility, the saving rate (SR) is defined as gross domestic savings divided by gross domestic product (GDP). Gross savings is calculated as the difference between the GDP and total consumption, where total consumption is calculated as the sum of government and household consumption. The investment rate (IR) is calculated as the ratio of gross domestic investment to GDP. Both the SR and IR are expressed as percentages. Real private consumption and GDP data are used to obtain the cross-country consumption and output correlations. The GDP deflator is used to compute these real time series.

The data obtained from IFS are seasonally adjusted. However, the analysis of the correlogram for the SR and IR series of Austria, up to 36 lags, indicated the presence of seasonality. This behavior is observed only for these two series. Before continuing with the analysis of the SR and IR relationship, we control for this seasonality for the SR and IR series of Austria. The seasonal difference, $y_t - y_{t-4}$, is taken for each series.[32] The seasonally differenced series are then used in the following analysis. The autocorrelation functions of SR and IR series for other countries do not indicate the presence of seasonality.

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[31] For the Netherlands these data are available in quarterly frequencies starting from 1977:Q1. Therefore, for the Netherlands the analysis of S-I and consumption correlations cover the period 1977:Q1 to 2004:Q3.

[32] For quarterly data, in the presence of the seasonality, the autocorrelation function (ACF) will show significant spikes at lags 4, 8, 12… A similar pattern was observed in the ACF function of SR and IR series for Austria (see Enders (2004) for a more detailed explanation on correcting for seasonality).

4.3 Non-Dynamic Traditional Measures of IFI

This section reports the saving-investment relationship and cross-country consumption and output correlations for Austria, France, Germany, Netherlands, the U.S. and the U.K. These measures are commonly used to measure the degree of international financial integration. While many studies analyze saving-investment correlations and consumption correlations, this literature is updated by including data from 1990 onward. It is in this period where it is believed that the degree of IFI has increased significantly. The results presented in this section are comparable with those presented by Batavia et al. (2004). Using quarterly data for the period 1985–2002 for 13 European countries and U.S., they compute (1) covered interest rate parity, (2) uncovered interest rate parity, and (3) real interest rate parity to measure the degree of international financial integration across these countries.[33]

Saving-Investment Correlations

The saving-investment relationship, proposed by Feldstein and Harioka (1980), is one of the earliest empirical methods used to determine the degree of capital mobility across countries. The main idea of this measure is if capital is highly mobile internationally, then the correlation between domestic savings and investment in a country should not be high. In estimating the following equation

$$\left(I/Y\right)_i = a + b\left(S/Y\right)_i + \varepsilon_i \qquad (1)$$

[33] Please note that the main motivation of including this section in this chapter is that we can then compare the IFI results from using the two traditional measures with the new time-varying measure of IFI presented in Section 4.4. Therefore the number of countries analyzed in this section is smaller than that considered by Batavia et al. (2004).

under perfect capital mobility, for a small country, one should find a value of b equal to zero. This implies that the investment decisions in a small domestic country do not depend on domestic savings. If a value of b equal to 1 is found, that would suggest that domestic investment relies totally on domestic savings. For large open economies, under perfect capital mobility the b coefficient should approximate the country's share of the world's capital stock. Changes in savings in large countries would affect the world interest rate, therefore leading to higher level of correlation between saving and investments.

Feldstein and Harioka (1980) tested regression (1) for 16 OECD countries for the period 1960 to 1974. They obtain a b-coefficient equal to 0.89 that is not significantly different from zero, using a cross-section regression involving the saving and investment rates of these 16 OECD countries. Their result indicated a relatively low level of capital mobility across these countries. Following Feldstein and Harioka (1980) (FH hereafter), different empirical studies used equation (1) to determine the degree of international financial integration.[34] The results of these studies appear to be robust, suggesting a low degree of capital mobility across countries.

Different explanations have been suggested to explain the observed high correlations between domestic savings and investment. Frankel (1992) and Eijffinger and Lemmen (1995), among others, argue that if the domestic real interest rate is not tied to the foreign interest rate, then we cannot expect a zero correlation between savings and investment. Thus the high correlations observed between savings and investment in the various empirical studies can be explained by the failure of real interest rate parity.

[34] See Coakley et al. (1998) for a review of the studies using Feldstein and Harioka's (1980) approach in measuring the degree of international capital mobility.

Summers (1988) and Bayoumi (1990) argue that current account targeting by the government can result in higher degrees of SR and IR correlation, regardless of capital mobility. Other factors such as population growth, productivity shocks (Summer, 1988 and Obstfeld, 1986) or the existence of non-traded goods and factor immobility (see Murphy, 1986 and Wong, 1990) will also induce higher correlations between savings and investments.

Another possible alternative in explaining the high saving-investment correlations is that most of the early empirical studies applied a cross sectional analysis. The FH approach of measuring capital mobility can be undertaken using a cross sectional or a time series approach. An important limitation of using cross sectional analysis (see Sinn, 1992; Jansen 1994, 1996a) is that the dynamics of saving–investment correlations and the differences in the economic structure between countries are ignored. The cross sectional approach also does not account for nonstationarity in the variables. Consequently, more attention has recently been placed on using time-series analysis. Time series studies are not as plentiful as cross sectional studies, and as of now the results are mixed.

Jansen (1996b) used an error correction model (ECM) to analyze the dynamic relationship between saving and investment

$$\Delta I_t = a + b\Delta S_t + c\left(S_{t-1} - I_{t-1}\right) + dS_{t-1} \qquad (2)$$

Jansen (1996b) argued that this model provides us with more than one way to detect capital mobility. The *b* coefficient measures the short-run correlation between saving and investment. Relatively low values of this coefficient would be an indication of capital mobility. $S_{t-1} - I_{t-1}$ presents the long-run relationship between saving and investment ratios. When $c \neq 0$, these two series are said to be cointegrated. Following Jansen (1996b) the cointegrating relationship between saving and investment is

$$a + c\left(\bar{S} - \bar{I}\right) + d\,\bar{S} = 0 \qquad (3)$$

where the bar indicates the long-run values. The cointegrating relationship implied by (3) is *(1 + d/c, -1)*. If $d = 0$, this would imply that the current account (S–I) is equal to $-a/c$. Therefore, the current account is stationary. If $a = d = 0$, then the current account fluctuates around zero. In both cases, the cointegrating relation assumed is *(1, -1)*, which corresponds to the standard steady state. If $c \neq 0$ and $d \neq 0$, then the cointegrating vector between saving and investments will not be *(1, -1)*, but *(1 + d/c, -1)*. In this case the current account is not stationary. Nonstationarity of current account would imply a certain degree of capital mobility.

As suggested by Jansen (1996b), in the case where $c \neq 0$ and $d = 0$, our attention is shifted to analyzing the value of the *b* coefficient. Low values of this coefficient are obtained in the situations when there is capital mobility. Jansen (1996b) cautions that although low values of the *b* coefficient are indicative of capital mobility, this coefficient does not give a quantitative measure of capital mobility. Jansen (1996b) suggests the following steps for detecting capital mobility when saving and investments are found to be cointegrated. *First*, if *d* is different from zero, the current account is not converging to a constant in the long run, and therefore there is capital mobility. *Second*, if *d* is zero and *c* is different from zero, then small values of *b* would imply that there is no short-run correlation between saving and investment. Applying this ECM to 23 OECD countries, he finds evidence of capital mobility. His results support the idea that S–I correlations are positively related to country size. Rensselaer and Copeland (2000) use the same ECM to investigate the S–I correlation between 15 Latin America countries and find evidence of capital mobility.

In this section, the results of the time series analysis of the saving-investment correlations for Austria, France, Germany, Netherlands, the U.S. and the U.K. are presented. First, the stationarity of the SR and IR series for each country is determined. If series are found to be stationary then equation (1) is used to determine the degree of capital mobility. In case of nonstationarity, each series is differenced *l* times to invoke stationarity. In this case, the series is said to be integrated of order *l*, that is I*(l)*. If SR and IR series are found to be integrated of the same order then a cointegration test is used to determine the existence of a long-run equilibrium. If cointegration is not found, that would imply that there is no long-run relationship between saving and investment, implying capital mobility. If cointegration is found, Jansen's (1996b) procedure presented in equation (2) is used to determine the degree of capital mobility.

Table 4.1 presents the unit root tests for the saving and investment series for each country. Two commonly used tests, the augmented Dickey-Fuller (ADF, 1979) and Kwiatkowski, Phillips, Schmidt, and Shin (KPSS, 1992) tests are conducted to test for the presence of unit root. The ADF test takes the unit root as the null hypothesis. The test regression used to test for the presence of a unit root where both an intercept and a trend are included is as follows:

$$\Delta y_t = a_0 + \gamma y_{t-1} + a_1 t + \sum_{i=2}^{p} \beta_i \Delta y_{t-i+1} + \varepsilon_t$$

Table 4.1: Unit Root Tests for Saving and Investment Ratios

		ADF							KPSS
Country	Variable	Lag	No Const.	Const.	Const &Trend	φ_1	φ_2	φ_3	H_0: Level Stationarity
Austria	SR	7	-3.89^{*}	-4.09^{*}	-4.64^{*}	0.79	2.10	2.34	0.29
	IR	8	-2.70^{*}	-2.79^{***}	-3.04	0.29	0.70	0.76	0.15
France	SR	7	-1.07	-2.12	-1.98	2.05	1.47	0.16	1.79^{*}
	IR	8	-0.80	-2.09	-3.79^{**}	2.06	4.79	4.97	2.79^{*}
Germany	SR	1	-1.16	-3.08^{**}	-2.89	4.26	2.94	0.18	1.16^{*}
	IR	4	-1.41	-2.82^{***}	-2.87	3.65	2.58	0.24	0.70^{**}
Netherlands	SR	8	1.23	-2.61^{***}	-0.72	3.76	2.56	0.08	2.13^{*}
	IR	8	-0.10	-2.48	-2.50	3.13	2.30	0.29	0.18
U.K.	SR	8	-0.98	-1.68	-3.72^{**}	1.25	4.52	5.42	2.52^{*}
	IR	4	-0.57	-2.33	-3.30^{***}	2.62	3.56	2.63	1.76^{*}
U.S.	SR	8	-1.22	-0.11	-2.12	0.00	1.99	2.99	2.42^{*}
	IR	5	0.00	-3.10^{**}	-3.28^{***}	4.83	3.61	0.58	0.71^{*}

Note: For ADF unit root test, the null hypothesis is the existence of a unit root. The lag length for this test was selected using the AIC and SIC criteria. The critical values for the test statistics φ_1, φ_2, and φ_3 at 5% significance level are 4.71, 4.88 and 6.49 respectively. For the KPSS test, under the null the series are assumed to be locally stationary. The lag length for KPSS test is equal to 5 and was selected using the level selection procedure proposed by Newey and West (1994). *, **, and *** indicate the significance level at 1%, 5% and 10% respectively.

where p is the lag level used using the Akaike information criteria (AIC) and Schwarz information criteria (SIC).[35] The null hypothesis of a unit root implies $\gamma = 0$. The test results are sensitive to the inclusion, or exclusion of an intercept, an intercept and a trend or neither in the test regression.

Dickey-Fuller (1981) provide three additional tests, φ_1, φ_2, and φ_3, to test the significance of the inclusion of an intercept and a trend in the above regression. The φ_1 statistic tests the null hypothesis of $\gamma = a_0 = 0$. The φ_2 statistic tests the null of $\gamma = a_0 = a_1 = 0$, and φ_3 statistic test the null of $\gamma = a_1 = 0$.[36] The ADF test results indicate that the null hypothesis of a unit root cannot be rejected for all these series except for the SR of Austria, and the IR of Austria. In addition, the φ_1, φ_2, and φ_3 test statistics indicate that no intercept and no trend are needed in the above regression.

The last two columns of Table 4.1 report the KPSS unit root test. This test differs from the ADF test in that the series are assumed stationary under the null. This test is conducted under the null of level stationarity. Based on these results we fail to reject the null hypothesis of stationarity of the SR and IR of Austria. This result confirms the ADF test results. Based on the KPSS test results we also fail to reject the null hypothesis of stationarity for the IR of Netherlands.[37] For all other SR and IR series the KPSS test

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[35] AIC and SIC are two common information criteria used to determine the lag length. However these criteria often lead to different optimal lag lengths. Practice has shown that AIC tends to overfit the lag length while SIC tends to underfit it. In such cases, where AIC and SIC lead to two different conclusions we report the optimal lag suggested by AIC. At the optimal lag length the residuals are white noise. However, even when the AIC and SIC suggest different optimal lag lengths, they give the same results for unit root tests.

[36] See Enders (2004) pp. 181–183 for a more detailed description of these tests.

[37] Although ADF and KPSS give contradictory results, an analysis of the plot for the Netherlands' IR series indicates that it is nonstationary. Therefore, in the following analysis we consider this series to be nonstationary.

reject the stationarity hypothesis. In addition, the ADF test and KPSS tests reject the null hypothesis of a second unit root.

As SR and IR series of Austria are stationary, in testing the degree of capital mobility equation (1) is used. The estimated results of equation (1) for Austria are

$$IR_t^{Au} = \underset{(0.11)}{0.00} + \underset{(0.12)}{0.39} SR_t^{Au},$$

where the numbers in parentheses give the White's (1980) robust standard errors. Although the estimated value of the b coefficient, 0.39, is significantly different from zero, it is also significantly different from 1. This indicates that there is a relatively high degree of capital mobility.

For France, Germany, Netherlands, the U.S. and U.K. the SR and IR series were integrated of order one. Therefore, the existence of cointegration between SR and IR for each country was tested. The lag length is selected using Sims' (1980) likelihood ratio corrected for small samples,

$$LR = (T-k) * \left(\log|\Omega_r| - \log|\Omega_u|\right).$$

Where T is the sample size, k is the number of coefficients in each equation, $\log|\Omega_r|$ is the log determinant of residual covariance matrix when the model is restricted and $\log|\Omega_u|$ is the log determinant of residual covariance matrix when the model is unrestricted.[38]

[38] The optimal lag length was selected using data in first differences. The maximum lag length tested was 10. The restricted model has a lag length one of k-1, while the unrestricted model has a lag of k. At the optimal lag length the residuals are white noise.

The cointegration results are sensitive to the assumptions made with respect to the deterministic components. There are five possible models within the Johansen procedure (Johansen, 1994):

(1) Series X have no deterministic trends and the cointegrating equations (CE) do not have intercepts;

(2) Series X have no deterministic trends and the CE have intercepts;

(3) Series X have a deterministic trends and the CE have intercepts;

(4) Both series X and CE have deterministic trends; and

(5) Series X have quadratic trends and the CE have trends

with the first model being the most restrictive model and the last model being the least restrictive model. [39]

In order to determine which of the above models better represent the data the following test statistic is proposed by Johansen (1991)

$$-T\sum_{i=r+1}^{n}\left[\ln\left(1-\hat{\lambda}_i^{*}\right)-\ln\left(1-\hat{\lambda}_i\right)\right]$$

where T is the sample size, $\hat{\lambda}_i^{*}$ and $\hat{\lambda}_i$ are the characteristic roots of the restricted and unrestricted model respectively, and r is the number of the nonzero characteristics root of the unrestricted model. Asymptotically, this test statistics has a χ^2 distribution with (n - r) degrees of freedom.[40] Rarely would economic series exhibit a quadratic trend, therefore in the empirical applications of the cointegration analysis, only models (1) through (4) are considered.

[39] Section 2.5 of this dissertation describes in more detail the Johansen (1991) cointegration procedure.

[40] See Enders (2004) pp 354–357 for a detailed description of this test.

The cointegration results are presented in Table 4.2. Johansen's (1992) cointegration test is used. For all the cointegration analysis, using the above test statistic we reject models (3) and (4). Excluding the existence of a trend in the VAR, this leaves us with two possible assumptions with regard to the CE: inclusion, or exclusion of a constant in CE, respectively model (1) and (2). The last column in Table 4.2 reports the p-value of the above test statistics for the hypothesis H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. The results of the model suggested by this test are underlined. For all countries, we reject the null hypothesis of no constant in the CE.

Table 4.2: Cointegration of SR and IR Series

SR and IR cointegration for	Lag	No Time Trend in Data: No Intercept in CE*	No Time Trend in Data: Intercept in CE	p-value
France	4	r=0	<u>r=0</u>	0.02
Germany	4	r=0	<u>r=1</u>	0.02
Netherlands	4	r=0	<u>r=0</u>	0.03
U.K.	4	r=0	<u>r=0</u>	0.03
U.S.	5	r=0	<u>r=0</u>	0.00

Note: CE stands for cointegrating equation, r gives the cointegration rank. The results reported above are for the case when no trend is included in the VAR. The lag length is selected using Sims' (1980) likelihood ratio test corrected for a small samples. The last column gives the p-value of the LR test statistic for the null hypothesis of H_0: *No constant in the CE* versus the alternative H_1: *There is an intercept in the CE*. The results of the model suggested by this LR test are underlined.

The results presented in Table 4.2 suggest that there is no evidence of cointegration between SR and IR series for France, Netherlands, U.K. and the U.S. For Germany, the SR and IR series show evidence of cointegration. Therefore, following Jansen (1996b), an error correction model (ECM) is used to analyze the dynamic relationship between saving and investment for Germany. The estimated results of equation (2) for Germany are as follows,

$$\Delta IR_t = -\underset{0.73}{0.33} + \underset{(0.12)}{0.04}\Delta SR_t + \underset{(0.03)}{0.07}\left(SR_{t-1} - IR_{t-1}\right) + \underset{(0.03)}{0.01}SR_{t-1}$$

where the numbers in parentheses give the White's (1980) robust standard errors. The estimated value of the coefficient on $\left(SR_{t-1} - IR_{t-1}\right)$ is equal to 0.07 and is significantly different from zero at the 1% level significance. This confirms the previous results on the presence of cointegration between SR and IR for Germany.

The presence of cointegration does not necessarily imply the lack of capital mobility. As suggested by Jansen (1996b), the second step in detecting the presence of capital mobility is to test the significance of the *d* coefficient on SR_{t-1}. A *d* coefficient not significantly different from zero would imply a stationary current account *(SR-IR)* and a *d* coefficient significantly different from zero would imply a nonstationary current account *(SR-IR)*. The nonstationarity of current accounts is evidence of capital mobility. For Germany, the estimated value of the *d* coefficient is 0.01, which is not significantly different from zero. This suggests that the current account *(SR-IR)* for Germany is constant over time; therefore, there is no evidence of capital mobility. As Jansen (1996b) suggests, in this situation, when there is a long-run equilibrium relationship between SR and IR, and the current account is constant over time, we need to look at the value of the *b* coefficient. This coefficient measures the short-run correlation between saving and investment. The estimated value for this coefficient is 0.04, which is not significantly different from zero. As Jansen (1996b) suggests this low value of *b* coefficient indicates a degree of capital mobility.

The above results with regard to Germany indicate strong evidence of a long-run relationship between saving and investment. Note that the saving-investment relationship of a country depends not only on the degree of the capital mobility. Country size also has an impact on this relationship. Changes in savings in large countries would affect the world interest rate, therefore leading to a higher level of correlation between saving and investment. Tobin (1983), Obstfeld (1986) and Baxter and Crucini (1993) among others, give evidence that the saving-investment correlations appear to be positively related with country size. The large country effect can be one possible explanation of the result found for Germany. Nevertheless, the large country effect cannot be used to explain the results for the U.S. This is consistent with the findings of Frankel (1986).

Cross-Country Consumption Growth Correlations

Table 4.3 presents the consumption correlations and output correlations for each country with Germany and U.S. These cross-country correlations are an alternative method used to measure the degree of international capital mobility. The main idea is that if individuals across countries have access to the same set of financial instruments, then under perfect capital mobility there should be a perfect co-movement of a country's consumption growth with world consumption growth (see Obstfeld, 1986). Bayoumi and MacDonald (1995) argue that the analysis of consumption growth correlations has the attractive feature that the underlying theory is stronger than that underlying the FH approach. Comparison of consumption growth correlation is different from other methods also because it does not require comparisons of dissimilar assets.

Table 4.3: Consumption, Output Correlations

	Correlations with Germany			Correlations with the U.S.		
	corr(C^j, C^{GM})	*corr*(Y^j, Y^{GM})	DIF	*corr*(C^j, C^{US})	*corr*(Y^j, Y^{US})	DIF
Austria	0.856*	0.031	0.825*	-0.020	-0.066	0.045
France	0.918*	0.037	0.881*	-0.004	0.074	-0.078
Germany	1.000	1.000	0.000	-0.016	0.031	-0.047
Netherlands	-0.154**	0.255*	-0.409*	-0.139***	0.228*	-0.367
U.K.	0.091	0.273*	-0.182	0.110***	0.270*	-0.160
U.S.	-0.016	0.031	-0.047	1.000	1.000	0.000

Note: *corr*(C^j, C^i) and *corr*(Y^j, Y^i) indicate the consumption and output correlations between countries. DIF gives the difference between *corr*(C^j, C^i) and *corr*(Y^j, Y^i). *, ** and *** indicate significance at 1%, 5% and 10% level respectively.

An increase in the degree of capital mobility allows countries to smooth their consumption and be less exposed to the fluctuations of domestic output. Therefore, a higher degree of capital mobility would imply that cross-country consumption correlations should be higher than cross-country output correlations:

$$corr\left(C^i C^j\right) > corr\left(Y^i Y^j\right).$$

Table 4.3 presents the cross-country consumption and output correlations. The first part of this table presents the consumption and output correlation of Austria, France, Netherlands, the U.K. and U.S. with German's consumption and output. The third column of Table 4.3 presents the difference between consumption and output correlations $\left[corr\left(C^i C^{GM}\right) - corr\left(Y^i Y^{GM}\right)\right]$. High values of this difference indicate high levels of consumption smoothing. Columns 4 and 5 in Table 4.3 present the consumption and output correlations of Austria, France, Germany, Netherlands and U.K. with U.S. consumption and output respectively. The last column presents the difference between these consumption and output correlations for each country.

In the previous chapters, we found evidence of a relatively high degree of IFI between Austria, France and Germany. The correlations presented in Table 4.3 for these countries confirm our previous results. These correlations are significantly higher than the corresponding output correlations indicating a high degree of capital mobility across these countries. The consumption correlations of the U.S. and U.K. with German consumption are relatively small; they are not significantly different from zero. In addition, even the corresponding output correlations are not significantly different from zero. Therefore, these results indicate a low degree of capital mobility of the U.S. and U.K. with Germany. The results of the consumption correlations of Netherlands with Germany are interesting. Contrary to what we found in the previous chapters, the small correlation value indicates no consumption smoothing between these countries.

The correlations of Austrian, French, German, the Netherlands and U.K. consumption and output with the U.S. consumption and output indicate a low level of consumption smoothing. Even for the U.S. and U.K., the consumption correlation, although it is the largest correlation coefficients in this group of countries, still does not indicate a high level of consumption smoothing. These results are inconsistent with our previous results. Overall, the correlations with respect to the U.S. indicate a low level of financial integration between the U.S. and other countries considered in the analysis.

In conclusion, the saving-investment correlations and cross-country correlations give mixed results about the degree of international financial integration. The results presented in this section suggest that Austria, France and Germany are almost fully integrated. These results confirm the finding of Batavia et al. (2004), where results of almost full integration between these countries are found. Different from what we found in the previous chapters, these results indicate a low level of integration between Germany and the Netherlands and between the U.S. and U.K. Batavia et al. (2004) find similar results with respect to Germany and the Netherlands. In their results, while the U.K. appears to be more integrated with the U.S. than with other European countries, these two countries are far from being fully integrated.

However, these measures are not able to show us the dynamics of financial integration between these countries. Using these measures, we are not able to measure how the degree of financial integration has changed over time. In the next section, we describe the model used to estimate a time varying measure of IFI and section 4.5 presents the new time-varying IFI estimation results.

4.4 Measuring Financial Integration

This section presents the relationship between time-varying conditional correlations and the beta coefficient in the international CAPM. Using this beta coefficient, we are able to obtain a time-varying measure of IFI. Let $Y_{i,t}$ denote the domestic bond return of country i. Let $r_{i,t}$ be the deviation of $Y_{i,t}$ from its conditional mean and let H_t be the conditional variance-covariance matrix of r_i's, $i = 1, \ldots k$,

$$r_t | \Psi_{t-1} \sim N\left(0, H_t\right) \ . \tag{4}$$

Then, the traditional international CAPM implies that $r_{i,t}$ will depend on world market shocks, $r_{w,t}$, and on idiosyncratic shocks, $\varepsilon_{i,t}$

$$r_{i,t} = \beta_{i,t}^{w} r_{w,t} + \varepsilon_{i,t} \ . \tag{5}$$

The above model assumes that the shock can originate from two sources: the world shock and the internal (idiosyncratic) shocks. In an isolated market, the returns in domestic countries are driven only by the idiosyncratic shocks. As the domestic market becomes more financially integrated with world markets, the idiosyncratic shocks will become less relevant, while the impact of world shocks on domestic returns increases. The coefficient $\beta_{i,t}^{w}$ is a time-varying measure of the sensitivity of market *i* to the world shocks, which can be interpreted as a measure of financial integration. In contrast to the structural model of Bekaert, Harvey and Ng (2003), we use a pure time series approach to estimate the $\beta_{i,t}^{w}$'s.

Assuming that the world and idiosyncratic shocks are uncorrelated, the conditional variances and covariances are

$$h_{w,t} = E\left[r_{w,t}^{2} \mid I_{t-1}\right] \tag{6}$$

$$h_{i,t} = E\left[r_{i,t}^{2} \mid I_{t-1}\right] = \left(\beta_{i,t}^{w}\right)^{2} h_{w,t} + \sigma_{i,t}^{2} \tag{7}$$

$$h_{i,w,t} = E\left[r_{w,t} r_{i,t} \mid I_{t-1}\right] = \beta_{i,t}^{w} h_{w,t} \tag{8}$$

$$h_{ij,t} = E\left[r_{i,t} r_{j,t} \mid I_{t-1}\right] = \beta_{i,t}^{w} h_{w,t} \quad , \tag{9}$$

where $h_{i,j}$ and $h_{i,w}$ are the conditional covariances between countries, h_i and h_w are the conditional variances of market *i* and world market, and $\sigma_{i,t}^{2}$ is the conditional variance of the idiosyncratic shock of market *i*. Therefore, the conditional correlation between the world market returns and the returns in country *i* is given by

$$\rho_{i,w,t} = \frac{\beta_{i,t}^{w} \sqrt{h_{w,t}}}{\sqrt{h_{i,t}}} \ . \tag{10}$$

The expression above shows that an increase in conditional correlation does not necessarily imply that the markets are becoming more integrated. A simple transformation of (10) gives

$$\beta_{i,t}^{w} = \rho_{i,w,t} \frac{\sqrt{h_{i,t}}}{\sqrt{h_{w,t}}} . \tag{11}$$

To obtain our measure of IFI, $\hat{\beta}_{i,t}^{w}$, we need to estimate the conditional correlation between the world market and market *i*, and the conditional variance of each market.

Engle's (2002) DCC MV-GARCH model is used to estimate the conditional variances and covariances in (6)–(9). Bollerslev, Engle and Wooldridge (1988) originally presented the MV-GARCH model in the so called *vech* parameterization.[41] However, the number of parameters needed to be estimated for large models is very high. This brings into question the positive definiteness of the variance covariance matrix.[42] To reduce the number of parameters and ensure the positive definiteness of the variance-covariance matrix, Bollerslev (1990) proposed the constant correlation MV-GARCH model. Although the positive definiteness of variance-covariance matrix is ensured, different studies have shown that the constant conditional correlation assumption is not a plausible one (see Tsui and Yu, 1999 and Tse, 2000). Engle and Kroner (1995) proposed another class of MV-GARCH model, which is denoted the BEKK model. This parameterization reduces the minimum number of parameters in the original *vech* model of Bollerslev, Engle and Wooldridge (1988). The disadvantage of the BEKK model is that the parameters cannot be easily interpreted. Engle (2002) proposed a new class of multivariate GARCH models that allow correlations to change over time and at the same

[41] The term *vech* comes from the column-stacking operator VECH(.) applied to the upper triangle of a symmetric matrix.

[42] For example, in a system of three series, the unrestricted *vech* model will require the estimation of 78 parameters.

time has the flexibility of the univariate GARCH. The positive definiteness of the variance-covariance matrix is easily ensured.

Let the conditional variance-covariance matrix of the r_i's, $i = 1, ...k$, be

$$H_t \equiv D_t R_t D_t$$

where D_t is the k x k diagonal matrix of time-varying standard deviations from univariate GARCH, and R_t is the time-varying correlation matrix. Each diagonal element of D_t is specified as a univariate GARCH model,

$$h_{it} = \omega_i + \sum_{p=1}^{\mathscr{P}} \alpha_{ip} r_{i(t-p)}^2 + \sum_{q=1}^{\mathscr{Q}} \beta_{iq} h_{i(t-q)} , \qquad i = 1, ... k$$

where

$$\sum_{p=1}^{\mathscr{P}} \alpha_{ip} + \sum_{q=1}^{\mathscr{Q}} \beta_{iq} < 1 .$$

The structure of dynamic conditional correlation as suggested by Engle (2002) is,

$$\Phi_t = \left(1 - \sum_{m=1}^{M} \alpha_m - \sum_{n=1}^{N} \beta_n \right) \bar{\Phi} + \sum_{m=1}^{M} \alpha \left(\varepsilon_{t-m} \varepsilon_{t-m}^{'} \right) + \sum_{n=1}^{N} \beta_n \Phi_{t-n} \qquad (12)$$

$$R_t = diag\{\Phi\}^{-1} \Phi_t diag\{\Phi\}^{-1}$$

where $\varepsilon_t = D_t^{-1} r_t$ are the standardized residuals, Φ_t and $\bar{\Phi}_t$ are the matrices of conditional and unconditional covariances of standardized residuals. This model is mean reverting as long as

$$\sum_{m=1}^{M} \alpha_m + \sum_{n=1}^{N} \beta_n < 1$$

When $N = M = 1$ and $\alpha + \beta = 1$, the process followed by conditional covariances have a random walk structure. In this case, the covariances are integrated and (12) can be written as

$$\Phi_t = (1-\lambda)\left(\varepsilon_{t-1}\varepsilon_{t-1}^{'}\right) + \beta\Phi_{t-1} \tag{13}$$

where (1 - λ) = α and λ = β.

Engle's DCC model allows for two-stage estimation. The log likelihood of the DCC model is

$$L = -\frac{1}{2}\sum_t \left(k\log(2\pi) + \log(|H_t|) + r_t^{'} H_t^{-1} r_t\right)$$

$$= -\frac{1}{2}\sum_t \left(k\log(2\pi) + \log(|D_t R_t D_t|) + r_t^{'} D_t^{-1} R_t^{-1} D_t^{-1} r_t\right)$$

$$= -\frac{1}{2}\sum_t \left(k\log(2\pi) + 2\log(|D_t|) + r_t^{'} D_t^{-1} r_t - \varepsilon_t^{'}\varepsilon_t + \log|R_t| + \varepsilon_t^{'} R_t^{-1}\varepsilon_t\right) \tag{14}$$

Let θ denote the parameters in D_t and let ϕ be the parameters in R_t. We can then write the log-likelihood function in (14) as

$$L(\theta,\phi) = L_v(\theta) + L_c(\theta,\phi) \tag{15}$$

where the first term in (15) gives the volatility term

$$L_v(\theta) = -\frac{1}{2}\sum_t \left(k\log(2\pi) + \log\left(|D_t|^2 + r_t^{'} D_t^{-1} r_t\right)\right) \tag{16}$$

and the second term is the correlation part

$$L_c(\theta,\phi) = -\frac{1}{2}\sum_t \left(\log|R_t| + \varepsilon_t^{'} R_t^{-1}\varepsilon_t\right).$$

In the first stage, $L_v(\theta)$ is maximized by estimating univariate GARCH models for each residual series. In the second stage, taking the estimated parameters of univariate models as given, the DCC parameters are estimated by maximizing $L_c(\theta,\phi)$. Engle and Sheppard (2001) show that this two-step procedure yields consistent and asymptotically normal estimates.

4.5 Time-Varying Measures of IFI

In this section, the integrated form of the DCC MV-GARCH model is used to estimate the time-varying conditional correlations, which are then used to obtain time-varying measures of IFI. Although the DCC MV-GARCH model does not constrain the number of series included in the system, it assumes that the structure of time varying correlations is the same. Therefore, taking into consideration the results presented in the previous chapters of this dissertation, we will consider two systems of international financial integration. The first group will consist of the bond returns of Austria, France, Germany and Netherlands, while the second group will include the bond returns of the U.S. and U.K.

The DCC model requires the series to have an expected value equal to zero. Therefore, autocorrelation is removed from each international bond return.[43] Table 4.4 presents Engle's (1982) LM test for the presence of ARCH. All the series show evidence of a (G)ARCH effect. Table 4.5 reports the estimated parameters for each univariate GARCH(1,1) model and the Wald test H_0: $\alpha_1 + \beta_1 = 1$ against the alternative of H_1: $\alpha_1 + \beta_1 < 1$. In addition, the Q-statistics for the standardized residuals and squared standardized residuals at lag 4, 8 and 12 are reported. These Q-statistics are insignificant, indicating no autocorrelation in the residuals.

Figure 4.1 presents the estimated conditional variance for each market index. The estimated conditional variances show considerable variation over time. For the estimated conditional volatilities of Austria, France, Germany and the Netherlands we see an increase in the conditional variance during the late 1980s up to the mid-1990s. This is

[43] This process is explained in Section 4.2 and in Chapter 2 of this dissertation.

Table 4.4: The Univariate ARCH LM Tests

	Austria	France	Germany	Netherlands	U.K.	U.S.
ARCH(2)	7.29**	14.73*	13.15*	15.55*	11.52*	5.13***
ARCH(4)	12.26*	21.97*	17.93*	24.06*	12.65*	17.70*

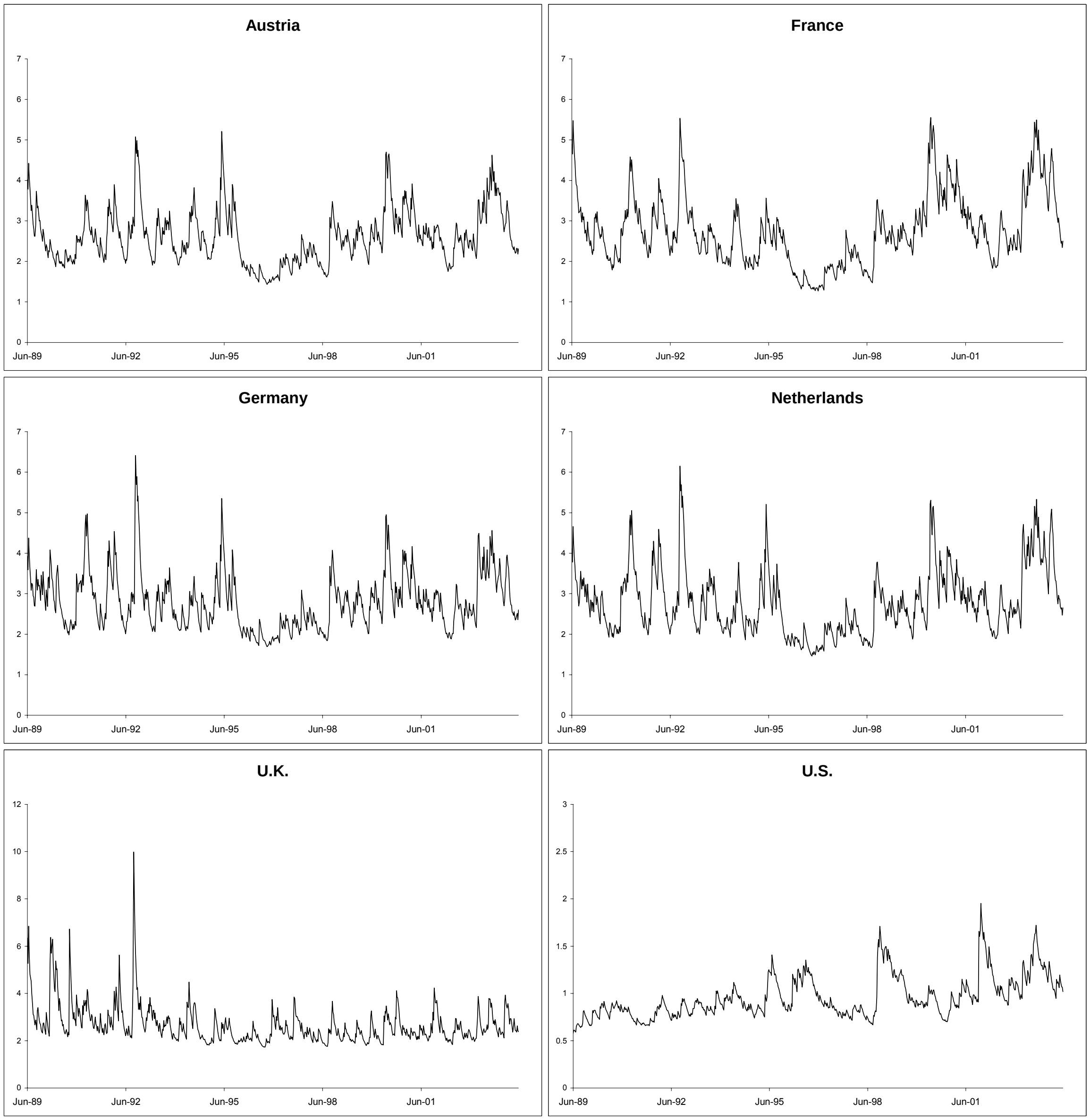
Note: *, ** and *** denote 1%, 5% and 10% significance level respectively.

Table 4.5: Univariate GARCH (1,1)

	Austria	France	Germany	Netherlands	U.K.	U.S.
GARCH(1,1): $h_t^2 = a_0 + a_1\xi_{t-1}^2 + \beta_1 h_{t-1}^2$						
α_0	0.13***	0.09***	0.23***	0.16***	0.41***	0.04***
	(0.08)	(0.05)	(0.13)	(0.09)	(0.27)	(0.02)
α_1	0.06*	0.07*	0.07*	0.07*	0.09***	0.04*
	(0.02)	(0.02)	(0.03)	(0.03)	(0.05)	(0.01)
β_1	0.89*	0.90*	0.85*	0.87*	0.76*	0.92*
	(0.05)	(0.03)	(0.07)	(0.05)	(0.14)	(0.03)
Wald Test						
$H0: \alpha_1 + \beta_1 = 1$	2.86***	2.81***	3.21***	3.02***	2.81***	2.62***
Q-Statistics						
Q(4)	3.17	3.21	3.02	3.97	1.47	2.39
Q(8)	7.58	6.75	6.26	7.42	12.91	10.06
Q(12)	15.03	10.61	11.30	12.43	15.15	11.68
$Q^2(4)$	0.84	0.80	0.78	0.84	1.84	2.16
$Q^2(8)$	2.67	4.30	2.59	3.05	2.63	4.47
$Q^2(12)$	6.47	5.69	7.07	7.20	4.37	8.15

Note: *, ** and *** indicate significance at 1%, 5% and 10% respectively.

Figure 4.1: Conditional Volatilities of Bond Returns Obtained from the GARCH (1,1) Model



followed by a relatively quiet period in the mid to late 1990s. Then there is an increase in conditional volatility at the end of the sample period. For the U.K., a high volatility characterizes the first 5–6 years of our sample period. For the U.S., the first part of our sample period can be characterized as a quiet period, compare to the second part where conditional volatility appears to have increased.

Before estimating the integrated DCC model, the null of constant conditional correlation is tested against the alternative of dynamic conditional correlations. Following Engle and Sheppard (2001), the outer product of standardized residuals from the univariate GARCH processes is regressed on constant and lagged outer products of these residuals. If the conditional correlations are constant, these residuals should be II with a variance-covariance matrix equal to an identity. Therefore, the constant and the parameters of lagged regressors should be zero. Different numbers of lags are used, ranging from 1 to 8. The null of constant correlations is rejected for both groups.

Using the standardized residuals of each univariate GARCH(1,1) process, the DCC model presented by equation (12) is estimated. The model fitted is a simple DCC(1,1) MVGARCH. The estimated DCC(1,1) parameters for the first group, that consists of the standardized residuals of Austria, France, Germany and the Netherlands, are

$$\Phi_t = (1 - 0.07 - 0.91)\bar{\Phi} + 0.07\left(\varepsilon_{t-1}\varepsilon_{t-1}^{'}\right) + (0.91)\Phi_{t-1}\,.$$

For the second group, consisting of the standardized residuals of the U.S. and U.K. bond returns, the estimated parameters of the DCC(1,1) model are

$$\Phi_t = (1 - 0.03 - 0.93)\bar{\Phi} + 0.03\left(\varepsilon_{t-1}\varepsilon_{t-1}^{'}\right) + (0.93)\Phi_{t-1}\,,$$

where $\varepsilon_t = D_t^{-1} r_t$ are the standardized residuals, Φ_t and $\bar{\Phi}_t$ are the matrices of conditional and unconditional covariances of standardized residuals.

The first column of graphs in Figure 4.2 presents the estimated DCCs for the first group of bond returns, which is the DCCs between German bond returns and each of the bond returns of Austria, France and Netherlands. Using the estimated conditional correlations between German bond returns and each of the other countries' bond returns in this group, and the estimated univariate conditional volatilities, the beta coefficients in (11) are obtained. These estimated beta coefficients are presented in the second column of graphs in Figure 4.2.

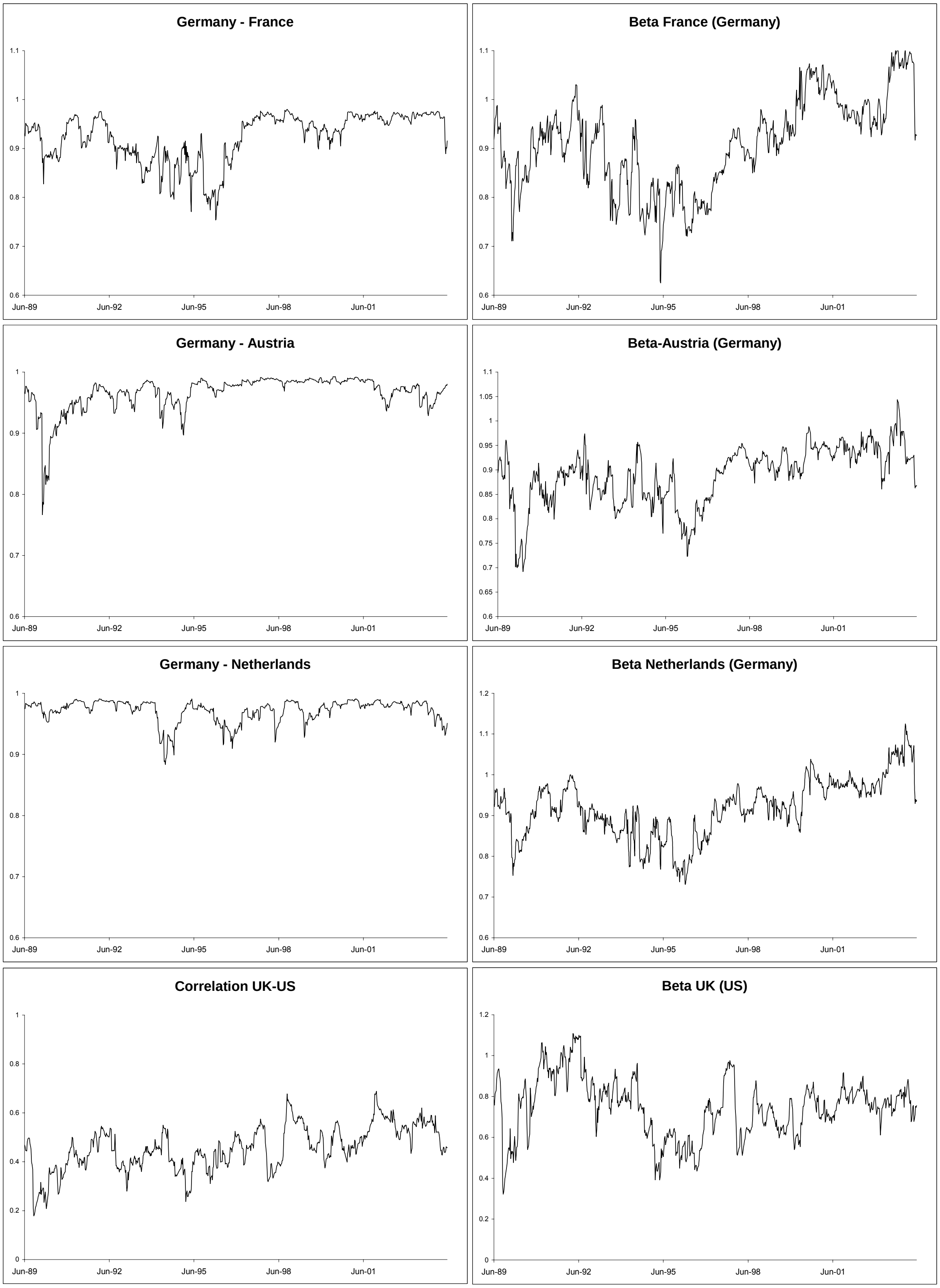
The beta coefficients, which represent the time varying measures of international financial integration of Austria, France and Netherlands bond returns with respect to the German bond returns clearly indicate an increase in the level of integration across these countries. Since the mid-1990s there appears to be a positive trend in the level of financial integration between Austria, France, Germany and Netherlands. With regards to the U.S. and U.K., the results do not indicate any clear evidence of a positive trend in the level of financial integration between these two countries. This would suggest that the increase in the level of financial integration between the U.S. and U.K. might be less than expected.

The analysis of the saving-investment relationship as well as the consumption correlations obtained in Section 4.3 of this chapter indicated a strong level of financial integration between Austria, France and Germany for the period 1970–2004. Based on these results we concluded that these countries are almost fully integrated. The time varying measures of IFI presented in Figure 4.2 indicates that the process of international financial integration is an ongoing process and that an increase in the level of IFI has become more evident during the last 7–8 years of our sample size.

While the saving-investment correlations and consumption correlations analyses failed to find strong evidence of financial integration between Germany and Netherlands, the time-varying measure of IFI presented in Figure 4.2 clearly indicates an upward trend

Figure 4.2: Estimated DCC and Time-Varying Measures of International Financial Integration

Correlations | Measures of IFI



in the level of integration between these countries starting from the mid-1990s. With respect to Germany and U.K., consistent with the results found in Section 4.3 of this chapter and the results presented in previous chapters, the degree of financial integration between these two countries is lower than the level of IFI found between Austria, France, Netherlands and Germany.

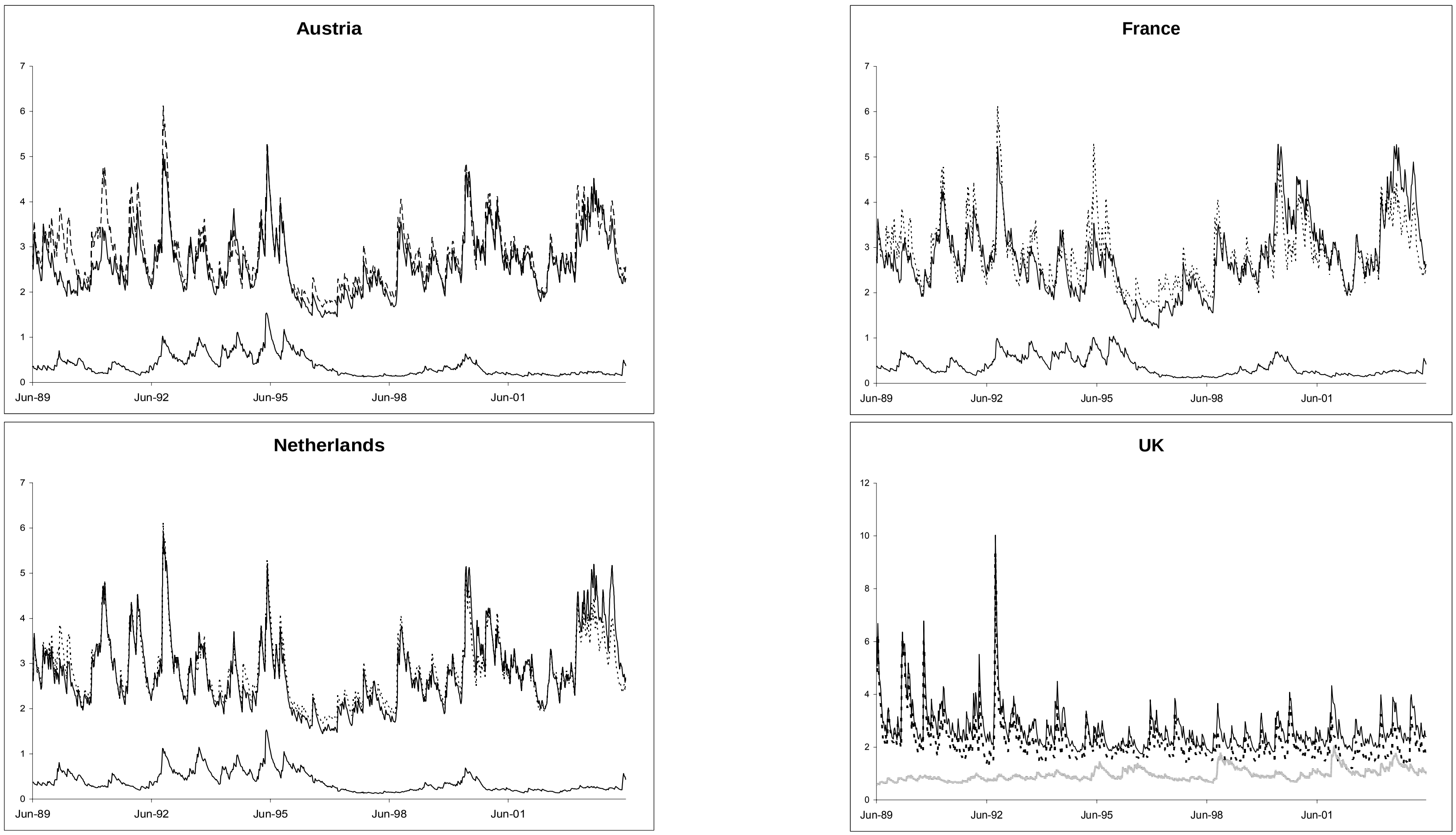
Having obtained time-varying estimates of the beta coefficients, the variances of the idiosyncratic shocks can now be calculated and their relative impact on the volatility of the bond return for each market analyzed. From (7), the variances of the idiosyncratic shocks for each market can be obtained using

$$\sigma_{i,t}^{2} = h_{i,t} - \left(\beta_{i,t}^{w}\right)^{2} h_{w,t} . \tag{17}$$

These idiosyncratic shocks are estimated for both groups of bond returns and are presented in Figure 4.3. For Austria, France and Netherlands the information presented in each of these graphs consists of the estimated conditional volatility ($h_{i,t}$), the German integration volatility (${\beta_{i,t}^{w}}^{2} h_{w,t}$), that in this case would present the global volatility impact in this group of bond markets, and their idiosyncratic volatility ($\sigma_{i,t}^{2}$). For the U.K. the information presented consist of the estimated conditional volatility of the U.K. bond returns, the U.S. integration volatility that in this case would present the global volatility impact on the U.K. bond market returns, and the U.K. idiosyncratic volatility.

For the bond market returns of Austria, France and the Netherlands, the shocks originating from the German bond market appear to be the driving force. A different picture is depicted with respect to the U.S. and U.K. bond markets. Shocks originating from the U.S. bond market do not appear to have a very strong impact on U.K. bond

Figure 4.3: Idiosyncratic Shock(s)



- - - - Idiosyncratic shocks, _______ Conditional volatility, ------- German (U.S) variance markets. For the U.K. bond market, the idiosyncratic shocks appear to be the driving force. In the literature, the common believe is that the U.K. is becoming less financially integrated with the European countries and becoming more integrated with the U.S. The results presented here indicate that the U.K. does not have a high level of integration with European countries and with the U.S. Overall the U.K. appears to have a relatively low level of integration with the rest of the world.

In conclusion, the time varying measures of international financial integration indicate a high degree of integration among the bond market returns of Austria, France, Germany, and the Netherlands. On the other hand, this time varying measure indicates that the level of integration between the U.S. and U.K. bond market returns appear to be less than expected. We should keep in mind the fact that the CAPM model that is used to obtain the time varying measure of international financial integration includes only one world factor. We should look further into this model and separate the world from regional factors. This way we will be able to analyze separately the importance of these factors.

4.6 Conclusions

The analysis of the degree of financial integration has been the focus of this dissertation. Using different measures, we have tried to determine the degree of IFI across international bond markets. In this chapter, the dynamics of IFI using a new methodology was analyzed. The DCC MV-GARCH model of Engle (2002) is used to obtain a time-varying beta coefficient in the CAPM without relying on highly parametric CAPM. In addition, the IFI results obtained from using this new method for measuring the time variation of international financial integration were compared with the results obtained from two traditional approaches of measuring IFI: saving investment correlations and cross-country consumption correlations.

The results presented in this chapter indicate a significant increase in the level of financial integration between Austria, France, Germany and Netherlands. Their level of financial integration has increased significantly, in particular beginning with the mid-1990s. The U.K. bond market does not appear to display increases in the level of integration with these European countries. The results obtained in the previous chapters indicate a relatively higher level of integration between the U.K. and the U.S. relative to the integration between U.K. and other European countries. In this chapter, although there is some evidence of a higher level of integration between these two countries, the results are not as strong. This suggests that the U.K. has a low level of integration not only with the European countries, but also with the rest of the world. This lack of integration is more evident with respect to the European countries, and one possible explanation might be the fact that these countries are undergoing a significant process of financial, economical and political integration.

CHAPTER 5

CONCLUSIONS

The empirical investigation of the degree of IFI has mainly focused on the analysis of international stock market integration. This extensive literature indicates a significant increase in the level of IFI across countries especially since the early 1990s. This literature indicates that the significant increase in the level of IFI is especially evident for European countries. The last decade has been a very important period in the integration process of the European Union. We saw the abolishment of the restrictions on capital movements between these countries and monetary policy coordination during the 1990s. This process culminated with the conduct of a single monetary policy across these countries in the late 1990s and the introduction of a common currency in January 2001. These important events have affected the degree of IFI observed across the European countries.

One important result emerges in most of the empirical research. There is no evidence of an increase in the level of IFI between U.K. stock markets and other EU stock markets. In addition, there is evidence that the U.K. stock market is becoming more integrated financially with the U.S. stock market. Therefore, the main conclusion drawn from this literature is a significant increase in the level of IFI across EU stock markets with the exception of the U.K. The U.K. differs from other EU countries and has become more integrated with the U.S.

The analysis of the degree of IFI ignores the examination of the degree of integration across international bond markets. This lack of the international bond market analysis is not because of a lesser importance of these markets relative to the international stock markets. In addition to reflecting financial and economic integration across countries, the aggregate behavior of international bond markets also reflects political coalitions. Financial integration, political integration and economic integration processes should be exhibited in the behavior observed across the international bond market, and there should be a causal interaction across these processes. However, the focus of this dissertation is to explain how financial, political and economic integration is displayed in the behavior observed across international bond market returns.

In this dissertation, the degree of international financial integration is examined by mainly focusing on the behavior of the international bond markets. While the analysis of the correlations across international stock markets has indicated a high degree of IFI, we ask the question: What does the analysis of the rolling correlations across the international bond market returns tells us with regard to the level of IFI? The analyses of these correlations indicate a significant increase in the degree of IFI. For most of the European Union countries, this increase has become significant particularly since the mid-1990s. The rolling correlation analysis indicates that most EU bond markets are almost fully integrated and that the U.S. and U.K. forms another group of financially integrated economies.

In addition to the correlations across international bond market returns, we also analyze the degree of cointegration, i.e. the existence of common trends across these market returns. While rolling correlations enable us to detect any possible linear relationship between international bond market returns, the cointegration analysis enables to detect if there is a long-run equilibrium among the returns of these international bond markets.

Do the international market returns tend to move together over time? The U.S. and U.K. bond market returns results indicate the existence of a long-run relationship confirming the findings of stock market analysis of these two countries. With regard to the EU countries, we fail to find strong evidence that all countries are fully integrated. In fact, among the EU bond markets, only the bond markets of Austria, France, Germany and the Netherlands are fully cointegrated.

We then analyze whether the relationship we find across these international bond market returns is reproduced in the output correlations of these countries. Is the higher degree of financial integration observed across international bond market returns also reflected in higher output synchronization across these countries? To our surprise, our results indicate that we do not have strong evidence of business cycles synchronization across countries.

Why do our two approaches lead to different conclusions? Krugman (1993) argues that if the increase in the financial integration is associated with a higher level of inter-industry specialization, then the production structure may actually become more vulnerable to idiosyncratic shocks. This leads to lower business synchronization across countries. Although this may serve as a possible explanation, we doubt its applicability in this case since we are dealing with some of the world’s largest economies such as the U.S., U.K., Germany, France, Austria, and the Netherlands. Krugman’s (1993) argument will typically hold for those economies that do not have diversified production structures.

However, it will be unrealistic to assume that economies such as those of the U.S., U.K., Germany, or France do not have a diversified production structure. Hence, we need another explanation for the inconsistencies we find in analyzing the correlations and cointegration across international bond market returns versus the business cycle synchronizations. This suggests that we need to look more deeply into finding a more appropriate measure of IFI.

One alternative approach to measuring IFI is to analyze the volatility across international bond market returns. Volatility per se, a measure of the risk level across these countries, has become an important tool in the analysis of international asset markets. Therefore, we test whether there is any information with regard to the level of IFI contained in the volatility of international bond market returns. As these markets become more financially integrated we would expect to see increases and decreases in volatility simultaneously across markets that are financially integrated. Therefore, while in the second chapter we considered the bond market returns and attempted to understand the degree of IFI suggested by analyzing the dependence structure between these returns, in the third chapter we asked, what is the degree of IFI suggested by the analysis of the risk level across these international bond market returns?

The common feature methodology introduced by Engle and Kozicki (1993) is used to test whether international bond markets share the same volatility process. This parallels the analysis presented earlier when we analyzed cointegration in the first moment. While Johansen's (1988) cointegration methodology analyzes the existence of a common trend in the first moments of the international bond returns, Engle and Kozicki's (1993) common feature methodology analyzes whether a common variance (cointegration in the second moment) is observed across these bond markets. We fail to find the existence of a common volatility process across all international bond markets, in particular across all EU bond markets. However, as with the results presented in Chapter 2, among EU countries, the bond markets of Germany, France, Austria and the Netherlands emerge as a group of countries sharing the same volatility process. In addition, there is evidence that the U.K. and U.S. bond markets share the same volatility process.

In Chapters 2 and 3, our approach to testing for IFI in international bond markets consisted of three methodologies: rolling correlations, cointegration of the first moment (Johansen (1988) methodology), and the cointegration of the second moment (common feature methodology). All three approaches seem to suggest that, mostly, IFI is present.

We are nonetheless doubtful that the IFI process is complete for several reasons. One, the results with respect to the output synchronization does not point to full IFI. Two, while the cointegration analyses (with respect to the first and second moments) indicates that there are two cointegrated regions, we lack details because of the missing dynamics in these approaches. IFI could be rising over time or falling over time and these methodologies do not allow us to observe these trends. These trends would have important implications for concluding the actual degree of IFI across countries. Third, while the rolling correlations allow us to observe the dynamics in IFI and how it has changed over time, this methodology is suspect. Trends in the degree of correlations over time and the changes in the volatility are confounded. Therefore, we need to contrast our results with yet another dynamic approach that can truly detect IFI.

In Chapter 4, we presented a dynamic measure of IFI, which differs from the correlation measure in that it corrects for the changes in dynamic dependence across international bond markets due to higher volatilities. While previous measures of IFI presented in Chapters 2 and 3 analyzed either the mean or the variance of the international bond market returns, by using the time-varying measure of IFI presented in Chapter 4, we were able to capture both the mean and variance of these returns as well as their offsetting values. Therefore, the time-varying measure results of IFI presented in Chapter 4 gave us a more general and comprehensive measure of financial integration.

Our findings suggest that, while there is an increase in the level of financial integration across the bond market returns of Austria, France, Germany and the Netherlands, these markets are not fully integrated. We also failed to find strong financial integration between U.K. and U.S. bond market returns as previously suggested by the cointegration and common variance analysis. These results would suggest the degree of financial integration between these countries is lower than commonly believed.

An interesting fact emerges in the analysis of the time-varying measure of IFI plots in Chapter 4 indicating a possible decline in the level of integration across EU's bond market returns during the last 2 to 3 years of our sample period. Although we cannot determine the significance of this possible decline, the changes in trends might reflect the economic developments of the EU countries. During the last three years, we have seen a slowing of the most important economies in the EU, particularly in France and Germany. These economies have found it difficult to fulfill the Maastricht Treaty criteria, especially the fiscal criteria. In addition, the entry of new members in the EU and the skepticism associated with EU enlargement might reflect the trends in IFI that we observe during the last 2 to 3 years of the sample period. On the other hand, this possible downward trend might reflect the overall recession that world economies have undergone recently, and that stagnation or decline in the level of IFI is only temporary. It will be the focus of future research to determine the trends of financial integration across EU countries, what role the U.K. economy plays within EU countries and with the U.S., and whether the role of the U.S. economy is changing over time and in what direction.

In conclusion, the results presented in this dissertation, while indicating an increase in the level of IFI across some countries, indicate that this process is not complete. The increase in the financial integration observed in business cycles synchronizations is not revealed. The results presented here indicate that there has not been a significant increase in the nominal and real convergence across countries, at least not to the extent that it is commonly believed.

How will the results presented in this dissertation impact investors' decisions in building diversified international portfolios? The argument behind building internationally diversified portfolios is that by doing so investors will not be prone to country specific shock. These investors will be able to create portfolios with lower levels of volatility and therefore these portfolios will have relative lower risk level. A common argument in the literature is that high a degree of IFI will reduce the benefits of international portfolio diversification. High level of IFI implies high levels of correlations between international assets, reducing the benefits of international portfolio diversification.

The analysis of rolling correlations between international bond market returns, in Chapter 2 of this dissertation, will lead us to believe that indeed the benefits of the international portfolio diversification has decreased significantly, especially since the mid-1990s. However, by looking at the time varying measure of IFI it is clear that the effectiveness of international portfolio diversification has not been reduced. International diversification can still bring benefits in reducing the overall risk level of international asset portfolios.

We considered the discrepancy found between the business cycle synchronization results and the degree of IFI suggested by using rolling correlations and cointegration analysis (in the first and second moments) as an indicator that these measures might not be appropriate ones in accurately measuring the degree of IFI. The underlying assumption behind this is that financial integration can have a significant impact on output markets. Therefore, we would expect that an increase in the degree of IFI enhances the synchronization of business cycles across countries However we should recognize the possibility that synchronization of business cycles may not be closely related with dynamic changes of the degree of financial integration. Whether and to what degree fluctuations in business cycles are related with dynamic changes in the international financial integration process remains an empirical question. It will be very interesting empirical research to understand how changes in output markets impact the process of IFI and vice versa.

However, while analyzing different economic factors that can have an impact on the financial integration process, we should not overlook other factors such as sovereignty and national identity. These factors can have an important impact on the process of financial integration, and one can argue that they have become more important especially between European countries. Economic and political developments in Europe indicate that as the integration process has marched forward, some have begun to view it as a threat to national values. The most recent rejection of the new EU constitution by French and Dutch voters, to a certain extent, reflects the fears of losing national identity during this process of integration. A strong financial integration can be achieved if countries do not see the process of integration as compromising their national identity, and instead they view it as an instrument in strengthening national identities. In conclusion it is appropriate to cite Timothy F. Geithner's remark, president and chief executive officer at the Federal Reserve Bank of New York, this last April at the European Commission Conference in New York:

> *As in national defense, nations are likely to want to retain a significant degree of authority over the terms and conditions that shape financial intermediation in their economy. This is appropriate, and financial integration can go a long way among nations that have very different financial systems and different preferences to balance competing regulatory objectives. But to be effective, national frameworks for financial stability are going to have to be complemented in the future with a more intensive process of cooperation.*

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