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Capital Market Integration: U.S. and Japan Equity and Debt Markets

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Abstract

This paper studies the long-run comovement of Japanese and U.S. stock and bond markets using two different cointegration tests. Unlike the previous studies, we use both the Engle-Granger cointegration test and the Canonical Cointegration Regression (CCR) method to test the long-run comovement of asset returns. The Engle-Granger cointegration tests indicate that there is little evidence on cointegration between the bond and stock markets of the two countries, which is consistent with the results found in the previous studies. Using the CCR method, however, we find more favorable evidence of comovement between the asset returns. Tests with monthly data show some evidence of cointegration between the asset return series, while with quarterly data we find that most of the time series of asset returns are cointegrated. Our empirical study presents indirect evidence of the effects of cross investments leading to more integration of asset markets.

JEL classifications: G15; C32

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1. Introduction

Stocks and bonds are two classes of assets that are available for investors interested in diversifying their portfolios. With the relaxation of cross-border investment restrictions and the resultant availability of stocks and bonds issued by firms and governments of different countries, the set of assets available for investment has increased dramatically in recent years. Based on the classic portfolio optimization models of Markowitz (1952), the benefit of diversification depends on the correlation between the returns of these two classes of assets. Even though the portfolio optimization models look at the correlation between the asset returns, with the time-varying nature of asset return correlations, it is necessary to also investigate the long run comovement of asset returns.

In this paper we investigate the long run comovement of the following categories of assets from Japan and the U.S. – equities, short and long-term government bonds and corporate bonds.[2] Japan and the U.S. are the largest two economies of the world, and the equity and bond markets of these two countries are two of the largest in their respective categories. There is considerable cross listing of Japanese equities in U.S. markets; at the same time, Japanese are one of the largest investors in the U.S. Treasury bond market. These two factors can lead to the assumption that there must be a certain degree of comovement between the asset markets of these two countries. Therefore, the main thrust of this paper is to see whether there is long run comovement between the asset returns within the national market and across the markets of the two countries.

[2] Campbell and Hamao (1992) also test the integration between the U.S. and Japanese capital markets using the predictability of monthly excess returns on U.S. and Japanese equity portfolios over the U.S. Treasury bill rate. See Campbell and Hamao (1992) for details.

From a practitioner's perspective, understanding the nature of the comovement of asset returns can result in better asset allocations.

One possible explanation for the comovement of stock and bond returns is based on the concept of "flight-to-quality."[3] When investors perceive that equity markets are going to be riskier in the future, they tend to sell their equity holdings and move the money into less risky government bond markets. If demand in the bond market increases, bond prices will go up and the bond yields will go down. On the other hand, when investors sell, both the stock prices and stock returns will go down. According to this popular notion, during times of increased uncertainty in the stock market, investors will shift their funds to bonds. Bond prices will increase, relative to the stock prices, and comovement between the stock and bond returns will become less positively correlated. In fact, this flight to quality can result in sustained negative correlation between the stock and government bond returns.

There are two main channels through which information affects the relationship between the stock and bond markets: (1) common sources of information influencing the expectations in both stock and bond markets simultaneously, and (2) sources of information that not only alter expectations in one market but spill over into another market.[4] For example, downgrading of the debt of a firm can affect both the stock price and bond price of the firm, while an increase in interest rates by the central bank may have more pronounced effect on the bond market and less so on the stock market. However, a shock in one market may cause asset rebalancing, which then spills over into another market; in this sense the information affecting one market can affect the other market and hence increase the comovement between the two markets.

[3] Barsky (1989)
[4] Kim et al. (2006)

The next step is to see if there is any comovement between the asset markets in two different countries. A possible explanation for comovement of asset markets in two countries is liberalization of regulations with respect to cross-country investments.[5] With the increase in liberalization of regulations and the consequent globalization of the capital markets, it can be perceived that comovement between equity markets must also have increased during the past twenty years. Most of the evidence is anecdotal, where certain dramatic events in a country or a group of countries create ripples that move through the entire world markets. Detailed empirical studies are required to verify this co-movement of markets.

In this paper we use the cointegration methodology to test whether there is any comovement between the stock and bond markets in Japan and the U.S. A practical reason for this study is also to test indirectly whether the sizable presence of Japanese investors in the U.S. Treasury bond market and the cross-listing of Japanese stocks in the U.S. market have contributed to long-term comovement between the equity and bond markets of these two countries. Our empirical study finds that there is systematic long-run comovement between the asset returns of Japanese and U.S. markets with quarterly data, which presents indirect evidence of the effect of cross investments leading to more integration of asset markets of the two countries.

The rest of the paper is organized as follows. Section 2 provides literature review. Section 3 describes the econometric methodology used in this study. The details of the data are described in section 4. In section 5 the results of the study are discussed, and section 6 concludes this paper.

2. Literature review

[5] Berben and Jansen (2005)

There are a few studies on the linkages of international bond markets, while there are a fairly large number of papers on the linkages of stock markets. In a largely deregulated international bond market, bond yields, which are analogous to asset prices, should to a certain degree move together. On the other hand, if bond yields are largely the resultant of the monetary policies of a given country, then these should exhibit comovements to a lesser degree. Within the bond markets it is possible to have differences between the short and long-term covariances of bond returns from different countries. The long-term covariances between the two bond markets may be high, indicating that the bond markets are integrated and hence the benefit of diversification is lower. On the other hand, the short-term covariances may be lower, which gives a higher diversification benefit.

The empirical studies on comovement of asset returns can be divided into two categories. The first group of studies looks into the comovement and linkages between the bond markets of various countries. The comovement of stock and bond markets within a country as well as across a group of countries are studies by the second group. Among the studies that belong to the first group, Perignon et al. (2007) find the U.S. bond returns share only one common factor of change in the level of domestic term structure with German and Japanese bond returns. On the other hand, Driessen et al. (2003) identify the five factors in bond return from the U.S., Japan and Germany and find that these factors are related to level and steepness of the term structure in the respective countries. Clare et al. (1995) find that the bond markets of major industrial countries have unique characteristics. They attribute this uniqueness to taxation differences, maturity structures, investment cultures, issuance patterns and institutional arrangements.[6] Mills and Mills (1991) find that the bond yields are not cointegrated, and that in the long run they are determined by the domestic fundamentals.[7] Bredin et al. (2010) find that bond returns primarily react to domestic as compared to foreign monetary policy surprises. As in Driessen et al. (2003), a possible explanation for this is that many investors prefer to invest in long-term government bonds issued by their own governments. There is a potential for reducing the portfolio risk by holding a diversified portfolio of bonds issued by different countries, which may not be fully utilized by the domestic investors.

Among the studies that look into the comovement of stock and bond returns within a country as well as across a group of countries, Shiller (1982) finds little evidence on comovement between equity prices, bond and land prices, while Campbell and Shiller (1988) show that dividend price ratios are uncorrelated with subsequent real interest rates.[8] Furthermore, Kim et al. (2006), Li (2002), Schulman and Miller (1999), Panchenko and Wu (2009), and Berben and Jansen (2009) study the relationship between the stock and bond returns of selected countries. Kim et al. (2006) find that the bond markets are largely segmented within the context of the countries studied. Li (2002) reports that major trends in stock and bond correlation are determined by the uncertainty about expected inflation. Schulman and Miller (1999) find that the correlation between the stock and bond returns are abnormally high during a period of high inflation. Panchenko and Wu (2009) find an unambiguous and robust link between emerging

[6] They further presume that the uniqueness of the bond markets might be due to the fact that the macroeconomic policies of the governments are, in general, not coordinated and also that different economies may be at different points in the economic cycles.
[7] Both Clare et al. (1995) and Mills and Mills (1991) use cointegration methodology to test the interlinkages between government bond markets of the U.K., the U.S., Germany and Japan.
[8] Both Shiller (1982) and Campbell and Shiller (1988) study the U.S. market only.

stock market integration and stock-bond return decoupling. Berben and Jansen (2009) show strong evidence of greater comovement across the board for both stock markets and government bond markets.[9]

There are two theoretical models for movement of stock and bond prices/returns to move together. One is the standard consumption based asset pricing model and the other is VAR framework. Barsky (1989) employs the standard model and reports that the changes in risk and real economic productivity growth affect the joint movement of the stock and bond prices.[10] With VAR framework, Campbell and Ammer (1993) find that the news about future excess returns account for most of the variation in excess stock returns, while the bond returns are mostly affected by news of future inflation. This explains the low positive correlation between the stock and bond returns in the post World War II era in the U.S.[11] Some studies further examine the relationship between stock and bond markets by investigating the volatility linkages of those markets while others examine the relationship between stock markets and other factors such as yield curves and a set of economic variables.[12] From the studies cited above it is unclear

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[9] See Kim et al. (2006), Li (2002), Schulman and Miller (1999), Pancehnko and Wu (2009), and Berben and Jansen (2009) for details.

[10] Similar conclusions are reached by Bekaert and Granadier (2001).

[11] Engsted and Tanggaard (2001) use the same methodology of Campbell and Ammer (1993) to Danish stock and bond markets. The results indicate that excess stock and bond returns are negatively correlated for the Danish market.

[12] DeGoeij and Marquering (2001) and Flemming et al. (1998) investigate the volatility linkages of stock, bond and money markets. The relationship of yield curves and risk premiums of stocks of eight industrialized countries is examined by McCowan (2001). Nasseh and Strauss (2000) look at the long-run relationship between the stock prices and domestic and international economic activity such as industrial production, business surveys of manufacturing orders, short- and long-run interest rates, foreign stock prices and production of six European economies. Cappiello et al. (2006) employ the AG-DCC model to analyze the behavior of international equities and government bonds. See DeGoeij and Marquering (2001), Flemming et al. (1998), McCowan (2001), Nasseh and Strauss (2000), and Cappiello et al. (2006) for details.

whether there is any real long-term comovement between the bond market returns as well as between the bond market and the stock market returns.

3. Econometric methodology

3.1. Unit root and the Engle-Granger cointegration tests

As a preliminary step, we test for a unit root in variables to check stationarity of the variables concerned. To test for a unit root (or the difference stationary process), we employ both the Augmented Dickey-Fuller (ADF) test (1979) and the Phillips-Perron (PP) test (1988) based on the following regressions:

(a) Augmented Dickey Fuller regression : $\Delta x_t = a_0 + \rho x_{t-1} + \sum_{i=1}^{m} \beta_i \Delta x_{t-i} + \eta_t$,

(b) Phillips Perron regression : $x_t = \mu + \alpha x_{t-1} + \varepsilon_t$

The difference between the two unit root tests lies in their treatment of any 'nuisance' serial correlation. The PP test tends to be more robust to a wide range of serial correlation and time-dependent heteroscedasticity. The null hypothesis is that a series is nonstationary: $\rho = 0$ in the ADF test, and $\alpha = 1$ in the PP test.

After testing for stationarity of the variables, we employ the following Engle-Granger cointegration test to estimate the long-run equilibrium relationship between two variables.

$$y_t = \alpha + \beta x_t + \varepsilon_t \qquad (1)$$

If both y_t and x_t are *I(1)*, and ε_t, the bivariate spreads between *y* and *x*, is stationary, it indicates that *y* and *x* are cointegrated order of *(1,1)*. The residual series are the estimated values of the deviations from the long-run relationship.

3.2. Canonical cointegration regression (CCR) method

We further employ Park's (1992) Canonical Cointegration Regression (CCR) method to estimate a cointegrating vector and to test cointegration between the variables in which we are interested. One reason for using CCR is that Monte Carlo simulations in Park and Ogaki (1991) have shown that the CCR estimators have better small sample properties in terms of mean squared error than Johansen's (1988) Maximum Likelihood (ML) estimators when the sample size is small and even when the Gaussian VAR structure assumed by Johansen is true. Kahn and Ogaki (1991) find that Park's tests for the null of cointegration have reasonable small sample properties.

Consider a cointegrated system,

$$y_t = X_t'\gamma + \varepsilon_t \tag{2}$$

$$\Delta X_t = \nu_t \tag{3}$$

where y_t and X_t are difference stationary, and ε_t and ν_t are stationary with zero mean. Here, y_t is a scalar and X_t is a $(n\text{-}1)\times 1$ random vector. Let

$$w_t = (\varepsilon_t, \nu_t') \tag{4}$$

Define $\Phi(I) = E(w_t w_{t-I}')$, $\Sigma = \Phi(0)$, $\Gamma = \sum_{i=0}^{\infty}\Phi(i)$, and $\Omega = \sum_{i=-\infty}^{\infty}\Phi(i)$. Here Ω is the long run covariance matrix of w_t. Partition Ω as

$$\Omega = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{bmatrix} \tag{5}$$

and partition Γ conformably. Define $\Omega_{11,2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21}$ and $\Gamma_2 = (\Gamma_{12}', \Gamma_{22}')$. The CCR procedure assumes that Ω_{22} is positive definite, implying that X_t is not itself cointegrated. This assumption ensures that $(1,-\gamma)$ is the unique cointegrating vector. The OLS estimator in equation

(2) is super-consistent because the estimator converges to γ at the rate of T (sample size) even when $\Delta x(t)$ and $u(t)$ are correlated. However, the OLS estimator is not asymptotically efficient in this case. To obtain an asymptotically efficient OLS estimator, Park suggests a transformed model:

$$y_t^* = y_t + \Pi_y' w_t \tag{6}$$

$$X_t^* = X_t + \Pi_x' w_t \tag{7}$$

Because w_t is stationary, y_t^* and X_t^* are cointegrated with the same cointegrating vector $(1, -\gamma)$ as y_t and X_t for any Π_y and Π_x. The idea of CCR is to choose Π_y and Π_x, so that the OLS estimator is asymptotically efficient when y_t^* is regressed on X_t^*. This requires

$$\Pi_y = \Sigma^{-1}\Gamma_2\gamma + (0, \Omega_{12}\Omega_{22}^{-1})', \tag{8}$$

where $\Pi_y = \Sigma^{-1}\Gamma_2$,

In practice, the long-run covariance parameters in these formulas are estimated, and the estimated Π_y and Π_x are used to transform y_t and X_t. As long as these parameters are estimated consistently, the CCR estimator is asymptotically efficient.

The CCR estimators have asymptotic distributions that can be essentially considered as normal distributions, implying that their standard errors have the usual interpretation. The *H(p,q)* tests basically apply Park's *G(p,q)* tests to CCR residuals for the null of stationarity to OLS regressions [see Park (1992) for more explanation]. The *H(p,q)* statistic converges in distribution to a χ^2_{p-q} random variable under the null hypothesis of cointegration. In particular, the *H(0,1)* statistic tests the deterministic cointegrating restriction and the *H(1,q)* statistic tests stochastic cointegration.

3.3. Bivariate Granger causality tests

We employ Granger causality tests to gain more insight into the dynamic relationship between two variables. Causality tests can provide useful information on whether knowledge of past security price movements improve forecasts of current and future movements in the other security prices, and vice versa. Formally, if the prediction of y using past x is more accurate than without using past x in the mean square error sense [i.e., if $\sigma^2(y_t|I_t) < \sigma^2(y_t|I_t - x_t)$, where I_t is the information set], we say that x Granger causes y, denoted by $x \xrightarrow{G.C} y$.

The causal relations are based on a bivariate causality test between one security performance and another security index. Therefore, to determine whether a market index Granger-cause another asset returns, or vice versa, the following system of equation is estimated:

$$\Delta y_t = a + \sum_{i=1}^{m} \alpha_i \Delta y_{t-i} + \sum_{i=1}^{m} \beta_i \Delta x_{t-i} + \varepsilon_t \qquad (9)$$

where Δy (Δx) represents the first-differenced series of a log variable. The first-differenced series (the usual definition of returns) are stationary variables, and the disturbance terms, ε_t, in equation (9) is assumed to have zero means, constant variances, and be individually serially uncorrelated. The null hypothesis is that x does not Granger-cause y if $\beta_i = 0$ for all i in (9).

4. Data

The data for this study include 312 monthly and 104 quarterly observations, covering a 27-year period from January 1984 to December 2009.[13] The major market indexes under study are stock indices, 10-year government bond and 3-month government bill total return indexes (and yields) of the United States (US) and Japan (JP) obtained from Datastream International.

[13] Japanese 3 month Treasury bill index (JP3MIL) and U.S. 3 month Treasury bill (US3MI) indexes begin from January 1986.

Most of the previous studies (e.g., Clare et al., 1995; Sutton, 2000; Smith, 2002) use monthly observations for the cointegration tests. Since cointegration tests work better with low frequency data, we also use quarterly data to examine the long-run relationship between the financial markets of the U.S. and Japan.

For the Japanese market, Nikkei 225 Stock Average Price Index is used as a proxy for the aggregate Japanese stock market performance. We use the total return indexes and yields of Japanese 10-year government bond and 3-month Treasury bill. These indexes represent the total return (including reinvested coupon payments) to the investor from a representative portfolio government bonds with the above maturities. The indexes are obtained both in local currency and U.S. dollar terms, with returns measured in both Japanese yen and U.S. dollars. In general, the results in local currency terms are more relevant if exchange rate risk is fully hedged in international investment (Barr and Priestley, 2004). By contrast, the results in U.S. dollar terms reflect the possible benefits of international bond diversification to U.S. investors (Smith, 2002). Yields on Japanese AAA and B corporate bonds are used to test the long run comovement with the U.S. corporate bond yields.[14]

For the U.S. market, Standard & Poor's 500 Price Index is used as a proxy for the aggregate U.S. stock market performance. We use the total return indexes and yields of the U.S. 10-year Treasury bond and 3-month Treasury bill as well as the yields on AAA and BAA corporate bonds.

[14] Japanese AAA and B corporate bond yields are available beginning from January 1997. For corporate bonds, we use yields on corporate bonds of the U.S. and Japan because corporate bond market indexes are not available. U.S. and Japanese AAA corporate bonds, and U.S. BAA and Japanese B rated corporate bonds are paired for this study, respectively given lack of Japanese BAA corporate bond data.

5. Empirical results

Table 1 provides a summary listing of the variables used in the study. Yields on debt securities of both Japan and the U.S. have been declining over the sample period.

Japanese and U.S. stock, government bond and bill markets behaved very differently over the sample period. The Japanese stock market peaked in the year 1989 and has been through a prolonged period of decline. . On the other hand, the U.S. stock market had a dramatic drop in 1987 and a more prolonged decline after the bursting of the internet bubble in the year 2001 and the global financial crisis in 2008. Both markets have rebounded since 2009. Japanese short term interest rates have remained close to zero for the past ten years, while the U.S. short term interests have declined to near zero in the past year.

Before we test the comovement of the markets, it is necessary to conduct more formal tests for non-stationarity of the time series. Two standard procedures that we employ for this purpose are the ADF test and the PP test. The optimal lags are selected by minimizing Akaike's (1974) information criterion (AIC) and Schwarz's (1978) Bayesian information criterion (SBIC). The null hypothesis for both procedures is that a

Table 1. Variable names and descriptions

Variable name	Variable description
NIKKEI	Nikkei 225 stock average price index

NIKKEIS	Nikkei 225 stock average price index (U.S. dollar denominated)
JP10I	Japan 10 year government bond total return index (U.S. dollar denominated)
JP10IL	Japan 10 year government bond total return index (Japanese Yen denominated)
JP3MI	Japan 3 month government bill total return index (U.S. dollar denominated) (1/3/86-12/31/09)
JP3MIL	Japan 3 month government bill total return index (Japanese Yen denominated) (1/3/86-12/31/09)
JP10Y	Japan 10 year government bond yield
JPAAA	Japan AAA corporate bond yield (1/3/97-12/31/09)
JPB	Japan B corporate bond yield (1/3/97-12/31/09)
SP	S&P 500 composite price index
US10I	U.S. 10 year government bond total return index
US3MI	U.S. 3 month government bill total return index (1/3/86-12/31/09)
US10Y	U.S. 10 year government bond yield
US3MY	U.S. 3 month government bill yield
USAAA	U.S. AAA corporate bond yield
USBAA	U.S. BAA corporate bond yield

- The sample period is from January 1984 to December 2009 (unless specified above).
- The frequencies of the variables are monthly and quarterly.

- U.S. dollar denominated Japanese indexes are obtained by dividing them by the prevailing exchange rates between the U.S. and Japan.

unit root exists and in Table 2 we report the results for these tests. The stock index return series, as well as the long-term government bond index returns for Japan and the U.S., are likely to be non-stationary on the level at the 5% significant level. On the other hand, we

Table 2. Unit root tests

Augmented Dickey Fuller [ADF] Test : $\Delta x_t = a_0 + \rho x_{t-1} + \sum_{i=1}^{m} \beta_i \Delta x_{t-i} + \eta_t$

Phillips Perron [PP] Test : $x_t = \mu + \alpha x_{t-1} + \varepsilon_t$

Variables	Monthly		Quarterly	
	ADF	PP	ADF	PP
NIKKEI	-1.5032	-1.5036	-1.6466	-1.6656
NIKKEIS	-2.4187	-2.5178	-2.2815	-2.3274
JP10I	-2.1319	-2.1202	-1.9061	-1.9295
JP10IL	-1.3873	-1.4347	-1.2947	-1.3105
JP3MI	-3.0488**	-3.0523**	-2.6119*	-2.6475*
JP3MIL	-2.7137*	-6.7555***	-2.3163	-7.6074***
JP10Y	-1.4178	-1.2682	-1.2658	-1.2813
JPAAA	-1.4873	-1.5032	-1.6076	-1.6621
JPB	-1.4861	-1.5020	-1.7732	-1.8334

SP	-1.2698	-1.2712	-1.3901	-1.4071
US10I	-2.0013	-2.0093	-2.2105	-2.2376
US3MI	-2.9009**	-4.6639***	-1.9216	-5.9459***
US10Y	-1.7362	-1.5624	-1.8241	-1.8464
US3MY	-1.6197	-1.0999	-2.6269*	-1.4257
USAAA	-1.5627	-1.3344	-1.6799	-1.7005
USBAA	-1.5006	-1.3379	-1.6057	-1.7005

- The numbers in this table represent the *t*-statistics of the estimates of ρ and α, respectively.
- Critical values of *t*-statistic with 100 (250) observations are 10%, -2.58(-2.57); 5%, -2.89(-2.88); and 1%, -3.51(-3.46), respectively. [Fuller (1976), Table 8.5.2, pp. 371-373]
- Lag (*m*) in the Phillips-Perron (PP) test is the number of lags included in the calculation of autocovariances of ε_t. The details of the adjusted *t*-statistic are referred to Phillips and Perron (1988).
- *, **, and *** denote significance at the 10%, 5% and 1% levels, respectively.
- Log level variables are used for the unit root tests.

reject the null of unit roots for the series of the three-month Treasury bill returns of Japan and the U.S. at the 5% level of significance. One of the possible explanations for the stationarity of the Japanese short-term bill returns can be the low interest rates that prevailed in Japan for most of the latter half of the period covered in this study. All of the returns series seem to be stationary in their first difference (these results are available on request) at the 5% significant level. These results are similar to those of previous studies (e.g., DeGennaro et al., 1994; Clare et. al., 1995; Smith, 2002).

The results of the Engle-Granger cointegration tests are given in Table 3.[15] When two variables are non-stationary, it is necessary to test the cointegration between the two variables to find whether these two variables exhibit any long run comovement. The test is based on testing the (non)stationarity of the residuals from a cointegrating regression. If there were no cointegration, there would be no long-run relationship binding the series together, so that the series could wander apart without bound. The null of no cointegration cannot be rejected in most cases for the whole sample period at the 5% (10%) significant level with the exception of quarterly Japanese 10-year government bond and the U.S. 10-year government bond returns, Nikkei stock index and Japanese 3-month government bill return index in the Japanese markets, and the long term government bond total return and yield in both U.S. and Japanese markets. The result from the Engle-Granger test for cointegration is consistent with DeGennaro et al. (1994) and Clare et al. (1995) in showing that international government bond return indices are

Table 3. Engle-Granger cointegration tests

Engle-Granger cointegration tests between *y* and *x* are performed. The bivariate spreads between *y* and *x* are represented by ε_t. If ε_t is stationary, it indicates that *y* and *x* are cointegrated.

$$y_t = \alpha + \beta x_t + \varepsilon_t$$

Dept. Var.	Indep. Var.	Monthly		Quarterly	
		ADF	PP	ADF	PP
JAPAN-U.S.					

[15] For the regression to be a cointegrating regression, the time series variables of interest need to be unit root nonstationary in this study. It seems, however, more reasonable to assume that they are stationary but have autoregressive roots near one. In that case, as Elliot (1998) points out, the point estimates can still be expected to be fairly precise and we can obtain more efficient cointegrating estimators than OLS because these estimators are asymptotically consistent and have smaller biases than OLS.

NIKKEI	SP	-2.2582	-2.1779	-2.2432	-2.2636
NIKKEIS	SP	-2.2422	-2.2375	-2.1188	-2.1448
JP10I	US10I	-2.3588	-2.3682	-3.2340*	-2.4405
JP10IL	US10I	-2.3838	-2.4034	-2.5362	-2.5673
JP3MI	US3MI	-2.4715	-2.4824	-2.4348	-2.3864
JP3MIL	US3MI	-2.7211	-0.7823	-2.9729	-0.7859
JP10Y	US10Y	-2.9771	-2.9891	-3.1147*	-3.1529*
JPAAA	USAAA	-1.9411	-1.9618	-1.7557	-1.8153
JPB	USBAA	-1.7179	-1.7363	-1.9042	-1.9688
JAPAN					
JP10IL	NIKKEI	-3.0065	-3.0065	-2.7838	-2.8180
JP3MIL	NIKKEI	-4.3646***	-4.3839***	-3.7144**	-3.7649**
JP10IL	JP10Y	-2.5903	-2.9249	-3.0423*	-3.0797*
JP10Y	NIKKEI	-2.5666	-2.5768	-2.6307	-2.6630
U.S.					
US10I	SP	-2.0543	-2.0626	-2.2948	-2.3230
US3MI	SP	-1.7033	-1.7108	-1.8472	-1.8724
US10Y	SP	-2.5307	-2.5408	-2.9700	-3.0064
US10I	US10Y	-2.9225	-2.9343	-3.1221*	-3.1604*

- For the cointegration test of $\varepsilon_{t,}$ critical values with 100(200) observations are 10%, -3.03(-3.02); 5%, -3.37(-3.37); and 1%, -4.07(-4.00), respectively [Engle and You (1987) Table pp.157]
- ε_{it} denote the residual sequences from the above long-run relationships between *y* and *x* variables.
- *,**, and *** denote significance at the 10%, 5% and 1% levels, respectively.
- Log level variables are used for cointegration tests.
- NIKKEIS is obtained by dividing NIKKEI by EXRATE.

not cointegrated, but contradicts the results of Barassi et al. (2001) and Smith (2002).[16] As shown by Clare et al. (1995), the lack of a long-run relationship may be due to the existence of many barriers to market access in international bond markets, such as heterogeneous taxation and maturity structure, investment culture, and international arrangements. Furthermore, this also can be explained by a small sample bias and a lack of power of the Engle-Granger test because it is known that the Engle-Granger method suffers from a low power in a small sample and a possible simultaneous equation bias.

The CCR method is further conducted as one more robustness check because the CCR estimators have better small sample properties than the Engle-Granger test estimators. Tables 4 and 5 represent the CCR results using monthly and quarterly data, respectively. The results present a somewhat different picture. With monthly data, the deterministic cointegrating restriction is not rejected in 3 out of the 9 cases for Japanese and the U.S. pairs at the 5% significant level. In the Japanese market, only the null hypothesis of deterministic cointegration is not rejected in 2 out of 4 cases, while in the U.S. market, only one out of 4 cases rejects the null hypothesis at the 5% level of significance. On the other hand, the null of stochastic cointegration is not rejected in 3 out of the 9 cases for Japanese and U.S. pairs at the 5% level. In the Japanese market, the stochastic cointegrating restriction is not rejected in 3 out of 4 cases, and the same is true for the U.S. market at the 5% level of significance.[17]

[16] The different findings on cointegration between this study and previous studies might also be due to the difference in sample periods, different sets of markets under consideration, and different proxies for bond markets.
[17] For the terms of stochastic and deterministic cointegration, refer Park (1990).

It is very interesting to see that the CCR results are more favorable with quarterly data than with monthly data. With quarterly data, the deterministic cointegrating

Table 4. CCR results (Monthly)

$Y_t = \alpha + \beta X_t + \varepsilon_t$ for CCR

		CCR			
Dep. Var.	Ind. Var.	H(0,1)	H(1,2)	H(1,3)	H(1,4)
JAPAN-U.S.					
NIKKEI	SP	0.3403 (0.5597)	1.4673 (0.2258)	8.2376 (0.0163)**	8.3312 (0.0396)**
NEKKEIS	SP	0.0002 (0.9895)	1.4972 (0.2211)	3.5724 (0.1676)	3.5834 (0.3101)
JP10I	US10I	12.7161 (0.0004)**	0.5733 (0.4489)	19.707 (0.0001)**	23.294 (0.0000)**
JP10IL	US10I	8.2392 (0.0041)**	0.0129 (0.9096)	10.912 (0.0043)**	12.637 (0.0055)**
JP3MI	US3MI	5.0485 (0.0246)**	0.0095 (0.9224)	1.7993 (0.4067)	2.5232 (0.4711)
JP3MIL	US3MI	8.3043 (0.0040)**	2.2669 (0.1322)	15.025 (0.0005)**	15.085 (0.0017)**
JP10Y	US10Y	22.984 (0.0000)**	2.2103 (0.1371)	19.265 (0.0001)**	20.973 (0.0001)**
JPAAA	USAAA	14.226 (0.0002)**	7.0309 (0.0080)**	7.1878 (0.0275)**	8.4665 (0.0373)**
JPB	USBAA	0.6922 (0.4054)	0.0500 (0.8231)	0.1981 (0.9057)	0.7829 (0.8535)
JAPAN					
JP10IL	NIKKEI	5.0607	02719	11.444	13.558

		(0.0245)**	(0.6021)	(0.0033)**	(0.0036)**
JP3MIL	NIKKEI	0.0607	1.7983	4.6059	4.6340
		(0.8054)	(0.1799)	(0.1000)	(0.2006)
JP10IL	JP10Y	5.7415	0.4701	1.2743	1.4590
		(0.0166)**	(0.4929)	(0.5288)	(0.6918)
JP10Y	NIKKEI	1.6561	0.5821	0.6264	0.7269
		(0.1981)	(0.4455)	(0.7311)	(0.8669)
U.S.					
US10I	SP	1.7031	0.1120	3.7938	3.8158
		(0.1919)	(0.7379)	(0.1500)	(0.2821)
US3MI	SP	0.0417	1.3256	2.8215	2.8264
		(0.8382)	(0.2496)	(0.2440)	(0.4192)
US10Y	SP	2.9164	3.8382	9.2249	16.168
		(0.0877)**	(0.0501)**	(0.0099)**	(0.0010)**
US10I	US10Y	4.4107	0.2059	0.9367	1.6139
		(0.0357)**	(0.6500)	(0.6260)	(0.6562)

Note: Numbers in parenthesis are *p*-values. For instance, when *p*-value is less than 0.05, we reject the null hypothesis of cointegration at 5% level of significance.

** denotes significance at the 5% level.

The *H(0,1)* statistic tests the deterministic cointegrating restriction and the *H(1,q)* statistic tests stochastic cointegration.

Table 5. CCR results (Quarterly)

$Y_t = \alpha + \beta X_t + \varepsilon_t$ for CCR

		CCR			
Dep. Var.	Ind. Var.	H(0,1)	H(1,2)	H(1,3)	H(1,4)
JAPAN-U.S.					
NIKKEI	SP	9.7782 (0.0018)**	1.6373 (0.2007)	6.5567 (0.0377)**	10.120 (0.0176)**
NEKKEIS	SP	0.5202 (0.4708)	0.3637 (0.5464)	0.5289 (0.7676)	0.6089 (0.8944)
JP10I	US10I	4.2238 (0.0399)**	02660 (0.6060)	0.4517 (0.7978)	5.7989 (0.1218)
JP10IL	US10I	0.0485 (0.8256)	0.5298 (0.4667)	0.5504 (0.7594)	1.4449 (0.6950)
JP3MI	US3MI	0.9214 (0.3371)	0.7979 (0.3717)	0.9018 (0.6370)	1.1672 (0.7609)
JP3MIL	US3MI	0.8488 (0.3569)	0.3836 (0.5357)	0.3909 (0.8225)	0.5900 (0.8987)
JP10Y	US10Y	7.2919 (0.0069)**	3.0253 (0.0820)**	11.965 (0.0025)**	13.990 (0.0029)**
JPAAA	USAAA	0.4401 (0.5071)	0.0673 (0.7953)	0.0866 (0.9576)	0.0961 (0.9923)
JPB	USBAA	5.0543 (0.0246)**	2.0326 (0.1540)	2.0355 (0.3614)	2.4125 (0.4913)
JAPAN					
JP10IL	NIKKEI	10.716 (0.0011)**	1.3211 (0.2504)	3.4422 (0.1789)	4.7065 (0.1946)
JP3MIL	NIKKEI	8.6238 (0.0033)**	1.9346 (0.1643)	6.3415 (0.0420)**	9.5866 (0.0224)**

JP10IL	JP10Y	3.3848 (0.0658)**	0.4447 (0.5049)	0.4919 (0.7820)	1.0087 (0.7991)
JP10Y	NIKKEI	0.4415 (0.5064)	0.4591 (0.4980)	0.6392 (0.7264)	0.6915 (0.8752)
U.S.					
US10I	SP	0.0093 (0.9230)	0.4818 (0.4876)	1.2095 (0.5462)	3.6459 (0.3023)
US3MI	SP	0.1607 (0.6885)	0.6901 (0.4061)	0.9952 (0.6080)	1.0619 (0.7863)
US10Y	SP	0.7939 (0.3729)	0.3475 (0.5555)	0.3767 (0.8283)	0.4507 (0.9296)
US10I	US10Y	0.8298 (0.3623)	0.4171 (0.5184)	0.4446 (0.8007)	0.5626 (0.9049)

Note: Numbers in parenthesis are *p*-values. For instance, when *p*-value is less than 0.05, we reject the null hypothesis of cointegration at 5% level of significance.

** denotes significance at the 5% level.

The *H(0,1)* statistic tests the deterministic cointegrating restriction and the *H(1,q)* statistic tests stochastic cointegration.

restriction is not rejected in 5 out of 9 cases for Japanese and U.S. pairs. In the Japanese market, the null hypothesis of deterministic cointegration is not rejected in 2 out 4 cases. Furthermore, we fail to reject the null in all cases in the U.S. market. On the other hand, the null hypothesis of stochastic cointegration is not rejected in all cases. According to this result, as shown in Mark (1995), it is likely that we have more favorable results in terms of cointegration between stock and bond markets in the U.S. and Japan at the longer horizon rather than at the shorter horizon data.

The results of the bivariate Granger causality tests are presented in Table 6. The causality model requires the determination of the appropriate lag structure in the equation. We use both

AIC and SBIC in conjunction with analyzing the model's residuals to select the appropriate lag structure.

The results show that there is some causal linkage between the monthly U.S. and Japanese bond markets at the 5% significant level, which is consistent with the result of the Engle-Granger causality test. That is, U.S. 10-year Treasury bond returns Granger-cause Japanese 10-year government bond returns, but not vice versa. There is a bi-directional causation between the U.S. and Japan's 10-year government bond yields in monthly basis. Nikkei stock average returns help predict future Japanese 10-year government bond returns (and yields). In the U.S., the causal relation is found from quarterly 10-year government bond returns (and yields) to the S&P 500 stock returns. However, the overall results suggest the Granger causal relationships are not pronounced. One possible explanation for this is that these tests may not be efficient in dealing with small samples. Another is that Granger-causality analysis may fail to find stronger causal relationships because the appropriate time interval over which to investigate causality

Table 6. Granger causality tests

$$\Delta y_t = a + \sum_{i=1}^{m} \alpha_i \Delta y_{t-i} + \sum_{i=1}^{m} \beta_i \Delta x_{t-i} + \varepsilon_t$$

$H_0 : \beta_i = 0$, for $\forall\, i$ (x does not Granger-cause y.)

Dept. Var.	Indep. Var.	Monthly			Quarterly		
		F-statistic	*p*-value	$x \xrightarrow{G.C.} y$	*F*-statistic	*p*-value	$x \xrightarrow{G.C.} y$
JAPAN-U.S.							
NIKKEI	SP	0.3200	0.5781	No	0.1764	0.6765	No
SP	NIKKEI	0.0159	0.8995	No	0.4577	0.5007	No

JP10I	US10I	8.8098	0.005***	Yes	0.0179	0.8940	No
US10I	JP10I	0.1596	0.6899	No	0.0131	0.9091	No
JP3MI	US3MI	0.4437	0.7220	No	0.9967	0.3215	No
US3MI	JP3MI	0.0480	0.9860	No	0.0147	0.9039	No
JP10Y	US10Y	5.0158	0.0260**	Yes	0.0943	0.9596	No
US10Y	JP10Y	4.5426	0.0340**	Yes	0.7526	0.3883	No
JPAAA	USAAA	0.2867	0.5936	No	1.1973	0.2835	No
USAAA	JPAAA	1.1128	0.2943	No	0.2193	0.6433	No
JPB	USBAA	0.0150	0.9027	No	1.8725	0.1825	No
USBAA	JPB	0.7049	0.4033	No	0.1246	0.7269	No
JAPAN							
JP10IL	NIKKEI	9.7572	0.0019***	Yes	6.7189	0.0114***	Yes
NIKKEI	JP10IL	1.4131	0.2357	No	0.5699	0.4525	No
JP3MIL	NIKKEI	2.0359	0.1098	No	0.5449	0.4628	No
NIKKEI	JP3MIL	0.2995	0.8258	No	0.0068	0.9345	No
JP10IL	JP10Y	0.4687	0.4942	No	0.3935	0.5323	No
JP10Y	JP10IL	1.4000	0.2378	No	0.3997	0.5291	No
JP10Y	NIKKEI	12.239	0.0005***	Yes	5.3841	0.0229**	Yes
NIKKEI	JP10Y	0.6484	0.4215	No	0.0018	0.9659	No
U.S.							
US10I	SP	2.1027	0.1483	No	0.1220	0.7278	No
SP	US10I	2.4726	0.1171	No	3.6773	0.0589**	Yes
US3MI	SP	0.9175	0.4332	No	0.0942	0.7598	No
SP	US3MI	0.2353	0.8717	No	0.2048	0.6523	No

US10Y	SP	0.4849	0.4868	No	0.8363	0.3632	No
SP	US10Y	1.0866	0.2982	No	2.8234	0.0968*	Yes
US10I	US10Y	2.2012	0.1391	No	4.4759	0.0375**	Yes
US10Y	US10I	12.743	0.0004***	Yes	5.6161	0.0202**	Yes

- *, **, and *** denote significance at the 10%, 5% and 1% levels, respectively.
- The first difference in the log level variables are used for Granger-causality tests. "Yes (No)" indicates presence (absence) of causality with a p-value of equal or less than 0.10.

may be shorter than a month. If investors respond more quickly to asset returns, it may not be possible to observe Granger-causality using monthly or quarterly data. It is notable, however, that we still find evidence of Granger-causality using monthly and quarterly data, which suggests that monthly and quarterly time intervals may not be inappropriate for some asset categories.

6. Conclusions

The long run comovement of asset returns is a topic of considerable interest to both academicians and practitioners. Previous studies have looked into the linkages between asset markets within a country as well as across countries. In this study we test the long run comovement of stock and bond returns of Japan and the U.S. using both the Engle-Granger cointegration test and the Canonical Cointegration Regression method. The Engle-Granger cointegration tests show limited cointegration between the asset markets of the two countries, which is consistent with some of the earlier studies. But, unlike the previous studies, tests using the CCR method show evidence of comovement of asset returns of the two countries. The evidence of cointegration is especially strong when we use quarterly data. With small samples, as in this study, the CCR method is known as the better method for testing cointegration than the standard Engle-Granger method. Hence, it is reasonable to follow the results of the CCR method to draw conclusions of this study.

One of the key results of this study is the evidence of comovement between the asset returns of the two markets during the time period covered. The implication of this result is that static correlation assumptions used in many portfolio optimization models can still be valid as the asset returns are showing comovement. The second key result is that the comovement of asset returns seems to be more pronounced with the quarterly data as compared to the monthly data. As in Mark (1995), a possible explanation for this result is that while monthly asset returns tend to be dominated by noise, this noise is apparently averaged out over time. Thus systematic comovements between asset returns are likely to be more predictable with longer horizon data. Overall the results of this study can be construed as indirect evidence of the effect of cross investments leading to more integration of asset markets of the two countries.

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