Stocks and Bonds: Eggs in the Same or Different Baskets. A Cointegration Analysis

Niall O' Sullivan*

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Abstract
The Johansen cointegration testing and estimation procedure is applied to examine the relationships among the stock markets, government bond markets and credit bond markets of the US, UK, Europe and Japan over the period 1985M1:2002M4. Asset class relationships are examined with returns denominated in dollars, sterling, euro and yen to determine whether long run diversification gains were achievable by international investors with these as base currencies. Cointegrating relations among currency hedged returns are also investigated. Cointegration findings, and by inference long run diversification opportunities, are found to be highly sensitive to the choice of currency in which returns are denominated and to whether currency risk is hedged, revealing the important role of exchange rates in international portfolio diversification.

Keywords: Cointegration, diversification, asset classes

JEL Classification: G11, G12, G15

* Department of Economics, University College Cork, Cork, Ireland. Email: niall.osullivan@ucc.ie

1. Introduction

The globalisation of capital markets has intensified the search for diversification opportunities among financial assets by portfolio managers worldwide. Many studies have examined the level of cointegration between domestic and international markets within the same asset class such as stock, bonds, currencies and futures. A finding of cointegration is used to infer that there are no diversification gains to be made between the assets tested.[1] The long run nature of inter-asset class relationships has attracted comparatively less attention. Furthermore, the nature of inter-asset class relationships may be sensitive to the base currency in which returns are denominated and to whether the currency risk faced by investors is hedged.

This paper attempts to shed some light on diversification opportunities by testing for cointegration between the international stock markets, government bond markets and credit bond markets of the United States (US), United Kingdom (UK), Europe and Japan[2]. In addition, cointegration between these three asset classes within each of the four regions is investigated. Further insight into the possible diversification benefits obtainable by investors with different base currencies is gained by re-examining each of these relationships separately with all returns redenominated in each of the four currencies: dollar, euro, sterling and yen. Finally, as many investors opt to hedge the currency risk associated with international investments in government bonds and credit bonds, we retest for cointegration between the four government bond markets and between the four credit bond markets with local currency returns hedged into dollars, euro, sterling and yen.

Empirical evidence indicates that strategic asset allocation, rather than asset selection or market timing, is the dominant determinant of portfolio returns in the long run, (Bogle (1994) and Blake Lehmann and Timmermann (1999)). Therefore, insight into the nature

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1 Cointegration among financial assets is also often used as evidence against market efficiency owing to the predictability implied by an error correction mechanism (ECM) representation. However, unless the asset returns under consideration are risk adjusted this may not be a valid inference.

2 For ease of ease description throughout, markets are described in terms of these four ‘regions’. However, the credit bond data include eurobonds and foreign government bonds.

of the long run relationships between asset classes is particularly important for fund managers with long term investment horizons who often operate under regulatory restrictions regarding the level of risk to which their funds can be exposed. This insight is also important for financial institutions which hold bonds for the long term as part a capital requirement with a regulatory authority and/or their own balance sheet purposes.

Studies examining long term relationships between international stock indices have produced mixed findings.[3] Using the Engle and Granger (1987) cointegration methodology, Taylor and Tonks (1989) find that the UK stock market became cointgerated with that of Germany, Japan and the Netherlands after the abolition of UK exchange controls in 1979, providing evidence that there are no long run diversification gains for the UK investor in these markets. Cotsomitis, Kwan and Sim (1995) and Kasa (1992) also find evidence in support of long run integration between major world stock markets. Andrade, Clare and Thomas (1991) test a multivariate Johansen system comprising the sterling adjusted returns of the UK, US, German and Japanese stock markets but fail to find evidence of a cointegrating vector. This contrast in findings between Andrade, Clare and Thomas (1991) and Taylor and Tonks (1989) is further examined by Allen and MacDonald (1995) who cite the use of the Engle and Granger and Johansen cointegration testing procedures respectively by these two studies as the source of the contrasting findings. Allen and MacDonald (1995) and Richards (1995) apply both the Engle and Granger two-step procedure and the Johansen procedure in testing for cointegration between 16 of the world's major stock market indices but find little evidence to support a cointegration finding. Hall (1991) along with Richards (1995) caution that Johansen cointegration test results are sensitive to the VAR lag length specification and, linked to this, the frequent use by researchers of asymptotic rather than small sample critical values in the procedure, which quickly uses up degrees of freedom.

The conclusions from studies of cointegration between government (sovereign) bond markets tend on the whole to be more consistent. Clare, Maras and Thomas (1995) and

3 Direct comparison between studies is complicated by the use of different sample periods, different frequencies of returns, the currency denominations of returns and perhaps most importantly different cointegration testing procedures such as Engle and Granger (1987) and Johansen (1988).

DeGennaro et al (1994) fail to find supporting evidence of cointegration between international bond markets while Levi and Lerman (1988) conclude that bonds provide an important diversification opportunity for a US stock market investor. However, Burik and Ennis (1990) question whether the currency risk in a foreign bond portfolio is compensated in bond returns and argue that hedging the return in the forward market negates against the diversification benefit.

In addition to examining the gains to a sterling based investor from diversifying between domestic and international stock markets and between domestic and international bond markets, Andrade, Clare and Thomas (1991) test for cointegration between stocks and government bonds within the domestic markets of the US, Germany and Japan. The authors find no evidence of cointegration when the Johansen procedure is applied to either the 4 stock markets, the 4 bond markets or the 8 variable system of both stocks and bonds. In bivariate tests of domestic stocks and bonds, the authors find evidence of cointegration in the case of Japan only. Thus overall, Andrade, Clare and Thomas (1991) find much evidence in support of diversification opportunities between international stock markets and international government bond markets and between stocks and bonds within each domestic market.

2. Data Description

The stock indices used in this study are Morgan Stanley Capital International monthly total return indices with reinvested gross dividends. The indices cover at least 60% of each market in terms of capitalisation and track various sectors so as to be as representative of the total market as possible. The bond indices are Schroder Salomon Smith Barney capitalisation weighted (one year or more to maturity) monthly total return indices, with cashflows reinvested at spot money market rates until the end of each month. Government bonds are domestic sovereign bonds. Credit bonds include corporate bonds (industrial, utilities, financials), agency bonds and foreign government bonds and include eurobonds. The euro area is treated as a single unit. Euro indices combine member state indices, weighted by member state capitalisation and are measured in ecu pre-January 1999. All indices used in this study cover the period 1985M1:2002M4. All variables are in natural logs. Variable mnemonics in the tables of results are illustrated by the following examples: ECD = *European Credit* bonds denominated in *Dollars*, UKSY = *UK Stock* returns denominated in *Yen*, USGE = *US Government* bond returns denominated in *Euro* etc. Hedged returns in government bond and credit bond markets are calculated as follows: at the start of each month the cashflow to a bond portfolio in local currency terms for that month is predicted and converted to base currencies by one-month forward currency contracts.[4] The effect of hedging returns in stock markets is not examined as predictions of monthly cashflows may not be sufficiently accurate for a meaningful exercise.

3. The Johansen Procedure

This work applies the Johansen (1988) methodology to test for cointegration between the asset class returns described above

Generalising the multivariate augmented Dickey-Fuller test to allow for a higher order autoregressive process a VAR system of n variables of lag order p may be written as

(1) $$x_t = \sum_{i=1}^{p} A_i x_{t-1} + \varepsilon_t$$

where x_t is an (nx1) vector of I(1) variables and A_i are (nxn) matrices of parameters. The Johansen method is based on the Error Correction Representation of this VAR(p) model with equation (1) written as:

(2) $$\Delta x_t = \sum_{i=1}^{p-1} \pi_i \Delta x_{t-i} + \pi x_{t-p} + \mu + \delta D_t + \varepsilon_t$$

with $\pi = -\left[I - \sum_{i=1}^{p} A_i \right]$ and $\pi_i = -\left[I - \sum_{j=1}^{i} A_j \right]$

5. We assume a relatively short holding period when choosing to use one-month forward currency contracts in calculating hedged returns. Holding periods longer than one-year may be an unreasonable assumption to impose while alternative forward contracts between one-month and one-year are unlikely to substantially alter the conclusions drawn here.

where, I is an (nxn) identity matrix, μ is an (nx1) vector of constants to model a possible drift term, π is an (nxn) matrix of the parameters describing the cointegrating relations of x_t, where cointegration exists. The rank of π, or the number of non-zero characteristic roots of π, equals the number of cointegrating vectors of the system of variables x_t. If rank $\pi = 0$ then π is the null matrix and the system reduces to a VAR in first differences. If π has full rank then all n variables in x_t are stationary (and all linear combinations are therefore also stationary). The Johansen ML procedure estimates equation (2) subject to the hypothesis that rank $\pi = k$, $0<k<n$, where π can be factorised as two (nxk) matrices, $\pi = \alpha\beta'$. β is the matrix of cointegrating parameters and each row of α contains the weights attaching to each cointegrating vector in each of the n equations of the VAR. D_t represents possible dummy variables.

As a first step in the cointegration analysis the Augmented Dickey-Fuller (ADF) unit root test was applied to ascertain the order of integration of the bond and stock market indices. Where the plot of a variable suggested the presence of a linear trend component, a trend term was specified in the ADF test in order to allow for trend stationarity in the alternative hypothesis of the test. Lag orders of up to 12 were tested in each ADF equation in order to whiten the error terms, as independent homoscedastic errors are required by the test. The ADF test results for a total of 73 series indicate that the null hypothesis of a unit root cannot be rejected in all indices, except the index of Japanese stock market returns denominated in dollars, euro and sterling.

4. Empirical Results

In this section the Johansen procedure is applied to the VAR model in Eq (2) and results are reported in table 1 to table 10. In all estimations a constant was included in the VAR to allow for a linear trend component in the data. The Akaike Information Criterion (AIC), Schwarz Bayesian Criterion (SBC) and likelihood ratio tests using small sample adjusted critical values were used to determine appropriate lag lengths in the VAR. These were further determined by the requirement of serial independence in the system's residuals. However, each cointegration test was conducted with a number of alternative

VAR lag lengths in order to examine the robustness of findings. Both the maximum eigenvalue and trace cointegration test statistics are presented in the tables.

4.1 International Stock Markets

Table 1 presents the results of the Johansen testing procedure applied to the stock markets of the US, UK, Europe and Japan with returns denominated in local currencies. The null hypothesis of zero cointegrating vectors fails to be rejected by both the maximum eigenvalue and trace statistics at the 95% critical value for a VAR (1) specification. This no cointegration finding proved robust with respect to higher order VAR specifications. The lack of support for a cointegrating relationship between these four major world stock markets initially suggests that portfolio diversification benefits were available to international investors over the period. However, a markedly different conclusion emerges when returns are redenominated from local currencies, into sterling, into dollars, into euro and into yen to reflect the view point of investors with one of these as their base currencies.

Evidence in support of a contegrating relationship between the international stock markets is found in the case of all four separate currency denominations. Table 2 presents the results with the returns of the four stock markets denominated in sterling. The series of Japanese stock market returns expressed in sterling, i.e. JSS, is an I(0) series. In this case, the JSS variable is omitted from the cointeration analysis but is initially specified in the VAR as an exogenous variable and its significance tested in order to determine its final inclusion or omission. In this case, the Japanese stock market variable proved significant, although only results pertaining to the Error Correction Mechanism are presented in table 2.[5] The results indicate the existence of a single cointegrating vector and this conclusion proved robust with respect to VAR lag length. This finding for the period 1985M1:2002M4 contrasts with that of Andrade, Clare and Thomas (1991) for the period 1972M12:1990M9, with cointegration possibly arising here from a globalisation effect in the later period. Single cointegrating vectors between the four stock markets

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5 The Japanese stock market returns expressed in dollars and in euro, i.e. JSD and JSE respectively, are also I(0) series and a similar procedure was followed when testing for cointegration between the four stock markets with returns denominated in dollars an euro.

were also found with returns denominated in dollars, in euro and in yen (results not shown). As no cointegration was found with returns expressed in local currencies, this finding clearly underlines the importance of the role of exchange rates in assessing the gains from international portfolio diversification in stock markets. Table 2 also shows the constant terms and speed of adjustment parameters in the ECM along with the constant term in the cointegration space.

4.2 International Government Bond Markets

The government bond markets of the US, UK, Europe and Japan are also tested for cointegrating relations with results presented in table 3 for returns expressed in local currencies. Similar to the results from stock markets, no evidence is found in support of cointegration, in this case for a VAR(4) system. This finding proved robust for alternative VAR length specifications. However, once again there is some evidence of cointegration when the four government bond markets are re-examined separately with returns adjusted for investors with base currencies in dollars, in euro and in sterling. Table 4 presents the results of the cointegration findings for the four government bond markets with returns denominated in dollars. The maximum eigenvalue and trace statistics are in disagreement but the former, with the sharper alternative hypothesis which is preferred in identifying the exact number of cointegrating vectors, narrowly rejects the null of zero cointegrating vectors in favour of a single vector. This was not a robust finding, however, where lower order VAR specifications indicated no cointegration. These results for dollar adjusted government bond returns are broadly consistent with those found when returns are redenominated in sterling and in euro (results not shown), ie a single cointegrating vector is found at higher but not lower order VAR lags. No evidence of cointegration was found with government bond returns expressed in yen. The findings for sterling adjusted government bond market returns conflicts with that of Andrade Clare and Thomas (1991) for the period 1978M1:1990M4 who report no cointegration.

To further explore the role of currency risk and hedging this risk, the above exercise was repeated with the four series of government bond returns in local currencies terms hedged into dollars, euro, sterling and yen using one month forward currency contracts in each case. This is to examine portfolio diversification gains from the perspective of the many investors who hedge against currency risk in returns. No evidence of cointegration was found among the four government bond markets with returns hedged into any of the four currencies. Results for returns hedged from local currencies into dollars are displayed in table 5. This contrast relative to the unhedged returns further emphasises the important role of exchange rates in identifying diversification gains.

4.3 International Credit Bond Markets

A VAR (2) Johansen system was also estimated for the four credit bond markets in local currency returns. The results presented in table 6 point to a single cointegrating vector, although this finding was less robust at higher order VAR specifications. This test was repeated separately for the four credit bond markets with returns expressed in dollars, in sterling, in euro and in yen (results not shown). In the case of all four currency adjusted returns, no robust evidence of cointegration is found indicating that the four credit bond markets provide long run portfolio diversification gains for investors with any of these four base currencies. In an identical testing procedure to that described above for government bonds, the effect of hedging local currency credit bond returns into dollars, euro, sterling and yen was also examined. No evidence of cointegration was found among hedged returns except in the case of the euro where a single cointegrating vector existed. Thus diversification gains among hedged and unhedged returns in international credit bond markets are once again somewhat dependent on the base currency into which returns are hedged.

4.4 Domestic Markets.

Cointegration between the three asset classes of stocks, government bonds and credit bonds within the UK, US, Europe and Japan is now examined. In the case of the UK, the results in table 7 provide some evidence by the maximum eigenvalue statistic of a single cointegrating vector between the three asset class returns denominated in sterling. This result was relatively consistent over alternative VAR lag specifications and also proved robust when these UK asset markets were retested with returns redenominated in both dollars and euro. This suggests that long run diversification benefits were not achievable from these UK asset markets for investors with sterling, dollar or euro as base currencies. These returns adjusted for the yen based investor showed no sign of cointegration, however.

In the case of both the US and European markets, returns in the stock, government bond and credit bond markets, measured in local currency, were found to be cointegrated for a VAR(2) and VAR(1) model respectively with the single cointegrating vector in each case described in table 8 and table 9. Higher order VAR specifications were found to be less robust with respect to a cointegration finding, however. The three asset classes in both the US and European markets showed no signs of cointegration when expressed in any currency other than local currency, however, and this highlights once more the importance of currency effects on portfolio diversification. In the case of the three asset classes in the Japanese market, no evidence of cointegration is found using the maximum eigenvalue statistic over a range of VAR specifications. Results for a VAR(1) are presented in table 10 for returns expressed in yen. The no cointegration finding in this case was unaltered by redenominating returns in dollar, sterling or euro.

Conclusion

In this paper the Johansen methodology is applied to test for cointegration between the returns in stock markets, government bond markets and credit bond markets of the US, UK, Europe and Japan and between these three asset class returns within each of the four regions. This is done with a view to identifying long-run portfolio diversification gains for international investors. Integration between international markets and between asset classes within each market is found to be sensitive to the currencies in which returns are denominated and to whether currency risk in returns is hedged. In separate tests of international stock markets and international government bond markets, no evidence of cointegration is found when returns are expressed in local currency terms but international markets are cointegrated when returns are redenominated in dollars, in euro and in sterling. Hedged returns show no sign of cointegration. While local currency returns among international credit bond markets were found to be cointegrated, no robust evidence of cointegration is found when returns are expressed in dollars, sterling, euro or yen suggesting that long run diversification benefits were achievable to these base currency investors. Within the US and European markets the three asset classes are found to be integrated with returns expressed in local currency only, but not otherwise. This again indicates that diversification opportunities are dependent on the choice of base currency. More robust evidence of cointegration is found among the three UK asset classes while the Japanese assets are not cointegrated in any of the four currency denominations.

Appendix of Tables: Johansen Estimation Results

Table 1: Stock Indices Denominated in Local Currencies, VAR (1).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.105	r = 0	23.07(27.42)	40.27(48.88)
0.054	r ≤ 1	11.69(21.12)	17.19(31.54)
0.024	r ≤ 2	5.11(14.88)	5.50(17.86)
0.0018	r ≤ 3	0.389(8.07)	0.389(8.07)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

Table 2: Stock Indices Denominated in Sterling, VAR (1).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.134	r = 0	29.79(21.12)	40.30(31.54)
0.046	r ≤ 1	9.80(14.88)	10.50(17.86)
0.0034	r ≤ 2	0.70(8.07)	0.70(8.07)

ECM	Constant	Adjustment Parameter[3]	Serial Correlation[4]
ΔUSSS	-0.049	0.027(0.035)	5.099
ΔUKSS	0.007	0.096(0.028)	6.91
ΔESS	0.004	0.143(0.030)	5.18

Cointegrating Vector[5]: USSS + 0.185 - 0.387*UKSS - 0.888*ESS
(0.274) (0.270)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

3,5 Standard Errors in parentheses.

4 LM Test of Residual Serial Correlation, $\chi^2(12)$, 5% critical value = 21.0261

Table 3: Government Bond Markets Denominated in Local Currencies, VAR (4).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.094	r = 0	20.24(27.42)	40.35(48.88)
0.068	r ≤ 1	14.45(21.12)	20.10(31.54)
0.017	r ≤ 2	3.61(14.88)	5.65(17.86)
0.009	r ≤ 3	2.03(8.07)	2.03(8.07)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

Table 4: Government Bond Markets Denominated in Dollars, VAR (4).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.129	r = 0	28.23(27.42)	47.71(48.88)
0.053	r ≤ 1	11.11(21.12)	19.48(31.54)
0.040	r ≤ 2	8.37(14.88)	8.37(17.86)
0.00024	r ≤ 3	0.0049(8.07)	0.0049(8.07)

ECM	Constant	Adjustment Parameter[3]	Serial Correlation[4]
ΔUSGD	0.005	0.014(0.012)	9.68
ΔUKGD	0.001	0.126(0.030)	9.46
ΔJGD	0.00004	0.093(0.033)	13.17
ΔEGD	-0.0019	0.132(0.025)	13.24

Cointegrating Vector[5]: USGD – 0.129 – 0.857*UKGD – 0.192*JGD +0.152*EGD
(0.142) (0.154) (0.217)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

3,5 Standard Errors in parentheses.

4 LM Test of Residual Serial Correlation, $\chi^2(12)$, 5% critical value = 21.0261

Table 5: Government Bond Markets Denominated in Dollars - Hedged, VAR (4).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.084	r = 0	17.97(27.42)	35.19(48.88)
0.054	r ≤ 1	11.36(21.12)	17.22(31.54)
0.026	r ≤ 2	5.34(14.88)	5.85(17.86)
0.0025	r ≤ 3	0.510(8.07)	0.510(8.07)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

Table 6: Credit Bond Markets Denominated in Local Currencies, VAR (2).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.138	r = 0	30.71(27.42)	53.36(48.88)
0.062	r ≤ 1	13.36(21.12)	22.64(31.54)
0.030	r ≤ 2	6.46(14.88)	9.27(17.86)
0.0135	r ≤ 3	2.81(8.07)	2.81(8.07)

ECM	Constant	Adjustment Parameter[3]	Serial Correlation[4]
ΔUSCD	0.005	-0.049(0.018)	4.58
ΔUKCS	0.006	0.060(0.023)	15.77
ΔJCY	0.003	-0.014(0.016)	17.25
ΔECE	0.003	-0.024(0.014)	10.08

Cointegrating Vector[5]: USCD –1.975 – 1.72*UKSC – 0.468*JCY + 1.573*ECE
(0.314) (0.578) (0.663)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

3,5 Standard Errors in parentheses.

4 LM Test of Residual Serial Correlation, $\chi^2(12)$, 5% critical value = 21.0261

Table 7: UK Domestic Market in Sterling, VAR (2).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.099	r = 0	21.49(21.12)	26.7(31.54)
0.023	r ≤ 1	4.89(14.88)	5.20(17.86)
0.0015	r ≤ 2	0.316(8.07)	0.316(8.07)

ECM	Constant	Adjustment Parameter[3]	Serial Correlation[4]
ΔUKSS	-0.015	-0.099(0.034)	14.85
ΔUKGS	0.062	0.022(0.014)	14.49
ΔUKCS	0.022	0.006(0.014)	12.64

Cointegrating Vector[5]: UKSS – 1.190 – 3.612*UKGS + 2.298*UKCS
(0.802) (0.772)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

3,5 Standard Errors in parentheses.

4 LM Test of Residual Serial Correlation, $\chi^2(12)$, 5% critical value = 21.0261

Table 8: US Domestic Market in Dollars, VAR (2).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.117	r = 0	25.73(21.12)	35.19(31.54)
0.025	r ≤ 1	5.24(14.88)	9.46(17.86)
0.020	r ≤ 2	4.21(8.07)	4.21(8.07)

ECM	Constant	Adjustment Parameter[3]	Serial Correlation[4]
ΔUSSD	0.0008	0.018(0.012)	8.72
ΔUSGD	-0.0009	0.010(0.0039)	6.002
ΔUSCD	-0.0009	0.005(0.003)	5.00

Cointegrating Vector[5]: USSD + 11.398 – 26.669*USGD + 23.338*USCD
(7.93) (7.45)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

3,5 Standard Errors in parentheses.

4 LM Test of Residual Serial Correlation, $\chi^2(12)$, 5% critical value = 21.0261

Table 9: European Domestic Market in Euro, VAR (1).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.101	r = 0	22.10(21.12)	27.71(31.54)
0.021	$r \leq 1$	4.44(14.88)	5.61(17.86)
0.006	$r \leq 2$	1.17(8.07)	1.17(8.07)

ECM	Constant	Adjustment Parameter[3]	Serial Correlation[4]
ΔESE	-0.001	-0.049(0.018)	11.56
ΔEGE	0.004	-0.01(0.003)	18.41
ΔECE	0.003	-0.013(0.003)	19.35

Cointegrating Vector[5]: ESE + 0.619 – 5.167*EGE + 3.941*ECE
(2.02) (2.08)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

3,5 Standard Errors in parentheses.

4 LM Test of Residual Serial Correlation, $\chi^2(12)$, 5% critical value = 21.0261

Table 10: Japanese Domestic Market in Yen , VAR (1).

Eigenvalues	Null	$(\lambda_{max})^1$	$(\lambda_{trace})^2$
0.088	r = 0	19.005(21.12)	34.11(31.54)
0.064	$r \leq 1$	13.75(14.88)	15.11(17.86)
0.006	$r \leq 2$	1.3636(8.07)	1.3636(8.07)

1. Maximum Eigenvalue Statistic (95% critical values in parentheses).
2. Trace Statistic (95% critical values in parentheses).

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