Essays on the Arbitrage Pricing Theory

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Abstract

Essays on the Arbitrage Pricing Theory

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In essay one, an APT is estimated in which the risk premia vary in proportion to the conditional volatilities of macroeconomic innovations which follow an autoregressive specification [as in Davidian and Carroll (1987)]. The conditional variances of the macroeconomic innovations exhibit strong January seasonality. For size-ranked portfolios of all the shares traded on the Toronto Stock Exchange over the period from March 1962 through March 1988, six macrofactors (namely, the lag of industrial production, the Canadian index of 10 leading indicators, the U.S. composite index of 12 leading indicators, exports, the exchange rate and the residual market index) have time-varying and priced risk premia. The small-firm effect is absent in the risk-adjusted returns, and a significant portion of the observed January seasonality is explained by the model with time-varying risk premia.

Essay two modifies the arbitrage pricing theory (APT) and its international version (IAPT) to encompass the hypotheses that the Canadian and global North American equity markets are completely or partly integrated (segmented). The exchange rate determination literature is used to identify observable and potentially priced international (binational) macroeconomic factors. The findings indicate that the Canadian equity market is only partly integrated (segmented) with the American equity market [as Jorion and Schwartz (1986) find

using the CAPM and ICAPM]. Canadian stock returns are influenced by the purely domestic components of the term structure and lagged industrial production, and by the purely international components of the differential in the Canada-U.S. leading indicators and the interest rate of eurodeposits.

In essay three, the APT model with constant and time-varying and pre-specified macroeconomic factors is used to derive predicted returns. These predicted returns are then compared with actual future returns, in order to determine how much of the observed variability in stock returns is explained by the variability in the macroeconomic factors. For the period from April 1967 to December 1987, when the constant risk premia (time-varying risk premia) model is used to derive predicted returns on the equally-weighted portfolio, about 80% (83.5%) of the observed return variability of the equally-weighted portfolio is explained.

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This thesis is dedicated to my family.

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CHAPTER ONE: INTRODUCTION

The task of explaining the structure of stock prices has occupied a central and important position in the research agenda of the financial economics literature for at least the last three decades. The Capital Asset Pricing Model (CAPM), developed simultaneously by Sharpe (1964), Lintner (1965), and Mossin (1966), is the first general equilibrium model of capital asset pricing to rest on the main postulate of modern portfolio theory that the only risk priced in capital markets is that portion of total risk that cannot be diversified away. This risk is called systematic risk.

The main prediction of the CAPM, that all economic agents hold the market portfolio, along with Roll's (1977) critique regarding the unobservability of the market portfolio, has given rise to the Arbitrage Pricing Theory (APT) developed by Ross (1976). The APT predicts that the expected returns on assets are linearly related to one or several systematic factors that may be priced risks. The original development of the APT by Ross (1976) relies on a no-arbitrage argument and holds only as an approximation. Connor (1984), using a general equilibrium framework, shows that the APT holds exactly, if a well-diversified market portfolio exists.

Two different statistical methods have been used to test the APT. The first method is factor analysis and does not require the identification of the systematic factors influencing returns. Several problems have been documented concerning the application of factor analysis in the asset pricing literature. For example, Brown (1989) demonstrates that factor analysis understates the number of significantly priced systematic factors beyond the first and, as shown by Anderson (1984) and Conway and Reinganum (1988), factor loadings are not unique and their statistical properties are unknown when the errors fail to follow a joint normal distribution. Chen et al. (1986) developed a second method to test the APT, by assuming that the unobserved systematic factors influencing returns can be measured by observable macroeconomic factors. As shown by McElroy and Burmeister (1987), this methodology along with Gallant's (1985) method of non-linear seemingly unrelated regressions (NSUR), avoids the problems associated with factor analysis.

Another body of literature documents a number of seemingly unexplainable patterns in asset returns. These are referred to as market anomalies. Banz (1981) and Reinganum (1983) find that small-capitalization firms listed on the NYSE and ASE have risk-adjusted returns that significantly exceed those of large market value firms. This pattern of stock returns is the so-called small-firm effect. In addition, Keim (1983) finds that more than fifty percent of the excess returns for small firms is concentrated in the first week of January. This pattern of stock returns is the so-called January or turn-of-the-year effect. Finally, evidence provided by Shiller (1981a, 1981b) and LeRoy and Porter (1981) (and subsequently many other researchers) shows that prices are generally too volatile to be determined in an efficient capital market. Numerous studies, using stock data for the U.S., Canada, the U.K, Japan, and other countries, have failed adequately to explain these three effects. Moreover, the persistency of these effects across different countries suggests some or even complete integration of capital markets. The general failure to explain anomalies and the excessive variability of stock returns implies either that investors in these capital markets are using information about economic fundamentals inefficiently or that the asset pricing models used to measure risk are misspecified or both. Before the verdict that these capital markets are inefficient is reached, a more careful examination of the specification of the asset pricing models is required.

In the light of the above discussion, there are two principal aims to this thesis: (i) to use the APT to explain the market anomalies and the apparent excess variability of stock returns; and (ii) to modify the APT to test if the Canadian and global North American equity markets are integrated or segmented.

The integrating theme of chapters two, three and four is the APT. Chapters two and four use the APT with time-varying risk premia and both domestic (Canadian) and international (North American) macroeconomic factors to examine (i) above. In chapter three the APT, with only domestic macrofactors, and the international APT (IAPT) with only international factors and constant risk premia, are used to examine (ii) above. In essence, in chapters two and four it is assumed that the Canadian and American equity markets are partly integrated (based on the findings presented in chapter three) since the Canadian/U.S. exchange rate is included directly in the factor structure as a proxy for common international (binational) factors [see Solnik (1983)]. In chapter three the exchange rate variable is replaced with international macrofactors [see Frankel (1979)] in order to test the hypotheses of complete segmentation or complete integration of the Canadian/U.S.

equity markets. This methodology might seem to create a contradiction, because the significant factor structure in chapter two and four is different from that in chapter three. In fact such a statement is seen to be false because the different factor structure arise from wholly different hypotheses. As a first step the constant (unconditional) risk premia model is used to test for integration or segmentation of the Canadian and U.S. equity markets. A natural extension of chapter three is to allow for time-varying (conditional) risk premia in further research.

A word of caution is in order at this point since all three essays are affected by what is known, in the empirical asset pricing literature, as the errors-in-variables problem (EIV). The EIV problem arises from the need to estimate variables in the first step that are then used to estimate the asset pricing models in the second step. This can bias the individual coefficients towards significance. As shown in Shanken (1992), this problem is alleviated by using maximum-likelihood procedures and joint, as opposed to individual, coefficient tests of significance. Maximum-likelihood estimation methods and joint tests are reported when the important hypotheses to this theses are tested. It follows that individual t-statistics have to be interpreted with some caution.

The thesis is organized as follows:

In chapter two, the APT model is first extended to account for the possible influence of the conditional volatilities of macroeconomic variables on the time-variation of risk premia and, consequently, on expected stock returns; and second, this extended APT model is used to investigate whether the January seasonality and small firm effects can be explained by the changing conditional volatilities of the macroeconomic variables.

In chapter three, the IAPT and the domestic APT are first generalized into two separate comprehensive forms in which the competing hypotheses of complete and partial market integration and segmentation are nested. Next, the priced North American macroeconomic factors relevant to these models are identified. Finally, the hypothesis that the Canadian equity market is fully integrated is tested against the alternative that it is segmented. Unlike previous studies, the binational factors from the Frankel (1979) exchange rate determination model are used to identify global North American macroeconomic factors.

The APT with identifiable macroeconomic factors is used to construct predicted stock returns in chapter four. The variability of this measure is then compared with the variability of actual returns to determine how much of observed stock return variability is explained (and justified) by time-variation in risk premia and the variation in economic fundamentals.

The major findings of this thesis and their implications are summarized in chapter five. This is followed by a discussion regarding possible directions for future research.

CHAPTER TWO: MACROFACTOR CONDITIONAL VOLATILITIES, TIME-VARYING RISK PREMIA AND STOCK RETURN BEHAVIOR

2.1 INTRODUCTION

The study of the behavior of asset prices has been at the forefront of research in financial economics for at least the last three decades. Linking movements in asset returns to the macroeconomy has been and still is a fascinating and controversial subject. Although it is commonly believed that unanticipated changes in macroeconomic variables do affect stock returns, less agreement exists on which macrovariables significantly affect returns and how these variables influence the pricing of stocks. Since there is no widely accepted economic theory that links the stock market to various economic variables, economic intuition is used to determine the appropriate macrovariables. Studies by Chen, Roll and Ross (1986), Brown and Otsuki (1989) and Kryzanowski and Zhang (1992) use the Arbitrage Pricing Theory (APT) and economic intuition to relate the American, Japanese and Canadian stock market returns, respectively, to unanticipated changes in various macroeconomic variables. Since these studies find that macroeconomic variables influence stock returns significantly (as predicted by the APT), they lend support to the commonly held notion that economic fundamentals determine asset returns.

Another body of literature documents a number of seemingly unexplainable patterns in asset returns. The lack of any generally acceptable explanation and the persistence of these patterns are the main reasons why they are referred to as market anomalies. Two puzzling anomalies are the small-firm effect and calendar seasonality (especially, the abnormal returns exhibited by stocks during the month of January). Banz (1981) and Reinganum (1983) find that small capitalization firms listed on the NYSE and ASE have risk-adjusted returns that significantly exceed those of large market value firms. Keim (1983) finds that more than fifty percent of the excess returns for small firms is concentrated in the first week of January. This pattern of stock returns is the so-called January or turn-of-the-year effect. Numerous studies, using stock data for the U.S., Canada, U.K, Japan, and other countries, have failed to adequately explain these effects. All of these studies have used either the Capital Asset Pricing Model (CAPM) or the Arbitrage Pricing Theory (APT) to risk-adjust returns so that meaningful comparisons between the returns of small and large firms can be made. A similar procedure is followed for comparisons between returns obtained during the month of January and the other months.[1]

The general failure to explain these anomalies implies either that investors in these capital markets are utilizing economic fundamentals inefficiently or that the asset pricing models used to measure risk are misspecified (i.e., economic fundamentals are not incorporated correctly). Before the verdict that these capital markets are inefficient is reached, a more careful examination of the specification of the asset pricing models used is required.

The restrictive assumption that asset returns are multivariate normal or preferences are quadratic, together with Roll's (1977) critique about the use of the market index as a valid proxy for the market portfolio when estimating the CAPM, seriously challenges all the studies that use the CAPM to adjust for risk. The APT is a more general model that allows for more than one factor to influence returns, and does not require any restrictive assumptions on preferences. However, many studies that use the APT to test for anomalies do not specify the factors affecting returns. Instead, they use factor analysis, which has all the flaws revealed by for example Brown (1990). Using the APT with identified macroeconomic factors for 70 individual U.S. securities, Burmeister and McElroy (1988) find that the January seasonal is strongly significant and that the residual market factor (RMF) is a priced risk. Similarly, in a study of the Japanese equity market, Brown and Otsuki (1989) find that the RMF has a significant risk premium. In contrast, for portfolios of U.S. securities, Chen, Roll and Ross (1986) find that the RMF has an insignificant risk premium. Using the APT with specified macrofactors for portfolios of Canadian securities, Kryzanowski and Zhang (1992) cannot explain the January effect [as was the case for Burmeister and McElroy (1988)].[2] However, like Chen, Roll and Ross, Kryzanowski and Zhang find an insignificantly priced risk premium for the RMF. Kryzanowski and Zhang attribute this result to their use of returns on size-ranked portfolios, unlike Burmeister and McElroy who use returns on randomly selected stocks. This explanation is not totally satisfactory because Brown and Otsuki obtain similar results for three types of portfolios (size-ranked, industry-ranked and line-of-business-ranked), and Kryzanowski and Zhang use a full information statistical method to estimate the prices of risk unlike Chen, Roll and Ross. Moreover, no theoretical reason exists to justify different risk premia for individual stocks and portfolios.

Another possible explanation may exist in the recent theoretical and empirical works by Merton (1973, 1980), Cox, Ingersoll and Ross (1985), Breeden (1986), Campbell (1987), French, Schwert and Stambaugh (1987), Lauterbach (1989) and Ferson and Harvey (1991) that relate the risk premium associated with the pervasive influence of a factor on stock returns and the conditional volatility of that factor. This relationship causes the risk premia to vary over time with the time-varying volatilities of the factors. Therefore, the previous studies may obtain different results because they do not account explicitly for the possible time variation in the risk premia due to the changing volatilities of the factors. Moreover, the time variation of the risk premia may explain the observed January seasonality and the small-firm effect.

The purpose of this chapter is two-fold: first, to extend the APT model to account for the possible influence of the conditional volatilities of the macroeconomic variables on the time-variation of risk premia and, consequently, on expected stock returns; and second, to investigate whether the January seasonality and small firm effects can be explained by the changing conditional volatilities of the macroeconomic variables.[3]

The chapter is organized as follows. In the next section, the extended APT model is presented. In section three, the data used to estimate the model are described. In section four, tests of the extended APT model and two anomalies are conducted and analyzed. In the fifth and last section, some concluding remarks are offered.

2.2 THE APT WITH CONDITIONAL MACROFACTOR VOLATILITIES

The basic APT hypothesis states that the difference between the actual and expected returns on the i'th asset at time t is determined by the following linear factor model (LFM) with K factors:

$$r_i(t)=E_t[r_i(t)]+\sum_{j=1}^{K} b_{ij}F_j(t)+\varepsilon_i(t) \tag{2.2.1}$$

where E_t is the rational expectation of the return on asset i conditioned on information at the beginning of period t; $r_i(t)$ is the total return on asset i in period t; $F_j(t)$ is the unanticipated realization of factor j in period t; b_{ij} is the sensitivity of asset i to factor j; and $\varepsilon_i(t)$ is the idiosyncratic disturbance of firm i in period t. $F_j(t)$ satisfies:

$$E_t[F_j(t)] = 0, \text{ and}$$

$$\begin{aligned} E_t[F_i(t)F_j(s)] &= v_{ij}, & t &= s \\ &= 0. & t &\neq s \end{aligned}$$

$\varepsilon_i(t)$ satisfies:

$$\begin{aligned} E_t[\varepsilon_i(t)] = 0, \qquad E_t[\varepsilon_i(t)\varepsilon_j(s)] &= \sigma_{ij} & t &= s \\ &= 0, & t &\neq s \end{aligned}$$

and $\quad E_t[\varepsilon_i(t)F_j(s)] = 0 \qquad$ for all i, j, t and s.

By assuming that the LFM holds true, that no arbitrage profits exist and that certain other regularity conditions are satisfied, Ross (1977) is able to derive the following approximate pricing relationship for the n assets:

$$E_t[r_i(t)] \approx \lambda_o(t) + \sum_{j=1}^{K} b_{ij}\lambda_j(t) \tag{2.2.2}$$

where $\lambda_o(t)$ is the return on a risk-free asset for period t, which is known at the beginning of the period. Given that b_{ij} can be interpreted as the risk associated with asset i due to the systematic influence from factor j, $\lambda_j(t)$ can be interpreted as the risk premium associated with the pervasive influence of factor j on all the assets. Therefore, $\sum_{j=1}^{K} b_{ij}\lambda_j(t)$ can be interpreted as the risk premium expected from holding asset i in an environment where the K factors are not diversifiable risks. The pricing relationship given by (2.2.2) is what Shanken (1985) has labelled the arbitrage APT. It holds only as an approximation. Dybvig (1983) estimates a conservative monthly error bound of .015%.

McElroy and Burmeister (1988) extend the empirical APT by assuming that an unobserved residual market factor (RMF) is included in the LFM. Specifically, they assume that the J≡K-1 factors are observed but that the Kth factor is unobserved. Therefore, the Kth factor enters the LFM through the error term:

$$\varepsilon_i(t) \equiv b_{iK}F_K(t) + u_i(t) \tag{2.2.3}$$

The RMF captures the pervasive macrofactors omitted from the specification of $F_j(t)$, J = K-1. A market index $r_m(t)$ must also be represented by the same linear factor model:

$$r_m(t) = E_t[r_m(t)] + \sum_{j=1}^{J} b_{mj}F_j(t) + b_{mK}F_K(t) + \varepsilon_m(t) \qquad (2.2.4)$$

where $\varepsilon_m(t) = \sum_{i=1}^{N} w_i\varepsilon_i(t)$, and w_i is the portfolio weights summing to one. Therefore, the variance of $\varepsilon_m(t) = 0$, and $\varepsilon_m(t)$ (which is the unsystematic risk of the market portfolio) approaches zero for well diversified portfolios such as the market portfolio. With no economic consequences, normalizing $b_{mK} = 1$ yields:

$$r_m(t) = E_t[r_m(t)] + \sum_{j=1}^{J} b_{mj}F_j(t) + F_K(t) \qquad (2.2.5)$$

In practice, since $r_m(t)$ can be approximated by a broad market index, the residual market factor $F_K(t)$ can be estimated.[4] As shown in Connor (1984) and Wei (1988), the RMF can enter the APT pricing equation given by (2.2.2) in order for market clearing to obtain in the securities market. This version of the APT is called the equilibrium APT due to the assumption of market clearing.

The preceding extensions allow us to rewrite the APT pricing equation (2.2.2) as:

$$E_t[r_i(t)] = \lambda_0(t) + \sum_{j=1}^{J} b_{ij}\lambda_j(t) + b_{iK}\lambda_K(t) \qquad (2.2.6)$$

The APT pricing counterpart of (2.2.5) can be written as:

$$E_t[r_m(t)] = \lambda_o(t) + \sum_{j=1}^{J} b_{mj}\lambda_j(t) + \lambda_K(t) \tag{2.2.7}$$

Clearly, the system of equations (2.2.1), (2.2.5), (2.2.6) and (2.2.7) can be estimated using a nonlinear system estimation techniques. McElroy and Burmeister (1988) first estimate (2.2.5) to get the fitted values $\hat{F}_K(t)$, and then estimate a variant of the following system:

$$r_i(t) - \lambda_o(t) = \sum_{j=1}^{J} b_{ij}\lambda_j(t) + \sum_{j=1}^{J} b_{ij}F_j(t)$$
$$+ b_{iK}\lambda_K(t) + b_{iK}\hat{F}_K(t) + u_i(t) \tag{2.2.8}$$

McElroy and Burmeister estimate (2.2.8) with no time subscripts for λ_j because the risk premia are assumed to be constant over time. Although in this paper equation (2.2.8) is interpreted as an APT pricing equation, Chamberlain (1988) has shown that if the market portfolio is well-diversified then equation (2.2.8) can also be interpreted as Merton's (1973) Intertemporal CAPM (ICAPM). In Chamberlain's framework the two pricing models are not testably distinct.

Theoretical work by Merton (1973), Cox, Ingersoll and Ross (1985) and Breeden (1986) relate the risk premium of factor j to its volatility and a constant proportionality factor. Empirical studies by Merton (1980), French, Schwert and Stambaugh (1987) and Campbell (1987) support these theoretical results because they find that the expected risk premium on the stock market is positively correlated with the predictable volatility of stock returns. Lauterbach (1989) documents a relation between the expected returns on U.S. Treasury bills and the conditional volatilities of consumption, the spot interest rate, and industrial production. Therefore, to complete the asset pricing model the following relationship is postulated:

$$\lambda_j(t) = a_j + R_j \sigma_j^2(t) \tag{2.2.9}$$

where R_j is a proportionality coefficient; $\sigma_j^2(t)$ is the conditional volatility of factor j; and a_j is a parameter. If the conditional volatilities of the macrofactors change over time, then the risk premia will be time-varying. Substituting equation (2.2.9) into the system (2.2.8) yields the following complete intertemporal asset pricing model:

$$r_i(t) - \lambda_o(t) = \sum_{j=1}^{K} b_{ij} a_j + \sum_{j=1}^{J} R_j \, b_{ij} \sigma_j^2(t) + \sum_{j=1}^{J} b_{ij} F_j(t)$$

$$+ b_{iK} R_K \sigma_K^2(t) + b_{iK} \hat{F}_K(t) + u_i(t) \tag{2.2.10}$$

Since the first term in equation (2.2.10) is a constant, equation (2.2.10) can be rewritten as:

$$r_i(t) - \lambda_o(t) = c_i + \sum_{j=1}^{J} R_j \, b_{ij} \sigma_j^2(t) + b_{iK} R_K \sigma_K^2(t) + \sum_{j=1}^{J} b_{ij} F_j(t)$$

$$+ b_{iK} \hat{F}_K(t) + u_i(t) \tag{2.2.11}$$

where $u_i(t)$ is normally and independently distributed.

This expanded asset pricing system can be used to test whether the predictable volatility of the macrofactors are significant sources of risk. This model allows for a test of the null hypothesis of the existence of one or more significant risk premia and for a test of the hypothesis that the risk premia are jointly time-varying. To test for a January effect, the additional term $\Phi_i D_t$ is included in equation (2.2.11), where $D_t = 1$ in January and zero otherwise, and Φ_i represents asset i's sensitivity to the January effect. A significant Φ_i implies that the January seasonality cannot be explained by the extended APT model.

If $\lambda_j(t) = a_j + R\sigma_j^2(t)$ (i.e., the proportionality factor, R, is the same for each macrofactor), then the following restricted variant of equation (2.2.11) obtains:

$$r_i(t)-\lambda_o(t) = c_i + R \sum_{j=1}^{J} b_{ij}\, \sigma_j^2(t) + R\, b_{iK}\sigma_K^2(t) + \sum_{j=1}^{J} b_{ij}F_j(t) + b_{iK}\hat{F}_K(t) + u_i(t) \qquad (2.2.12)$$

This restrictive formulation can be used to test whether the proportionality factor is the same for the various macrofactors. Under this formulation, R, can be interpreted as a measure of relative risk aversion.[5] Since the innovations of the macroeconomic factors can only be observed monthly, their conditional variances can be estimated using the method developed by Davidian and Carroll (1987) who argue that the variances of the factors estimated in this fashion yield more robust estimates than other variance specifications. Specifically the following autoregression is estimated for each factor j:

$$\sigma_j^2(t) = \sum_{j=1}^{12} \alpha_j D_j(t) + \sum_{i=1}^{12} \theta_i \, \sigma_j^2(t-i) + V_j(t) \qquad (2.2.13)$$

where $\sigma_j^2(t) = (\pi/2)\ F_j(t)^2$ is the unconditional variance of factor j at time t, since $E[F_j(t)] = 0$ [$F_j(t)$ are innovations]. Since $F_j(t)^2$ are single point variance estimates they have to be adjusted by the term $(\pi/2)$ [see Schwert (1989) and Schwert and Seguin (1990)]. The fitted variances $\hat{\sigma}_j^2(t)$ are estimates of the conditional variances of factor j. This is similar to the ARCH specification for modelling time-varying volatilities, since the monthly dummy variable $D_j(t)$ allows for different monthly standard deviations. Schwert (1989) and Schwert and Seguin (1990) use this specification to estimate the standard deviations of monthly returns conditional on information up to the present time.

A separate test to determine whether or not the conditional volatilities have monthly seasonality can also be performed. To do so, $\hat{F}_j(t)^2$ is regressed on 11 monthly dummies as follows:

$$\hat{F}_j(t)^2 = \delta_j + \sum_{i=2}^{12} \delta_i D_i(t) + w_j(t) \qquad (2.2.14)$$

where δ_j estimates the average volatility for January of factor j, and δ_i measures the difference in the estimates of the conditional variances between January and month i.[6] An F-statistic under the null hypothesis of $\delta_2=\delta_3=\ldots=\delta_{12}= 0$ indicates how significantly the conditional volatilities from February to December jointly differ from that in

January.

2.3 DATA

This study uses all the stocks traded on the Toronto Stock Exchange (TSE) from January 1961 through March 1988. The return data are obtained from the TSE/Western Monthly Data Base. The decision to use size-ranked portfolios is based on two considerations. First, since a test of the small-firm effect is conducted in this study, such size-ranked portfolio returns are required. Second, most studies which test asset pricing models use these portfolios to obtain a maximum dispersion in returns since small capitalization (cap) firms generally have higher returns (and risks) than large cap firms.

The construction of the size-ranked portfolios requires that stocks first be ranked monthly according to their annual December-end outstanding market values, and then cutoff points equal to the desired number of portfolios be determined. The returns on stocks contained within each of the cutoff points are used to calculate the average equally-weighted portfolio returns. In this study, twenty and fifty size-ranked portfolios are used.

All of the macroeconomic variables whose innovations could enter the LFM are taken from Statistics Canada's CANSIM Mini Base. (See Table 2.1 for a list of these variables.) All the series have been converted to real values, and are seasonally adjusted. The first differences in the logarithms (growth rates) of most of the macroeconomic variables are used herein.

The choice of the macroeconomic variables was dictated by several factors. Since no generally accepted theory exists for linking stock returns to the economy, none could be used to derive unique and universally accepted macrovariables. As a result, general economic theory and intuition were the main inputs used in the selection process. Macroeconomic factors that have been found to influence stock returns in past studies and data availability were also important inputs affecting the selection decision. Whether or not the macroeconomic variables appeared in the popular financial media was also an important consideration in the final stage of the selection process.

Variables designed to capture the real sector of the economy include the Canadian composite index of leading indicators and industrial production. The variable designed to capture the monetary and financial sector is the money supply. Since Canada's economy is highly related to the performance of the U.S. economy, variables designed to capture the influence of the foreign sector on stock returns include the Canada/U.S. exchange rate, total exports, and the U.S. composite index of leading indicators.[7]

As defined in Chen, Roll and Ross (1985), Chan, Chen and Hsieh (1985) and McElroy and Burmeister (1988), the risk premium is given by:

$$PREM(t) = CBOND(t) - LBOND(t)$$

where CBOND is the average yield for ten industrial bonds that constitute the McLeod, Young and Weir bond index, and LBOND is the yield on Government of Canada long-term bonds with maturities of ten years and over.[8] The shape of the term structure is defined as:

$$TERM(t) = LBOND(t) - TBILL(t-1)$$

where TBILL is the average monthly yield on Government of Canada 91-day

Treasury Bills. Although these two variables are not mean zero (since the term structure is usually upward sloping and the risk premium for holding the more risky corporate bonds instead of the less risky government bonds is always positive), most of the previous U.S. studies have used these variables directly as innovations. In this study, the innovations of these variables will be used to ensure that they are both zero-mean and serially uncorrelated, as innovations should be.

The innovations of all the macroeconomic variables are equal to the forecast errors that are obtained by fitting each of the macroeconomic series using Akaike's (1976) state-space procedure. Brown and Otsuki (1989) and Kryzanowski and Zhang (1992) use this procedure. The t-statistics for the null hypothesis of zero mean, the Kolmogorov D-statistics for normality, and the Fisher Kappa and the Ljung-Box Q-statistics for the null hypothesis of white noise are reported in Table 2.2.

All the variables have t-statistics that do not allow for rejection of the null hypothesis of zero mean. The D-statistics indicate that the null hypothesis, that the innovations are normally distributed, cannot be rejected for all but the TERM and PREM macroeconomic variables. Fisher's Kappa test statistics for the innovations of industrial production, term structure and the risk premium indicate that these variables are not white noise. The Ljung and Box Q-statistics indicate that none of the innovations are autocorrelated at the one percent level.

To ensure that each macroeconomic innovation does not contain information embodied in any of the other innovations a series of sequential regressions is performed. Specifically, the second innovation is regressed against the first innovation and the residual from this regression is used as the second innovation. Then, the third innovation is the residual of a regression of itself on the first and second innovation. The same procedure is followed for the rest of the macroeconomic innovations to get orthogonalized innovations by construction.

2.4 EMPIRICAL RESULTS

2.4.1 Linear factor model estimation

The linear factor models (LFM) for the three market indices on the macroeconomic innovations are reported in Table 2.3. The returns on the TSE300, equally-weighted, and value-weighted market indices are significantly affected by the innovations of the U.S. composite index (USINDEX), the exchange rate (EX), the Canadian leading index (CINDEX), the lag of the industrial production index (LINDUS), Canadian exports (EXPORTS) and the term structure (TERM). The t-values indicate that these six variables are significant at the .01 significance level. Also the F-values indicate that the overall regressions are significant at the .01 significance level. The D-W statistics are close to two for all three LFMs, indicating no apparent misspecification due to serial correlation. Also, these results are generally consistent with the findings of Chen, Roll and Ross (1986), McElroy and Burmeister (1988) and Kryzanowski and Zhang (1992).

2.4.2 Seasonality in the conditional volatilities

The macroeconomic innovations that enter the LFMs significantly are used to test whether or not the conditional variance of each macrovariable exhibits January seasonality. First, the autoregressions given by equation (2.2.13) are estimated for the innovation series for each of the macroeconomic variables. The fitted values (the conditional variances) are regressed against 11 monthly dummies and a constant [see equation (2.2.14)] to test for monthly seasonality. The results from these regressions are reported in Table 2.4. The F-statistics, which test the null hypothesis of no difference between the conditional variances in the month of January and the other months jointly, reject the null at the .01 significance level for all of the macrovariables with exception of the term structure variable. It seems that economic agents expect the variances of the macroeconomic innovations to be consistently different in January from the other months. If agents react according to their expectations, then the finding of a January effect in all of the past studies should not be a surprise. However, the issue of why economic agents expect the conditional volatilities of the innovations to be significantly different in January than those of the other months jointly has to be resolved.

2.4.3 Asset pricing model estimations

Nonlinear ordinary least squares (NOLS) and nonlinear seemingly unrelated regression (NSUR) parameter estimates of the extended (unrestricted version) APT model given by the system of equations in (2.2.11) are reported in Table 2.5. Both methods accommodate the cross-equation restrictions implied by the theory behind the model. The restricted NOLS presumes cross-equation covariances are zero and the restricted NSUR presumes these cross-equation covariances are non-zero. Thus NSUR is a more general estimation method and, moreover if the equations exclude common variables that should be included, then the non-zero error covariances go some way to allowing for this misspecification. NOLS does not allow for these possibly excluded variables in this way. The NSUR standard errors account for the cross-equation covariances and are correctly calculated from this point of view. The NOLS standard errors do not account for the cross-equation covariances and hence are not properly calculated. Therefore, NSUR is more reliable. The NOLS estimates are reported for completeness because these have to be obtained as part of the NSUR procedure. The estimates of the mean factor sensitivities, b_{ij}s, and the mean of the constant portion of the factor risk premia, c_i, are reported in Panels A and B when the fifty and twenty portfolios, respectively, are used for estimating the system of equations in (2.2.11). Estimates of the risk premium proportionality coefficients, R_j [j = LINDUS, CINDEX, TERM, USINDEX, EXPORTS, EX and the residual market factor (RMF)] for the fifty and twenty portfolio models are reported in Panels C and D, respectively.

In general, the estimates of the factor sensitivities for the twenty and fifty portfolios do not differ significantly. However, the variances of these parameter estimates tend to increase as the number of portfolios increases. This is evident from the drop in the value of the t-statistics as more portfolios are used to estimate the extended APT model (see Panels A and B in Table 2.5). This result may be explained by the fact that the number of stocks in each portfolio diminishes as the number of portfolios increases from twenty to fifty. Each of the fifty portfolios are on average less diversified than each of the twenty portfolios. Decreased portfolio diversification adds more non-systematic risk to the APT system which increases the variances of the regressions and the factor betas. The NOLS and NSUR mean factor loading estimates are very similar with the exception of the mean beta for the exchange rate (b_{EX}). It is significant when twenty portfolios are used to estimate the system of equations in (2.2.11), and is insignificant when the fifty portfolios are used to estimate the same system.

The NOLS and NSUR estimates for the risk premia proportionality coefficients when the twenty and fifty portfolios are used in equation (2.2.11) differ significantly. This indicates the importance of accounting for the full residual covariance matrix (see Panels C and D in Table 2.5). For the fifty portfolio model, the term structure proportionality coefficient (R_{TERM}) is significant with NOLS and insignificant with NSUR. The opposite occurs for the exports proportionality coefficient ($R_{EXPORTS}$). When the twenty portfolios are used to estimate (2.2.11), the R_{TERM} is significant with the NOLS and insignificant with the NSUR estimation methods, respectively. The $R_{EXPORTS}$ coefficient is insignificant for both the NOLS and NSUR methods. The proportionality coefficient of the residual market factor (RMF) is significant in the fifty portfolio model for both the NOLS and NSUR estimation methods and in the twenty portfolio model for the NOLS method. To summarize, the following macrofactors appear to have time-varying risk premia: the lag of industrial production (LINDUS), the

Canadian leading indicators index (CINDEX), the U.S. index of leading indicators (USINDEX), EXPORTS, the exchange rate (EX) and the RMF. Since the RMF has a time-varying risk premium, the contradiction between the results of Burmeister and McElroy (1988), Chen et al. (1986) and Kryzanowski and Zhang (1992) is likely to be resolved.

The mean coefficient estimates and the estimate of the single proportionality coefficient for the restricted model given by the system of equations in (2.2.12) are reported in Table 2.6. The mean beta estimates for the fifty portfolio version of equations (2.2.12) are reported in Panel A and do not differ significantly from the beta estimates obtained for the twenty portfolio model reported in Panel B. However, as with the system of equations (2.2.11), the estimated variances of these betas increase as the number of portfolios increases. The estimate of the proportionality coefficient (R) is significant for the fifty portfolio model when the NOLS and NSUR estimation methods are used. For the twenty portfolio model, R is significant only with the NOLS method.

The appropriate differences of the relevant quadratic forms obtained when different interesting restrictions are imposed on equations (2.2.11) and (2.2.12) for the twenty and fifty portfolio models are reported in Table 2.7. These differences are asymptotically chi-square distributed with degrees of freedom equal to the number of retrictions. The null hypothesis that the risk premia are insignificant [i.e., $c_i=R_j=0$ $\forall$ $i=1,\dots,n$ and $j=1,..J$, and K in system (2.2.11)] is rejected at the .01 significance level for the both the twenty and fifty portfolio models. The same null hypothesis for the restricted system

(2.2.12) [i.e., $c_i=R=0$ $\forall$ $i=1,\dots,n$ in system (2.2.12)] is rejected at the .01 significance level for the fifty portfolio model and at the .06 significance level for the twenty portfolio model. The null hypothesis that the risk premia are jointly time-invariant [i.e., $R_j=0$ $\forall$ $j=1,\dots,J$, and K in system (2.2.11)] is rejected at the .01 significance level for both the twenty and fifty portfolio models. The restriction on system (2.2.12) that R=0 cannot be rejected at the .05 significance level for both the twenty and fifty portfolio models. Therefore, the unrestricted APT model described by system (2.2.11) is superior to the restricted version described by (2.2.12), and the risk premia for at least six variables (namely, LINDUS, CINDEX, USINDEX, EXPORTS, EX and RMF) are time-varying.

2.4.4 Tests of the January and Firm-size Effects

The NOLS and NLSUR parameter estimates of the extended APT model given by equation (2.2.11), which includes different proportionality coefficients and a January dummy variable, are reported in Table 2.8.[9] The mean coefficient of the January dummy is insignificant at traditional significance levels for both the fifty and twenty portfolio models and both estimation methods. More specifically, in the fifty and twenty portfolio models, sixteen (thirty-three in the LFM) and seven (seventeen in the LFM) portfolios exhibit a significant January seasonality at the .10 significance level. This indicates that the unrestricted version of the extended APT with time-varying risk premia is a tremendous improvement over the LFM (i.e., an improvement of 51.5% and 58.8% for the fifty and twenty portfolio models, respectively) in explaining observed stock return behavior. The inclusion of the dummy variable does not affect the estimated beta coefficients, which confirms the findings of Kryzanowski and Zhang (1992) and Burmeister and McElroy (1988). Moreover, the estimates of the proportionality coefficients are also very robust when the January seasonal is included.

The estimates of the extended APT model with the same proportionality coefficient for all of the risk premia [i.e. equation (2.2.12)] and a January dummy variable are reported in Table 2.9. The mean coefficient and statistical significance of the January dummy for both the fifty and twenty portfolio models are virtually identical to those of the LFM. This indicates the superiority of the unrestricted version of the extended APT model over the restricted version in explaining observed stock return behavior.

Joint tests of the January seasonal are reported in Table 2.10. The unadjusted and Bartlett's small-sample adjusted χ^2-statistics are reported for the system of equations (2.2.11) and (2.2.12) for both the twenty and fifty portfolio models. The null hypothesis that there is no January seasonality in stock returns cannot be rejected for the unconstrained APT model [i.e. equation (2.2.11)] with twenty portfolios at the .05 significance level using either the unadjusted or small-sample adjusted χ^2-statistics. When fifty portfolios are used to estimate system (2.2.11) the null hypothesis of no January seasonality is rejected at the .01 significance level if the unadjusted χ^2-statistic is used. If the small-sample adjusted χ^2-statistic is used as in Burmeister and McElroy (1988), then the null hypothesis of no January seasonality cannot be rejected at the .01 significance level. This result differs from the Burmeister and McElroy result where the same null hypothesis is rejected overwhelmingly.

The null hypothesis of no January seasonality is rejected clearly for the restricted version of the extended APT given by system (2.2.12) for both the fifty and twenty portfolio models using either the unadjusted or small-sample adjusted χ^2-statistics.

The difference in raw returns between the firms with market values in the top and the bottom 20% of those listed on the TSE from 1963 to 1988 is equal to 4.5%. The associated t-statistic of 1.73 indicates that this difference is significantly different than zero at the 0.10 level. The corresponding risk-adjusted return of -.00006%, which is computed from the residuals of equation (2.2.11) for the fifty portfolio model, has an associated t-statistic of -.0033.[10] This result indicates that the difference in the risk-adjusted returns of small and large capitalization firms is insignificantly different from zero at the 0.01 level of significance.

2.5 CONCLUDING REMARKS

An APT with time-varying conditional volatilities of macroeconomic innovations (or equivalently time-varying risk premia) was first estimated herein. This was followed by tests for the presence of a January seasonal and a small-firm effect in the risk-adjusted returns of the extended APT model. The empirical results show that the conditional variances of six macrofactors have time-varying risk premia. These are the lag of industrial production, the Canadian index of 10 leading indicators, the U.S. composite index of 12 leading indicators, exports, the exchange rate and the residual market factor. The asset pricing model with time-varying risk premia is able to explain from 51.5 percent (for the fifty portfolio model) to 58.8 percent (for the twenty portfolio model) of the observed January seasonality. No small-firm effect is detected in the risk-adjusted returns.

CHAPTER THREE: INTEGRATION OR SEGMENTATION OF THE CANADIAN STOCK MARKET: EVIDENCE BASED ON THE APT

3.1 INTRODUCTION

National economies are becoming more internationalized through increased trade and the mutual cooperation of national governments to lower (and eventually eliminate) all impediments to the free flow of goods, services and financial, physical and human capital. For example, the twelve members of the European community are currently attempting to integrate their economies completely into one large market economy. In a similar vein, Canada and the United States (U.S.) signed a free trade agreement in 1988 to lower their already low barriers to bilateral trade. Casual empiricism suggests that more countries are losing full control over their macroeconomic policy instruments (such as interest rate levels) as internationalization increases. The consequences of internationalization have also spread to specific sectors of the economy, such as capital markets. The implications for financial decisions, whether or not financial assets are priced in a national or international context, are well known and need not be addressed further.

According to Jorion and Schwartz (1986, p.604): "... integration imposes restrictions on the pricing of assets, by ruling out relationships between expected returns and purely domestic factors .... On the other hand, complete segmentation implies that only national factors ... should enter the pricing of assets." Jorion and Schwartz tested the Capital Asset Pricing Model (CAPM) and the International Capital Asset Pricing Model (ICAPM) using monthly returns for a sample containing only Canadian stocks. They concluded that the Canadian equity market is not completely integrated (or segmented) relative to a global North American equity market. Not only were these two asset pricing models poor descriptions of how Canadian stocks were priced for the period studied by Jorion and Schwartz but they also required the identification of their respective market portfolios. However, as noted by Roll (1977), the market proxies used are unreliable because the domestic and world market portfolios are unobservable. In addition, due to differences in consumption baskets across countries and in exchange rate uncertainties, the ICAPM predicts that investors will hold different optimal portfolios, especially hedge portfolios [Adler and Dumas (1983)]. Since single factor models are used in the derivations of the CAPM and the ICAPM, these two models used by Jorion and Schwartz do not allow for explicit and unambiguous tests of the very plausible hypotheses of partial integration and partial segmentation. This led Jorion and Schwartz (p. 613) to conclude their paper with the following statement: "The methodology could be extended to more general multifactor asset pricing models, and it would be interesting to see whether purely national factors also lead to rejection of integration."

The deficiencies inherent in earlier empirical tests of capital market integration and segmentation can be rectified by modifying the more general and less restrictive International Arbitrage Pricing Theory (IAPT). This model was developed by Solnik (1983),[11] and was used in its unmodified form by Gultekin et al. (1989) to examine whether the Japanese and U.S. capital markets were integrated or segmented before and after the enactment of the Foreign Exchange and Foreign Trade Law in December of 1980. The IAPT can be modified to encompass the complete integration and the partial integration hypotheses, if the macrofactors of the whole of North America can be identified. Specifically, to test whether or not the Canadian equity markets are completely or partly segmented relative to a North American equity market, Ross's (1976) APT model with purely domestic factors needs to be modified to allow for the possible influence and pricing of purely international factors.

The purpose of this chapter is three-fold: first, to generalize the IAPT and the domestic APT into two separate comprehensive forms in which the competing hypotheses of complete and partial market integration and segmentation are nested within each model; second, to identify the priced North American macroeconomic factors relevant to these models; and third, to test empirically whether or not the Canadian equity market is integrated or segmented relative to one complete North American market.[12] Unlike previous studies, the binational factors from the Frankel (1979) exchange rate determination model are used herein.

The remainder of the chapter is organized as follows: Section two reviews the existing literature on international market integration versus segmentation. Section three derives the model and the hypotheses to be tested. Section four describes the data. Section five presents and discusses the empirical results. Section six concludes the chapter.

3.2 REVIEW OF THE LITERATURE

The general definition of capital market integration that emerges from the existing literature is that capital markets are completely integrated if investors can earn the same expected return on investments in each market after adjusting for risk and currency differences. For example, the capital markets situated in New York, Chicago, San Francisco and Los Angeles are expected to satisfy the above definition, since they are completely integrated within the U.S. capital market and economy. However, whether capital markets situated within different national boundaries are integrated or segmented, and to what degree, is less obvious. To test for integration or segmentation of capital markets, three asset pricing models (APM's) that account for systematic and exchange rate risks are used in the literature. Each is discussed in turn.

3.2.1 ICAPM studies

Solnik (1974a) developed an intertemporal equilibrium model of an international capital market in the framework of the Sharpe-Lintner-Mossin CAPM. In an international setting, the market portfolio is not mean-variance efficient since investors will hold different portfolios due to the need to hedge against foreign exchange risk. In an exploratory analysis using the ICAPM, Solnik (1974b) found that both a strong domestic and international factor influence on stock returns.

Solnik (1977) was unsuccessful in discriminating statistically between his (1974a) model and the Grauer et al. (1976) model. The two models have different asset pricing equations due to different exchange risk specifications. Exchange risk stems from differences in consumption tastes between countries in the Solnik model, and from uncertain monetary inflation in the Grauer et al. model.

Using the Fama-MacBeth two-step procedure and a capital market equilibrium model based on a multiperiod logarithmic utility function, Stehle (1977) tested if the U.S. equity market was completely integrated or segmented from other stock markets. His statistical tests provided no support for either the domestic or the international versions of his model. Using a two-factor return-generating model with a domestic and international factor [as in Stehle (1977)], Errunza and Losq (1985) tested the segmentation versus the integration hypothesis. They found support for the hypothesis of segmentation in the world market.

Jorion and Schwartz (1986) examined whether the Canadian equity market was integrated or segmented relative to a global North American market over the period 1968-1982. Using the two-factor ICAPM methodology of Stehle, they found that the Canadian market was segmented. Like Stehle (1977), Jorion and Schwartz assumed that all investors had a logarithmic utility function. Unlike Stehle (1977), Jorion and Schwartz used the maximum-likelihood estimation approach to minimize measurement error.

3.2.2 ICCAPM studies

Stulz (1981) constructed (but did not test) an intertemporal consumption-based capital asset pricing model (ICCAPM) which admits differences in consumption opportunity sets. The model yields an asset pricing equation which states that the real expected return of a risky asset is proportional to the covariance of the return of that asset with changes in the world real consumption rate. Using a discrete version of Stulz's consumption-based model, Wheatley (1988) could not reject the hypothesis that equity markets are integrated internationally for the U.S. and for 17 other countries for the period 1960-1985.

3.2.3 IAPT studies

Cho et al. (1986) tested the IAPT using factor analysis. Although they found between one and five common factors, they conducted no explicit test of the market segmentation hypothesis against the alternative of integration.

Unfortunately, there are various methodological problems associated with the use of factor analysis for testing the segmentation hypothesis using the IAPT, because identification of the revealed factors is ad hoc. Moreover, the number of factors identified as being significant in factor analysis generally exceeds the number of priced factors (usually one). The likelihood ratio test tends to overstate the significant number of factors, and the χ^2-test not only reduces the marginal impact of a particular factor as the number of factors increases, but does not provide any indication about the exactness of the model [Dhrymes et al. (1985)]. Chen (1983) reports that the number of priced factors is not invariant to rotation. If the factor structure contains k factors, then the number of priced factors may be lower than k. Based on Monte Carlo simulations, Brown (1989) demonstrates that factor analysis understates the significance of the risk premia on factors beyond the first one (i.e., it is biased towards choosing one factor even if an exact k-factor model is prespecified). This confirms the Kryzanowski and To (1983) and Trzcinka (1986) findings that at least one priced factor (mimicking portfolio) exists in the factor structure. Measurement error in the factor loadings results from the well-known two-pass procedure required for the construction of mimicking portfolios. As Burmeister and McElroy (1988) show, avoiding this source of measurement error positively affects the robustness of the resulting estimations. Kryzanowski and To (1983) demonstrate that the number of statistically significant factors increases as the number of securities in the test increase [also, see Dhrymes et al. (1985)].

As noted by Anderson (1984) and Conway and Reinganum (1988), the properties of estimated factor loadings are unknown when the errors are not jointly normal, the estimates of the factor loadings are not unique; moreover, the estimated factor scores (prices of risk) are not invariant to the way assets are apportioned to portfolios, and the first factor obtained from the first portfolio may not be the same as the first factor derived from the second portfolio. Furthermore, the use of factor analysis as a statistical tool without any simultaneous grounding in economic theory (intuition) makes interpretation difficult and hence renders any findings of little operational significance.

Gultekin et al (1989) examine whether the Japanese and U.S. capital markets are integrated or segmented. The authors use a multi-index APM with identified factors (as well as factor analysis with unidentified factors) to test whether or not the enactment of the Foreign Exchange and Foreign Trade Control Law in December of 1980 had a significant impact on eliminating capital controls, and thus assisted in integrating the financial markets of the two countries. Their research findings support the segmentation hypothesis prior to the enactment of the December 1980 law and the integration hypothesis thereafter.

Concurrently with this thesis, Mittoo (1992) examined whether the Canadian and U.S. equity markets are integrated or segmented. Using the

APT she found evidence of segmentation for the period from 1977-81 and integration for the period from 1982-86. Given her sample size of only twenty-one Canadian stocks and her methodological difficulties [she does not define and test an international (integrated) APT model against the alternative of a domestic (segmented) APT model of the Canadian and U.S. stock returns] her results are questionable.

3.3 THEORETICAL DERIVATION OF THE IAPT-BASED TESTS FOR INTEGRATION AND SEGMENTATION

The IAPT can be derived as follows. Suppose that the returns on the set of n assets follow the following international linear factor model (LFM):

$$\tilde{R}_{it} = E(\tilde{R}_{it}) + b_{i1}\tilde{I}_{1t} + \ldots + b_{iK}\tilde{I}_{Kt} + \tilde{\varepsilon}_{it}, \quad \begin{array}{l} i = 1,\ldots,n \\ t = 1,\ldots,T \end{array} \qquad (3.3.1)$$

where $\tilde{R}_{it}$ is the random return of the i'th asset for the period ending at time t; $E(\tilde{R}_{it})$ is the rational expectation of the random return of asset i given information at the beginning of time t; $\tilde{I}_j$ is the zero-mean international common factor j; b_{ij} is the sensitivity of asset i to the jth factor; $\tilde{\varepsilon}_i$ is the idiosyncratic risk of asset i, where $E(\tilde{\varepsilon}_{it}/\tilde{I}_{js}) = 0$ for $\forall$ i,j and $\forall$ t and s and $E(\tilde{\varepsilon}_i^{\,2}) = \sigma_i^{\,2} < \infty$; and $E(\varepsilon_{it}\varepsilon_{vs}) = \sigma_{iv}$ for i = v and t = s and zero otherwise. The returns can be expressed in terms of a given numeraire currency, such as the Canadian dollar.

If investors have homogeneous beliefs about the LFM, then nonarbitrage arguments dictate the following pricing equation:

$$E(\tilde{R}_{it}) = \lambda_{ot} + \lambda_1 b_{i1} + \ldots + \lambda_K b_{iK}, \quad i = 1, \ldots, n \tag{3.3.2}$$

where λ_j is the price of risk for undertaking the systematic risk of the factor j associated with holding asset i, and λ_0 is the risk-free rate. Solnik (1983) demonstrates that equation (3.3.2) is valid irrespective of the chosen numeraire, provided that the exchange rates follow the k-factor model of equation (3.3.1). Substituting equation (3.3.2) into equation (3.3.1) yields the following pricing system of n-equations:

$$\tilde{R}_{it} = \lambda_{ot} + \sum_{j=1}^{K} b_{ij} (\tilde{I}_{jt} + \lambda_j) + \tilde{\varepsilon}_{it}. \qquad \begin{array}{l} i=1,\ldots,n \\ t=1,\ldots,T \end{array} \tag{3.3.3}$$

This pricing relationship is what Shanken (1985) has labelled the arbitrage APT and holds only as an approximation. However, Dybvig (1983) estimated a conservative error bound of .015% per month. As shown in Connor (1984), Wei (1988), and Burmeister and McElroy (1988), including the residual market factor (RMF) in (3.3.3) yields an exact APT. The residual market factor is intended to capture pervasive macrofactors (international and/or domestic) omitted from the approximate pricing system (3.3.3).

To test the integration hypothesis, the complete IAPT model described by equation (3.3.3) is augmented to allow for the possible influence and pricing of domestic factors, $\tilde{D}_g$. The augmented system becomes:

$$\tilde{R}_{it} = \lambda_{0t} + \sum_{j=1}^{K} b_{ij} (\tilde{I}_{jt} + \lambda_j) + \sum_{g=1}^{G} c_{ig} (\tilde{D}_{gt} + \theta_g) + \tilde{\varepsilon}_{it}, \qquad (3.3.4)$$

$$i=1,\ldots,n$$
$$t=1,\ldots,T$$

where c_{ig} is the sensitivity of asset i to the domestic factor, g, and θ_g is the price of risk associated with factor g.

The $E(\tilde{I}_{jt}\tilde{D}_{gt})$ may not all be equal to zero since one country's national factors may constitute some nonsignificant portion of the international factors. The correlation between the domestic and world market portfolios in tests of the segmented versus integrated hypothesis was first noted by Stehle (1977). His suggested solution was used by Stehle (1977), Jorion and Schwartz (1986), and Gultekin et al. (1989). The solution is to use projections to separate the pure national factors from the pure international factors.[13] To do this, regressions of the form:

$$\tilde{D}_{gt} = \alpha_o + \underline{\alpha}_1' \underline{\tilde{I}}_t + \tilde{V}_{gt}, \quad g = 1,\ldots,G \qquad (3.3.5)$$

are run to obtain the residual values, $\hat{V}_{gt}$, for every g=1,..., G. The fitted values, which are the components of the domestic factors that are orthogonal to the international factors can then be used as measures of the pure national factors. The $\hat{V}_{gt}$ replace the $\tilde{D}_{gt}$ in equation (3.3.4) to obtain:

$$\tilde{R}_{it} = \lambda_{0t} + \sum_{j=1}^{K} b_{ij} (\tilde{I}_{jt} + \lambda_j) + \sum_{g=1}^{G} c_{ig} (\hat{V}_{gt} + \theta_g) + \tilde{\varepsilon}_{it}, \qquad (3.3.6)$$

$$i=1,\ldots,n$$
$$t=1,\ldots,T$$

The hypothesis of complete integration can then be tested by examining whether the $\theta_g = 0$, $\forall$ g. Partial integration is easily tested by examining whether some of the θ's are significantly different from zero.

To test the segmentation hypothesis, the domestic APT (where only national factors enter the LFM) is augmented to allow for the possible influence and pricing of international factors as follows:

$$\tilde{R}_{it} = \lambda_{0t} + \sum_{g=1}^{G} c_{ig} (\tilde{D}_{gt} + \theta_g) + \sum_{j=1}^{K} b_{ij} (\hat{U}_{jt} + \lambda_j) + \tilde{\mu}_{it}, \qquad (3.3.7)$$

$$i=1,\ldots,n$$
$$t=1,\ldots,T$$

where the $\hat{U}_{jt}$ are the purely international factors. The following regressions:

$$\tilde{I}_{jt} = \delta_o + \underline{\delta}_1' \tilde{\underline{D}}_{gt} + \tilde{U}_{jt}, \quad j = 1,\ldots, K \qquad (3.3.8)$$

are run to obtain the residual values, $\hat{U}_{jt}$, $\forall$ j. The hypothesis of complete segmentation (i.e., H_0: $\lambda_j = 0$, $\forall$ j), and the hypothesis of partial segmentation (where some of the λ's are different than zero) is easily tested.

Theoretically, if the modified IAPT given by the system of equations (3.3.6) predicts that the Canadian equity market is completely integrated relative to a North American equity market (i.e. θ_g=0 for every g=1,...,G and some λ_j not equal to zero), then the domestic APT given by the system of equations (3.3.7) should reject both the complete and partial segmentation hypotheses (i.e. all or some of the λ_js should be significant and all θ_gs should be equal to zero). Conversely if the domestic APT predicts that the Canadian equity market is completely segmented relative to a North American equity market, then the IAPT should reject both the complete and partial integration hypotheses. However, if the IAPT predicts that the Canadian equity market is partly integrated relative to a North American equity market, then the domestic APT should also predict the partial segmentation of the Canadian equity market relative to a North American equity market. Thus, the two models (IAPT and the domestic APT) are competitors for modelling asset prices only if the extreme hypothesis of either complete market integration or segmentation holds. Otherwise, the two models are equivalent. Therefore, past studies that include variables (like the exchange rate and interest rates on eurodeposit rates) are implicitly assuming partial integration. For example, Brown and Otsuki (1989) modelled the Japanese equity market using these variables because of the internalization of the Japanese economy (primarily due to its large export market).[14] The international APT and domestic APT models given by the system of equations (3.3.6) and (3.3.7) respectively are estimated herein using both the iterated non-linear ordinary least squares (NOLS) and the iterated non-linear seemingly unrelated regression (NSUR) techniques.[15] In the NOLS method the error covariance matrix is assumed to be diagonal, whereas the NSUR method assumes a full error covariance matrix. As shown by Gallant (1985), these NOLS and NSUR estimators are strongly consistent and asymptotically normal, even in the absence of normally distributed errors. If errors are normally distributed, these estimators are also maximum-likelihood estimators. In our application, n=50 and T=228.

A test of any hypothesis based on a theoretical model, like an asset pricing model, is a joint test of at least three separate hypotheses. The first hypothesis is that the model is valid (which in our case is supported by considerable empirical evidence); the second is that the actual hypothesis being tested is valid; and the third is that the computed test statistic is drawn from the hypothesized distribution (which is a problem using factor analysis). In addition, since the APT does not specify which factors should enter the LFM (which is somewhat rectified herein by using those factors in the Frankel (1979) exchange rate determination model) the additional hypothesis of whether the chosen macrofactors are the true factors is also jointly tested. Therefore, while a specific set of tests may accept or reject segmentation or integration, the inference may actually be due to the failure of any one or any combination of the other implicitly maintained hypotheses. Unfortunately, no easy solution exists for this problem since testing of each hypothesis separately is virtually impossible. By performing multiple statistical tests and letting the data choose from many intuitively plausible macrofactors, the researcher hopes to minimize the possibility of failure coming from the falsity of the implicit maintained hypotheses.

To estimate the system of equations in (3.3.6) [and in (3.3.7)], the variables which constitute the international factors, $\tilde{I}_j$'s, [and the national factors, $\tilde{D}_g$,] must be determined. Since general factors are expected to come from the variables that affect the macroeconomy, some of the potential candidates for the domestic factors are output, interest rates, money supply, and inflation. To identify potential binational factors, recall that the IAPT is only valid if the exchange rates of the countries involved and the returns of the equities follow the same linear factor structure. Therefore, a priori information on the structure of exchange rates can be used to identify potential "international" (binational) common factors. Since the IAPT described by equations (3.3.1) and (3.3.2) is derived under the hypothesis of complete market integration, the common factors included in (3.3.1) and priced by (3.3.2) are "international" (or binational) factors, if segmentation of the Canadian equity market relative to the North American market is being examined.

These binational factors can be determined explicitly based on the literature on exchange rate determination. One general form of an exchange rate determination model is the following:

$$s_t = \sum_{q=1}^{M} d_q \; (X_{qt} - X^{*}_{qt}) + \omega_t, \tag{3.3.9}$$

where s_t is the log of the spot rate; X_{qt} and X^{*}_{qt} are the logs of the q'th domestic and foreign macroeconomic variables, respectively; d_q is the sensitivity coefficient of the spot rate to the difference of the q'th domestic and foreign macroeconomic variables; and ω_t is an equation error with zero mean and constant variance. Equation (3.3.9) encompasses all significant exchange rate models as different restrictions on the coefficients. In Frankel (1979), equation (3.3.9) takes the form:

$$s_t = (m_t - m^*_t) - \phi(y_t - y^*_t) - \alpha(r_t - r^*_t) + \beta(\pi_t - \pi^*_t) + U_t \qquad (3.3.10)$$

where s_t is the log of the spot rate; m_t and m_t^* are the logs of the money supply at home (Canada) and abroad (U.S.), respectively; y_t and y_t^* are the logs of output at home and abroad, respectively; r_t and r_t^* are the logs of one plus the domestic and foreign interest rates, respectively; Π_t and Π_t^* are the current rates of expected long-run inflation at home and abroad, respectively; and U_t is an equation error with zero mean and constant variance.

Taking conditional expectations of equation (3.3.9) on the available information set I at time t yields:

$$E_t(s_t/I_t) = \sum_{q=1}^{M} d_q \; E_t(X_{qt} - X^*_{qt}/I_t). \qquad (3.3.11)$$

Subtracting equation (3.3.11) from equation (3.3.9) yields:

$$s_t = E_t(s_t/I_t) + \sum_{q=1}^{M} d_q \; [(X_t - X^*_t) - E_t(X_t - X^*_t/I_t)] + \omega_t. \qquad (3.3.12)$$

Equation (3.3.12) is the LFM for the exchange rate of an integrated economy based on a general exchange rate determination model. Therefore, the binational common factors are the terms specified in the brackets. Clearly, these factors have zero means, and they conform to previous expectations of what kinds of variables should constitute binational factors. The need to use projections to filter out the purely (bi)national factors is now evident, since the national and binational factors are correlated.

Based on Frankel (1979), some of he binational factors, which can be used, are the following innovations:

$$F_{1t}=\left[(m_t-m_t^*)-E_t(m_t-m_t^* /I_t)\right],\quad F_{2t}=\left[(y_t-y_t^*)-E_t(y_t-y_t^* /I_t)\right],$$

$$F_{3t}=\left[(r_t-r_t^*)-E_t(r_t-r_t^* /I_t)\right],\quad F_{4t}=\left[(\pi_t-\pi_t^*)-E_t(\pi_t-\pi_t^* /I_t)\right].$$

The determination of the conditional expectations, $E_t(\ /I_t)$, is discussed in the next section of this paper.

3.4 DATA

The monthly rates of return on all stocks traded on the Toronto Stock Exchange (TSE) from January 1969 through March 1988 are used herein to calculate the returns on fifty size-ranked portfolios. These portfolios are used in order to gain the maximum dispersion in returns [Poon and Taylor (1991)], since it is well-known that small capitalization firms have higher betas and average returns than larger market value firms. Thus, using large number of portfolios tends to increase dispersion in betas, which, in turn, leads to more precise estimates.

The macroeconomic factors that are used to obtain the innovations which affect stock returns are all taken from Statistics Canada's CANSIM Mini Base. All the variables used herein are defined in Table 3.1. The growth rates of most variables (except for the interest rate variables) are calculated by taking the first differences in their logarithms. The innovations are estimated as the forecast errors of a multivariate model fitted for each of the macroeconomic series using Akaike's (1976) multivariate state-space procedure. Brown and Otsuki (1989) and Kryzanowski and Zhang (1992) have used this innovative method to produce forecast errors that have zero means and are serially uncorrelated.

Since no widely accepted theories exist for linking stock market performance and specific economic variables, other considerations are used to determine which macroeconomic variables might affect stock returns. Specifically, economic intuition, findings from past empirical studies, and frequently published macroeconomic variables in the popular business media provide the basis for the selection process.

Innovations for two groups of macroeconomic variables are estimated. The first group contains domestic (or national) factors; the second group contains "international" factors. Both groups consist of variables designed to capture the real, monetary and financial sectors of the domestic (Canadian) and international (North American) economies. National factors, which capture the real sector of the Canadian economy, include industrial production, gross domestic product and the Canadian composite leading index. The monetary sector of the Canadian economy is captured by the money supply and inflation (as measured by the change in the Consumer Price Index or CPI). The financial sector is captured by different interest rate variables, such as the yields on long-term Canada's, long-term corporates and 91-day Treasury bills. These yields are used to create two additional variables, the term structure variable, TERM, which is defined as the difference in the yields on long-term government bonds and 91-day Treasury bills, and the risk premium variable, RISK, which is defined as the difference between the yields on long-term corporate bonds and long-term government bonds.

Except for Kryzanowski and Zhang (1992), all previous studies have used the TERM and RISK variables directly as innovations. For example, Chen et al. (1986) assume that both of these variables are uncorrelated because they are differences in interest rates, and therefore can be considered as innovations. However, these variables do not have zero means because the term structure is upward sloping more often than it is downward sloping, and the risk premium for holding the more risky corporate bonds instead of the less risky government bonds is always positive. Therefore, these variables do not satisfy the mean-zero property of proper innovations. McElroy and Burmeister (1988) used the term structure variable directly as an innovation because the mean of this variable was equal to zero in their sample. They modified the risk premium variable by adding a constant to make the sample mean equal to zero. Both of these variables are fitted herein using the state-space procedure [Akaike (1976)] to determine their innovations.

The second group of variables contains "international" (binational) macroeconomic factors which are constructed from both Canadian and U.S. macroeconomic variables. The only variable that is truly international is the interest rate offered on 3-month U.S. dollar deposits in London. This eurorate, which is intended to capture term structure effects, was used by Gultekin et al. (1989). The other "international" variables are constructs similar to the ones used in tests of exchange rate determination. The differentials in the growth rates of the Canadian and U.S. composite leading indices and industrial production are binational factors that capture real economic activity in North America. The binational monetary sector is captured by the growth differential in the Canadian and U.S. money supplies and the corresponding growth differential for inflation rates. The financial sector is captured by the differentials in the term structure and risk premium variables of the two countries. These differentials in the macroeconomic variables and the eurorate are each fitted using the state-space procedure to obtain the binational innovations.

The interpretation of the binational innovations is straightforward as is demonstrated by the following example. Suppose X_t and X_t^* are the growth rates at time t of the Canadian and U.S. macroeconomic variables, respectively. The binational factor is $DX_t=(X_t-X_t^*)$. Fitting this variable with the state-space procedure yields the following binational innovation: $[DX_t-E_t(DX_t)]=(X_t-X_t^*)-E_t(X_t-X_t^*)=(X_t-E_t(X_t))-(X_t-E_t(X_t^*))$, where $E_t(\ \ )$ is the conditional expectation operator based on the information at time t. This binational innovation is just the differential innovation of X_t and X_t^* across the two countries.

Summary statistics for the innovations are reported in Table 3.2. The t-statistics for testing the null hypothesis that the innovations have zero means are not significant for each and every macroeconomic innovation. The Kolmogorov D-statistics for testing the null hypothesis of normally distributed innovations is only rejected for the following three variables: term structure, term structure differential and money differential. The Fisher kappa test statistics for testing the null hypothesis that the innovations are white noise is rejected for the following four variables: the risk premium on corporate bonds, the differentials in the money supply and industrial production and the Canadian leading indicators index. The Ljung and Box Q-statistics for testing the same null hypothesis of white noise innovations is not rejected for any of the variables.

To ensure that each domestic factor does not contain information embodied in any of the other domestic factors, a series of sequential regressions is performed. Specifically, the second domestic factor is regressed against the first domestic factor and the residual from this regression is used as the second factor. Then, the third domestic factor is the residual of a regression of itself on the first and second domestic factors. The same procedure is followed for the rest of the domestic factors to get orthoganalized domestic factors by construction. The same methodology is followed to orthogonalize each of the international factors to the other international factors. Then, equations (3.3.6) and (3.3.7) are estimated using the orthogonalized domestic and international factors to determine the purely domestic component of each of the domestic factors and the international component of each of the international factors. However, except for a few subperiods, the results are robust when the factors are not completely orthogonalized, because the innovations from using the state-space procedure have correlations close to zero.

3.5 EMPIRICAL RESULTS

3.5.1 Selection of the macroeconomic variables

To determine which macroeconomic innovations enter the international LFM, a series of linear regressions are performed. In these regressions, the dependent variables are total returns on the Toronto Stock Exchange 300 stock composite index (TSE300), the equally-weighted (EW), and value-weighted (VW) stock indices of all the stocks listed on the TSE, and the fifty equally-weighted size-ranked portfolios. The independent variables are the innovations of the international factors identified in Table 3.1. Although the LFM states that stock returns at time t should only be affected by the macroeconomic innovations at time t, lagged values are also used as independent variables in the stepwise regressions. As noted by Chen et al. (1986), this may capture unsynchronized information arrival, the information for some variables lagging equity returns by as much as one month. The variables that are significant for most of the portfolios and the market indices are selected as being the binational macroeconomic factors.

The selected international LFM's for the three market indices are reported in Table 3.3. The last column of the table reports the average parameter values of the LFM for the fifty size-ranked portfolios. The binational factors that enter the LFM's significantly for the sample period from March 1969 to March 1988 are: the lag of the differential in the industrial production innovations of Canada and the U.S (DLINDUS), the differential in the composite leading indices (DINDEX), and the interest rate innovation on eurodollar deposits (REURO). These factors together explain approximately fifteen to twenty percent of the variation in the returns on the market indices and ten percent of the variation in the returns on the fifty portfolios. The p-values indicate that these three variables are individually significant for conventional significance levels in the regressions. Also the F-values indicate that the overall regressions are significant at the .01 significance level.

The domestic LFM for the three indices and the mean parameter estimates of the LFM's for the fifty portfolios are reported in Table 3.4. Based on the mean t-values, the three significant domestic factors are: the lag of industrial production (LINDUS), the term structure (TERM), and the Canadian leading index (INDEX). These factors combine to explain approximately sixteen percent of the variation in the returns on the market indices and ten percent for portfolio returns.

Since an international factor may be significant in the international LFM because of its domestic component and a domestic factor may be significant in the domestic LFM because of its international component, the LFM's for the market indices and the fifty size-ranked portfolios are re-estimated with each factor decomposed into its pure component. The results from these regressions are reported in Table 3.5. These results indicate that the pure international component of the term structure is insignificant in the LFM's for the market indices and the fifty size-ranked portfolios. Kryzanowski and Zhang (1992) found that, while the term structure variable is priced in the Canadian equity market over the period from February 1956 to March 1988, its beta became insignificant when the LFM with the APT restrictions was estimated. One possible reason may be the fact that only the domestic component of the term structure is significant and priced, while its international component is irrelevant for pricing Canadian stocks. Inclusion of both components in the LFM may have an overall diluting effect on the significance of the term structure beta coefficient. The other variable, that is significant at the market level, but insignificant at the portfolio level, is the pure international component of the differential in industrial production. The D-W (Durbin-Watson) statistics of these regressions are close to two, indicating no apparent misspecification due to serial correlation. Judged by the high F-statistics, none of the regressions may be rejected at the .01 significance level, even though the overall fit, as judged by the R^2, never exceeds 16.04% of the total variation.

For all the models in Tables 3.3, 3.4 and 3.5, there is no evidence of first-order serial correlation or heteroscedasticity in the error terms at traditional significance levels based on the D-W statistics, the Ljung and Box Q-statistics, the Breusch-Pagan-Godfrey (B-P-G) test statistics and ARCH test statistics and their associated probability values.[16]

3.5.2 Tests of the segmentation and integration hypotheses

The estimated risk premia for the augmented IAPT given by equation (3.3.6) are reported in Table 3.6. The three international factors, REURO, DINDEX and DLINDUS and the domestic components of LINDUS, INDEX and TERM (denoted by LINDUSD, INDEXD and TERMD, respectively) are the independent variables of the model, and the fifty equally-weighted size-ranked portfolios are the dependent variables. The estimated risk premia using NOLS and NSUR are reported in panels A and B, respectively.

Unlike the NOLS method, the NSUR method uses the full covariance matrix of errors which yields more efficient estimates when across equation restrictions are present as is the case herein. This explains the differences in the size and significance of the estimated risk premia reported in the table. While the risk premia for the entire sample period associated with INDEXD and DLINDUS are significant and insignificant, respectively, with NOLS, they are insignificant with NSUR.[17] The NSUR estimates of the risk premia for the period from March 1969 to March 1988 are significant for the two international factors, REURO and DINDEX, and the domestic components of two domestic factors, LINDUSD and TERMD. Therefore, H_0, the complete integration hypothesis that some "international" factor(s) is (are) priced and all pure domestic factors are not priced, is rejected in favour of the partial integration hypothesis H_α: that some "international" factor(s) is (are) priced and at least one purely domestic factor is priced. Subperiod estimates of the augmented IAPT indicate that in all of these subperiods the Canadian equity markets were only partly integrated relative to a North American equity market since some of the binational and some of the domestic risks were priced indicating less than complete integration. Another interesting result is the time-variation of the risk premia. Overall, the partial integration hypothesis cannot be rejected for the Canadian equity markets.

The estimated risk premia for the augmented IAPT with the residual market included as an additional factor are reported in Table 3.7. The additional factor is intended to capture the influence of any omitted macroeconomic variables. Again, no significant differences exist between the NOLS and NSUR estimates. This indicates that the across-equation error correlations are insignificant. For the whole period, the NOLS estimates of the risk premia are all insignificant. The NSUR method yields significant estimates for the domestic component of the domestic factor, TERMD, and the international factors, DINDEX and REURO. This indicates that the partial integration hypothesis cannot be rejected for the Canadian equity market. The residual market factor (RMF) is found to be significant for the subperiod from March 1969 to March 1978.

In Tables 3.8 and 3.9, the mean factor loadings for the augmented IAPT (without and with the RMF, respectively) are reported. By comparing panels A and B, it can be seen that NOLS and NSUR yield similar beta estimates. Subperiod estimates of the factor loadings indicate that the betas are time-varying. This confirms what has already been found in the literature using factor analysis [e.g., Cho and Taylor (1987)].

The NOLS and NSUR estimated risk premia of the augmented domestic APT, given by equation (3.3.7), are reported in Panels A and B, respectively, of Table 3.10. The NOLS estimates indicate that the variables, TERM and REURO, are priced risks. The NSUR estimates of the risk premia indicate that the complete segmentation hypothesis, H_0, that some domestic factor(s) is (are) priced and all "international" factors are not priced, is rejected since the domestic factor TERM has priced risks and two of the "international" factors, REURO and the "international" component of the leading index (DINDEXI), are priced risks. However, the partial segmentation hypothesis cannot be rejected, since at least one of the domestic factors is priced.

The NOLS and NSUR estimates of the risk premia for the augmented APT with the RMF included as an independent variable are reported in Table 3.11. The NOLS estimates indicate partial segmentation since one domestic factor, TERM, and the binational component of the differential in the Canadian and U.S leading indicator indices, DINDEXI, are significant. Using the NSUR method, DINDEXI and RMF are priced risks for the entire period from March 1969 to March 1988. Since it is not known whether the RMF factor is capturing the influence of omitted domestic, binational or both factors, there is ambiguity regarding partial segmentation. However, for both subperiods, the estimated risk premia clearly indicate that the partial segmentation hypothesis cannot be rejected. Therefore, for the whole period, the partial segmentation hypothesis cannot be rejected, and the RMF factor has to include domestic omitted variables.

The mean factor loadings are reported in Tables 3.12 and 3.13 for the augmented APT without and with the RMF factor, respectively, for the whole sample, and for two different subperiods. The estimates for the two models are not markedly different. However, the time-variation of the betas is again confirmed.

3.6 CONCLUDING REMARKS

In this chapter, the hypotheses that the Canadian equity market is completely (partly) integrated and segmented relative to a complete North American equity market are tested using modified forms of the IAPT and APT models. For the period from March 1969 through March 1988, both models indicated that the Canadian equity market is only partly integrated (or equivalently, partly segmented) with the American equity market. The Canadian macroeconomic factors found to influence Canadian stock returns are the pure domestic component of the term structure and the lag of the industrial production index. The North American macroeconomic factors affecting Canadian returns are the pure "international" components of the differential in the Canada/U.S. leading indicators, and the interest rate on U.S. dollar deposits in London (i.e., eurodeposits).

CHAPTER FOUR: EXPLANATION OF STOCK RETURN VARIABILITY USING PRESPECIFIED ECONOMIC FACTORS WITH TIME-VARYING RISK PREMIA

4.1 INTRODUCTION

The issue of whether financial asset markets are informationally efficient has been one of the most important subjects of research in the field of financial economics during the last three decades. Numerous studies have followed the pioneering work of Fama (1970) where the notion of asset market efficiency was given a formal and widely accepted interpretation.

Fama (1970) defines informational efficiency for three different information sets. If financial asset prices fully reflect all past information, then the market is said to be efficient in the weak-form. If all current public information is reflected, the market is deemed to be efficient in the semi-strong form. Finally, if all public and private information is reflected in current asset prices, the market is deemed to be efficient in the strong-form. Fama (1991) groups tests for weak-form efficiency into the more general category of tests for return predictability; semi-strong-form tests fall under the category of event studies; and strong-form tests are referred to as tests for private information.

Evidence provided by Shiller (1981a, 1981b), LeRoy and Porter (1981) and subsequently many other researchers show that stock prices are too volatile to be determined in an informationally efficient market. According to West (1988b) and others, the basic idea behind these volatility tests is to compare two magnitudes, V and V^*. V measures the volatility of the market's forecasts of fundamental asset prices, and V^* is the volatility of the econometrician's measure of the same fundamental asset prices. If V is greater than V^*, then excess volatility is indicated.

To estimate V^*, an asset pricing model is specified. All past studies, which use the simple Discounted Cash Flow (DCF) model to create the econometrician's forecast of asset prices, cannot reject the hypothesis that financial asset markets are too volatile to be efficient. However, the DCF model is only valid if investors are assumed to be risk-neutral.[18] This assumption implies that economic agents are expected return maximizers who ignore the risk characteristics of different investments. This implication is contrary to the foundations of modern financial economics that assume a tradeoff between risk and return for risk averse agents. Risk averse investors require higher risk premia in order to hold assets with increasingly uncertain returns (i.e., higher risk). Merton (1973), Cox, Ingersoll and Ross (1985) and Breeden (1986) relate the risk premium on risk factor j to its own time-varying volatility.

Schwert (1989) departs from traditional volatility research that uses the DCF model to explain stock market volatility by examining the underlying systematic factors causing dividends to change, such change inducing variability in stock prices. In his exploratory analysis he associates time-varying stock return volatility with the time-varying volatility of different macroeconomic and financial variables. However he does not make any attempt to test for causes of stock price volatility as in the volatility literature using a specific asset pricing model. Roll (1988) associates asset price changes with (systematic) contemporaneous news events. He finds that less than 40% of the variance of price changes is explained by the regressions. Fama (1990) expands on Roll's methodology for explaining stock price changes by including both contemporaneous and leading variables in his regressions. He finds that about 58% of the variation of the NYSE value-weighted returns is explained by the regressions. Unlike the volatility tests, that are based on some specific asset pricing model, the exploratory analysis of Roll and Fama does not use an asset pricing model to explain observed stock return variability. Thus if a specific asset pricing model is used along with identified macroeconomic factors (unlike Roll and Fama), it may be possible to explain more of the observed stock return variability. It may also be possible to reject the hypothesis of market inefficiency implied by the volatility literature.

The Arbitrage Pricing Theory (APT) model [developed by Ross (1976)] with identified macroeconomic factors incorporates investor's risk aversion and rests on the direct relationship between risk and return. The systematic risk factors (measured as the innovations on various important macroeconomic variables) that influence asset returns are explicitly priced by the APT to provide the market determined risk premia to the risk averse investor. This makes investments with differing risk characteristics comparable. The APT model with constant risk premia within a given period, but varying across time-periods, and the APT model with time-varying risk premia within and across periods are used to generate an intertemporal sequence of predicted returns. These predicted returns allow for the hypothesis, put forward by Pindyck

(1984), that excess variability, a possible violation of the efficient market hypothesis, may be explained by changes in risk that induce changes in risk premia and consequently change expected returns. Grossman and Shiller (1981), Campbell and Shiller (1988) and West (1988a) allow for time-varying expected returns in the DCF model using a consumption-based asset-pricing model. In this study the APT is used to capture the possible time-variation of expected returns in the return-generating process (i.e., the linear factor model in the APT literature).[19]

The two principal aims of this chapter are: first, to use the APT with identifiable macroeconomic factors to construct the predicted stock returns; second, to compare the variability of this measure to the variability of actual returns to determine how much of the observed stock return variability is explained (and justified) by the time-variation in risk premia and the variation in economic fundamentals.

The remainder of the chapter is organized as follows. In section two, the use of the APT model to derive predicted returns is explained. In section three, the data used in the study are described. In section four, the empirical results are presented. Finally the results are summarized in section five, along with the main conclusions.

4.2 METHODOLOGY

The basic assumption underlying the APT is the existence of a Linear Factor Model (LFM) for returns, $R_i(t)$,

$$R_i(t) = E_t[R_i(t)] + \sum_{j=1}^{K-1} b_{ij}F_j(t) + b_{iK}F_K(t) + \varepsilon_i(t) \qquad (4.2.1)$$

where E_t is the rational expectation of asset i, conditional on information available at time t; b_{ij} is the sensitivity of asset i to factor j; $F_j(t)$ for $j=1,\ldots,K-1$ is the innovation in the j'th observable macroeconomic factor in period t; $F_K(t)$ is an unobservable factor capturing the effect of any relevant omitted variables [as in Burmeister and McElroy (1988)]; all the factors are assumed to have zero mean, and to be serially uncorrelated and uncorrelated with each other; and $\varepsilon_i(t)$ is the idiosyncratic risk of asset i at time t. $E_t[\varepsilon_i(t)] = 0$ and $E_t[\varepsilon_i(t)\varepsilon_j(t)] = \sigma_i^2$ for $i = j$, and is zero otherwise.

The market index (or any portfolio) $R_m(t)$ can also be represented by the same linear factor model:

$$R_m(t) = E_t[R_m(t)] + \sum_{j=1}^{K-1} b_{mj}F_j(t) + b_{mK}F_K(t) + \varepsilon_m(t)$$

where $\varepsilon_m(t) = \sum_{i=1}^{N} w_i\varepsilon_i(t)$, and the w_i are the proportional portfolio weights which sum to one. The variance of $\varepsilon_m(t) = 0$, and $\varepsilon_m(t)$ (which is the unsystematic risk of the market portfolio) approaches zero for well-diversified portfolios such as the market portfolio. With no economic consequences, normalizing $b_{mK} = 1$ yields:

$$R_m(t) = E_t[R_m(t)] + \sum_{j=1}^{K-1} b_{mj}F_j(t) + F_K(t) \tag{4.2.2}$$

In practice, since $R_m(t)$ can be approximated by a broad market index, the residual market factor (RMF), $F_K(t)$, can be estimated as the residual from the regression of the broad market index on a constant and the observable macroeconomic factors.

By assuming that the LFM holds true, that no arbitrage profits exist and that certain other regularity conditions are satisfied, the arbitrage pricing model (APT) for the n assets becomes:

$$E_t[R_i(t)] = \lambda_o(t) + \sum_{j=1}^{J-1} b_{ij}\lambda_j + b_{iK}\lambda_K \tag{4.2.3}$$

where $\lambda_o(t)$ is the return on a risk-free asset for period t, which is supposed to be known at the beginning of the period. Given that b_{ij} can be interpreted as the risk associated with asset i due to the systematic influence from factor j, λ_j can be interpreted as the risk premium associated with the pervasive influence of factor j on all the assets.

Substituting the system of equations given by (4.2.3) into the system of equations given by (4.2.1), the following non-linear system of equations is obtained:

$$R_i(t) - \lambda_o(t) = \sum_{j=1}^{K-1} b_{ij}\lambda_j + \sum_{j=1}^{K-1} b_{ij}F_j(t)$$
$$+ b_{iK}\lambda_K + b_{iK}F_K(t) + \varepsilon_i(t) \tag{4.2.4}$$

McElroy and Burmeister (1988) first estimate (4.2.2) to get the fitted values of $F_K(t)$, and then estimate (4.2.4) using the method of non-linear seemingly unrelated regressions (NSUR). Although equation (4.2.4) is interpreted as the constant risk premium APT pricing equation in this thesis, Chamberlain (1988) shows that it can also be interpreted as Merton's (1973) Intertemporal CAPM (ICAPM), if the market portfolio is well-diversified. In Chamberlain's framework, the two pricing models are not testably distinct.

Theoretical work by Merton (1973), Cox, Ingersoll and Ross (1985) and Breeden (1986) relate the risk premium of factor j to its volatility and a proportionality factor, which is interpreted as a measure of risk aversion [Roll and Ross (1980), and Merton (1980)]. Empirical studies by Merton (1980), French, Schwert and Stambaugh (1987) and Campbell (1987) support these theoretical results because they find that the expected risk premium on the stock market is positively correlated with the predictable volatility of stock returns. Lauterbach (1989) documents a relation between the expected returns on U.S. Treasury bills and the conditional volatilities of consumption, the spot interest rate, and industrial production. Therefore, the following relationship is postulated to complete the asset pricing model:

$$\lambda_j(t) = a_j + R_j\sigma_j^2(t) \tag{4.2.5}$$

where R_j is a proportionality coefficient; $\sigma_j^2(t)$ is the conditional volatility of factor j; and a_j is a parameter. If the conditional volatilities of the macrofactors change over time, then the risk premia will be time-varying. In Merton (1980), equation (4.2.5) is referred to as 'Model #1'. Since Merton examines the single factor model, $j = m$ (i.e., the market portfolio). Substituting equation (4.2.5) into the system (4.2.4) yields the following time-varying risk premia asset pricing model (APT):

$$R_i(t) - \lambda_o(t) = c_i + \sum_{j=1}^{K-1} R_j\, b_{ij}\sigma_j^2(t) + \sum_{j=1}^{K-1} b_{ij}F_j(t) + b_{iK}R_K\sigma_K^2(t) + b_{iK}F_K(t) + \varepsilon_i(t) \qquad (4.2.6)$$

where $c_i = \sum_{j=1}^{K} b_{ij}a_j$.

Since the innovations of the macroeconomic factors are only observed monthly, their conditional variances are estimated using the method developed by Davidian and Carroll (1987). Specifically, the following autoregression is performed for each factor j:

$$\sigma_j^2(t) = \sum_{j=1}^{12} \alpha_j D_j(t) + \sum_{i=1}^{12} \theta_i\, \sigma_j^2(t-i) + V_j(t) \qquad (4.2.7)$$

where $\sigma_j^2(t) = (\pi/2)\ F_j(t)^2$ is the unconditional variance of factor j at time t, since $E[F_j(t)] = 0$. Since the squared innovations, $F_j(t)^2$, are single point variance estimates, they are adjusted by the term $(\pi/2)$ [see Schwert (1989) and Schwert and Seguin (1990)]. The fitted variances $\hat{\sigma}_j^2(t)$ are estimates of the conditional variances of factor j.

This specification is similar to the ARCH specification for modelling time-varying volatilities, since the monthly dummy variable $D_j(t)$ allows for different monthly standard deviations. Schwert (1989) and Schwert and Seguin (1990) use this specification to estimate the standard deviations of monthly returns conditional on information up to the present time.

To construct the perfect foresight or predicted returns from the constant and time-varying risk-premia models, the system of equations given by (4.2.4) and (4.2.7) are estimated in turn using a subset of the available observations, T, say $t=1,\ldots,T_1$. This yields:

$$(R_i - \lambda_o) = \sum_{j=1}^{K} \hat{b}_{ij}(F_j + \hat{\lambda}_j) + \hat{\varepsilon}_i, \qquad i=1,\ldots,n \tag{4.2.8}$$

for the constant risk premia APT model, and

$$(R_i - \lambda_o) = \hat{c}_i + \sum_{j=1}^{K} \hat{b}_{ij}(F_j + \hat{R}_j\hat{\sigma}_j^2), \qquad i=1,\ldots,n \tag{4.2.9}$$

for the time-varying risk-premia APT model. These arbitrage asset pricing equations are two of the possible models available to determine the required rates of return. Therefore, these models of return determination are used to find the predicted returns R_C^* and R_{TV}^* from the constant and time-varying risk premia APT models, respectively. The variance of these ex post rational (and theoretically predicted) returns are compared to the variance of the actual returns.

Specifically, after each model is estimated using the first T_1 observations of the returns and macroeconomic factors, an ex post forecast of the returns for the periods T_1+1, T_1+2, ..., T_1+12 are obtained, since the macroeconomic factors are known with certainty for these periods.[20] To obtain the predicted returns for the twelve-month period after T_1+12, the models are re-estimated with T_1 (less the first twelve observations) to T_1+12 observations. Dropping the first twelve observations ensures that the models are always estimated with the same number of observations. A new updated set of parameters is obtained and used to construct the forecasted returns. This updating procedure continues until all the out-of-sample observations of the macroeconomic factors are exhausted. The end result is a vector of predicted returns for the period T_1 to T. The variances of these two series are the measures of variability of the predicted returns obtained from the two models. Then, these variances are compared to the variance of the actual returns for the same period. If the APT models capture the observed stock return variability successfully, the ratio of the variance of the predicted returns to the variance of the actual returns for each model should be equal to one for the population.[21]

4.3 DATA

All the stocks traded on the Toronto Stock Exchange (TSE) from January 1961 through December 1987 are used herein. The stock return data are obtained from the TSE/Western Monthly Data Base. Size-ranked portfolios are used because most studies which test asset pricing models use these portfolios to obtain a maximum dispersion in returns: small capitalization (cap) firms generally have higher returns (and risks) than large cap firms.

The construction of the size-ranked portfolios requires that stocks first be ranked monthly according to their annual December-end outstanding market values, and then cutoff points equal to the desired number of portfolios be determined. The returns on stocks contained within each of the cutoff points are used to calculate the average equally-weighted portfolio returns. In this study, fifty size-ranked portfolios are used.

All of the macroeconomic variables whose innovations could enter the LFM are taken from Statistics Canada's CANSIM Mini Base. All the series are converted to real values, and are seasonally adjusted. The first differences in the logarithms (growth rates) of most of the macroeconomic variables are used.

The choice of the macroeconomic variables is dictated by several factors. Since no generally accepted theory exists for linking stock returns to the economy, none is used to derive unique and universally acceptable macrovariables. As a result, general economic theory and intuition are the main inputs used in the selection process. Macroeconomic factors that influence stock returns in past studies and data availability are important inputs affecting the selection decision. Whether or not the macroeconomic variables appear in the popular financial media is also an important consideration in the final stage of the selection process.

Variables designed to capture the real sector of the economy include the Canadian composite index of leading indicators (CINDEX) and industrial production (INDUS). The variable designed to capture the monetary and financial sector is the money supply (M).[22] Since Canada's economy is highly related to the performance of the U.S. economy, variables designed to capture the influence of the foreign sector on stock returns include the Canada/U.S. exchange rate (EX), total exports (EXPORTS), and the U.S. composite index of leading indicators (USINDEX). Due to the lag of aggregate economic information of almost one month, a misalignment may exist among the stock returns and the innovations of the macroeconomic variables. Therefore, the lags, leads and current growth rates of the macroeconomic series are tried initially in the (unreported) regressions. Our findings, unlike Chen et al. (1985) who use future growth rates of industrial production and Chan et al. (1985) who use future growth rates of both industrial production and net business formation, support the use of lagged growth rates for industrial production (LINDUS).

As defined in Chen et al. (1985), Chan et al. (1985) and McElroy and Burmeister (1988), the risk premium is given by:

$$PREM(t) = CBOND(t) - LBOND(t)$$

where CBOND is the average yield for ten industrial bonds that constitute the McLeod, Young and Weir bond index, and LBOND is the yield on Government of Canada long-term bonds with maturities of ten years and over. The shape of the term structure is defined as:

$$TERM(t) = LBOND(t) - TBILL(t-1)$$

where TBILL is the average monthly yield on Government of Canada 91-day Treasury Bills. Although these two variables are not mean zero (since the term structure is usually upward sloping and the risk premium for holding the more risky corporate bonds instead of the less risky government bonds is always positive), most of the previous U.S. studies use these variables directly as innovations. In this study, the innovations of these variables are used to ensure that they are both zero-mean and serially uncorrelated, as innovations should be.

The innovations of all the macroeconomic variables are equal to the forecast errors that are obtained by fitting each of the macroeconomic series using Akaike's (1976) state-space procedure. Brown and Otsuki (1989) and Kryzanowski and Zhang (1992) use this procedure to obtain white noise innovations.

To ensure that each macroeconomic innovation does not contain information embodied in any of the other innovations a series of sequential regressions is performed. The second innovation is regressed against the first innovation and the residual from this regression is used as the second innovation. Then, the third innovation is the residual of a regression of itself on the first and second innovations. The same procedure is followed for the rest of the macroeconomic innovations to get orthogonalized innovations, by construction.

4.4 EMPIRICAL RESULTS

4.4.1 Linear factor model estimation

The linear factor models (LFM) for the return on the Toronto Stock Exchange (GTSE300) and the returns on the fifty portfolios are reported in Table 4.1. The returns on the TSE 300 market index are significantly affected by the innovations of the U.S. composite index (USINDEX), the exchange rate (EX), the Canadian leading index (CINDEX), the lag of the industrial production index (LINDUS), Canadian exports (EXPORTS) and the term structure (TERM).[23] The p-values indicate that these six variables are significant at the .01 significance level. Also the F-values indicate that the overall regressions are significant at the .01 significance level. The Residual Market Factor (RMF) is the residual from the LFM for GTSE300. Based on the mean t-values, the same macroeconomic variables (with the exception of EX which is significant at the 0.06 level) and the RMF are significant for the fifty portfolios. These factors as a group explain approximately twenty-four percent of the variation in the returns on the TSE 300 and the fifty portfolios. These results are generally consistent with the findings of Chen et al. (1986), McElroy and Burmeister (1988) and Kryzanowski and Zhang (1992).

4.4.2 Construction of the predicted returns

To construct the predicted returns from the constant risk premia model described by the system of equations (4.2.4), a series of nonlinear seemingly unrelated regressions (NSUR) are performed. The dependent variables are the excess (over the risk-free rate) returns on the fifty portfolios, and the dependent variables are the innovations of the U.S. composite index (USINDEX), the exchange rate (EX), the Canadian leading index (CINDEX), the lag of the industrial production index (LINDUS), Canadian exports (EXPORTS) and the term structure (TERM). For the time-varying risk premia model, the dependent variables are the excess portfolio returns, and the independent variables are those for the constant risk premium model plus their conditional variances. Each conditional variance is estimated using equation (4.2.7).

Each model is estimated initially for the period from March 1962 to March 1967 using the NSUR method. These estimated pricing equations (4.2.8) and (4.2.9) are used to derive the predicted (or perfect foresight) returns for the period from April 1967 to March 1968, since the macroeconomic innovations are known with certainty. Then, the first twelve observations are dropped from the sample, and the models are re-estimated from March 1963 to March 1968. That is, the new information set is updated to include the April 1967 to March 1968 information about the macroeconomic factors. This yields the new updated pricing equations (4.2.8) and (4.2.9) that are used now to construct the forecasted returns for the period from April 1968 to March 1969. The procedure is repeated until all the out-of-sample macroeconomic observations are used. Specifically, March 1982 to March 1987 is the last set of data for estimating equations (4.2.8) and (4.2.9), and the last set of forecasted returns are from April 1987 to December 1987. Combining all the forecasts obtained from the recursive estimation technique described above yields the predicted returns $R^*_{Ci}(t)$ and $R^*_{TVi}(t)$ for the constant and time-varying risk premia APT models, respectively, over the period from April 1967 to December 1987 for $i = 1,\ldots,50$ portfolios.[24]

Estimates for the risk premia for the two models for different subperiods are reported in Table 4.2.[25] As shown in panel A where subperiod estimates for the constant risk premium are reported, strong time variation of the risk premia exists across the different subperiods. The significance of the time-varying coefficients, R_j, (j = LINDUS, CINDEX, TERM, USINDEX, EXPORTS, EX, and RMF) reported in panel B indicates the existence of significant time-variation in the risk premia (to the point where the signs of some premia change from one subperiod to another) even within each subperiod.[26] These findings are consistent with the findings by haugen et al. (1991) that investors revise risk premia frequently and significantly. For the subperiod from March 1977 to March 1982, the constant risk premia APT model estimates for the risk premia are all insignificant (panel A), while those for the time-varying risk premia APT model are all significant (panel B). This indicates that, if strong variation in the risk premia exists and the model is estimated as if the risk premia are constant, then there is a high probability that the estimates of the risk premia will be insignificant.

4.4.3 Variance ratios

The monthly forecasted returns, R_i^* i=1,...,50, estimated from the constant risk premia APT model for the fifty portfolios are used to estimate the sample variance for each portfolio. These sample variances for the period from April 1967 to December 1987 are reported in Table 4.3. The actual portfolio returns R_i are used to estimate the actual variance for each of the fifty portfolios for the same period. The variance ratio, $Var(R^*)/Var(R)$, indicates that the constant APT model explains anywhere from 29.38% (for portfolio 41) to 73.01% (for portfolio 16) of the variance of the observed stock returns. The mean variance ratio, $Var(R^*)/Var(R)$, is equal to .4952. This indicates that at least half of the monthly portfolio return variability can be explained by the constant risk premia model. Based on Table 4.4, the variance ratio, $Var(R^*)/Var(R)$, indicates that the time-varying APT model explains anywhere from 30.71% (for portfolio 41) to 90.14% (for portfolio 2) of the variance of the observed stock returns. There is also an improvement in the mean variance ratio, $Var(R^*)/Var(R)$, from .4952 to .5118.[27] The 50 portfolios are pooled into one aggregate portfolio so that the results are more comparable to those of Fama (1990), who finds that 58% of the variance of the annual returns on the NYSE value-weighted index is explained. As discussed in Shiller (1989), our aggregate portfolio accounts for the covariances between portfolios. In turn, this may reduce the error in predicting aggregate returns, and thereby increase the probability of explaining the variability of aggregate stock returns.

To determine how much of the observed variability of the returns on the aggregate portfolio can be explained by the two APT models, the predicted returns for the individual portfolios are combined to form the predicted equally-weighted portfolio returns, RP^*. The actual returns are also combined to form the actual equally-weighted portfolio returns, RP. Then, the variance of the predicted and actual returns on the equally-weighted portfolio, $Var(RP^*)$ and $Var(RP)$, respectively, are estimated, and the variance ratio $Var(RP^*)/Var(RP)$ is constructed.

The sample $Var(RP^*)$ and $Var(RP)$ estimated annually, over five year subperiods, and over the entire time period, which are derived from the constant risk premia APT model, are reported in Table 4.5. When the variance ratio $Var(RP^*)/Var(RP)$ is computed annually, there are periods where the ratio is greater than one (i.e., above the maximum population value of the variance ratio of one). Most of these occurrences are during the 1976 to 1980 period. Since the APT model probably did not capture a structural change (such as the second oil shock or a recession) in the economic system during this period, the sample variance ratio is greater than one. The variance ratios computed for the five-year periods confirm what is found with the annual variance ratios for the 1976 to 1980 period. For the entire period from April 1967 to December 1987, the constant APT model explains 79.99% of the observed variability of the equally-weighted portfolio. When the variance ratio is computed for the period from April 1967 to March 1987 (i.e., October 1987 is omitted from the sample), the constant APT model explains almost 83% of the observed variability in the returns of the equally-weighted portfolio.

The two sample variances, $Var(RP^*)$ and $Var(RP)$, are reported in Table 4.6. These are derived from the time-varying risk premia APT model and are estimated annually, over five year subperiods, and over the entire period. The variance ratios $Var(R^*)/Var(R)$ for this model are similar to those of the constant risk premia model. Specifically, the five year variance estimates yield exactly the same inferences as the corresponding constant risk premia model. However, as expected over the entire time period from April 1967 to December 1987, the time-varying risk premia model explains a higher percentage (83.49%) of the observed variation in returns on the equally-weighted portfolio. Similarly, when the pre-October and October 1987 period is excluded, the explained variability for the conditional risk premia APT model rises to 86.78%. Based on the variance ratio values, the time-varying (conditional) risk premia model outperforms the constant (unconditional) risk premia model in explaining stock return variability.

The two versions of the APT model evidently explain significantly more of the observed stock market variability than has been previously reported in the literature.

4.5 CONCLUDING REMARKS

This chapter attempts to explain the observed variability of stock returns using two possible methods to model the risk premia of the APT model with pre-specified macroeconomic factors. In the first method, the risk premia are assumed to be time-invariant over a given period. This gives rise to the constant (unconditional) risk premia APT model. In the second method, the risk premia are assumed to be time-varying. This gives rise to the time-varying (conditional) risk premia APT model. These two APT model specifications are then used to derive predicted returns, and predicted returns are then compared with actual future returns, in order to determine how much of the observed variability in stock returns is explained by the variability in the macroeconomic factors.

For the period from April 1967 to December 1987, when the constant risk premia (time-varying risk premia) model is used to derive the predicted returns on the equally-weighted portfolio, about 80% (83.5%) of the observed return variability of the equally-weighted portfolio is explained. Since more than 83% of the observed variance in the returns on an equally-weighted portfolio using the time-varying risk premia model is explained by the variability of macroeconomic factors, this is clearly welcome news for those that believe in the conditional APT with identified macroeconomic factors and that markets are indeed efficient.

CHAPTER FIVE: CONCLUSION

This thesis has investigated the small-firm effect and January seasonality, the variability of stock returns and the integration of the Canadian equity market with a global North American equity market. This was done using modified forms of the Ross (1976), Connor (1984), and Solnik (1983) APT models. The major findings of this thesis can be summarized as follows:

(1) The APT with time-varying risk premia is able to explain from 51.5 percent to 58.8 percent of the observed January seasonality. No small-firm effect is detected in the risk-adjusted returns. The empirical results show that the conditional variances of six macrofactors have time-varying risk premia. These are the lag of industrial production, the Canadian index of 10 leading indicators, the U.S. composite index of 12 leading indicators, exports, the exchange rate and the residual market factor.

(2) For the period from March 1969 through March 1988, the Canadian equity market is only partly integrated (or equivalently, partly segmented) with the American equity market. The Canadian macroeconomic factors found to influence Canadian stock returns are the pure domestic component of the term structure and the lagged industrial production index. The North American macroeconomic factors affecting Canadian returns are the purely "international" components of the differential in the Canada/U.S. leading indicators, and the interest rate on U.S. dollar deposits in London (i.e., eurodeposits).

(3) For the period from April 1967 to December 1987, when the constant

risk premia (time-varying risk premia) APT model is used to derive predicted returns on the equally-weighted portfolio, about 80% (83.5%) of the observed return variability of the equally-weighted portfolio is explained. Since more than 83% of the observed variance in the returns on an equally-weighted portfolio using the time-varying risk premia model is explained by the variability of macroeconomic factors, this is welcome news for those that believe in the conditional APT with identifiable factors and that markets are indeed efficient.

Several directions for future research emerge from this thesis. First, the research methodology from all three essays could be applied to American stock data. This would provide some tests for the robustness of the results reported in this thesis. Another interesting avenue of research is to apply the methodologies presented to other asset markets such as the bond market. In all three essays the conditional variances are derived from a linear autregressive model of order 12. However, Cao and Tsay (1992) show that a non-linear variance specification may be more appropriate, since the volatilities series are non-linear. Extensions of all three essays could examine the effects of using different variance specifications to derive the conditional volatilities.

FOOTNOTES

1. The tax-loss-selling hypothesis attempts to expain the January effect by asserting that investors temporarily drive security prices below their equilibrium levels at year-end to realize capital losses for tax purposes. Evidence supporting this hypothesis includes Branch (1977), Dyl (1977) and Roll (1983). If this hypothesis is true then the risk premium estimated from the tradi'ional asset pricing models will not be able to explain the January seasonal. However, the presence of a January seasonal in the absence of tax-loss selling pressure in various international markets [Brown et al. (1983), Berges, McConnell and Schlarbaum (1984) and Tinic and Barone-Adesi (1988)] weaken the explanatory power of this hypothesis, and alternative expanations such as the time-variation of risk premiums have to be examined.

2. Similarly, Chang and Pinegar (1990) find that the factor risk premia and factor betas in the Chen, Roll and Ross (1986) model exhibit strong seasonal nonstationarity in January versus non-January months.

3. This is an extension of the papers by Chang and Pinegar (1989) who test if the first differences of the macrofactors are seasonal, and Kryzanowski and Zhang (1992) who test if the innovations of the macrofactors are seasonal. Neither study examines whether the conditional standard deviations of the innovations of the macrofactors are seasonal.

4. If one uses mimicking portfolios instead of identified macroeconomic factors to derive the RMF, then some ambiguity regarding the interpretation of the RMF is introduced. With identified factors the

interpretation of the RMF as the orthogonal component of the market index on the macrofactors is clearly understood. However, when a set of mimicking portfolios is used to get the RMF, it is not at all clear that this will have the same interpretation since one of the mimicking portfolios may actually be the market index. Then, it may be redundant to re-estimate the RMF. In addition, if one attempts to estimate the RMF from a regression of the market index on the mimicking portfolios, then it is possible that one may be estimating a 'residual macroeconomic factor' rather than the RMF.

5. The system of equations described by (2.2.9) is a special form of the equilibrium conditions derived from a portfolio optimization problem where individuals choose a consumption withdrawal plan and an optimal portfolio in order to maximize the discounted expected value of the utility of future consumption (which is a function of wealth and the current state of nature). In particular, if state dependencies are ignored, R will be independent of j (as shown in Roll and Ross, 1980). Otherwise, R will depend on j. In Merton (1980), this restricted form of equation (2.2.9) is referred to as 'Model #1', and R is interpreted as a measure of relative risk aversion. However, since Merton is examining the single factor model, j = m (i.e., the market portfolio).

6. It should be noted that in equation (2.2.13) the unconditional variances are the dependent variables. Since we are interested in testing if the conditional variances have monthly seasonality (since the conditional variances enter the asset pricing models described previously), equation (2.2.14) with the conditional variances as dependent variables has to be estimated.

7. Due to the lag of aggregate economic information of almost one month, there may be misalignment among the stock returns and the macroeconomic variables. Therefore, the lags, leads and current growth rates of the macroeconomic series were initially tried in the unreported regressions. Our findings, unlike Chen et al. (1985) who use the future growth rates of industrial production and Chan et al. (1985) who use the future growth rates of both industrial production and net business formation, support the use of the lagged growth rates for industrial production (LINDUS).

8. Chen, Roll and Ross (1985) and Chan, Chen and Hsieh (1985) use returns in defining PREM and TERM. As in McElroy and Burmeister (1988), innovations of yields are used herein.

9. The linear factor model (LFM) with a January dummy included for both the twenty and fifty portfolios was also estimated to verify the existence of January seasonality in our sample over the period from March 1963 to March 1988. The mean January dummy coefficients (mean t-statistics) for the LFM with the fifty and twenty portfolios are .0270 (2.13) and .0268 (2.87), respectively. More specifically, in the fifty and twenty portfolio LFMs, thirty-three and seventeen portfolios exhibited significant January seasonality at the .10 significance level, respectively. Clearly, there is significant January seasonality present in our stock return data set that cannot be explained by the LFM alone.

10. Chan et al. (1985) use this methodology to test for a size effect for firms listed on the NYSE.

11. The IAPT is more general because it allows for more factors to influence returns, and it is less restrictive since it is based on an

arbitrage argument and not investor preferences.

12. Testing whether the U.S. equity market is integrated or segmented relative to a North American market is an interesting topic which is left for further research. One possibility is that, while the Canadian equity market may be integrated relative to a North American market, the U.S. equity market may not be. This would be the case if the barriers that Canadian investors face within a North American market are not binding, while the barriers that U.S. investors face within the same North American market are binding.

13. As in any two-step procedure, orthogonalization in the first step results in an errors-in-variables problem in the implicit second-pass regression estimation. As a result, the reported t-statistics are biased towards significance [Bodurtha (1986)].

14. Other studies that include international factors in the APT include Bodurtha et al. (1989) for U.S. stocks and Hamao (1989) for Japanese stocks.

15. By estimating models (3.3.6) and (3.3.7) using NOLS and NLSUR, both the factor betas and prices of risk are jointly estimated. This avoids the classic two-step Fama-MacBeth procedure where the betas are first estimated and then the prices of risk are estimated. As is well known, the two-step procedure introduces an error-in-variables problem into the estimation.

16. To determine whether any of the models had some unknown form of heteroscedasticity that could not be detected by the B-P-G and ARCH tests, White heteroscedasticity-consistent standard errors are also computed. Since the corrected standard errors are virtually identical

to the unadjusted ones, they are not reported herein. This also indicates the absence of significant heteroscedasticity.

17. For a discussion of the relative merits of these two methods, see Chapter 2, page 16.

18. This assumption about risk-neutral economic agents implies constant expected returns. This is contrary to the findings by many researchers [see Fama (1991) for an extensive list of this literature]. For this reason, Cochrane (1991) and Fama (1991) do not view the volatility tests as being informative about market efficiency. Instead, they view the volatility tests as evidence that expected returns vary through time.

19. In this type of framework where an asset pricing model is specified the hypothesis is unavoidably a joint hypothesis that the asset pricing model is correct, and that asset markets are efficient.

20. The model is estimated using monthly data over a certain period, T_1, and then used to obtain the ex post forecasted returns for the next twelve months.

21. Ferson and Harvey (1991) use the same method of rolling regressions to compute a similar variance ratio to determine how much of the variation of the expected returns is captured by the asset-pricing model. In this study we examine how much of the variation of actual future returns is captured by return forecasts of the asset-pricing models.

22. Since 30-day Treasury bills are not issued in Canada, the risk-free rate is calculated as $\log(1+R_1)$ where $R_1=[(1+91/365)*\mathrm{lag}(\mathrm{TBILL90}/100)]^{30.4/91}-1$, and TBILL90 is the interest

on 90-day Treasury Bills from Korkie (1990).

23. The estimated intercept is significant for the LFMs of the returns on the TSE 300 market index and the returns on the fifty portfolios indicating positive expected returns [i.e., the intercept in equation (4.2.1) is positive].

24. This recursive estimation procedure is extremely computer intensive. It takes approximately 24-hours of CPU time to derive the ex post predicted returns for each model.

25. The models were initially estimated with an intercept term and all the macroeconomic variables for the entire period. For both models, the estimated intercept was not significantly different from zero, indicating that the APT pricing equation given by (4.2.3) captures all the positive expected return found when the LFM was estimated. In subsequent estimations, the intercept was suppressed to reduce the number of estimated coefficients.

26. Based on Table 4.2, the residual market factor (RMF), generally, has the largest risk premium. This confirms the finding by Ferson and Harvey (1991).

27. The reader is reminded that the constant APT model has constant risk premia for a given period but time-varying risk premia across periods. The time-varying risk premia model allows for time-varying risk premia both within a given period and across periods. Since both models allow for time-varying risk premia every 5-years, only a minor improvement was expected for the time-varying risk premia model over the constant risk premia model.

TABLE 2.1

Description of all macroeconomic and stock market variables.

Variable	Description
CINDEX	Index of 10 leading indicators
EX	Exchange rate (Cdn/US)
EXPORTS	Total exports
EW	Equally-weighted index of all the stocks on the TSE
INDUS	Industrial production
MONEY	Money supply (M1)
PREM	Unexpected change in the risk premium
RMF	The residuals from the market portfolio APT equation
TERM	Unexpected change in the term structure
TSE300	The Toronto Stock Exchange 300 stock index
USINDEX	U.S. composite index of 12 leading indicators
VW	Value-weighted index of all the stocks listed on the TSE
$\lambda_0(t)$	The one-month risk-free rate[a]

[a]Since in Canada there are no 30-day Treasury bills issued, the risk-free rate
is calculated as $\log(1+R_1)$ where
$R_1=[(1+91/365)*\text{lag}(\text{TBILL90}/100)]^{30.4/91}-1$
and TBILL90 is the interest on 90-day Treasury Bills from Korkie (1990).

TABLE 2.2

Summary statistics for the innovations are presented herein. Period covered is from March 1962 to March 1988.

Variable	Mean	Std.Dev	t:Mean=0[a]	D:normal[b]	Fisher's-k[c]	Q(6)[c]
CINDEX	-.000060	.009241	-.1159	.04687	4.6910	8.68
INDUS	-.000108	.011758	-.1624	.04368	9.2631	8.59
USINDEX	-.000122	.008131	-.2661	.04947	4.2093	4.89
EXPORTS	-.000317	.054147	-.1037	.03671	5.2704	7.67
EX	-.000155	.007566	-.3631	.09352	4.1303	.79
MONEY	.000111	.012156	.1622	.05908	6.8630	3.79
TERM	-.007408	.656494	-.1996	.11150	6.7701	10.92
PREM	.002396	.167375	.2536	.10892	7.9150	45.12

[a]The null hypothesis of zero mean cannot be rejected for all variables at the .10 significance level.

[b]The null hypothesis of normality cannot be rejected for all but the TERM and PREM variables at the .10 significance level.

[c]The null hypothesis of white noise cannot be rejected for all the variables except for INDUS and PREM at the .10 level (using the Fisher-k statistic). The Ljung and Box Q-statistics indicate that the null is only rejected for the PREM variable.

TABLE 2.3

OLS coefficient estimates for equation (2.2.5) are presented below. Period covered is from March 1963 to March 1988.

Variable	TSE300	EW	VW
CONSTANT	.0063	.0156	.0116
	(2.56)[a]	(7.05)	(5.09)
	[.0054][b]	[.0003]	[.0001]
USINDEX	1.7353	1.7587	1.5950
	(5.80)	(6.49)	(5.72)
	[.0004]	[.0003]	[.0004]
EX	-.8672	-.6312	-.7615
	(-2.68)	(-2.16)	(-2.53)
	[.0038]	[.0158]	[.0059]
CINDEX	1.6502	1.5296	1.4192
	(6.07)	(6.21)	(5.61)
	[.0003]	[.0003]	[.0004]
LINDUS	.5510	.7051	.5455
	(2.65)	(3.75)	(2.82)
	[.0042]	[.0001]	[.0026]
EXPORTS	-.1632	-.1892	-.1531
	(-3.46)	(-4.42)	(-3.47)
	[.0003]	[.0001]	[.0003]
TERM	-.2291	-.3186	-.2637
	(-2.81)	(-4.32)	(-3.47)
	[.0026]	[.0001]	[.0003]
F-value	15.739	20.345	15.473
R^2	.2358	.2852	.2328
D.W.	2.1367	1.9504	2.2132

[a]The t-statistics are given in the parentheses.

[b]The probability values are given in the brackets.

TABLE 2.4

Summary results for the regressions of the conditional standard deviations of the innovations on the monthly dummy variables are reported herein. The prefix 'C' indicates that the variable is the conditional variance of the innovation for this variable. Thus, CRMF is the conditional variance of the residual market factor (RMF). Period covered is from March 1962 to March 1988.

Variable	January	February	March	April	May	June	July	August	September	October	November	December	F-value	R^2
CINDUS	.0164	-.0022	-.0047	.0189	.0015	-.0030	.0013	.0286	.0171	.0043	.0013	-.0008	51.02	.6502
	(11.18)[a]	(-1.05)	(-2.27)	(9.12)	(0.72)	(-1.46)	(0.65)	(13.76)	(8.23)	(2.07)	(0.61)	(-0.37)		
	[.0001][b]	[.2940]	[.0234]	[.0001]	[.4739]	[.1449]	[.5164]	[.0001]	[.0001]	[.0388]	[.5425]	[.7088]		
CUSINDEX	.0146	-.0064	-.0062	.0059	-.0067	-.0043	-.0045	-.0032	-.0045	-.0074	-.0085	-.0039	12.47	.3171
	(13.38)	(-4.19)	(-4.09)	(3.84)	(-4.35)	(-2.89)	(-2.08)	(-2.94)	(-2.95)	(-4.85)	(-5.55)	(-2.59)		
	[.0001]	[.0001]	[.0001]	[.0002]	[.0001]	[.0051]	[.0381]	[.0035]	[.0034]	[.0001]	[.0001]	[.0001]		
CEXPORTS	.5674	-.2156	-.0901	-.2553	-.1239	-.1810	-.1762	.0184	-.1226	.1279	-.0702	-.3007	9.53	.2577
	(14.28)	(-3.87)	(-1.63)	(-4.54)	(-2.20)	(-3.22)	(-3.14)	(0.33)	(-2.18)	(2.27)	(-1.25)	(-5.35)		
	[.0001]	[.0001]	[.1041]	[.0001]	[.0282]	[.0014]	[.0019]	[.7431]	[.0300]	[.0235]	[.2125]	[.0001]		
CEX	.0082	.0034	.0034	-.0019	.0028	.0071	-.0035	.0029	-.0003	-.0034	-.0024	.0008	13.72	.3333
	(9.15)	(2.71)	(2.72)	(-1.52)	(2.26)	(5.68)	(-2.81)	(2.36)	(-0.28)	(-2.71)	(-1.94)	(0.71)		
	[.0001]	[.0071]	[.0069]	[.1303]	[.0241]	[.0001]	[.0053]	[.0189]	[.7774]	[.0071]	[.0537]	[.4776]		
CCINDEX	.0292	-.0209	-.0162	-.0128	-.0165	-.0195	-.0205	-.0144	-.0117	-.0181	-.0241	-.0137	69.55	.7170
	(39.3)	(-20.1)	(-15.6)	(-12.2)	(-15.7)	(-18.6)	(-19.6)	(-13.7)	(-11.2)	(-17.2)	(-23.0)	(-13.1)		
	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]	[.0001]		
CTERM	.1342	-.0068	.0033	.0194	.0167	.0251	.0423	.0329	.0157	.0030	-.0262	-.0022	0.16	.0058
	(2.85)	(-0.10)	(0.05)	(0.29)	(0.25)	(0.37)	(0.63)	(0.49)	(0.23)	(.045)	(-0.39)	(-0.03)		
	[.0047]	[.9179]	[.9598]	[.7703]	[.8024]	[.7065]	[.5261]	[.6213]	[.8141]	[.9640]	[.6941]	[.9738]		
CRMF	.0039	-.0024	-.0013	-.0017	-.0018	-.0021	-.0020	-.0004	-.0017	.0029	.0002	-.0020	28.23	.5078
	(13.75)	(-6.07)	(-3.19)	(-4.27)	(-4.38)	(-5.14)	(-5.02)	(-0.96)	(-4.24)	(7.23)	(0.68)	(-4.94)		
	[.0001]	[.0001]	[.0016]	[.0001]	[.0001]	[.0001]	[.0001]	[.3363]	[.0001]	[.0001]	[.4911]	[.0001]		

[a]The t-statistics are given in the parentheses.

[b]The probability values are given in the brackets.

TABLE 2.5

Coefficient estimates for the unrestricted pricing model given by equation (2.2.11) are presented below. Period covered is from March 1962 to March 1988.

$\hat{c}$	$\hat{b}_{LINDUS}$	$\hat{b}_{CINDEX}$	$\hat{b}_{TERM}$	$\hat{b}_{USINDEX}$	$\hat{b}_{EXPORTS}$	$\hat{b}_{EX}$	$\hat{b}_{RMF}$
Panel A: Estimates of the mean factor loadings with 50 portfolios							
NOLS							
-.0055	.5574	1.3911	-.2200	1.7059	-.1741	-.5814	.8008
(-0.89)[a]	(1.85)	(4.13)	(-1.73)	(4.04)	(-2.11)	(-1.37)	(9.01)
[.3762][b]	[.0659]	[.0001]	[.0855]	[.0001]	[.0358]	[.1715]	[.0001]
NSUR							
-.0042	.5659	1.4364	-.2005	1.7023	-.1704	-.0929	.8006
(-0.86)	(1.97)	(4.66)	(-1.87)	(4.67)	(-2.40)	(-0.62)	(8.84)
[.3909]	[.0500]	[.0001]	[.0626]	[.0001]	[.0171]	[.5374]	[.0001]
Panel B: Estimates of the mean factor loadings with 20 portfolios							
NOLS							
-.0039	.5488	1.3826	-.2255	1.7253	-.1796	-.5666	.8021
(-0.88)	(3.09)	(6.42)	(-2.66)	(6.28)	(-3.85)	(-1.73)	(12.14)
[.3818]	[.0022]	[.0001]	[.0081]	[.0001]	[.0358]	[.0844]	[.0001]
NSUR							
-.0049	.5557	1.4169	-.2142	1.7442	-.1802	-.5321	.8189
(-0.84)	(2.76)	(6.47)	(-3.02)	(4.53)	(-3.74)	(-1.67)	(9.99)
[.3997]	[.0062]	[.0001]	[.0028]	[.0001]	[.0002]	[.0959]	[.0001]

$\hat{R}_{LINDUS}$	$\hat{R}_{CINDEX}$	$\hat{R}_{TERM}$	$\hat{R}_{USINDEX}$	$\hat{R}_{EXPORTS}$	$\hat{R}_{EX}$	$\hat{R}_{RMF}$
Panel C: Coefficient estimates with 50 portfolios						
NOLS						
-.3677	.8283	.0375	-.1858	.0159	-.5799	1.4317
(-5.06)	(13.74)	(4.40)	(-4.08)	(1.30)	(-3.92)	(3.92)
[.0001]	[.0001]	[.0001]	[.0001]	[.1939]	[.0001]	[.0001]
NSUR						
-.3999	.8500	.0081	-.2746	.0482	-4.5424	1.6726
(-3.41)	(7.25)	(0.65)	(-3.03)	(2.59)	(-2.45)	(2.26)
[.0001]	[.0001]	[.5166]	[.0027]	[.0100]	[.0148]	[.0248]
Panel D: Coefficient estimates with 20 portfolios						
NOLS						
-.3786	.8198	0.0392	-.1702	.0116	-.4196	1.2508
(-4.24)	(11.64)	(3.86)	(-3.17)	(0.81)	(-2.37)	(2.86)
[.0001]	[.0001]	[.0001]	[.0017]	[.4207]	[.0045]	[.0045]
NSUR						
-.4015	.7871	.0105	-.2077	-.0023	-.6132	1.0052
(-2.47)	(5.45)	(0.62)	(-1.87)	(-0.08)	(-2.13)	(1.10)
[.0142]	[.0001]	[.5378]	[.0619]	[.9328]	[.0344]	[.2727]

[a] The mean t-statistics are reported in the parentheses.

[b] The probability values are reported in the brackets.

TABLE 2.6

Coefficient estimates for the restricted pricing model given by equation (2.2.12) are presented below. Period covered is from March 1962 to March 1988.

$\hat{c}$	$\hat{b}_{LINDUS}$	$\hat{b}_{CINDEX}$	$\hat{b}_{TERM}$	$\hat{b}_{USINDEX}$	$\hat{b}_{EXPORTS}$	$\hat{b}_{EX}$	$\hat{b}_{RMF}$	R
Panel A: Estimates of the mean factor loadings and RRA coefficient with 50 portfolios								
NOLS								
.0083	.5702	1.4193	-.2371	1.7255	-.1698	-.7464	.8019	.0364
(2.21)[a]	(1.78)	(4.06)	(-1.88)	(3.69)	(-2.23)	(-1.60)	(9.01)	(5.66)
[.0280][b]	[.0765]	[.0004]	[.0617]	[.0003]	[.0262]	[.1099]	[.0001]	[.0001]
NSUR								
.0072	.5764	1.4154	-.2328	1.7452	-.1744	-.7549	.8008	.0202
(1.91)	(1.88)	(4.32)	(-1.80)	(3.71)	(-2.18)	(-1.53)	(10.12)	(2.07)
[.0576]	[.0609]	[.0001]	[.0734]	[.0003]	[.0304]	[.1262]	[.0001]	[.0397]
Panel B: Estimates of the mean factor loadings and RRA coefficient with 20 portfolios								
NOLS								
.0090	.5652	1.3998	-.2405	1.7429	-.1738	-.7021	.8274	.0384
(3.21)	(2.58)	(5.46)	(-2.67)	(5.12)	(-3.46)	(-2.16)	(13.34)	(4.97)
[.0018]	[.0105]	[.0001]	[.0079]	[.0001]	[.0006]	[.0318]	[.0001]	[.0001]
NSUR								
.0074	.5741	1.3944	-.2323	1.7722	-.1799	-.7142	.8167	.0145
(2.87)	(2.64)	(5.53)	(-2.70)	(5.26)	(-3.55)	(-2.18)	(12.84)	(1.05)
[.0043]	[.0087]	[.0001]	[.0074]	[.0001]	[.0005]	[.0301]	[.0001]	[.2937]

[a]The mean t-statistics are reported in the parentheses.

[b]The probability values are reported in the brackets.

TABLE 2.7

Joint tests of various restrictions imposed on systems (2.2.11) and (2.2.12) are presented below. Period covered is from March 1962 to March 1988. NSUR is used in the estimation.

Restriction	The χ^2 test statistic[a] (degrees of freedom) 20	50
1. H_0: $R = \mathbf{c} = 0$; i.e., risk premia are jointly equal to zero in the APT system (2.2.12)	32 (21)[b] [.0585][c]	82 (51) [.0038]
2. H_0: $R_j = \mathbf{c} = 0$, $\forall$ j; i.e., risk premia are jointly equal to zero in the APT system (2.2.11)	72 (27) [.0001]	168 (57) [.0001]
3. H_0: $R = 0$; i.e., risk premia are jointly time-invariant in the APT system (2.2.12)	1 (1) [.3173]	3 (1) [.0833]
4. H_0: $R_j = 0$, $\forall$j ; i.e., risk premia are jointly time-invariant in the APT system (2.2.11)	39 (7) [.0001]	84 (7) [.0001]

[a]This test-statistic is calculated using the MODEL procedure of the SAS-ETS package, and is analogous to the likelihood ratio test. 20 and 50 refer to 20 and 50 portfolios, respectively.

[b]The degrees of freedom are in the parentheses

[c]The probability values are given in the brackets.

TABLE 2.8

Coefficient estimates for the unrestricted pricing model given by equation (2.2.11) with a January dummy are presented below. Period covered is from March 1962 to March 1988.

$\hat{c}$	$\hat{b}_{LINDUS}$	$\hat{b}_{CINDEX}$	$\hat{b}_{TERM}$	$\hat{b}_{USINDEX}$	$\hat{b}_{EXPORTS}$	$\hat{b}_{EX}$	$\hat{b}_{RMF}$	ϕ
Panel C: Estimates of the mean factor loadings with 50 portfolios and a January dummy								
NOLS								
-.0013	.5444	1.3739	-.2205	1.7042	-.1693	-.5657	.7929	.0146
(-0.42)[a]	(2.03)	(3.82)	(-1.76)	(3.86)	(-1.98)	(-0.95)	(8.85)	(1.22)
[.6774][b]	[.0435]	[.0001]	[.0791]	[.0001]	[.0482]	[.3432]	[.0001]	[.2290]
NSUR								
-.0007	.5529	1.4082	-.2025	1.7058	-.1637	-.1086	.7905	.0185
(-0.72)	(1.73)	(3.81)	(-1.84)	(4.30)	(-2.68)	(-0.43)	(8.82)	(1.43)
[.4737]	[.0840]	[.0002]	[.0662]	[.0001]	[.0077]	[.6669]	[.0001]	[.1442]
Panel D: Estimates of the mean factor loadings with 20 portfolios and a January dummy								
NOLS								
.0002	.5368	1.3609	-.2260	1.7227	-.1754	-.5557	.7745	.0141
(0.27)	(2.43)	(4.96)	(-2.67)	(5.43)	(-3.83)	(-1.64)	(9.69)	(1.47)
[.7845]	[.0155]	[.0001]	[.0081]	[.0001]	[.0002]	[.1013]	[.0001]	[.1356]
NSUR								
.0013	.5384	1.3930	-.2132	1.7421	-.1762	-.5315	.7824	.0166
(0.40)	(2.59)	(5.11)	(-2.51)	(5.97)	(-4.01)	(-1.70)	(11.82)	(1.58)
[.6874]	[.0101]	[.0001]	[.0125]	[.0001]	[.0171]	[.0905]	[.0001]	[.1152]

$\hat{R}_{LINDUS}$	$\hat{R}_{CINDEX}$	$\hat{R}_{TERM}$	$\hat{R}_{USINDEX}$	$\hat{R}_{EXPORTS}$	$\hat{R}_{EX}$	$\hat{R}_{RMF}$
Panel C: Coefficient estimates with 50 portfolios and a January dummy						
NOLS						
-.2643	.5511	.0379	-.1782	.0295	-.6115	1.1493
(-3.64)	(7.53)	(4.47)	(-3.91)	(2.32)	(-4.05)	(3.09)
[.0003]	[.0001]	[.0001]	[.0001]	[.0210]	[.0001]	[.0022]
NSUR						
-.2537	.4864	.0093	-.2509	.0623	-4.0813	1.3282
(-2.29)	(3.61)	(0.75)	(-2.79)	(3.23)	(-2.68)	(1.77)
[.0229]	[.0004]	[.4548]	[.0057]	[.0014]	[.0077]	[.0771]
Panel D: Coefficient estimates with 20 portfolios and a January dummy						
NOLS						
-.2758	.5567	.0395	-.1653	.0241	-.4265	.9654
(-3.08)	(6.46)	(3.90)	(-3.09)	(1.62)	(-2.38)	(2.17)
[.0022]	[.0001]	[.0001]	[.0022]	[.1063]	[.0179]	[.0306]
NSUR						
-.4015	.7871	.0105	-.2077	-.0023	-.6132	1.0052
(-2.47)	(5.45)	(0.62)	(-1.87)	(0.08)	(-2.13)	(1.10)
[.0142]	[.0001]	[.5378]	[.0619]	[.9328]	[.0344]	[.2727]

[a]The mean t-statistics are reported in the parentheses.

[b]The probability values are reported in the brackets.

TABLE 2.9

Coefficient estimates for the restricted pricing model given by equation (2.2.12) with a January dummy are reported below. Period covered is from March 1962 to March 1988.

$\hat{c}$	$\hat{b}_{LINDUS}$	$\hat{b}_{CINDEX}$	$\hat{b}_{TERM}$	$\hat{b}_{USINDEX}$	$\hat{b}_{EXPORTS}$	$\hat{b}_{EX}$	$\hat{b}_{RMF}$	ϕ	R
Panel A: Estimates of the mean factor loadings and RRA coefficient with 50 portfolios									
NOLS									
.0058	.5340	1.3736	-.2363	1.7094	-.1682	-.6375	.7849	.0268	.0328
(1.54)[a]	(1.67)	(3.16)	(-1.84)	(3.91)	(-2.29)	(-1.42)	(8.27)	(2.11)	(5.13)
[.1244][b]	[.0968]	[.0018]	[.0673]	[.0001]	[.0225]	[.1565]	[.0001]	[.0367]	[.0001]
NSUR									
.0051	.5373	1.3752	-.2324	1.7254	-.1702	-.6434	.7842	.0269	.0209
(1.38)	(1.96)	(3.42)	(-1.93)	(3.72)	(-2.79)	(-1.38)	(8.21)	(2.13)	(2.18)
[.1709]	[.0513]	[.0007]	[.0549]	[.0002]	[.0056]	[.1687]	[.0001]	[.0342]	[.0298]
Panel B: Estimates of the mean factor loadings and RRA coefficient with 20 portfolios									
NOLS									
.0066	.5290	1.3549	-.2404	1.7261	-.1719	-.5936	.7947	.0266	.0384
(2.31)	(2.60)	(5.03)	(-2.85)	(5.94)	(-4.18)	(-1.76)	(12.45)	(2.85)	(4.97)
[.0021]	[.0097]	[.0001]	[.0046]	[.0001]	[.0001]	[.0809]	[.0001]	[.0043]	[.0001]
NSUR									
.0051	.5350	1.3575	-.2315	1.7556	-.1752	-.6041	.8163	.0267	.0145
(1.78)	(2.37)	(5.79)	(-3.16)	(6.08)	(-4.05)	(-1.82)	(11.83)	(2.87)	(1.05)
[.0741]	[.0183]	[.0001]	[.0017]	[.0001]	[.0001]	[.0697]	[.0001]	[.0038]	[.2937]

[a] The mean t-statistics are reported in the parentheses.

[b] The probability values are reported in the brackets.

TABLE 2.10

Joint tests of the January effect for systems (2.2.11) and (2.2.12) are reported below. Period covered is from March 1962 to March 1988.

Restriction	The χ^2 statistic[a] (degrees of freedom)		Bartlett's small-sample adjusted χ^2 statistic (degrees of freedom)	
	20	50	20	50
1. H_0: $\Phi = 0$; i.e., The January dummy is jointly equal to zero in the APT system (2.2.11).	28[c] (20)[d]	78 (50)	27.02 (20)	71.53 (50)
2. H_0: $\Phi = 0$; i.e., The January dummy is jointly equal to zero in the APT system (2.2.12).	56 (20)	114 (50)	54.04 (20)	104.54 (50)

[a]This test-statistic is calculated using the MODEL procedure of the SAS-ETS package, and is analogous to the likelihood ratio test.

[b]The adjustment factors are (312-20/2-1)/312=.965 and (312-50/2-1)/312=.917 for the 20 and 50 portfolio models, respectively.

[c]For 20 and 50 degrees of freedom, the $\alpha = .05$ ($\alpha = .01$) critical values for the χ^2 distribution are 31.41 and 67.51 (37.57 and 76.15), respectively.

[d]The degrees of freedom are given in the parentheses.

TABLE 3.1

Description of all macroeconomic and stock market variables.

National[a]	Description
EW	Equally-weighted index of all stocks listed on the Toronto Stock Exchange (TSE)
GDP	Gross domestic product
INDEX	Index of 10 leading indicators
INDUS	Industrial production
INFLAT	Change in the Consumer Price Index (CPI)
LINDUS	The lag of industrial production
MONEY	Money supply (M1)
PREM	Yield on long-term corporate bonds less the yield on long-term government bonds
TERM	Yield on long-term government bonds less the yield on 91-day Treasury bills
TSE300	The Toronto Stock Exchange 300 stock index
International[b]	
DINDUS	Differential in Can/U.S. industrial production
DINDEX	Differential in Can/U.S. leading indices
DINFLAT	Differential in Can/U.S. inflation rates
DLINDUS	The lag of DINDUS
DMONEY	Differential in Can/U.S. money supply
DPREM	Differential in Can/U.S. risk premium
DTERM	Differential in Can/U.S term structure
EURO	Interest rate on eurodollar deposits in London
SP500	Standard and Poor's 500 stock index

[a]The 'pure' domestic component of each of these variables is indicated by the addition of a "D" to the designation for the domestic variable. For example, the 'pure' domestic component of INDEX is denoted by INDEXD.

[b]The 'pure' international component of each of these variables is indicated by the addition of an "I" to the designation for the international variable. For example, the 'pure' international component of DINDEX is denoted by DINDEXI.

TABLE 3.2

Summary statistics for the innovations based on the multivariate state-space procedure.

Variable	Mean	Std.Dev.	t:Mean=0[a]	D:normal[b]	k:Fisher[c]	Q(6)[c]
GDP	-4.73E-05	.005873	-.1218	.03397	6.7242	8.93
INDEX	8.06E-06	.010601	.0115	.03351	7.0368	4.53
INDUS	.000211	.012002	.0267	.05486	6.0326	7.19
INFLAT	-4.75E-05	.003269	-.2205	.04966	6.8520	9.66
MONEY	4.81E-06	.013579	.0054	.06081	6.6306	3.51
PREM	.000014	.002058	.1049	.08874	7.7895	8.38
TERM	.000210	.007439	.4287	.11153	4.4487	6.38
DINDEX	1.93E-06	.008655	.0034	.04771	5.7827	3.62
DINDUS	-4.30E-06	.011832	-.0055	.04580	7.3583	5.82
DINFLAT	.000010	.003663	.0435	.04302	6.0130	4.19
DMONEY	.000259	.036335	.1081	.13206	7.6717	5.72
DPREM	.000012	.001995	.0956	.06970	6.9337	1.62
DTERM	.000325	.009378	.5247	.10231	5.4066	6.95
REURO	-.000933	.003269	-.2205	.04966	5.0303	0.78

[a]The null hypothesis of zero mean cannot be rejected for all variables at the .10 significance level.

[b]The null hypothesis of normality cannot be rejected for all but the TERM, DTERM and DMONEY variables at the .10 significance level.

[c]The null hypothesis of white noise cannot be rejected for all the variables except for PREM, DMONEY, DINDUS and INDEX at the .10 significance level (using the Fisher k statistic). The Ljung and Box Q-statistics indicate that the null is not rejected for any of the variables.

TABLE 3.3

The selected international LFM for the TSE300, VW, EW and fifty portfolios for the total period March 1969 through March 1988.

Variable	TSE300	VW	EW	PR1-PR50
INTERCEPT	.0062	.0100	.0127	.0116
	(2.08)[a]	(3.19)	(4.01)	(2.73)[b]
	[.0190][c]	[.0008]	[.0001]	[.0034]
DLINDUS	.6655	.5836	.5760	.5944
	(2.66)	(2.20)	(2.14)	(1.65)
	[.0042]	[.0143]	[.0165]	[.0502]
DINDEX	1.6750	1.5705	1.6797	1.7636
	(4.89)	(4.33)	(4.57)	(3.24)
	[.0001]	[.0001]	[.0001]	[.0007]
REURO	-1.8571	-1.6018	-1.5394	-1.2869
	(-5.21)	(-4.25)	(-4.03)	(-2.36)
	[.0001]	[.0001]	[.0001]	[.0095]
R^2	.2059	.1566	.1569	.0983
F-value	19.441	13.928	13.959	8.152
D.W.	1.893	1.953	1.871	1.824
Q-TESTS FOR AUTOCORRELATION				
Q(1)	.6040	.1069	1.8079	.7072
	[.4370]	[.7437]	[.1787]	[.4003]
Q(6)	6.5500	2.5140	6.9660	5.2120
	[.3644]	[.8669]	[.3240]	[.5169]
BREUSCH-PAGAN-GODFREY (B-P-G) TEST FOR HETEROSCEDASTICITY				
chi-square (3 d.f.)	3.789	4.773	3.066	3.329
	[.2852]	[.1892]	[.2544]	[.3436]
ARCH(12) TEST				
chi-square	18.246	6.844	7.197	
(12 d.f)	[.1084]	[.8677]	[.8443]	
Percent of portfolios for which ARCH(12) is rejected at .05 level				60%

[a]The t-statistics are given in the parentheses.

[b] The average t-values are in the parentheses for the mean factor loading estimates.

[c]The probability values are given in the brackets.

TABLE 3.4

The selected national LFM for the TSE300, VW, EW and fifty portfolios for the total period March 1969 through March 1988.

Variable	TSE300	VW	EW	PR1-PR50
INTERCEPT	.0077	.0114	.0145	.0146
	(2.48)[a]	(3.58)	(4.53)	(3.13)[b]
	[.0068][c]	[.0002]	[.0001]	[.0010]
LINDUS	.6513	.6188	.7605	.7452
	(2.54)	(2.34)	(2.87)	(1.85)
	[.0058]	[.0099]	[.0022]	[.0328]
INDEX	1.6193	1.6980	1.6532	1.5607
	(4.93)	(5.01)	(4.86)	(3.24)
	[.0001]	[.0001]	[.0001]	[.0007]
TERM	-.2512	-.2348	-.2916	-.3150
	(-2.80)	(-2.54)	(-3.14)	(-2.66)
	[.0027]	[.0058]	[.0009]	[.0042]
R^2	.1562	.1450	.1604	.1005
F-value	13.249	12.176	15.491	8.385
D.W.	2.032	2.015	1.864	1.889
Q-TEST FOR AUTOCORRELATION				
Q(1)	.0680	.0169	1.1689	.8538
	[.7942]	[.8965]	[.2796]	[.3555]
Q(6)	5.6946	4.4319	5.6642	6.4867
	[.4582]	[.6184]	[.4618]	[.3709]
BREUSCH-PAGAN-GODFREY B-P-G TEST FOR HETEROSCEDASTICITY				
chi-square (3 d.f.)	5.397	5.955	.817	2.210
	[.1449]	[.1138]	[.8454]	[.5299]
ARCH TEST				
chi-square	19.516	8.229	19.018	
(12 d.f)	[.07680	[.7675]	[.0881]	
Percent of portfolios for which ARCH(12) is rejected at .05 level				62%

[a]The t-statistics are given in the parentheses.

[b] The average t-values are in the parentheses for the mean factor loading estimates.

[c]The probability values are given in the brackets.

TABLE 3.5

The complete LFM for the three market indices and the 50 size-ranked portfolios for the total period March 1969 through March 1988.

Variable	TSE300	VW	EW	PR1-PR50
INTERCEPT	.0058	.0098	.0122	.0120
	(2.14)[a]	(3.35)	(4.16)	(2.86)[b]
	[.0165][c]	[.0005]	[.0001]	[.0023]
DLINDUSI	1.3073	1.1479	1.1201	1.1651
	(2.85)	(2.36)	(2.28)	(1.31)
	[.0024]	[.0097]	[.0117]	[.0957]
LINDUSD	1.2147	1.1476	1.4368	1.8363
	(2.68)	(2.38)	(2.96)	(2.47)
	[.0039]	[.0089]	[.0017]	[.0071]
DINDEXI	1.6274	1.5278	1.6322	1.1295
	(5.14)	(4.54)	(4.83)	(3.25)
	[.0001]	[.0001]	[.0001]	[.0007]
INDEXD	1.5776	1.5287	1.6322	1.5379
	(5.39)	(5.34)	(5.20)	(3.09)
	[.0001]	[.0001]	[.0001]	[.0011]
DTERMI	-.1205	-.0941	-.1910	-.2369
	(-1.21)	(-.892)	(-1.79)	(-1.38)
	[.1128]	[.1868]	[.0369]	[.0845]
TERMD	-.3511	-.3472	-.3493	-.6457
	(-3.27)	(-3.04)	(-4.40)	(-2.86)
	[.0006]	[.0013]	[.0001]	[.0023]
REURO	-1.8652	-1.6104	-1.5483	-1.5905
	(-5.69)	(-4.62)	(-4.40)	(-2.78)
	[.0001]	[.0001]	[.0001]	[.0029]
R^2	.3390	.2906	.2987	.1679
F-value	16.192	12.935	13.446	6.053
D.W.	2.078	2.108	1.871	1.949
Q-TESTS FOR AUTOCORRELATION				
Q(1)	.3882	.7077	.8937	.2390
	[.5332]	[.4002]	[.3445]	[.6249]
Q(6)	7.5762	3.5200	7.3582	7.5670
	[.2708]	[.7413]	[.2889]	[.2715]
BREUSCH-PAGAN-GODFREY B-P-G TEST FOR HETEROSCEDASTICITY				
chi-square (7 d.f.)	10.375	8.538	9.824	4.607
	[.1683]	[.2875]	[.1987]	[.2029]
ARCH TEST				
chi-square	15.906	6.455	10.284	
(12 d.f)	[.1956]	[.8914]	[.5911]	
Percent of portfolios for which ARCH(12) is rejected at .05 level				49%

[a]The t-statistics are given in the parentheses.

[b] The average t-values are in the parentheses for the mean factor loading estimates.

[c]The probability values are given in the brackets.

TABLE 3.6

Estimates of the risk premia for the augmented IAPT given by equation (3.3.6) for the fifty size ranked portfolios for various time periods.

(Sub)period	LINDUSD	INDEXD	TERMD	DLINDUS	DINDEX	REURO	$\chi^2(3)$[c]
Panel A: NOLS Estimates							
Mar.1969-Mar.1988	-.0046**	-.0052*	-.0123*	.0014	.0011	-.0062**	13.52
	(-2.01)[a]	(-1.79)	(-1.81)	(0.48)	(0.57)	(-2.60)	[.0036]
	[.0456][b]	[.0753]	[.0723]	[.6335]	[.5702]	[.0100]	
Mar.1969-Mar.1978	.0017	.0167**	-.0110	-.0112	-.0062*	.0033	6.39
	(0.53)	(2.10)	(-1.40)	(-1.56)	(-1.73)	(1.43)	[.0941]
	[.6000]	[.0382]	[.1648]	[.1208]	[.0858]	[.1563]	
Mar.1978-Mar.1988	-.0139**	.0027	-.0154	.0032	-.0045	-.0145**	6.43
	(-2.06)	(0.77)	(-1.14)	(0.71)	(-1.11)	(-2.26)	[.0924]
	[.0421]	[.4428]	[.2562]	[.4778]	[.2708]	[.0259]	
Panel B: NSUR Estimates							
Mar.1969-Mar.1988	-.0022*	-.0003	-.0137**	.0010	-.0039**	-.0047**	8.61
	(-1.71)	(-0.26)	(-3.24)	(0.57)	(-2.41)	(-2.94)	[.0349]
	[.0888]	[.7971]	[.0014]	[.5668]	[.0166]	[.0036]	
Mar.1969-Mar.1978	-.0032*	.0066**	-.0177**	-.0046*	-.0077**	-.0004	14.25
	(-1.85)	(2.59)	(-3.23)	(-1.67)	(-4.04)	(-0.33)	[.0025]
	[.0670]	[.0109]	[.0016]	[.0975]	[.0001]	[.7442]	
Mar.1978-Mar.1988	-.0131**	.0013	-.0133	.0016	-.0029	-.0130**	8.07
	(-4.54)	(0.62)	(-1.60)	(0.65)	(-1.21)	(-3.75)	[.0446]
	[.0001]	[.5371]	[.1114]	[.5184]	[.2291]	[.0003]	

[a]The t-values are in the parentheses.
[b]The probability values are in the brackets.

[c]The χ^2-statistics have three degrees of freedom and test the null hypothesis that the Canadian equity market is integrated relative to a North-American equity market (i.e. the prices of risk for the domestic factors are jointly equal to zero).
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 3.7

Estimates of the risk premia for the augmented IAPT given by equation (3.3.6) with the residual market factor (RMF) for the fifty size-ranked portfolios for various time periods.

(Sub)period	LINDUSD	INDEXD	TERMD	DLINDUS	DINDEX	REURO	RMF	$\chi^2(3)^c$
Panel A: NOLS Estimates								
Mar.1969-Mar.1988	-.0092	-.0080	-.0437	-.0001	.0094	-.0162	-.0375	1.84
	(-1.48)[a]	(-1.17)	(-1.45)	(-0.02)	(1.20)	(-1.58)	(-1.24)	[.6063]
	[.1416][b]	[.2414]	[.1482]	[.9862]	[.2333]	[.1146]	[.2171]	
Mar.1969-Mar.1978	-.0039**	-.0038	.0135**	.0031*	-.0021	.0002	.0283**	11.91
	(-2.52)	(-1.51)	(2.90)	(1.17)	(-1.46)	(0.30)	(4.96)	[.0077]
	[.0134]	[.1350]	[.0046]	[.2462]	[.1481]	[.7665]	[.0001]	
Mar.1978-Mar.1988	-.0120**	.0010	-.0303*	.0033	-.0042	-.0176**	-.0181	7.03
	(-2.05)	(0.31)	(-1.69)	(0.74)	(-1.05)	(-2.36)	(-1.54)	[.0709]
	[.0423]	[.7571]	[.0929]	[.4586]	[.2962]	[.0198]	[.1270]	
Panel B: NSUR Estimate								
Mar.1969-Mar.1988	-.0017	.0001	-.0118**	.0012	-.0053**	-.0035**	.0047	5.22
	(-1.31)	(0.05)	(-2.70)	(0.72)	(-2.95)	(-2.22)	(0.99)	[.1564]
	[.1908]	[.9593]	[.0075]	[4741]	[.0035]	[.0273]	[.3222]	
Mar.1969-Mar.1978	-.0035**	-.0017	.0048*	.0028*	-.0037**	-.0007	.0274**	11.79
	(-3.39)	(-1.60)	(1.78)	(1.79)	(-3.73)	(-1.19)	(7.88)	[.0081]
	[.0010]	[.1123]	[.0776]	[.0758]	[.0003]	[.2353]	[.0001]	
Mar.1978-Mar.1988	-.0132**	.0009	-.0151*	.0015	-.0024	-.0143**	-.0120*	8.58
	(-4.38)	(0.40)	(-1.71)	(0.60)	(-0.92)	(-3.67)	(-1.79)	[.0354]
	[.0001]	[.6872]	[.0899]	[.5477]	[.3612]	[.0004]	[.0767]	

[a]The t-values are in the parentheses.
[b]The probability values are in the brackets.

[c]The χ^2-statistics have three degrees of freedom and test the null hypothesis that the Canadian equity market is integrated relative to a North-American equity market (i.e. the prices of risk for the domestic factors are jointly equal to zero).
*Significant at the 0.01 level.
**Significant at the 0.05 level.

TABLE 3.8

Mean estimates for the factor sensitivities for the augmented IAPT given by equation (3.3.6) for the fifty size-ranked portfolios for various time periods.

(Sub)period	LINDUSD	INDEXD	TERMD	DLINDUS	DINDEX	REURO
Panel A: NOLS Estimates						
Mar.1969-Mar.1988	.9346*	1.5265**	-.3101**	.6914*	1.6301**	-1.4850**
	(1.79)[a]	(3.61)	(-1.98)	(1.90)	(3.27)	(-3.07)
	[.0742][b]	[.0004]	[.0487]	[.0589]	[.0012]	[.0024]
Mar.1969-Mar.1978	.5256	1.2691**	-.4754	.6878	1.9931**	-2.4210**
	(0.67)	(2.06)	(-1.08)	(0.86)	(2.16)	(-2.34)
	[.5039]	[.0417]	[.2839]	[.3906]	[.0328]	[.0214]
Mar.1978-Mar.1988	.8538*	1.5189**	-.3526**	.3928	1.6773**	-1.1499**
	(1.67)	(2.77)	(-2.12)	(0.97)	(3.12)	(-2.29)
	[.0986]	[.0066]	[.0363]	[.3329]	[.0023]	[.0241]
Panel B: NSUR Estimates						
Mar.1969-Mar.1988	.9546*	1.4739**	-.3600**	. 7029*	1.4247**	-1.7103**
	(1.87)	(3.25)	(-2.33)	(1.87)	(3.04)	(-3.56)
	[.0634]	[.0013]	[.0204]	[.0634]	[.0026]	[.0005]
Mar.1969-Mar.1978	.7568	1.5655**	-.3936	.8109	1.9939**	-2.6844**
	(0.95)	(2.01)	(-1.33)	(1.45)	(2.29)	(-2.40)
	[.3426]	[.0473]	[.1910]	[.1510]	[.0243]	[.0184]
Mar.1978-Mar.1988	1.0532*	1.3749**	-.3822**	.4724	1.6660**	-1.3983*
	(1.97)	(2.41)	(-2.46)	(0.76)	(2.45)	(-1.81)
	[.0510]	[.0180]	[.0155]	[.4510]	[.0157]	[.0728]

[a]The average t-values are in the parentheses for the mean factor loading estimates.
[b]The probability values are in the brackets.
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 3.9

Mean estimates of the factor sensitivities for the augmented IAPT given by equation (3.3.6) with the residual market factor (RMF) for the fifty size-ranked portfolios for various time periods.

(Sub) period	LINDUSD	INDEXD	TERMD	DLINDUS	DINDEX	REURO	RMF
Panel A: NOLS Estimates							
Mar.1969-Mar.1988	.9445**	1.5445**	-.2552**	.4711	1.6598**	-1.5506**	.8859**
	(2.10)[a]	(4.05)	(-1.95)	(1.26)	(3.28)	(-3.95)	(9.56)
	[.0379][b]	[.0001]	[.0519]	[.2073]	[.0012]	[.0001]	[.0001]
Mar.1969-Mar.1978	.8183	1.2836**	-.3108	.4788	1.8878*	-2.5344**	.9496**
	(1.23)	(2.27)	(-1.10)	(1.03)	(2.16)	(-2.33)	(7.20)
	[.2224]	[.0251]	[.2755]	[.3078]	[.0330]	[.0218]	[.0001]
Mar.1978-Mar.1988	.8291*	.9807**	-.2104*	.2892	1.5431**	-1.0449**	.7616**
	(1.98)	(2.01)	(-1.86)	(0.92)	(3.94)	(-2.26)	(6.24)
	[.0503]	[.0472]	[.0656]	[.3608]	[.0001]	[.0259]	[.0001]
Panel B: NSUR Estimates							
Mar.1969-Mar.1988	.8932**	1.5844**	-.2844*	.5984*	1.6939**	-1.4991**	.8574**
	(2.11)	(3.79)	(-1.77)	(1.71)	(3.73)	(-2.68)	(9.47)
	[.0361]	[.0001]	[.0775]	[.0892]	[.0002]	[.0080]	[.0001]
Mar.1969-Mar.1979	1.1383*	1.4681**	-.2797	.2284	1.8972**	-1.9035**	.9472**
	(1.73)	(2.29)	(-0.95)	(0.56)	(2.20)	(-2.18)	(6.97)
	[.0875]	[.0241]	[.3467]	[.5845]	[.0303]	[.0318]	[.0001]
Mar.1978-Mar.1988	.7724*	1.2183**	-.3103*	.3146	1.5188**	-1.1921**	.7022**
	(1.76)	(2.13)	(-1.93)	(0.98)	(3.04)	(-2.48)	(7.51)
	[.0810]	[.0352]	[.0561]	[.3271]	[.0030]	[.0145]	[.0001]

[a]The average t-values are in the parentheses for the mean factor loading estimates.
[b]The probability values are in the brackets.
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 3.10

Estimates of the risk premia for the augmented APT given by equation (3.3.7) for the fifty size-ranked portfolios for various time periods.

(Sub)period	LINDUS	TERM	INDEX	REURO	DINDEXI	$\chi^2(2)$[c]
Panel A: NOLS Estimates						
Mar.1969-Mar.1988	-.0048*	-.0159**	-.0002	-.0055**	-.0012	5.58
	(-1.78)[a]	(-2.18)	(-0.14)	(-3.02)	(-0.75)	[.0614]
	[.0765][b]	[.0306]	[.8901]	[.0028]	[.4517]	
Mar.1969-Mar.1978	-.0031	-.0211**	.0056**	-.0007	-.0037**	5.21
	(-1.15)	(-3.26)	(2.63)	(-1.44)	(-2.88)	[.0739]
	[.2519]	[.0015]	[.0099]	[.5929]	[.0048]	
Mar.1978-Mar.1988	-.0072*	-.0269**	.0011	-.0124**	-.0056*	9.83
	(-1.71)	(-2.12)	(0.49)	(-3.03)	(-1.81)	[.0073]
	[.0895]	[.0358]	[.6236]	[.0030]	[.0727]	
Panel B: NSUR Estimates						
Mar.1969-Mar.1988	.0006	-.0294**	.0024	-.0035*	-.0096**	11.94
	(0.29)	(-3.52)	(1.41)	(-1.89)	(-3.52)	[.0025]
	[.7691]	[.0005]	[.1592]	[.0595]	[.0005]	
Mar.1969-Mar.1978	.0055	.0365**	-.0001*	-.0018*	.0057**	10.72
	(3.10)	(6.63)	(-0.02)	(-1.68)	(4.16)	[.0047]
	[.0025]	[.0001]	[.9804]	[.0961]	[.0001]	
Mar.1978-Mar.1988	-.0140**	-.0110	.0014	-.0165**	-.0019	6.49
	(-4.27)	(-1.22)	(0.75)	(-4.54)	(-0.90)	[.0388]
	[.0001]	[.2257]	[.4571]	[.0001]	[.3713]	

[a]The t-values are in the parentheses.
[b]The probability values are in the brackets.

[c]The χ^2-statistics have two degrees of freedom and test the null hypothesis that the Canadian equity market is segmented relative to a North-American equity market (i.e. the prices of risk for the international factors are jointly equal to zero).
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 3.11

Estimates of the risk premia for the augmented APT given by equation (3.3.7) with the residual market factor (RMF) for the fifty size-ranked portfolios for various time periods.

(Sub)period	LINDUS	TERM	INDEX	REURO	DINDEXI	RMF	$\chi^2(2)$[c]
Panel A: NOLS Estimates							
Mar.1969-Mar.1988	-.0001	.0347**	.0020	.0031	-.0063**	.0319**	5.08
	(-0.01)[a]	(2.07)	(0.76)	(1.13)	(-2.10)	(2.73)	[.0788]
	[.9914][b]	[.0395]	[.4498]	[.2593]	[.0367]	[.0068]	
Mar.1969-Mar.1978	-.0022	.0379**	-.0028	.0016	-.0019	.0424**	0.57
	(-0.63)	(2.06)	(-0.72)	(1.03)	(-0.83)	(2.94)	[.7520]
	[.5324]	[.0416]	[.4750]	[.3049]	[.4067]	[.0040]	
Mar.1978-Mar.1988	-.0042	-.0661**	.0015	-.0150**	-.0029	-.0307**	5.64
	(-0.90)	(-2.37)	(0.56)	(-2.54)	(-0.85)	(-2.15)	[.0596]
	[.3709]	[.0194]	[.5733]	[.0125]	[.3996]	[.0337]	
Panel B: NSUR Estimates							
Mar.1969-Mar.1988	.0025	-.0049	.0020	.0015	-.0104**	.0243**	5.82
	(1.08)	(-0.67)	(1.19)	(0.84)	(-3.62)	(3.48)	[.0545]
	[.2799]	[.5028]	[.2361]	[.4008]	[.0004]	[.0006]	
Mar.1969-Mar.1978	.0001	.0363**	-.0068**	.0006	-.0013	.0452**	0.46
	(0.01)	(2.85)	(-2.56)	(0.52)	(-0.74)	(5.00)	[.7945]
	[.9951]	[.0052]	[.0120]	[.6073]	[.4622]	[.0001]	
Mar.1978-Mar.1988	-.0129**	-.0346**	.0003	-.0217**	.0026	-.0383**	6.25
	(-3.10)	(-2.36)	(0.14)	(-3.61)	(0.90)	(-3.03)	[.0439]
	[.0024]	[.0201]	[.8883]	[.0005]	[.3715]	[.0030]	

[a]The t-values are in the parentheses.
[b]The probability values are in the brackets.

[c]The χ^2-statistics have two degrees of freedom and test the null hypothesis that the Canadian equity market is segmented relative to a North-American equity market (i.e. the prices of risk for the international factors are jointly equal to zero).
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 3.12

Mean estimates for the factor sensitivities for the augmented APT given by equation (3.3.7) for the fifty size-ranked portfolios for various time periods.

(Sub)period	LINDUS	TERM	INDEX	REURO	DINDEXI
Panel A: NOLS Estimates					
Mar.1969-Mar.1988	.9154**	-.2371*	1.5043**	-1.5185**	1.4458**
	(2.46)[a]	(-1.87)	(2.43)	(-2.82)	(2.99)
	[.0148][b]	[.0643]	[.0157]	[.0052]	[.0031]
Mar.1969-Mar.1978	.5678	-.2295	1.0487*	-1.5109**	1.7952**
	(1.63)	(-1.06)	(1.88)	(-2.01)	(2.25)
	[.1057]	[.2928]	[.0627]	[.0470]	[.0263]
Mar.1978-Mar.1988	.9583*	-.2736*	1.3071**	-1.1796*	1.3667**
	(2.08)	(-1.93)	(2.27)	(-2.56)	(2.58)
	[.0395]	[.0564]	[.0248]	[.0118]	[.0111]
Panel B: NSUR Estimates					
Mar.1969-Mar.1988	.8935**	-.2156*	1.5319**	-1.5462**	1.6338**
	(2.26)	(-1.79)	(2.49)	(-3.13)	(3.55)
	[.0246]	[.0754]	[.0137]	[.0020]	[.0005]
Mar.1969-Mar.1978	.4869	-.1657	.8495	-1.6538*	1.8823**
	(0.84)	(-0.62)	(1.16)	(-1.71)	(2.68)
	[.4006]	[.5375]	[.2477]	[.0903]	[.0084]
Mar.1978-Mar.1988	1.5821**	-.2152	1.1065*	-1.0190**	1.5673**
	(2.59)	(-1.55)	(1.97)	(-2.13)	(2.19)
	[.0107]	[.1244]	[.0512]	[.0349]	[.0303]

[a]The average t-values are in the parentheses for the mean factor loading estimates.
[b]The probability values are in the brackets.
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 3.13

Mean estimates of the factor sensitivities for the augmented APT given by equation (3.3.7) with the residual market factor (RMF) for the fifty size-ranked portfolios for various time periods.

(Sub)period	LINDUS	TERM	INDEX	REURO	DINDEXI	RMF
Panel A: NOLS Estimates						
Mar.1969-Mar.1988	.8159**	-.2493**	1.5534**	-1.3448**	1.6721**	.8613**
	(2.44)[a]	(-2.21)	(3.39)	(-2.90)	(3.83)	(8.87)
	[.0155][b]	[.0280]	[.0008]	[.0041]	[.0002]	[.0001]
Mar.1969-Mar.1978	.7622*	-.3109**	1.8416**	-1.9538**	1.7871**	.9438**
	(1.25)	(-2.78)	(2.82)	(-2.09)	(2.51)	(6.64)
	[.2134]	[.0065]	[.0058]	[.0395]	[.0136]	[.0001]
Mar.1978-Mar.1988	.8154**	-.3022**	1.4958**	-1.2639**	1.5223**	.8042**
	(2.65)	(-3.18)	(3.78)	(-3.13)	(2.57)	(8.26)
	[.0086]	[.0017]	[.0002]	[.0020]	[.0107]	[.0001]
Panel B: NSUR Estimates						
Mar.1969-Mar.1988	.8922**	-.3261**	1.5837**	-1.4049**	1.6857**	.7853**
	(2.34)	(-2.79)	(4.10)	(-3.23)	(4.25)	(9.24)
	[.0204]	[.0057]	[.0001]	[.0014]	[.0001]	[.0001]
Mar.1969-Mar.1978	.7781	-.3343*	1.8618**	-2.1225**	1.7931**	.8766*
	(1.46)	(-1.79)	(2.93)	(-2.61)	(2.52)	(8.15)
	[.1483]	[.0766]	[.0042]	[.0104]	[.0133]	[.0001]
Mar.1978-Mar.1988	.7764**	-.2888**	1.6830**	-1.5105**	1.6542**	.7853**
	(2.56)	(-2.71)	(4.58)	(-3.31)	(3.77)	(8.82)
	[.0133]	[.0072]	[.0001]	[.0011]	[.0002]	[.0001]

[a]The average t-values are in the parentheses for the mean factor loading estimates.
[b]The probability values are in the brackets.
*Significant at the 0.10 level.
**Significant at the 0.05 level.

TABLE 4.1

Regression results for the selected linear factor model (LFM) for the TSE300 and the fifty portfolios for the total period, March 1962 through December 1987, are presented herein. EX is the innovation in the exchange rate (Cdn/US), USINDEX is the innovation in the U.S. composite index of 12 leading indicators, EXPORTS is the innovation in total exports, TERM is the innovation in the term structure, CINDEX is the innovation in the Canadian index of 10 leading indicators, LINDUS is the lagged innovation in industrial production, RMF is the residual market factor, GTSE300 is the return on the Toronto Stock Exchange 300 stock index, and PR1-PR50 are the returns on the fifty size-ranked portfolios.

DEPENDENT VARIABLE	*INTERCEPT*	*EX*	*USINDEX*	*EXPORTS*	*TERM*	*CINDEX*	*LINDUS*	*RMF*	*F-VALUE*	R^2
GTSE300	0.0064	-0.8966	1.7201	-0.1626	-0.2312	1.6485	0.5502	N.A.	15.674	0.2346
	(2.58)[a]	(-2.74)	(5.75)	(-3.44)	(-2.82)	(6.09)	(2.65)	N.A.		
	[0.0051][b]	[0.0032]	[0.0001]	[0.0003]	[0.0025]	[0.0001]	[0.0042]	N.A.		
PR1-PR50	0.0143[c]	-0.7364	1.8881	-0.1899	-0.3068	1.6675	0.6523	0.8109	15.792	0.2363
	(3.98)	(-1.56)	(4.40)	(-2.77)	(-2.62)	(4.25)	(2.14)	(9.81)		
	[0.0001]	[0.0593]	[0.0001]	[0.0029]	[0.0046]	[0.0001]	[0.0165]	[0.0001]		

[a]The t-statistics are given in the parentheses.

[b]The probability values are given in the brackets.

[c]The mean coefficient estimates are given in this row. The average t-values are in the parentheses in the next row for the mean coefficient estimates.

TABLE 4.2

Selected estimates for the risk premia for the fifty portfolios are reported below. All variables are defined in Table 1.

PANEL A: Constant risk premia estimates

Period[a]	LINDUS	CINDEX	TERM	USINDEX	EXPORTS	EX	RMF
1962.3-1967.3	-0.0287	0.0088	-0.0084	-0.0086	-0.0020	0.0112	0.0313
	(-2.61)[b]	(2.09)	(-1.59)	(-1.95)	(-0.11)	(2.56)	(2.25)
	[0.0152][c]	[0.0468]	[0.1128]	[0.0611]	[0.3947]	[0.0171]	[0.0338]
1967.3-1972.3	0.0133	-0.0099	-0.0315	-0.0028	-0.0149	-0.0089	-0.0101
	(4.55)	(-6.55)	(-6.32)	(-1.96)	(-3.43)	(-3.91)	(-1.76)
	[0.0001]	[0.0001]	[0.0001]	[0.0549]	[0.0012]	[0.0003]	[0.0837]
1972.3-1977.3	-0.0263	0.0154	-0.0407	0.0065	0.0739	0.0339	0.0375
	(-1.90)	(1.91)	(-1.77)	(1.35)	(1.77)	(2.06)	(2.84)
	[0.0630]	[0.0618]	[0.0819]	[0.1817]	[0.0826]	[0.0441]	[0.0063]
1977.3-1982.3	-0.1069	0.0119	0.3418	0.0860	-0.1408	0.1118	0.0352
	(-0.98)	(0.93)	(1.02)	(1.00)	(-0.91)	(0.96)	(0.65)
	[0.3327]	[0.3574]	[0.3145]	[0.3210]	[0.3680]	[0.3391]	[0.5158]
1982.3-1987.3	-0.0095	-0.0057	0.0102	0.0020	0.0014	-0.0039	0.0214
	(-7.34)	(-5.39)	(2.47)	(1.87)	(0.31)	(-3.29)	(5.63)
	[0.0001]	[0.0001]	[0.0168]	[0.0670]	[0.7575]	[0.0017]	[0.0001]

PANEL B: Time-varying risk premia estimates

1962.3-1967.3	-0.7690	0.2795	-0.1298	-0.5160	0.0121	8.8106	-1.0570
	(-8.57)	(5.41)	(-6.60)	(-10.40)	(1.24)	(2.63)	(-3.29)
	[0.0001]	[0.0001]	[0.0001]	[0.0001]	[0.2203]	[0.0113]	[0.0018]
1967.3-1972.3	-0.1411	0.6953	0.0392	0.0455	0.0097	-0.6925	1.9271
	(-2.19)	(19.06)	(2.61)	(1.20)	(4.68)	(-20.30)	(5.03)
	[0.0330]	[0.0001]	[0.0119]	[0.2356]	[0.0001]	[0.0001]	[0.0001]
1972.3-1977.3	-0.6742	0.4662	2.0067	0.0112	-0.0522	-1.3987	3.8627
	(-10.49)	(6.28)	(5.34)	(0.60)	(-7.48)	(-8.72)	(9.70)
	[0.0001]	[0.0001]	[0.0001]	[0.5492]	[0.0001]	[0.0001]	[0.0001]
1977.3-1982.3	-0.5157	0.7939	0.0521	-0.2782	0.0335	7.5030	1.2097
	(-13.70)	(12.01)	(13.87)	(-5.86)	(5.73)	(7.83)	(1.95)
	[0.0001]	[0.0001]	[0.0001]	[0.0001]	[0.0001]	[0.0001]	[0.0569]
1982.3-1987.3	-0.4890	1.3742	-0.0447	-0.1146	0.0289	-0.3838	0.9729
	(-20.24)	(9.96)	(-11.63)	(-2.56)	(3.95)	(-4.48)	(2.61)
	[0.0001]	[0.0001]	[0.0001]	[0.0134]	[0.0002]	[0.0001]	[0.0117]

[a] x.y refers to the year, followed by the month of the year.

[b] The t-statistics are given in the parentheses.

[c] The probability values are given in the brackets.

TABLE 4.3

The variance ratio VAR(R*)/VAR(R) for each of the fifty portfolios when the constant risk premia model is used to derive R^* is reported below. Period covered is from April 1967 to December 1987. R* and R are the predicted and actual portfolio returns, respectively.

PORTFOLIO	VAR(R*)	VAR(R)	VAR(R*)/VAR(R)
1	0.0023	0.0056	0.4101
2	0.0027	0.0051	0.5243
3	0.0029	0.0052	0.5550
4	0.0026	0.0069	0.3737
5	0.0034	0.0061	0.5527
6	0.0028	0.0070	0.4073
7	0.0031	0.0054	0.5800
8	0.0029	0.0055	0.5393
9	0.0027	0.0057	0.4700
10	0.0021	0.0051	0.4110
11	0.0021	0.0037	0.5744
12	0.0017	0.0041	0.4178
13	0.0026	0.0057	0.4638
14	0.0024	0.0062	0.3953
15	0.0021	0.0045	0.4703
16	0.0037	0.0050	0.7301
17	0.0029	0.0074	0.3938
18	0.0033	0.0063	0.5333
19	0.0026	0.0064	0.4067
20	0.0034	0.0064	0.5372
21	0.0033	0.0063	0.5255
22	0.0030	0.0079	0.3797
23	0.0021	0.0051	0.4185
24	0.0032	0.0059	0.5409
25	0.0027	0.0058	0.4616
26	0.0027	0.0066	0.4139
27	0.0022	0.0054	0.4158
28	0.0049	0.0085	0.5788
29	0.0028	0.0050	0.5626
30	0.0023	0.0054	0.4288
31	0.0027	0.0056	0.4870
32	0.0027	0.0061	0.4460
33	0.0035	0.0073	0.4768
34	0.0036	0.0083	0.4378
35	0.0035	0.0056	0.6247
36	0.0036	0.0067	0.5334
37	0.0023	0.0047	0.4866
38	0.0028	0.0057	0.4901
39	0.0037	0.0073	0.4997
40	0.0035	0.0071	0.4995
41	0.0021	0.0071	0.2938
42	0.0031	0.0055	0.5738
43	0.0032	0.0054	0.5835
44	0.0041	0.0076	0.5437
45	0.0043	0.0077	0.5637
46	0.0031	0.0065	0.4731
47	0.0038	0.0069	0.5460
48	0.0031	0.0064	0.4856
49	0.0044	0.0067	0.6599
50	0.0029	0.0049	0.5854
MEAN	0.0030	0.0061	0.4952

TABLE 4.4

The variance ratio VAR(R*)/VAR(R) for each of the fifty portfolios when the time-varying risk premia model is used to derive R^* are reported below. Period covered is from April 1967 to December 1987. R* and R are the predicted and actual portfolio returns, respectively.

PORTFOLIO	VAR(R*)	VAR(R)	VAR(R*)/VAR(R)
1	0.0028	0.0056	0.5036
2	0.0046	0.0051	0.9014
3	0.0029	0.0052	0.5570
4	0.0027	0.0069	0.3875
5	0.0034	0.0061	0.5535
6	0.0032	0.0070	0.4522
7	0.0030	0.0054	0.5546
8	0.0025	0.0055	0.4604
9	0.0025	0.0057	0.4370
10	0.0022	0.0051	0.4421
11	0.0022	0.0037	0.5890
12	0.0019	0.0041	0.4593
13	0.0026	0.0057	0.4639
14	0.0025	0.0062	0.4024
15	0.0021	0.0045	0.4673
16	0.0037	0.0050	0.7362
17	0.0026	0.0074	0.3562
18	0.0038	0.0063	0.6121
19	0.0027	0.0064	0.4122
20	0.0034	0.0064	0.5246
21	0.0034	0.0063	0.5394
22	0.0034	0.0079	0.4276
23	0.0024	0.0051	0.4603
24	0.0031	0.0059	0.5251
25	0.0024	0.0058	0.4179
26	0.0030	0.0066	0.4569
27	0.0023	0.0054	0.4322
28	0.0047	0.0085	0.5498
29	0.0028	0.0050	0.5617
30	0.0024	0.0054	0.4380
31	0.0040	0.0056	0.7107
32	0.0030	0.0061	0.4850
33	0.0036	0.0073	0.4886
34	0.0033	0.0083	0.4014
35	0.0033	0.0056	0.5928
36	0.0038	0.0067	0.5621
37	0.0025	0.0047	0.5353
38	0.0028	0.0057	0.4852
39	0.0037	0.0073	0.5031
40	0.0036	0.0071	0.5014
41	0.0022	0.0071	0.3071
42	0.0032	0.0055	0.5760
43	0.0027	0.0054	0.5016
44	0.0043	0.0076	0.5654
45	0.0049	0.0077	0.6401
46	0.0027	0.0065	0.4092
47	0.0037	0.0069	0.5381
48	0.0034	0.0064	0.5306
49	0.0043	0.0067	0.6443
50	0.0026	0.0049	0.5312
MEAN	0.0031	0.0061	0.5118

TABLE 4.5

The variance ratio VAR(RP*)/VAR(RP) for the equally-weighted market portfolio when the constant risk premia model is used to derive RP* are reported below. RP* and RP are the predicted and actual returns on an equally-weighted portfolio, respectively.

PERIOD	VAR(RP*)	VAR(RP)	VAR(RP*)/VAR(RP)
1967.4-1967.12	0.0006	0.0007	0.7988
1968.1-1968.12	0.0008	0.0016	0.4991
1969.1-1969.12	0.0016	0.0029	0.5447
1970.1-1970.12	0.0021	0.0041	0.5179
1971.1-1971.12	0.0017	0.0031	0.5585
1972.1-1972.12	0.0010	0.0026	0.3980
1973.1-1973.12	0.0024	0.0034	0.7007
1974.1-1974.12	0.0051	0.0045	1.1321
1975.1-1975.12	0.0043	0.0082	0.5227
1976.1-1976.12	0.0032	0.0011	2.8208
1977.1-1977.12	0.0013	0.0004	3.7296
1978.1-1978.12	0.0021	0.0007	3.1302
1979.1-1979.12	0.0031	0.0020	1.5513
1980.1-1980.12	0.0062	0.0034	1.8285
1981.1-1981.12	0.0016	0.0019	0.8347
1982.1-1982.12	0.0021	0.0035	0.6046
1983.1-1983.12	0.0004	0.0022	0.2012
1984.1-1984.12	0.0009	0.0009	1.1054
1985.1-1985.12	0.0006	0.0008	0.7887
1986.1-1986.12	0.0004	0.0005	0.8330
1987.1-1987.12	0.0030	0.0061	0.4957
1967.4-1972.3	0.0014	0.0029	0.4838
1972.4-1977.3	0.0035	0.0041	0.8543
1977.4-1982.3	0.0029	0.0021	1.4038
1982.4-1987.3	0.0009	0.0015	0.6035
1967.4-1987.3	0.0022	0.0026	0.8288
1967.4-1987.12	0.0022	0.0028	0.7999

TABLE 4.6

The variance ratio VAR(RP*)/VAR(RP) for the equally-weighted market portfolio when the time-varying risk premia model is used to derive RP^* are reported below. RP* and RP are the predicted and actual returns on an equally-weighted portfolio, respectively.

PERIOD	VAR(RP*)	VAR(RP)	VAR(RP*)/VAR(RP)
1967.4-1967.12	0.0007	0.0007	0.9565
1968.1-1968.12	0.0008	0.0016	0.5062
1969.1-1969.12	0.0015	0.0029	0.5098
1970.1-1970.12	0.0016	0.0041	0.3999
1971.1-1971.12	0.0018	0.0031	0.5719
1972.1-1972.12	0.0011	0.0026	0.4104
1973.1-1973.12	0.0029	0.0034	0.8421
1974.1-1974.12	0.0043	0.0045	0.9559
1975.1-1975.12	0.0043	0.0082	0.5227
1976.1-1976.12	0.0058	0.0011	5.0993
1977.1-1977.12	0.0007	0.0004	1.9957
1978.1-1978.12	0.0028	0.0007	4.1397
1979.1-1979.12	0.0032	0.0020	1.6366
1980.1-1980.12	0.0073	0.0034	2.1582
1981.1-1981.12	0.0016	0.0019	0.8389
1982.1-1982.12	0.0019	0.0035	0.5454
1983.1-1983.12	0.0005	0.0022	0.2500
1984.1-1984.12	0.0010	0.0009	1.1786
1985.1-1985.12	0.0008	0.0008	0.9791
1986.1-1986.12	0.0005	0.0005	0.8733
1987.1-1987.12	0.0030	0.0061	0.4886
1967.4-1972.3	0.0014	0.0029	0.4838
1972.4-1977.3	0.0038	0.0041	0.9284
1977.4-1982.3	0.0030	0.0021	1.4567
1982.4-1987.3	0.0010	0.0015	0.6400
1967.4-1987.3	0 0023	0.0026	0.8678
1967.4-1987.12	0.0023	0.0028	0.8349

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