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RISK AVERSION, MULTIVARIATE PROXIES AND THE BEHAVIOR OF ASSET RETURNS*

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An asset pricing model with constant relative risk aversion (CRRA) is tested with data from Sweden for the period 1977–1990. As a proxy for consumption growth, we employ a return based mimicking portfolio. We find that significant structural shifts in the model parameters occur between 1977–83 and 1984–90. The results indicate that the goodness-of-fit for the model in the two subperiods is fairly good and the estimates of CRRA seem reasonable. However, contrary to the cross-sectional implications of the model, CRRAs implied from individual assets differ a lot from each other. (JEL G12)

1. Introduction

Following the seminal work of Lucas (1978) and Hansen and Singleton (1982), consumption based rational expectations models for pricing risky assets have been the subject of extensive empirical testing. In these models, equilibrium expected returns on assets are determined by the expected growth in aggregate consumption and the covariability between returns and the intertemporal marginal rate of substitution of consumption. Thus, the fluctuations in asset returns should be consistent with swings in consumption growth for economically plausible levels of risk aversion. However, the empirical performance of the consumption based capital asset model (CCAPM) derived under time-additive, isoelastic, utility has been unsatisfactory. The model seems to fit the data poorly and the implied estimates of relative risk aversion are often implausible (e.g., see Hansen and Singleton, 1982, Ferson and Constantinides, 1991, Ferson, 1989, Bossaerts and Dammon, 1989, and Ferson and Harvey, 1992). In addition, the classic CCAPM also seems to be unable to explain the magnitude of the risk premium on equities for any reasonable parameter values (e.g., see Mehra and Prescott, 1985).

However, as emphasized in e.g. Ferson and Constantinides (1991) and Ferson and Harvey (1992), the measurement of consumption is quite problematic. Although several extensions of the basic model have been proposed to reconcile the apparent smoothness in consumption series with the high volatility in asset returns (e.g., time-nonseparable utility in Epstein and Zin, 1991, durability in Ferson and Constantinides, 1991, and seasonality in Ferson and Harvey, 1992), the analytical tractability of the classic model makes it attractive. Thus, to circumvent the problems associated with reported series on consumption without loosing the tractability of the model, investigations of asset return based versions of the classic intertemporal model appear worthwhile. Indeed, Breeden, Gibbons and Litzenberger (1989) documented an improved data fit for a linear CCAPM using an asset return based mimicking portfolio for consumption growth. Moreover, Brown and Gibbons (1985) and Ferson (1989) presented more reasonable estimates of relative risk aversion (RRA) with univariate stock market prox-

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ies for consumption. For instance, Brown and Gibbons (1985) found support for logarithmic utility, while Ferson (1989) reported RRAs on the order of 7.0 to 8.4. Finally, when investigating a more general conditional multiperiod asset pricing model in which asset returns are determined both by the growth in aggregate consumption and in wealth (in empirical tests a stock market index), Epstein and Zin (1991) documented that the latter source of risk almost completely dominated the former in the pricing of stocks and bonds.

This paper empirically tests restrictions on the asset prices imposed by an intertemporal CAPM with a multivariate, return based, pricing benchmark. Swedish data is used. We extend previous empirical analyses of return oriented nonlinear Euler equations (e.g., see Ferson, 1989) by considering a multi-factor benchmark and by conducting some additional tests. In developing the asset pricing model examined, we adopt the working hypothesis that the Swedish capital markets are partially segmented. Indeed, the in- and outflows of capital were regulated in Sweden during the period examined and results by Bergström, Rydqvist and Sellin (1993) render support for this approach.

This paper is organized as follows. Section 2 presents the model, Section 3 examines the data and Section 4 presents the empirical results. Section 5 contains the conclusions.

2. *Testing the asset pricing model*

2.1 The pricing model

The basic setting is a real, competitive, frictionless, discrete time multiperiod economy with complete markets and a single nonstorable consumption good. It is populated by optimizing agents with identical beliefs and time- and state-separable isoelastic (i.e. constant relative risk aversion or CRRA) utility of consumption. Following Hansen and Singleton (1982) and Lucas (1978), the intertemporal optimality condition for the representative consumer at all times t-1 can (by the "envelope condition") be stated as

$$(1) \qquad E\left[\rho\left[\frac{f_t W_t}{f_{t-1} W_{t-1}}\right]^{-\alpha} (1 + \mathbf{r}_t) \mid \Omega_{t-1}\right] = \mathbf{1}$$

where $m_t = \rho\left[\frac{f_t W_t}{f_{t-1} W_{t-1}}\right]^{-\alpha} = \rho\left[R_{ct}\right]^{-\alpha}$

is the pricing benchmark or the intertemporal marginal rate of substitution, W_t is wealth at time t, f_t is the marginal propensity to consume out of wealth at time t, α is the CRRA coefficient (or the inverse of the intertemporal elasticity of consumption substitution) of the representative agent, $\rho < 1$ is his time-preference parameter, $\mathbf{R}_t = (1 + \mathbf{r}_t)$ is a N + 1 vector of gross real asset returns over time t and Ω_{t-1} is the information set available to consumers at time t–1.

Risk averse or concave preferences imply $\alpha > 0$, whereas the myopic logarithmic utility investor has $\alpha = 1$. For more (less) risk averse investors than the log utility agent, $\alpha > (<) 1$. In equilibrium, the optimality condition in (1) must hold for all securities in the economy. Thus, all assets should provide the same expected real return (i.e., marginal rate of transformation) properly adjusted for the comovements between asset returns and the growth in the marginal utility of wealth. The higher α is, the lower m_t is and the higher are the expected returns on securities in equilibrium, ceteris paribus.

As the constrains on in- and outflow of capital have been quite stringent in the case of Sweden (e.g., stocks were divided into restricted and unrestricted classes with legal restrictions on the permitted proportion of unrestricted shares that were available to foreign investors during the period examined), we adopt the working hypothesis that Swedish asset markets are not perfectly integrated to world capital markets. On partially segmented markets, both local and global systematic risks are likely to influence asset prices in equilibrium (e.g., see Bergström, Rydqvist and Sellin, 1993). Based on the model in (1), the economic risks that are rewarded in equilibrium must covary with the marginal utility of wealth, $(f_t W_t)^{-\alpha}$. Thus, on one hand, from Merton (1973) and Cox, Ingersoll and Ross (1985b), the "market portfolio of domestic wealth" and the short-term interest rate are likely to be rewarded local factors. On the other, from global pricing models (e.g., see Adler and Dumas, 1983), risks associated with the "world market portfolio" are also prone to be priced.[1] Consequently, the growth in $f_t W_t$ is

[1] *Moreover, with deviations from purchasing power parity, real exchange rate risks are also likely to matter in the pricing of assets in globally integrated capital markets. In this case, exchange rate risks come into play since investors from different nations evaluate the expected real return on a given asset differently. This additional source of uncertainty could give rise to a "hedging demand" for portfolios whose returns are correlated with deviations from PPP.*

here mimicked with the real own-currency returns on the "local market portfolio", a local short-term bond and a proxy for "international wealth". However, there is no claim that these three variables uniquely capture the actual variation in consumption growth.

As a measure of national wealth, we use a value-weighted stock market index at time t (R_{smt}). To capture interest rate risk, we employ the holding period return on a one-month Treasury bill based on the rate available at time t (R_{ft}). In Ferson and Harvey (1991), these two domestic risk factors were found to be the most important sources of time-variation in expected returns on US assets. The results in Harvey (1991) indicate that world market risks are significant, though insufficient, determinants of Swedish stock returns. To account for international influences on local asset prices, we use the Standard and Poor index for US industrial stocks as a rough measure of global wealth at time t (R_{ust}). Of course, the SP industrials index is a very incomplete proxy for the "world market portfolio". However, it is should quite accurately reflect global wealth movements since the US stock market is historically the largest component e.g. in the Morgan Stanley Capital International (MSCI) world stock index. The US market stands for more than 30 % of the MSCI world index and the correlation between US and world equity returns is 86 % (e.g., see Harvey, 1991). The SP index is an exogenous measure of global influences while the "world index" is not (e.g., the MSCI world equity index contains the Swedish stock market with a weight of 1 %). Since we take the perspective of the Swedish investor, the dollar denominated returns from t–1 to t are converted to Swedish crowns (SEK) with the SEK/USD exchange rate at time t.[2] Employing the resulting mimicking portfolio $R_{ct} = (1 - x^1 - x^2)R_{smt} + x^1R_{ust} + x^2R_{ft}$, in (1) yields

$$(2) \quad E[\kappa[(1 - x^1 - x^2)R_{smt} + x^1R_{ust} + x^2R_{ft}]^{-\alpha}\mathbf{R}_t|\Omega_{t-1}] = \mathbf{1}$$

where κ is a utility parameter.[3] However, it is no longer valid to interpret the constant κ as measure of pure time preference ρ (although ρ is a function of κ).

The three-factor pricing model in (2) is fairly general. For instance, it imposes no explicit distributional assumptions on asset returns (e.g., multivariate normality). On the other hand, if consumption growth rates are independently and identically distributed, the model in (2) can be viewed as a static version of the intertemporal model (1) with the "true market portfolio" being a combination of the three reference assets (e.g., see Rubinstein, 1976, and Shanken, 1987). However, in essence, the model in (2) implies that a portfolio of the three marketable factors is generalized efficient. This means that the reference portfolio is efficient when allowing for taste induced tradeoffs over "all return moments", not just e.g. over the two first moments (e.g., see Ingersoll, 1987).

2.2 *Testable restrictions imposed by the pricing model*

To test the model in (2), we have to specify a subset of the available information Ω_{t-1}. This so called instrument set $\mathbf{F}_{t-1}$ is a h × 1 vector which also includes the unit vector that places unconditional restrictions on the return moments. Consistent with isoelastic tastes, the utility parameters (κ, α) should be time-invariant. In empirical tests, the 1 × 2 vector of spanning portfolio weights x is also restricted to be constant over time.[4] Thus, the set of constants to be implied when estimating the model in (2) consists of (κ, α, x).

[2] *Hence, the total risk in the global factor R_{ust} is composed of "international wealth risk", "exchange rate risk" and "uncertainty about changes in the local price level". In (2), the consumers are "incrementally unconcerned" about exchange rate changes, i.e. they have no desire or possibility to hold separate "hedge portfolios" against adverse movements in exchange rates.*

[3] *In (2), we formulate an exact representation of m_t. In other words, the return on the mimicking portfolio is implictly assumed to be perfectly correlated with the growth in consumption. On one hand, while this assumption may seem restrictive, it is standard in testing return based versions of the model in (1) [e.g., see Brown and Gibbons (1985) and Ferson (1989)]. On the other hand, the assumption of "a perfect proxy" is consistent with the assumption of (effectively) complete markets that also underlies the model in (1).*

[4] *For the model in (2), the spanning weights for time t generally depend on Ω_{t-1} whereas our assumption of constant weights restricts the latent mimicking portfolio to be time-invariant. While the use of time-varying mimicking weights would in principle be preferable, many additional parameters would have to be implied in model estimation. Indeed, this could rather severly deteriorate the small sam-*

For ex post real returns, innovations from the model in (2) can now be stated as $\mathbf{u}_t = \kappa[(1 - x^1 - x^2)R_{smt} + x^1R_{ust} + x^2R_{ft}]^{-\alpha}\mathbf{R}_t - \mathbf{1}$, where $\mathbf{u}_t$ is a N + 1 vector of disturbance terms. Rational expectations imply that $E[\mathbf{u}_t | \mathbf{F}_{t-1}] = \mathbf{0}$. A subset of this general restriction corresponding to linear expectations (orthogonality between $\mathbf{u}_t$ and $\mathbf{F}_{t-1}$, $E[\mathbf{u}_t \otimes \mathbf{F}_{t-1}] = \mathbf{0}$, where $\otimes$ is the Kronecker product) is the restriction studied by Hansen and Singleton (1982) and many others thereafter. As in Hansen and Singleton (1982), we test for model error orthogonality or H_0: $E[\mathbf{u}_t \otimes \mathbf{F}_{t-1}] = \mathbf{0}$ with the generalized methods of moments (GMM). Unbiasedness in model errors is tested with a multivariate, asymptotically χ^2-distributed goodness-of-fit test statistic developed in Hansen (1982). The test statistic is derived from the minimized value of a quadratic norm of the deviations from the sample counterpart of $E[\mathbf{u}_t \otimes \mathbf{F}_{t-1}] = \mathbf{0}$ evaluated at the minimizing values of (κ, α, x). There are as many degrees of freedom as there are moment restrictions or dim $(\mathbf{u}_t \otimes \mathbf{F}_{t-1})$ to be satisfied in excess of the number of implied model parameters in the set (κ, α, x). As in e.g. Ferson and Foerster (1992), we use the iterated approach for estimating the model to achieve more accurate results.

Under the null, the expected return on the reference portfolio should be mirrored in the anticipated returns on all assets: evaluated at the calibrating values for (κ, α, x), changes in the intertemporal marginal rate of substitution should at all times be consistent with the revisions in expected asset returns (adjusted for the movements in covariances). On the other hand, the vague null $E[\mathbf{u}_t \otimes \mathbf{F}_{t-1}] = \mathbf{0}$ imposes both cross-sectional and time-series restrictions. In empirical tests, the null is a joint hypothesis of (a) optimality, (b) the taste specification (contant utility parameters (κ, α)) and (c) the assumption of constant portfolio weights x (e.g., see the criticism in Fama, 1991). Of course, restriction (a) is economically fundamental, whereas restrictions (b)–(c) are mainly of technical nature. Hence, to examine if the model fails because the implied intertemporal marginal rates of substitution differ across assets or because the model parameters (κ, α, x) are not structural, more specific alternative hypotheses are needed.

To assess if the model parameters (κ, α, x) are structural or stable across time, we investigate the out-of-sample predictions of the model. That is, we evaluate how well the moment restrictions in one sample (2) are satisfied with the parameters estimated in a disjoint sample (1). The model null implied by the time-series restrictions is H_0: $E[\mathbf{u}_{1t}(\kappa, \alpha, x) \otimes \mathbf{F}_{t-1}] = E[\mathbf{u}_{2t}(\kappa, \alpha, x) \otimes \mathbf{F}_{t-1}] = \mathbf{0}$, whereas the alternative hypothesis is that the model is valid in sample (1) but fails in sample (2) or H_1: $E[\mathbf{u}_{1t}(\kappa, \alpha, x) \otimes \mathbf{F}_{t-1}] = \mathbf{0}$, $E[\mathbf{u}_{2t}(\kappa, \alpha, x) \otimes \mathbf{F}_{t-1}] \neq \mathbf{0}$. Stability of the model parameters is tested with an asymptotically χ^2-distributed predictive test statistic developed and applied in Ghysels and Hall (1990a, b). For the predictive test statistic, the degrees of freedom equal the number of moment restrictions in the second sample or dim $(\mathbf{u}_{2t}(\kappa, \alpha, x) \otimes \mathbf{F}_{t-1})$.

When testing the cross-sectional restriction that the marginal rates of transformation or risk-return tradeoffs ("prices of risk") really are the same for all assets as implied by the model, we allow the CRRAs to be asset specific constants. Hence, α is an unrestricted vector of "prices of risks" for the assets. The null derived from the cross-sectional restrictions of the model is H_0: $\tilde{\alpha} = \alpha\mathbf{1}$. Finally, to see if other sources of risk than (nonlinear transformations of) "domestic wealth" contribute to the movements and cross-sectional dispersion in expected returns, we test restrictions on the mimicking weights in x. These evaluations correspond to the widely employed tests of which risk factors are rewarded in linear multi-beta models. For instance, $x^1 = 0$ is broadly consistent with a complete segmentation of Swedish capital markets from international asset markets (i.e. capital outflows are zero and all local assets are held by domestic investors). This is at odds with models of international portfolio choice and it suggests that all Swedish assets (i.e. restricted and unrestrictred alike) are priced based on local systematic risks. The validity of nested restrictions is investigated with a generalized likelihood ratio test outlined in Newey and West (1987). The asymptotically χ^2-distribut-

ple properties of the GMM estimator in testing the Euler equations and thus nullify the gains obtained from relaxing the assumption of time-invariant portfolio weights [e.g., see Ferson and Foerster (1992) and Tauchen (1986)]. However, we examine the empirical validity of "constant portfolio weights" when conducting joint tests of the stability of the model parameters.

ed test statistic is calculated as the difference in the goodness-of-fit test statistics between the restricted and unrestricted models using the optimal weighting matrix for the latter. The degrees of freedom for the test are equal to the number of restricted parameters.

3. Summary statistics

The time period examined is January 1977 to December 1990. Major changes in the economic environment have occurred during the sample period (e.g., the deregulation of financial markets was initiated in 1983, the Swedish central bank started to pursue a low inflation monetary policy in the mid 1980s and global influences in the Swedish economy have increased due to growth in foreign trade and the deregulation of capital movements in and out of Sweden). Based on this rather vague prior, it seems that the mid-point of the sample or January 1984 is a reasonable candidate event for revealing potential structural changes in the data. Therefore, we will also examine the data using a breakdown of the entire period into two 7-year subperiods, i.e. 1977–83 and 1984–90.[5]

3.1 Asset returns

Monthly, continuously compounded, holding period returns on stocks and bonds are used in the model estimations. All asset returns are measured in real terms: the log change in the consumer price index (CPI) is subtracted from the nominal returns. Eight industry sorted stock portfolios and two long-term bonds are used as test assets, while the multivariate pricing benchmark consists of the returns on a value-weighted Swedish stock market index, the Standard and Poor's index for US industrial stocks, and the Swedish one-month Treasury bill. All returns are computed in Swedish crowns (SEK). Data on the Swedish long-term bond yields, inflation rates and all used international variables are taken from "OECD: Main Economic Indicators". The overall quality of the Swedish interest rate data is not, however, very high prior to the early 1980s since the secondary markets for bonds were not well developed at the time.

The industry portfolios are equally-weighted and they are updated each month based on the information available at the beginning of the month. The industry classification is updated annually but the portfolio forming strategy only requires that a given share is listed for less than two consecutive months (i.e. changes in listings are accounted for on a monthly basis). One stock series, the most actively traded one, was chosen for each firm. The set of test assets contains both restricted and unrestricted stocks which is also consistent with the use of local and global pricing factors. Overall, the selection criteria is not stringent and the data should not be subject to any significant survivorship bias. The considered industrial sectors are: (1) banking and finance, (2) insurance and investment, (3) trade and transportation, (4) consumer goods and services, (5) construction and real estate, (6) forestry and paper, (7) basic industries and utilities and (8) capital goods and durables. All domestic stock market return series are constructed at the Swedish School of Economics and Business Administration. The local stock returns are adjusted for splits, issues and dividends. The bond data consists of a 10-year government and a 5-year corporate bond, respectively. Due to unavailability of quoted prices, we have to estimate the monthly holding period returns on bonds. Approximate bond prices are computed from the yields under the empirically reasonable assumptions that (a) all bonds are of bullet type (i.e. principals are paid off in full on maturity) and (b) the coupon rate is a moving average of the bond yields in the previous 12 months.[6]

Some descriptive statistics for the real return data for the 1977–90 period are reported in Table 1. As seen in Table 1, the dispersion in mean

[5] *Although the subperiod split used broadly conforms to the external changes in the market environment, it is evident that it does not match them perfectly. Indeed, one of the problems with handling the changes that have taken place in Sweden during this time period is that the development has been gradual without a single breakpoint. On the other hand, subperiod analyses are certainly warranted for general robustness and stability reasons even without explicit reference to the environmental changes.*

[6] *A limited sensitivity analysis using alternative coupon rates produced very similar bond return series as those used in the tests. Hence, despite the fact that the employed bond series are only approximate, changing the unknown coupon rate does not seem to materially affect the obtained bond returns.*

Table 1

Summary statistics for monthly asset returns (in real terms).

Variable	Mean	Std.dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_{12}
		1977/1 – 1990/12:				
Industry 1	0.8463	5.4778	0.340**	0.138	0.069	–0.047
Industry 2	0.7508	6.4253	0.263**	0.052	–0.042	0.017
Industry 3	0.2580	6.3777	0.308**	0.140	–0.016	0.046
Industry 4	0.3579	6.1436	0.253**	0.103	–0.031	0.050
Industry 5	0.8858	6.2412	0.126	0.048	0.059	–0.010
Industry 6	0.5291	8.1289	0.182*	0.104	–0.004	0.101
Industry 7	0.7502	6.1428	0.226*	0.023	–0.038	0.075
Industry 8	0.6808	5.9788	0.211*	0.021	–0.027	0.107
Gvt. bond	0.1710	2.1163	0.232**	0.013	0.024	0.010
Corp. bond	–0.1283	1.3380	0.239**	0.224*	0.067	0.214*
Market return	0.8198	5.9572	0.249**	0.059	–0.028	–0.006
SP Industrials return	0.3429	4.5670	0.277**	–0.107	–0.115	–0.014
T-bill return	0.2778	0.6211	0.216*	0.150	0.216*	0.430**
Inflation rate (CPI)	0.6466	0.5828	0.123	0.051	0.123	0.364**

Note: The asymptotic standard error for autocorrelation by $\sqrt{1/(\tau-1)}$, where τ is the number of observations and 1 is the lag length, under the hypothesis of a stationary normal distribution. *denotes significant at the 5 %-level and** denotes significant at the 1 %-level. Monthly returns are multiplied 100. All returns are continuously compounded (log) and they are expressed in real terms (the rate of inflation (CPI) is subtracted from raw returns).

returns and standard deviations is quite substantial across the assets. Indeed, this is useful for obtaining more powerful model tests. Among stocks, banking and finance has the lowest standard deviation in returns whereas the returns from forestry and paper have the highest volatility. Not surprisingly, the two bonds are the least volatile test assets and they also have the lowest mean returns (the corporate bond has a negative average real return). Overall, the stocks carry a notable "real return premium" over the bonds in the present sample.

As seen in Table 1, The Swedish stock index has also performed well. For the entire 1977–90 period, the Swedish and US markets have mean returns in SEK of 0.82 % and 0.34 %, respectively. The average real return on one-month T-bill is also positive: the average real return is 0.28 % per month or roughly 3.3 % p.a. Finally, there is typically quite strong first-order serial correlation in Swedish real asset returns.[7]

3.2 The conditioning variables

The forecasting instruments in $\mathbf{F}_{t-1}$ known at time $t-1$ should be chosen in such a way that the information is able to detect deviations from the asset pricing model. In principle, useful instruments should be able to capture changing expectations about asset returns and the growth in marginal utility of wealth. To represent changing expectations, we employ the following, easily observable, instrumental variables: (a) the change in the nominal 10-year government bond yield (Δy), (b) the change in the

[7] *Frennberg and Hansson (1993) have analyzed the autocorrelation in Swedish stock returns. Their findings clearly suggest that the strong autocorrelation in monthly asset returns is not induced by infrequent trading. In other words, the strong serial dependence in stock returns is thus likely to be real. This suggests that past market returns are potentially useful forecasting variables.*

Table 2

Summary statistics for the monthly instrumental variable data.

Variable	Mean	Std.dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_{12}
		1977/1 – 1990/12:				
Δy_{t-1}	0.0189	0.3310	0.185*	–0.008	0.013	–0.028
Δr_{t-1}	0.0249	0.4200	0.336**	0.076	0.158*	0.144
$r(m)_{t-1}$	1.4656	5.9015	0.256**	0.065	–0.007	0.002
$r(sp)_{t-1}$	0.7036	3.7423	0.303**	–0.071	–0.131	–0.127
ip_{t-1}	0.9265	5.5667	0.548**	0.372**	0.466**	–0.096

Correlations:

	Δy_{t-1}	Δr_{t-1}	$r(m)_{t-1}$	$r(sp)_{t-1}$	ip_{t-1}
Δy_{t-1}	1.000				
Δr_{t-1}	0.116	1.000			
$r(m)_{t-1}$	–0.166*	–0.078	1.000		
$r(sp)_{t-1}$	–0.112	–0.099	0.381**	1.000	
ip_{t-1}	0.047	0.155*	–0.017	–0.060	1.000

Note: The asymptotic standard error for autocorrelation by $\sqrt{1/(\tau-1)}$, where τ is the number of observations and 1 is the lag length, under the hypothesis of a stationary normal distribution. *denotes significant at the 5 %-level and** denotes significant at the 1 %-level.

nominal German 3-month Treasury bill (Δr), (c) the raw return on the Swedish value-weighted stock index (r(m)), (d) the raw USD return on the SP industrials index (r(sp)) and (e) the 12-month log growth rate in Swedish industrial production (ip). Our ip measure is calculated as in Chen (1991) and it broadly measures "cyclical components" in industrial production net of a stochastic trend and seasonal effects. To ensure that the conditioning information used was available to agents in calculating returns over period t, all market variables are used with a one-month lag and the ip variable with a two-month lag to account for reporting delays (all variables should be available at time t–1).

The motivations for considering these conditioning variables are straightforward. First, motivated by the forcing variable structure in the empirical model, we use wealth and interest rate related instrumental variables (a) – (d) to capture movements in conditional asset return moments. As Sweden is a small open enonomy and we assume partial segmentation, we consider both domestic and international instruments. The wealth variables (c) – (d) are proxies for the future health of the domestic and world economies, respectively, whereas changes in the interest rates (a) – (b) presumably capture movements in the productivity of domestic and world investments and inflation expectations, respectively (e.g., see Chen, 1991, and Breeden, 1986). Reasonably enough, the US stock market is a leading indicator of the global economy, whereas German monetary policy is the most influential determinant of interest rates in Europe. Second, following Chen (1991), we expect actual RRA to be negatively related to the current health of the economy as proxied by (e). Indeed, Ferson and Merrick (1987) found that asset returns behave differently during recessions and non-recessions, while Balvers, Cosimano and MacDonald (1990), Chen (1991) and Daniel and Torous (1991) reported that business cycle measures of industrial production are able to predict asset returns. In addition, the forecasting power of market based variables similar to ours is documented in e.g. Campbell (1987), Harvey (1989) and Shanken (1990). To summarize, while (e) is foremost a measure of the prevailing economic conditions, the market

Table 3

Predictive tests of structural stability of the Euler equations with a multivariate proxy.

	Instruments					
	All		Domestic		International	
	χ^2	*p*-value	χ^2	*p*-value	χ^2	*p*-value
	Sample split: 1977–83 and 1984–90					
8 industry portfolios and 2 bonds	67.91	0.226	59.73*	0.023	49.36*	0.014
8 industry portfolios	56.37	0.190	56.47**	0.005	39.29*	0.025
2 bonds	21.10*	0.049	15.52*	0.050	13.71*	0.033

Note: The P-value is the probability under the null hypothesis of realizing a higher χ^2 value than the observed sample value; $^+$denotes significant at the 10 %-level,* significant at the 5 %-level and** significant at the 1 %-level. All instruments consist of $\{1, \Delta y, \Delta r, r(m), r(sp), ip\}_{t-1}$, the domestic instruments of $\{1, \Delta y, r(m), ip\}_{t-1}$ and the international instruments of $\{1, \Delta r, r(sp)\}_{t-1}$.

variables (a) – (d) primarily measure anticipations about future economic activity and inflation. Table 2 gives summary statistics for the conditioning variables.

As seen in Table 2, the autocorrelation coefficients for the variables are quite moderate with the exception of ip. However, although the serial correlation coefficients for ip are admittedly high, they appear to decay toward zero at longer lags which is suggestive of stationarity.[8] Also, the pairwise correlations among the forecasting variables are usually low, although there are some significant correlations (e.g., among the Swedish and US market returns). However, the pairwise correlations tend to be higher for the 1984–90 period. For instance, in the 1984–90 period, the strongest correlation is not surprisingly among r(m) and (r(sp): it is as high as 0.558. In addition, the correlation between r(m) and Δy is fairly high: 0.252 (not reported). Indeed, these results are consistent with the presumed effects of the economic changes that have occurred in Sweden. In summary, the employed forecasting variables appear to quite comfortably satisfy the requirements for stationarity and they are not mutually very highly correlated.

[8] *In the entire 1977–90 period, an augmented Dickey-Fuller test for a unit root in ip produces a F-value of 15.19 with (2, 162) degrees of freedom. Hence, there is support for the stationarity of ip at a less than 1 % level of significance.*

4. Empirical evidence

4.1 Structural stability of the model

In this section, we examine the stability of the model parameters. In particular, the theory predicts (or assumes) that the utility parameters (ρ, α) are constant through time. Also, to test the model with a multivariate proxy for consumption, we assume that the reference portfolio is time-invariant or that (x^1, x^2) are constants. Hence, to infer something about optimality, we should attempt to control for violations of constant parameters in (κ, α, x). As discussed previously, we have a rough prior knowledge about the regime switch: the structural stability (model validity) is investigated with a break at January 1984. Ghysels and Hall (1990b) stress the importance of using informationally efficient instruments with respect to the break prior for discriminating between the null and structural instability when applying the predictive test. Indeed, since the power of the predictive test appears somewhat sensitive to the dimension and nature of the moment restrictions imposed on the estimation (see Ghysels and Hall, 1990b, and Epstein and Zin; 1991), we consider here three instrument sets: (a) all instruments $\{1, \Delta y, \Delta r, r(m), r(sp), ip\}_{t-1}$, (b) domestic instruments $\{1, \Delta y, r(m), ip\}_{t-1}$ and, finally, (c) international instruments $\{1, \Delta r, r(sp)\}_{t-1}$. Referring to the discussion on the deregulation of the financial markets, deregulated capital flows and the increased degree of global influences in the Swedish economy, this infor-

Table 4

Likelihood ratio tests of the implied reference portfolio composition
($R_{ct} = (1 - x^1 - x^2)R_{smt} + x^1R_{ust} + x^2R_{ft}$).

Test of	$x^1 = 0$		$x^2 = 0$		$x^1 = x^2 = 0$	
	χ^2	*p*-value	χ^2	*p*-value	χ^2	*p*-value
1977/1 – 1983/12:						
8 industries and 2 bonds	10.57**	0.001	22.45**	0.000	28.19**	0.000
8 industries	18.34**	0.000	9.07**	0.003	19.81**	0.000
2 bonds	0.19	0.666	14.81**	0.000	15.44**	0.000
1984/1 – 1990/12:						
8 industries and 2 bonds	20.29**	0.000	8.48**	0.004	21.08**	0.000
8 industries	15.22**	0.000	5.08*	0.024	15.54**	0.000
2 bonds	0.40	0.529	7.33**	0.007	9.54**	0.008

Note: The P-value is the probability under the null hypothesis of realizing a higher χ^2 value than the observed sample value; +denotes significant at the 10 %-level, *significant at the 5 %-level and** significant at the 1 %-level. The instrument set consists of $\{1, \Delta y, \Delta r, r(m), r(sp), ip\}_{t-1}$.

mation partition seems to be reasonable. The results from predictive tests of the model are reported in Table 3.

Not surprisingly, the results reported in Table 3 suggest that the model is not valid for the 1977–90 period. Although rejections of parameter stability are not overwhelming, it seems apparent that the model parameters (κ, α, x). are not invariant across time and "policy regimes". With all assets, the *p*-values for the null being true range from 22.6% to 1.4 %. However, the tests are generally more powerful when using only returns from stock portfolios and bonds, respectively. Interestingly, international instruments seem slightly more able to reject the null with bonds only while domestic instruments are more successful in detecting parameter instability with industry portfolios. Apparently, the tests are weak for industry portfolios when augmenting the domestic forecasting variables with global ones. This suggests that using all instruments jointly does not permit discrimination among the hypotheses for large moment systems due to problems with the small sample properties of the test statistic with "many degrees of freedom" or the potential use of "redundant information" which reduces the power of the tests.

To summarize, there is evidence of structural instability in the model parameters with confidence on the order of 77.4 % to 99.5 %. Given the profoundness of the changes that have taken place in the economic environment over the 1977–90 period, these rejections can be regarded as intuitive. However, as pointed out in Ghysels and Hall (1990b), it is not straightforward to interpret rejections of stability as implying that model parameters have changed e.g. during or after the deregulation of the financial markets. Rather, the evidence suggests that, consistent with the changes in the economic environment, the proposed consumer preferences are misspecified and/or the laws of motion for returns and consumption are nonstationary thereby invalidating the model for the entire period.

4.2 Tests of model fit and the preferences

This section presents empirical results for eight industry portfolios and two bonds using data for the nonoverlapping periods 1977–83 and 1984–90. The empirical results from testing various restrictions are summarized in Tables 4 and 5, respectively.

In Table 4, we provide some initial information about the model by reporting the results from examining the implied composition of the reference portfolio, $R_{ct} = (1 - x^1 - x^2)R_{smt} + x^1R_{ust} + x^2R_{ft}$. Generalized "likelihood ratio" tests of restrictions on the pricing proxy, namely $x^1 = 0$ and/or $x^2 = 0$, are presented. In essence, these analyses serve the purpose of checking if international stock returns and domestic bond returns are incrementally rewarded (i.e. affect the cross-sectional dispersion in expected asset returns) in a nonlinear pricing model in addition to the Swedish stock market return. If the cross-sectionally restricted weights x are jointly zero (i.e. $x^1 = x^2 = 0$), then (a non linear transformation of) the local stock market factor alone is sufficient for describing the variation in expected returns. This corresponds to the case examined by e.g. Brown and Giggons (1985) and Ferson (1989). Also, if $x^1 = 0$, the evidence suggests that Swedish capital markets are completely segmented.

The results in Table 4 render support for the multivariate nature of the proxy: both the US market factor and Swedish bill factor appear to receive a non zero weight in the reference portfolio both individually and jointly.

As seen in Table 4, the local bill factor appears to be an important component both in the pricing of bonds and stocks: in the 1977–83 period, the *p*-values for a zero weight range from 0.3 % to less than 0.1 % whereas the *p*-values for the 1984–90 period are on the order of 2.4 % to 0.4% across tests. Thus, perhaps surprisingly, the domestic bill return factor seems more significant in the 1977–83 period. However, this finding could emerge from the fact that the stock market better impounds interest rate variation in the 1984–90 period. This conjecture seems consistent with the stronger correlation between bond yield changes and stock market returns for the 1984–90 period (see section 3.2).

Not surprisingly, the international stock market factor seems to contribute significantly to the pricing of stocks only, whereas it seems to lack any axplanatory power for the variation in bond returns. For the industry portfolios, the influence of the US stock market factor appears to be highly significant for both periods (the *p*-values are less than 0.1 %). Hence, although capital flows into and out from Sweden were fairly stringently regulated particularly in the 1977–83 period, foreign influences seem to be present in Swedish stock returns also for this period. On the whole, our results are suggestive of some degree of global stock market integration for Sweden and they are consistent with the findings in Harvey (1991) and Ferson and Harvey (1993).

However, joint tests reveal that the international stock market factor and the domestic bill factor are "partial substitutes" in the pricing of stocks. Although being significant at usual levels, the joint χ^2 test statistics for industry portfolios are clearly lower than the sum of the individual test statistics for both periods. In contrast, tests of the joint factor restriction are slightly stronger for bonds than the partial tests for both periods. In all, the own-currency US stock market and the Swedish bill returns seem to explain the common variation in asset returns in addition to the Swedish stock market return.

In Table 5, we display information about model performance. First, the calibrating estimates for the concavity parameter α are reported. Second, results from testing the logarithmic utility hypothesis or $\alpha = 1$ are displayed. In the present context, this test can be view as an indirect way of assessing whether agents behave in a myopic fashion. More specifically, log utility restricts the expected returns on all assets relative to the benchmark return to move perfectly together over time (more precisely, for $\alpha = 1$, we have

$$E[\kappa[\frac{R_{jt}}{(1-x^1-x^2)R_{smt}+x^1R_{ust}+x^2R_{ft}}|F_{t-1}]=1$$

for all j assets). In this static and nonlinear (or "multi-moment") logarithmic utility model, the "aggregate wealth" is represented by a 3-dimensional proxy of the reference assets. Third, the fit of the orthogonality conditions of the model to the data (goodness-of-fit) is reported.

The cross-sectionally restricted estimates of relative risk aversion reported in Table 5 appear quite plausible both in sign and magnitude. Consistent with risk averse tastes, all concavity coefficients are positive and more than two standard errors away from zero. In the 1977–83 (1984–90) period, the risk aversion estimates for the representative agent lie in the range from 2.269 to 3.448 (3.396 to 5.202) across the asset classes. Interestingly, industry portfolios yield higher estimates than bonds for both periods. The evidence also indicates that logarithmic utility or "myopia" is consistently at odds with the data: the *p*-values for log utility range from 6.2 % to less than 0.1 % across tests. Hence, while the representative agent seems to be less

Table 5

Tests of restrictions imposed by the Euler conditions.

Test of			$\alpha = 1$ (log utility)		Goodness-of-fit	
	α	α $(s.e.)$'	χ^2	p-value	χ^2	p-value
		1977/1 – 1983/12:				
8 industries and 2 bonds	2.269	0.1496	51.60**	0.000	71.27$^+$	0.082
8 industries	3.448	0.9960	6.23*	0.013	43.22	0.505
2 bonds	2.441	0.6234	5.54*	0.019	14.18$^+$	0.077
		1984/1 – 1990/12:				
8 industries and 2 bonds	5.078	1.4878	8.02**	0.005	52.88	0.594
8 industries	5.202	1.8989	5.18*	0.023	44.34	0.457
2 bonds	3.396	1.3098	3.49$^+$	0.062	10.20	0.251

Note: The P-value is the probability under the null hypothesis of realizing a higher χ^2 value than the observed sample value; $^+$denotes significant at the 10 %-level,* significant at the 5 %-level and** significant at the 1 %-level. The instrument set consists of $\{1, \Delta y, \Delta r, r(m), r(sp), ip\}_{t-1}$. The standard errors for the concavity parameter estimates are corrected upwards by the "traditional adjustment" described in Ferson and Foerster (1992). The adjustment factor used for s.e.: equals $\sqrt{84/80}$.

risk tolerant in the more volatile 1984–90 period, the evidence suggests that he is more risk averse than the log investor for both periods. Heuristically, this suggests that the "income effect" dominates in aggregate preferences (e.g., "a higher expected or less uncertain future wealth increases current consumption due to intertemporal consumption smoothing"). On the other hand, our estimates of CRRA are lower than those documented in e.g. Ferson (1989).

As seen in Table 5, based on examining the overidentifying restrictions, the model seems to perform well although there is evidence against its restrictions at the 10 % significance level in the first subperiod. Overall, the tests with bonds only are clearly stronger than the tests with stocks alone. With all test assets, the rejection of the model for the 1977–83 period seems to be driven by the poor fit of the moment restrictions for the bonds (from Table 5, the p-value for bonds in the first period is 7.7 %). Across the test asset groupings, the p-values from the goodness-of-fit tests range from 50.5 % to 8.2 % and from 59.4 % to 25.1 % for the 1977–83 and 1984–90 periods, respectively.

Finally, to supplement the multivariate goodness-of-fit tests, we regressed the fitted model errors for each test asset on the instrumental variables. These regressions are conducted for the model errors obtained from the estimations for the two asset classes separately. In the 1977–83 period, the model fit for consumer goods and services (#4) and forestry and paper (#6) are the worst (with adjusted R^2:s of 11.89 % and 7.47 %, respectively) among the eight industry portfolios. Likewise, in the 1984–90 period, forestry and paper exhibits the poorest model fit (with an adjusted R^2 of 2.58 %). Indeed, it appears quite plausible that global risk factors are generally more important for forestry and paper than for other industries due to its strong economic reliance on foreign markets. In testing the model with bonds only, the corporate bond exhibits the worst fit to the model: for the 1977–83 (1984–90) period, the explained proportion of the variation in model errors is 3.02 % (12.97 %). Thus, although the multivariate χ^2 tests are fairly weak, significant departures from unbiasedness in the model innovations appear to be present for some assets which is at variance with the model's implications.

4.3 Tests of cross-sectional restrictions

Table 6 presents the results from examining the validity of the cross-sectional restriction of an equal relative risk aversion coefficient across

Table 6

Likelihood ratio tests of cross sectional restrictions imposed by the Euler conditions.

Test of	$\tilde{\alpha} = \alpha 1$	
	χ^2	*p*-value
1977/1 – 1983/12:		
8 industries and 2 bonds	339.99**	0.000
8 industries	41.53**	0.000
2 bonds	10.02**	0.002
1984/1 – 1990/12:		
8 industries and 2 bonds	77.06**	0.000
8 industries	33.93**	0.000
2 bonds	7.34**	0.007

Note: The P-value is the probability under the null hypothesis of realizing a higher χ^2 value than the observed sample value; +denotes significant at the 10 %-level,* significant at the 5 %-level and** significant at the 1 %-level. The instrument set consists of $\{1, \Delta y, \Delta r, r(m), r(sp), ip\}_{t-1}$.

the assets (H_0: $\tilde{\alpha} = \alpha \mathbf{1}$). Indeed, while the model performed well in the goodness-of-fit tests, the implied asset specific "prices of risk" seem to display significant dispersion which is in conflict with the implications of the model. In fact, the concavity parameters are rather different even across the industry portfolios and the bonds, respectively, let alone across the two asset classes. Thus, an equal "price of risk" across the assets is rejected for both periods.

In the 1977–83 period, the implied asset specific relative risk aversion parameters vary from 0.542 to 1.216 for the industry portfolios while being 0.275 and 0.055 for the government and corporate bonds, respectively. Among industry portfolios, the lowest CRRA is reported for banking and finance (0.542) and the highest CRRA for forestry and paper (1.216). Interestingly, the former sector also has the lowest and the latter the highest dispersion in returns. Similarly, government bond returns have a larger variance than corporate bond returns (see Table 1). This suggests that there is some relationship between ex post return volatilities and the implied "prices of risk". Indeed, an equal CRRA across the assets is violated in all tests at the 1 % level of significance.

Overall, the data fit is more reasonable for all assets individually and jointly in the 1984–90 period. Although the equality of the "prices of risk" within and across asset categories is rejected at the 1 % level, the corresponding χ^2 test statistics are now lower than for the 1977–83 period. In the 1984–90 period, the relative risk aversion parameters are generally also higher than for the 1977–83 period. As noted earlier, this seems fairly intuitive since the 1984–90 period was characterized by more turbulent asset markets than the 1977–83 period. For industry portfolios, the CRRA estimates are on the order of 4.357 to 6.299. The highest CRRA is documented for banking and finance, while consumer goods and services has the lowest implied CRRA. For the government and corporate bonds, the fitted "prices of risk" are 4.176 and 3.715, respectively. To interpret the cross-sectional dispersion in the implied CRRAs, we evidently have to rely on heterogeneity in tastes and the existence of out-of-model "preferred asset habitats".[9]

[9] *The dispersion in implied relative risk aversion coefficients across assets with different consumption risk characteristics can more readily be understood e.g. in a framework with heterogenous tastes. The induced preferences for the representative agent exhibit in that case different properties than CRRA tastes [e.g., see Benninga and Mayshar (1993)]. Consistent with our findings, Mankiw and Zeldes (1991) estimate that the risk aversion of non-stockholders is substantially higher than that of stockholders.*

In summary, the implied risk-return tradeoffs vary significantly across the assets for both periods and asset classes. Thus, model rejections appear to be related to industry characteristics as well as to differences in the risk-return characteristics for stocks and bonds, respectively. In addition, the cross-sectionally restricted estimates of relative risk aversion for the 1977–83 period (2.269 to 3.448) are higher than any of the estimates obtained from the test assets (0.055 to 1.216). In contrast, despite model rejections, the point estimates of CRRA are consistent in magnitude across test constructions for the 1984–90 period. All in all, these findings agree with the the documented rejections of model stability and differences in the goodness-of-fit for the model for the two periods.

5. *Conclusions*

The present paper tests restrictions imposed on asset prices by a return based version of the Lucas (1978) pricing model with constant relative risk aversion (CRRA) tastes. As test assets, we use eight industry sorted stock portfolios and two long-term bonds. The conditioning information consists of Swedish and US stock market returns, changes in Swedish and German interest rates and the growth in Swedish industrial production. As a proxy for consumption growth, we employ a multivariate portfolio of the Swedish and US stock market and one-month Swedish Treasury bill factors.

In monthly data from 1977–90, this study discovered a number of empirical regularities of interest. Consistent with the profound changes in the Swedish market environment (e.g., deregulation of the financial markets), tests of structural stability with a break at January 1984 reject constant model parameters over the 1977–90 period. Thus, we proceeded to study the nonoverlapping subperiods 1977–83 and 1984–90. Our results indicate that the US stock market factor has explanatory power in the pricing of stocks after controlling for local pricing factors for both periods. This is suggestive of partial integration between the Swedish and US stock markets. Also, we document that the domestic Treasury bill return is an important factor in the pricing of stocks and bonds.

In general, model errors appear to be fairly unbiased and the cross-sectionally restricted estimates of representative CRRA implied in the tests of model unbiasedness seem plausible: they support risk aversion, but with a lower tolerance than for the log utility investor. The CRRA estimates range from 2.269 to 5.202 across periods and test assets used. This supports the presence of a strong "income effect" in tastes, i.e. intertemporal consumption smoothing in response to changes in available investment opportunities. However, the equality of the implied CRRA coefficients across stock portfolios and bonds is violated in the tests. Indeed, the evidence suggests that a common CRRA is not a correct specification of risk-return tradeoffs across stocks and bonds, respectively, and even less so across the two asset categories. Interestingly, the estimates of relative risk aversion implied from bond returns are significantly lower than those inferred from stocks returns. This suggests that bond and stock markets may attract different "consumer clienteles" due to heterogeneity in preferences, market frictions, liquidity constraints or other reasons for why investors may have customized values for different assets. Thus, despite yielding interesting empirical insights, the estimated pricing model with CRRA preferences seems to be an unsatisfactory approximation of reality.

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