INTERNATIONAL JOURNAL OF BUSINESS, 4(2), 1999 ISSN:1083-4346

Asset Pricing Model without Consumption Data: An Empirical Study of Pacific Basin Equity Markets

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This study examines whether the IAPM can explain the variation in Pacific Basin equity markets. We use an asset-pricing model without consumption data, which is developed by Campbell [5], to avoid the measurement errors of consumption data in these markets. The overidentifying restrictions of the IAPM are not rejected at a conventional significance level; however, the null hypothesis of the IAPM is marginally rejected against the model that include a constant term to each market to account for the cross-sectional variations. This result implies that Campbell's IAPM may explain some behaviors of the Pacific Basin equity markets. However, there must exist some other risk factors omitted in our formulation, especially the news of state variables on future world wealth. There also exist other potential sources of misspecification that might explain our results, such as the omission of transaction cost, poor proxy of world market portfolio, heterogeneity of consumers, and invalid assumption of full market integration.

I. INTRODUCTION

The Pacific Basin stands out as the world's fastest growing region. This spectacular development success is convincing because of open economies, growing regional integration, and trade and capital flows. The capital markets have accompanied this economic wave. The investment benefits of these markets have attracted the attention of scholars and practitioners around the world. However, the relationship between the expected asset return and the investment risk has not been fully explored.

Understanding the relationship between expected Pacific Basin equity returns and investment risk can help international investors benefit from investing in these markets. If the time-series and cross-sectional patterns of Pacific Basin equity returns can be explained using existing models, then one

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can easily identify the risk factors, and construct a desirable portfolio. The purpose of this paper is to examine whether the behavior of Pacific Basin equity returns can be explained using the Intertemporal Asset Pricing Model (IAPM) without consumption data, which is developed by Campbell [5].

Financial economists typically attribute the expected asset return to the covariance between the asset return and the typical investor's consumption, because the consumer wants to maximize his lifetime utility, which depends on the consumption. The Consumption Capital Asset Pricing Model of Lucas [22] and Breeden [3] has been the subject of extensive empirical studies relating the behavior of expected asset returns to the covariance between realized returns and consumption growth. Most studies use a time- and state- separable power utility function of a nondurable consumption good. However, this approach has led to the so-called equity premium puzzle of Mehra and Prescott [24] and the risk-free rate puzzle of Weil [28]. Moreover, the direct test of their Euler equations is also rejected; see, e.g., Hansen and Singleton [14].

Researchers have attempted to resolve these puzzles by using the non-expected utility functions of Kreps and Porteus [19,20], which allow the separation of the coefficient of relative risk aversion and the elasticity of intertemporal substitution. This class of recursive, but not necessarily expected utility, preference over intertemporal consumption is developed by Epstein and Zin [8] and Weil [29]. Thus, researchers have integrated into the IAPM this kind of preference to examine the linkage between asset returns and consumption growth. However, the results are mixed. Giovannini and Jorion [11] reject the restrictions implied by the Euler equation. Weil [28] shows that relaxing the parametric restriction between the coefficient of relative risk aversion and the elasticity of intertemporal substitution do not suffice to solve the equity premium puzzle. The results of Epstein and Zin [9] provide some evidences against their own model for nondurable goods and services. Combining non-expected utility functions and modification of market fundamentals, however, Hung [17] shows that the non-expected utility is capable of exactly matching the means of the risk-free rate and the risk premium.

The problems of using consumption data to test the restriction of the Euler equation are that there may exist measurement errors, e.g., Wheatley [30], and poor proxy of the agent's consumption by aggregate consumption, e.g., Mankiw and Zelds [23] and Campbell and Kyle [7]. Recently, Campbell [5] presents an IAPM without consumption data, which is based on the Epstein and Zin [8] and Weil [29] type of preference. Campbell uses log-linear approximation of the budget constraint to eliminate consumption data, resulting in a tractable and empirically implementable IAPM. Although this model resembles the equilibrium model of Merton [25], the risk factors are defined as the impact of news of economic variables on future total wealth.

Three previous papers, as far as we know, have used Campbell's IAPM to examine the U.S. stock market returns. Campbell [6] attempts to apply his own model to examine the coefficient of relative risk aversion with the introduction of human capital. Hardouvelis *et al.* [15] also use the same model to compare model with explicitly pricing consumption risk. Li [21] proposes a time-varying conditional covariance's approach to test the Campbell's IAPM and the IAPM without constraining the parameters of factor risk prices. All the results show that Campbell's IAPM fit the data slightly better than IAPM without imposing restriction on risk prices; however, the goodness-of-fit test and likelihood ratio test of the model are generally rejected.

The measurement error of consumption data in Pacific Basin markets must be more serious than that of the U.S. market. This is the reason why we use the IAPM without consumption data to examine the behavior of Pacific Basin equity markets. Ours results show that the model cannot be rejected at a conventional significance level. However, when we test the null hypothesis of Campbell's IAPM against the model by adding a constant term to each market to account for cross-sectional variation in the Pacific Basin equity markets, Campbell's IAPM is marginally rejected. Using the non-expected utility function does not help the IAPM without consumption data to explain the risk premium in these markets. There must exist some other risk factors that are omitted in our formulation affecting the risk premium of Pacific Basin equity markets. Therefore, the distinct risk factors that exist in each market must be identified to fully explore the risk premium in the Pacific Basin equity markets.

The next section reviews the properties of the Campbell's IAPM and presents an approach to test the model. Section 3 describes the data and summary statistics. Section 4 presents empirical results from estimating and testing Campbell's IAPM. Concluding remarks are found in Section 5.

II. Methodology

The asset pricing model of Campbell [5] lies in two features: one is to log-linearize dynamic budget constraint and substitute covariance between asset return with consumption growth out the Euler equation of non-expected utility model; another is to use a first-order vector autoregression (VAR) of Campbell [4] to capture the revisions in discounted innovation of forecast future market return. If the asset return and consumption growth are jointly log-normal, following the studies of Hardouvelis *et al*. [15] and Li [21], the expected excess return becomes:

$$E_t(R_{i,t+1}/\Omega_t) - R_{f,t+1} \cong [\gamma + (\gamma-1)\lambda_1]V_{im} + (\gamma-1)\sum_{k=2}^{K}\lambda_k V_{ik}, \quad (1a)$$

$$E_t(R_{i,t+1}/\Omega_t) - R_{f,t+1} \cong [\gamma + (\gamma-1-\delta)\lambda_1]V_{im,t} + (\gamma-1-\delta)\sum_{k=2}^{K}\lambda_k V_{ik,t}. \quad (1b)$$

In the above equations, $E_t(\bullet/\Omega_t)$ represents expectation given the information set Ω_t at time t. $R_{i,t+1}$ and $R_{f,t+1}$ denote the simple compounding return on risky asset i and risk-free asset from t to $t+1$. V_{ab} represents covariance of a with b. γ is the coefficient of relative risk aversion (CRRA). σ is the elasticity of intertemporal substitution. $\delta = \psi(\gamma - 1/1 - \sigma)$, where ψ is a constant parameter that relate the covariance of the market return with consumption growth to expected log real market return $r_{m,t+1}$.[1] $\lambda \equiv e1\rho A(I-\rho A)^{-1}$, where ρ can be interpreted as a constant ratio of invested wealth to total wealth, A is a kxk coefficient matrix of a VAR model, and $e1$ is a k-element vector whose first element is unity and whose other elements are all zero. λ is proposed to measure the impact of each news of state variables on the future market return.

Equations (1a) and (1b) are the Campbell's K-factor asset pricing model under the assumptions that asset return and consumption growth are conditional homoskedasticity and heteroskedasticity, respectively. The first factor is the innovation in the market return, while the other factors are innovations in variables that consumers use to forecast the return on the market.

To investigate the asset-pricing model without consumption data for Pacific Basin equity markets, we combine a first-order VAR model with Equations (1):

$$\begin{aligned} &\varepsilon_{t+1} = z_{t+1} - Az_{t+1}, \\ &e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - [\gamma + (\gamma-1)\lambda_1]r_{i,t+1}\varepsilon_{1,t+1} - (\gamma-1)\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{k,t+1}, \end{aligned} \quad (2)$$

in the case of homoskedasticity, and

$$\begin{aligned} &\varepsilon_{t+1} = z_{t+1} - Az_{t+1}, \\ &e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - [\gamma + (\gamma-1-\delta)\lambda_1]r_{i,t+1}\varepsilon_{1,t+1} - (\gamma-1-\delta)\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{k,t+1}, \end{aligned} \quad (3)$$

in the case of heteroskedasticity, where ε_{t+1} and $e_{i,t+1}$ are innovations, and z_t, a subset of Ω_t, is the information consumers used to form their expectations.

Let $h_{t+1} = (\varepsilon_{t+1}, e_{t+1})$. Under rational expectations, (2) and (3) imply that $E_t(h_{t+1} / z_t) = 0$. Therefore, Hansen's [13] generalized method of moment (GMM) estimation can be used to estimate the system equations.[2] If there are n portfolios to be examined, there will be $(k+1)(k+n)$ orthogonality conditions, where we add a constant to the lagged k-element state vector z_t to form instrumental variables. With k^2+1 and k^2+2 parameters, there will be $n(k+1)+k-1$ and $n(k+1)+k-2$ overidentifying restrictions in the case of homoskedasticity and heteroskedasticity, respectively. Hence, we can test the adequacy of Equations (2) and (3) using Hansen's J statistic, which is asymptotically distributed as chi-squared (χ^2) with the number of degrees of freedom equal to the number of overidentifying restrictions.

If Campbell's IAPM can describe the behavior of the Pacific Basin equity markets exactly, there should be no other factors can provide more explanations to the variation of these markets. However, the different markets in the Pacific Basin may have cross-sectional variations, which cannot be accounted for by Campbell's IAPM, such as effects of market segmentation or political risk. If there are another factors that can affect the expected return of these markets, these factors can be added to the model to get a more explanatory power. Therefore, we add a constant term to each market asset pricing equations (2) and (3) to account for the cross-section variations, i.e.,

$$\begin{aligned} &\varepsilon_{t+1} = z_{t+1} - Az_{t+1}, \\ &e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - k_i - [\gamma + (\gamma-1)\lambda_1] r_{i,t+1}\varepsilon_{1,t+1} - (\gamma-1)\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{k,t+1}, \end{aligned} \tag{4}$$

and

$$\begin{aligned} &\varepsilon_{t+1} = z_{t+1} - Az_{t+1}, \\ &e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - k_i - [\gamma + (\gamma-1-\delta)\lambda_1] r_{i,t+1}\varepsilon_{1,t+1} - (\gamma-1-\delta)\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{k,t+1} \end{aligned} \tag{5}$$

for the homoskedasticity and heteroskedasticity model, where k_i are constant parameters for each market.

If Campbell's IAPM can explain the behavior of Pacific Basin equity market returns correctly, all the constant terms should be equal to zero. Thus, we can use the likelihood ratio test of Newey and West [27] to test the null hypothesis that all the constant terms in Equations (4) and (5) are all equal to zero. This test is conducted as follows. First, we estimate Equations (4) and (5), which are the unconstrained model, and hold the weighting matrix. Second, Equations (2) and (3), which constrain all the constant terms equal to zero, are estimated using the unconstrained weighting matrix. The difference between the J statistic for these two models is distributed as χ^2 with the number of degrees of freedom equal to the number of parameters been constrained.

To get a more powerful test of Campbell's IAPM, Hardouvelis, *et al.* [15] also test the adequacy of Equations (4) and (5). However, they only use the Wald test. Li [21] uses the same likelihood ratio test of Newey and West [26] to verify the correctness of Campbell's IAPM. Nevertheless, the work of Li focuses on the multiple comparison among Campbell's homoskedasticity model, heteroskedasticity model, and IAPM without imposing restriction on the parameters of risk price.

III. DATA AND SUMMARY STATISTICS

We examine six equity markets: Hong Kong, Japan, South Korea, Malaysia, Thailand, and the U.S. The data of the Asian Pacific markets are drawn from the Pacific-Basin Capital Markets (PACAP) databases, while the U.S. data come from the German and International Financial (GERFIN) database. The market returns from the PACAP database are value weighted market returns with cash dividend reinvested, and the market return for the U.S. is the monthly change rate of Standard and Poor's 500 composite Share Index (S&P500). Available monthly data cover the period from April 1980 to December 1993, which contains 165 observations.

All the market returns are converted into numeraire currency, the U.S. dollar. The excess equity market return is the return on that market converted into dollars minus the U.S. one-month nominally risk-free rate, which is the Fama/Bliss risk-free rate from the Center for Research in Security Prices (CRSP). The real equity market return is the return on that market converted into dollars minus the inflation rate, which is the monthly growth rate of the U.S. Consumer Price Index.[3]

The state variables are the information consumers use to forecast the future world market portfolio. Many researchers have found that some economic variables can capture the dynamics of the world market portfolio. Following Campbell [6], Hardouvelis *et al.* [15], and Li [21], we choose the lagged Morgan Stanley Capital International (MSCI) world log real market return, the lagged S&P500 dividend yield (USD), and the lagged U.S. term premium (TB3) which is the difference between the return on 3-month treasury bill and 1-month treasury bill as state variables.[4]

Table 1 presents the summary statistics for the excess equity returns.

The means, standard deviations, reward to risk ratios, and autocorrelation are shown in Panels A for excess returns. All the excess returns of the five Asian Pacific markets are higher than those of the U.S. and MSCI world market. The highest return appears in Hong Kong, while the U.S. shows the lowest return. However, the higher return takes the price of the higher standard deviation. When the average returns are divided by the standard deviations, the reward to risk ratios of the five Asian Pacific markets are still above those of the U.S. and the MSCI world market.

The excess returns of Malaysia and Thailand show a significant first-order autocorrelations. The development of these two equity markets is slower than the others. Moreover, Bessembinder and Chan [2] have shown that technical analysis could gain abnormal returns in Malaysia and Thailand. Consequently, it is not surprising that these two markets have a significant first-order autocorrelation. While the Korean return has a very low first-order autocorrelation, it exhibits a significant second order autocorrelation.

The unconditional correlation matrix for excess market returns is shown in Panel B of Table 1. Except for Japan, the equity returns from the other four Asian Pacific countries are less correlated with the world market than the U.S. is. The U.S. and Japan show a 0.58 and 0.73 correlation with the MSCI world return, respectively. Among the markets, only the U.S. and Japan are included in the calculation of the MSCI world index. Hence, these two markets display high correlation with the MSCI world index.

Table 2 displays some statistics for state variables. The S&P500 dividend yield and term premium exhibit significant first-order autocorrelation. In addition, the higher order autocorrelations of the S&P500 dividend yield decrease exponentially, suggesting an autoregressive process pattern. We use the Phillips-Perron [27] test to examine the null hypothesis of the existence of a unit root. Although the tests of the term premium and log real world return are rejected when constant and trend are included with truncation at lag 1, the test for the S&P500 dividend yield is not rejected at a 5% significance level. The lack of rejection in the S&P500 dividend yield might result from its method of construction, which is a moving average.

Table 1

Summary statistics for the market return

Panel A: Excess returns

Market[a]	Mean	Std. Dev	Reward /Risk	Autocorrelation					
				ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
HK	0.0158	0.094	0.168	0.094	-0.064	-0.018	-0.108	-0.015	0.030
JP	0.0085	0.069	0.123	0.088	-0.006	0.018	0.048	0.067	-0.002
KO	0.0105	0.074	0.142	-0.000	0.174*	-0.063	0.076	0.099	0.104
MY	0.0115	0.081	0.142	0.179*	0.114	-0.031	0.010	-0.019	0.032
TH	0.0176	0.078	0.225	0.207*	0.143	0.004	-0.113	0.024	-0.028
U.S.	0.0038	0.035	0.110	0.016	0.020	-0.100	-0.020	-0.119	0.025
World	0.0047	0.043	0.109	0.053	0.031	-0.022	-0.005	0.017	0.131

Panel B: Unconditional correlation for excess market return

Market	HK	JP	KO	MY	TH	U.S.	World
HK	1	0.272	0.141	0.493	0.336	0.302	0.416
JP		1	0.337	0.291	0.168	0.293	0.729
KO			1	0.176	0.032	0.136	0.251
MY				1	0.467	0.357	0.432
TH					1	0.322	0.339
U.S.						1	0.577
World							1

[a] Six markets are examined: Hong Kong (HK), Japan (JP), South Korea (KO), Malaysia (MY), Thailand (TH) and the United States (U.S.). Meanwhile, the Morgan Stanley Capital International (MSCI) world index is included as a efficiency world market portfolio. Available monthly data cover the period from 1980:4 to 1993:12, which contains 165 observations.

* Significance at the 5% level.

Table 2

Summary statistics for the state variables

Panel A: State variables

			Autocorrlation					
Variable[a]	Mean	Std. Dev	ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	τ^b
LNWR	0.0047	0.043	0.063	0.021	-0.024	-0.010	0.015	-154.8*
USD	0.0337	0.008	0.948*	0.886*	0.828*	0.792*	0.506*	-18.6
TB3	0.0055	0.007	0.209*	0.107	-0.116*	-0.161*	0.153	-132.3*

Panel B: Unconditional correlation of state variables

Variable	LNWR	USD	TB3DIF
LNWR	1	-0.016	-0.040
USD		1	0.037
TB3DIF[c]			1

[a] The state variables contain the log real MSCI world market return (LNWR), the S&P500 dividend yield (USD), and the term premium (TB3), which is calculated by the return on 3-month treasury bill less the return on 1-month treasury bill.

[b] The unit root test is the Phillips-Perron [27] test, which includes constant and trend term with truncation at lag 1.

[c] To avoid the inclusion of a second highly first order series besides the S&P500 dividend yield series, we follow the suggestion of Hardouvelis *et al*. [15] to use the first difference of term premium (TB3DIF).

* Significance at the 5% level.

Table 2 also shows the correlation for state variables in Panel B. The state variables will form a VAR model to capture the news of state variables. Avoiding the inclusion of a second highly first order series besides the S&P500 dividend yield series, we follow the suggestion of Hardouvelis, *et al.* [15] to use the first difference in the term premium. Panel B shows a very low contemporaneous correlation between these state variables, so each variable may contain an independent prediction about future world market return.

Table 3 presents the results of estimating the VAR model by Seeming Unrelated Regression (SUR). The investment/wealth ratio, ρ, is set at 0.99 monthly in the estimation, which is similar to those in Campbell (0.94 annually) and Hardouvelis, *et al*. (0.985 quarterly). For the log real world return, a positive log real world return and dividend yield predict a positive log real world return for the next month. The result is supported by many researchers, e.g., Harvey [16]. A negative first difference of term premium anticipates higher log real world return for the next period. Hardouvelis *et al*. [15] attribute this phenomenon to risk premium. When risk increased, stock prices would go down to gain a higher premium. The lowering of stock price pushes treasury bill prices up, resulting in a lower yield. Although the results are consistent with previous studies, only the coefficient of dividend yield reaches the 10% significance level. The corresponding coefficient of determination (R^2) is only 0.8%. However, the dividend yield and first difference of term premium have nice explanatory power. For dividend yield, the coefficient of log real world return and lag dividend yield reach significance at the 1% level. The associated R^2 is 94.8%. For the first difference of term premium, the coefficient of its own lag reaches a 1% significance level, resulting in R^2 of 21.6%.

The first element of λ_k, which displays the importance of the news of each state variable affecting the future world market return, is a shock of the lagged log real world market return. This shock is negative, implying that the world market return has negative correlation with future world market return. The next element, which is an impact of the dividend yield shock, shows a large influence on the future world market return. The influence of term premium, which is the last element of λ_k, also shows a positive but very low impact. The impacts of these state variables on the world market return can be explained by financial theory, but all the impacts do not reach any significance level.[5] The U.S. financial variables may be a poor predictor for future MSCI world return. The MSCI world index also cannot predict its future very well. Although these state variables do not work well for predicting future MSCI world return, they still provide some clue for future MSCI world return. Thus, we will use these variables to form a VAR model in our empirical study.

Table 3

Summary statistics of VAR model*

Dependent variable	Coefficient				
	LNWR	USD	TB3DIF	λ_k	R^2
LNWR	0.062	0.171	-0.052	-0.094	0.008
(t-statistics)	(0.80)	(1.74)	(-0.14)	(-0.04)	
USD	-0.014	0.995	0.008	10.499	0.948
(t-statistics)	(-4.74)	(260.6)	(0.57)	(0.56)	
TB3DIF	-0.011	-0.011	-0.438	0.028	0.216
(t-statistics)	(-0.80)	(-0.64)	(-6.67)	(0.003)	

* Three state variables: log real MSCI world return (LNWR), S&P500 dividend yield (USD), and the first difference of term premium (TB3DIF), are included in the VAR model, $z_{t+1} = Az_t + \varepsilon_t$, where z_t is the economical information by which consumers use to forecast future total wealth, A is coefficient matrix, and ε_t is innovation. The heteroskedasticity consistent t-statistics are displayed in parentheses. $\lambda \equiv e1\rho A(I - \rho A)^{-1}$, which measure the impacts of each news on the forecast world market return. R^2 is the coefficient of determination.

IV. EMPIRICAL RESULTS

A. Intertemporal asset pricing model under homoskedasticity

Table 4 reports the results of estimating Campbell's IAPM under homoskedasticity. There are four instrumental variables, which include a constant term and three state variables. This results in $(3+1)\times(3+6)=36$ moment conditions. In Panel A, there are 10 parameters in the model, leaving 26 overidentifying restrictions to be tested. The value of Hansen's J statistic is 33.09, leading to a p-value of 16.0%, which does not reach any conventional significance level. We could not reject this model within the Pacific Basin equity markets. Therefore, Campbell's IAPM might have explanatory power for these markets.

The estimate of CRRA (γ) is 9.292, which is significant at the 10% level but is high. Mehra and Prescott [24] claim that the CRRA must be high enough to explain the equity premium puzzle of the Consumption Capital Asset Pricing Model. Kandel and Stambaugh [18] also show a high CRRA to match the first moment of the asset returns. For Pacific Basin equity markets, it is plausible to require a high CRRA to explain the observed high equity premium in these markets. Although we cannot compare our results to those of other studies owing to the different markets which are examined, the value of our CRRA is close to the Back-of-The Envelope calculation of Campbell [6], CRRA is 7.8 when the human wealth to total wealth ratio is set at two-thirds. However, the estimate in our study is insignificant at the 5% level. The imprecision in the estimate of the CRRA is also a feature of other studies, e.g., Giovannini and Jorion [12], Attanasio and Weber [1] and Hardouvelis *et al*.[15]. The attitude toward risk might change over time to reflect economic growth or democratization in the Pacific Basin region. Furthermore, investors from different markets may have different attitudes toward risk. As a result, the changes and differences in risk aversion undermine our estimate of CRRA.

The estimates of the coefficients in the VAR model are not exactly equal to the summary statistics in the unconstrained VAR model in Table 3. In particular, there is a discrepancy between Table 3 and Table 4 for the impact of news of the state variables on the future MSCI world return, λ. The lag MSCI world return also shows a negative impact on its future, but is larger than that of Table 3. The U.S. dividend yield still has a positive but increased impact on the future MSCI world return. On the other hand, the first difference of term premium shows a larger impact in the opposite direction. It is disappointing that the estimates of λ are not different from zero at any conventional significance level.

Table 4

Estimation and test of IAPM under homoskedasticity

In Panel A, the following equations are estimated by GMM

$$\varepsilon_{t+1} = z_{t+1} - Az_{t+1},$$

$$e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - [\gamma + (\gamma - 1)\lambda_1]r_{i,t+1}\varepsilon_{1,t+1} - (\gamma - 1)\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{i,t+1}.$$

In Panel B, a constant term is added to each asset pricing equation,

$$\varepsilon_{t+1} = z_{t+1} - Az_{t+1},$$

$$e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - k_i - [\gamma + (\gamma - 1)\lambda_1]r_{i,t+1}\varepsilon_{1,t+1} - (\gamma - 1)\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{i,t+1},$$

where $k_i, i = 1,2,...,6$, are constant parameters. Six markets: Hong Kong, Japan, South Korea, Malaysia, Thailand, and the United States, are examined for period from 1980:4 to 1993:12, which contain 165 monthly observations.

Panel A: IAPM without consumption data[a]

	VAR Coefficient				APM Coefficient
	LNWR	USD	TB3DIF	λ_k	γ
LNWR[b]	-0.177	0.232	-0.420	-0.265	9.292
	(-4.89)	(2.71)	(-2.95)	(-0.04)	(1.76)
USD	-0.009	0.998	0.004	14.738	
	(-21.4)	(331.6)	(1.60)	(0.67)	
TB3DIF	-0.002	-0.001	-0.463	-0.167	
	(-0.27)	(-0.09)	(-8.96)	(-0.13)	

$\chi^2(26)$= 33.090; p-value=0.160

Table 4 (continued)

Panel B: IAPM with adding a constant term

	VAR Coefficient				APM Coefficient			
	LNWR	USD	TB3DIF	λ_k		k_i		γ
LNWR	-0.230	0.329	-0.260	-0.233	HK	0.018	(2.76)	0.799
	(-6.12)	(3.66)	(-1.59)	(-0.06)	KO	0.008	(1.55)	(0.39)
USD	-0.003	0.994	-0.011	16.117	JP	0.013	(2.05)	
	(-3.30)	(315.8)	(-1.45)	(0.80)	MY	0.008	(1.26)	
TB3DIF	0.011	-0.000	-0.539	-0.244	TH	0.010	(2.09)	
	(1.27)	(-0.001)	(-8.89)	(-0.17)	US	0.006	(2.40)	

$\chi^2(20)$= 25.626; p-value=0.179

Panel C: Testing the null hypothesis that all constant terms are equal to zero.

	IAPM with constant terms	IAPM without constant terms	difference	p-value
χ^2	25.626	38.1103	12.477	0.052

[a] The heteroskedasticity consistent t-statistics are reported in parentheses.

[b] The VAR models are formed by three state variables: log real MSCI world return (LNWR), S&P500 dividend yield (USD), and the first difference of the term premium (TB3DIF).

In order to account for the cross-sectional variations, we add a constant term to the asset-pricing model for each market. The results of estimating the modified model are displayed in Panel B of Table 4. With 16 parameters and 20 overidentifying restrictions, the specification test yields a J statistic of 25.63 and a p-value of 17.9%. The modified model is not rejected by the data. The estimate of CRRA is 0.8, which is much more reasonable but is not significantly different from zero. The difference between this CRRA value and the value in Panel A may be explained by the constant terms, which have almost the same order as the means of the market returns displayed in Panel A of Table 1.

To compare these two models, which are not rejected by the data, we use likelihood ratio test of Newey-West [27] to test the null hypothesis that all the constant terms are equal to zero. The result is shown in Panel C. With six degrees of freedom and $\chi^2 = 12.48$, the Campbell's IAPM is rejected at the 10% level but not rejected at the 5% level. It is hard to distinguish these two models when the p-value is 5.2%, but from the standpoint of CRRA, the model with constant terms may be closer to the existing asset pricing theory. Consequently, there must exist some unique impacts on each market, which cannot be described by the MSCI world return or Campbell's model.

B. Intertemporal asset pricing model under heteroskedasticity

The results of estimating the IAPM under heteroskedasticity are shown in Panel A of Table 5. There are 11 parameters, which imply 25 overidentifying restrictions. The associated J statistic equals 29.33, leading to a p-value of 25%. The Campbell's IAPM cannot be rejected by the data. The estimate of the CRRA, however, is 11.46, which is larger than in the homoskedastic model. The estimate of δ, defined as $\psi(\gamma - 1/1 - \sigma)$, is 2.61. The former is significant, but the latter is insignificant. The estimates of the VAR coefficient are the same as in the homoskedastic model, implying the same impact of news of state variables on the future MSCI world return.

Table 5

Estimation and test of IAPM under heteroskedasticity

In Panel A, the following equations are estimated by GMM

$$\varepsilon_{t+1} = z_{t+1} - Az_{t+1},$$

$$e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - [\gamma + [(\gamma - 1) - \delta]\lambda_1]r_{i,t+1}\varepsilon_{1,t+1} - [(\gamma - 1) - \delta]\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{i,t+1},$$

In Panel B, a constant term was added to each asset pricing equation,

$$\varepsilon_{t+1} = z_{t+1} - Az_{t+1},$$

$$e_{i,t+1} = R_{i,t+1} - R_{f,t+1} - k_i - [\gamma + [(\gamma - 1) - \delta]\lambda_1]r_{i,t+1}\varepsilon_{1,t+1} - [(\gamma - 1) - \delta]\sum_{k=2}^{K}\lambda_k r_{i,t+1}\varepsilon_{i,t+1},$$

where $k_i, i = 1,2,...,6$, are constant parameters. Six markets: Hong Kong, Japan, South Korea, Malaysia, Thailand, and the United States, are examined for period from 1980:4 to 1993:12, which contain 165 monthly observations.

Panel A: IAPM without consumption data [a]

	VAR Coefficient				APM Coefficient	
	LNWR	USD	TB3DIF	λ_k	γ	δ
LNWR [b]	-0.182	0.240	-0.370	-0.270	11.460	2.605
	(-4.55)	(2.66)	(-2.73)	(-0.04)	(2.01)	(1.47)
USD	-0.009	0.999	0.008	15.482		
	(-20.2)	(335.1)	(2.75)	(0.66)		
TB3DIF	0.001	-0.004	-0.485	-0.099		
	(-0.18)	(-0.29)	(-9.37)	(-0.07)		

$\chi^2(25)$=29.328; p-value=0.250

Table 5 (continued)

Panel B: IAPM with adding a constant term to each asset pricing equation

	VAR Coefficient				APM Coefficient			
	LNWR	USD	TB3DIF	λ_k		k_i	γ	δ
LNWR	-0.230	0.331	-0.266	-0.233	HK	0.018(2.55)	1.097	0.632
	(-5.93)	(3.57)	(-1.63)	(-0.06)	KO	0.007(1.43)	(0.46)	(0.26)
USD	-0.003	0.994	-0.010	16.204	JP	0.012(1.84)		
	(-3.26)	(316.4)	(-1.32)	(0.80)	MY	0.007(1.06)		
TB3DIF	0.011	-0.000	-0.537	-0.240	TH	0.010(2.02)		
	(1.24)	(-0.01)	(-8.82)	(-0.17)	US	0.005(1.62)		

$\chi^2(19)=25.393$; p-value=0.148

Panel C: Testing the null hypothesis that all constant terms are equal to zero

	IAPM with constant terms	IAPM without constant terms	difference	p-value
χ^2	25.393	36.474	11.081	0.086

[a] The heteroskedasticity consistent t-statistics are reported in parentheses.

[b] The VAR model are formed by three state variables: log real MSCI world return (LNWR),S&P500 dividend yield (USD), and the difference of the term premium (TB3DIF).

Panel B shows the results of adding a constant term to each market to account for the cross-sectional variations. With J statistic equal to 29.328, we cannot reject the model, which has 17 parameters and 19 overidentifying restrictions. The estimate of the CRRA declines to a reasonable value, 1.1, but does not reach any conventional significance level. The estimate of δ is 0.63 and is also insignificant. The cross-sectional variations are captured by the constant terms. The order of the estimates of the constant terms is the same as in the homoskedastic model, and is almost the same as the order for the unconditional means of the market returns.

No matter whether constant terms are added, the heteroskedastic IAPM without consumption data is not rejected. Hence, we use the likelihood ratio test of Newey and West [27] again to test the null hypotheses that all the constant terms are equal to zero. The result is displayed in Panel C. With 6 degrees of freedom and a J statistic of 11.1, the model with all the constant terms equal to zero is also rejected at the 10% level but not at the 5% level. Under either homoskedasticity or heteroskedasticity, Campbell's IAPM is marginally rejected against the model with adding a constant term to each market to account for the cross-sectional variations. Campbell's IAPM may explain some characteristics of the Pacific Basin equity return, but some risk factors which are important in describing these markets are omitted in our formulation of the IAPM.

V. CONCLUSION

This study examines whether the IAPM can explain the variation in Pacific Basin equity markets. We use an asset-pricing model without consumption data, which is developed by Campbell [5], to avoid the measurement errors of consumption data in these markets. The overidentifying restrictions of the IAPM are not rejected at a conventional significance level; however, the null hypothesis of the IAPM is marginally rejected against the model that include a constant term to each market to account for the cross-sectional variations. This result implies that Campbell's IAPM may explain some behaviors of the Pacific Basin equity markets. However, there must exist some other risk factors omitted in our formulation, especially the news of state variables on future world wealth. There also exist other potential sources of misspecification that might explain our results, such as the omission of transaction cost, poor proxy of world market portfolio, heterogeneity of consumers, and invalid assumption of full market integration.

Moreover, the estimate of the coefficient of relative risk aversion associated with the IAPM is implausibly high. This high estimate may result from the high equity premium of Pacific Basin markets and the fluctuation of market returns. However, when we incorporate into the IAPM a constant term for each market, the estimate of the coefficient of relative risk aversion comes down to a reasonable value. Still, the high equity premium and the fluctuation of market returns could be interpreted by some other risk factors, which are omitted in our formulation.

NOTES

1. Specifically, $\mu_{m,t+1} = \mu_0 + \psi E_t(r_{m,t+1} / \Omega_t)$, where $\mu_{m,t+1}$ is a covariance function of the market return and consumption growth, μ_0 and ψ are constant parameters.
2. We use iterated GMM, which has been proven to be superior to the two-step GMM in a finite sample by Ferson and Foerster [10].
3. The U.S. Consumer Price Index comes from the INTLINE database.
4. The U.S. dividend yield and term structure are all drawn from the U.S. Financial database.
5. The t-statistics were calculated on asymptotic standard errors, which are computed from the matrix, $[\partial f/\partial a]' V [\partial f/\partial a]$, where V is the variance-covariance matrix of the estimated VAR parameters, a, and $[\partial f/\partial a]$ is the matrix of derivatives of the elements of λ with respect to each element of the vector, a.

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