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A best choice among asset pricing models? The Conditional Capital Asset Pricing Model in Australia

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Abstract

We use Australian data to test the Conditional Capital Asset Pricing Model (Jagannathan and Wang, 1996). Our results are generally supportive: the model performs well compared with a number of competing asset pricing models. In contrast to the study by Jagannathan and Wang, however, we find that the inclusion of the market for human capital does not save the concept of the time-independent market beta (it remains insignificant). We find support for the role of a small-minus-big factor in pricing the cross-section of returns and find grounds to disagree with Jagannathan and Wang's argument that this factor proxies for misspecified market risk.

Key words: Asset pricing; Australia; Conditional CAPM; APT

JEL classification: G12

1. Introduction

Motivated by Fama and French's (1992) study finding that the 'relation between market beta and average return is flat' (p. 427), Jagannathan and Wang (1996) revisited the performance of the Capital Asset Pricing Model (CAPM) with US data but, unlike Fama and French (1992), made specific allowance for the possibility that firm betas vary through time. The resulting tests rejected the

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standard CAPM (when firm betas were assumed stationary) but tests of the conditional CAPM (where the effects of time-varying betas were accounted for) performed well. This conditional CAPM outperformed the empirically derived version of the Arbitrage Pricing Theory (APT) by Chen *et al.* (1986) and displayed a level of explanatory power similar to that of the Fama-French three-factor model (Fama and French, 1992, 1993).

Despite the success of the analysis by Jagannathan and Wang (1996), no study has examined the conditional CAPM outside of the USA. This study utilizes Australian data and closely follows the methodology in Jagannathan and Wang (discussed in Section 3). Following Jagannathan and Wang, the standard CAPM and conditional CAPM are tested (Section 4 of the present paper). The CCAPM is then tested against Fama and French's (1993) three-factor model (Section 5.1) and against the APT (Section 5.2). Consistency with the Black (1972) version of the CAPM is also investigated (Section 5.4). To this point we have a complete replication of the work of Jagannathan and Wang but, to extend the investigation into the Australian context, further tests are conducted to evaluate the influence of US market movements over Australian stock returns (Section 5.3). We provide an overview of pertinent published literature in Section 2. Section 6 concludes the present paper.

In keeping with Jagannathan and Wang's findings for the US market, we find that the explanatory power of the CAPM is poor (with an R^2 of only 7.25 per cent), but a marked improvement is noted when the time-variation in betas is accounted for (the R^2 is raised to 65.31 per cent). The results imply that the assumption of time-invariant firm betas leads to a poorly specified test of the CAPM. In contrast to Jagannathan and Wang, we find that extending the market portfolio from the value-weighted index of stocks to include a measure of the return to human capital does little to improve the explanatory power of the model. Our analysis does not support Jagannathan and Wang's argument, regarding Fama and French's three-factor model, that the size and book-to-market variables may simply proxy for the risks associated with time-series beta variation and the return on human capital: in Australia, small-minus-big (SMB) retains high levels of significance in all the cross-sectional tests reported in the present study. Like Jagannathan and Wang, we find that the conditional CAPM outperforms realizations of the APT. The conditional CAPM also outperforms models that hypothesize that returns are driven by US market forces. When conditional CAPM variables are added to these models, however, the resulting explanations are significantly better than those based on the conditional CAPM alone.

2. Asset pricing tests and the conditional CAPM

Early tests of the static CAPM (Sharpe, 1964; Lintner, 1965; Mossin, 1966) were broadly supportive of the predicted linear relationship of returns to systematic risk (Fama and MacBeth, 1973; Black *et al.*, 1972). In Australia, Ball, Brown, and Officer (1976) found results consistent with the zero-beta CAPM.[1] In the US, subsequent research uncovered empirical regularities, such as the size effect (Banz, 1981), that were, on the face of it, difficult to reconcile with the CAPM paradigm.[2] Fama and French (1992) found no evidence that market β was priced. Fama and French (1993) argue that the cross-section of returns can be modelled using three factors: a market factor, factors based on size, and the ratio of book-to-market value of equity. It is safe to say that the three-factor model has succeeded the CAPM as the paradigm within which asset prices are analysed.

Support for the three-factor model's applicability in Australia is provided in Fama and French (1998), although the number of firms available for inclusion in their study was far fewer than those listed in Australia.[3] Halliwell *et al.* (1999) found the size factor to be significant. Their results, however, were sensitive to the stocks that were included in the sample and the portfolio formation technique,[4] and the book-to-market factor was found to be statistically insignificant. In contrast to Halliwell *et al.* (1999), in Faff (2001) the book-to-market effect was found to be statistically significant but a negative risk premium was estimated for the size factor. It should be noted that the sample used by Halliwell *et al.* ended in June 1991, while Faff used data from the period January 1991 to April 1999, and, rather than constructing factors, Faff proxies high-minus-low (HML) and SMB using commercially available measures.

Perhaps the failure of the CAPM is driven by the time-varying nature of asset betas[5] and market and risk-free premiums? To address this question, Jagannathan and Wang (1996) developed the conditional CAPM as a model of the cross-section of returns in which the value of a firm's beta is conditional on the state of the economy. Readers wishing for a detailed understanding of the model should refer Jagannathan and Wang (1996), but the testable model ultimately derived, and the focus of this paper, is referred to as the Premium-Labour (PL) model (equation (1)) and takes the form:

$$E[R_{it}] = c_0 + c_{VW}\beta^{VW} + c_{prem}\beta^{prem} + c_{labour}\beta^{labour}. \tag{1}$$

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[1] After Roll (1977), the focus in published Australian literature shifted to tests of mean-variance efficiency of market proxies. Stokie (1982) rejects mean-variance efficiency for a number of samples while later analyses provide inconclusive evidence (Faff, 1991; Wood, 1991).

[2] Australian studies have also found evidence of empirical regularities (Officer, 1975; Brown *et al*., 1983; Brailsford and Easton, 1991; Gaunt *et al*., 2000).

[3] Maroney and Protopapadakis (2002), in a similar multicountry study, find less clear support for the three-factor model.

[4] Their results differ for value or weighted portfolios.

[5] Bollerslev *et al*. (1988); Harvey (1989), Ferson and Harvey (1991, 1993) and Ferson and Korajczyk (1995) provide US evidence in support of this. Australian evidence is provided by Faff *et al*. (1992); Brooks *et al*. (1992); and Brooks *et al*. (1997).

R_{it} is the return of asset i in period t,[6] β^{VW} is the market beta based on a value-weighted portfolio of all stocks, β^{prem} captures the systematic changes in a firm's beta with variation in the market premium, and β^{labour} is the beta for the human capital market. Whereas the conventional CAPM prices assets to compensate only for levels of systematic risk, the addition of the extra variable in equation (1) postulates that asset prices will be determined not only by an asset's systematic risk, but by the predictable component of the security's change in systematic risk when there are shifts in the state of the economy. Following Mayers (1972), the market portfolio of wealth is assumed to include not only stocks, but also human capital. The CAPM is nested as a special case (when the cross-sectional coefficients of β^{prem} and β^{labour} are equal to zero) allowing direct comparisons of the conditional CAPM with the standard CAPM.[7]

Using monthly US returns from July 1963 to December 1990, Jagannathan and Wang (1996) found that, when the CAPM was tested, c_{VW} was not significantly different from zero, and the R^2 value was 1.35 per cent when only β^{VW} was included in the model. However, adding β^{prem} resulted in a statistically significant c_{prem} coefficient and increased the R^2 to 29.32 per cent. Adding β^{labour} further increased the R^2 to 55.21 per cent. The PL model captured the size effect. In addition, Jagannathan and Wang compared the *PL* model to other prominent asset pricing models. Their results suggest that the *PL* model outperforms the APT model of Chen *et al.* (1986) and performs at least as well as Fama & French's three-factor model (suggesting that 'the two Fama and French (1993) factors may proxy for the risk associated with the return on human capital and beta instability' – Jagannathan and Wang (1996), p. 32). Although the data would suggest that the *PL* model is a significant improvement over the CAPM, there is still the somewhat contrary finding that the c_{VW} coefficient is statistically insignificant across all regressions. On the surface, this suggests that systematic risk is unimportant, whereas the predictable change in systematic risk is important (i.e. the systematic risk of systematic risk is priced). That the c_{labour} term is often significant, however, is taken as an indication that the market may simply be better characterized by the return on human capital than by the return on common stocks.

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[6] Following Jagannathan and Wang (1996), returns are taken as price relatives, P_t/P_{t-1}, where P_t is the price of the security at time t adjusted for dividends and capitalization changes.

[7] Jagannathan and Wang (1996) note that, although a multiple-beta model is presented, the model is fundamentally different from either Merton's (1973) multi-beta intertemporal models, or Ross's (1976) Arbitrage Pricing Theory. Whereas Merton shows equilibrium to hold when returns are a linear function of several conditional betas, Jagannathan and Wang's model remains conditional upon only one beta; the market beta. The second beta is necessarily incorporated into the model when moving from a conditional to an unconditional formulation of the model. Jagannathan and Wang (1996, pp. 9–10 and pp. 37–40) provide a full discussion of this matter. In contrast to Ross (1976), returns are not assumed to have a linear factor structure.

3. Data and methodology

To test equation (1), share price data for all listed Australian stocks were taken from the Share Price and Price Relative Database (SPPR) from the Centre for Research in Finance at the Australian Graduate School of Management. The database provides the monthly returns on all Australian listed stocks used in the present study, as well as market capitalization figures at the end of each month, used as a measure of firm size. The All Ordinaries Accumulation Index was used as the value-weighted stock index (yielding R_t^{VW}),[8] and as this index begins in 1980, the sample period extends from 31 January 1980 to 31 December 2001.[9] Because of the need to estimate betas prior to portfolio formation, however, the test period begins on 31 January 1982. We follow the portfolio formation methodology of Jagannathan and Wang (1996) and construct 49 portfolios. First, seven size-based portfolios were formed. Each of these portfolios was then sorted into 7 further portfolios based on their companies' equity market betas.[10] We adopted this septile approach because forming portfolios based on deciles, as Jagannathan and Wang did, is impractical in the present study given the smaller number of stocks listed in the Australian market.[11] The number of equities available for inclusion in the portfolios, whose returns form our dependent variable, at each portfolio formation date, may be found in Table 1.

Obtaining measures of the second and third exogenous variables, β^{labour} and β^{prem}, is problematic outside the USA. Monthly macroeconomic data is not available in Australia so we relied on quarterly data. This resulted in a time series of 80 return observations for each of the 49 portfolios. Our data for the return on human capital, used to calculate β^{labour}, is from the 'Average Weekly Earnings of Employees: Dollars: Seasonally Adjusted' time-series available from the Australian Bureau of Statistics, and, following Jagannathan and Wang, a 2-month moving average was used to smooth the series. The spread between long-term AAA and BAA-rated commercial bonds is well supported as a forward indicator of economic conditions in the US.[12] In

[8] On 1 April 2000, the composition of the index was broadened to include 500 stocks, as opposed to 266 before this date. Analysis of the time series of monthly returns after this date, and also the relationship of the index to the American index (see sub-section 5.3 below) provides no evidence of a statistically significant variation in the index which might flow through to the present study.

[9] Whether to include the final quarter of 1987 given the substantial market crash in October is a moot point: inclusion of observations from this time period does not materially affect the conclusions we draw from our analysis.

[10] Based on a minimum of 2 years (and maximum of 5 years) of monthly data prior to portfolio formation.

[11] We also analysed the data forming portfolios on the basis of quartiles (with 16 portfolios in total) and quintiles (yielding 25 portfolios). Our results are robust to the formation of fewer, though more densely populated, portfolios.

[12] See Jagannathan and Wang (1996, p. 11) for relevant published US literature.

Table 1
Sample size at each portfolio formation

Portfolio formation year (31 December)	Stocks available for portfolio formation	Portfolio formation year (31 December)	Stocks available for portfolio formation
1981	415	1991	642
1982	525	1992	619
1983	590	1993	651
1984	589	1994	651
1985	630	1995	718
1986	655	1996	783
1987	749	1997	762
1988	762	1998	764
1989	818	1999	795
1990	658	2000	792

Australia, the illiquidity of the domestic bond market means that data on investment grade bonds is available only from 1997 on, and data on speculative bonds is nonexistent. Given Australia's integration with the global economy, and evidence that there is a strong link with US markets,[13] we believe expectations for the US economy should also act as a leading indicator for Australian conditions. Consequently, we utilize the Fisher effect by means of the following equation

$$r_{AU} = \left[\frac{(1 + i_{AU})(1 + r_{US})}{1 + i_{US}}\right] - 1 = R_{t-1}^{prem}, \tag{2}$$

where r_{AU} and r_{US} refer to the 10-year government bond yields at the end of period $t - 1$, while i_{AU} and i_{US} denote the realized inflation rates in period t. The analysis reported in the present paper indicates that this variable seems to function for Australian returns much as the yield spread functions in the US market.[14] Having found β_i^{prem} for portfolio i, $i = 1, 2, \ldots, 49$, by regressing portfolio returns on R_{t-1}^{prem}, β_i^{prem} is then modified to retain only the proportion of the estimate that is orthogonal to β_i^{VW}. This is done by regressing β_i^{prem} on β_i^{VW}, and taking the regression residuals as the new measure of β_i^{prem}. A similar calculation is performed for the labour beta. In this case, however, the modified value is orthogonal to both β_i^{VW} and β_i^{prem}.

Each equation is estimated using both ordinary least square (OLS) and generalised method of moments (GMM) specifications. In our OLS calculations, we

[13] See Ragunathan *et al*. (1999), Durand *et al*. (2001) and Section 5.3 below.

[14] Further tests of this proxy would be a useful course for future research. The 10-year government bond rates and consumer price index (CPI) levels in the USA are obtained from the Federal Reserve website and CPI levels in Australia are provided by the Australian Bureau of Statistics (ABS) website.

follow Jagannathan and Wang (1996) in adjusting *t*-statistics for sampling error inherent in the estimated betas.[15] In our GMM analyses, we follow Jagannathan and Wang in using the weighting matrix, $A = (E[R_t R_t^T])^{-1}$, for testing various restricted models, as it remains constant across competing models, and allows direct comparison of the pricing errors across each specification by means of the Hansen-Jagannathan (*HJ*) distance (Hansen and Jagannathan, 1994, 1997).[16] In a departure from Jagannathan and Wang, we utilize a bootstrapping procedure, based on 10,000 samples, to derive the empirical distribution of the *HJ*-distance statistic.[17] The bootstrapping technique is used to construct not only a sample distribution for the *HJ*-distance, but also a unique sampling distribution for each of the δ estimates appearing in each equation estimated using GMM.

4. Tests of the conditional CAPM and the CAPM

Table 2 reports summary statistics for the data used in the analysis.[18] Perhaps the most striking aspect relates to market capitalization. The data shows that, by way of contrast to the USA, the size-effect in Australia is very large indeed.[19]

Tests of the CAPM and the CCAPM, including the return on human capital are reported in Table 3.[20] We express *p*-values as percentages in order to be consistent with the approach by Jagannathan and Wang (1996) in presenting results. The OLS tests of the CAPM (Panel A) are consistent with the US versions: the evidence that beta is priced is, at best, marginal (the corrected *t*-statistic has a *p*-value of 7.19 per cent) and the value of R^2 is relatively low (7.25 per cent). In contrast to Jagannathan and Wang, the *t*-statistics for the coefficients (corrected for sampling errors in β from the Fama-MacBeth estimation) result in a marked difference in the interpretation of our results: there appears to be more noise in our data set.[21]

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[15] The technique is discussed in detail in Appendix B of Jagannathan and Wang (1996).

[16] See Jagannathan and Wang (1996, pp. 16–18) for a detailed discussion.

[17] In this analysis, the bootstrap provided a tractable alternative to the analytic solution provided by Jagannathan and Wang (1996) in Appendix C. It is well known that the bootstrap results in a consistent estimate for the distribution of the test statistic. See, for example, Shao and Tu (1995) for a detailed discussion (especially Chapter 3).

[18] Additional summary statistics may be downloaded from http://www.ecom.uwa.edu.au/information_about/staff/durand_robert.

[19] Such high returns should not be surprising to researchers familiar with recent Australian evidence. Gaunt *et al*. (2000, p. 40, Table 2) find average monthly returns as high as 11.26 per cent per month for portfolios of small stocks.

[20] We also tested the CAPM with human capital but the results, which may be downloaded from http://www.ecom.uwa.edu.au/information_about/staff/durand_robert, do not materially affect our analysis.

[21] We do not report uncorrected *t*-statistics and their *p*-values in subsequent tables. In all analyses we subsequently report, correcting the bias in the estimated *t*-statistics has a material effect on the conclusions that may be drawn from the analysis. The uncorrected values appear redundant.

Table 2
Basic characteristics of the 49 portfolios

	$\beta - L$	$\beta - 2$	$\beta - 3$	$\beta - 4$	$\beta - 5$	$\beta - 6$	$\beta - H$
Panel A: Time-series averages of returns							
Size – S	16.38	14.57	20.22	20.67	24.21	19.41	22.43
Size – 2	8.27	8.93	9.30	9.46	8.87	9.69	8.52
Size – 3	1.83	4.99	5.73	4.30	4.72	5.63	5.12
Size – 4	4.48	5.25	3.25	2.60	4.29	1.03	−0.17
Size – 5	3.79	3.56	3.18	2.92	1.68	3.71	1.79
Size – 6	3.27	3.41	2.99	3.38	3.68	1.29	0.29
Size – B	4.04	5.70	4.07	3.61	1.99	2.07	1.06
Panel B: The estimated β_i^{VW}'s							
Size – S	−0.03	0.86	1.36	1.10	1.67	1.48	1.56
Size – 2	0.69	0.53	1.19	1.15	1.42	1.41	1.11
Size – 3	0.50	0.76	0.61	1.00	1.18	1.34	1.57
Size – 4	0.51	0.74	0.74	0.71	1.00	1.10	1.23
Size – 5	0.48	0.58	0.72	0.86	0.96	1.31	1.12
Size – 6	0.68	0.61	0.73	0.90	1.04	1.31	1.43
Size – B	0.89	0.89	0.88	0.96	1.11	1.04	1.24
Panel C: The time-series averages of size (log million Australian dollars)							
Size – S	2.25	2.26	2.32	2.32	2.33	2.35	2.35
Size – 2	2.65	2.66	2.67	2.63	2.64	2.65	2.65
Size – 3	2.89	2.92	2.93	2.91	2.90	2.91	2.89
Size – 4	3.19	3.21	3.19	3.18	3.18	3.17	3.16
Size – 5	3.53	3.54	3.54	3.54	3.51	3.54	3.50
Size – 6	3.93	3.99	3.98	4.02	4.03	4.01	3.98
Size – B	4.80	4.84	4.90	4.90	4.91	4.99	4.79

Panel D: The estimated β_i^{prem} *that is orthogonal to* β_i^{VW}							
Size – S	−3.13	−1.67	−1.86	−4.37	−1.75	−4.23	−3.77
Size – 2	−1.98	−3.86	0.90	0.30	1.88	−0.39	−0.03
Size – 3	0.19	−0.79	−0.29	−0.16	1.37	−0.32	1.53
Size – 4	−0.67	1.14	1.24	1.05	0.84	−0.10	0.84
Size – 5	0.96	0.71	0.42	−0.58	1.04	0.74	0.06
Size – 6	0.93	0.57	1.41	0.46	0.62	0.08	1.76
Size – B	1.53	1.83	1.55	0.98	1.22	1.52	0.25
Panel E: The estimated β_i^{labour} *that is orthogonal to* β_i^{prem} *and* β_i^{VW}							
Size – S	−1.43	−3.24	−2.38	4.05	1.03	−0.48	0.28
Size – 2	−1.40	0.24	−1.09	−3.49	−2.57	−0.27	0.62
Size – 3	0.32	−0.18	0.92	−1.00	−1.02	−3.61	0.16
Size – 4	0.00	−0.93	−0.45	−0.63	−0.18	−0.20	−0.04
Size – 5	−0.04	1.19	−0.21	0.84	−0.62	0.33	2.16
Size – 6	−0.06	0.85	0.74	1.49	0.73	1.20	1.41
Size – B	1.58	−0.11	0.78	0.71	0.94	0.18	2.91

Using publicly listed, non-financial Australian firms, the 49 portfolios are formed in a similar manner to Fama and French (1992). For every calendar year, starting from 1982, firms are first sorted into size septiles based on their market value at the end of December. For each size category, each firm's pre-beta is estimated by the slope coefficient in the regression of the 24 to 60 months of past-return data on a constant and the All Ordinaries Accumulation Index of the corresponding months. Firms within each size septile are then sorted into beta septiles based on their pre-betas. This gives 49 portfolios, and the return on each of these portfolios for the next 12 calendar months is computed by equally weighting the returns on stocks in the portfolio. This procedure is repeated for each calendar year. This gives a time series of monthly returns, which is then converted to quarterly returns, as the inflation data used in the calculation of R_t^{prem} is only available quarterly. The time series extends from March 1982 to December 2001; that is, 80 observations. β_i^{VW} is the slope in the regression of portfolio i's return on the All Ordinaries Accumulation Index and a constant for the entire 80-quarter period. A portfolio size is calculated as the equally weighted average of the logarithm of the market value (in millions of Australian dollars) of the stocks in the portfolio. β_i^{prem} and β_i^{labour} are calculated in a similar way. The numbers given in Panel D are that part of β_i^{prem} which is orthogonal to a constant and β_i^{VW}, and the numbers in Panel E are that part of β_i^{labour} which is orthogonal to a constant, to β_i^{VW} and to β_i^{prem}.

Table 3
Evaluation of various Capital Asset Pricing Model (CAPM) specifications

Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{size}	R-square
Panel A: The static CAPM without human capital						
Estimate:	1.02	0.05				7.25
t-value:	79.19	2.01				
p-value:	0.00	4.99				
Corrected-t:	70.85	1.84				
Corrected-p:	0.00	7.19				
Estimate:	1.19	0.04			−0.05	45.85
t-value:	34.42	1.63			−4.44	
p-value:	0.00	10.92			0.01	
Corrected-t:	32.25	1.56			−4.16	
Corrected-p:	0.00	12.66			0.01	
Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	c_{size}	HJ-dist
Estimate:	−3.21	3.99				1.89
p-value:	65.74	50.20				0.20
Panel B: The Conditional CAPM without human capital						
Estimate:	1.02	0.05	−0.03			65.31
t-value:	79.19	2.01	−5.64			
p-value:	0.00	5.01	0.00			
Corrected-t:	33.90	0.87	−2.51			
Corrected-p:	0.00	38.92	1.57			
Estimate:	1.08	0.04	−0.02		−0.02	68.59
t-value:	41.07	1.93	−6.66		−2.26	
p-value:	0.00	6.03	0.00		2.87	
Corrected-t:	20.61	0.98	−3.62		−1.14	
Corrected-p:	0.00	33.12	0.07		26.23	
Estimate:	−4.23	3.70	55.90			1.79
p-value:	64.90	48.96	48.59			0.55

Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	c_{size}	*HJ*-dist
Panel C: The conditional CAPM with human capital						
Estimate:	1.02	0.05	−0.03	−0.01		68.56
t-value:	79.19	2.01	−5.64	−2.62		
p-value:	0.00	5.02	0.00	1.20		
Corrected-*t*:	27.47	0.71	−2.03	−0.92		
Corrected-*p*:	0.00	48.44	4.83	36.04		
Estimate:	1.06	0.04	−0.02	0.00	−0.01	69.45
t-value:	43.97	2.01	−7.12	−1.84	−1.45	
p-value:	0.00	5.04	0.00	7.24	15.36	
Corrected-*t*:	18.02	0.84	−3.16	−0.77	−0.60	
Corrected-*p*:	0.00	40.73	0.29	44.33	55.40	
Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	c_{size}	*HJ*-dist
Estimate:	−8.55	3.79	54.03	4.22		1.79
p-value:	47.76	49.20	50.34	46.27		0.33

This table gives the estimates for the cross-sectional regression model

$$E[R_{it}] = c_0 + c_{size} \log (ME_i) + c_{VW}\beta_i^{VW} + c_{prem}\beta_i^{prem} + c_{labour}\beta_i^{labour}$$

and for the model for the moments

$$E[R_{it}(\delta_0 + \delta_{VW}R_t^{VW} + \delta_{prem}R_{t-1}^{prem} + \delta_{labour}R_t^{labour})] = 1$$

with either a subset or all of the variables. Here, R_{it} is the price relative (P_t/P_{t-1}) for portfolio i ($i = 1, 2, \ldots, 49$) in quarter t (March 1982–December 2001), R_t^{VW} is the return on the value-weighted index of stocks, R_{t-1}^{prem} is the expected 10-year government bond yield in Australia based on US yields and inflation differentials between the two countries, and R_t^{labour} is the growth rate in per capita labour income. β_i^{VW} is the slope coefficient in the OLS regression of R_{it} on a constant and R_t^{VW}. The other betas are estimated in a similar way. The portfolio size, log (ME_i), is calculated as the equally weighted average of the logarithm of the market value (in millions of Australian dollars) of the stocks in portfolio i. The regression models are estimated using the Fama-MacBeth procedure. The corrected t- and p-values take sampling errors in the estimated betas into account. The models for the moments are estimated using the Generalized Method of Moments (GMM) with the Hansen-Jagannathan (HJ) weighting matrix. The minimized value of the GMM criterion function is the first item under the HJ-distance, with the associated p-value immediately below it. All the R-squares and p-values are reported as percentages, with all GMM p-values derived from a bootstrapping procedure.

Panels B and C of Table 3 present the results for the conditional CAPM. OLS analyses of the PL model without (Panel B), and with (Panel C) returns to human capital, do not save the market beta; the corrected *t*-statistics indicate that, in no instance, can the null hypothesis that c_{VW} equals zero be rejected. In Panels B and C, we see that c_{labour} is not significant. Rather, it is c_{prem} that is found to have a negative and statistically significant relationship to the cross-section of returns in Panels B and C. The negative relationship is perhaps surprising but consistent with the consumption smoothing approach in Breedon, 1979).[22] Overall, using the value of R^2 as a guide, the conditional CAPM reported in Panels B and C provide a better explanation of the cross section of returns than the CAPM.[23]

In contrast to Jagannathan and Wang (1996), who found that the value of the PL intercept in their data was too high to be consistent with a correctly specified model of returns, the estimate of 1.02 (equivalent to 8 per cent *per annum*) reported in the present paper is in keeping with the average risk-free rate over the period, which was 9.52 per cent *per annum* (the standard deviation of the average risk-free rate is estimated to be 4.3 per cent). All the intercepts calculated in the OLS regressions we report (Tables 3–6) are consistent with this analysis.

Despite the success of the OLS analyses, none of the GMM analyses reported in Table 3 provides support for the hypothesis that the factors we study function as the stochastic discount factor. In each case, the estimated coefficients for δ are not significantly different from zero. The measures of the *HJ*-distance indicate that the pricing errors are significantly different from zero. In contrast to the OLS test, the GMM approach tests strict arbitrage-free equilibrium conditions. Clearly, none of the models we study conforms to these strict conditions. Therefore, while our OLS analyses suggest that our models provide a good explanation for the cross-section of returns, we cannot conclude that these relationships are consistent with a no-arbitrage equilibrium determining the cross-section of returns.

In view of the prima facie evidence for the importance of size, we augment each of the OLS analyses in Table 3 with ME_i – the market value of firm *i*'s outstanding equity (in millions of Australian dollars). If the models are to be good descriptions of the cross-section of returns, the estimate of c_{size} should be insignificantly different from zero. Panel A shows that c_{size} is negative and significant when augmenting the CAPM. In both specifications of the PL model

[22] Consumption CAPM have not been extensively studied in Australia. Faff (1998) finds evidence for significant consumption betas using quarterly data, but utilizing monthly data, and a maximum correlation portfolio methodology, the results were inconclusive. Durand (1999), however, found no support for the inverse relationship of consumption to returns hypothesized by such models.

[23] Like Jagannathan and Wang (1996) we produced plots of the models reported in Table 3 to illustrate the goodness of fit. These plots may be obtained from http://www.ecom.uwa.edu.au/information_about/staff/durand_robert.

Table 4
Comparison with the factors used by Fama and French (1993)

Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{SMB}	c_{HML}	R-square
Estimate:	0.99	−0.01			0.13	0.01	71.70
Corrected-t:	102.96	−1.05			4.43	0.62	
Corrected-p:	0.00	30.09			0.01	53.68	
Coefficient:	δ_0^*	δ_{VW}^*	δ_{prem}^*	δ_{labour}^*	δ_{SMB}^*	δ_{HML}^*	HJ-distance*
Estimate:	6.80	5.80			−2.95	−8.19	2.11
p-value:	24.96	53.19			8.36	40.10	0.00
Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{SMB}	c_{HML}	R-square
Estimate:	0.99	−0.02	0.00	0.004	0.14	0.01	76.26
Corrected-t:	79.84	−1.09	−0.55	2.38	3.19	0.37	
Corrected-p:	0.00	28.12	58.75	2.18	0.27	71.04	
Coefficient:	δ_0^*	δ_{VW}^*	δ_{prem}^*	δ_{labour}^*	δ_{SMB}^*	δ_{HML}^*	HJ-distance*
Estimate:	57.99	5.23	42.82	−51.14	−3.36	−7.45	2.09
p-value:	34.44	54.21	34.14	36.54	12.91	41.27	0.00

This table gives the estimates for the cross-sectional regression model:

$$E[R_{it}] = c_0 + c_{VW}\beta_i^{VW} + c_{prem}\beta_i^{prem} + c_{labour}\beta_i^{labour} + c_{SMB}\beta_i^{SMB} + c_{HML}\beta_i^{HML}$$

and for the model for the moments

$$E[R_{it}(\delta_0 + \delta_{VW}R_t^{VW} + \delta_{prem}R_{t-1}^{prem} + \delta_{labour}R_t^{labour} + \delta_{SMB}SMB_t + \delta_{HML}HML_t)] = 1$$

with either a subset or all of the variables. Here, R_{it} is the price relative (P_t/P_{t-1}) for portfolio i ($i = 1, 2, \ldots, 49$) in quarter t (March 1990–December 2001), R_t^{VW} is the return on the All Ordinaries Accumulation Index, R_{t-1}^{prem} is the expected 10-year government bond yield in Australia based on US yields and inflation differentials between the two countries, R_t^{labour} is the growth rate in per capita labour income, and SMB_t and HML_t denote the respective Fama and French (1993) factors that are designed to capture the risks related to firm size and book-to-market equity. β_i^{VW} is the slope coefficient in the ordinary least squares (OLS) regression of R_{it} on a constant and R_t^{VW}. The other betas are estimated in a similar way. The regression models are estimated using the Fama-MacBeth procedure. The corrected t- and p-values take sampling errors in the estimated betas into account. The models for the moments are estimated using the Generalized Method of Moments (GMM) with the Hansen-Jagannathan (HJ) weighting matrix. The minimized value of the GMM criterion function is the first item under the HJ-distance, with the associated p-value immediately below it. All the R-square and p-values are reported as percentages, with all GMM p-values derived from a bootstrapping procedure. *Results are from quintile portfolio formation. As only 48 time intervals are available for SMB and HML construction, the number of portfolios under septile formation (49) exceeds time periods. In this situation the computer program for the exact GMM solution and bootstrapping fails.

Table 5
Comparison with the factors used by Chen *et al*. (1986)

Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{UTS}	c_{GDP}	c_{UI}	R-square
Estimate:	1.01	0.07			0.01	−0.01	0.00	38.41
Corrected-t:	26.34	1.10			1.19	−2.80	−0.56	
Corrected-p:	0.00	27.66			24.08	0.76	57.98	
Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	δ_{UTS}	δ_{GDP}	δ_{UI}	HJ-distance
Estimate:	−12.46	4.01			95.41	8.97	18.85	1.83
p-value:	41.52	49.85			47.03	40.04	46.85	0.17
Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{UTS}	c_{GDP}	c_{UI}	R-square
Estimate:	1.04	0.04	−0.03	−0.01	0.00	−0.01	0.00	78.15
Corrected-t:	26.75	0.65	−2.44	−1.00	−0.69	−1.93	0.16	
Corrected-p:	0.00	52.14	1.88	32.54	49.13	5.98	87.20	
Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	δ_{UTS}	δ_{GDP}	δ_{UI}	HJ-distance
Estimate:	−74.62	3.69	63.47	22.83	120.60	46.49	3.32	1.67
p-value:	36.44	48.21	52.47	40.36	45.04	37.81	46.28	0.83

This table gives the estimates for the cross-sectional regression model:

$$E[R_{it}] = c_0 + c_{VW}\beta_i^{VW} + c_{prem}\beta_i^{prem} + c_{labour}\beta_i^{labour} + c_{UTS}\beta_i^{UTS} + c_{GDP}\beta_i^{GDP} + c_{UI}\beta_i^{UI}$$

and for the model for the moments

$$E[R_{it}(\delta_0 + \delta_{VW}R_t^{VW} + \delta_{prem}R_{t-1}^{prem} + \delta_{labour}R_t^{labour} + \delta_{UTS}UTS_t + \delta_{GDP}GDP_t + \delta_{UI}UI_t)] = 1$$

with either a subset or all of the variables. Here, R_{it} is the price relative (P_t/P_{t-1}) for portfolio i ($i = 1, 2, \ldots, 49$) in quarter t (March 1982–December 2001), R_t^{VW} is the return on the All Ordinaries Accumulation Index, R_{t-1}^{prem} is the expected 10-year government bond yield in Australia based on US yields and inflation differentials between the two countries, R_t^{labour} is the growth rate in per capita labour income, UTS_t is the return spread between long-term government bonds and Treasury bills, GDP_t is the growth rate in Australia's Gross Domestic Product, and UI_t is the change in inflation rate. β_i^{VW} is the slope coefficient in the ordinary least squares (OLS) regression of R_{it} on a constant and R_t^{VW}. The other betas are estimated in a similar way. The regression models are estimated using the Fama-MacBeth procedure. The corrected t- and p-values take sampling errors in the estimated betas into account. The models for the moments are estimated using the Generalized Method of Moments (GMM) with the Hansen-Jagannathan (HJ) weighting matrix. The minimized value of the GMM criterion function is the first item under the HJ-distance, with the associated p-value immediately below it. All the R-square and p-values are reported as percentages, with all GMM p-values derived from a bootstrapping procedure.

Table 6
Is the USA a priced factor?

Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	$c_{VW\text{-}US}$	$c_{USD/AUD}$	R-square
Estimate:	1.04	−0.05			0.11	0.61	26.30
Corrected-t:	29.65	−0.76			1.60	2.25	
Corrected-p:	0.00	45.35			11.63	2.91	
Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	$\delta_{VW\text{-}US}$	$\delta_{USD/AUD}$	HJ-distance
Estimate:	1.64	6.30			−3.24	−2.81	1.83
p-value:	49.83	57.12			61.60	62.12	0.30
Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	$c_{VW\text{-}US}$	$c_{USD/AUD}$	R-square
Estimate:	1.03	−0.02	−0.03	−0.01	0.10	0.26	73.78
Corrected-t:	24.57	−0.29	−1.65	−0.71	1.27	1.21	
Corrected-p:	0.00	77.57	10.54	48.44	21.20	23.37	
Coefficient:	δ_0	δ_{VW}	δ_{prem}	δ_{labour}	$\delta_{VW\text{-}US}$	$\delta_{USD/AUD}$	HJ-distance
Estimate:	11.34	5.64	51.79	−11.42	−3.39	−1.75	1.77
p-value:	58.09	56.99	49.21	58.06	61.53	63.30	0.37

This table gives the estimates for the cross-sectional regression model

$$E[R_{it}] = c_0 + c_{VW}\beta_i^{VW} + c_{prem}\beta_i^{prem} + c_{labour}\beta_i^{labour} + c_{VW\text{-}US}\beta_i^{VW\text{-}US} + c_{USD/AUD}\beta_i^{USD/AUD}$$

and for the model for the moments

$$E[R_{it}(\delta_0 + \delta_{VW}R_t^{VW} + \delta_{prem}R_{t-1}^{prem} + \delta_{labour}R_t^{labour} + \delta_{VW\text{-}US}R_t^{VW\text{-}US} + \delta_{USD/AUD}e_t^{USD/AUD})] = 1$$

with either a subset or all of the variables. Here, R_{it} is the price relative (P_t/P_{t-1}) for portfolio i $(i = 1, 2, \ldots, 49)$ in quarter t (March 1982–December 2001), R_t^{VW} is the return on the All Ordinaries Accumulation Index, R_{t-1}^{prem} is the expected 10-year government bond yield in Australia based on US yields and inflation differentials between the two countries, R_t^{labour} is the growth rate in per capita labour income, $R_t^{VW\text{-}US}$ is the return on the Standard and Price 500 Composite Index in the USA, and $e_t^{USD/AUD}$ is the Australian/US exchange rate, expressed as a direct quote in Australia. β_i^{VW} is the slope coefficient in the ordinary least squares (OLS) regression of R_{it} on a constant and R_t^{VW}. The other betas are estimated in a similar way. The regression models are estimated using the Fama-MacBeth procedure. The corrected t- and p-values take sampling errors in the estimated betas into account. The models for the moments are estimated using the Generalized Method of Moments (GMM) with the Hansen-Jagannathan (HJ) weighting matrix. The minimized value of the GMM criterion function is the first item under the 'HJ-dist', with the associated p-value immediately below it. All the R-square and p-values are reported as percentages, with all GMM p-values derived from a bootstrapping procedure.

(Panels B and C), the null hypothesis that ME_i does not influence the cross-section of returns cannot be rejected.

5. Does the conditional CAPM outperform competing asset pricing models?

5.1. Comparison with factors used by Fama and French (1993)

To test Fama and French's (1993) three-factor model, time series returns are required for the SMB and HML factors. The SMB factor is constructed with stock returns from the Securities Industry Research Centre of Asia-Pacific's Core Research Database, and market capitalization figures from the Australian Graduate School of Management's SPPR Database. The sample of stocks is then ranked according to market capitalization, with the SMB measure defined as the discrete monthly return to a value-weighted portfolio of the smallest 30 per cent of stocks, minus the discrete monthly return to a value-weighted portfolio of the largest 30 per cent of stocks.[24] These portfolios are rebalanced each month, on the last trading day before the beginning of the successive month. The HML measure is constructed in a similar manner, from a similar dataset. Stocks are ranked according to each firm's ratio of book equity to market equity, with book value data (specifically, the book value of shareholders' equity less deferred taxes and the book value of deferred stock) provided by the Company Analysis Database from Datastream (and market equity values obtained from the SPPR database). As with the SMB measure, the portfolios used to *calculate* HML are rebalanced monthly.[25] Both SMB and HML cover the period March 1990 to December 2001 (as reliable data does not extend as far back as 1982, which marks the commencement of the sample for all other tests).[26]

The results of the analysis are reported in Table 4. In keeping with Jagannathan and Wang (1996), we first report the results for the three-factor model and then add the PL factors to the three-factor model. The insignificant coefficients and significant pricing errors estimated using GMM, are consistent with the results reported in Table 3. Therefore, the remaining discussion focuses on the OLS estimates. In both the OLS regressions, the estimated value for the intercept is statistically significant and consistent with our findings in Table 3.

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[24] Although this differs from the 50/50 split adopted in Fama and French (1993, pp. 8–9), the grouping we have adopted is common in the published literature (eg., Halliwell *et al.*, 1999) and is arguably more appropriate when comparing against the value-growth grouping.

[25] In contrast to Fama and French (1993, p. 8) we adopt this more active approach to facilitate comparison with SMB. To reduce the influence of any look-ahead bias, when we rank stocks at month t we use book values at time $t - 1$.

[26] SMB was found to have a monthly average price relative of 1.0412 and HML 1.0097. The correlation between SMB and HML was 0.42. Further summary statistics on these variables has been included with the analysis that may be downloaded from http://www.ecom.uwa.edu.au/information_about/staff/durand_robert.

In the analysis of the three-factor model, the hypothesis that the estimated coefficients for the market and HML betas are equal to zero cannot be rejected. c_{SMB} is positive and statistically significant. Therefore, our findings are consistent with those of Halliwell *et al*. (1999) and inconsistent with those of Faff (2001), although the time period we examine is coincident with the period examined by Faff rather than the period examined by Halliwell *et al*. It may be that the discrepancy is a result of differences in factor construction: Faff, as noted in Section 2, proxies HML (and SMB) through commercially available factors. The sensitivity of such analyses to the construction of HML is clearly a useful area for future research. The adjusted *t*-statistics indicate that SMB remains significant when the additional two variables, β^{prem} and β^{labour}, are added. In this case, rather than c_{prem}, c_{labour} is positive and statistically significant although the effect is small.

Table 4 provides no support for the three-factor model per se, but given the significance of c_{SMB} and the values of R^2 for both equations, there is strong support for the role of SMB as a determinant of the cross-section of returns. Using the value of R^2 as a guide, the three-factor model outperforms the PL model. The inclusion of c_{labour} is found to be statistically significant in the second OLS regression and there is an increase in the value of R^2, indicating that a combination of the three-factor model and the PL model supplies useful incremental explanatory power in explaining the cross-section of returns.

5.2. Comparison with arbitrage pricing theory factors

As with the early CAPM studies, development of the APT in Australia has followed a similar path to the research in the USA. Early work in Australia utilized data reduction techniques to determine the number of priced factors (Sinclair, 1984; Faff, 1988, 1992). Later, Groenewold and Fraser (1997) followed an approach similar to Chen *et al*. (1986) through an investigation of whether pre-specified variables were priced.

Results comparing the PL model with the factors of Chen *et al*. are reported in Table 5.[27] The OLS intercept values and findings for the GMM analysis are consistent with previously reported analyses. Hence, we again focus on the results of the OLS analyses.[28]

There is little evidence that the Chen *et al*. factors are priced in Australia as they are in the USA. Table 5 indicates that only the gross domestic product (GDP) beta exhibits a significant coefficient (at the 0.76 per cent level after

[27] An analysis comparing the PL model with Groenewold and Fraser's factors may be downloaded from http://www.ecom.uwa.edu.au/information_about/staff/durand_robert. We find that the PL model is superior to Groenewold and Fraser's realization of the APT.

[28] Australian inflation rates are calculated from the CPI time-series provided by the ABS (the same time-series used in the calculation of R_t^{prem}), 10-year government bonds and 13-week Treasury Bills are obtained from the Reserve Bank of Australia (RBA) website, and seasonally adjusted GDP figures are available from the Australian Bureau of Statistics.

correcting for beta sampling errors). The model R^2 is lower than that reported for the PL models in Table 3. The addition of the premium and labour betas increases the p-value of c_{GDP} to 5.98 per cent and the model R^2 increases to 78.15 per cent. Once again there is a significant negative loading on β_i^{prem}. The negative coefficients for both c_{GDP} and c_{prem} are consistent with the consumption-smoothing hypothesis discussed in Section 4.

5.3. Is the USA a priced factor?

There is evidence that movements in the US markets influence returns in other markets around the world.[29] Ragunathan *et al.* (1999) found that Australian and US returns are related, but that the relationship is sensitive to the stage of the business cycle. Durand *et al.* (2001) found that variations in the US market explain over 20 per cent of the daily variance of the Australian market and that Australian returns are Granger-caused by movements in the US market. Durand and Scott (2003) suggest that the relationship of the Australian to the US market may be consistent with investors overreacting to US market movements. Given these findings, we examine the influence of US market movements over the cross-section of returns in Australia. Table 6 reports the analysis, including both US value-weighted index movements and the US/Australian exchange rate in the cross-sectional regression.[30,31] The OLS intercepts and findings for the GMM analysis are consistent with previously reported analyses.

The results reported in Table 6 show that only the exchange rate is significant when added to the CAPM. When premium and labour betas are added, the exchange rate effect disappears. As might be expected, the explanatory variables are highly correlated: the correlation between R_t^{VW} and $R_t^{VW\text{-}US}$ is 0.60. The correlation between β_i^{VW} and $\beta_i^{VW\text{-}US}$ is 0.93. Given that the Australian market beta is found to have no role in explaining the cross-section of returns, the finding that the strongly related US beta also has no role in explaining the cross-section of returns should not be surprising.

5.4. The cross-section of excess returns

A final test conducted by Jagannathan and Wang (1996) was to determine how well the PL model accords with an important outcome of the Black version

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[29] See, for example, Eun and Shim (1989), Theodossiou and Lee (1993), Phylaktis (1997), and Ghosh *et al.* (1999).

[30] Loudon (1993) and Di Iorio and Faff (2002) provide evidence that the exchange rate has had a statistically significant relationship to Australian equity returns.

[31] A second analysis includes only the value-weighted index of US stocks, but the measure is expressed in Australian dollars. The findings are consistent with those reported in this paper. The analysis may be downloaded from http://www.ecom.uwa.edu.au/information_about/staff/durand_robert.

of the conditional CAPM, which predicts an intercept term equal to the return on a zero-beta portfolio (or equal to the risk free rate r_f if borrowing and lending is available at such a rate). This zero-beta portfolio return should lie somewhere between riskless lending and borrowing rates. Hence a test is conducted to determine whether the empirically derived intercept term lies between what could plausibly be a riskless lending and borrowing rate.

The results of the analysis are reported in Table 7. The intercept term is statistically insignificantly different from zero in the OLS estimate of the PL model. Not only does this test suggest that the zero-beta and risk-free rates are similar, but that one is not statistically distinguishable from the other. In the OLS estimation using the Fama-French factors, the null hypothesis that the intercept equals zero can be rejected at a significance level of 5.44 per cent and the results for c_{SMB} are also in keeping with those reported in Table 4. The marginal evidence of significance of a negative intercept suggests that the risk-free rate of return may be slightly higher than the zero beta return.

The results for the GMM estimates for both models provide a clear contrast to the analyses reported in Tables 3–6. The models are clearly sensitive to using excess returns rather than raw returns and, in these cases, the strict no-arbitrage restrictions are satisfied by both formulations to within the accuracy of the sample estimation. In both the PL and the three-factor estimates, pricing errors are insignificantly different from zero, although the PL model is to be preferred with a *HJ*-distance of 0.01 (compared to a *HJ*-distance of 0.05 for the Fama-French model). The null hypothesis that the coefficients for the three estimated values of delta equal zero can be rejected with confidence for the PL model. In the Fama-French formulation, the estimate of β_{HML} is also significantly different from zero but the hypothesis that β_{SMB} equals zero cannot be rejected. Such findings are seemingly at odds with those reported for the Fama/French model in Table 4 and suggest that further research into the role of HML is warranted with Australian data.

By using excess returns, the variation in the risk-free rate through time is implicitly incorporated into the specification, as the returns at each time period are calculated using the corresponding risk-free rate for each particular time period. Hence, the GMM acceptance of the stochastic discount factor for excess returns, and rejection of the stochastic discount factor for gross returns, could well be a reflection of the non-stationarity in parameters. When the risk-free rate is permitted to vary through time, the strict no-arbitrage equilibrium of the GMM is satisfied by the data, but when the non-stationary parameters are forced to assume a constant value (through the tests in all tables other than Table 7) the equilibrium relation is rejected.

6. Conclusion

Jagannathan and Wang (1996) develop and test a conditional CAPM that makes specific allowance for the fact that portfolios' betas may vary systematically.

Table 7
Tests using the time series of quarterly excess returns on the 49 size-beta sorted portfolios

Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{SMB}	c_{HML}	R-square
Estimate:	−0.01	0.05	−0.03	−0.01			68.56
Corrected-t:	−0.15	0.71	−2.03	−0.92			
Corrected-p:	88.22	48.44	4.83	36.04			
Coefficient:	$\tilde{\delta}_{VW}$	$\tilde{\delta}_{prem}$	$\tilde{\delta}_{labour}$	$\tilde{\delta}_{SMB}$		$\tilde{\delta}_{HML}$	HJ-distance
Estimate:		−0.02	0.45	−0.98			0.01
P-value:		0.69	0.00	0.00			30.14
Coefficient:	c_0	c_{VW}	c_{prem}	c_{labour}	c_{SMB}	c_{HML}	R-square
Estimate:	−0.02	−0.02			0.13	0.01	72.13
Corrected-t:	−1.98	−1.37			4.43	0.54	
Corrected-p:	5.44	17.80			0.01	58.93	
Coefficient:	$\tilde{\delta}^*_{VW}$	$\tilde{\delta}^*_{pren}$	$\tilde{\delta}^*_{labou}$	$\tilde{\delta}^*_{SMB}$		$\tilde{\delta}^*_{HML}$	HJ-distance
Estimate:		0.29			−0.07	−0.89	0.05
p-value:		14.55			18.00	0.00	52.05

This table gives the estimates for the following two regression models:

$$E[\tilde{R}_{it}] = c_{VW}\tilde{\beta}_i^{VW} + c_{prem}\tilde{\beta}_i^{prem} + c_{labour}\tilde{\beta}_i^{labour}$$

$$E[\tilde{R}_{it}] = c_{VW}\tilde{\beta}_i^{VW} + c_{SMB}\tilde{\beta}_i^{SMB} + c_{HML}\tilde{\beta}_i^{HML}$$

and for the two models for the moments

$$E[\tilde{R}_{it}(1 + \tilde{\delta}_{VW}R_t^{VW} + \tilde{\delta}_{prem}R_{t-1}^{prem} + \tilde{\delta}_{labour}R_t^{labour})] = 0$$

$$E[\tilde{R}_{it}(1 + \tilde{\delta}_{VW}\tilde{R}_t^{VW} + \tilde{\delta}_{prem}R_{t-1}^{prem} + \tilde{\delta}_{labour}R_t^{labour})] = 0$$

Here, $\tilde{R}_{it} = R_{it} - R_t^{TBill}$ where R_{it} is the price relative (P_t/P_{t-1}) on portfolio i ($i = 1, 2, \ldots, 49$) in quarter t (March 1982–December 2001 for the 1st model, and March 1990–December 2001 for the second) and R_t^{Tbill} is the return on the T-bill. R_t^{VW} is the return on the All Ordinaries Accumulation Index and $\tilde{R}_t^{VW} = R_t^{VW} - R_t^{TBill}$. R_{t-1}^{prem} is the expected 10-year government bond yield in Australia based on US yields and inflation differentials between the two countries, R_t^{labour} is the growth rate in per capita labour income, and SMB_t and HML_t denote the respective Fama and French (1993) factors that are designed to capture the risks related to firm size and book-to-market equity. $\tilde{\beta}_i^{VW}$ is the slope coefficient in the ordinary least squares (OLS) regression of $\tilde{R}_{it}$ on a constant and R_t^{VW}. The other $\tilde{\beta}$s are estimated in a similar way. $\tilde{\beta}_i^{VW}$ is the slope coefficient in the OLS regression of $\tilde{R}_{it}$ on a constant and $\tilde{R}^{VW}$. The regression models are estimated using the Fama-MacBeth procedure. The corrected t- and p-values take sampling errors in the estimated betas into account. The models for the moments are estimated using the Generalized Method of Moments (GMM) with the Hansen-Jagannathan (HJ) weighting matrix. The minimized value of the GMM criterion function is the first item under the HJ-distance, with the associated p-value immediately below it. All the R-square and p-values are reported as percentages, with all GMM p-values derived from a bootstrapping procedure.
*Results are from quintile portfolio formation. As only 48 time intervals are available for SMB and HML construction, the number of portfolios under septile formation (49) exceeds time periods. In this situation the computer program for the exact GMM solution and bootstrapping fails.

Their empirical examination of the model using US data is supportive of this hypothesis. We have followed Jagannathan and Wang, with some relatively minor modifications and extensions, in testing the (PL) formulation of the conditional CAPM with Australian data. Our OLS analyses provide evidence that the PL model successfully describes the cross-section of returns, although it is clear that it may be improved upon by adding other variables that have been shown to have explanatory power. Our GMM analyses find the model to be inconsistent with the strict arbitrage-free conditions required by a stochastic discount model of returns when the risk-free rate is assumed to be constant. In our analysis of excess returns (Section 5.4), however, both the PL model and the three-factor model are consistent with stochastic-discount explanations for returns.

In contrast to Jagannathan and Wang's analysis of US data, we do not save the market beta through the inclusion of a proxy for the returns to human capital. When it does achieve significance in our analysis of the PL model, the human capital effect is negative rather than positive, suggesting that it functions as a state variable for current consumption as postulated by Breeden (1979) rather than as a market risk premium (in this latter case we would expect the effect to be positive). This variable is sensitive to the inclusion of other explanatory variables: its significance disappears when we add β^{prem} to the analysis. It is β^{prem}, derived from the US long-term interest rate adjusted for inflation differentials between the two countries, that drives the explanatory power of the model in almost all the situations we study.

Using the value of R^2 as a guide, we find that the PL model outperforms the APT, although it slightly underperforms the three-factor model. The latter finding is driven by the significance of SMB: we do not find a role for HML, although our analysis of zero-beta formulations suggests that the role of this factor requires further research for its full explication. When comparing the PL and three-factor models, Jagannathan and Wang argue that SMB and HML proxy for the multiple sources of market risk overtly specified in the PL model. The evidence in the present paper points to a role for SMB in an augmented PL model. It appears that more influence can be ascribed to SMB than as just acting as a proxy for mispriced market risk. We depart from Jagannathan and Wang's analysis in that we also consider a role for the US market in determining the cross-section of returns. We find that the PL model outperforms our US-market variables although, given the strong role of β^{prem}, we suggest that future research is required to address whether the Australian market is merely a vassal of the dominant American market.

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