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Asymmetric adjustment of stock prices to their fundamental value and the predictability of US stock returns

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Abstract

Using momentum threshold autoregressive (MTAR) models, we explore the adjustment mechanism between US stock prices and fundamentals. Next, we derive non-linear error-correction models and examine whether they can improve forecasts of stock returns. Results support MTAR model only for data at annual frequency.
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1. Introduction

In a precedent work (Boucher, 2006), we estimated the common long-term trend in the US earning–price ratio and inflation and found that the *linear* transitory deviations from this common trend exhibit substantial out-of-sample forecasting abilities for excess returns at short and intermediate horizons. Further, some recent studies (e.g. Coakley and Fuertes, 2006) have documented non-linearities in the U.S. equity ratios. Then, we investigate whether a non-linear model can improve forecasts of stock returns.

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We apply a momentum threshold autoregressive (MTAR) procedure designed to detect asymmetric short-run adjustments to the long run equilibrium between aggregate stock prices and fundamentals. This specification is especially relevant when the adjustment is such that the series exhibits more "momentum" in one direction than the other as it is the case when stock prices are affected by speculative bubbles. US Annual data from 1871 and quarterly data from 1953 are used to explore the adjustment mechanism between stock prices and fundamentals.

This paper proceeds as follows: Section 2 describes the MTAR model. Data description and symmetric/asymmetric cointegration test results are provided in Section 3. Section 4 presents estimates of the symmetric and asymmetric error-correction models. Section 5 reports the out-of-sample forecasting exercises and some simple trading strategies. Section 6 offers some concluding remarks.

2. The MTAR model

Enders and Siklos (2001) extend the Engle and Granger (1987) framework to test for non-linear cointegration. The MTAR model is given as:

$$\Delta\hat{\mu}_t = M_t\rho_1\hat{\mu}_{t-1} + (1-M_t)\rho_2\hat{\mu}_{t-1} + \varepsilon_t, \tag{1}$$

$$M_t = \begin{cases} 1, & \text{if } \Delta\hat{\mu}_{t-1} \geq \tau \\ 0, & \text{if } \Delta\hat{\mu}_{t-1} < \tau \end{cases} \tag{2}$$

where $\hat{\mu}$ are the estimated residuals from the estimated cointegration relationship. M_t is the zero–one Heaviside indicator function, τ denotes the value of the threshold.[1]

The no cointegration hypothesis (H_0: $\rho_1 = \rho_2 = 0$) is tested using specifically derived critical values provided by ES (2001). The statistic testing this null hypothesis is denoted Φ. If the null hypothesis of no cointegration is rejected, the null hypothesis of symmetric adjustment (H_0: $\rho_1 = \rho_2$) can be tested using a standard F-test.

3. Data and cointegration evidence

We use two sets of data. The first consists of annual observations spanning 1871–2002 and the second consists of quarterly observations from 1953:Q2 to 2003:Q2. Stock prices, dividends per share, and quarterly earnings per share all correspond to the Standard & Poor's (S&P) Composite Index.[2] The inflation rate, i_t, is the percentage change in the Consumer Price Index (All Urban Consumers) published by the BLS and prior to 1913 by Robert Shiller. The T-bill rate is available from the FRED II database of

[1] We follow ES (2001), who used Chan's (1993) grid search method to find a consistent estimate of the threshold.

[2] The S&P data are available from Robert Shiller's home page at http://www.econ.yale.edu/~shiller. The complete documentation for the data sources is also provided here. Data are updated from the Standard & Poor's web site (S&P 500 Earnings and Estimate Report).

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the Federal Reserve Bank of St. Louis Real. Let r_t denote the nominal return on the S&P index. The three month T-bill rate, $r_{f,t}$, is used to construct the log excess return $(r_t - r_{f,t})$. Log price, p_t, is the natural logarithm of the S&P price level in year (quarter) t. Log dividends, d_t, are the natural logarithm of dividends per share in year (quarter) t. Log earnings, e_t, are the natural logarithm of earnings per share in year (quarter) t.

Preliminary unit root tests indicate that p_t, d_t and e_t as well as $e_t - p_t$, $d_t - p_t$ and i_t can be characterized as I(1) processes. However, we assume that the non-stationarity of the (log) dividend–price and earning–price ratios reflects the highly persistent behavior of expected returns due to an inflation conditional risk premium rather than the presence of an explosive rational bubble (see Boucher, 2006).

We apply linear and non-linear cointegration tests to test the present value model with a long-term risk premium and investigate the relation between valuation ratios and the inflation rate. We used not only dividends as a primary measure of fundamentals but also earnings since dividends can be affected by shifts in corporate financial policy, that it creates several difficulties for empirical work.[3] Moreover, we also estimate non-restricted versions of the present value models that do not constrained the elasticity of stock prices with respect to dividends or earnings to be equal to one. Specifically, we consider the following six models:

$$d_t - p_t = \alpha_0 + \mu_t \tag{3}$$

$$e_t - p_t = \alpha_0 + \mu_t \tag{4}$$

$$d_t - p_t = \alpha_0 + \alpha_1 i_t + \mu_t \tag{5}$$

$$e_t - p_t = \alpha_0 + \alpha_1 i_t + \mu_t \tag{6}$$

$$p_t = \alpha_0 + \alpha_1 i_t + \alpha_2 d_t + \mu_t \tag{7}$$

$$p_t = \alpha_0 + \alpha_1 i_t + \alpha_2 e_t + \mu_t \tag{8}$$

We use the estimated residuals of each model, $\hat{\mu}_t$, to estimate a MTAR model and implement the ADF t-test of the Engle–Granger procedure.[4] Tables 1 and 2 consider the linear and MTAR models respectively for the annual and quarterly samples. The estimated long-run equilibrium relationships indicates that (i) the inflation rate coefficient is positive for the models (3) to (6) and negative for

[3] See, e.g. Campbell (2000).
[4] The lag-length is determined via Hall's (1994) general to specific approach starting with a maximum lag-length and using the 5% significance level.

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Table 1
Symmetric and asymmetric cointegration tests (1871–2002)

Model	(3)	(4)	(5)	(6)	(7)	(8)
	d_t-p_t	e_t-p_t	$d_t-p_t\ i_t$	$e_t-p_t\ i_t$	$p_t\ i_t\ d_t$	$p_t\ i_t\ e_t$
$[\alpha_i]$	[−3.11]	[−2.60]	[−3.11 −0.20]	[−2.64 −1.78]	[3.11 0.93 −1.17]	[2.62 −2.11 1.05]
Engle Granger						
ADF *t*-test	−1.04	−3.26*	−1.08	−3.99***	−2.87	−4.61***
MTAR						
$\hat{\tau}$	0.0937	0.0980	−0.1608	−0.2533	0.1640	0.2607
ρ_1	0.02	−0.37	−0.08	−0.16	0.025	−0.59
	(−0.16)	(−4.60)	(−2.10)	(−2.16)	(0.26)	(−4.45)
ρ_2	−0.05	−0.04	0.09	−0.45	−0.20	−0.16
	(−1.21)	(−0.52)	(1.34)	(−4.06)	(−3.52)	(−1.77)
Φ	0.75	10.7***	3.09	10.55***	6.24*	10.29***
$\rho_1=\rho_2$	0.37	9.48***	4.97**	4.74**	4.03**	8.94***

Note: $[\alpha_i]$ denote estimated long-run coefficients from models (4.1) to (4.6). ADF *t*-test indicates *t*-statistics of the augmented Dickey–Fuller test. MTAR is defined by Eqs. (1) with (2). $\hat{\tau}$ denotes the (consistently) estimated threshold (Chan, 1993); ρ_1 and ρ_2, the estimated parameters of the MTAR model with *t*-statistics in parentheses. Φ and $\rho_1=\rho_2$ denote the *F*-statistics for the null hypothesis of no cointegration and symmetry respectively. The lag lengths are selected using the general to specific procedure (Hall, 1994). *, ** and *** denote significance at 10%, 5% and 1% significance level (MacKinnon, 1991; Enders and Siklos, 2001).

models (7) and (8) as expected and; (ii) stock prices move more than one-for-one with dividends or earnings.[5]

For the annual sample, only models with earnings and the inflation risk premium (6) and (8) appear to be cointegrating relationships according to the Engle–Granger test at a 1% level. The *F* statistic rejects the null hypothesis of symmetric adjustment toward the long-run equilibrium at a 5% significance level for models (4), (6) and (8). However, the *t*-value of $\hat{\rho}_2$ in model (4) strongly exceeds the 10% critical value. Comparison of the absolute values of the speeds of adjustment support the hypothesis that stock prices exhibit run-ups followed by crashes in the short-run.

For the quarterly sample, the Engle–Granger test indicates that only the models (6) and (8) are cointegrating relationships. The value of the Φ-statistic confirms that stock prices, earnings and inflation are cointegrated with and without the elasticity constraint. The Φ-statistic also indicates that stock prices, dividends and inflation are cointegrated (model (7)). Comparisons of the absolute values of $\hat{\rho}_1$ and $\hat{\rho}_2$ in the models (7) and (8) suggest convergence such that the speed of adjustment is faster for positive than for negative discrepancies from the threshold estimated. However, the *F* statistic rejects the null hypothesis of symmetric adjustment toward the long-run equilibrium only at a 10% level. For the model (6), while the comparison of the absolute values of speeds of adjustment suggests asymmetric convergence, the *F* statistic is less than the 10% critical value. Thus, for the quarterly sample the asymmetric adjustment mechanism is less convincing.

[5] Wald tests reject the hypothesis that the elasticity of stock prices with respect to dividends or earnings is equal to one, a feature that is broadly consistent with a model in which investors extrapolate short-run trends when they estimate the time-varying long-run dividend/earning growth rate (see, Barsky and De Long, 1993).

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Table 2
Symmetric and asymmetric cointegration tests (1953:Q2–2003:Q2)

Model	(3)	(4)	(5)	(6)	(7)	(8)
	$d_t - p_t$	$e_t - p_t$	$d_t - p_t\ i_t$	$e_t - p_t\ i_t$	$p_t\ i_t\ d_t$	$p_t\ i_t\ e_t$
$[\alpha_i]$	[−3.45]	[−2.75]	[−3.59 15.78]	[−2.98 24.21]	[3.17 −17.85 1.28]	[2.61 −26.91 1.17]
Engle Granger						
ADF *t*-test	−1.53	−2.16	−2.01	−4.07**	−3.62	−4.97***
MTAR						
$\hat{\tau}$	0.0561	0.0160	0.0075	0.0257	0.1475	0.1941
ρ_1	0.03	−0.02	−0.00	−0.06	−0.19	−0.27
	(0.85)	(−0.76)	(−0.02)	(−1.00)	(−2.51)	(−2.98)
ρ_2	−0.04	−0.05	−0.09	−0.13	−0.06	−0.11
	(−2.20)	(−2.23)	(−2.65)	(−2.67)	(−1.55)	(−2.02)
Φ	2.79*	2.76*	3.51*	3.8**	4.23**	5.89***
$\rho_1 = \rho_2$	3.22*	0.87	2,92*	1.22	2.60*	2.65*

Note: See Table 1.

We next consider the MTAR model using the annual sample 1951–2002 period.[6] For the models (6) and (8), $\hat{\tau}$ is estimated to be −0.22 and 0.18 respectively. The sample value of Φ is 8.53 and 11.08 and the *F*-statistic for the null hypothesis $\rho_1 = \rho_2$ is 9.58 and 8.94 with a *p*-value of nearly one percent. The hypotheses $\rho_1 = \rho_2 = 0$ and $\rho_1 = \rho_2$ are both soundly rejected. The point estimates for $\rho_2 = -0.10$ and $\rho_2 = -0.59$ (respectively −0.84 and −0.24) suggest that negative (positive) discrepancies from long-run equilibrium are eliminated rather quickly but that others are allowed to persist. The adjustment mechanism in then asymmetric on the annual sample 1951–2002. This suggests that, for U.S. stock prices, an accumulation of changes during *several quarters* in the disequilibrium relationship between stock prices and fundamentals above the threshold is followed by a sharp movement back to the equilibrium position both on the 1871–2002 and 1951–2002 samples.

4. Estimated error-correction models

The positive findings of cointegration with MTAR adjustment between stock prices, earnings and inflation, for at least the annual sample, justify estimation of error-correction models. Since the coefficient on earnings is statistically different from one, we do not impose any restriction on this coefficient. The estimated error-correction equations with MTAR adjustment respectively for annual (1871–2002) and quarterly (1953:Q2–2003:Q2) data are (with *t*-statistics in parentheses):

$$\Delta p_t^a = \underset{(2.736)}{0.041} - \underset{(-2.151)}{0.171}\, M_t \hat{\mu}_{t-1}^a - \underset{(-3.120)}{0.168}\,(1 - M_t)\hat{\mu}_{t-1}^a + \varepsilon_t, \quad \overline{R}^2 = 0.104 \tag{9}$$

$$\Delta p_t^q = \underset{(3.235)}{0.017} - \underset{(-2.026)}{0.072}\, M_t \hat{\mu}_{t-1}^q - \underset{(-3.226)}{0.064}\,(1 - M_t)\hat{\mu}_{t-1}^q + \underset{(1.384)}{0.097}\,\Delta p_{t-1}^q + \underset{(-1.550)}{1.195}\,\Delta i_{t-1}^q + \varepsilon_t, \quad \overline{R}^2 = 0.077 \tag{10}$$

[6] Not shown on tables.

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where $\hat{\mu}^a_{t-1} = p^a_t - 1.05\ e^a_t + 2.11\ i^a_t - 2.62$ and $\hat{\mu}^q_{t\text{-}1} = p_t - 1.17\ e^q_t + 26.91\ i^q_t - 2.61$. a and q denote respectively annual and quarterly samples. The Heaviside indicator, M_t, is set according to Eq. (2) and the consistent estimates of τ that appear in Tables 1 and 2.

The error correction terms are statistically significant and the associated coefficients are negative. However, adjustment coefficients, ρ_1 and ρ_2, do not appear noticeably different.

The symmetric error correction models estimated are quite similar ($\bar{R}^2$ statistics and estimated coefficients) to the asymmetric ones on both samples (Eqs. (11) and (12)). The error correction models with MTAR adjustment does not improve stock prices forecasts.

$$\Delta p^a_t = \underset{(2.846)}{0.041} - \underset{(-3.875)_t}{0.169}\ \hat{\mu}^a_{t-1} + \varepsilon_t, \quad \bar{R}^2 = 0.104 \tag{11}$$

$$\Delta p^q_t = \underset{(3.247)}{0.017} - \underset{(-3.729)}{0.066}\ \hat{\mu}^q_{t-1} + \underset{(1.384)}{0.096}\ \Delta p^q_{t-1} + \underset{(-1.544)}{1.185}\ \Delta i^q_{t-1} + \varepsilon_t, \quad \bar{R}^2 = 0.076. \tag{12}$$

We report next estimates from symmetric and asymmetric error-correction models that include control variables known to contain predictive power for stock returns, namely the consumption–wealth ratio, cay_t, and the relative T-bill rate, $rrel_t$:[7]

$$\begin{aligned} \Delta p^q_t = {} & \underset{(3.592)}{0.01} - \underset{(-2.026)}{0.077}\ M_t\hat{\mu}^q_{t-1} - \underset{(-2.196)}{0.045}\ (1-M_t)\hat{\mu}^q_{t-1} + \underset{(2.604)}{1.105}\ cay^q_{t-1} - \underset{(-2.563)}{1.229}\ rrel^q_{t-1} \\ & + \underset{(1.198)}{0.082}\ \Delta p^q_{t-1} + \underset{(-2.117)}{1.665}\ \Delta i^q_{t-1} + \varepsilon_t, \quad \bar{R}^2 = 0.146 \end{aligned} \tag{13}$$

$$\begin{aligned} \Delta p^q_t = {} & \underset{(3.518)}{0.018} - \underset{(-2.799)}{0.052}\ \hat{\mu}^q_{t-1} + \underset{(2.611)}{1.107}\ cay^q_{t-1} - \underset{(-2.449)}{1.447}\ rrel^q_{t-1} + \underset{(1.211)}{0.083}\ \Delta p^q_{t-1} \\ & + \underset{(-2.032)}{1.584}\ \Delta i^q_{t-1} + \varepsilon_t, \quad \bar{R}^2 = 0.143. \end{aligned} \tag{14}$$

The points estimate on cay_t and $rrel_t$ are strongly significant, the $\bar{R}^2$ statistic increases to 15% and, noticeably, the spread between error correction coefficients increases. However, the predictive power of the asymmetric error-correction model as measured by the $\bar{R}^2$ statistic is still similar to the symmetric one.

5. Check the robustness: out-of-sample forecasts and economic values of market timing

Goyal and Welch (2004) cast doubt on the in-sample evidence of stock return predictability because they find that the predictive variables used by the early authors have negligible out-of-sample predictive power. Ferson, Sarkissian, and Simin (2003) have also cautioned about spurious regression and data mining problems. To address these issues, we conduct out-of-sample analysis. We denote, pei_t, the deviation of (log) stock prices from its predicted value based on the cointegrating relationship between stock prices, earnings and price inflation estimated recursively.[8]

[7] Lettau and Ludvigson (2001) find that they are strong predictors of stock market returns and subsume the information content of many others macro-financial variables.

[8] To estimate the coefficients of the cointegration relationship, we employ the Dynamic Ordinary Least Squares (DOLS) developed by Stock and Watson (1993), a method that generates optimal estimates of the cointegrating parameters in a multivariate setting.

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Table 3
One-quarter-ahead forecasts of excess returns: nested comparisons

Row	Comparison unrestricted vs. restricted	MSE_u/ MSE_r	*ENC-NEW* Statistic	*MSE-F* Statistic
Panel A: Stock returns (excluding dividends)				
1	pei_{t-1} vs. *const*	0.9883	3.606**	1.462*
2	pei_{t-2} vs. *const*	0.9866	5.215***	1.677*
3	pei_{t-1} vs. *AR*	0.9818	4.736***	2.393**
4	pei_{t-2} vs. *AR*	0.8768	4.975***	3.239**
Panel B: Excess returns				
5	pei_{t-1} vs. *const*	0.9853	4.147**	1.959**
6	pei_{t-2} vs. *const*	0.9869	5.769***	1.688*
7	pei_{t-1} vs. *AR*	0.9771	5.362***	3.024**
8	pei_{t-2} vs. *AR*	0.9771	5.111***	3.230**

Note: The nested comparisons are made by alternately augmenting the benchmark with either the one-period lagged value of pei_t, or the two-period lagged value. The ENC-NEW statistic is used to test the null hypothesis that restricted model forecasts encompass the unrestricted model forecasts. The MSE-F statistic is used to test the null hypothesis that the MSE for the restricted model forecasts is less than or equal to the MSE for the unrestricted model forecasts. We estimate the cointegration parameters recursively. Observations from the period 1953:Q2 to 1968:Q1 are used to obtain the initial in-sample estimation, and the forecasting error is calculated for the remaining period 1968:Q2 to 2003:Q2, recursively. A * (**) (***) denotes significance at the ten (five) (one) percent level.

Table 4
Annualized returns and switching strategies

	Buy and hold	Switching strategy			
		without Transaction Costs		with Transactions Costs	
		pei_{t-1}	pei_{t-2}	pei_{t-1}	pei_{t-2}
Panel A: 1968:Q2–2003:Q2					
Mean	0.1052	0.1207	0.1119	0.1196	0.1105
SD	0.0769	0.0691	0.0669	0.0691	0.0671
Panel B: 1968:Q2–1979:Q4					
Mean	0.0444	0.0444	0.0444	0.0444	0.0444
SD	0.0833	0.0833	0.0833	0.0833	0.0833
Panel C: 1980:Q1–1989:Q4					
Mean	0.1735	0.1987	0.1907	0.1981	0.1901
SD	0.0787	0.0649	0.0641	0.0649	0.0642
Panel D: 1990:Q1–2003:Q2					
Mean	0.0981	0.1213	0.1035	0.1189	0.1003
SD	0.0701	0.0567	0.0504	0.0568	0.0508

Note: The table reports returns on switching strategies, which require holding stocks if the predicted excess return is positive and holding bonds otherwise. We assume that investors have to pay a proportional transaction cost of 25 basis points when they switch from stocks to bonds or vice versa.

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We use three statistics to compare the out-of-sample performance of our forecasting model with a benchmark of constant expected returns and a benchmark of autoregressive returns: the mean squared forecasting error (MSE) ratio, the Clark and McCracken's (2001) encompassing test (ENC-NEW) and the McCracken's (2004) equal forecast accuracy test (MSE-F). Moreover, we present out-of-sample predictability results using the two-period lagged value of pei_t because this variable is available with a one-quarter delay. We use the first one-third observations for the initial in-sample estimation and form the out-of-sample forecast recursively in the remaining sample. Since symmetric and asymmetric error-correction regressions yield similar qualitative results at a one-quarter horizon, we only consider linear forecasting models.

We report results of the out-of-sample one-quarter-ahead nested forecast comparisons of both stock returns (excluding dividends) and excess returns in Table 3. We find that the unrestricted model (which include pei_t) has smaller MSE than the constant restricted model or the autoregressive restricted model. Regardless whether the one- or two-period lagged value of pei_t is used as a predictive variable, both *ENC-NEW* and *MSE-F* tests reject the null hypothesis that pei_t provides no information about future stock returns at better than the five percent level in almost every case.

We also investigate whether we can exploit the forecasting ability of pei_t using a simple trading strategy, i.e., hold stocks if the predicted excess return is positive and holding bonds otherwise. We consider a proportional transaction cost of 25 basis points. Table 4 reports that over the period 1968:Q2 to 2003:Q2, managed portfolios have returns of higher mean and lower standard deviation than those of the buy-and-hold strategy.[9]

6. Conclusion

In this paper, we applied symmetric and asymmetric (MTAR) cointegration tests to the present value model with an inflation conditional time-varying long-term risk premium for U.S. stocks. MTAR test results indicate that the adjustment mechanism is asymmetric on annual samples (1871–2002 and 1951–2002) and provide support for the hypothesis that, in the short-run, stock prices exhibit run-ups followed by crashes. For the quarterly sample (1953:Q2–2003:Q2), however, the asymmetric adjustment mechanism is less convincing. This suggests that an accumulation of changes *during several quarters* in the disequilibrium relationship between stock prices and fundamentals above the threshold is followed by a sharp movement back to the equilibrium position. A promising direction of future research would be to investigate whether an asymmetric model with a long difference threshold variable, such as Caner and Hansen (2001) proposes, instead of a first difference one can identify the timing of the adjustment mechanism.

In a second step, we used symmetric and asymmetric error-correction models to predict stock market returns (excluding dividends). We found that *linear* deviations from the common trend in stock prices, earnings and realized inflation predict quarterly stock price fluctuations at a one-quarter horizon even if we include control variables as proxies for time-varying expected returns. To assess the robustness of our result, we show that the out-of-sample predictability of stock market returns is both statistically and economically significant.

[9] Except for the period 1968:Q2–1979:Q4 for which the forecasting model does not recommend to rebalance the managed portfolio.

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