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A contribution to event study methodology with an application to the Dutch stock market

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A Contribution to Event Study
Methodology with an Application
to the Dutch Stock Market

by

Frank de Jong, Angelien Kemna

and

Teun Kloek

Reprinted from Journal of Banking
and Finance,
Vol. 16, No. 1, 1992

Reprint Series
no. 86

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A contribution to event study methodology with an application to the Dutch stock market*

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This paper proposes an extended market model for event studies based on daily stock returns. For actual data the assumptions of the simple market model are violated. The return distribution is not normal and neither the variance of the error term nor the risk parameter β are constant. Our model incorporates the generalized autoregressive conditional heteroskedasticity (GARCH) model with t -distributed errors and a time-dependent β . We test for anomalies by adding dummy variables in the regression equation. Our model is fairly general and could be used in a wide variety of event study situations. We illustrate the model by an analysis of the weekend and the option-expiration effect. We use return data from the Dutch stock market. The weekend effect on stock returns is significant, but no expiration effect could be detected.

1. Introduction

Much of the empirical work on event studies rests on simple econometric models with strong statistical assumptions. Since Roll (1977) most research on market efficiency is based on the market model, which relates the return on an individual asset to the return on a market index and an asset-specific

*We have benefitted from discussions with Ray Ball, Phelim Boyle, Theo Nijman, Peter Schotman, Stephen Taylor and Christian Wolff, and workshop participants at the Erasmus University Rotterdam, Tilburg University, EIASM Brussels, the International DTB-AMEX Colloquium in Frankfurt and the University of Manchester. We thank an anonymous referee for his comments.

constant. The estimation and testing is usually carried out under the assumption that the error term and hence the returns follow a normal distribution with constant variance.

In this paper we confirm results of Connolly (1989) which indicate that the assumptions of this simple market model are violated with actual data. For instance, for daily data the returns are not normally distributed with constant variance. Furthermore, we reject the hypothesis of a constant beta for the individual stocks. We, therefore, propose modifications that capture the deviations from normality and allow for a time-dependent beta. The resulting extended market model incorporates the approach of Bos and Newbold (1984) to stochastic betas and the method of Engle and Bollerslev (1986) based on conditional variance and fat-tailed error distributions. These modifications spring from empirical considerations. In the early seventies it was reported that stock return distributions are fat-tailed [e.g. Blattberg and Gonedes (1974)]. More recently, the GARCH model has been frequently used in studies on stock return behaviour [e.g. Chou (1988), Connolly (1989) and Akgiray (1989)].¹

By simply adding dummy variables in the regression equation the extended market model allows us to test for any periodic event, e.g. the weekend, the day-of-the-week and the option-expiration date.² Our model is quite general and can be applied to a variety of empirical investigations of interest to researchers in finance and accounting. Indeed the naive market model has been the workhorse for much of the market based research in this area. It seems evident that future work will require more attention to the concerns we itemize in this paper and use techniques similar to ours.³

In this paper we apply our model for the weekend and the option-expiration effect. We use daily return data from the Dutch stock market. We analyze the 13 major stocks on which options are listed on the European Options Exchange in Amsterdam during the sample period.

For both events we test for effects on the stock return and the stock return volatility, since both effects have been reported in studies on U.S.-markets [e.g. Connolly (1989) and French and Roll (1986) for the weekend effect and Stoll and Whaley (1986) for the expiration effect]. For the weekend effect it turns out that there is a significant negative return effect, but no effect on the return volatility. For the expiration effect we compare the results using our extended GARCH- t model with the results using a normal homoskedastic market model. It is clear that the test results for the abnormal returns and return volatilities change with the model specification. This dependence of

¹Applications of GARCH to exchange rates can be found in Engle and Bollerslev (1986) and Bollerslev (1987).

²Stock options expire every third Friday in January, April, July and October.

³Of course these deviations from the market model assumptions have been pointed out by other researchers [Connolly (1989) among others].

the results on the model specification could have dramatic consequences for some event studies. For the expiration event the consequences were clearly visible but small in magnitude.

The organisation of this paper is as follows. In section 2 we develop the extended market model with a time-dependent beta and a GARCH- t error specification. We also incorporate the dummies for the weekend effect. In section 3 we briefly discuss the data and present the empirical results of our extended model. Section 4 analyzes the expiration effect in detail. Section 5 provides a brief conclusion. In the Appendices we give an outline of the estimation and testing methods.

2. Model

2.1. The basic model

The usual model for analyzing stock returns is the simple market model

$$r_{it} = \alpha_i + \beta_i r_t^M + e_{it}, \quad (1)$$

where r_{it} is the one-period return on asset i at time t and r_t^M is the return on the market index at time t . In this model it is assumed that e_{it} is a temporally uncorrelated normally distributed error term on asset i at time t . However, there is strong evidence that successive returns on individual stocks are correlated [see e.g. Lo and McKinlay (1988)]. The correlation is often negative, especially for daily data on individual stock returns [see e.g. Jennings and Starks (1986)]. The correlation may be caused by infrequent trading or by measurement errors, e.g. due to the bid-ask spread, see French and Roll (1986) and Glosten (1987).⁴ To capture the correlated structure of the returns, we choose an ARMA(1, 1) model as in Taylor (1986). This leads to our basic model

$$r_t = \beta r_t^M + \gamma_0 + \gamma_1 r_{t-1} + e_t + \gamma_2 e_{t-1}, \quad (2)$$

where the subscript i has been dropped.

2.2. A time-dependent systematic risk parameter

In this section we explore different ways of modelling time-dependent betas. Several authors [Bos and Newbold (1984) and Collins, Ledolter and

⁴If the bid-ask price error is proportional to the stock price and this error is random and serially uncorrelated, the observed logged price equals the true (logged) price plus an uncorrelated random error. This implies that the observed returns are equal to $\tilde{r}_t = \ln(p_t/p_{t-1}) + \epsilon_t - \epsilon_{t-1}$, and, assuming that the true returns are uncorrelated, the first-order autocorrelation of the observed returns is equal to $-\sigma_\epsilon^2$.

Rayburn (1987) among others] have presented empirical evidence that the beta of the market model is not constant, but varies through time. A frequently applied method to overcome this problem is to estimate beta for relatively short periods, in which it is assumed to be constant. The price of this method is loss of efficiency due to the smaller number of observations used. A more attractive way to model a time-dependent beta is given by the random parameter models, where beta is considered to be a random variable with a specific distribution.

Bos and Newbold (1984) allow beta to follow a mean reverting AR(1) process

$$\beta_t - \beta = \phi(\beta_{t-1} - \beta) + \xi_t, \quad (3)$$

where β is the mean, ϕ the adjustment parameter and ξ_t white noise with zero mean and variance σ_ξ^2 . Rearrangement of the Bos–Newbold specification leads to

$$\begin{aligned} \beta_t &= \beta + \delta_t \\ \delta_t &= \phi\delta_{t-1} + \xi_t, \end{aligned} \quad (4)$$

where δ_t are serially correlated disturbances. The disturbances in this specification carry over to the next period; the magnitude of this effect depends on the value of ϕ . If $\phi=0$, one obtains the Hildreth–Houck (1968) random coefficients model, where β_t equals a constant mean plus a serially uncorrelated random disturbance. Another special case of the AR(1) model is the random walk, where $\phi=1$ and the mean β is not identified. In the random walk each disturbance has a non-vanishing effect in all subsequent periods. Bos and Newbold estimate the simple market model (1) with a time-dependent beta with monthly data for 464 U.S. stocks and test for the presence of a stochastic beta. For most stocks considered a constant beta was rejected, but in only six cases a significant autoregressive parameter was found.

Another model for a time-dependent beta was proposed by Collins, Ledolter and Rayburn (1987). They add a serially uncorrelated disturbance to the Bos and Newbold model

$$\begin{aligned} \beta_t &= \beta + \delta_t + \varepsilon_t \\ \delta_t &= \phi\delta_{t-1} + \xi_t. \end{aligned} \quad (5)$$

This specification allows a distinction between transitory (ε_t) and correlated (δ_t) random shocks to β_t . Collins et al. (1987) show that the estimate of ϕ is

biased towards zero if $\sigma_\varepsilon > 0$, which may explain the results of Bos and Newbold who found a significant ϕ for only a few series.⁵ The disadvantage of the latter specification is that it is difficult to estimate. If $\phi = 0$, δ_t cannot be distinguished from ε_t and one of the variances (σ_ε^2 and σ_δ^2) is not identifiable.

For simplicity we start with the AR(1) random beta specification. In the empirical results we find rejection of constant betas and significant autoregressive parameters. The differences in results with Bos and Newbold (1984) can be due to the fact that we use daily instead of monthly data. Furthermore, a test against the Collins et al. (1987) specification did not reject the AR(1) model.

2.3. Generalized autoregressive conditional heteroskedasticity (GARCH)

Having discussed the non-stationarity of the betas we now consider the behaviour of the variance through time. There is ample evidence that clustering of large price changes takes place, suggesting that the variance of future returns is partly predictable from the past. Since Engle's seminal work [Engle (1982)] it is common to use an ARCH or similar variance specification for financial time series including exchange rates, foreign currency futures and stock prices. In an ARCH model the variance of the current error, h_t , is predicted from the square of past errors

$$h_t \equiv E_{t-1}(e_t^2) = \alpha_0 + \sum_{i=1}^p \alpha_i e_{t-i}^2, \quad (6)$$

where E_{t-1} denotes expectation conditional upon all information available in the model at time $t-1$. To keep each conditional variance positive, the parameters α_i , $i=0, 1, \dots, p$, are bounded below by zero. This model specification has been extended by Engle and Bollerslev (1986), in order to capture a large number of different variance patterns without having to estimate too many parameters. Engle and Bollerslev use a generalised autoregressive conditional heteroskedasticity (GARCH(p, q)) model which can be defined as follows:

$$h_t = \alpha_0 + \sum_{i=1}^p \alpha_i e_{t-i}^2 + \sum_{i=1}^q \alpha_{i+p} h_{t-i}. \quad (7)$$

⁵The estimate ϕ from the AR(1) formulation is a consistent estimate of $\phi\sigma_\delta^2/(\sigma_\delta^2 + \sigma_\varepsilon^2)$ but not of ϕ .

Besides the conditional variance, h_t , the errors in a GARCH model have an unconditional variance, σ^2 . This variance can be obtained by taking unconditional expectations in (7),

$$\begin{aligned}\sigma^2 &= \alpha_0 + \sum_{i=1}^{p+q} \alpha_i \sigma^2 \\ &= \alpha_0 / \left(1 - \sum_{i=1}^{p+q} \alpha_i \right).\end{aligned}\tag{8}$$

In order to provide an economic interpretation of the GARCH specification we consider a GARCH(1,1) and eliminate α_0 in eq. (7) by using eq. (8). Rewriting (7) results in

$$h_t - \sigma^2 = \alpha_1(e_{t-1}^2 - h_{t-1}) + (\alpha_1 + \alpha_2)(h_{t-1} - \sigma^2).\tag{9}$$

The first term of this equation causes the conditional variance to depend on the previous surprise. This captures clustering of large price changes. The second term is a mean reversion process for the conditional variance with σ^2 as mean and $\alpha_1 + \alpha_2$ as the adjustment parameter.

If $\alpha_1 = \alpha_2 = 0$, the variance is constant, or stated differently, the errors are homoskedastic; the null hypothesis of homoskedasticity may be tested against the GARCH alternative ($\alpha_1, \alpha_2 \neq 0$). The specification tests indicate a strong rejection of homoskedasticity of the error terms. Furthermore, diagnostic tests show that there is no need to incorporate additional lags to the GARCH (1,1) specification.

If $\alpha_1 + \alpha_2$ equals 1, the unconditional variance σ^2 does not exist, and only the conditional variance can be estimated. In this case, the GARCH(1,1) model given in (9) becomes

$$h_t = h_{t-1} + \alpha_1(e_{t-1}^2 - h_{t-1}).\tag{10}$$

The value of h_1 is unknown and must be estimated from the data, Engle and Bollerslev call this model Integrated GARCH (IGARCH), for there is a 'unit root' in the conditional variance. For more details on IGARCH and its statistical properties we refer to Engle and Bollerslev (1986).

Although GARCH gives a specification for the conditional variance of the model's errors, the error distribution is not determined by this specification. In a GARCH model with a normal conditional error distribution, the unconditional error distribution has fatter tails than the normal. Often this specification is sufficient to account for the observed unconditional kurtosis in the data. However, it is possible that the conditional error distribution has fatter tails than the normal. Weiss (1986) shows that assuming normality

while the true distribution has fat tails renders consistent but inefficient estimates.

An adjustment has been used by Bollerslev (1987) and others, who allow the errors to be conditionally t -distributed. The t -distribution is more likely to generate larger errors than the normal distribution. Using the t -distribution implies that the outliers are given smaller weights in the estimates and test statistics [see Box and Draper (1987, pp. 83–90)]. In Appendix A the log likelihood function of a model with normal or t -distribution is derived and the estimation and testing procedures are presented. From our specification tests we find that normality of the conditional error distribution is strongly rejected. We, therefore, propose a GARCH- t model specification for our extended market model.

2.4. The extended market model and the weekend effect

The extended market model is the basic model with an AR(1) time-dependent beta and a GARCH(1, 1) variance structure:

$$r_t = \beta_t r_t^M + \gamma_0 + \gamma_1 r_{t-1} + e_t + \gamma_2 e_{t-1} \quad e_t \sim (0, h_t) \quad (11a)$$

$$\beta_t - \beta = \phi(\beta_{t-1} - \beta) + \xi_t \quad \xi_t \sim (0, \sigma_\xi^2) \quad (11b)$$

$$h_t - \sigma^2 = \alpha_1(e_{t-1}^2 - h_{t-1}) + (\alpha_1 + \alpha_2)(h_{t-1} - \sigma^2). \quad (11c)$$

The parameters to be estimated are β , γ_0 , γ_1 , γ_2 , σ^2 , α_1 , α_2 , ϕ and σ_ξ^2 . We do not assume a particular distribution for e_t and ξ_t but rather assume that the prediction errors in the Kalman filter (Appendix B) are Student t distributed (with ν degrees of freedom). The degrees of freedom parameter ν is also to be estimated.

The extended market model can easily be applied for various periodic events by adding dummies to the return eq. (11a) and the variance eq. (11c). This also makes it possible to differentiate between different events. For example, in testing for an expiration effect a separate dummy for the weekend effect should be included, because the expiration effect will otherwise be influenced by the weekend effect. In section 4 we test for the expiration effect in addition to the weekend effect.

When presenting the results for our extended market model in the next section we also test for the weekend effect. From various studies on U.S. data a significant negative weekend effect on the stock return is reported. From the paper by French and Roll (1986) it also may be expected that the variance in the weekend differs from that on trading days. We, therefore, test for the weekend effect on both the stock return and the stock return variance. We introduce a weekend dummy, M_t , which takes the value 1 if

day t is a Monday (or a Tuesday after a holiday on Monday) and 0 elsewhere. We add this dummy variable to eq. (11a) as follows:

$$r_t = \beta_1 r_t^M + \gamma_0 + \gamma_1 r_{t-1} + e_t + \gamma_2 e_{t-1} + \gamma_3 M_t \quad e_t \sim (0, h_t) \quad (12a)$$

The adjustment for the weekend effect on the variance is by means of an intervention dummy described by Box and Tiao (1975). This results in adding a dummy variable to the unconditional variance, changing (11c) into

$$h_t - (\sigma^2 + \alpha_3 M_t) = \alpha_1 (e_{t-1}^2 - h_{t-1}) + (\alpha_1 + \alpha_2)(h_{t-1} - (\sigma^2 + \alpha_3 M_{t-1})). \quad (12b)$$

The additional parameters to be estimated are γ_3 and α_3 .

3. Empirical results

3.1. Data

In this section we describe the data used to estimate and test our extended market model including the weekend effect. In the next section we use the same data to test for the expiration effect. The sample consists of daily closing prices of 13 major Dutch stocks listed on the Amsterdam Stock Exchange. These stocks were selected because options are traded on these stocks during the whole sample period: January 3, 1984 to August 31, 1987. This makes a total of 921 observations for each stock. The period contains 15 quarterly option expiration dates on which options on all listed stocks expire. (It does not contain the world-wide crash in October 1987). The data were taken from official stock exchange lists.

The market return is the return on the ANP-CBS general stock index. The value of this index is a weighted average of several group indices. The weights are based on real sales of the individual firms in each group in a base year. Roughly speaking, this results in a 50% weight for the group internationals, 40% for industry and 10% for the other groups. Each group index is an equally weighted index. The ANP-CBS index does not correct for dividend payments on the stock. Although this index has theoretical shortcomings, it is the only reasonable stock index available for the Dutch stock market during the whole sample period.

Similarly, the return on an individual stock, given by the first difference of the log price, is not corrected for dividend payments.⁶ Table 1 shows the names of the 13 series, the abbreviations we use and some descriptive

⁶Dividend payments are usually concentrated in April and October.

statistics.⁷ The mean and standard deviation of the returns are shown in percentages per day. SK is the skewness of the standardised residuals and KU the kurtosis.

3.2. Specification and diagnostic tests

In this section we justify our extended market model with a GARCH(1,1)- t distribution and a time-dependent beta. We subsequently present some specification tests in table 2 and diagnostic tests in table 3. In Appendix A technical details on the tests are given.

The specification tests are used to test the normality of the error distribution, the constancy of the variance and the risk parameter beta. In the first two columns of table 2 results are presented for the test on normality of the errors of the return distribution. The test is based on the third and fourth moment from the residuals of the models with an assumed normal distribution.⁸ The test is computed from the residuals of the constant variance model and from the standardised residuals of the GARCH model. For all stocks, the tests firmly reject normality.

There are several ways to test the constancy of the variance. We use the likelihood ratio test that compares the value of the log likelihood under the assumption of a constant variance with the log likelihood under a GARCH(1,1) variance specification. The test is computed under the normal and the t -distribution; the results are shown in the third and fourth column of table 2. The likelihood ratio test rejects constancy of the variance in all cases except one. Since normality is also rejected, we prefer the GARCH(1,1)- t specification.

Finally, the Lagrange Multiplier (LM) test for a constant beta is presented in the last column. The null hypothesis is $\sigma_\epsilon^2 = 0$. Note that ϕ is not identified under the null. The LM test's alternative is the Hildreth-Houck random coefficient model, which corresponds to $\phi = 0$ in the AR(1) random beta model. More powerful tests as given by Watson and Engle (1985) have not been fully developed for models with heteroskedastic disturbances. The results show that a constant beta is rejected in most cases.

Before presenting parameter estimates for our extended market model we discuss a number of diagnostic tests to which the model has been subjected.

⁷There is an outlier in the HOO series in August 15, 1986, caused by the announcement of unexpectedly bad quarterly earnings. In table 1 this observation was deleted. In the subsequent analysis a correction for this outlier is made by intervention modelling, described by Box and Tiao (1975), in which the outlying observation is not deleted, but rather adjusted by means of an intervention dummy.

⁸Define $\hat{\mu}_3$ and $\hat{\mu}_4$ as the estimated third and fourth moment of the standardised residuals. Then the test statistic is $\text{NORM} = T(\hat{\mu}_3^2/6 + (\hat{\mu}_4 - 3)^2/24) \chi^2(2)$. This test measures the deviations from the third and fourth moment of the standardised normal distribution, which are zero and 3, respectively.

These are shown in table 3. The first diagnostic test checks the need for more lags in the variance equation; we report the LM test against GARCH(2,1) in the first column. It is clear that there is no need for more lags in the variance equation, a result in line with other literature, e.g. Bollerslev (1987), Chou (1988) and McCurdy and Morgan (1988). The second column gives the LM test for heteroskedasticity, obtained by adding the squared market return to the variance equation; this may be seen as a test against the Collins, Ledolter and Rayburn random beta specification (5). The AR(1) random beta specification is not rejected against this alternative, except for UNI.

The appropriateness of the lag structure is tested by adding the two-period lagged own return, r_{t-2} , or the two-period lagged error term, e_{t-2} , to the model and computing the LM test for exclusion of this variable. The outcomes are reported in the third and fourth column. The exclusion of the two-period lagged variables is sometimes rejected, but we do not conclude that the lag structure is incorrect.

Finally, the last column shows an LM test against the GARCH-in-mean model [Engle, Lilien and Robins (1987)]. In this model the conditional variance at day t , h_t , is added as an explanatory variable for the stock return.⁹ Because h_t measures the diversifiable (unsystematic) risk, it should have no explanatory power. Only for NEDL this test statistic is significant, so that we conclude that the conditional error variance is not a significant explanatory variable for stock returns.

Summarizing the above specification and diagnostic tests we conclude that our extended market model is a reasonable specification for the stock return and stock return volatility.

3.3. *Parameter estimates for the extended market model*

In table 4 we report the estimation results of our extended market model under the null hypothesis that no expiration effect exists. The parameters are estimated by the method of Maximum Likelihood (see Appendix A) under the assumption of a conditional t -distribution for the prediction errors. We subsequently discuss the results concerning the dynamic structure of the model, the time-dependent β_t , the GARCH- t model and the weekend effect. In the next section we present the results for the expiration effect.

The dynamic structure of the model is given by the AR-parameter γ_1 and the MA-parameter γ_2 . Both are significantly different from zero (γ_1 negative and γ_2 positive) in most cases, but their magnitude is roughly the same ($\gamma_1 + \gamma_2 = 0$), indicating that there may be a common root. However, omitting γ_1 or γ_2 (or both) introduces first-order serial correlation in the error term.

⁹Economic theory suggests that risk averse investors require a risk premium for volatile assets. The conditional variance is a measure for the risk of stock investment. Geweke (1989) shows how a risk averse investor could use the conditional variance to maximise his utility.

Table 1
Data description.

Name	Abbr.	Mean	St. dev.	SK	KU
Algemene Bank Nederland	ABN	0.031	1.084	-0.257	5.34
Ahold	AH	0.062	1.434	-0.453	7.83
Akzo	AKZO	0.045	1.400	-0.496	6.52
AMRO Bank	AMRO	0.018	1.395	0.072	6.87
Gist Brocades	GIS	0.038	1.456	-0.870	10.75
Heineken	HEI	0.052	1.360	0.204	7.32
Hoogovens	HOO	-0.043	2.508	-2.801	33.36
Hoogovens*	HOO	-0.009	2.287	-0.995	11.21
Koninkl. Luchtvaart Mij	KLM	0.006	1.808	-0.134	5.65
Nationale Nederlanden	NATN	0.029	1.338	-0.884	26.92
NedLloyd	NEDL	0.048	1.466	-0.297	13.31
Philips	PHI	0.025	1.355	-0.988	13.76
Koninklijke Olie	RD	0.068	1.110	-0.074	4.71
Unilever	UNI	0.107	0.968	-0.067	6.30
Index	ANP-CBS	0.061	0.881	0.003	6.95

*With correction for an outlier (observation 659, August 15, 1986, deleted). In all other tables we only report the corrected series.

Table 2
Specification tests.*

Series	NORM1	NORM2	LR1	LR2	Constant β LM
ABN	742**	745**	56.96**	50.72**	12.33**
AH	3,728**	2,798**	94.42**	46.12**	0.54
AKZO	1,354**	1,425**	22.14**	17.96**	28.50**
AMRO	408**	394**	19.64**	17.02**	13.12**
GIS	5,552**	3,435**	101.44**	45.12**	4.56*
HEI	1,238**	599**	111.06**	59.88**	2.71
HOO	1,608**	2,153**	154.28**	82.94**	11.28**
KLM	500**	621**	9.34**	16.80**	143.09**
NATN	927**	1,674**	15.22**	17.42**	186.87**
NEDL	7,783**	11,110**	145.18**	55.40**	5.13*
PHI	10,073**	13,161**	2.84	12.36**	8.89**
RD	544**	572**	27.76**	21.20**	12.31**
UNI	1,185**	1,687**	29.90**	24.72**	3.36

*NORM1: test for normality based on assumptions of normality and constant variance, $\chi^2(2)$;

NORM2: test for normality based on assumptions of normality and GARCH, $\chi^2(2)$;

LR1: likelihood ratio test for GARCH, based on normality, $\chi^2(2)$;

LR2: likelihood ratio test for GARCH, based on Student t distribution, $\chi^2(2)$.

LM: Lagrange-Multiplier test for constant beta, $\chi^2(1)$.

Table 3
Diagnostic checking of the GARCH(1,1)- t specification.^a

	GARCH(2,1) LM	Hetero LM	r_{1-2} LM	e_{1-2} LM	GARCH-M LM
ABN	0.14	0.42	1.60	0.01	2.57
AH	1.43	0.36	0.01	0.37	0.01
AKZO	0.06	1.48	1.98	2.78	0.44
AMRO	0.59	3.02	10.53**	9.17**	0.85
GIS	0.44	0.68	2.03	1.71	0.51
HEI	0.07	0.86	4.45*	5.16*	0.94
HOO	0.15	0.10	0.61	1.81	2.74
KLM	0.63	0.12	1.22	0.15	0.04
NATN	1.69	1.40	5.70*	4.75*	2.44
NEDL	0.82	2.39	2.74	0.95	5.23*
PHI	1.99	0.02	7.95**	2.70	0.01
RD	0.01	3.38	3.56	6.30*	0.09
UNI	0.42	8.62**	3.96*	7.61**	2.69

^aAll tests are asymptotically $\chi^2(1)$ distributed.

Since the sum is close to the zero, this can be interpreted as a Koyck effect with respect to all other variables [see, for instance, Judge et al. (1985), p. 379], in particular, the market return. This implies that the stock returns are influenced by the lagged market return. In our empirical results the estimated Koyck parameter γ_1 is systematically negative and of the order -0.2 , so that second and higher order powers are negligible. A tentative explanation could be that stocks overreact to news as reflected by the market return, which is systematically corrected the next day.

The estimates of the time-dependent β_t model are the mean β , the autoregressive parameter ϕ and the standard deviation σ_β , computed as $\sigma_\beta = \sqrt{(\sigma_\epsilon^2/(1-\phi^2))}$. The results show that the patterns differ among the various stocks. Some β_t have a short memory (PHI, UNI and NEDL), while others (AH, AKZO, HEI and RD) show a long delay in returning to the mean, if they return to the mean at all.¹⁰ To illustrate the behaviour of β_t for the latter group the β_t of AKZO is plotted in fig. 1. The standard deviation of β_t is low for AH and high for KLM, compared with the mean β . The result for the behaviour of β_t is illustrated in fig. 2 for KLM.

The t -ratio of $\alpha_1 + \alpha_2$ proves that the GARCH parameters are jointly significant for all stocks. The slow decay of the conditional variance after sharp rises is clear from the plot of AH's conditional variance (fig. 3). The conditional variance structure has a long memory for many stocks (i.e. $\alpha_1 + \alpha_2$ is close to one), but it tends to be mean reverting: the conditional variances are temporary deviations from the unconditional variance σ^2 . We

¹⁰To the best of our knowledge formal tests for ϕ equals one have not yet been developed. There is an analogy with the Dickey-Fuller test for 'unit-root', but it is not clear whether we can use their critical values [see Dickey and Fuller (1979)].

Table 4
Parameter estimates for the extended market model.

$$r_t = \beta r_t^M + \gamma_0 + \gamma_1 r_{t-1} + e_t + \gamma_2 e_{t-1} + \gamma_3 M_t \quad e_t \sim (0, h_t)$$

$$\beta_t - \beta = \phi(\beta_{t-1} - \beta) + \xi_t \quad \xi_t \sim (0, \sigma_\xi^2)$$

$$h_t - (\sigma^2 + \alpha_3 M_t) = \alpha_1 (e_{t-1}^2 - h_{t-1}) + (\alpha_1 + \alpha_2)(h_{t-1} - (\sigma^2 + \alpha_3 M_t)).$$

	β	γ_0	γ_1	γ_2	γ_3	σ^2	α_1	$\alpha_1 + \alpha_2$	α_3	ϕ	σ_β	ν	κ	LL
ABN	0.722 (0.036)*	0.042 (0.027)	-0.169 (0.031)	0.125 (0.054)	-0.089 (0.052)	0.936 (0.356)	0.391 (0.115)	0.765 (0.106)	0.011 (0.073)	0.647 (0.159)	0.315	4.20	4.03	3,115
AH	0.877 (0.057)	0.076 (0.034)	-0.111 (0.037)	0.092 (0.049)	-0.091 (0.072)	1.272 (0.541)	0.028 (0.010)	0.992 (0.005)	0.204 (0.186)	0.966 (0.039)	0.187	3.87	8.95	2,868
AKZO	0.849 (0.176)	0.117 (0.035)	-0.180 (0.028)	0.126 (0.046)	-0.199 (0.064)	0.910 (0.116)	0.072 (0.031)	0.889 (0.060)	-0.091 (0.110)	0.990 (0.008)	0.282	5.41	6.73	2,962
AMRO	1.006 (0.050)	0.067 (0.038)	-0.217 (0.033)	0.168 (0.052)	-0.178 (0.073)	0.967 (0.255)	0.015 (0.010)	0.992 (0.005)	0.056 (0.129)	0.416 (0.165)	0.526	6.97	3.49	2,899
GIS	0.884 (0.043)	0.103 (0.038)	-0.039 (0.040)	0.078 (0.056)	-0.184 (0.070)	1.387 (0.366)	0.054 (0.018)	0.977 (0.012)	-0.152 (0.150)	0.484 (0.371)	0.285	4.87	10.31	2,827
HEI	0.906 (0.090)	0.078 (0.036)	-0.216 (0.033)	0.180 (0.049)	-0.213 (0.068)	1.119 (0.234)	0.255 (0.057)	0.819 (0.060)	0.160 (0.107)	0.988 (0.011)	0.228	6.90	3.90	2,907
HOO	1.401 (0.069)	0.110 (0.057)	-0.290 (0.031)	0.220 (0.051)	-0.289 (0.093)	5.983 (4.497)	0.289 (0.085)	0.918 (0.065)	-0.371 (0.315)	0.647 (0.170)	0.658	3.52	6.17	2,474
KLM	1.008 (0.066)	0.095 (0.059)	-0.204 (0.043)	0.270 (0.060)	-0.444 (0.100)	2.928 (1.138)	0.081 (0.032)	0.966 (0.029)	-0.100 (0.266)	0.348 (0.197)	0.706	5.14	4.29	2,536
NATN	0.741 (0.042)	0.066 (0.032)	-0.081 (0.037)	0.058 (0.059)	-0.182 (0.069)	0.872 (0.107)	0.193 (0.075)	0.450 (0.194)	0.133 (0.131)	0.436 (0.259)	0.387	4.81	8.37	2,949
NEDL	0.672 (0.038)	0.021 (0.037)	-0.171 (0.044)	0.210 (0.061)	-0.241 (0.069)	2.396 (1.106)	0.344 (0.089)	0.876 (0.069)	0.068 (0.157)	-0.157 (0.280)	0.404	3.52	19.08	2,821
PHI	0.864 (0.039)	0.102 (0.034)	-0.186 (0.031)	0.199 (0.052)	-0.233 (0.062)	1.025 (0.160)	0.233 (0.097)	0.292 (0.113)	-0.012 (0.146)	0.202 (0.250)	0.444	3.69	25.97	2,941
RD	0.691 (0.074)	0.107 (0.035)	-0.224 (0.042)	0.207 (0.056)	-0.148 (0.060)	0.771 (0.121)	0.045 (0.019)	0.961 (0.023)	-0.051 (0.089)	0.969 (0.020)	0.272	7.24	3.86	3,021
UNI	0.777 (0.027)	0.114 (0.024)	-0.160 (0.026)	0.102 (0.049)	-0.039 (0.042)	0.454 (0.068)	0.159 (0.062)	0.742 (0.108)	-0.087 (0.048)	0.235 (0.259)	0.265	4.96	8.69	3,293

*Asymptotic standard errors are in parentheses.

Fig. 1. Random beta AKZO.

Fig. 2. Random beta KLM.

Fig. 3. Conditional variance AH.

computed several test statistics for IGARCH, which tend to reject the 'unit-root' hypothesis.¹¹

The estimates of ν are fairly low, about 4, with a low 3.52 for HOO and NEDL and a high 7.24 for RD. The estimates of ν under the assumption of a constant variance¹² are even lower than the estimates under GARCH, from which we conclude that the heteroskedasticity accounts for some of the kurtosis, but not for all. Even the conditional error distributions are heavily fat tailed. This can also be seen from the next column, which represents the estimated kurtosis of the standardised residuals. If the conditional error distribution is normal, κ should be equal to 3; a larger κ indicates leptokurtosis.

Finally, the weekend effect on the stock return and the stock return variance are given by γ_3 and α_3 . For most cases the weekend effect on the return is small but significantly negative (with a maximum of -0.4% per

¹¹Engle and Bollerslev (1986) comment on the problems regarding 'unit-root' test. They report a Monte Carlo study, where the test statistics are relatively well behaved. Our evidence points in a different direction. The standard Wald test rejects IGARCH in 5 out of 13 cases, the LM test in 12 cases and the LR test in all cases. This suggests that the distribution of the log likelihood in the region close to $\alpha_1 + \alpha_2$ is far from normal.

¹²These estimates are not presented here, but are available from the authors.

day for KLM).¹³ The weekend effect on the variance, however, is negligible. For none of the stocks considered the effect is significant, nor is the sign of α_3 similar for the various stocks.

4. Option expiration effect

The influence of exchange-traded options on stock returns and stock return volatility has received much attention, especially since the October 1987 stock market crash. Recent studies by Conrad (1989), Skinner (1989) and Harris (1989) concentrate on the effect of option introduction on the stock volatility. In this study, however, we investigate the impact of the option expirations on the stock return and stock return volatility. Although the discussion in the U.S. seems to be closed with the study of Stoll and Whaley (1986) it is still very much alive in smaller and less liquid European markets, e.g. Pope and Yadav (1988) and Van den Bergh and Kemna (1988).

Using the same data set, Van den Bergh and Kemna (1988) find a positive excess return before option expiration for several stocks and a decrease in variance on the expiration date.¹⁴ Options on the stocks in our sample are traded on the European Options Exchange (EOE) in Amsterdam. The EOE started in 1978 with options on 9 major Dutch stocks. The trading and clearing system is similar to the system of the CBOE. At the start of our dataset, January 1984, options on 13 stocks were traded. In 1989 the EOE confirmed its position as a leading options market in Europe with a total trading volume of 13.4 million contracts. The stock options (on 25 Dutch stocks) contributed 75% of this volume with Philips as largest with 1.76 million contracts. On a monthly basis we find trading volumes of 300,000 contracts for Philips and 10,000 for Heineken in a lively expiration month like October 1989. This trading volume drops with 50% in a month without expiration.

The period of observation contains fifteen quarterly option expirations for each stock. To test significance of the excess return on certain days before or after the expiration dates, we define dummy variables, which take the value 1 for those days are zero elsewhere, and re-estimate the model with these dummies as additional explanatory variables. In this study we use six dummy variables, one for each day from two days before to three days after each expiration date.

We define these dummies D_{jt} , $j = -2, \dots, 3$ with D_{-2t} equals one on the

¹³In the computer program, all return data were multiplied by 100. Consequently, the returns are measured as percent per day, which affects the estimates of the parameters γ_3 and σ^2 . Of course, the significance (*t* ratio) of these parameters is scale-independent and not affected by this transformation.

¹⁴They used an approach similar to Klemkosky (1978) and Officer and Trennepohl (1981).

Table 5a
Excess returns around option expirations; normal, homoskedastic errors.

Series	LM	LR	Wald	Effect on return day from expiration						Effect on variance day from expiration		
				-2	-1	0	+1	+2	+3	-1	0	+1
ABN	5.01	5.69	6.52									
AH	6.53	2.77	0.66							-	-	-
AKZO	7.98	5.84	3.98									
AMRO	7.86	5.18	2.91									
GIS	5.96	4.65	3.15								-	-
HEI	6.69	5.59	2.96									
HOO	6.64	5.03	4.25								-	-
KLM	2.62	2.88	3.44									
NATN	4.03	5.53	7.21									
NEDL	25.34**	28.72**	33.11**		-	-	-			-	-	-
PHI	8.51	7.89	8.19									
RD	10.02	11.00	13.46*		-	-					-	-
UNI	8.82	7.38	6.71									

Wednesday before and $D_{3,t}$ equals one on the Wednesday after expiration. With these eqs. (12a) is changed to

$$r_t = \beta_t r_t^M + \gamma_0 + \gamma_1 r_{t-1} + e_t + \gamma_2 e_{t-1} + \gamma_3 M_t + \sum_{j=-2}^3 \psi_j D_{jt} \quad e_t \sim (0, h_t), \quad (13a)$$

where h_t is given by eq. (12b).

The significance tests of these dummy variables are reported in tables 5a and 5b: the former gives the results for assumed normal, homoskedastic errors, and the latter gives the estimates for the more general GARCH(1,1)- t error specification. In both cases we consider the time-dependent β_t and the weekend effect. The first three columns give LM, LR and Wald tests of joint significance of the six dummy variables. Asymptotically, these three tests are equivalent and have a $\chi^2(6)$ distribution. The next six columns show the sign of the estimated parameter of each individual dummy variable (i.e. the estimated excess return); the number of + and - signs indicates the level at which the dummies are significant: 5% or 1%.

In the GARCH- t model, there is no clear expiration effect on stock returns. There are only a few significant dummy variables, and none of the tests indicates that the six dummies are jointly significant. Inspection of the residuals shows why: the residuals around different expiration dates do not have the same sign. The number of significant dummy variables is not larger than might be expected at the 5% significance levels. These results differ from

Table 5b
Excess returns around option expirations; GARCH-*t*.

Series	LM	LR	Wald	Effect on return day from expiration						Effect on variance day from expiration		
				-2	-1	0	+1	+2	+3	-1	0	+1
ABN	7.81	8.42	7.45					+				
AH	5.62	3.51	2.46									--
AKZO	6.34	6.08	4.98									
AMRO	9.26	6.83	4.19									
GIS	1.72	1.16	0.71									
HEI	7.00	7.63	6.46					+				--
HOO	8.81	8.18	4.38									
KLM	4.60	5.51	6.88									
NATN	4.08	5.52	6.65									
NEDL	12.27	10.92	7.15				-					
PHI	7.17	8.11	8.30		+							
RD	7.58	5.79	4.04									
UNI	11.73	10.70	9.64		-							

the results when the model is estimated under the usual assumptions of normally distributed errors with constant variance. In that case, NEDL and RD exhibit significant excess returns.

In order to test for variance effects around option expirations we use an approach similar to the weekend effect. We add the dummy variables to eq. (12b) resulting in¹⁵

$$\begin{aligned}
 h_t - (\sigma^2 + \alpha_3 M_t + \sum_{j=-1}^1 \psi_j D_{jt}) &= \alpha_1 (e_{t-1}^2 - h_{t-1}) \\
 &+ (\alpha_1 + \alpha_2) (h_{t-1} - (\sigma^2 + \alpha_3 M_{t-1} + \sum_{j=-1}^1 \psi_j D_{j,t-1})).
 \end{aligned} \quad (13b)$$

Note that we only consider one day before, on and one day after expiration. For simplicity we do not use eq. (13a) for the return specification but (12a) without the expiration effect. Furthermore, we do not test the dummy variables simultaneously but one by one. In that case testing on significant parameters can be performed by an LM test given in Appendix A, eq. (A.10).

In the last three columns of tables 5a and 5b the LM tests for variance effects are presented. The significance is denoted in the same way as for the

¹⁵Estimating equation (13c) is difficult, because the ψ_j tend to be negative, which may cause negative h_t 's. Introducing restrictions makes it even more complicated. A possible solution is a generalisation of a logarithmic model as given by Taylor (1989).

dummies in the return equation. If there is any effect on the conditional variance visible in table 5b, it is likely to be negative. This effect could be caused by controlled convergence of the stock price towards the nearest-by exercise price, thereby temporally reducing the volatility of the stock price. This pressure may be in positive or in negative direction, leading to an average effect close to zero. This could explain the lack of significance of excess returns.

Future research may concentrate on studying price movements towards exercise prices based on the open interest on the Thursday before expiration and transaction prices of the stocks on the Friday of expiration.

5. Conclusion

In this paper we developed an extended market model for event studies based on daily stock returns. With respect to the statistical specification of the market model we draw the following conclusions:

- (a) The systematic risk parameter (beta) is not constant over time; an AR(1) random coefficient model seems an appropriate alternative.
- (b) The market model's errors are conditionally heteroskedastic, which is a property frequently found in financial data. The GARCH(1,1) specification provides an adequate, yet parsimonious representation of the conditional variance.
- (c) Both the unconditional and the conditional error distribution are fat-tailed; normality is rejected for all stock price series considered. A Student t distribution with (relatively) low degrees of freedom fits the data and much better.

We illustrated our extended model by an analysis of the weekend and the option-expiration effect. We confirmed the results from Connolly (1989) that irrespective of the estimation method, the weekend effect remains significantly negative. With respect to the expiration effect only a few systematic excess returns were found, but there is evidence of a slight reduction in variance around expirations, possibly caused by controlled convergence of the stock price towards the nearest-by exercise price. A comparison of these results with the results obtained under the usual assumptions on the error process (homoskedastic, normal distribution) shows that ignoring the fat tails and the heteroskedasticity may lead to spurious results.

Appendix A

Maximum likelihood estimation and testing

In this appendix we discuss the estimation and testing procedures.

Parameter estimates of the extended market model in section 2 can be obtained by maximisation of the log likelihood function of the parameter vector θ , $\ln L(\theta)$. The likelihood function given a sequence of T observation on the dependent variable, is the joint density of the observations given the value of θ : $L(\theta) = f(y_1, \dots, y_T | \theta)$. By the prediction error decomposition [Harvey (1981)],

$$\ln L(\theta) = \sum_{i=2}^T \ln f(y_i | Y_{i-1}, \theta) + \ln f(y_1 | \theta), \quad (\text{A.1})$$

which is obtained by repeatedly writing the joint distribution as the product of the conditional distribution of y_i and the marginal distribution of $Y_{i-1} \equiv (y_{i-1}, \dots, y_1)$. The initial condition for this likelihood is the distribution of y_1 .

In the GARCH- t model without random coefficients, the prediction error of y_i is the error of the market model, e_i . The prediction errors in a model with random coefficients are computed from the Kalman filter equations given in Pagan (1980) adapted for this model in Appendix B. Let v_i denote the prediction error and f_i its variance, then the conditional probability density function under a normal distribution is

$$n(0, f_i) = \frac{1}{\sqrt{2\pi f_i}} e^{-\frac{1}{2} v_i^2 / f_i}. \quad (\text{A.2})$$

The pdf under a t distribution with degrees of freedom ν is

$$t_{f_i, \nu} = \frac{\Gamma\left(\frac{\nu+1}{2}\right)}{\Gamma\left(\frac{\nu}{2}\right) \sqrt{\pi}} \frac{1}{\sqrt{(\nu-2)f_i}} \left(1 + \frac{v_i^2}{(\nu-2)f_i}\right)^{-(\nu+1)/2}. \quad (\text{A.3})$$

Hence, the log likelihood under a normal distribution can be written as

$$\ln L(\theta) = \sum_{i=1}^T \left(-\frac{1}{2} \ln f_i - \frac{1}{2} v_i^2 / f_i\right) - \frac{T}{2} \ln 2\pi. \quad (\text{A.4})$$

In the case of a t distribution the log likelihood becomes

$$\ln L(\theta) = T \left[\ln \Gamma\left(\frac{\nu+1}{2}\right) - \ln \Gamma\left(\frac{\nu}{2}\right) - \frac{1}{2} \ln \pi(\nu-2) \right]$$

$$+ \sum_{i=1}^T \left[-\frac{1}{2} \ln f_i - \left(\frac{\nu+1}{2} \right) \ln \left(1 + \frac{v_i^2}{(\nu-2)f_i} \right) \right]. \quad (\text{A.5})$$

The normal distribution is a limiting case of the t distribution for $\nu \rightarrow \infty$, or, equivalently, $1/\nu \rightarrow 0$. The difference between the normal and the t distribution is clear from these formulas. The log likelihood function of the normally distributed errors declines linearly in the squared standardised errors, while the log likelihood of the t distributed errors has a slower, logarithmic decline. Hence, under a normal distribution, large errors do have a greater impact on the log likelihood and therefore on the maximum likelihood estimates than under a t distribution.

The initial conditions are given in Appendix B. To eliminate any disturbing effect of these conditions, we use the first 21 observations only to start the recursion. These first 21 observations do not contribute to the likelihood function, which leaves 900 observations to calculate its value.

Following Engle (1982) we use the BHHH algorithm [Berndt et al. (1974)], which is a modified scoring algorithm, for maximisation of the log likelihood function. For each iteration, the parameter estimates are given by

$$\theta^{i+1} = \theta^i + \lambda_i (S^T S)^{-1} S^T \mathbf{1}, \quad (\text{A.6})$$

where S is the scoring, matrix evaluated at θ^i , with elements

$$s_{ij} = \frac{\partial \ln f(y_i | Y_{t-1}, \theta)}{\partial \theta_j}, \quad (\text{A.7})$$

where θ_j denotes the j th element of the parameter vector. Due to the complexity of the derivatives, especially in the GARCH-in-mean model, numerical derivatives are used. The gradient of the log likelihood function is $S^T \mathbf{1}$ ($\mathbf{1}$ being a vector of unit elements) and the information matrix, $I(\theta)$, is estimated by $1/T$ times the product of the first derivatives, $S^T S$. The directional vector, β , is easily obtained by solving the equations $S^T S \beta = S^T \mathbf{1}$, which avoids inversion of $S^T S$, and the step length λ_i is found by maximising the likelihood in this direction. The iteration process is stopped if convergence is satisfactory, for example if $\mathbf{1}^T S (S^T S)^{-1} S^T \mathbf{1}$ is smaller than a certain value, for which we chose 0.05.

Asymptotic standard errors can be obtained as the square roots of the diagonal elements of $1/T$ times the inverse information matrix. The elements of the inverse information matrix (times $1/T$) can be consistently estimated by the inverse of the outer product of the gradients, $(S^T S)^{-1}$, evaluated at the maximum likelihood values θ_{ML} . Assuming asymptotic normality, one can use the standard errors for t tests of parameter significance. A generalisation

for non-linear models of the common t and F tests is the Wald test. For a k -dimensional restriction vector $h(\theta) = 0$ the value of the Wald test is

$$W = Th(\theta)^T [h'(\theta)^T I^{-1}(\theta) h'(\theta)]^{-1} h(\theta) \sim \chi^2(k), \quad (\text{A.8})$$

evaluated at θ_{ML} ¹⁶ (primes denote first derivatives). If it is possible to estimate a model both under the null and under a more general alternative hypothesis, with ML parameter estimates θ_0 and θ_a , the Likelihood Ratio (LR) test can be computed:

$$LR = 2 [\ln L(\theta_a) - \ln L(\theta_0)]. \quad (\text{A.9})$$

The LR statistic is approximately χ^2 distributed with degrees of freedom equal to the number of restrictions (k). Sometimes, when it is difficult or costly to compute ML estimates under the alternative hypothesis, the Lagrange Multiplier test is attractive. The LM test can be obtained from the first step of the BHHH algorithm for the model under the alternative hypothesis, with starting values for the parameters given by the estimates under the null, see Chou (1988). The LM test statistic is asymptotically $\chi^2(k)$ distributed and its value is given by

$$LM = t^T S(S^T S)^{-1} S^T t. \quad (\text{A.10})$$

The LM test is frequently used to test exclusion restrictions: a variable that is excluded from the original specification is added to the model, and one step in the BHHH algorithm is performed for the extended model. This procedure is referred to as an LM variable addition test.

In a recent paper, Calzolari and Panattoni (1988) showed that in small samples the choice of the estimator of the variance-covariance matrix of the ML parameter estimates is important. Different estimators are likely to give very different results, although they can be asymptotically equivalent. For example, the outer-product-of-gradient matrix (which we use) tends to give larger standard errors than the Hessian (matrix of second derivatives of the likelihood function) in their examples. We did some experimentation with estimating the Hessian, but the resulting standard errors did not differ too much from the original ones; therefore, we do not present them.

¹⁶For example, the Wald test of joint significance of a subset θ_1 of the parameter vector θ is $W = \theta_1 [E(S^T S)^{-1} E]^{-1} \theta_1$, where E is a $\dim(\theta)$ square matrix with diagonal elements 1 for the parameters included in θ_1 and zeros elsewhere.

Appendix B

The Kalman filter

Pagan (1980) discusses the identification and estimation of models with time-dependent coefficients. It is convenient to rewrite the extended market model with a GARCH variance specification and stochastic beta, given in formula (12) and repeated here as (B.1) in the state space form.

$$\begin{aligned} r_t &= \beta_t r_t^M + \gamma_0 + \gamma_1 r_{t-1} + e_t + \gamma_2 e_{t-1} + \gamma_3 M_t \\ \beta_t - \beta &= \phi(\beta_{t-1} - \beta) + \xi_t \\ h_t - (\sigma^2 + \alpha_3 M_t) &= \alpha_1 (e_{t-1}^2 - h_{t-1}) + (\alpha_1 + \alpha_2)(h_{t-1} - (\sigma^2 + \alpha_3 M_{t-1})). \end{aligned} \quad (\text{B.1})$$

Define

$$\begin{aligned} y_t &= r_t - \beta r_t^M - \gamma_0 - \gamma_1 r_{t-1} - \gamma_2 e_{t-1} - \gamma_3 M_t \\ x_t &= r_t^M \\ z_t &= \beta_t - \beta, \end{aligned}$$

where y_t contains all the predetermined parts of the model. Using these definitions, one can rewrite the first two lines of (B.1) as follows:

$$\begin{aligned} y_t &= x_t z_t + e_t \quad e_t \sim (0, h_t) \\ z_t &= \phi z_{t-1} + \xi_t \quad \xi_t \sim (0, \sigma_\xi^2), \end{aligned} \quad (\text{B.2})$$

where h_t is recursively determined by the GARCH specification (12c). The Kalman filter provides a way of composing the likelihood function of a model in state space form. The filter consists of a set of prediction and updating equations from which one can obtain the prediction errors conditionally on x_t

$$v_t = y_t - E(y_t | x_t, J_{t-1}), \quad (\text{B.3})$$

where J_{t-1} is the information set available at $t-1$, and their variance

$$f_t = E(v_t^2 | x_t, J_{t-1}), \quad (\text{B.4})$$

for all observations but the first. Assuming that the prediction errors follow a

normal or a t -distribution, one obtains the likelihood function by using the expressions in (A.4) and (A.5).

The Kalman filter equations are, Pagan (1980)

- (1) The prediction equations, where $a_{t|t-1}$ denotes $E_{t-1}(a_t)$

$$z_{t|t-1} = \phi z_{t-1|t-1}$$

$$p_{t|t-1} = \phi^2 p_{t-1|t-1} + \sigma_\xi^2$$

$$v_t = y_t - x_t z_{t|t-1} \quad (\text{B.5})$$

$$f_t = x_t^2 p_{t|t-1} + h_t$$

$$h_t = (\sigma^2 + \alpha_3 M_t) + \alpha_1 (e_{t-1|t-1}^2 - h_{t-1}) + (\alpha_1 + \alpha_2)(h_{t-1} - (\sigma^2 + \alpha_3 M_{t-1})).$$

- (2) The Kalman gain

$$k_t = p_{t|t-1} x_t / f_t. \quad (\text{B.6})$$

- (3) The updating equations, incorporating the new knowledge at time t

$$z_{t|t} = z_{t|t-1} + k_t v_t$$

$$p_{t|t} = (1 - k_t x_t) p_{t|t-1} \quad (\text{B.7})$$

$$e_{t|t} = y_t - x_t z_{t|t}$$

- (4) The initial conditions

$$h_1 = \sigma^2$$

$$e_{1|1} = 0$$

$$z_{1|1} = 0$$

$$p_{1|1} = \sigma_\varepsilon^2 / (1 - \phi^2).$$

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