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Version: Accepted Version

Article:

Dias, G.F. (2017) The time-varying GARCH-in-mean model. *Economics Letters*, 157. pp. 129-132. ISSN: 0165-1765

<https://doi.org/10.1016/j.econlet.2017.06.005>

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The time-varying GARCH-in-mean model

Gustavo Fruet Dias^{a,1}

^a*CREATES and Aarhus University, Fuglesangs Allé 4, 8210 Aarhus V, Denmark.
E-mail address: gdias@econ.au.dk*

Abstract

1 I propose an estimation strategy for the stochastic time-varying risk premium
2 parameter in the context of a time-varying GARCH-in-mean (TVGARCH-in-
3 mean) model. A Monte Carlo study shows that the proposed algorithm has good
4 finite sample properties. Using monthly excess returns on the CRSP index, I
5 document that the risk premium parameter is indeed time-varying and shows
6 high degree of persistence.

JEL classification numbers: C13, C15, C22, G12

Keywords: risk-return tradeoff, time-varying coefficients, iterative estimators,
GARCH-type models

¹I thank George Kapetanios, Cristina Scherrer and an anonymous referee for constructive suggestions. I acknowledge support from CREATES - Center for Research in Econometric Analysis of Time Series (DNRF78), funded by the Danish National Research Foundation.

7 **1. Introduction**

8 Asset pricing theories suggest that riskier assets should demand higher ex-
 9 pected returns. Using Merton’s (1973) theoretical framework, the conditional
 10 expectation of the market excess returns reads

$$\mathbb{E}(r_{t+1}^m | \mathcal{F}_t) - r_t^f = \lambda_t \text{Var}(r_{t+1}^m | \mathcal{F}_t), \quad (1)$$

11 where r_{t+1}^m and r_t^f are the returns on the market portfolio and risk-free asset,
 12 \mathcal{F}_t is the market-wide information available at time t , and λ_t is the coefficient
 13 of relative risk aversion defined as the elasticity of marginal value with respect
 14 to wealth. Most studies assume the risk-return trade-off is constant over time
 15 and linear in the variance, which is usually associated with the reasons behind
 16 mixed empirical evidences when estimating the risk-return trade-off (Linton
 17 and Perron (2003), Brandt and Wang (2010), Christensen, Dahl, and Iglesias
 18 (2012), among others). To address this issue, I adopt the time-varying GARCH-
 19 in-mean (TVGARCH-in-mean) model in the spirit of Anyfantaki and Demos
 20 (2016) which allows λ_t to be a time-varying stochastic process and put forward
 21 a feasible estimation strategy for λ_t (see references in Anyfantaki and Demos
 22 (2016) for variants of the TVGARCH-in-mean models). Specifically, I com-
 23 bine Giraitis, Kapetanios, and Yates’s (2013) time-varying kernel least squares
 24 estimator with Linton and Perron’s (2003) semiparametric iterative approach
 25 to estimate the time-varying risk premium coefficient. A Monte Carlo study
 26 shows that the proposed algorithm has good finite sample properties. Using the
 27 excess returns of the Center for Research on Security Prices (CRSP) index, I
 28 document that the risk premium parameter is indeed time-varying, alternating
 29 positive (statistically significant) and nonsignificant values over time.

30 **2. The time-varying GARCH-in-mean**

31 The generic TVGARCH-in-mean(p,q) is defined as:

$$r_t = \lambda_t \sigma_t + \epsilon_t, \quad (2)$$

$$\epsilon_t = \sigma_t \eta_t, \quad (3)$$

$$\sigma_t^2 = \omega + \sum_{i=1}^p \alpha_i \epsilon_{t-i}^2 + \sum_{i=1}^q \beta_i \sigma_{t-i}^2, \quad (4)$$

$$\epsilon_t^2 = \psi_0 + u_t + \sum_{i=1}^{\infty} \psi_i u_{t-i}, \quad (5)$$

32 where η_t is an independent and identically distributed (*iid*) zero mean process
 33 with unit variance; σ_t is a latent conditional standard deviation; (5) is the
 34 $MA(\infty)$ representation of the conditional variance equation; $u_t = \epsilon_t^2 - \sigma_t^2$ is a
 35 martingale difference sequence process; $\phi = (\omega, \alpha_1, \dots, \alpha_p, \beta_1, \dots, \beta_q)'$ collects the
 36 free parameters in (4); and $\psi_i := \varrho_i(\phi)$ $i = 1, 2, \dots$ are deterministic functions

37 of the elements in ϕ . Similarly as in Giraitis, Kapetanios, and Yates (2013),
 38 the time-varying risk premium parameters are assumed to evolve smoothly over
 39 time, so that it satisfies a local stability condition in the form of $\sup_{s: \|s-t\| \leq h} \|\lambda_t + \lambda_s\|_2 = O_p(h/t)$.
 40

41 Estimating the free parameters in (2) and (4) by maximum-likelihood is not
 42 a feasible alternative, as the class of TVGARCH-in-mean(p,q) models involves
 43 two unobserved processes: λ_t and ϵ_t . Anyfantaki and Demos (2016) address
 44 this issue in the context of the time-varying EGARCH(1,1)-in-mean model.
 45 Specifically, their work differs from mine in two ways. First, they parameterize
 46 the conditional variance as an EGARCH(1,1) model and, most importantly, λ_t
 47 as a stationary AR(1) process. By contrast, λ_t in (2) is assumed to satisfy
 48 $\sup_{s: \|s-t\| \leq h} \|\lambda_t + \lambda_s\|_2 = O_p(h/t)$, which encompasses the case of the driftless
 49 random walk process considered in Chou, Engle, and Kane (1992). Second,
 50 while I propose a kernel-based nonparametric method to estimate the time-
 51 varying risk premium parameter, Anyfantaki and Demos's (2016) estimation
 52 strategy is based on Bayesian methods (Markov chain Monte Carlo (MCMC)
 53 likelihood based estimation procedure).

54 I combine Linton and Perron's (2003) iterative semiparametric estimator
 55 with Giraitis, Kapetanios, and Yates's (2013) kernel-based least squares frame-
 56 work to estimate the free parameters $\theta = (\underline{\lambda}, \phi)'$, where $\underline{\lambda} = (\lambda_1, \dots, \lambda_T)'$. This
 57 method consists of recursively updating estimates of σ_t and u_t on each iteration,
 58 and then computing estimates of $\underline{\lambda}$ and ϕ . To this end, consider moment
 59 conditions based on (2) and (5),

$$\mathbb{E}[\sigma_t (r_t - \lambda_t \sigma_t)] = 0, \quad \text{for each } t = 1, 2, \dots, T, \quad (6)$$

$$\mathbb{E}[z_t u_t] = 0, \quad \text{with } z_t := \frac{\partial \left(\psi_0 + \sum_{i=1}^{\bar{q}} \psi_i u_{t-i} \right)}{\partial \phi}, \quad (7)$$

60 where (7) is truncated at some lag-order \bar{q} with $\bar{q} > p + q + 1$. Notably, (7) holds
 61 because u_t is a martingale difference sequence and z_t is a function of lagged
 62 values of u_t . It follows that estimating θ by the standard generalized method of
 63 moments (GMM) using the moments defined in (6) and (7) is not operational,
 64 as z_t and σ_t are latent variables. Using Linton and Perron's (2003) approach,
 65 rewrite (6) and (7) using estimates of σ_t and u_t obtained at some j iteration,

$$\mathbb{E}[\sigma_{j,t} (r_t - \lambda_{j+1,t} \sigma_{j,t})] = 0, \quad \text{for each } t = 1, 2, \dots, T, \quad (8)$$

$$\mathbb{E}[z_{j,t} u_{j+1,t}] = 0, \quad (9)$$

66 where $\sigma_{j,t}$ and $z_{j,t}$ denote the filtered estimates of σ_t and z_t based on $\widehat{\theta}_j$, and
 67 $u_{j+1,t} = \epsilon_{j,t}^2 - \psi_{j+1,0} - \sum_{i=1}^{\bar{q}} \psi_{j+1,i} u_{j,t-i}$ with $\epsilon_{j,t}^2 = (r_t - \lambda_{j+1,t} \sigma_{j,t})^2$. While the
 68 finite sample counterpart of (9) is given by the usual sample mean, computing
 69 the sample counterpart of (8) is less obvious. The work of Giraitis, Kapetanios,
 70 and Yates (2013) suggests the use of local kernels to construct operational

71 sample counterparts of (8). In turn, a feasible moment condition based on (8)
 72 reads

$$K_t^{-1} \sum_{\tau=1}^T k_{t,\tau} \sigma_{j,\tau} \left(r_\tau - \widehat{\lambda}_{j+1,t} \sigma_{j,\tau} \right) = 0, \quad \text{for each } t = 1, 2, \dots, T, \quad (10)$$

73 where $k_{t,\tau} = K((t-\tau)/H)$ denotes a kernel function such that $K(x) \geq 0$ for
 74 any $x \in \mathbb{R}$ is a continuous bounded function with a bounded first derivative and
 75 $\int K(x) dx = 1$; H is the bandwidth parameter satisfying $H = o(T/\ln(T))$ as
 76 $H \rightarrow \infty$; and $K_t = \sum_{\tau=1}^T k_{t,\tau}$. Notably, writing the moment conditions as in
 77 (10) is consistent with previous studies in the time-varying parameter literature
 78 which maximizes kernel weighted log-likelihood functions (see Robinson (1989),
 79 Giraitis, Kapetanios, Wetherilt, and Žikeš (2016), among others).

80 I use the fact that (10) is exactly identified for each t , and hence estimates
 81 of λ_t can be obtained independently of ϕ . In turn, estimates of θ are computed
 82 iteratively by a two-step procedure. The first step consists of solving (10) for
 83 each t , while the second step mimics the work of Linton and Perron (2003) and
 84 consists of estimating ϕ using the sample counterpart of (9). In practise, the
 85 kernel-based iterative estimator is as follows:

86 **Step 1:** Choose starting values $\widehat{\lambda}_0$ and $\widehat{\phi}_0$, such that $\widehat{\phi}_0$ satisfies the second-
 87 order stationarity conditions of the GARCH(1,1) model. Using $\widehat{\theta}_{0,t} =$
 88 $(\widehat{\lambda}_0, \widehat{\phi}_0)'$, compute recursively $\{\sigma_{0,t}^2\}_{t=1}^T$, and $\{u_{0,t}\}_{t=1}^T$ from (2)-(5).

89 **Step 2:** Given $\{\sigma_{0,t}^2\}_{t=1}^T$, calculate

$$\widehat{\lambda}_{1,t} = \left(\sum_{\tau=1}^T k_{t,\tau} \sigma_{0,\tau}^2 \right)^{-1} \sum_{\tau=1}^T k_{t,\tau} \sigma_{0,\tau} r_\tau, \quad \text{for each } t = 1, 2, \dots, T. \quad (11)$$

90 **Step 3:** Solving the sample counterpart of (9) is equivalent to estimate $\widehat{\phi}_1$ by
 91 nonlinear least squares. Calculate

$$\widehat{\phi}_1 = \arg \min_{\widehat{\phi}_1} \sum_{t=1}^T \left\{ \left(r_t - \widehat{\lambda}_{1,t} \sigma_{0,t} \right)^2 - \widehat{\psi}_{1,0} - \sum_{i=0}^{\bar{q}} \widehat{\psi}_{1,i} u_{0,t-1-i} \right\}^2. \quad (12)$$

92 **Step 4:** Update recursively $\{\sigma_{1,t}^2\}_{t=1}^T$ and $\{u_{1,t}\}_{t=1}^T$ based on $\widehat{\theta}_1$.

93 Repeat steps 2-4 j times until $\widehat{\theta}_j$ converges. Convergence occurs when
 94 $\|\widehat{\lambda}_j - \widehat{\lambda}_{j-1}\|_2 \leq \varepsilon$ and $\|\widehat{\phi}_j - \widehat{\phi}_{j-1}\|_2 \leq \varepsilon$, with ε set to 10^{-5} . Parameters on

95 the j^{th} iteration are given by:

$$\hat{\lambda}_{j,t} = \left[\sum_{\tau=1}^T k_{t,\tau} \sigma_{j-1,\tau}^2 \right]^{-1} \sum_{\tau=1}^T k_{t,\tau} \sigma_{j-1,\tau} r_{\tau}, \quad \text{for each } t = 1, 2, \dots, T, \quad (13)$$

$$\hat{\phi}_j = \arg \min_{\hat{\phi}_j} \sum_{t=1}^T \left[r_t - \hat{\lambda}_{j,t} \sigma_{j-1,t} \right]^2 - \hat{\psi}_{j,0} - \sum_{i=0}^{\bar{q}} \hat{\psi}_{j,i} u_{j-1,t-1-i} \Bigg]^2. \quad (14)$$

96 Finally, three inputs are still necessary to implement the above algorithm: the
 97 kernel function, the bandwidth parameter H , and the truncation lag \bar{q} . As in
 98 Giraitis, Kapetanios, and Yates (2013), three kernel functions are used: the
 99 Epanechnikov, Gaussian, and flat kernels. The choice of H follows from the
 100 Monte Carlo study conducted in Section 3.1, while \bar{q} is chosen to be proportional
 101 to $\ln(T)$, (Dufour and Jouini (2005)).

102 Asymptotic theory for the Quasi-Maximum Likelihood (QMLE) estimator in
 103 the GARCH-in-mean models is yet to be fully established. Conrad and Mammen
 104 (2016) give an important step forward and prove the asymptotic distribution
 105 of the QMLE estimator for the simple GARCH(1,1)-in-mean. As discussed in
 106 Linton and Perron (2003), the semiparametric GARCH-in-mean models offer
 107 additional complications compared to the standard GARCH-in-mean models,
 108 and hence rigorous inference is still not available. Similar difficulties arise in
 109 the TVGARCH-in-mean specification. In turn, this note follows Linton and
 110 Perron's (2003) approach as it briefly discusses the general conditions required
 111 for consistency and asymptotic normality; uses the wild bootstrap to conduct
 112 inference; and adopts a Monte Carlo study to assess the finite sample properties
 113 of the proposed iterative estimator.

114 The concept of asymptotic contraction mapping (ACM) developed in Do-
 115 minitz and Sherman (2005) is useful to guide the discussion on the asymptotic
 116 properties of the kernel iterative estimator. If a collection is an ACM, then it
 117 will have a unique fixed point that depends on the sample characteristics and
 118 hence the iterative procedure converges.² While the two-step procedure given
 119 in (13) and (14) is seen as the sample mapping, (8) and (9) are their popula-
 120 tion counterpart. Consistency and asymptotic normality require the population
 121 mapping to be an ACM, which implies, under some uniform convergence condi-
 122 tions, that the sample mapping is also an ACM and hence has an unique fixed
 123 point (regardless of the initial values). Combining Theorem 4 in Dominitz and
 124 Sherman (2005) with Giraitis, Kapetanios, and Yates's (2013) results, estimates
 125 of λ_t are expected to be \sqrt{H} consistent and asymptotically normally distributed,
 126 and estimates of ϕ are expected to be consistent and asymptotically normally
 127 distributed at the usual \sqrt{T} rate.

²See Dominitz and Sherman, 2005, p. 841 for a formal definition of the ACM.

128 **3. Numerical illustrations**

129 *3.1. Monte Carlo*

130 I simulate data from (2)-(4) where $p = q = 1$, η_t is normally distributed with
131 zero mean and variance equal to one, and λ_t follows a bounded random walk
132 process (see detailed discussion in the online Supplement). The sample size and
133 the number of replications are set to 2,000 and 1,000, respectively.

134 Table 1 displays the results for the kernel-based iterative estimator computed
135 with alternative bandwidth choices and the Epanechnikov, Gaussian, and flat
136 kernels. Results are reported in terms of the root mean squared error (RMSE)
137 and pointwise correlation between λ_t and the kernel-based estimates. The best
138 choices of bandwidth parameters, in terms of minimizing the RMSE, are $H =$
139 $T^{1/2}$ and $H = T^{6/10}$. These are also the bandwidths that deliver the highest
140 pointwise correlation (about 0.85) between the kernel-based estimates and the
141 true latent time-varying risk premium parameter. All combinations of kernel
142 methods and bandwidth parameters deliver unbiased estimates of $\phi = (\omega, \alpha, \beta)'$
143 (apart from $H = T^{2/10}$). Finally, convergence rates are greater than 98% for all
144 specifications, suggesting that (8) and (9) are ACMs.

145 *3.2. Empirical results*

146 I use excess returns of the CRSP value-weighted index aggregated on a
147 monthly basis. Figure 1 plots monthly estimates of λ_t and their 90% wild
148 bootstrap confidence bands from a TVGARCH(1,1)-in-mean model with band-
149 width $H = T^{6/10}$ (see the Supplement material for a Monte Carlo study showing
150 that the wild bootstrap produces valid inference). Not surprisingly, likewise the
151 semiparametric GARCH-in-mean models, the empirical confidence bands are
152 relatively wide, which reflects the difficulties associated with estimating the risk-
153 return trade-off (Linton and Perron (2003) and Christensen, Dahl, and Iglesias
154 (2012)). I find that the risk premium parameter is indeed time-varying, with $\hat{\lambda}_t$
155 assuming both positive (generally significant) and negative (insignificant) val-
156 ues. This finding sheds light on the mixed evidence on the risk-return literature
157 regarding the sign and significance of the risk premium parameter. Addition-
158 ally, in periods where $\hat{\lambda}_t$ is statistically significant, market volatility is low. On
159 contrary, when $\hat{\lambda}_t$ is not statistically significant, market volatility is high. This
160 indicates that identification of the risk premium parameter is problematic in
161 periods of high volatility (Rossi and Timmermann (2010)).

162 **4. Conclusion**

163 I introduce a kernel-based iterative estimator that combines the estimators in
164 Giraitis, Kapetanios, and Yates (2013) and Linton and Perron (2003) to estimate
165 the stochastic time-varying risk premium parameter in the TVGARCH(1,1)-in-
166 mean model. The Monte Carlo study shows that the kernel-based estimator
167 presents a good finite sample performance. I investigate the time-varying risk
168 premium for the CRSP index and find strong evidence that λ_t is indeed time-
169 varying.

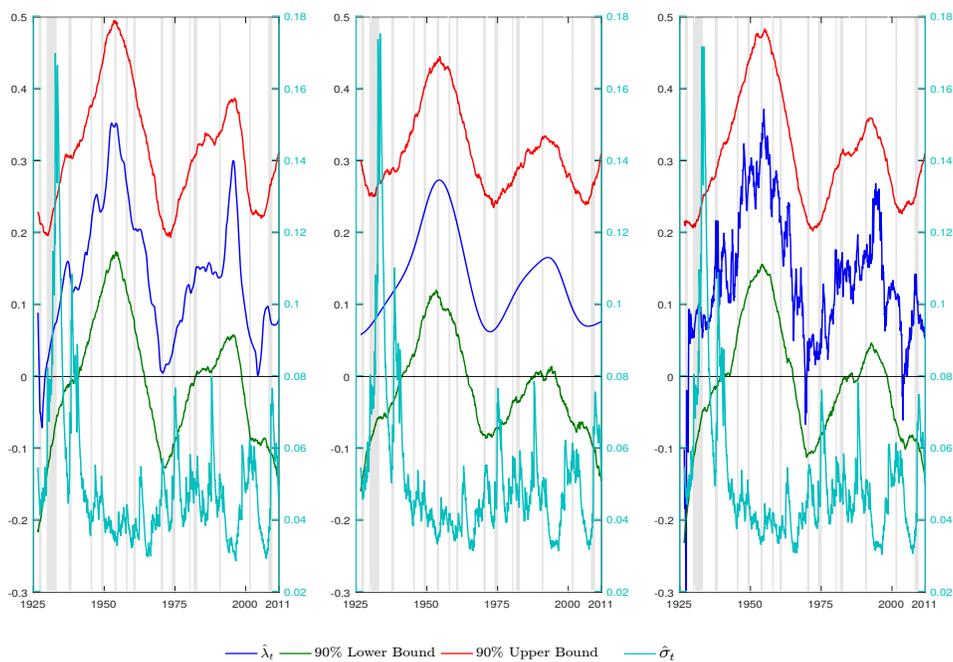
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208 versity of California, San Diego.

Table 1: Finite sample performance of the kernel-based iterative estimator

Bandwidth - H	Kernel	λ_t		RMSE			Mean		
		RMSE	Corr	ω	α	β	ω	α	β
$T^{2/10}$	Epanechnikov	0.39	0.60	0.01	0.03	0.10	0.02	0.07	0.84
$T^{2/10}$	Gaussian	0.25	0.71	0.01	0.02	0.06	0.01	0.05	0.88
$T^{2/10}$	Flat	0.36	0.61	0.01	0.02	0.07	0.01	0.06	0.86
$T^{3/10}$	Epanechnikov	0.26	0.72	0.01	0.02	0.05	0.01	0.05	0.88
$T^{3/10}$	Gaussian	0.18	0.81	0.01	0.02	0.05	0.01	0.05	0.89
$T^{3/10}$	Flat	0.24	0.73	0.01	0.02	0.05	0.01	0.05	0.89
$T^{4/10}$	Epanechnikov	0.18	0.80	0.01	0.02	0.05	0.01	0.05	0.90
$T^{4/10}$	Gaussian	0.13	0.86	0.01	0.02	0.06	0.01	0.05	0.90
$T^{4/10}$	Flat	0.17	0.81	0.01	0.02	0.05	0.01	0.05	0.90
$T^{5/10}$	Epanechnikov	0.13	0.86	0.01	0.02	0.05	0.01	0.05	0.90
$T^{5/10}$	Gaussian	0.11	0.88	0.01	0.02	0.05	0.01	0.05	0.90
$T^{5/10}$	Flat	0.13	0.86	0.01	0.02	0.05	0.01	0.05	0.90
$T^{6/10}$	Epanechnikov	0.12	0.88	0.01	0.02	0.06	0.01	0.05	0.89
$T^{6/10}$	Gaussian	0.13	0.86	0.01	0.02	0.06	0.01	0.05	0.89
$T^{6/10}$	Flat	0.12	0.86	0.01	0.02	0.06	0.01	0.05	0.89
$T^{7/10}$	Epanechnikov	0.13	0.85	0.01	0.02	0.05	0.01	0.05	0.90
$T^{7/10}$	Gaussian	0.16	0.79	0.01	0.02	0.05	0.01	0.05	0.90
$T^{7/10}$	Flat	0.15	0.81	0.01	0.02	0.05	0.01	0.05	0.90
$T^{8/10}$	Epanechnikov	0.17	0.77	0.01	0.02	0.05	0.01	0.05	0.89
$T^{8/10}$	Gaussian	0.21	0.68	0.01	0.02	0.05	0.01	0.05	0.89
$T^{8/10}$	Flat	0.19	0.67	0.01	0.02	0.05	0.01	0.05	0.89

RMSE accounts for root mean squared error; Corr is the pointwise correlation between λ_t and $\widehat{\lambda}_t$; λ_t is defined as a bounded random walk process with upper and lower bounds given by 0.90 and -0.90, respectively; and $\phi = (0.01, 0.05, 0.9)'$.

Figure 1: Time-varying risk premium estimation and conditional standard deviation



The left-hand-side, center and right-hand-side graphs display estimates of λ_t with the Epanechnikov, Gaussian, and the flat kernel functions, respectively. Estimates of λ_t and the 90% confidence intervals are on the left axis. The bandwidth parameter is equal to $H = T^{6/10}$. The conditional standard deviation is in light blue on the right axis. I perform 1000 replications in the bootstrap algorithm. The shaded areas account for the recession periods from the National Bureau of Economic Research (NBER).