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TESTING THE COMPOUNDING STRUCTURE OF THE CP-INARCH MODEL

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Abstract: A statistical test to distinguish between a Poisson INARCH model and a Compound Poisson INARCH model is proposed, based on the form of the probability generating function of the compounding distribution of the conditional law of the model.

The normality of the test statistics' asymptotic distribution is established, either in the case where the model parameters are specified, or when such parameters are consistently estimated. As the test statistics law involves the moments of inverse conditional means of the Compound Poisson INARCH process, the analysis of their existence and calculation is performed by two approaches.

A simulation study illustrating the finite-sample performance of this test methodology in what concerns its size and power concludes the paper.

Keywords: Count-data time series; compound Poisson distribution; INGARCH model; diagnostic tests; inverse moments.
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1. Introduction

The *INGARCH models*, which constitute an *in*teger-valued counterpart to the conventional *g*eneralized *a*uto*r*egressive *c*onditional *h*eteroskedasticity models, were introduced by Heinen (2003); Ferland et al. (2006). Instead of considering the conditional variances as in the conventional GARCH model, they assume the conditional means $M_t := E[X_t \mid X_{t-1}, \ldots]$ to satisfy a linear recursion,

$$M_t \;=\; \alpha_0 \;+\; \sum_{i=1}^{p} \alpha_i\, X_{t-i} \;+\; \sum_{j=1}^{q} \beta_j\, M_{t-j}, \tag{1}$$

where $\alpha_0 > 0$ and $\alpha_1, \ldots, \alpha_p, \beta_1, \ldots, \beta_q \geq 0$. Having specified the conditional mean, the most common type of conditional distribution is the Poisson one, i. e., $X_t \sim \mathrm{Poi}(M_t)$, leading to the *Poisson* INGARCH model, where existence and strict stationarity with finite first and second order moments can be shown under the condition $\sum_{i=1}^{p} \alpha_i + \sum_{j=1}^{q} \beta_j < 1$ (Ferland et al., 2006). The Poisson INGARCH model was further investigated by several authors including Fokianos et al. (2009); Weiß (2009); Neumann (2011). But also different choices for the conditional distribution have been considered in the

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literature, see, e. g., Xu et al. (2012); Zhu (2012); Gonçalves et al. (2015a,b) and the discussion below. The INGARCH models exhibit an ARMA-like autocorrelation structure, and they are particularly well-suited for time series of counts showing overdispersion, i. e., which have a variance larger than the mean. In particular, the case $q = 0$, referred to as an INARCH(p) model, has the same autocorrelation structure as a usual AR(p) model. So the INARCH model, which is the main focus of the present work, constitutes a count-data type of autoregressive model.

The standard INARCH model has a conditional Poisson distribution and is therefore conditionally equidispersed. Its unconditional distribution, however, exhibits overdispersion, where the degree of overdispersion depends on the dependence parameters $\alpha_1, \ldots, \alpha_p$. To overcome this limitation, Xu et al. (2012) proposed the family of *dispersed INARCH models* (DINARCH), which again assume a linear relationship for the conditional mean, but with an additional scaling factor $\theta > 0$ for the conditional variance:

$$M_t \;:=\; E[X_t \mid X_{t-1}, \ldots] \;=\; \alpha_0 + \textstyle\sum_{i=1}^{p} \alpha_i X_{t-i}, \qquad V[X_t \mid X_{t-1}, \ldots] \;=\; \theta\, M_t. \tag{2}$$

So the standard Poisson INARCH model is an instance of the DINARCH model with $\theta = 1$. A more comprehensive instance of the DINARCH model is obtained from a family of INGARCH models that was recently developed by Gonçalves et al. (2015a), who proposed to use a conditional compound Poisson (CP) distribution (Johnson et al., 2005). The CP-INARCH model to be considered in the sequel is defined by the conditional probability generating function (pgf)

$$\mathrm{pgf}_{X_t|X_{t-1},\ldots}(z) \;=\; \exp\Big(\tfrac{M_t}{H'(1)}\,\big(H(z)-1\big)\Big) \qquad \text{with } M_t \text{ according to (2)}, \tag{3}$$

where $H(z)$ denotes the pgf of the compounding distribution (assumed to be normalized to $H(0) = 0$ for uniqueness). From Theorem 5 in Gonçalves et al. (2015a), we know that the above condition $\sum_{i=1}^{p} \alpha_i < 1$ again guarantees the existence of a strictly stationary and ergodic solution to the CP-INARCH model (3), and this solution has finite first and second order moments. The CP-INARCH model constitutes an instance of the DINARCH model, where

$$V[X_t \mid X_{t-1}, \ldots] \;=\; M_t\, \underbrace{\big(1 + H''(1)/H'(1)\big)}_{=\theta}. \tag{4}$$

1.1 Example (Special CP-INARCH Models) Choosing $H(z) = z$, we obtain the standard Poisson INARCH model. But also the NB-INARCH(p) model (*negative binomial*) proposed by Xu et al. (2012) is a special type of CP-INARCH(p) model, where the compounding distribution is a log-series distribution (Johnson et al., 2005),

$$H(z) \;=\; 1-\frac{\ln\big(\theta+(1-\theta)z\big)}{\ln\theta} \qquad \text{with } H'(1) = -\frac{1-\theta}{\ln\theta}, \quad H''(1) = \frac{(1-\theta)^2}{\ln\theta}. \tag{5}$$

Hence, we simply have $1 + H''(1)/H'(1) \;=\; \theta$. Further examples include the INARCH model proposed by Zhu (2012) having a conditional generalized Poisson (GP) distribution, and the one by Gonçalves et al. (2015b) having a conditional *Neyman type-A* (NTA) distribution. The latter has a Poisson compounding structure: the NTA($\mu/\phi,\ \phi$)-distribution is defined by the pgf (Johnson et al., 2005)

$$\mathrm{pgf}(z) \;=\; \exp\left(\frac{\mu}{\phi}\left(e^{\phi(z-1)}-1\right)\right) \;=\; \exp\left(\mu\,\frac{1-e^{-\phi}}{\phi}\left(\frac{e^{\phi z}-1}{e^{\phi}-1}-1\right)\right), \tag{6}$$

and for the NTA-INARCH model, the mean parameter μ is replaced by M_t. The compounding pgf, $H(z) \;=\; (e^{\phi z}-1)/(e^{\phi}-1)$ with $H^{(k)}(z) \;=\; \phi^k\, e^{\phi z}/(e^{\phi}-1)$, is the one from the zero-truncated Poisson distribution and therefore satisfies the normalization constraint $H(0) = 0$. In particular, we have $1 + H''(1)/H'(1) \;=\; 1 + \phi$. In the sequel, we shall consider the problem of distinguishing between the simple Poisson INARCH model and true CP-INARCH model, i. e., we are confronted with the following hypotheses:

$$\begin{aligned} &H_0: \;\; (X_t)_{\mathbb{Z}} \text{ is a Poisson INARCH process} \quad (\text{i. e., } H(z) = z); \\ &H_1: \;\; (X_t)_{\mathbb{Z}} \text{ is a true CP-INARCH process} \quad (\text{i. e., } H(z) \neq z). \end{aligned} \tag{7}$$

For this purpose, in Section 2, we develop a general approach for analyzing the compounding structure of a CP-INARCH model. This approach is then used in Section 3 to develop a test procedure for the first-order INARCH model, where the test statistic involves the factorial moment of order r of X_t. The case of specified parameters is briefly described, and the normality of the test statistics' asymptotic distribution under the null hypothesis (7) is established either in this case, or in that one, important in practice, where such parameters are consistently estimated. As the test statistics law involves the moments of inverse conditional means of the Compound Poisson INGARCH process, the analysis of their existence and calculation is performed by two approaches. In Section 4, a simulation study is presented illustrating the finite-sample performance of this test methodology in what concerns its size and power for different values of r. Section 5 concludes, and Appendix A includes the detailed derivations.

2. Analyzing the Compounding Structure of CP-INARCH Models

Given the past observations $X_{t-1}, \ldots$, the conditional CP model in (3) implies that first a stopping count N_t is generated according to $\mathrm{Poi}\big(M_t/H'(1)\big)$, and then (independently) the N_t i.i.d. counts $Y_{t,1}, \ldots, Y_{t,N_t}$ according to the compounding model having the pgf $H(z)$, also see Johnson et al. (2005). The next observation is obtained as $X_t = Y_1 + \ldots + Y_{N_t}$.

To distinguish between the null hypothesis H_0 and the alternative hypothesis H_1 according to (7), information about $H(z)$ is required, the unique pgf of the $Y_{t,i}$. In fact, it suffices to check if the mean $H'(1)$ of the compounding distribution is equal to 1 (H_0) or larger than 1 (H_1). Hence, the mean statistic

$$\frac{1}{T}\sum_{t=1}^{T} \frac{Y_{t,1} + \ldots + Y_{t,N_t}}{N_t} \;=\; \frac{1}{T}\sum_{t=1}^{T} \frac{X_t}{N_t}$$

would be a reasonable candidate to infer $H'(1)$. But we do not observe N_t in practice, we only know that it has mean $M_t/H'(1)$. Therefore, we may consider a slightly modified version,

$$\frac{1}{T}\sum_{t=1}^{T} \frac{X_t}{M_t},$$

which we expect to give values close to 1. Note that the summands X_t/M_t are just the residuals ε_t as defined in Zhu & Wang (2010). To be more precise, for an underlying INARCH(p) model structure, the statistic

$$\widehat{C}_p \;:=\; \frac{1}{T-p}\sum_{t=p+1}^{T} \frac{X_t}{M_t} \;=\; \frac{1}{T-p}\sum_{t=p+1}^{T} \frac{X_t}{\alpha_0 + \sum_{i=1}^{p} \alpha_i X_{t-i}} \tag{8}$$

could be computed from the available data $X_1, \ldots, X_T$ and from the parameters $\alpha_0, \ldots, \alpha_p$ of the null model. Does this statistic allow to distinguish between H_0 and H_1?

Since the conditional mean $E[X_t \mid X_{t-1},\ldots] = M_t$ for *any* CP-INARCH process according to (2), we necessarily have

$$E\left[\frac{X_t}{M_t}\right] \;=\; 1, \qquad Cov\left[\frac{X_t}{M_t}, \frac{X_{t-k}}{M_{t-k}}\right] \;=\; 0 \quad \text{for } k \geq 1,$$

which immediately follows by applying the laws of total expectation and covariance. For the variance, we obtain

$$\begin{aligned} V\left[\frac{X_t}{M_t}\right] \;&=\; V\left[\frac{E[X_t \mid X_{t-1},\ldots]}{M_t}\right] \;+\; E\left[\frac{V[X_t \mid X_{t-1},\ldots]}{M_t^2}\right] \\ &=\; 0 \;+\; \left(1 + \frac{H''(1)}{H'(1)}\right) E\left[\frac{1}{M_0}\right] \end{aligned}$$

because of (4) and because of stationarity. Here, $E\big[M_0^{-1}\big]$ is an inverse moment with $0 < M_0^{-1} \leq 1/\alpha_0$. Altogether, the summands in (8) are always uncorrelated such that we finally obtain:

$$E\big[\widehat{C}_p\big] \;=\; 1, \qquad V\big[\widehat{C}_p\big] \;=\; \frac{1}{T-p}\left(1 + \frac{H''(1)}{H'(1)}\right) E\left[\frac{1}{M_0}\right]. \tag{9}$$

(9) implies that the variance of $\widehat{C}_p$ is inflated by $1+H''(1)/H'(1)$ (compared to the null model with $H''(1) = 0$). But the mean of $\widehat{C}_p$ is always 1, independent of the type of CP-INARCH(p) model.

Therefore, we consider a higher-order extension of the test statistic $\widehat{C}_p$ from (8) such that also its mean is affected if violating H_0. Considering that the r^{th} factorial moment ($r \in \mathbb{N}$) of the Poisson distribution Poi(μ) just equals μ^r (Johnson et al., 2005), it follows that

$$E\big[(X_t)_{(r)} \mid X_{t-1},\ldots\big] \;=\; M_t^r,$$

where $x_{(r)} = x\cdots(x-r+1)$ denotes the falling factorial. So we define

$$\widehat{C}_{p;\,r} \;:=\; \frac{1}{T-p}\sum_{t=p+1}^{T}\frac{(X_t)_{(r)}}{M_t^r} \;=\; \frac{1}{T-p}\sum_{t=p+1}^{T}\frac{X_t(X_t-1)\cdots(X_t-r+1)}{\big(\alpha_0 + \sum_{i=1}^{p}\alpha_i X_{t-i}\big)^r}, \tag{10}$$

where $\widehat{C}_p = \widehat{C}_{p;\,1}$. If $(X_t)_{\mathbb{Z}}$ is Poisson INARCH(p) with given parameter values for $\alpha_0, \alpha_1, \ldots, \alpha_p$ (i. e., if H_0 holds), we obtain with analogous computations as for (9) that

$$E\left[\frac{(X_t)_{(r)}}{M_t^r}\right] \;=\; 1, \qquad Cov\left[\frac{(X_t)_{(r)}}{M_t^r}, \frac{(X_{t-k})_{(r)}}{M_{t-k}^r}\right] \;=\; 0 \quad \text{for } k \geq 1.$$

To compute the variance, we need the following identity for falling factorials:

$$x_{(r)}^2 \;=\; \textstyle\sum_{k=0}^r \binom{r}{k}^2 k!\, x_{(2r-k)}.$$

Then we obtain

$$\begin{aligned}
V\left[\frac{(X_t)_{(r)}}{M_t^r}\right] \;&=\; V\left[\frac{E\big[(X_t)_{(r)} \mid X_{t-1},\ldots\big]}{M_t^r}\right] \;+\; E\left[\frac{V\big[(X_t)_{(r)} \mid X_{t-1},\ldots\big]}{M_t^{2r}}\right]\\
&=\; 0 + E\left[\frac{E\big[(X_t)_{(r)}^2 \mid X_{t-1},\ldots\big]}{M_t^{2r}} - 1\right] = \left(\sum_{k=0}^r \binom{r}{k}^2 k!\, E\big[M_t^{-k}\big]\right) - 1\\
&=\; \sum_{k=1}^r \binom{r}{k}^2 k!\, E\big[M_t^{-k}\big]. \qquad (11)
\end{aligned}$$

Overall, under H_0, i. e., for a Poisson INARCH(p) model, we obtain that

$$E\big[\widehat{C}_{p;\,r}\big] \;=\; 1, \qquad V\big[\widehat{C}_{p;\,r}\big] \;=\; \frac{1}{T-p}\sum_{k=1}^r \binom{r}{k}^2 k!\, E\big[M_t^{-k}\big]. \qquad (12)$$

The following example considers the case of the alternative H_1.

2.1 Example (Second Order Statistic) Let us consider the second order statistic $\widehat{C}_{p;\,2}$, i. e., the case $r = 2$. Under H_0, (12) implies

$$E\big[\widehat{C}_{p;\,2}\big] \;=\; 1, \qquad V\big[\widehat{C}_{p;\,2}\big] \;=\; \frac{1}{T-p}\Big(4\,E\big[M_0^{-1}\big] + 2\,E\big[M_0^{-2}\big]\Big).$$

If, in contrast, the Poisson assumption is violated (H_1), then also the mean becomes sensitive to such a violation. For an underlying CP-INARCH(p) model, we have

$$\begin{aligned}
E\big[(X_t)_{(2)} \mid X_{t-1},\ldots\big] \;&=\; V[X_t \mid X_{t-1},\ldots] + E[X_t \mid X_{t-1},\ldots]^2 - E[X_t \mid X_{t-1},\ldots]\\
&\overset{(4)}{=}\; M_t \underbrace{\left(1 + \frac{H''(1)}{H'(1)}\right)}_{=\theta} + M_t^2 - M_t \;=\; M_t^2 + (\theta - 1)\, M_t,
\end{aligned}$$

such that

$$E\big[\widehat{C}_{p;\,2}\big] \;=\; E\left[\frac{(X_t)_{(2)}}{M_t^2}\right] \;=\; 1+(\theta-1)\,E\big[M_0^{-1}\big].$$

Therefore, $\widehat{C}_{p;\,2}$ might be a useful statistic to distinguish between H_0 and H_1 in practice.

3. Testing the CP-INARCH's Compounding Structure

From now on, we concentrate on the case of first-order autoregression, i. e., on the case $p = 1$. According to (7), H_0 assumes the two-parametric *Poisson INARCH(1) model* given by

$$X_t \;\big|\; X_{t-1}, X_{t-2}, \ldots \quad \sim\; \mathrm{Poi}(\alpha_0 + \alpha_1 \cdot X_{t-1}). \tag{13}$$

Though being a rather simple model, it has already found a number of real applications, e. g., to monthly claims counts (Weiß, 2009), to download counts (Zhu & Wang, 2010), to counts of iceberg orders (Jung & Tremayne, 2011), and to monthly strike counts data (Weiß, 2010). A Poisson INARCH(1) process is a stationary, ergodic Markov chain (Ferland et al., 2006; Zhu & Wang, 2011) with simple Poisson probabilities as the transition probabilities. According to Neumann (2011), it is β-mixing (and hence also α-mixing) with exponentially decreasing weights. All moments of a Poisson INARCH(1) process exist (Ferland et al., 2006), and they can be determined according to the recursive scheme provided by Weiß (2009, 2010), see equation (22) below.

For the first-order version of the DINARCH model (2), unconditonal mean and variance are given by (Xu et al., 2012, 4.3)

$$\mu \;=\; \frac{\alpha_0}{1-\alpha_1} \qquad \text{and} \qquad \sigma^2 \;=\; \frac{\theta}{1-\alpha_1^2}\cdot\frac{\alpha_0}{1-\alpha_1}. \tag{14}$$

So θ allows to control the degree of overdispersion independently of α_1.

3.1. Case of Specified Parameters. Let us investigate the distribution of the statistics $\widehat{C}_{1;\,r}$ introduced in the previous Section 2 under H_0, in the case of the Poisson INARCH(1) model with specified parameter values for α_0, α_1. We denote the (inverse) moments

$$q_{k,l} := q_{k,l}(\alpha_0, \alpha_1) \;:=\; E\Big[\frac{X_0^k}{(\alpha_0 + \alpha_1\, X_0)^l}\Big] \qquad \text{for } k, l \geq 0. \tag{15}$$

The moments $q_{k,l}$ from (15) are just the stationary marginal moments for $l = 0$, and for $l > 0$, they are easily computed numerically from the stationary marginal distribution of the Poisson INARCH(1) process $(X_t)_{\mathbb{Z}}$, see Section 3.3 below. The $q_{k,l}$ allow us to rewrite (12) as

$$E\big[\widehat{C}_{p;\,r}\big] \;=\; 1, \qquad V\big[\widehat{C}_{p;\,r}\big] \;=\; \tfrac{1}{T-1}\,\textstyle\sum_{k=1}^{r}\,\binom{r}{k}^{2}\,k!\,q_{0,k}. \tag{16}$$

As stated above, we know that the null model, the Poisson INARCH(1) model, is α-mixing with exponentially decreasing weights and has existing moments up to any order. So we apply the central limit theorem of Ibragimov (1962) to obtain that the statistics $\widehat{C}_{1;\,r}$ are even asymptotically normally distributed. Hence, one could test the null of a Poisson INARCH(1) model against the alternative of a true CP-INARCH(1) model based on the resulting approximate normal distribution for $\widehat{C}_{1;\,r}$.

These asymptotics, however, only hold for the case of specified H_0 parameters, since these are required to compute the statistics $\widehat{C}_{1;\,r}$. In practice, however, one usually has to estimate these parameters. Plugging-in these estimators into the definition of $\widehat{C}_{1;\,r}$, we obtain a statistic with a different asymptotic distribution than the one mentioned before. So to make the test applicable in practice, further investigations are required.

3.2. Case of Estimated Parameters. To derive an asymptotic approximation to the distribution of $\widehat{C}_{1;\,r}$ under H_0 *but in the presence of estimated parameters*, say, $\hat{\alpha}_0$ and $\hat{\alpha}_1$, we shall look at the first-order Taylor approximation of

$$\widehat{C}_{1;\,r}(\alpha_0,\alpha_1) \;=\; \frac{1}{T-1}\,\sum_{t=2}^{T}\frac{(X_t)_{(r)}}{(\alpha_0+\alpha_1\,X_{t-1})^{r}},$$

which has the partial derivatives

$$\tfrac{\partial}{\partial\alpha_0}\,\widehat{C}_{1;\,r} \;=\; \tfrac{1}{T-1}\,\sum_{t=2}^{T}\frac{-r\,(X_t)_{(r)}}{(\alpha_0+\alpha_1\,X_{t-1})^{r+1}}, \qquad \tfrac{\partial}{\partial\alpha_1}\,\widehat{C}_{1;\,r} \;=\; \tfrac{1}{T-1}\,\sum_{t=2}^{T}\frac{-r\,(X_t)_{(r)}\,X_{t-1}}{(\alpha_0+\alpha_1\,X_{t-1})^{r+1}}. \tag{17}$$

By conditioning, it follows that

$$E\Big[\frac{(X_t)_{(r)}}{(\alpha_0+\alpha_1\,X_{t-1})^{r+1}}\Big] \;=\; q_{0,1}, \qquad E\Big[\frac{(X_t)_{(r)}\,X_{t-1}}{(\alpha_0+\alpha_1\,X_{t-1})^{r+1}}\Big] \;=\; q_{1,1}, \tag{18}$$

where we used the abbreviation from (15). So we approximate $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ by

$$\widetilde{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1) \;:=\; \widehat{C}_{1;r}(\alpha_0, \alpha_1) \;-\; r\, q_{0,1}\,\big(\hat{\alpha}_0 - \alpha_0\big) \;-\; r\, q_{1,1}\,\big(\hat{\alpha}_1 - \alpha_1\big), \tag{19}$$

and an approximation of the distribution of $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ is obtained by deriving the distribution of $\widetilde{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$.

From now on, we shall use the usual moment estimators $\hat{\alpha}_0 := \bar{X}\big(1-\hat{\rho}(1)\big)$ and $\hat{\alpha}_1 := \hat{\rho}(1)$, the asymptotic distribution of which is studied in Weiß & Schweer (2016). Using the bias approximations for $\hat{\alpha}_0, \hat{\alpha}_1$ given there, it immediately follows that

$$\begin{aligned} E\big[\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)\big] \;&\approx\; E\big[\widehat{C}_{1;r}(\alpha_0, \alpha_1)\big] - r\, q_{0,1}\, E\big[\hat{\alpha}_0 - \alpha_0\big] - r\, q_{1,1}\, E\big[\hat{\alpha}_1 - \alpha_1\big] \\ &\approx\; 1 \;-\; r\,\tfrac{q_{0,1}}{T-1}\, \Big(\tfrac{1+3\alpha_1}{1-\alpha_1}\,\alpha_0 + \tfrac{2\alpha_1^2(1+2\alpha_1^2)}{1-\alpha_1^3}\Big) \;+\; r\,\tfrac{q_{1,1}}{T-1}\, \Big(1+3\alpha_1 + \tfrac{\alpha_1}{\alpha_0}\,\Big(1+\tfrac{2\alpha_1(1+2\alpha_1^2)}{1+\alpha_1+\alpha_1^2}\Big)\Big). \end{aligned} \tag{20}$$

The derivation of the asymptotic variance of the approximate quantity (19), however, is more demanding, see Appendix A.1 for the details. We finally obtain the approximate variance $\sigma_{1;r}^2/(T-1)$ with

$$\begin{aligned} \sigma_{1;r}^2 \;=\; &\textstyle\sum_{k=1}^{r} \binom{r}{k}^2 k!\, q_{0,k} \;-\; 2r^2\, q_{0,1} \;+\; r^2\, q_{0,1}^2\, \tfrac{\alpha_0}{1-\alpha_1}\big(\alpha_0(1+\alpha_1) + \tfrac{1+2\alpha_1^4}{1+\alpha_1+\alpha_1^2}\big) \\ &+ r^2\, q_{1,1}^2\,(1-\alpha_1^2)\big(1+\tfrac{\alpha_1(1+2\alpha_1^2)}{\alpha_0(1+\alpha_1+\alpha_1^2)}\big) \;-\; 2r^2\, q_{0,1}q_{1,1}\,\big(\alpha_0(1+\alpha_1) + \tfrac{(1+2\alpha_1)\alpha_1^3}{1+\alpha_1+\alpha_1^2}\big). \end{aligned} \tag{21}$$

So the test statistics $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ can now be applied in practice by choosing the critical values from a normal distribution with mean and variance given according to (20) and (21), respectively.

3.3. Inverse Moments. Before investigating the finite-sample performance of the proposed test, some background on the (numerical) computation of the Poisson INARCH(1)'s inverse moments is required. Equation (15) defines the moments

$$q_{k,l} \;=\; E\Big[\frac{X_0^k}{(\alpha_0+\alpha_1\, X_0)^l}\Big] \qquad \text{for } k,l \geq 0,$$

which are just the stationary marginal moments μ_k for $l = 0$. These can be computed exactly in two steps. First, the marginal cumulants κ_k are calculated according to the scheme provided by Weiß (2009, 2010). Denoting the Stirling numbers of the first kind (Douglas, 1980, Appendix 9.1) by $s_{k,j}$, it holds that

$$\kappa_1 \;=\; \tfrac{\alpha_0}{1-\alpha_1}, \qquad \kappa_k \;=\; -(1-\alpha_1^k)^{-1}\cdot \textstyle\sum_{j=1}^{k-1} s_{k,j}\cdot\kappa_j \quad \text{for } k\geq 2. \tag{22}$$

In the second step, these cumulants are transformed into the moments μ_k via Smith (1995)

$$\mu_k \;=\; \textstyle\sum_{j=0}^{k-1}\binom{k-1}{j}\,\kappa_{k-j}\,\mu_j \qquad \text{for } k\geq 1. \tag{23}$$

So it remains to consider the case $l>0$. Applying the binomial sum formula to $X_0^k \;=\; \alpha_1^{-k}\,\big((\alpha_0+\alpha_1 X_0)-\alpha_0\big)^k$, we obtain

$$q_{k,l} \;=\; \sum_{j=0}^{k}\binom{k}{j}(-1)^{k-j}\,\frac{\alpha_0^{k-j}}{\alpha_1^k}\cdot E\big[(\alpha_0+\alpha_1\,X_0)^{j-l}\big], \tag{24}$$

$$\text{where} \qquad E\big[(\alpha_0+\alpha_1\,X_0)^{j-l}\big] \;=\; \begin{cases} q_{0,l-j} & \text{if } j<l,\\ \sum_{i=0}^{j-l}\binom{j-l}{i}\,\alpha_0^{j-l-i}\alpha_1^i\,\mu_i & \text{if } j\geq l,\end{cases}$$

where the last expression again follows from the binomial sum formula. So equation (24) implies that $q_{k,l}$ can be traced back to either the usual moments μ_k or to purely inverse moments of the form $q_{0,l}$. So it suffices to discuss how to obtain the $q_{0,l} \;=\; E\big[(\alpha_0+\alpha_1\,X_0)^{-l}\big]$ for $l\geq 1$, the value of which is obviously bounded by $0<q_{0,l}<\alpha_0^{-l}$.

If only being interested in the numerical computation of $q_{0,l}$ (as required for applying the proposed $\widehat{C}_{1;r}$-test), the Markov chain approximation (Weiß, 2010) can be used: we compute the Poisson INARCH(1)'s transition probabilities

$$p_{r|s} := P(X_t=r \mid X_{t-1}=s) \;=\; \exp\left(-\alpha_0-\alpha_1\,s\right)(\alpha_0+\alpha_1\,s)^r\,/\,r!$$

for all $0\leq r,s\leq M$ (with M sufficiently large), define the matrix $\mathbf{P}_M := (p_{r|s})_{r,s=0,\ldots,M}$, and numerically solve the eigenvalue problem $\mathbf{P}_M\,\boldsymbol{p}=\boldsymbol{p}$ (invariance equation) in $\boldsymbol{p}$. The normalized eigenvector $\boldsymbol{p}$ (i. e., with nonnegative entries summing up to one) is used as an approximation for the marginal probabilities $\big(P(X_t=0),\ldots,P(X_t=M)\big)^\top$, and $q_{0,l}$ is approximated by the sum

$$q_{0,l} \;\approx\; \sum_{r=0}^{M}\frac{1}{(\alpha_0+\alpha_1\,r)^l}\cdot p_r. \tag{25}$$

The calculation of $q_{0,l}$ may also be performed following the method provided in Adell et al. (1996) to calculate negative moments of nonnegative random variables, and taking into account that the distribution of X_t given all the past is Poisson with mean $M_t = \alpha_0 + \alpha_1 X_{t-1}$. Let us begin by stating a result relating the radius of convergence of the moment generating function of M_1 with the values of the coefficient α_1.

3.1 Lemma *If the moment generating function of* M_1, $\mathrm{mgf}_{M_1}(u) = E\big[\exp(uM_1)\big]$, *is defined for every* $u \in (u_1; u_2)$, *where* $u_1 < 0 < u_2$ *with* $\min\{-u_1, u_2\} = b$, *then* $\alpha_1 < \frac{\ln(b+1)}{u}$ *for all* $0 < u < b$.

Proof: For $u \in (-b; b)$, we have

$$\begin{aligned}
\mathrm{mgf}_{M_1}(u) &= E\big[\exp(uM_1)\big] = E\Big[E\big[\exp\big(u\,(\alpha_0 + \alpha_1 X_0) \mid X_{-1}\big)\big]\Big] \\
&= \exp(u\alpha_0)\, E\big[\exp\big(M_0\,(\exp(u\alpha_1) - 1)\big)\big] \\
&= \exp(u\alpha_0)\, \mathrm{mgf}_{M_0}\big(\exp(u\alpha_1) - 1\big).
\end{aligned}$$

Then

$$-b < \exp(u\alpha_1) - 1 < b \quad \Leftrightarrow \quad -\infty < u\alpha_1 < \ln(b+1)\,,$$

and for all $0 < u < b$, we obtain that $\alpha_1 < \frac{\ln(b+1)}{u}$. ■

To find $q_{0,l} = E[M_t^{-l}]$ for $l \geq 1$, with $M_t = \alpha_0 + \alpha_1 X_{t-1} = E[X_t \mid X_{t-1}]$, we note that

$$\begin{aligned}
q_{0,l} = E\left[\frac{1}{M_t^l}\right] &= \frac{1}{\alpha_0^l}\, E\left[E\left[\left(\frac{\frac{\alpha_0}{\alpha_1}}{\frac{\alpha_0}{\alpha_1} + X_{t-1}}\right)^l \mid X_{t-2}\right]\right] \\
&= \frac{1}{\alpha_1^l}\, E\left[\sum_{n=0}^{+\infty} \frac{(-1)^n\, M_{t-1}^n}{n!} \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1} + j\right)^l}\right].
\end{aligned}$$

Let us now consider that the moment generating function of M_1, $\mathrm{mgf}_{M_1}(u) = E\big[\exp(uM_1)\big]$, is defined for every $u \in (u_1; u_2)$ where $u_1 < 0 < u_2$ such that the radius of convergence satisfies $\min\{-u_1, u_2\} = b > 2$ (also see Lemma 3.1). With these conditions, we have (see Appendix A.2)

$$q_{0,l} = E\left[\frac{1}{M_t^l}\right] = \frac{1}{\alpha_1^l} \sum_{n=0}^{+\infty} \frac{(-1)^n}{n!}\, E[M_{t-1}^n] \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1} + j\right)^l}, \tag{26}$$

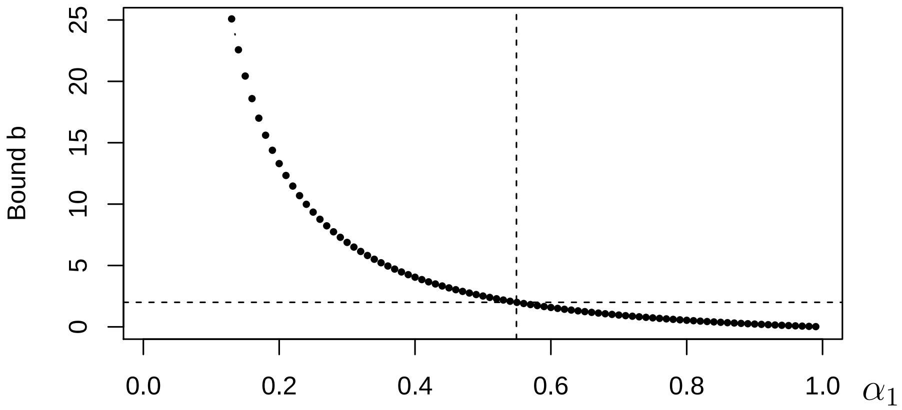

FIGURE 1. Solution b of equation $\ln(b+1)/b = \alpha_1$ against α_1.

that is, the change between the expectation and the infinite sum is allowed. So according to the previous Lemma 3.1, $\alpha_1 < \frac{\ln(b+1)}{u}$ for all $0 < u < b$. Thus, if $\alpha_1 \geq \frac{\ln(b+1)}{u} > \frac{\ln(b+1)}{b}$ with $b > 2$, the equality (26) may not be true. In the Figure 1, we plot b according to the equation $\ln(b+1)/b = \alpha_1$ (lower bound for the radius of convergence) against α_1. This value b decreases with increasing α_1 and falls below 2 for $\alpha_1 = \ln(3)/2 \approx 0.549$; the dashed line refers to the above condition $b > 2$.

In Table 1, we present the values for $q_{0,l}$ with $l = 1, 2, 3, 4$ obtained with the two approaches (25) and (26). In the latter case, the summation in n was stopped if the difference between successive summands felt below 10^{-8}, or if 100 summands were reached. In the left block, the marginal mean is 2.5, in the right, it is 5.0. We note the non-convergence of the approach (26) only for $\alpha_1 > 0.6 > \frac{\ln(3)}{2}$.

4. Simulation Study

To analyze the quality of the approximate distribution (20), (21) of the statistics $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ as well as the finite-sample performance of the corresponding test (size, power), a simulation study has been done with 10 000 replications per scenario. The results shown in Table 2 refer to simulated Poisson INARCH(1) processes (13) (upper half: $\mu = 2.5$; lower half: $\mu = 5.0$). They show mean and standard deviation as computed according to the approximate formulae (20), (21), and compare these values with the corresponding sample counterparts obtained from simulations. The simulated means are

α_0	α_1	l	$q_{0,l}$ by (25)	n. s.	$q_{0,l}$ by (26)	α_0	α_1	l	$q_{0,l}$ by (25)	n. s.	$q_{0,l}$ by (26)
2	0.2	1	0.4064081	11	0.4064081	4	0.2	1	0.2016350	11	0.2016350
		2	0.1676993	11	0.1676993			2	0.0409823	12	0.0409823
		3	0.0702093	12	0.0702093			3	0.0083949	12	0.0083949
		4	0.0298009	12	0.0298009			4	0.0017328	12	0.0017328
1.5	0.4	1	0.4299554	18	0.4299554	3	0.4	1	0.2075920	20	0.2075920
		2	0.1980567	19	0.1980567			2	0.0447126	21	0.0447126
		3	0.0972296	20	0.0972296			3	0.0099853	22	0.0099853
		4	0.0505194	20	0.0505194			4	0.0023098	22	0.0023098
1	0.6	1	0.4973967	48	0.4973967	2	0.6	1	0.2238847	56	0.2238847
		2	0.3046319	52	0.3046320			2	0.0563205	60	0.0563205
		3	0.2212899	55	0.2212899			3	0.0159104	62	0.0159104
		4	0.1815225	57	0.1815225			4	0.0050165	64	0.0050165
0.5	0.8	1	0.8060558	100	$2.247 \cdot 10^{26}$	1	0.8	1	0.2940770	100	$2.091 \cdot 10^{27}$
		2	1.1167550	100	$1.706 \cdot 10^{27}$			2	0.1322086	100	$1.269 \cdot 10^{28}$
		3	1.9693380	100	$7.069 \cdot 10^{27}$			3	0.0845151	100	$4.043 \cdot 10^{28}$
		4	3.7735840	100	$2.155 \cdot 10^{28}$			4	0.0672318	100	$9.050 \cdot 10^{28}$

TABLE 1. Approximations for $q_{0,l}$ with $l = 1, 2, 3, 4$ with approaches (25) and (26), where "n. s." is the number of summands used for (26).

below the theoretical value $C_{1;\,r} = 1$ under the null, but the approximate formula (20) accounts for the negative bias to some degree. The approximation (21) of the standard deviation works rather well especially for the second-order statistic $\widehat{C}_{1;\,2}$; for higher orders $r = 3, 4$, the quality of approximation deteriorates with increasing α.

The most important criterion for the practitioner are the true rejection rates if $\widehat{C}_{1;\,r}(\hat{\alpha}_0, \hat{\alpha}_1)$ is used as a test statistic. So from each simulated time series, upper-sided tests on the nominal level 5 % were designed and executed. The fraction of rejections among the replications was computed for each scenario, which expresses the empirical size under the null of the Poi-INARCH(1) model, and the empirical power otherwise. As the alternative model, the NB-INARCH(1) model by Xu et al. (2012) with different levels of the dispersion parameter $\theta > 1$ was used, see (5) in Example 1.1, i. e., the counts were generated according to the recursive scheme

$$X_t \mid X_{t-1}, X_{t-2}, \ldots \quad \sim \text{NB}\left(\frac{\alpha_0 + \alpha_1\, X_{t-1}}{\theta - 1},\ \frac{1}{\theta}\right) \quad \text{with } \theta > 1. \tag{27}$$

			$E[\widehat{C}_{1;2}(\cdot)]$		$\sqrt{V}[\widehat{C}_{1;2}(\cdot)]$		$E[\widehat{C}_{1;3}(\cdot)]$		$\sqrt{V}[\widehat{C}_{1;3}(\cdot)]$		$E[\widehat{C}_{1;4}(\cdot)]$		$\sqrt{V}[\widehat{C}_{1;4}(\cdot)]^{1/2}$	
α_0	α_1	T	appr	simul	appr	simul	appr	simul	appr	simul	appr	simul	appr	simul
2	0.2	100	0.999	0.992	0.058	0.060	0.999	0.976	0.186	0.184	0.998	0.953	0.444	0.426
		250	1.000	0.997	0.037	0.037	1.000	0.993	0.118	0.115	0.999	0.985	0.280	0.273
		500	1.000	0.999	0.026	0.026	1.000	0.995	0.083	0.082	1.000	0.991	0.198	0.195
		1000	1.000	0.999	0.018	0.018	1.000	0.997	0.059	0.058	1.000	0.995	0.140	0.139
1.5	0.4	100	0.996	0.990	0.064	0.066	0.994	0.981	0.206	0.207	0.992	0.964	0.501	0.504
		250	0.998	0.997	0.041	0.041	0.997	0.992	0.130	0.130	0.997	0.985	0.316	0.313
		500	0.999	0.998	0.029	0.029	0.999	0.995	0.092	0.093	0.998	0.990	0.223	0.225
		1000	1.000	0.999	0.020	0.020	0.999	0.998	0.065	0.066	0.999	0.998	0.158	0.161
1	0.6	100	0.982	0.987	0.088	0.091	0.973	0.972	0.269	0.277	0.964	0.946	0.698	0.685
		250	0.993	0.996	0.056	0.057	0.989	0.994	0.170	0.174	0.986	0.994	0.440	0.464
		500	0.996	0.998	0.039	0.040	0.995	0.994	0.120	0.120	0.993	0.989	0.311	0.310
		1000	0.998	0.999	0.028	0.028	0.997	0.997	0.085	0.086	0.996	0.996	0.220	0.234
0.5	0.8	100	0.876	0.958	0.219	0.180	0.813	0.942	0.616	0.840	0.751	0.926	1.933	3.401
		250	0.951	0.984	0.138	0.136	0.926	0.977	0.388	0.418	0.901	0.967	1.219	1.472
		500	0.975	0.992	0.098	0.098	0.963	0.995	0.274	0.320	0.951	1.010	0.861	1.311
		1000	0.988	0.995	0.069	0.069	0.982	0.994	0.194	0.197	0.975	0.996	0.609	0.643
4	0.2	100	1.000	0.995	0.029	0.030	0.999	0.988	0.089	0.088	0.999	0.977	0.196	0.192
		250	1.000	0.999	0.018	0.019	1.000	0.997	0.056	0.057	1.000	0.994	0.124	0.124
		500	1.000	0.999	0.013	0.013	1.000	0.998	0.040	0.040	1.000	0.996	0.087	0.086
		1000	1.000	1.000	0.009	0.009	1.000	0.998	0.028	0.028	1.000	0.996	0.062	0.062
3	0.4	100	0.998	0.996	0.030	0.032	0.997	0.986	0.094	0.093	0.996	0.973	0.207	0.199
		250	0.999	0.998	0.019	0.020	0.999	0.995	0.059	0.059	0.999	0.990	0.131	0.128
		500	1.000	0.999	0.014	0.014	0.999	0.997	0.042	0.042	0.999	0.995	0.092	0.092
		1000	1.000	1.000	0.010	0.010	1.000	0.999	0.030	0.029	1.000	0.998	0.065	0.064
2	0.6	100	0.993	0.994	0.037	0.039	0.990	0.987	0.108	0.114	0.986	0.977	0.242	0.276
		250	0.997	0.997	0.023	0.023	0.996	0.995	0.068	0.068	0.994	0.991	0.152	0.151
		500	0.999	0.998	0.016	0.016	0.998	0.997	0.048	0.048	0.997	0.995	0.108	0.108
		1000	0.999	0.999	0.012	0.012	0.999	0.998	0.034	0.034	0.999	0.998	0.076	0.076
1	0.8	100	0.960	0.984	0.077	0.070	0.940	0.978	0.191	0.215	0.919	0.971	0.455	0.531
		250	0.984	0.994	0.048	0.047	0.976	0.992	0.120	0.154	0.968	0.997	0.287	0.772
		500	0.992	0.996	0.034	0.034	0.988	0.995	0.085	0.089	0.984	0.996	0.203	0.238
		1000	0.996	0.999	0.024	0.025	0.994	0.998	0.060	0.063	0.992	0.999	0.143	0.169

TABLE 2. Mean and standard deviation of $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$: approximation by (20), (21) vs. simulated values.

The obtained results are summarized in Table 3.

If we look at the size values (highlighted in gray) in Table 3, we see that the empirical size usually agrees quite well with the nominal level 0.05. An exception is the fourth-order statistic for large α and small T, where the empirical size values are visibly smaller than 0.05. So up to now, there is not much difference between the orders $r = 2, 3, 4$ under the null (except for large α). Hence, the crucial question is about the power of these tests

			$\widehat{C}_{1;2}(\hat{\alpha}_0, \hat{\alpha}_1)$; $\theta =$				$\widehat{C}_{1;3}(\hat{\alpha}_0, \hat{\alpha}_1)$; $\theta =$				$\widehat{C}_{1;4}(\hat{\alpha}_0, \hat{\alpha}_1)$; $\theta =$			
α_0	α_1	T	1	1.2	1.4	1.6	1	1.2	1.4	1.6	1	1.2	1.4	1.6
2	0.2	100	0.051	0.354	0.720	0.901	0.051	0.328	0.667	0.874	0.051	0.272	0.561	0.786
		250	0.049	0.636	0.966	0.999	0.053	0.581	0.947	0.997	0.056	0.478	0.878	0.985
		500	0.051	0.874	0.999	1.000	0.052	0.829	0.998	1.000	0.057	0.717	0.988	1.000
		1000	0.049	0.989	1.000	1.000	0.055	0.975	1.000	1.000	0.056	0.927	1.000	1.000
1.5	0.4	100	0.053	0.337	0.691	0.893	0.054	0.305	0.644	0.855	0.049	0.244	0.532	0.755
		250	0.053	0.608	0.956	0.999	0.055	0.561	0.929	0.996	0.053	0.448	0.848	0.979
		500	0.051	0.848	0.999	1.000	0.058	0.805	0.997	1.000	0.060	0.680	0.983	1.000
		1000	0.052	0.984	1.000	1.000	0.060	0.969	1.000	1.000	0.061	0.905	1.000	1.000
1	0.6	100	0.063	0.305	0.617	0.830	0.050	0.266	0.579	0.805	0.036	0.193	0.448	0.668
		250	0.061	0.522	0.910	0.991	0.060	0.487	0.888	0.987	0.047	0.353	0.760	0.942
		500	0.059	0.748	0.993	1.000	0.057	0.722	0.989	1.000	0.048	0.555	0.944	0.997
		1000	0.056	0.944	1.000	1.000	0.060	0.932	1.000	1.000	0.058	0.803	0.999	1.000
0.5	0.8	100	0.054	0.206	0.443	0.642	0.040	0.189	0.412	0.617	0.019	0.111	0.277	0.456
		250	0.056	0.336	0.695	0.891	0.052	0.325	0.696	0.894	0.025	0.186	0.493	0.745
		500	0.057	0.522	0.891	0.985	0.062	0.503	0.896	0.989	0.032	0.294	0.696	0.923
		1000	0.056	0.729	0.990	1.000	0.061	0.735	0.993	1.000	0.042	0.457	0.917	0.996
4	0.2	100	0.049	0.362	0.737	0.918	0.048	0.344	0.707	0.905	0.047	0.307	0.647	0.866
		250	0.053	0.647	0.972	0.999	0.053	0.621	0.964	0.998	0.057	0.557	0.934	0.995
		500	0.050	0.886	1.000	1.000	0.052	0.858	0.999	1.000	0.053	0.801	0.997	1.000
		1000	0.048	0.990	1.000	1.000	0.050	0.984	1.000	1.000	0.054	0.967	1.000	1.000
3	0.4	100	0.059	0.358	0.712	0.914	0.050	0.338	0.693	0.901	0.049	0.296	0.627	0.850
		250	0.056	0.632	0.968	0.999	0.051	0.610	0.960	0.999	0.054	0.539	0.927	0.995
		500	0.053	0.867	0.999	1.000	0.049	0.846	0.999	1.000	0.055	0.788	0.997	1.000
		1000	0.053	0.989	1.000	1.000	0.050	0.981	1.000	1.000	0.053	0.960	1.000	1.000
2	0.6	100	0.067	0.334	0.661	0.870	0.054	0.317	0.657	0.867	0.048	0.269	0.574	0.811
		250	0.056	0.568	0.942	0.996	0.058	0.565	0.942	0.997	0.053	0.488	0.897	0.991
		500	0.051	0.813	0.998	1.000	0.056	0.809	0.998	1.000	0.058	0.730	0.992	1.000
		1000	0.058	0.969	1.000	1.000	0.056	0.973	1.000	1.000	0.057	0.941	1.000	1.000
1	0.8	100	0.065	0.239	0.486	0.708	0.054	0.242	0.520	0.760	0.037	0.187	0.421	0.667
		250	0.062	0.364	0.747	0.930	0.058	0.410	0.826	0.969	0.043	0.318	0.724	0.927
		500	0.056	0.540	0.934	0.996	0.059	0.634	0.974	0.999	0.048	0.508	0.927	0.996
		1000	0.060	0.763	0.995	1.000	0.063	0.863	1.000	1.000	0.055	0.745	0.997	1.000

TABLE 3. Simulated rejection rates for upper-sided test $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$, nominal level 5 %, under H_0: Poi-INARCH(1) model ($\theta = 1$), and H_1: NB-INARCH(1) model ($\theta > 1$).

with respect to the alternative (27). From Table 3, it can be seen that the power values quickly increase with increasing T, and the power is generally better for lower values of the dependence parameter α_1. It can also be seen that the respective power values are larger in the lower half of the table, where we have a larger marginal mean. Comparing the power among the different orders $r = 2, 3, 4$, Table 3 shows a rather clear picture. The fourth-order test is always worse than the second-order test, and with very few exceptions ($\alpha_0 = 1$, $\alpha_1 = 0.8$), the same conclusion also holds between the third- and second-order test. This desirable increase in the rejection rates with increasing θ is caused by increases in both the mean and the standard deviation of $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ (the actual values are omitted in Table 3). Taking these power results together with the described properties under the null, we give a recommendation for using the second-order test $\widehat{C}_{1;2}(\hat{\alpha}_0, \hat{\alpha}_1)$ in practice.

5. Conclusions

The INGARCH models have known, since their introduction by Heinen (2003); Ferland et al. (2006), great extension and development namely through the assumption of new conditional distributions in alternative to the Poisson one, initially considered by those authors. Recently, Gonçalves et al. (2015a) introduced a wide class of this type of models, the CP-INGARCH with compound Poisson conditional distribution, which includes the main INGARCH models present in literature and, particularly, the simple Poisson INGARCH ones.

In order to contribute to the distinction between a simple Poisson INARCH model and a true CP-INARCH, we proposed in this paper a test for such hypotheses based on the form of the probability generating function of the compounding distribution related to the model conditional law. The normality of the test statistics' asymptotic distribution, for the particular case of a INARCH(1) process, was established either in the case, where the model parameters are specified, or when such parameters are consistently estimated. This involves the moments of inverse conditional means of CP-INARCH process, the analysis of their existence and calculation was conducted using two methods. The generalization of the proposed test to higher-order models is an open subject and deserves additional studies.

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Appendix A. Derivations

A.1. Derivation of Formula (21). To obtain the asymptotic variance of the approximate quantity $\widetilde{C}_{1;r}(\hat{\alpha}_0,\hat{\alpha}_1)$ from (19), we start by defining the vectors

$$\boldsymbol{Y}_t^{(r)} \ :=\ \Big(\frac{(X_t)_{(r)}}{(\alpha_0+\alpha_1\,X_{t-1})^r}-1,\ X_t-f_1,\ X_t^2-f_2-f_1^2,\ X_tX_{t-1}-\alpha_1\,f_2-f_1^2\Big)^\top \tag{A.1}$$

with mean $\mathbf{0}$, and by deriving a central limit theorem for $(\boldsymbol{Y}_t^{(r)})_{\mathbb{Z}}$.

A.1 Lemma *Let $(X_t)_{\mathbb{Z}}$ be a stationary INARCH(1) process, define $\boldsymbol{Y}_t^{(r)}$ as in formula (A.1). Denote $f_k := \alpha_0/\prod_{i=1}^k(1-\alpha_1^i)$ such that $\mu=f_1$ and $\sigma^2=f_2$: Then*

$$\begin{aligned}
&\tfrac{1}{\sqrt{T}}\textstyle\sum_{t=1}^T \boldsymbol{Y}_t^{(r)} \xrightarrow{\mathcal{D}} \mathrm{N}\big(\mathbf{0},\boldsymbol{\Sigma}^{(r)}\big) \qquad \textit{with } \boldsymbol{\Sigma}^{(r)}=\big(\sigma_{ij}^{(r)}\big) \textit{ given by}\\
&\sigma_{ij}^{(r)} \ =\ E\big[Y_{0,i}^{(r)}Y_{0,j}^{(r)}\big]+\textstyle\sum_{k=1}^\infty\Big(E\big[Y_{0,i}^{(r)}Y_{k,j}^{(r)}\big]+E\big[Y_{k,i}^{(r)}Y_{0,j}^{(r)}\big]\Big),
\end{aligned}\tag{A.2}$$

where $Y_{k,i}^{(r)}$ denotes the i-th entry of $\boldsymbol{Y}_k^{(r)}$, and where the entries $\sigma_{ij}^{(r)}$ of the symmetric matrix $\boldsymbol{\Sigma}^{(r)}$ are given as follows:

$$\sigma_{11}^{(r)} \ =\ \textstyle\sum_{k=1}^r \binom{r}{k}^2 k!\,q_{0,k} \qquad \textit{(remember (15))},$$

$$\sigma_{12}^{(r)} = \tfrac{r}{1-\alpha_1},\quad \sigma_{13}^{(r)} = \tfrac{2r\,f_1}{1-\alpha_1}+\tfrac{r^2}{1-\alpha_1^2}+\tfrac{r\,\alpha_1}{(1-\alpha_1)(1-\alpha_1^2)},\quad \sigma_{14}^{(r)} = \tfrac{2\,r\,f_1}{1-\alpha_1}+\tfrac{r^2\,\alpha_1}{1-\alpha_1^2}+\tfrac{r\,\alpha_1^2}{(1-\alpha_1)(1-\alpha_1^2)},$$

and

$$\begin{aligned}
\sigma_{22}^{(r)} &= \tfrac{f_1}{(1-\alpha_1)^2},\qquad \sigma_{23}^{(r)} = \tfrac{1+\alpha_1+2\alpha_1^2}{(1-\alpha_1)(1-\alpha_1^2)}\,f_2+\tfrac{2\,f_1^2}{(1-\alpha_1)^2},\qquad \sigma_{24}^{(r)} = \tfrac{\alpha_1(2+\alpha_1+\alpha_1^2)}{(1-\alpha_1)(1-\alpha_1^2)}\,f_2+\tfrac{2\,f_1^2}{(1-\alpha_1)^2},\\
\sigma_{33}^{(r)} &= \tfrac{1+2\alpha_1+8\alpha_1^2+9\alpha_1^3+4\alpha_1^4+6\alpha_1^5}{(1-\alpha_1^2)^2}\,f_3+\tfrac{2(3+4\alpha_1+7\alpha_1^2+4\alpha_1^3)}{1-\alpha_1^2}\,f_2^2+\tfrac{4\,f_1^3}{(1-\alpha_1)^2},\\
\sigma_{34}^{(r)} &= \tfrac{\alpha_1(2+5\alpha_1+8\alpha_1^2+10\alpha_1^3+3\alpha_1^4+2\alpha_1^5)}{(1-\alpha_1^2)^2}\,f_3+\tfrac{2(1+6\alpha_1+6\alpha_1^2+4\alpha_1^3+\alpha_1^4)}{1-\alpha_1^2}\,f_2^2+\tfrac{4\,f_1^3}{(1-\alpha_1)^2},\\
\sigma_{44}^{(r)} &= \tfrac{\alpha_1(1+3\alpha_1+8\alpha_1^2+8\alpha_1^3+8\alpha_1^4+2\alpha_1^5)}{(1-\alpha_1^2)^2}\,f_3+\tfrac{1+8\alpha_1+16\alpha_1^2+8\alpha_1^3+3\alpha_1^4}{1-\alpha_1^2}\,f_2^2+\tfrac{4\,f_1^3}{(1-\alpha_1)^2}.
\end{aligned}$$

Proof: With the same arguments as in Section 2 of Weiß & Schweer (2016), Theorem 1.7 of Ibragimov (1962) is applicable. Furthermore, the expressions for $\sigma_{kl}^{(r)}$ with $k,l\geq 2$ are already known from Theorem 2.2 in Weiß & Schweer (2016), and $\sigma_{11}^{(r)}$ was derived before in the context of formula (11). Hence, to prove Lemma A.1, it remains to compute the entries $\sigma_{12}^{(r)}$, $\sigma_{13}^{(r)}$ and $\sigma_{14}^{(r)}$ of the asymptotic covariance matrix $\boldsymbol{\Sigma}^{(r)}$.

We start with some auxiliary expressions. We have

$$Q_1^{(r)} \ :=\ E\Big[\tfrac{(X_t)_{(r)}\,X_t}{M_t^r}\Big] \ =\ E\Big[\tfrac{E[(X_t)_{(r+1)}+r\,(X_t)_{(r)}\ \mid\ X_{t-1},\ldots]}{M_t^r}\Big] \ =\ E[M_t+r] \ =\ f_1+r. \tag{A.3}$$

Similarly, using that

$$E[M_t^2] \ =\ \alpha_0^2+2\alpha_0\alpha_1\,f_1+\alpha_1^2\,(f_2+f_1^2) \ =\ (\alpha_0+\alpha_1\,f_1)^2+\alpha_1^2\,f_2 \ =\ f_1^2+\alpha_1^2\,f_2,$$

it follows that

$$\begin{aligned} Q_2^{(r)} &:= E\Big[\tfrac{(X_t)_{(r)}\,X_t^2}{M_t^r}\Big] \;=\; E\Big[\tfrac{E[(X_t)_{(r+2)}+(2r+1)\,(X_t)_{(r+1)}+r^2\,(X_t)_{(r)}\ \mid\ X_{t-1},\ldots]}{M_t^r}\Big] \\ &= E[M_t^2+(2r+1)\,M_t+r^2] \\ &= r^2+f_1^2+\alpha_1^2\,f_2+(2r+1)f_1 \;=\; r^2+2r\,f_1+f_2+f_1^2. \end{aligned} \tag{A.4}$$

Finally,

$$\begin{aligned} Q_{1,1}^{(r)} &:= E\Big[\tfrac{(X_t)_{(r)}\,X_tX_{t-1}}{M_t^r}\Big] \;=\; E\Big[\tfrac{X_{t-1}\,E[(X_t)_{(r+1)}+r\,(X_t)_{(r)}\ \mid\ X_{t-1},\ldots]}{M_t^r}\Big] \\ &= E\big[X_{t-1}\,(M_t+r)\big] \;=\; (r+\alpha_0)\,f_1+\alpha_1\,(f_2+f_1^2) \\ &= r\,f_1+\alpha_1\,f_2+f_1\,(\alpha_0+\alpha_1\,f_1) \;=\; r\,f_1+\alpha_1\,f_2+f_1^2. \end{aligned} \tag{A.5}$$

Now we can start with computing $\sigma_{1j}^{(r)}$ for $j=2,3,4$. For $k\geq 1$, we always have

$$E[Y_{k,1}^{(r)}Y_{0,j}^{(r)}] \;=\; E\big[E[Y_{k,1}^{(r)}Y_{0,j}^{(r)}\mid X_{k-1},\ldots]\big] \;=\; E\big[Y_{0,j}^{(r)}\,\underbrace{E[Y_{k,1}^{(r)}\mid X_{k-1},\ldots]}_{=0}\big] \;=\; 0. \tag{A.6}$$

Let us compute $\sigma_{12}^{(r)}$ first. For $k\geq 1$, by conditioning and using that $M_k=\alpha_0+\alpha_1\,X_{k-1}$, we have

$$\begin{aligned} E[Y_{0,1}^{(r)}Y_{k,2}^{(r)}] &= E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_k\big] \;-\; f_1 \;=\; \alpha_1\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_{k-1}\big]+\alpha_0 \;-\; f_1 \\ &= \;\ldots\; = \alpha_1^k\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_0\big]+\alpha_0\,(1+\alpha_1+\ldots+\alpha_1^{k-1}) \;-\; f_1 \\ &= \alpha_1^k\,Q_1^{(r)}+\alpha_0\,\tfrac{1-\alpha_1^k}{1-\alpha_1} \;-\; f_1 \;=\; \alpha_1^k\,(Q_1^{(r)}-f_1) \;\overset{(A.3)}{=}\; \alpha_1^k\,r, \end{aligned}$$

which also holds for $k=0$. Together with (A.6), it follows that

$$\sigma_{12}^{(r)} \;=\; \sum_{k=0}^{\infty} E[Y_{0,1}^{(r)}Y_{k,2}^{(r)}] \;=\; \sum_{k=0}^{\infty} r\,\alpha_1^k \;=\; \frac{r}{1-\alpha_1}.$$

Concerning $\sigma_{13}^{(r)}$, first note that the 2nd non-central moment of the Poisson distribution implies

$$E[X_t^2\mid X_{t-1},\ldots] \;=\; M_t^2+M_t \;=\; \alpha_1^2\,X_{t-1}^2+\alpha_1(2\alpha_0+1)\,X_{t-1}+\alpha_0(\alpha_0+1).$$

Then we compute by successive conditioning that

$$\begin{aligned} E[Y_{0,1}^{(r)}Y_{k,3}^{(r)}] &= \alpha_1^2\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_{k-1}^2\big]+\alpha_1(2\alpha_0+1)\,(r\,\alpha_1^{k-1}+f_1)+\alpha_0(\alpha_0+1) \;-\; f_2-f_1^2 \\ &= \alpha_1^2\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_{k-1}^2\big]+(2\alpha_0+1)\,r\,\alpha_1^k+f_1\,\big(1+f_1(1-\alpha_1^2)\big) \;-\; f_2-f_1^2 \\ &= \;\ldots\; = \alpha_1^{2k}\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_0^2\big]+(2\alpha_0+1)\,r\,\alpha_1^k(1+\alpha_1+\ldots+\alpha_1^{k-1}) \\ &\qquad +f_1\,\big(1+f_1(1-\alpha_1^2)\big)(1+\alpha_1^2+\ldots+\alpha_1^{2(k-1)}) \;-\; f_2-f_1^2 \\ &= \alpha_1^{2k}\,Q_2^{(r)}+(2\alpha_0+1)\,r\,\alpha_1^k\,\tfrac{1-\alpha_1^k}{1-\alpha_1}+(f_2+f_1^2)(1-\alpha_1^2)\,\tfrac{1-\alpha_1^{2k}}{1-\alpha_1^2} \;-\; f_2-f_1^2 \\ &= \alpha_1^{2k}\,\big(Q_2^{(r)}-r\,\tfrac{2\alpha_0+1}{1-\alpha_1}-f_2-f_1^2\big)+r\,\alpha_1^k\,\tfrac{2\alpha_0+1}{1-\alpha_1} \\ &\overset{(A.4)}{=} r\,\alpha_1^{2k}\,(r-\tfrac{1}{1-\alpha_1})+r\,\alpha_1^k\,(2f_1+\tfrac{1}{1-\alpha_1}). \end{aligned}$$

So it follows that

$$\sigma_{13}^{(r)} \;=\; r\,(2f_1+\tfrac{1}{1-\alpha_1})\,\sum_{k=0}^{\infty}\alpha_1^k \;+\; r\,(r-\tfrac{1}{1-\alpha_1})\,\sum_{k=0}^{\infty}\alpha_1^{2k} \;=\; \frac{2r\,f_1}{1-\alpha_1}+\frac{r^2}{1-\alpha_1^2}+\frac{r\,\alpha_1}{(1-\alpha_1)(1-\alpha_1^2)}.$$

Finally, combining the previous derivations, we compute $\sigma_{14}^{(r)}$ as

$$\begin{aligned}
E[Y_{0,1}^{(r)}Y_{k,4}^{(r)}] \;&=\; \alpha_1\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_{k-1}^2\big]+\alpha_0\,E\big[\tfrac{(X_0)_{(r)}}{M_0^r}\,X_{k-1}\big] \;-\; \alpha_1\,f_2-f_1^2\\
&=\; \alpha_1\Big(r\,\alpha_1^{2(k-1)}\,(r-\tfrac{1}{1-\alpha_1})+r\,\alpha_1^{k-1}\,(2f_1+\tfrac{1}{1-\alpha_1}) \;+\; f_2+f_1^2\Big)\\
&\quad +\alpha_0\,(r\,\alpha_1^{k-1}+f_1) \;-\; \alpha_1\,f_2-f_1^2\\
&=\; \tfrac{r}{\alpha_1}\,(r-\tfrac{1}{1-\alpha_1})\,\alpha_1^{2k}+r\,\alpha_1^k\,\big(\tfrac{1}{1-\alpha_1}+f_1+\tfrac{f_1}{\alpha_1}\big)
\end{aligned}$$

for $k\geq 1$, while

$$E[Y_{0,1}^{(r)}Y_{0,4}^{(r)}] \;=\; Q_{1,1}^{(r)}-\alpha_1\,f_2-f_1^2 \;\overset{(A.5)}{=}\; r\,f_1.$$

Therefore,

$$\begin{aligned}
\sigma_{14}^{(r)} \;&=\; r\,\big(\tfrac{1}{1-\alpha_1}+f_1+\tfrac{f_1}{\alpha_1}\big)\,\textstyle\sum_{k=0}^{\infty}\alpha_1^k \;+\; \tfrac{r}{\alpha_1}\big(r-\tfrac{1}{1-\alpha_1}\big)\,\textstyle\sum_{k=0}^{\infty}\alpha_1^{2k}\\
&\quad -\tfrac{r}{\alpha_1}\,\big(r-\tfrac{1}{1-\alpha_1}\big)-r\,\big(\tfrac{1}{1-\alpha_1}+\tfrac{f_1}{\alpha_1}\big)\\
&=\; r\,\big(\tfrac{1}{(1-\alpha_1)^2}+\tfrac{f_1(1+\alpha_1)}{\alpha_1(1-\alpha_1)}\big) \;+\; \tfrac{r}{\alpha_1(1-\alpha_1^2)}\,\big(r-\tfrac{1}{1-\alpha_1}\big) \;-\; \tfrac{r^2}{\alpha_1}+\tfrac{r}{\alpha_1}-\tfrac{r\,f_1}{\alpha_1}\\
&=\; \tfrac{2\,r\,f_1}{1-\alpha_1} \;+\; \tfrac{r^2\,\alpha_1}{1-\alpha_1^2} \;+\; \tfrac{r\,\alpha_1^2}{(1-\alpha_1)(1-\alpha_1^2)}.
\end{aligned}$$

This completes the proof. ■

In the next step, we apply the Delta method to derive the joint distribution of $(\widehat{C}_{1;r},\hat{\alpha}_0,\hat{\alpha}_1)^\top$.

A.2 Corollary *Let $(X_t)_{\mathbb{Z}}$ be a stationary INARCH(1) process. Then the distribution of $(\widehat{C}_{1;r},\hat{\alpha}_0,\hat{\alpha}_1)^\top$ is asymptotically approximated by a normal distribution with mean vector $(1,\alpha_0,\alpha_1)^\top$ and covariance matrix $\frac{1}{T-1}\,\tilde{\boldsymbol{\Sigma}}^{(r)}$, where*

$$\tilde{\boldsymbol{\Sigma}}^{(r)} \;=\; \begin{pmatrix}
\sum_{k=1}^{r}\binom{r}{k}^2 k!\,q_{0,k} & r & 0\\
r & \frac{\alpha_0}{1-\alpha_1}\big(\alpha_0(1+\alpha_1)+\frac{1+2\alpha_1^4}{1+\alpha_1+\alpha_1^2}\big) & -\alpha_0(1+\alpha_1)-\frac{(1+2\alpha_1)\alpha_1^3}{1+\alpha_1+\alpha_1^2}\\
0 & -\alpha_0(1+\alpha_1)-\frac{(1+2\alpha_1)\alpha_1^3}{1+\alpha_1+\alpha_1^2} & (1-\alpha_1^2)\big(1+\frac{\alpha_1(1+2\alpha_1^2)}{\alpha_0(1+\alpha_1+\alpha_1^2)}\big)
\end{pmatrix}.$$

Proof: Define the function $\boldsymbol{g}:\mathbb{R}^4\to\mathbb{R}^3$ by

$$g_1(\boldsymbol{y}) \;:=\; y_1,\quad g_2(\boldsymbol{y}) \;:=\; y_2\,\frac{y_3-y_4}{y_3-y_2^2},\quad g_3(\boldsymbol{y}) \;:=\; \frac{y_4-y_2^2}{y_3-y_2^2}. \tag{A.7}$$

Note that $g_2\big(\cdot,f_1,f_2+f_1^2,\alpha_1\,f_2+f_1^2\big) \;=\; \alpha_0$ and $g_3\big(\cdot,f_1,f_2+f_1^2,\alpha_1\,f_2+f_1^2\big) \;=\; \alpha_1$.

From the proof of Theorem 4.2 in Weiß & Schweer (2016) (see p. 13 in Appendix B.4), we know that the Jacobian of $\boldsymbol{g}$ equals

$$\mathbf{J}_{\boldsymbol{g}}(\boldsymbol{y}) \;=\; \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & \dfrac{(y_3-y_4)\big(y_3+y_2^2\big)}{\big(y_3-y_2^2\big)^2} & \dfrac{y_2\big(y_4-y_2^2\big)}{\big(y_3-y_2^2\big)^2} & \dfrac{-y_2}{y_3-y_2^2} \\ 0 & \dfrac{2y_2(y_4-y_3)}{\big(y_3-y_2^2\big)^2} & \dfrac{y_2^2-y_4}{\big(y_3-y_2^2\big)^2} & \dfrac{1}{y_3-y_2^2} \end{pmatrix},$$

such that $\mathbf{D} := \mathbf{J}_{\boldsymbol{g}}\big(1, f_1, f_2+f_1^2, \alpha_1\, f_2+f_1^2\big)$ is given by

$$\begin{aligned} \mathbf{D} \;&=\; \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & \dfrac{(1-\alpha_1)(f_2+2f_1^2)}{f_2} & \dfrac{\alpha_1\, f_1}{f_2} & -\dfrac{f_1}{f_2} \\ 0 & -\dfrac{2(1-\alpha_1)\, f_1}{f_2} & -\dfrac{\alpha_1}{f_2} & \dfrac{1}{f_2} \end{pmatrix} \\ &=\; \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & (1-\alpha_1)\big(1+2(1-\alpha_1^2)\, f_1\big) & \alpha_1(1-\alpha_1^2) & -(1-\alpha_1^2) \\ 0 & -2(1-\alpha_1)(1-\alpha_1^2) & -\dfrac{\alpha_1}{f_2} & \dfrac{1}{f_2} \end{pmatrix}. \end{aligned}$$

Now, let us look at

$$\tilde{\mathbf{\Sigma}}^{(r)} = \big(\tilde{\sigma}_{ij}^{(r)}\big) \;:=\; \mathbf{D}\mathbf{\Sigma}^{(r)}\mathbf{D}^\top,$$

where $\mathbf{\Sigma}^{(r)}$ is the covariance matrix from Lemma A.1 above. The components $\tilde{\sigma}_{22}^{(r)}, \tilde{\sigma}_{23}^{(r)}, \tilde{\sigma}_{33}^{(r)}$ are already known from formula (11) in Weiß (2010) (or from Theorem 4.2 in Weiß & Schweer (2016)), and $\tilde{\sigma}_{11}^{(r)} = \sigma_{11}^{(r)}$ obviously holds.

So it remains to compute $\tilde{\sigma}_{12}^{(r)} \;=\; \sum_{j=2}^4 d_{11}d_{2j}\, \sigma_{1j}^{(r)}$ and $\tilde{\sigma}_{13}^{(r)} \;=\; \sum_{j=2}^4 d_{11}d_{3j}\, \sigma_{1j}^{(r)}$. We get

$$\begin{aligned} \tilde{\sigma}_{12}^{(r)} \;&=\; (1-\alpha_1)\big(1+2(1-\alpha_1^2)\, f_1\big)\, \sigma_{12}^{(r)} + \alpha_1(1-\alpha_1^2)\, \sigma_{13}^{(r)} - (1-\alpha_1^2)\, \sigma_{14}^{(r)} \\ &=\; r + 2r\,(1-\alpha_1^2)\, f_1 + 2r\, f_1\, \alpha_1(1+\alpha_1) + r^2\, \alpha_1 + \tfrac{r\, \alpha_1^2}{1-\alpha_1} - 2\, r\, f_1\, (1+\alpha_1) - r^2\, \alpha_1 - \tfrac{r\, \alpha_1^2}{1-\alpha_1} \\ &=\; r, \end{aligned}$$

as well as

$$\begin{aligned} \tilde{\sigma}_{13}^{(r)} \;&=\; -2(1-\alpha_1)(1-\alpha_1^2)\, \sigma_{12}^{(r)} - \tfrac{\alpha_1}{f_2}\, \sigma_{13}^{(r)} + \tfrac{1}{f_2}\, \sigma_{14}^{(r)} \\ &=\; -2r\,(1-\alpha_1^2) - 2r\, \alpha_1(1+\alpha_1) - \tfrac{r^2\, \alpha_1}{f_1} - \tfrac{r\, \alpha_1^2}{f_1\,(1-\alpha_1)} + 2r\,(1+\alpha_1) + \tfrac{r^2\, \alpha_1}{f_1} + \tfrac{r\, \alpha_1^2}{f_1\,(1-\alpha_1)} \\ &=\; 0. \end{aligned}$$

This completes the proof. ∎

Using Corollary A.2, we are able to approximate the variance of $\widehat{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ by the asymptotic variance $\frac{1}{T-1}\,\sigma_{1;r}^2$ of $\widetilde{C}_{1;r}(\hat{\alpha}_0, \hat{\alpha}_1)$ according to (19):

$$\begin{aligned}
\sigma_{1;r}^2 &= \tilde{\sigma}_{11}^{(r)} + r^2\, q_{0,1}^2\, \tilde{\sigma}_{22}^{(r)} + r^2\, q_{1,1}^2\, \tilde{\sigma}_{33}^{(r)} - 2r\, q_{0,1}\, \tilde{\sigma}_{12}^{(r)} + 2r^2\, q_{0,1}q_{1,1}\tilde{\sigma}_{23}^{(r)} \\
&= \textstyle\sum_{k=1}^{r} \binom{r}{k}^2 k!\, q_{0,k} \;-\; 2r^2\, q_{0,1} \;+\; r^2\, q_{0,1}^2\, \frac{\alpha_0}{1-\alpha_1}\big(\alpha_0(1+\alpha_1) + \frac{1+2\alpha_1^4}{1+\alpha_1+\alpha_1^2}\big) \\
&\quad + r^2\, q_{1,1}^2\,(1-\alpha_1^2)\big(1 + \frac{\alpha_1(1+2\alpha_1^2)}{\alpha_0(1+\alpha_1+\alpha_1^2)}\big) \;-\; 2r^2\, q_{0,1}q_{1,1}\,\big(\alpha_0(1+\alpha_1) + \frac{(1+2\alpha_1)\alpha_1^3}{1+\alpha_1+\alpha_1^2}\big).
\end{aligned}$$

So the proof of formula (21) is complete.

A.2. Derivation of Equality (26). First, we note that if the random variable Z follows a Poisson distribution with mean λ, and if $a > 0$, we have for $k = 1, 2, \ldots$

$$\begin{aligned}
E\left[\left(\frac{a}{a+Z}\right)^k\right] &= \int_0^1 \exp\big(-\lambda\,(1-s)\big)\,\frac{a^k}{(k-1)!}\, s^{a-1}\, \log^{k-1}\left(\frac{1}{s}\right)\, ds \\
&= \frac{a^k}{(k-1)!}\sum_{n=0}^{+\infty}\frac{(-1)^n\,\lambda^n}{n!}\int_0^1 (1-s)^n\, s^{a-1}\,\log^{k-1}\left(\frac{1}{s}\right)\, ds \\
&= a^k\sum_{n=0}^{+\infty}\frac{(-1)^n\,\lambda^n}{n!}\sum_{j=0}^{n}\binom{n}{j}\frac{(-1)^j}{(a+j)^k},
\end{aligned}$$

using the Dominated Convergence Theorem and the following result (formula 16 on page 552 of Gradshteyn & Ryzhic (2007))

$$\int_0^1 \left(\log\frac{1}{x}\right)^n (1-x^q)^m\, x^{p-1}\, dx \;=\; n!\sum_{k=0}^{m}\binom{m}{k}\frac{(-1)^k}{(p+kq)^{n+1}} \qquad \text{with } p, q > 0.$$

We note that for $k = 1$, the expression may be replaced by the equivalent one

$$E\left[\frac{a}{a+Z}\right] \;=\; \Gamma(a+1)\sum_{n=0}^{+\infty}\frac{(-1)^n}{\Gamma(a+n+1)}\,\lambda^n,$$

since

$$\frac{\Gamma(a+1)}{\Gamma(a+n+1)} \;=\; \frac{a}{n!}\sum_{j=0}^{n}\binom{n}{j}\frac{(-1)^j}{a+j}$$

as may be proved by recurrence.

Let us now consider that the moment generating function of M_1, $\mathrm{mgf}_{M_1}(u) = E\left[\exp(uM_1)\right]$, is defined for every $u \in (u_1; u_2)$, where $u_1 < 0 < u_2$ such that $\min\{-u_1, u_2\} = b > 2$. With these conditions, we will prove that

$$E\left[\frac{1}{M_t^l}\right] \;=\; \frac{1}{\alpha_1^l}\sum_{n=0}^{+\infty}\frac{(-1)^n}{n!}\,E[M_{t-1}^n]\sum_{j=0}^{n}\binom{n}{j}\frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l},$$

that is, the change between the expectation and the infinite sum is allowed. For this purpose, let us consider s such that $0 < s < \frac{1}{2}\min\{-u_1, u_2\}$ and the function

$$H(t) = \int \sum_{n=0}^{+\infty} \frac{(-1)^n (tx)^n}{n!} \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l} dP_{M_1}(x), \qquad t \in (-s; s).$$

Considering the functions

$$h_k(x) = \sum_{n=0}^{k} \frac{(-1)^n (tx)^n}{n!} \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l} \qquad \text{with } k \in \mathbb{N}_0,$$

and $h(x) := h_\infty(x)$, we have for every x and for $k = 1, 2, \ldots$

$$|h_k(x)| \le \sum_{n=0}^{k} \frac{|tx|^n}{n!} \sum_{j=0}^{n} \binom{n}{j} \frac{1}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l} \le \left(\frac{\alpha_1}{\alpha_0}\right)^l \sum_{n=0}^{k} \frac{(2|tx|)^n}{n!} \le \left(\frac{\alpha_1}{\alpha_0}\right)^l \exp(2s|x|),$$

since $|t| < s$, and also $\lim_{k\to\infty} h_k(x) = h(x)$. Moreover,

$$\begin{aligned}\int \exp(2s|x|)\, dP_{M_1}(x) &\le \int_{-\infty}^{+\infty} \exp(2sx)\ dP_{M_1}(x) + \int_{-\infty}^{+\infty} \exp(-2sx)\ dP_{M_1}(x) \\ &= \mathrm{mgf}_{M_1}(2s) + \mathrm{mgf}_{M_1}(-2s) < +\infty.\end{aligned}$$

So, we may apply the Dominated Convergence Theorem and we obtain

$$H(t) = \int h(x)\, dP_{M_1}(x) = \lim_{k\to\infty} \sum_{n=0}^{k} \frac{(-1)^n\ t^n}{n!} \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l} \int x^n dP_{M_1}(x),$$

that is,

$$E\left[\sum_{n=0}^{+\infty} \frac{(-1)^n\ (tM_{t-1})^n}{n!} \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l}\right] = \sum_{n=0}^{+\infty} \frac{(-1)^n}{n!} E[t^n M_{t-1}^n] \sum_{j=0}^{n} \binom{n}{j} \frac{(-1)^j}{\left(\frac{\alpha_0}{\alpha_1}+j\right)^l},$$

for $t \in [-s; s]$. The result is valid for $t = 1$ if and only $s > 1$, which is possible as $\min\{-u_1, u_2\} > 2$, and so (26) follows.

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