Let (X, Y) be a bivariate random vector with joint distribution function $H(x, y)$ and continuous marginal distribution function $F(x)$ and $G(y)$. Its associated copula C is defined by $H(x, y) = C\{F(x), G(y)\}$. Since F and G are continuous, the copula C is unique and we can write $C(u, v) = H(F^{\leftarrow}(u), G^{\leftarrow}(v))$ for $0 \leq u, v \leq 1$ and where $F^{\leftarrow}(u) = inf\{v \in \mathbb{R} | F(v) \geq u\}$ and $G^{\leftarrow}(u) = inf\{v \in \mathbb{R} | G(v) \geq u\}$ are the generalized inverse functions of F and G respectively. At each $t \in \{1, \dots, T\}$, we suppose that one of both entries may be missing. The observations consist of a sample of independent, identically distributed quadruples

$$(I_t, J_t, I_t X_t, J_t Y_t) \quad t \in 1, \dots, T$$

The indicator variables I_t (respectively J_t) is equal to 1 or 0 according to wheter X_t or Y_t is observed or not. We suppose that the indicator functions I_t and J_t are independent. The probability of observing a realisation partially or completely is denoted by $p_X = \mathbb{P}(I_t = 1) > 0$ and $p_Y = \mathbb{P}(J_t = 1) > 0$.

We define :

$$C(u, v) = H(F^{\leftarrow}(u), G^{\leftarrow}(v)) = \varphi(H)(u, v)$$

and,

$$Z_T(u, v) = \sqrt{T}\{\hat{C}_T(u, v) - C(u, v)\} \tag{1}$$

where $\hat{H}_T$ corresponds to the empirical distribution function of the sample $(X_1, Y_1), \dots, (X_T, Y_T)$

$$\hat{H}_T(u, v) = \frac{\sum_{t=1}^{T} 1_{\{X_t \leq u, Y_t \leq v\}} I_t J_t}{\sum_{t=1}^{T} I_t J_t}$$

We define also the corresponding empirical distribution functions in the case of missing data :

$$\hat{F}_T(u) = \frac{\sum_{t=1}^{T} 1_{\{X_t \leq u\}} I_t}{\sum_{t=1}^{T} I_t}$$

$$\hat{G}_T(v) = \frac{\sum_{t=1}^{T} 1_{\{Y_t \leq v\}} J_t}{\sum_{t=1}^{T} J_t}$$

Condition 1. *We suppose for all $t \in \{1, \dots, T\}$, the pairs (I_t, J_t) and (X_t, Y_t) are independent, the data are missing completely at random. Furthermore, we suppose that there exist at least one $t \in \{1, \dots, T\}$ such that $I_t J_t \neq 0$.*

Proposition 1. *Under hypothesis 1, $\hat{H}_T, \hat{F}_T, \hat{G}_T$ are consistant estimators of H, F, G.*

Démonstration. We check the consistency for $\hat{H}_T$. By independence, we have

$$\mathbb{E}[T^{-1} \sum_{t=1}^{T} I_t J_t] = T^{-1} \sum_{t=1}^{T} \mathbb{E}[I_t]\mathbb{E}[J_t] = p_X p_Y$$

So, by applying the law of large numbers, we have :

$$T^{-1} \sum_{t=1}^{T} I_t J_t \longrightarrow p_X p_Y \quad a.s. \quad as \quad T \to \infty$$

Then, we now use the first hypothesis to get :

$$T^{-1} \sum_{t=1}^{T} \mathbb{E}[1_{\{X_t \leq u, Y_t \leq v\}} I_t J_t] = T^{-1} \sum_{t=1}^{T} \mathbb{E}[1_{\{X_t \leq u, Y_t \leq v\}}]\mathbb{E}[I_t J_t] = H(x, y) p_X p_Y$$

By applying again the law of large numbers, we derive :

$$\sum_{t=1}^{T} 1_{\{X_t \leq u, Y_t \leq v\}} I_t J_t \longrightarrow H(x, y) p_X p_Y \quad a.s. \quad as \quad T \to \infty$$

We can now apply the continuous mapping theorem to the function $f : (x, y) \mapsto \frac{x}{y}$ which are continuous on $\mathbb{R}_+ \times \mathbb{R}_+ \setminus 0$ to conclude that :

$$\hat{H}_T(x, y) \longrightarrow H(x, y) \quad a.s. \quad as \quad T \to \infty$$

□

Condition 2. *1. The bivariate distribution function H has continuous margins F, G and copula C.*

2. The first order partial derivatives $\dot{C}_1(u,v) = \frac{\partial C}{\partial u}(u,v)$ and $\dot{C}_2(u,v) = \frac{\partial C}{\partial v}(u,v)$ exists and is continuous on the set $\{(u,v) \in [0,1]^2, 0 < u, v < 1\}$

Condition 3. *There exists $\gamma_t > 0$ and $r_t > 0$ such that $r_t \longrightarrow \infty$ as $t \to \infty$ such that in the space $l^\infty(\mathbb{R}^2) \otimes (l^\infty(\mathbb{R}), l^\infty(\mathbb{R}))$ equipped with the topology of uniform convergence, we have the joint weak convergence*

$$(r_t(\hat{H}_T - H); r_t(\hat{F}_t - F), r_t(\hat{G}_t - G)) \rightsquigarrow (\alpha \circ (F,G), \beta_1 \circ F, \beta_2 \circ G)$$

The stochastic processes α and β_j take values in $l^\infty([0,1]^2)$ and $l^\infty([0,1])$ respectively, and are such that $\alpha \circ F$ and $\beta_j \circ F_j$ have continuous trajectories on $[-\infty, \infty]^2$ and $[-\infty, \infty]$ almost surely.

Theorem 1 (Theorem 2.3 and example 3.5 in [Seg14])**.** *If conditions 2.1 and 2.2 holds, then uniformly in $u \in [0,1]^2$,*

$$\begin{aligned} r_T\{\hat{C}_T(u,v) - C(u,v)\} &= r_T\{\hat{H}_T((F,G)^{\leftarrow}(u,v) - C(u,v)\} && (2)\\ &- \dot{C}_1(u,v) r_T\{\hat{F}_T(F^{\leftarrow}(u)) - u\} 1_{(0,1)}(u) && (3)\\ &- \dot{C}_2(u,v) r_T\{\hat{G}_T(G^{\leftarrow}(v)) - v\} 1_{(0,1)}(v) + o_{\mathbb{P}}(1) && (4)\end{aligned}$$

as $T \to \infty$. Hence in $l^\infty([0,1]^2)$ equipped with the supremum norm, as $T \to \infty$,

$$(r_T\{\hat{C}_T(u,v) - C(u,v)\})_{u,v \in [0,1]^2} \rightsquigarrow (\alpha(u,v) - \dot{C}_1(u,v)\beta_1(u) - \dot{C}_2(u,v)\beta_2(v))_{u,v \in [0,1]^2}$$

We denote by $S_C(u,v)$ the process defined on the right-hand side in the weak convergence from above.

Remark 1. *If we consider the empiric copula, theorem 1 gives us the weak convergence of this process to a brownian bridge $N_C(u,v)$ defined by, $\forall (u,v) \in [0,1]$*

$$N_C(u,v) = B_C(u,v) - \dot{C}_1(u,v) B_C(u,1) - \dot{C}_2(u,v) B_C(1,v) \tag{5}$$

where B_C is a brownian bridge in $[0,1]^2$ with covariance function

$$\mathbb{E}[B_C(u,v) B_C(u',v')] = C(u \wedge u', v \wedge v') - C(u,v)C(u',v')$$

for each $0 \leq, u, v, u', v' \leq 1$. This result is well known since 2004 due to [FRW04].

In our specific case with missing data, (ref) shows that $r_T\{\hat{C}_T(u,v) - C(u,v)\}$ is weakly convergent toward $\alpha(u,v) - \dot{C}_1(u,v)\beta_1(u) - \dot{C}_2(u,v)\beta_2(v)$ where $\beta_1(u) = p_X^{-1}\mathbb{G}(1_{X \leq F^{\leftarrow}(u)} - u 1_{I=1})$, $\beta_2(v) = p_Y^{-1}\mathbb{G}(1_{Y \leq G^{\leftarrow}(v)} - v 1_{J=1})$ and $\alpha(u,v) = (p_X p_Y)^{-1}\mathbb{G}(1_{X \leq F^{\leftarrow}(u)} 1_{Y \leq G^{\leftarrow}(v)} - C(u,v) 1_{I=1} 1_{J=1})$. Furthermore, by these expressions, we can detail the structure of the covariance matrix between the three processes.

Definition 1. *Let $(X_1, Y_1), \dots, (X_T, Y_T)$ a T bivariate random vectors with unknown margins F and G. A λ-FMadogram is the quantity defined by :*

$$\nu(\lambda) = \frac{1}{2}\mathbb{E}[|F^\lambda(X) - G^{1-\lambda}(Y)|] \tag{6}$$

We estimative the λ-FMadogram with the following quantity :

$$\hat{\nu}_T(\lambda) = \frac{1}{2\sum_{t=1}^T I_t J_t} \sum_{t=1}^T |\hat{F}_T^\lambda(X_t) - \hat{G}_T^{1-\lambda}(Y_t)| I_t J_t \tag{7}$$

Proposition 2 (Proposition 3 of [NGCD09])**.** *Suppose that conditions 2 holds and that $\sum_{t=1}^T I_t J_t = T$ (no missing data). Let J be a function of bounded variation, continuous from above and with discontinuities of the first kind. Then*

$$T^{-1/2} \sum_{t=1}^T (J(\hat{F}_T(X_t), \hat{G}_T(Y_t)) - \mathbb{E}[J(F(X), G(Y))])$$

converges in distribution to $\int_{[0,1]} N_C(u,v) dJ(u,v)$ where $N_C(u,v)$ is defined by equation (5) and the integral is well defined as a Lebesgue-Stieltjes integral (see section 5.2 of [Bre20]). The special case, $J(x,y) = \frac{1}{2}|x^\lambda - y^{1-\lambda}|$ provide the weak of convergence of the λ-Madogram estimator defined by (7) :

$$T^{1/2}\{\hat{\nu}_T(\lambda) - \frac{1}{2}\mathbb{E}[|F^\lambda(X) - G^{1-\lambda}(Y)|]\}$$

converge in distribution to $\int_{[0,1]^2} N_C(u,v)dJ(u,v)$ *where the latter integral satisfies :*

$$\int_{[0,1]^2} f(x,y)dJ(x,y) = \frac{1}{2}\int_0^1 f(0, y^{1/(1-\lambda)})dy + \frac{1}{2}\int_0^1 f(x^{1/\lambda}, 0)dx - \int_0^1 f(x^{1/\lambda}, x^{1/(1-\lambda)})dx \quad (8)$$

for all bounded-measurable function $f : [0,1]^2 \mapsto \mathbb{R}$.

We add some elements in order to prove the identity (8).

Lemma 1. *For all bounded-measurable function* $f : [0,1]^2 \mapsto \mathbb{R}$*, if* $J(s,t) = |s^\lambda - t^{1-\lambda}|$*, then the following integral satisfies :*

$$\int_{[0,1]^2} f(x,y)dJ(x,y) = \frac{1}{2}\int_0^1 f(0, y^{1/(1-\lambda)})dy + \frac{1}{2}\int_0^1 f(x^{1/\lambda}, 0)dx - \int_0^1 f(x^{1/\lambda}, x^{1/(1-\lambda)})dx$$

Démonstration. Let A a element of $\mathcal{B}([0,1]^2)$. We can pick an element of the form $A = [0,s] \times [0,t]$, where $s,t \in [0,1]$ and $\lambda \in [0,1]$. Let us introduce the following indicator function :

$$f_{s,t}(x,y) = 1_{\{(x,y)\in[0,1]^2, 0\leq x\leq s, 0\leq y\leq t\}}$$

Then, for this function, we have in one hand :

$$\int_{[0,1]^2} f_{s,t}(x,y)dJ(x,y) = J(s,t) - J(0,0) = |s^\lambda - t^{1-\lambda}|$$

in other hand, using the equality $\frac{|x-y|}{2} = \frac{x}{2} + \frac{y}{2} - min(x,y)$, one has to show

$$\begin{aligned}\frac{1}{2}|s^\lambda - t^{1-\lambda}| &= \frac{s^\lambda}{2} + \frac{y^{1-\lambda}}{2} - min(s^\lambda, t^{1-\lambda}) \\ &= \int_0^1 f_{s,t}(x^{\frac{1}{\lambda}}, 0)dx + \int_0^1 f_{s,t}(0, y^{\frac{1}{1-\lambda}})dy - \int_0^1 f_{s,t}(x^{\frac{1}{\lambda}}, x^{\frac{1}{1-\lambda}})dx\end{aligned}$$

Notice that the class :

$$\mathcal{E} = \{A \in \mathcal{B}([0,1]^2) : \int_{[0,1]^2} 1_A(x,y)dJ(x,y) = \int_0^1 1_A(x^{\frac{1}{\lambda}}, 0)dx + \int_0^1 1_A(0, y^{\frac{1}{1-\lambda}})dy - \int_0^1 1_A(x^{\frac{1}{\lambda}}, x^{\frac{1}{1-\lambda}})dx\}$$

contain the class $\mathcal{P}$ of all closed pavements of $[0,1]^2$. It is otherwise a monotone class (or $\lambda - system$). Hence as the class $\mathcal{P}$ of closed pavement is a $\pi - system$, the class monotone theorem ensure that $\mathcal{E}$ contains the sigma-field generated by $\mathcal{P}$, that is $\mathcal{B}([0,1]^2)$.

This result holds for simple $f(x,y) = \sum_{i=1}^n \lambda_i 1_{A_i}$ where $\lambda_i \in \mathbb{R}$ and $A_i \in \mathcal{B}([0,1]^2)$ for all $i \in \{1,\ldots,n\}$. We then can prove the identity for positive measurable function by approximation with an increasing sequence of simple function using Beppo-Levy theorem. This identity is then extended to measurable function $f : [0,1]^2 \mapsto \mathbb{R}$ considering $f = f_+ - f_-$ with $f_+ = max(f,0)$ and $f_- = min(-f,0)$. We take the function bounded-measurable in order that the left hand size of the equality is well defined as a Lebesgue-Stieljes integral. □

Furthermore, as the limiting process is the linear transformation of a tight gaussian process, we know from [vdVW96] that it is Gaussian. Before going further, we want to detail the structure of the variance of the limiting process. Doing that, we introduce the following lemma :

Lemma 2. *Let* $(B_C(u,v))_{u,v\in[0,1]^2}$ *a brownian bridge with covariance function defined by :*

$$\mathbb{E}[B_C(u,v)B_C(u',v')] = C(u\wedge u', v\wedge v') - C(u,v)C(u',v')$$

for each $0 \leq, u, v, u', v' \leq 1$*. Let* $a,b \in [0,1]$ *fixed, if* $a = 0$ *or* $b = 0$*, then wet get the following equality :*

$$\mathbb{E}[\int_0^1 B_C(u,a)du \int_0^1 B_C(b,u)du] = 0$$

Démonstration. Without loss of generality, suppose that $a = 0$ and $b \in [0,1]$. Using the linearity of the integral, we obtain :

$$\begin{aligned}\mathbb{E}[\int_0^1 B_C(u,0)du \int_0^1 B_C(b,u)du] &= \mathbb{E}[\int_0^1\int_0^1 B_C(u,0)B_C(b,v)dudv] \\ &= \int_0^1\int_0^1 \mathbb{E}[B_C(u,0)B_C(b,v)]dudv\end{aligned}$$

We then use the definition of the covariance function of our Brownian bridge, we have

$$\mathbb{E}[B_C(u,0)B_C(b,v)] = C(u\wedge v,0) - C(u,0)C(b,v)$$

We recall that, by definition, a copula satisfy $C(u,0) = C(0,u) = 0$ for every $u \in [0,1]$. Then, the equation below is equal to 0. Our conclusion directly follows. □

Using this lemma, we can infer the following proposition :

Proposition 3. *Let $N_C(u,v)$ the process defined in equation (5) and $a, b \in [0,1]$ fixed. If $a = 0$ or $b = 0$, then :*

$$\mathbb{E}[\int_0^1 N_C(u,a)du \int_0^1 N_C(b,u)du] = 0$$

With this proposition, we can infer a better form of the variance of our limiting process :

Theorem 2. *Let $N_C(u,v)$ the process defined in equation (5) and $J(x,y) = |x^\lambda - y^{1-\lambda}|$, then :*

$$Var(\int_{[0,1]^2} N_C(u,v)dJ(u,v)) = Var(\int_0^1 N_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})du) \tag{9}$$

Démonstration. Recall that, with J defined as in the statement that :

$$\int_{[0,1]^2} N_C(u,v)dJ(u,v) = \frac{1}{2}\int_0^1 N_C(0,v^{1/(1-\lambda)})dv + \frac{1}{2}\int_0^1 N_C(u^{1/\lambda},0)du - \int_0^1 N_C(u^{1/\lambda},u^{1/(1-\lambda)})du$$

Taking the variance and using the proposition 3 gives that only the variance of the third term matters. □

We are now able to detail precisely the variance of the limiting process with a given Copula. This is the purpose of the following proposition :

Proposition 4. *Under the framework of theorem 2 and if we take $C(u,v) = uv$, the independent copula, then the variance of the lambda FMadogram has the following form*

$$Var(\int_{[0,1]^2} N_C(u,v)dJ(u,v)) = \left(\frac{\lambda(1-\lambda)}{1+\lambda(1-\lambda)}\right)^2\left(\frac{1}{1+2\lambda(1-\lambda)} - \frac{1-\lambda}{1+\lambda+2\lambda(1-\lambda)} - \frac{\lambda}{2-\lambda+2\lambda(1-\lambda)}\right)$$

with $\lambda \in [0,1]$

Démonstration. With direct computing and using the same techniques used is lemma 1, we obtain that :

$$\begin{aligned}
Var\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})du\right] &= \left(\frac{\lambda(1-\lambda)}{1+\lambda(1-\lambda)}\right)^2\left(\frac{1}{1+2\lambda(1-\lambda)}\right)\\
Var\left[\int_0^1 u^{\frac{1}{1-\lambda}} B_C(u^{\frac{1}{\lambda}},1)du\right] &= \left(\frac{\lambda(1-\lambda)}{1+\lambda(1-\lambda)}\right)^2\left(\frac{1-\lambda}{1+\lambda+2\lambda(1-\lambda)}\right)\\
Var\left[\int_0^1 u^{\frac{1}{\lambda}} B_C(1,u^{\frac{1}{1-\lambda}})du\right] &= \left(\frac{\lambda(1-\lambda)}{1+\lambda(1-\lambda)}\right)^2\left(\frac{\lambda}{2-\lambda+2\lambda(1-\lambda)}\right)\\
cov\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})du, \int_0^1 u^{\frac{1}{1-\lambda}} B_C(u^{\frac{1}{\lambda}},1)du\right] &= \left(\frac{\lambda(1-\lambda)}{1+\lambda(1-\lambda)}\right)^2\left(\frac{1-\lambda}{1+\lambda+2\lambda(1-\lambda)}\right)\\
cov\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}), du \int_0^1 u^{\frac{1}{\lambda}} B_C(1,u^{\frac{1}{1-\lambda}})du\right] &= \left(\frac{\lambda(1-\lambda)}{1+\lambda(1-\lambda)}\right)^2\left(\frac{\lambda}{2-\lambda+2\lambda(1-\lambda)}\right)\\
cov\left[\int_0^1 u^{\frac{1}{1-\lambda}} B_C(u^{\frac{1}{\lambda}},1)du, \int_0^1 u^{\frac{1}{\lambda}} B_C(1,u^{\frac{1}{1-\lambda}})du\right] &= 0
\end{aligned}$$

Using the identity $Var(X-Y) = Var(X) + Var(Y) - 2cov(X,Y)$ gives the desired result. □

Combining theorem 1 and proposition 2 gives the following result :

Proposition 5. *Suppose that the assumption of theorem 1 holds. Let J be a function of bounded variation, continuous from above and with discontinuities of the first kind. Then*

$$T^{1/2}\left(\frac{\sum_{t=1}^{T} J(\hat{F}_T(X_t), \hat{G}_T(Y_t)) I_t J_t}{\sum_{t=1}^{T} I_t J_t} - \mathbb{E}[J(F(X), G(Y))]\right)$$

converges in distribution to $\int_{[0,1]^2} S_C(u,v) dJ(u,v)$ where $S_C(u,v)$ is defined in theorem 1. The special case, $J(x,y) = \frac{1}{2}|x^\lambda - y^{1-\lambda}|$ provide the weak of convergence of the λ-Madogram estimator defined by (7) :

$$T^{1/2}\left(\hat{\nu}_T(\lambda) - \frac{1}{2}\mathbb{E}[|F^\lambda(X) - G^{1-\lambda}(Y)|]\right)$$

converge in distribution to $\int_{[0,1]^2} S_C(u,v) dJ(u,v)$ where the latter integral satisfies :

$$\int_{[0,1]^2} f(x,y) dJ(x,y) = \frac{1}{2}\int_0^1 f(0, y^{1/(1-\lambda)}) dy + \frac{1}{2}\int_0^1 f(x^{1/\lambda}, 0) dx - \int_0^1 f(x^{1/\lambda}, x^{1/(1-\lambda)}) dx$$

for all bounded functions f.

We now consider the bivariate extreme value copula which can be written in the following form (See Segers extreme value copulas)

$$C(u,v) = (uv)^{A(log(v)/log(uv))} \tag{10}$$

for all $u, v \in [0,1]$ and where $A(\cdot)$ is the pickhands dependence function. With this copula, we want to compute the following integral :

$$\begin{aligned} Var(\int_{[0,1]} N_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) du) = Var(&\int_0^1 B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) du + \int_0^1 B_C(u^{\frac{1}{\lambda}}, 1) \frac{\partial C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})}{\partial u} du \\ &+ \int_0^1 B_C(1, v^{\frac{1}{1-\lambda}}) \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial v}) \end{aligned}$$

We have for the following

$$Var\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) du\right] = \left(\frac{\lambda(1-\lambda)}{A(1-\lambda) + \lambda(1-\lambda)}\right)^2 \left(\frac{A(1-\lambda)}{A(1-\lambda) + 2\lambda(1-\lambda)}\right)$$

$$Var\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, 1) \frac{\partial C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})}{\partial u} du\right] = \left(\frac{\lambda(1-\lambda)}{A(1-\lambda) + \lambda(1-\lambda)}\right)^2 \left(\frac{\kappa^2(1-\lambda)}{2A(1-\lambda) - (1-\lambda) + 2\lambda(1-\lambda)}\right)$$

$$Var\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, 1) \frac{\partial C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})}{\partial u} du\right] = \left(\frac{\lambda(1-\lambda)}{A(1-\lambda) + \lambda(1-\lambda)}\right)^2 \left(\frac{\zeta^2 \lambda}{2A(1-\lambda) - \lambda + 2\lambda(1-\lambda)}\right)$$

Where $\kappa := A(1-\lambda) + A'(1-\lambda)(1-\lambda)$ and $\zeta = A(1-\lambda) - A'(1-\lambda)(1-\lambda)$. We now compute the covariance :

$$\begin{aligned} &cov\left[\int_0^1 B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) du, \int_0^1 B_C(u^{\frac{1}{\lambda}}, 1) \frac{\partial C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})}{\partial u} du\right] = \int_0^1 \int_0^1 \mathbb{E}[B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) B_C(v^{\frac{1}{\lambda}}, 1)] \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial u} dudv \\ &= \int_0^1 \int_0^u \mathbb{E}[B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) B_C(u^{\frac{1}{\lambda}}, 1)] \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial u} dudv + \int_0^1 \int_v^1 \mathbb{E}[B_C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) B_C(u^{\frac{1}{\lambda}}, 1)] \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial u} dudv \end{aligned}$$

for the first one, we have :

$$\int_0^1 \int_0^u (C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) - C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) v^{\frac{1}{\lambda}}) \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial u} dudv = \frac{\kappa}{2}\left(\frac{\lambda(1-\lambda)}{A(1-\lambda) + \lambda(1-\lambda)}\right)^2 \left(\frac{\lambda}{2A(1-\lambda) + \lambda(1-\lambda}\right)$$

For the second part, using Fubini, we have :

$$\int_0^1 \int_0^u (C(v^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) - C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})v^{\frac{1}{\lambda}}) \frac{\partial C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})}{\partial u} dv du$$

for the right hand side of the "minus" sign, we may compute :

$$\int_0^1 \int_0^u C(u^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}})v^{\frac{1}{\lambda}} \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial u} dv du = \frac{\kappa}{2} \left(\frac{\lambda(1-\lambda)}{A(1-\lambda) + \lambda(1-\lambda)} \right)^2$$

The last one still difficult to handle

$$\int_0^1 \int_0^u C(v^{\frac{1}{\lambda}}, u^{\frac{1}{1-\lambda}}) \frac{\partial C(v^{\frac{1}{\lambda}}, v^{\frac{1}{1-\lambda}})}{\partial u} dv du \tag{11}$$

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