Latent variable analysis indicates that seasonal anisotropy accounts for the higher prevalence of left-handedness in men

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ABSTRACT

According to the Geschwind Galaburda theory of cerebral lateralization, high intrauterine testosterone levels delay left brain hemisphere maturation and thus promote left handedness. Human circulating testosterone levels are higher in the male fetus and also vary with length of photoperiod. Therefore, a higher prevalence of left handedness, coupled with seasonal anisotropy (i.e., a non uniform distribution of handedness across birth months or seasons), may be expected among men. Prior studies yielded inconsistent evidence for seasonal anisotropy and suffered from confounding and a number of shortcomings affecting statistical power. This study examined hand preference and associations of handedness with sex, age, and season of birth in independent discovery ($n = 7658$) and replication ($n = 5062$) samples from Central Europe with latent class analysis (LCA). We found clear evidence of a surplus of left handed men born during the period November–January, which is consistent with predictions from the Geschwind Galaburda theory. Moreover, seasonal anisotropy fully accounted for the higher prevalence of left handedness among men, relative to women. Implications of these findings with regard to seasonal anisotropy research and handedness assessment and classification are discussed.

1. Introduction

Limb preference and handedness are phenomena of many biological species. In humans, a dominant preference of the right limb for manual tasks is culturally universal (with a prevalence rate of right handedness of about 90%; [Coren, 1993](#); [Peters, Reimers, & Manning, 2006](#)) and has been linked

to brain lateralization. Whilst the genetic bases of handedness are still under debate ([McManus, Davison, & Armour, 2013](#); [Medland et al., 2009](#)), there have been numerous accounts investigating its associations with human capacities and traits.

However, research into handedness has also produced some fickle and elusive results. Whilst a higher prevalence of left handedness among men appears to be a robust finding

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(Papadatou Pastou, Martin, Munafo, & Jones, 2008), the evidence for some assumed biological underpinnings of left handedness appear inconclusive. The cerebral lateralization theory of Geschwind Galaburda (Geschwind & Galaburda, 1985a, 1985b, 1985c, 1987) posits that high intrauterine levels of testosterone promote the development of left handedness because of differential effects of testosterone on brain hemisphere maturation; specifically, left brain hemisphere maturation is delayed. Male brains are exposed to substantially higher testosterone levels than female brains during prenatal development; hence, a sex effect for handedness may be expected. Furthermore, the secretion of sex hormones is higher during seasons with a long photoperiod, i.e., extended day length (Macchi & Bruce, 2004). As a consequence, the theory also suggests a birth seasonality effect, i.e., seasonal anisotropy, with regard to handedness.

There is direct (Lombardo et al., 2012) and indirect (Beaton, Rudling, Kissling, Taurines, & Thorne, 2011) evidence for associations of intrauterine testosterone levels with brain development and left handedness in men. In contrast, evidence for seasonal anisotropy is highly inconsistent: studies have reported anisotropy for widely different periods of the year, alternately affecting either men, or women, or both sexes likewise, or have found no effect at all (Stoyanov, Nikolova, & Pashalieva, 2011). One meta analysis (Jones & Martin, 2008), based on 48581 participants from 11 primary studies, has concluded that in the Northern hemisphere more left handed men are born during the period March–July. This meta analysis has been criticized with regard to the handedness classification in the primary studies (Beaton, 2008). Indeed, its conclusion runs counter to expectation, because in the Northern hemisphere the photoperiod is the longest during May–July. However, the critical period of testosterone induced human sexual differentiation does not occur perinatally, but much earlier prenatally, during gestational weeks 8–24; afterwards, plasma testosterone levels are generally low (Baron Cohen, Lutchmaya, & Knickmeyer, 2004). The peak of embryonic neuronal migration also occurs during gestational week 12–20 (Tau & Peterson, 2010). Testosterone induced effects on brain lateralization thus likely take place early during embryonic development. For the Northern hemisphere one would therefore expect more left handed men to be born during November–January, not during March–July, as suggested by the meta analysis. Some studies indeed have reported a higher prevalence of left handedness or non right handedness among men born in winter and/or autumn (Badian, 1983; Preti, Lai, Serra, & Zurrida, 2008; Stoyanov et al., 2011). However, in light of the accumulating inconsistent findings in general, more evidence is clearly needed. Moreover, studies on this theme need to take into account specific validity threats and statistical power considerations, as elaborated below.

Whilst the effects of sex hormones and sex chromosomes on brain differentiation are likely more complex than previously thought (Lentini, Kasahara, Arver, & Savic, 2012; Lust et al., 2011), available research in this area suffers from at least three different sources of bias and confounding which all threaten reliability and decrease statistical power. First, sample size is a pervasive problem in all handedness related research. The skewed 10:90 handedness ratio lowers the

power of statistical tests by about 50%, as compared to an optimally powered 50:50 ratio. Alternatively, oversampling left handers may result in dissimilarity of sampled populations and thus introduce additional bias.

Second, handedness classification is often problematic. Hand preference in writing is molded by social and cultural pressure (Preti et al., 2011), which might lead to an underestimation of the true associations of handedness with other variables, when handedness classification is solely based on the hand used for writing (Papadatou Pastou et al., 2008). More reliable multi item inventories, like the Edinburgh Handedness Inventory (EHI; Oldfield, 1971) or the Lateral Preference Inventory (LPI; Coren, 1993), still use arbitrarily adopted criteria and either a dichotomy (right/left, right/non right) or a trichotomy (right/mixed/left) for handedness classification. This may prohibit direct comparisons between studies and thus renders meta analytical results questionable (Beaton, 2008). Moreover, such varying classification criteria may even lead to different conclusions derived from the same data (Denny, 2008; Kelley, 2012). What is more, scoring and classification typically assigns equal weight to all items within an inventory. However, psychometrically this may be inappropriate (Dragovic, 2004). Existing taxometric (Dragovic, Milenkovic, & Hammond, 2008) and latent class analyses (Büsch, Hagemann, & Bender, 2009; Dragovic & Hammond, 2007; see Methods) suggest that handedness in actual fact is discrete and thus support classification approaches. However, studies on handedness associations with other variables seldom make use of latent variable analyses that would allow for a differential weighting of indicators and the direct modeling of handedness.

Third, photoperiod in the Northern hemisphere is longest during May–July, shortest during November–January, and intermediate during February–April and August–October. Some studies have based their analyses either on the four meteorological seasons (winter: December–February; spring: March–May; summer: June–August; autumn: September–November; e.g., Stoyanov et al., 2011) or on the astronomical seasons (winter: December 21/22–March 19/20; spring: March 20/21–June 20; summer: June 21–September 21/22; autumn: September 22/23–December 20/21; e.g., Tonetti, Adan, Caci, Fabbri, & Natale, 2012). Other studies have contrasted half year periods (meteorological autumn and winter versus spring and summer; Jones & Martin, 2008) or have used a monthly resolution for analysis (e.g., Milenković, Rock, Dragović, & Janca, 2008). However, the meteorologically or astronomically defined seasons match with the seasons, as defined by photoperiod variation, by at most about two thirds, and monthly analyses neglect the non independence of consecutive months.

In the present study an attempt was made to overcome all of the above shortcomings. Given the inconsistency of past research, we investigated handedness and its seasonal anisotropy in a discovery sample as well as in an independent replication sample, in order to assess the replicability of our own initial findings. Independent discovery and replication samples within the same study are now considered best practice in genome wide association studies, in order to guard against false positive findings and to demonstrate the robustness of an effect, if the replication is successful

(McCarthy et al., 2008). More generally, there is also an increased awareness of widespread replication failures and the irreproducibility of initially reported research findings across the medical and social sciences, which likewise has recently led to recommendations to intensify replication efforts (Asendorpf et al., 2013). Recommendations for incentivizing replication efforts (Button & Munafò, 2014) and to reward high power replication research (Wagenmakers & Forstmann, 2014) have also recently been published in this journal. Our study design thus conformed to these recent recommendations to counteract potentially false positive and thus irreproducible research findings.

We utilized latent class analysis (LCA) for the classification of handedness and examined the associations of handedness with sex, age, and season of birth. Specifically, we expected a surplus, i.e., seasonal anisotropy, of left handed men born during the period November–January.

2. Methods

2.1. Participants

In designing the study, independent discovery and replication samples were implemented. Descriptive statistics for the two samples are given in Table 1. Overall, both samples in the majority comprised Austrian and German participants and slightly more women than men. Participants from other countries were almost exclusively Caucasians; the further nationalities with the highest proportions were Italian (1.4%), Turkish (.6%), and Romanian (.4%). Range of participant age was broad.

2.2. Procedure and measures

The data were collected by a multitude of data collectors in the course of larger projects on individual differences variables (see Tran, Koller, et al., 2014, Tran, Stieger, & Voracek, 2014). The two data collection waves, yielding the discovery and the replication sample, were temporally separated by about one half year and independent from each other regarding the data collectors involved. The data collectors were geographically diverse and, mutually independent from each other, collected the data for course credit. In addition to basic demographic information and a number of further individual

differences measures, all participants provided hand preference ratings according to the German translation of four items selected from the LPI (Coren, 1993). Participants were approached on a personal basis, utilizing personal contacts and word of mouth advertising. Participants had to be fluent in German, as this was the survey language. Apart from insufficient language proficiency, there were no further exclusion criteria. Study participation was voluntary, anonymous, and participants were not remunerated for participation.

For survey economy, four items from of the full version, 12 item LPI handedness scale were selected for this research: hand preference with regard to writing (Item 1), throwing a ball to hit a target (throwing; Item 3), using a knife to cut something without simultaneously using a fork (knife; Item 6), and using a hammer to drive a nail into something (hammer; Item 7). Response options were *left*, *either*, and *right* (in this order and reading direction). For handedness classification, the consistency of the preferences across this item set may be utilized (thereby distinguishing consistent from inconsistent right handers), or, alternatively, a total score may be computed, with a coding of *left* = 1, *either* = 0, and *right* = +1 applied, and then summing up across all items. The sign and magnitude of the total score across the items is then indicative of the direction and strength of hand preference, respectively. Coren (1998) reported that self reported hand preferences according to the LPI were 96% concordant with respondents' behavioral performance on the same motor actions. As mentioned above, handedness was assessed in addition to further variables for a larger individual differences project. For the sake of efficiency, this assessment was thus deliberately kept brief. Apart from assessing the adult writing hand (Item 1), items were chosen with respect to broad trait coverage (vs redundancy), balance of fine motor versus gross motor skills represented in the specific item contents, and beneficial item performance indicators in particular. The items probing the writing hand, the throwing hand, and holding a knife had excellent psychometric properties (i.e., high squared multiple correlations of items with test scores) in a recent psychometric study of Dragovic (2004) as well as in a replication study of Milenkovic and Dragovic (2013), who developed a refined version of the 10 item EHI (Oldfield, 1971). The EHI shares six items with the LPI, and the refined 7 item version of Milenkovic and Dragovic (2013) includes three of the items (writing hand, throwing hand, and holding a knife) selected for the present study. Sample reliability figures (Cronbach α) for the full 12 item LPI frequently approaches .95 (e.g., Voracek, Reimer, Ertl, & Dressler, 2006). In the present study, Cronbach α values for the abbreviated 4 item scale were .91 and .92 in the discovery and replication samples, respectively, and .91, when both samples were combined. That is, reliability of the 4 item selection administered here was nearly as high as the reliability typically observed for the complete (12 item) LPI.

2.3. Analysis

LCA was used to investigate the preference ratings in the LPI. LCA is based on a latent variable model, wherein associations between observed variables (indicators) are explained by a

Table 1 – Sample descriptive statistics.

	Discovery sample	Replication sample
n	7658	5062
Women, n (%)	4456 (58.2%)	2749 (54.3%)
Age, range (years)	18–89	18–92
Interquartile range	22–35	22–48
Mean (SD)	30.24 (12.62)	35.74 (16.05)
Nationality, n (%) ^a		
Austria	5339 (70.0%)	3415 (67.7%)
Germany	1773 (23.2%)	1310 (26.0%)
Other	514 (6.7%)	316 (6.3%)

^a ns = 7626 and 5041 because of partially missing data.

Table 2 – Frequencies of hand preference across the administered LPI items.

Item	Sample	Hand preference			Missing values	χ^2
		Right	Left	Either		
Writing	Discovery	7006 (91.6%)	582 (7.6%)	60 (.8%)	10	.17
	Replication	4624 (91.4%)	394 (7.8%)	41 (.8%)	3	
Throwing	Discovery	6836 (89.4%)	553 (7.2%)	257 (3.4%)	12	8.41*
	Replication	4501 (89.1%)	414 (8.2%)	135 (2.7%)	12	
Knife	Discovery	6805 (89.0%)	604 (7.9%)	239 (3.1%)	10	9.62**
	Replication	4525 (89.4%)	423 (8.4%)	113 (2.2%)	1	
Hammer	Discovery	6853 (89.6%)	581 (7.6%)	214 (2.8%)	10	3.77
	Replication	4514 (89.3%)	419 (8.3%)	121 (2.4%)	8	

Note. $\chi^2 = \chi^2$ test of independence, $df = 2$. Note that the statistical tests were overpowered, owing to the large samples involved; observed between sample differences are negligible. * $p < .05$, ** $p < .01$.

number of discrete latent classes (Collins & Lanza, 2010). Formally, the statistical model of LCA may be expressed as a subtype of structural equation modeling and therefore shares many of its advantageous features. Specifically, LCA allows for a differential weighting of indicators with regard to class assignment (just like in factor analysis with regard to the estimation of the continuous latent trait).

We evaluated the fit of a number of latent class models, determining empirically the smallest number of latent classes that optimally accounted for the data observed in the discovery and the replication sample. In these analyses, items were treated as nominal, because of the evidence for a discrete nature of handedness (Büsch et al., 2009; Dragovic & Hammond, 2007; Dragovic et al., 2008). Model fit was assessed with the log likelihood value (LL) and the Bayesian information criterion, BIC(LL), which is based on the LL. Higher LL and lower BIC(LL) values indicate better model fit. Furthermore, we applied the likelihood ratio goodness of fit statistic (L^2) to compare the observed item response frequencies with the expected frequencies, as predicted by the model, and calculated the percentages of classification error. Lower (and preferably statistically not significant) L^2 values and lower percentages of classification error indicate better model fit. LCA was conducted with Latent GOLD 4.5 (Vermunt & Magidson, 2005). Sixty four (.5%) participants had 1 missing value and one participant 2 missing values for the four LPI items, which cases were included in the analysis. The latent handedness classes identified were used in the subsequent seasonal anisotropy analyses of handedness.

The hand preference observed in the LPI was investigated with χ^2 tests of independence, and the associations of hand edness with sex, age, and season of birth with logistic regression analyses and analyses of variance (ANOVA). Replicability across samples was assessed in combined samples analyses by adding a main effect and interactions with regard to type of sample (discovery vs replication) in logistic regression models and analyses of variance. In these tests, significant interaction effects would indicate non replicability of findings, i.e., effects not emerging in both the discovery and the replication sample. In the tests for birth month distribution and seasonality of handedness, we applied corrections for baseline differences regarding the total number of births per month in the samples; that is, the observed frequencies were differentially weighted to account

for these baseline differences. This adjustment also corrected for baseline differences due to the unequal length of months.

3. Results

3.1. Sample differences in demographics

The discovery sample comprised about 4% more women than the replication sample [$\chi^2(1) = 18.69, p < .001$]. Participants in the replication sample were on average 5.5 years older [$t(12718) = 21.56, p < .001, d = .39$]. Owing to the large samples, these inferential statistical tests clearly were overpowered and thus both nominally significant, albeit these differences would seem practically negligible.

3.2. Hand preference in the LPI

Table 2 lists the observed response frequencies on the four handedness items from the LPI. Endorsement of response categories were similar in the discovery and the replication sample. Between sample differences, wherever these were nominally significant, overall were negligible ($\leq 1\%$). Right hand preference ranged from 89.0% to 91.6% across samples and items, left hand preference from 7.2% to 8.4%. The prevalence of mixed handed writing was low (.8%), whereas a preference for both hands was reported by 2.2%–3.4% for the other three LPI items.

3.3. Latent structure of handedness

Table 3 displays the fit of 1, 2, 3, and 4 cluster models in the discovery and the replication sample. According to the BIC(LL) and the percentages of classification error, in both samples a 3 cluster solution fitted the data best. The mean posterior assignment probabilities of participants to clusters were high for Class 1 (99.4% and 99.4% in the discovery and the replication sample, respectively) and Class 2 (96.9% and 96.9%), whereas somewhat lower for Class 3 (89.4% and 90.0%), but overall indicating a high classification certainty. Fig. 1 depicts the estimated response probabilities for the 3 cluster model in the discovery and the replication sample.

Class 1, subsuming right handed individuals, had class sizes of 89.5%, with a 95% confidence interval of [88.7–90.3%]

Table 3 – Fit of latent class models for handedness classification.

Model	LL	Npar	BIC(LL)	L^2	df	p	Classification error, %
Discovery sample							
1 cluster	11665.52	8	23402.59	8441.22	184	<.001	.00
2 cluster	7908.81	17	15696.66	927.79	175	<.001	.52
3 cluster	7551.28	26	15335.10	212.74	166	.008	1.03
4 cluster	7511.84	35	15336.69	133.84	157	.910	3.82
Replication sample							
1 cluster	7652.41	8	15373.05	5871.32	176	<.001	.00
2 cluster	4988.65	17	10122.31	543.80	167	<.001	.49
3 cluster	4805.94	26	9833.65	178.38	158	.130	1.05
4 cluster	4782.74	35	9864.01	131.98	149	.840	3.58

Note. LL = log likelihood; Npar = number of parameters; BIC(LL) = Bayes information criterion based on LL; L^2 = likelihood ratio test statistic.

in the discovery sample and of 89.0% [88.0–90.0%] in the replication sample. Consistency of hand preference was high: response probabilities of right hand preference were $\geq 97\%$ on any item, and $\leq 2\%$ for mixed hand and left hand preference.

Class 2, subsuming left handed individuals, had class sizes of 7.7% [7.1–8.3%] and 8.4% [7.6–9.3%], respectively. Consistency of hand preference in Class 2 was lower than in Class 1: response probabilities of left hand preference were only $\geq 74\%$, with the lowest values for the writing hand item (80% in the discovery and 74% in the replication sample, respectively), and the highest values for the hammer item (91% and 92%). In return, response probabilities of right hand preference were highest for writing, 19% and 25% in the discovery and the replication samples, but lowest for hammer, 6% and 6%, respectively. The higher probability of right handed writing in the replication sample did not differ significantly from the discovery sample, as the respective 95% confidence intervals overlapped. Comparable to Class 1, mixed hand response probabilities were $\leq 4\%$ on any item in Class 2.

Class 3 had a class size of 2.8% [2.3–3.4%] and 2.6% [1.9–3.3%], respectively, and subsumed mixed handed

participants. Members of Class 3 had response probabilities of using either hand in throwing, knife, and hammer of $\geq 48\%$, which consistently were highest across both samples for the hammer item, 59% and 58% in the discovery and the replication sample, respectively. Mixed handed writing had response probabilities of 21% and 20%. Otherwise, members of Class 3 preferred the right over the left hand: right handed writing had response probabilities of 66% and 56%, as compared to 13% and 24% for left handed writing. With the exception of the knife item (25% in the discovery sample), across both samples right hand preference had response probabilities of at least 33% on all other items, as compared to a maximum of 24% (throwing) with regard to left hand preference. None of the response probabilities differed significantly between the samples in Class 3 (all respective 95% confidence intervals overlapped).

Overall, Item 4 (hammer) of the LPI contributed strongest in differentiating the latent classes. In the 3 cluster model, the loading of Item 4 was .85 and .87 in the discovery and the replication sample, respectively, whereas loadings of the other three items ranged between .74 and .81 in the two

Fig. 1 – Estimated response probabilities in the 3-cluster model. Black lines depict probabilities in the discovery sample, gray lines the respective probabilities in the replication sample.

Table 4 – Distribution of class membership within participant sex.

Sample		Class 1	Class 2	Class 3	χ^2
Discovery	Men	2854 (89.1%)	260 (8.1%)	88 (2.7%)	2.69
	Women	4019 (90.2%)	335 (7.5%)	102 (2.3%)	
Replication	Men	2038 (88.1%)	227 (9.8%)	48 (2.1%)	9.24**
	Women	2486 (90.4%)	204 (7.4%)	59 (2.1%)	

Note. $\chi^2 = \chi^2$ test of independence, $df = 2$. ** $p < .01$.

samples. Accordingly, prediction of class membership through Item 4 yielded the highest classification accuracy of all items: across both samples, the positive predictive value of Item 4 (hammer) was 98.8% for Class 1 (right handedness), 92.9% for Class 2 (left handedness), and 60.0% for Class 3 (mixed handedness).

Also, class membership reliably discriminated participants with regard to their LPI scores. Class membership explained 91% of total score variability in the two samples [ANOVA with factors class and sample; class: $F(2, 12649) = 64070.11, p < .001, \eta_p^2 = .910$; sample: $F(1, 12649) = 2.52, p = .112, \eta_p^2 < .001$; interaction: $F(2, 12649) = 4.38, p = .013, \eta_p^2 = .001$], with members of Class 1 (right handers) scoring highest ($M = 3.90, SD = .40$), Class 3 (mixed handers) scoring intermediate ($M = .83, SD = 1.15$), and Class 2 (left handers) scoring lowest ($M = 2.84, SD = 1.43$). Given the fact that the internal consistency (Cronbach α) of the four LPI items was .91, it follows that essentially all reliable LPI score variance was explained by class membership (as indicated by the above effect size: $\eta_p^2 = .910$). Members of Class 1 also scored somewhat lower in the discovery than in the replication sample (significant interaction); however, this difference was practically negligible ($\eta_p^2 < .001$).

3.4. Associations of handedness with sex, age, and season of birth

The distribution of class membership with regard to participant sex is presented in Table 4. In both samples, the proportion of Class 2 members (left handers) was slightly elevated among men, relative to women. This effect was significant in the replication sample, but not in the discovery sample (see Table 4). In the logistic regression analysis, no sex effect was obtained overall; however, the test for non replicability (interaction effect of sex with sample) also failed to reach nominal significance ($p = .094$; see Table 5, column 1). When this interaction term was omitted from the model, the sex effect became significant ($OR = 1.19 [1.05–1.36], p = .007$). We accepted this result as suggestive of an overall sex difference in Class 2 proportions that amounted to a 1.3% [4–2.3%] difference (overall proportions of Class 2 were 8.8% among men and 7.5% among women).

No association of class membership with age could be observed [ANOVA with factors class and sample; class: $F(2, 12714) = .61, p = .545, \eta_p^2 < .001$; sample: $F(1, 12714) = 72.73, p < .001, \eta_p^2 = .006$; interaction: $F(2, 12714) = .01, p = .993, \eta_p^2 < .001$]. However, older age among members of Class 2 (left handers) specifically predicted right handed writing (see Table 5, column 2). Members of Class 2 had their odds raised by a factor of 3.12 [2.55–3.81] per 10 years increase of age to

report right hand writing. No further association of hand preference with age was observed for any of the other three LPI items in Class 2 ($ps \geq .152$ for the overall models; results omitted for brevity).

The birth month distribution of class membership is depicted in Fig. 2, collapsing the discovery and the replication samples and omitting Class 3 (mixed handers) for which at most 20 observations were available for each month. Consequently, the respective estimates of the proportions were deemed unreliable and therefore not used in Fig. 2. Among men, the monthly proportions of Class 2 (left handers) exceeded that of Class 1 (right handers) consecutively from November to January by 1.4%–1.9%, whereas no such pattern was observable for women. Compared to the remainder of the year, among men (but not women) the proportion of left handedness was significantly higher during the period November–January (see Table 5, columns 3 and 4). Overall, the difference in Class 2 proportions among men with regard to the period November–January (10.5%) and the remainder of the year (8.2%) amounted to 2.3% [4–4.1%]. Moreover, seasonality fully accounted for the difference in the proportions of left handedness between men and women (see Table 5, column 5). Of all female Class 2 members, 21.5% were born during November–January, whereas among men the corresponding figure was 29.9%.

In seasonality analyses (Table 5, columns 3–5), left handedness was contrasted with the mixed handedness and right handedness groups combined. In additionally calculated analyses, combined groups of left handedness and mixed handedness were contrasted with the right handedness group. This alternative analysis however left the above findings essentially intact (details omitted for brevity). In particular, the sex by season interaction remained significant and was of comparable magnitude as in the above analysis, $OR = 1.45 [1.08–1.95], p = .014$. In similar vein, conducting the seasonality analyses separately for the two samples did not yield results contradicting those of the above analysis. The sex by season interaction was clearly significant in the replication sample, $OR = 1.97 [1.22–3.19], p = .006$, and only closely missed nominal significance in the discovery sample, $OR = 1.46 [.99–2.14], p = .054$, which may be attributed to the relative loss of statistical power in these separate sample analyses, as compared to the combined sample analysis. Of note, confidence intervals of these sex by season effect estimates from the separate sample analyses overlapped considerably, thus suggesting comparable effects. Furthermore, in the separate sample analyses there also emerged a main effect of season in the replication sample, $OR = .65 [.45–.95], p = .025$, but not in the discovery sample, $OR = .91 [.70–1.19], p = .499$. However, in the presence of the significant

Table 5 – Associations of left-handedness (Class 2) with sex, age, and season of birth according to logistic regression analyses.

	Sex: Class 2 versus classes 1 and 3	Age: Right handed writing among members of Class 2 ^a	Season of birth: Class 2 versus classes 1 and 3 ^b		
			Men	Women	Sexes combined
Sex ^c	1.09 [.92 1.29]				1.05 [.91 1.22]
Age		1.12 [1.10 1.14]***			
Season ^d			1.33 [1.01 1.76]*	.91 [.70 1.19]	.81 [.66 1.01]
Sample ^e	.99 [.82 1.18]	.58 [.16 2.12]	1.27 [1.01 1.58]*	1.06 [.86 1.29]	1.09 [.94 1.25]
Sex × season					1.53 [1.09 2.15]*
<i>Tests of nonreplicability</i>					
Sex × sample	1.25 [.96 1.62]				
Age × sample		1.00 [.97 1.04]			
Season × sample			.96 [.64 1.46]	.72 [.45 1.13]	
Sex × season × sample					1.13 [.78 1.64]
Model fit χ^2 (df)	12.40 (3)**	371.86 (3)***	12.01 (3)**	5.90 (3)	20.02 (5)**
Nagelkerke R ²	.002	.478	.005	.002	.004

Note. Cells entries are odds ratios with 95% confidence intervals. * $p < .05$, ** $p < .01$, *** $p < .001$.

^a $n = 1013$ of whom 213 reported right handed writing.

^b Corrected for baseline differences in numbers of births per month. Coding:

^c 1 = men, 0 = women.

^d 1 = November–January, 0 = remainder of the year.

^e 1 = replication sample, 0 = discovery sample.

sex by season interaction (see above), the significant main effect of season in the replication sample simply mirrored the relatively stronger overall effect of sex on handedness in this sample (see above).

Notably, comparing the meteorological winter with the remainder of the year, a seasonality effect could not be verified (logistic regression model as in Table 5, column 5; sex by season interaction: OR 1.27 [.90–1.79], $p = .183$; further details omitted for brevity). Likewise, no association of season of birth with Class 2 membership was observed by comparing the half year periods August–February (autumn and winter) and March–July (spring and summer; sex by season interaction: OR 1.06 [.79–1.42], $p = .691$; further details omitted for brevity).

4. Discussion

This study set out to reassess predictions for seasonality of handedness, following the theory of Geschwind Galaburda, using two large and independent samples and LCA of self reported handedness data. Our results suggest that, first, three classes of handedness need to be considered – right handedness, left handedness, and mixed handedness. Second, handedness appeared to be independent of age, but not of sex and season of birth: men were more often left handed than women, and more left handed men were born during the period November–January than during the remainder of the year, which consorts well with predictions from the Geschwind–Galaburda theory. Third, in the present data this

Fig. 2 – Radial plots of the birth month distribution of class membership among men (left) and women (right), comparing the Classes 1 and 2, with the discovery and replication samples combined. Proportions are corrected for baseline differences in numbers of births per month. Gray lines signify the constant proportion of 8.3% that would be expected for each handedness class under the assumption of uniform distribution of handedness across birth months.

seasonal anisotropy fully explained the higher prevalence of left handedness among in men, relative to women.

Thus, even though previous results on seasonal anisotropy were inconsistent, this study provides clear evidence in support of one prediction derived from the theory of Geschwind–Galaburda and therefore adds to the prior positive evidence of Badian (1983), Preti et al. (2008), and Stoyanov et al. (2011). Furthermore, because according to our data seasonal anisotropy specifically affected men born during the period November–January, as expected, one may infer that the critical period of testosterone induced effects on brain lateralization effectively is limited to gestational weeks 14–18. Otherwise, individuals born in both the precedent and subsequent months should have been affected to some extent as well. Seasonal anisotropy was specifically confined to men, as in the reports of Badian (1983) and Stoyanov et al. (2011).

The proposed short critical period corresponds to the magnitude of the effects observed in this study: men's prevalence rate of left handedness exceeded women's by 1.3% (which broadly corresponds to previous estimates; Peters et al., 2006), and among men the overall prevalence of left handedness of birth cohorts of November–January exceeded that of the remainder of the year by only 2.3%. In contrast, the difference in the proportions of left handedness among men and women for the period November–January, 29.9% vs 21.5%, appears sizable. Nonetheless, given the overall low prevalence of left handedness, large samples and high precision of handedness classification are mandatory to spot an effect of this size. For these combined reasons, previous studies could not provide a clear picture on this seasonal anisotropy effect. In support of these arguments is the fact that we were unable to obtain the seasonality effect with our own data, when we, by way of trial, analyzed half year periods (Jones & Martin, 2008) or seasons not aligned to photoperiod variation (Tonetti et al., 2012).

Even though in our data seasonal anisotropy fully accounted for the sex difference in left handedness, there conceivably are a number of further prenatal, perinatal, and postnatal candidate factors to consider for refined explanation, such as maternal age, birth order, birth stress, or cultural factors (Badian, 1983; Bailey & McKeever, 2004; Papadatou Pastou et al., 2008). Also, regional effects and latitude need to be taken into account vis à vis seasonal variation (Badian, 1983). For example, the null finding of the large ($N = 15390$) study of Cosenza and Mingoti (1995) may have been caused by the only small seasonal variation in photoperiod at their study locale, which was situated close to the equator (Minas Gerais in Brazil; latitude 19.54° S). In contrast, sampling in the present study primarily took place in Central Europe (Austria and Germany), with the latitudes for these two countries extending from 46° to 55° N. We suggest that the seasonal variation in photoperiod – not merely meteorologically or astronomically defined seasons – together with the moderating effect of latitude on the extent of this seasonal variation might be beneficially taken into account in future research.

Regarding handedness assessment, our study provided prevalence estimates of right handedness (about 89%) that are in line with many previous results. Our data are further indicative of a trichotomy of handedness, comprising right handedness, left handedness, and mixed handedness, which is consistent with recent taxonomic and latent

class analyses (Dragovic & Hammond, 2007; Dragovic et al., 2008). Overall, the prevalence of left handedness was about 8%, whereas those of mixed handedness less than 3%.

Analysis revealed that consistency of handedness was highest among right handers, but less so among non right handers. Inconsistent use of the left hand was more or less the norm among left handers, who frequently stated to prefer their right hand as well. Concerning the writing hand, we found strong evidence for a decreasing social and cultural pressure against left handedness among the more recent birth cohorts: older, but not younger, left handers frequently reported writing with their right hand (McManus, Moore, Freegard, & Rawles, 2010). For this reason, hand preference for writing appears to be a relatively weak and moreover confounded indicator of handedness (Preti et al., 2011), which therefore should probably omitted from handedness classification.

Even though the group of mixed handers was only small and right hand preference was also frequent among these individuals, this was the only group which consistently and most often indicated no hand preference regarding the four manual tasks of the LPI. There have been accounts of mixed handedness alternately being described as detrimental (Crow, Crow, Done, & Leask, 1998; Kelley, 2012; Nettle, 2003; Peters et al., 2006) or beneficial (Christman & Propper, 2001; Denny, 2008; Niebauer, 2004) for various psychological abilities and health conditions. Our study suggests that inconsistent results such as these might partly stem from mingling mixed handers, who are characterized by no clear hand preference, with left handers, who are generally characterized by an inconsistent preference of their left hand. Clearly, more research is needed to shed more light on this specific topic.

Furthermore, we found evidence that a bimanual task which involves the risk of hurting oneself, if executed faultily (using a hammer to drive a nail into something), had the highest reliability of classifying handedness on its own. Bimanual tasks are assumed to confer a higher classification reliability (Papadatou Pastou et al., 2008). However, our study suggests that it may need to be a task which requires accuracy and, if executed faultily, exerts a direct negative or hurtful consequence to the individual performing it. This idea ought to be followed up, as it might improve the assessment and classification of handedness from inventories of hand preference.

Summing up, utilizing a discovery and replication sample approach, our study provided clear evidence for a seasonal anisotropy of left handedness among men, which effect is consistent with predictions derived from the Geschwind–Galaburda theory. Future research along these lines may need to take more closely into account the validity threats and statistical power issues elaborated on here. Likewise, the moderating effects of latitude on the amount of seasonal variation of photoperiod should be further investigated vis à vis seasonal anisotropy.

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