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Benevolent sexism and hostile sexism across the ages

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Authors' Note. MPlus syntax for the models reported here are available on the New Zealand Attitudes and Values Survey website (<http://www.psych.auckland.ac.nz/uoa/NZAVS>).

Abstract

Ambivalent sexism theory states that prejudice toward women comprises two interrelated ideologies. Endorsement of hostile sexism—aggressive and competitive attitudes toward women—is linked with endorsement of benevolent sexism—paternalistic and patronizing attitudes toward women. We conduct the first systematic tests of how endorsement of sexism differs across age and across time, using six waves of a nationally representative panel sample of New Zealand adults ($N = 10,398$). Results indicated U-shaped trajectories for men's endorsement of hostile sexism, women's hostile sexism, and women's benevolent sexism across the lifespan. However, over time, endorsement of these sexist attitudes tended to decrease for most ages. In contrast, men's benevolent sexism followed a positive linear trajectory across age, and tended not to change over time. These results provide novel evidence of how ambivalent sexism differs across age and highlight that benevolent sexism is particularly tenacious.

Keywords. Benevolent Sexism, Hostile Sexism, Ambivalent Sexism, Age Differences, Latent Growth Model.

Benevolent sexism and hostile sexism across the ages

Ambivalent sexism theory (Glick & Fiske, 1996) states that prejudice toward women is distinctive because it comprises two related ideologies with ambivalent tones. *Hostile sexism* is an ideology which characterizes women as incompetent, overly emotional and attempting to manipulate men to gain power. However, hostile sexism is not an effective way to maintain men's power because it impedes the fulfilment of heterosexual men's intimacy needs (Glick & Fiske, 1996). *Benevolent sexism* is an ideology that tempers these relational costs, while maintaining men's power, by characterizing women as delicate, pure and in need of men's protection and care. Existing research primarily focuses on the consequences of endorsing hostile sexism and benevolent sexism, particularly as a predictor of gender inequalities (for reviews see Connor, Fiske, & Glick, 2016; Hammond & Overall, 2016; Hammond & Overall, 2017), but only theorizes about how endorsement of sexism differs across the lifespan (see Glick & Fiske, 1996; Glick & Fiske, 2001; Glick & Hilt, 2000). The current research utilizes a nationally representative longitudinal sample to model (1) how the endorsement and interrelation of sexist attitudes differ across age, and (2) age-cohort changes in the endorsement of sexism over a five-year span.

Ambivalent Sexism and the Structure of Sexist Attitudes

A fundamental principle of ambivalent sexism theory is that sexist attitudes encompass two positively associated ideologies that both function to maintain gender inequality (Glick & Fiske, 1996). Men and women who endorse *hostile sexism*, derogatory attitudes toward women who could challenge men's power, also tend to endorse *benevolent sexism*, subjectively positive expressions toward women in traditional roles (Glick & Fiske, 1996). Sexist attitudes have three theorized sources that are rooted in the societal and biological differences between men and women (Glick & Fiske, 1996). First, sexist attitudes justify men's power by expressing that men are dominant over women (hostile sexism) and

simultaneously responsible to protect and provide for women (benevolent sexism). Second, sexist attitudes justify distinctiveness of gender identity by prescribing that women should be warm and docile, whereas men should be men competent and composed. Third, sexist attitudes encourage intimacy by idealizing women's relational capacity to "complete" men (benevolent sexism) but temper women's interpersonal power by vilifying women who might manipulate men's relational needs (hostile sexism).

In tandem, sexist ideologies function to maintain gender inequality across societies (Brandt, 2011; Glick & Fiske, 1996; Glick et al., 2000), such as by discriminating against or intimidating women in career or political domains (Connor et al., 2016) and rewarding and praising women who adopt traditional gender roles (Hammond & Overall, 2016). Attention is often directed to the ways that benevolent sexism undermines women's resistance to societal gender inequalities (e.g., Becker & Wright, 2011; Hammond & Sibley, 2011), and encourages women to invest in their relationships rather than pursue career or educational goals (e.g., Chen, Fiske, & Lee, 2009; Montañés, de Lemus, Moya, Bohner, & Megías, 2013). However, none of the evidence linking sexist attitudes to markers of gender inequality have investigated how sexist attitudes differ across the lifespan or over time. Thus, the literature on sexist attitudes lacks a contextual map of the ages for which endorsement of sexism (and its effects) are most prominent or the ages for which endorsement of sexism *changes*. In the current research we begin to build this map by modelling differences in the endorsement of benevolent sexism and hostile sexism across age, and further, modelling the five-year change in endorsement of sexist attitudes.

The Expected Trajectory of Sexist Attitudes across Age

We established exploratory predictions how sexist attitudes might differ across the average lifespan based on the *normative age-graded* concerns that are relevant to the theorized sources of sexism we reviewed above—power, identity, and relational needs.

Normative age-graded events are influences on development that occur in similar ways and at similar life stages for people in a given culture, including beginning a career or committing to a serious relationship (Baltes, Reese, & Lipsitt, 1980; also see Elder, 1995, Heckhausen, 2006, for similar paradigms). In younger adulthood, normative goals include the attainment of personal status and power, establishment of a social identity and career, and serious intimate relationships (Arnett, 2001; Baltes et al., 1980; Casey, Jones, & Somerville, 2011; Takahashi, 2005). By contrast, in middle adulthood, people's traits and roles are less likely to follow societal norms (Bedford & Turner, 2006; O'Neil & Egan, 1992) and their romantic relationships tend to be stable but less central to the self (Chopik, Edelstien, & Fraley, 2013). However, in older ages, traditional and conservative values are relatively high (Gouveia, Vione, Milfont, & Fischer, 2015), and people experience high dependence on romantic partners (Wrzus, Hänel, Wagner, & Neyer, 2013). In sum, across the lifespan, concerns over power, identity and dependence follow a high-low-high course. We therefore developed tentative expectations of U-shaped trajectories of sexism across age: Relatively high endorsement of sexism in early adulthood, relative low endorsement in middle-adulthood, and relatively high endorsement in late adulthood.

A U-shaped pattern of endorsement of hostile sexism and benevolent sexism across age is consistent with demographic data from two large Spanish samples. Endorsement of hostile sexism and benevolent sexism were relatively high in the youngest and oldest age-groups relative to middle-aged age-groups (Fernández, Castro, & Lorenzo, 2004, $N = 1003$; Gariagordobil & Aliri, 2013, $N = 5110$). We extend this prior research by testing whether this pattern is curvilinear, therefore providing the first formal test of whether endorsement of sexism follows a U-shaped trajectory across age. Furthermore, we employ novel methodologies to examine further characteristics of sexist attitudes: We (1) conduct age-cohort based longitudinal tests over a 5-year span to examine how endorsement of sexism

changes longitudinally, (2) model the relationship *between* hostile sexism and benevolent sexism across age (discussed next), and (3) estimate exploratory cohort effects by comparing the age-based trajectory of sexism with the model of longitudinal change in sexism.

The Expected Ambivalence of Sexist Attitudes across Age

We examined age-related differences in the association between benevolent sexism and hostile sexism because this association is at the heart of ambivalent sexism theory; sexism is proposed to be ambivalent *because* of the interrelatedness of benevolent sexism and hostile sexism (Glick & Fiske, 1996; Glick et al., 2000). Benevolent sexism is a “crucial complement to hostile sexism” because subjectively positive praise for women in traditional roles is used to rationalize aggression toward non-traditional women (Glick & Fiske, 2001, p. 111). However, prior research has illustrated variability in this association. Benevolent sexism and hostile sexism appear to be more strongly related for women than for men, and for younger samples than for older samples (Glick & Fiske, 1996; Glick et al., 2000). However, no research has specifically tested differences in the association between benevolent sexism and hostile sexism across age. Following the prediction that people’s sexist beliefs become more individuated and idiosyncratic across the lifespan (Glick & Fiske, 1996, p. 508), we expected that the association between benevolent sexism and hostile sexism will weaken across age.

Current Research

Our study tested a nationally representative, longitudinal panel sample in an egalitarian country to examine the age-differences, relationships between, and changes in endorsement of benevolent sexism and hostile sexism. We conducted four sets of analyses. First, age-based latent growth models estimated the linear and curvilinear trajectories of benevolent sexism and hostile sexism across age for men and for women. Second, we estimated the standardized relationship between benevolent sexism and hostile sexism,

adjusting for gender, across age. Third, we estimated the 5-year change in sexism in multi-group cohort-sequential latent growth models. Finally, we conducted tentative estimation of cohort effects by comparing effects between the age-based model and cohort-based model of sexism.

Method

Participants

Eligible participants were 10,398 people (62.5% women) who responded to at least three out of the first six waves of the New Zealand Attitudes and Values Study (NZAVS), an annual longitudinal panel survey. Average age at Wave 1 (2009) was 49.45 (SD = 14.98; see Table 1). At least three waves of completion were necessary for participants' data to contribute information about the rate of change in sexism and meet the requirements to fit latent growth models. The majority of participants (85.7%) were New Zealand European. See the Online Supplementary Materials and Satherly et al. (2015) for detailed demographics of the sampling procedures, sample characteristics, and how these characteristics relate to attrition.

Table 1. The number of participants in each of the sequential 5-year birth cohorts organized for the multi-group latent growth models.

Birth Cohort	<i>N</i> women	<i>N</i> men	Ages Sampled (T1 to T5)
1940 and older	198	197	69 – 74
1941 to 1945	363	330	64 – 69
1946 to 1950	637	509	59 – 64
1951 to 1955	744	492	54 – 59
1956 to 1960	803	497	49 – 54
1961 to 1965	794	488	44 – 49
1966 to 1970	721	366	39 – 44
1971 to 1975	644	284	34 – 39
1976 to 1980	491	206	29 – 34
1981 to 1985	356	104	24 – 29
1986 to 1990	381	165	19 – 24
1991 and younger	186	63	18 – 23

Measures

Benevolent sexism (e.g., “Women should be cherished and protected by men”) and hostile sexism (e.g., “Women are too easily offended”) were assessed at each wave using a ten-item short form of the Ambivalent Sexism Inventory (Glick & Fiske, 1996; 1 = *Strongly Disagree*; 7 = *Strongly Agree*).

Attrition

Over 60% of participants were retained from Wave 1, and approximately 80% of participants were retained wave-to-wave (also see Satherly et al. 2015). We examined attrition by comparing eligible participants who had completed Wave 1 with participants who did not meet the inclusion criterion. Eligible participants were more likely to be women (61.1% vs. 55.5%) and tended to be older ($M = 49.51$, $SD = 14.97$) than those not selected ($M = 44.66$, $SD = 17.04$, $d = .30$). Eligible participants’ endorsement of benevolent sexism ($M = 4.07$, $SD = 1.17$) and endorsement of hostile sexism at Wave 1 ($M = 3.38$, $SD = 1.25$) were both relatively lower than those not selected (Benevolent sexism $M = 4.40$, $SD = 1.16$, $d = .285$; Hostile sexism $M = 3.69$, $SD = 1.31$, $d = .29$).

Results

We first examined descriptive statistics for endorsement of sexism and tested measurement invariance. We then tested our predictions in four sets of multilevel models using *MPlus 7.4* (Muthén & Muthén, 1998-2015), with full-information maximum likelihood (robust standard error) estimation. These models also account for missing data by weighting model parameters so that slopes for participants who completed more waves contribute more to the overall parameter estimates.

Measurement

Descriptive Statistics. Descriptive statistics, bivariate correlations and reliabilities for the measures are presented in Table 2. The test-retest correlations for hostile sexism and

Table 2. Bivariate correlations and descriptive statistics for Hostile Sexism (HS) and Benevolent Sexism (BS) across the 6 repeated assessments from Wave 1 (W1) to Wave 6 (W6); and age and gender (assessed at Wave 1).

	1	2	3	4	5	6	7	8	9	10	11	12	13	14
1. HS_W1														
2. HS_W2	.733**													
3. HS_W3	.700**	.751**												
4. HS_W4	.682**	.734**	.765**											
5. HS_W5	.673**	.721**	.756**	.774**										
6. HS_W6	.661**	.699**	.733**	.745**	.768**									
7. BS_W1	.430**	.366**	.346**	.341**	.355**	.349**								
8. BS_W2	.366**	.415**	.353**	.361**	.364**	.348**	.746**							
9. BS_W3	.378**	.371**	.396**	.368**	.383**	.370**	.728**	.751**						
10. BS_W4	.343**	.352**	.354**	.417**	.383**	.371**	.709**	.752**	.755**					
11. BS_W5	.351**	.349**	.373**	.382**	.431**	.394**	.703**	.733**	.752**	.764**				
12. BS_W6	.342**	.343**	.371**	.378**	.392**	.429**	.681**	.716**	.728**	.740**	.770**			
13. Gender	.252**	.226**	.262**	.261**	.236**	.222**	.118**	.108**	.134**	.129**	.141**	.147**		
14. Age	-.012	.012	.019	.022	.050**	.052**	.116**	.104**	.118**	.136**	.140**	.148**	.115**	
Mean	3.317	3.212	3.021	3.014	3.023	3.057	4.073	4.030	3.863	3.826	3.791	3.811	—	49.450
SD	1.252	1.227	1.222	1.213	1.204	1.209	1.694	1.153	1.168	1.168	1.190	1.192	—	15.979
Reliability	.807	.807	.821	.819	.819	.825	.722	.728	.742	.737	.746	.757	—	—

Note. ** $p < 0.01$. Bolded associations denote associations relevant to assessing the stability of sexism across time.

benevolent sexism (presented in bold font) indicated relatively high stability for both hostile sexism and benevolent sexism, and measurements at each time point demonstrated consistently acceptable internal reliability.

Measurement Invariance of Sexism. We first tested the possibility that a particular five-year age cohort (at Wave 1) answered the ambivalent sexism inventory in substantially different ways to other cohorts, or if men answered differently to women. We estimated three increasingly restrictive latent-variable models (i.e., configural, metric, and scalar). Configural invariance was specified by restricting item loadings so that manifest indicators for each sexist ideology loaded on the corresponding latent sexist ideology across all of the age cohorts or across gender. Metric invariance was specified by restricting the item loadings to equality across all of the age cohorts or across gender. Scalar invariance was specified by restricting both item loadings and item intercepts to equality across all of the age cohorts or across gender. As displayed in Table 3, the relative model fit indices demonstrated reasonably good fit for all measurement specifications across age groups and between genders, but minor violations when specifying scalar invariance ($\Delta CFI > .01$; see Cheung & Rensvold, 2002). These results generally indicated that metric measurement invariance best fit the data, suggesting that people generally answered the ambivalent sexism inventory in similar ways, but that any mean differences in sexism across age or gender did not necessarily reflect group-differences in the latent variables.

Temporal Measurement Invariance of Sexism. We next utilized all available waves of data to strict measurement invariance in latent-growth models across time. We tested models in three parts, displayed in Table 3, imposing configural, metric, and scalar invariance for each ideology (1) over time, (2) across the 5-year age-cohorts (see Table 1) over time, and (3) across gender over time. The measurement specifications identified minor violations of temporal measurement invariance for both benevolent sexism and hostile sexism across time

Table 3. Tests of strict measurement invariance. Fit indices are displayed for benevolent sexism and hostile sexism, modelled simultaneously at Wave 1 (*Cross-Sectional* model), and separately to test temporal measurement invariance (*Over Time* models).

Model	χ^2	<i>df</i>	RMSEA	CFI	TLI	SRMR
Cross-Sectional Model						
1. Between age groups	1843.10	476	.079	.897	.863	.059
2. Between age groups <i>metric</i>	1978.53	580	.072	.894	.885	.064
3. Between age groups <i>scalar</i>	2268.62	684	.071	.880	.890	.068
4. Between gender <i>configural</i>	1217.89	68	.072	.907	.877	.049
5. Between gender <i>metric</i>	1253.18	76	.069	.905	.887	.050
6. Between gender <i>scalar</i>	1594.75	84	.074	.878	.869	.058
Benevolent Sexism Over Time						
1. <i>Configural</i> over time	1756.77	29	.076	.891	.831	.072
2. <i>Metric</i> over time	2349.49	35	.080	.854	.813	.087
3. <i>Scalar</i> over time	2467.83	39	.078	.847	.824	.089
4. Between age groups	2578.50	341	.086	.854	.788	.090
5. Between age groups <i>metric</i>	2732.48	425	.079	.849	.824	.096
6. Between age groups <i>scalar</i>	3075.09	520	.075	.833	.841	.104
7. Between gender <i>configural</i>	2136.08	62	.081	.867	.807	.082
8. Between gender <i>metric</i>	2218.95	74	.075	.863	.833	.084
9. Between gender <i>scalar</i>	3095.96	88	.082	.807	.803	.103
Hostile Sexism Over Time						
1. <i>Configural</i> over time	576.978	29	.043	.966	.948	.027
2. <i>Metric</i> over time	1171.20	35	.056	.930	.911	.052
3. <i>Scalar</i> over time	1251.49	39	.055	.926	.914	.053
4. Between age groups	1408.85	341	.060	.934	.904	.056
5. Between age groups <i>metric</i>	1601.29	425	.056	.927	.915	.064
6. Between age groups <i>scalar</i>	1866.82	520	.054	.916	.920	.071
7. Between gender <i>configural</i>	1054.73	62	.056	.936	.907	.048
8. Between gender <i>metric</i>	1252.36	74	.056	.924	.907	.058
9. Between gender <i>scalar</i>	2265.92	88	.070	.859	.856	.113

(see Cheung & Rensvold, 2002). Assessing measurement invariance in these longitudinal models places high power demands on the data and, accordingly, should be considered conservative tests of measurement invariance. Nonetheless, these results again indicated metric measurement invariance best fit the data, again highlighting the need for caution in interpreting mean differences in sexism across time or between groups because these differences may not necessarily reflect effects in the latent variables.

Differences in Sexist Attitudes across Age

Model Estimation. We estimated an age-based cohort-sequential latent growth model (see Milojev & Sibley, 2016; Preacher, Wichman, MacCallum, & Briggs, 2008) based on participants' age at each measurement wave, utilizing exact age at survey completion as an individually-varying time indicator, allowing us to utilize data from all waves while accounting for any variation in time between survey completions.¹ We tested for gender differences by estimating a multi-group model that allowed the intercepts and trajectories for men and women to differ. The slopes reported here are a relatively conservative test of the age-related trajectories of sexism because they account for the cohort-by-cohort change across time that we report in the next section. We modelled the differences in benevolent sexism and hostile sexism as a polynomial growth function including linear, quadratic and cubic components (non-significant higher-order interaction terms were retained in the models to most accurately model the trajectories across age).

Results. Parameter estimates for models estimating overall changes in the trajectories of benevolent sexism and hostile sexism from ages 19 to 74 are presented in Table 4 and Figure 1 (dark line). For men, a significant linear effect indicated that the trajectory of benevolent sexism increased across age. For women, a significant quadratic effect of age

¹ Analyses included all participants, but we conservatively estimated the model-implied growth trajectories between ages 19-74 given the relatively lower sample sizes for extreme values of age. See OSM for more detail.

Table 4. Estimates from Latent Growth Models estimating the age-based trajectories of endorsement of benevolent sexism (upper model) and hostile sexism (lower model).

<i>Model</i>	<i>Estimate</i>	<i>se</i>	<i>z</i>	<i>95% CIs</i>		<i>Variances</i>
				<i>Low</i>	<i>High</i>	
Benevolent Sexism						
<i>Men</i>						
Intercept	3.924	.024	164.962*	3.878	3.971	1.020*
Age – Linear	.073	.017	4.371*	.040	.100	.394*
Age – Quadratic	.016	.009	1.812	-.001	.030	.038
Age – Cubic	.002	.003	0.890	-.003	.007	.002
<i>Women</i>						
Intercept	3.672	.018	199.213*	3.636	3.708	1.020*
Age – Linear	.010	.014	0.742	-.017	.038	.394*
Age – Quadratic	.034	.007	5.111*	.021	.047	.038
Age – Cubic	-.003	.002	-1.076	-.007	.002	.002*
Hostile Sexism						
<i>Men</i>						
Intercept	3.437	.027	128.691*	3.385	3.490	1.099*
Age – Linear	-.067	.019	-3.603*	-.103	-.031	.446*
Age – Quadratic	.033	.010	3.413*	.014	.052	.051
Age – Cubic	-.002	.003	-0.531	-.008	.004	.003
<i>Women</i>						
Intercept	2.766	.018	151.758*	2.731	2.802	1.099*
Age – Linear	-.034	.014	-2.407*	-.062	-.006	.446*
Age – Quadratic	.040	.006	6.226*	.028	.053	.051
Age – Cubic	.001	.002	0.212	-.004	.005	.003

Note. $N = 10398$ ($Men = 3899$; $Women = 6499$) for estimation of Benevolent Sexism. $N = 10397$ ($Men = 3899$; $Women = 6498$) for estimation of Hostile Sexism. For ease of displaying parameters, age (grand-mean centered) was scaled so that each unit represented 10 years. See OSM for more detail. * $p < .05$.

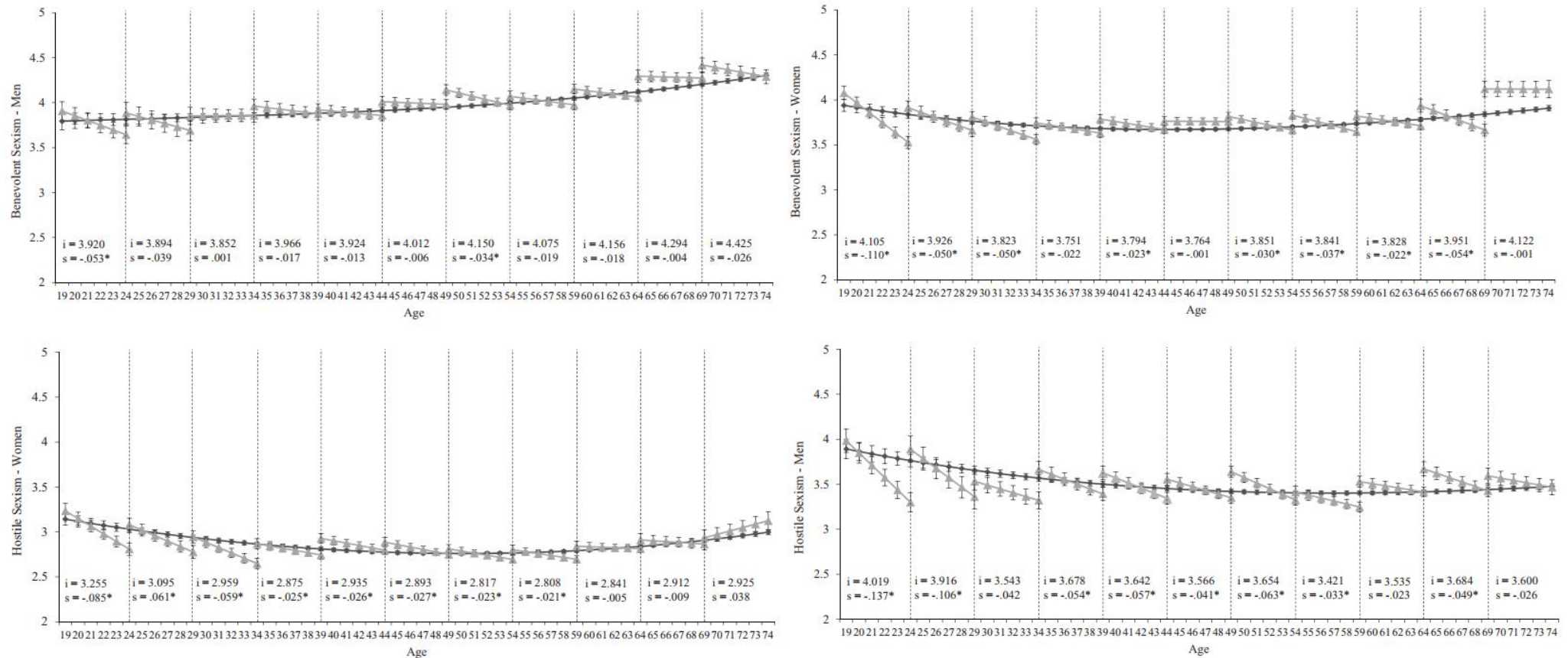


Figure 1. Change trajectories for men's and women's endorsement of benevolent sexism (displayed as a dark line in the upper panels) and hostile sexism (displayed as a dark line in the lower panels) from ages 19 to 74. The lines within each 5-year cohort represent longitudinal change in endorsement of benevolent sexism and hostile sexism estimating latent intercepts (i) and slopes (s). * $p < .05$.

suggested that endorsement of benevolent sexism followed a U-shaped trajectory across the lifespan. Women's endorsement of benevolent sexism was high in early adulthood, low in middle adulthood, and relatively high again in older ages.

Endorsement of hostile sexism followed U-shaped trajectories across age for both men and women. The trajectory of hostile sexism was initially high in early adulthood but decreased into middle adulthood, at which point the trajectory recovered and increased into late adulthood (these results were similar when estimating benevolent sexism and hostile sexism simultaneously as latent variables at Wave 1; see OSM).

The Association between Benevolent Sexism and Hostile Sexism across Age

Model Estimation. We estimated whether the association between benevolent sexism and hostile sexism differed across age using data from the first wave, in which hostile sexism was predicted by benevolent sexism (modeled as latent scores), moderated by gender and age (modeled as latent interactions), estimated at every age from 18 to 80. Parameters were fully standardized based on standard deviations of the predictor and the outcome, so that the standardized effect of benevolent sexism on hostile sexism represented the overall relationship between the two constructs, and any significant interactions indicated that this relationship varied by age or gender.

Results. The significant standardized effect of benevolent sexism displayed in Table 5 indicated that, holding age and gender constant, benevolent sexism and hostile sexism were strongly related. The effect of gender indicated that the relationship between benevolent sexism and hostile sexism was generally stronger for women ($\beta = .528$, 95% CI = .482, .575, $p < .001$) than for men ($\beta = .496$, 95% CI = .434, .559, $p < .001$). However, for men and women equally, the significant negative effect of age indicated that the association between benevolent sexism and hostile sexism decreased from age 18 to 80 years old (Figure 2). The average association between benevolent sexism and hostile sexism was strongest at 18 years

old ($\beta = .600$, 95% CI = .530, .670), weakened linearly as age increased, and was lowest at 80 years old ($\beta = .452$, 95% CI = .378, .526). These results indicated that the typically strong link between benevolent sexism and hostile sexism is not uniform across age: Ambivalent attitudes were strongest in young adulthood and relatively weakest, although still strongly related, in older ages.

Table 5. Standardized estimates of the relationship between latent benevolent sexism and latent hostile sexism, statistically adjusting for the random effects of gender, age and the interaction between gender and age.

	β	<i>se</i>	<i>t</i>	<i>p</i>	95% Confidence Interval	
					<i>Low</i>	<i>High</i>
<i>Benevolent Sexism</i>	.528	.024	22.175	< .001	.482	.575
<i>Gender</i>	-.032	.015	-2.123	.034	-.062	-.002
<i>Age</i>	-.038	.014	-2.659	.008	-.065	-.010
<i>Age*Gender</i>	-.002	.003	-0.521	.602	-.007	.004

Note. Gender was coded -.5 = *Women*, .5 = *Men*. Model Fit: RMSEA = .072 [.069, .075]; CFI = .883; SRMR = .053. *N* = 6466. Model fit estimates did not include the random-effect interaction term.

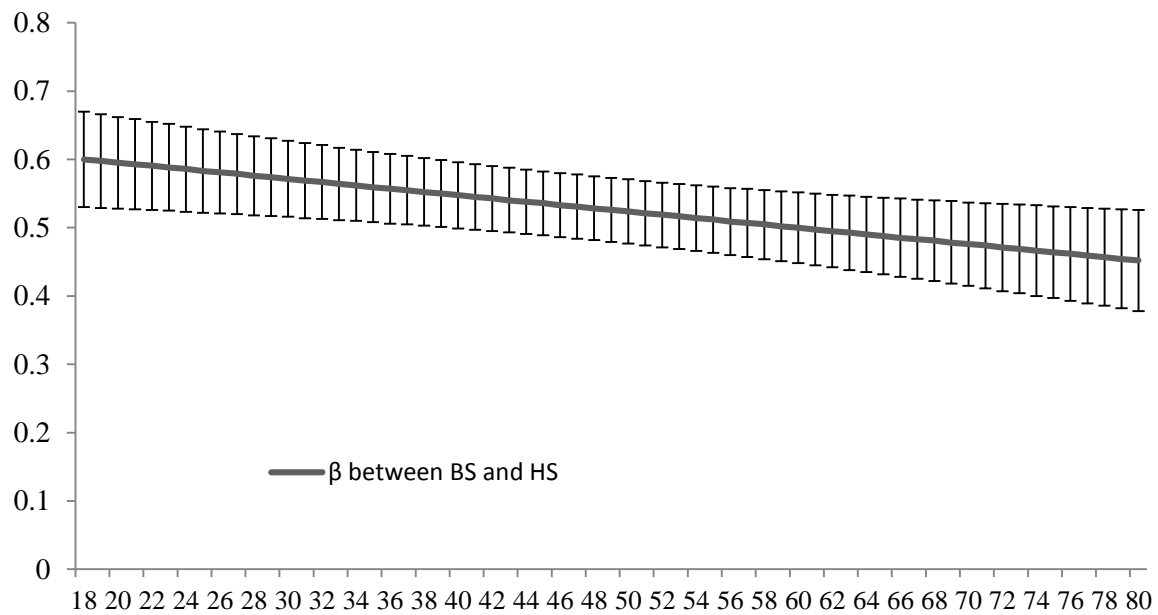


Figure 2. The association between benevolent sexism and hostile sexism from age 18 to 80, modelled as the latent standardized coefficient at measurement Wave 1, statistically adjusting for gender. Error bars represent 95% confidence intervals plotted around stability point estimates.

Longitudinal Change in Sexist Attitudes within Cohorts

Model Estimation. We conducted multi-group cohort-sequential latent growth models by organizing the sample into 12 sequential 5-year birth cohorts (see Table 1) to match the first five waves of annual assessment (see Lucas & Donnellan, 2011). Models utilized the time of survey completion as individually-varying time indicators. We estimated a latent intercept and latent linear slope for each of the 5-year age cohorts. Variances and covariance of intercepts and slopes were constrained to equality across birth cohorts.

We organized birth cohorts into consecutive 5-year bands to match the consecutive 5-year spans of the multi-group models, meaning we could arrange the latent growth models in sequence (e.g., the 1986–1990 birth-cohort represented change from 19 to 24 years-old; the 1981–1985 cohort represented change from 24 to 29 years old, and so on). For each cohort, the youngest age was taken as an indicator of age. This approach simultaneously estimates cross-sectional cohort differences in the latent intercepts, change trajectories within each cohort, and cohort differences in the rate of change, as well as the overall pattern of change that may be observed across the adult life span. We approximated effect sizes by calculating the change as a percentage of the entire scale range (*SR%*; see OSM for individual cohorts).

Results. Figure 1 displays the multi-group cohort-sequential latent growth models. Endorsement of benevolent sexism declined across time for men in the youngest cohort (*SR%* = 0.8%). As age increased, the steepness of the declines in benevolent sexism decelerated and benevolent sexism was relatively stable across all remaining cohorts in the lifespan, except for a decline in the 49- to 54-year-old cohort (*SR%* = 0.5%). Men's endorsement of hostile sexism also decreased across time most sharply in the youngest cohorts (*SR%* = 1.5% to 2%), but longitudinal declines in endorsement of hostile sexism were present in most cohorts (*SR%* = 0.6% to 0.9%).

For women, endorsement of benevolent sexism declined over time in the age cohorts from 19 to 34, 39 to 44, and 49 to 69 ($SR\% = 0.3\%$ to 1.6%), and was otherwise stable. Women's endorsement of benevolent sexism was high in young adulthood but decreased over time, with the rate of decrease decelerating into middle adulthood. Women's endorsement of hostile sexism also declined across time in the young-adult cohorts ($SR\% = 0.8\%$ to 1.2%). In older cohorts, these declines decelerated but the decreases across time remained significant up to the 54 to 59 cohort. In older cohorts, there was no significant change across time, and a significant increase across time in the 69 to 74 cohort ($SR\% = 0.5\%$).

In sum, endorsement of sexism tended to decrease most prominently in early adulthood. However, unlike the age-based trajectory model, the longitudinal cohort model indicated that endorsement of sexist attitudes tended to decline for most of the age cohorts. This inconsistency indicated the presence of cohort effects which we tested next.

Estimation of Cohort Effects

Model Estimation. We examined indicators for the presence of cohort effects by running models that compared estimates from the age-based trajectory model with the cohort-based longitudinal model (also see Milojev & Sibley, 2016). We re-analysed the cohort-based longitudinal model adding (1) constraints that equalized the latent intercepts and latent slopes across the cohorts, and (2) constraints that *further* equalized these effects across gender. We then compared the fit of these different models. If the constrained models have substantially worse fit, then this would indicate (1) strong cohort effects, and/or (2) between-gender differences in cohort effects. Second, we examined any cohort effects more closely by estimating the discrepancy between the mean levels of sexist attitudes based on the five-year latent change trajectory with the mean levels of sexist attitudes based on participants' age at first measurement. This discrepancy identifies ages at which there are the differences

between the model-implied values that are due to change over time and model-implied values that are due to between-cohort differences.

Results. First, fit estimates for the free vs. constrained cohort-based models of sexist attitudes are displayed in Table 6. When effects were free to vary across gender but were constrained across age cohorts, there was very little difference in the fit of the model. This indicated that cohort effects were relatively weak overall. However, model fit worsened when effects were also constrained across gender, indicating relatively greater between-cohort variability for men than for women.

Second, estimation of cohort effects are displayed in Figure 3. For men's endorsement of benevolent sexism, significant differences (where the error bars do not overlap zero) emerged in ages 49, 59, and 64. These findings indicated only few cases of inconsistency between the age-group trajectories and the linear trajectory across age. In contrast, for men's hostile sexism, cohort effects were observed for ages 24, 34 to 49, 59, and 64, indicating relatively consistent cohort effects across age in which the within-cohort trends of hostile sexism decreased relative to the U-shape of the age-trajectory model.

Figure 3 displays significant cohort effects for women's endorsement of benevolent sexism at ages 24, 34, 39, and from 54 to 69. These effects indicated that women's benevolent sexism decreased more in early ages (vs. a shallower decline in the age-trajectory model) and was relatively stable in late age (vs. an increase in benevolent sexism in the age-trajectory model). For women's endorsement of hostile sexism, the discrepancies between the cohort-based model and age-based model were less pronounced and restricted to ages 24, 34, 39, and 54. Thus, the longitudinal changes in the cohorts of women's hostile sexism largely matched the U-shaped trajectory across age, except for a relatively steeper decline in hostile sexism in younger age cohorts.

Sexism across Age

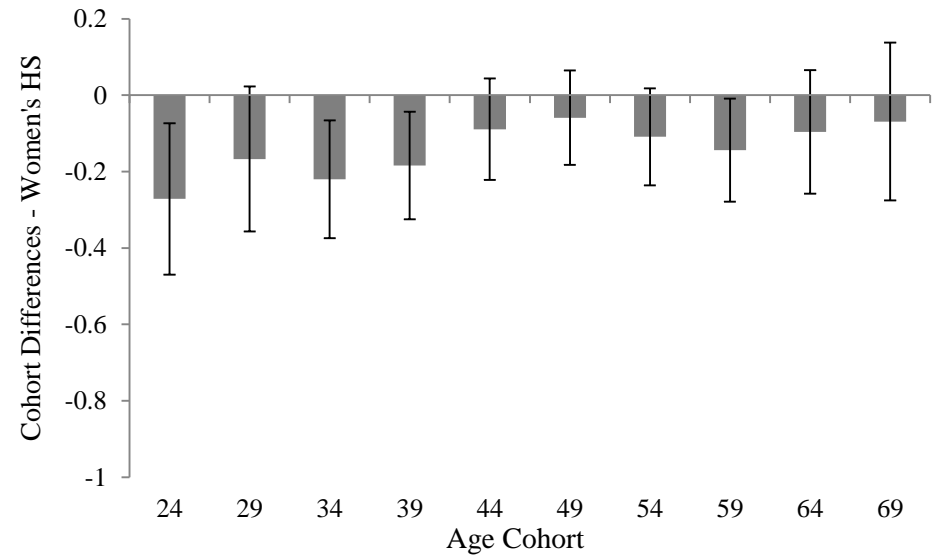
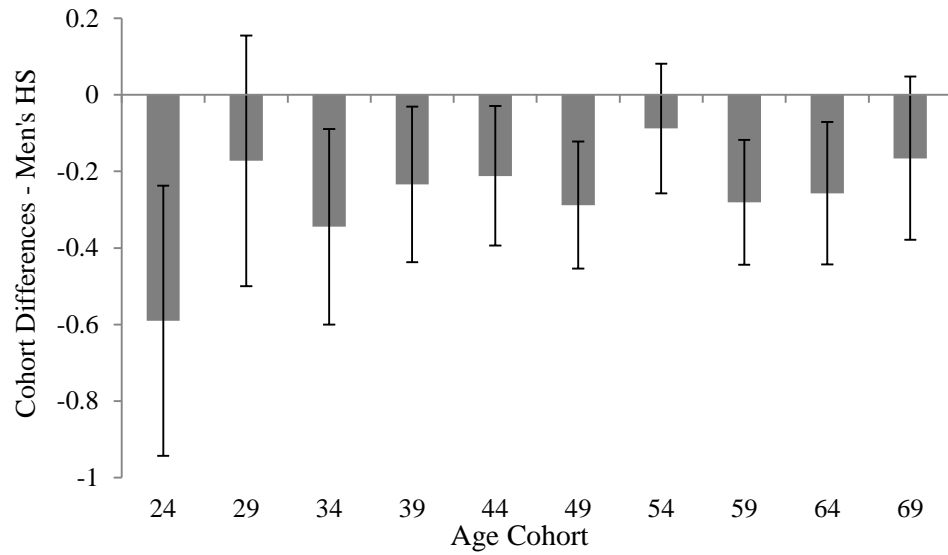
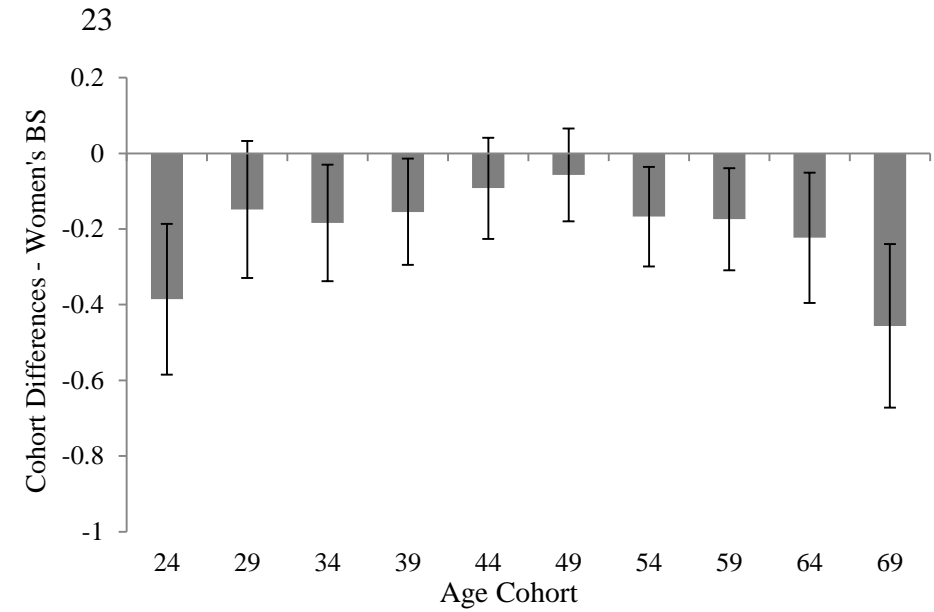
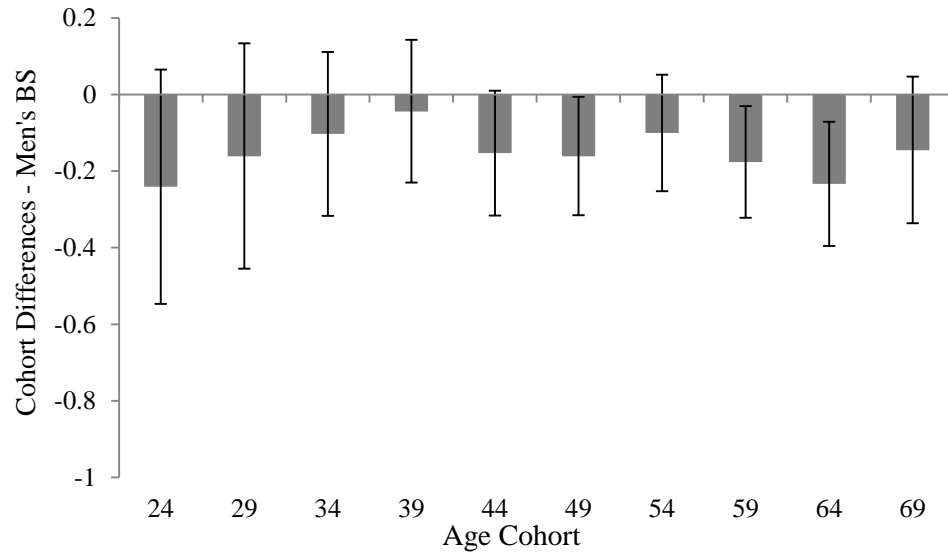


Figure 3. Estimated cohort effects for benevolent sexism (BS) and hostile sexism (HS) obtained via comparing the model-implied estimated values from the age-based models of endorsement of sexist attitudes with values from the five-year latent change trajectory models. Negative effects indicate that model-estimated values were lower at a given age for the younger cohort (i.e., born in a *later* year) relative to the subsequent cohort.

Table 6. Model Fit indices comparing the Multi-Group Cohort Sequential Growth models allowed for cohort differences (*Free Model*), constrained the latent intercepts and latent slopes to equality across age-cohorts while allowing variation across gender (*Constrained Model-Gender Free*), or imposed equality across gender (*Fully Constrained Model*).

	Free Model	Constrained Model (Gender Free)	Fully Constrained Model
Benevolent Sexism			
AIC	78347.403	78446.805	78600.825
aBIC	78557.196	78479.081	78625.032
Hostile Sexism			
AIC	80461.231	80491.153	81281.206
aBIC	80671.019	80523.428	81305.413

Note. AIC = Akaike Information Criterion; aBIC = Sample-size adjusted Bayesian Information Criterion.

Discussion

We utilized a national panel sample in an egalitarian country to conduct four models of age-related differences in endorsement of benevolent sexism and hostile sexism. Endorsement of sexism was generally stable, but significant differences emerged across time and age. First, endorsement of sexism generally followed U-shaped curvilinear trajectories across age: Trajectories for men's hostile sexism, women's benevolent sexism, and women's hostile sexism were relatively high in late adolescence, lower through middle adulthood, and high again in older ages. In contrast, men's endorsement of benevolent sexism followed a positive linear trajectory across age. Second, the 'ambivalence' of sexist attitudes was consistently present, but weakened slightly, across age: The association between hostile sexism and benevolent sexism was strongest in late adolescence and decreased linearly as age increased. Third, we employed cutting-edge statistical analyses to model the changes in benevolent sexism and hostile sexism over five years. Endorsement of sexism declined over time in 26 of 44 age-based cohorts, declines that were pronounced in the youngest cohorts (ages 19-24 and 25-29), and did not differ over time in the 18 remaining cohorts. Finally, tentative estimation of possible cohort effects indicated the presence of weak differences that were relatively consistent across age, but were most prominent in the youngest cohort.

Two intriguing findings emerged that extend current understanding of ambivalent sexism theory. First, although sexism was high in the youngest cohorts, the youngest cohorts also reported pronounced decreases in endorsement of sexist attitudes *over time*. This is consistent with prior research across age indicating that attitudes toward political and racial issues also tend to be highly susceptible to change in early adulthood, particularly for subjectively relevant issues (Visser & Krosnick, 1998). Finding similar flexibility in sexism during early adulthood is consistent with the perspective that life experiences prompt individuated beliefs about gender (see Glick & Fiske, 1996). Indeed, endorsement of hostile

sexism was less closely related to benevolent sexism as age increased; the ambivalence of attitudes exhibited a small but consistent decline across age, perhaps signaling the development of more idiosyncratic beliefs about gender. The sharp declines in men's hostile sexism across young adulthood were consistent with this expectation; men's hostile sexism is highly incompatible with attainment of relationship goals (Hammond & Overall, 2016) which are a priority in young adulthood (e.g., Baltes et al., 1980; Takahashi, 2005).

Our results also provide a contextual map for understanding benevolent sexism. First, the positive linear trajectory, and relative stability, for *men's* benevolent sexism, was in contrast to the U-shaped trajectory (and longitudinal declines) found for men's hostile sexism and women's ambivalent sexism, as well as traditional values or prejudice more generally (e.g., Gouveia et al., 2015; Visser & Krosnick, 1998). In sum, men's endorsement of benevolent sexism was unexpectedly persistent over time. These results indicate a consistent appeal of benevolent sexism to men at all ages, perhaps because of its romantic appearance (e.g., Barreto & Ellemers, 2005) or the ways that benevolent sexism subtly maintains men's power (e.g., Hammond & Overall, 2015; Hammond, Overall, & Cross, 2016; Shnabel, Bar-Anan, Kende, Bareket & Lazar, 2016).

We also provide context for effects linked with *women's* endorsement of benevolent sexism. Women's benevolent sexism is relatively higher in young adulthood than mid-adulthood. Accordingly, the orientation toward relationship-oriented goals (rather than career-oriented goals) that accompanies women's benevolent sexism (e.g., Chen et al., 2009; Montañés et al., 2013) are relatively higher at a life-stage when women are making decisions about higher education, careers and establishing serious relationships.

Caveats and Future Directions

Results from our models of the trajectories of sexism across age were consistent with age-group differences found in Spanish samples (see Fernández et al., 2004; Gariagordobil &

Aliri, 2013), and our models approximating cohort effects did not reveal large differences. Nonetheless, the current models cannot speak to the mechanisms underlying the links between sexism and age, completely distinguish age-related differences from cohort effects, or test potential moderators without exceeding computational limits. Our current findings do however provide a map for investigating the moderators of lifespan changes in sexism. For example, endorsement of sexism tended to be relatively low in middle-aged cohorts, perhaps reflecting transitioning gender-role expectations as men take on parenting or alloparenting roles (e.g., Bedford & Turner, 2006). The results also point toward potential *historical* moderators. Analyses indicated the presence of small cohort effects for ages 59-64 and 65-69 (but not consistently for adjacent cohorts). These cohorts would have developed through adolescence to adulthood in a normative context of egalitarianism that occurred post-World War II (see Mason, Czajka, & Arber, 1976), a theorized critical age for people's receptivity to sexist attitudes (e.g., Glick & Hilt, 2000). A good direction for future research would be to test these potential personal-, historical- and societal-level moderators.

Finally, data were collected in New Zealand, a country in which endorsement of sexism and objective indices of gender inequality are relatively low compared to other countries (see Glick et al., 2000; United Nations Development Programme, 2016). In less egalitarian countries, individualistic motivations to endorse sexism (that would correspond to age-graded developmental changes) may be overshadowed by societal pressures to endorse sexism. For example, people may endorse benevolent sexism as a means of attaining security in response to threat (Fischer, 2006; Glick et al., 2000) or because of societal pressure to uphold traditional beliefs and customs (see Schwartz & Rubel-Lifschitz, 2009), rather than to justify personally-held hostile beliefs toward women (Glick et al., 2000). Thus, in less egalitarian contexts, it is possible that the trajectories of individuals' sexist attitudes are more closely related to shifts in societal and cultural norms.

Conclusion

We examined differences in sexism across age in a nationally representative sample. The trajectories of women's benevolent sexism, women's hostile sexism, and men's hostile sexism were relatively high in early adulthood, lower in middle adulthood, and high in late adulthood. In addition, longitudinal models indicated that endorsement of hostile sexism and women's benevolent sexism tended to decline in over half of the age cohorts, and did not increase over time in any age cohort, suggesting that people are generally becoming more egalitarian over time. However, men's benevolent sexism followed a positive linear trajectory across age and was also relatively stable over time, highlighting the resilience of men's subjectively romantic sexist attitudes amongst the trend toward egalitarianism. These results provide the first map of differences in the trajectory of sexism over the lifespan and the development of sexism over time.

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