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Looks and gaming: Who and why?

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A B S T R A C T

Americans spend 2.5 percent of their waking hours video-gaming. Using the American Add Health Study, we show that adults who are better-looking have more close friends. Gaming being costlier for them, they engage in less of it. Physically attractive teens are less likely than others to game at all. Attractive adults are less likely than others to spend any time gaming; if they do, they spend less time on it than other gamers. The reverse is true among teens and adults for some other non-market activities—sports and hobby groups. Using the longitudinal nature of the Study, we find that these relationships may be causal for adults: good looks decrease gaming time, not vice-versa. The results provide new evidence on how looks affect non-market time use and perhaps indicate the role that they play in personal development.

1. Introduction and motivation

A now immense literature has examined the impact of human beauty on labor market outcomes (Deryugina and Shurchkov, 2015; Hamermesh and Biddle, 1994; Harper, 2000; Scholz and Sicinski, 2015); well-being (Hamermesh and Abrevaya, 2013); criminal and other risky activities (Chung and Zhang, forthcoming; Green et al., 2023; Mocan and Tekin, 2010); and success in eSports (Babin et al., 2024). In terms of pre-market impacts, several studies (Babin *et al.*, 2020; Hamermesh and Parker, 2005; Mehic, 2022), have shown effects on the education process and its outcomes. Even longevity has been shown to be affected by physical attractiveness (Sheehan and Hamermesh, 2024).

Some effort has been spent to determine the sources of these effects of beauty (Möbius and Rosenblat, 2006; Stinebrickner et al., 2019). What has not been examined formally is how beauty affects our daily activities—how people of different physical attractiveness spend their time. Due to the absence of information on respondents' looks in any dataset worldwide that describes time use generally, this has not been possible. We can, however, examine how attractiveness affects the time devoted to several related activities. This study thus goes

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behind the question of “What” differences in looks do to economic outcomes to examine “How” those outcomes might result.

We examine how looks affect time engaged in video-gaming, an activity which, as we show, occupies substantial percentages of young people and even of adults. The worldwide popularity of vide-gaming cannot be overstated. As a recent example, within just three days in 2023, Nintendo sold 10 million copies of *The Legend of Zelda: Tears of the Kingdom* (BBC News, 2023; Tassi, 2023). Some reviews of this popular new game belabored the point that today’s gamers are not only teenage boys but also adult men and women, who likely became *Zelda* fans and gamers during their formative years.[1]

Despite the ubiquity and popularity of gaming as a leisure activity, several stereotypes persist about the average gamer. For instance, the sociology literature consistently documents that being physically unattractive is ascribed to gamers as a norm (e.g., Kowert *et al*., 2012; Stone, 2019). In the economics literature, early studies on video-gaming predominantly centered on its impact on crime or violent behavior (Cunningham *et al*., 2016; Ward 2010, Ward, 2011). Others have considered how screen time generally affects teens, focusing on class attendance and time spent on homework (Ward, 2018) or their long-term development (Jürges and Khanam, 2021; Malamud and Pop-Eleches, 2011; Orben and Przybylski, 2019). Taking a different approach, Aguiar *et al*. (2021) fit the choice to game within a standard theoretical framework that features trade-offs, both contemporaneous and dynamic, among labor supply, gaming, and other leisure or productive activities.

Especially relevant to our research, some recent studies found that better-looking people have an advantage in social networks (e.g., O’Connor and Gladstone, 2018). For instance, using a school-based survey conducted in China, Zhai *et al*. (2022) found that children with a more attractive physical appearance are more popular in their friendship groups.

Building upon this literature, we propose a simple economic hypothesis. Video-gaming involves very few, if any, face-to-face interactions. Given that physical attractiveness confers advantages in face-to-face interactions in social or leisure activities, individuals deemed more physically attractive will face a higher opportunity cost of engaging in video-gaming. Consequently, we hypothesize a negative relationship between beauty and gaming time, with more attractive individuals likely to spend less time gaming. In other words, good-looking gamers will be relatively scarce because of its higher cost.

Although our empirical analyses focus specifically on estimating the effect of physical attractiveness on gaming time, our hypothesis has broader implications. More attractive individuals face higher opportunity costs when engaging in activities that lack face-to-face interaction, suggesting that beauty should negatively influence time spent in such activities, an extension that we examine briefly. To our knowledge, this is the first study to propose and test this hypothesis. Accordingly, we define gaming broadly to include not only traditional console gaming but also computer gaming and other leisure activities on computers, where face-to-face interactions are generally limited.[2]

We test this hypothesis first using teenagers from Wave I of the Add Health dataset. Our analysis reveals supportive evidence that is robust to a variety of alternative specifications of the measure of beauty and to different sets of covariates. The evidence holds, however, only for the extensive margin of time spent gaming by teens. We then examine the same hypothesis among adults and observe that those deemed more physically attractive are less likely to spend time gaming, and if they do game, they dedicate less time to it.

We also explore whether the estimated relationships between gaming and beauty can be interpreted causally. One concern might be that gaming affects beauty rather than vice-versa: Perhaps gaming time replaces activities that can bolster physical attractiveness, such as grooming and regular exercise. We exploit the longitudinal dimension of the Add Health study and find that our estimates of beauty’s effect on gaming time are robust to accounting for this concern.

Section 2 introduces the Add Health dataset and describes how we estimate the relationship between beauty and video-gaming among teenagers and among adults. Section 3 presents and discusses the baseline results for teens, and Section 4 presents the results for adults. Section 5 examines robustness tests, considers impacts on a few other non-market uses of time, and proposes a causal interpretation.

2. Data and empirical specification

To investigate the impact of beauty on time spent video-gaming, we employ the public-use sample of the Add Health study, a widely used longitudinal American dataset (UNC Carolina Population Center. n.d.). Add Health comprises a representative sample of American adolescents spanning grades 7 through 12 (generally ages 12–18) during the 1994–95 school year, with five follow-up waves, the most recent collected between 2022 and 2025. We focus on Wave I to examine teenage behavior and Wave IV, conducted in 2008, when respondents had reached adulthood (generally ages 26–32).[3]

In these waves, at the end of each interview, the field interviewer rated the physical attractiveness of the respondent according to

[1] *Tears of the Kingdom* is not the only modern game that appeals to different age groups and genders. As many as four in ten people worldwide were video gamers in 2022 (The Economist, 2023). An estimated 212 million people were gamers in the United States that year, with 48 percent being female and only 24 percent below age 18 (Clement, 2022-23). In the U.K., 55 percent and 56 percent of adult men and women, respectively, play some form of video games (Clement, 2022-23).

[2] Our definition of gaming can also include mobile phone gaming. However, as our dataset was collected before mobile gaming became widespread, including mobile gaming in our definition does not affect our analysis.

[3] We do not use data from Wave V, collected between 2016 and 2018, as that Wave lacks pertinent information on physical attractiveness.

the following question:

> *How physically attractive is the respondent?* with the following options: 1. "very unattractive", 2. "unattractive", 3. "about average", 4. "attractive", and 5. "very attractive".[4]

Figs. 1a and 1b depict the distributions of interviewers' responses describing the male and female teenage respondents, while Figs. 2a and 2b show the distributions for adult respondents. These distributions of looks among teenagers and among adults are quite similar.[5] As observed in previous studies (Mocan and Tekin, 2010; Stinebrickner *et al.*, 2019), females have a slightly higher average attractiveness score and standard deviation than males.

We combine the two sparse responses, "very unattractive" and "unattractive" into one category, constructing a 0–1 indicator. Similarly, we define another 0–1 indicator, "attractive or very attractive," based on those responses, so that "about average" is the excluded category for both indicators. Table 1 provides summary statistics describing these two variables, with a smaller fraction of adults rated as attractive/very attractive and a slight larger fraction rated unattractive/very unattractive.

Since our hypothesis stems from the notion of a beauty advantage in social networking, as described in the sociological literature, we examine the presence of this advantage among the Add Health study members using responses to the following question, asked of adults (only in Wave IV):

> *How many close friends do you have? (Close friends include people whom you feel at ease with, can talk to about private matters, and can call on for help.)* with categories as: 1. "none", 2. "1 or 2 friends", 3. "3 to 5 friends", 4. "6 to 9 friends", and 5. "10 or more friends".

(Regrettably this question on the number of close friends was not asked of teenagers.) The summary statistics in Table 1 show that "3 to 5 close friends" is both the median and the modal response, although 13 percent of the respondents say that they have 10 or more close friends. The response "no close friends" is rare, at 3 percent.

We first regress an indicator of "no close friends" on the two indicators of perceived physical attractiveness, with the results presented in the first three columns of Table 2. The estimated effects on the absence of close friendships among the 44 percent of adult respondents deemed "attractive or very attractive" are consistently negative and significant, at least at the 5 percent level, compared, to the 48 percent of respondents with average looks. There is no significant evidence that being among the 8 percent of adults rated as "unattractive or very unattractive" increases the chances of having no close friends, compared to being "about average" in looks, although the coefficient estimate is positive as expected.

Columns (4)-(6) of Table 2 report OLS estimates describing the relationship between the number of close friends claimed and the two beauty indicators, with the dependent variable taking the midpoints of each classification as its values, and with the top-coded classification set equal to 12.[6] Attractive/very attractive adults have about 0.4 more close friends (on a mean of 4.9) than those of average attractiveness. Conversely, the small fraction of adults whose looks are rated unattractive/very unattractive claim about 0.7 fewer close friends than those of average attractiveness. The gap in the average number of close friends between the attractive and unattractive adults is thus 22 percent of the overall mean. Taken together, these findings imply that individuals perceived as physically attractive are likely to have more close friends, which aligns with the documented beauty advantage in the social networking literature and provides the basis for our economic hypothesis on the impact of looks on gaming time.[7]

To test our central hypothesis, we consider the time spent video-gaming based on the interviewees' responses to the following questions :[8]

> *How many hours a weekday do you play video or computer games [which we convert to hours per week]?* (Wave I – teenagers)
>
> *In the past seven days, how many hours did you spend playing video or computer games, or using a computer? Do not count internet use for work or school.* (Wave IV – adults)

The responses were coded as 0 for non-gamers, positive integers for those reporting some gaming.

Data from Waves I and IV were collected before the widespread popularity of mobile phone gaming, with internet use still limited in 1994–1995 and smartphones being relatively rare in 2008. Thus, both data points are consistent with our definition of gaming. The Wave IV question suggests a broad definition of gaming time for adults, encompassing not only traditional console gaming but also home computer gaming and general leisure activities on computers. These provided only very limited face-to-face interactions, since

[4] One might be concerned that interviewers' ratings have been affected by the interviewees' responses during a face-to-face interview. In the German ALLBUS surveys (see Hamermesh and Abrevaya, 2013), in which interviewers rated subjects' looks at both the start and end of the interview, there is an extremely high correlation between the two ratings, one that is unaffected by any of the observables in the survey.

[5] For example, Sheehan and Hamermesh (2024, Supplemental Table S1) demonstrated the very strong correlation of beauty ratings across waves in the Add Health dataset.

[6] The appropriate ordered probit yields qualitatively similar results to those shown in Table 2. Indeed, the ordered probits yield implicit movements between the cut points that match very well the linear estimates.

[7] If we replace the indicators of attractiveness as an adult with those as teenagers, then the results of these regressions are similar. The results are available upon request from the corresponding author.

[8] Calculations using data from the American Time Use Survey (ATUS) show that the average daily hours spent on playing games and leisure computer use by individuals aged 25-34 is 0.4 hours, equivalent to 2.8 hours per week in 2008 (U.S. Bureau of Labor Statistics, n.d.). This figure aligns with our sample average of 3.03 hours per week (0.405 x 7.483).

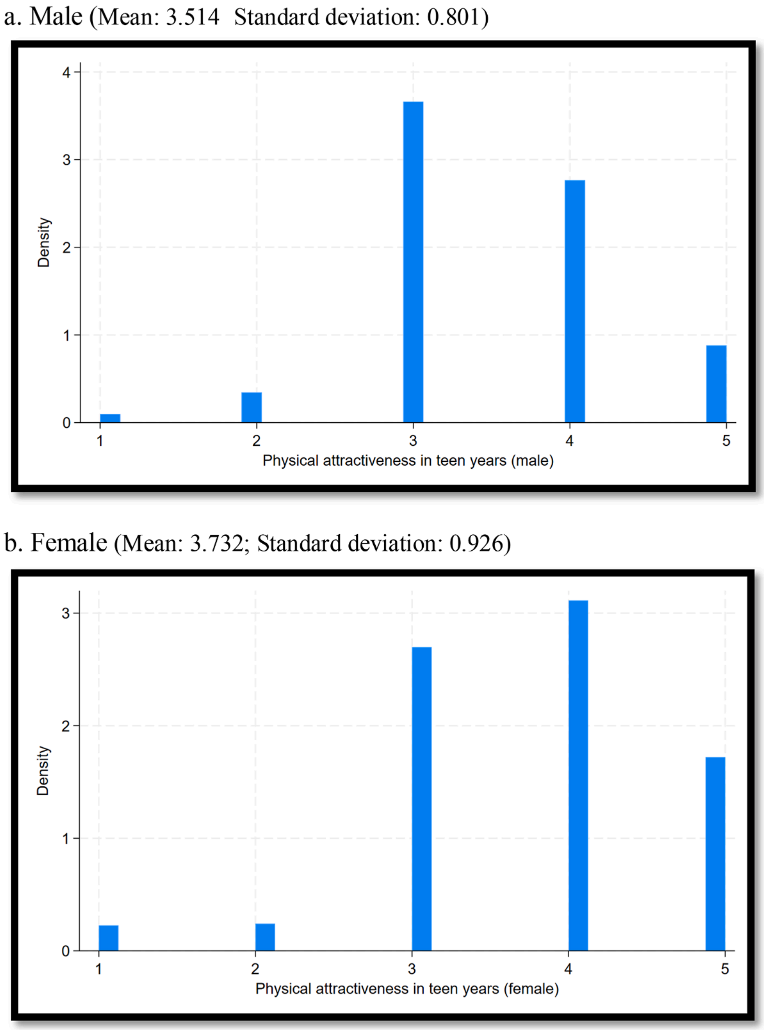


Fig. 1. Distributions of physical attractiveness ratings in teen years. a. Male (Mean: 3.514 Standard deviation: 0.801). b. Female (Mean: 3.732; Standard deviation: 0.926).

that period was more the era of the now defunct MSN Messenger and Bebo, not today's more dynamic platforms like TikTok and Twitch.

As the descriptive statistics in Table 1 show, slightly more than half of teens engaged in gaming in 1994/5, with gamers averaging 5 h per week in the activity. The incidence of gaming was lower in adulthood in 2008; however, the 40 percent of adults who did game spent more time doing this, 7.5 h per week (nearly 7 percent of their likely waking hours).

Figs. 3a and 3b show histograms of time spent gaming by gender among teenagers, while Figs. 4a and 4b show the distributions among adults. The means and variances are lower among girls and women than boys and men. Fig. 3a and, especially, Fig. 4a show that there are small minorities of extreme male gamers.

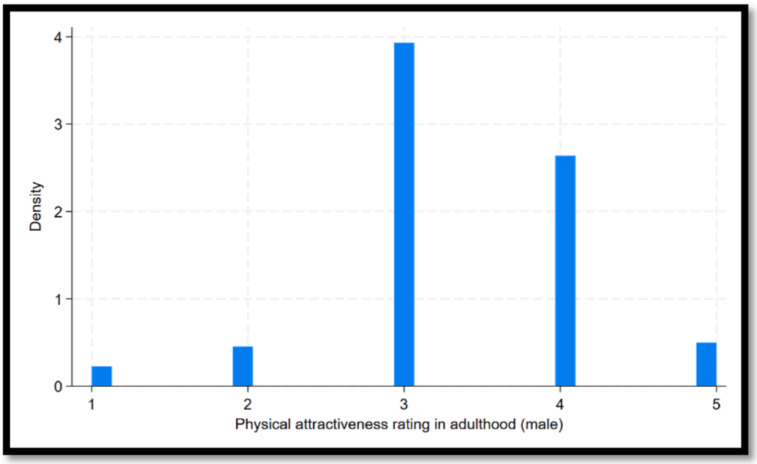
Because of the large concentrations at zero of the time spent gaming, we examine the determinants of the extensive and intensive margins of gaming time separately. In models describing the extensive margin, we estimate the relationship between physical attractiveness and whether the teen (adult) games at all; in those describing the intensive margin, we estimate the conditional relationship between looks and gaming time among gamers.

The Add Health dataset has additional measures of non-market time use that might be related to respondents' attractiveness in ways like time spent video-gaming. Among teenagers (Wave I) information is available on a yes/no basis on whether the respondent is a member of a hobby or sports group. Wave IV provides information on the number of times that the respondent participated in team sports in the previous week. It also contains a similar question about participation in individual sports. As a check on our results on gaming, we also examine how looks relate to time spent in these activities.

All the estimates in the next two sections are based on variants of the following model:

$$G_i = \alpha + \beta_1 A_i + \beta_2 U_i + \boldsymbol{\theta} \boldsymbol{X}_i + \gamma_{v(i)} + \varepsilon_i \tag{1}$$

a. Male (Mean: 3.352; Standard deviation: 0.806)



b. Female (Mean: 3.456; Standard deviation: 0.895)

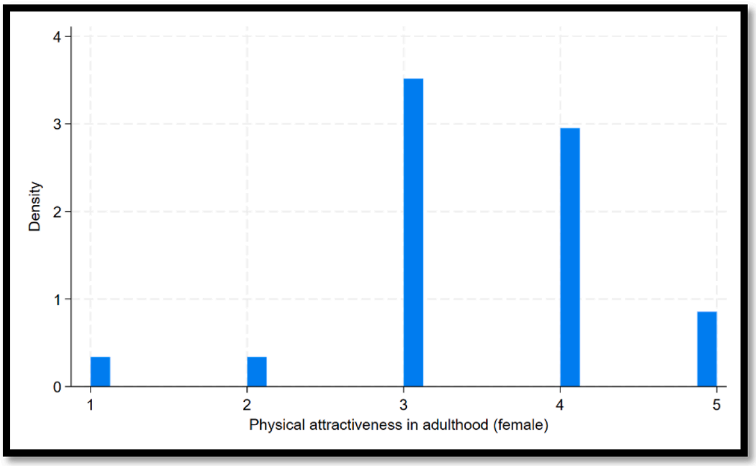


Fig. 2. Distributions of physical attractiveness ratings in adulthood. a. Male (Mean: 3.352; Standard deviation: 0.806). b. Female (Mean: 3.456; Standard deviation: 0.895).

Table 1
Descriptive statistics of key variables—add health waves I and IV.

	Obs	Mean	Std. dev.	Min.	Max.
Attractive or very attractive teen	3229	0.521			
Unattractive or very unattractive teen	3229	0.069			
Attractive or very attractive adult	3228	0.440			
Unattractive or very unattractive adult	3228	0.084			
Close friends as adult: 0	3202	0.029			
Close friends as adult: 1–2	3202	0.213			
Close friends as adult: 3–5	3202	0.463			
Close friends as adult: 6–9	3202	0.170			
Close friends as adult: 10 or more	3202	0.125			
Video-gaming last week? teen	3228	0.544			
Hours gaming last week: teen	1726	5.283	8.779	1	99
Video-gaming last week? adult	3147	0.405			
Hours gaming last week: adult	1236	7.481	10.395	1	105

Notes: Authors' calculations from Add Health. The calculations here and in all tables use sampling weights.

Table 2
The effects of physical attractiveness on having close friends in adulthood.

Dep. variable:	No close friends, {Yes,No}={1,0}			Number of close friends		
	All	Male	Female	All	Male	Female
Ind. variable:	(1)	(2)	(3)	(4)	(5)	(6)
Attractive/very attractive	−0.023	−0.025	−0.019	0.433	0.639	0.298
	(0.007)	(0.011)	(0.009)	(0.144)	(0.227)	(0.181)
Unattractive/very unattractive	0.011	−0.012	0.034	−0.654	−0.364	−0.920
	(0.015)	(0.018)	(0.024)	(0.232)	(0.365)	(0.281)
Adj. R^2	0.005	0.003	0.008	0.008	0.009	0.010
N	3199	1416	1783	3199	1416	1783

Notes: A constant is contained in all specifications. The excluded category is "about average" attractiveness. Robust standard errors, clustered on interviewers, are in parentheses.

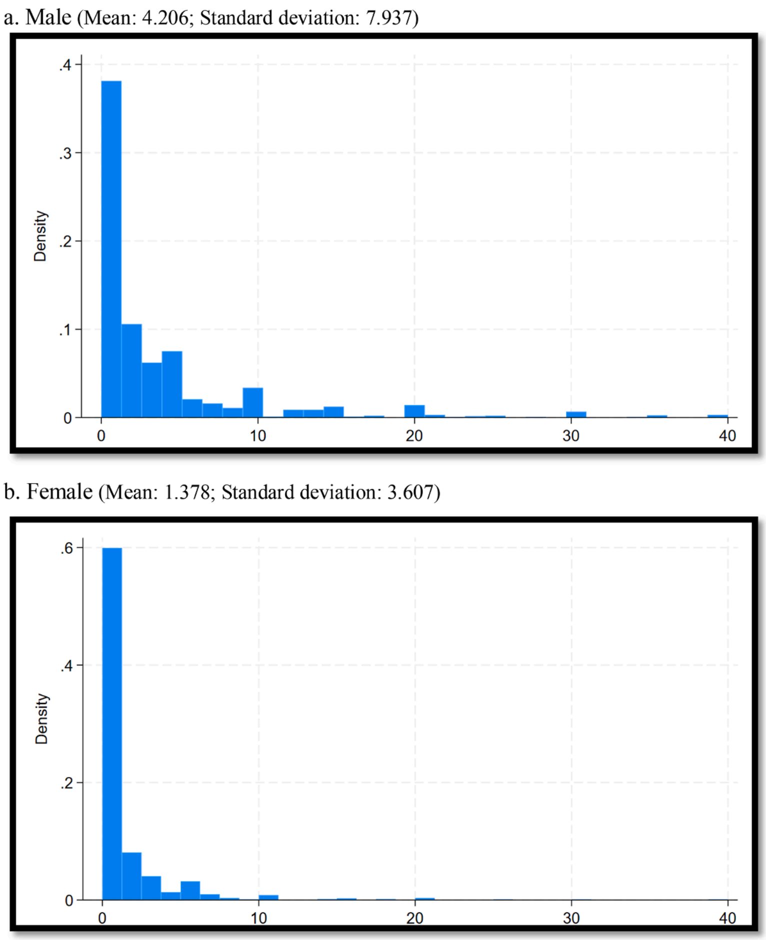
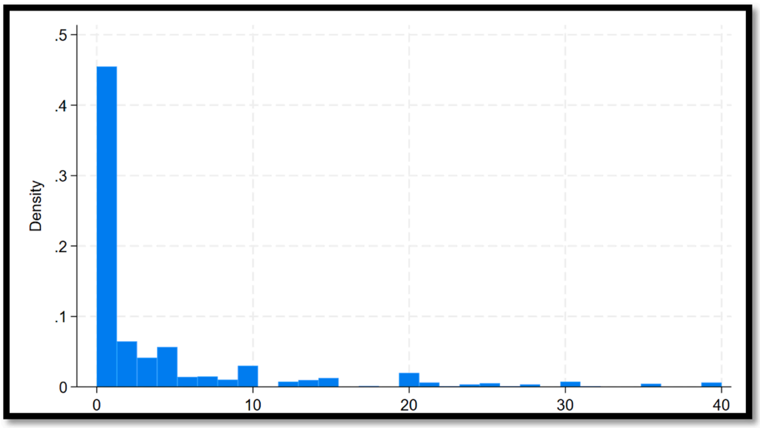


Fig. 3. Distributions of gaming hours in the previous week among teenagers. a. Male (Mean: 4.206; Standard deviation: 7.937). b. Female (Mean: 1.378; Standard deviation: 3.607).

where G_i measures gaming activity (any gaming, or the number of hours per week if any, as a teenager or an adult). A_i and U_i are the beauty indicators, "attractive or very attractive" and "unattractive or very unattractive". $\boldsymbol{X}_i$ is a set of control variables, with the row vector of coefficients $\boldsymbol{\theta}$. Descriptive statistics (not presented in the Table) indicate that our estimation sample is generally representative of the Add Health dataset and thus of the U.S. population in their age cohorts. In any case, we use the Add Health cross-sectional sampling weights in all the within-wave calculations and model estimation, as well as the longitudinal weights when looking at individuals across the teenage and adult waves.

The parameters of interest in (1), β_1 and β_2, represent either the average effects on the probability of gaming or the conditional number of hours spent gaming compared to people whose looks are "about average." Broadly speaking, our hypothesis implies testing

a. Male (Mean: 4.589; Standard deviation: 9.935)



b. Female (Mean: 1.793l Standard deviation: 5.584)

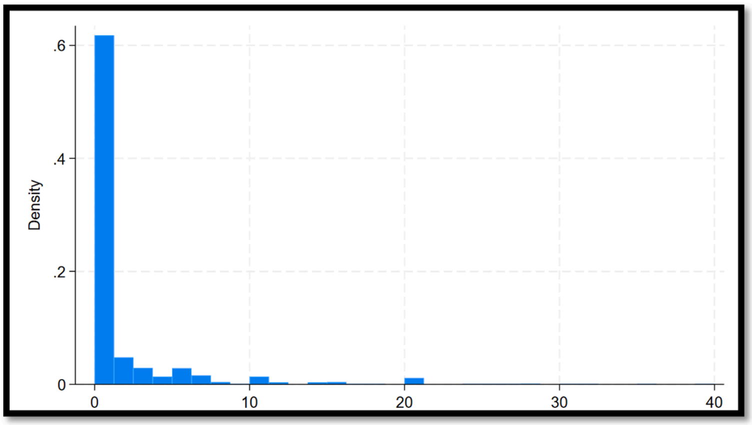


Fig. 4. Distributions of gaming hours in the previous week among adults. a. Male (Mean: 4.589; Standard deviation: 9.935). b. Female (Mean: 1.793l Standard deviation: 5.584).

that $\beta_1 < 0$, and/or $\beta_2 > 0$. Like other studies (e.g., Green *et al.*, 2023; Hamermesh and Biddle, 1994; Sheehan and Hamermesh, 2024), tests of our hypothesis do not impose symmetric effects of being attractive or unattractive. We thus test for the symmetry of the estimated effects by examining the restriction $\beta_1 + \beta_2 = 0$.

In all our estimates, we adjust for interviewer-specific fixed effects, $\gamma_{v(i)}$, to account for the possibility that some interviewers are more generous than others in their ratings of the Add Health respondents' looks. To the extent that interviewers' relative leniency/harshness is correlated with the incidence and/or extent of gaming time, adding these fixed effects obviates this potential source of bias in the estimates of β_1 and β_2. There were 390 interviewers of the adolescents and 263 of the adults. We have no information on possible overlaps in these two sets. We also estimate standard errors that are robust to interviewer-level clusters, which for the teenage model is approximately equivalent to school-level clustering, given the design of the Add Health study.[9]

3. Beauty and teen gaming

Table 3 presents linear probability estimates of (1) for the determinants of the incidence of teen gaming. In Column (1), we regress the gaming indicator on the two indicators of looks, gender and some basic controls, including whether the respondent was born in the U.S., age, and birth weight. Consistent with our hypothesis, the estimated coefficient for being "attractive or very attractive as a teen" is negative and statistically significant (at the 10 percent level), while the coefficient for "unattractive or very unattractive as a teen" is negative but essentially zero statistically. The effects of looks are not negligible: The difference in the incidence of gaming between attractive and average teens is 4.4 percentage points.

[9] We do not include person fixed effects due to the limited variation in beauty from adolescence onward (Sheehan and Hamermesh 2024, Supplemental Table S1), as doing so would likely absorb any effect of appearance. Our current approach, however, allows us to distinguish the differential impact of looks on gaming incidence and intensity between teens and adults.

Table 3
Estimated effects for teenagers of physical attractiveness on the extensive margin of hours playing video/computer games.

	Dep. variable: Gaming, {Yes,No}={1,0}			
Ind. variable:	(1)	(2)	(3)	(4)
Attractive/very attractive (β_1)	−0.044	−0.046	−0.046	−0.050
	(0.024)	(0.023)	(0.023)	(0.031)
Unattractive/very unattractive (β_2)	−0.015	−0.011	−0.011	0.005
	(0.054)	(0.053)	(0.054)	(0.074)
Female	−0.309	−0.306	−0.307	−0.309
	(0.023)	(0.023)	(0.024)	(0.035)
Attractive/very attractive × female (β_3)				0.007
				(0.043)
Unattractive/very unattractive × female (β_4)				−0.035
				(0.095)
Suspended from school			−0.018	−0.018
			(0.027)	(0.027)
Expelled from school			0.059	0.059
			(0.063)	(0.063)
p-value on H_0: $\beta_1 + \beta_2 = 0$	0.348	0.358	0.361	
p-value on H_0: $\beta_3 = \beta_4 = 0$				0.901
Adj. R^2	0.153	0.156	0.156	0.155
N	3060	3060	3060	3060

Notes: A constant is contained in all specifications, and interviewer fixed effects are included.. In specification (1), other control variables include teenage ln(household income), born in the U.S., age, age squared, birth weight, birth weight squared. In specifications (2) to (4), African American, Native American/Aleut/Pacific Islander, Asian American, other races, Hispanic origin, college graduate mother/father are also controlled for. Clustered standard errors at the interviewer level (390 interviewers) are in parentheses.

In Column (2) of Table 3 we introduce additional controls, including race/ethnicity and parents' educational attainment. Despite the inclusion of these variables, the coefficient estimates on the looks indicators remain largely unchanged compared to those without this additional set of covariates. Moving to Column (3), we add indicators of a teen having difficulties in school (having been suspended or expelled). Since a teenager who is present in school has less free time for gaming, the estimated effect of bad looks may be driven solely by absence from school rather than the hypothesized social network mechanism.[10] With these additional controls, the estimated effect of being unattractive remains essentially zero, while the significant effect of being attractive is nearly unchanged. In Column (3), which presents our best estimates, the difference in gaming incidence between the attractive and average teens is 4.6 percentage points, i.e., 8 percent of the mean incidence. The null hypothesis of symmetric beauty effects cannot be rejected, as the first set of hypothesis tests near the bottom of Table 3 show.

While our hypothesis does not predict any gender difference in the effect of physical attractiveness on gaming, the above results show that females are much less likely to game. To examine whether the impacts of physical attractiveness differ by gender, we estimate the model in Column 3 with additional terms interacting gender and both attractiveness indicators. Column (4) of Table 3 shows the results. The coefficients on these interaction terms are insignificant, both individually and jointly. As such, there is no evidence of gender differences in the effects of physical attractiveness on gaming.

These results are exactly what we would expect from the motivating estimates presented in Table 2 (although those estimates are available only for adults). Those results showed that being good-looking implies a substantial reduction in the probability of having no close friends and a significant increase in the number of close friends, compared to average- or unattractive people. Arguably the differences demonstrated in Table 3 arise because the good-looking teens have more friends with whom to engage in other leisure activities.

The estimates in Table 4 test whether the relationship between beauty and the incidence of teen gaming carries over to the intensity of gaming among those teens who do game. It regresses gamers' time spent gaming sequentially on the same variables included in Columns (1)-(3) of Table 3. None of the effects of beauty on gamers' time spent gaming is statistically significant.

The controls for possible extra time available because a teen is one of the few who is not in school (due to behavioral problems) do not affect the intensity of gaming. Similarly, we cannot reject the hypothesis that the effects of being attractive or unattractive on conditional gaming time are symmetric. Finally, Column (4) shows the results with gender and looks interaction terms included. Again, the coefficient estimates on these terms are insignificant both individually and jointly.

[10] Given the construction of the Add Health survey, essentially all the respondents in Wave I must be enrolled in school. Thus while Ward (2018) shows a relation between video-gaming and missing class, the tiny fraction of Wave I respondents who have been suspended or expelled makes it unsurprising that this control variable does not matter.

Table 4
Estimated effects for teenagers of physical attractiveness on the intensive margin of playing video/computer games.

	Dep. variable: Hours non-zero gaming			
Ind. variable:	(1)	(2)	(3)	(4)
Attractive/very attractive (β_1)	0.554	0.552	0.561	0.740
	(0.647)	(0.618)	(0.597)	(0.925)
Unattractive/very unattractive (β_2)	1.029	1.156	1.153	1.473
	(1.389)	(1.384)	(1.386)	(2.065)
Female	−2.986	−3.074	−3.026	−2.676
	(0.669)	(0.680)	(0.725)	(0.702)
Attractive/very attractive × female (β_3)				−0.540
				(1.188)
Unattractive/very unattractive × female (β_4)				−0.986
				(2.483)
Suspended from school			0.372	0.391
			(0.954)	(0.946)
Expelled from school			−0.555	−0.607
			(1.811)	(1.840)
p-value on H_0: $\beta_1 + \beta_2 = 0$	0.332	0.285	0.283	
p-value on H_0: $\beta_3 = \beta_4 = 0$				0.874
Adj. R^2	0.161	0.167	0.166	0.165
N	1643	1643	1643	1643

Notes: A constant is contained in all specifications, and interviewer fixed effects are included.. In specification (1), other control variables include teenage ln(household income), born in the U.S., age, age squared, birth weight, birth weight squared. In specifications (2) to (4), African American, Native American/Aleut/Pacific Islander, Asian American, other races, Hispanic origin, college graduate mother/father are also controlled for. Clustered standard errors at the interviewer level (328 interviewers) are in parentheses.

4. Beauty and adult gaming

As Table 1 and Figs. 3,4 showed, gaming is more concentrated among adults than it is among teenagers. We first estimate the effect of looks at the extensive margin (Table 5), then at the intensive margin (Table 6), analogous to the estimates for teens in Tables 3 and 4. The control variables are the same as for teens (excluding the measures of suspension or expulsion from school), except that the most expanded specifications also include measures of the respondent's completed educational attainment. Since additional education is a good indicator of a greater value of time, including this vector provides a further dimension to testing for the importance of the opportunity cost of engaging in gaming.

The results in Table 5 strongly support our central hypothesis. As with teens, attractive adults are less likely to game than others, while the few adults rated as unattractive are more likely to engage in gaming (although not significantly so). The results are almost invariant to the inclusion of either of the expanded vectors of covariates. The effects are approximately symmetric around those of the middle group. Moving from the unattractive 8 percent of adults to the attractive 44 percent reduces the likelihood of gaming by over 10 percentage points, about 26 percent of the average incidence. Educational attainment appears to make little difference in the likelihood of an adult gaming. Finally, while women are much less likely than otherwise identical men to engage in gaming, Column (4) shows that the relationship between gaming and looks does not differ at the extensive margin between men and women.

Table 6 shows that attractive adults who do game spend significantly less time doing so than average-looking adults, whose gaming time is, albeit insignificantly, less than that of the small group of unattractive adults. The effect of good looks is somewhat reduced when we include a vector of indicators of educational attainment, which is expected since the latter also proxy for the value of the person's time, and since educational attainment is correlated with attractiveness.[11]

Despite the inclusion of educational attainment, the impact of good looks remains significantly negative. It is noteworthy that, while there was no evidence of a relationship between education and the incidence of gaming, the intensity of gaming generally decreases with education, albeit not significantly. This suggests that whether adults game is largely independent of the value of time that might be spent working, while gamers do consider the implicit value of the time they spend gaming (or that adults' gaming time results from addictive behavior à la Becker and Murphy, 1988). As at the extensive margin, the effects of differences in looks on adults' gaming hours are substantial: Compared to unattractive adults, attractive adults who game spend, on average, 2.05 h fewer doing so per week, i.e., 27 percent of the mean conditional gaming time.

One might be concerned that those who spend more time gaming do so because they have fewer other opportunities, which may also be affected by their looks. Of course, all uses of time beyond gaming will be affected by the same variables, including beauty, that

[11] If we do not interpret the impact of educational attainment as reflecting the opportunity cost of time, then including the vector of educational indicators may mean that we understate the total effect of looks on gaming incidence and conditional time. In the sample of adults, those who are rated as attractive/very attractive are 10 percentage points more likely than those rated average to have at least a college degree (0.095, s.e.=0.023); those rated unattractive/very unattractive are 4 percentage points less likely than average to have finished college (0.041, s.e.=0.042).

Table 5
Estimated effects for adults of physical attractiveness on the extensive margin of playing video/computer games.

	Dep. variable: Gaming, {Yes,No}={1,0}			
Ind. variable:	(1)	(2)	(3)	(4)
Attractive/very attractive (β_1)	−0.061	−0.061	−0.061	−0.048
	(0.023)	(0.023)	(0.023)	(0.035)
Unattractive/very unattractive (β_2)	0.035	0.038	0.043	0.098
	(0.040)	(0.040)	(0.040)	(0.056)
Female	−0.187	−0.187	−0.189	−0.168
	(0.022)	(0.022)	(0.022)	(0.029)
Attractive/very attractive × female (β_3)				−0.027
				(0.043)
Unattractive/very unattractive × female (β_4)				−0.112
				(0.074)
Completed high school or voc/tech training			−0.010	−0.012
			(0.052)	(0.052)
Some college			0.071	0.071
			(0.048)	(0.048)
Completed college			0.035	0.034
			(0.050)	(0.051)
Completed a master's degree			−0.003	−0.005
			(0.063)	(0.063)
Completed a doctoral degree			−0.138	−0.140
			(0.092)	(0.093)
Post-baccalaureate education			0.010	0.006
			(0.081)	(0.080)
p-value on H_0: $\beta_1 + \beta_2 = 0$	0.618	0.660	0.721	
p-value on H_0: $\beta_3 = \beta_4 = 0$				0.295
Adj. R^2	0.100	0.101	0.105	0.105
N	3116	3116	3116	3116

Notes: A constant is contained in all specifications, and interviewer fixed effects are included.. In specification (2), other control variables include born in the U.S., age, age squared, birth weight, birth weight squared. In specification (3) to (4), African American, Native American/Aleut/Pacific Islander, Asian American, other races, Hispanic origin, high school, some college, college, masters, doctorate, and professional qualifications are also controlled for. Clustered standard errors at the interviewer level (263 interviewers) are in parentheses.

influence gaming time. Therefore, gaming time should be thought of as part of a complete system of demand equations describing all uses of time, both market and non-market.

Ignoring this problem, we re-estimated the models in Tables 5 and 6, adding measures of time spent in market work and the presence of children (to reflect the incentive to engage in additional household production). When we add these measures, the estimated effects of looks become slightly smaller in absolute value on the extensive margin but are essentially unchanged along the intensive margin. At both margins, work time has the expected negative effect, although these impacts are not statistically significant. The presence of children, however, has a significant negative relationship with gaming along both margins.[12]

5. Alternative estimators, other non-market activities, specification, and causality

5.1. Alternative estimators

Taken together, the estimates in Tables 3 and 4 showed that beauty matters for gaming among teenagers. Indeed, when combining the estimates and ignoring the differences in the impacts of looks on the extensive and intensive margins (i.e., the unconditional hours spent gaming) by specifying a Tobit model with a lower limit of zero, the effect of bad looks on gaming time is positive and statistically significant compared to the reference group—average-looking teens. As with the estimates for teens, Tobit estimation of the determinants of gaming at both margins among adults yields the same inferences as do the LPM and OLS estimates reported in Tables 5 and 6. This is unsurprising, since the crucial variables—the measures of attractiveness—have effects in similar directions along both margins.

Another concern is suggested by inspection of Figs. 3 and 4, which suggest substantial skewness in the distributions of nonzero gaming time. If the extreme values are not random with respect to attractiveness, the OLS estimates shown in Tables 4 and 6 will give an incorrect impression of the relationship of gaming time and looks at the intensive margin.

To examine this possibility, we re-estimated all the models shown in those tables using a quantile (median) estimator, including the

[12] Here and in the previous section we went beyond the most expanded models by adding measures of interviewers' perceptions of the respondents over/underweight status, and an indicator of how outgoing the respondent seemed. None of these additions made major changes in the estimated impacts of looks, either here or among teens.

Table 6
Estimated effects for adults of physical attractiveness on the intensive margin of playing video/computer games.

Dep. variable: Hours non-zero gaming				
	(1)	(2)	(3)	(4)
Ind. variable:				
Attractive/very attractive (β_1)	−1.849	−1.829	−1.590	−1.357
	(0.716)	(0.712)	(0.677)	(1.033)
Unattractive/very unattractive (β_2)	0.915	0.844	0.461	0.510
	(1.578)	(1.611)	(1.555)	(1.957)
Female	−2.843	−2.918	−2.632	−2.380
	(0.691)	(0.711)	(0.710)	(0.902)
Attractive/very attractive × female (β_3)				−0.574
				(1.439)
Unattractive/very unattractive × female (β_4)				−0.095
				(2.060)
Completed high school or voc/tech training			2.524	2.513
			(1.666)	(0.166)
Some college			0.103	0.094
			(1.580)	(1.578)
Completed college			0.046	0.026
			(1.579)	(1.581)
Completed a master's degree			−1.649	−1.644
			(2.149)	(2.153)
Completed a doctoral degree			0.739	0.828
			(3.758)	(3.773)
Post-baccalaureate education			−1.462	−1.540
			(1.783)	(1.788)
p-value on H_0: $\beta_1 + \beta_2 = 0$	0.621	0.610	0.547	
p-value on H_0: $\beta_3 = \beta_4 = 0$				0.923
Adj. R^2	0.088	0.089	0.097	0.095
N	1191	1191	1191	1191

Notes: A constant is contained in all specifications, and interviewer fixed effects are included.. In specification (2), other control variables include born in the U.S., age, age squared, birth weight, birth weight squared. In specifications (3) to (4), African American, Native American/Aleut/Pacific Islander, Asian American, other races, Hispanic origin, high school, some college, college, masters, doctorate, and professional qualifications are also controlled for. Clustered standard errors at the interviewer level (218 interviewers) are in parentheses.

same controls as before along with interviewer fixed effects.[13] Compared to the estimates of the coefficients on the attractive and unattractive indicators in Table 4, Column (3), the estimates using median regressions are very similar. For the intensive margin of adult gaming time, that is also the case.[14] These, the Tobit, and the Poisson estimates are all listed in the Online Appendix Table A1.

All the estimates presented in Tables 3–6 include interviewer fixed effects, as might be necessary to the extent that interviewers differ in their average generosity in rating interviewees' looks and that these differences are correlated with the outcomes. We re-estimated all the models excluding the fixed effects. The estimates for the indicators "attractive/very attractive" and "unattractive/very unattractive," comparable to those in Columns (3) of Tables 3 and 4 respectively, are −0.051 (s.e.=0.020) and −0.025 (s.e.=0.042); 0.538 (s.e.=0.573) and 0.737 (s.e.=0.988). For the estimates for adults, they are −0.046 (s.e.=0.021) and 0.042 (s.e.=0.038); and −1.153 (s.e.=0.626) and −0.482 (s.e.=1.193), comparable to those in Columns (3) of Tables 5 and 6 respectively. Where the estimates in the tables were statistically significant (the extensive margin among teens, both margins among adults), the estimated impact of good looks here are quite similar (although the estimate at the intensive margin among adults become only marginally significant). Including interviewer fixed effects makes sense, but it hardly alters our inferences.

5.2. Other non-market uses of time

The additional reports on time spent included in the Add Health dataset could, for the same reasons as gaming, be related to the respondents' looks. We thus regress the indicator of a teen's membership in a hobby or sports group on the two indicators of attractiveness (and the same set of controls underlying the results in Columns (3) of Table 3). Noting that all we have is information on the extensive margin, the results, shown in the first column of Table 7, are opposite those that we demonstrated for gaming. Those teens who are attractive/very attractive are 3.4 percentage points more likely than average-looking teens to be members of such groups, while those who are unattractive/very unattractive are 6.1 percentage points less likely than the average-looking to belong to these groups. The 9.5 percentage-point difference between the two groups' membership rates is over half the average rate of

[13] The STATA routine is *mmqreg*.

[14] Replacing the LPM estimators of incidence with probits and the OLS estimators of intensity with Poisson estimates does not alter the inferences made in Sections 3 and 4 about the directions and significance of the relationships of gaming to attractiveness.

Table 7
Estimated effects of physical attractiveness on other non-market uses of time.

Dep. variable:	(1) Teen hobby clubs (yes/no)	(2) Adult team sports (yes/no)	(3) Adult team sports (times/week)	(4) Adult individual sports (yes/no)	(5) Adult individual sports (times/week)
Ind. variable:					
Attractive/very attractive	0.034	0.009	0.315	0.040	0.350
(β_1)	(0.019)	(0.016)	(0.268)	(0.024)	(0.179)
Unattractive/very	−0.061	0.018	0.009	−0.001	0.252
unattractive (β_2)	(0.031)	(0.027)	(0.296)	(0.030)	(0.500)
Adj. R^2	0.064	0.086	0.213	0.066	0.004
N	3017	3199	315	3199	780

Notes: A constant is contained in all specifications, and interviewer fixed effects are included.. Column (1) includes the same controls as in Column (3) of Table 3. Columns (2)-(5) include the same controls as in Column (3) of Table 5. Clustered standard errors at the interviewer level are in parentheses.

participation, a difference that is significantly nonzero.

The additional information for adults on their non-market time is also useful for analyzing the impacts of looks on behavior. As shown in Column (2) of Table 7, regressing an indicator for whether the respondent participated in individual sports during the previous week reveals only small and statistically insignificant differences by perceived attractiveness. Column (3) presents results for the intensive margin of participation in team sports. Although both coefficients for "attractive" and "unattractive" respondents are statistically insignificant at the 10 percent level, the coefficient for attractive individuals is positive with a t-ratio exceeding 1, whereas that for unattractive individuals is close to zero.

Column (4) reports results from a regression like that in Column (2), but describing an indicator of whether the respondent participated in individual sports in the previous week as the dependent variable. Comparing the results by the respondents' looks, LPM estimates show that the attractive adults are 4 percentage points more likely to engage in individual sports than average-looking adults on a mean incidence of 25.7 percent.

Similar results exist for behavior at the intensive margin of participation in individual sports, with results shown in Column (5) of Table 7. Conditional on the large vector of covariates included in Column (3) of Table 6, those with above-average looks participate 0.35 times more per week than the average-looking respondent, with no significant difference in behavior between adults of average- or below-average attractiveness.

The estimated impacts on teens' participation in clubs and adults' participation in individual sports convey the strong impression that attractiveness matters in additional non-market activities, with its effects being generally opposite those on teens' and adults' gaming time. Since even individual sports involve interactions with other people (think of tennis or golf), the findings for these activities show that where good looks might pay, better-looking adults engage in more of them. These results for both teens and adults thus mirror the impacts of looks on gaming.

5.3. *Specification*

Our main results could be sensitive to the specification of physical attractiveness—specifically, dividing it into only three categories from the five available to the Add Health study interviewers. Despite the paucity of respondents rated very unattractive, either as teens or adults, we re-specify the equations to include all four possible indicators of looks. All the estimates, shown in the Online Appendix Table A2, are comparable to our preferred specifications, which include all the covariates used in Columns (3) of Tables 3–6. Among both teens and adults, and along both the extensive and intensive margins, they show that the results arise similarly from the behavior of both the attractive and the very attractive. The distinction is greatest when we examine the intensive margin among teen gamers. These results, which use all the available information on attractiveness, do not alter the conclusions from Tables 3–6.

5.4. *Reverse causality*

With the effect of looks at both margins being greater in adulthood, one might be concerned that what we have documented is explained by reverse causality—i.e., adults' looks have been at least partly determined by the incidence and amount of gaming as a teenager. This argument hinges on the potential endogeneity of beauty. We find it difficult to construct mechanisms by which looks are affected by gaming time, although perhaps being isolated from others as a teen leads to slovenliness that affects subsequent perceptions of an adult's looks.

We cannot simply dismiss the notion that gaming can affect physical attractiveness; but we can use the longitudinal aspect of the Add Health study to explore its validity. If the results in the previous section were largely driven by the effect of time spent gaming on looks, rather than vice-versa, then we would anticipate that teenage gaming will predict lower adult physical attractiveness, conditional on teenage physical attractiveness. This could be especially the case if teenage gaming, conditional on teenage looks, also predicts persistent gaming into adulthood.

To explore the potential reverse causality of gaming on looks, we estimate variants of the following linear regression models:

Table 8
Estimated effects of gaming time and physical attractiveness as teens on the gaming time of adults.

Dep. variable:	Hours of gaming	Gaming, {Yes,No}={1,0}	Hours non-zero gaming
Ind. variables from Wave I:	(1)	(2)	(3)
Gaming hours (ρ)	0.137		0.393
	(0.059)		(0.130)
Gaming, {Yes,No}={1,0} (ρ)		0.092	
		(0.028)	
Attractive/very attractive (β_1)	−0.675	−0.061	0.103
	(0.466)	(0.026)	(1.144)
Unattractive/very unattractive (β_2)	0.815	0.055	1.489
	(0.902)	(0.057)	(1.379)
Female	−2.351	−0.159	−1.234
	(0.410)	(0.030)	(1.063)
Interviewer fixed effects, Wave I	Yes	Yes	Yes
p-value on H_0: $\beta_1 + \beta_2 = 0$	0.899	0.920	0.449
Adj. R^2	0.114	0.126	0.114
N	2059	2059	426

Notes: See Eq. (2). A constant is contained in all specifications. Clustered standard errors at the Wave I interviewer level are in parentheses.

$$G_{i,4} = \alpha + \rho G_{i,1} + \beta_1 A_{i,1} + \beta_2 U_{i,1} + \boldsymbol{\theta} \boldsymbol{X}_{i,1} + \gamma_{v(i,1)} + \varepsilon_{i,4} \tag{2}$$

$$B_{i,4} = \pi + \sigma G_{i,1} + \delta_1 A_{i,1} + \delta_2 U_{i,1} + \widetilde{\boldsymbol{\theta}} \boldsymbol{X}_{i,1} + \xi_{r(i,4)} + \varphi_{v(i,1)} + \nu_{i,4} \tag{3}$$

where the subscript $\{i,4\}$ indicates information about an Add Health respondent in Wave IV, as an adult, and $\{i,1\}$ indicates information for that same respondent from Wave I, as a teenager. Eq. (2) is similar to our main regression model, except that it specifically regresses the measure of gaming in adulthood at Wave IV of Add Health, $G_{i,4}$, on the equivalent measure when the individual was a teenager at Wave I, $G_{i,1}$, with ρ thus estimating the persistence of gaming conditional on the looks of the person recorded at Wave I, $A_{i,1}$ and $U_{i,1}$.

Eq. (3) regresses a measure of physical attractiveness as an adult at Wave IV on some measure of gaming as a teen at Wave I (either unconditional time, incidence, or conditional gaming time), the attractiveness indicators from Wave I, and interviewer-specific fixed effects at both Waves I and IV. Thus, σ indicates whether teenage gaming, which may or may not be persistent according to Eq. (2), predicts the physical attractiveness of adults conditional on how their looks were judged as teens.

Table 8 presents the estimates of Eq. (2), using three different measures of gaming, with only gender included in the vector of controls, $\boldsymbol{X}_{i,1}$. In Column (1) we consider all adult gaming, including zeroes. Column (2) examines the extensive margin of adult gaming, depending on whether the person gamed as a teenager or not, and Column (3) estimates the intensive marginAs Column (1) demonstrates, the net effect is that an attractive adult spends less time in this activity than other adults, even conditional on the time spent gaming as a teenager, while the small number of unattractive adults do more gaming than the average-looking adult, but not significantly so.[15] As a comparison of the estimates in Columns (2) and (3) demonstrates, this finding arises more from behavior on the extensive margin.

A person's gaming behavior as a teen is positively related to their incidence of gaming as an adult and the time spent in this activity. We cannot infer from these estimates whether this dynamic relationship results from an underlying preference for gaming, or instead from addictive behavior that began in a person's teen years or earlier. Suffice it to say that there is substantial persistence of gaming from adolescence onward.

We can even use these data to provide an admittedly weak direct test of our hypothesis that the role of beauty is causal, resulting from unattractiveness leading to having fewer friends. Respondents in Wave I were asked to "nominate" either 1 male and 1 female or (in a small sub-sample) up to 5 male and 5 female friends, with some nominating nobody.[16] Creating an indicator variable equaling one if no friends are listed and replacing the beauty indicators in Column (1) of Table 8 with it yields a t-statistic on this indicator of +1.31. Even controlling for teen gaming time, we observe that those who were "friendless" as teens also spend more time gaming as adults than those with at least some teen friends.

Table 9 shows OLS estimates of Eq. (3), which tests directly for the reverse causality of teen gaming and gaming time on beauty as assessed during adulthood. The estimates in Column (1) show that the unconditional amount of time spent gaming as a teen has a negative but statistically insignificant impact on adult looks. Column (2) shows no indication that adults' beauty ratings are related to whether they spent time gaming as a teenager: The incidence of teen gaming has a tiny and 3statistically insignificant negative relation to adult beauty, conditional on looks as a teenager. Similarly, Column (3) shows that the impact of time spent gaming on adult beauty

[15] We also estimate Equation (2) separately for males and females, with the results presented in the Online Appendix Tables A3 and A4. The beauty effect is predominantly observed among females, conditional on teenage gaming time.

[16] This is different from the question in Wave IV used in Table 2. That variable is from the direct question on "close friends".

Table 9
Estimated effects of gaming time and physical attractiveness as teens on the physical attractiveness of adults.

Dep. variable:	Physical attractiveness in Wave IV, {unattractive/very unattractive, average, attractive/very attractive}={0,1,2}		
Ind. variables from Wave I:	(1)	(2)	(3)
Gaming hours (σ) (Hours in Col. 1; hours non-zero gaming in Col. 3)	−0.002		−0.002
	(0.002)		(0.008)
Gaming, {Yes,No}={1,0} (σ)		−0.011	
		(0.036)	
Attractive/very attractive (δ_1)	0.200	0.200	0.151
as teen	(0.037)	(0.037)	(0.113)
Unattractive/very unattractive (δ_2)	−0.256	−0.255	−0.537
as teen	(0.077)	(0.078)	(0.272)
Female	0.051	0.054	−0.010
	(0.038)	(0.039)	(0.103)
Interviewer fixed effects: Waves I & IV	Yes	Yes	Yes
p-value on H_0: $\delta_1 + \delta_2 = 0$	0.535	0.537	0.185
Adj. R^2	0.216	0.216	0.237
N	2076	2076	333

Notes: See Eq. (3). A constant is contained in all specifications. Clustered standard errors at the Wave I and IV interviewer levels are in parentheses.

among those teens who do game is also essentially zero. In all three specifications the estimated coefficients δ_1 and δ_2 unsurprisingly show that those who are rated (un)attractive as teens are more likely to be rated as (un)attractive as adults.[17]

We have analogized testing for causality here to panel VAR models. The estimates of Eqs. (2) and (3) suggest that, while looks affect the incidence and amount of adult gaming, even accounting for teen gaming, prior gaming has essentially no effect on adult's looks. One might summarize this exploration by concluding that adult gaming nerds appear to be born, not made (by prior gaming).

6. Conclusion

We expand the literature on the impact of human beauty by examining its effects on how people spend their time. The results provide compelling evidence, both among teenagers and adults, but especially the latter, that, at least during the gaming environments of the 1990s and 2000s, more physically attractive individuals spent less time on video-gaming. They also spent more time in clubs and in sports, especially individual sports. We offer suggestive evidence that these relationships arise because the better-looking have more friends with whom to socialize. The relationship between looks and gaming does not arise because gaming harms people's looks: the causation appears to go from looks to gaming, not vice-versa.

Whether the changing technology of gaming might have altered the impacts of looks on the incidence/intensity of gaming since the late 2000s is unclear. One can easily make arguments either way. Suffice it however, to note, that we have provided at least an indication of habit persistence/addiction to gaming, so that the results that we have shown for earlier periods almost surely persist among some part of today's population.

No doubt one could construct non-economic “stories” to explain our results, linking the “why”—the greater number of close friends among the good-looking—to the “who”—the finding of less gaming among the better-looking. Nonetheless, taken together, these results suggests that a simple economic explanation is quite consistent with what we observe—the better-looking have a higher opportunity cost of gaming as they have a comparative advantage in social interactions in alternative leisure activitiesy.

The evidence of a negative effect of looks on gaming among teenagers suggests its importance in affecting development and thus longer-term outcomes. The negative effect among both teenagers and adults underscores its potential importance in shaping non-market time use and, more generally, lifestyle choices. Understanding how beauty influences gaming behavior and other time use can provide valuable insights into the broader economic and social mechanisms underlying the consumption of leisure in the digital age. The finding that the activity is more likely to be undertaken by the unattractive suggests the need to be particularly concerned for those who are otherwise least advantaged by their looks.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

[17] The results for the separate samples of males and females are presented in the Online Appendix Tables A5 and A6. Consistently, neither the incidence of teen gaming nor the time spent gaming shows a statistically significant effect on adult beauty, regardless of gender.

Supplementary materials

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.jebo.2025.107340.

Data availability

The corresponding author can share the employed dataset with replicators, provided they are faculty, staff, or students at an ICPSR member institution.

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