PNAS

Main Manuscript for

The Parenthood Paradox and the Hedonic Costs of Life-Course Misalignment

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Author Contributions: Saurabh Bhargava was responsible for the conception, development, and writing of the paper.

Competing Interest Statement: I have no competing interests to declare.

Classification: Social Sciences (Psychological and Cognitive Sciences)

Keywords: Subjective well-being, parenthood, experience sample, happiness

This PDF file includes:

Main Text
Figures 1 to 5

Abstract

A large literature finds that parenthood is weakly or negatively associated with well-being, a result widely regarded as paradoxical. Analyzing over one million momentary reports from 3,775 U.S. adults, I show that parenthood reshapes daily experience in ways conventional evaluative measures obscure. Motherhood is associated with higher momentary mood, greater arousal, and more frequent peak emotional states, whereas fatherhood is characterized by increased emotional volatility and negative emotion. Custodial parenthood also compresses the distribution of emotional experience and social time, rendering parents' daily lives more similar to one another. These effects are organized by life-course preparedness: parenthood premiums are largest among adults whose circumstances most strongly predict parenthood, driven not by differential benefits of parenting but by well-being penalties among similarly prepared adults who remain childless. I replicate these core patterns— the motherhood premium, distributional compression, and preparedness heterogeneity— using U.S. data from the Gallup World Poll. This framework has a critical boundary condition: the motherhood premium is erased for Black women, primarily because Black non-mothers do not exhibit the penalties observed among non-Black counterparts. Using data from the American Time Use Survey, I show that life satisfaction measures predictably attenuate these experiential effects, functioning as negativity-weighted compressors that discount positive experience. Together, these findings reframe the paradox: the weak empirical signature of parenthood reflects both a reliance on evaluative measures ill-suited to capturing how parenthood reorganizes daily life and an analytic focus on mean parental gains that obscures the costs of childlessness among those whose circumstances most predict parenthood.

Significance Statement

How parenthood affects well-being is central to social science and policy, especially amid sustained fertility decline. Using over one million momentary reports from 3,775 U.S. adults, this study shows that parenthood reshapes daily experience in ways conventional measures obscure. Motherhood is associated with higher mood, greater arousal, and more frequent peak states; fatherhood with elevated negative emotion and volatility. Custodial parenthood also compresses between-person dispersion in emotional and social experience. The largest gaps emerge among adults whose circumstances most predict parenthood, driven by well-being penalties among non-parents rather than benefits of parenting— a pattern absent for Black women. Life satisfaction measures systematically attenuate these experiential differences, helping explain why prior research has understated parenthood's hedonic consequences.

Introduction

How parenthood affects well-being is among the most consequential questions in the social sciences, with growing relevance as societies confront sustained fertility decline (Doepke et al., 2023). Children are widely believed to be a primary source of joy and meaning (Nelson et al., 2013; Hansen, 2012), and most adults express a desire to become parents. Yet decades of research have found that the average association between parenthood and well-being—typically measured as life satisfaction—is small, null, or negative, albeit with pronounced heterogeneity across subgroups and institutional settings (Nomaguchi & Milkie, 2020; Stanca, 2012). This disconnect—sometimes termed the "parenthood paradox"—raises a foundational question: Why does one of the most consequential transitions in the human life course appear to leave such weak and variable traces in reported well-being?

Three broad explanations have been advanced. The first emphasizes contextual heterogeneity. A large literature documents systematic variation in the parenthood–well-being association by gender, marital status, financial resources, child age, and institutional environment (Nomaguchi & Milkie, 2020; Nelson et al., 2014). Marriage, higher income, and older age at first birth often attenuate negative associations (Margolis & Myrskylä, 2011; Myrskylä & Margolis, 2014; Pollmann-Schult, 2014; Herbst & Ifcher, 2016), while cross-national comparisons reveal apparent parenthood penalties in contexts characterized by intensive parenting norms and weaker family policy (Glass, Simon, & Andersson, 2016; Aassve, Mencarini, & Sironi, 2015). These moderators clarify which parents fare better or worse on average, but they do not offer an integrated account of why average effects remain modest despite the scale of the changes parenthood introduces into daily life—nor do they explain why the same observable conditions are sometimes associated with positive and sometimes negative effects across studies.

A second explanation concerns measurement. Studies relying on evaluative measures such as life satisfaction often find weak average associations with parenthood, whereas analyses using measures of experienced well-being—capturing how people feel in the moment—sometimes find positive associations (Deaton & Stone, 2014; Musick, Meier, & Flood, 2016; Negraia & Augustine, 2020). These approaches capture partially overlapping but conceptually distinct dimensions of well-being (Kahneman & Krueger, 2006). Evaluative judgments draw on values, social comparisons, and reflective goals alongside recent experience (Schwarz, 1999), and may be disproportionately influenced by salient or negative aspects of parenting (Smith et al., 2006). By collapsing lived experience into a single retrospective assessment, they may also obscure higher-order features of daily life—variability, emotional composition, contextual organization—that are potentially central to understanding how parenthood alters everyday experience. At the same time, experiential measures may miss dimensions such as purpose and meaning that evaluative assessments capture and that parents frequently endorse (Nelson et al., 2013; Baumeister et al., 2013). Understanding the parenthood–well-being association thus requires both approaches and close attention to how they map onto one another.

A third, less examined explanation concerns the counterfactual to parenthood. Most research relies on cross-sectional comparisons that implicitly treat non-parenthood as a neutral baseline. This framing obscures a potentially critical asymmetry: the well-being consequences of parenthood may depend not only on having children, but on what it means to remain childless when one's life-course circumstances—age, partnership, economic stability—strongly predict family formation. Life-course theory emphasizes that major transitions carry different consequences depending on their alignment with normative timing and social expectations (Elder, 1998; Settersten & Mayer, 1997). When circumstances align with parenthood, remaining childless may itself carry experiential costs, whether through unmet expectations, social stigma, or the absence of the coordinating structure that parental roles provide (Dykstra & Hagestad, 2007; Koropeckyj-Cox, 1998; McQuillan et al., 2012). Standard comparisons leave unresolved whether weak or variable parenthood effects reflect limited gains among parents, unmeasured penalties among the childless, or both.

This study engages all three explanations within a unified empirical framework, drawing on a large-scale experience-sampling dataset that captures well-being as it is lived in daily life. The data consist of over one million in-the-moment reports of mood, emotion, and time use collected from 3,775 U.S. adults over a ten-day diary period (Bhargava, 2024; Chin et al., 2017). By eliciting repeated assessments of momentary mood throughout each waking day, the design provides a close empirical analogue to experienced utility (Kahneman, Wakker, & Sarin, 1997). Unusually high compliance, detailed emotional and contextual information at each prompt, and the high-frequency longitudinal structure of the data enable analyses that extend beyond mean comparisons to examine distributional features of experienced well-being and the organization of daily social life. The analyses address three interrelated questions: How is parenthood associated with the average structure of emotional experience? How do these associations vary systematically with life-course preparedness? And beyond mean effects, how does parenthood shape the distribution of emotional and social experience across individuals?

Several findings emerge. First, parenthood is associated with a sharply gendered affective profile: motherhood corresponds with meaningfully higher average mood and arousal, while fatherhood is characterized by elevated negative emotion and greater emotional volatility. Second, these associations are systematically organized by life-course preparedness. The parent–non-parent well-being gap widens substantially among adults whose circumstances most strongly predict parenthood—seemingly driven not by differential enjoyment of children or favorable time reallocation, but by broad well-being declines among similarly prepared adults who remain childless. Third, custodial parenthood compresses the distribution of both emotional experience and social time use across individuals, with parents leading more similar daily lives than observationally comparable non-parents. Finally, analyses using contemporaneous national survey data indicate that conventional life satisfaction measures systematically attenuate the experiential associations documented here, consistent with a lossy and asymmetrically weighted mapping between experienced and evaluative well-being. Collectively, these findings—which replicate in nationally representative Gallup data spanning two decades—reframe the parenthood paradox as partly an artifact of how well-being is typically measured and partly a reflection of how the counterfactual to parenthood is conventionally understood.

Data and Analytic Overview

The primary analyses draw on USA Touchpoints (TP), a large-scale experience-sampling study in which a diverse set of 3,775 U.S. adults, aged 21 to 65, completed a 10-day mobile diary between 2010 and 2014, yielding over one million momentary reports (Bhargava, 2024; Chin et al., 2017). Participants were recruited from a nationally representative commercial panel and were unaware of the present research purpose. Every waking half hour, participants used a mobile diary application to record their current activity, social context (whom they were with), and location. At each prompt, respondents rated their momentary mood on a five-point scale anchored from "bad mood" to "good mood," rated their arousal from "relaxed" to "alert," and indicated the presence or absence of 14 specific discrete emotions (e.g., anger, happy, lonely, love, sad) via emoji-labeled indicators. The latter was used to calculate a measure of net emotion as the net count of positive minus negative emotions reported. Generous incentives and the provision of dedicated mobile devices produced an estimated 97% compliance rate, defined as at least 7 days of qualified data (i.e., at least 16 half-hour reports).

Extensive background surveys administered during recruitment provide demographic and socioeconomic information, including age, gender, race and ethnicity, education, marital status, employment, household income, and household net worth. Parenthood is defined using self-reported household composition and self-reported indications of non-custodial and adult children, with the baseline definition capturing any parenthood exclusive of grandparents. Alternative definitions that vary custodial status and the treatment of grandparents are examined in robustness checks (SI Appendix).

The analyses proceed in several stages. I first estimate the association between parenthood and experienced well-being using participant-level regressions that adjust for observed demographic and socioeconomic characteristics, reporting results separately by gender for average mood, arousal, discrete emotions, and within-person emotional volatility (mood variance). To examine heterogeneity across the life course, I construct gender-specific indices of life-course preparedness based on the predicted probability of parenthood given observable characteristics. These indices summarize the degree to which each respondent's demographic and socioeconomic profile aligns with the conditions under which parenthood is normatively expected. The resulting scores are not used as control variables but as an organizing dimension along which I examine heterogeneity in the parenthood–well-being association.

To decompose potential mechanisms, I use within-person fixed-effects models to estimate the affective returns to child-related time and Gelbach (2016) decompositions to assess the contribution of time-use shifts to observed well-being differences. To examine distributional effects, I estimate heteroskedastic regression models that allow the conditional variance of experienced well-being to differ by parenthood status and use Shannon entropy measures to assess convergence in the social organization of daily life. Throughout, standard errors are clustered at the participant level.

Supplementary analyses draw on two nationally representative datasets. The U.S. segment of the Gallup World Poll (GWP), spanning roughly two decades, provides a population-level benchmark for evaluating the generalizability of the experience-sampling findings across both evaluative (life satisfaction) and remembered emotional well-being. And the American Time Use Survey (ATUS), which in 2012–2013 paired time-use diaries with both episode-level affect ratings and a global life satisfaction measure from the same respondents, permits examination of the empirical mapping between experienced and evaluative well-being. Details on measurement, variable construction, sample selection, and all estimation procedures are provided in Methods and the SI Appendix.

Results

Experienced Well-Being and Parenthood. Figure 1 summarizes the relationship between parenthood and experienced well-being estimated from participant-level regressions adjusting for observed demographic and socioeconomic characteristics, reported separately by gender. In pooled specifications, the association between parenthood and average mood is positive but statistically indistinguishable from zero, consistent with the aggregate null widely reported in the literature. This null, however, masks substantial gender heterogeneity.

Among women, motherhood is associated with a significant increase in average momentary mood ($b = 0.086$, $p = 0.020$; 95% CI [0.014, 0.158]), roughly 80 percent as large as the within-person weekday–weekend difference and one-half of the marital association (SI Appendix Table S1). Motherhood is also associated with a 4.1 percentage point increase in the probability of reporting the highest mood level — a peak-affect measure capturing the frequency of qualitatively positive moments — equivalent to 84 percent of the weekday–weekend difference in peak-mood incidence (and 82 percent of the marital association). Motherhood additionally corresponds to significantly higher arousal ($b = 0.119$, $p = 0.017$; 0.10 SD), by a margin exceeding the weekday–weekend difference, suggesting not only higher experienced well-being but a more engaged daily emotional life. Corresponding associations for men are smaller and statistically insignificant across all three measures. These patterns are robust to alternative definitions of parenthood, including specifications that vary custodial status and the treatment of grandparents (SI Appendix Table S2).

Beyond average mood and arousal, Figure 1 characterizes the emotional correlates of parenthood (Panel B). Here again, the results are sharply gendered. Motherhood is not associated with significant shifts in aggregate positive or negative emotion, nor with increased emotional volatility. Instead, it is characterized by offsetting changes across specific emotions: significant increases in love and reductions in loneliness and worry, counterbalanced by a decline in hope. This pattern of affective reweighting, rather than a uniform shift in valence, implies that the motherhood association operates through the composition of emotional life rather than through a simple uplift in positive affect.

Fatherhood presents a qualitatively different profile. Rather than shifts in average mood, fatherhood is characterized by elevated negative emotion ($b = 0.036$, $p = 0.001$; 0.20 SD of non-fathers' negative emotion) driven primarily by higher reported exhaustion, frustration, and boredom, and by increased mood volatility ($b = 0.042$, $p = 0.044$), concentrated among unmarried fathers ($b = 0.081$, $p = 0.032$).

While these estimates rely on cross-sectional comparisons, under the assumption that parenthood status is conditionally independent of potential well-being outcomes given the observed covariates— an assumption supported by the richness of the covariate set and assessed formally through a subsequent sensitivity analysis— these estimates admit a causal interpretation.

Life-Course Alignment and the Apparent Parenthood Premium. The preceding results document pronounced gender differences in the average association between parenthood and experienced well-being. I next examine how these associations vary with life-course preparedness — the degree to which an individual's observable demographic and socioeconomic circumstances align with the conditions under which parenthood is normatively expected. To characterize this alignment, I estimate gender-specific logistic regressions of parenthood status on the full set of baseline covariates. The resulting predicted probabilities serve as a continuous index of preparedness that summarizes, in a single dimension, the joint configuration of age, partnership, employment, education, and economic position that distinguish typical parents from non-parents within each gender. The index is strongly predictive: among women, 93% of individuals in the highest tercile are parents compared with 27% in the lowest; among men, the corresponding figures are 91% and 16%. These propensity scores are used to stratify the sample rather than to adjust outcome regressions, which retain the full covariate set throughout.

Figure 2 plots locally smoothed average mood, adjusted for covariates, as a function of preparedness, separately for parents and non-parents. A striking asymmetry emerges. For both women and men, parents' experienced well-being is relatively flat across the preparedness distribution. Among non-parents, by contrast, experienced well-being declines markedly at high levels of predicted parenthood. As a result, the parent–non-parent gap widens substantially among the most prepared individuals — those whose life circumstances most strongly predict family formation but who remain childless.

Tercile-based estimates reinforce this pattern. Pooling across genders, the parenthood association increases by 0.211 points ($p = 0.009$) from the lowest to highest preparedness tercile, with a similarly significant gradient for peak-mood incidence ($b = 0.091$, $p = 0.02$). Gender-specific estimates reveal consistent gradients of similar magnitude for both women ($T3-T1 = 0.218$, $p = 0.054$) and men ($T3-T1 = 0.204$, $p = 0.075$) — individually less precise than the pooled estimate, as expected, but directionally aligned. Among women, the parenthood association is small and insignificant at low preparedness ($b =$

0.047, $p = 0.48$) but large and significant at high preparedness ($b = 0.265$, $p = 0.004$). Among men, the association shifts from negative ($b = -0.109$, $p = 0.16$) to modestly positive ($b = 0.095$, $p = 0.26$). Across genders, the gradient is generated by declines in non-parent well-being at high preparedness rather than by increases in parental well-being.

To identify which variables contribute to this gradient, I sequentially exclude each covariate from the propensity score and re-estimate the gradient while retaining the full covariate set in the outcome regression. A sharp gender asymmetry emerges (SI Appendix Table S3). Among women, excluding age collapses the gradient entirely (T3−T1 shifts from +0.218 to −0.050), with spousal and own employment contributing secondarily; excluding marriage leaves it unchanged (0.217), and excluding income and wealth modestly increases it (0.273). Among men, the pattern reverses: excluding marriage produces near-complete attenuation (from 0.204 to 0.036), with age contributing secondarily. The dimensions along which preparedness structures heterogeneity thus differ markedly by gender—age-normative timing and employment structure for women, and partnership and age for men. In neither case does it reduce to financial resources.

Decomposing mood into positive and negative emotion clarifies the affective structure of this gradient. Across genders, the widening parent–non-parent gap at high preparedness is mirrored by a sharp divergence in positive emotion, while negative emotion remains comparatively flat across the preparedness distribution (SI Appendix Figure S1). This pattern — declining positive affect with stable negative affect among non-parents — contributes to a significant preparedness gradient in net emotion (pooled T3−T1 = 0.163, $p = 0.043$), though the gradient is substantially stronger among women (T3−T1 = 0.231, $p = 0.032$) than men (T3−T1 = 0.094, $p = 0.431$). The affective profile among highly prepared non-parents — characterized by the absence of positive engagement rather than the presence of distress — is suggestive of an anhedonic pattern consistent with normative misalignment rather than acute distress. Gradients for individual discrete emotions are reported in the SI Appendix.

Decomposing Mechanisms: Time Use, Affective Returns, and the Non-Parent Penalty. I consider three broad classes of explanation for the observed heterogeneity in parental returns.[1] First, some groups may derive greater affective returns from time spent with children. Second, parenthood may differentially reallocate daily time toward more or less rewarding activities. Third, parenthood may alter the affective quality of non-child contexts — either because parents experience ordinary activities more positively, or because the counterfactual experience of remaining childless is worse for some groups than others.

Affective Returns to Child Time. Using individual fixed-effects models that exploit within-person variation in child presence across the diary window, I estimate how mood changes for the same parent when they are with a child versus not. Both mothers ($b = 0.058$, $p < 0.001$) and fathers ($b = 0.128$, $p < 0.001$) report significantly higher mood during child-present episodes, confirming that time with children is affectively rewarding on average. However, these within-person returns do not parallel the broader heterogeneity. The return is substantially larger for fathers than for mothers (difference = 0.070, $p <$

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[1] Formally, average experienced well-being for group g can be written as $E[W \mid g] = \sum_{c \in C} \pi_{cg}\ \mu_{cg}$, where c indexes mutually exclusive daily contexts, π_{cg} is the share of time group g spends in context c, and μ_{cg} is average mood within that context. Partitioning contexts into child-present (C^{ch}) and non-child (C^{nc}) sets, differences in well-being across groups must arise from some combination of (i) differences in within-context mood during child-present episodes (μ_{cg} for $c \in C^{ch}$), (ii) differences in time allocation across contexts (π_{cg}), or (iii) differences in within-context mood during non-child episodes (μ_{cg} for $c \in C^{nc}$).

0.001)—the reverse of the gender asymmetry in average experienced well-being. And affective returns are effectively identical across preparedness terciles (difference = 0.002, $p = 0.917$). Differential enjoyment of children therefore cannot account for the observed heterogeneity in the parenthood premium.

Time-Use Reallocation. Parenthood is associated with large shifts in the composition of daily time. Pooling across genders, parenthood corresponds with increases of 16.2 percentage points in dyadic child time and 15.5 percentage points in triadic time with child and partner (both $p < 0.001$), offset primarily by a 15.2-percentage-point reduction in partner-only time ($p < 0.001$) and smaller reductions in self-maintenance and domestic work. These compositional shifts are sharply gendered: mothers experience substantially larger increases in dyadic child time than fathers (21.8 vs. 9.7 percentage points), while fathers' gains are concentrated in triadic time (17.0 vs. 14.1 percentage points). Reductions in partner-only time are similar in magnitude across genders (~15 percentage points), but mothers absorb larger reductions in self-maintenance time (13.2 vs. 8.7 percentage points).

Gelbach (2016) decompositions indicate that these compositional shifts cannot mechanically account for the observed well-being differences (SI Appendix Table S4). Among women, the hedonic contributions of increased child-centered time are substantial — dyadic child time contributes +0.223 points to the motherhood mood premium and triadic time contributes +0.185 points (both $p < 0.001$ — but these gains are almost entirely offset by hedonic losses from reduced partner time (−0.205, $p < 0.001$), leisure (−0.161, $p < 0.001$), and domestic time (−0.052, $p < 0.05$). Among men, a similar offsetting pattern obtains at smaller magnitudes (net contribution = −0.018, $p > 0.10$). Among high-preparedness individuals, where the parenthood premium is largest, the net time-use contribution is likewise non-positive and insignificant (−0.066, $p > 0.10$). In all three subgroups, the time-use-adjusted parenthood coefficient is at least as large as the unadjusted coefficient indicating that the compositional shifts induced by parenthood, if anything, work slightly against the observed premium. Time-use reallocation is thus a significant consequence of parenthood that does not, on net, explain the observed well-being differences.

Hedonic Returns to Non-Child Contexts: The Non-Parent Penalty. The evidence instead points to a third channel: systematic differences in hedonic returns to non-child contexts. Figure 4 plots covariate-adjusted average mood across core non-child activities — including work, leisure, and domestic tasks—by parenthood status, gender, and preparedness. A notable pattern emerges that mirrors the aggregate preparedness gradient. Among parents, mood across non-child contexts is relatively stable across preparedness; among non-parents, it declines sharply at high preparedness. This asymmetry is not confined to a single activity domain—highly prepared non-parents report systematically lower mood across a wide range of everyday activities, generating the parent–non-parent gap from the non-parent side. Quantitatively, the parent–non-parent gap in non-child well-being widens by approximately 0.17 to 0.19 points between the lowest and highest preparedness terciles for both genders, driven in each case by declining non-parent mood rather than rising parental mood.

A formal decomposition of the aggregate preparedness gradient reinforces this conclusion. Expressing the T3−T1 contrast in the parenthood association as the difference between the preparedness slope among parents and the preparedness slope among non-parents—an accounting identity that holds mechanically—reveals that the gradient is generated overwhelmingly from the non-parent side. Among women, covariate-adjusted mean mood among mothers is virtually identical at low and high preparedness (3.86 at both T1 and T3), while non-mothers' adjusted mood declines from 3.79 to 3.66 — a drop of 0.13 points that accounts for the entirety of the widening gap. Among men, both groups decline with preparedness, but non-fathers account for approximately 87 percent of the gradient, declining 0.28 points compared with 0.04 among fathers. These patterns are consistent with a broad, context-general well-being deficit among highly prepared non-parents.

Interpreting the decomposition causally—as evidence that the gradient reflects a non-parent penalty rather than a parental gain—requires the assumption that, absent parenthood, preparedness would shift well-being similarly for individuals who become parents and those who do not, analogous to a parallel trends assumption in difference-in-differences designs. The near-flat preparedness slope among parents and the invariance of within-person affective returns to child presence across preparedness terciles are both consistent with this assumption, but it is not directly testable, and alternative explanations — including selection on unobserved traits — cannot be definitively excluded. These identification issues are addressed more fully in the Discussion.

If preparedness-linked differences reflect life-course misalignment, such penalties should be most pronounced when institutional coordination demands are most binding. The evidence is directionally consistent with this interpretation: among highly prepared adults, non-parents exhibit a larger weekend mood rebound than parents (within-person weekend effects of 0.14 vs. 0.10), partially narrowing the parent–non-parent gap as institutional constraints relax. This attenuation is consistent across alternative weekend definitions, though statistical precision is limited by the small number of weekend days in the diary window (SI Appendix Table S5).

To further examine whether the non-parent penalty reflects time-limited normative misalignment rather than stable personality differences, I decompose the motherhood association by age, estimating covariate-adjusted mean mood separately for mothers and non-mothers across five-year age bands (Figure 3). Two features are immediately apparent. First, maternal mood is remarkably stable: adjusted mean mood varies by only 0.07 points across the entire age range from 28 to 58, indicating that motherhood provides a relatively constant hedonic floor regardless of age. Second, non-mothers' mood traces a pronounced valley during the normative fertility window, declining from 3.66 in the early twenties to a trough of 3.28 at age 38 before recovering to 3.62 by age 48. The motherhood premium is thus generated almost entirely by this non-mother dip rather than by elevated well-being among mothers. The concentration of the deficit within the fertility window, and its resolution as normative expectations relax, is consistent with a time-limited misalignment process and difficult to reconcile with stable personality selection— which predicts a constant gap across ages— or with chronic reproductive grief, which predicts persistence rather than recovery at older ages. The hump-shaped age profile is robust to including grandparents in the parent definition, which produces a smoother but qualitatively identical pattern.

Racial Heterogeneity as a Boundary Condition. The preceding analyses document a robust motherhood premium in experienced well-being and trace it to the non-parent side of the comparison. A natural question is whether this pattern generalizes across social groups. The sample includes 170 Black women with clearly defined parental status (110 mothers, 60 non-mothers), permitting an exploratory examination of racial heterogeneity in the motherhood–well-being association. Given the modest subsample size, these analyses should be interpreted as suggestive.

A pronounced boundary condition emerges (SI Appendix Table S6). Among non-Black women, motherhood is associated with a large and statistically significant increase in average experienced mood ($b = 0.112$, $p < 0.01$). Among Black women, the corresponding association is negative but imprecisely estimated ($b = -0.100$, $p = 0.293$). The Black × Parenthood interaction is large and statistically significant

($b = -0.212$, $p < 0.05$), indicating that the motherhood premium observed among non-Black women is effectively absent for Black women.

To identify where this interaction originates, I decompose the parent–non-parent contrast into covariate-adjusted mean well-being separately for Black and non-Black mothers and non-mothers, expressing the interaction as the difference between the racial gap among mothers and the racial gap among non-mothers. This decomposition reveals that the interaction in average mood is driven primarily by differences in the counterfactual experience of non-parenthood: Black and non-Black mothers exhibit similar adjusted mood levels ($b = 0.04$, $p = 0.54$), whereas Black non-mothers report substantially higher mood than non-Black non-mothers. Roughly four-fifths of the −0.21 interaction is attributable to this baseline difference among non-mothers. The disappearance of the motherhood premium for Black women thus reflects the absence of a large non-parent well-being penalty rather than lower well-being among Black mothers themselves. At the same time, the emotional correlates of motherhood differ across groups: among non-Black women, motherhood is associated with significant reductions in worry and loneliness and increases in love, while these affective shifts are largely absent among Black women. That the boundary condition is generated primarily by the absence of a non-parent penalty, rather than by lower maternal well-being, is consistent with the possibility that alternative social structures sustain well-being among Black women who remain childless, a possibility examined in the Discussion.

Distributional Compression of Affect and Social Structure. The analyses thus far have focused on how parenthood is associated with the central tendency of experienced well-being and its variation across life-course circumstances. I now shift attention from the mean to the distribution, asking whether parenthood also compresses the cross-sectional dispersion of emotional experience and social organization across individuals. If parenthood operates as a coordinating institution—channeling adults into a more constrained set of roles and routines—it may narrow the range of daily experience, rendering the lives of parents more similar to one another than those of non-parents. The analysis focuses on custodial parents with minor children, the margin along which daily routines and responsibilities are most plausibly reorganized.

Emotional Compression. Heteroskedastic regression models in which custodial parenthood enters both the conditional mean and the conditional variance of participant-level well-being reveal significant compression (SI Appendix Table S7). For average mood, custodial parenthood is associated with a variance ratio of approximately 0.90 ($p = 0.065$) — a 10 percent reduction in between-person variance, roughly comparable to marriage and employment. Compression is substantially larger for net emotion, with a variance ratio of 0.63 ($p < 0.001$) — a 37 percent reduction that exceeds both marriage and employment. These distributional shifts carry welfare implications. Evaluated using an Atkinson inequality index ($\varepsilon = 0.5$) — a measure that captures welfare-relevant inequality rather than dispersion alone — custodial parenthood is associated with a 43.1 percent reduction in inequality in net emotion, exceeding educational attainment (6.3 percent), employment (15.6 percent), and marriage (28.4 percent). That parenthood rivals or exceeds marriage and employment as a compressor of emotional experience — and that this compression translates into large reductions in welfare inequality — underscores that a conventional focus on averages may miss a central feature of how parenthood reshapes adult well-being.

Social Compression. To assess whether parenthood compresses the structure of daily social life, I measure the diversification of each individual's time allocation across social contexts using Shannon entropy — a summary statistic, borrowed from information theory, that captures how evenly time is spread across categories. An individual who divides waking hours equally among many social contexts (time alone, with partner, with friends, with coworkers) has high entropy; one whose social life is dominated by a single context has low entropy. Compression is then measured by the degree to which custodial parents cluster more tightly around a common entropy level than observationally comparable non-parents, with negative coefficients indicating that parents' social structures are more similar to one another than non-parents' social structures are to one another.

This analysis reveals a sharply domain-specific pattern (SI Appendix Table S8). Custodial parenthood is not associated with reduced dispersion in activities or spatial movement, but is associated with large and robust compression in the social domain ($b = -0.039$, $p < 0.001$), with significant effects among both mothers ($b = -0.030$, $p < 0.01$) and fathers ($b = -0.041$, $p < 0.001$). To put the magnitude in perspective, the pooled coefficient represents approximately 10 percent of the non-parent standard deviation in social entropy. Notably, parents exhibit *higher* mean social entropy than non-parents (1.52 vs. 1.27), indicating more diversified social portfolios on average— they interact with a wider range of social partners—but the between-person dispersion in entropy is substantially compressed (SD = 0.27 among parents vs. 0.39 among non-parents). Parenthood thus channels adults toward a common, more diversified social structure rather than a uniform one. This compression is not a mechanical artifact of increased child exposure. When child interactions are excluded from the entropy measure, compression persists ($b = -0.023$, $p < 0.05$), concentrated among fathers ($b = -0.035$, $p < 0.05$); excluding both children and partners yields a similar pattern ($b = -0.025$, $p < 0.05$), again driven by fathers ($b = -0.039$, $p < 0.05$). Social convergence thus extends beyond caregiving to reflect broader constraints on the social opportunity set, concentrated in discretionary and relational domains — time alone, with partners, and with friends — while structurally heterogeneous domains such as work exhibit expansion (SI Appendix Table S9). The gendered sources map onto distinct coordination mechanisms: motherhood regularizes the child-centered core of family life, while fatherhood constrains the periphery of male social interaction, particularly friendships.

Coupling of Compression. Person-level analyses suggest that these two forms of compression are coupled, with a pronounced gender asymmetry. Among women, the interaction between social regularization and motherhood is positive and significant ($b = 0.147$, $p = 0.016$): a one-standard-deviation increase in social regularization is associated with a 0.085 SD reduction in residual emotional volatility among mothers ($p = 0.038$), an association that persists when social distance is anchored relative to other mothers ($b = 0.093$, $p = 0.028$) and that is absent among non-mothers. Among men, the pattern reverses: the interaction is negative and significant ($b = -0.132$, $p = 0.027$), and father-specific slopes are small and insignificant ($b = -0.056$, $p = 0.174$). This asymmetry helps explain the gendered volatility findings reported earlier: motherhood stabilizes emotional life in part through the social reorganization it imposes, whereas fatherhood imposes comparable social structure without a corresponding emotional dividend — suggesting that fathers' elevated volatility arises from sources that structured daily routines cannot reach. Results are robust to the use of momentary mood rather than net emotion (SI Appendix). Parenthood thus functions less as a uniform hedonic shock than as a coordinating institution that narrows the range of daily experience for both mothers and fathers, but through different channels: for mothers, social and emotional regularization are tightly linked; for fathers, emotional compression occurs through mechanisms not captured by the social reorganization of daily life.

Generalizability: Gallup World Poll Replication. A natural question is whether the patterns documented using momentary experience sampling replicate under different measurement paradigms, at population scale, and during more recent time periods. To assess external validity, I replicate the paper's three central empirical signatures — the motherhood premium, distributional compression, and preparedness heterogeneity — using roughly two decades of U.S. data from the Gallup World Poll (GWP), which elicits life satisfaction on a 0–10 scale (Cantril ladder) and binary indicators for whether the respondent experienced specific emotions (including happiness and sadness) during the prior day. Custodial parenthood is proxied by the presence of a child under 15 in the household, with conservative adjustments based on household size and respondent age to reduce misclassification of parents whose children have aged out of the roster (SI Appendix). The GWP relies on single-shot retrospective assessments — the very measures this paper argues attenuate experiential effects — but its scale, national representativeness, and two-decade span provide a demanding benchmark for generalizability. The sample is restricted to adults aged 21–55 to further limit age-dependent misclassification. Net affect is computed as the happiness indicator minus the sadness indicator, yielding scores of +1, 0, or −1.

Gendered mean effects. Custodial motherhood is associated with a statistically significant increase in life satisfaction ($b = 0.142$, $p = 0.05$), while fatherhood shows no corresponding increase ($b = -0.022$, $p = 0.73$), reproducing the gender asymmetry observed in the experience-sampling data (SI Appendix Table S10). The motherhood association is approximately one-half the magnitude of the marriage association for women in GWP, paralleling the proportional relationship observed in the primary data. A similar asymmetry appears for remembered net affect, with custodial motherhood associated with higher net affect ($b = 0.067$, $p = 0.01$) while the corresponding estimate for fathers is small and insignificant ($b = 0.022$, $p = 0.33$).

Preparedness heterogeneity. Paralleling the primary analysis, I construct a propensity-based preparedness index from gender-specific logistic regressions of custodial parenthood on the available GWP covariates and stratify the sample into terciles. The parenthood–life satisfaction association increases monotonically across terciles, yielding a 0.25-point increase from the lowest to highest ($p = 0.057$). The gradient's direction and approximate magnitude are consistent with the experience-sampling results, reinforcing that parent–non-parent contrasts are systematically organized by life-course alignment across both measurement paradigms.

Distributional compression. Finally, heteroskedastic models indicate significant compression of life satisfaction among custodial parents equivalent to a variance ratio of 0.92 ($p = 0.01$), similar to the compression of mood from the main data. A qualitatively similar though weaker pattern exists for remembered net affect (ratio: 0.94, $p = 0.10$).

The persistence of all three signatures across a nationally representative sample using conceptually distinct measures supports the generalizability of the core findings beyond the primary dataset.

The Lossy Mapping Between Experiential and Evaluative Well-Being. The analyses thus far document meaningful associations between parenthood and experienced well-being that are substantially larger than those typically reported using evaluative measures. A natural question is whether this divergence can be explained empirically. To investigate, I examine the mapping between experiential and evaluative well-being using data from the American Time Use Survey (ATUS), which uniquely in 2012 and 2013 — a period contemporaneous with the primary dataset — elicited both a global life satisfaction measure and prior-day experienced well-being from the same respondents. I summarize experienced well-being using indices of positive experience (happiness and meaning) and negative experience (sadness, stress, tiredness, and pain).

Figure 5 (Panel A) plots average standardized life satisfaction against deciles of experienced well-being. A distinctive feature of the figure is that gradients are shallow, falling below the 45-degree line: large differences in daily experience translate into comparatively modest differences in life satisfaction. A regression of life satisfaction on positive and negative experience yields an R^2 of 0.18. Restricting the sample to days the respondent rated as "typical" raises the R^2 only marginally, to 0.19, indicating that the shallow pass-through is not an artifact of experiential reports drawn from unrepresentative days. This limited overlap is expected given that life satisfaction incorporates inputs beyond recent affect — including values, social comparisons, and future expectations — but it establishes that the pass-through from experience to evaluation is low, so that sizable experiential shifts can leave weak evaluative traces. Regressions indicate that the mapping is asymmetric: negative experience receives significantly greater weight ($\beta = -0.30$, $p < 0.01$) than positive experience ($\beta = 0.24$, $p < 0.01$; $p < 0.001$). This asymmetry implies that diffuse positive experiential gains, of the kind documented for motherhood above, are particularly attenuated in their evaluative translation.

Panel B applies this estimated transfer function to the covariate-adjusted experiential shift associated with parenthood. As reported in SI Appendix Table S11, parenthood is associated with a large and highly significant increase in positive experienced well-being ($\Delta z = 0.233$, $p < 0.001$) with no corresponding change in negative experience ($\Delta z = -0.016$, $p = 0.47$). Passing these experiential shifts through the Panel A transfer function yields an implied life satisfaction increase of 0.061 standard deviations— a statistically significant but modest effect ($p = 0.003$) that closely matches the observed evaluative estimate ($b = 0.063$, $p = 0.003$). This quantitative correspondence suggests that the moderate evaluative signal in ATUS is not evidence of a small experiential change; it is the predictable consequence of a large experiential change compressed through a shallow, asymmetrically weighted mapping. In this sense, the near-null evaluative effects widely reported in the literature need not reflect an absence of experiential gains, they are quantitatively compatible with the more substantial experiential associations documented in this paper.[2]

Robustness. Given the non-experimental identification in this paper, I assess the sensitivity of the primary findings to unobserved confounding, alternative selection accounts, and reporting artifacts. Although the covariate set is unusually rich for a well-being study— encompassing household net worth and spousal employment alongside standard demographics— the cross-sectional design cannot rule out selection into parenthood on unobserved traits that also affect experienced well-being. To assess how severe such confounding would need to be, I compute coefficient stability bounds following Oster (2019), which compare the degree of coefficient movement and R^2 change produced by observed covariates to the movement that would be required from unobserved confounders to eliminate the estimated effect. These bounds indicate that unobserved selection would need to be approximately 1.6 times as strong as the full observed covariate set to eliminate the motherhood mood premium, and more than seven-fold to eliminate the peak-mood association (SI Appendix Table S12). The peak-mood result is highly robust to unobserved confounding; the average mood result is more sensitive but would still require selection on unobservables substantially exceeding selection on observables — a threshold that is demanding given the breadth of the covariate set.

[2] I refer to this mapping as lossy, borrowing from information theory where the term describes a transformation that discards information from the input signal. The characterization is empirical rather than normative; it describes the documented properties of the mapping without implying that evaluative measures are deficient.

Several features of the data argue against simple positive selection, in which inherently happier individuals disproportionately become parents. Under this account, the preparedness–mood gradient should be steeper among parents. Instead, the gradient is statistically indistinguishable across parents and non-parents among women (difference in slopes = −0.01, $p = 0.95$) and is significantly steeper among non-parents than parents among men (difference = 0.23, one-sided $p = 0.04$). Moreover, the age decomposition reported in Figure 3 shows that the motherhood premium is confined to the normative fertility window and absent among older women, a pattern more consistent with time-limited misalignment than with stable personality differences, though the concurrent operation of these mechanisms cannot be fully excluded without direct measures of fertility intentions.

Finally, measurement non-invariance is unlikely to account for the documented demographic heterogeneity. Gender differences in momentary mood remain stable after conditioning on emotional expression patterns and individual baselines (Bhargava, 2024), arguing against reporting artifacts as a primary driver of the gendered parenthood associations. The racial results pose an even steeper challenge for a reporting account. Baseline differences in scale use between Black and non-Black respondents are absorbed by the race main effect; the race × parenthood interaction would require the more specific assumption that motherhood differentially shifts reporting behavior by race. Moreover, because the interaction is driven predominantly by the non-mother side— Black non-mothers report higher mood than non-Black non-mothers, while reports among mothers are similar across race— a reporting artifact would need to explain this specific pattern rather than a general racial difference in scale use.

Discussion

Using over one million in-the-moment reports of mood, emotion, and time use, this study revisits the parenthood paradox from the perspective of experienced well-being. The analyses show that parenthood is associated with substantial changes in lived experience—including a sizable increase in average momentary mood among mothers and greater emotional volatility among fathers—patterns that are largely obscured in conventional evaluative measures of well-being. Two further signatures extend beyond these mean effects: systematic heterogeneity organized by life-course preparedness and distributional compression of emotional experience and social time use. All three patterns replicate in nationally representative Gallup data. I discuss each in turn before addressing limitations and implications.

Parenthood as Gendered Reorganization of Experience. These findings represent, to my knowledge, the first characterization of the parenthood–well-being association using high-frequency momentary mood — the closest empirical analogue to experienced utility as defined by Kahneman, Wakker, and Sarin (1997). Prior experience-sampling studies have relied on episode-level or end-of-day affect reconstructions and have generally reported small or context-dependent average effects (Nelson et al., 2013; Negraia & Augustine, 2020; Musick, Meier, & Flood, 2016). By measuring well-being continuously across daily contexts, the present data reveal that parenthood is associated with distinct and sharply gendered experiential signatures.

The emotional composition of these effects is theoretically informative. Motherhood operates through affective reweighting — more love, less loneliness and worry, offset by less hope — a compositional pattern more consistent with identity-based accounts in which caregiving restructures emotional priorities (Nomaguchi & Milkie, 2020) than with hedonic accounts that treat children as a simple source of pleasure or displeasure. Fatherhood's signature — elevated exhaustion, frustration, and volatility, particularly among unmarried men — extends evidence of paternal stress during caregiving (Offer & Schneider, 2011; Musick, Meier, & Flood, 2016) by demonstrating that these costs permeate the broader texture of daily emotional life rather than remaining confined to child-related episodes.

Life Course Alignment and the Non-Parent Counterfactual. A central contribution of this paper is demonstrating that the parent–non-parent well-being gap is systematically organized by life-course preparedness and that this gradient is driven primarily by the non-parent side of the comparison. This finding has several implications for how parenthood effects are conceptualized and estimated.

First, it reframes the heterogeneity widely documented in the literature. The well-being consequences of parenthood appear to be structured by the same life-course conditions that shape the propensity for parenthood itself—such as age-normative timing and partnership status—so that the average effect estimated in any given study reflects how that study samples across conditions of experiential surplus and deficit. From this perspective, the search for a single "parenthood effect" is conceptually misleading, and cross-study variability is better understood as evidence of differing degrees of life-course alignment rather than as statistical noise.

Second, the preparedness gradient identifies a mechanism that standard designs typically overlook. The three-channel decomposition shows that the gradient does not arise from differential enjoyment of children or from favorable time reallocations. Instead, it reflects broad differences in non-child experienced well-being. Highly prepared non-parents report lower mood across work, leisure, and domestic activities—not only in moments when children might be salient—indicating a diffuse experiential deficit rather than episodic distress. The affective profile of this deficit—declining positive emotion alongside relatively stable negative emotion—resembles what the mental health literature terms languishing (Keyes, 2002), a state characterized by diminished positive engagement rather than the presence of acute suffering. This interpretation is consistent with life-course misalignment accounts (Elder, 1998; Settersten & Mayer, 1997) and with evidence that childlessness can carry psychological costs when it conflicts with internalized expectations (McQuillan et al., 2012), though alternative explanations—including involuntary childlessness, relationship dissatisfaction, or unobserved personality traits—cannot be ruled out.

The age decomposition provides suggestive evidence for this interpretation. The motherhood premium emerges, peaks, and dissipates in close correspondence with the normative fertility window—absent at age 23, maximal around 38, and largely resolved by the late forties. This temporal pattern places meaningful constraints on competing explanations. Stable dispositional accounts would predict a relatively constant gap across ages, while chronic reproductive grief would imply persistence or intensification at older ages. Instead, the hump-shaped profile aligns closely with life-course theory: the experiential costs of remaining childless appear to rise as the biological and social window for parenthood narrows and to attenuate once it closes. Notably, the temporal pattern is generated almost entirely by movements in non-mothers' mood, while maternal well-being remains comparatively stable across age. This suggests that what appears as a motherhood premium may in practice reflect a childlessness penalty concentrated during the years when the absence of children is most counter-normative.

Finally, the gradient reveals that the sources of life-course misalignment differ across gender. For women, the preparedness gradient is primarily structured by age-normative timing: being off-schedule relative to the biological and cultural clock appears to carry experiential costs independent of partnership or material resources. For men, the gradient is organized largely by partnership status. The absence of a committed partnership—which both predicts childlessness and may independently reduce well-being—appears to be the principal dimension of misalignment. This distinction suggests that the social meaning of childlessness is itself gendered. The factors that structure misalignment for women differ from those that structure it for men, even though both ultimately generate similar gradients in the parent–non-parent gap. Recognizing this distinction may help explain why the moderating role of marriage has appeared so inconsistent across prior studies.

A Racial Boundary Condition. A suggestive but theoretically important boundary condition emerges for Black women, among whom the motherhood premium is absent. This divergence arises primarily from differences among non-mothers. Black non-mothers do not exhibit the well-being deficit observed among non-Black non-mothers, while Black and non-Black mothers report similar levels of average mood. The racial asymmetry therefore appears to reflect the absence of a non-parent penalty rather than diminished well-being among Black mothers.

One interpretation is that the mechanisms linking life-course alignment to well-being operate differently across social contexts. Scholarship on Black family life emphasizes that adult identity and social integration are often sustained through extended kin networks and community institutions that do not depend on nuclear-family parenthood (Stack, 1974; Burton & Tucker, 2009; Cross, 2018). If so, the experiential costs of remaining childless may be attenuated in settings where kinship, caregiving, and social belonging are less tightly organized around biological motherhood. At the same time, structural conditions under which many Black women parent—including economic precarity, health disparities, and institutional stressors (Geronimus, 1992; Dow, 2019)—may constrain the affective returns to motherhood itself. The combination of weaker childlessness penalties and reduced caregiving returns would jointly erase the premium observed among other groups.

This interpretation is necessarily tentative given the modest subsample (n = 170) and the complexity of racial variation in family meaning. Nevertheless, the pattern highlights an important implication of the life-course alignment framework: the psychological consequences of alignment depend on the social institutions that structure both parenthood and childlessness. Findings derived from racially undifferentiated samples may therefore obscure meaningful variation in how life-course expectations translate into lived experience.

Parenthood as a Coordinating Institution. Beyond mean effects, the distributional analyses reveal that parenthood compresses the cross-sectional variation in both emotional experience and social organization across individuals, with welfare-relevant effects on inequality that rival or exceed those of marriage. This finding adds a dimension largely absent from the parenthood–well-being literature, which has focused almost exclusively on central tendency. Parenthood does not simply shift where people are in the well-being distribution; it narrows the distribution itself, rendering the daily lives of parents more similar to one another than those of observationally comparable non-parents.

That this compression is concentrated in the social domain—particularly in discretionary and relational time—rather than in activities or spatial movement is consistent with classic accounts of social roles as coordinative frameworks (Goffman, 1959). Parenthood constrains the social opportunity set, especially for fathers, whose peer interactions are most compressed. The coupling between social and emotional regularization observed among mothers—but not fathers—suggests that maternal well-being may be partially stabilized through the very social constraints that parenthood imposes, while paternal emotional instability arises from sources less tightly linked to social organization.

The Evaluative-Experiential Divergence. Deaton and Stone (2014) were among the first to demonstrate that evaluative and hedonic measures yield divergent conclusions about parenthood but left open the question of whether this divergence could be explained by the functional properties of the evaluative mapping itself. The ATUS analysis addresses this directly, showing that the mapping from daily experience to global evaluation is lossy— shallow in gradient and asymmetrically weighted toward negative states— so that diffuse positive experiential gains associated with motherhood are substantially attenuated in evaluative translation. The quantitative calibration confirms that a modest evaluative signal is consistent with meaningful experiential improvement passed through this mapping. Evaluative and experiential measures are thus best understood as complementary rather than competing: each captures distinct aspects of parenthood's consequences, and the apparent paradox arises partly because the literature has relied disproportionately on evaluative measures whose functional form systematically attenuates the specific experiential changes parenthood produces.

Limitations. Several limitations warrant consideration. First, the design is observational and cross-sectional with respect to parenthood status. Causal interpretation therefore rests on conditional independence given observed covariates—an assumption supported by the richness of the covariates and by Oster (2019) sensitivity bounds, but not directly testable. The preparedness gradient requires an additional assumption that preparedness would shift well-being similarly for future parents and non-parents in the absence of children—a parallel trends condition supported by several diagnostics but not definitively distinguishable from alternative explanations such as involuntary childlessness, relationship distress, or unobserved personality differences. Direct measures of fertility intentions, relationship quality, and reproductive constraints would substantially strengthen this interpretation. In addition, some covariates used to construct the preparedness index—particularly partnership status and employment—may themselves be influenced by parenthood, meaning the alignment analyses may partially reflect post-parenthood structural changes rather than selection alone. The racial analyses should likewise be interpreted as exploratory given the modest subsample.

Second, although participants were recruited from a nationally representative panel, the demands of a high-compliance diary protocol may influence sample composition. The data were also collected between 2010 and 2014, prior to major shifts in parenting culture, labor markets, and remote work arrangements. While the structural patterns documented here likely reflect relatively stable features of family life, and the Gallup replication spanning two decades provides reassurance, replication using contemporary experience-sampling data would strengthen external validity.

Finally, the five-point mood scale imposes ceiling constraints that could contribute mechanically to the observed compression in mood variance. However, the substantially larger compression observed for net emotion—measured across a wider effective range—suggests that the distributional patterns documented here reflect genuine convergence in emotional experience rather than scale artifacts alone.

Implications. These findings carry implications for both family policy and the measurement of well-being in policy evaluation. With respect to family policy, the time-use decomposition indicates that simply increasing the quantity of parent–child time is unlikely to raise experienced well-being. Hedonic gains from child-centered activities are largely offset by reductions in leisure and partner time. Policies that reduce the structural costs of these tradeoffs—flexible scheduling, affordable childcare, paid leave, and support for relationship maintenance—are therefore more likely to improve daily well-being without requiring parents to do more parenting.

The distributional results introduce a second dimension. Custodial parenthood compresses the dispersion of experienced well-being, reducing inequality in daily emotional experience even when mean effects are modest. In this sense, parenthood operates as a partial hedonic equalizer: individuals' emotional lives become more similar once coordinated around shared caregiving roles and routines. Policy environments that impose additional unpredictability on parents—unstable work schedules, unreliable childcare, or housing instability—may erode these coordinating benefits by increasing volatility and inequality in lived experience.

The preparedness gradient further implies that part of the observed parent–non-parent contrast reflects experiential costs borne by adults whose circumstances predict family formation but who remain childless—whether by choice or constraint. In the context of sustained fertility decline, this suggests that pronatalist policies focused exclusively on reducing the costs of parenthood address only one side of the well-being equation. Reducing stigma around diverse family trajectories, strengthening alternative sources of social integration and purpose for childless adults, and expanding access to fertility treatment for those who desire children may attenuate the gap from the non-parent side. The racial boundary condition underscores that these dynamics are embedded in broader social structures: communities that sustain adult identity and belonging through extended kin networks and non-nuclear family ties may naturally buffer against the costs of childlessness, pointing toward structural rather than purely individual policy approaches.

Finally, the lossy mapping between experienced and evaluative well-being implies that policies generating diffuse experiential gains may be systematically attenuated when evaluated using life satisfaction alone. This argues for incorporating experiential measures alongside evaluative metrics in policy contexts where the texture of daily life is the target of intervention.

Conclusion. The parenthood paradox—the apparent disconnect between the centrality of reproduction in human life and the weak well-being effects reported in decades of empirical research—arises in part from how well-being is measured and in part from how the counterfactual to parenthood is conceptualized. When well-being is observed at the resolution of daily experience, a more structured pattern emerges than evaluative measures alone suggest: parenthood reorganizes emotional life in gendered ways, apparent premiums are often generated by experiential deficits among highly prepared non-parents, and custodial parenthood compresses the distribution of lived experience across individuals. These findings suggest that the consequences of parenthood cannot be understood solely by asking whether moments with children are joyful—often they are—but by examining how institutions, norms, and social supports organize the adult role of raising them.

Methods

For additional details please refer to the SI Appendix. *Baseline Associations.* The association between parenthood and experienced well-being is estimated using participant-level linear regressions:

$$w_i = \alpha + \gamma \, \text{Parent}_i + X_i'\phi + \varepsilon_i,$$

where w_i is average experienced well-being for individual i (mean momentary mood across the ten-day diary window), $Parent_i$ indicates parenthood exclusive of grandparenthood, and X_i includes age (entered flexibly), education, marital status, employment status, race and ethnicity, household income, household net worth, spousal employment, and state and wave fixed effects. Analogous specifications are estimated for arousal, peak mood (an indicator for reporting the maximum mood value), a net count of positive and negative emotion, individual discrete emotions, and within-person emotional volatility (the variance of momentary mood across prompts). All models are estimated separately by gender unless otherwise noted, with standard errors clustered at the participant level. Alternative parenthood definitions are examined in SI Appendix Table S2.

Life-Course Preparedness. The preparedness index is derived from gender-specific logistic regressions of parenthood on the full covariate set X_i: $\hat{p}_i = Pr(Parent_i = 1 \mid X_i)$. The resulting predicted probabilities are used to stratify the sample into gender-specific terciles across which I examine heterogeneity in the parenthood–well-being association. Analyses are restricted to observations in the estimation support of the gender-specific propensity models. The index is strongly predictive: 92% of individuals in the highest tercile are parents, compared with 22% in the lowest. To identify which dimensions organize heterogeneity, I sequentially exclude each covariate from the propensity model and re-estimate the preparedness gradient while retaining the full covariate set in the outcome regression.

Within-Person Affective Returns. To estimate the hedonic return to child presence, I use individual fixed-effects models estimated among parents:

$$w_{it} = \alpha_i + \beta \,\text{WithChild}_{it} + \pi\,(\text{WithChild}_{it} \times \text{Father}_i) + \theta Z_{it} + \varepsilon_{it},$$

where α_i are individual fixed effects and Z_it includes hour-of-day and day-of-week indicators. The coefficient β captures within-person mood changes associated with child presence for mothers; π captures the differential return for fathers. This specification is extended by interacting child presence with preparedness terciles to test whether affective returns vary with life-course alignment. Standard errors are clustered at the participant level.

Time-Use Decomposition. I summarize daily time use into six composite categories designed to isolate nuclear-family social configurations while grouping all non-family time by activity type: child-only time (any activity), triadic time with child and partner (any activity), partner-only time (any activity), non-family work time, non-family domestic time, and non-family leisure/alone time (see SI Appendix for explicit categorization rules). To assess whether time-use shifts account for observed well-being differences, I implement Gelbach (2016) decompositions that partition the change in the parenthood coefficient between a baseline and time-use-adjusted specification into additive contributions from each category. To characterize non-child experienced well-being, I recover covariate-adjusted predicted mood within each non-child context for parents and non-parents separately by gender and preparedness tercile.

Weekend and Age Decompositions. To test whether the non-parent penalty intensifies under institutional coordination demands, I estimate within-person fixed-effects models comparing day-level mood across weekdays and weekends separately for parents and non-parents in the highest preparedness tercile. To examine the temporal profile of the motherhood premium, I estimate the baseline specification with a fully saturated age-band × motherhood interaction, yielding age-specific parenthood effects and covariate-adjusted mean mood for mothers and non-mothers at each age midpoint, evaluated at a common covariate profile. Alternative weekend definitions and the inclusion of grandparents in the age decomposition are reported in the SI Appendix.

Racial Heterogeneity. To examine racial differences in the parenthood–well-being association, I interact parenthood with an indicator for Black race:

$$w_i = \alpha + \gamma_1 \text{ Parent}_i + \gamma_2 \text{ Black}_i + \gamma_3(\text{Parent}_i \times \text{Black}_i) + X_i'\phi + \varepsilon_i.$$

The coefficient γ_3 captures the Black–non-Black difference in the parent–non-parent contrast. To locate this interaction in the conditional mean structure, I decompose it into the difference between the racial gap among mothers and the racial gap among non-mothers — an accounting identity that holds mechanically given the estimated model.

Emotional Compression. To test whether custodial parenthood compresses the between-person distribution of experienced well-being, I estimate heteroskedastic regression models:

$$y_i = \alpha + \pi \text{ Parent}_i + X_i'\phi + \varepsilon_i, \varepsilon_i \sim N(0, \sigma_i^2),$$

$$\log(\sigma_i^2) = \delta + \theta \text{ Parent}_i.$$

The conditional mean and variance are estimated jointly by maximum likelihood. The exponentiated coefficient exp(θ) yields a variance ratio; values below one indicate compression. To benchmark magnitudes, I compute the proportional reduction in a standard Atkinson inequality index ($\varepsilon = 0.5$) associated with custodial parenthood and compare it to other life-course institutions (marriage, employment, educational attainment). Models are estimated in pooled samples and separately by gender.

Social Convergence. To assess convergence in daily social structure, I compute Shannon entropy for each individual across social context categories:

$$H_i = -\sum_{k=1}^{K} p_{ik} \log p_{ik},$$

where p_ik is the share of waking intervals spent in social category k. Convergence is measured by regressing the absolute deviation of each individual's entropy from a covariate-predicted baseline (excluding parenthood) on custodial parenthood status; negative coefficients indicate that parents cluster more tightly around a common social structure. Entropy is computed over the full social vector and re-computed excluding child and partner categories to assess whether convergence extends beyond nuclear-family interactions. Category-level decompositions are reported in the SI Appendix.

Supplementary Analyses. Methods for the Gallup World Poll replication, the ATUS evaluative-experiential mapping, the coupling between social and emotional regularization, and all robustness procedures (including Oster (2019) sensitivity bounds and measurement invariance diagnostics) are described in the SI Appendix.

Acknowledgments

I thank Jim Spieth for the data and Karim K. Kassam, George Loewenstein, and Carey Morewedge for insight into an earlier version of this project.

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Figures

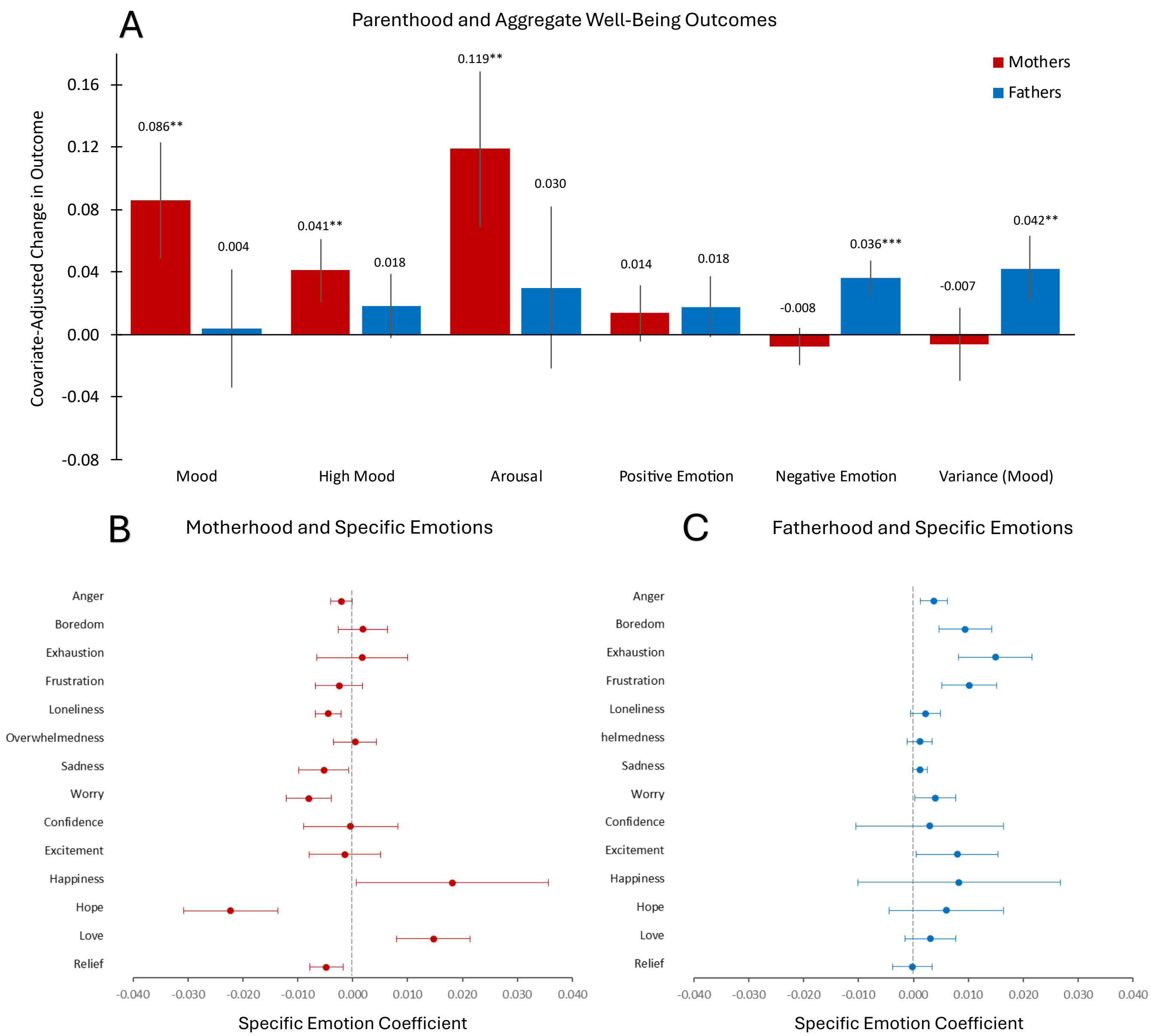

Fig. 1. Parenthood and Experienced Well-Being
This figure plots well-being coefficients from regressions of well-being outcomes on parenthood, conditioned on demographic covariates, estimated separately for mothers and fathers. Error bars represent +/- one standard error. Statistical significance is denoted as (***, p < 0.01; **, p < 0.05). Panel A depicts mood, arousal, and aggregate emotion outcomes. Panels B and C depict specific emotion outcomes for women and men, respectively.

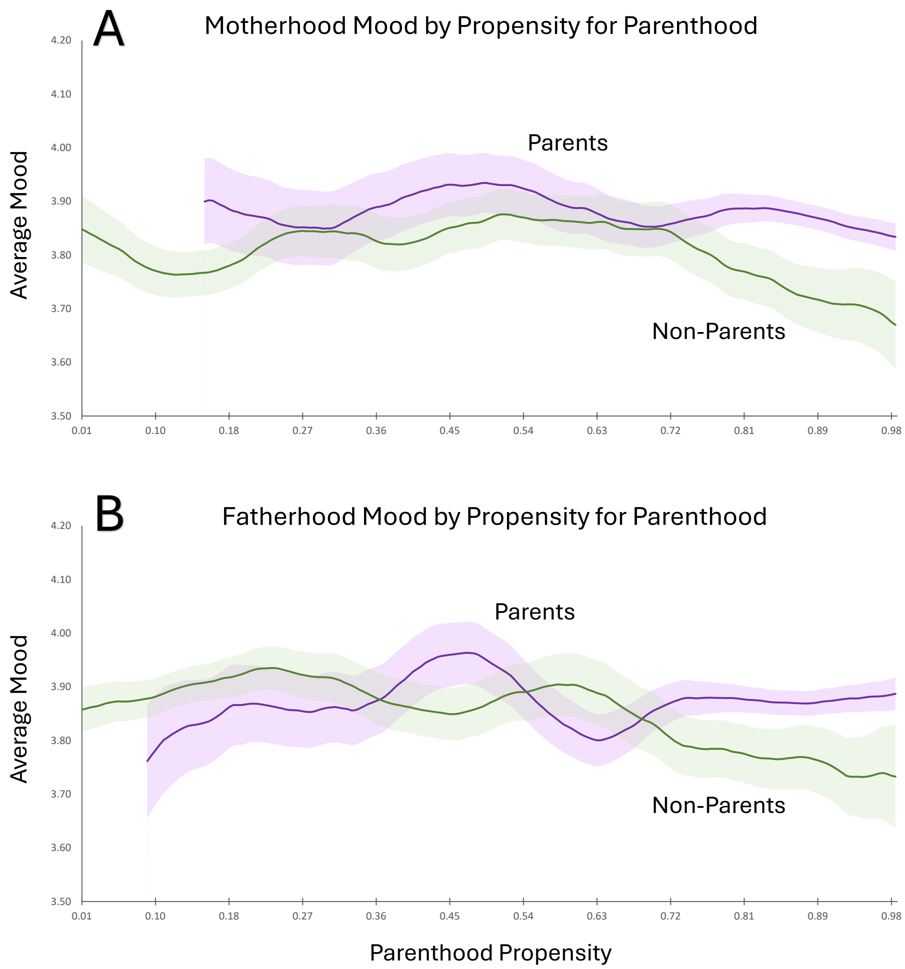

Figure 2. Parenthood and Mood by Parental Propensity
This figure depicts kernel-smoothed local means of covariate-adjusted experienced mood by parenthood propensity across women (Panel A) and men (Panel B) and parenthood. The gender-specific propensity score is estimated from pre-parenthood demographic and socioeconomic characteristics, and mood is residualized on the same covariates and recentered at the gender-specific mean. Local means are computed using a triangular kernel with bandwidth, h=0.15; curves are displayed only where there is sufficient support (i.e., at least 40 observations within the kernel window), and shaded regions denote +/- 1 standard error around the local means.

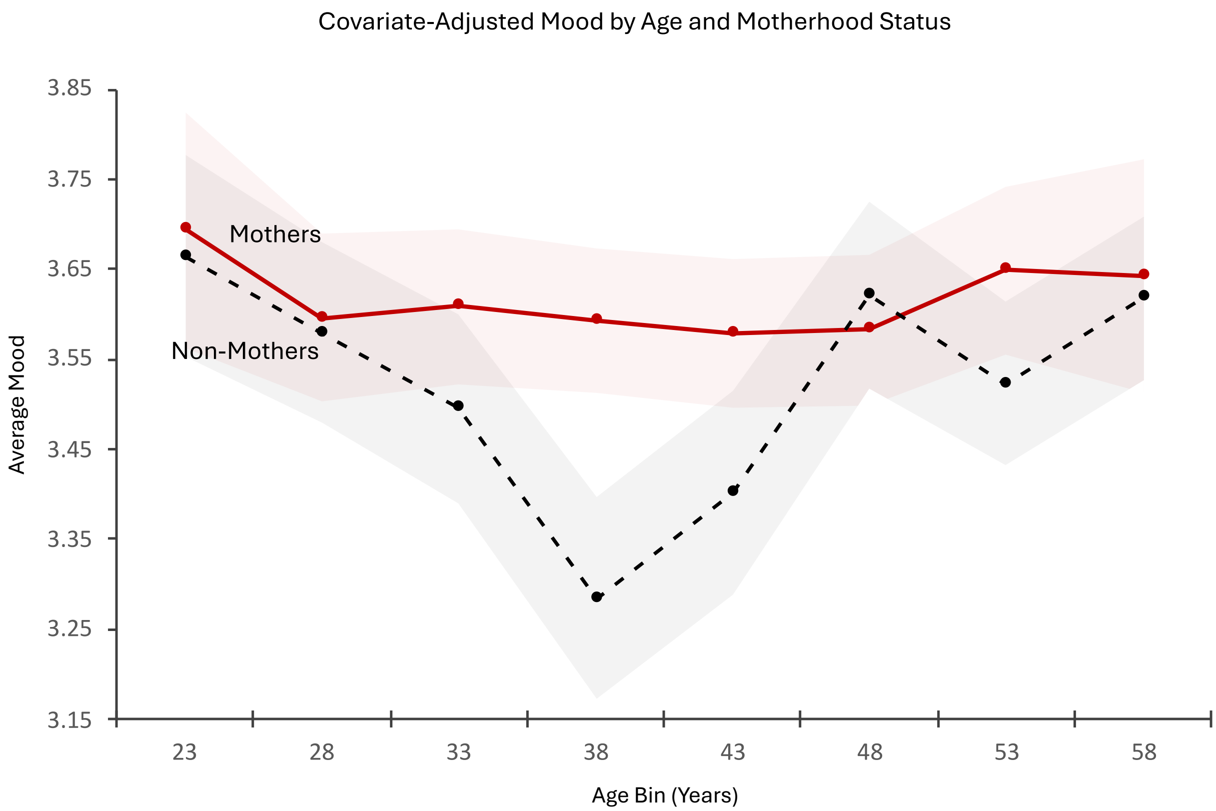

Figure 3. Covariate-adjusted mood by age and motherhood status.
The figure plots predicted mean experienced mood for mothers (solid red) and non-mothers (dashed black) across age bins (midpoints shown on the x-axis), estimated from a fully saturated age-band × motherhood regression among women ages 21–58 (with the 63-year bin pooled into the 58-year bin). Mood is adjusted for demographic and socioeconomic covariates, with predictions evaluated at a common covariate profile derived from the estimation sample. Shaded regions denote +/- one standard error around the predicted mean. The vertical distance between the two curves corresponds to the estimated motherhood effect at each age. The midlife premium peaks in the late 30s and is driven primarily by a pronounced dip in mood among non-mothers rather than a sharp increase among mothers.

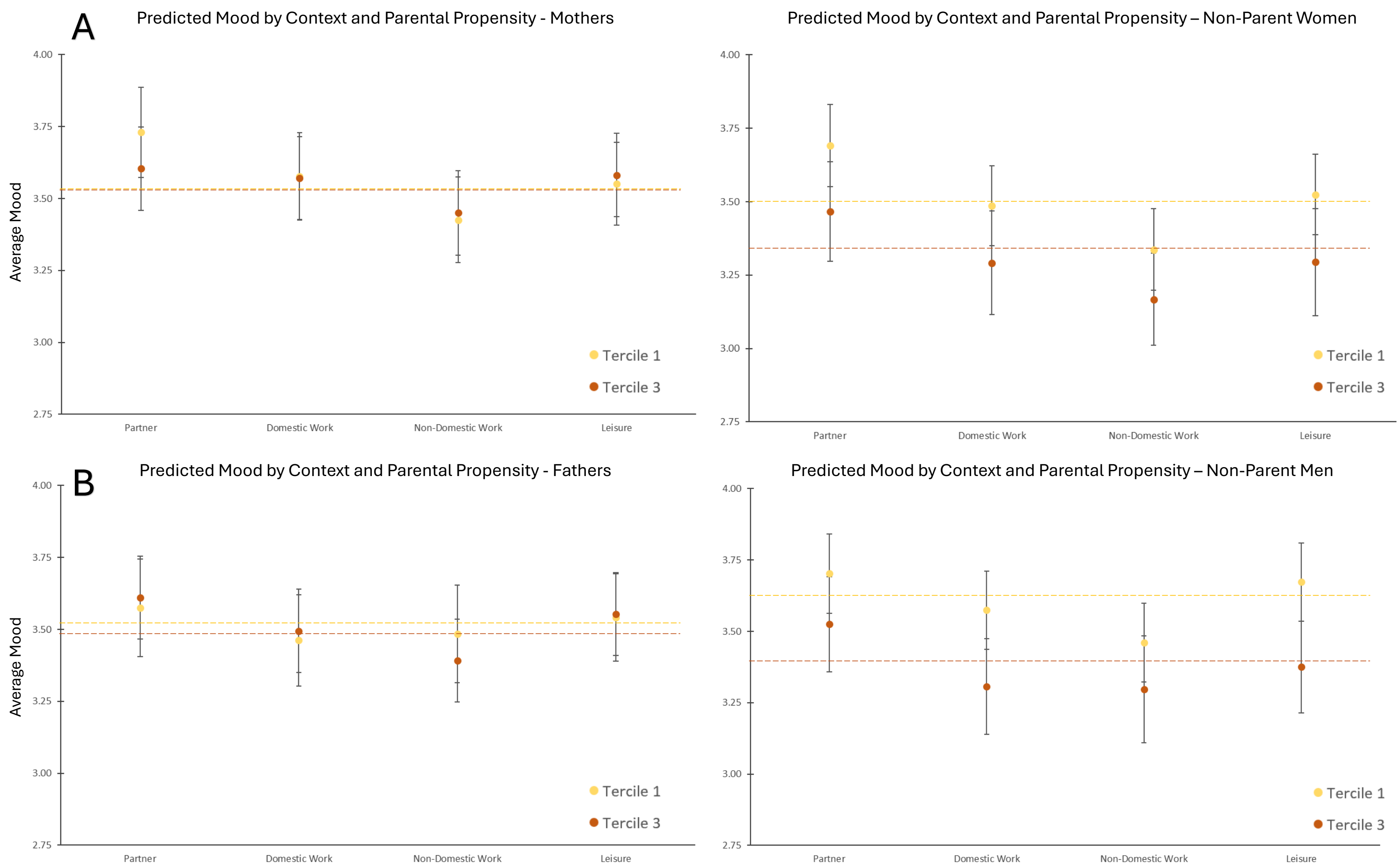

Figure 4. Predicted Mood by Time-Use Context and Parental Propensity
Panels A (women) and B (men) plot predicted experienced mood by non-child time-use context—partner time, domestic work, non-domestic paid work, and leisure—separately for parents and non-parents, and for the lowest (T1) and highest (T3) terciles of a gender-specific parenthood propensity score. Predictions are obtained from saturated interaction regressions of momentary mood on context, parenthood status, gender, and preparedness tercile, controlling for demographic, socioeconomic, temporal, and state fixed effects. Points denote context-specific predicted mood levels; error bars denote ±1 standard error. Horizontal dashed lines indicate the average predicted mood across non-child contexts within each group and preparedness tercile, weighted by the group-specific distribution of time use across contexts.

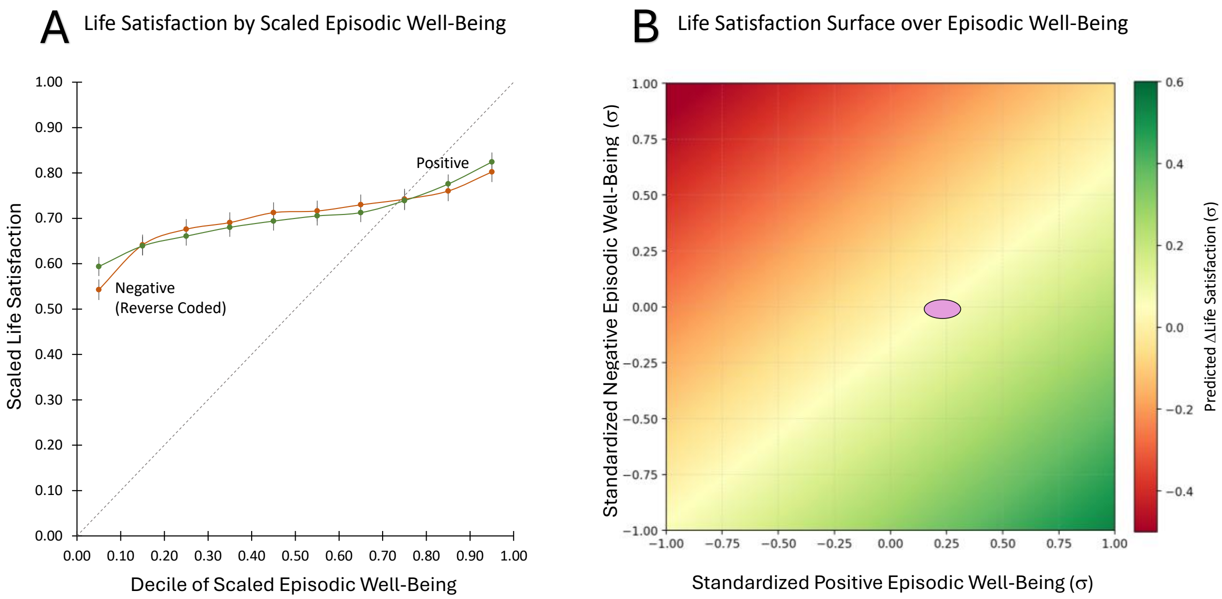

Fig. 5. Life Satisfaction as Lossy Compressor of Episodic Experience
This figure illustrates how life satisfaction functions as a lossy, asymmetric aggregator of experienced well-being using data from the 2012 and 2013 ATUS. Panel A plots average scaled life satisfaction against decile bins of positive and negative experienced well-being on the prior day; error bars represent ±1 standard error. Panel B presents the model-based evaluative surface mapping standardized life satisfaction over a wide range of standardized positive and negative experienced well-being. The purple ellipse denotes the joint 95% confidence region for the covariate-adjusted parental shift in experienced well-being, with coordinates given by the estimated effects of parenthood on standardized positive ($\beta = 0.233$) and negative ($\beta = -0.016$) experienced well-being. Projecting this region through the evaluative surface yields a modest predicted increase in life satisfaction, reflecting substantial attenuation and a negativity bias in evaluative aggregation.

PNAS

Supporting Information for
The Parenthood Paradox and the Hedonic Costs of Life-Course Misalignment

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This PDF file includes:

Supporting text
Figure S1
Tables S1 to S12
Exhibit S1
SI References

This appendix provides supplementary details on data collection, variable definitions, estimation procedures, and additional results referenced in the main text. Analysis code is available on OSF.

A. Data Source and Collection Procedures

The primary data, known as USA Touchpoints (TP), were collected between 2010 and 2014 through a collaboration between the Coalition for Innovative Media Measurement and the Media Behavior Institute. The data were originally collected to measure cross-platform media consumption, daily activities, and the social and environmental contexts in which they occur. Participants were unaware of the present research purpose.

Data were collected using a mobile diary application that prompted participants to record their circumstances every 30 minutes during waking hours over a 10-day period. At each prompt, participants completed a sequence of screens recording: (i) location from a predefined set of categories; (ii) current activities; (iii) social context, indicating who was present (alone, spouse/partner, children, siblings/other family, coworkers, friends, parents, pets, or others); and (iv) emotional experience and well-being. The application also passively recorded media and online activity data that are not analyzed in the present study. An end-of-day module captured purchasing behavior and advertising exposure and was likewise excluded from analysis.

Participants could retroactively complete diary entries for up to two preceding intervals. The data do not distinguish between contemporaneous and retrospective entries, although the short recall window limits potential recall bias. Data were retained for participant-days with at least 16 half-hour diary entries and a completed end-of-day survey. Participants who failed to meet these criteria for at least 7 of the 10 diary days were excluded.

B. Sample Recruitment

Participants were recruited from a nationally representative commercial panel maintained by GfK MRI comprising approximately 26,568 U.S. adults. The analytic dataset contains 3,775 English-speaking adults aged 21–65 residing in the contiguous United States who had previously completed a GfK MRI personal interview and associated questionnaires. Recruitment occurred in four waves of approximately 1,000 participants each: August 2011–January 2012 (wave 1), February–July 2012 (wave 2), July 2012–January 2013 (wave 3), and July 2013–March 2014 (wave 4). Participants received incentives of $100–$150 and were provided with dedicated mobile devices where necessary.

Compliance was high. Based on recruitment targets, approximately 97% of recruited participants produced qualifying diary data. The average participant completed 292 diary reports across 9.83 days, yielding more than one million momentary observations. The final sample reflects broad demographic diversity: 49% female; mean age 44.7 (SD = 11.8); 67% non-Hispanic White, 13%

Black, 9% Hispanic, and 11% other race/ethnicity; 58% married or engaged; 72% employed; and median household income category $60,000–$75,000.[1]

C. Variable Definitions

C.1 Well-Being Measures

At each diary prompt, participants rated their current mood on a 5-point scale ranging from 1 ("bad mood") to 5 ("good mood"). For participant-level analyses, mood is averaged across all diary prompts within the observation window. Peak mood is defined as an indicator for reporting the maximum value (5) at a given prompt; the participant-level measure is the proportion of prompts at which peak mood occurs.

Participants also rated arousal on a 5-point scale ranging from 1 ("relaxed") to 5 ("alert"), elicited on the same screen as the mood measure.

Participants indicated the presence of specific emotions by selecting from a set of emoji-labeled indicators. Fourteen emotions were consistently elicited across all waves: anger, bored, confident, excited, exhausted, frustrated, happy, hopeful, lonely, loving, overwhelmed, relieved, sad, and worried. Options for *indifferent*, *interested*, and *content* were excluded due to ambiguous valence or inconsistent interpretation. Additional emotions introduced in later waves were excluded to maintain cross-wave comparability.

Using these responses, I construct prompt-level emotion indices. Positive emotion equals the number of positive emotions reported (happy, confident, excited, hopeful, loving, relieved), while negative emotion equals the number of negative emotions reported (anger, bored, exhausted, frustrated, lonely, overwhelmed, sad, worried). *Net emotion* is defined as the difference between positive and negative counts at each prompt and is averaged across prompts for participant-level analyses.

Within-person mood volatility is measured as the variance of momentary mood across prompts within the diary window.

C.2 Parenthood Definitions

Parenthood status is constructed using two data sources: (i) self-reported household composition from recruitment questionnaires, which includes the presence and ages of children residing in the household, and (ii) supplementary survey items indicating the presence of non-custodial children and grandparental status. Using these sources, approximately 75% of respondents can be clearly classified as either parents or non-parents, with most remaining cases identified as grandparents.

The baseline definition used throughout the paper codes parenthood = 1 for any respondent reporting biological, adopted, or step-children, excluding respondents identified as grandparents.

[1] Age reflects the average for participants in the first three waves as wave 4 age is recorded categorically.

This definition captures both custodial and non-custodial parents while avoiding the misclassification of grandparents as parents.

Alternative definitions examined in Table S2 vary along two dimensions. The first concerns custodial status. *Custodial minor parenthood* restricts the definition to respondents with at least one child under age 18 residing in the household. Additional specifications distinguish partnered custodial parents from solo custodial parents based on marital or cohabiting status.

The second dimension concerns the treatment of grandparents. In supplementary analyses, grandparenthood is examined as a separate category, and an inclusive definition combining parenthood and grandparenthood is also estimated.

For the distributional compression analyses, parenthood is defined as custodial parenthood of at least one minor child, the margin along which daily routines and responsibilities are most plausibly reorganized.

C.3 Demographic and Socioeconomic Covariates

Background characteristics were obtained from extensive questionnaires administered during panel recruitment. The covariate set used in the primary regressions includes demographic, socioeconomic, household, and geographic characteristics.

Age is entered flexibly using indicators corresponding to approximately five-year age bands. Gender is binary (male/female) based on self-reported sex; non-binary categories were not available in the dataset. Race and ethnicity are classified as non-Hispanic White, Black, Hispanic, Asian, and Other (including American Indian/Alaska Native and respondents reporting multiple races).

Educational attainment is captured using indicators for less than high school, high school graduate, some college, college graduate, and post-college education. A continuous measure of years of education is also available. Marital status distinguishes married or engaged, separated or divorced, and never married. Employment status is classified as full-time employed, part-time employed, unemployed, student, retired, and not in the labor force. Household income is reported in categorical ranges from less than $2,500 to $250,000+ and entered as category indicators in regressions. Household net worth is similarly measured in four categories ranging from less than $100,000 to $500,000+. Spousal employment is captured by a binary indicator for whether the respondent’s spouse or partner is employed.

Geographic and temporal variation are captured using state-group indicators and survey wave indicators.

C.4 Time-Use Classification

Each diary prompt records both social context (who is present) and activity indicators. Because respondents may report multiple activities and social partners simultaneously, each episode is assigned to a single mutually exclusive category using a hierarchical classification that prioritizes family configurations. This hierarchy ensures that episodes involving children and partners are consistently identified even when additional activities or social contacts are reported.

The classification proceeds in two stages. First, episodes are sorted by social context. Episodes in which both a partner and children are present are classified as triadic time, regardless of activity or other social contacts. Among the remaining episodes, those with children present but no partner are classified as child-only time, and those with a partner present but no children are classified as partner-only time. In both cases, the presence of additional contacts (e.g., friends or relatives) is permitted.

Second, episodes involving neither a partner nor children are classified by activity type using a fixed priority ordering. Episodes involving paid employment or formal study are classified as non-family work time. Among the remaining episodes, those involving unpaid household production, caregiving, household management, or errands are classified as non-family domestic time. All remaining episodes—including socializing with non-family others, entertainment, hobbies, exercise, relaxing, self-maintenance, and time spent alone—are classified as non-family leisure/alone time.

The resulting six categories—child-only, triadic, partner-only, non-family work, non-family domestic, and non-family leisure/alone—are exhaustive and mutually exclusive across waking diary prompts and form the basis of the time-use decomposition and within-person analyses reported in the main text. For the social convergence analyses, entropy is computed using a broader set of non-hierarchical social categories described in Section D.7.

D. Supplementary Estimation Details

D.1 Life-Course Preparedness Index

The preparedness index is constructed from gender-specific logistic regressions of the baseline parenthood indicator on the full covariate set described above. Models are estimated separately for women and men to allow the determinants of parenthood to differ by gender. The predicted probability from this model summarizes the degree to which each individual's observable demographic and socioeconomic characteristics align with conditions under which parenthood is normatively expected.

Predictive performance. The index strongly predicts parenthood status. Among women, 93% of individuals in the highest preparedness tercile are parents compared with 27% in the lowest tercile. Among men, the corresponding figures are 91% and 16%.

Common support. Preparedness distributions overlap substantially between parents and non-parents within each gender. Tercile cutpoints are defined within gender to ensure approximately equal-sized groups.

Sequential exclusion procedure. To identify which covariate dimensions organize the preparedness gradient, each covariate block (age, marriage, education, income/wealth, race, spousal employment, own employment, state, and survey wave) is sequentially excluded from the propensity model and the preparedness gradient (T3 − T1 parenthood effect on mood) is re-estimated. The outcome regression retains the full covariate set throughout; thus changes in the gradient reflect the reassignment of individuals across preparedness terciles rather than changes in covariate adjustment. Results are reported in Table S3.

D.2 Within-Person Affective Returns to Child Presence

Within-person affective returns to child presence are estimated using individual fixed-effects models among parents. The dependent variable is momentary mood at each diary prompt. The baseline specification includes an indicator for the presence of children, interacted with gender to allow different returns for mothers and fathers, along with controls for hour of day and day of week. Individual fixed effects absorb all time-invariant respondent characteristics.

To test whether affective returns vary with life-course alignment, the model is extended by interacting the child-presence indicator with preparedness tercile indicators. Standard errors are clustered at the individual level in all specifications.

D.3 Gelbach (2016) Decomposition

The Gelbach (2016) decomposition partitions the change in the parenthood coefficient between a baseline specification (conditioning only on demographic covariates) and a time-use-augmented specification (additionally conditioning on time-use category indicators) into additive contributions from each time-use category.

Formally, the contribution of category j equals the product of (i) the coefficient on category j in the augmented regression and (ii) the coefficient from an auxiliary regression of time in category j on parenthood and demographic covariates. The decomposition is exact and invariant to the order in which mediators are entered.

Decompositions are estimated separately for men, all women, and high-preparedness individuals (top tercile of the gender-specific parenthood propensity score). Time-use categories include dyadic child time, triadic time, dyadic partner time, non-domestic work, domestic work, and leisure including solo time. Standard errors are obtained via nonparametric bootstrap with 300 replications. Results are reported in Table S4.

D.4 Weekend and Age Decompositions

Weekend analysis. Within-person fixed-effects models compare day-level mood across weekdays and weekends separately for parents and non-parents in the highest preparedness tercile, pooling men and women. The dependent variable is person-day average mood. Three weekend definitions are examined: (i) Saturday–Sunday, (ii) Friday–Sunday, and (iii) Saturday–Sunday relative to Monday–Thursday (excluding Friday).

Gap attenuation is defined as the difference between non-parent and parent weekend effects. Positive values indicate that non-parents experience a larger weekend mood rebound, partially narrowing the parent–non-parent gap on weekends. Complementary between-person models condition on the full covariate set and include a parenthood × weekend interaction. Results are reported in Table S5.

Age decomposition. To examine the temporal profile of the motherhood premium among women, the baseline specification is estimated with a fully saturated age-band × motherhood interaction. Age bands correspond to five-year intervals centered at midpoints from 23 to 58, with the 63-year bin pooled into the 58-year bin due to small cell sizes.

The model yields age-specific parenthood effects and covariate-adjusted mean mood for mothers and non-mothers at each age midpoint. Predicted values are evaluated at a common covariate profile derived from estimation-sample means for continuous variables and modal values for categorical variables.

D.5 Racial Heterogeneity Decomposition

Racial heterogeneity is examined among women by estimating a regression of average mood on parenthood, a Black race indicator, their interaction, and the full covariate set (Table S6). The interaction coefficient captures the Black–non-Black difference in the parent–non-parent contrast.

To locate the source of this interaction within the conditional mean structure, the estimate is decomposed into the difference between the racial gap among mothers and the racial gap among non-mothers—an accounting identity implied by the regression specification. Covariate-adjusted mean mood is therefore computed separately for four groups: Black mothers, non-Black mothers, Black non-mothers, and non-Black non-mothers.

The interaction can then be expressed as: (Black mother − non-Black mother) − (Black non-mother − non-Black non-mother). The relative contribution of each component is reported as the share of the total interaction attributable to differences among mothers versus differences among non-mothers.

D.6 Individual Discrete Emotions by Preparedness Tercile

Among women, preparedness is associated with higher reported hope ($T3 - T1 = 0.043$, $p = 0.09$) and lower reported worry ($T3 - T1 = -0.028$, $p = 0.08$) among parents relative to non-parents. Across genders, the preparedness gradient widens the parent–non-parent gap for excitement (average $T3 - T1 = 0.030$, $p = 0.03$) and narrows it for anger (average $T3 - T1 = -0.008$, $p = 0.08$). Preparedness is additionally associated with relatively greater worry among fathers compared with non-fathers ($T3 - T1 = 0.017$, $p = 0.055$). No other discrete emotions exhibit statistically meaningful differences across preparedness terciles.

D.7 Preparedness Gradient for Positive and Negative Emotion

SI Appendix Figure S1 plots covariate-adjusted positive and negative emotion as functions of preparedness, separately for parents and non-parents. The widening parent–non-parent gap at high preparedness corresponds primarily to divergence in positive emotion, while negative emotion remains comparatively flat across the preparedness distribution. Among women, the preparedness gradient for net emotion (T3 − T1 = 0.231, p = 0.032) is at least as large as the corresponding gradient for average mood and exhibits the same sequential exclusion structure. Among men, the gradient attenuates for net emotion (T3 − T1 = 0.094, p = 0.431).

D.8 Heteroskedastic Regression and Atkinson Index

Heteroskedastic models. The conditional mean and conditional variance of participant-level well-being are estimated jointly by maximum likelihood, with custodial parenthood entering both equations. The variance equation is specified as $\log(\sigma^2_i) = \delta + \theta \cdot \text{Parent}_i$, where $\exp(\theta)$ represents the variance ratio; values below one indicate compression of the well-being distribution among parents. All models condition on the full covariate set and are estimated both in pooled samples and separately by gender. Results are reported in Table S7.

Atkinson inequality index. To benchmark the magnitude of distributional compression, I compute the Atkinson inequality index (Atkinson, 1970) with inequality-aversion parameter $\varepsilon = 0.5$ using covariate-adjusted participant-level well-being, separately for parents and non-parents. Because the Atkinson index requires strictly positive inputs, covariate-adjusted residuals are recentered at the sample mean and a small constant is added to ensure strict positivity for all observations. The proportional reduction in inequality associated with parenthood is computed as 1 − A(parents) / A(non-parents), where A(·) denotes the Atkinson index. The same procedure is applied to other life-course institutions—marriage, employment, and educational attainment—to provide benchmark magnitudes. These calculations assume that momentary mood and net emotion satisfy the continuity and interpersonal comparability conditions required for Atkinson-based welfare comparisons.

D.9 Shannon Entropy and Social Convergence

Entropy computation. For each participant, Shannon entropy is computed over the distribution of waking half-hour intervals across social-context categories:

$$H_i = -\sum_k p_{ik} \log{(p_{ik})},$$

where p_{ik} denotes the proportion of intervals spent in social category k. Higher entropy indicates more diversified social time use. Entropy is first computed using the full social-context vector (alone, partner, children, other family, friends, coworkers, pets). It is then recomputed excluding child interactions and excluding both child and partner interactions to assess whether convergence extends beyond nuclear-family contact.

Convergence measure. To evaluate whether custodial parents converge toward a more similar social organization, I regress the absolute deviation of each individual's entropy from a covariate-predicted baseline on custodial parenthood status. The baseline is obtained from a regression of entropy on the full covariate set excluding parenthood; the predicted value represents each individual's demographically expected entropy. Negative coefficients on the parenthood indicator therefore indicate that parents cluster more tightly around a common social structure than observationally comparable non-parents.

Category-level decomposition.
To identify which social categories drive compression, I compute the absolute deviation of each individual's time share in each category from the covariate-predicted baseline share and regress these deviations on custodial parenthood. Results are reported in Table S9.

D.10 Social–Emotional Regularization and Within-Person Coupling

The analyses in the main text document social compression and hedonic compression as group-level phenomena. A natural question is whether these patterns are linked within individuals—specifically, whether individuals whose social lives are most regularized by parenthood also exhibit more stable emotional experience. To examine this possibility, I construct participant-level measures of emotional volatility and social structural deviation from a covariate-adjusted template. All analyses use the full ten-day diary window for each participant and are conducted separately by gender.

Emotional Regularization. Emotional regularization is operationalized using both total and context-adjusted volatility. For each participant i, total emotional volatility is defined as the within-person standard deviation of momentary net emotion across all diary prompts. Let w_{it} denote net emotion at time t and $\bar{w}_i$ the participant's mean across prompts. Total volatility is defined as:

$$\sigma_i^{total} = \sqrt{\frac{1}{T_i - 1}\sum_t (w_{it} - \bar{w}_i)^2}$$

where T_i is the number of prompts observed for participant i. Note that this measure uses the standard deviation rather than the variance employed for mood volatility in Section C.1, as the coupling analyses rely on standardized variables for which the distinction is immaterial.

To isolate volatility net of observable context, I estimate pooled participant fixed-effects regressions of momentary net emotion on time-of-day indicators, day-of-week indicators, and mutually exclusive activity categories corresponding to the primary time-use partitions used throughout the paper (child time, triadic interaction, partner-only time, non-family work, non-family domestic time, and solitary leisure). Specifically:

$$w_{it} = \alpha_i + \delta_{h(t)} + \gamma_{d(t)} + \theta' A_{it} + \varepsilon_{it}$$

where α_i are participant fixed effects, $\delta_{h(t)}$ and $\gamma_{d(t)}$ capture temporal structure, and A_{it} are activity indicators.

Residual emotional volatility is defined as the within-person standard deviation of the estimated residuals:

$$\sigma_i^{resid} = \sqrt{\frac{1}{T_i - 1}\sum_t \hat{\varepsilon}_{it}^2}$$

This measure captures emotional fluctuation not attributable to common contextual structure, reflecting emotional idiosyncrasy conditional on occupying similar contexts. Both volatility measures are standardized within gender prior to regression analysis.

Social Regularization. Social regularization is constructed using the distribution of mutually exclusive time-use categories across the diary window. For each participant, I compute the share of prompts spent in each social category, s_{ik}, where shares sum to one across categories.

To construct a reference social template that varies across demographic and socioeconomic characteristics but excludes parenthood, I estimate gender-specific regressions of each time-use share on the full set of background covariates (age group, education, marital status, employment status, race/ethnicity, household income, household net worth, spousal employment, state of residence, and survey wave). Parenthood is excluded from these regressions so that predicted shares, $\hat{s}_{ik}$, reflect expected social structure conditional on covariates but not on parental status.

Social distance is defined as the L_1deviation between observed and predicted shares:

$$D_i = \sum_k \mid s_{ik} - \hat{s}_{ik} \mid$$

This metric captures the total absolute deviation of a participant's observed social structure from the covariate-adjusted template. For interpretability, social regularization is defined as the negative of this distance ($-D_i$), so that higher values indicate greater alignment with the predicted template. The resulting measure is standardized within gender.

Coupling Specification. The primary coupling specification links emotional volatility to social regularization at the participant level. For each gender, I estimate:

$$\sigma_i = \alpha + \beta_1 SocialReg_i + \beta_2 Parent_i + \beta_3 (SocialReg_i \times Parent_i) + X_i' \phi + \varepsilon_i$$

where $\sigma_i$is standardized residual emotional volatility, $SocialReg_i$is standardized social regularization, $Parent_i$indicates custodial parenthood (excluding grandparents), and $X_i$contains the demographic and socioeconomic controls.

The coefficient $\beta_1$captures the association between social and emotional regularization among non-parents, while $\beta_3$captures the differential association among parents. I also estimate within-parent models restricted to the subsample of custodial parents to test whether parents whose social lives are more regularized relative to demographically similar peers exhibit greater emotional stability.

Parallel analyses using momentary mood rather than net emotion confirm the gender-asymmetric moderation pattern reported in the main text (women: $b = 0.125, p = 0.047$; men: $b = -0.145, p = 0.015$), though within-parent tests for mood are not statistically significant for either gender.

E. Gallup World Poll Replication

E.1 Data Overview

The Gallup World Poll (GWP) provides nationally representative survey data for the United States spanning 2006–2024. The survey elicits life satisfaction using the 0–10 Cantril ladder and asks respondents whether they experienced specific emotions during the prior day, including happiness and sadness. Net affect is defined as the happiness indicator minus the sadness indicator, yielding values of +1, 0, or −1. Respondents reporting both happiness and sadness receive a score of 0, as do those reporting neither emotion.

Custodial parenthood is proxied using the GWP indicator for the presence of a child under age 15 in the household. Because this definition is narrower than the baseline definition in the primary data, I implement an age-based reclassification procedure designed to limit age-dependent misclassification. Respondents whose household size exceeds the number implied by marital status alone— greater than two for married respondents, greater than one for unmarried respondents— but who report no children under 15 are retained as non-parents below age 35 (where such configurations most likely reflect non-parental households), treated as missing between ages 35 and 39 (where classification is ambiguous), and reclassified as parents at age 40 and above (where such configurations most plausibly reflect parents of older children). The analytic sample is restricted to respondents aged 21–55 to further limit misclassification arising from empty-nest households.

E.2 Estimation

All models condition on a covariate set paralleling the TP specification to the extent permitted by GWP: age (flexibly specified), gender, marital status, education, employment status, household income, and race/ethnicity.

I estimate gender-specific mean effects using separate regressions of life satisfaction and net affect on custodial parenthood and covariates. Preparedness heterogeneity is assessed using a propensity-based preparedness proxy constructed analogously to the TP preparedness index using the available GWP covariates. The sample is stratified into preparedness terciles, and the gradient in the parenthood–life satisfaction association (T3 − T1) is estimated across terciles.

To assess distributional compression, I estimate heteroskedastic regression models for life satisfaction and net affect in which custodial parenthood enters both the conditional mean and conditional log-variance equations. Results are reported in Table S10.

F. ATUS Evaluative-Experiential Mapping

F.1 Data Overview

The American Time Use Survey (ATUS) Well-Being Module, fielded in 2012 and 2013, uniquely elicited both global life satisfaction and episode-level affect ratings from the same respondents. The analytic sample is restricted to individuals aged 21–55 to parallel the prime-age focus of the main analyses and to reduce misclassification of parents whose children have left the household as non-parents. The final sample consists of 13,883 respondents.

Life satisfaction is measured using the ATUS ladder question on a 0–10 scale. Experienced well-being is summarized using two composite indices. Positive experience is defined as the mean of episode-level happiness and meaning ratings (0–6 scale), rescaled to the unit interval. Negative experience is defined as the mean of episode-level sadness, stress, tiredness, and pain ratings (0–6 scale), also rescaled to the unit interval. All variables are standardized for regression analyses.

F.2 Evaluative Mapping

To estimate the mapping from experienced well-being to evaluative well-being, I estimate the following linear model:

$$z(LS_i) = \beta_+ z(Pos_i) + \beta_- z(Neg_i) + \varepsilon_i$$

where $z(\cdot)$denotes standardization. Negativity bias is assessed by testing the restriction $|\ \beta_-\ | = \beta_+$, which corresponds to symmetric weighting of positive and negative experience. This specification provides a reduced-form approximation of how evaluative well-being aggregates underlying experiential states and therefore allows the experiential parenthood effects estimated in the main analysis to be translated into predicted changes in life satisfaction. As a robustness check, the sample is restricted to respondents who rated their diary day as typical. Results are reported in Table S11.

F.3 Structural Calibration

To assess whether the observed parenthood effects on life satisfaction are consistent with the estimated evaluative mapping, I estimate covariate-adjusted effects of parenthood on standardized positive experience, negative experience, and life satisfaction using pooled regressions that condition on demographic and socioeconomic controls paralleling the TP specification.

The implied change in life satisfaction predicted by the evaluative mapping is then computed as

$$\widehat{\Delta LS} = \hat{\beta}_{+} \cdot \Delta Pos + \hat{\beta}_{-} \cdot \Delta Neg$$

and compared to the directly estimated parenthood coefficient on life satisfaction.

G. Robustness and Sensitivity

G.1 Oster (2019) Sensitivity Bounds

To assess the sensitivity of the primary parenthood associations to unobserved confounding, I implement the procedure proposed by Oster (2019). For each outcome, I report the uncontrolled and controlled parenthood coefficients, the corresponding R^2values, and the proportional degree of selection on unobservables relative to observables (δ^*) required to drive the estimated coefficient to zero.

The assumed maximum $R^2$is set to $1.3 \times R^2_{controlled}$, following common practice in applications of the method. Bias-adjusted coefficients are also reported under assumed selection ratios of $\delta = 1$and $\delta = 2$. Results are presented separately for women (average mood, peak mood, arousal) and men (negative emotion, mood volatility) in Table S12.

G.2 Measurement Invariance

Observed gender differences in experienced well-being may partly reflect differences in reporting behavior rather than differences in underlying experience. Evidence from related analyses in Bhargava (2024) addresses this concern by conditioning momentary mood reports on emotional expression patterns and individual baselines. Gender differences in mood remain stable under those adjustments, suggesting that reporting artifacts are unlikely to account for the gendered parenthood associations documented here.

For the racial analyses, baseline differences in reporting are absorbed by the race main effect in the regression specification. The race × parenthood interaction would therefore require the stronger assumption that motherhood differentially shifts reporting behavior by race. The empirical pattern is inconsistent with such an explanation: the interaction arises primarily from differences among non-mothers (Black non-mothers reporting higher mood than non-Black non-mothers), while Black and non-Black mothers report similar levels of mood. In addition, directionally consistent interactions across binary emotion indicators—which are less sensitive to scale-use differences than the 5-point mood scale—provide further support for a substantive interpretation of the racial heterogeneity.

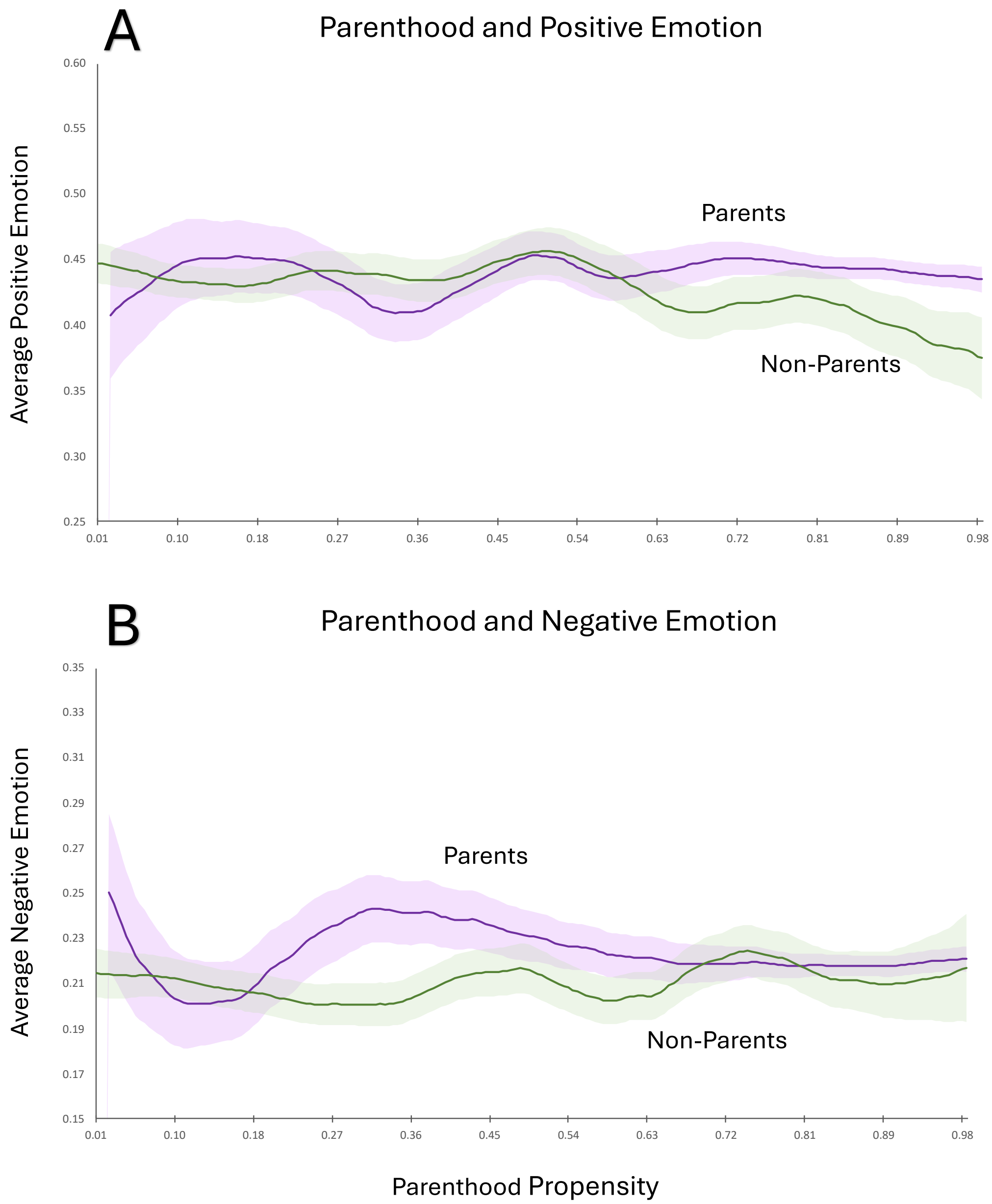

Fig. S1. Parenthood and Emotion by Parental Propensity

This figure displays kernel-smoothed local means of covariate-adjusted positive (Panel A) and negative (Panel B) emotion by estimated parenthood propensity for parents and non-parents, pooled across gender. The propensity score is estimated separately within gender using demographic and socioeconomic covariates and then combined for pooled analysis. Positive and negative emotion are residualized on the same covariates within gender and recentered at the corresponding gender-specific mean. Local means are computed using a triangular kernel with bandwidth h = 0.15. Curves are shown only where there is sufficient local support (at least 40 observations within the $\pm h$ window). Shaded regions denote ±1 standard error, computed using a Kish effective sample size under kernel weighting.

Table S1.
Covariate-Adjusted Reference Contrasts for Mood, Emotion, and Variance

Reference Comparison	Δ Mood	High Mood	Arousal	Δ Positive Emotion	Δ Negative Emotion	Δ (Within-Person) Mood Volatility	Variance Ratio - Mood	Variance Ratio - Net Emotion
Women								
Weekend (weekday)	0.107	0.049	-0.106	0.045	-0.032	-0.018	1.058	1.095
	(0.007)	(0.004)	(0.010)	(0.004)	(0.003)	(0.005)	(0.013)	(0.022)
Married/Engaged (never married)	0.170	0.050	-0.035	0.034	-0.036	-0.012	0.888	0.795
	-(0.041)	-(0.023)	-(0.055)	-(0.021)	-(0.013)	-(0.025)	(0.037)	(0.033)
Employed (unemployed)	0.043	-0.003	0.128	0.009	-0.032	-0.021	0.886	0.819
	-(0.062)	-(0.033)	-(0.091)	-(0.030)	-(0.019)	-(0.037)	(0.062)	(0.057)
Men								
Weekend (weekday)	0.107	0.043	-0.125	0.043	-0.020	-0.015	1.026	1.034
	(0.007)	(0.003)	(0.010)	(0.004)	(0.003)	(0.005)	(0.012)	(0.017)
Married/Engaged (never married)	0.035	-0.008	0.000	-0.023	-0.032	-0.044	0.913	0.829
	-(0.039)	-(0.021)	-(0.052)	-(0.019)	-(0.012)	-(0.021)	(0.035)	(0.032)
Employed (unemployed)	0.111	0.051	-0.008	0.020	0.014	0.036	0.905	0.847
	-(0.056)	-(0.029)	-(0.078)	-(0.028)	-(0.018)	-(0.029)	(0.059)	(0.055)

Notes: This table reports covariate-adjusted associations between selected reference comparisons and experienced well-being outcomes, estimated separately by gender using the USA Touchpoints data. Columns report estimated effects on average mood, probability of reporting the highest mood level (High Mood), arousal, aggregate positive and negative emotion, within-person mood volatility, and between-person variance ratios for mood and net emotion. Reference categories are shown in parentheses. For the weekend contrast, mean and volatility effects are estimated using individual fixed-effects models comparing the same individual across weekend and weekday observations. All other contrasts are estimated from participant-level regressions conditioning on the full set of demographic and socioeconomic controls described in the text, with employment indicators excluded from employment contrasts to avoid mechanical collinearity. Within-person mood volatility is the standard deviation of momentary mood across observations within each individual. Variance ratios are defined as the ratio of the between-person variance in the comparison group to that in the reference group, computed from covariate-adjusted person-level measures; values below one indicate compression (reduced dispersion) in the comparison group. Standard errors (in parentheses) are clustered at the individual level.

Table S2.
Association Between Parenthood and Average Mood Across Alternative Parenthood Definitions

	Covariate-Adjusted Change in Mood		
Parental Definition	Female	Male	Difference
Baseline Parenthood (excl. grandparents)	0.086	0.004	-0.082
	(0.037)	(0.038)	(0.053)
Custodial Minor Parenthood (excl. grandparents)	0.086	-0.021	-0.107
	(0.039)	(0.040)	(0.056)
Partnered Custodial Parenthood (excl. grandparents)	0.063	0.028	-0.035
	(0.054)	(0.047)	(0.071)
Solo Custodial Parenthood (excl. grandparents)	0.073	-0.118	-0.192
	(0.051)	(0.064)	(0.082)
Grandparenthood	0.086	0.058	-0.028
	(0.049)	(0.050)	(0.070)
All Parenthood (incl. grandparents)	0.082	0.018	-0.064
	(0.034)	(0.034)	(0.048)

Notes: This table reports the estimated association between parenthood and average momentary mood across alternative definitions of parental and family-structure status, estimated separately by gender using the USA Touchpoints data. Each row corresponds to a different parenthood definition. For each definition, participant-level regressions of average mood on the relevant parenthood indicator are estimated conditioning on the full set of demographic and socioeconomic controls described in the text. The Female and Male columns report the estimated difference in average mood between individuals in the specified category and observationally comparable non-parents of the same gender. The Difference column reports the male coefficient minus the female coefficient; its standard error is computed as the square root of the sum of the squared gender-specific standard errors, reflecting estimation on non-overlapping samples. Standard errors (in parentheses) are clustered at the individual level.

Table S3.
Sequential Exclusion Analysis of the Preparedness Gradient

Excluded Component	Mothers	Fathers
Full Model	0.218	0.204
Age	−0.050	0.152
Marriage	0.217	0.036
Education	0.224	0.208
Income/Wealth	0.273	0.113
Race	0.166	0.17
Spouse Employment	0.096	0.217
Own Employment	0.120	0.226
State/Wave Fixed Effects	0.126	0.222

Notes: This table reports the preparedness gradient — defined as the difference in the estimated parenthood association with average mood between the highest and lowest preparedness terciles (T3 − T1) — when each component block is sequentially excluded from the gender-specific propensity score used to construct preparedness terciles. The outcome regression retains the full set of demographic and socioeconomic controls described in the text throughout. The Full Model row reproduces the primary gradient reported in the main text. Each subsequent row reports the gradient when the indicated component is removed from the propensity score. Larger attenuation relative to the full model indicates that the excluded component is a primary organizing dimension of the preparedness gradient. Results are reported separately for women (Mothers column) and men (Fathers column).

Table S4.
Gelbach Decomposition of the Parenthood–Well-Being Association Across Time-Use Categories

	Women		Men		High Preparedness Tercile	
	Mood	Net Emotion	Mood	Net Emotion	Mood	Net Emotion
Unadjusted Gap	0.086**	0.022	0.004	-0.019	0.154**	0.096
	(0.037)	(0.037)	(0.035)	(0.035)	(0.063)	(0.060)
Time-Use Adjusted Gap	0.108*	0.015	0.022	-0.008	0.220**	0.116
	(0.055)	(0.051)	(0.046)	(0.051)	(0.085)	(0.084)
Dyadic Time with Children (no Partner)	0.223***	0.261***	0.011	0.025	0.063	0.050
	(0.076)	(0.072)	(0.033)	(0.031)	(0.062)	(0.047)
Triadic Time with Children and Partner	0.185***	0.191***	0.095*	0.136***	0.160	0.144*
	(0.049)	(0.047)	(0.053)	(0.051)	(0.100)	(0.081)
Dyadic Time with Partner (no Children)	-0.205***	-0.198***	-0.070	-0.096**	-0.220*	-0.170*
	(0.053)	(0.049)	(0.046)	(0.044)	(0.113)	(0.093)
Non-Domestic Work Time (no Children/Partner)	-0.013	-0.014	-0.002	-0.003	-0.014	-0.006
	(0.008)	(0.009)	(0.005)	(0.005)	(0.015)	(0.010)
Domestic Work Time (no Children/Partner)	-0.052**	-0.071***	-0.007	-0.020**	-0.026	-0.0270*
	(0.024)	(0.023)	(0.009)	(0.009)	(0.019)	(0.016)
Leisure and Solo Time (no Children/Partner)	-0.161***	-0.163***	-0.046*	-0.052*	-0.029	-0.011
	(0.049)	(0.050)	(0.027)	(0.027)	(0.034)	(0.027)
Total Time Use Contribution	-0.022	0.007	-0.018	-0.010	-0.066	-0.020
	(0.039)	(0.033)	(0.028)	(0.032)	(0.060)	(0.060)

Notes: This table reports a Gelbach (2016) decomposition of the association between parenthood and average momentary well-being into additive contributions from mutually exclusive time-use categories. Results are shown separately for men, women, and high-preparedness respondents, and for two outcomes: average mood and net emotion. The Unadjusted Gap row reports the estimated parenthood coefficient from a baseline regression conditioning on the full set of demographic and socioeconomic controls described in the text. The Time-Use Adjusted Gap row reports the parenthood coefficient after augmenting the specification with time-use indicators for each category. Subsequent rows report the contribution of each time-use category to the difference between the unadjusted and adjusted coefficients, representing the portion of the parenthood association attributable to differential time allocation in that category. The Total Time Use row sums these category-specific contributions. Time-use categories are mutually exclusive: dyadic child time (exclusive of partner), triadic time with child and partner, dyadic partner time (exclusive of child), non-domestic work, domestic work, and leisure including solo time. Standard errors are in parentheses, clustered at the individual level. Statistical significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table S5.
Differential Weekend Effects in Mood Among High-Preparedness Individuals

	Within-Person Estimates					Between-Person Estimates		
Weekend Definition	Non-Parent Gap	Parent Gap	Δ Gap Attenuation	SE	p-value	Δ Gap Attenuation	SE	p-value
Sat to Sunday	0.137	0.096	0.042	0.029	0.147	0.060	0.039	0.127
Fri to Sunday	0.141	0.086	0.056	0.027	0.037	0.093	0.036	0.010
Sat to Sunday (vs. Mon to Thurs)	0.161	0.107	0.054	0.031	0.081	0.083	0.042	0.050

Notes: This table reports differential weekend effects in average day-level mood among individuals in the top preparedness tercile, pooling men and women, using the USA Touchpoints data. Within-Person Estimates report the weekend mood effect estimated from individual fixed-effects regressions separately for non-parents (Non-Parent Gap) and parents (Parent Gap). Between-Person Estimates report the attenuation in the parent–non-parent mood gap from weekdays to weekends, estimated from pooled regressions including parenthood, weekend, and their interaction, conditioning on the full set of demographic and socioeconomic controls described in the text. Δ Gap Attenuation is the difference between the non-parent and parent weekend effects; positive values indicate that non-parents exhibit a larger weekend mood rebound, partially narrowing the parent–non-parent gap. Weekend definitions vary as indicated: "Sat–Sun" includes Saturday and Sunday; "Fri–Sun" includes Friday through Sunday; "Mon–Thu vs. Sat–Sun" excludes Friday and contrasts Monday–Thursday with Saturday–Sunday. All specifications use person-day mean mood as the dependent variable. Standard errors are clustered at the individual level.

Table S6.
Racial Heterogeneity in the Association Between Parenthood and Well-Being

	Mood		Variance(Mood)		Positive Emotion		Negative Emotion	
	Women	Men	Women	Men	Women	Men	Women	Men
Parent	0.112***	0.000	-0.012	0.033	0.024	0.017	-0.013	0.036***
	-(0.039)	-(0.039)	-(0.025)	-(0.021)	-(0.019)	-(0.020)	-(0.013)	-(0.011)
Black x Parent	-0.212**	0.039	0.042	0.098*	-0.083*	0.014	0.043*	0.006
	-(0.099)	-(0.112)	-(0.053)	-(0.055)	-(0.048)	-(0.054)	-(0.025)	-(0.031)
Net Effect of Black Parent	-0.100	0.039	0.030	0.131**	-0.059	0.031	0.030	0.042
p-value	0.293	0.720	0.553	0.018	0.207	0.562	0.206	0.165

Notes: This table reports heterogeneous associations between parenthood and experienced well-being by race, estimated separately by gender using the USA Touchpoints data. Outcomes are average mood, within-person mood variance, aggregate positive emotion, and aggregate negative emotion. The Parent row reports the covariate-adjusted association between parenthood and the outcome for non-Black respondents. The Black × Parent row reports the differential association for Black respondents relative to non-Black respondents. The Net Effect of Black Parent row reports the total estimated association for Black parents, computed as the linear combination of the Parent and Black × Parent coefficients; the corresponding p-value tests whether this net effect differs from zero. All regressions include a Black race indicator and the full set of demographic and socioeconomic controls described in the text. Standard errors (in parentheses) are clustered at the individual level. Statistical significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table S7.
Hedonic Compression Associated with Custodial Parenthood

	Mood			Net Emotion		
Dispersion Statistic	All	Women	Men	All	Women	Men
theta	-0.106	-0.047	-0.149	-0.470	-0.458	-0.434
	(0.057)	(0.085)	(0.082)	(0.057)	(0.084)	(0.082)
Variance Ratio	0.900	0.954	0.862	0.625	0.633	0.648
p-value of equivalence test	0.065	0.577	0.069	0.000	0.000	0.000

Notes: This table reports estimates of between-person compression in experienced well-being associated with custodial parenthood (at least one minor child in the household), estimated using the USA Touchpoints data. Outcomes are average mood and net emotion. Estimates are obtained from heteroskedastic regression models in which custodial parenthood enters both the conditional mean and the conditional log-variance of participant-level well-being, conditioning on the full set of demographic and socioeconomic controls described in the text. The coefficient θ captures the estimated difference in the log between-person variance for parents relative to non-parents; negative values indicate reduced dispersion among parents. The variance ratio is computed as exp(θ); values below one indicate compression. Reported p-values test the null hypothesis of equal dispersion (θ = 0). Results are reported for the pooled sample and separately by gender. Standard errors (in parentheses) are robust.

Table S8.
Structural Compression in Daily Time Use Associated with Custodial Parenthood

Domain	All	Mothers	Fathers
Activity	-0.002	0.006	-0.001
	(0.007)	(0.009)	(0.012)
Location	-0.0197**	-0.0153	-0.0222
	(0.009)	(0.013)	(0.014)
Social	-0.0392***	-0.0303***	-0.0406***
	(0.009)	(0.012)	(0.014)
Social excluding Children	-0.0228**	-0.0029	-0.0352**
	(0.010)	(0.014)	(0.015)
Social excluding Children / Partner	-0.0251**	-0.0122	-0.0393**
	(0.011)	(0.015)	(0.017)

Notes: This table reports estimates of structural compression in daily life across activity, spatial, and social domains associated with custodial parenthood (at least one minor child in the household), estimated using the USA Touchpoints data. For each domain, Shannon entropy is computed over the relevant category set to measure diversification in time allocation. The dependent variable is the absolute deviation of each individual's entropy from a covariate-predicted baseline estimated excluding parenthood; negative coefficients indicate that custodial parents cluster more tightly around a common time-use structure than observationally comparable non-parents. In the social domain, entropy is additionally computed over restricted category sets that progressively exclude child and partner presence to assess whether compression extends beyond nuclear-family interactions. All regressions condition on the full set of demographic and socioeconomic controls described in the text. Results are reported for the pooled sample and separately by gender. Standard errors (in parentheses) are clustered at the individual level. Statistical significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table S9.

Structural Compression by Social Time Use Category Associated with Custodial Parenthood

Domain	All	Mothers	Fathers
Alone	-0.0265***	-0.0324***	-0.0198**
	(0.006)	(0.008)	(0.010)
Partner	-0.0190***	-0.0182***	-0.0172***
	(0.004)	(0.007)	(0.006)
Other Family	-0.0073***	-0.0089**	-0.0058
	(0.003)	(0.004)	(0.004)
Friends	-0.0113***	-0.0058*	-0.0131***
	(0.002)	(0.003)	(0.003)
Coworkers	-0.0066**	-0.0041	-0.0087*
	(0.003)	(0.004)	(0.005)
Pets	-0.0395***	-0.0560***	-0.0197***
	(0.005)	(0.007)	(0.005)

Notes: This table reports estimates of structural compression in social time use across specific social-context categories associated with custodial parenthood (at least one minor child in the household), estimated using the USA Touchpoints data. For each category, the dependent variable is the absolute deviation of each individual's time share in that category from the covariate-predicted baseline estimated excluding parenthood; negative coefficients indicate reduced between-person heterogeneity among custodial parents. Categories are: time alone, with partner, with other family (parents and siblings combined), with friends, with coworkers, and with pets. All regressions condition on the full set of demographic and socioeconomic controls described in the text. Results are reported for the pooled sample and separately by gender. Standard errors (in parentheses) are clustered at the individual level. Statistical significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table S10.
Association Between Parenthood and Well-Being Across Datasets and Outcome Measures

	Women	Men	Scale
USA Touchpoints			
Mood	+0.086**	+0.004	Momentary [0 to 5]
High Mood	+0.041**	+0.018	Momentary [0,1]
Net Emotion	+0.022	-0.019	Momentary [1,0,-1]
Gallup World Poll			
Life Satisfaction	+0.142*	-0.022	Evaluative [0 to 10]
Net Emotion	+0.067**	+0.022	Remembered Affect [1,0,-1]

Notes: This table reports covariate-adjusted associations between parenthood and well-being across two datasets and multiple outcome measures, estimated separately by gender. USA Touchpoints outcomes are derived from experience-sampling data collected between 2010 and 2014. Mood is mean momentary mood (1–5 scale). High Mood is an indicator for reporting the maximum mood value at a given prompt. Net Emotion is computed at each prompt as the count of positive emotions reported minus the count of negative emotions reported, then averaged at the participant level. Gallup World Poll outcomes are derived from U.S. survey data spanning 2006 to 2024, restricted to respondents aged 21 to 55; custodial parenthood is defined by the presence of a child under 15 in the household. Life Satisfaction is measured on a 0–10 evaluative scale. Gallup Net Emotion is computed as remembered happiness minus remembered sadness based on binary indicators for each. All regressions condition on the full set of demographic and socioeconomic controls described in the text. Positive coefficients indicate higher well-being among parents relative to non-parents. The Scale column indicates the measurement type and range for each outcome. Statistical significance: * $p < 0.05$, ** $p < 0.01$.

Table S11.
Evaluative Well-Being as a Function of Experienced Well-Being in the ATUS

	Full Sample	Typical Sample
Panel A. Evaluative Mapping		
Positive Experience	0.239	0.249
	(0.008)	(0.010)
Negative Experience	-0.299	-0.324
	(0.008)	(0.010)
R-Squared	0.179	0.186
Panel B. Covariate Adjusted Effects of Parenthood		
Positive Experience	0.233	--
	(0.022)	
Negative Experience	-0.016	--
	(0.022)	
Life Satisfaction	0.063	--
	(0.021)	
Panel C. Structural Calibration		
Implied Change in Life Satisfaction	0.061	--
Observed Change in Life Satisfaction	0.063	--
	(0.021)	

Notes: This table examines the mapping from experienced to evaluative well-being using data from the 2012–2013 American Time Use Survey (ATUS) Well-Being Module, restricted to respondents aged 21 to 55. Panel A reports linear regressions of standardized life satisfaction on standardized positive and negative experienced well-being. Positive experience is the mean of episode-level happiness and meaning ratings (0–1 scale); negative experience is the mean of episode-level sadness, stress, tiredness, and pain ratings (0–1 scale). All variables are standardized. The Full Sample includes all respondents; the Typical Sample restricts to respondents who rated their diary day as typical. Panel B reports covariate-adjusted associations between parenthood and standardized positive experience, negative experience, and life satisfaction, estimated from pooled regressions conditioning on the full set of demographic and socioeconomic controls described in the text. Panel C applies the Panel A transfer function to the covariate-adjusted parenthood shifts in experienced well-being from Panel B to compute the implied change in life satisfaction ($\beta_+ \times \Delta$Positive + $\beta_- \times \Delta$Negative) and reports the corresponding observed change for comparison. Standard errors are in parentheses.

Table S12.
Sensitivity of Primary Parenthood Associations to Unobserved Confounding

	Women			Men	
	Mood	High Mood	Arousal	Neg. Emotion	Mood Volatility
Coefficients					
Uncontrolled	0.118	0.038	0.071	0.014	0.028
Controlled	0.086	0.041	0.119	0.036	0.042
R-squared					
Uncontrolled	0.058	0.050	0.039	0.107	0.086
Controlled	0.089	0.078	0.061	0.141	0.107
Max	0.115	0.101	0.079	0.183	0.139
δ*	1.638	7.378	-3.535	-1.153	-5.029
Bias Adjusted Coefficient					
δ = 1	0.044	0.046	0.176	0.077	0.088
δ = 2	-0.049	0.053	0.309	-0.100	-0.100

Notes: This table reports the sensitivity of estimated parenthood associations to potential bias from unobserved confounding, following the method of Oster (2019). Each column corresponds to a different experienced well-being outcome. For women, outcomes are average mood, probability of highest mood (High Mood), and arousal. For men, outcomes are aggregate negative emotion and within-person mood volatility. The top panel reports uncontrolled coefficients (no covariates) and controlled coefficients (full covariate set described in the text), with corresponding R^2 values. The assumed maximum R^2 is set to 1.3 × R^2_controlled. The coefficient δ* reports the proportional degree of selection on unobservables, relative to selection on observables, that would be required to drive the parenthood coefficient to zero. Values of |δ*| greater than one indicate that unobserved confounding would need to exceed the explanatory contribution of the full covariate set to eliminate the estimated association. Positive δ* values indicate that observed controls attenuate the raw coefficient; negative δ* values indicate that observed controls increase the coefficient's magnitude, implying that unobservables would need to operate in the opposite direction to eliminate the result. Bias-adjusted coefficients report the implied parenthood association under assumed ratios of unobserved-to-observed selection of δ = 1 and δ = 2.

Exhibit S1. Mobile App Screenshots from Half-Hour Diary (initial screens)

SI References

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