

Meta-Analysis of Prospect Theory Parameters*

Taisuke Imai¹
Jilong Wu³

Salvatore Nunnari²
Ferdinand M. Vieider^{3,†}

¹Institute of Social and Economic Research, The University of Osaka

²Department of Economics, Bocconi University

³RISL $\alpha\beta$, Department of Economics, Ghent University

[†]Corresponding author: ferdinand.vieider@ugent.be

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Abstract

We conduct a meta-analysis of the parameters of prospect theory (PT) using 812 estimates from 166 papers. Since these parameters capture risk-taking propensities, their relevance extends beyond PT itself. To synthesize the evidence, we develop an inverse-variance weighted approach that accounts for correlations among PT parameters and imputes missing standard errors. The average estimates align with the stylized facts of PT, including diminishing sensitivity to outcomes and probabilities. However, three additional findings qualify this baseline. First, parameter estimates exhibit substantial between-study variation. Second, this heterogeneity is poorly explained by observable study characteristics. Third, the strongest predictors of variation are experimental and measurement features, indicating systematic violations of procedure invariance. These results suggest that unexplained differences in risk-taking are not random but systematically linked to how preferences are elicited. This highlights the potential of cognitive accounts of behavior to organize and explain such variation.

JEL Classification Codes: C11, D81, D91

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1 Introduction

Economists have traditionally explained risk-taking behavior through the lens of diminishing marginal utility of money, an idea that dates back to the resolution of the St. Petersburg paradox by Daniel Bernoulli (1738/1954). A substantial body of literature, beginning with Preston and Baratta (1948), has further demonstrated systematic likelihood-dependence in risk-taking. Both outcome-dependence and likelihood-dependence in risk-taking have been formally integrated into prospect theory (PT; Kahneman and Tversky, 1979; Tversky and Kahneman, 1992; Wakker, 2010). A large number of studies have quantified the parameters governing the PT functionals. Here, we systematically investigate the consensus emerging from this large set of measurements. Prospect theory parameters summarize systematic tendencies in risk-taking that extend well beyond the specific model, capturing how behavior changes with the stakes and probabilities of a decision. The results we present here thus hold interest beyond prospect theory itself.

Here, we conduct a systematic review of estimates of PT functionals by quantitatively meta-analyzing all existing studies that estimate PT parameters. This allows us not only to summarize aggregate parameters derived from the underlying studies but also to describe heterogeneity across studies and determine whether it can be explained by observable study characteristics. Our meta-analysis includes 812 parameter estimates from 166 papers, summarizing the decisions taken by 52,000 subjects across 69 countries. While we analyze all PT parameters jointly, we focus on utility curvature and probability weighting, given that loss aversion has been examined elsewhere (Brown et al., 2024).

Why meta-analysis. Meta-analysis constitutes a powerful tool to aggregate estimates from an entire literature, and has been gaining traction in economics in recent years (Havránek, 2015; Imai, Rutter and Camerer, 2021; Brown et al., 2024; Nunnari and Pozzi, forthcoming; Cala et al., forthcoming; Cheung, Tymula and Wang, forthcoming). Meta-analysis can thereby quantify the accumulative wisdom of an entire field, providing aggregate level parameter estimates that can serve as priors for future empirical studies, or that can be used to calibrate parameter values in applications. Beyond aggregate summary estimates, meta-analysis can also reveal the consistency of a given field and help uncover inconsistencies across studies that are often not apparent when examining them one at a time. It is thus highly complementary to narrative reviews, which tend to focus on conceptual evolutions within a field, but risk missing the big quantitative picture.

Measurement errors in correlated parameter estimates. Meta-analysis treats estimates reported in the literature as the object of study. The core model takes the

form of a measurement-error model: reported parameter means are treated as noisy estimates of the true, but latent, parameters, with the estimation noise assumed to be proportional to the parameter’s standard errors. The latent true parameters are then modeled as being drawn from a common mean, resulting in adjustments of the true study-level parameters towards the aggregate mean in proportion to 1) the standard error surrounding the estimate; and 2) the distance from the aggregate mean. This allows inferences about the underlying true effect size in the absence of sampling variation, as well as providing an assessment of true heterogeneity between studies once sampling variation is taken into account. Further, allowing estimates to systematically differ by study characteristics through meta-regression enables assessment of systematic differences between studies.

Applying meta-analysis to multiple parameters estimated to describe risk-taking data presents several challenges. One such challenge arises from the lack of standardization in statistical reporting: many papers report parameter estimates but fail to provide sufficient statistical information to infer their standard errors. Another challenge arises because PT parameters tend to be correlated both for mechanical (cf. [Zeisberger, Vrecko and Langer, 2012](#)) and substantive reasons (cf. [Vieider, 2024b](#)). Finally, parameterizations tend to differ across PT estimations, raising non-trivial questions about how to jointly analyze the parameters. We tackle these issues by developing a novel Bayesian meta-analysis approach that 1) imputes missing standard errors organically within the model structure; 2) analyzes all PT parameters estimated in a given study *jointly*; and 3) leverages the correlation structures both in the estimation of latent parameters and in the imputation of standard errors. This allows us to explicitly model the data-generating process underlying the estimates we analyze—a crucial element in hierarchical modeling (see [Gelman et al., 2014](#), ch. 8).

Meta-analytic parameter averages support stylized PT patterns. The meta-analytic average parameters across all 812 estimates we obtain—a “collective best guess” of what the PT parameters may be—provide strong support for the stylized PT patterns of decreasing sensitivity towards changes in wealth and probabilities. The average utility curvature (constant relative risk aversion, CRRA) coefficient for gains is 0.31 (95% credible interval: 0.28–0.33), indicating decreasing sensitivity towards increases in wealth. We also document decreasing sensitivity towards decreases in wealth (convex utility over losses). Sensitivity towards losses, however, decreases significantly more slowly than towards gains (with a mean of 0.27 and a credible interval of 0.23 to 0.30). The average elevation parameter of the probability weighting function is 0.99 (95% credible interval: 0.96–1.03), and it does not differ by domain. Overall, we do not find much support for either probabilistic optimism or pessimism.

We also find clear evidence for likelihood-insensitivity, the observation that relative risk aversion systematically increases in the probability of winning (decreases in the probability of losing). The mean likelihood-sensitivity parameter across outcome domains is 0.68 (95% credible interval: 0.66–0.70). Some individual studies in the literature have reported *over*-sensitivity to probability, leading to an S-shaped probability weighting function. Surprisingly, however, such patterns receive no support in our latent-parameter estimates: the relatively large standard errors of the raw estimates indicating S-shapes combined with their outlying nature relative to the bulk of the evidence results in their meta-analytic “true effects” being invariably corrected to fall below the perfect sensitivity cutoff.¹

We furthermore document significant correlations between PT parameters. The first set of correlations concerns the same parameters for gains and losses, showing reflection of risk attitudes between gains and losses (e.g., [Schoemaker, 1990](#)). Utility curvature parameters show a positive correlation, whereas the parameters governing the elevation of the probability weighting function show no correlation. Taken together, these correlations suggest that, at least at the level of aggregate estimates, risk attitudes are indeed reflected: subject populations that are more risk averse for gains tend to be more risk seeking for losses. Furthermore, populations that exhibit greater sensitivity toward probabilities in the gain domain also show higher likelihood-sensitivity in the loss domain. Beyond these gain-loss relationships, we also observe correlations among different PT parameters: for instance, utility curvature and likelihood-sensitivity are significantly correlated.

Meta-regression reveals violations of procedural invariance. The single most important predictor of the estimated PT parameters in our meta-regressions is the experimental method employed to elicit or measure risk attitudes.² Utility curvature is significantly less pronounced when a choice list design is used instead of a binary choice design. It is also attenuated in studies that employ bisection procedures to pinpoint indifference points. In contrast, lottery menu methods tend to yield greater utility curvature relative to binary choice. Choice list formats also influence probability weighting: they are associated with more elevated probability weighting functions and reduced likelihood sensitivity. Not all choice lists are, however, created equal. Notably, [Holt and Laury \(2002\)](#) type choice lists, where probabilities vary within the list, produce higher likelihood sensitivity than certainty equivalent lists, with a fair share even indicating

¹Note that we purely focus on given probabilities or *risk* in our meta-analysis. S-shaped probability weighting has, for instance, been documented by many studies when PT parameters are obtained from experience-based choice, in which probabilities must be discovered by sampling. Such studies are not included in our meta-analysis, given that they have been meta-analyzed recently by [Wulff, Mergenthaler-Canseco and Hertwig \(2018\)](#).

²Dummies for the elicitation method employed in the experiment account for 39.7%, 35.3%, and 20.8% of the overall heterogeneity we can explain for utility curvature, likelihood sensitivity, and elevation, respectively.

likelihood *over*-sensitivity (similar effects have been documented for probability equivalents; see [Feldman and Ferraro, 2023](#)). Direct matching methods, typically implemented via willingness-to-pay or willingness-to-accept tasks, yield more depressed probability weighting functions (indicating pessimism for gains, optimism for losses).

These findings point to violations of *procedure invariance*—the principle that measured preference parameters should not depend on the elicitation method, a tenet implicitly assumed by PT. Such results raise concerns for PT, which makes no predictions about how measurement methods might impact behavior. This pinpoints another strength of meta-analysis: even while analyzing parameters estimated within a given theoretical framework, the methods allow us to pinpoint potential weaknesses in said framework.

Large unexplained variation in parameters. The PT literature has largely emphasized agreement in parameter estimates across studies, with only a few exceptions, most notably concerning the S-shaped probability weighting function discussed above. A surprising insight from our meta-analysis is the sheer extent of heterogeneity in these estimates. This is particularly surprising, given that our meta-analytic measurement-error model explicitly accounts for heterogeneity in estimates arising from sampling error. If differences across experiments reflected only random draws from a common population, posterior estimates of the true parameter values ought to converge toward the population mean. This is not the case in our data, which instead indicates substantial true heterogeneity across studies.

One reason for such true heterogeneity may be that studies differ in subject pools, measurement methods, or other design elements. To account for this, we control for a large set of observable characteristics of the subject pool, the experimental study, and the estimation techniques using meta-regression. Even after controlling for such characteristics, however, unexplained heterogeneity across estimates remains very large. Figure 1 shows our best possible predictions about the true utility function (Panel A) and the true probability weighting function (Panel B). Even though we control for a wide set of characteristics—including functional utility specifications, probability weighting specifications, elicitation method, reward types, data types, and subject pool—the heterogeneity across studies in this extended-specification estimate is not much reduced relative to that under the baseline specification, which controls only for utility specification.

The finding of considerable heterogeneity, even after controlling for observable characteristics of the subject pool and the experimental implementation, suggests that more subtle contextual factors likely exert first-order influence on risk-taking. This implies that the combined force from all existing PT estimates does still not allow us to make precise quantitative predictions about future estimates, suggesting that important aspects of what determines risk-taking in experiments remain poorly understood.

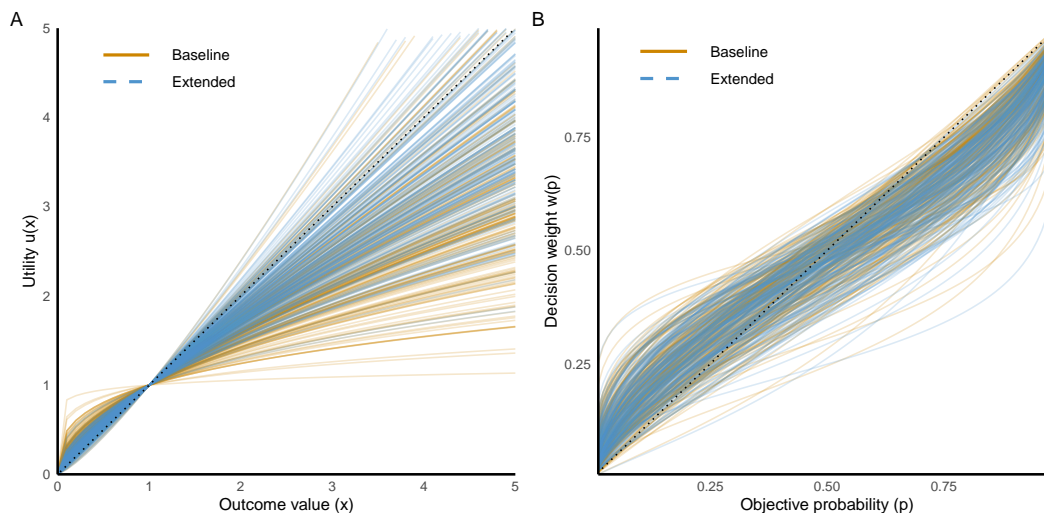


FIGURE 1: Predictive function shapes implied by posterior predictive coefficients: (A) utility functions under the power specification; (B) probability-weighting functions under the linear-in-log-odds specification. *Notes:* Solid orange lines depict predictions from the baseline specification, which controls only for the utility specification. Dashed blue lines depict predictions from the extended specification, which additionally accounts for a broad set of observable study characteristics. These extended-specification predictions correspond to a study with linear-in-log-odds probability weighting over gains, CRRA utility, binary-choice elicitation with real incentives, aggregate data, and a sample of university students in a laboratory setting in Europe.

Wider implications for accounts of risk-taking. Both the systematic violation of procedure invariance and the substantial unexplained heterogeneity pose problems for a model like PT, which explains risk-taking as a preference-based response to objectively perceived choice primitives. Our findings, in particular, imply that risk-taking behavior may be influenced by more subtle experimental cues, or details of the estimation procedure or data collection that ought to be irrelevant according to the model used to produce the estimates. One possibility is that reactions to these subtle cues reveal underlying cognitive mechanisms that generate PT-like behavior. For instance, [Oprea and Vieider \(2025\)](#) show that even subtle variations in visual displays or the numerical magnitudes in which identical stakes are denominated can generate first-order changes in likelihood-dependence and average risk-taking levels.

Several recent studies propose that risk-taking patterns as described in PT arise from cognitive frictions in the processing of choice attributes rather than from stable preferences ([Khaw, Li and Woodford, 2021; 2023; Vieider, 2024b; Frydman and Jin, 2023](#)). [Enke and Graeber \(2023\)](#) show how likelihood-sensitivity in choice lists can be predicted by survey measures of cognitive uncertainty. [Oprea \(2024\)](#) shows that the type of likelihood-insensitivity obtained under risk can be reproduced when removing the risk entirely, while maintaining the complexity of the choice situation. [Oprea and Vieider \(2024\)](#) show that providing redundant information in the form of samples from fully described choice op-

tions makes likelihood-insensitivity disappear, resulting in broadly neoclassical behavior.

Accounts of risk-taking as arising from cognitive frictions in the representation of choice stimuli, and optimal ways of dealing with such frictions by leveraging information about the distribution of choice stimuli in the environment, hold the promise to provide a generative account of the origin of the PT violations we document in this paper (Vieider, 2025). They are also consistent with accounts that have shown behavior such as sensitivity to probabilities and to losses to systematically depend on cognitive ability (Choi et al., 2022; Chapman et al., 2025; Bouchouicha, Li and Vieider, 2025). We thus see them as a promising way of explaining some of the puzzles we document in our meta-analysis—a point to which we will return in the discussion.

2 Prospect Theory

We start by providing a succinct overview of PT and by outlining the functional forms of utility and probability weighting used in the literature. For the sake of clarity, we will focus on binary lotteries $(x, p; y)$, where outcome x occurs with probability p and outcome y with a complementary probability $1 - p$. Such simple binary lotteries make up the lion’s share of tasks used in the studies we meta-analyze (over 80%). For simple binary lotteries like this, original prospect theory, rank-dependent expected utility theory, dual-expected utility theory, and disappointment aversion are all special cases of PT (see Wakker, 2010, Section 7.11). This allows us to cast a wide net and include studies that estimate the functionals of these models in our analysis.

A lottery is deemed non-mixed if both outcomes are either positive or negative ($x > y \geq 0$ or $x < y \leq 0$), and mixed when it includes both a positive and a negative outcome, so that $x > 0 > y$. In PT, the value of a non-mixed lottery $(x, p; y)$ is given by:

$$w^s(p) \cdot v(x) + (1 - w^s(p)) \cdot v(y), \tag{1}$$

where v is the utility function (typically assumed to have a fixed point at 0, $v(0) = 0$), and w^s represents the probability weighting function, which satisfies $w^s(0) = 0$ and $w^s(1) = 1$; $s = + (-)$ denotes the gain (loss) domain.³ For mixed lotteries $(x, p; y), x > 0 > y$, the PT value is represented as

$$w^+(p) \cdot v(x) + w^-(1 - p) \cdot v(y). \tag{2}$$

³Some papers have estimated separable versions of probability weighting (e.g., Camerer and Ho, 1994), i.e., $w^s(p) \cdot v(x) + w^s(1 - p) \cdot v(y)$. Our dataset excludes these articles since they cannot be easily mapped to separable estimates.

Mixed lotteries are used to estimate loss aversion—the kink of the utility at the origin. We exclude loss aversion from our meta-analysis since it has recently been analyzed elsewhere (Brown et al., 2024).

Empirically studying PT typically requires assuming a specific functional form (see Abdellaoui, 2000 and Bleichrodt and Pinto, 2000 for nonparametric estimates of PT; see Gonzalez and Wu, 1999 for a semi-parametric approach). These assumptions enable researchers to summarize patterns of choice across diverse tasks—differing in stakes and probabilities—using relatively few parameters. In our meta-analysis, we code up to three parameters per *estimate*: two for probability weighting and one for utility. A *study* may report separate parameters for gains and losses, which we treat as distinct estimates. Since few studies report a complete set of parameters for both domains, we code gains and losses separately and control for the outcome domain in all meta-regressions. While crucial for estimating PT parameters, the stochastic choice model has typically received little attention in the PT literature, and reports of the error terms are too sparse to allow us to include an additional noise term in our analysis.

Utility functions. Panel A of Table 1 presents the functional forms commonly used in the literature, along with their frequencies in our dataset of 812 collected estimates. For further details on the data, see Section 3. We use ρ^s throughout to represent the utility curvature parameter, capturing attitudes towards stakes, where $s \in \{+, -\}$ indicates the corresponding payoff sign. Functions from the power utility family constitute the large majority of utility functions, accounting for 81.7% of the total (e.g., Tversky and Kahneman, 1992). For power utility functions, the parameter ρ^s is also known as the relative risk aversion coefficient.⁴ The linear function represents 6.2% of all estimates, amounting to not estimating utility curvature at all (which is a special case of power utility with $\rho^s = 0$). The remaining category, with 12.2%, consists of exponential functions (e.g., Köbberling and Wakker, 2005).⁵

Utility functions from the power and exponential families are challenging to compare directly. Mappings from one into the other—while possible in principle—are only valid locally, and are very sensitive to the stake levels assumed for the approximation. We will thus consistently control for functional form in parametric analysis, and describe results for the two functional families separately. Panel A of Figure 2 illustrates three power

⁴Note that simple power functions (where $u(x) = x^r$) coexist in the literature with CRRA functions. To make the parameters comparable, we consistently code ρ^s as the CRRA parameter for utility functions from the power family, i.e., $\rho^s = 1 - r^s$.

⁵Some PT estimations we encountered did not report which functional forms were used, and we thus felt compelled to exclude these estimates ($N = 46$). A few estimates also used less commonly used utility function specifications, such as the logarithmic function ($N = 5$) and the expo-power function ($N = 2$). Once again, we excluded these estimates from our analysis. Note, however, that papers reporting these estimates typically also reported alternative specifications, which we included, so we did not exclude the studies as a whole but only the specific estimates using these non-standard functional forms.

TABLE 1: Utility and probability weighting function specifications.

Function	Functional form	Freq.	%
A: Utility			
Power (TK)	$v(x) = \begin{cases} \frac{1}{1-\rho^s} x^{1-\rho^s} & \text{if } x \geq 0 \\ \frac{-\lambda}{1-\rho^s} (-x)^{1-\rho^s} & \text{if } x < 0 \end{cases}$	663	81.7
Exponential	$v(x) = \begin{cases} \frac{1}{\alpha^s} (1 - \exp(-\alpha^s x)) & \text{if } x \geq 0 \\ \frac{-\lambda}{\alpha^s} (1 - \exp(-\alpha^s (-x))) & \text{if } x < 0 \end{cases}$	99	12.2
Linear	$v(x) = \begin{cases} x & \text{if } x \geq 0 \\ -\lambda x & \text{if } x < 0 \end{cases}$	50	6.2
Total number of estimates		812	100
B: Probability weighting			
Tversky-Kahneman (TK)	$w(p) = \frac{p^\gamma}{(p^\gamma + (1-p)^\gamma)^{1/\gamma}}$	281	34.6
Prelec II	$w(p) = \exp\left(-\frac{1}{\delta}(-\ln p)^\gamma\right)$	219	27.0
LLO	$w(p) = \frac{\delta p^\gamma}{\delta p^\gamma + (1-p)^\gamma}$	199	24.5
Prelec I	$w(p) = \exp\left(-(-\ln p)^\gamma\right)$	102	12.6
Power	$w(p) = p^{\frac{1}{\delta}}$	10	1.2
Gul	$w(p) = \frac{p}{1+(1-p)^{\frac{1}{\delta}}}$	1	0.1
Total number of estimates		812	100

utility functions obtained from different coefficient values. Specifically, the function is convex for $\rho^+ < 0$, linear for $\rho^s = 0$, and concave for $\rho^+ > 0$ (while not shown, it is concave for $\rho^- < 0$, and convex for $\rho^- > 0$). Assuming a linear probability weighting function, a convex utility function implies risk-seeking, while a concave utility function reflects risk aversion. Under more general, nonlinear probability weighting functions, risk attitudes are jointly captured by probability weighting and utility curvature—one of the reasons we will analyze the different model parameters *jointly*.

Probability weighting functions. We use γ^s to denote *likelihood-sensitivity*, and δ^s to designate the elevation of the function. The elevation has different meanings in the gain and loss domains: δ^+ captures *optimism*, while δ^- captures *pessimism*. This derives from the convention of attaching the decision weight $w(p)$ to the best outcome for gains, but to the worst outcome for losses (i.e., to the largest loss). There are six probability weighting functions underlying the estimates in our dataset. Unlike utility functions,

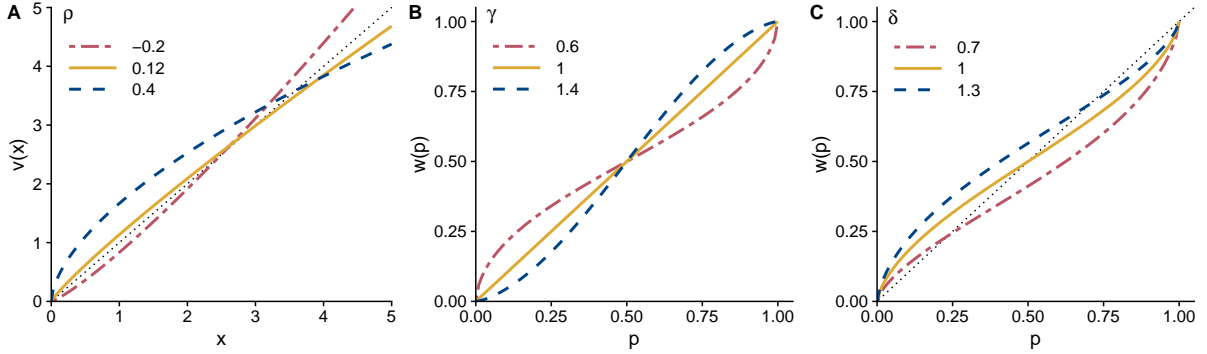


FIGURE 2: Examples of prospect theory functions. (A) The CRRA utility function with three levels of curvature in the gain domain. This is the specification used by [Tversky and Kahneman \(1992\)](#), $v(x) = \frac{x^{1-\rho}}{1-\rho}$ for $x \geq 0$. (B) The LLO function proposed by [Goldstein and Einhorn \(1987\)](#), $w(p) = \frac{\delta p^\gamma}{\delta p^\gamma + (1-p)^\gamma}$, with three levels of likelihood insensitivity (γ) and a neutral elevation coefficient ($\delta = 1$). (C) The LLO function with three levels of elevation (δ) and a neutral likelihood insensitivity level ($\gamma = 1$).

however, it is relatively straightforward to map parameters across several functional forms to one another. While some mappings rely on approximations, this approach enables us to analyze the parameters collectively. We will also control for functional form in meta-regression to determine the extent to which the functional form assumptions affect parameter estimates.

We take the linear in log-odds (LLO; [Gonzalez and Wu, 1999](#)) specification as our baseline. In probability space, the function takes the form $w(p) = \frac{\delta p^\gamma}{\delta p^\gamma + (1-p)^\gamma}$. This specification accounts for 24.5% of all estimates. The parameter $\gamma > 0$ captures likelihood-sensitivity—the phenomenon whereby risk attitudes systematically vary across probabilities. A value of $\gamma = 1$ captures perfect sensitivity (the EUT case), $\gamma < 1$ likelihood-insensitivity, and $\gamma > 1$ likelihood-oversensitivity. The parameter $\delta > 0$ captures the elevation of the function, with $\delta > 1$ capturing optimism for gains (pessimism for losses), and $\delta < 1$ capturing pessimism for gains (optimism for losses).⁶ Panel B and C of Figure 2 respectively illustrate how likelihood-sensitivity and elevation affect the shape of a weighting function. Gul’s disappointment aversion function ([Gul, 1991](#)) is a special case of the LLO function with the sensitivity parameter fixed to 1, and its pessimism parameter derived as the inverse of the LLO optimism parameter, δ^{-1} .

The [Prelec \(1998\)](#) 2-parameter function (Prelec II) does not have a direct mapping into

⁶This interpretation of the parameters is particularly intuitive in the log-odds version of this weighting function, which takes the form $\log\left(\frac{w(p)}{1-w(p)}\right) = \gamma \log\left(\frac{p}{1-p}\right) + \log(\delta)$. Values of $\gamma < 1$ —which takes the form of a power to the odds—will “compress” the odds towards 1, since odds smaller than 1 (i.e., a probability smaller than 0.5) will be uplifted towards 1, and odds larger than 1 will be reduced (with the opposite effect for $\gamma > 1$). Values $\gamma < 1$ thus produce a sort of regression to the mean of 1. This mean, however, is further affected by the value of δ , which acts as the intercept of the function in log-odds space.

the LLO function. To obtain a mapping, we approximate its anti-elevation parameter by the inverse of the LLO elevation parameter to have a common interpretation. We also equate the sensitivity parameters across the two functional forms. While this constitutes an approximation, the two functions are only distinguishable for extreme probabilities, which are rarely included in experiments for practical reasons. The Prelec I function is a special case of the Prelec II with $\delta = 1$. Another special case of the Prelec II function is the power function, with $w(p) = p^{\frac{1}{\delta}}$, where $\gamma = 1$. This is the case because $\exp(-\frac{1}{\delta}(-\ln(p))) = p^{\frac{1}{\delta}}$. This functional family makes up 40.6% of all estimates in our data. The most important function still missing is the [Tversky and Kahneman \(1992\)](#) (TK) function, used in 34.6% of all estimates, which cannot be reduced to any of the others. Nevertheless, the sensitivity parameter will typically be similar to those of other functions, and we will thus analyze them jointly.⁷

3 Data

3.1 Identification and Selection of Relevant Studies

We identified and selected papers that estimated PT parameters using clearly specified inclusion criteria. The primary criterion was to include “all empirical papers that estimate PT parameters.” Under this criterion, we included papers that used choice data from both laboratory and field/online experiments, as well as from surveys conducted by mail or phone. Papers only estimating utility but not probability weighting were not included (see [Online Appendix A.1](#) for precise search terms). Our search for relevant papers was conducted in the Web of Science database.

The initial search, conducted in the summer of 2022, yielded 2,034 papers. In the initial phase of paper identification, we scrutinized titles and abstracts, setting aside 1,453 papers that were evidently irrelevant to our study. Subsequently, we thoroughly examined the remaining papers, applying our inclusion criteria based on content, and proceeded to code the relevant information. Additionally, we employed IDEAS/RePEc and Google Scholar to explore unpublished working papers. The comprehensive search and selection procedure is outlined in [Figure A.2](#) in the [Online Appendix](#). By the end of this process, we identified 166 papers, 12 of which remained unpublished as of the initial data compilation in winter 2023.⁸

⁷Apart from the estimates reported in [Table 1](#), three estimates adopted the Karmarkar specification ([Karmarkar, 1978](#)), $w(p) = \frac{p^\gamma}{(p^\gamma + (1-p)^\gamma)^\delta}$, which is a generalization of the TK. Given the few occurrences ($N = 3$) and the difficulty of linking them to other forms, we exclude them from our analysis.

⁸Once a first version of this paper is completed, we will share it with the community to determine whether we missed any papers. We are also committed to updating our database and results close to the publication date to make sure the results are as up-to-date as possible.

3.2 Data Coding

Our meta-analysis dataset is assembled by encoding estimates of PT parameters and their associated standard errors (SEs), along with details about the experiment, subject pool, and the estimation procedures underlying the parameters. For meta-analyses, SEs play a critical role in computing weighted averages. In situations where SEs are not explicitly reported, we reconstructed them using available information, such as standard deviations (SD) and p -values, or approximated them using the interquartile range (IQR). A detailed overview of our calculation methodology can be found in the Online Appendix [A.2](#). We also coded variables detailing the location of the experiment (e.g., lab, field, classroom, online), types of rewards (e.g., real money or hypothetical money, and other consequences), subject population (e.g., university population, general population, and other population like farmers and athletes), functional forms of probability weighting (e.g., Prelec I, Prelec II, TK, and LLO) and utility (Power, Exponential, and Linear) which will be described in Section 2, among other characteristics. Online Appendix [A.3](#) contains a comprehensive list of all the variables coded in the study. A random subset of 10% of the data was coded independently by at least two co-authors to ensure data quality and coding consistency.

3.3 Descriptive Statistics

We identified a total of 166 papers for inclusion in our meta-analysis (see the Online Appendix [A.4](#) for a list). Out of these, 154 articles were published across 61 journal outlets, and the rest remained unpublished. The dataset encompasses papers from diverse disciplines, including economics, management, psychology, neuroscience, medicine, psychiatry, agriculture, environment, transportation, and operations research. Moving forward, we shift our focus to the primary variable of interest—the estimated parameters of the utility and probability weighting functions. The dataset comprises a total of 812 estimates (see Table 2). Of these estimates, 54.7% pertain to aggregate or “representative agent” estimates, which pool data from all subjects. Some 24.4% report means of individual-level estimates, and 20.9% report medians. There are 61 cases in which both the mean and median of the distribution of PWF parameters estimated at the individual level are available. 113 of the 812 estimations do not report SEs or any statistical information that can be used to calculate SEs, including the seminal estimation of [Tversky and Kahneman \(1992\)](#).

We also determined the country where the data were collected. The majority of papers report estimates from data collected in a single country. Four collected data across multiple countries/regions, and two of these four ([Rieger, Wang and Hens, 2017](#); [L’Haridon and Vieider, 2019](#)) conducted large-scale cross-country studies, gathering data from 30 or

TABLE 2: Types of PT estimates.

	All estimates		With SE	
	Freq.	%	Freq.	%
Aggregate-level	444	54.7	390	55.8
Individual-level mean	198	24.4	180	25.8
Individual-level median	170	20.9	129	18.5
Total	812	100.0	699	100.0

TABLE 3: Characteristics of studies estimating PT parameters.

	Freq.	%		Freq.	%
Total number of studies	812	100			
<i>Continent type</i>			<i>Data type</i>		
Europe	484	59.6	Lab	586	72.2
North America	133	16.4	Class	120	14.8
Asia	104	12.8	Online	58	7.1
Central/South-America	44	5.4	Field	48	5.9
Africa	26	3.2	<i>Reward type</i>		
Oceania	21	2.6	Real money	464	57.1
<i>Elicitation type</i>			Hypo money	307	37.8
Binary	256	31.5	Other	41	5.0
List	254	31.3	<i>Subject type</i>		
Matching	170	20.9	University	652	80.3
Bisection	110	13.5	Other	85	10.5
Lottery Menu	22	2.7	General	75	9.2

Notes: Regarding *Data type*, the category of “online” includes experiments conducted online as well as other special cases, such as survey data collected via phone calls or mail. For *Subject type*, the “other” category mainly includes farmers ($N = 20$), athletics ($N = 20$), health professionals ($N = 10$), businessman ($N = 11$). Last, the other reward types contain health and time.

more countries. In total, the estimates included in our dataset originate from 69 countries (see Figure A.3 for a global map). Some of the main characteristics of interest to our analysis include the type of reward, the type of subject pool, the data type (e.g., lab vs. field), the elicitation method, and the domain of stimuli. Table 3 reports the descriptive statistics of these characteristics.

Our dataset primarily comprises studies conducted in Europe, which represent 59.6% of the total. The second largest proportion, accounting for 16.4%, originates from North America. Additionally, 30 studies (12.8%) were carried out in Asia, while a few others were conducted across various other continents. Binary choices emerged as the most popular data collection method, utilized in 256 estimates (31.5%). The choice list format was the preferred method for 31.3% of the estimates. Among these choice lists, certainty

equivalents were the most widely used tool, accounting for 217 out of 254 estimates. The third most common approach was the matching method (20.9%), which involves directly asking participants to provide a matching value, usually in the form of a willingness-to-pay or willingness-to-accept measure. The bisection method accounts for a smaller share, while the lottery menu category accounts for 2.7% of estimates.

The majority of our data comes from laboratory experiments (72.2%), augmented by studies conducted in field settings or online.⁹ Monetary rewards, whether real or hypothetical, were the most common form of incentive, accounting for over 90% of all rewards offered. Non-monetary rewards, such as health and time, were less common. A significant portion of the subjects were university students (80.3%) or the general public (10.5%), and the remaining studies focused on niche demographics, including businessmen, health practitioners, and farmers.

4 Methodological Framework

We begin our analysis within a conventional meta-analytic framework, treating the parameter estimates reported in the literature as noisy measurements of underlying true effect sizes (parameter values). However, applying this standard approach to the parameters of Prospect Theory (PT) presents several challenges. Most important, perhaps, is the interdependence of the parameters: because PT parameters *jointly* characterize risk attitudes and are econometrically identified from choice data, assuming independence among them can lead to biased inferences. To address this issue, we extend standard meta-analytic methods to a joint modeling framework that explicitly accounts for the correlation structure among parameters. Moreover, the presence of missing standard errors and the variability in which subsets of parameters are reported across studies requires further adaptations. These modifications allow us to incorporate as many estimates as possible in the meta-analysis while maintaining methodological rigor.

4.1 Joint Meta-Analysis Model

Our core model closely follows the standard meta-analytic framework but extends it to accommodate multiple, potentially correlated parameters. We begin by outlining this multivariate generalization of conventional meta-analysis methods.

Measurement error model. Let $\boldsymbol{\theta}_i = (\rho_i, \gamma_i, \delta_i, \lambda_i)$ denote the vector of parameters for estimate i (which, for now, we assume to be complete). Let $\mathbf{se}_i = (se(\rho)_i, se(\gamma)_i, se(\delta)_i, se(\lambda)_i)$

⁹Online experiments primarily refer to those conducted on online platforms, such as Prolific and Amazon MTurk. This category also includes a small proportion of studies conducted by mail or phone.

denote the corresponding vector of standard errors. We assume that the encoded parameters $\boldsymbol{\theta}_i$ constitute a noisy measure of the true but unobserved underlying effect sizes, designated by $\widehat{\boldsymbol{\theta}}_i$. The noise may arise from sampling error, weak parameter identification, or limitations in econometric procedures. We capture this relationship using the following measurement error model:

$$\boldsymbol{\theta}_i \sim \mathcal{N}\left(\widehat{\boldsymbol{\theta}}_i, \text{diag}(\mathbf{se}_i^2)\right),$$

where the diag operator transforms the vector of squared standard errors into a variance-covariance matrix, placing the squared standard errors on the main diagonal and zeros on the off-diagonal. Each parameter in $\boldsymbol{\theta}_i$ is thus modeled as providing a measure of the true latent parameter in $\widehat{\boldsymbol{\theta}}_i$, with potential measurement error proportional to \mathbf{se}_i^2 .

Meta-analytic aggregation. Assuming that the parameter estimates are drawn from an underlying distribution of parameters, one source of variation is sampling error. Additional noise may arise from weak parameter identification, stemming from limitations in the data, the econometric methods employed, or both—each contributing to large standard errors around individual estimates. The impact of such errors can be mitigated by aggregating across studies. This approach effectively pools individual parameter estimates toward the meta-analytic mean, with greater weight assigned to estimates with lower uncertainty:

$$\widehat{\boldsymbol{\theta}}_i \sim \mathcal{N}(\boldsymbol{\mu}, \boldsymbol{\Sigma}), \tag{3}$$

where $\boldsymbol{\mu}$ denotes a vector of parameter-specific meta-analytic means, and $\boldsymbol{\Sigma}$ is a covariance matrix containing the variances σ^2 along its main diagonal and the covariances between parameters in the off-diagonal elements. This covariance structure explicitly allows for correlations among the true effect sizes, capturing the extent to which the different parameters covary. By modeling the covariance structure of the true (rather than observed) effect sizes, we can account for substantive reasons underlying potential parameter correlations (see, for example, [Vieider, 2024b](#), for model-based predictions of such dependencies). More importantly, this approach enables us to assess these correlations while controlling for observable study characteristics, as described below.

Meta-regression. Given the wide variation in measurement methods, estimation techniques, and study locations in our dataset, it is highly unlikely that measurement error alone accounts for the observed differences in parameter estimates $\boldsymbol{\theta}_i$. This highlights the importance of controlling for study-level characteristics, both to better explain between-study variance and to obtain more accurate estimates of the aggregate effect sizes. More critically, when study-level differences that may influence effect sizes are known *ex ante*, the assumption of exchangeability—that is, the notion that effects are drawn from a

common distribution—is violated. Exchangeability is a foundational assumption in meta-analysis (Gelman et al., 2014), and it is clearly violated in our data. For example, exponential utility parameters are expressed on a different scale than CRRA parameters, and our dataset includes separate entries for gains and losses. To address these issues, all meta-regressions include controls for the functional form of utility and for the outcome domain.¹⁰

This is achieved by replacing μ with $\mathbf{X}_i\boldsymbol{\beta}$ in (3), where \mathbf{X}_i is a $1 \times K$ vector of study characteristics for estimate i (including a column of ones to capture the intercept), and $\boldsymbol{\beta}$ is a $K \times 4$ matrix of regression coefficients. This specification allows us to examine, inter alia, the extent to which the variance in each parameter can be explained by the observable characteristics captured in \mathbf{X} .¹¹ Let θ_i denote a generic (scalar) parameter, and let τ^2 represent its variance. We then obtain a measure of the explained variance as $R^2 = 1 - \tau_1^2/\tau_0^2$, where τ_0^2 is the variance from a model without covariates (i.e., the baseline model), and τ_1^2 is the variance from the regression model whose explanatory power is being assessed.

4.2 Model Refinements

The model described above follows a conventional approach for scalar parameters, with its primary innovation being the joint analysis structure, which facilitates the assessment of parameter covariances. We now introduce refinements that enable the simultaneous incorporation of all coded parameters into the meta-analysis, ensuring that each parameter both contributes to and is informed by the endogenously estimated meta-analytic averages.

Error imputation. Several studies estimating PT functionals report parameter estimates but lack sufficient statistical information to recover standard errors. This includes the seminal work by Tversky and Kahneman (1992). In some cases, studies provide statistical information for certain parameters but not for others. Rather than discarding such studies and losing valuable data, we treat the missing standard errors as missing variables and impute them within our meta-analytic model. Note that this approach assumes that missing statistical information is “missing at random,” i.e., there are no

¹⁰An alternative modeling approach would be to include parameters for gains and losses within a single parameter vector. While this would allow for the exploitation of additional correlations, it would preclude a straightforward meta-analytic assessment of *differences* between gain and loss parameters. We therefore chose to treat losses as separate estimates, especially given the relatively small number of studies that jointly estimate the full set of gain and loss parameters.

¹¹The benchmark model controls for exponential utility form and gain/loss domain for all four parameters. The meta-regression extends this by incorporating the study characteristics in Section 3. Given limited observations with loss aversion estimates ($N = 375$), we apply the full covariate set only to sensitivity, elevation, and utility, while restricting loss aversion to the benchmark specification to avoid over-parameterization.

systematic differences between parameters from studies that report standard errors and those that do not. Comparisons between parameters with and without reported standard errors support this assumption in our dataset. As a robustness check, we also conduct the meta-analysis using only the subset of data with complete observations. The results are consistent with those presented in the main text (see Online Appendix B).

To impute standard errors, we construct an error-prediction model. Let $\log(\mathbf{se}_i)$ denote the element-wise natural logarithm of the standard errors, which ensures non-negativity. We model this transformed vector as:

$$\log(\mathbf{se}_i) \sim \mathcal{N}(\mathbf{Z}_i \boldsymbol{\xi}, \boldsymbol{\Omega}), \quad (4)$$

where \mathbf{Z}_i is a $1 \times M$ vector of characteristics of study i predictive of its standard errors, $\boldsymbol{\xi}$ is a $M \times 4$ matrix of coefficients, and $\boldsymbol{\Omega}$ is a covariance matrix with variances on the diagonal and covariances on the off-diagonal. The predictors in \mathbf{Z}_i include the characteristics such as elicitation methods and the square root of the number of subjects—which, as expected, is strongly negatively correlated with the standard errors—and a constant term to capture the intercept.

The vector $\log(\mathbf{se}_i)$ has a distinctive structure: it includes the logarithm of observed standard errors for studies where these are available, and the logarithm of *predicted* standard errors for cases where the standard errors are missing and must be inferred. The model thus fulfills two roles: (1) it estimates the regression coefficients $\boldsymbol{\xi}$ from the observed study characteristics in \mathbf{Z} ; and (2) it uses these characteristics, along with the estimated coefficients $\boldsymbol{\xi}$, to *predict* (or impute) standard errors where they are not observed. The correlation structure encoded in $\boldsymbol{\Omega}$ ensures that the dependencies among errors are properly accounted for during the imputation process.

Figure 3 illustrates the impact of error imputation on the utility curvature parameter, ρ . The effects on other parameters are qualitatively similar and are presented in Online Appendix B. Panel A shows that the imputed standard errors, depicted as red diamonds, fall within the range of the observed standard errors. The vertical dashed lines highlight a key feature of imputed errors: as variables imputed within the model, they are treated as uncertain quantities themselves. This modeling approach has the advantage of directly incorporating second-order uncertainty into the estimation process. Panel B demonstrates the correlation between the standard errors of the utility parameter and the likelihood-sensitivity parameter, again highlighting imputed values in red diamonds. The figure shows two things. First, the standard errors are strongly correlated: the correlations between parameters estimated from $\boldsymbol{\Omega}$ are 0.61 between ρ and γ , 0.52 between ρ and δ , and 0.75 between γ and δ .¹² Second, this correlation is clearly mirrored in the imputed

¹²The correlations of the standard errors of loss aversion with those of ρ , γ , and δ are 0.34, 0.34, and

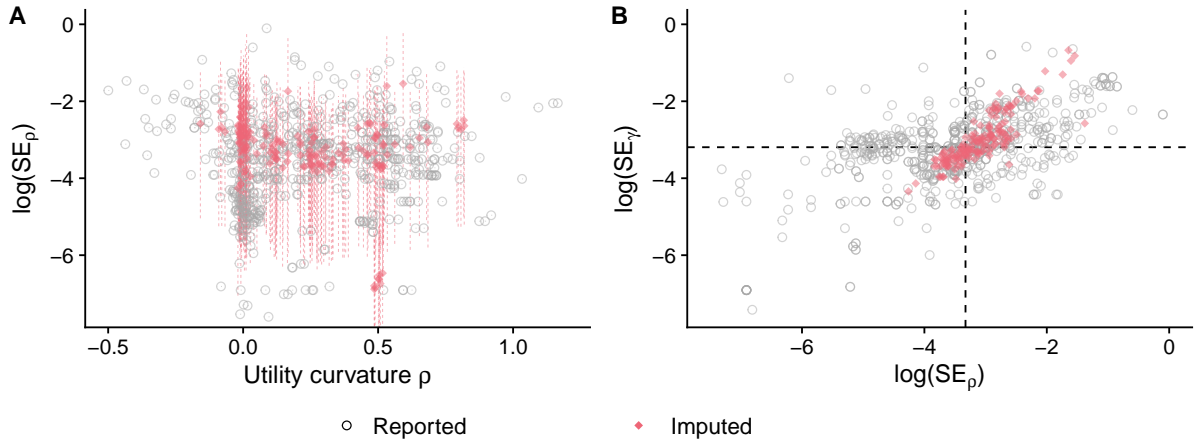


FIGURE 3: Missing standard error imputation and correlation of standard errors for utility curvature and likelihood sensitivity. (A) Scatter plot showing the relationship between power utility curvature (ρ) and the logarithm of its standard errors (SEs). (B) Scatter plot showing the correlation between the logarithms of the SEs for utility curvature (SE_ρ) and likelihood sensitivity (SE_γ). *Notes:* In both panels, observations with reported standard errors (SEs) are displayed as gray circles, while those with imputed SEs are highlighted with red diamonds. The x -axis is truncated for improved visualization. In Panel A, dashed red lines represent posterior uncertainty intervals for the imputed values. In Panel B, dashed gray lines indicate the medians of the respective variables.

standard errors. In fact, the correlation structure has a stronger influence than sample-size effects alone, as evidenced by the clustering of imputed standard errors around the 45-degree line.

Missing parameters. Of the 812 estimates in our dataset, only 418 include a complete set of three parameters of interest, and among these, 150 also include an estimate of the loss aversion coefficient. An additional 383 estimates contain both a utility and a sensitivity parameter (225 of them include an estimate of the loss aversion coefficient), while the remaining 11 estimates include a utility and elevation parameter (those cases assuming a Gul or power probability weighting function). To enable a unified joint estimation across all available data, we construct sub-models tailored to each parameter combination. Importantly, the overall model is structured so that estimates with missing parameters still contribute to the estimation of relevant regression coefficients and their associated variance and covariance components.

Consider a general case involving a two-parameter model, with parameters indexed by j and k . The imputation model is specified as:

$$\log(\mathbf{se}_{\{j,k\}_i}) \sim \mathcal{N}(\mathbf{Z}_i \boldsymbol{\xi}_{\{j,k\}}, \boldsymbol{\Omega}_{\{j,k\}}),$$

where $\log(\mathbf{se}_{\{j,k\}})$ denotes a vector comprising the log standard errors for parameters j and k , respectively.

and k , and the subscript $\{j, k\}$ indicates subsetting of vectors to elements $\{j, k\}$ (and of matrices to the corresponding rows and columns). The modeling of measurement error and the hierarchical aggregation of latent parameters proceed analogously:

$$\begin{aligned}\boldsymbol{\theta}_{\{j, k\}_i} &\sim \mathcal{N}\left(\widehat{\boldsymbol{\theta}}_{\{j, k\}_i}, \text{diag}(\mathbf{se}_{\{j, k\}_i}^2)\right), \\ \widehat{\boldsymbol{\theta}}_{\{j, k\}_i} &\sim \mathcal{N}(\mathbf{X}_i \boldsymbol{\beta}_{\{j, k\}}, \boldsymbol{\Sigma}_{\{j, k\}}),\end{aligned}$$

where $\widehat{\boldsymbol{\theta}}_{\{j, k\}_i}$ represents the latent effects corresponding to the pair of parameters j and k , and the remaining notation follows the same subsetting convention described above.

Performing this iteratively across all parameter combinations ensures that: (1) all parameter estimates can be analyzed jointly, even when the number of parameters differs; and (2) all parameters contribute to the estimation of *the same* regression and covariance parameters, conditional on study characteristics in \mathbf{X} .

4.3 Implementation

We estimate our model using Bayesian hierarchical techniques in Stan (Carpenter et al., 2017), launched from R (R Core Team, 2023) through CmdStanR (Gabry et al., 2023). Meta-analysis is inherently Bayesian, since the endogenously estimated parameters in $\boldsymbol{\mu}$ serve as priors for the latent parameters in $\boldsymbol{\theta}$. The key distinction from frequentist approaches lies in whether explicit priors are specified for the aggregate-level parameters—when they are not, such methods are often labeled “empirical Bayes,” regardless of the estimation technique used. In our model, we specify diffuse hyperpriors to avoid affecting our conclusions in any way, serving solely to make the exploration of the posterior parameter space more efficient.

The coding in the Stan framework provides several advantages. It allows us to hand-code a model precisely tailored to our data and directly quantify the probability mass in favor of a given hypothesis or parameter range, consistent with Bayesian interpretability. Importantly, the Bayesian framework allows us to directly assess the plausibility that a parameter lies within a region near a neutral benchmark (e.g., perfect likelihood-sensitivity) by quantifying the posterior probability mass in that region. This enables us to both support and challenge benchmark-based hypotheses using the posterior distribution. For a practical guide to estimating (hierarchical) decision models in Stan, including sections on measurement error models and meta-analysis, see Vieider (2024a).

5 Results

We structure the results section by presenting findings for each parameter individually, while emphasizing that all estimates are derived from a joint analysis model. We conclude with a discussion of parameter correlations and an examination of potential publication bias. Our meta-analytic estimates control for the use of exponential utility, due to its distinct properties relative to power utility. However, when reporting results for the two weighting function parameters, we do not control for the utility specification, as it showed no significant effect. All reported effects of study characteristics, including outcome domains, are based on meta-regressions that simultaneously include all relevant study-level covariates.

5.1 Utility Curvature ρ

Descriptive statistics. The coded CRRA utility coefficients have a mean of 0.30 across domains, a median of 0.27, and an interquartile range (IQR) of [0.08, 0.51]. This indicates that most studies report a ρ value between 0 and 1, although a minority (16.7%) of estimates are zero or negative—predominantly in the loss domain (see below). For exponential utility, the IQR is [0.00, 0.05], with a median of 0.03 and a mean of 0.08. However, 24.2% of these estimates are negative. These observations align with the stylized patterns discussed in the PT literature, where the utility function is typically concave for gains and convex for losses, exhibiting decreasing sensitivity towards changes in wealth.

Parameter distributions. Panel A of Figure 4 shows the density of the estimated latent parameter values ($\hat{\rho}_i$), represented by the blue dashed line. This density is notably narrower than that of the raw data values, ρ_i , reflecting the effect of meta-analytic pooling: individual estimates that diverge substantially from the overall mean are adjusted toward more credible values. The degree of this adjustment, often referred to as shrinkage, is inversely related to the precision of the estimates; that is, only outliers with large standard errors are substantially pulled toward the mean. We estimate the overall mean of the CRRA functions at 0.29, with a Bayesian 95% credible interval (CrI) of [0.27, 0.31]. In contrast, the mean for the exponential function is much lower at 0.06, with a CrI of [0.01, 0.10], as reflected by the peak around 0 in the density plot. These results confirm the expected pattern: individuals generally exhibit a concave utility function for gains and a convex one for losses.

Gain-loss comparison. An important question is whether PT parameter estimates differ across domains—specifically, whether individual utility curvature differs between gains and losses. Panel B of Figure 4 illustrates the empirical cumulative distribution functions (ECDFs) for encoded ρ^+ (gains) and ρ^- (losses). The plot reveals that, on

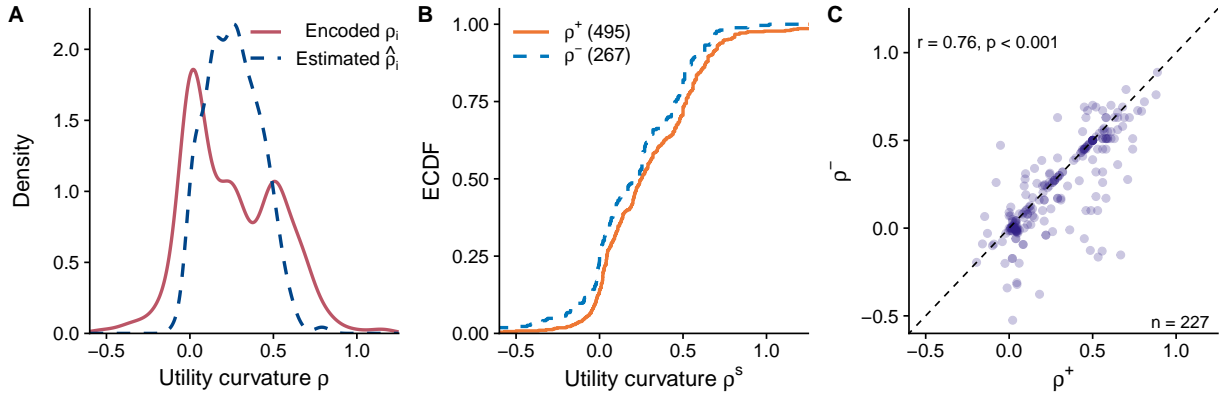


FIGURE 4: (A) Density plot comparing encoded ρ_i values to their corresponding estimates $\hat{\rho}_i$. (B) Empirical CDFs of the power utility curvature, separated by gain and loss domains. (C) Scatter plot of parameter estimates for studies reporting both gain and loss values. *Notes:* The x -axis is truncated for improved visualization. However, density estimations include all observations, including those beyond the displayed range.

average, individuals exhibit a concave utility function for gains (mean, 0.30; median, 0.25) and a convex utility function for losses (mean, 0.21; median, 0.23). Moreover, the ECDF for ρ^+ lies almost entirely below that for ρ^- , suggesting that utility functions tend to be more linear for losses than for gains. Additionally, negative values of ρ are more prevalent for losses (24.3%) than for gains (14.1%), suggesting that convex utility for losses coexists with a relatively larger share of concave cases.

It is important to emphasize that differences across outcome domains should not be interpreted causally, as they may partially reflect variation across studies. To more precisely assess the causal role of outcome sign, we focus on the 227 observations reporting estimates of ρ for both gains and losses. The scatter plot in Panel C shows that observations lie below the diagonal line more frequently than above it. We observe a strong correlation in ρ across domains ($r = 0.76, p < 0.01$), and a statistically significant difference between gains and losses based on a nonparametric test ($p < 0.01$), in line with patterns in the full dataset. This finding is further supported by a meta-regression that includes a loss-domain dummy together with other study characteristics: the estimated coefficient is -0.04 , which is significantly different from zero (see Table 4 for full results).

Procedure invariance. An important question concerns whether elicited choice patterns vary across different measurement methods. In our dataset, four commonly used elicitation methods are binary choice, bisection procedures, choice lists, and direct matching. As shown in Panel A of Figure 5 and reported in Table 4, utility curvature tends to be more pronounced when measured using binary choices than with the bisection method, direct matching, or certainty equivalent (CE) choice lists—the latter showing the strongest difference. However, curvature estimates from binary choices are still smaller than those

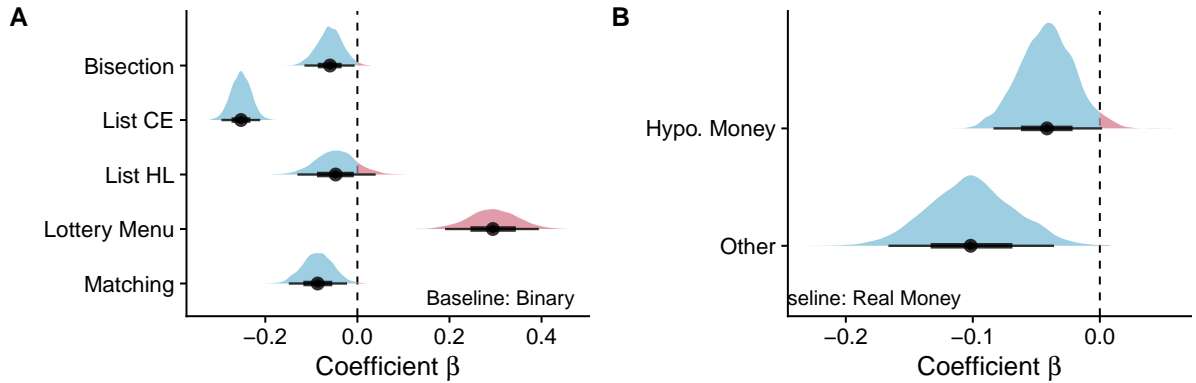


FIGURE 5: Posterior distributions of meta-regression coefficients. (A) Different elicitation methods compared to binary choices. (B) Two alternative reward types compared to real monetary rewards. *Notes:* The posterior distributions of the Bayesian random-effects meta-regression coefficient(s) β , along with the posterior medians (represented by a black dot), 66% credible intervals (indicated by thick solid lines), and 95% credible intervals (shown as thin solid lines), are displayed.

obtained from methods that primarily rely on various types of lottery menus.

The most pronounced effects are associated with CE choice lists, which tend to yield power utility curvature estimates that are 0.25 lower than those obtained using the baseline method of binary choices (see [Bouchouicha et al., 2024](#) for a discussion of implications for risk aversion). Bisection and direct matching methods also yield significantly lower estimates, with reductions of 0.06 and 0.09, respectively. In contrast, lottery menus produce substantially higher estimates—about 0.29 above those from binary choices and notably higher than all other methods. Closer examination shows that this outlier pattern is primarily driven by [Cheung and Johnstone \(2017\)](#), which accounts for 12 of the 22 lottery menu estimates and employs a distinct experimental design and estimation approach.¹³ These findings highlight the considerable impact that methodological choices can have on the estimation of utility curvature.

Explained versus unexplained heterogeneity. Additional observable study characteristics that may help explain between-study variance are summarized in Table 4. Panel B of the figure further illustrates that utility curvature is less pronounced for hypothetical outcomes (marginally significant) and for outcomes other than money. However, we caution against interpreting these associations as causal. These categories often involve substantially higher stakes than incentivized experiments. Meta-analytically examining experiments specifically studying incentive effects in individual decisions, [Li and Vieider \(2025\)](#) conclude that incentives do not materially impact such decisions.

¹³[Cheung and Johnstone \(2017\)](#) used an investment task in which earnings depend on both risk and relative skill placement. The authors estimated PT parameters while incorporating an additional subjective belief parameter, q , that reflects perceived task skill.

TABLE 4: Meta-regression analysis of utility curvature.

Category	Variable	Median	2.5%	97.5%	Category	Variable	Median	2.5%	97.5%
Sign	Loss	-0.04	-0.07	-0.01	Subject	Univ. pop.	Baseline		
U	Expo	-0.17	-0.22	-0.12		General	0.02	-0.05	0.08
PWF	LLO	Baseline				Other	-0.09	-0.15	-0.02
	Gul/Power	0.07	-0.05	0.19	Reward	Real Money	Baseline		
	Prelec I	0.17	0.11	0.22		Hypo. Money	-0.04	-0.08	0.00
	Prelec II	0.10	0.05	0.15		Other	-0.10	-0.17	-0.04
	Elicitation	TK	0.07	0.02	0.11	Data	Lab	Baseline	
Binary choice		Baseline			Class		0.26	0.18	0.34
Bisection		-0.06	-0.12	-0.01	Field		0.12	0.04	0.20
List CE		-0.25	-0.30	-0.21	Online		-0.10	-0.18	-0.02
List HL		-0.05	-0.13	0.04	Continent	Europe	Baseline		
Matching		-0.09	-0.15	-0.02		Africa	0.19	0.12	0.27
Other		0.29	0.19	0.39		Asia	0.00	-0.04	0.05
Estimate	Aggregate	Baseline			C/S-America	0.11	0.05	0.18	
	Ind. Mean	-0.02	-0.06	0.02	North America	0.12	0.07	0.16	
	Ind. Median	-0.03	-0.07	0.01	Oceania	0.14	0.01	0.27	

Even after accounting for these study-level characteristics, more than 30% of the variation in parameter estimates remains unexplained. Notably, a substantial portion of the explained variation is attributable to the elicitation method, highlighting its central role in the analysis. To assess its relative importance, we compare the full meta-regression model with an alternative specification that excludes the elicitation method. This comparison indicates that the elicitation method alone accounts for 39.7% of the explained variation.

5.2 Likelihood Sensitivity γ

For the likelihood-sensitivity parameter, we exclude 11 observations corresponding to the Gul and Power functions, as these specifications set γ to 1. This leaves 801 parameter estimates that inform our analysis of likelihood sensitivity.

Descriptive statistics. We begin by examining the overall distribution of the likelihood sensitivity parameter, γ^s , depicted in Panel A of Figure 6 (red solid curve). The mean and median are 0.72 and 0.68, respectively, with an IQR of [0.55, 0.83]. These values are consistent with an inverse-S shaped probability weighting function, a well-established pattern in the literature. Notably, approximately 7.7% of the raw estimates exceed 1, suggesting S-shaped weighting in a subset of cases.

Parameter distributions. Panel A illustrates how the meta-analysis adjusts raw parameter estimates to infer latent, true effect sizes. The density of the estimated values, $\hat{\gamma}_i$, shows a clear shift relative to the reported values, with increased concentration between 0.5 and 1.0 and diminished density outside this range. As before, this pattern reflects

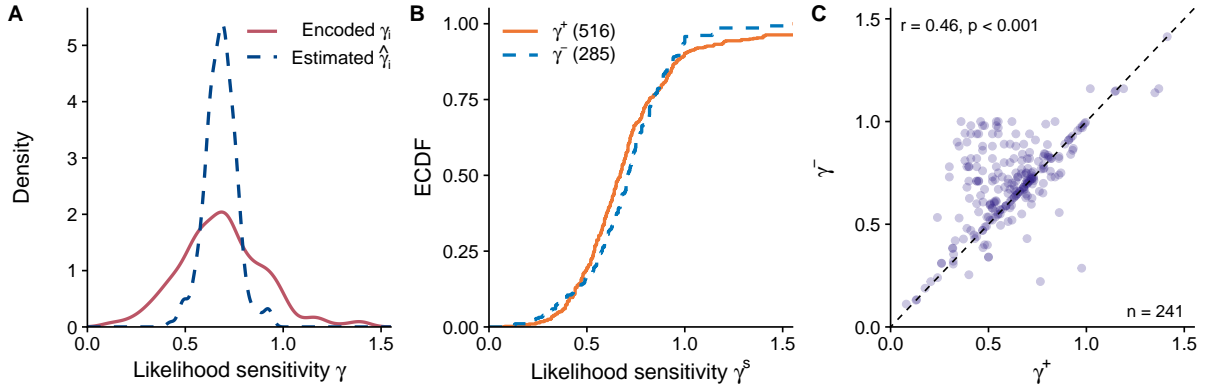


FIGURE 6: (A) Density plot comparing encoded γ_i values to their corresponding estimates $\hat{\gamma}_i$. (B) Empirical CDFs of the likelihood sensitivity parameter, separated by gain and loss domains. (C) Scatter plot of parameter estimates for studies reporting both gain and loss values. *Notes:* The x -axis is truncated for improved visualization. However, density estimations include all observations, including those beyond the displayed range.

the meta-analytic pooling, which systematically down-weights outliers. For likelihood-sensitivity in particular, pooling is notably strong, largely because extreme estimates tend to be measured with lower precision. As a result, the meta-analytic mean is 0.68, with a 95% CrI of [0.66, 0.70], substantially lower than a simple average of the raw estimates would suggest. The narrow credible interval highlights the high precision of our meta-analytic estimate, driven by both the large number of observations for likelihood-sensitivity and the fact that outliers tend to be especially noisy.

Gain-loss comparison. Panel B of Figure 6 compares the encoded γ estimates across outcome domains but reveals no consistent pattern. The mean and median values of γ^+ are 0.73 and 0.67, respectively, while for γ^- , both values are 0.71. To control for potential confounds, we turn to the meta-regression results in Table 5, which suggest weaker probability distortions in the loss domain than in the gain domain, as indicated by a positive coefficient. Finally, we analyze studies that report likelihood sensitivity parameter estimates for both outcome domains. First, we observe a significant correlation between likelihood sensitivity across domains ($r = 0.46$, $p < 0.01$). Panel C of Figure 6 shows that γ^- exceeds γ^+ in 136 cases, compared to just 50 cases where the reverse holds. A nonparametric test confirms that likelihood sensitivity is significantly higher for losses than for gains ($p < 0.01$). This highlights the value of isolating subsets of data where causal interpretations are more defensible: in this instance, the causally interpretable subset reveals an effect direction opposite to that observed in the aggregate analysis.

Functional forms. It is informative to explore whether the choice of probability weighting function specification influences the estimated likelihood sensitivity parameter. The

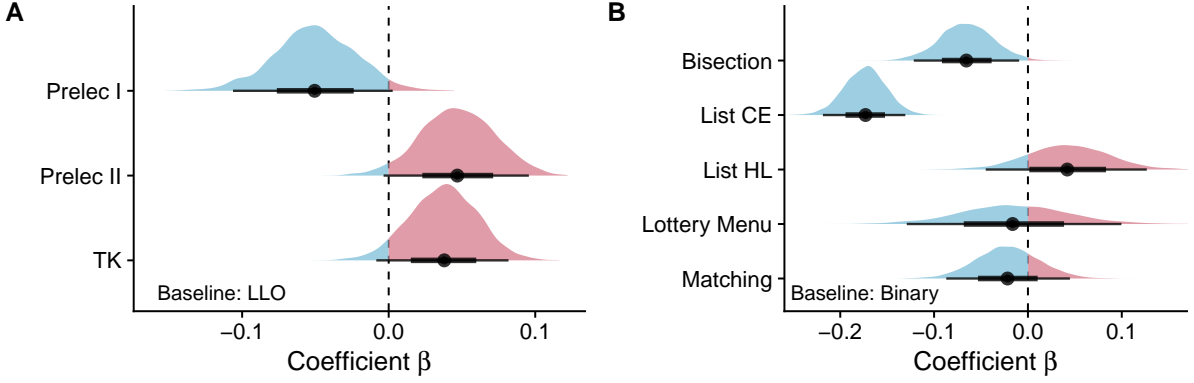


FIGURE 7: Posterior distributions of meta-regression coefficients. (A) Different PWF forms against LLO. (B) Different elicitation methods compared to binary choices. *Notes:* The posterior distributions of the Bayesian random-effects meta-regression coefficient(s) β , along with the posterior medians (represented by a black dot), 66% credible intervals (indicated by thick solid lines), and 95% credible intervals (shown as thin solid lines), are displayed.

meta-regression results, presented in Panel A of Figure 7, indicate that, relative to the baseline LLO form, the Prelec I specification is associated with lower γ values by approximately 0.05. In contrast, Prelec II and TK yield slightly higher γ estimates by 0.05 and 0.04, respectively. However, these effects are only marginally supported regarding the 90% CrI. We emphasize that these associations should not be interpreted causally: the selection of functional form may be endogenous to the elicitation method, or to other experimental features, implying that causal claims would require conditioning on the data set used to produce the estimates.

Procedure invariance. As demonstrated earlier, utility curvature tends to be less pronounced when responses are elicited using CE choice lists and bisection methods compared to experiments with binary choices. Panel B of Figure 7 shows that estimates of likelihood sensitivity are also influenced by the measurement method. Specifically, likelihood sensitivity is significantly lower when elicited via CE choice lists and bisection methods than via binary choices, but it is similar to binary choice when elicited via direct matching. Overall, these findings indicate that methodological choices in preference elicitation can substantially impact the reported values of likelihood sensitivity. While our inference cannot be interpreted causally, [Bouchouicha et al. \(2024\)](#) report an experiment that provides evidence supporting a causal interpretation of the difference between binary choices and certainty equivalents.

Explained versus unexplained heterogeneity. Table 5 presents the meta-regression results. Despite the factors discussed above, only 37.2% of the heterogeneity in likelihood sensitivity is explained. This indicates that a significant portion of the variation in estimates remains unexplained, highlighting the complexity and diversity of the factors

TABLE 5: Meta-regression analysis of likelihood sensitivity.

Category	Variable	Median	2.5%	97.5%	Category	Variable	Median	2.5%	97.5%
Sign	Loss	0.03	0.00	0.06	Subject	Univ. pop.	Baseline		
Utility	Expo	0.11	0.05	0.16		General	-0.06	-0.14	0.02
PWF	LLO	Baseline				Other	-0.05	-0.11	0.02
	Gul/Power				Reward	Real Money	Baseline		
	Prelec I	-0.05	-0.11	0.00		Hypo. Money	-0.10	-0.14	-0.06
	Prelec II	0.05	0.00	0.10		Other	-0.16	-0.24	-0.10
	Elicitation	TK	0.04	-0.01	0.08	Data	Lab	Baseline	
Binary Choice		Baseline			Class		-0.02	-0.10	0.05
Bisection		-0.07	-0.12	-0.01	Field		0.19	0.09	0.27
List CE		-0.17	-0.22	-0.13	Online		0.04	-0.04	0.12
List HL		0.04	-0.04	0.13	Continent	Europe	Baseline		
Matching		-0.02	-0.09	0.04		Africa	-0.09	-0.19	0.01
Other		-0.02	-0.13	0.10		Asia	-0.02	-0.07	0.02
Estimate	Aggregate	Baseline			C/S-America	0.00	-0.08	0.07	
	Ind. Mean	0.03	-0.01	0.07	North America	0.09	0.05	0.14	
	Ind. Median	0.01	-0.03	0.05	Oceania	0.09	-0.06	0.23	

influencing γ . As with other parameters, the elicitation method emerges as the most significant predictor of variation across studies in our dataset, accounting for 35.3% of the explained variation.

5.3 Elevation δ

For the elevation parameter δ , we exclude 383 observations associated with the two one-parameter PWF specifications, Prelec I and TK. The Prelec I specification fixes δ at 1, while the TK function does not include this parameter. Additionally, due to the limited number of estimates for the Power and Gul functions and their similarity, we combine them for further heterogeneity analysis.

Descriptive statistics. We begin by examining the overall distribution of the elevation parameter. For gains, this parameter reflects optimism when $\delta > 1$ and pessimism when $\delta < 1$ (with the opposite interpretation for losses). As shown in Panel A of Figure 8 (solid red curve), the mean and median values are 1.14 and 1.00, respectively, with the IQR of [0.83, 1.23]. The distribution is skewed to the right due to a small number of exceptionally high estimates (26 values exceed 2, with the highest reaching 5.9).

Parameter distributions. Panel A compares the distributions of the raw data parameters to those of the estimated latent parameters. Due to shrinkage effects, the estimated densities, $\hat{\delta}_i$, exhibit a noticeable shift, with higher density concentrated around 1. The analysis shows that the mean of δ is 0.99, with a 95% CrI of [0.96, 1.03], suggesting only a limited degree of elevation, if any.

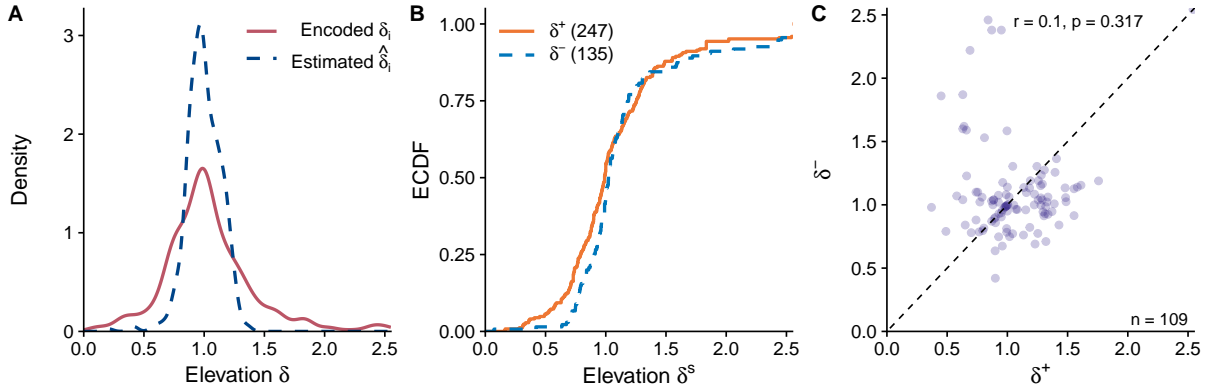


FIGURE 8: (A) Density plot comparing encoded δ_i values to their corresponding estimates $\hat{\delta}_i$. (B) Empirical CDFs of the elevation parameter, separated by gain and loss domains. (C) Scatter plot of parameter estimates for studies reporting both gain and loss values. *Notes:* The x -axis is truncated for improved visualization. However, density estimations include all observations, including those beyond the displayed range.

Gain-loss comparison. Figure 8 compares encoded δ estimates across outcome domains. As shown in Panel B, the empirical CDFs for gains and losses are nearly indistinguishable. This is reflected in their respective central values: for δ^+ , the mean and median are 1.13 and 0.99; for δ^- , they are 1.17 and 1.02. Our meta-regression analysis (Table 6) likewise reveals no evidence of a sign effect for δ , with the coefficient being both economically and statistically insignificant. We also examine studies that report elevation parameter estimates for both gain and loss domains. The pattern in Panel C reinforces our earlier findings: the mean of δ^+ is 1.09, while the mean of δ^- is 1.13 ($p = 0.34$), indicating no significant difference. Additionally, the correlation between gain and loss parameters is low and statistically insignificant ($r = 0.097, p = 0.317$).

Functional forms and procedure invariance. Panel A of Figure 9 shows that parameter estimates from the two most widely used two-parameter PWFs are statistically indistinguishable. In addition, estimates based on the power or Gul weighting functions also suggest similar probability distortion. As shown in Panel B of Figure 9, our meta-regression results indicate that elevation parameters are unaffected by the elicitation method. Table 6 presents the full meta-regression results, including all covariates.

Explained versus unexplained heterogeneity. The proportion of explained heterogeneity for elevation is 18.8%, notably lower than that for utility curvature (66.8%). This finding highlights the complexity of the factors shaping individual probability perception, suggesting that a broad range of influences, beyond the study characteristics considered here, may drive variation in individual optimism/pessimism.

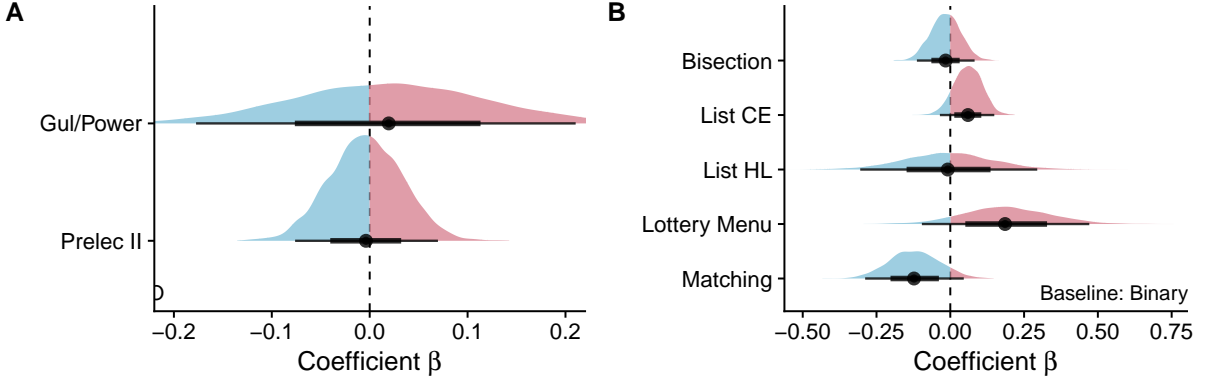


FIGURE 9: Posterior distributions of meta-regression coefficients. (A) Different PWF forms against LLO. (B) Various elicitation methods against Binary Choices. *Notes:* The posterior distributions of the Bayesian random-effects meta-regression coefficient(s) β , along with the posterior medians (represented by a black dot), 66% credible intervals (indicated by thick solid lines), and 95% credible intervals (shown as thin solid lines), are displayed.

TABLE 6: Meta-regression analysis of elevation.

Category	Variable	Median	2.5%	97.5%	Category	Variable	Median	2.5%	97.5%
Sign	Loss	0.02	-0.05	0.09	Subject	Univ. Pop.	Baseline		
Utility	Expo	0.12	0.02	0.21		General	0.16	0.02	0.31
PWF	LLO	Baseline				Other	0.13	0.01	0.25
	Gul/Power	0.02	-0.18	0.21	Reward	Real Money	Baseline		
	Prelec I					Hypo. Money	-0.08	-0.16	0.02
	Prelec II	0.00	-0.08	0.07		Other	0.14	-0.03	0.30
	Elicitation	TK				Data	Lab	Baseline	
Binary Choice		Baseline			Class		0.07	-0.13	0.28
Bisection		-0.02	-0.11	0.08	Field		-0.19	-0.37	-0.01
List CE		0.06	-0.04	0.15	Online		-0.20	-0.40	0.01
List HL		-0.01	-0.31	0.29	Continent	Europe	Baseline		
Matching		-0.12	-0.29	0.05		Africa	0.31	0.11	0.50
Other	0.19	-0.10	0.47	Asia		0.06	-0.05	0.16	
Estimate	Aggregate	Baseline			C/S-America	0.08	-0.10	0.25	
	Ind. Mean	0.06	-0.03	0.14	North America	0.05	-0.05	0.16	
	Ind. Median	0.01	-0.08	0.10	Oceania	0.08	-0.20	0.35	

5.4 Parameter Correlations

Figure 10 displays the correlations among the three PT parameters. Utility curvature (or outcome insensitivity) is positively correlated with likelihood sensitivity ($r = 0.26$, $p < 0.01$), but does not correlate with elevation ($r = -0.06$, $p < 0.23$). We further find a weak but statistically significant correlation between the two weighting function parameters ($r = 11$, $p = 0.03$).

Correlations between PT parameters have received relatively little attention in the literature. One potential source of such correlations is measurement and econometric noise. Specifically, when measurement error is present, parameters can be difficult to disentangle (Zeisberger, Vrecko and Langer, 2012); for instance, elevation and utility curvature may

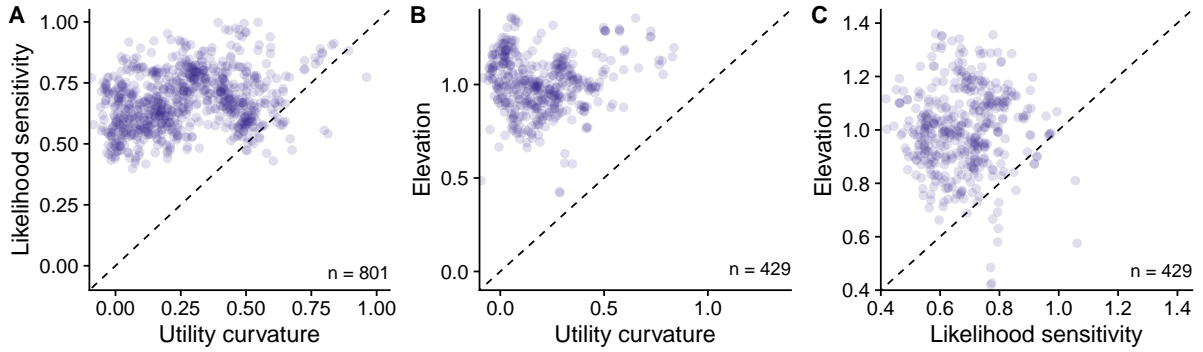


FIGURE 10: Scatter plots of the parameter estimates. Spearman’s correlation coefficient: (A) $r = 0.26$, $p < 0.01$, (B) $r = -0.06$, $p = 0.23$, (C) $r = 0.11$, $p = 0.03$. *Notes:* The x - and y -axes are truncated for improved visualization.

reflect overlapping motivational constructs, and utility curvature may be interdependent with likelihood sensitivity. Correlations may also arise from uncontrolled between-study heterogeneity or from substantive, structural relationships rooted in the underlying psychological or behavioral processes that drive the observed parameter values (Vieider, 2024b).

5.5 Testing for Publication Bias

Meta-analysis offers a powerful means to quantitatively synthesize findings from the literature. However, its inferences are only as reliable as the data fed into the model. A key challenge arises when certain types of results are more likely to be reported by authors, or published by editors, than others.¹⁴ For example, the early focus on inverse-S shaped probability weighting may have discouraged the reporting or publication of findings suggesting S-shaped functions. Whether this occurred in the PT literature remains unclear; indeed, at some point, results deviating from standard findings may have become more publishable than yet another replication of common patterns. Moreover, many PT estimates appear in papers with different primary objectives, and the presence of multiple parameters makes it unclear which, if any, are subject to publication bias (to wit, convex utility for gains and concave utility for losses are both fairly common in the literature). These considerations make it all the more essential to test for potential publication bias in the estimates included in our dataset.

Figure 11 presents funnel plots for the three PT parameters, a standard visual tool for

¹⁴This phenomenon is often referred to as “publication bias” or the “file-drawer problem.” In the context of PT parameters, two main sources of such bias are possible. First, journals may favor parameter estimates that align with canonical values, such as the utility curvature of 0.12 reported in Tversky and Kahneman (1992), and exhibit skepticism toward deviations from this benchmark. For bias to occur, the favored estimate must systematically diverge from the “true” value that would emerge from a broader population of studies (Borenstein et al., 2009). A second form of bias involves editorial preference for statistically significant results, often defined by a p -value below 0.05, which signals rejection of a null hypothesis (Andrews and Kasy, 2019; Brodeur, Cook and Heyes, 2020; Chopra et al., 2024).

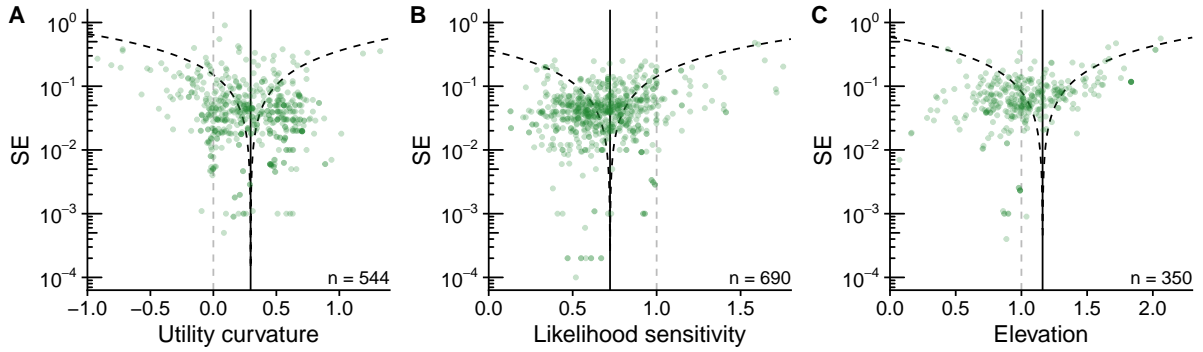


FIGURE 11: Relationship between reported parameter estimates and associated standard errors. *Notes:* Only estimates accompanied by standard errors are included. The solid vertical line represents the aggregate mean θ , while the dashed gray line indicates the neutrality benchmark ($x = 0$ or $x = 1$). The two dashed curves mark the boundaries for statistically significant deviations from the mean parameter value. The x -axis is truncated, and the y -axis is displayed on a log scale for improved visualization.

assessing potential publication bias (Borenstein et al., 2009; Stanley and Doucouliagos, 2012). The plots include all complete observations with both parameter estimates and associated standard errors. In the absence of publication bias, the data points should be symmetrically distributed around the “true” effect size, indicated by the vertical solid line. Less precise studies (i.e., those with larger standard errors) are expected to scatter more widely due to sampling variability but should still do so symmetrically around the true value. Asymmetry among these less precise estimates—appearing higher on the y -axis—is commonly interpreted as evidence of publication bias. In such cases, the average estimate may no longer reflect the true underlying effect size.

Egger’s test reveals no significant funnel plot asymmetry for utility curvature ($p = 0.38$).¹⁵ For likelihood sensitivity, the test similarly fails to detect significant asymmetry ($p = 0.16$). However, when retaining the top 1% of outliers, the test indicates a significant bias in favor of larger γ estimates. Such estimates are relatively common in our dataset and tend to be associated with larger standard errors, possibly due to specific measurement methods. To investigate this, we re-ran Egger’s test while controlling for study-level characteristics. Although the magnitude and statistical significance of the asymmetry are reduced, the effect remains significant at conventional levels. Finally, we observe strong asymmetry in Panel C for the elevation parameter δ ($p < 0.01$). While selection based on optimism or pessimism is arguably less likely, the observed asymmetry may stem from the truncated distribution of estimates, i.e., the failure of the normality assumption implicit in both the funnel plot and Egger’s test.

¹⁵Egger’s test is a parametric method that formally assesses funnel plot asymmetry (Egger et al., 1997). It involves a weighted regression of effect size estimates on their precision (typically the inverse of the standard error or its logarithm). To enhance robustness, we trimmed the top 1% of outliers before conducting the analysis.

6 Discussion

Prospect theory has proved a remarkably popular tool for quantifying risk-taking, as evidenced by the 166 empirical papers included in this meta-analysis. Our study offers a rigorous quantitative synthesis of its core model parameters. On average, utility curvature estimates reveal diminishing sensitivity to increases in wealth (gains) and, to a somewhat lesser extent, to decreases in wealth (losses). The elevation parameter of the probability weighting function centers around neutral values, suggesting a general absence of optimism or pessimism. Importantly, the meta-analysis provides clear evidence of likelihood insensitivity—the tendency for relative risk aversion to increase systematically with the probability of winning or losing. These central tendencies strongly support the stylized behavioral patterns that originally motivated the development of prospect theory.

At the same time, our findings raise several challenges to prospect theory. Chief among them is the influence of the measurement or elicitation method, which emerged as the most significant predictor of variation in parameter estimates. Specifically, we observed notable differences in utility curvature and sensitivity parameters across methods, such as choice lists versus binary choices, bisection, and direct matching, as well as between choice lists varying outcomes versus probabilities. These inconsistencies violate the principle of procedure invariance, which prospect theory implicitly assumes—namely, that preference functionals should yield stable responses regardless of how choices are presented. However, these findings should be interpreted with caution: measurement methods may correlate with unobserved study characteristics that we did not code or control for, precluding causal inference. A more appropriate interpretation is that these results call for rigorously controlled experiments to identify underlying causal mechanisms.

Another mystery from our meta-analysis is the substantial unexplained heterogeneity in parameter estimates. Significant variability persists even after accounting for a wide range of study characteristics, including outcome domain, functional forms, measurement methods, study populations, and incentive structures. This suggests that PT parameter estimates may be sensitive to subtle experimental details. From the perspective of a preference-based model such as prospect theory, it would seem desirable to specifically investigate what might be driving these differences across studies. Elements such as the use of visual aids to represent probabilities, the size of the urn used to convey risk, or even the numerical scale of outcomes may help account for the observed variability (Oprea and Vieider, 2025).

Since prospect theory does not explicitly incorporate the effects of such contextual factors, a promising direction for future research is to endogenize PT-like parameters. Rather than seeing parameters governing choice processes as “preferences” (or at least as exogenous parameters) as in prospect theory, these models depict observed choice regularities as an

outgrowth of cognitive frictions affecting the decision process and optimal ways of dealing with them. Several studies have thus explained decreasing sensitivity towards changes in wealth (e.g., Robson, 2001; Netzer, 2009; Khaw, Li and Woodford, 2021), probability weighting (e.g., Zhang and Maloney, 2012; Steiner and Stewart, 2016; Enke and Graeber, 2023; Herold and Netzer, 2023; Frydman and Jin, 2023; Oprea and Vieider, 2024), or both (e.g., Khaw, Li and Woodford, 2023; Vieider, 2024b). Some of these models furthermore make predictions that are specific to the choice context: while Vieider (2024b) presents a model of probability distortions in binary choice, the models proposed by Khaw, Li and Woodford (2023) and Bouchouicha et al. (2024) are specific to valuations or choice lists. We hope that the continued development of such models will help illuminate the puzzling degree of true between-study heterogeneity we have documented in this paper.

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Online Appendix

Meta-Analysis of Risk-Taking Propensities

Taisuke Imai Salvatore Nunnari
Jilong Wu Ferdinand M. Vieider

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A Data

A.1 Paper Search and Screening

We conducted a search for pertinent literature within the Web of Science, a scientific citation indexing database. After several rounds of trial and error to refine our approach, we settled on the following combination of query terms for our search.

$$\left(\begin{array}{l} \text{"prospect theory"} \\ \text{OR "probability weighting" OR "probability distortion"} \\ \text{OR ("risk preference" AND ("risk attitudes" OR "ambiguity attitudes"))} \end{array} \right) \\ \text{AND} \\ (\text{estimat* OR measur* OR experiment* OR survey})$$

FIGURE A.1: Keywords used in the search.

The initial search, conducted in the spring of 2023, yielded 2,034 papers. In the first phase of paper identification, we reviewed the titles and abstracts, eliminating 1,453 papers that were evidently not relevant to our study. Subsequently, we thoroughly examined the remaining papers, applied our inclusion criteria focusing on content, and proceeded to code the information. Additionally, we utilized IDEAS/RePEc and Google Scholar to locate unpublished working papers.

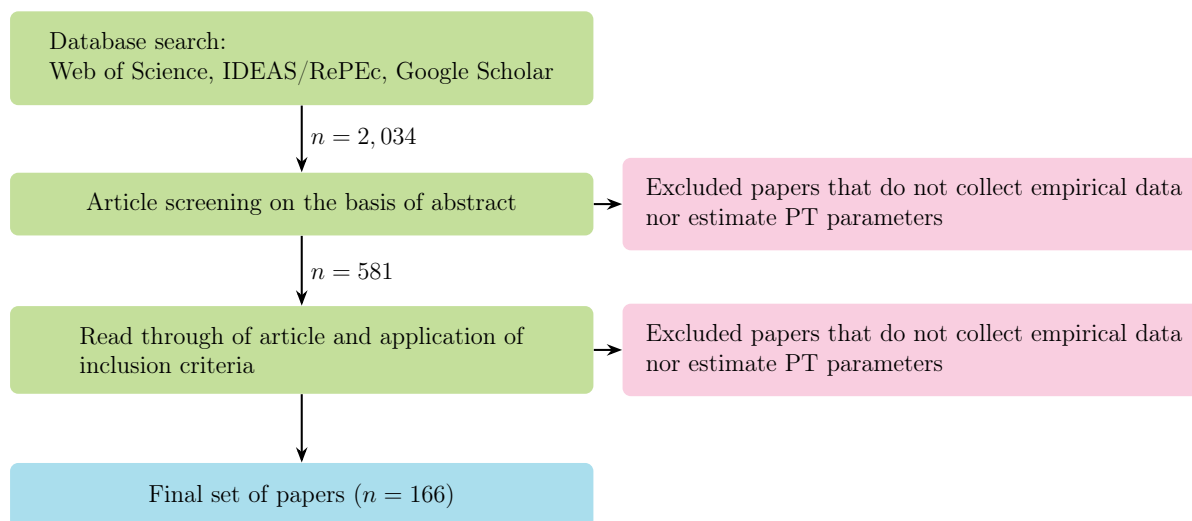


FIGURE A.2: Paper search and data construction.

A.2 Approximation and Imputation of Missing Standard Errors

There are different standard error information sources related to the parameter estimates. We explain below how we calculate SEs using different sources.

- 387 estimates are reported with exact SEs values in the paper.
- 170 estimates are accompanied by standard deviations. We derived $se = SD/\sqrt{n}$.
- 32 SEs were derived from 95% confidence intervals $[lb, ub]$ as $se = (ub - lb)/3.92$.
- 12 SEs are calculated from the effect size (ES) and t -value: $se = |ES|/t$.
- Interquartile ranges are reported for 52 estimates, for which we approximated SEs using $se \approx IQR/(1.35 \times \sqrt{n})$.
- 95% credible intervals or 95% highest density intervals are reported for 31 estimates. We treated these intervals as confidence intervals and followed the above-mentioned approach, $se \approx (ub - lb)/3.92$.
- 12 observations have provided p -value, from which we calculated SEs according to $se \approx |ES|/\Phi^{-1}(1 - p)$, where Φ^{-1} is the quantile function of the standard normal distribution.
- For 3 estimates that provide the maximal and minimal values, $se \approx (Max - Min)/(4 \times \sqrt{n})$.

Note that these approximation formulas are deemed valid when the parameters exhibit a normal distribution within the population. We acknowledge this is a strong assumption for our dataset. Despite this assumption, opting for an “approximated” standard error was considered preferable to discarding the observation altogether or resorting to alternative, potentially stronger assumptions to retain the observation.

A.3 Coded Variables

TABLE A.1: List of coded variables.

Variable	Description
<i>Article meta data</i>	
title	Title of the paper
author_lastnames	Last names of the authors
author_firstnames	First names of the authors
published	1 = published paper
journal	Journal
year	Year of publication
num_subject	Number of subjects
num_choice	Number of choices each subject made
num_list	Number of choices list each subject made
<i>Estimates</i>	
res_est_u1	Reported utility function curvature
res_est_u2	Reported extra utility function curvature
res_est_pwf_alpha	Reported PWF parameter γ
res_est_pwf_beta	Reported PWF parameter δ
res_est_la	Reported loss aversion coefficient
res_err_u1	SE of utility function curvature
res_err_u2	SE of extra utility function curvature
res_err_pwf_alpha	SE of PWF parameter γ
res_err_pwf_beta	SE of PWF parameter δ
res_err_la	SE of loss aversion coefficient
<i>Model features</i>	
u_form	Utility function specification adopted
u_common_gain_loss	1 = common u is assumed for gains and losses
pwf_form	PWF specification adopted
pwf_common_gain_loss	1 = common PWF is assumed for gains and losses
pwf_num_parameters	Number of parameters of the PWF
<i>Type of data</i>	
exp_lab	1 = the data is collected in lab
exp_field	1 = the data is collected in field
exp_class	1 = the data is collected in classroom
exp_online	1 = the data is collected online or via survey

Continued on next page.

Variable	Description
<i>Type of elicitation</i>	
choice_bisection	1 = the bisection setup is used
choice_binary	1 = the binary choice setup is used
choice_list	1 = the choice list setup is used
choice_matching	1 = the direct matching is used
choice_menu	1 = the lottery menu is used
<i>Level of measurement</i>	
est_aggregate	1 = aggregate level estimate
est_aggregate_median_data	1 = aggregate level (“median subject”) estimate
est_individual	1 = individual level estimate
est_mixture	1 = mixture model estimation
<i>Subject pool</i>	
subject_uni	1 = university subjects are recruited
subject_general	1 = general subjects are recruited
subject_other	1 = special subjects are recruited
subject_other_type	specify the population when subject_other = 1
<i>Reward type</i>	
reward_money	1 = real monetary reward
reward_hypothetical	1 = hypothetical monetary choices
reward_other	1 = non monetary reward (other)
<i>Location of the experiment/survey</i>	
location_country	Country location of the experiment
location_continent	Continent location of the experiment
<i>Estimation Strategy</i>	
est_strategy	Description of estimation strategy
est_loss	1 = parameter estimate for loss domain
<i>Estimation Strategy</i>	
interface_visual_aids	1 = lotteries represented with visual aids
interface_iconic_express	1 = lotteries represented with iconic visual aids

A.4 Collected Studies for Meta-Analysis

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A.5 Journals

Table A.2 provides a summary of the journals in which the papers included in our dataset were published. Journal categories are based on the classification provided by The Master Journal List (<https://mjl.clarivate.com/home>).

TABLE A.2: List of journals.

Journal	Category
1 Agricultural Economics	Economics
2 American Economic Review	Economics
3 American Journal of Agricultural Economics	Economics
4 Attention, Perception & Psychophysics	Psychology
5 BioPsychoSocial Medicine	Psychology, Multidisciplinary
6 Cognition	Psychology, Experimental
7 Cognitive Psychology	Psychology
8 Decision	Management
9 Decision Support System	Operations Research & Management Science
10 Ecological Economics	Economics
11 Econometrica	Economics
12 Economic Development and Cultural Change	Economics
13 Economic Inquiry	Economics
14 Economica	Economics
15 Ekonomický časopis	Economics
16 eNeuro	Neurosciences
17 Environmental and Resource Economics	Economics
18 European Review of Agricultural Economics	Economics
19 Experimental Economics	Economics
20 Frontiers In Psychology	Psychology, Multidisciplinary
21 Games and Economic Behavior	Economics
22 Geneva Risk and Insurance Review	Economics
23 Health Economics	Economics
24 Healthcare	Health Policy & Services
25 International Economic Review	Economics
26 Journal of Banking & Finance	Economics
27 Journal of Behavioral and Experimental Economics	Economics
28 Journal of Behavioral Decision Making	Psychology, Applied
29 Journal of Behavioral Finance	Economics
30 Journal of Development Economics	Economics
31 Journal of Development Studies	Economics
32 Journal of Econometrics	Economics
33 Journal of Economic Behavior & Organization	Economics
34 Journal of Economic Theory	Economics
35 Journal of Experimental Psychology: General	Psychology, Experimental
36 Journal of Experimental Psychology: Learning	Psychology
37 Journal of Experimental Social Psychology	Psychology, Social
38 Journal of Mathematical Psychology	Psychology, Mathematical
39 Journal of Money and Economy	N.A.
40 Journal of Neuroscience	Neurosciences
41 Journal of Risk and Uncertainty	Economics
42 Journal of the Economic Science Association	Economics
43 Judgment and Decision Making	Psychology, Multidisciplinary
44 Journal of Behavioral Decision Making	Psychology, Applied

Continued on next page.

Journal	Category
45 Management Science	Management
46 Medical Decision Making	Health Policy & Services
47 New Zealand Economic Papers	Economics
48 Operations Research	Management
49 Organizational Behavior and Human Decision Processes	Management
50 PeerJ	Multidisciplinary Sciences
51 PLOS Neglected Tropical Diseases	Parasitology
52 PLOS ONE	Multidisciplinary Sciences
53 Psychological Medicine	Psychology
54 Psychological Science	Psychology, Multidisciplinary
55 Psychonomic Bulletin & Review	Psychology, Experimental
56 Quantitative Economics	Economics
57 Review of Behavioral Economics	Economics
58 Review of Economics and Statistics	Economics
59 Revista de Administração de Empresas	Management
60 Scientific Reports	Multidisciplinary Sciences
61 Social Choice and Welfare	Economics
62 Social Cognitive and Affective Neuroscience	Neurosciences
63 Social Science & Medicine	Social Sciences, Biomedical
64 Southern Economic Journal	Economics
65 Spanish Journal of Finance and Accounting	Business, Finance
66 Theory and Decision	Economics
67 Transportation Research Part A	Economics
68 Transportation Research Part B	Operations Research & Management Science
69 Transportation Research Part C	Transportation Science & Technology
70 Water Resource and Economics	Economics

A.6 Global Map

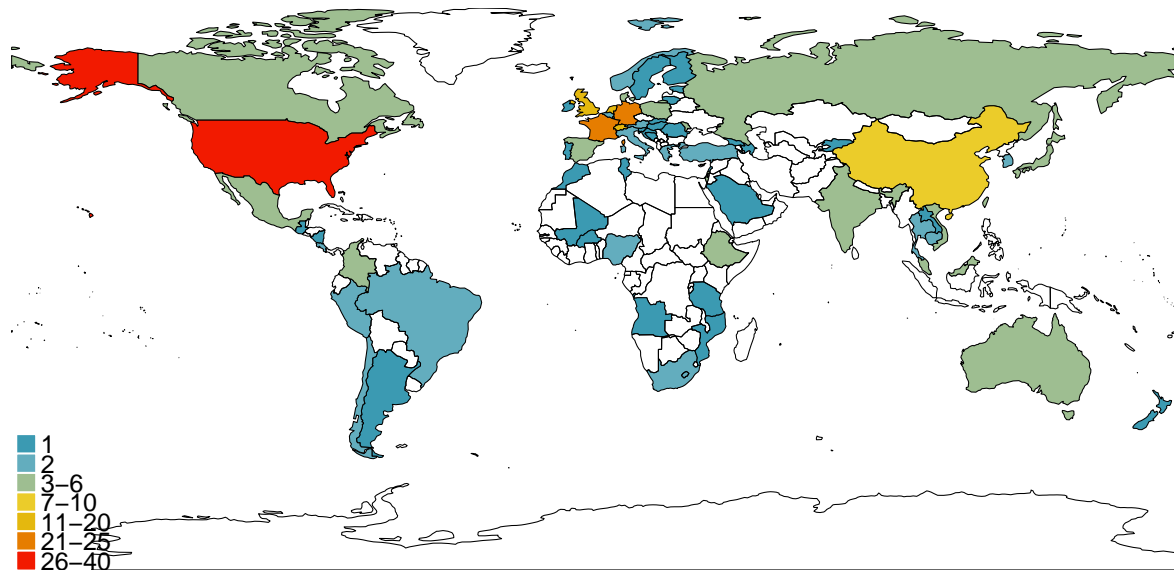


FIGURE A.3: Study location. *Notes:* This map was created using R (<https://www.r-project.org/>) on a base world map obtained from Natural Earth (<https://www.naturalearthdata.com/>).

B Standard Error Imputation

In our dataset, 113 out of 812 estimates lack associated standard errors. To address those missing SEs, the fundamental approach involved estimating the parameters characterizing their distribution in the data, represented as $\log(\mathbf{se}_i) \sim \mathcal{N}(\mathbf{Z}_i \boldsymbol{\xi}, \boldsymbol{\Omega})$. Here, $\log(\mathbf{se}_i)$ be a vector of element-wise natural logarithms of the standard errors, which serve to enforce non-negativity in the definition of standard errors, \mathbf{Z}_i is a $1 \times M$ vector of characteristics of study i predictive of its standard errors, $\boldsymbol{\xi}$ is a $M \times 4$ matrix of coefficients, and $\boldsymbol{\Omega}$ is a covariance matrix with variances on its main diagonal, and covariances in its off-diagonal cells. In terms of predictors in \mathbf{Z} , we include characteristics—the square root of the number of subjects, experiment location, parameter sign, continent, utility function forms, and probability weighting function forms—in addition to a column of 1s to capture the intercept. In total, these characteristics accommodate standard error variations well: 31.5%, 21.6%, 41.3%, and 42.5% of the total variation explained for the four parameters, respectively.

Moreover, the vector $\log(\mathbf{se}_i)$ has a distinctive structure: it includes the logarithm of observed standard errors for studies where these are available, and the logarithm of *predicted* standard errors for cases where the standard errors are missing and must be inferred. The model thus fulfills two roles: (1) it estimates the regression coefficients $\boldsymbol{\xi}$ from the observed study characteristics in \mathbf{Z} ; and (2) it uses these characteristics, along with the estimated coefficients $\boldsymbol{\xi}$, to *predict* (or impute) standard errors where they are not observed. The correlation structure encoded in $\boldsymbol{\Omega}$ ensures that the dependencies among errors are properly accounted for during the imputation process. As Figure B.1 shows, the mutual correlation of these three PT parameter estimates' standard errors exists, which supports the validity of our approach.

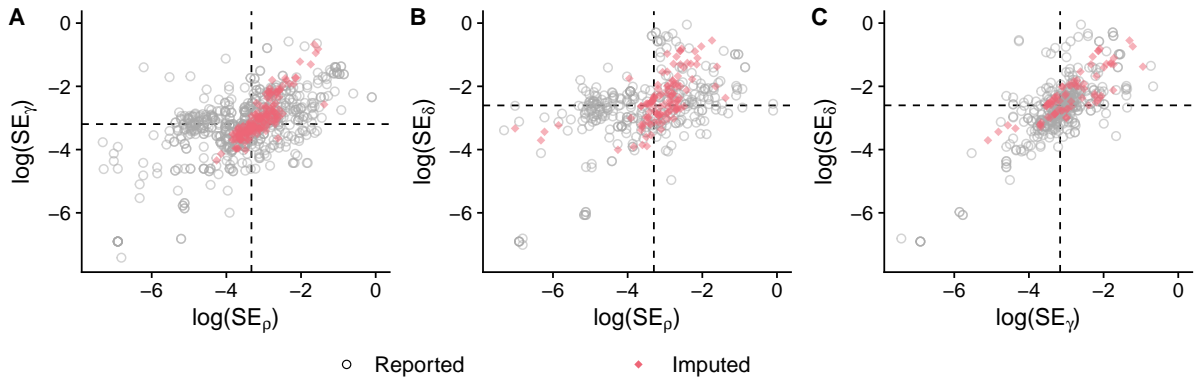


FIGURE B.1: Scatter plot of reported PT parameter and $\log(\text{SE})$. *Notes:* Gray open circles represent reported standard errors (SEs), while red diamonds indicate imputed SEs. Dashed lines denote the medians of the corresponding variables.

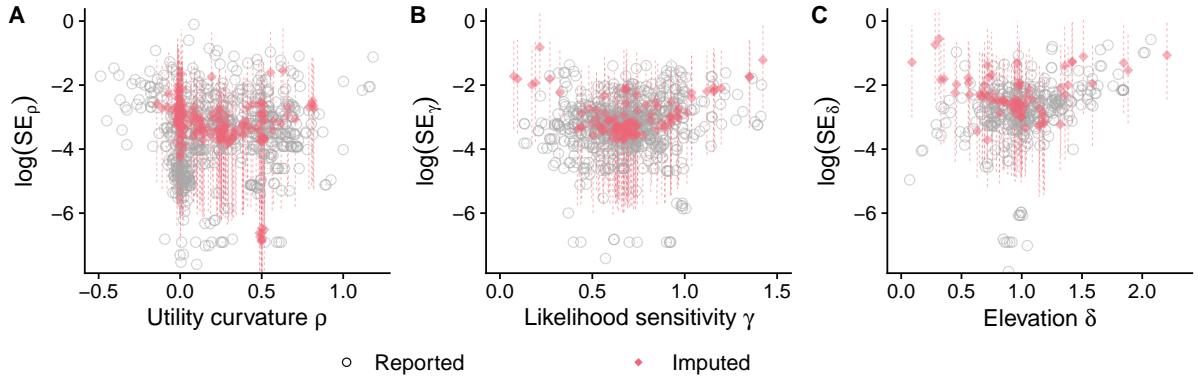


FIGURE B.2: Scatter plot of reported PT parameter and $\log(SE)$. *Notes:* Gray open circles represent complete observations, while red diamonds mark observations with imputed standard errors. Dashed red lines represent the percentile range from 2.5% to 97.5% across the 4,000 posterior draws for SE_i^{imputed} .

B.1 Balance Check

A key question is whether studies that report estimates' standard errors differ systematically from those that do not. Figure B.2 presents a scatter plot of the PT parameter estimates and their associated standard errors. Gray open circles denote complete observations, while red diamonds indicate incomplete observations with imputed standard errors. A visual inspection suggests that the incomplete observations are distributed evenly across the full range of parameter values. Comparing central tendencies, the differences between complete and incomplete observations are generally small: 0.26 vs. 0.22 for ρ , 0.72 vs. 0.73 for γ , and 1.16 vs. 1.05 for δ .¹⁶ The Wilcoxon test indicates a marginally significant difference in reported estimates of elevation ($p = 0.06$), but no significant differences for likelihood sensitivity ($p = 0.34$) or utility curvature ($p = 0.16$). These differences may reflect heterogeneity in study characteristics, as documented in the main text. To partially address this concern, we examine the subsample in which a power utility function is assumed. In this subsample, the differences in all three key prospect theory parameters become insignificant, including that for δ ($p = 0.45$).

To eliminate concerns about potential imbalance, the next subsection presents meta-analysis results based solely on complete observations. The results are essentially unchanged.

¹⁶In the t -test, we find only a marginally significant difference for utility curvature ρ ($p = 0.07$).

B.2 Robustness Check

To ensure that our results are not biased by our standard-error imputation practice, this subsection reports the results of the meta-analysis, which include only estimates with associated standard errors. Table B.1 reports the weighted average of parameter estimates across gains and losses, while Tables B.2, B.3, and B.4 reports corresponding meta-regression results.

TABLE B.1: Meta-analysis of complete PT estimates.

	ρ		γ		δ	
	Mean	95% CrI	Mean	95% CrI	Mean	95% CrI
Gains	0.33	[0.31, 0.36]	0.68	[0.66, 0.71]	0.99	[0.94, 1.04]
Losses	0.28	[0.24, 0.32]	0.70	[0.66, 0.73]	0.98	[0.91, 1.04]

TABLE B.2: Robustness of meta-regression analysis of utility curvature (cf. Table 4).

Category	Variable	Median	2.5%	97.5%	Category	Variable	Median	2.5%	97.5%
Sign	Loss	-0.08	-0.12	-0.05	Subject	Univ. pop.	Baseline		
Utility	Expo	-0.10	-0.16	-0.03		General	0.05	-0.04	0.15
PWF	LLO	Baseline				Other	-0.02	-0.11	0.07
	Gul/Power	0.08	-0.08	0.23	Reward	Real money	Baseline		
	Prelec I	0.08	0.01	0.14		Hypo. money	-0.05	-0.11	0.00
	Prelec II	0.07	0.01	0.14		Other	-0.11	-0.19	-0.02
	Elicitation	TK	-0.02	-0.08	0.04	Data	Lab	Baseline	
Binary choice		Baseline			Class		0.38	0.28	0.48
Bisection		-0.08	-0.15	-0.01	Field		0.16	0.06	0.27
List CE		-0.27	-0.33	-0.22	Online		-0.10	-0.19	0.00
List HL		-0.03	-0.13	0.06	Continent	Europe	Baseline		
Matching		-0.09	-0.17	-0.01		Africa	0.15	0.07	0.25
Other	0.37	0.23	0.52	Asia		0.02	-0.04	0.08	
Estimate	Aggregate	Baseline				C/S-America	0.12	0.03	0.20
	Ind. mean	0.04	-0.01	0.09	North America	0.12	0.07	0.18	
	Ind. median	0.01	-0.05	0.06	Oceania	0.18	0.05	0.31	

TABLE B.3: Robustness of meta-regression analysis of likelihood sensitivity (cf. Table 5).

Category	Variable	Median	2.5%	97.5%	Category	Variable	Median	2.5%	97.5%
Sign	Loss	0.06	0.01	0.10	Subject	Univ. pop.	Baseline		
Utility	Expo	0.08	0.01	0.14		General	-0.28	-0.38	-0.16
PWF	LLO	Baseline				Other	-0.15	-0.25	-0.05
	Gul/Power				Reward	Real money	Baseline		
	Prelec I	-0.04	-0.12	0.03		Hypo. money	-0.09	-0.14	-0.03
	Prelec II	0.06	-0.01	0.13		Other	-0.27	-0.37	-0.18
	TK	0.04	-0.02	0.11	Data	Lab	Baseline		
Elicitation	Binary choice	Baseline				Class	-0.07	-0.17	0.02
	Bisection	-0.05	-0.12	0.03		Field	0.50	0.37	0.63
	List CE	-0.24	-0.30	-0.18		Online	0.15	0.04	0.25
	List HL	-0.03	-0.14	0.09	Continent	Europe	Baseline		
	Matching	-0.04	-0.13	0.04		Africa	-0.09	-0.19	0.02
	Other	0.16	0.02	0.31		Asia	-0.04	-0.10	0.03
Estimate	Aggregate	Baseline				C/S-America	-0.04	-0.13	0.05
	Ind. mean	0.00	-0.05	0.05		North America	0.13	0.07	0.19
	Ind. median	-0.06	-0.12	0.01		Oceania	0.07	-0.07	0.21

TABLE B.4: Robustness of meta-regression analysis of elevation (cf. Table 6).

Category	Variable	Median	2.5%	97.5%	Category	Variable	Median	2.5%	97.5%
Sign	Loss	-0.03	-0.13	0.07	Subject	Univ. pop.	Baseline		
Utility	Expo	0.20	0.07	0.33		General	0.24	0.01	0.48
PWF	LLO	Baseline				Other	0.16	-0.03	0.36
	Gul/Power	-0.11	-0.40	0.15	Reward	Real money	Baseline		
	Prelec I					Hypo. money	-0.06	-0.19	0.07
	Prelec II	-0.14	-0.25	-0.02		Other	0.22	-0.05	0.48
	TK				Data	Lab	Baseline		
Elicitation	Binary choice	Baseline				Class	0.23	-0.09	0.54
	Bisection	-0.10	-0.26	0.06		Field	-0.35	-0.67	0.00
	List CE	-0.15	-0.31	0.00		Online	-0.33	-0.63	-0.01
	List HL	0.06	-0.29	0.41	Continent	Europe	Baseline		
	Matching	-0.22	-0.49	0.05		Africa	0.21	-0.03	0.45
	Other	0.23	-0.21	0.70		Asia	0.10	-0.06	0.27
Estimate	Aggregate	Baseline				C/S-America	0.20	0.01	0.41
	Ind. mean	-0.02	-0.16	0.11		North America	0.11	-0.03	0.24
	Ind. median	-0.14	-0.29	0.01		Oceania	0.10	-0.27	0.45

C Loss Aversion Results

In our raw data, the mean and median of the loss aversion coefficient λ are 1.53 and 1.36, respectively, when the utility function is a power function. When the utility function is exponential one, the mean and median of λ are 1.64 and 1.53, respectively. In our joint Bayesian hierarchical estimation, the posterior mean is 1.46 with a 95% CrI of [1.41, 1.51]. All these results support the idea that “Losses loom larger than gains.”

As demonstrated in Figure C.1, loss aversion is negatively correlated with utility curvature and likelihood sensitivity, with Spearman correlation coefficients of $r = -0.25$ and $r = -0.46$, respectively (both $p < 0.01$). In contrast, it is positively correlated with elevation ($r = 0.22$, $p < 0.01$), indicating that higher loss aversion is associated with greater elevation of probability weighting.

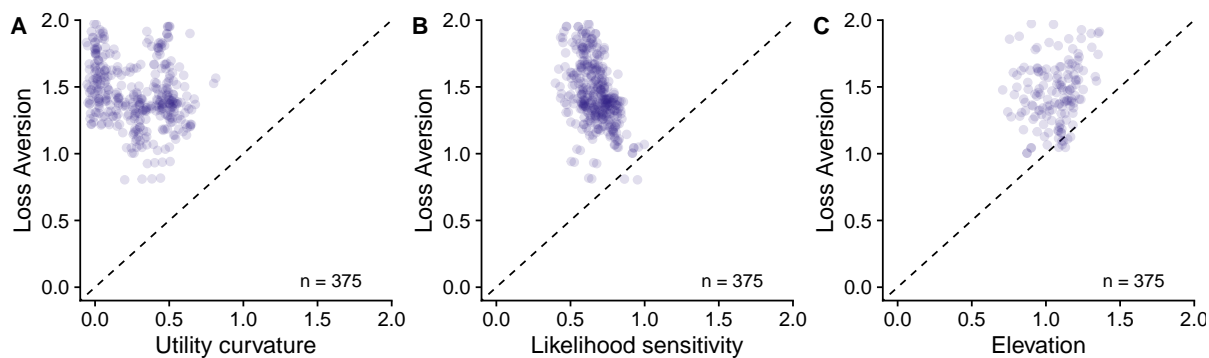


FIGURE C.1: Scatter plots of the parameter estimates. Spearman’s correlation coefficient: (A) $r = -0.25$, $p < 0.01$, (B) $r = -0.46$, $p < 0.01$, (C) $r = 0.22$, $p < 0.01$. *Notes:* The x - and y -axes are truncated for improved visualization.