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Who Converges Toward Whom? The Direction of Risk-Attitude Convergence within Couples

Nicolas Frémeaux
Université de Rouen Normandie (LERN)
and IZA@LISER

SeEun Jung
Inha University

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Abstract

Individual risk attitudes shape household financial decisions, yet most models treat these attitudes as fixed individual traits. Studies of spousal risk attitudes have documented that partners' preferences are correlated, but whether preferences converge within marriage, and *who converges toward whom*, remains an open question. We fill this gap using a balanced panel of 2,331 Korean married couples (KLIPS, 2004 and 2007). We first document that spouses' risk attitudes are positively matched, with a within-couple correlation of about 0.2 in both waves. We estimate the husband-to-wife and wife-to-husband transmission channels separately and show that their relative strength is systematically related to the wife's economic position. When the wife works, her risk attitudes stop responding to her husband's baseline while his begin to respond to hers; when she out-earns him, her responsiveness to his baseline falls sharply. This opposite-sign pattern is difficult to generate through common shocks or differential mean reversion alone. The results are consistent with a bargaining account in which the same economic variables that the collective-household literature identifies as determinants of resource allocation also govern the direction of preference convergence.

JEL classification

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couple, marriage, risk attitudes, risk aversion, household

Corresponding author

SeEun Jung
jse@inha.ac.kr

1 Introduction

Individual risk attitudes are central to household financial decisions (Barsky et al., 1997; Browning and Chiappori, 1998; Guiso and Paiella, 2008). While traditional models treat the household as a single decision-maker with common preferences (Becker, 1981), a large body of evidence documents that spouses' risk attitudes, though positively correlated, are far from identical (Kimball et al., 2009; Dohmen et al., 2012; Mazzocco, 2007; Lundberg and Pollak, 2008).

The source of this similarity, however, remains disputed. One strand of the literature attributes it to sorting at the marriage market: couples match on risk preferences at the outset. Cross-sectional evidence of positive assortative mating on risk preferences is robust across contexts and methods (Kimball et al., 2009; Dohmen et al., 2012; Arrondel and Frémeaux, 2016; Nikoloulopoulos and Moffatt, 2019), and Chiappori and Reny (2016) provide theoretical conditions under which positive or negative assortative matching on risk aversion can arise, depending on the household production technology.

A second strand argues that spouses become similar *within* marriage. Bacon et al. (2014) estimate a spousal correlation of +0.285 in risk-taking willingness using the German SOEP and find that it rises with marriage duration, consistent with either socialization or differential exit of mismatched couples. Di Falco and Vieider (2018) document that risk preferences assimilate over the course of a marriage in ways not fully attributable to sorting, while others document persistent gaps linked to asymmetric economic roles (Bernasek and Shwiff, 2001; Jianakoplos and Bernasek, 2008). These two accounts are not mutually exclusive, and distinguishing between them remains an active challenge.

Even among couples who remain together, the question of *how* preferences evolve is far from settled. Serra-Garcia (2022) shows that mismatched couples are more likely to dissolve, so that observed convergence among survivors partly reflects differential exit; but conditional on staying married, the existing convergence literature has a notable blind spot. Studies that document within-marriage convergence typically estimate a single couple-level

distance measure and ask whether it shrinks over time. They do not, in general, identify *who is converging toward whom*. This leaves the direction of adjustment anonymous: we know that spouses become more similar, but not whether one spouse systematically moves more than the other, or what economic variables predict that asymmetry. Identifying the direction matters because it speaks directly to the mechanism: if convergence is driven by household bargaining, as collective models suggest, then the direction should depend on each spouse's economic leverage within the household.

Our main contribution is to identify the direction of within-couple convergence in risk aversion and to show that it is governed systematically by the wife's economic bargaining position. Using a balanced panel of 2,331 married couples from the Korean Labor and Income Panel Study (KLIPS, 2004 and 2007), we estimate first-difference regressions that separately identify the husband-to-wife and wife-to-husband transmission channels. We then interact these channels with two proxies for the wife's bargaining power: employment status and relative earnings.

The Korean setting offers a particularly sharp test of this hypothesis. Korea's gender employment and wage gaps are among the largest in the OECD, and the participation gap is driven almost entirely by married women (Stansbury et al., 2024). In earlier work using the same panel, we document that despite very high educational assortativeness, spousal earnings are negatively correlated in Korea, reflecting low labor-force participation among married women (Frémeaux et al., 2024). For our purposes, this environment generates wide variation across couples in the wife's economic position: a working or higher-earning wife represents a much larger shift in relative economic standing than in settings where dual-earner couples are the norm. If bargaining position governs the direction of preference adjustment, Korea in the mid-2000s is a setting where the asymmetry should be clearly visible. At the same time, Korean couples do not appear unusual in their risk attitudes themselves: our within-couple correlation is close to estimates from European panels (Bacon et al., 2014), suggesting that what is distinctive about the setting is the distribution of bargaining power, not preferences.

We first document that spouses' risk attitudes are positively matched, with a within-couple correlation of about 0.2 in both waves, and that couples converge substantially over the panel window. The key empirical finding is that the wife's economic position systematically weakens the husband-to-wife channel; *for the employment proxy, the same variable is simultaneously associated with an opening of the wife-to-husband channel. When the wife has stronger bargaining power*, her risk attitudes become less responsive to her husband's baseline, while his, if anything, become more responsive to hers. This opposite-sign pattern across equations is difficult to reconcile with common household shocks or pure demographic heterogeneity, and is most naturally interpreted as evidence that bargaining weights govern the direction of preference adjustment.

Two additional findings reinforce this conclusion. First, convergence is substantial overall: 81.7% of couples who are initially discordant in risk attitudes become concordant within three years, and the continuous within-couple distance in risk aversion falls by 35% over the panel window ($\beta = -0.94$, $p < 0.001$). The relatively short three-year window is conservative: if anything, it biases against finding convergence, since preference adjustment is likely to be gradual. Second, a full specification that adds the spouse's contemporaneous change (ΔRA) reveals that the spouse-baseline convergence coefficient, which appears insignificant in the standard specification, becomes large and highly significant in both equations ($+0.133^{***}$ and $+0.249^{***}$), resolving an apparent asymmetry caused by omitted-variable bias.

From a theoretical perspective, our findings contribute to the collective household literature (Chiappori, 1988; Browning and Chiappori, 1998; Mazzocco, 2007). These models predict that bargaining weights allocate *resources*; we present evidence that they also shape *preference formation*. This connects to the cooperative bargaining framework of Lundberg and Pollak (1993), in which threat points shaped by outside options determine the distribution of household outcomes, and to Jianakoplos and Bernasek (2008), who show that household financial risk-taking shifts when the wife earns more. We extend the domain of collective models from the allocation of given preferences to the formation of preferences

themselves.

The remainder of the paper is organized as follows. Section 2 describes the data and measurement. Section 3 presents the empirical strategy. Section 4 reports results. Section 5 provides robustness checks. Section 6 concludes.

2 Data and Measurement

2.1 The Korean Labor and Income Panel Study

We use the Korean Labor and Income Panel Study (KLIPS), a nationally representative longitudinal survey of Korean households conducted annually since 1998 by the Korea Labor Institute. KLIPS interviews all household members aged 15 and older on a wide range of economic and demographic topics, including employment, earnings, assets, and, in selected waves, risk attitudes. Our analysis draws on waves 7 (2004) and 10 (2007), the two survey rounds in which an identical five-item lottery battery was administered to both spouses.[1]

We restrict the sample to married couples in which both spouses responded to all five lottery items in both waves, yielding a balanced panel of 2,331 couples. The sample construction proceeds as follows: we begin with all married couples in which both spouses are aged 18–65 and matched within the same household in 2004 (3,056 couples), retain those present in both waves (2,599 balanced couples), and then keep the 2,331 couples for whom all five lottery items, including a directly recorded response on the key item (r4), are available for both spouses in both waves. The final sample therefore uses only raw, self-reported lottery responses.[2]

[1] This question regarding risk aversion was only asked in these two waves of the survey. We are therefore unable to observe the couple over a longer period or to provide a more recent measurement.

[2] As a check for selection bias from this last restriction, we compared the within-couple RA4 correlation in our analysis sample (0.208 in 2004) with that in the full balanced panel (N = 2,599; correlation 0.239). The difference is small and, if anything, our stricter sample shows a *lower* correlation, suggesting that this restriction does not inflate our convergence estimates.

2.2 Measuring risk aversion

The KLIPS lottery battery presents respondents with five hypothetical gambles, each offering a choice between a risky lottery and a safe cash amount. The five items vary in their expected values (EV) relative to the safe option:

Item	Lottery	Cash (KRW)	EV (KRW)	EV vs. Cash
r1	50% chance of 15,000	10,000	7,500	EV < Cash
r2	50% chance of 20,000	10,000	10,000	EV = Cash
r3	40% chance of 20,000	10,000	8,000	EV < Cash
r4	**60% chance of 200,000**	**100,000**	**120,000**	**EV > Cash**
r5	20% chance of 500,000	100,000	100,000	EV = Cash

We focus on item r4 as our baseline risk-aversion measure. On this item, the expected value of the lottery (KRW 120,000) strictly exceeds the safe cash option (KRW 100,000) by 20%. A risk-neutral individual should prefer the lottery; choosing cash therefore requires forgoing a positive expected-value premium and is difficult to rationalize as ordinary financial prudence. We define $\text{RA4}_i = 1$ if individual i chose cash on r4, and $\text{RA4}_i = 0$ otherwise. By this measure, 94.1% of wives and 84.0% of husbands are classified as risk-averse in 2004, rising to 95.8% and 90.2% in 2007.

The remaining four items do not share this identification advantage. On items r1 and r3, the lottery's expected value is *below* the cash amount, so choosing cash is consistent with risk neutrality or even mild risk-seeking. On items r2 and r5, expected values are equal, making cash selection ambiguous between risk aversion and indifference. Composite indices that sum across all five items (e.g., $\text{RAcount} = \sum_{j=1}^{5} \mathbf{1}[\text{cash}_j]$, ranging from 0 to 5) therefore mix genuine risk aversion with rational conservatism. We use RAcount as a robustness measure in Section 5 and find that it yields qualitatively identical but quantitatively stronger results, consistent with the composite reducing measurement error while introducing some noise from items with weaker identification content.

2.3 Bargaining-power proxies

To examine whether the direction of spousal convergence depends on intra-household bargaining power, we use two proxies of the wife's economic position measured at baseline (2004):

1. *Wife working*: an indicator for whether the wife is employed (= 1 for 46.1% of couples).[3]

2. *Wife earns more*: an indicator for whether the wife's monthly labor income exceeds the husband's (= 1 for 9.7% of all couples, or 23.3% of dual-earner couples).[4]

These proxies are measured at baseline and are therefore predetermined relative to the 2004–2007 change in risk attitudes. We do not claim that they are exogenous to risk preferences in general, but the consistency of the results across two distinct dimensions of economic position strengthens the interpretation. We note that KLIPS does not contain information on parental wealth, inheritance, or social background in the waves we use, limiting our ability to capture non-labor dimensions of bargaining power.

3 Empirical Strategy

3.1 Baseline convergence specification

Our main specification is a first-difference regression estimated separately for wives and husbands:

$$\Delta \mathrm{RA4}_i^{\mathrm{own}} = \alpha + \beta \cdot \mathrm{RA4}_i^{\mathrm{spouse},2004} + \gamma \cdot \mathrm{RA4}_i^{\mathrm{own},2004} + \mathbf{X}_i'\delta + \varepsilon_i \tag{1}$$

[3] A closely related proxy, dual-earner status (both spouses employed, 41.7%), is nearly collinear with wife working (correlation = 0.915) because 88.2% of husbands work. We verified that using dual-earner status in place of wife working produces virtually identical results; these are available upon request.

[4] We also examined relative education, which proxies potential rather than realized earnings: indicators for the wife being strictly more educated than her husband (9.1% of couples) or at least as educated (60.4%). Neither moderates the transmission channels (interaction p-values between 0.23 and 0.86, with point estimates of inconsistent sign), and both are essentially uncorrelated with the employment and earnings proxies ($r \approx 0.03$). Within-household influence appears to respond to realized economic position rather than to potential earnings.

where $\Delta \text{RA4}_i^{\text{own}} = \text{RA4}_i^{2007} - \text{RA4}_i^{2004}$ is the change in own risk aversion, $\text{RA4}_i^{\text{spouse},2004}$ is the spouse's baseline risk aversion, $\text{RA4}_i^{\text{own},2004}$ is own baseline (included to absorb mechanical mean reversion in a bounded outcome), and $\mathbf{X}_i$ is a vector of demographic controls (age and years of schooling for both spouses). In extended specifications, we add household income and dual-earner status. Throughout the paper, we report heteroskedasticity-robust (Huber–White) standard errors.

The coefficient of interest is β: a positive β in the wife equation indicates that wives whose husbands were more risk-averse in 2004 became more risk-averse themselves by 2007, consistent with husband-to-wife convergence. A positive β in the husband equation indicates wife-to-husband convergence. Ex ante, we expect $\beta \geq 0$ in both equations: if spouses converge, each should move toward the other's baseline position. We are agnostic, however, as to whether the two channels are symmetric. A unitary household with common preferences would imply symmetric adjustment, whereas bargaining frameworks predict that the relative strength of the two channels varies with each spouse's economic position; the next subsection develops this prediction.

3.2 Bargaining-power interactions

To test whether the direction of convergence depends on bargaining power, we augment (1) with an interaction between the spouse's baseline and a bargaining-power moderator M_i:

$$\Delta \text{RA4}_i^{\text{own}} = \alpha + \beta_1 \cdot \text{RA4}_i^{\text{sp}} + \beta_2 \cdot (\text{RA4}_i^{\text{sp}} \times M_i) + \gamma_1 \cdot \text{RA4}_i^{\text{own}} + \gamma_2 \cdot (\text{RA4}_i^{\text{own}} \times M_i) + \theta M_i + \mathbf{X}_i' \delta + \varepsilon_i \quad (2)$$

The key coefficient is β_2. Under the bargaining hypothesis, we expect $\beta_2 < 0$ in the wife equation (the husband-to-wife channel weakens when the wife has more bargaining power) and $\beta_2 > 0$ in the husband equation (the wife-to-husband channel strengthens). This opposite-sign pattern across the two equations is the key prediction that distinguishes a

bargaining mechanism from common shocks.

The opposite-sign prediction also distinguishes bargaining from a non-bargaining exposure channel. The wife's employment may reduce her responsiveness to her husband simply through independent exposure: working wives face labor-market risk, financial information, and social networks outside the household, which could anchor their attitudes to outside reference points. Independent exposure alone, however, predicts only a weakening of the husband-to-wife channel ($\beta_2 < 0$ in the wife equation), with no implication for the wife-to-husband channel. Bargaining predicts, in addition, that the reverse channel opens ($\beta_2 > 0$ in the husband equation). We treat the two accounts as complementary, since both reflect a reduction in the wife's economic and informational dependence on her husband, but the reverse channel is what allows the data to speak to bargaining specifically.

A critical feature of this specification is the inclusion of the own-baseline interaction $\gamma_2 \cdot (\text{RA4}_i^{\text{own}} \times M_i)$. Without this term, a significant β_2 could reflect differential mean reversion across groups (e.g., if working wives have higher baseline RA4 and therefore more room to revert). Including both interactions ensures that β_2 captures the moderating effect of M_i on the *spouse's* influence, net of any group-specific mean-reversion pattern.

3.3 Full specification with contemporaneous co-movement

The baseline specification (1) uses temporal precedence for identification: the spouse's 2004 level predicts own 2004–2007 change. A complementary approach adds the spouse's contemporaneous change to the right-hand side:

$$\Delta\text{RA4}_i^{\text{own}} = \alpha + \beta \cdot \text{RA4}_i^{\text{spouse,2004}} + \phi \cdot \Delta\text{RA4}_i^{\text{spouse}} + \gamma \cdot \text{RA4}_i^{\text{own,2004}} + \mathbf{X}_i'\delta + \varepsilon_i \qquad (3)$$

This full specification serves a diagnostic purpose. The spouse's contemporaneous change $\Delta\text{RA4}^{\text{spouse}}$ is mechanically negatively correlated with the spouse's own baseline ($r \approx -0.74$

in our data, a consequence of the bounded 0/1 outcome). Omitting $\Delta\text{RA4}^{\text{spouse}}$ from (1) therefore induces a predictable downward bias on β through the standard omitted-variable formula. When we estimate (3), we expect β to increase, and we can assess whether the apparent weakness of the spouse-baseline coefficient in (1) is an artifact of this bias.

We retain (1) as our main specification because it preserves temporal precedence: the spouse's 2004 level is determined before the outcome period. The full specification (3) captures contemporaneous co-movement (ϕ), which conflates preference transmission with common household shocks and therefore has a less clean causal interpretation. We report both and show that all convergence-related coefficients are in the predicted direction across specifications.

4 Results

4.1 Cross-sectional correlations

Before turning to dynamics, we document the cross-sectional pattern. Table 1 reports descriptive statistics. Wives are substantially more likely than husbands to be classified as risk-averse by our measure (94.1% vs. 84.0% in 2004), and both shares rise by 2007 (95.8% and 90.2%), reflecting a population-wide drift toward greater risk aversion over this period. In terms of socio-economic characteristics, wives are on average 3 years younger than their husbands (42.33 years in 2004 compared with 45.53). They are significantly less educated, as more than 76% of them have at most a high-school education, compared with 65.5% of married men. Finally, wives are less likely to be in work, with fewer than half of them in employment, compared with 88.2% of men.

The within-couple correlation in RA4 is 0.208 in 2004 and 0.206 in 2007 among the full sample of couples (Table 2), indicating modest but stable positive assortative similarity. This correlation is positive for every subgroup we examine, ranging from 0.117 (wife $>$ 40, in 2007) to 0.370 (wife earns more, in 2007). Crucially, partial correlations that control for

both spouses' age, years of schooling, and children are nearly identical to the raw correlations (0.195 vs. 0.208 for the full sample in 2004). This implies that the within-couple match is on risk preferences themselves, not a byproduct of demographic or socio-economic sorting.

Two subgroup patterns are worth noting. First, couples in which the wife out-earns her husband display higher assortativeness in both waves ($r = 0.294$ and 0.370), significantly so in 2007, suggesting that couples with more balanced economic power are better matched on risk preferences. The point estimate rises between waves, but with 226 couples the change is not statistically distinguishable from zero. Second, renters show substantially stronger assortativeness than homeowners (0.300 vs. 0.159 in 2004); the partial correlations controlling for age, schooling, and children are nearly identical (0.293 vs. 0.140), ruling out a pure age or demographic composition effect. This gap may reflect tighter economic constraints that make preference alignment more salient.

4.2 Convergence in risk attitudes

We begin with couple-level transitions, which describe convergence without imposing a regression structure. Table 3 classifies each couple into one of four types based on each spouse's RA4 in a given wave: *Both RA* (both spouses risk-averse), *W:RA, H:NR* (wife risk-averse, husband not), *W:NR, H:RA* (husband risk-averse, wife not), and *Both NR* (neither risk-averse), and reports where each 2004 type is found in 2007. Among the 382 couples who are initially discordant (one spouse risk-averse, the other not), 81.7% become concordant by 2007. The dominant transition is toward joint risk aversion: 77.6% of W:RA, H:NR couples become Both RA, while 87.8% of W:NR, H:RA couples also move to Both RA.

We note that part of this pattern reflects the population-wide drift toward greater risk aversion between 2004 and 2007 rather than couple-specific adjustment. The regression-based convergence estimates to which we now turn net out the common time trend through the intercept and are not affected by this aggregate drift.

Table 4 reports estimates of equation (1). In the wife equation (columns W1–W3), the

spouse's baseline RA4 enters positively: +0.025 ($p < 0.10$) without controls (W1), declining to +0.016 (not significant) with demographic controls (W2) and +0.017 with full controls (W3). The own-baseline coefficient is large and negative throughout (−0.983 to −0.992, all $p < 0.01$), confirming strong mechanical mean reversion in this bounded outcome.

In the husband equation (columns H1–H3), the spouse's baseline coefficient is small and negative (−0.015 to −0.029, none significant). Taken at face value, this would suggest that convergence operates in only one direction: husbands' risk attitudes predict wives' subsequent change, but not vice versa. However, as we show next, this apparent asymmetry is misleading.

Columns Wfull and Hfull of Table 4 report estimates of equation (3), which adds the spouse's contemporaneous change $\Delta\text{RA4}^{\text{spouse}}$ to the right-hand side. The results shift dramatically. The spouse-baseline coefficient rises from +0.016 to +0.133 ($p < 0.01$) in the wife equation and from −0.028 to +0.249 ($p < 0.01$) in the husband equation. The coefficient on $\Delta\text{RA4}^{\text{spouse}}$ is itself positive and highly significant in both equations (+0.129 and +0.279, both $p < 0.01$).

All six convergence-related coefficients in the full specification are in the predicted direction: own baselines enter negatively (mean reversion), spouse baselines enter positively (convergence toward the spouse's 2004 position), and spouse changes enter positively (contemporaneous co-movement). The strengthening of the spouse-baseline coefficient is consistent with the omitted-variable bias discussed in Section 3.3: the spouse's ΔRA4 is mechanically negatively correlated with the spouse's baseline ($r \approx -0.74$), so omitting it biases β downward. This bias is especially severe in the husband equation because husbands have lower baseline RA4 (84% vs. 94%), leaving more room for mean reversion and a stronger mechanical correlation.

We retain the baseline specification as our primary estimate because it preserves temporal precedence. The full specification serves as a diagnostic that reveals bilateral convergence operating in both directions, once the omitted-variable bias is accounted for.

4.3 The role of bargaining power

Table 5 reports estimates of equation (2), in which the spouse's baseline RA4 is interacted with each of the two bargaining-power proxies.

In Panel A (wife equation), the interaction coefficient $\hat{\beta}_2$ is negative and statistically significant for both bargaining-power proxies: wife working (-0.049, $p < 0.10$) and wife earns more (-0.055, $p < 0.01$). The interpretation is that when the wife has stronger economic bargaining power, the husband-to-wife transmission channel weakens. For example, among couples where the wife does not work, the spouse-baseline coefficient is $+0.038$ ($p < 0.10$); among couples where she works, it falls to $+0.038 - 0.049 = -0.011$ (not significant). The channel essentially closes. The wife-earns-more interaction is the strongest result in the panel, although the identifying cell is small ($N = 226$ couples in which the wife out-earns her husband, of which only 26 have a husband who is not risk-averse at baseline), so we interpret its magnitude with caution.

In Panel B (husband equation), the interaction coefficient is positive for wife working ($+0.045$), directionally consistent with the wife-to-husband channel opening when the wife's bargaining position improves, but imprecisely estimated. For wife earns more, the Panel B interaction is essentially zero (-0.017): for this proxy we detect a weakening of the husband-to-wife channel without a corresponding strengthening of the reverse channel, with the small cell again limiting power. None of the Panel B coefficients reaches conventional significance levels, reflecting the lower precision of the husband equation documented in Section 4.2.

The opposite-sign pattern across panels is the paper's central finding. For wife working, the same variable that weakens the husband-to-wife channel (Panel A, $\hat{\beta}_2 < 0$) is associated with an opening of the wife-to-husband channel (Panel B, $\hat{\beta}_2 > 0$). This pattern is difficult to generate through common household shocks, which would push both interaction coefficients in the same direction. It is also not an artifact of differential mean reversion: the own-baseline interaction $\hat{\gamma}_2 \cdot (\text{RA4}^{\text{own}} \times M)$ is small and insignificant in every specification, confirming that the moderating effect operates through the *spouse's* channel, not through group differences

in own mean reversion. A formal cross-equation test, obtained by stacking the two equations and clustering standard errors at the couple level, supports this reading for the employment proxy: the difference between the husband-equation and wife-equation interaction coefficients is +0.094 (SE = 0.061), significant in the predicted direction at the 10% level (one-sided $p = 0.06$).

The opposite-sign pattern also helps distinguish our findings from two alternative explanations. First, if the results were driven by selection (mismatched couples divorcing, leaving a similar-looking pool), we would expect both interaction coefficients to move in the same direction, since differential exit reduces heterogeneity symmetrically. Instead, for the employment proxy, the same moderator weakens one channel while strengthening the other, a pattern that selection cannot generate. Second, common household shocks (e.g., a job loss or health event affecting both spouses) would also push both coefficients in the same direction.[5] The opposite-sign pattern is more naturally consistent with a reallocation of influence within the couple, where the spouse with greater economic leverage exerts a stronger pull on the partner's risk attitudes.

As anticipated in Section 3.2, independent exposure to the labor market could by itself explain the Panel A pattern; it is the opening of the reverse channel for the employment proxy that points toward bargaining specifically, and we view the two accounts as complementary.

We emphasize that our moderators are predetermined but not exogenous, and we frame the results as associations rather than causal effects. The consistency of the negative Panel A interaction across both economic proxies, however, makes it less likely that a single omitted variable drives the finding: such a confounder would need to correlate with both employment and relative earnings in precisely the right way.

Before turning to specification robustness, we examine whether specific life events accel-

[5]KLIPS records employment status and labor income for both spouses in both waves. Adding indicators for job loss between 2004 and 2007 (5.6% of husbands, 9.5% of wives) to the interaction specification leaves the wife-working interaction unchanged (−0.049 in both cases). The husband's job loss does not predict the wife's change in risk aversion, while the wife's own job loss is associated with a small increase in her own risk aversion (+0.022, $p < 0.05$). Health events are not available in our extract, though self-reported health is collected in the KLIPS person files.

erate convergence.

4.4 Life events

We examine three life events that may trigger joint financial decisions and thereby accelerate convergence: homeownership transitions, debt acquisition, and childbearing.

For homeownership, we estimate whether becoming a first-time buyer (renter in 2004, owner by 2007) predicts a larger change in risk aversion, using the subsample of 2004 renters. Table 6 reports the results. The new-buyer coefficient is small and insignificant for both wives (+0.017) and husbands (−0.025). Similarly, among couples with no debt in 2004, acquiring debt by 2007 does not significantly predict changes in either spouse's risk aversion.

For children, we examine three definitions: an indicator for having any children in 2004 (73.4% of couples), the number of children, and an indicator for having a newborn during the panel window (N = 76). None of these variables significantly predicts the direction or magnitude of spousal convergence.

These null findings are informative: they suggest that convergence in risk attitudes is a broad, ongoing process within marriage rather than a response triggered by specific financial or demographic events.

5 Robustness Checks

This section examines the sensitivity of our main results to alternative risk-aversion measures, specification choices, subgroup heterogeneity, potential confounds from age-related trends, ceiling effects, and continuous distance.

All robustness results refer to Tables 7 and 8.

5.1 Alternative risk-aversion measures

Our main analysis uses RA4, a single binary item. A natural concern is that results may be sensitive to this particular measure. We replicate the baseline specification (equation 1, with demographic controls) using three alternatives: RAhigh (an indicator equal to one if the respondent chose cash on at least four of the five items, i.e. RAcount ≥ 4), the r5 item alone (a high-variance lottery with EV equal to cash), and RAcount (the sum of all five cash choices, ranging from 0 to 5).

The wife-equation spouse coefficient is positive across all four measures and the husband-equation coefficient is consistently near zero or negative in the baseline specification. When the full specification (equation 3) is used, all coefficients become strongly positive. Notably, RAcount yields *larger* coefficients than RA4 in the full specification (spouse baseline: +0.146 vs. +0.133 in the wife equation; +0.328 vs. +0.249 in the husband equation, all $p < 0.01$). This is consistent with the composite measure reducing attenuation bias from measurement error, while our choice of the more conservative single-item measure (RA4) understates, if anything, the true convergence effect.

5.2 Ceiling effects

A potential concern is that 94.1% of wives are already classified as risk-averse in 2004, leaving little room for upward movement. Two observations mitigate this concern. First, the own-baseline coefficient in all specifications is very large (-0.99), confirming that the regression fully absorbs the mechanical mean reversion inherent in a bounded outcome. Second, RAcount (0–5 scale) has a less severe ceiling than the binary RA4 (78% of husbands at maximum vs. 84%), yet produces stronger results. If the ceiling were driving our findings, we would expect weaker results with a less compressed measure; we observe the opposite.

5.3 Subgroup stability

We re-estimate the baseline wife equation across several subsamples to check whether convergence is concentrated in a particular group: dual-earner couples, male-sole-earner couples, younger couples (wife ≤ 40), older couples (wife > 40), high-income households, and low-income households. The spouse-baseline coefficient is positive in every subgroup except dual-earner couples, where it is essentially zero (-0.010). This exception is itself consistent with the bargaining results of Section 4.3, which imply that the husband-to-wife channel closes when the wife works. Among the remaining subgroups, the coefficient ranges from $+0.009$ (older couples) to $+0.055$ (wife ≤ 30), with the male-sole-earner estimate ($+0.039$) significant at the 10% level. No subgroup produces a significant negative coefficient.

As an additional check motivated by the concern that a three-year window may be too short to detect preference change, we examine younger couples separately: wife ≤ 35 (N = 682, approximately corresponding to marriages of less than ten years) and wife ≤ 30 (N = 302, marriages of less than five years).[6] In the baseline specification, coefficients are positive but imprecise; in the full specification, they are strongly significant ($+0.153^{***}$ for wife ≤ 35; $+0.146^{**}$ for wife ≤ 30), virtually identical to the full-sample estimate of $+0.133^{***}$. Convergence is not restricted to long-established couples.

5.4 Age-related confounds

A natural concern is that our convergence estimates may reflect a joint-aging confound: if risk aversion increases with age and both spouses age together, then one spouse's baseline may spuriously predict the other's change simply because both are getting older. We address this with four progressively aggressive age controls:

1. Linear age of both spouses (our baseline).

2. Quadratic age (adding age^2 for both spouses).

[6] KLIPS does not record the year of first marriage. We approximate marriage duration using the wife's age, given that the average first-marriage age for Korean women was approximately 26–27 in 2004.

3. Couple-level age-bin fixed effects, constructed from the full interaction of five-year age bins for both spouses (47 cells).

4. Residualized spouse RA4: we first regress the spouse's baseline RA4 on both spouses' ages, obtain the residual, and use this age-purged measure as the regressor.

The spouse-baseline coefficient is virtually unchanged across all four specifications, ranging from +0.014 to +0.016 in the wife equation. This stability indicates that the convergence signal is not an artifact of correlated aging.

5.5 Continuous within-couple distance

As a complement to the binary analysis, we construct a continuous measure of within-couple risk-attitude distance using RAcount: $d_i = |\text{RAcount}_{w,i} - \text{RAcount}_{m,i}|$, ranging from 0 to 5. We then regress the change in distance ($\Delta d_i = d_{i,2007} - d_{i,2004}$) on baseline distance:

$$\Delta d_i = \alpha + \lambda \cdot d_{i,2004} + \mathbf{X}_i'\delta + \varepsilon_i$$

The estimated $\hat{\lambda}$ is -0.943 ($p < 0.001$), indicating strong reversion of within-couple distance toward its mean. Because a coefficient of this magnitude is also consistent with classical measurement error in the underlying items, we place more weight on the level evidence: the average within-couple distance falls from 0.653 in 2004 to 0.426 in 2007, a reduction of 35%, and at the couple level convergence outnumbers divergence (19.9% of couples converge, 11.8% diverge, and 68.3% show no change). This provides a continuous-measure counterpart to the binary transition analysis in Section 4.2 and confirms that convergence is not an artifact of the discrete measurement scale.

6 Conclusion

The literature on spousal risk attitudes has established that partners' preferences are correlated and become more similar over time, but it has treated convergence as a symmetric, couple-level phenomenon. This paper asks a sharper question: *who converges toward whom*, and what determines the direction? Using a balanced panel of 2,331 Korean couples observed in 2004 and 2007, we estimate the husband-to-wife and wife-to-husband transmission channels separately and show that their relative strength is systematically related to the wife's economic position. When the wife works, her risk attitudes stop responding to her husband's baseline while his begin to respond to hers; when she out-earns him, her responsiveness to his baseline falls sharply. A formal cross-equation test supports the opposite-sign pattern for the employment proxy, and the pattern survives controls for differential mean reversion, making it difficult to attribute to common household shocks, demographic heterogeneity, or selective divorce.

Alongside this directional evidence, we document substantial convergence overall: 81.7% of initially discordant couples become concordant within three years (a figure that partly reflects a population-wide drift toward risk aversion over this period), the continuous within-couple distance falls by 35%, and a full specification that accounts for the mechanical correlation between a spouse's baseline and her contemporaneous change reveals positive and significant convergence coefficients in both equations. Convergence in this sample is bilateral; what the wife's economic position governs is the *balance* between the two directions.

These findings extend the collective household model in a specific way. Collective models treat preferences as primitives and ask how bargaining weights allocate resources given those preferences. Our results suggest that the same variables that the literature uses to proxy bargaining weights also predict how preferences themselves evolve within the household. If preference formation responds to the intra-household distribution of economic power, then treating spousal preferences as fixed primitives may understate the long-run effects of policies that change relative earnings or female employment, and empirical work that measures only

one spouse's risk attitude may mismeasure the household's effective risk preference in ways that vary with the wife's economic position.

Several limitations qualify these conclusions. First, our measure is a single binary lottery item observed in two waves three years apart, and its individual-level persistence is low; classical measurement error of this kind attenuates the convergence coefficients and works against our findings rather than for them, and the stability of the cross-sectional within-couple correlation across waves (0.208 and 0.206), together with the stronger results obtained from the five-item composite, indicates that the couple-level signal is not noise. Second, our bargaining proxies are predetermined but not exogenous, and we accordingly interpret the estimates as systematic associations rather than causal effects. Third, the setting is Korea in the mid-2000s, a labor market with pronounced gender asymmetries; whether the direction of preference adjustment responds to the wife's economic position in more gender-equal settings is an open question. Longer panels, incentivized preference elicitation, and extensions to other preference domains such as time preferences are natural next steps.

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Tables

Table 1: Descriptive Statistics

	2004		**2007**	
	Mean	SD	Mean	SD
Wife				
RA4 (r4 cash choice)	0.941	0.236	0.958	0.200
Age	42.33	9.81	45.33	9.81
High-school or less	0.764	0.424	0.764	0.424
Working	0.461	0.499	0.461	0.499
Husband				
RA4 (r4 cash choice)	0.840	0.366	0.902	0.297
Age	45.53	10.16	48.53	10.16
High-school or less	0.655	0.475	0.655	0.475
Working	0.882	0.322	0.882	0.322

Notes: RA4 = 1 if respondent chose cash over a 60% chance of KRW 200,000 (EV = 120,000 > 100,000). High-school or less = KLIPS education level ≤ 5. Schooling and employment status are measured at baseline (2004) and repeated across columns for reference. *** $p < 0.01$.

Table 2: Within-Couple RA4 Correlations by Subgroup

		Raw		**Partial**	
Subgroup	N	2004	2007	2004	2007
Full sample	2,331	+0.208***	+0.206***	+0.195***	+0.191***
Dual earner	972	+0.188***	+0.203***	+0.179***	+0.191***
Male sole earner	1,257	+0.207***	+0.213***	+0.194***	+0.198***
Wife earns more	226	+0.294***	+0.370***	+0.279***	+0.345***
Wife ≤ 40	1,046	+0.206***	+0.223***	+0.208***	+0.223***
Wife > 40	1,285	+0.188***	+0.117***	+0.181***	+0.106***
Both high school or less	1,447	+0.165***	+0.222***	+0.148***	+0.195***
Neither high school or less	469	+0.251***	+0.213***	+0.242***	+0.208***
Homeowner 2004	1,501	+0.159***	+0.146***	+0.140***	+0.130***
Renter 2004	830	+0.300***	+0.271***	+0.293***	+0.263***
Has children	1,711	+0.186***	+0.205***	+0.178***	+0.198***
No children	620	+0.289***	+0.161***	+0.282***	+0.139***

Notes: Partial correlations control for age (both), schooling (both), and has_children. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 3: Couple-Type Transition Matrix (2004 $\rightarrow$ 2007)

	2007				
2004	Both RA	W:RA, H:NR	W:NR, H:RA	Both NR	N
Both RA	89.7% (1,690)	6.7% (127)	2.1% (40)	1.5% (28)	1,885
W:RA, H:NR	77.6% (239)	15.9% (49)	3.9% (12)	2.6% (8)	308
W:NR, H:RA	87.8% (65)	5.4% (4)	6.8% (5)	0.0% (0)	74
Both NR	78.1% (50)	15.6% (10)	3.1% (2)	3.1% (2)	64

Notes: Row percentages. N in parentheses. Among 382 initially discordant couples, 81.7% become concordant by 2007.

Table 4: First-Difference Main Regressions

	Wife eq (ΔRA4$_w$)				**Husb eq (ΔRA4$_m$)**			
	(W1)	(W2)	(W3)	(Wfull)	(H1)	(H2)	(H3)	(Hfull)
Spouse baseline RA4	+0.025*	+0.016	+0.017	+0.133***	−0.015	−0.028	−0.029	+0.249***
	(0.014)	(0.014)	(0.014)	(0.027)	(0.028)	(0.028)	(0.028)	(0.056)
Spouse ΔRA4	—	—	—	+0.129***	—	—	—	+0.279***
				(0.025)				(0.050)
Own baseline RA4	−0.983***	−0.991***	−0.992***	−0.987***	−0.894***	−0.908***	−0.907***	−0.912***
	(0.022)	(0.022)	(0.022)	(0.022)	(0.021)	(0.021)	(0.021)	(0.021)
Demographics	N	Y	Y	Y	N	Y	Y	Y
Full controls	N	N	Y	N	N	N	Y	N
N	2,331	2,331	2,331	2,331	2,331	2,331	2,331	2,331
R^2	0.571	0.578	0.579	0.593	0.554	0.562	0.563	0.578

Notes: OLS on first-differenced panel. Wfull/Hfull add spouse's ΔRA4. Heteroskedasticity-robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 5: Heterogeneous Convergence by Bargaining Power

	Wife working	Wife earns more
Panel A: Wife equation ($\Delta RA4_w$)		
Spouse RA4	+0.038*	+0.020
	(0.021)	(0.015)
Spouse RA4 × M	−0.049*	−0.055***
	(0.026)	(0.019)
Panel B: Husband equation ($\Delta RA4_m$)		
Spouse RA4	−0.049	−0.025
	(0.038)	(0.030)
Spouse RA4 × M	+0.045	−0.017
	(0.055)	(0.068)
N	2,331	2,331

Notes: M = column-header moderator. All specs include own baseline, own × M, M, and demographics. Heteroskedasticity-robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 6: Life Events and Changes in Risk Aversion

	Wife eq (ΔRA4$_w$)		**Husb eq (ΔRA4$_m$)**		
	Coef	SE	Coef	SE	N
New buyer (rent → own)	+0.017	(0.017)	−0.025	(0.025)	830
Debt acquisition (no → yes)	+0.011	(0.014)	−0.020	(0.020)	1,071
Had newborn	+0.014	(0.019)	+0.034	(0.027)	2,331

Notes: Each row reports the coefficient on the indicated life-event variable from a separate OLS regression of ΔRA4 on the event indicator, spouse's baseline RA4, own baseline RA4, and demographic controls (age and years of schooling, both spouses). New buyer: subsample restricted to 2004 renters. Debt acquisition: subsample restricted to couples with no debt in 2004. Had newborn: full sample; 76 couples experienced a birth during the panel window. Heteroskedasticity-robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 7: Robustness: Alternative Measures, Continuous Distance, and Age Controls

	Wife eq		**Husb eq**		
	Spouse coef	SE	Spouse coef	SE	N
Panel A: Alternative risk-aversion measures					
RA4 (baseline)	+0.016	(0.014)	−0.028	(0.028)	2,331
RAhigh	+0.009	(0.013)	−0.022	(0.030)	2,331
r5 only	+0.009	(0.011)	+0.009	(0.034)	2,331
RAcount (0–5)	+0.012	(0.014)	−0.039	(0.031)	2,331
Panel B: Continuous within-couple distance					
Δdistance on baseline distance	−0.943***		(0.021)		2,331
Panel C: Age-effect robustness (wife equation only)					
Linear age (baseline)	+0.016	(0.014)			2,331
+ age^2 both spouses	+0.015	(0.014)			2,331
Couple age-bin FE (47 cells)	+0.015	(0.014)			2,331
Residualized spouse RA4	+0.014	(0.014)			2,331

Notes: All Panel A specifications include own baseline RA4 and demographic controls (age, schooling, both spouses). Panel B regresses the change in $|\text{RAcount}_w - \text{RAcount}_m|$ on its 2004 level. Panel C varies the age controls in the wife equation; the spouse coefficient is virtually unchanged. Heteroskedasticity-robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 8: Robustness: Subgroup Stability (Wife Equation)

Subsample	Spouse coef	SE	N
Full sample	+0.016	(0.014)	2,331
Dual earner	−0.010	(0.017)	972
Male sole earner	+0.039*	(0.021)	1,257
Wife ≤ 40	+0.017	(0.022)	1,046
Wife > 40	+0.009	(0.015)	1,285
Wife ≤ 35 (married $\sim$<10yr)	+0.021	(0.029)	682
Wife ≤ 30 (married $\sim$<5yr)	+0.055	(0.049)	302
High HH income	+0.019	(0.022)	1,069
Low HH income	+0.015	(0.017)	1,262

Notes: Each row re-estimates the baseline wife equation (equation 1 with demographic controls) on the indicated subsample. Heteroskedasticity-robust standard errors in parentheses. Marriage duration is approximated using wife's age (average first-marriage age for Korean women $\approx$ 26–27 in 2004). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.