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Socioeconomic and Housing Conditions, Perceived Landslide Hazard, and Stated Adaptive Intentions in Two Hillside Sectors of Quito, Ecuador

Bethsabe Galarza Poveda, Elinor Mercedes Flores Ramirez, Karen Dayana Recalde Cortez and Belizario Amador Zárate Torres *

Departamento de Ingeniería Civil, Arquitectura y Geociencias, Universidad Técnica Particular de Loja, Loja 110101, Ecuador; bgalarza@utpl.edu.ec (B.G.P.); emflores25@utpl.edu.ec (E.M.F.R.); kdrecalde1@utpl.edu.ec (K.D.R.C.)
* Correspondence: bazarate@utpl.edu.ec

Abstract

Landslides affect low-income households on urban hillsides, yet household-level evidence from Andean cities is scarce. This exploratory cross-sectional study examined how socioeconomic and housing indicators and perceived soil hazard relate to two stated adaptive intentions, willingness to relocate and to modify the dwelling structurally, in two hillside sectors of Quito, Ecuador, with reported landslide events (Pomasqui and Guápulo). Interviewers surveyed a convenience sample of 408 adults, with interviewer teams nested within sectors. Analyses used non-parametric tests, false-discovery-rate correction, and logistic regression. The sectors differed in length of residence, tenure, and housing; 68.0% of Guápulo households, versus 32.6% in Pomasqui, reported no better housing alternative, whereas most risk-perception items differed by less than $|r| = 0.20$, excluding moderate, not small, differences. Seventy percent were willing to relocate; in an exploratory model with post-hoc pooled hazard categories, willingness was 18 percentage points higher (95% CI 5–31) at low than at medium perceived hazard, but this rested on 54 households (44 in Pomasqui) and a poorly calibrated model. Associations with willingness to modify were borderline and fragile to clustering. Explanatory power was low (optimism-corrected area under the curve ≈ 0.60), and sector was confounded with length of residence. The findings are hypothesis-generating.

Keywords: landslide risk; perceived hazard; housing tenure; relocation intention; structural mitigation; Andean cities

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1. Introduction

Landslides are among the most damaging natural hazards for human settlements, particularly in mountainous regions undergoing rapid urban growth. Between 2004 and 2016, non-seismic landslides caused more than 55,000 fatalities worldwide, concentrated in low- and middle-income countries where informal urbanization on steep terrain is common [1]. Built-up areas on large inhabited landslides have grown faster than urban expansion overall during the past 3 decades, exposing an estimated half a million people to direct landslide hazard [2]. Climate-driven increases in rainfall intensity and unplanned urban sprawl onto steepening hillslopes have been identified as joint drivers of this trend [3]. Understanding why households settle and remain on such slopes therefore requires attention

not only to terrain characteristics but also to the socioeconomic conditions that shape residential decisions.

This pattern is particularly acute in Latin America, where informal settlement has been the main route through which low-income households access urban land, as documented for Bogotá, Lima, and Santiago [4]. Reviews of informal settlements in the Global South identify inadequate and insecure housing as a structural driver of hazard and climate vulnerability [5,6], and multi-source assessments show how physical housing conditions in such settlements can be characterized at scale [7]. In this article, housing conditions refer to the dwelling characteristics captured by the questionnaire: tenure, dwelling type, builder, basic services, and extensions (Section 2.3). Standard housing-deficit indicators were not measured.

Evidence from Latin America links social disadvantages to hazard exposure. In Mexico City, neighborhoods with higher social lag are concentrated in zones of elevated landslide susceptibility [8]. In Medellín, a multitemporal assessment documented how the landslide exposure and vulnerability of informal settlements changed over a 24-year period [9]. Low-cost, community-based early-warning systems have been proposed there as a complement where formal relocation is not feasible [10].

Ecuador's Metropolitan District of Quito (DMQ) exemplifies these dynamics. Its ravine-dissected topography has repeatedly pushed formal and informal development toward steep, landslide-susceptible slopes [11]. Municipal landslide risk-reduction policy has evolved since 2011, yet a large share of high-susceptibility areas lies outside officially designated risk zones [12]. Satellite-based analyses indicate that recent urban expansion occurred mainly on flatter terrain, whereas future scenarios point to growing encroachment on high-susceptibility zones [13]. Pomasqui, in the north of the DMQ, and Guápulo, on a ravine slope of the historic eastern sector, were selected as two hillside sectors with reported landslide events (Section 2.2).

Social vulnerability describes the differential capacity of households to anticipate, cope with, resist, and recover from a hazard [14]. It is produced by underlying conditions such as income, employment, tenure security, and access to housing alternatives, and not by physical exposure alone [14,15]. The MOVE framework (Methods for the Improvement of Vulnerability Assessment in Europe) places such conditions among the societal factors that shape susceptibility and the responses available to households [16], and the territorial-vulnerability approach developed by researchers of the Institut de Recherche pour le Développement working in Quito examines how vulnerable elements are distributed within the city [17]. This study does not operationalize either framework as an index. It borrows from them the selection of household-level indicators: income, income earners, education, occupation, subsidies, and social security as socioeconomic conditions, and tenure, dwelling type, builder, basic services, and extensions as housing conditions. Canton-level and subnational indices show that area averages obscure such household-level conditions [18,19].

Risk perception links these conditions to household responses. Protection Motivation Theory (PMT) holds that protective intentions depend on threat appraisal, the perceived probability, and severity of harm, and on coping appraisal, the perceived efficacy of a response, self-efficacy, and response costs [20]; it has been applied to precautionary behavior against natural hazards [21]. Reviews show that risk perception is at best weakly related to protective behavior, and that coping appraisal, trust and resources often matter more [22,23]. Where economic and tenure constraints coincide with a shortage of safe and affordable housing, remaining in a hazardous sector may therefore reflect limited alternatives and response costs rather than ignorance of the risk, a possibility discussed in the relocation literature [24]. In Türkiye, willingness to relocate from landslide- and flood-prone dwellings

depended on individual characteristics, self-efficacy, and institutional support, as well as on risk perception [25], and relocation is increasingly discussed as managed retreat, whose acceptability depends on how costs, housing alternatives, and ties to place are handled [26].

In Loja Province, Ecuador, a Protection Motivation Theory survey found that institutional trust and hazard knowledge were the strongest predictors of household protective behavior, and it documented a gap between stated willingness and implemented measures [27]. That study was conducted by the corresponding author of the present article; the present study adapted part of its instrument but used new data, collected in a different city and urban setting, and different outcomes (relocation and structural modification instead of protective behavior). A similar perception–behavior gap has been reported in Bangladesh [28].

Most landslide research in Andean cities relies on geotechnical and remote-sensing approaches at parcel or pixel scale. In Quito, such work has mapped landslide susceptibility across the urban area [12] and related urban expansion to susceptible terrain [13], and in Medellín, it has quantified the exposure of informal settlements [9]. These studies show where susceptibility is highest, but they say little about why households settle, remain, or consider leaving. This article is not a systematic review, and the literature cited was selected narratively; the statement below therefore reflects the authors' reading of the literature and not an exhaustive search. To our knowledge, and within that limit, household-level evidence that jointly captures socioeconomic conditions, housing conditions, and perceived hazard in hillside sectors of the DMQ is scarce.

In this article, adaptive intentions are stated, as well as hypothetical intentions to reduce exposure (willingness to relocate) or physical vulnerability (willingness to carry out recommended structural modifications); they are not implemented behaviors. This study therefore compares the socioeconomic profile, housing conditions, and risk-related indicators of households in Pomasqui and Guápulo, and it examines which indicators are statistically associated with the two intentions.

RQ1. *How do socioeconomic and housing conditions differ between the sectors?*

RQ2. *Do the sectors differ in risk perception, hazard knowledge, and preparedness?*

RQ3. *Which socioeconomic, housing, and risk-related indicators are associated with the two intentions?*

From PMT and the vulnerability frameworks, four directional hypotheses were stated for RQ3:

H1. *Higher perceived soil hazard is associated with greater willingness to relocate and to modify.*

H2. *Higher income, with greater willingness to modify.*

H3. *Ownership, with greater willingness to modify and lower willingness to relocate.*

H4. *Longer residence, with lower willingness to relocate.*

H5. *Economic and tenure constraints would be more strongly associated with both intentions than hazard perception.*

H1–H4 were formulated during the revision, after the data had been analyzed; they are used to organize the comparison of the results with PMT and the vulnerability literature, not as confirmatory tests, and none was pre-registered. The design is cross-sectional; the analysis identifies associations, not causes.

The scope is limited to two purposively selected sectors surveyed with a non-probability convenience sample, in which interviewer teams were nested within sectors. The findings describe the sampled households and cannot be generalized to the DMQ, in the same spirit as the non-probabilistic survey in Loja Province [27]. This study contributes in two ways. First, it provides household-level evidence on socioeconomic and housing indicators, perceived hazard and stated intentions in two hillside sectors of the DMQ, a scale for which the literature we reviewed offers little evidence. Second, it shows how strongly such comparisons depend on design: it quantifies the confounding of sector with length of residence and the sensitivity of each association to clustering, and it reports effect sizes and optimism-corrected model performance. It does not establish causes or measure behavior. Data and scripts are archived (see Data Availability Statement). The results are relevant to Sustainable Development Goal 11, Targets 11.1, and 11.5 [29], as hypotheses for programs that combine physical risk reduction with housing and tenure support.

2. Materials and Methods

2.1. Study Design and General Objective

This study used a quantitative, non-experimental, cross-sectional design with an exploratory, descriptive, and correlational scope. The general objective was to examine how household socioeconomic and housing indicators and perceived soil hazard are associated with two adaptive intentions (willingness to relocate and willingness to undertake structural modification, hereafter willingness to modify) among households in two hillside sectors of the DMQ, Pomasqui, and Guápulo. The descriptive component characterized the socioeconomic profile, housing conditions, and risk perception of households in each sector. The correlational component examined associations among these variables and the two intentions, guided by the social-vulnerability framework [14] and PMT [20].

No variable was manipulated, and all data were reported by respondents at a single point in time. Residential permanence, that is, the decision to stay, was not measured and is not an outcome of this study; reasons for settling and the economic items are reported only as context. The relocation item is hypothetical and conditional on institutional support and access to safe housing. It therefore removes part of the economic constraint and measures an intention, not behavior.

2.2. Study Area and Sample

The study was conducted in Pomasqui, in the north of the DMQ, and in Guápulo, on the ravine slope of the historic eastern sector. The two sectors were purposively selected because both are hillside areas where landslides or slope erosion have been reported in recent years; they are not a sample of DMQ hillside sectors. The units also differ in kind: Pomasqui is a rural parish of the DMQ, with 36,883 inhabitants in the 2022 census [30], whereas Guápulo is a barrio within the urban area for which we found no published census count. No household listing existed for the surveyed areas, so the number of households per sector and the sampling fraction could not be determined. Differences in length of residence and housing between the sectors are results of this survey, not selection criteria. Hazard is described from institutional documents and reports of events. The Pomasqui development and land-use plan 2023–2027 describes a mountainous topography susceptible to erosion and landslides, especially on steep slopes and under heavy rain [31]; a municipal risk diagnosis exists for the Monjas River basin [32], where mitigation works were reported in January 2026 [33], and reported events include a hillside collapse in La Pampa in November 2024 [34] and river erosion that left 15 families at risk in barrio Santa Clara del Común in April 2025 [35]. In Guápulo, a landslide after heavy rain on 6 April, 2025, led to the evacuation of three families [36,37], another slide affected traffic 2 days earlier [38], and

in March 2026, residents reported landslides and new hillside construction in areas they considered risky, claims that were not independently verified [39]. These reports refer to specific barrios and streets, whereas the sample covered each sector more widely, and because no geolocation was collected, we do not know how close respondents lived to the events; for the same reason, the susceptibility map available for urban Quito [12] could not be linked to the dwellings, and Figure 1 shows the location of both sectors without susceptibility classes. Exposure is represented only by self-reported topographic location and by perceived soil hazard, both of which are subjective measures (Section 4.6).

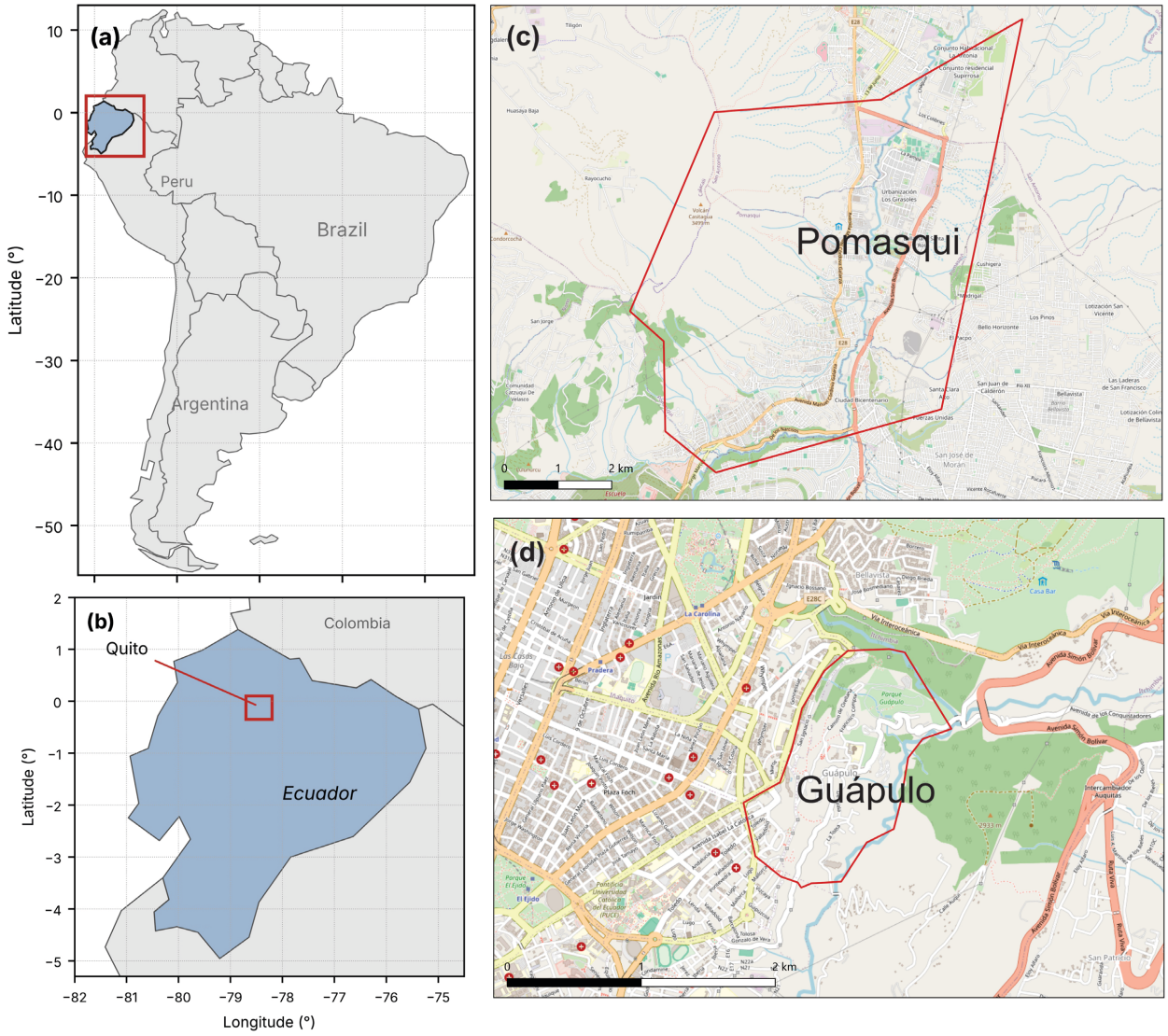

Figure 1. Study area. (**a**) Location of Ecuador (red box) in South America; (**b**) location of the Metropolitan District of Quito (red box) in Ecuador; (**c**) Pomasqui and (**d**) Guápulo, with the limit of each sector (red line). North is up. Panels (**a**,**b**) were prepared in R 4.3.3 and panels (**c**,**d**) in QGIS V 4.2.3. Country outlines: Natural Earth (public domain). Basemap in (**c**,**d**): © OpenStreetMap contributors; data available under the Open Database License (ODbL).

The survey was conducted among residents aged 18 years or older between 26 June, and 23 July 2026. Participants were selected by non-probability convenience sampling. Trained interviewers visited households in the two sectors, and each interviewer was assigned to a defined area within a sector, so that no household was approached twice. This approach suits exploratory comparisons when no sampling frame exists, but it restricts generalization

and can introduce selection bias [40,41]. Because interviewer teams were assigned to sectors, interviewer effects cannot be separated from sector effects [42], and between-sector differences are interpreted with caution (Section 4.6). Interviewer identifiers were not recorded with the responses.

The final analytical sample comprised 408 respondents (236 in Pomasqui, 57.8%; 172 in Guápulo, 42.2%), each answering on behalf of one household. Inclusion criteria were age of 18 years or older, residence in either sector at the time of the survey, and voluntary consent. Exclusion criteria were incomplete core sociodemographic or housing items and duplicate records. Refusals and incomplete questionnaires were not recorded, so only the analyzed numbers are reported, and no response rate is calculated. Residence in the stated sector was established by the interviewers during the household visit, and no geolocation was collected. Education and occupation refer to the household head and were reported by proxy, so some misclassification is possible. The sample composition could not be compared with census data by sector because census tabulations are not published for the barrio of Guápulo, and the census age bands do not match the questionnaire categories.

No a priori sample-size calculation was performed. A post-hoc sensitivity analysis showed that, with $N = 408$ and $\alpha = 0.05$, chi-square tests of independence had 80% power to detect effects of Cohen's $w \geq 0.139$ (1 degree of freedom) to $w \geq 0.183$ (6 degrees of freedom). Power to detect a small effect ($w = 0.10$) ranged from 52.4% (1 df) to 27.5% (6 df). Associations weaker than these thresholds may therefore have gone undetected.

2.3. Data Collection Instrument

Data were collected by face-to-face interview with a structured questionnaire implemented in Google Forms, after interviewers had been trained in its administration. The digital format allowed standardized item presentation, mandatory-response validation for key variables, and direct export to a spreadsheet. The items were adapted from the instrument of a previous Protection Motivation Theory study in Loja Province [27] and were reviewed by two disaster-risk-management experts from the National Risk Management Secretariat for clarity, comprehensiveness, and relevance. No pilot test with residents was conducted, so problems of item wording and response format could not be detected before fieldwork; the instrument requires cognitive pretesting and validation before it is reused. The questionnaire was administered in Spanish; the English wording used in this article was translated by the authors for reporting, and both versions are provided as Supplementary File S1.

The questionnaire comprised 47 variables organized in the following blocks:

- Sociodemographic characteristics: sector of residence, length of residence, sex, age, educational level, and main occupation of the head of household, main reason for residing in the sector, approximate monthly household income, household size, and number of household members generating income.
- Housing conditions and tenure: government subsidies, social-security affiliation, type of housing, housing tenure, builder of the dwelling, availability of basic services, extensions built with or without permits, and self-reported topographic location of the dwelling (flat ground, gentle slope, moderate slope, steep slope or ravine edge). Topographic location was used as an ordinal proxy of exposure, not as a measured hazard class.
- Hazard knowledge and experience: perceived soil hazard (five levels, from very low to very high), risk training received, prior awareness of landslide risk before moving in, and having witnessed ground instability (three yes/no items).
- Economic-vulnerability items: seven statements rated on a five-point agreement scale (1 = strongly disagree to 5 = strongly agree), addressing economic constraints, land or

housing cost, absence of better alternatives, proximity to employment, and the difficulty of relocating. Item E3 refers to the past decision to settle, and item E7 (continuing to live in the sector out of economic necessity despite awareness of the risks) is double-barreled and presupposes risk awareness.

- Risk-perception items: seven items scored 1–5, addressing familiarity with the landslide concept, awareness of historical instability, perceived probability of damage, concern in the rainy season, adequacy of institutional information, adequacy of mitigation works, and family preparedness. Four of them (familiarity, awareness, probability, and concern) are worded as questions but were answered on the same agreement scale as the others (1 = strongly disagree to 5 = strongly agree). Because this response format does not match the question stem, the validity of these four items is doubtful; their scores are reported only as ordinal indicators, and no conclusion rests on them alone (Section 4.6).
- Signs of instability, mitigation, and preparedness: three multiple-response items on signs of ground instability observed in the preceding 12 months, physical mitigation measures implemented, and non-physical preparedness measures adopted.
- Spending-priority ranking: a forced ranking (1 = most important to 4 = least important, without ties) of aesthetics and finishes, structural safety, space expansion, and equipment.
- Adaptive intentions and behavior: insurance against natural hazards; willingness to carry out the structural modifications recommended by a geotechnical report (the item does not state who would pay or the cost, and it is less applicable to households that cannot alter the dwelling); the maximum share of annual household income the household would invest in stabilization works (four categories); and willingness to relocate if institutional support and access to safe housing were available [25].

Perceived soil hazard is conceptually related to, but distinct from, the risk-perception block: it is a single site-specific judgment, whereas the block captures attitudes toward information, mitigation, and preparedness. The two are reported separately.

Socioeconomic conditions were described through household income, income earners, the head's education and occupation, subsidies, social security and the seven economic items, and housing conditions through tenure, housing type, builder, basic services, and extensions without permits; all of these were compared between sectors (RQ1 and RQ2). For the primary models (RQ3), only income and tenure were retained as socioeconomic and housing indicators because they correspond directly to the response-cost and tenure hypotheses (H2 and H3) and have adequate cell sizes; the remaining indicators enter the block comparison (Section 2.5). Building materials and the number of rooms were not collected, so housing-deficit indicators such as overcrowding could not be derived. All indicators were analyzed individually because no composite index showed acceptable reliability (Section 3.3).

2.4. Data Cleaning and Processing

The Google Forms export was imported into R as character data and converted explicitly: nominal variables to unordered factors, hierarchical variables (for example, length of residence, income, and perceived soil hazard) to ordered factors, Likert items to integers from 1 to 5, and ranking items to integers from 1 to 4.

A data-quality protocol was applied to the full dataset: (i) detection of exact duplicated records (identical answers to all items); (ii) verification that the four spending-priority ranks summed to 10 per respondent; (iii) verification that Likert responses lay within 1–5; (iv) a logical check that the minimum plausible number of income earners did not exceed the maximum plausible household size; and (v) detection of straight-lining, defined as

identical responses to all 14 Likert items and, more strictly, as the same response to at least 12 of the 14 items, identical responses to all seven items of a block, or a within-person standard deviation below 0.5. The Google Forms export used for analysis contained no submission timestamps or interviewer identifiers, so response times and interviewer-level response patterns could not be examined. The R scripts are deposited with the data in the same repository.

The text category "None" for the number of income earners and for basic services was retained as a valid response because it denotes a household with no earner and a dwelling with no basic services, which are informative categories. Missing values arising from skip logic or non-response (one case for one Likert item and one for the physical-mitigation item) were excluded listwise for the corresponding analysis; no imputation was applied. The analytic N is stated for every model in Section 3.6.

2.5. Statistical Analysis in R

All analyses were performed with the archived R scripts in R (Version 4.3.3) [43], and the reported results were independently re-computed in Python 3.13.16 (statsmodels 0.15.0 and SciPy 1.18.1), with agreement for all checked estimates. Package versions are listed in Supplementary Table S1, generated by the archived script. The packages were readxl 1.4.3, dplyr 1.1.4, tidyr 1.3.1, stringr 1.5.1, forcats 1.0.0, and purr 1.0.2 for data handling; psych 2.4.1 [44] for reliability, polychoric correlations, the Kaiser–Meyer–Olkin (KMO) measure, Bartlett's test, and exploratory factor analysis; FactoMineR 2.9 [45] for multiple correspondence analysis (MCA); effectsize 0.8.6 [46] for rank-biserial correlations; car 3.1.2 for Type II tests and generalized variance-inflation factors [47], and performance 0.10.8 for diagnostics; MASS 7.3.60.0.1 for proportional-odds models; pROC 1.18.5 [48] for receiver operating characteristic (ROC) analysis; brglm2 0.9.2 [49] for penalized regression; pwr 1.3.0 for power analysis; ggplot2 3.4.4, within the tidyverse [50], for graphics; and openxlsx 4.2.5.2 for export.

Descriptive statistics were computed by sector. Internal consistency of the seven-item economic-vulnerability block and the seven-item risk-perception block was assessed with Cronbach's alpha [51], using the conventional threshold of 0.70 for scale acceptability [52]; blocks below the threshold were analyzed item by item. Because the items are ordinal, ordinal alpha and omega total were also computed from polychoric correlations [53]. No item was reverse-coded. Dimensionality was examined with exploratory factor analysis (EFA) of the 14 items using polychoric correlations, the KMO measure, and Bartlett's test. The number of factors was set by parallel analysis, and factors were extracted by maximum likelihood with varimax rotation. Oblimin rotation and a three-factor solution were examined as sensitivity checks, and communalities were inspected for Heywood cases.

Differences between sectors on the 14 Likert items were tested with the Mann–Whitney U test [54], with rank-biserial correlations (negative values indicate lower scores in Pomasqui) and percentile bootstrap 95% confidence intervals (CI; 5000 resamples) as effect sizes. Equivalence was assessed with a two one-sided tests (TOST) logic [55]: a difference was considered smaller than a moderate effect when the 90% bootstrap interval fell within ± 0.20; this margin excludes moderate or larger differences, but not small ones, which corresponds to a probability of superiority between 0.40 and 0.60. No anchor-based smallest effect size of interest was available, so this margin was set post-hoc, and because the 95% intervals had half-widths of about 0.10, the stricter margins of ± 0.15 and ± 0.10 are also reported [56]. The percentage scoring 4–5, 3, and 1–2 was reported for each item. The sector effect on each item was also estimated with proportional-odds models adjusted for length of residence, with the assumption checked by the Brant test [57], and with Mann–Whitney tests within strata of length of residence (5 years or less, and more than 5 years).

Associations between sector and 23 categorical variables were tested with Pearson chi-square tests, with Cramér's V, and its bias-corrected version for sparse tables [58]. Insurance against natural hazards was excluded because it was measured inconsistently (Section 3.5). The expected-count rule was applied to every table: when more than 20% of cells had expected counts below 5, or any expected count was below 1, the *p*-value was obtained by Monte Carlo simulation (10,000 replicates); the tables concerned are identified in Section 3.2. Adjusted standardized residuals identified the category deviating most from independence. For ordinal variables, order was also taken into account with a Mann–Whitney test on the ordinal scores and a Mantel–Haenszel linear-by-linear test. Because the 14 Mann–Whitney tests and 23 chi-square tests are 37 bivariate comparisons on the same sample, the Benjamini–Hochberg procedure [59] controlled the false discovery rate (FDR), and both unadjusted and adjusted *p*-values are reported (Table 1).

Spending-priority rankings were analyzed with the Friedman test [60] and are reported in the Supplementary Materials (Table S11 and Figure S5). The maximum share of income households would invest was described by sector and also analyzed as an ordinal outcome with a proportional-odds model that included the five primary covariates. The multiple-response items were described by sector, and each option was compared with chi-square tests with FDR correction.

MCA of 24 categorical variables (sector supplementary; insurance excluded) was used only as a descriptive summary of household profiles [45], with inertia reported as raw and as Benzécri- and Greenacre-adjusted percentages [61]. Because variables strongly associated with sector were active, the separation of sectors was re-examined in a restricted MCA that excluded the three intention variables and the ten variables with a Cramér's V of 0.30 or more (age, occupation, reason for residence, length of residence, household size, tenure, housing type, builder, topographic location, and prior risk awareness). Methods and results are given in Supplementary Table S2 and Figure S6.

Two binary logistic regression models were fitted for willingness to relocate and for willingness to modify. The primary models included sector, household income, tenure, perceived soil hazard, and length of residence. The covariates were chosen from the social-vulnerability framework [14] and from the studies in Türkiye and Loja [25,27], and their number was limited by the sample size. Sector is the design variable. Income represents economic capacity (H2). Tenure is the housing-condition indicator with adequate cell sizes (H3). Perceived soil hazard is a site-specific hazard judgment (H1). Length of residence represents cohort of arrival and ties to place (H4) and is a likely confounder of sector. The covariate set was specified before the models were fitted and was not modified afterwards; it was not pre-registered and was not derived from a directed acyclic graph. Respondent age and household size, which also differed by sector, were added in a sensitivity analysis.

Because this set omits other indicators, four blocks of covariates (sector and socioeconomic profile; housing conditions; hazard perception and knowledge, including the risk-perception items; and economic-vulnerability items) were compared on the 407 complete cases, each evaluated alone, added cumulatively, and removed from the full model with a likelihood-ratio (LR) test. With few events per parameter and blocks of different sizes, this comparison is exploratory; it is reported in Supplementary Table S12 with the LR statistic per degree of freedom.

Table 1. Association of sectors with 23 categorical variables (insurance excluded; Section 3.5). χ^2 (df) is Pearson chi-square; cells < 5 is the percentage of cells with expected counts below 5; † *p*-value by Monte Carlo simulation (10,000 replicates). V is the bias-corrected Cramér's V; *p*/*p* (FDR) gives the unadjusted *p*-value and the dBenjamini–Hochberg-adjusted *p*-value over 37 comparisons. Trend gives the Mantel–Haenszel linear-by-linear *p*-value and the rank-biserial r (negative = lower in Pomasqui) for ordinal variables. The last two columns show the category with the largest adjusted standardized residual (ASR; positive = more frequent in Pomasqui) and its percentage within Pomasqui (n = 236) and Guápulo (n = 172).

Variable	χ^2 (df)	Cells < 5 (%)	V	*p*/*p* (FDR)	Trend *p* [r]	Category with Largest ASR	%Pomasqui vs. %Guápulo (ASR)
Sex	2.766 (1)	0.0	0.072	0.096/0.146	–	Female	51.3 vs. 42.4 (+1.763)
Age	50.574 (4) †	20.0	0.338	<0.001/<0.001	<0.001 [−0.372]	18 to 25 years	28.4 vs. 5.8 (+5.755)
Education level	23.834 (5)	8.3	0.215	<0.001/<0.001	0.739 [−0.019]	Secondary education	34.3 vs. 18.6 (+3.503)
Occupation	56.138 (6)	0.0	0.351	<0.001/<0.001	–	Private-sector employee	13.1 vs. 36.0 (−5.447)
Reason for residence	56.580 (5)	0.0	0.356	<0.001/<0.001	–	Lack of another housing alternative	12.3 vs. 34.3 (−5.339)
Length of residence	167.678 (3)	0.0	0.636	<0.001/<0.001	<0.001 [−0.700]	More than 10 years	5.1 vs. 47.7 (−10.088)
Household income	13.622 (3)	0.0	0.161	0.003/0.007	0.003 [+0.155]	More than USD 1000	28.4 vs. 16.3 (+2.858)
Household size	36.805 (3)	0.0	0.288	<0.001/<0.001	0.012 [−0.163]	1–2 people	28.0 vs. 8.1 (+4.981)
Income earners	22.596 (3)	12.5	0.219	<0.001/<0.001	<0.001 [−0.165]	None (no income earner)	8.9 vs. 0.0 (+4.017)
Tenure	52.505 (3) †	25.0	0.349	<0.001/<0.001	–	Borrowed (informal, no lease)	7.2 vs. 35.5 (−7.169)
Housing type	44.007 (3)	0.0	0.317	<0.001/<0.001	–	Mixed-use dwelling	4.2 vs. 20.9 (−5.264)
Builder	77.768 (3)	0.0	0.429	<0.001/<0.001	–	Foreman/builder (maestro de obra)	22.0 vs. 64.5 (−8.655)
Basic services	15.118 (3)	0.0	0.173	0.002/0.004	0.090 [−0.073]	None (no basic services)	5.1 vs. 0.0 (+3.002)
Extensions/additions	12.885 (2)	0.0	0.163	0.002/0.003	–	No	51.7 vs. 34.9 (+3.373)
Topographic location	72.689 (4)	0.0	0.411	<0.001/<0.001	<0.001 [−0.384]	On a moderate slope	30.5 vs. 60.5 (−6.033)
Perceived soil hazard	28.571 (4)	0.0	0.246	<0.001/<0.001	0.694 [+0.009]	Medium	33.1 vs. 54.7 (−4.363)
Willingness: structural modification	2.732 (1)	0.0	0.071	0.098/0.146	–	Yes	64.4 vs. 55.8 (+1.755)
Willingness to relocate	0.000 (1)	0.0	0.000	0.988/0.988	–	No	29.7 vs. 30.2 (−0.125)
Government subsidy	0.071 (1)	0.0	0.000	0.790/0.860	–	Yes	18.2 vs. 19.8 (−0.394)
Social security	1.927 (1)	0.0	0.055	0.165/0.235	–	No	61.4 vs. 68.6 (−1.493)
Risk training received	5.189 (1)	0.0	0.107	0.023/0.040	–	No	67.4 vs. 55.8 (+2.382)
Prior risk awareness	48.857 (1)	0.0	0.348	<0.001/<0.001	–	No	70.8 vs. 35.5 (+7.091)
Witnessed instability	5.773 (1)	0.0	0.114	0.016/0.030	–	Yes	47.9 vs. 35.5 (+2.504)

Ordered predictors were entered as ordered factors with orthogonal polynomial contrasts in the primary models (Supplementary Tables S3 and S4), and inference rests on omnibus Type II likelihood-ratio tests per predictor, which do not depend on coding. Departure from linearity was tested by comparing linear-score and categorical coding. For interpretation, a simplified-coding model, defined after the observed proportions had been inspected and therefore exploratory, used income as a linear score, three tenure groups, perceived soil hazard collapsed to three levels (medium as reference), and length of residence in four categories (less than 1 year as reference). Households in borrowed dwellings and in legalization ($n = 5$) were merged as the tenure reference, a heterogeneous group of 83, and all models were re-estimated without the households in legalization.

Because events per parameter were limited [62], models were also fitted with Firth's penalized likelihood [49,63]. Model performance was summarized with McFadden's pseudo-R^2, the global likelihood-ratio test, the Hosmer–Lemeshow test (interpreted with caution because of few covariate patterns), the generalized variance inflation factor (GVIF), the apparent area under the ROC curve (AUC) with a DeLong 95% confidence interval, and the optimism-corrected AUC and calibration slope from 1000 bootstrap resamples [64].

Because sector and length of residence were strongly associated, three post-hoc analyses were added: the change in the sector odds ratio (OR) as covariates were introduced; sector × length-of-residence interactions (linear score, two categories and four categories); and the sector effect within strata of length of residence. Marginal predicted probabilities by sector, tenure, perceived soil hazard, and length of residence were obtained by standardization with percentile bootstrap intervals. A sensitivity analysis restricted the models to owners. No neighborhood or interviewer identifier was collected, so cluster-robust errors or random effects could not be used (Section 4.6). Instead, a design-effect (deff) sensitivity analysis was performed [65]: each test statistic was divided by deff, a first-order correction, to find the deff at which the result would lose significance at $\alpha = 0.05$, and results were re-examined for deff = 1.5, 2 and 3, which correspond, for clusters of 20 households, to intra-cluster correlations of about 0.03, 0.05, and 0.11 ($\text{deff} = 1 + (m - 1)\rho$). The significance level was $\alpha = 0.05$; coefficients within models were not adjusted for multiplicity, which is why omnibus tests are emphasized.

During the preparation of this manuscript and its revision, Claude (Anthropic, San Francisco, CA, USA; Sonnet 4.6, Sonnet 5.5 and Opus 5.5 models) was used to assist with language editing, with drafting revised text in response to the reviewers, with writing and running the R code for the additional analyses requested in the revision and the independent re-computation in Python, with the preparation of the Supplementary Tables and Figures from the study data, with the English translation of the questionnaire and with checking references against their sources. The study design and data collection were carried out by the authors. All code, outputs, and text were reviewed and verified by the authors, who take full responsibility for them.

2.6. Ethical Considerations

The study was conducted in accordance with the Declaration of Helsinki and the Ecuadorian Organic Law on Personal Data Protection (Ley Orgánica de Protección de Datos Personales). No name, national identification number, address, or geolocation was collected. Under Article 27(h) of the Internal Regulation of the Human Research Ethics Committee of the Universidad Técnica Particular de Loja (CEISH-UTPL) (code DGMU_RG_1_2023_V1; University Council Resolution No. 870.955.2023), which the Ministry of Public Health endorsed on 20 November, 2023 (code DNIVS-CEISH-11-UTPL-30), the study falls within "Research Without Risk", and, under Article 41, is exempt from full committee review; no study-specific approval code was issued. The questionnaire opened with a statement of

compliance with the Declaration of Helsinki and the Ecuadorian Organic Law on Personal Data Protection, of the anonymity of participants and the exclusive use of aggregated data, and of the study's classification as "Research Without Risk". A mandatory first item stated that participation was voluntary and free of risk, that the survey was anonymous and that no name, identification number or address would be collected, and asked whether the person agreed to participate. Consent was therefore recorded electronically, and only respondents who answered "Yes, I agree to participate" continued; those who declined were taken to a closing page. Because respondents were interviewed at home and reported income and tenure status, the data are better described as de-identified than fully anonymous. Results are reported in aggregate.

3. Results

3.1. Sample Characteristics

The analytical sample comprised 408 households (one respondent per household): 236 in Pomasqui (57.8%) and 172 in Guápulo (42.2%). The full sociodemographic and housing profile by sector is given in Table S5. The data-quality checks described in Section 2.4 found no exact duplicated records, no inconsistent household counts, no out-of-range values, and no invalid rankings. No respondent gave identical answers to all 14 Likert items; 20 respondents met at least one of the stricter straight-lining criteria, and excluding them did not change the results (Table S17). All 408 respondents contributed to the primary logistic models; the 407 with complete Likert data contributed to the factor analysis and the block comparison.

3.2. Socioeconomic and Housing Differences Between Sectors (RQ1)

Of the 23 categorical variables tested, 18 were associated with sector (unadjusted $p < 0.05$), and all 18 remained significant after FDR correction (Table 1). The strongest association was with length of residence (Cramér's V = 0.641; bias-corrected V = 0.636). A total of 183 of 236 households in Pomasqui (77.5%) had lived there for 5 years or less, against 27 of 172 in Guápulo (15.7%), and 47.7% of Guápulo households had lived there for more than 10 years, against 5.1% in Pomasqui (Figure S1). In the design-effect analysis, 15, 12, and 10 of the 18 associations remained significant after FDR correction at deff = 1.5, 2, and 3, including length of residence, tenure, builder, and topographic location at deff = 3 (Table S13).

The direction of the differences is shown in the last two columns of Table 1. Pomasqui households were younger (28.4% aged 18–25 years, versus 5.8%), had higher incomes (28.4% above USD 1000 per month, versus 16.3%), and more often owned their dwelling (57.6% versus 37.2%). Guápulo households more often borrowed their dwelling (35.5% versus 7.2%), lived in mixed-use dwellings (20.9% versus 4.2%), reported a foreman as the builder (64.5% versus 22.0%), and described their dwelling as on a moderate slope (60.5% versus 30.5%). In Guápulo, 34.3% gave lack of another housing alternative as the main reason for residing there, against 12.3% in Pomasqui (Figures 2 and 3).

Hazard knowledge and experience also differed. Prior awareness of landslide risk was lower in Pomasqui (70.8% reported no prior awareness, versus 35.5%), and fewer Pomasqui households had received risk training (32.6% versus 44.2%). More Pomasqui households had witnessed ground instability (47.9% versus 35.5%). The difference in prior awareness was present among both newer residents (67.2% versus 22.2% without prior awareness) and longer-term residents (83.0% versus 37.9%), so it is not explained by length of residence alone. Perceived soil hazard was distributed differently without a monotonic trend (trend $p = 0.694$): Pomasqui had more households at the extremes, with 18.6% perceiving very low

or low hazard (versus 5.8%) and 48.3% high or very high hazard (versus 39.5%), whereas 54.7% of Guápulo households chose the medium level (versus 33.1%).

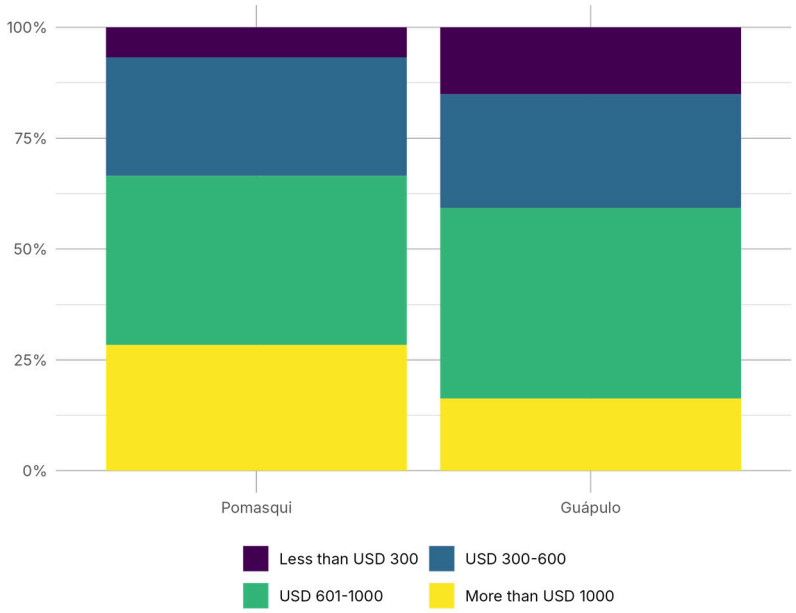

Figure 2. Distribution of approximate monthly household income by sector (percentage of households within each sector; Pomasqui n = 236, Guápulo n = 172; income in USD).

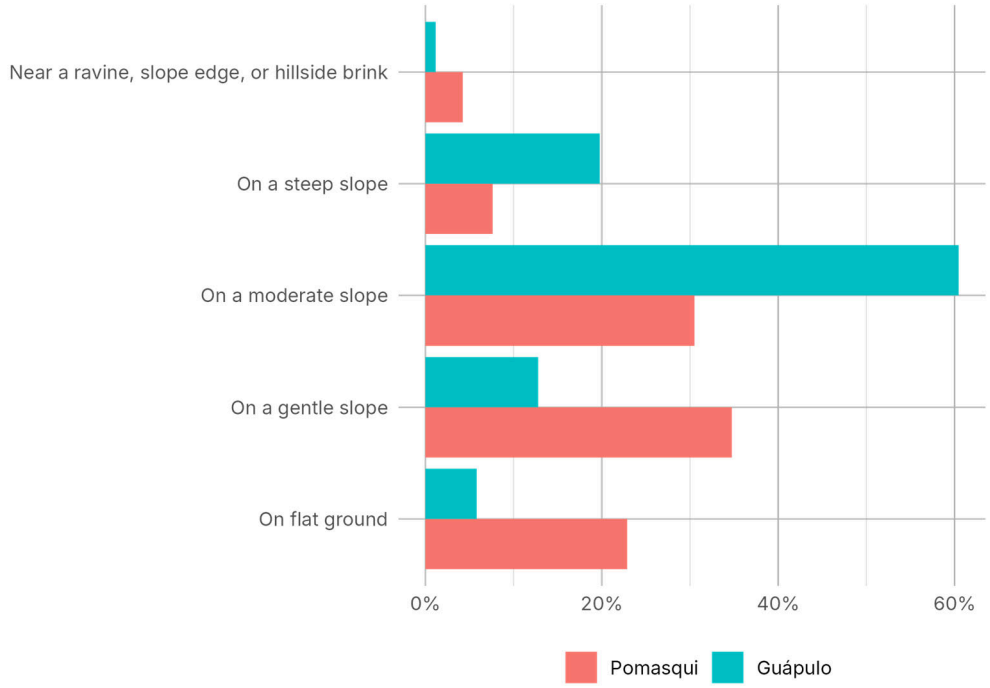

Figure 3. Self-reported topographic location of the dwelling by sector (percentage of households within each sector; Pomasqui n = 236, Guápulo n = 172).

An MCA of 24 categorical variables summarized the same pattern. Dimension 1 (5.05% of raw inertia; 30.0% after Benzécri adjustment) separated the sectors (r = 0.890), mainly through length of residence, topographic location, housing type, and reason for residence. Because these variables were themselves strongly associated with sector, the separation

largely reflects them: in the restricted MCA without the intention variables and without variables with V ≥ 0.30, it became small (r = 0.123, $p = 0.034$; Table S2; Figure S6).

3.3. Reliability and Dimensionality of the Likert Item Blocks

Internal consistency of the two seven-item blocks was assessed with Cronbach's alpha (Table 2). The economic-vulnerability block yielded α = 0.386 and the risk-perception block α = 0.446, both below the threshold of 0.70. Ordinal alpha (0.447 and 0.492) and omega total from polychoric correlations (0.462 and 0.498) were similarly low, and the mean inter-item polychoric correlations were 0.10 and 0.12. The 14 items were therefore analyzed individually, and no composite index was constructed.

Table 2. Internal consistency of economic-vulnerability and risk-perception item blocks: Cronbach's alpha with Feldt 95% CI, and ordinal alpha and omega total computed from polychoric correlations (407 complete cases).

Block	Cronbach's Alpha	95% CI (Feldt)	Ordinal Alpha	Omega Total (Ordinal)
Economic vulnerability	0.386	0.291–0.473	0.447	0.462
Risk perception	0.446	0.356–0.522	0.492	0.498

The EFA of the 14 items showed inadequate sampling adequacy (KMO = 0.581; Bartlett's $\chi^2 = 643.2$, df = 91, $p < 0.001$). Parallel analysis suggested five factors (39.7% of variance), three of them single-item factors, two with Heywood-type communalities of 0.995; the only interpretable factor grouped items R5–R7 (loadings 0.41–0.59), and oblimin and three-factor solutions were not interpretable (Table S6). Item–rest correlations ranged from 0.08 to 0.32. We therefore do not claim that "economic vulnerability" and "risk perception" were measured as latent scales, and the items are treated as individual indicators.

3.4. Comparison of Likert Items Between Sectors (RQ2)

Responses clustered around the midpoint: 10 of the 14 items had a median of 3 in both sectors, so Table 3 reports the share of respondents scoring 4–5, 3, and 1–2 for every item, together with effect sizes and equivalence results. After FDR correction, three items differed between sectors. First, "no better housing alternatives" was endorsed (scores 4–5) by 68.0% of Guápulo households and 32.6% of Pomasqui households (r = −0.486, 95% CI −0.573 to −0.393). Second, "proximity to work" influenced location choice for 31.8% in Pomasqui and 7.6% in Guápulo (r = 0.362, 0.258 to 0.461). Third, "my family knows how to act in the event of a landslide" was endorsed by 64.5% in Guápulo and 33.9% in Pomasqui (r = −0.432, −0.529 to −0.341; Figure 4; all 14 items in Figure S2).

Table 3. Likert items by sector: percentage scoring 4–5, 3, and 1–2 (Pomasqui/Guápulo), rank-biserial effect size r with bootstrap 95% CI (negative = lower scores in Pomasqui), unadjusted and Benjamini–Hochberg-adjusted *p*-values (37 comparisons), and equivalence within |r| < 0.20 (90% bootstrap CI), which corresponds to a probability of superiority between 0.40 and 0.60. Pomasqui n = 236; Guápulo n = 172 (171 for R2). E1–E7 are economic-vulnerability items; R1–R7 are risk-perception items.

Item	%4–5 (P/G)	%3 (P/G)	%1–2 (P/G)	r [95% CI]	*p*/*p* (FDR)	Equivalence
E1 Lack of economic resources influenced decision to live here	31.8/26.2	33.1/38.4	35.2/35.5	0.025 [−0.084, 0.130]	0.648/0.743	Equivalent
E2 Low cost of land or house was decisive	31.8/29.1	33.1/34.9	35.2/36.0	0.031 [−0.073, 0.140]	0.577/0.712	Equivalent
E3 No better housing alternatives	32.6/68.0	33.5/29.7	33.9/2.3	−0.486 [−0.573, −0.393]	<0.001/<0.001	Not equivalent
E4 Proximity to work influenced location choice	31.8/7.6	31.4/36.0	36.9/56.4	0.362 [0.258, 0.461]	<0.001/<0.001	Not equivalent

Table 3. *Cont.*

Item	%4–5 (P/G)	%3 (P/G)	%1–2 (P/G)	r [95% CI]	*p*/*p* (FDR)	Equivalence
E5 Changing residence would be economically difficult	61.0/56.4	29.7/40.7	9.3/2.9	0.059 [−0.051, 0.167]	0.285/0.391	Equivalent
E6 Need to own a home outweighed sector risks	34.7/34.3	26.7/33.1	38.6/32.6	−0.035 [−0.141, 0.072]	0.533/0.679	Equivalent
E7 Continue living here out of economic necessity despite risk awareness	33.1/29.1	31.8/37.8	35.2/33.1	0.013 [−0.096, 0.121]	0.821/0.868	Equivalent
R1 Familiarity with the landslide concept	36.9/44.8	29.7/29.7	33.5/25.6	−0.097 [−0.202, 0.008]	0.081/0.130	Equivalent
R2 Awareness of historical soil-instability records	28.4/34.5	33.9/37.4	37.7/28.1	−0.115 [−0.216, −0.007]	0.039/0.065	Not equivalent
R3 Perceived probability of house damage (5 years)	36.4/33.1	28.8/40.7	34.7/26.2	−0.024 [−0.129, 0.081]	0.662/0.743	Equivalent
R4 Concern for house safety in rainy season	40.7/37.8	23.3/26.2	36.0/36.0	0.004 [−0.104, 0.114]	0.942/0.968	Equivalent
R5 Authorities adequately informed residents	33.5/34.9	28.4/30.8	38.1/34.3	−0.025 [−0.132, 0.083]	0.649/0.743	Equivalent
R6 Existing mitigation works are sufficient	31.4/35.5	35.6/33.7	33.1/30.8	−0.045 [−0.154, 0.063]	0.418/0.552	Equivalent
R7 Family knows how to act in a landslide	33.9/64.5	36.0/30.2	30.1/5.2	−0.432 [−0.529, −0.341]	<0.001/<0.001	Not equivalent

Two economic items concern staying in the sector. Changing residence would be economically difficult for 61.0% of Pomasqui and 56.4% of Guápulo households (E5), and fewer than 10% disagreed in either sector. Continuing to live in the sector out of economic necessity despite awareness of the risks (E7) was endorsed by 33.1% and 29.1%, and 35.2% and 33.1% disagreed (r = 0.013, 95% CI −0.096 to 0.121). Because E7 is double-barreled and presupposes risk awareness, its lower endorsement may partly reflect respondents who did not regard their dwelling as at risk; E5 is the more direct indicator of an economic constraint on moving, and it was endorsed by a majority in both sectors.

Ten of the 14 items differed between sectors by less than $|r| = 0.20$ (probability of superiority 0.40–0.60); nine did so within ±0.15, but only one (concern in the rainy season) within ±0.10, so moderate but not small differences are excluded. The item on awareness of historical instability records had a 95% interval that excluded zero (r = −0.115, −0.216 to −0.007) but did not survive FDR correction (p-FDR = 0.065). The effect of sector on three differing items persisted after adjustment for length of residence (proportional-odds OR = 7.32, 95% CI 4.48–11.95, for "no better alternatives"; 0.197, 0.123–0.314, for proximity to work; 3.76, 2.36–6.02, for family preparedness) and, for the first two, within both length-of-residence strata (Table S7). Odds for sector were not proportional for the first two items (Brant $p < 0.001$ and $p = 0.046$) but were for family preparedness ($p = 0.332$), so those adjusted ORs average effects that vary across thresholds; the rank-based and stratified results point in the same direction (Tables S7 and S14), and all three differences remained significant after FDR correction at a design effect of 3 (Table S13). Family preparedness differed among household residents more than 5 years (r = −0.563), but not among newer residents (r = −0.115; only 27 in Guápulo). Familiarity with the landslide concept was higher in Guápulo only after adjusting for length of residence (OR = 1.76, 1.12–2.76; p-FDR = 0.048 within the adjusted set); because this adjustment uses a nearly collinear covariate and a different correction family, the result is exploratory.

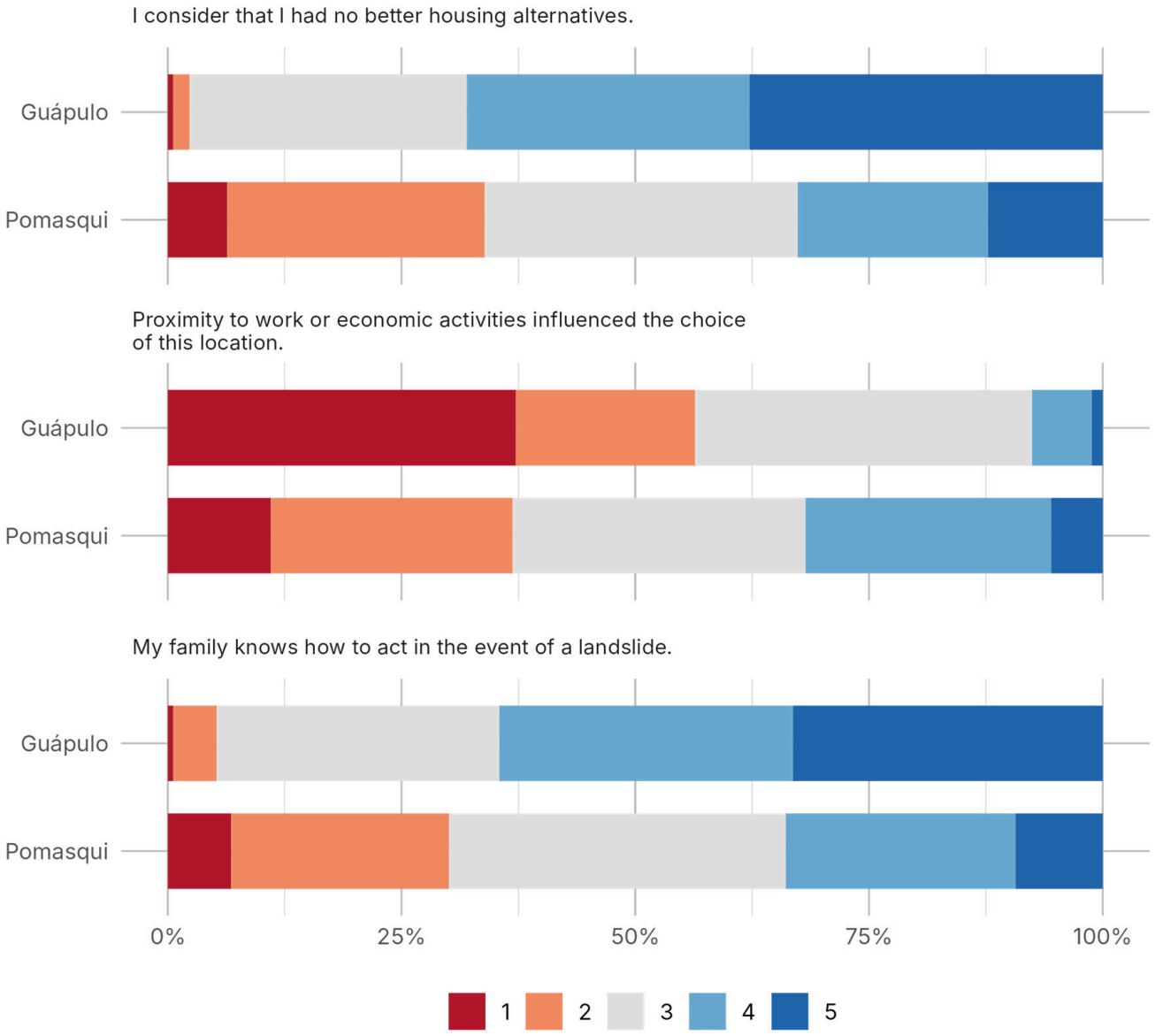

Figure 4. Distribution of Likert responses (1 = lowest to 5 = highest; percentage of households within each sector; Pomasqui n = 236, Guápulo n = 172) for the three items that differed between sectors after FDR correction.

3.5. Willingness to Invest and Reported Mitigation

Spending-priority rankings, in which aesthetics and finishes were ranked above structural safety (mean ranks 1.94 and 2.23) with low agreement among respondents (Kendall's W = 0.151), are reported in Table S11 and Figure S5.

Nearly half of the households (47.3%, 193/408) would invest none of their annual income in stabilization works. This share was higher in Guápulo (52.9%) than in Pomasqui (43.2%), and fewer Guápulo households would invest more than 10% (5.2% versus 13.6%; $p = 0.004$). Bivariately, willingness to invest also varied with tenure ($p = 0.019$) and increased with income (Spearman $\rho = 0.12$, $p = 0.015$; Table S8). In a proportional-odds model with the five primary covariates (global LR $\chi^2 = 38.80$, df = 13, $p < 0.001$), willingness to invest was associated with sector ($p = 0.001$), income ($p = 0.027$), perceived soil hazard ($p = 0.004$), and length of residence ($p = 0.035$), but not with tenure ($p = 0.087$). Guápulo households (OR = 0.42, 95% CI 0.25–0.71), owners (0.55, 0.32–0.93), and households perceiving very low or low hazard (0.53, 0.29–0.98) were willing to invest smaller shares, and

higher-income households were willing to invest larger shares (OR = 1.39 per category, 1.12–1.73). Odds for sector were not proportional (Brant $p = 0.003$; Table S15).

Reported signs of instability differed in two of four types: an uneven or tilted floor (55.8% in Guápulo versus 22.9% in Pomasqui; p-FDR < 0.001) and doors or windows that stick (37.2% versus 24.2%; p-FDR $= 0.017$). Overall, 79.4% of households (323/407) reported at least one physical mitigation measure (one Pomasqui household left the item blank and was treated as missing), with no difference by sector (78.7% in Pomasqui, 80.2% in Guápulo), but the types differed: land filling or leveling was reported by 41.3% in Guápulo and 11.9% in Pomasqui, and retaining walls were reported by 20.4% in Pomasqui and 10.5% in Guápulo. Emergency kits (52.9% versus 37.3%) and home insurance listed among preparedness measures (22.7% versus 9.3%) were more frequent in Guápulo (both p-FDR ≤ 0.008; Table S8). The separate question on insurance against natural hazards gave a different picture (41.3% of Guápulo households and 33.9% of Pomasqui households answered yes), so insurance was measured inconsistently and was excluded from Table 1 and the MCA (Section 4.6). Having implemented any physical measure was not associated with the predictors (omnibus $p \geq 0.097$) or with stated willingness to modify (78.9% among willing households, 80.0% among the others).

3.6. Logistic Regression Models (RQ3)

Overall, 70.1% of households (286/408) were willing to relocate (Pomasqui 70.3%, Guápulo 69.8%), and 60.8% (248/408) were willing to undertake structural modifications (64.4% and 55.8%). Table 4 gives the observed proportions by predictor level with the number of households. Perceived soil hazard showed a non-monotonic pattern (Figure 5): willingness to relocate was highest among households perceiving low hazard (89.7%, n = 29) and very high hazard (76.8%, n = 56) and lowest among those perceiving medium or high hazard (66.9% and 65.9%), and willingness to modify was lowest among households perceiving very low hazard (32.0%, n = 25).

Table 4. Observed willingness to relocate and to undertake structural modification by predictor level. Cells show n (%) [Wilson 95% CI]. n (P/G) is the number of households in Pomasqui and Guápulo. † Tenure categories 3 and 4 were merged in the models.

Predictor	Level	n (P/G)	Willing to Relocate	Willing to Undertake Structural Modification
Sector	Pomasqui	236/0	166 (70.3) [64.2–75.8]	152 (64.4) [58.1–70.2]
	Guápulo	0/172	120 (69.8) [62.5–76.1]	96 (55.8) [48.3–63.0]
Perceived soil hazard	Very low	17/8	19 (76.0) [56.6–88.5]	8 (32.0) [17.2–51.6]
	Low	27/2	26 (89.7) [73.6–96.4]	19 (65.5) [47.3–80.1]
	Medium	78/94	115 (66.9) [59.5–73.5]	107 (62.2) [54.8–69.1]
	High	79/47	83 (65.9) [57.2–73.6]	85 (67.5) [58.9–75.0]
	Very high	35/21	43 (76.8) [64.2–85.9]	29 (51.8) [39.0–64.3]
Tenure	Borrowed (informal, no lease)	17/61	51 (65.4) [54.3–75.0]	53 (67.9) [57.0–77.3]
	In legalization process †	4/1	4 (80.0) [37.6–96.4]	3 (60.0) [23.1–88.2]
	Owned	136/64	138 (69.0) [62.3–75.0]	114 (57.0) [50.1–63.7]
	Rented	79/46	93 (74.4) [66.1–81.2]	78 (62.4) [53.7–70.4]
Length of residence	Less than 1 year	81/8	64 (71.9) [61.8–80.2]	63 (70.8) [60.6–79.2]
	1 to 5 years	102/19	76 (62.8) [53.9–70.9]	68 (56.2) [47.3–64.7]
	6 to 10 years	41/63	73 (70.2) [60.8–78.1]	61 (58.7) [49.0–67.6]
	More than 10 years	12/82	73 (77.7) [68.2–84.9]	56 (59.6) [49.5–68.9]
Household income	Less than USD 300	16/26	33 (78.6) [64.1–88.3]	20 (47.6) [33.4–62.3]
	USD 300–600	63/44	74 (69.2) [59.9–77.1]	66 (61.7) [52.2–70.3]
	USD 601–1000	90/74	117 (71.3) [64.0–77.7]	101 (61.6) [54.0–68.7]
	More than USD 1000	67/28	62 (65.3) [55.3–74.1]	61 (64.2) [54.2–73.1]

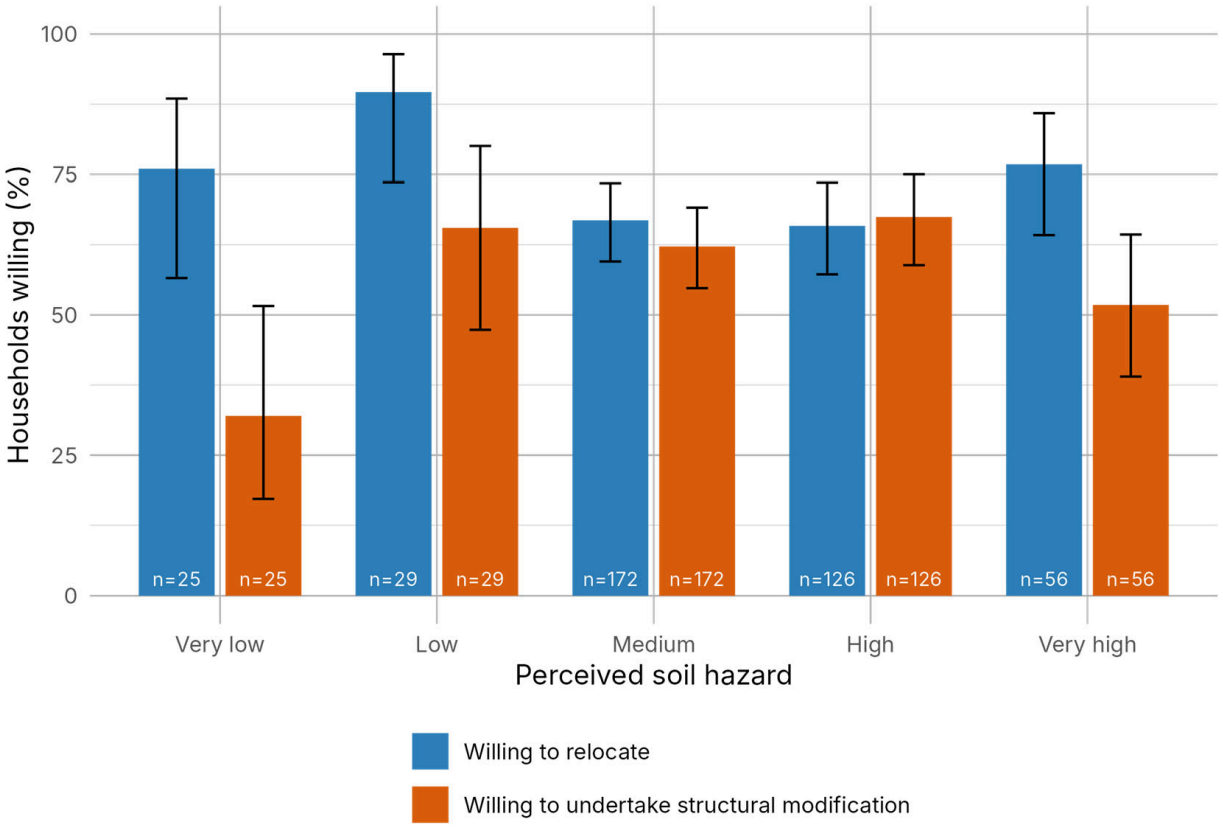

Figure 5. Observed percentage of households willing to relocate and to undertake structural modification by perceived soil hazard level (error bars: Wilson 95% CI; n shown within bars; N = 408).

Table 5 reports omnibus likelihood-ratio tests and model fit for the primary models (N = 408 for each). For willingness to relocate, perceived soil hazard ($\chi^2 = 12.88$, df = 4, $p = 0.012$) and length of residence ($\chi^2 = 11.04$, df = 3, $p = 0.012$) were associated with the outcome, whereas sector, income, and tenure were not ($p \geq 0.215$). For willingness to modify, perceived soil hazard ($\chi^2 = 14.66$, df = 4, $p = 0.005$), tenure ($\chi^2 = 6.76$, df = 2, $p = 0.034$), and sector ($\chi^2 = 3.96$, df = 1, $p = 0.047$) were associated with the outcome, whereas income ($p = 0.181$) and length of residence ($p = 0.084$) were not. The relationship with perceived soil hazard and length of residence departed from linearity ($p \leq 0.038$ in both models), consistent with the significant higher-order polynomial terms in Tables S3 and S4; income showed no departure from linearity ($p \geq 0.349$). These associations would lose significance at design effects between 1.03 (sector) and 1.55 (perceived soil hazard for modification), that is, with intra-cluster correlations below 0.03 for clusters of 20 households (Table S13).

Both models explained little variation (McFadden's $R^2 = 0.053$ and 0.060). The apparent AUC was 0.661 (95% CI 0.604–0.718) and 0.653 (0.598–0.708), but the optimism-corrected AUC was 0.603 and 0.600, with calibration slopes of 0.625 and 0.702, indicating overfitting. The models therefore have little predictive capacity, and the associations below describe differences in average willingness between groups, not an ability to predict individual intentions. The Hosmer–Lemeshow test did not indicate lack of fit ($p = 0.575$ and 0.712) but has little power with few covariate patterns. The adjusted GVIF was at most 1.39 (Table S9). Firth-penalized estimates differed from maximum-likelihood estimates by 0.10 log-odds or less (Table S9).

In the simplified-coding models (Table 6 and Figure 6), which are exploratory because the hazard categories were pooled after inspection of the observed proportions, households perceiving very low or low hazard were more willing to relocate than those perceiving medium hazard (OR = 2.81, 95% CI 1.24–6.35) corresponding to marginally standardized

probabilities of 84.0% (95% CI 72.6–93.3) and 65.8% (58.6–73.2), a difference of 18.2 percentage points (4.8–30.6; observed 83.3%, 45/54, versus 66.9%, 115/172) but less willing to undertake structural modification (OR = 0.47, 0.24–0.90). Of the 54 households in the very low or low category, 44 lived in Pomasqui; within Pomasqui alone the relocation OR was 4.38 (1.60–11.97), and a sector × hazard interaction was not significant ($p = 0.270$). Owners were less willing than households in borrowed dwellings or in legalization to undertake structural modification (OR = 0.46, 0.25–0.86), whereas renters did not differ from that group (OR = 0.73, 0.39–1.36). Guápulo households were less willing than Pomasqui households to modify (OR = 0.54, 0.31–0.94). Income entered as a linear score was associated with willingness to modify (OR = 1.29 per category, 1.01–1.64), but the omnibus test of the primary model was not significant ($p = 0.181$; Table 5), so income is not treated as associated with willingness to modify. The Hosmer–Lemeshow test indicated poor calibration for the simplified relocation model ($p = 0.017$; 0.112 for the structural model).

Table 5. Primary logistic models: Type II likelihood-ratio tests per predictor and model performance. Both models include sector, household income, tenure (three groups), perceived soil hazard, and length of residence (ordered factors). Departure from linearity: LR test of categorical versus linear coding (relocation/structural). Events per parameter use the smaller outcome group. AUC-corrected and calibration slope are bootstrap optimism-corrected (1000 resamples).

	Relocation: Statistic	***p***	**Structural Modification: Statistic**	***p***
Sector	$\chi^2 = 1.534$ (df 1)	0.215	$\chi^2 = 3.958$ (df 1)	0.047
Household income	$\chi^2 = 4.168$ (df 3)	0.244	$\chi^2 = 4.881$ (df 3)	0.181
Housing tenure	$\chi^2 = 1.530$ (df 2)	0.465	$\chi^2 = 6.759$ (df 2)	0.034
Perceived soil hazard	$\chi^2 = 12.876$ (df 4)	0.012	$\chi^2 = 14.664$ (df 4)	0.005
Length of residence	$\chi^2 = 11.036$ (df 3)	0.012	$\chi^2 = 6.651$ (df 3)	0.084
Departure from linearity				
Household income	$\chi^2 = 2.106$ (df 2)	0.349	$\chi^2 = 1.198$ (df 2)	0.549
Perceived soil hazard	$\chi^2 = 12.671$ (df 3)	0.005	$\chi^2 = 13.164$ (df 3)	0.004
Length of residence	$\chi^2 = 7.042$ (df 2)	0.03	$\chi^2 = 6.546$ (df 2)	0.038
Model performance				
N (households)	408		408	
Willing/not willing	286/122		248/160	
Parameters (events per parameter)	13 (9.4)		13 (12.3)	
Global LR test	$\chi^2 = 26.269$ (df 13)	0.016	$\chi^2 = 32.971$ (df 13)	0.002
McFadden's R^2	0.053		0.060	
AUC, apparent (95% CI)	0.661 (0.604–0.718)		0.653 (0.598–0.708)	
AUC, optimism-corrected	0.603		0.600	
Calibration slope, corrected	0.625		0.702	
Hosmer–Lemeshow		0.575		0.712

Because the primary models omit other covariates, four blocks were compared on 407 complete cases (Table S12). Removing the hazard-perception and knowledge block worsened the relocation model most ($\chi^2 = 47.38$, df = 12, $p < 0.001$; 3.95 per degree of freedom), and removing the housing-conditions block worsened the structural-modification model most ($\chi^2 = 46.57$, df = 12, $p < 0.001$; 3.88 per degree of freedom). With 2.9–3.8 events per parameter, calibration slopes of 0.53–0.56 and blocks of different sizes, this ordering is exploratory and is not a stable ranking of explanatory importance.

Table 6. Simplified-coding logistic models (maximum likelihood): odds ratios (OR) with 95% confidence intervals. Reference categories: sector = Pomasqui; tenure = borrowed or in legalization; perceived soil hazard = medium; length of residence = less than 1 year. Income is a linear score (OR per category). N = 408; relocation 286 willing and 122 not willing; structural modification 248 willing and 160 not willing. Hazard categories were pooled after inspection of the observed proportions, so these estimates are exploratory.

Term	Relocation: OR (95% CI)	*p*	Structural Modification: OR (95% CI)	*p*
Sector: Guápulo (vs. Pomasqui)	0.739 (0.406–1.345)	0.322	0.538 (0.307–0.944)	0.031
Household income (per category)	0.844 (0.653–1.091)	0.196	1.289 (1.014–1.639)	0.038
Tenure: Owned (vs. borrowed/legalization)	1.366 (0.729–2.559)	0.331	0.464 (0.251–0.858)	0.014
Tenure: Rented (vs. borrowed/legalization)	1.574 (0.821–3.015)	0.172	0.729 (0.390–1.363)	0.322
Perceived soil hazard: Very low/low (vs. Medium)	2.806 (1.240–6.348)	0.013	0.466 (0.241–0.903)	0.024
Perceived soil hazard: High/very high (vs. Medium)	1.219 (0.765–1.940)	0.405	0.895 (0.570–1.404)	0.629
Length of residence: 1 to 5 years (vs. <1 year)	0.665 (0.363–1.218)	0.186	0.561 (0.309–1.018)	0.057
Length of residence: 6 to 10 years (vs. <1 year)	1.180 (0.570–2.444)	0.655	0.635 (0.320–1.261)	0.195
Length of residence: More than 10 years (vs. <1 year)	2.028 (0.890–4.620)	0.093	0.850 (0.398–1.814)	0.674

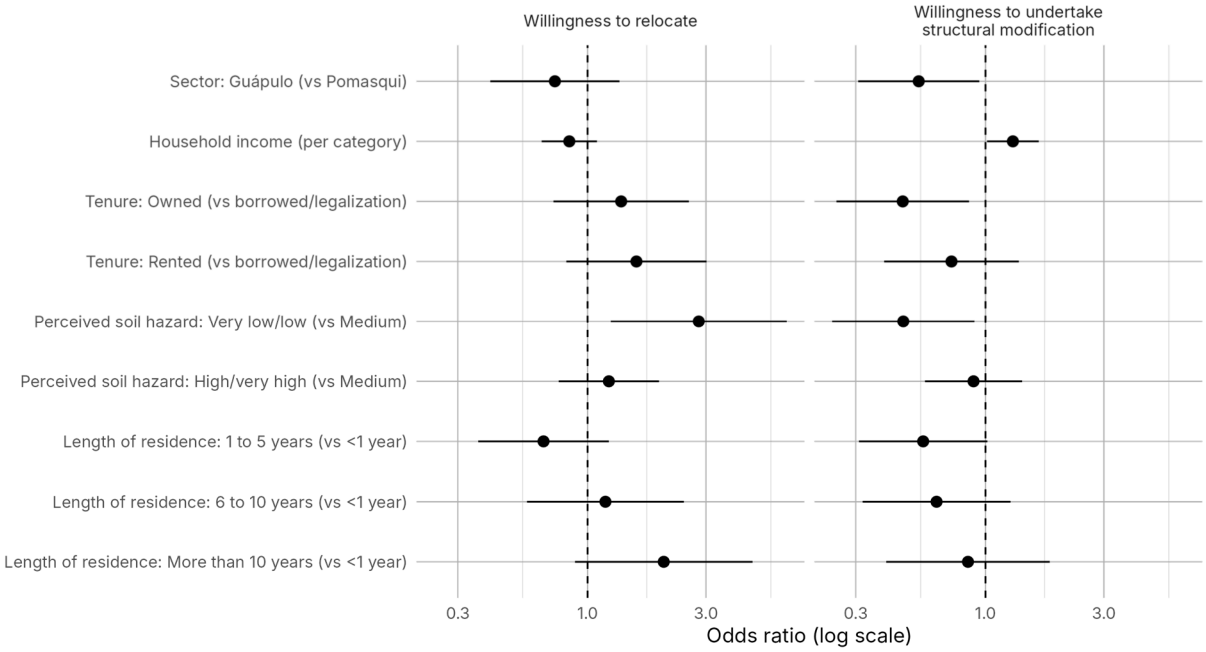

Figure 6. Odds ratios (log scale) and 95% confidence intervals from the simplified-coding models; the dashed line marks OR = 1 (N = 408). Reference categories as in Table 6.

The sector effect on structural modification was sensitive to the covariates. The crude OR was 0.70 (95% CI 0.47–1.04, $p = 0.080$); it changed little after adding length of residence (0.64) or income (0.73) but fell to 0.59 ($p = 0.016$) after adding tenure, and to 0.56 in the full primary model (Wald $p = 0.048$; likelihood-ratio $p = 0.047$; Table S9). Of three post-hoc codings of the sector × length-of-residence interaction, none was significant for relocation (linear score $p = 0.055$; two categories $p = 0.311$; four categories $p = 0.735$), and only the four-category coding was significant for structural modification ($\chi^2 = 11.81$, df = 3, $p = 0.008$; linear score $p = 0.386$; two categories $p = 0.181$). Among household residents of more than 5 years (Pomasqui n = 53, Guápulo n = 145), Guápulo households were less willing to

modify (55.2% versus 69.8%; adjusted OR = 0.34, 95% CI 0.16–0.74); among newer residents (n = 183 and 27), the difference was small (OR = 0.76, 0.31–1.82). Given the crude $p = 0.080$, the dependence on adjustment for tenure, the single significant coding, and the small cells, this sector effect is treated as hypothesis-generating.

Table 7 summarizes the robustness of the main associations. Adding respondent age and household size or excluding the households in legalization changed the estimates little (Table S16). Among owners only (n = 200), the sector and low-hazard estimates for structural modification kept their direction but were not significant, whereas the low-hazard association with relocation persisted. All model-based associations lost significance at design effects between 1.03 and 1.60. Marginal predicted probabilities (Table S10; Figure S3) were 73.1% in Pomasqui and 65.5% in Guápulo for relocation, and 66.1% and 53.1% for structural modification. By tenure, predicted willingness to modify was 69.9% for borrowed or legalizing households, 53.9% for owners and 65.2% for renters. Odds ratios of the primary structural-modification model are plotted in Figure S4.

Table 7. Robustness of the main associations from the simplified-coding logistic models: odds ratios (95% CI) in the full sample, with respondent age and household size added, and among owners only, and the design effect (deff) above which p would exceed 0.05. The OR for income is per category. Estimates excluding the five households in legalization are given in Table S16. – = not estimable (tenure is constant among owners); n.s. = not significant without correction.

Outcome and Association	Full Sample (N = 408)	+Age and Household Size	Owners Only (n = 200)	Tipping-Point deff
Relocation: very low/low vs. medium hazard	2.81 (1.24–6.35)	2.87 (1.26–6.53)	4.03 (1.33–12.19)	1.60
Relocation: more than 10 years vs. less than 1 year	2.03 (0.89–4.62)	2.09 (0.91–4.77)	2.03 (0.64–6.39)	n.s.
Structural modification: Guápulo vs. Pomasqui	0.54 (0.31–0.94)	0.56 (0.31–1.00)	0.43 (0.18–1.06)	1.22
Structural modification: owned vs. borrowed or in legalization	0.46 (0.25–0.86)	0.47 (0.25–0.87)	–	1.56
Structural modification: very low/low vs. medium hazard	0.47 (0.24–0.90)	0.50 (0.25–0.97)	0.41 (0.17–1.00)	1.33
Structural modification: income (per category)	1.29 (1.01–1.64)	1.30 (1.02–1.67)	1.22 (0.83–1.79)	1.12

4. Discussion

4.1. Main Findings

Three results frame the interpretation. First, the sectors differ strongly in who lives there (length of residence, tenure, housing, hazard knowledge), but differences on most risk-perception items were smaller than $|r| = 0.20$, which excludes moderate but not small differences. Second, the results were only partly consistent with the expectations derived from PMT and the vulnerability frameworks (H1–H4; because these were formulated after the analysis, this is a descriptive comparison, not a test). Perceived soil hazard was associated with both intentions, but not monotonically (H1); income was not associated with willingness to modify in the primary model (H2); owners were less, not more, willing to modify, and tenure was unrelated to relocation (H3); and willingness to relocate did not decline with length of residence (H4). H5 was not supported for relocation and only partly for structural modification, but this assessment rests on an exploratory block comparison and on economic items of low reliability, so it is not a fair test. Third, all associations were weak and fragile: the optimism-corrected AUC was about 0.60, and every model-based association would lose significance at a design effect of 1.6 or less. Conjectures

C1–C5 below were formulated after the analysis and are hypotheses for future studies, not tested explanations.

4.2. Sector Contrasts and Their Confounding with Length of Residence

Sector and length of residence are nearly collinear ($V = 0.641$), so sector differences cannot be separated from cohort of arrival in this design, and because interviewer teams were nested within sectors, neither can be separated from interviewer effects [42]. Every sector difference reported here is therefore a difference between two groups of households that also differ in length of residence and interviewer team, not an effect of the sectors themselves. The contrasts could describe two stages of urban consolidation, a recently urbanized periphery and a long-established ravine settlement (C1). This is consistent with accounts of incremental, path-dependent consolidation in Latin American informal settlements [4,66] and with the uneven expansion documented for the DMQ [11,13], but with two sectors C1 cannot be tested. The sector difference in "no better housing alternatives" persisted after adjusting for length of residence, within both strata and at a design effect of 3, so it is unlikely to be only a cohort effect.

4.3. Risk Perception: Partial Convergence and the Role of Limited Alternatives

Ten of the 14 items, including five of the seven risk-perception items (R1, R3, R4, R5, and R6), differed between sectors by less than $|r| = 0.20$, among them the perceived probability of damage and concern in the rainy season. This margin is lenient: only one item met a margin of ± 0.10, so small differences cannot be excluded, and four of these items have a response format that does not match their wording. With these caveats, the similarity of perceived probability and concern despite large differences in tenure, housing, and hazard knowledge is consistent with reviews reporting that risk perception is weakly related to hazard experience and protective behavior [22,23], and with a study in which residents who had experienced floods or landslides were less willing than the general public to adopt mitigation measures [67] (C2). Convergence is partial: family preparedness, awareness of historical records and several knowledge and experience variables differed (Section 3.2). Shared institutional narratives could contribute to the similarity of items on information and mitigation works (C3), but no data here address this.

Two economic items bear on staying. A majority in both sectors (61.0% and 56.4%) agreed that moving would be economically difficult (E5), whereas about one-third endorsed living there out of economic necessity despite the risks (E7), a double-barreled item that presupposes risk awareness. Economic constraints on moving therefore appear widespread in both sectors, and the lower endorsement of E7 cannot be read as evidence that economic necessity matters little. Two-thirds of Guápulo households also reported having no better housing alternative. That perception coincides with a larger share of borrowed dwellings in Guápulo (35.5%) and with lack of alternatives being named as the main reason for residence (34.3%), and it is consistent with housing constraints being salient in that sector. Because residential permanence was not measured, these data cannot show whether such constraints keep households in place.

4.4. Adaptive Intentions: Comparison with Previous Studies

For willingness to relocate, the association with perceived soil hazard does not agree in form with PMT or with the Loja survey, where hazard-related variables were positively associated with protective behavior [27]: households perceiving very low or low hazard were the most willing to relocate (84% versus 66% at medium hazard). This group was small ($n = 54$) and concentrated in Pomasqui (44 households), its pooling was chosen after the pattern had been seen, and the simplified relocation model was poorly calibrated. The

association persisted within Pomasqui and among owners, so sector and tenure do not explain it; a low rating may refer to the soil under the dwelling rather than the surrounding slope, residents who see little reason to stay may downplay local hazard, or response styles may differ between interviewer teams. We treat it as an anomaly for future study, not as evidence against threat appraisal. The agreement with the Türkiye study is also partial [25]: there, individual characteristics, self-efficacy, and institutional support mattered, whereas here neither income nor tenure was associated with relocation, and self-efficacy and institutional support were not measured. The higher willingness after more than 10 years of residence rests on a non-significant contrast (OR = 2.03, $p = 0.093$) in a group of whom 82 of 94 lived in Guápulo, so it is not evidence against place attachment [24,68], and cumulative exposure to instability (C4) remains a conjecture. Finally, 70% of households were willing to relocate although a majority considered moving economically difficult (E5). The two answers are compatible because the item assumes institutional support and safe housing, but a hypothetical, socially approved option asked face to face is prone to acquiescence and social-desirability bias [69,70], so the 70% indicates conditional openness to supported relocation [26], not demand.

For structural modification, owners were less willing than households in borrowed dwellings or in legalization, contrary to the expectation in H3. That comparator is small (n = 83) and heterogeneous, and the item does not state who would pay, so owners may have assumed that they would bear the whole cost, whereas households in borrowed dwellings may have expected the owner or an institution to pay (C5). Owners were also willing to invest smaller shares of their own income in stabilization works (OR = 0.55), which makes a pure wording effect less likely and is consistent with response-cost considerations. The owners-only analysis is central: among owners, the sector and low-hazard estimates kept their direction but were not significant (Table 7), so these results depend partly on the tenure composition of the sample. The sector coefficient emerged mainly after adjustment for tenure, was significant in only one of three interaction codings and lost significance at a design effect of 1.03, so it is hypothesis-generating. Income was associated with willingness to invest but not, in the primary model, with willingness to modify. Forced rankings placed aesthetics above structural safety (Table S11), a weak proxy for spending that needs confirmation; households in borrowed dwellings ranked safety like owners, which does not support an explanation based on tenure insecurity [71]. Reported physical mitigation (79.4%) was unrelated to stated willingness to modify, resembling the intention–behavior gap documented in Loja and Bangladesh [27,28].

4.5. Explanatory Power and Model Reliability

The relevant fact for interpretation is not the number of significant coefficients but the low explanatory power. McFadden's R^2 of 0.053 and 0.060, a corrected AUC of about 0.60, and calibration slopes of 0.63–0.70 indicate that the five-predictor models explain little and that apparent discrimination is optimistic; they describe group differences in average willingness and have little capacity to predict an individual household's intention. Adding blocks raised the apparent AUC to 0.79, but with 2.9–3.8 events per parameter and calibration slopes near 0.55, this gain is largely overfitting since the corrected AUC reached only 0.69–0.70 [62,64]. The Hosmer–Lemeshow test did not reject the primary models ($p = 0.575$ and 0.712) but did reject the simplified relocation model ($p = 0.017$), so predicted probabilities are approximate. Moreover, modest clustering by interviewer or neighborhood would remove the significance of every model-based association, whereas the larger descriptive differences between sectors are robust to it. Most variation in both intentions therefore remains unexplained by the measured variables.

4.6. Limitations and Future Research

Design and sampling. Participants were selected by convenience sampling, without a sampling frame, so the sample cannot be assumed to represent the households of either sector, and its composition could not be compared with census data. Convenience samples can over-represent households that are at home and willing to talk and under-represent those that are absent or refuse, and the direction of the resulting bias on perceived hazard and intentions is unknown [40,41]. Interviewer teams were assigned to sectors, so interviewer effects may be confounded with sector [42]. Interviewer allocation by area was intended to prevent repeated interviews with the same household, but one respondent per household was not verified independently, and education and occupation were reported by proxy. Households are probably clustered by neighborhood or interviewer area, and no cluster identifier was collected; the design-effect analysis indicates which results are robust to such clustering (Table S13). The cross-sectional design cannot establish temporal order, and both intentions are hypothetical.

Measurement. The Likert blocks had low reliability (alpha, ordinal alpha and omega below 0.50), four risk items have a response format that does not match their wording, E7 is double-barreled, and the EFA produced Heywood-type solutions, so the constructs were not measured as scales. Because the questionnaire was not pilot-tested with residents, problems of wording and format, and not only the multidimensionality of the concepts, may explain the low reliability. Eight of the 14 items had about one-third of responses in each of the ranges 1–2, 3, and 4–5, with few extreme answers, which may reflect satisficing or acquiescence [69]; the export has no timestamps or interviewer identifiers to examine this. Housing conditions are not a housing-deficit measure. No GIS-based or geotechnical assessment of susceptibility, hazard, or dwelling exposure was made, so the study cannot show whether perceived hazard corresponds to actual spatial risk, and perceived hazard is not a hazard assessment; a susceptibility map exists for urban Quito [12], but the dwellings were not geolocated, and the other documents reviewed describe hazards qualitatively [11]. Both intentions were measured with single, hypothetical, binary items; the modification item does not specify who pays, and the relocation item removes part of the economic constraint. Insurance was measured inconsistently and was excluded from the comparisons. Face-to-face administration may have induced social-desirability bias on items such as income, tenure, and the two intentions [70].

Analysis. The 13 coefficients per model were not adjusted for multiplicity and several p-values lie close to 0.05; the simplified coding, the interaction codings, and the owners-only and age and household-size models were post hoc; the equivalence margin was set post hoc and is lenient; and the sample could reliably detect only associations of $w \geq 0.14$–0.18, so non-significant results are not evidence of no association. Many analyses were exploratory, and the results should be replicated in independent samples.

Future work should use probability or stratified sampling with recorded interviewer and neighborhood identifiers; add objective susceptibility at the dwelling or block level (for example by geolocating dwellings and linking them to the susceptibility map for Quito [12] and the municipal Monjas basin diagnosis [32], whose findings were not used here); measure standard housing-deficit indicators, actual mitigation behavior, and who bears the cost; validate multi-item scales for Andean settings after cognitive pretesting with residents; and extend the comparison to more sectors and to longitudinal designs.

5. Conclusions

Findings. In this convenience sample of 408 households in two hillside sectors of Quito with reported landslide events, Pomasqui and Guápulo differed markedly in length of residence, tenure, housing conditions, and hazard knowledge, but differences on most risk-

perception items were below $|r| = 0.20$, so moderate differences are excluded but small ones are not. About two-thirds of Guápulo households, versus one-third in Pomasqui, reported having no better housing alternative, and a majority in both sectors considered moving economically difficult. Willingness to relocate was associated non-monotonically with perceived soil hazard and with length of residence; willingness to modify was associated with tenure, perceived soil hazard, and, in a hypothesis-generating way, sector. These model-based associations were weak and would not withstand modest clustering.

Contribution and limitations. For research on socio-environmental vulnerability, the study shows, with household-level data from Andean hillside settlements, that socioeconomic and housing indicators, perceived hazard and stated intentions are only loosely coupled, and that the expected dominance of economic and tenure constraints was not observed. It also shows how a two-sector comparison is confounded with residential history and interviewer allocation, and how a design-effect analysis separates robust descriptive contrasts from fragile model-based associations. The associations are exploratory and cross-sectional; explanatory power was modest (optimism-corrected AUC ≈ 0.60), the Likert blocks were unreliable, exposure was self-reported, and the findings cannot be generalized beyond the sampled households.

Implications. The data do not support direct policy recommendations, but they suggest proposals that risk-management programs in the DMQ could test, in line with Sustainable Development Goal 11, Targets 11.1, and 11.5 [29]: (i) when geotechnical reports recommend reinforcement, state who pays, and offer cost-sharing schemes that distinguish owners from households in borrowed or rented dwellings; (ii) direct hazard information to Pomasqui, where most households had arrived within 5 years and 70.8% had not known of the landslide risk before moving in, linked to the municipal Monjas basin diagnosis [32]; (iii) frame relocation offers around safe and affordable housing alternatives, since willingness was conditional on institutional support and most households considered moving economically difficult; and (iv) record dwelling location and objective susceptibility in surveys so that perceived and actual hazard can be compared. Each proposal should be evaluated with probability samples and designs that measure actual behavior, exposure and costs.

Supplementary Materials: The following supporting information can be downloaded at https://doi.org/10.5281/zenodo.22262309 (accessed on 10 October 2026): Table S1: Software environment (R and package versions); Table S2: Multiple correspondence analysis with 24 active variables (insurance excluded): eigenvalues (raw, Benzécri- and Greenacre-adjusted), active variables, variable-level eta-squared, category contributions, sector separation on Dimension 1 and restricted-MCA sensitivity; Table S3: Polynomial-contrast coefficients of the primary model for willingness to relocate; Table S4: Polynomial-contrast coefficients of the primary model for willingness to modify; Table S5: Sociodemographic and housing profile of surveyed households by sector; Table S6: Exploratory factor analysis: summary, loadings (varimax and oblimin), communalities and item–rest correlations; Table S7: Likert items: distributions, equivalence at three margins, sector effects adjusted for length of residence and stratified comparisons; Table S8: Willingness to invest, signs of instability, mitigation and preparedness measures, and model for any physical mitigation; Table S9: Primary-model diagnostics and sensitivity analyses (exclusion of the households in legalization, Firth estimates), sector odds-ratio adjustment, sector × length-of-residence interactions and stratified analyses; Table S10: Marginal predicted probabilities from the primary models by sector, tenure, perceived soil hazard and length of residence; Table S11: Spending-priority rankings: Friedman test, mean ranks, post hoc comparisons and comparisons by sector, tenure and income; Table S12: Comparison of covariate blocks; Table S13: Design-effect sensitivity analysis; Table S14: Brant tests of the proportional-odds assumption; Table S15: Proportional-odds model for willingness to invest; Table S16: Simplified-coding logistic models with respondent age and household size, for owners only and excluding the households in legalization; Table S17: Likert items: full response frequencies by sector, E1 × E2 cross-tabulation, inter-item correlations and straight-lining checks; Figure S1: Length of residence by sector; Figure S2:

Distribution of the 14 Likert items by sector; Figure S3: Marginal predicted probabilities of willingness to relocate and to modify from the primary models; Figure S4: Odds ratios of the primary model for willingness to modify; Figure S5: Mean rank of four spending categories; Figure S6: Individual factor map of the multiple correspondence analysis; Supplementary File S1: Questionnaire (original Spanish version and English translation).

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Institutional Review Board Statement: Ethical review and approval were waived for this study because it was classified as "Research Without Risk" under Article 27(h) of the Internal Regulation of the Human Research Ethics Committee of the Universidad Técnica Particular de Loja (CEISH-UTPL; code DGMU_RG_1_2023_V1; University Council Resolution No. 870.955.2023), and is therefore exempt from full committee review under Article 41. The survey was anonymous, collected no personal identifiers and was conducted in accordance with the Declaration of Helsinki and the Ecuadorian Organic Law on Personal Data Protection.

Informed Consent Statement: Informed consent was obtained from all subjects involved in the study.

Data Availability Statement: The datasets presented are publicly available at https://doi.org/10.5281/zenodo.22262309.

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Conflicts of Interest: The authors declare no conflicts of interest.

Abbreviations

The following abbreviations are used in this manuscript:

ASR	Adjusted standardized residual
AUC	Area under the receiver operating characteristic curve
CEISH-UTPL	Human Research Ethics Committee of the Universidad Técnica Particular de Loja
CI	Confidence interval
deff	Design effect
df	Degrees of freedom
DMQ	Metropolitan District of Quito
EFA	Exploratory factor analysis
FDR	False discovery rate
GVIF	Generalized variance inflation factor
KMO	Kaiser–Meyer–Olkin
LR	Likelihood ratio
MCA	Multiple correspondence analysis
MOVE	Methods for the Improvement of Vulnerability Assessment in Europe

OR	Odds ratio
PMT	Protection Motivation Theory
ROC	Receiver operating characteristic
RQ	Research question
TOST	Two one-sided tests
UTPL	Universidad Técnica Particular de Loja

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