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Measuring Poverty Dynamics in Iran Using a Synthetic Panel Approach

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Abstract

Economic sanctions have become a routine instrument of foreign policy, yet their distributional consequences for households in target countries remain poorly understood. This paper examines how successive waves of international sanctions reshaped poverty dynamics and income mobility in Iran over 2010–2019. Using six rounds of nationally representative Household Income and Expenditure Surveys, we construct synthetic panels from repeated cross-sections and estimate non-parametric and parametric bounds on poverty transitions, validating the approach against the 2014–2016 household panel and a set of robustness checks.

We distinguish three episodes—sanctions escalation (2010–2013), the JCPOA-related relief (2014–2016), and the "maximum pressure'' campaign (2017–2019). Chronic poverty rises from roughly 7.5–11% in the first episode to 8–15.5% under relief and then surges to 18–22% under maximum pressure. Poverty entry responds sharply to sanction intensity, increasing by about 8 percentage points between 2014–2016 and 2017–2019, while exits remain persistently low even when sanctions are eased. These asymmetric transitions signal deepening long-term deprivation and growing vulnerability among near-poor households.

The effects are far from uniform. Female-headed, rural, low-education, and informally employed households face the highest risks of remaining or becoming poor, and chronic poverty becomes increasingly concentrated in peripheral provinces, whereas central regions are more insulated. Overall, sanctions not only raise aggregate poverty but also amplify regional and socio-economic divides, and short periods of partial relief, without strong domestic reforms, appear insufficient to reverse these dynamics.

Keywords: economic sanctions; poverty dynamics; income mobility; chronic poverty; synthetic panels; Iran; regional inequality

JEL Classifications: I32, O15, F51

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Introduction

In 1919, U.S. President Woodrow Wilson described economic sanctions as "something more tremendous than war," an instrument that "does not cost a life outside the nation boycotted, but…brings pressure upon that nation which…no modern nation could resist." A century later, sanctions have shifted from rare, exceptional measures to a routine tool of statecraft. The United States, and increasingly other major powers, now rely on a broad repertoire of economic restrictions whose frequency and intensity have risen sharply over the past three decades (Drezner, 2021; Niyazova et al., 2025).

Economic sanctions are a coercive instrument imposed by states and international organizations to induce policy change in target governments, often framed as a less violent and more "surgical" alternative to military intervention. For the sending country, sanctions may appear relatively low-cost, at least in the short run. For the target country, however, they can reshape macroeconomic trajectories and distributive outcomes. By depressing investment, reducing long-run growth, and squeezing public and private budgets, protracted sanctions can erode both human capital and infrastructure and impose substantial welfare losses (Allen and Lektzian, 2013; Zamani, 2024). These losses are rarely borne evenly: the welfare of ordinary households is typically more exposed than that of politically connected groups, with distributional effects that depend on domestic institutions and coalition structures.

Iran has been one of the most heavily sanctioned economies in the world since the 1979 Islamic Revolution and, especially, over the last two decades. Early sanctions in the 1980s and 1990s largely sought to constrain Iran's regional influence. From the early 2000s onward, however, U.S., EU, and UN measures were increasingly justified by concerns over Iran's nuclear program. Successive rounds of UN, U.S., and EU restrictions progressively targeted Iran's core oil sector, banking, shipping, and insurance, sharply curtailing oil revenues and external financial access (Hinz, 2017). As a result, Iran has become a central case in debates over both the effectiveness and the humanitarian consequences of economic sanctions.

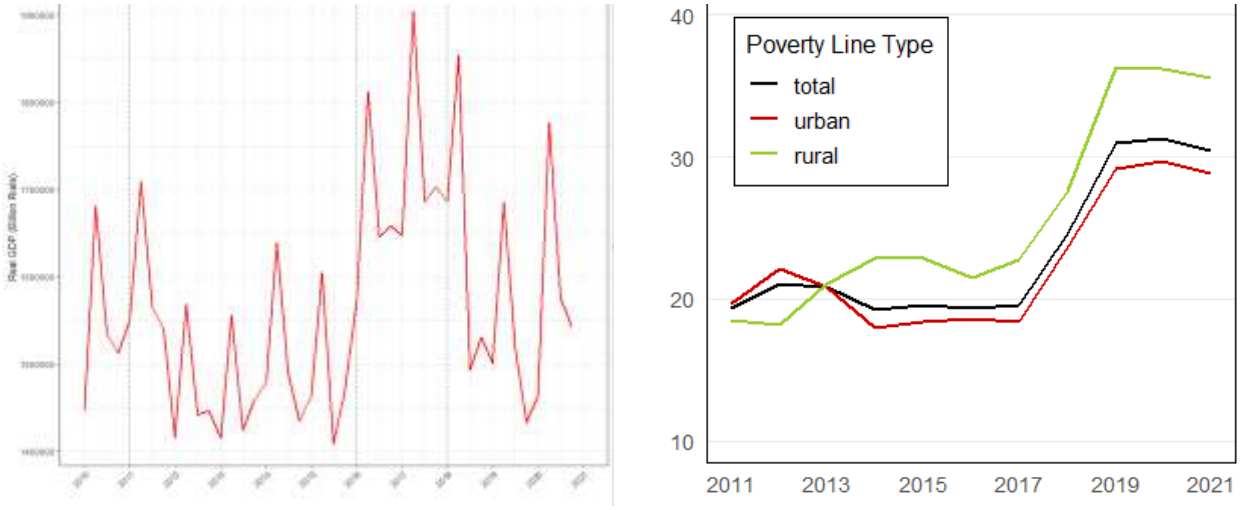
The ethical and welfare implications of sanctions were thrown into sharp relief by the experience of Iraq in the 1990s, where comprehensive UN sanctions are estimated to have doubled child mortality (Ali and Shah, 2000). While sanctions typically do not entail the immediate physical destruction associated with war, they can generate comparable long-run welfare losses. Recent

empirical work has documented large effects of sanctions on output, macroeconomic volatility, and trade patterns, including for Iran. By contrast, much less attention has been paid to how sanctions reshape poverty and income distribution within target countries, particularly along the intensive margin of movements into and out of poverty.[1]

This paper examines how poverty and income mobility evolved in Iran during the 2010s, a decade marked by the escalation, partial suspension, and aggressive re-imposition of economic sanctions. We focus on the period 2010–2019, deliberately excluding the post-2019 years to avoid conflating sanctions with the Covid-19 shock. In the early 2010s, successive UN and especially U.S. and EU measures broadened into far-reaching restrictions on Iran's oil exports and banking system. The Joint Comprehensive Plan of Action (JCPOA), agreed in July 2015 and implemented from January 2016, suspended many nuclear-related sanctions in exchange for constraints on Iran's nuclear program, creating a brief window of relief. This trajectory was reversed after the U.S. withdrawal from the JCPOA in May 2018 and the re-imposition of secondary sanctions on oil, banking, shipping, and related services by November 2018, inaugurating the "maximum pressure" campaign. Between 2017 and 2019, Iran's GDP fell from 457.5 to 435 billion U.S. dollars, while the national headcount poverty rate rose by almost ten percentage points in less than two years (see Figure 1).

Beyond these aggregate trends, sanctions operate through multiple channels. Direct effects arise from lower oil export revenues, restricted access to foreign exchange and finance, and disruptions to trade. Indirect effects include higher transaction costs, shortages of critical imports (including medicines and medical equipment despite formal exemptions), and the fiscal and administrative costs of mitigation and circumvention. Quantifying the burden of these combined channels on household welfare is therefore of particular relevance. Yet the poverty-monitoring literature on Iran has largely focused on changes in poverty headcount ratios over time, devoting relatively little attention to poverty *dynamics* at the household level. Observed persistence—where individuals or households remain poor over long periods—would be consistent with theoretical accounts in which sanctions reinforce poverty-trap mechanisms (Neuenkirch and Neumeier, 2015; Afesorgbor and Mahadevan, 2016).

(A) **(B)**



Iran's Real GDP Iran's poverty rate

Figure 1. (A) *Quarterly Real GDP in billion rials. The dotted vertical lines mark the beginning of the Obama-era sanctions, the JCPOA-related relief, and the maximum-pressure campaign. Source: Central Bank of Iran.* **(B)** *Poverty rates in Iran throughout the 2010s. Source: Parliamentary Research Center of Iran.*

Studying poverty mobility requires longitudinal data that track the same households over time. Such panel datasets are rare in developing countries, expensive to collect, and often limited by small samples and attrition[2]. In their absence, a large literature has relied on pseudo-panels constructed from repeated cross-sections, grouping individuals into cohorts to approximate dynamics.[3] More recently, synthetic-panel methods have been developed that model households' outcomes as functions of time-invariant characteristics, allowing the estimation of bounds on poverty transitions using only repeated cross-sections.

Building on this latter approach, we construct synthetic panels from six rounds of Iran's Household Income and Expenditure Surveys (HIES) for 2010, 2013, 2014,2016, 2017, and 2019 to analyze poverty dynamics under three sanctions episodes: the escalation of sanctions in the early 2010s, the JCPOA-related relief, and the subsequent maximum-pressure campaign. We estimate non-parametric and parametric bounds on poverty transitions, validate the synthetic-panel approach against the limited genuine panel data available for 2014–2016, and compare our findings with earlier work on Iran's poverty trajectories (e.g., Ghomi, 2022). Relative to existing studies, our analysis (i) extends the time window to cover the maximum-pressure episode, (ii) exploits a richer

set of covariates, and (iii) draws on more survey waves, allowing more precise and robust estimates of mobility.

A second contribution is to document how poverty dynamics vary across socio-economic groups and space. Iran's "lost decade" (2011–2020) was characterized by growing disparities between central and peripheral regions, with poverty increasingly concentrated in the Southeast and Northwest while more central provinces fared relatively better (World Bank, 2023).[4] National trends or broad urban–rural breakdowns can conceal substantial provincial heterogeneity. We therefore estimate chronic and transitory poverty, and movements into and out of poverty, across all 31 provinces and by household characteristics such as education, gender of the household head, sector of employment and urban–rural residence. This allows us to identify the groups and regions most exposed to the deep macroeconomic shocks associated with sanctions.

The rest of the paper is organized as follows. Section 2 outlines the evolution of international sanctions on Iran and reviews the related literature. Section 3 describes the data and empirical strategy, including the synthetic-panel methodology and the measurement of poverty dynamics. Section 4 presents the main results on chronic and transitory poverty and examines heterogeneity across socio-economic groups and provinces. Section 5 reports robustness checks. Section 6 concludes.

2 History of Sanctions and Economic Adjustment in Iran

2.1 Chronology and Categorization of U.S. Economic Sanctions on Iran

In this section, we briefly review the three phases of sanctions on Iran during the 2010s: the escalation of U.S. sanctions under the Obama administration, the partial relief following the JCPOA, and the renewed intensification of sanctions under the Trump administration's "maximum pressure" campaign. This chronology provides the institutional and policy context for analyzing how successive sanctions episodes shaped the structure and dynamics of Iran's economy.

Phase I: Escalation and Secondary Sanctions (2010–2015)

Although U.S. sanctions on Iran date back to 1979, the period 2010–2015 marks a sharp escalation. Following the breakdown of nuclear talks in 2009, Washington and its allies moved toward far-reaching secondary sanctions that penalized non-U.S. entities doing business with Iran. UN Security Council Resolution 1929 (2010), the Comprehensive Iran Sanctions, Accountability, and

Divestment Act (CISADA, 2010), and subsequent U.S. legislation extended restrictions to the Central Bank of Iran, major commercial banks, and the oil sector, while the EU's 2012 decision to disconnect Iranian banks from SWIFT deepened Iran's isolation from the global financial system (Torbat, 2005).

By 2012–2013, these measures had produced a pronounced economic downturn: synthetic-control estimates suggest that Iran's real GDP fell to roughly 17% below its counterfactual level, alongside sharp declines in oil exports (Gharehgozli, 2017). Rising inflation, currency depreciation, and fiscal strain set the stage for the poverty dynamics discussed elsewhere in this paper (World Bank, 2023).

Table 1 – Phase I: Escalation and Secondary Sanctions (2010–2015)

Sanctions Category	Key Measures and Enforcement	Economic Impact (summary)
Global financial and banking isolation	UNSC Resolution 1929; CISADA; sanctions on the Central Bank; EU SWIFT cutoff	Severe output contraction; sharp fall in oil exports
Targeted nuclear and military entities	Designations under E.O. 13382 and related orders targeting proliferation and missile-related entities	Freezing of reserves; constraints on foreign-exchange inflows

Phase II: The JCPOA Respite (2016–May 2018)

Phase II began with the implementation of the Joint Comprehensive Plan of Action (JCPOA) in January 2016. In exchange for verifiable nuclear constraints, the U.S., EU, and UN suspended or lifted a wide set of nuclear-related sanctions. The United States waived many secondary sanctions on Iran's oil, gas, petrochemical, shipping, and automotive sectors and delisted several hundred individuals and entities, allowing non-U.S. firms to re-enter the Iranian market without facing U.S. penalties. However, core primary sanctions, such as the U.S. trade embargo, dollar-clearing bans, and terrorism-related designations, remained in place, leaving major international banks cautious (Farzanegan & Batmanghelidj, 2023; Nelson & Rosen, 2024). Sanctions relief generated a short but visible economic rebound: oil exports and GDP recovered, and poverty indicators temporarily improved, particularly in rural areas (World Bank, 2023). Yet lingering legal and political uncertainty constrained foreign investment and made the recovery fragile.

Table 2 – Phase II: The JCPOA Respite (2016–May 2018)

Sanctions Category	Key Measures and Relief	Economic Outcomes (2016–2018, summary)
Nuclear-related secondary sanctions	Suspension of many secondary sanctions on oil, petrochemicals, shipping, autos; delisting of many SDN entities	Oil exports and GDP rebound; renewed trade prospects
Financial and banking	Partial re-integration into SWIFT; limited banking channels opened for non-U.S. trade	Improved access to FX; continued risk aversion by major banks
Remaining core sanctions	U.S. primary embargo and terrorism-related measures remain in force	Structural legal risk persists; incomplete normalization

Phase III: The Maximum Pressure Campaign (2018–2020)

On 8 May 2018, the United States withdrew from the JCPOA and launched the Maximum Pressure campaign. All nuclear-related waivers were revoked, and sanctions were expanded to encompass nearly all major sectors of Iran's economy. Executive Orders 13846, 13871, 13876, and 13902 re-imposed and broadened sanctions on the financial system, the leadership office, metals, automotive, construction, mining, and the entire oil and petrochemical value chain. Shipping companies, ports, and numerous state-owned enterprises were also targeted (Nelson & Rosen, 2024). The resulting shock led to a collapse in oil exports, renewed recession, and a surge in inflation and currency depreciation (Laudati & Pesaran, 2023). During this phase, household-level evidence points to large welfare losses and a substantial increase in the share of the population at risk of falling into poverty (Salehi-Isfahani, 2023; World Bank, 2023).

Table 3 – Phase III: Maximum Pressure Campaign (2018–2020)

Sanctions Category	Key Target Sectors and Executive Orders	Economic Outcomes (2018–2020, summary)
Financial and banking	E.O. 13846, 13902; sanctions on major banks and financial services; designation of leadership office	FX shortages; high inflation; and deep currency depreciation
Energy and industrial	E.O. 13871, 13902; sanctions on oil, gas, petrochemicals, metals, autos, construction, mining	Oil exports drop sharply; a collapse in oil revenues
Shipping & transportation	Sanctions on IRISL, NITC, ports, and logistics companies	Constrained foreign trade; higher transaction costs

Post-2020 Developments: Sanctions Under the Biden Administration

With the inauguration of President Biden in January 2021, U.S. rhetoric shifted and diplomatic efforts to revive the JCPOA resumed, but no comprehensive agreement had been restored by the

mid-2020s. The administration maintained the core architecture of late-2010s sanctions on Iran's financial and energy sectors, while modestly expanding humanitarian channels, such as mechanisms for food, medicine, and COVID-19 vaccines, and intermittently relaxing enforcement of oil sanctions, enabling Iran to increase exports mainly to China (Nelson & Rosen, 2024; Farzanegan & Batmanghelidj, 2023).

2.2 Literature Review

Although economic sanctions are formally intended to change the policies of target governments, a large theoretical and empirical literature shows that they have often failed to do so while imposing substantial economic and humanitarian costs on the population, making it crucial to quantitatively assess their effects on the poverty gap and wider social and economic outcomes, particularly through the deepening of poverty. Hufbauer et al. (2007) show that sanctions succeed in changing target-state behavior in only about one-third of cases, despite imposing substantial economic and humanitarian costs. Eaton and Engers (1992, 1999) provide the core theoretical framework for understanding the strategic dynamics of sanctions in repeated interactions, while Dizaji and Van Bergeijk (2013) offer empirical evidence that output and revenue losses are typically largest in the early years of a sanctions episode and diminish as the target economy adapts. In the Iranian context, studies such as Esfahani, Mohaddes, and Pesaran (2013) and Laudati and Pesaran (2023) document that external shocks, especially disruptions to oil exports, generate sharp but relatively short-lived macroeconomic fluctuations, highlighting both the economy's vulnerability to sanctions and its capacity for adjustment over time.

Within this broader debate, a first strand of work quantifies the aggregate output and trade costs of sanctions. Using cross-country panel data, Neuenkirch and Neumeier (2015) find that UN and U.S. sanctions significantly reduce real GDP per capita growth in target economies. Country-specific work on Iran using synthetic control by Gharehgozli (2017) suggest that real GDP was more than 17% lower over 2011–2014 than in a no-sanctions counterfactual, with the deepest contraction in 2012. Trade-focused studies show more reorientation than collapse: Torbat (2005) argues that unilateral crude-oil bans are partly circumvented because Iran can redirect exports to alternative buyers, whereas financial sanctions that restrict borrowing and access to capital goods have more persistent effects. Gravity-based work documents substantial trade destruction and diversion for Iran and other sanctioned economies, with export shortfalls for both the target and its

main partners (Haidar, 2017; Crozet & Hinz, 2020). Recent survey evidence highlights that the growing reliance on financial and secondary sanctions magnifies macroeconomic costs and raises adjustment frictions for firms and households (Farzanegan & Batmanghelidj, 2023).

A second strand shifts the focus from aggregate growth to distributional outcomes, poverty, and inequality. Cross-country evidence indicates that sanctions have regressive welfare effects: Drury and Peksen (2014) document that women's economic and social vulnerability increases during sanction episodes, while Ali and Shah (2000) find that UN sanctions on Iraq were associated with a doubling of infant and child mortality. Peksen (2016), using panel data, argues that sanctions can exacerbate economic and political discrimination against ethnic minorities by shrinking the resource base and incentivizing governments to adopt exclusionary policies. Taken together, these contributions suggest that sanctions tend to widen distributional gaps and harm those with the least political and economic power.

The Iranian experience is consistent with these distributional patterns. Using a CGE model, Gharibnavaz and Waschik (2018) estimate that international sanctions cut overall welfare by about 15%, with rural households facing nearly twice the losses of urban households and the poorest deciles hit hardest. At the micro level, Salehi-Isfahani (2023) shows that real per capita household consumption fell by roughly 20% between 2010 and 2020 and that the national poverty rate almost doubled, with rural, female-headed, larger, and less-educated households significantly more likely to fall into poverty.

Recent work connects these welfare losses to food security and diet quality. Hejazi and Emamgholipour (2022) find that, one year after the U.S. withdrawal from the JCPOA and the onset of "maximum pressure," prices of most food items had risen by more than 50%, pushing the cost of a healthy diet beyond the reach of many households, especially in lower expenditure groups. Using three decades of household survey data, Roustaee et al. (2024) document declining milk and yogurt consumption, particularly among poorer households, and show that income and socioeconomic characteristics strongly shape dairy intake.

Despite these advances, the existing literature has several limitations for the purposes of this study. First, most empirical work focuses on earlier sanction episodes up to the mid-2010s and does not cover the post-2018 period, when the U.S. withdrawal from the JCPOA and "maximum pressure" sanctions sharply intensified external constraints. Second, poverty is largely treated in a static way,

through consumption or income levels at specific dates, rather than as a dynamic process of movements into and out of poverty. Third, most studies rely on national or, at best, broad urban–rural aggregates and pay limited attention to the geographical dimension of vulnerability within Iran, such as provincial differences in exposure to sanctions and in poverty trajectories. Finally, relatively few contributions systematically distinguish between employment segments, public versus private and formal versus informal, when analyzing how sanctions affect household welfare. By exploiting repeated cross-sectional household expenditure data and synthetic panel methods, and by explicitly incorporating both provincial and labor-market heterogeneity, the present study addresses these gaps and documents how sanctions reshape poverty dynamics and income mobility in a regionally unequal middle-income economy.

3 Data and Methodology

3.1 Synthetic panel method

This section provides a summary of the methodology proposed by Dang et al. (2014). We consider two rounds of repeated cross-sectional surveys, $i = 1,2$. We are primarily interested in identifying households that were poor in the first round and exited poverty in the second or vice versa, captured in Equation (1), denoting per capita income or expenditure by y_i and the poverty line by z_i.

$$P(y_{i1} < z_1 \; and \; y_{i2} > z_2) \tag{1}$$

Since the same unit i is not followed over time in repeated cross-sectional surveys, it is not possible to directly estimate the probabilities required for analyzing dynamic poverty transitions. However, synthetic panels can be constructed, which provide both upper and lower bounds for each of the four potential outcomes of the dynamic process.

For this purpose, we apply a linear decomposition of consumption as shown in the equation below, where x_{it} is a vector of time-invariant characteristics of households, such as language, religion, or location, and ε_{it} is the error term.

$$y_{it} = \beta_1' x_{it} + \varepsilon_{it} \quad t = 1,2 \tag{2}$$

Observations from the second round allow us to estimate first-round consumption or income by applying the first-round OLS parameter estimates (β_1') to the corresponding time-invariant household characteristics in round 2 (x_{i2}).

The model, as outlined by Dang et al. (2014), relies on two critical assumptions that must be satisfied for the bounds on poverty transitions to be credible. First, it is assumed that the underlying population sampled is the same in both rounds. It may not hold if the population shifts over time because of demographic changes, such as births, deaths, or migration out of the sample, and significant economic or natural events between the surveys.[5] Annex A, Figures A, depicts Iran's overall demographic dynamics, including births, deaths, and migration, over the 2010-2019 period. The data indicate that these variables remained relatively constant throughout these years. Furthermore, Annex A, Table A1 presents t-test results for the time-invariant characteristics of households across the two survey rounds. Except for the literacy of the household head and the household's place of residence, all other characteristics show no significant differences at the 95% confidence level. Second, the correlation between the error terms of the consumption model in the two rounds is non-negative or positive quadrant dependent.[6]

3.1.1 Non-parametric bounds

We are able to estimate both upper and lower bounds of mobility by considering two extreme cases of the assumed relationship between the error terms.

Assuming full independence of the error terms across survey rounds ($Corr(\varepsilon_{i1}, \varepsilon_{i2}) = 0$) produces the upper-bound estimate of mobility, by randomly drawing with replacement for each household *i* in the second round from the empirical distribution of first round estimated residuals (denoted by $\widehat{\widetilde{\varepsilon_{i1}}}$) as follows:

$$\hat{y}_{i1}^{2U} = \hat{\beta}_1' x_{i2} + \widehat{\widetilde{\varepsilon_{i1}}} \tag{3}$$

Based on the predicted consumption levels obtained from equation (3), we estimate the transitions into and out of poverty. Given the stochastic nature of the estimation, the simulation is performed R times, and the mean of equation 3 across all replications provides the upper-bound estimate of mobility.[7] If we assume perfect correlation between the error terms ($Corr(\varepsilon_{i1}, \varepsilon_{i2}) = 1$) We can obtain the lower-bound estimate of poverty mobility. In this case, the residuals estimated from the second round, adjusted by λ, can be directly used to produce predicted values of first-round consumption or income, as specified below.[8]

$$\hat{y}_{i1}^{2L} = \hat{\beta}_1' x_{i2} + \gamma \hat{\varepsilon}_{i2} \tag{4}$$

3.1.2 Parametric bounds

The non-parametric approach discussed earlier has the advantage of requiring only a few assumptions while still producing reasonably promising results. However, its main limitation lies in the fact that achieving more precise estimates and narrower bounds requires a rich set of time-invariant covariates. In the absence of such comprehensive data, the resulting bounds may be excessively wide and of limited practical use. To overcome this, Dang et al. (2014) propose a parametric framework which, despite its stronger assumptions, yields more consistent and robust estimates. This specification further assumes that ε_{i1} and ε_{i2} have a bivariate normal distribution with a non-negative correlation coefficient ρ and standard deviations δ_{ε_1} and δ_{ε_2} respectively.

Using the estimated regression parameters from equation 2 and the predicted standard errors $\hat{\delta}_{\varepsilon_1}, \hat{\delta}_{\varepsilon_2}$ for the error terms, the probability that a household escapes poverty is calculated as follows, where $\emptyset_2$ refers to the cumulative density function of a binormal distribution.

$$\hat{p}\ (y_{i1} < z_1\ and\ y_{i2} > z_2) = \ \emptyset_2\left(\frac{z_1-\beta_1' x_{i2}}{\delta_{\varepsilon_1}}\ , -\frac{z_2-\beta_2' x_{i2}}{\delta_{\varepsilon_2}}\ , -\rho\right) \tag{5}$$

The upper bound of the above probability is determined by the maximum possible value of ρ, while the lower bound is obtained from the minimum value.

3.2 Data

To measure poverty dynamics in Iran, we draw on the Household Income and Expenditure Surveys (HIES) conducted by the Statistical Center of Iran (SCI). The HIES is a nationally representative survey of Iranian households that provides detailed information on household expenditures, incomes, and the socioeconomic characteristics of household members. Since 1968, the SCI has surveyed approximately 40,000 households annually in both urban and rural areas.

In the surveys, households are selected based on a three-stage stratified random cluster sampling method. The census areas are classified and selected at the first stage. At the second stage, the urban and rural blocks are selected, and the selection of households is done at the third stage.[9] To analyze how sanctions have influenced poverty dynamics over Iran's "lost decade", we employ household survey data from the years 2010, 2013, 2014, 2016, 2017, and 2019, all of which fall within the 2010–2020 decade. Corresponding poverty lines are obtained from the official poverty

rates, reported in Figure 1, which are calculated annually by the Research Center of the Iranian Parliament (MRC).[10]

In order to satisfy the model's fundamental assumption that variables remain invariant over time, we incorporate three distinct categories of explanatory variables. The first category of covariates captures the characteristics of the household head, including gender, age, marital status, educational attainment, and economic activity.[11] The second category focuses on household composition, comprising the total number of household members, the number of children aged 0–6 years in 2017, and the number of school-aged children (7–18 years) in 2017. The third category encompasses residential and household asset characteristics, including urban or rural residence, dwelling size, home ownership, vehicle ownership, and the availability of household amenities.[12] Since the set of residential variables, covering access to utilities and living conditions, was both extensive and highly correlated, we employed Multiple Correspondence Analysis (MCA) to reduce their dimensionality to two components, capturing over 80% of the variation in the original dataset.

Although certain variables in our specification, such as the household head's occupation, household size, and urban/rural status, may not be entirely time-invariant, they can be regarded as approximately stable over time under reasonable assumptions. The HIES follows a rotating panel structure that allows identifying common households every three years, with roughly one-quarter of them replaced in each of the subsequent two years. Between the 2018–2019 and 2019–2020 rounds, 18,114 households were observed in both surveys, and as illustrated in Annex B, the distribution of their key characteristics indicates that these variables can be considered time-invariant in short intervals. Finally, we limit the sample to households with heads aged between 25 and 60 to reduce life-cycle heterogeneity and ensure the comparability of the underlying population.

4 The Impact of sanctions on poverty Dynamics

This section documents how Iran's poverty dynamics evolved under different sanction regimes during the 2010s. We begin by estimating poverty transitions during 2017–2019, the period corresponding to the U.S. "maximum pressure" campaign. We then place these results in historical context by comparing them with two earlier episodes—the pre-JCPOA tightening of sanctions (2010–2013) and the subsequent phase of partial sanctions relief (2014–2016). Finally, we

examine the heterogeneity of these effects, analysing how the intensity and persistence of poverty vary across different household groups and across provinces. Together, these results provide a comprehensive picture of how external economic pressure reshaped income mobility and welfare across Iranian society.

4.1 Overall Results from the Synthetic Panel

In this section, we present the parametric and non-parametric results from applying the synthetic-panel methodology to Iran's Household Income and Expenditure Survey (HIES) for 2017 and 2019. We regress log per capita real expenditure on a set of time-invariant characteristics (location, age, education, and gender of the household head, household composition, amenities, assets and labor-market status), using three specifications.[13] As shown in Tables C1 and C2 of Annex C, the third and richest specification, which expands the set of covariates, provides the best overall fit.

Drawing on non-parametric synthetic-panel estimates for Iran, the analysis examines poverty transitions over a brief but turbulent interval running from the twilight of JCPOA-era into "maximum pressure". Rather than single point estimates, the method yields, an interval with lower and upper bounds.

Under model 3 illustrated in table 4, the implied ranges for the four transition states are approximately between 18–27.6 percent for households that are poor in both 2017 and 2019. These intervals point to a substantial group of *chronically poor* households: Even when using the first specification, whose intervals are relatively loose, close to one-fifth of households are poor in both years, while a more conservative reading pushes this share towards one-quarter or slightly more. The shift from the JCPOA period to the maximum-pressure phase did not represent a temporary disturbance but rather the continuation of a regime of sustained deprivation.

By contrast, downward mobility is sizeable: the probability that initially better-off households fell into poverty lies between roughly 14.6 and 29 percent, indicating that the re-imposition of sanctions and the associated macroeconomic pressures pushed a significant fraction of near-poor and middle-income families below the line. While the share of chronically poor households has risen markedly, opportunities for upward mobility (move from poor to non-poor status) remain scarce. The lower bound for exits from poverty is essentially zero, and even the upper bound suggests that, at best, fewer than one in five poor households (17.5 percent) manage to escape poverty between 2017 and 2019.

A combination of high entry and low exit generates a net accumulation of poor households within the most vulnerable subgroups.

Finally, between about 46.4 and 57.7 per cent of households remain above the poverty line in both years. This pattern suggests that groups with structural insulation from the shock—such as those in secure public-sector positions, those benefiting from economic rents or foreign-currency income, and those holding assets or income streams that appreciated with inflation—were least affected and were able to stay non-poor across both periods.

Taken together, the interval estimates across all three models depict a pattern of double pressur**e** in Iran during the late 2010s sanctions era: a sizeable, persistent core of chronic poverty at the bottom of the distribution; high vulnerability and downward mobility among segments of the initially non-poor, and only modest upward mobility to offset these losses. This configuration also implies an erosion of the middle class, pushing many of them into transitory or sustained poverty.

Table 4. Transitions matrices - repeated cross-sections using the nonparametric model 2017-2019

Status in 2017, 2019	Nonparametric lower bound			Nonparametric upper bound		
	Model 1	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**
Poor–poor	28.5	28.9	27.6	16.8	17.4	18.1
Poor–non-poor	0	0	0	17.7	17.9	17.5
Non-poor–poor	13.6	13.3	14.6	29.2	29.1	29.4
Non-poor– non-poor	57.7	57.7	57.7	43.8	44	46.4
# Observations	28079	28079	28079	28079	28079	28079

Data source: Household Income and Expenditure Survey (HIES), Statistical Center of Iran
Results are constrained to the panel sample of households whose heads are between 25 and 65 years old. Results are weighted using household-level survey-sampling weights. Number of replications for the estimates is 500.

In order to sharpen our bounds on mobility, we also computed income mobility and poverty transitions using the parametric version of the synthetic-panel approach. As discussed earlier, this specification assumes that the two error terms, ε_{i1} and ε_{i2}, have a bi-normal distribution with correlation coefficient $\rho > 0$. To compute poverty mobility using the parametric approach, we need to estimate the correlation coefficient of the error terms, ρ. Following the recommendation of Dang and et al (2014), when genuine panel data are available for a subset of years one can estimate ρ directly from those panels and then use the resulting values to calibrate the bounds in the synthetic-panel model. In their applications to six countries, they report values of ρ typically in the range 0.4–0.6.[14] In the case of the Iranian HIES, panel data are available for certain years with smaller

sample sizes, and this study exploits those real panel data to estimate the error-term correlation. Table 5 reports the resulting correlation coefficients for selected years of the 2010s, which lie in an approximate interval from 0.1 to 0.5. In the empirical analysis, we therefore adopt two sets of constraints for ρ: a broader interval (0.2, 0.6) in the first set of bounds, and a more conservative interval (0.3, 0.5) in the second.

Table 5. Estimated correlation of error terms (ρ) for panel data in the 2010s.

Year	Correlation coefficient (ρ)	Number of observations
2018-2019	0.119	18114
2016-2017	0.112	20136
2015-2017	0.387	9139
2015-2016	0.494	19708
2014-2016	0.411	8378
2014-2015	0.471	20132

Source: Calculations based on Household Income and Expenditure data in the 2010s, Statistical Center of Iran

Table 6 reports poverty mobility using the parametric version of the synthetic-panel model. As expected, the parametric approach yields substantially tighter intervals than the non-parametric bounds, and for each transition state the parametric ranges lie fully inside the corresponding non-parametric intervals. Under the first correlation constraint for the error terms ($\rho \in (0.2, 0.6)$), between 17 and 21 percent of households are classified as poor in both periods, indicating a sizeable group experiencing chronic poverty over the 2017–2019 interval. Over the same period, only about 6 to 11 percent of households are estimated to move out of poverty, suggesting that upward mobility out of poverty is relatively limited. Among households initially above the poverty line, 18-23 percent fall into poverty, while roughly half remain non-poor in both years.

Because the parametric bounds rely on a bivariate normality assumption for the regression errors, we also assess the plausibility of this assumption. Formal Anderson–Darling tests[15] reject exact bivariate normality, even though the empirical distributions of the error terms for 2017 and 2019 (see Figures 1 and 2 in Annex D) against the standard normal distribution show a close visual correspondence.

To address this concern, we conduct a robustness experiment in which we simulate a bivariate normal distribution for the error terms using the moments of the estimated residuals, and then repeat the parametric estimation for different values of ρ. The resulting upper and lower bounds

for poverty transition probabilities are virtually identical, up to two decimal places, to those presented in Table 6 (The full set of simulation results is presented in Annex D, Table D1).

Table 6. Transition matrices - repeated cross-sections using the parametric model. 2017- 2019

Status in 2017, 2019	Nonparametric	parametric lower bound		parametric upper bound		Nonparametric
		Constraint 1	Constraint 2	Constraint 1	Constraint 2	
Poor–poor	28.5	21.7	20.4	17	18.1	16.8
Poor–non-poor	0	6.5	7.8	11.3	10.2	17.7
Non-poor–poor	13.6	17.8	18.7	22.6	22	29.2
Non-poor– non-poor	57.7	53.8	52.5	49	50.1	43.8
# Observations	28079	28079	28079	28079	28079	28079

Data source: Household Income and Expenditure Survey (HIES), Statistical Center of Iran
Results are constrained to the panel sample of households whose heads are between 25 and 65 years old. Results are weighted using household-level survey-sampling weights. ρ = (0.2, 0.6) in the first constraint, and ρ = (0.3, 0.5) in the second constraint.

4.2 Poverty mobility under changing sanction regimes

In the previous subsection, we showed that between 2017 and 2019, poverty in Iran increased sharply: a large share of households fell into poverty, and relatively few managed to escape it. A natural question that arises is whether sanctions themselves are responsible for this pattern, or whether it can be explained entirely by domestic mismanagement, rent-seeking, corruption, and climate shocks. To assess rigorously whether sanctions have in fact had an independent impact, we extend the analysis by computing poverty mobility across the different sanction episodes of the 2010s. Three sanction regimes of the 2010s are large, externally driven shocks. The intensity of sanctions fluctuated markedly over this decade, creating a useful opportunity to examine how tightening and relaxation of external pressure affect movements into and out of poverty. While the domestic factors mentioned earlier undoubtedly played a role, our results indicate that the dominant force shaping these poverty dynamics has been the sanctions themselves.

Table 7 compares poverty transitions across three such regimes.[16] The share of chronically poor households rises from about 7.5–11 percent in the Obama episode to around 8–15.5 percent in the JCPOA episode, and then jumps to roughly 18–22 percent under maximum pressure. Thus, even in the period when sanctions were partially relaxed, and macroeconomic indicators briefly improved, there is no evidence that this easing translated into meaningful gains for the most vulnerable households. The share of chronic poverty did not fall; if anything, it increased, and it surged further once sanctions were tightened again.

These patterns should also be interpreted in light of the very different macroeconomic starting points of the two main sanction waves. The first round of nuclear-related sanctions in 2010–2013 hit an economy coming off almost a decade of relatively strong performance, with annual growth of around 6–8 per cent, inflation broadly contained at 10–12 per cent, and exceptionally high oil revenues between 2005 and 2010. By contrast, when the maximum-pressure sanctions were re-imposed in 2018, Iran had already experienced several years of weak growth, eroded foreign-exchange buffers, lower investment, and strained public finances. Thus, even though the basic legal architecture of sanctions in 2010–2013 and 2017–2019 was similar, the latter shock struck a far more fragile economy, which helps to explain why chronic poverty and downward mobility intensified so sharply in the late 2010s.

Table 7 Income mobility in Iran in the 2010s

Status in the 2010s	Obama-era sanctions		sanctions relief		maximum pressure Regime	
	lower bound	upper bound	lower bound	upper bound	lower bound	upper bound
Poor–poor	11.2	7.4	11.5	7.9	21.7	18.1
Poor–non-poor	9.9	13.8	9.1	12.8	6.5	10.2
Non-poor–poor	9.7	13.5	8.4	12	17.8	22
Non-poor– non-poor	69	65	70.9	67.3	53.8	50.1
# Observations	28079	28079	28079	28079	28079	28079

Data source: Household Income and Expenditure Survey (HIES), Statistical Center of Iran
Results are constrained to the panel sample of households whose heads are between 25 and 65 years old.
Results are weighted using household-level survey-sampling weights.

The probability of entering poverty reacts even more sharply to changes in sanction intensity. Between 2010–2013 and 2014–2016, the risk of falling into poverty rises only modestly, by about 2.5 percentage points. between 2014–2016 and 2017–2019, however, it increases by around 8 percentage points on average, with the bounds in the Trump period reaching around18–22 percent. Put differently, under "maximum pressure" the likelihood of falling into poverty rose considerably, especially for near-poor and middle-income households.

By contrast, the chances of exiting poverty do not increase when sanctions are relaxed: the share of households moving out of poverty is very similar in the Obama and JCPOA periods. However, during the Trump "maximum pressure" period, this probability in fact declines by roughly two percentage points on average. This asymmetry (sharp increases in the probability of entering poverty when sanctions intensify, but only small and slow improvements in the chances of

escaping poverty when sanctions are eased) indicates that sanctions relief, in the absence of strong domestic reforms, has limited power to reduce poverty. Once households have been pushed into chronic poverty, modest improvements in macroeconomic conditions appear insufficient to lift them out again.

These patterns also help to explain why the JCPOA period looks so similar to the earlier Obama episode in terms of poverty exits and overall income mobility. The nuclear agreement was in force for a relatively short time, and the complex secondary sanctions architecture built up since 2010 was only partially unwound, so that a number of restrictions remained in place even after implementation day. Moreover, many key U.S. measures were merely suspended through executive orders rather than permanently removed, keeping perceived investment risk in Iran high. As a result, much of the temporary improvement in macroeconomic indicators reflected a rebound in oil exports rather than a broad-based revival of private investment and productivity growth. This helps to explain why, despite initial optimism, the period of partial sanctions relief did not translate into meaningful gains for ordinary households, in terms of lower poverty, greater economic security, and helping vulnerable families move out of deprivation.

A similar story emerges for households that remain non-poor in both periods. Under Obama-era sanctions, 65–69 percent of households stay above the poverty line; this range increases slightly to 67–71 percent in the subsequent sanctions-relief period, consistent with modest gains in income security when external pressure is eased. Under maximum pressure, however, the share of households that remain above the poverty line falls to around 50–54 per cent. In fact, around 17 percentage points of households that were non-poor during the JCPOA years fell into poverty by 2017–2019. In many other studies on developing countries that use similar methods, researchers conclude that "the likelihood of falling into poverty has decreased considerably" over time;[17] in Iran, the opposite is true. Sanctions, regardless of whether they are being tightened or partially relaxed, appear to widen inequalities, hollow out the middle, and deepen the divide between a growing group of chronically poor households and a shrinking core of securely non-poor households, while leaving a large share increasingly vulnerable at the margin of poverty.

Figure 2 and 3 reinforce this interpretation by reporting, for a range of subgroups, the upper bounds on the probabilities of remaining poor and of entering poverty across these three critical sanction episodes. In virtually all subgroups, both persistence and entry probabilities are highest in the

Trump period. Moreover, some groups display high risks of chronic poverty in all three regimes, suggesting that they faced difficulties escaping poverty even before sanctions were tightened and that sanctions then amplified pre-existing structural vulnerabilities. We return to these heterogeneous patterns in more detail in the next section.

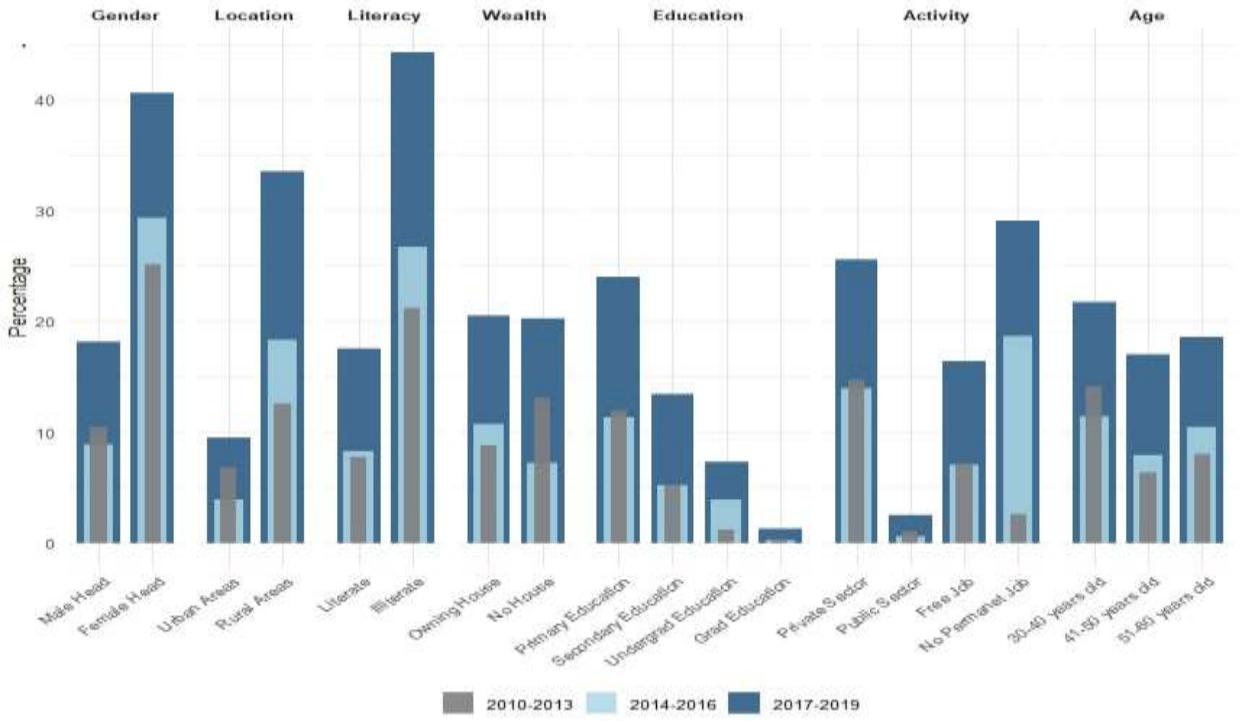


Figure 2. *Comparison of the upper bound of the poverty persistence rate for different subgroups across the three periods.*

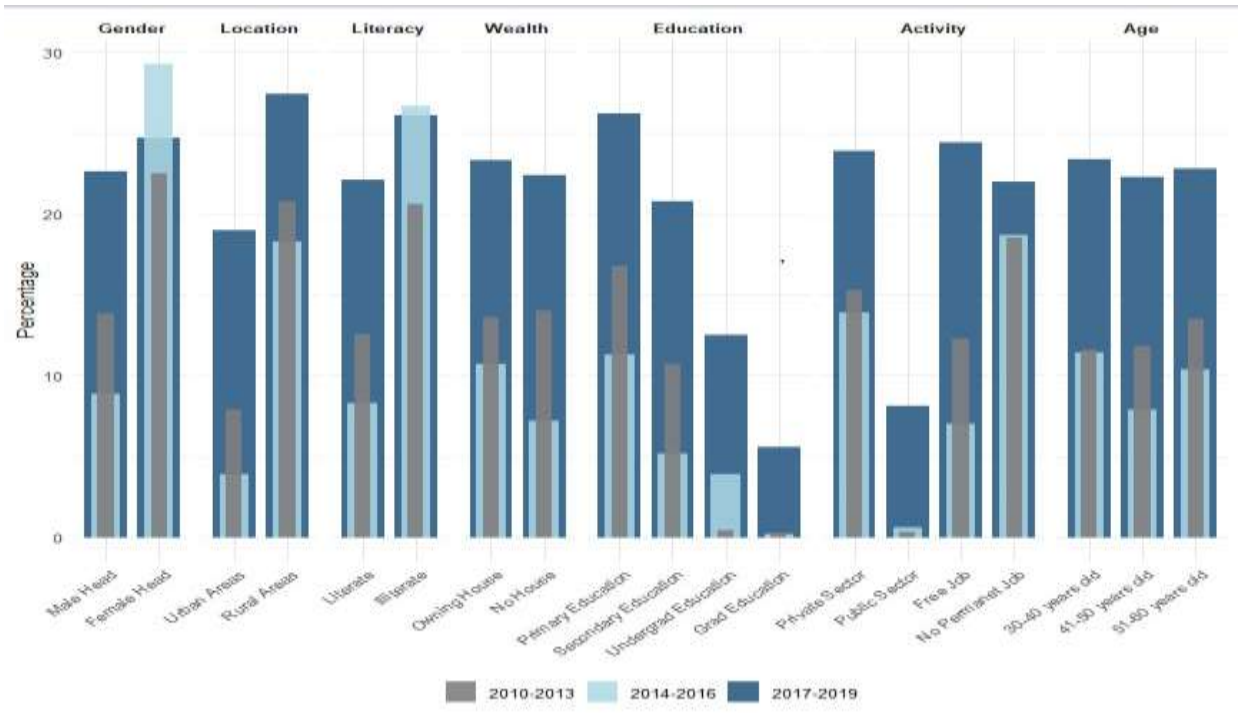


Figure 3. *Comparison of the upper bound of the poverty entry rate for different subgroups across the three periods*

4.3 Heterogeneity and inequality across subgroups

The aggregate results in Sections 4.1 and 4.2 show that, across the three major sanction episodes of the 2010s, chronic poverty increased and income security deteriorated for the population as a whole. Yet, such aggregate patterns may obscure significant differences across subgroups, where the effects of sanctions are far from uniform. Identifying which groups are more or less likely to remain in, enter, or escape poverty is crucial for designing effective anti-poverty policies. In this section, we compare the distribution of households across poverty transition categories by subgroup, focusing first on chronic poverty and then downward mobility. Figure 4 summarizes these patterns for 2017–2019.

The evidence on chronic poverty points to strong heterogeneity by education, gender, location, and labor-market position. Households with an illiterate head record the highest persistence in poverty, with a chronic-poverty rate of 44 percent on average, while the probability of remaining poor declines sharply with higher levels of education. Female-headed households also face a very high risk of chronic poverty, around 41 percent, compared with only about 18 percent for male-headed households, indicating a large gender gap in vulnerability. Urban households have a much lower

probability of remaining poor (around 9.5 percent) than rural households (about 33.5 percent), reflecting the spatial concentration of persistent deprivation in rural areas. Labor-market status is equally important: households whose head is employed in the public sector have the lowest chronic-poverty rate (about 2.5 percent), whereas those with no clearly defined job or in unstable employment face persistence rates close to 29 percent. These patterns suggest that secure, formal employment and higher education provide powerful protection against long-term poverty generated by externally imposed shocks such as sanctions, while female-headed households, rural residency, low level of schooling and labor-market informality are strongly associated with being trapped in poverty.

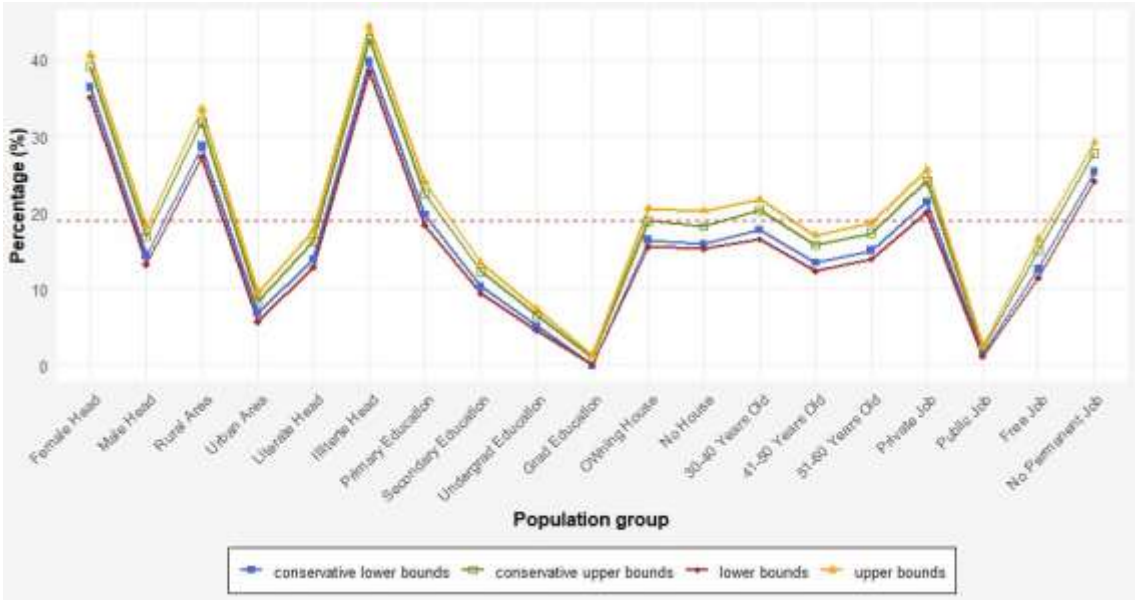


Figure 4. *Poverty persistence rate in specific subgroups, 2017-2019*

By contrast, some commonly used targeting criteria appear to be much less informative. The probability of chronic poverty is very similar for owner-occupiers and non-owners (around 19 and 18 percent, respectively), and differences across age groups of the household head are modest (less than 4 percentage points), although younger heads face somewhat greater difficulty escaping poverty. This implies that simple asset tests, such as homeownership, are a poor proxy for chronic poverty risk and that one-dimensional eligibility rules may misclassify both poor and non-poor households. A multi-dimensional perspective that combines information on education, gender, location and employment status is likely to be more effective in identifying those at highest risk of

remaining poor over time, and is therefore essential for designing targeted and impactful policy interventions.

The results on entry into poverty across subgroups, which is shown in Figure 5, closely aligned with the corresponding patterns of poverty persistence shown in Figure 4. Taken together, the evidence on entries and persistence indicates that sanctions not only trap those who are already poor but also push vulnerable non-poor households below the poverty line, reinforcing the aggregate finding that a specific set of groups has struggled to escape poverty under all three sanction regimes and that sanctions have deepened pre-existing structural inequalities rather than shifting all households proportionally.

The corresponding results on upward mobility and on households that remain non-poor in both periods are reported in Annex E.

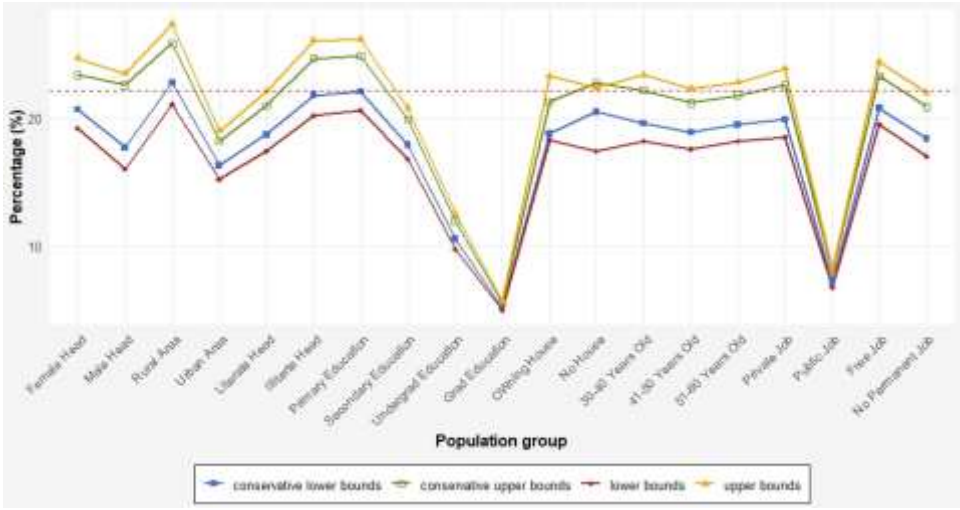


Figure 5. *Poverty entry rate in specific subgroups, 2017-2019*

4.4 Provincial poverty maps under three sanction regimes

The previous subsections showed that sanctions in the 2010s were associated with heterogeneity across household subgroups at the national level. However, given Iran's pronounced regional inequalities, national averages may conceal sharply diverging provincial poverty paths. Using synthetic-panel estimates of chronic and transitory poverty for three key sanction episodes in Iran's so-called "lost decade" of the 2010s, we map the geography of poverty dynamics across the country's 31 provinces; in Figure 6 (comprising sub-figures 1-1 through 1-6) we present these

results, distinguishing between chronic and transient poverty for all provinces.

Across all three periods, a clear and persistent core–periphery pattern emerges. Provinces in the east and south-east, notably Sistan and Baluchestan, Kerman, Golestan, Hormozgan, North Khorasan and Lorestan, consistently record some of the highest rates of chronic poverty. In 2010–2013, chronic poverty in Sistan and Baluchestan is already very high at about 36.8 per cent of households, with Golestan, Hormozgan, Kerman and Lorestan all in the 13–20 per cent range. By 2014–2016, deprivation in the periphery intensifies further: chronic poverty reaches around 32.8 per cent in Kerman and remains above 20 per cent in Sistan and Baluchestan and Golestan. The picture worsens dramatically in 2017–2019, at the height of the "maximum pressure" episode.

In that interval, more than half of households in Kerman (51.8 percent) and nearly one-half in Sistan and Baluchestan (48.7 percent) are estimated to be chronically poor, while provinces such as Semnan, North Khorasan, West Azerbaijan, Lorestan and Ilam record rates in the 22–28 per cent range. By contrast, more central and economically diversified provinces such as Tehran, Alborz, Isfahan, Yazd, Mazandaran, and Qazvin, consistently occupy the lower end of the chronic-poverty distribution. For example, even in 2017–2019, the estimated chronic-poverty rate in Tehran is around 11 percent and in Alborz about 12.7 percent, substantially below the figures observed in the eastern and south-eastern periphery. These large metropolitan and industrial centers host a disproportionate share of political and economic elites with preferential access to rents and public resources, and their more diversified economic structure appears to shield a greater fraction of households from sliding into long-term poverty when national conditions deteriorate.

These spatial patterns interact closely with the timing and intensity of successive sanctions episodes. The fact that the same peripheral provinces remain at the top of the chronic-poverty ranking in all three episodes, and that their situation worsens most sharply during 2017–2019, strongly suggests that external shocks from sanctions hit already disadvantaged provinces disproportionately hard, rather than affecting all regions in the same way.

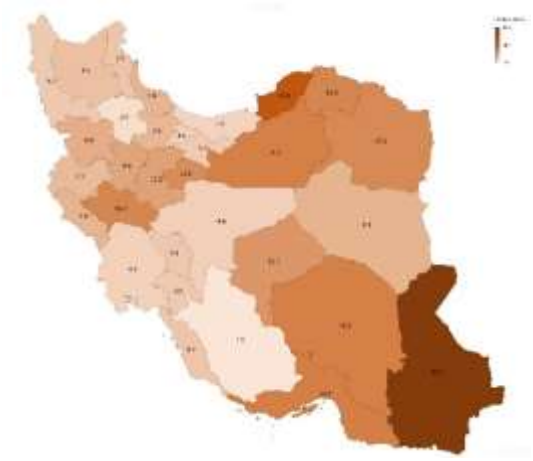


Fig. 6A *Chronic Poverty Map, 2010–2013*

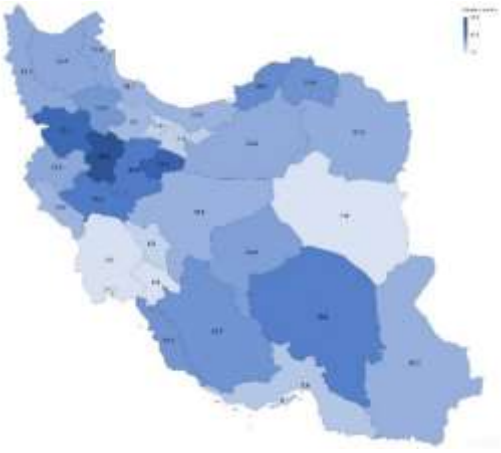


Fig. 6B *Transitory Poverty Map, 2010–2013*

Fig. 6C *Chronic Poverty Map, 2014–2016*

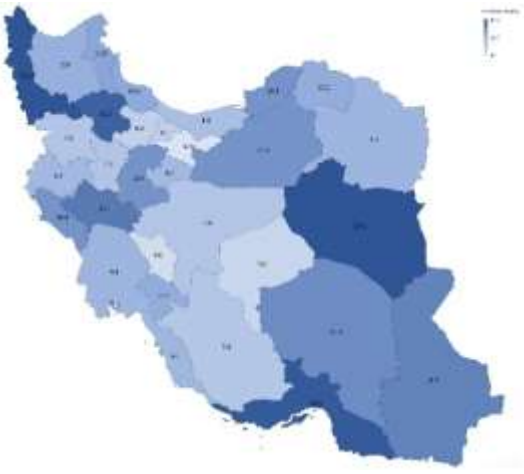


Fig. 6D *Transitory Poverty Map, 2014–2016*

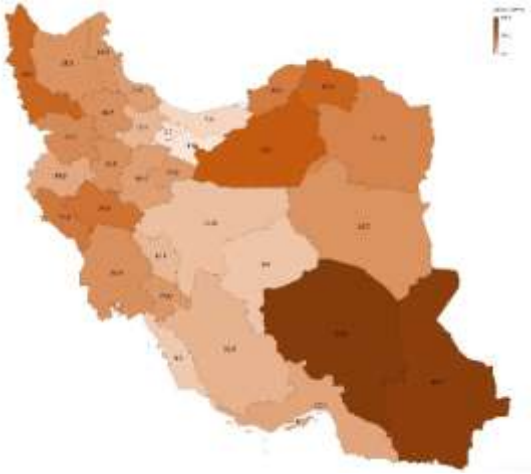


Fig. 6E *Chronic Poverty Map, 2017–2019*

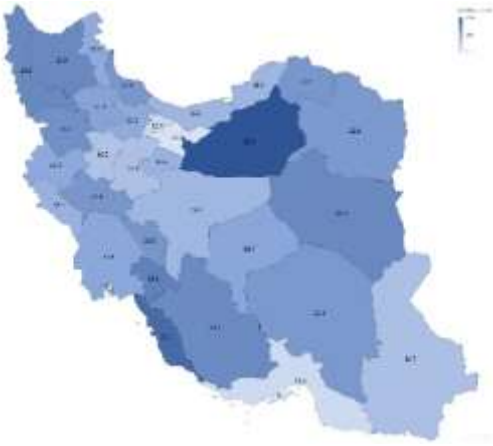


Fig. 6F *Transitory Poverty Map, 2017–2019*

Figure 6. ***Iran's Poverty Dynamic Maps***

Provinces with weaker productive bases, heavier reliance on low-productivity agriculture and informal cross-border trade, and historically under-provided infrastructure appear least able to absorb exchange-rate collapse, inflation, and cuts in public spending. From an institutional and political-economy perspective, these patterns are consistent with theories of extractive institutions and limited access orders. In Acemoglu and Robinson's (2012) framework, countries may be governed by broadly extractive institutions, but the intensity of extraction and exclusion is not uniform: core regions that are closely connected to political and economic elites tend to receive a disproportionate share of public investment and protection, while peripheral areas suffer from chronic under provision of infrastructure, weak public services and a persistent reliance on low-productivity activities. North, Wallis and Weingast (2009) similarly characterize "limited access orders" as systems in which ruling coalitions control entry to lucrative markets and political positions and concentrate rents among insiders, leaving marginal regions with little access to state resources or effective social protection. Viewed through this lens, Iran as a whole can be seen as a limited access order, but historically marginalised provinces bear the brunt of its extractive and exclusionary features. When sanctions tighten and the overall pool of rents and fiscal resources shrinks, core provinces with stronger ties to the governing coalition are better able to shield themselves, whereas peripheral provinces, already disadvantaged in terms of infrastructure, productive capacity and access to formal markets, experience disproportionately large welfare losses.

The maps of transitory poverty reveal a somewhat different, yet complementary, geography of disparities. Under the Obama-era sanctions regime, high rates of entry into poverty are observed not only in some poorer provinces (e.g. Kurdistan, Lorestan, and Kerman) but also in provinces such as Hamedan and Qom, where a sizeable share of households cross the poverty line in at least one of the two years. By 2014–2016, the pattern shifts: South Khorasan, West Azerbaijan, Hormozgan, and Zanjan record some of the highest transitory-poverty rates, with markedly elevated probabilities of falling into poverty. In 2017–2019, at the height of "maximum pressure", Semnan, Kohgiluyeh and Boyer-Ahmad, and the two Azerbaijan provinces feature prominently among the provinces with the highest transitory-poverty rates, whereas Tehran, Alborz, and a small group of central provinces display much lower entry probabilities. Strikingly, even Bushehr, a province that had previously ranked relatively well in terms of chronic poverty, now records a

transitory-poverty rate above 27 percent in this third episode, suggesting that the Trump-era sanctions gradually exposed new regions to intense short-run poverty risks.

Two insights follow from comparing the chronic and transitory maps. First, some provinces,[17] are doubly disadvantaged: they combine very high chronic poverty with non-negligible transitory poverty, indicating that large shares of their population are poor in both periods and that many of the remaining households are at constant risk of falling into poverty when shocks hit. Second, other provinces, notably resource-rich or trade-exposed regions such as Bushehr, Semnan, and some smaller provinces, have more moderate chronic-poverty rates but very high transitory-poverty rates in certain episodes. In these provinces, sanctions that disrupt oil exports, external trade or public-investment projects translate into large swings in employment and incomes, pushing many previously non-poor households temporarily below the line.

5 Robustness checks

This subsection uses two complementary exercises to assess the robustness of the poverty-transition estimates that we obtained in the previous sections using the synthetic-panel method. First, we compare the synthetic-panel bounds with "true" transition probabilities calculated from an actual household panel for a sub-period (2014–2016) in which a genuine panel, but with limited data, is available. Second, we examine whether our conclusions are sensitive to the choice of poverty line by recalculating bounds over a wide range of alternative thresholds.

For the first check, we exploit the fact that the Household Income and Expenditure Survey (HIES) has a genuine panel component for a subset of households between 2014 and 2016. For roughly a few thousand households observed in both years, we can directly compute the true transition matrix (chronic poverty, movements out of poverty, movements into poverty, and remaining above the poverty line) simply by following the same families over time and averaging the observed transitions. We then pretend that panel information is not available and apply our parametric synthetic-panel estimator, using constraint 1 described in section 4.1, to the same households treated as two independent cross-sections. This yields lower and upper bound estimates for each cell of the transition matrix. Table 8 reports the true point estimates and the corresponding synthetic-panel bounds for the 2014–2016 period. In every case, the true transition probability lies

comfortably inside the estimated interval, and the bounds are relatively tight. For instance, the share of households that are poor in both years is 10.1 percent, while the parametric bounds range from 7.5 to 11.3 percent; for those remaining above the poverty line in both years, the true value of 71.8 percent is bracketed by a lower bound of 67.2 percent and an upper bound of 72.0 percent. Similar patterns hold for the two off-diagonal cells (poor–non-poor and non-poor–poor). This close alignment between the "truth" from the real panel and the intervals generated by the parametric synthetic-panel model provides validation that the method is able to recover the underlying poverty transitions in the Iranian data.

Table 8 Transitions matrices - repeated cross-sections vs. panel data Iran 2014- 2016

Status in 2014, 2016	Lower bound estimates	Truth	Upper bound estimates
Poor, poor	7.5	10.1	11.3
Poor, non-poor	8.73	8.8	12.6
Non-poor, poor	8.7	9.1	12.4
Non-poor, non-poor	67.2	71.8	72
# Observations	5652	5652	5652

Data source: Household Income and Expenditure Survey (HIES), Statistical Center of Iran.
Results are constrained to the panel sample of households whose heads are between 25 and 65 years old. Results are weighted using household-level survey-sampling weights.

The second robustness exercise addresses sensitivity to the choice of poverty line. Up to this point, the analysis has been based on a single benchmark poverty threshold. To investigate this, we recomputed the parametric bounds and the corresponding "true" transitions for a continuum of poverty lines that span the entire feasible range of base-year headcount rates from 0 to 100 percent. In practical terms, we vary the poverty line so that, in the base year, essentially no one is poor at the lower extreme and almost everyone is poor at the upper extreme, and at each step we calculate the fraction of the population that manages to escape poverty.

The resulting curve for Iran (Fig. 7) displays a pattern consistent with the broader poverty-dynamics literature. When the base-year poverty line is very low, so that only a small minority of households are classified as poor, the share of the population that can potentially escape poverty is mechanically small, and the estimated mobility out of poverty is correspondingly limited. As the line rises and more households fall below it, the fraction of households moving out of poverty increases. Beyond a certain point, however, further increases in the poverty line bring a growing

share of the population into the "poor" category, while the proportion of this enlarged group that manages to exit poverty begins to decline; once the poverty line is set so high that almost everyone is poor, virtually nobody can escape. Throughout this exercise, our parametric synthetic-panel bounds continue to bracket the true transition probabilities at each poverty line. Moreover, the gap between the upper and lower bounds is widest when roughly half of the base-year population is classified as poor, precisely where there is most uncertainty about who is just above or just below the line, and becomes much narrower at very low or very high poverty lines, where mobility out of poverty is inherently limited.

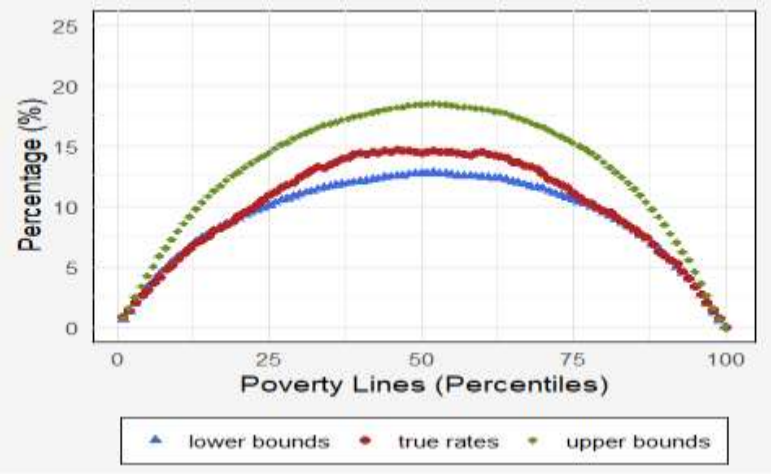


Figure 7. *Estimates of mobility out of poverty for alternative poverty lines, Iran 2014-2016*

Taken together, these two robustness checks, (i) benchmarking against actual panel data and (ii) varying the poverty line over a wide range, indicate that the parametric synthetic-panel approach delivers internally consistent and empirically credible estimates of poverty transitions.

6 Conclusion

Drawing on repeated cross-sections from Iran's Household Income and Expenditure Survey between 2010 and 2019 and a synthetic-panel approach validated against limited genuine panel data, this study documents the evolution of poverty dynamics across the escalation of sanctions in the early 2010s, the JCPOA-related respite, and the subsequent maximum-pressure campaign. The results show that chronic poverty and transitions into poverty both deepened during periods of sanctions tightening, with only modest and short-lived improvements during JCPOA relief. A large portion of households remained persistently poor across the survey waves, exits from poverty were

few, and downward mobility among initially non-poor households increased during the latter part of the decade. Taken together, these patterns suggest that the incidence of poverty traps was rising and vulnerability near the poverty line was increasing throughout Iran's "lost decade."

Further analysis reveals significant disparities in terms of the distribution and space. Households headed by females and rural inhabitants, households led by individuals with low educational attainment, and households with informal or insecure employment have higher risks of chronic poverty and downward mobility on a consistent basis than better-educated, urban, and public-sector households. In fact, at the subnational level, chronic poverty is concentrated in peripheral provinces characterized by weaker productive bases, greater dependence on low-productivity agriculture or informal cross-border trade, and historically inadequate infrastructure. Some provinces with relatively lower chronic poverty experience increased rates of transitory poverty and poverty entry during this period of maximum pressure. Sanctions interact with these structural weaknesses through mechanisms such as reduced oil revenues, restricted external finance, exchange-rate depreciation, and inflation, in a manner that tends to accentuate inequalities rather than producing a uniform national impact.

These findings have several policy implications. First, national anti-poverty strategies cannot be spatially neutral. Provinces such as Sistan and Baluchestan, Kerman, Golestan, Lorestan, and parts of the western border regions require sustained, geographically targeted interventions that go beyond short-term relief to build productive capacity, infrastructure, and human capital. By contrast, provinces characterized by high transitory but lower chronic poverty-such as Bushehr, Semnan, Kohgiluyeh and Boyer-Ahmad, and some north-western regions-would most benefit from stabilization and insurance mechanisms-for example, unemployment insurance, shock-responsive cash transfers, and countercyclical local support-designed to prevent temporary income shocks from turning into persistent deprivation.

Secondly, the coexistence of sizeable chronic poverty and frequent downward transitions implies that policy must simultaneously tackle structural traps and vulnerability near the poverty line. In the case of chronically poor households, long-term income support combined with employment programs, access to quality education, and basic services is required in order to relax binding constraints that short-lived growth episodes or partial sanctions relief have not removed. Priority for the near-poor should be given instead to instruments for smoothing consumption and protecting

assets during macro-economic shocks-such as indexed cash transfers, subsidized health insurance, and credit schemes preventing distress sales-so that temporary negative shocks do not permanently displace households into poverty.

Third, the sharper protection observed among households connected with the formal and especially the public sector underlines the role of domestic institutions in mediating external shocks. In a sanctions-constrained environment, reforms that extend social insurance and labor protection to informal and precarious workers are key to limiting the regressive impact of external restrictions. At the same time, the results emphasize the limitations of relying on aggregate growth or episodic sanctions relief to reduce poverty: without explicit redistribution and spatially targeted investment, the gains are likely to bypass the most vulnerable groups and regions.

This paper has focused on a single-country case, while the synthetic-panel methodology it employs clearly applies to other sanction-hit or shock-prone economies where only data from repeated cross-sections are available. Work taking the analysis further could be done by incorporating in an explicit way measures of provincial exposure to sanctions, combining dynamics of poverty with labor-market and health outcomes, and assessing the effectiveness of particular social protection reforms. For Iran, however, the central conclusion is clear: sanctions have coincided with a decade of deteriorating poverty dynamics and deepening spatial and socio-economic divides, and mitigating these effects requires an anti-poverty strategy that is explicitly dynamic, vulnerability-aware, and territorially differentiated.

End notes

1 - A significant body of literature has examined the effects of sanctions on the economy of the target country. More recently, Zamani (2024) estimates that the total cost of sanctions to Iran over 2011–2022 amounts to approximately US$1.2 trillion, exceeding even the economic cost of the Iran-Iraq war. Other studies have examined the effects of international sanctions on various aspects of Iran's economy, including government expenditures and revenues (Dizaji 2014; Farzanegan 2011), trade (Haidar 2017; Shirazi et al. 2016), and finance (Torbat 2005).

2 - This is particularly relevant for long-term panel datasets that follow the same units of observation for more than five to ten years. In practice, tracking households that relocate or drop out of surveys is often challenging, and such attrition can introduce substantial bias, typically resulting in an underestimation of true mobility within the population.

3 - See Deaton (1985) for the initial setup of the approach. Deaton and Paxson (1994) and Banks et al. (2001) use pseudo-panels constructed from birth and age-based cohorts to study lifecycle and long-term patterns of consumption and income.

4 - For example, West Azerbaijan's poverty rate more than tripled from 13.6% to 44% between 2011 and 2020, the largest increase nationally, whereas a few provinces (e.g., Bushehr) even saw declines.

5 - Assumption 1 may also be violated if the sampling survey is modified across different rounds. For example, changes in sampling methodology from one round to the next. There is no indication that this has been the case in the HIES data in Iran between 2010 and 2020.

6- Given that consumption shocks may persist over time or that the model contains household fixed effects, the error terms are expected to be positively correlated. Indeed, studies using household-level data provide empirical support for this assumption (Khor and Pencavel, 2006; Dang et al., 2014; Jenkins, 2011).

7 - In this paper, we use 500 repetitions (R = 500).

8 - $\gamma = \frac{\hat{\delta}_{\varepsilon_1}}{\hat{\delta}_{\varepsilon_2}}$ where $\hat{\delta}_{\varepsilon}$ stand for the standard error of residuals.

9 - To enhance the representativeness of annual estimates, data collection is distributed evenly across the twelve months of the year.

10 - The poverty lines reported by the MRC are calculated using the repeated basic needs cost approach. In this method, a reference food basket, providing 2,100 kilocalories per person, is selected. The cost of obtaining this basket for households below the poverty line is extracted from the survey data, yielding the food poverty line. The total poverty line is then estimated using the Engel coefficient, which assumes that households whose food expenditure is at the poverty line spend other items proportionally.

11 - Educational attainment is classified into five levels: illiterate, primary education, secondary education, tertiary education, and postgraduate education. Economic activity is categorized into four groups: unemployed, public-sector employee, private-sector employee, and self-employed.

12 - Education level, age, and the number of children were measured with respect to a specific reference year, reflecting the household's characteristics at that particular point in time.

13. The initial model includes the variables of place of residence, age of the household head, gender of the household head, and an education-category variable for the household head with five groups (illiterate, primary education, secondary education, university education, and higher education). In the next model, household amenities, the number of children aged 0–6 in 2017, and the number of school-age children (7–18) in 2017 are added. The final model is estimated by adding variables that examine household assets, marital status of the household head, and an economic-activity sector category with four groups (unemployed, public-sector job, private-sector job, and employer). This model is the best one because it has high explanatory power.

14. Dang et al. (2014) note that, in the absence of panel data, ρ can be approximated using cohort-based methods, but Herault et al. (2019) show that such estimates are highly sensitive to cohort choice. For this reason, the present study relies on panel-based estimates of ρ and treats the (0.2, 0.6) and (0.3, 0.5) as cautious, empirically grounded assumptions.

15. The Anderson–Darling test is a statistical test of whether a given sample of data is drawn from a given probability distribution. In its basic form, the test assumes that there are no parameters to be estimated in the distribution being tested, in which case the test and its set of critical values are distribution-free.

16. The model estimates in this table are computed under Constraint 1 of the parametric specification.

17. See Rongen et al. (2023) for Malaysia, Cruces et al. (2015) for Chile, and Dang and Dabalen. (2019) for African countries.

18. particularly Sistan and Baluchestan, Kerman, Golestan, and parts of the western periphery (e.g., Lorestan, Kurdistan, Ilam)

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Annex A. Demographic Changes in Iran from 2010 to 2020

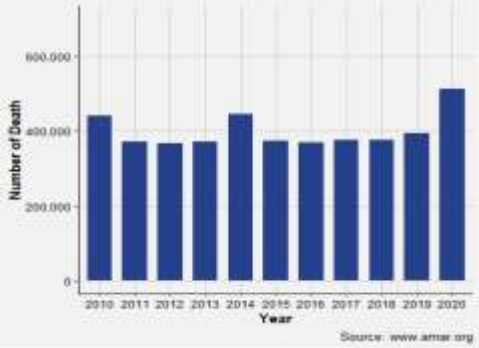


Fig A1. *Number of Births in Iran*

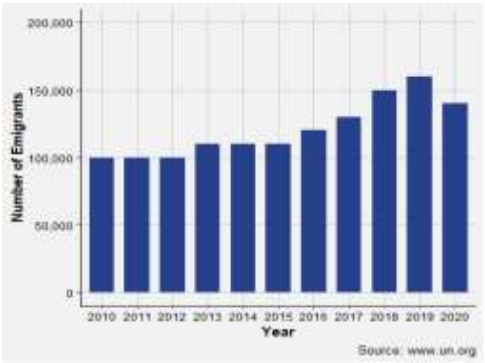


Fig A2. *Number of Deaths in Iran*

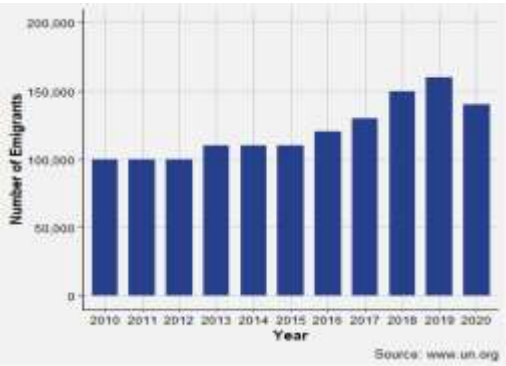


Fig A3. *Number of Emigrants from Iran*

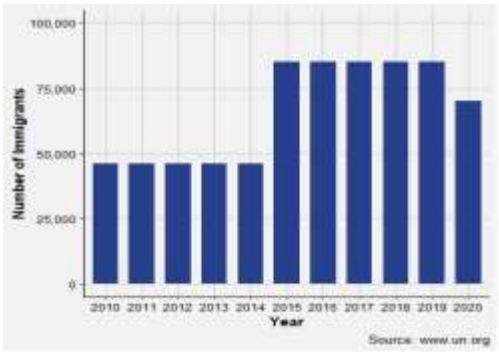


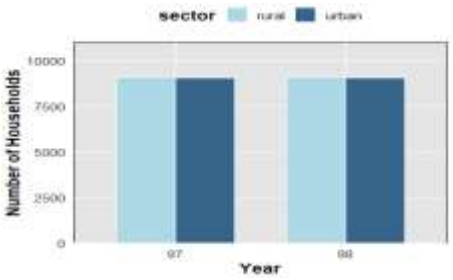
Fig A4. *Number of Immigrants to Iran*

Table A1. **Equality-of-means tests for selected HIES variables**

Variable	1396		1398		(p-value)
	Mean	Standard error	Mean	Standard error	
Age of household head	0.859	0.112	0.856	0.112	0.240
Gender of household head	51.69	0.002	51.51	0.002	0.101
Marital status of household head	0.868	0.002	0.870	0.002	0.409
Housing tenure status (e.g., owner, renter)	0.791	0.002	0.78	0.002	0.001
Place of residence (urban/rural)	0.480	0.003	0.507	0.003	0.00
Literacy status of the household head	0.727	0.003	0.755	0.003	0.00

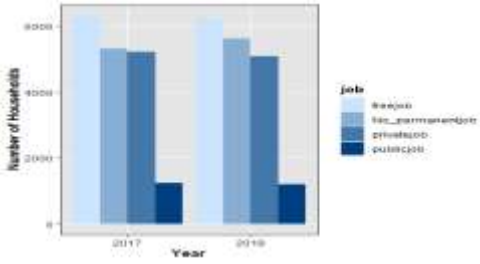
Source: Statistical Center of Iran, Household Income and Expenditure Survey (HIES) and authors' calculations.

Annex B. Comparison of the distributions of time-varying variables in 2018 and 2019

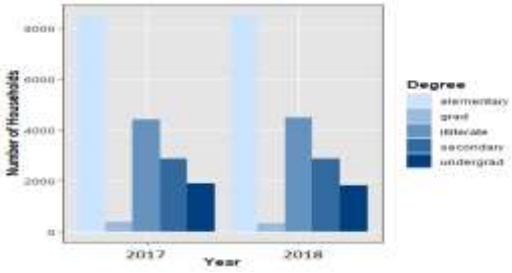


B1. *Urban–Rural Household Composition*

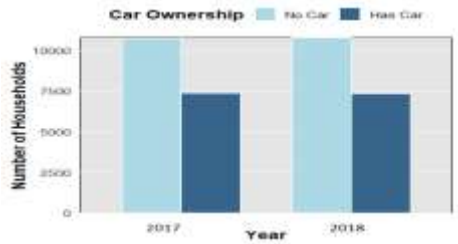
B2. *Marital Status of the Household Head*



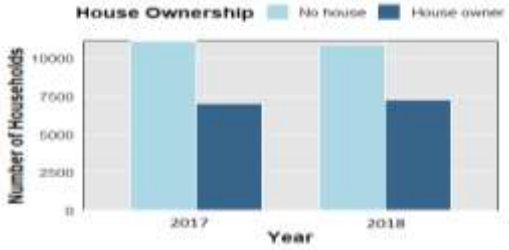
B3. *Employment Status of the Household Head*



B4. *Educational Degree of the Household Head*



B5. *Household Vehicle Ownership Status*



B6. *Household Homeownership Status*

Annex C

Table C1. Regression results based on HIES 2017 data

Variable	Model 1	Model 2	Model 3
Gender (male)	0.325***	0.307***	0.329
	(0.013)	(0.028)	(0.019)
Age	0.015***	0.015***	0.778***
	(0.000)	(0.000)	(0.000)
Region of residence (urban)	0.355***	0.342***	0.319***
	(0.008)	(0.008)	(0.003)
Primary education / illiterate	0.326***	0.293***	0.0187***
	(0.033)	(0.013)	(0.001)
Secondary education / illiterate	0.599***	0.568***	0.374***
	(0.014)	(0.014)	(0.013)
University education / illiterate	0.784***	0.759***	0.491***
	(0.015)	(0.015)	(0.014)
Postgraduate education / illiterate	1.172***	1.13***	0.739***
	(0.021)	(0.021)	(0.021)
Amenities 1		-0.189***	-0.182***
		(0.007)	(0.006)
Amenities 2		0.011	0.013
		(0.007)	(0.007)
Number of young children		-0.004	-0.088***
		0.008	(0.008)
Number of adolescent children		0.088***	-0.041***
		(0.007)	0.007
Marital status			0.024***
			(0.018)
Vehicle owner			0.345***
			(0.007)
Homeowner			-0.019***
			(0.007)
Dwelling floor area (m²)			0.002***
			(0.000)
Household size			0.067***
			(0.003)
Public-sector job / unemployed			0.305***
			(0.014)
Private-sector job / unemployed			0.107***
			(0.011)
Employer / unemployed			0.171***
			(0.010)
Intercept	18.14***	18.17	18.20***
	(0.025)	(0.028)	(0.028)
# Observations	26855	26855	26855
R-squared	0.286	0.307	0.421

Source: Findings of the study based on Household Income and Expenditure Survey data
Note: $p^{*} < 0$ .1, $p^{**} < 0.05$, and $p^{***} < 0.01$

Table C2. Regression results based on HIES 2019 data

Variable	Model 1	Model 2	Model 3
Gender (male)	0.327*** (0.025)	0.317*** 0.027	0.432* (0.018)
Age	0.013*** (0.000)	0.012*** (0.012)	0.713*** (0.000)
Region of residence (urban)	0.351*** (0.008)	0.341*** (0.008)	0.330*** (0.008)
Primary education / illiterate	0. 323*** (0.013)	0.290*** (0.013)	0.209*** 0.012
Secondary education / illiterate	0.546*** (0.014)	0.516*** (0.0014)	0.367*** (0.014)
University education / illiterate	0.799*** (0.016)	0.773*** (0.015)	0.535*** (0.015)
Postgraduate education / illiterate	1.165*** 0.021	1.13*** (0.021)	0.802*** (0.021)
Amenities 1		-0.134*** (0.007)	-0.124*** (0.006)
Amenities 2		-0.034*** (0.008)	-0.022** (0.008)
Number of young children		0.011 (0.008)	-0.107*** (0.007)
Number of adolescent children		0.088*** (0.007)	-0.043*** (0.007)
Marital status			0.036* (0.017)
Vehicle owner			0.325*** (0.007)
Homeowner			-0.010 (0.007)
Dwelling floor area			0.002*** (0.000)
Household size			0.078*** (0.003)
Public-sector job / unemployed			0.249*** (0.014)
Private-sector job / unemployed			0.058*** (0.011)
Employer / unemployed			0.119*** (0.010)
Intercept	18.10*** (0.025)	18.14*** (0.028)	18.10*** (0.028)
# Observations	28079	28079	28079
R-squared	0.271	0.284	0.421

Source: Findings of the study based on Household Income and Expenditure Survey data
Note: $p^{*} < 0.1$, $p^{**} < 0.05$, and $p^{***} < 0.01$

Annex D

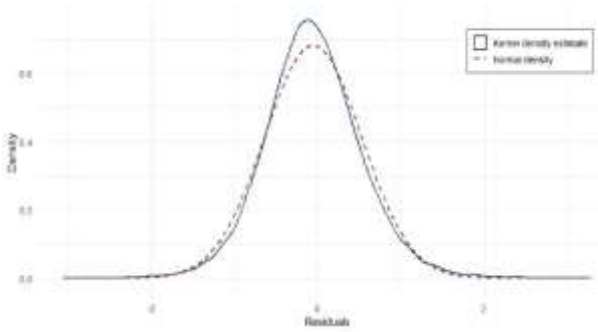


Figure D1. *Distribution of regression error terms for 2019 vs. the standard normal distribution*

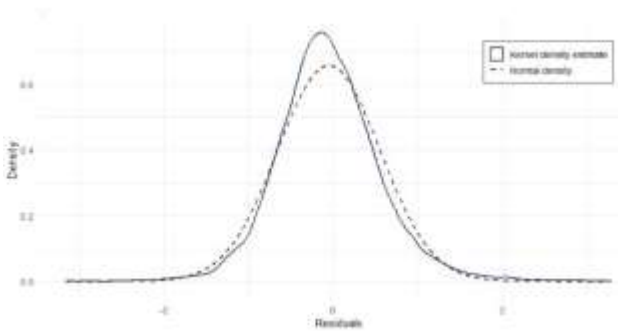


Figure D2. *Distribution of regression error terms for 2017 vs. the standard normal distribution*

Table D1. Income mobility for simulated data

Status in 2014, 2016	Simulated lower bound			Simulated upper bound		
	Nonparametric	Constraint 1	Constraint 2	Constraint 1	Constraint 2	Nonparametric
Poor-poor	28.5	21.8	29.5	17.2	18.1	16.8
Poor-non-poor	0	6.5	7.8	11.3	10.2	17.7
Non-poor-poor	13.6	17.8	19.1	22.6	21.5	29.2
Non-poor-poor	57.7	53.7	52.4	49	50.1	43.8
# Observations	28079	28079	28079	28079	28079	28079

Source: Statistical Center of Iran, Household Income and Expenditure Survey (HIES) and authors' calculations.

Annex E

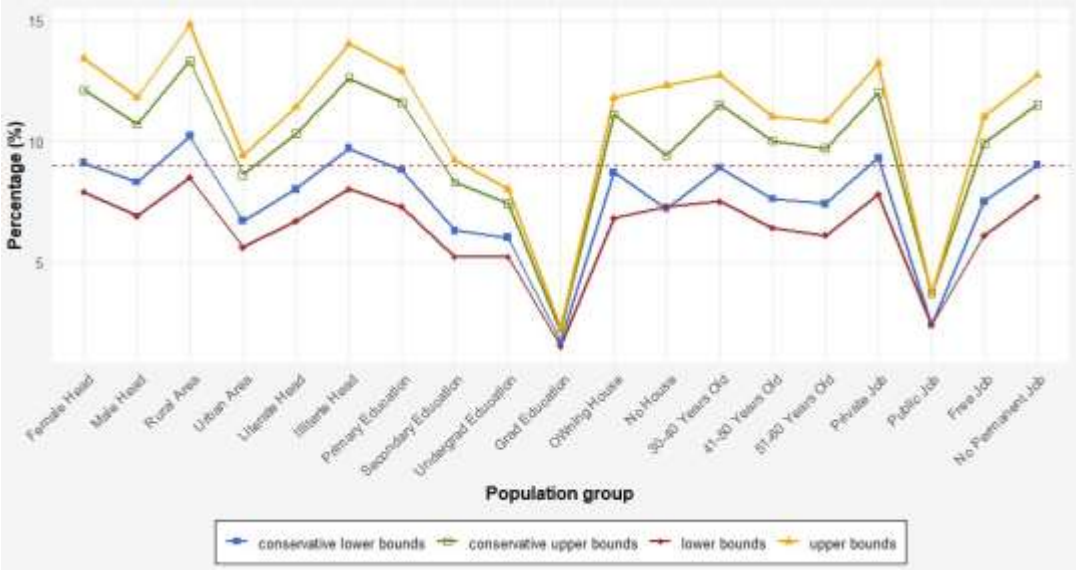


Figure E1. *poverty exit rate in specific subgroups, 2017-2019*

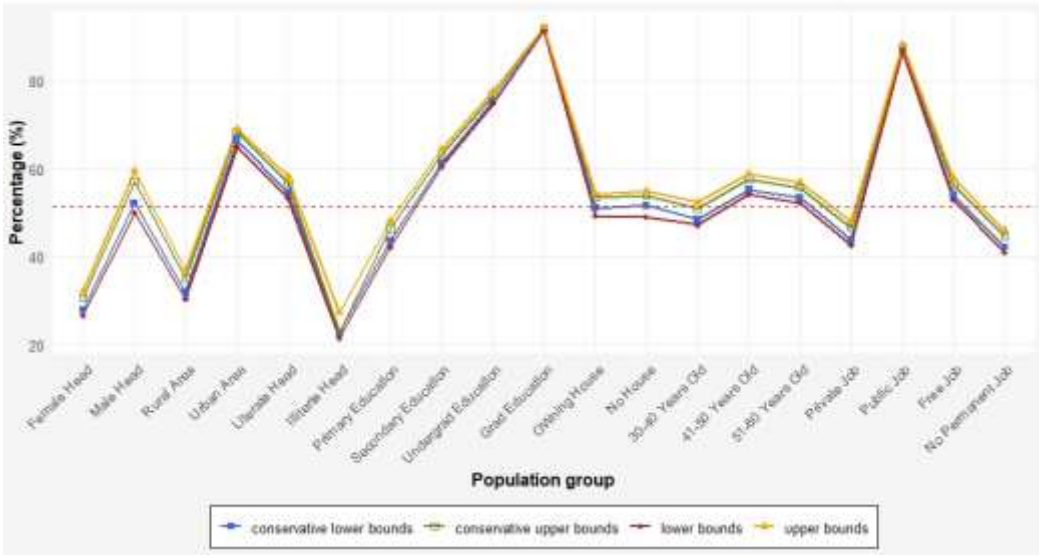


Figure E2. *Persistence rate in the non-poor stratum, 2017-2019*