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RESEARCH ARTICLE

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The Contribution of Women's Education and Health to Economic Development: Macroeconomic Transmission Mechanisms in Türkiye

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Abstract

This study examines the relationships between women's education and health, gender inequality, and economic development in Türkiye using annual data for 1990–2023. Given the limited number of observations, it employs an ARDL approach and evaluates the proposed nonlinear relationship between income and inequality through turning points, confidence intervals, the observed data range, and stability tests. The findings indicate persistent gender inequality and a generally negative association between income growth and the Gender Inequality Index across much of the observed development trajectory. However, the cubic income term is weak, the second turning point lies outside the observed income range, and the baseline model provides insufficient evidence of cointegration. The results therefore offer limited support for a long-run S-shaped Gender Kuznets Curve in Türkiye. Women's education remains positively associated with development, but gains in education and health do not translate proportionately into economic participation.

Keywords: Gender Inequality, Development, Gender Kuznets

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ARAŞTIRMA MAKALESİ

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Kadın Eğitimi ve Sağlığının Ekonomik Kalkınmaya Katkısı: Türkiye'de Makroekonomik Aktarım Mekanizmaları

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Öz

Bu çalışma, Türkiye'de kadınların eğitimi ve sağlığı, toplumsal cinsiyet eşitsizliği ve ekonomik kalkınma arasındaki ilişkileri 1990–2023 yıllık verileriyle incelemektedir. Sınırlı gözlem sayısı nedeniyle ARDL yaklaşımı kullanılmış; gelir ile eşitsizlik arasındaki doğrusal olmayan ilişki dönüm noktaları, güven aralıkları, gözlemlenen veri aralığı ve istikrar testleriyle değerlendirilmiştir. Bulgular, toplumsal cinsiyet eşitsizliğinin kalıcı olduğunu ve gözlemlenen kalkınma sürecinin büyük bölümünde gelir artışıyla eşitsizlik endeksi arasında negatif bir ilişki bulunduğunu göstermektedir. Ancak kübik gelir terimi güçlü değildir, ikinci dönüm noktası veri aralığının dışındadır ve temel modelde eşbütünleşme için yeterli kanıt yoktur. Dolayısıyla Türkiye'de uzun dönemli S-biçimli bir Toplumsal Cinsiyet Kuznets Eğrisi güçlü biçimde desteklenmemektedir. Kadınların eğitimi kalkınmayla pozitif ilişkili olsa da eğitim ve sağlıktaki ilerleme ekonomik katılıma aynı ölçüde yansımamaktadır.

Anahtar Kelimeler: Toplumsal Cinsiyet Eşitsizliği, Kalkınma, Kuznets Eğrisi

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INTRODUCTION

Women's education, health, and economic participation are central components of contemporary development analysis because they influence both individual capabilities and aggregate economic performance. Human-capital accumulation improves productivity, expands occupational choice, affects fertility and household investment decisions, and can generate intergenerational gains in health and education. When these processes are gendered, the exclusion of women from education, formal employment, or decision-making can reduce the effective stock of human capital available to an economy. For this reason, the development literature increasingly treats gender inequality not merely as a social outcome but as a macroeconomic constraint that may interact with income, structural transformation, demographic change, and institutional quality (Duflo, 2012; Kabeer, 2016; Klasen & Lamanna, 2009; Santos Silva & Klasen, 2021).

The role of female education is especially important in this debate. Cross-country and historical research has repeatedly associated improvements in women's schooling with higher labour productivity, lower fertility, stronger child human-capital investment, and improved long-run development outcomes (Barro & Lee, 2013; Knowles et al., 2002; Tembon & Fort, 2008). Yet these relationships are not mechanically identical across countries. Institutional arrangements, sectoral employment structures, unpaid care responsibilities, informality, and social norms may weaken the translation of educational progress into paid employment and income. Consequently, the macroeconomic return to female education can depend on whether labour markets, public services, and social institutions allow women to convert educational gains into economic agency (Kabeer, 2016; Seguino, 2020).

Türkiye provides a particularly informative setting for examining these relationships. Female educational attainment has improved markedly over recent decades, but women's labour-force participation remains substantially below men's participation. The manuscript's descriptive evidence shows a persistent gap in labour-market participation, employment, income, and representation even as women's schooling and life expectancy have improved. This combination creates an analytically important divergence between progress in human-capital indicators and incomplete economic inclusion. Earlier Türkiye-focused studies likewise emphasize that female labour-market participation is shaped not only by economic development but also by household structure, urbanization, social norms, and employment opportunities (Gündüz-Hoşgör & Smits, 2008; Tansel, 2001).

A second motivation concerns the possibility that the relationship between economic development and gender inequality is nonlinear. Standard development narratives often imply that rising income gradually improves education, health, and labour-market opportunities. However, the empirical literature does not support a universally monotonic trajectory. (Goldin, 1994) documents a U-shaped relationship between economic development and female labour-force participation, while (Eastin & Prakash, 2013) propose a gender Kuznets curve in which the income–equality relationship can evolve nonlinearly across stages of development. These perspectives are highly relevant to emerging economies, where structural

transformation may initially displace women from traditional sectors before later expansion of education and service-sector employment creates new opportunities.

At the same time, nonlinear terminology must be used cautiously. A U-shaped, inverted-U, or S-shaped relationship cannot be established solely from the signs of polynomial coefficients. Economically meaningful nonlinearities require that implied turning points be calculated, that those turning points fall within the observed support of the data, and that statistical uncertainty around the turning points be reported. This issue is especially important in a small annual sample because higher-order polynomials can create severe multicollinearity and unstable coefficient magnitudes. Accordingly, the revised empirical design treats nonlinear patterns as hypotheses to be tested rather than as relationships assumed in advance.

A third motivation concerns measurement. Composite gender-development indicators summarize several dimensions of women's welfare, but they also embed conceptual choices about weighting and aggregation. The UNDP Gender Inequality Index (GII), for example, combines reproductive health, empowerment, and labour-market participation. This makes it informative for broad gender disadvantage, but also means that empirical interpretations must distinguish movements in the composite index from direct measurement of any single mechanism. Critiques of UNDP gender indices emphasize issues of conceptual overlap, absolute versus relative deprivation, and the sensitivity of rankings to aggregation choices (Dijkstra & Hanmer, 2000; Klasen, 2017; Permanyer, 2013). These concerns are incorporated explicitly into the revised interpretation of the models.

The empirical challenge is therefore to study the Turkish experience without asking more of the data than the annual sample can support. The original version of the manuscript combined several nonlinear ARDL specifications with DCC-GARCH models. Review feedback correctly noted that such a design was overly ambitious for approximately three decades of annual observations, particularly when cubic income terms and multiple lags sharply reduce effective degrees of freedom. In response, the revised framework prioritizes parsimony, transparent variable construction, formal nonlinear diagnostics, and cautious interpretation. DCC-GARCH is no longer treated as the core identification device; the central emphasis is on a more tightly specified ARDL framework appropriate to annual macroeconomic data.

This revised strategy also changes the language of inference. ARDL bounds testing can identify whether a stable level relationship is supported by the data when regressors are a mixture of I(0) and I(1), but it does not by itself establish causal direction. This limitation is substantive in the gender-development nexus because development can expand women's opportunities while gender inequality can simultaneously affect human-capital accumulation, fertility, labour supply, and productivity. The literature therefore describes a potentially bidirectional process rather than a simple one-way causal chain (Duflo, 2012; Ferrant, 2015; Santos Silva & Klasen, 2021). The revised paper consequently uses terms such as "association," "relationship," and "long-run co-movement" unless the underlying empirical test warrants stronger language.

Against this background, the study investigates two closely related questions. First, how are women's educational attainment, female labour-force participation, and capital formation associated with economic development in Türkiye over time? Second, does the relationship between income and gender inequality display a statistically and economically meaningful nonlinear pattern once small-sample constraints, polynomial multicollinearity, turning points, confidence intervals, and model stability are explicitly addressed? These questions integrate the human-capital and gender-development literatures while maintaining a narrower empirical focus than the original specification.

The contribution of the study is threefold. First, it provides a Türkiye-specific time-series assessment linking female education and women's economic position to macroeconomic development, a setting in which national time-series evidence remains more limited than cross-country evidence. Second, it evaluates nonlinear income–gender relationships with explicit turning-point and confidence-band diagnostics rather than inferring curve shapes from coefficient signs alone. Third, the revised analysis places unusually strong emphasis on specification discipline in a small annual sample: polynomial terms are correctly constructed after logging income, mean-centring is used to reduce nonessential multicollinearity, lag structures are kept parsimonious, and results are interpreted as associations rather than causal estimates. In this sense, the contribution is not simply to add another nonlinear model, but to test how much of the previously claimed nonlinearity survives a reviewer-driven robustness exercise.

The remainder of the paper is organized as follows. The next section reviews the literature on female education, labour-force participation, gender inequality, composite gender indicators, and nonlinear development patterns. The data and methodology section then define the variables and revised ARDL strategy in detail, including the construction of polynomial income terms and the procedures for turning-point and stability analysis. The empirical section reports the results and associated robustness checks. The final section discusses the implications while distinguishing clearly between what is supported directly by the estimates and what remains an interpretation informed by the broader literature.

LITERATURE REVIEW

Human-capital theory provides the most direct starting point for understanding why female education may be related to economic development. Education increases skills, raises the productivity of labour, facilitates technological adaptation, and improves the quality of household decision-making. (Barro & Lee, 2013) document the central role of educational attainment in comparative development, while (Knowles et al., 2002) show that gender gaps in education can act as a brake on economic performance. In this framework, the relevant question is not only whether average schooling rises, but whether women obtain comparable access to education and are able to use acquired skills productively.

The development consequences of female education operate through several channels. More educated women tend to have greater labour-market attachment, lower fertility, better

knowledge of health practices, and stronger bargaining power within households. These effects can raise the quality of child health and education, thereby generating intergenerational human-capital gains. (Tembon & Fort, 2008) emphasize that girls' education is therefore not only a schooling policy but also a development policy with demographic, health, labour-market, and empowerment effects. The long-run nature of these channels is particularly relevant for annual macroeconomic analysis because improvements in female schooling may affect income gradually rather than immediately.

Cross-country evidence has often found that gender gaps in education are associated with weaker growth. (Klasen, 2002) and (Klasen & Lamanna, 2009) argue that gender inequality in education and employment can lower the average quality of human capital, distort the allocation of talent, and reduce potential growth. (Seguino, 2000), however, shows that the gender–growth relationship can be more complicated in economies where export-oriented industrialization has historically relied on low-cost female labour. This literature cautions against treating gender equality as a single homogeneous mechanism: education, wages, labour-force participation, and sectoral employment can affect growth through different channels and at different horizons.

Evidence from individual developing countries is likewise mixed. Studies such as (Dhar & Sarker, 2024; Fatima, 2011; Gebre, 2020) illustrate that positive associations between female education and economic performance are common but not universal. The magnitude of the relationship depends on labour-market opportunities, the quality of schooling, social constraints, and the ability of women to enter productive employment. The implication for Türkiye is that educational expansion should not automatically be interpreted as proof of equivalent labour-market integration.

Female labour-force participation (FLFP) is a key transmission channel linking human capital to aggregate production, but its relationship with development is theoretically ambiguous. (Goldin, 1994) describes a U-shaped pattern: in low-income economies, women often participate in family agriculture and household production; during early industrialization, measured female participation may fall as production shifts away from family-based work; at higher levels of development, education, fertility decline, white-collar employment, and service-sector expansion can increase women's participation again.

Beyond economic and institutional factors, gender outcomes are also embedded in broader social norms and perceptions. Evidence from Türkiye indicates that attitudes towards gender-related issues may vary across educational groups, further suggesting that improvements in female human capital operate within a wider social and cultural environment (Tuncer et al., 2021). Recent research also reinforces the importance of interpreting human-capital accumulation within its institutional and gendered context. Using comparative evidence for Türkiye and Greece, (Papaoikonomou, 2026) finds that improvements in human capital cannot be explained by educational expenditure alone and highlights gender equality and the effectiveness of institutional delivery as important components of human-capital development. From a longer historical perspective, Foukas (2023) documents

the gradual institutional development of women's education and the teaching profession in Türkiye and neighbouring countries. Taken together, these studies underline that the expansion of educational opportunities for women is embedded in broader institutional and social transformations, an issue that remains relevant when assessing whether advances in female education are translated into economic participation.

The U-curve hypothesis has particular relevance for Türkiye because structural transformation has occurred alongside rapid urbanization, informal employment, and persistent household-care responsibilities. Tansel (2001) provides early time-series and regional evidence that Turkish female labour-force participation is strongly related to the development process, while Gündüz-Hoşgör & Smits (2008) show that marital status, education, household context, and regional conditions help explain participation differences among women. These studies suggest that income alone cannot summarize the determinants of female economic activity.

The broader institutional literature reinforces this point. (Kabeer, 2003, 2016) emphasizes that women's economic agency depends on rights, access to resources, social norms, and the distribution of unpaid care. (Duflo, 2012) similarly argues that economic development and women's empowerment reinforce one another but that development alone does not automatically eliminate gender inequality. These insights are important when interpreting a model that fails to establish a stable long-run relationship for FLFP: such a result does not itself identify "culture" as the cause, but it is consistent with a literature in which participation depends on a wider set of institutional and household mechanisms than macro income alone.

The second major strand of literature treats gender inequality as both an outcome and a potential constraint on development. Klasen, (2000, 2002) and Klasen & Lamanna (2009) show that unequal access to education and employment can reduce aggregate productivity by preventing economies from allocating talent efficiently. (Santos Silva & Klasen, 2021) synthesize theoretical models in which gender inequality influences growth through fertility, human-capital investment, labour supply, and bargaining structures. These models generally imply that macroeconomic outcomes and gender institutions are jointly determined.

Ferrant (2015) explicitly frames gender inequality and development as a vicious cycle. Low income can restrict women's access to schooling, health services, credit, and formal employment, while gender disadvantage can in turn lower long-run income through underinvestment in female human capital and inefficient labour allocation. This bidirectional interpretation is methodologically important. A time-series association between income and GII should not be read as proof that income "causes" gender equality or that GII "causes" income without a separate identification strategy.

Duflo (2012) reaches a similar conclusion from the empowerment literature: economic development can improve women's position, and women's empowerment can support development, but neither relationship is automatic. Institutional reform, political representation, legal rights, and public-service provision can alter the strength of the feedback. For Türkiye,

where progress in education and health coexists with lower female labour participation and persistent gender gaps, a bidirectional framework is more plausible than a deterministic modernization story.

Nonlinear approaches arise from the observation that structural transformation changes the economic and social position of women in stages rather than smoothly. (Eastin & Prakash, 2013) develop the idea of a gender Kuznets curve and report a nonlinear relationship between economic development and gender equality in cross-country data. Their results suggest that equality can improve at some stages of development, stagnate at others, and resume improvement at higher income levels. The study has therefore become an important reference point for empirical work using quadratic or cubic income terms.

However, the transfer of a cross-country nonlinear pattern to a single-country annual time series requires care. Polynomial coefficients are sample-specific and can be highly collinear, especially when the observed income range is narrow and the sample is small. A positive linear term, negative squared term, and positive cubic term do not by themselves prove an S-shape. The implied derivative must have real roots, the turning points must lie inside the observed income support, and their uncertainty must be assessed. Otherwise, a visually appealing curve can be driven by extrapolation beyond the observed data.

This distinction is central to the revised study. The (Eastin & Prakash, 2013) framework is treated as a theoretical and empirical benchmark, not as an outcome that the Turkish data are required to reproduce. The revised specification therefore evaluates whether nonlinear terms remain informative after centring and parsimony adjustments, and whether any implied turning points are economically meaningful for Türkiye's observed income range. In this way, the nonlinear hypothesis is falsifiable rather than imposed.

The same caution applies to the U-shaped literature on female labour-force participation. (Goldin, 1994)'s historical and cross-country evidence provides a useful development narrative, but a single-country ARDL model must still demonstrate that the relevant curvature is statistically supported within the sample. The appropriate empirical question is thus whether Türkiye exhibits nonlinearities consistent with these frameworks, not whether it must necessarily follow them.

Empirical conclusions also depend on how gender inequality is measured. (Gaye et al., 2010) introduced the Gender Inequality Index to summarise disadvantages in reproductive health, empowerment, and labour-market participation. Its multidimensional nature is useful because it captures more than wage or employment gaps, but it also means that changes in the index can reflect several underlying components at once.

This feature has generated an extensive measurement debate. (Dijkstra & Hanmer, 2000) criticize earlier UNDP gender measures for mixing absolute development with relative gender inequality, while Permanyer (2013) highlights conceptual and aggregation problems in the GII. (Klasen, 2017) likewise argues that UNDP gender indices require greater transparency and conceptual consistency. These critiques do not make the indices unusable; rather, they imply that regression results involving composite indices should be interpreted as

relationships with a summary measure of disadvantage, not as direct effects on a single social mechanism.

The same concern is even more important for female-specific HDI measures because life expectancy, education, and income-related information are direct components of the index. Regressing such a composite measure on its own constituent indicators can generate mechanically strong associations. The revised paper therefore treats these relationships cautiously and does not use them as independent evidence of a structural mechanism. This directly addresses the conceptual overlap between explanatory variables and composite dependent indicators.

Türkiye combines rapid improvements in educational attainment and health with a slower transformation of women's labour-market outcomes. Tansel (2001) documents the link between development and female labour-force participation using both time-series and cross-province evidence. Gündüz-Hoşgör & Smits (2008) show that married women's labour-market participation varies systematically with education, household characteristics, regional context, and social structure. These findings support the idea that macro income is only one part of the explanation for women's economic participation.

The Turkish case is also important because it sits at the intersection of middle-income structural transformation and persistent gendered institutions. Rising education can increase the potential supply of skilled female labour, yet occupational segregation, care burdens, informal employment, and uneven regional opportunities may prevent full utilization of that human capital. This makes Türkiye an appropriate case for testing whether the standard development patterns identified in cross-country studies are visible within one national time series.

Despite the relevance of these issues, the Türkiye literature remains more fragmented than the cross-country literature. Studies often focus separately on education, participation, fertility, or employment. Fewer studies integrate female education, gender inequality, health, and macroeconomic development within a common time-series framework while also testing whether nonlinear income relationships are supported within the observed Turkish data range. This is the empirical space in which the present study is positioned.

The ARDL bounds-testing framework developed by (Pesaran et al., 2001) is widely used when regressors may be a mixture of I(0) and I(1). The approach is attractive for annual macroeconomic applications because it estimates short-run dynamics and tests for a level relationship within a single framework. (Narayan, 2005) further emphasizes the importance of finite-sample considerations in bounds testing, while Nkoro & Uko (2016) provide a practical discussion of ARDL specification and interpretation.

Nevertheless, ARDL is not automatically immune to small-sample problems. A model with many regressors, several lags per regressor, and higher-order polynomial terms can quickly exhaust degrees of freedom. This is especially problematic in annual samples of roughly three decades, where a few additional lags may consume a large share of the usable

observations. Parsimony is therefore not merely a stylistic preference but a condition for interpretable estimation.

The revised empirical strategy adopts this principle explicitly. Rather than maximizing methodological breadth, the paper uses a smaller set of theoretically motivated regressors, restrained lag structures, centred polynomial terms, formal multicollinearity checks, and turning-point diagnostics. HAC-robust inference and stability tests are used as complements rather than substitutes for specification discipline. This strategy follows the logic of the reviewer recommendation that a tighter, better-diagnosed ARDL analysis is more informative for the available data than combining several highly parameterized models.

The decision to de-emphasize DCC-GARCH follows the same reasoning. DCC models are designed for time-varying conditional covariances and are most informative in higher-frequency settings with many observations (Engle, 2002). With annual data, crisis periods are represented by only one or a few observations, limiting the credibility of detailed volatility and correlation narratives. The revised paper therefore gives priority to the ARDL framework and uses the available annual information for questions it can more plausibly answer.

The literature reviewed above reveals four related gaps. First, much of the strongest evidence on gender inequality and development comes from cross-country panels. Such studies are valuable for identifying broad regularities, but they do not show whether the same nonlinear patterns characterize Türkiye's own historical trajectory. Second, Türkiye-focused studies have produced important evidence on female labour-force participation and social constraints, yet fewer contributions jointly connect women's education, income, health, and composite gender inequality within a disciplined macro time-series framework.

Third, nonlinear gender-development studies often infer U-, inverted-U-, or S-shaped relationships from polynomial signs without demonstrating whether the implied turning points fall inside the observed data range. This is particularly problematic in small samples. The present study directly addresses this gap by combining centred polynomial specifications with explicit turning-point calculations, confidence intervals, graphical confidence bands, and observed-support checks.

Fourth, the literature sometimes uses language that moves too quickly from long-run association to causality. Given the bidirectional mechanisms emphasized by Ferrant (2015), (Duflo, 2012), and Santos Silva & Klasen (2021), the present study deliberately avoids causal interpretation. Its contribution is instead to identify which relationships are empirically stable, which nonlinear patterns are supported by the Turkish sample, and which claims disappear when the specification is simplified.

Accordingly, the revised paper contributes a more conservative but more credible empirical account. It treats the Gender Kuznets Curve and female-development hypotheses as propositions to be tested, not assumptions to be confirmed; it recognizes the measurement limits of composite gender indices; it aligns model complexity with the effective sample size; and it distinguishes statistically supported findings from interpretations that remain

grounded in the broader literature. This reviewer-responsive design strengthens the connection between theory, data construction, econometric specification, and substantive inference.

Table 1. Selected Literature and Its Relevance to the Revised Study

Study	Context / Method	Main Insight	Relevance to This Study
Klasen (2002)	Cross-country / panel	Gender gaps in education are associated with slower development.	Human-capital channel and gender gap motivation.
Klasen & Lamanna (2009)	Cross-country panel	Education and employment inequality are linked to lost growth potential.	Connects female human capital to macro performance.
Goldin (1994)	Historical and cross-country evidence	Female labour-force participation may follow a U-shaped development path.	Benchmark for nonlinear FLFP hypothesis.
Eastin & Prakash (2013)	Cross-country mixed-effects / cubic income	Proposes a nonlinear Gender Kuznets Curve.	Benchmark for testing, not assuming, nonlinear income–GII patterns.
Duflo (2012)	Review	Development and women's empowerment are mutually reinforcing but not automatic.	Supports bidirectional and non-causal interpretation.
Ferrant (2015)	Cross-country 3SLS	Gender inequality and development can form a vicious cycle.	Highlights endogeneity and two-way relationships.
Gaye et al. (2010)	Index construction	GII combines reproductive health, empowerment, and labour market dimensions.	Clarifies what LGII measures.
Permanyer (2013)	Measurement critique	Composite gender indices involve conceptual and aggregation limitations.	Supports cautious interpretation of index regressions.
Tansel (2001)	Türkiye time series / regional	Development and FLFP are related but not mechanically.	Türkiye-specific foundation.
Gündüz-Hoşgör & Smits (2008)	Türkiye micro/regional analysis	Education, household, and regional conditions shape women's participation.	Explains why income alone may not determine FLFP.
Pesaran et al. (2001)	Time-series econometrics	ARDL bounds testing accommodates mixed I(0)/I(1) regressors.	Core econometric framework.
Narayan (2005)	Finite-sample cointegration	Small samples require careful critical values and parsimonious specification.	Supports reviewer-responsive simplification.

Note. The table summarizes the literature streams directly informing the revised hypotheses, measurement strategy, and econometric design.

DATA AND METHODOLOGY

The empirical analysis uses annual data for Türkiye covering 1990-2023, which is the longest common period for which the variables required by the retained specifications are jointly available. The revised sample period is stated consistently throughout the manuscript. The

focal nonlinear ARDL equation uses 33 effective observations over 1991-2023 after the inclusion of one lag of the dependent variable. This deliberately parsimonious structure responds to the small-sample constraint highlighted in the review process and avoids estimating a large number of dynamic parameters from approximately three decades of annual data.

The study draws on internationally comparable macroeconomic and gender-development indicators. Gross national income per capita at purchasing-power-parity prices (GNIPC) and gross fixed capital formation (GFCF) are obtained from the World Bank's World Development Indicators. Female labour-force participation is based on ILO estimates disseminated through the World Bank. Female life expectancy is drawn from United Nations population data disseminated through the World Bank. The Gender Inequality Index (GII) is taken from the UNDP framework developed to summarize gender disparities in reproductive health, empowerment and labour-market participation (Gaye et al., 2010). Female schooling indicators are based on UNESCO/UNDP series. Table 2 reports the variables retained in the revised empirical strategy and, importantly, corrects the construction of the polynomial income terms.

Table 2. Variables, Transformations and Data Sources

Variable	Symbol used in estimation	Definition / transformation	Main source
Gross national income per capita, PPP	$GNIPC$	GNI per capita at PPP; natural log used in estimation	World Bank, WDI
Centered log income	$CLGNIPC$	Mean-centered natural log of GNIPC	Derived from WDI
Squared centered log income	$CLGNIPC^2$	Square of centered log income	Derived
Cubed centered log income	$CLGNIPC^3$	Cube of centered log income	Derived
Gender Inequality Index	$LGII$	Natural logarithm of the UNDP Gender Inequality Index	UNDP
Female life expectancy	LLE_F	Natural logarithm of female life expectancy at birth	UN Population Division / World Bank
Female mean years of schooling	$LMYS_F$	Natural logarithm of female mean years of schooling	UNESCO / UNDP
Female labour-force participation	$LFPR_F$	Female labour-force participation rate, ages 15+	ILO / World Bank
Gross fixed capital formation	$LGFCF$	Natural logarithm of gross fixed capital formation	World Bank, WDI

As shown in Table 2, the nonlinear terms are formed after taking the logarithm of GNIPC. Thus, the valid polynomial terms are $[\ln(GNIPC)]^2$ and $[\ln(GNIPC)]^3$, rather than $\ln\left(GNIPC^2\right)$ and $\ln\left(GNIPC^3\right)$. The latter expressions would be exact multiples of $\ln(GNIPC)$ and would therefore be perfectly collinear. In the revised specification, log income is additionally mean-centred before constructing the polynomial terms. Centring does not change the fitted polynomial relationship, but it substantially reduces non-essential multicollinearity among the linear, squared and cubed terms.

Before estimating the dynamic models, the order of integration is assessed using the Augmented Dickey-Fuller (ADF) test (Dickey & Fuller, 1979, 1981) and the Phillips-Perron (PP) test (Phillips & Perron, 1988). The ARDL bounds-testing framework of Pesaran et al. (2001) is suitable when regressors are a mixture of I(0) and I(1), provided that none is I(2). Given the short annual sample, the empirical strategy prioritizes parsimonious dynamics rather than extensive lag structures. This is also consistent with the small-sample motivation discussed by Narayan (2005).

Table 3. Selected Unit-Root Results for Variables Retained in the Revised Analysis

Variable	ADF level (constant)	ADF first difference (constant)	PP level (constant)	PP first difference (constant)	Main classification
LGNIPC	0.6198	-5.776***	2.0285	-6.492***	I(1)
$LGNIPC^2$	0.7507	-5.679***	2.3065	-6.262***	I(1)
$LGNIPC^3$	0.8861	-5.578***	2.5981	-6.003***	I(1)
LLE_F	-1.9927	-7.285***	-2.4350	-7.450***	I(1)
$LFPR_F$	-1.3388	-6.058***	-1.3100	-6.060***	I(1)
LGII	-2.764**	-4.321***	-2.639*	-4.258***	I(0) / stationary at level in several specifications
$LMYS_F$	-0.6630	-3.435***	-0.5840	-3.338**	I(1)
LGFCF	-2.9756**	-6.064***	-2.715*	-6.075	I(0)-I(1), depending on test/specification

Notes: Three, two and one asterisks denote significance at 1%, 5% and 10%, respectively. The table reports the constant specification to keep the presentation compact; trend and no-constant specifications were also examined in the original workfile. No retained series exhibits evidence of I(2) behaviour.

Table 3 indicates a mixture of level-stationary and first-difference-stationary variables, with no evidence that any retained variable is integrated of order two. The ARDL approach is therefore admissible for the retained equations. Importantly, the bounds test is interpreted as evidence about a stable level relationship, not as a causal identification strategy. Accordingly, the revised manuscript uses the language of *association, relationship* and *dynamic adjustment* rather than causal "effects" unless the design directly supports such language. This distinction is particularly important in the gender-development literature, where development and gender inequality are plausibly jointly determined (Klasen & Lamanna, 2009; Santos Silva & Klasen, 2021).

The first retained equation links per-capita income to female human capital and capital accumulation. Its general ARDL representation is

$$LGNIPC_t = \alpha_0 + \sum_{i=1}^{p} \phi_i LGNIPC_{t-i} + \sum_{j=0}^{q_1} \beta_j LMYS_{F,t-j} + \sum_{j=0}^{q_2} \gamma_j LFPR_{F,t-j} + \sum_{j=0}^{q_3} \delta_j LGFCF_{t-j} + u_t. \quad (1)$$

The corresponding error-correction representation is

$$\Delta LGNIPC_t = \alpha_0 + \sum_i \psi_i \Delta LGNIPC_{t-i} + \sum_j \omega_j \Delta X_{t-j} + \lambda ECT_{t-1} + \varepsilon_t, \tag{2}$$

where $X_t = (LMYS_F, LFPR_F, LGFCF)'$ and $\lambda < 0$ indicates adjustment toward the estimated long-run relation. This baseline specification retains the human-capital mechanism emphasized in the literature: female education can be associated with development through productivity, fertility, labour-market participation and intergenerational human-capital channels (Barro & Lee, 2013; Klasen & Lamanna, 2009).

The principal reviewer concern relates to the earlier nonlinear gender-inequality models, which combined a cubic income polynomial with several lags in a very small annual sample. The revised analysis therefore estimates a deliberately parsimonious dynamic equation. Let

$$CLGNIPC_t = \ln(GNIPC_t) - \overline{\ln(GNIPC)}. \tag{3}$$

The focal specification is

$$LGII_t = \alpha_0 + \rho LGII_{t-1} + \beta_1 CLGNIPC_t + \beta_2 CLGNIPC_t^2 + \beta_3 CLGNIPC_t^3 + \gamma LLE_{F,t} + u_t. \tag{4}$$

This focal equation corresponds to an ARDL(1,0,0,0,0) representation. The specification is intentionally restrictive: only one lag of the dependent variable is retained, while current values of the centred income polynomial and female life expectancy enter without additional distributed lags. This approach gives priority to degrees of freedom, coefficient stability and interpretability over a mechanically richer lag structure.

The cubic term is retained in this specification only to evaluate the previously asserted S-shaped relationship in a transparent way. Following the logic of the gender Kuznets-curve literature (Eastin & Prakash, 2013), an S-shaped pattern requires more than alternating coefficient signs. Economically meaningful turning points must exist, must fall inside the observed support of GNIPC, and must be estimated with reasonable precision.

For a cubic polynomial, the slope with respect to centred log income is

$$\frac{@LGII}{@CLGNIPC} = \beta_1 + 2\beta_2 CLGNIPC + 3\beta_3 CLGNIPC^2. \tag{5}$$

The candidate turning points are the real roots of this derivative:

$$CLGNIPC^*_{1,2} = \frac{-2\beta_2 \pm \sqrt{(2\beta_2)^2 - 12\beta_3\beta_1}}{6\beta_3}. \tag{6}$$

Each turning point is converted back to the original income scale using

$$GNIPC^{*} = \exp\left(CLGNIPC^{*} + \overline{\ln\left(GNIPC\right)}\right). \quad (7)$$

Confidence intervals for the turning points are obtained from repeated simulation draws from the estimated coefficient covariance matrix. Because the bounds test in the revised parsimonious model does not provide robust evidence of cointegration, ratios of contemporaneous coefficients to $(1 - \rho)$ are reported only as *implied dynamic ratios* for diagnostic completeness; they are not interpreted as cointegrating long-run causal or equilibrium effects.

Heteroskedasticity- and autocorrelation-consistent (HAC/Newey-West) standard errors are used for the focal regression. Polynomial multicollinearity is assessed using variance-inflation factors (VIFs) before and after mean-centring. Residual diagnostics include Breusch-Godfrey LM tests for serial correlation, the Breusch-Pagan test for heteroskedasticity, the Ramsey RESET test for functional-form misspecification and the Jarque-Bera test for residual normality. Parameter stability is evaluated with a recursive CUSUM process. A formal multiple-break analysis was also attempted; however, the full dynamic Bai-Perron specification could not be estimated reliably with the available degrees of freedom. Rather than forcing a break structure from an underpowered sample, the revised manuscript reports this limitation explicitly and relies on the successful CUSUM stability diagnostic.

Finally, the DCC-GARCH component is removed from the core empirical strategy. DCC models are designed for conditional covariance dynamics and are generally much better supported by high-frequency samples than by approximately three decades of annual observations (Engle, 2002). Removing this component substantially reduces methodological breadth but increases coherence between the available data, the estimable parameters and the substantive claims.

EMPIRICAL FINDINGS

The baseline human-capital equation continues to provide evidence of a stable association between female schooling, capital accumulation and per-capita income. Table 4 summarizes the reported long- and short-run coefficients from the retained ARDL specification.

Table 4. Baseline ARDL Estimates for Per-Capita Income

Variable	Horizon	Coefficient	Standard error	Test statistic / p-value
$LMYS_F$	Long run	0.8203	0.0795	10.31 (0.0000)
	Short run	0.2252	0.1014	2.22 (0.0349)
$LFPR_F$	Long run	0.0036	0.0056	1.85 (0.0524)
$LGFCF$	Long run	0.6626	0.2187	3.03 (0.0053)
$\Delta LGFCF$	Short run	0.4245	0.0596	7.13 (0.0000)
ECT_{t-1}	Adjustment	-0.2745	0.1245	-2.21 (0.0362)

As shown in Table 4, female mean years of schooling is positively associated with per-capita income in both the reported long-run relation and the short-run dynamics. The long-run coefficient of 0.8203 is statistically strong, while the short-run coefficient is 0.2252. Female labour-force participation is positive but only marginally significant at conventional levels. Gross fixed capital formation is positive in both horizons. The error-correction coefficient is negative and statistically significant, implying that approximately 27% of a deviation from the estimated equilibrium relation is corrected within one year. These findings are consistent with the broad human-capital literature, but the revised text avoids treating the ARDL coefficients as causal estimates (Barro & Lee, 2013; Klasen & Lamanna, 2009).

Table 5. Baseline Bounds Test and Residual Diagnostics

Diagnostic	Statistic	p-value / critical value	Interpretation
Bounds F-statistic	7.3189	5% I(1) = 4.194	Evidence of a level relationship
Breusch-Godfrey serial correlation	F = 0.6607	0.5253	No evidence of residual serial correlation
Heteroskedasticity test	F = 0.2227	0.6404	No evidence of heteroskedasticity

Table 5 shows that the baseline equation passes the principal residual checks reported in the original analysis. The F-statistic of 7.3189 is above the reported upper critical bound at the 5% level, while the serial-correlation and heteroskedasticity tests do not reject their respective null hypotheses. Consequently, the baseline education-development relation remains a defensible component of the paper after simplification.

The revised nonlinear analysis first addresses the polynomial multicollinearity concern directly. Table 5 compares the VIFs from the original uncentered polynomial with those from the centred specification.

Table 6. Variance-Inflation Factors Before and After Centring

Term	VIF: original polynomial	VIF: centred polynomial
Linear log income	16,869,775.33	24.22
Squared log income	67,231,165.40	3.00
Cubed log income	16,759,688.73	6.60
Female life expectancy	18.38	18.38

Table 6 demonstrates why the earlier raw polynomial coefficients were unstable and unusually large. The VIFs of the uncentered linear, squared and cubed log-income terms are extreme. Mean-centring reduces the VIF of the squared term to approximately 3 and the cubed term to approximately 6.6. Although the VIF for the linear centred term remains elevated (24.22), the magnitude of the polynomial collinearity is reduced by several orders of magnitude. The VIF for female life expectancy remains 18.38, indicating that some common trending information between income and female longevity remains in the annual data. The revised interpretation therefore emphasizes direction and robustness rather than large

polynomial coefficient magnitudes. Table 7 reports the focal ARDL(1,0,0,0,0) estimates using HAC standard errors. The sample contains 33 effective observations over 1991-2023.

Table 7. Parsimonious Dynamic Gender-Inequality Equation With HAC Standard Errors

Variable	Estimate	HAC standard error	t-statistic	p-value
Intercept	3.5610	2.4279	1.4667	0.1540
$LGII_{t-1}$	0.7537	0.1175	6.4142	<0.001
$CLGNIPC$	-0.2003	0.0979	-2.0468	0.0505
$CLGNIPC^2$	-0.2462	0.1335	-1.8439	0.0762
$CLGNIPC^3$	0.2130	0.1779	1.1976	0.2415
LLE_F	-0.8688	0.5699	-1.5244	0.1390

The dominant feature in Table 7 is persistence: the lagged Gender Inequality Index has a coefficient of 0.7537 and is highly significant. The centred linear income term is negative and lies almost exactly at the 5% threshold, while the squared term is only weakly significant at the 10% level. Most importantly for the paper's previous S-shape claim, the cubic term is statistically insignificant ($p = 0.2415$). Female life expectancy retains a negative coefficient, consistent with lower gender inequality at higher levels of female longevity, but the coefficient is not statistically significant in this more restrictive specification. These results show that the earlier substantive claims were sensitive to model complexity and should be stated more cautiously.

The actual-versus-fitted series are shown in Figure 1. The fitted line tracks the broad downward movement in LGII closely; however, this high fit should not be interpreted as causal explanatory power because the dependent variable is strongly persistent and its first lag is included in the regression.

Figure 1 is consistent with the very high in-sample fit of the dynamic equation, but the economically more important question is whether the nonlinear component displays the previously claimed S-shape. Figure 2 therefore plots the estimated income-GII relation while holding the remaining regressors at their sample means, together with the 95% confidence band and the observed data.

The turning-point calculation provides the formal test requested by the reviewers. Table 8 reports the two roots implied by the fitted cubic polynomial and the simulation-based 95% confidence intervals.

Table 8. Implied Turning Points and 95% Confidence Intervals

Turning point	Point estimate (GNIPC, PPP)	95% CI lower	95% CI upper	Within observed GNIPC range?
Turning point 1	14,382.93	72.04	18,020.46	Yes
Turning point 2	55,994.93	13,365.88	21,199,878.65	No

The observed GNIPC range in the estimation sample is approximately 12,584-34,507 PPP dollars. Table 7 shows that the first implied turning point, around 14,383, lies within this sup-

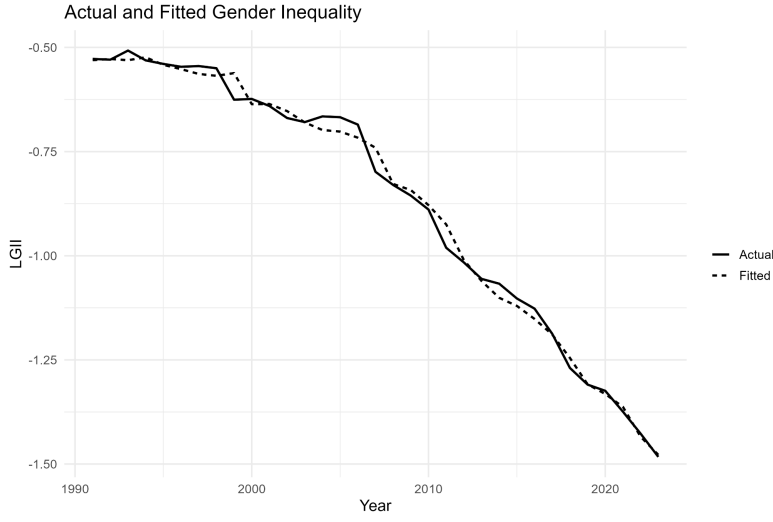


Figure 1. Actual and fitted Gender Inequality Index in the parsimonious dynamic model. The dashed fitted series follows the broad decline in observed LGII, while short-lived deviations remain around several crisis and adjustment years

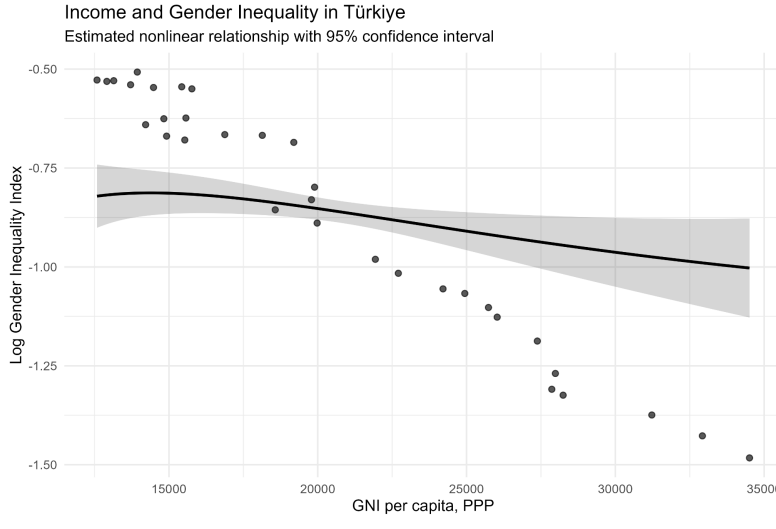


Figure 2. Estimated nonlinear association between GNI per capita and gender inequality. The solid line is the fitted nonlinear relation from the parsimonious specification; the shaded area is the 95% confidence interval and the points are observed annual values.

port. The second turning point, approximately 55,995, lies far outside the observed range. Its confidence interval is extremely wide, extending to values that are economically implausible for the sample. Consequently, the revised analysis does not provide robust evidence of an S-shaped relationship in Türkiye over the observed period. This is an important substantive correction to the earlier manuscript. The cubic specification is retained as a transparent diagnostic of the previous claim, but its second reversal is not empirically supported.

This conclusion also places the cross-country findings of (Eastin & Prakash, 2013) in appropriate context. Their gender Kuznets-curve result is a useful theoretical benchmark, but a nonlinear pattern established in a broad cross-country design need not reproduce itself within one country's relatively short annual time series. For Türkiye, the supported portion of the fitted curve is better described as a predominantly negative income-inequality association after a low-income turning region, rather than a complete S-shaped trajectory.

The central distinction in the revised interpretation is between a well-fitting dynamic regression and a statistically supported cointegrating long-run relation. Table 9 reports the relevant tests.

Table 9. Bounds Test and Diagnostics for the Parsimonious Gender-Inequality Equation

Test	Statistic	p-value	Revised interpretation
ARDL bounds F-test	3.3548	0.1281	No robust evidence of cointegration in the parsimonious specification
Breusch-Godfrey LM(1)	0.1198	0.7292	No residual serial correlation
Breusch-Godfrey LM(2)	0.1509	0.9273	No residual serial correlation
Breusch-Pagan	9.6243	0.0866	Heteroskedasticity not rejected at 5%
Ramsey RESET	0.5324	0.5937	No evidence of major functional-form misspecification
Jarque-Bera	7.2621	0.0265	Residual normality rejected

Table 9 shows that the package-reported bounds test does not support a level relationship at conventional significance levels ($F = 3.3548$, $p = 0.1281$). Given the very small sample, this result is treated conservatively rather than as definitive proof that no long-run relation can exist. It does, however, mean that the revised manuscript should not report the polynomial coefficients as cointegrating long-run effects. The dynamic regression itself performs well on serial correlation, heteroskedasticity and functional-form diagnostics. Residual normality is rejected, which provides an additional reason to rely on HAC inference and to avoid overinterpreting small differences in marginal significance.

Figure 3 presents the recursive CUSUM process. The empirical fluctuation process remains within the 5% critical boundaries throughout the estimation interval, providing no evidence of systematic parameter instability in the parsimonious equation.

The residual sequence is shown in Figure 4. Several relatively large one-year residuals appear around the late 1990s, the global financial-crisis period and the early 2010s, but there is no prolonged cluster of residuals with one sign. This visual evidence is consistent with the Breusch-Godfrey results and the CUSUM test.

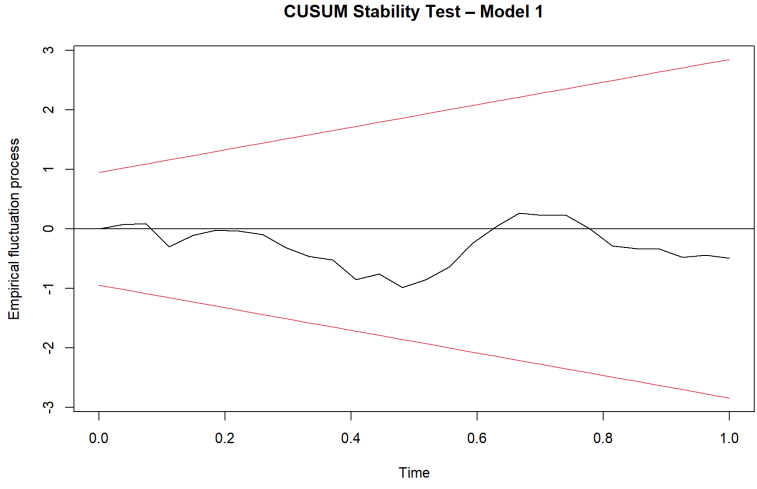


Figure 3. CUSUM stability test for the parsimonious gender-inequality equation. The empirical fluctuation process remains within the critical boundaries throughout the sample.

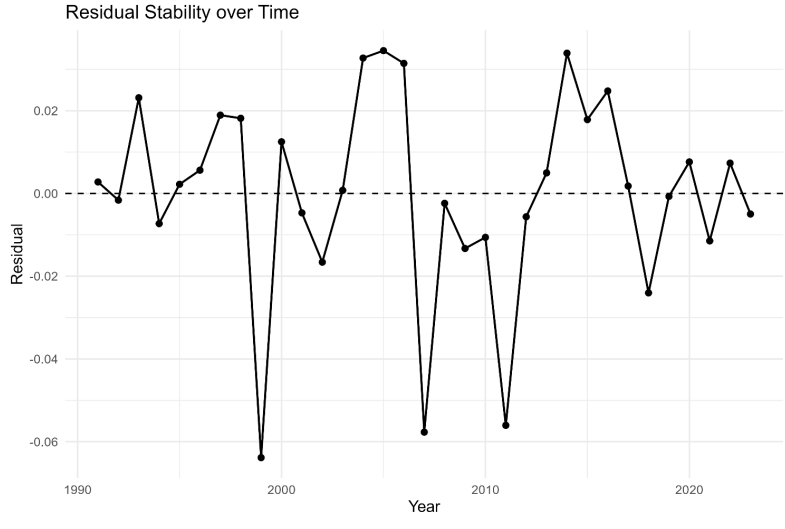


Figure 4. Residual stability over time. Residuals fluctuate around zero, with several isolated annual deviations but no sustained drift.

CONCLUSION AND POLICY IMPLICATIONS

This study examined the relationship between women's human capital, gender inequality, and economic development in Türkiye, with particular emphasis on the roles of female education, health, labour-market participation, and the possibility of nonlinear income–gender dynamics. The analysis was motivated by an important feature of Türkiye's development experience: substantial improvements in women's educational attainment and health have occurred alongside persistent gaps in labour-force participation, employment, income, and economic representation. This combination makes Türkiye particularly relevant for evaluating whether improvements in conventional development indicators are automatically translated into greater gender equality and economic inclusion.

The revised empirical analysis also responds directly to an important methodological issue arising from the available data. With only approximately three decades of annual observations, highly parameterized dynamic models can generate unstable coefficients, consume substantial degrees of freedom, and create apparently strong nonlinear patterns that may not survive more parsimonious specifications. The revised analysis therefore places greater emphasis on specification discipline. The ARDL structure is simplified, polynomial income variables are correctly defined as powers of logged income, mean-centring is employed to substantially reduce polynomial multicollinearity, and nonlinear relationships are evaluated not merely from coefficient signs but through explicit turning-point calculations, confidence intervals, observed-support checks, and graphical evidence.

The resulting evidence provides a more nuanced picture than the original specification. Gender inequality exhibits considerable persistence over time, while rising income is generally associated with declining gender inequality across much of Türkiye's observed development range. Nevertheless, the evidence does not provide robust support for a stable S-shaped Gender Kuznets Curve. In the parsimonious specification, the cubic income coefficient loses statistical significance. More importantly, although the first implied turning point lies within the observed income range, the second occurs far beyond the income levels observed in the sample and is surrounded by substantial statistical uncertainty. The nonlinear confidence-band figure similarly indicates a predominantly declining relationship over the empirically relevant income range rather than a clearly identified S-shaped trajectory.

This finding has an important substantive implication. Türkiye's experience should not be interpreted as mechanically following a predetermined sequence in which gender inequality necessarily rises and falls at particular stages of economic development. The results instead suggest that economic development has generally coincided with improvements in gender equality, but the shape and strength of that association are conditional rather than deterministic. Consequently, the Gender Kuznets Curve is better regarded as a hypothesis against which Türkiye's historical experience can be evaluated than as a universal law of development.

The revised bounds-testing results reinforce the need for caution. The parsimonious central specification does not provide sufficiently strong evidence of cointegration. Accord-

ingly, the estimated income–gender relationship should not be interpreted as establishing a structural long-run equilibrium, and the study does not claim that income growth causally reduces gender inequality. This distinction is especially important because the theoretical literature points to bidirectional relationships: economic development may expand women's opportunities, while improvements in women's education, health, labour-force participation, and economic agency may themselves support development. Without a separate causal identification strategy, the present results are therefore best interpreted as dynamic macroeconomic associations. At the same time, the diagnostic evidence provides reassurance regarding the revised specification itself. The parsimonious model displays no meaningful evidence of residual serial correlation or functional-form misspecification, and the CUSUM statistic remains within conventional critical boundaries. Mean-centering also dramatically reduces the extreme multicollinearity generated by uncentered polynomial income terms. These improvements demonstrate why parsimony is particularly important when nonlinear hypotheses are tested with relatively short annual time series. They also suggest that some of the stronger conclusions obtained from the original, more highly parameterized specifications were sensitive to model complexity.

The findings concerning women's human capital should likewise be interpreted in this broader context. Improvements in female schooling and health represent important dimensions of Türkiye's development process, but the descriptive and econometric evidence does not imply that such improvements automatically translate into proportional gains in women's economic participation. Türkiye has achieved substantial progress in female education and longevity while female labour-force participation continues to lag behind that of men. This divergence is consistent with a broader literature emphasizing that education is a necessary but not always sufficient condition for women's economic inclusion. The economic return to female human capital also depends on labour-market opportunities, care responsibilities, occupational structures, institutional arrangements, and women's ability to convert acquired capabilities into paid and productive employment.

The revised analysis also warrants caution when composite gender-development indicators are used. Measures such as the Gender Inequality Index summarize several dimensions of disadvantage and are useful for tracking broad development patterns, but their multidimensional construction complicates structural interpretation. Female-specific human-development indicators present an even more direct concern when education, longevity, or income variables used as regressors are themselves components of the dependent index. Strong associations in such models may partly reflect mechanical overlap rather than an independent economic transmission mechanism. For this reason, the revised study does not treat these relationships as separate causal evidence and places greater weight on models whose dependent and explanatory variables are conceptually distinct.

The policy message emerging from these findings is therefore broader than simply increasing women's years of schooling. Türkiye has already demonstrated that female educational attainment and health outcomes can improve substantially during economic development. The remaining challenge is to strengthen the institutional and labour-market

mechanisms through which these capabilities are converted into economic participation, employment quality, earnings, and decision-making power.

Education policy should consequently be complemented by policies that reduce barriers to women's entry into and retention in the labour market. Affordable and accessible childcare, stronger reconciliation of work and family responsibilities, measures supporting formal employment, reduced occupational segregation, and improved school-to-work transitions for educated women are particularly relevant. Policies encouraging female participation in high-productivity, technology-intensive, professional, and managerial occupations may also increase the macroeconomic return to investments already made in women's education.

Health policy remains an integral part of this strategy. Improvements in women's health and longevity expand capabilities and productive potential, but their economic benefits are more likely to materialize when accompanied by opportunities for sustained participation in education, employment, entrepreneurship, and economic decision-making. Human-capital policy should therefore be conceived as an integrated framework linking education and health investments with labour-market institutions rather than as a collection of isolated interventions.

More generally, the results caution against assuming that aggregate income growth will automatically eliminate gender inequality. Economic development can create resources and opportunities conducive to greater equality, but the conversion of those opportunities into actual economic inclusion depends on institutional and social conditions. Policies explicitly addressing barriers to women's participation may therefore be required even during periods of sustained income growth.

Several limitations should frame the interpretation of the results. Most importantly, the annual sample is small, limiting the complexity of dynamic specifications that can be estimated reliably. The study therefore deliberately prioritizes parsimonious models over more parameter-intensive alternatives. Second, ARDL methods provide evidence concerning dynamic and level relationships but do not by themselves identify causal effects. Third, composite gender indicators inevitably involve measurement and aggregation choices. Finally, institutional and social mechanisms—including unpaid care, household bargaining, occupational segregation, informality, and gender norms—cannot be directly identified from the aggregate annual variables used here.

Future research could address these limitations using longer datasets as they become available, provincial or regional panel data for Türkiye, micro-level labour-force and household surveys, or identification strategies capable of separating causal mechanisms from macroeconomic co-movement. Regional evidence would be particularly valuable because national averages may conceal substantial differences in women's educational opportunities, employment structures, and social constraints across Türkiye. Future work could also distinguish more explicitly among education quantity, education quality, field of study, formal employment, wages, and occupational mobility.

The revised evidence leads to a more restrained but empirically stronger conclusion. Türkiye's development experience is associated with substantial improvements in women's human capital and a broad decline in measured gender inequality, but the data do not robustly establish an S-shaped long-run Gender Kuznets Curve or a simple causal pathway from economic growth to gender equality. The central policy implication is therefore not that development alone will resolve gender disparities, but that the benefits of women's education and health are most likely to be realized when economic development is accompanied by institutions and labour-market arrangements that enable women to convert human-capital gains into effective and sustained economic participation.

Conflict of Interest

The authors declare that there is no conflict of interest regarding the publication of this study.

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Ethical Approval

This study is based exclusively on publicly available aggregate secondary data and does not involve human participants, personal data, interviews, surveys, or experimental procedures. Therefore, ethics committee approval was not required.

Author Contribution

The author is solely responsible for the conceptualization, methodology, data collection, formal analysis, interpretation of the findings, and writing and revision of the manuscript. The author contribution rate is 100%.

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Declaration of Generative AI and AI-Assisted Technologies

During the preparation and revision of this manuscript, the authors used generative artificial intelligence (AI) tools solely to improve the grammar, linguistic clarity, readability, and academic expression of the text. AI-assisted tools were also used to support language editing, sentence restructuring, and the presentation of the manuscript in a clearer and more concise academic style. The conceptual framework, research design, data selection, econometric methodology, empirical analyses, interpretation of the findings, and conclusions remain the intellectual responsibility of the authors. AI-generated content was critically reviewed, verified, and revised by the authors before inclusion in the manuscript.

Generative AI was not used as a substitute for scientific judgment, data analysis, or authorship, and no AI tool is listed as an author. The authors take full responsibility for the accuracy, integrity, and final content of the manuscript.

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