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Power at Home, Power at the Polls:
Gender Turnout Gap in Argentine Presidential Elections[1]

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Abstract

This paper examines whether gender differences in the labor market explain the gender turnout gap in Argentine congressional elections. Using a unique dataset of actual voting records disaggregated by gender, derived from Argentina's historical system of gender-segregated polling stations, in use until 2007, we construct a provincial panel covering 23 districts and eight biennial congressional elections from 1993 to 2007. Unlike the overwhelming majority of studies in this literature, our data come not from surveys but from official electoral records, which avoids the well-known biases associated with self-reported turnout. The key explanatory variable is women's income as a share of total household income, a continuous measure of female economic autonomy at the provincial level. Estimated using two-way fixed effects with province and year effects and clustered standard errors and confirmed by an instrumental variables strategy that addresses endogeneity, we find that a higher female income share is robustly and positively associated with a larger female turnout advantage. The result is consistent with the resource and autonomy model of political participation: economic independence raises women's capacity and motivation to engage electorally.

Keywords: Gender turnout gap, Women's earnings, Compulsory voting, Argentina

JEL Classification: D72, J16, P16

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I. Introduction

Since women received the right to vote in most Western democracies during the first half of the twentieth century, a voluminous literature has documented, debated, and sought to explain why women and men participate in elections at different rates. The direction of the gap has reversed over time: whereas men historically turned out more than women, in most contemporary democracies, women now participate at rates equal to or exceeding those of men (Desposato and Norrander 2009; Dåssonneville and Kostelka 2021).

The dominant explanations for gender differences in turnout fall into two broad categories. The first, associated with the resource model of political participation (Verba et al., 1995; Wolfinger and Rosenstone 1980), holds that participation requires time, money, and civic skills, the type of resources that were historically unequally distributed across genders. As women's educational attainment, labor force participation, and economic independence have converged toward those of men, their political participation rates have converged as well. The second, associated with the opportunity cost model, holds that the cost of voting, particularly the time cost, is increasing in income, so that higher-earning individuals are more likely to evade the voting obligation when it is compulsory, or to abstain when it is not. These two models generate opposite predictions under voluntary voting but interact in nuanced ways under compulsory voting institutions.

This paper contributes to the turnout gender gap literature with two distinctive features. First, we exploit a unique dataset of gender-disaggregated actual voting records from Argentine congressional elections, derived from the country's historical practice of maintaining separate polling stations and electoral rolls for men and women. This practice, in use from the first election in which women participated (1951) until 2007, generates administrative data on turnout by gender at the provincial level that are free from the biases inherent in survey-based measures of political participation. Actual voting records provide a qualitatively different and superior source of evidence compared with self-reported turnout data, which are subject to social desirability bias and sample limitations. Second, we focus on the within-province, over-time relationship between women's economic autonomy, measured as women's income as a share of total household income, and the gender turnout gap in congressional elections.

Our main finding is that a higher female income share is robustly positively associated with a larger female turnout advantage. Provinces in which women contribute a greater share of household income in one period subsequently exhibit a larger female turnout advantage. This result survives an instrumental variables correction for endogeneity.

Our results support the resource and autonomy model of participation over the opportunity cost model. The degree to which women independently control household resources translates into political autonomy, raising women's capacity and motivation to comply with the compulsory voting obligation. This interpretation connects directly to Bercoff and Meloni (2023), which found that the percentage of women who are heads of household, an indicator of economic autonomy, reduces the gender gap in vote choice for the incumbent in Argentine presidential elections. The present paper extends that logic upstream: the same economic autonomy mechanism that shapes how women vote also shapes whether they vote.

The remainder of the paper is organized as follows. Section II reviews the related literature. Section III describes gender differences in voter turnout in Argentina, presenting the institutional context and descriptive evidence from the unique dataset. Section IV presents the theoretical framework and empirical specification. Section V discusses the estimation results and robustness checks. Section VI concludes.

II. Related Literature

The study of gender differences in political participation has evolved substantially since the mid-twentieth century. Early studies across Western Europe and the Americas documented consistent male turnout advantages, attributed to women's exclusion from civic and professional life, their greater religiosity and deference to male authority, and their shorter history of enfranchisement (Duverger 1955; Tingsten 1975). As these structural conditions changed, so did the gap. By the 1980s the male advantage had disappeared in the United States and most of Western Europe, and by the 1990s a female turnout advantage had emerged that has persisted and, in some cases, widened (Wolfinger and Rosenstone 1980; Verba, Schlozman, and Brady 1995).

Recent comparative work has extended this picture to developing countries and to institutional contexts beyond the voluntary voting systems predominant in the earlier literature. Desposato and Norrander (2009) analyze seventeen Latin American countries finding that gender gaps in participation in the region are shaped by socioeconomic and cultural factors similar to those documented in developed countries, though economic development per se provides no stimulus for equalization. Córdova and Rangel (2017) find that compulsory voting significantly reduces gender gaps in turnout by providing women with opportunities and incentives to engage with the electoral process that they would not otherwise receive. Cox and Morales (2022, 2023), working with administrative census data from Chile, highlight the methodological superiority of actual voting records over survey data for studying gender gaps in turnout, noting that survey-based estimates are systematically distorted by differential overreporting across genders.

The measurement advantage of administrative records is central to our study. Most existing research on gender turnout gaps relies on opinion surveys. As Givens (2004) and Karp and Brockington (2005) note, survey respondents may give socially desirable answers, sample sizes are often insufficient to estimate gender-specific participation rates precisely at subnational levels, and samples may include individuals who do not actually vote. Argentina's historical system of gender-segregated polling stations, that was in use from 1951 to 2007, generates administrative data on turnout by gender at the provincial level that are free from these limitations. Lewis (1971) was the first to exploit these records for the period 1958–1965, noting that female turnout already exceeded male turnout in that early period, contrary to the common assumption of the time. Our paper extends that tradition to the congressional elections of 1993-2007.

The connection between labor market position and political participation is well established in the resource model tradition. Verba et al. (1995) show that employment exposes workers to politically relevant civic skills that raise participation. Brady et al. (1995) decompose the participation gap into resource, engagement, and recruitment components, finding resources to be the dominant factor.

For gender differences, the labor market channel operates through several reinforcing mechanisms. First, labor force participation directly raises women's access to the resources that underpin participation such as income, civic skills, and information networks. Cebula and Alexander (2017) find, using a state-level panel for five US presidential elections, that a one percentage point increase in the female labor force participation rate is associated with a 0.61% increase in overall voter turnout, consistent with the resource mechanism. Second, labor market position shapes women's economic autonomy within the household, which in turn affects their capacity for independent political decision-making. Storm (2014), using two British cohort studies with detailed household income data, finds that women vote according to their own income rather than their husband's income only when they are the primary earner in the household or work full-time. Eisenberg and Ketcham (2004), analyzing US presidential elections at county level for 1992–2000, find that the gender gap in economic voting is driven largely by the differential proportion of women who are heads of single-parent households meaning that women facing analogous economic incentives to men vote analogously.

This economic autonomy hypothesis is directly tested and confirmed in Bercoff and Meloni (2023), which uses Argentina's gender-segregated actual voting records for five presidential elections from 1989 to 2007 to show that the gender gap in vote choice for the incumbent narrows as the proportion of women who are heads of household increases. That paper establishes that labor market incentives, specifically the economic independence associated with head-of-household status, are the key driver of the gender gap in how Argentine women vote. The present paper asks whether those same incentives determine whether they vote at all.

The labor market channel is particularly salient in the Latin American context, where female economic participation has expanded rapidly but unevenly across countries and subnational units, generating cross sectional and temporal variation in the degree to which women's labor market integration translates into political engagement. Espinal and Zhao (2015), in a comprehensive review of civic and political participation across the region, find that female labor force participation is among the strongest predictors of women's political engagement at both the country and province level, operating through the resource and civic skills mechanisms identified in the broader literature. Morgan (2015) extends this finding by showing that women's access to formal employment changes not only their civic capacities but their material interest and political orientation, driving them towards greater engagement with economic policy outcomes and, ultimately, with electoral choices. Formally employed women develop stronger economic stakes in the decisions of incumbents, greater exposure to workplace civic skills, and more autonomous political identities, with raise the probability of compliance with electoral obligations. These findings are related to the present paper's key variable: the female income share that captures the degree to which women are economically integrated into the formal economy and the extent to which they independently control household resources, both of which the resources and autonomy model predict to be positively associated with the electoral participation.

Under mandatory voting the relevant margin of behavior shifts from the participation decision to the compliance decision: voters choose whether to comply with the legal obligation or evade it. According to Carreras (2018), there are notable gender differences in the perception of civic duty regarding voting, which lead to asymmetric compliance. Women demonstrate a stronger sense of duty to vote compared to men across various national contexts. This asymmetry serves as a direct mechanism by which compulsory voting can result in a higher turnout among women, despite the legal obligation being equally applicable to both genders.

Our results are consistent with this framework but extend it by showing that the size of the female compliance advantage varies systematically with the female income share. Provinces where women command a larger share of household income exhibit larger female turnout advantages, suggesting that economic autonomy amplifies the female compliance advantage rather than substituting for it. This finding is novel in the literature on compulsory voting and gender, which has focused primarily on cross-national variation in the existence of mandatory voting laws rather than on within-country variation in the economic conditions that modulate its gender effects.

III. Gender turnout differences in Argentina

Argentina is a middle-income developing country organized as a federal republic with 24 districts, the Autonomous City of Buenos Aires (known as CABA, its acronym in Spanish) which is the national capital, plus 23 provinces. Voting has been compulsory for citizens between 18 and 70 years of age since 1912, with sanctions for non-compliance that include fines and, historically, restrictions on access to public employment. In practice, enforcement has been uneven and the monetary fines relatively modest, so compulsory voting should be understood as raising the cost of non-compliance rather than eliminating discretionary evasion (Meloni, 2026a, 2026b). Although the democratic tradition of Argentina dates back to the nineteenth century with regular elections for the President and provincial governors beginning in 1862, it was not until 1916 that the first presidential election was held using a secret ballot system. During the twentieth century, the constitutional order was disrupted on six occasions, only to be reinstated following each military regime. The last military rule ended in 1983, after seven years of cruel dictatorship.

The Chamber of Deputies is elected by proportional representation from multi-member provincial districts, with each province constituting a single electoral district. Deputies serve four-year terms and the Chamber renews by halves every two years, so congressional elections are held every two years. In alternating cycles, congressional elections coincide with presidential elections (concurrent elections) or occur without a concurrent presidential race (midterm elections). Our sample of eight elections from 1993 to 2007 includes both types, a feature we exploit to test whether electoral salience affects the gender turnout gap, defined throughout the paper as female minus male turnout.

Argentina maintained separate polling stations and electoral rolls for men and women from the first election in which women participated—the 1951 presidential election—until the 2007 elections. Mixed rolls were introduced in the 2011 presidential and congressional elections. During this period, the electoral authority (*Dirección Nacional Electoral*) recorded turnout separately for male and female voters at each polling station. This practice generates administrative panel data on gender-disaggregated turnout at the provincial level that are unique in the Latin American context and rare globally. Only a handful of studies exploit similarly structured actual voting records: Panzer and Paredes (1991) and Cox and Morales (2022, 2023) for Chile, Lewis (2004) for Santiago de Chile, and Köppl-Turyna (2020) for Vienna. Our dataset for Argentine congressional elections extends this tradition.

The period under study covers some of the most dramatic macroeconomic episodes in Argentine history, such as the Tequila crises of 1994-95, the convertibility recession of 1999-2001, the catastrophic collapse of 2001-2002 and the recovery from 2003 onward. These economic shocks impacted male and female markets in diverse manners and at different speeds, creating significant within-province variations in the female income share across two-year intervals. Male employment is concentrated in tradable and construction sectors, which are sensitive to the business cycle, while female employment is more concentrated in services and the public sector which is more stable abut also more affected by fiscal adjustment. The gender composition of income therefore shifts meaningfully across biennial election cycles in ways that are not trivial.

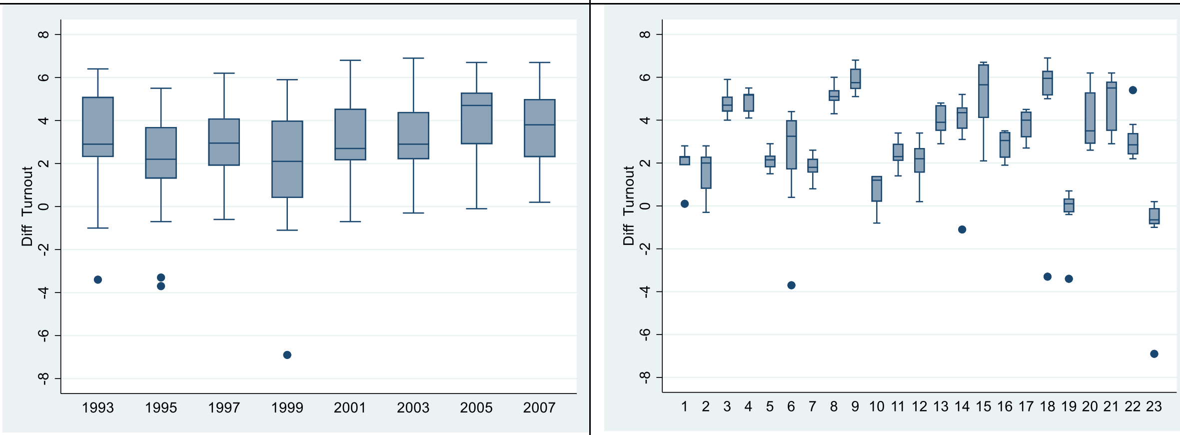
In most of the election analyzed, women turn out to vote at higher rates than men in Argentine congressional elections. Panels A and B of figure 1 display the gender turnout gap across the eight representatives' elections for each of the 23 subnational districts, as box plots. In every election year from 1993 to 2007, the provincial average gender turnout gap is positive, confirming that the female turnout advantage documented by Lewis (1971) for the 1958–1965 period has persisted through the democratic era.

The female turnout advantage is not constant over time. It narrows noticeably in the elections of 1995 and 1999, the years coinciding with the Tequila crisis aftermath and the onset of the deep recession that culminated in the 2001 political and economic crisis. This pattern is consistent with the resource model: macroeconomic shocks that depress female income and employment compress the female turnout advantage[2]. Conversely, the advantage widens in 2003 and 2007, the post-crisis recovery years in which female labor force participation and income shares began recovering. This temporal pattern motivates our empirical hypothesis that provincial variation in women's economic autonomy drives cross-sectional and time-series variation in the female turnout gap.

There is also substantial cross-provincial variation in the female turnout advantage that ranges from near zero in some interior provinces to over three percentage points in more urbanized, higher-income provinces. This cross-sectional variation in the gap correlates with cross-sectional variation in female income shares and labor force participation rates, providing preliminary motivating evidence for our empirical strategy.

Figure 1. **Turnout in congressional elections, 1993-2007 (%)**

Panel A. Gender Turnout Gap by election year	Panel B. Gender Turnout Gap by district



Source: Own calculations based on *Dirección Nacional Electoral*

Notes: Panels A and B: Gender Turnout Gap defined as female minus men turnout. Panel B (X-axis): 1. Buenos Aires, 2. Catamarca, 3. Chaco, 4. Chubut, 5. Córdoba, 6. Corrientes, 7. Entre Ríos, 8. Formosa, 9. Jujuy, 10. La Pampa, 11. La Rioja, 12. Mendoza, 13. Misiones, 14. Neuquén, 15. Río Negro, 16. Salta, 17. San Juan, 18. San Luis, 19. Santa Cruz, 20. Santa Fe, 21. Santiago del Estero, 22. Tierra del Fuego, 23. Tucumán, 24. CABA.

It is important to note what this female turnout advantage does and does not imply in the context of compulsory voting. Under voluntary voting, a female turnout advantage would reflect women's greater intrinsic motivation to participate. Under compulsory voting, it reflects differential compliance with a legal obligation. The relevant question is not why women choose to participate more but why men are more likely to evade compliance. Our empirical

[2] Only in 15 out of 163 elections the gender turnout gap is negative (men outnumber women).

framework addresses this question by relating the provincial gender turnout gap to provincial measures of gender labor market inequality, controlling for province and year fixed effects.

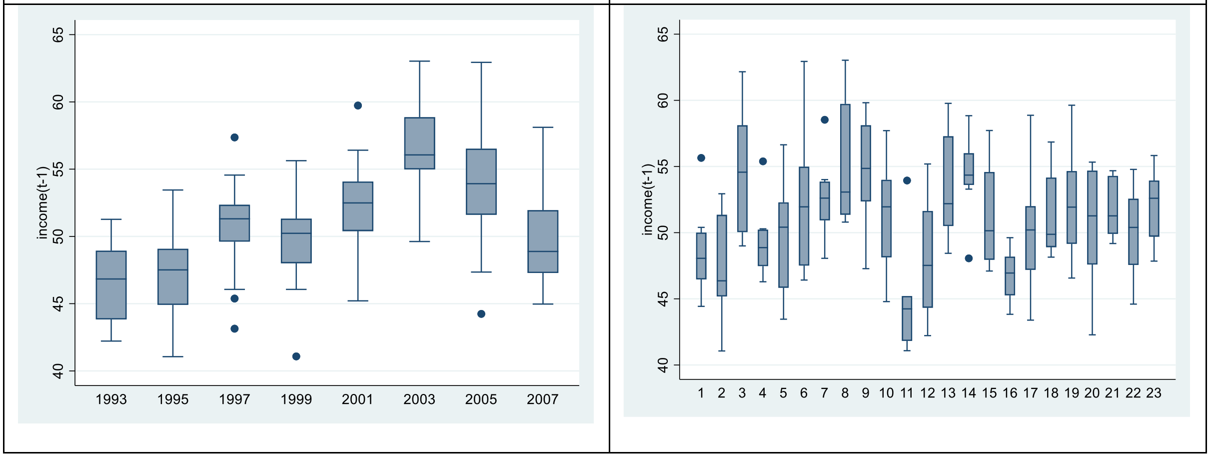
Figure 2 displays the distribution of women's labor income as a share of total household labor income across the sample period and across provinces. Panel A reveals three substantively important features of the temporal evolution of income. First, the variable exhibits a pronounced upward trend over most of the sample period: the provincial median income share rises from 47% in 1993 to a peak of 57% in 2003, reflecting the secular increase in female formal employment and the gradual convergence of male and female wage rates that characterized The Argentine labor market over this period. Second, this trend is sharply reversed after 2003: the median declines to 54% in 2005 and falls further around 51% in 2007, suggesting that the post-crisis economic recovery disproportionally benefited male relative to female employment. Third, the interquartile range widens considerably during the 2003-2005 elections, indicating that the post-crisis recovery was uneven across provinces and generated greater cross-provincial dispersion in women's relative income position that was observed during the stable convertibility period of the 1990s. The lower values and extreme observations below 45% in the late 1990s and early 2000s, particularly preceding the 2001 and 2003 elections, correspond to provinces where the severe macroeconomic downturn and the 2001-2002 crisis had a particularly heavy toll on female labor force participation and wage income. Taken together, Panel A confirms that our variable of interest, *income*, exhibits meaningful temporal variation within the sample, which is a necessary condition for the within-province identification strategy employed in section V.

Panel B shows equally significant cross-provincial heterogeneity in income share. Districts in the Northwest and Northeast regions, particularly Catamarca (2) and Formosa (8), exhibit some of the lowest median income shares, consistent with traditional household structures, lower female formal employment rates, and greater dependence on informal or public-sector activities. By contrast, provinces such as La Rioja (11) and San Luis (18) stand out with high median female income shares and remarkably tight interquartile ranges, indicating not only elevated but also relatively uniform female economic participation across electoral cycles. The capital city, CABA (24), similarly maintains high baseline values.

Several provinces including Neuquén and Santa Cruz display wide interquartile ranges reflecting the high internal heterogeneity of resources reach Patagonian provinces whose labor markets are segmented between capital-intensive extractive industries which are male dominated and the broader service economy. This cross-sectional heterogeneity in *income* across provinces is absorbed entirely by the province fixed effects in our empirical specification, so our identification of the impact of lagged income on gender turnout gap relies predominantly on the within-province temporal variation shown in Panel A, not on the structural differences across provinces visible in Panel B.

Figure 2. **Women's labor income as a share of total household labor income**

Panel A. Income (t-1) by election year	Panel B. Income (t-1) by district



<u>Source</u>: Own calculations based on Encuesta Permanente de Hogares (EPH), INDEC

Notes: Panels A and B: Gender Turnout Gap defined as female minus men turnout. Panel B (X-axis): 1. Buenos Aires, 2. Catamarca, 3. Chaco, 4. Chubut, 5. Córdoba, 6. Corrientes, 7. Entre Ríos, 8. Formosa, 9. Jujuy, 10. La Pampa, 11. La Rioja, 12. Mendoza, 13. Misiones, 14. Neuquén, 15. Río Negro, 16. Salta, 17. San Juan, 18. San Luis, 19. Santa Cruz, 20. Santa Fe, 21. Santiago del Estero, 22. Tierra del Fuego, 23. Tucumán, 24. CABA.

Data sources

Electoral data come from the Bureau of National Elections (Dirección Nacional Electoral) and comprise the total number of registered voters and actual votes cast, disaggregated by gender, for 23 provinces and 8 congressional elections from 1993 to 2007. The province of Río Negro is excluded due to missing socioeconomic data. Female and male turnout rates are computed as the ratio of actual votes cast to registered voters of the respective gender.

Labor market variables come from the Argentine National Household Survey (Encuesta Permanente de Hogares, EPH), conducted by the Argentine Bureau of Statistics (Instituto Nacional de Estadística y Censos -INDEC). The EPH provides provincial-level data on labor market outcomes disaggregated by gender for the urban population of each province. Our key variable, women's income as a share of total household income (coded *Income*), is computed from EPH data on average individual labor income by gender and province.

Provincial GDP estimations come from each province's activity accounts and the breakdown of Gross Value-Added computations by the Economic Commission for Latin America and the Caribbean (ECLAC). Population estimates were obtained from INDEC.

Given that these elections cover biennial periods and our labor market data are available at annual frequency, we use the value of each labor market variable in the year immediately preceding each election as our main specification (denoted with the subscript t−1). This lagged specification serves two purposes: it reflects the labor market conditions that voters experienced and that shaped their political engagement in the period leading up to the election, and it mitigates reverse causality from current electoral outcomes to current labor market conditions.

Table 1 presents descriptive statistics for the variables used in the empirical analysis.

Table 1. **Descriptive Statistics**

Variable	Obs	Mean	Std dev.	Min	Max
Turnout Gap	163	2.9859	2.2855	-6.9	6.9
Income (t-1)	177	50.9409	4.4919	41.06	63.03
GDP pc (t-1)	184	389.1276	263.946	95.28	1272.94
Unemployment (t-1)	184	11.3136	5.1083	1.2	25.5
Education Gap	184	-2.2373	1.7293	-7.1	2.94
Pop70 Gap	184	4.7414	2.0868	0.8	14.1
Concurrent	184	0.1793	0.3847	0	1
Rep candidate	184	3.9620	7.0993	0	40
Pres candidate	184	0.0489	0.2402	0	2

IV. Model and data.

We ground our empirical analysis in a resource and autonomy model of political participation under compulsory voting. The central insight of the resource model (Verba et al., 1995) is that political participation requires resources and that groups with fewer resources participate less. We adapt this framework to the compulsory voting context, where the relevant resource is not the capacity to choose to participate but the capacity to comply with the legal obligation and resist the temptation to evade it.

Consider a province with a continuum of voters. Each voter faces a legal obligation to vote, with a sanction $\boldsymbol{p}$ for non-compliance that is common to both genders. Voting requires a cost of compliance, which includes the time and effort of going to the polling station, gathering information about candidates, and making a voting decision. Let $\boldsymbol{c}$ denote the net compliance cost: the gross time and information cost of voting net of the intrinsic utility from participation

Formally, the compliance cost for a voter of gender *g* in province *i* at time *t* is:

$$c_{it}^{g} = \theta - \gamma \cdot s_{it}^{g}$$

where s_{it}^{g} is the individual's income share within the household, g can take two values F (female) and M (male), $\boldsymbol{\gamma > 0}$ captures the extent to which economic autonomy reduces the compliance cost and θ is the time cost of voting parameter (travel, queuing, etc.). As θ goes up, the compliance cost increases. Under compulsory voting, this cost is offset by the sanction for evasion, so that voters with $c_{it}^{g} > p$ evade the obligation; those with $c_{it}^{g} \leq p$ comply, either because the sanction binds or because participation is intrinsically valued (when $c_{it}^{g} \leq 0$). The parameter θ may be negative for voters with strong intrinsic civic motivation, capturing the fact that some individuals would participate even if legal obligations were absent.

The key insight is that the compliance cost $\boldsymbol{c}$ is not constant across voters. It includes an information cost component that decreases with income (higher-income individuals have greater access to information about candidates and parties, reducing the informational burden of voting) and an autonomy component that depends on the degree to which the individual controls household resources. Women who earn a larger share of household income are more economically autonomous: they make independent decisions about their time allocation, are less dependent on male household members for information and motivation about electoral choices and have stronger material stakes in economic policies that affect their personal income. These factors reduce the effective compliance cost for economically autonomous women relative to economically dependent women.

Notice that within a two-person household, income shares sum to unity, so $s^{M} = 1 - s^{F}$ therefore female compliance costs decrease in the female income share while male compliance costs increase in the female income share.

Aggregating these individual-level optimal decisions across households within a province yields the macro-level empirical prediction. As the average female income share rises, the provincial turnout rate for women increases due to reduced compliance costs. Conversely, male turnout remains largely unresponsive (or responds negatively) to shifts in female income. Aggregating across both groups, the provincial gender turnout gap expands as a direct function of the female income share:

ΔT_{it}=$T_{it}^{F} - T_{it}^{M} = f(income_{it}, X_{it}, \alpha_i, \lambda_t + \varepsilon_{it})$

where ΔT_{it} is the gender turnout gap in province i at election t, $income_{it}$ is women's income as a share of household income, X_{it} is a vector of control variables, α_i is a province fixed effect, λ_t is an election-year fixed effect, and ε_{it} is an error term. The female income share therefore affects the turnout gap through two reinforcing channels simultaneously: as women's income share rises, female compliance costs fall (∂T_{it}^{F} / ∂Income$_{it} \geq 0$) and male compliance costs rise (∂T_{it}^{M} / ∂Income$_{it} \leq 0$). Consequently, the resource and autonomy model yields an unambiguous theoretical prediction for the gender turnout gap: $\partial(\Delta T)/\partial Income > 0$. This is, a higher female income share is associated with a larger female turnout advantage.

We treat the provincial average female income share $\bar{s}_{it}^{F}$as a sufficient statistic for the distribution of individual income shares, under the assumption that an increase in the provincial average represents a first-order stochastic dominance shift in that distribution. Under this assumption, the aggregate compliance rates inherit the comparative statics of the individual model.

The female income share therefore affects the turnout gap through two reinforcing channels simultaneously: as women's income share rises, female compliance costs fall and male compliance costs rise. Both channels predict $\partial(\Delta T)/\partial s^{F} > 0$

This prediction differs from what a pure opportunity cost model would generate. Under the opportunity cost model (Becker 1965), higher income raises the shadow value of time, increasing the effective cost of spending time at the polling station. If men earn more than women on average, men face higher opportunity costs of voting, which would generate a female turnout advantage that increases with the income gap. This prediction runs in the same direction as our resource model for the dependent variable but through a different mechanism and with different implications for the coefficients on other variables. We discuss the distinction between these mechanisms in Section V in light of the estimated coefficients.

Empirical specification

Our baseline empirical specification is a two-way fixed effects model:

$$\Delta T_{ie} = \alpha_i + \lambda_e + \beta_1 \text{Income}_{i,t(e)-1} + \gamma' E_{i,t(e)-1} + \delta' Z_{ie} + \phi' P_{ie} + \varepsilon_{ie}$$

where i indexes provinces ($i = 1, \dots, 23$), e indexes congressional elections ($e \in \{1993, 1995, \dots, 2007\}$), and $t(e)$denotes the calendar year of election e. Therefore, the subscript $t(e) - 1$ refers to the calendar year immediately preceding each election. For instance, for the 1995 election the lagged variables are measured in 1994, not in 1993. The error term ε_{ie} has standard errors clustered by province.

The province fixed effects α_i absorb permanent differences across provinces in gender norms, political culture, geographic isolation, and institutional quality. Province fixed effects α_i also may account for any correlation that arises when provinces with sustained female turnout advantages also adopt pro-female labor market policies, thereby increasing female income shares.

The election fixed effects λ_e absorb national shocks common to all provinces in a given election year, including macroeconomic cycles, national political events, and any Argentina-wide shifts in gender norms. Election-year fixed effects λ_e may also capture the impact of national-level pro-female labor market policies adopted after high-female-turnout elections, such as minimum wage increases or federal anti-discrimination legislation, since these policies affect all provinces simultaneously.

The key explanatory variable is $\text{Income}_{i,t(e)-1}$, defined as women's income as a percentage of total household income in province i in the calendar year immediately preceding election e. The one-year lag serves two purposes: it reflects the labor market conditions that shaped voters' economic positions in the period leading up to the election, and it mitigates reverse causality from current electoral outcomes to current labor market conditions. The vectors $E_{i,t(e)-1}$and Z_{ie} contain economic and social controls respectively, with economic controls lagged one calendar year and social controls measured contemporaneously with the election. The vector P_{ie} contains political controls, all measured contemporaneously.

Notice that the identification of β_1 comes exclusively from within-province changes in the female income share across elections, conditional on national election-year effects.

Instrumental variables strategy

The contemporaneous female income share is potentially endogenous. Three sources of endogeneity are plausible. First, reverse causality: a larger female turnout advantage may induce politicians to adopt policies favoring women, including labor market regulations that raise female wages relative to male wages. Although our baseline specification uses the lagged income share, which addresses contemporaneous reverse causality, we also estimate specifications with the contemporaneous income share instrumented by its lagged value to address any remaining endogeneity concerns. Second, time-varying omitted variables:

provincial trends in gender norms may simultaneously drive the income share and the turnout gap through channels not captured by province and year fixed effects. Third, measurement error: provincial income data from the EPH are subject to sampling error, which creates classical attenuation bias in OLS estimates.

We instrument the contemporaneous income share $Income_{it}$ with its lagged value $Income_{i,t-1}$. The relevance condition requires that the lagged income share strongly predicts the contemporaneous income share, which is plausible given the persistence of labor market structures within provinces. The exclusion restriction requires that the lagged income share affects the current turnout gap only through the current income share, not directly. This assumption is defended on the grounds that province fixed effects absorb time-invariant factors (including persistent gender norms) and year effects absorb national norm shifts, so the remaining variation in the lagged income share is driven by idiosyncratic province-level economic fluctuations that plausibly affect the current turnout gap only through the current income channel.

We assess instrument quality using the Kleibergen-Paap rk Wald F statistic, comparing it against the Stock-Yogo 10% maximal IV size distortion critical value of 16.38. We inform the Cragg-Donald Wald F statistics for reference and conduct the endogeneity test to determine whether OLS-FE or IV-FE is the appropriate preferred specification.

V. Discussion of results

Table 2 reports the results of the two-way fixed effects estimation. The full preferred specification, including all controls (Column I), yields a positive and statistically significant coefficient on $Income_{t-1}$ of 0.109, significant at the 1% level. This is our core finding: within provinces, a one percentage point increase in women's income share is associated with a 0.109 percentage point increase in the gender turnout gap. Given that the female income share varies across provinces and elections in our sample by approximately 15 percentage points from the 25th to the 75th percentile, the interquartile range implies a predicted variation in the female turnout advantage of roughly 1.6 percentage points, which is a substantively meaningful effect relative to the mean female turnout advantage in our sample.

The coefficient on $Income_{t-1}$ is remarkably stable across specifications. Columns II through V present partial specifications that include only economic controls, only social controls, only political controls, and only the key variable without any controls, respectively. Across these specifications, the coefficient on $Income_{t-1}$ ranges from 0.071 to 0.090, all positive and significant at conventional levels (1% to 10% depending on the specification). This stability is strong evidence that the result is not driven by the inclusion or exclusion of any particular control group and is not a statistical artifact of omitted variable correlation with the controls.

To address the standard policy feedback concern, we implement Granger-causality tests and find that past turnout gaps do not Granger-cause current provincial income shares. This evidence suggests that the policy feedback mechanism is not operative in our setting

Among the control variables, the difference in incomplete secondary schooling between women and men (*Education*) is positive and significant at the 1% level in the full specification. A larger female disadvantage in secondary schooling completion is associated with a larger female turnout advantage. As discussed in Section II, this seemingly counterintuitive result is coherent under compulsory voting through the opportunity cost channel: women with incomplete secondary education are less likely to be in formal employment with real time constraints on election day and therefore face lower effective costs of compliance with the compulsory voting obligation. The GDP per capita and unemployment controls are not individually significant in the full specification, though unemployment becomes significant when year effects are excluded, suggesting that the year dummies absorb much of the macroeconomic cycle variation in unemployment.

The election-year fixed effects are jointly significant, and their pattern is substantively informative. The estimated year effects are consistently negative relative to the omitted base year (2007), indicating that the female turnout advantage was generally smaller in earlier elections than in 2007. The effects for 1999 and 1995 are the most negative, consistent with the interpretation that the *Tequila* crisis aftermath (1995) and the onset of the 1999-2002 recession compressed the female turnout advantage during those years. The 2001 midterm election, held in the midst of the political and economic crisis, shows a less negative year effect than 1999, which may reflect the extraordinary mobilization of protest voting (including blank and invalid ballots) that characterized that election and affected both genders similarly[3].

As an additional diagnostic, we apply the Wooldridge (2002) test for first-order serial correlation in panel data, as implemented by Drukker (2003). The null hypothesis specifies no first-order autocorrelation in the idiosyncratic error term. The test yields $F(1, 20) = 0.630$ ($p = 0.437$), failing to reject the null at any conventional significance level. The absence of serial

[33] Table 1A in the Appendix shows the estimated coefficients of year effects

correlation in the residuals confirms that the static two-way fixed effects specification is appropriate and that habit formation in the gender turnout gap does not represent a significant source of bias in our estimates. Province-clustered standard errors, which are robust to arbitrary within-province serial correlation by construction, therefore represent a conservative rather than necessary correction in this application.

Table 2. **Two-way fixed effects with standard errors clustered by province.**

Variables	Full Model	Only key var	Only Econ vars	Only Political vars	Only social vars	No time effects
Income (t-1)	0.1086***	0.0708*	0.0918**	0.0736*	0.0898**	0.1495***
	0.0361	0.0351	0.0408	0.0371	0.0338	0.0409
GDP pc (t-1)	-0.0018		-0.0015			-0.0007
	0.002		0.0017			0.0015
Unemployment (t-1)	-0.0834		-0.0852			-0.1102***
	0.0573		0.0581			0.0298
Education (t-1)	0.1265*				0.1143**	0.1591**
	0.0725				0.054	0.0684
Pop70 (t-1)	0.1286				0.1506	0.1572
	0.2114				0.1706	0.1676
Concurrent	0.1619			0.2488		-0.2668
	0.4624			0.4601		0.4598
Lists	-0.0089			-0.0044		-0.0043
	0.014			0.0146		0.0155
Candidate	-0.2023			-0.0777		-0.465
	0.2259			0.2117		0.3185
Constant	-0.7558	-0.0433	0.384	-0.2099	-1.5756	-3.4251*
	2.4051	1.858	2.5752	1.9697	1.9148	1.6849
R-sq within	0.2327	0.1944	0.2143	0.1982	0.2084	0.1155
R-sq between	0.0522	0.1459	0.2125	0.1670	0.0668	0.0390
R-sq overall	0.1079	0.1144	0.2008	0.1229	0.0137	

Note: Subscript $(t-1)$ denotes the calendar year preceding the election year. For example, for the 1995 election, $(t-1)$ refers to 1994.

Instrumental variables estimations

Table 3 reports the instrumental variables estimates. The instrument is strong across all specifications: the Kleibergen-Paap F statistics are 17.49 (full model), 18.17 (no political controls), and 31.69 (no economic controls), all above the Stock-Yogo 10% critical value of 16.38. The underidentification test rejects the null of underidentification at $p < 0.002$ in all specifications. These diagnostics confirm that the lagged income share is a relevant instrument for the contemporaneous income share in the within-province, within-year dimension required by the fixed effects estimator.

The endogeneity test rejects exogeneity of the income share at the 5% level in all specifications. This implies that the OLS-FE estimates are inconsistent and that the IV estimates are the appropriate causal estimates. Reassuringly, the IV and OLS-FE estimates share the same sign and direction, so the qualitative conclusion from the OLS-FE analysis is not overturned. However, the IV coefficients are approximately twice the OLS-FE coefficients: 0.254 (full model), 0.248 (no political controls), and 0.197 (no economic controls), compared with the OLS-FE estimate of 0.109. This amplification is consistent with classical measurement error in the provincial income share data creating attenuation bias in the OLS-FE estimates, which IV corrects.

Table 3. Instrumental variables with fixed effects and robust standard errors

Variables	**full model**	**No political**	**no Econ**	**no time**
Income (t)	0.2540***	0.2477***	0.1972***	0.4361***
	0.0979	0.0959	0.0714	0.1584
GDP pc (t-1)	-0.0013	-0.0013		-0.0011
	0.002	0.0021		0.0025
Unemployment (t-1)	-0.1106	-0.1097		-0.1639**
	0.0789	0.0779		0.0681
Education (t-1)	0.1003	0.111	0.1037	0.1217
	0.0808	0.0758	0.0753	0.0919
Pop70 (t-1)	0.0115	0.0765	0.0768	0.1151
	0.2026	0.1759	0.1812	0.2033
Concurrent	0.4942		0.4339	-0.1026
	0.4281		0.4403	0.5267
Lists	-0.0031		-0.0019	0.0064
	0.015		0.0147	0.0199
Candidate	0.1357		-0.01	0.3601
	0.5706		0.5558	0.8237
Under id chi-sq(1) p-val	0.0012	0.0014	0.0001	0.0020
Weak id Cragg-Wald F	22.958	23.722	30.502	18.254
Kleinberger-Paap rk -F	17.488	18.169	31.689	16.258
Endogeneity test chi-sq(1) p-val	0.0153	0.0133	0.00153	0.0027

Note: Subscript $(t-1)$ denotes the calendar year preceding the election year. For example, for the 1995 election, $(t-1)$refers to 1994.

Note that these models are exactly identified (one endogenous variable and one instrument) so the overidentification test (Hansen J statistic) has zero degrees of freedom and is not informative.

The income share coefficient remains positive and significant at 1% in all IV specifications. Among the controls, the difference in incomplete secondary schooling is positive but loses statistical significance in the IV estimates, likely reflecting the efficiency cost of IV estimation for a variable that was borderline significant in the OLS-FE. The political controls are uniformly insignificant across all specifications, consistent with the null result for these variables in the OLS-FE.

Interpretation of results: resource model versus opportunity cost model

The positive sign of the income share coefficient unambiguously supports the resource and autonomy model of participation over the opportunity cost model. Under the resource model, a higher female income share raises women's economic autonomy, which reduces compliance costs and raises the female turnout advantage. Under the opportunity cost model, a higher female income share would raise the shadow value of women's time, increasing their compliance costs and reducing the female turnout advantage—generating a negative coefficient. The consistently positive and significant coefficient on the income share across all specifications rejects the pure opportunity cost model.

This finding complements and extends Bercoff and Meloni (2023), which found that female head-of-household status reduces the absolute value of the gender gap in vote choice for the incumbent in Argentine presidential elections. That paper shows that economic autonomy shapes how women vote; the present paper shows that it also shapes whether they vote. Together, the two papers provide comprehensive evidence that labor market incentives, rather than intrinsic gender differences in values or preferences, are the fundamental driver of electoral gender gaps in Argentina.

The result also contributes to the broader literature on compulsory voting and gender. Córdova and Rangel (2017) show at the cross-national level that compulsory voting reduces gender gaps in turnout by providing structural incentives that benefit groups with lower baseline participation motivation. Our results show that, even within a compulsory voting system, the size of the remaining gender gap varies systematically with economic conditions. Economic autonomy amplifies women's compliance with the mandatory voting obligation, suggesting that the equalizing effects of compulsory voting institutions are contingent on the economic empowerment of the disadvantaged group.

Robustness checks

We conduct three additional robustness checks beyond the alternative specifications reported in Tables 2 and 3. First, we drop the 2001 election from the sample. The 2001 congressional election was held at the height of Argentina's political and economic crisis and was characterized by an unprecedented level of protest voting, including record rates of blank and invalid ballots. Excluding 2001 does not materially change the coefficient on the income share (significant at 5%), confirming that the main result is not driven by this extraordinary election.

Second, we drop the City of Buenos Aires from the sample. CABA is a structural outlier in the Argentine provincial panel: the wealthiest province, the most urbanized, with the highest female labor force participation and the highest female income share. Its undue leverage on the income share coefficient is a legitimate concern. Excluding CABA reduces the coefficient somewhat (0.089, significant at 5%) but the result survives, confirming it is not driven by the Buenos Aires outlier.

Third, we replace the gender turnout difference with the female-to-male turnout ratio as the dependent variable. This alternative gap measure is scale-invariant with respect to the overall level of turnout and therefore less sensitive to election-year level shifts in aggregate participation. The income share coefficient in this specification is positive and significant, consistent with the main result[4].

VI. Conclusions

This paper investigates whether gender differences in labor market economic autonomy explain the gender gap in voter turnout in Argentine congressional elections. We exploit a unique dataset of gender-disaggregated actual voting records derived from Argentina's historical system of separate polling stations for men and women, which was in use until 2007. This administrative data source is free from the social desirability bias and sample limitations that affect the survey-based turnout measures used in the overwhelming majority of existing studies.

Using a provincial panel of 23 Argentine districts across 8 biennial congressional elections from 1993 to 2007 and estimating two-way fixed effects models with province and election-year effects, we find that women's income as a share of total household income is robustly and positively associated with gender turnout gap. The result survives instrumental variables

[4] Table 2A in the Appendix shows the robustness checks results.

correction for endogeneity, with IV coefficients approximately twice the OLS-FE estimates, consistent with attenuation bias from measurement error being corrected by the IV strategy.

The finding supports the resource and autonomy model of political participation. Economic independence raises women's capacity and motivation to comply with the compulsory voting obligation, generating a larger female turnout advantage in provinces where women contribute more to household income. This interpretation is inconsistent with a pure opportunity cost model, which would predict that higher female income raises the time cost of voting and reduces the female turnout advantage.

Our results complement and extend Bercoff and Meloni (2023), which showed that the gender gap in vote choice for the incumbent in Argentine presidential elections narrows as women acquire head-of-household status. Together, these two papers establish that labor market incentives—not intrinsic gender differences in values, preferences, or civic motivation—are the fundamental driver of electoral gender gaps in Argentina. Economic autonomy shapes both whether women vote and how they vote, with the same mechanism operating consistently across two distinct electoral outcomes and two different dimensions of political behavior.

A secondary finding concerns the temporal pattern of year effects. The female turnout advantage narrowed significantly during the elections of 1995 and 1999, coinciding with the Tequila crisis aftermath and the onset of Argentina's deep recession. This pattern, visible in the estimated year fixed effects, is consistent with the resource model operating at the aggregate level: macroeconomic downturns that compress female income shares reduce the female compliance advantage nationally, independently of province-level variation in economic autonomy.

The policy implications are direct. Interventions that increase women's economic autonomy—including labor market anti-discrimination enforcement, childcare provision that enables female labor force participation, and programs that raise female educational attainment and labor market access—are predicted to reduce the gender turnout gap. In a compulsory voting system, this reduction takes the form of improved female compliance with the voting obligation. In voluntary voting systems, the resource model predicts that the same interventions would have even larger effects on female participation by raising both the capacity and the motivation to vote.

Several limitations of the present study suggest directions for future research. Our income share variable is measured at the provincial level for the urban population covered by the EPH, which may not fully capture gender income dynamics in rural areas of large provinces. The panel is limited to eight elections across 23 provinces, which constrains the power of tests that require substantial within-province time variation. Future work could extend the analysis to other Latin American countries with administrative gender-disaggregated turnout data, or to other elections types in Argentina, to assess the generalizability of the resource and autonomy mechanism identified here.

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Appendix

Table 1A. Estimated coefficients of year effects

Method: Two-way fixed effects with standard errors clustered by province

Variables	Full Model	Only key var	Only Econ vars	Only Political vars	Only social vars
t93	-0.2380362	-0.2260898	-0.5100594	-0.1767911	0.0979466
	0.5603757	0.3722673	0.534243	0.3905078	0.4712347
t95	-1.251614	-1.257351*	-1.343638*	-1.36807**	-1.041074
	0.7905466	0.6528436	0.6845591	0.6602555	0.71367
t97	-0.2416852	-0.570132***	-0.3568048	-0.5286252**	-0.4416693*
	0.3560827	0.2012469	0.360968	0.2009311	0.2469915
t99	-1.396708**	-1.487104**	-1.4372**	-1.495023**	-1.412639**
	0.5773058	0.5520959	0.6128068	0.5515606	0.5208085
t01	-0.1329084	-0.3880976*	-0.1650973	-0.3730101*	-0.3102177
	0.3057673	0.2068005	0.3173874	0.2153932	0.2330435
t03	-0.6346472	-1.061217***	-0.6011449	-1.05358***	-1.077408***
	0.5349389	0.2769137	0.5626089	0.2801554	0.2619099
t93	-0.2380362	-0.2260898	-0.5100594	-0.1767911	0.0979466
	0.5603757	0.3722673	0.534243	0.3905078	0.4712347

Table 2A. Robustness checks

Variable	Dep. variable: Difference women minus men turnout		Dep. variable: Ratio of women to men turnout
	Excluding election 2001	Excluding CABA	Full model
Income (t-1)	0.1145217***	0.1163066***	0.00151***
	0.0392453	0.0350225	0.000488
GDP pc (t-1)	-0.0018235	-0.0004751	-3.2E-05
	0.0022007	0.0017808	0.000027
Unemployment (t-1)	-0.0975984	-0.0939612	-0.0009
	0.0638903	0.0574981	0.000769
Education (t-1)	0.128915	0.1044755	0.001772*
	0.0769293	0.0657088	0.000944
Pop70 (t-1)	0.1335348	-0.0516676	0.000882
	0.2203364	0.1782152	0.002806
Concurrent	0.071613	0.2638618	0.001425
	0.4296678	0.4758652	0.006273
Lists	-0.011772	-0.0102222	-0.0002
	0.0139246	0.0129746	0.000196
Candidate	-0.1915589	-0.0655556	-0.00282
	0.2532524	0.6637564	0.002898
Constant	-0.8646221	-0.5728977	0.997751
	2.70241	2.299037	0.030554