Partner's Income Shock and Female Labor Supply. Evidence from the Repeal of Argentina's Convertibility Law

Abstract

Female employment is an important vector of economic development. This paper revisits the implicit assumption in the literature that in high-income developing countries with women-friendly work-related gender norms, an increase in female labor force will be explained by the combination of higher education, lower fertility, and structural change. Using data on couples in urban Argentina from 1996 to 2007, I show that in the short and medium term subsistence shapes female participation and employment at the extensive and intensive margins. More specifically, I study how women's labor supply reacts to negative income shocks affecting their partner. In order to assess the causal impact, I exploit the unexpected evolution of the economic environment triggered by the repeal of the convertibility law in January 2002 to instrument men's job loss. I find that women's probability of participating and finding a job is multiplied by 2 upon their partner's displacement. Turning to the dynamics of their labor supply, contrary to expectations, however, women do not symmetrically withdraw from the labor market once their partner finds a job. Evidence on repeated cross-sections confirms that the labor supply response persists long after the economic recovery. My findings are among the first attempts to evaluate the participation effects of temporary shocks in the medium term.

Keywords: female labor force, female employment, intra-household allocation, subsistence

JEL: D10, J22, O12, O54

1 Introduction

One of the most striking differences across developing regions relates to the participation of females in the labor force. Between 1990 and 2019 in Latin America and the Caribbean, the participation of women aged 15-64 increased sharply from 44.2 to 57.9%. Over the same period, it stagnated at a high level in Sub-Saharan Africa (63.6 to 62.8%), and at a much lower level in the Middle-East and North Africa (18.4 to 21.2%), while it decreased by 5-6 percentage points in Asia (from 71.8 to 66.2% in East Asia and the Pacific, and from 30.5 to 25.2% in South Asia).[1] As revealed by Duflo [2012], a strong body of evidence suggests that female employment matters for women's well-being and empowerment, and ultimately for many layers of development. In consequence, a large and growing literature is recording female participation in the labor force of the developing world, with the aim of capturing its diversity, and designing suitable policies.

In this paper, I show that subsistence can be an important driver for female participation in developing countries, even when they are at an advanced stage of development. Recently, the literature on female participation has put particular emphasis on the interplay between the structural transformation of an economy, on the one hand, and, on the other, the enduring gender norms against women's working, deeply rooted in preindustrial economic conditions. A number of recent surveys compile the existing empirical evidence on both aspects. Heath and Jayachandran [2017] explore the determinants of increased female employment, notably through a sectoral shift from industries to services, accompanied by policies promoting girls' education. Giuliano [2017] and Klasen [2019] document the historical determinants of contemporaneous differences in gender roles, and Jayachandran [2020] provides a comprehensive overview of the most recent studies on social norms as cultural barriers to female participation and employment. In essence, this flourishing literature suggests that, in the development process, structural change can be accompanied by a feminization of the workforce under two conditions. First, in line with the conventional U-shaped hypothesis [Goldin, 1995], the feminization of the workforce should accelerate with the advent of the service sector, that is, after a certain level of development. Then, for a given level of development, feminization should be more pronounced in regions where the gender norms against women's working are least restrictive. This important realization has triggered fundamental contributions on the malleability of gender norms [Bursztyn et al., 2020], and on the design of policies to foster female participation in regions with relatively low female employment [Field et al., 2021].

Of course, this growing strand of literature also has its blind spots. First, it essentially

1. Source: ILO [2021] – modeled ILO estimates.

focuses on the determinants of the *medium term* dynamics of women's labor force participation. Therefore, it indirectly overlooks the effect of *temporary* negative income shocks on women's choosing to work, notably the possibility that women may also enter the labor force by necessity. Women working for their subsistence would mostly live in low-income developing countries, where female attachment to labor is weak. Yet women's participation is necessary when household income falls too far to meet the most basic needs. This literature seeks to reconcile the apparent contradiction between the rapid growth of education and job opportunities for women in most developing regions, and the parallel stagnation of their labor force participation, with the hypothesis that gender norms are binding. Latin America happens to be the developing region where female participation grew most between 1990 and 2019, and where the correlation between education, fertility, and participation is highest [Klasen, 2019], suggesting that gender norms may in fact not be as stringent as in other parts of the world. As a result, the academic focus tends to shift to countries where these norms are particularly stringent. However, there is no particular reason to infer from these correlations that fertility and education are the sole drivers of female participation in Latin America.

This paper demonstrates that subsistence is an important determinant of women's participation in the labor market in an upper-middle-income economy like Argentina. First, I rely on labor force panel surveys representing urban Argentina over the period 1996-2007, and I explore the within-household correlation patterns between partners' labor outcomes, controlling for local labor market time-specific effects. Women whose partner becomes unemployed are 12% more likely than other women to enter the labor force, 8.5% more likely to be employed, and 6.4% more likely to work more than 20 hours a week. At the intensive margin, working women are 23% more likely to declare that they would like to work more, though they do not manage to increase their working time.

In the core of the paper, I set up an instrumental variable strategy to pin down the causal impact of a partner's job loss on female participation. In March 1991, the highly popular convertibility law began to peg the Argentine peso to the dollar, and successfully contained hyperinflation, stimulated growth, and also restored trust in financial institutions. When the Argentinean economy slid from a depression into a recession in 2001, jeopardizing the sustainability of this exchange regime, 'the authorities were unwilling even to consider the possibility of an exit: neither the government nor the public were prepared to take such a drastic course until it was forced upon them by events' [Daseking et al., 2005b]. Still, the unsustainable convertibility law finally had to be repealed in January 2002, after two months of unprecedented economic, social and political turmoil. I measure the unexpected decline in male employment following the sudden collapse of the convertibility regime for each of the 88 industry-occupation pairs in the data, and use it as an exogenous source of variation for the actual job loss experienced by male partners between 2001 and 2002. In support of this empirical strategy, I first document how a series of events built up to bring convertibility to an unexpected end, and I show that the differential effect of this shock on male unemployment across the different industry-occupation pairs could not be anticipated by looking at the employment trends before the collapse. I also show that because industries and occupations tend to be segregated by gender, the shocks experienced by men and by women do not correlate within couples. Using data on Argentinean couples between October 2000 and October 2002, I confirm that my measure for the shock intensity on male employment is positively and closely correlated with the probability that male partners will actually lose their job. As a main result, I find that an unanticipated shock on male employment multiplies by 2 the probability that their partner will join the workforce for at least one hour per week. Unraveling this result, I show that half of this effect is driven by the expansion of a workfare program which relaxed the constraints on labor demand.

Of course, displacement is not the only source of the negative income shocks incurred by households. In the face of inflation, real wages were also affected in different ways across industries and occupations, because the differential rise in unemployment across industries and occupations translated into a differential power to bargain over the nominal wages for workers. I empirically confirm the link between the magnitude of the employment shock and the decline in real monthly wages. I find evidence that a negative income shock on male workers also generated a positive labor supply response from their partner, concentrated on the bottom part of the pre-crisis household income distribution.

Finally, within the limits of my data, I question further the dynamics of female participation in the presence of negative shocks on male employment. An interesting and new finding is that female participation reacts to job *losses*, not job *entries*: women do not withdraw from the labor market once their partner finds a job. Furthermore, I find that the participation response is for most part contemporaneous with the shock, but may also be partly delayed to the following period. Corroborating these two results on repeated cross-sections, I estimate that the female participation response to the 2002 economic shock persisted for at least three years – and possibly even to the end of my period of observation, albeit on a smaller scale. Taken together, these results indicate that a shock coping response can have persistent effects on female participation, possibly through the channel of labor market attachment.

This paper makes a twofold contribution. First, causal estimates of the response of female workers to negative income shocks affecting their partners in developing countries are lacking. A number of empirical studies investigate the timing of partners' transitions from employment to unemployment in the US [Stephens, 2002, Juhn and Potter, 2007], as well as in Europe [Bredtmann et al., 2018, Halla et al., 2020]. Overall, they find women's employment response to male unemployment to be moderate at most. In the US, according to Stephens [2002], women would be discouraged from working, since shocks are likely to be correlated within households. Following Cullen and Gruber [2000], the generosity of social programs would crowd out intra-family insurance. Andersen et al. [2021] compare alternative coping mechanisms in response to male job loss in Denmark, and find that adjustment occurs mainly through reductions in savings and spending. Given the fundamental institutional, social, and economic differences between developed and developing countries, for instance, in terms of access to a social safety net, or to the credit market, the external validity of these results is essentially confined to the developed world. Through being set in the context of an emerging economy, this article contributes to the literature by offering an original empirical investigation of women's labor supply response to negative shocks to their partners' income. So far, the literature on female labor market responses to negative income shocks in developing countries has essentially been descriptive. Using the Indonesia Family Life panel, Frankenberg et al. [2003] document the immediate effects of the Asian crisis on the well-being of Indonesian families. They provide detailed descriptive evidence that the household labor supply increased at both the extensive and intensive margins during the Asian financial crisis, but they do not show that this response correlates with their local measure of the magnitude of the economic shock. Studying the Mexican peso crisis with a pseudo-panel, McKenzie [2003] finds that the coping strategy of adding more household members to the labor force, or increasing the labor hours of members already working, is not widely used. However, due to the cross-sectional nature of the data, he cannot rule out that the labor response may in fact vary with the intensity of the shock. Within this literature, a few papers focus more specifically on Argentina. McKenzie [2004] investigates whether households are able to mitigate the adverse income effect brought by the 2002 financial crisis. Among his findings, he shows that labor supply alone does not allow us to compensate for the loss in income. However, he notes that females entering the workforce contribute as much as males entering it, or more. Closest to my paper, Cerrutti [2000] establishes a direct connection between the growth in female labor force participation in Buenos Aires between 1991 and 1994, and male job instability within households. Using a first difference estimation, she documents that women living with partners whose labor force status changed are twice as likely to enter the labor force as those who live in households where the partner is continuously either employed, or out of the labor force. However, she remains inconclusive about whether her results are driven by job entries, job exits, or job instability in general. More to the point, the strength of her results is greatly hampered by the lack of an instrument to account for the endogeneity of partners' labor supply decisions. Using data on the Argentinean labor market over a longer time-frame (1996-2007) and a wider geographical area (the whole of urban Argentina), I am able to provide such an instrument, which allows for a causal measure of the female labor market's response to a partner's unemployment or income shock, as well as for a broad discussion on the mechanisms in play. To the best of my knowledge, Ayhan [2018] proposes the only other attempt to capture the causal effect of male unemployment on female participation in the context of a developing country. Using EU-SILC panel data from Turkey in the period 2007-2010, she instruments the probability of men's losing their job with variations in the output of male-dominated sectors during the 2008 crisis. She finds that the participation of Turkish married women increased by 24 percentage points in response to their partner's unexpected job loss. However, she does not study other margins, such as employment, or the hours worked, since she concentrates on the subsample of couples where the women were inactive before the Great Recession. By contrast, my paper focuses on all the women in couples, and provides evidence on a wider range of female labor outcomes, in a very different cultural area, where female participation is already high. As in her paper, my empirical strategy also makes use of the gender segregation across industries; however, while her source of exogenous variation for male unemployment stems from the variation in output from seven broad economic sectors, my instrument captures labor demand shocks at a fine industry-by-occupation level (22×4 cells), and my results are robust to the inclusion of province-by-period fixed effects, as well as to clustering the standard errors at the industry-by-occupation level.

With this paper, I also contribute to the broader literature on female labor supply in developing countries. First, in a comparative study of eight developing countries since 2000, Klasen et al. [2020] observe that in two high-income countries (South Africa and Brazil), households' economic conditions, and notably household income, do not correlate with married women's participation. They conclude that the labor supply choices of urban women in the richest developing countries are not very sensitive to income effects, following a pattern also observed for women in the US [Blau and Kahn, 2007]. My main results mitigate this conclusion: I find that women respond strongly to unexpected shocks affecting their partner's labor income. In addition, beyond this core result, I collect a range of evidence showing that female participation in response to temporary income shocks may in fact be long-lasting. I conjecture that, to a certain extent, even temporary labor supply adjustments can develop an attachment to the labor market. This conclusion is original, since it implies that temporary shocks need not necessarily be contrasted with underlying trends when one seeks to understand the evolution of the female labor supply. My paper also adds to the existing evidence on female labor force participation in Latin America. Using macro ILO estimates for female participation in various developing regions and countries over the period 1990-2015, Klasen [2019] notes that the decline in fertility in this developing region and the bridging of the gender gap in education correlate particularly strongly with female participation. In terms of female participation, Latin American countries are generally considered good performers when compared to other developing regions, which may explain why they are relatively understudied [Gasparini and Marchionni, 2015]. My paper contributes to filling this gap.

The paper is organized as follows. Section 2 describes the data, and presents the baseline results. Section 3 discusses the identification strategy. Section 4 presents my core findings. Section 5 discusses the robustness of my findings, and investigates the underlying mechanisms. Section 6 questions the dynamics of female participation in the presence of temporary shocks. Section 7 concludes. The Appendix contains further data details and robustness checks.

2 Data & Baseline Results

The data stem from the *Encuesta Permanente de Hogares (EPH)* collected by the *Instituto Nacional de Estadisticas y Censos (INDEC)*. Since 1996, the *EPH* has been a representative survey of households living in the main urban areas of Argentina. The survey provides detailed information on various labor outcomes and the usual socioeconomic characteristics at the household and individual level. Each household is observed at most four times, then rotated out. Before 2003, the survey was conducted in May and October of each year, and households were followed over 18 months. In 2003, the *EPH Continua* replaced the *EPH Puntual*. The survey became quarterly, with households observed for two years in succession, for the same two consecutive quarters of each year.

Main sample – I use all the *EPH* waves surrounding the 2002 crisis, from the first nation-wide survey to the Great Recession (1996-2007).[2] My main population of interest comprises women and men living together as household head and partner.[3] Bearing in mind legal age restrictions regarding marriage and retirement, I consider couples where women and men are aged 18-60 and 18-65 respectively (81.4% of observations). Because

2. Note that the information corresponding to the third quarter of 2007 is not available, since the provinces Mar del Plata-Bátan, Bahía Blanca-Cerri, Gran La Plata and the Greater Buenos Aires area were not surveyed due either to administrative reasons, or a strike of the *EPH* personnel.

3. This living arrangement represents 64% of the total sample of households. The remaining 36% are one-person households (16%), and single adults living with children or extended family (20%).

I am concerned with female labor market responses to shocks on male labor outcomes, I focus on active men (93.7%) with duly reported labor outcomes (96%). Since my empirical approach exploits within-household variations in time, I drop singletons (5,585 observations). My final sample consists of 243,240 observations on 81,956 unique couples, observed at least twice over the period 1996-2007.

Descriptive statistics – Table 1 documents the labor outcomes of women and men of my sample over three periods.[4] Female participation and employment increase sharply over time, from 47% to 58%, and from 41% to 52% respectively. In line with expectations, unemployment for men peaks between 2000 and 2002. At the intensive margin, employed individuals report the number of hours that they worked per week and state whether they would like to work more. This latter question is my measure for underemployment, equal to 1 if they answer positively to this question, 0 otherwise. As shown in Table 1, at the intensive margin, working women do not increase their hours of work, and underemployment decreases by one third between 2000-2002 and 2003-2007. For both genders, but mostly for men, monthly real labor income[5] decreases over time, and the simultaneous decrease in hours worked does not fully account for this decline. As a consequence, the hourly wage of women and men tend to converge. Tables A1 and A2 in the Appendix also display standard statistical information on individuals and households, and document their other income sources beyond the labor income of both partners. Table A1 shows that, on average, the women are younger than the men, and educated longer. The level of education increases in time for both sexes, while the number of children per couple decreases. As shown in Table A2, the primary additional source of income for one in every five couples is labor income from other household members. Pensions, unemployment benefits, or capital income also represent non negligible amounts, but concern very few households.

Baseline specification – My objective is to measure a woman's response in terms of labor supply to a shock in her partner's employment. The empirical specification is given by:

4. The survey provides sampling weights, which allow for population labor market statistics to be computed in the cross-section. However, these weights vary for the same household across waves, and they do not account for attrition in the panel, so they are not suited to longitudinal analysis. I prefer to use weighted data when I rely on repeated cross-sections to compute descriptive statistics, and my instrumental variable. In turn, I use unweighted data whenever the analysis relies on the panel dimension of the data.

5. Income is adjusted for inflation using the national consumer price index of Argentina provided by the Federal Reserve Bank of St. Louis (Index = 100 in 1995).

Table 1: Summary Statistics: Labor Outcomes

	1996-1999	2000-2002	2003-2007
Female labor outcomes			
Participating	0.47	0.49	0.58
	(0.50)	(0.50)	(0.49)
Working	0.41	0.42	0.52
	(0.49)	(0.49)	(0.50)
Working for 20 hours and more	0.28	0.28	0.32
	(0.45)	(0.45)	(0.47)
Observations	89,475	53,751	100,014
Conditionnal on working			
Underemployed	0.31	0.34	0.22
	(0.46)	(0.48)	(0.41)
Hours worked per week	33.53	32.33	29.97
	(19.66)	(19.25)	(19.28)
Monthly wage	516.09	467.72	394.30
	(543.67)	(548.30)	(607.57)
Observations	36,522	23,186	52,394
Male labor outcomes			
Unemployed	0.09	0.13	0.05
	(0.28)	(0.33)	(0.22)
Observations	89,475	53,751	100,014
Conditionnal on working			
Monthly wage	824.86	710.50	613.67
	(820.85)	(763.42)	(774.12)
Hours worked	48.16	45.96	45.85
	(18.58)	(19.45)	(19.58)
Observations	83,111	47,767	95,832

Note: The sample used is the main sample described in Section 2, divided into three periods. The core analysis focuses on the period 2000-2002.

$$Y^f_{h(r),t} = \gamma Unemployment^m_{h,t} + X_{h,t}\boldsymbol{\beta} + c_h + \mu_{r,t} + \epsilon_{h(r),t} \tag{1}$$

where $Y^f_{h(r),t}$ is the labor market participation of a female partner f, in household h living in region r at period t, and $Unemployment^m_{h,t}$ is a dummy equal to one if her male partner m is unemployed. The specification controls for household fixed effects c_h. This ensures that unobserved differences across households, such as the cognitive skills or soft skills of both spouses, which can affect employment and participation, do not bias the estimates. I therefore identify the parameter of interest γ on variations in labor market outcomes *within* rather than *across* couples.[6] The specification also includes a set of period-by-

6. 12.5% of the male partners in the panel changed employment status ($Unemployment^m_{h,t}$) at least once, while 85.8% (resp. 1.7%) remained employed (resp. unemployed). More precisely, 35% of them lost their

province fixed effects $\mu_{r,t}$, which accounts non-parametrically for any common variations in the female participation in the labor market within provinces over time. I therefore fully absorb any shock to female participation at the province level, such as local wage shocks, variations in the local cost of living, or local changes in gender norms.[7] The specification also controls for a series of time-varying covariates at the household level. A first set of controls measures the evolution in the household composition: the number of infants (0-2), pre-school children (3-5), children (6-15), enrolled children over 15, adults aged 16-64 and adults over 65, as listed in the bottom part of Table A1. I also include as controls the labor income of the household (excluding partners'), as well as its non-labor income, as listed at the bottom of Table A2; i.e. pensions, capital income, unemployment benefits, remittances, and other non labor income. Note that obvious reasons exist for believing that these covariates may be endogenous to female participation. For instance, the presence of a grand-parent can foster the participation of the main female caregiver in the household, but we can equally imagine that a grand-parent moves in precisely to take over child care when the parent decides to enter the labor market. But omitting these variables would create a bias, if they correlate with the unemployment status and decision of the couple to participate in the labor market. For instance, in the case of an unemployment shock to the main provider, the labor income of other adult members is more than possibly a substitute for female participation. Failing to control for household labor income would result in a downward biased estimation of γ. I therefore choose to control for the full set of covariates. $\epsilon_{h(r),t}$ is clustered by household, since unemployment, my variable of interest, varies at the household level.[8]

Baseline results – Column 1 of Table 2 presents the results of the baseline specification as described above. Columns 2 and 3 further investigate the extensive margin of participation, and look at female employment and full-time employment (defined as working 20 hours per week or more). Columns 4 and 5 document the responses at the

job, 36.6% found a new job, and 28.3% both lost and found a job during the observation period. These unstably employed male partners live in significantly larger households, with more children and other adults. Uneducated and primary school graduates are over-represented (70.6%, as against 52.5% in the stably employed group).

7. Provinces are the smallest geographic unit available in the survey. There are 29 provinces in the survey. The largest province is over 100,000 square mile, which may seem large to account for the changing local economic environment. However, in concrete terms, the survey concentrates on one, or sometimes two urban centers per province. Province fixed effects are therefore rather convincing proxies for local urban labor markets.

8. The results displayed in Table 2 are robust in sign and magnitude to the exclusion of the time-varying covariates listed in $X_{h,t}$. Their significance is not affected by the use of alternative standard errors, i.e., standard errors clustered at the province level, or two-way clustered at the province and period levels.

intensive margin on the subsample of working women.[9]

As seen in Column 1, women are 6 percentage points more likely to enter the labor market when their partner loses his job. This is a 12% increase from the 50.3% baseline participation indicated in the bottom part of the table. As displayed in Columns 2 and 3, for these women, employment increases by 4 percentage points and full-time employment (over 20 hours per week) by 2 percentage points. So out of three new female entrants in response to their partner's job loss, two actually find a job, and one even starts working full-time within 3-6 months.[10]

Looking at the correlation with other covariates, women's labor market decisions depend very little on changes in the household composition. Notably, they do not respond to changes in the number of children.[11] In line with expectations, female participation is a substitute for most other non-labor sources, such as pensions, unemployment benefits, or remittances. Such income is however a complement to other sources of labor income. This may be because the family business is expanding, or because the decision to participate also concerns other family members beyond the female partner.

Finally, I turn to the intensive margin. Columns 4 and 5 indicate how underemployment and weekly hours vary with the job loss of a male partner. As can be seen in Column 4, the partner's job loss is associated with a 25% increase in the probability that a woman wishes to work more. However, Column 5 shows that it is not correlated with the actual number of hours worked.

Interpreting γ – The suggestive evidence presented in Table 2 points towards a sizable correlation between unemployment and the labor market outcomes of life partners. When I estimate a comparable specification on the pooled cross-sectional data, excluding household fixed effects (not reported here), γ is two to three times higher, even when I control for the richest possible set of covariates. The panel dimension is thus crucially important for avoiding a large upward omitted variable bias. The period-by-province fixed effects account for the combined effect of local labor market opportunities for both

9. Appendix Table A3 compares women working to other women in couples, and also compares women continuously working throughout the panel to women changing employment status over time. As expected, working women are more educated and have fewer children. Interestingly, women with interrupted employment are very similar to inactive/unemployed women in terms of education and household composition.

10. Waves are spaced 6 months apart in the *EPH* (1996-2002), and 3 months apart in the *EPH continua* (2003-2007).

11. Note that the time span between consecutive observations in the panel is only three to six months, so this apparent absence of correlation may in fact be explained by anticipated withdrawals form the labor market.

Table 2: Baseline Specification

	Extensive Margin (All Women)			Intensive Margin (Working Women)	
	Particip.	Empl.	Full T.	Wish	Hours
Male partner is unemployed	0.0592***	0.0390***	0.0203***	0.0495***	0.142
	(0.004)	(0.004)	(0.004)	(0.008)	(0.239)
Composition					
N children 0-2	0.00535	-0.000534	0.00397	0.00995	1.464***
	(0.004)	(0.004)	(0.003)	(0.008)	(0.267)
N children 3-5	-0.00672	-0.00696*	-0.00332	0.00260	0.612**
	(0.004)	(0.004)	(0.004)	(0.008)	(0.282)
N children 6-15	0.000163	0.00123	-0.00307	0.00239	0.150
	(0.004)	(0.004)	(0.003)	(0.006)	(0.243)
N children 15+ studying	-0.000136	0.00302	0.00156	0.000623	0.479**
	(0.003)	(0.003)	(0.003)	(0.006)	(0.218)
N adults 16-64	-0.00629**	-0.00602**	-0.00456*	-0.000474	0.120
	(0.003)	(0.003)	(0.003)	(0.006)	(0.200)
N adults 65+	0.0232**	0.0187*	0.00782	-0.0327*	0.105
	(0.010)	(0.010)	(0.010)	(0.018)	(0.624)
Household Income					
Other Labor Income	0.00616***	0.00721***	0.00447***	-0.00220**	0.0370
	(0.001)	(0.001)	(0.001)	(0.001)	(0.040)
Pensions	-0.0125***	-0.0128***	-0.00809***	0.00212	-0.195**
	(0.001)	(0.001)	(0.001)	(0.002)	(0.090)
Capital	0.000731	0.000273	-0.000876	-0.00129	-0.00997
	(0.001)	(0.001)	(0.001)	(0.002)	(0.075)
Unemployment	-0.00451**	-0.0112***	-0.00803***	0.00230	-0.0545
	(0.002)	(0.002)	(0.002)	(0.004)	(0.127)
Remittances	-0.00305*	-0.00876***	-0.00990***	0.0101***	-0.170
	(0.002)	(0.001)	(0.001)	(0.003)	(0.104)
Other	-0.0116***	-0.0154***	-0.00483***	0.00684***	-0.245***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.067)
Household FE	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes
R^2	0.762	0.792	0.755	0.585	0.755
$\bar{Y}$ if Husb. Employed	0.503	0.457	0.308	0.209	32.383
Observations	243,240	243,240	243,240	101,332	101,332
Unique households	81,956	81,956	81,956	36,614	36,614

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by households are shown in parentheses. Particip. is a female participation dummy, Empl. is a female employment dummy, Full T. is a binary variable equal to 1 if women report working 20 hours or more, Wish is a binary variable equal to 1 if women declare that they are seeking to work longer hours, Hours is a continuous variable measuring the weekly working hours. The sample used is the main sample described in Section 2. In the last two columns, the sample is restricted to couples where the women are working.

partners, so γ captures the effect of male unemployment, net of the changing economic environment. Therefore, I consider the results summarized in Table 2 to be informative *per se*, and an important step in underpinning a causal effect of a man's job loss on the labor outcomes of his female partner.

Even so, γ still need not be interpretable as the causal effect of male job loss on female participation. In general, household members – notably, life partners – make their labor supply decision jointly. This has been shown in a variety of empirical applications. For instance, labor supply interdependencies between spouses have been identified in the context of an exogenous workweek reduction reform [Goux et al., 2014]. The participation of female partners over the lifecycle happens to be considerably influenced by joint taxation schemes [Groneck and Wallenius, 2021]. The retirement eligibility of a partner causally impacts the decision of the other partner to retire [Lalive and Parrotta, 2017]. Even more directly related to this paper, the literature provides theoretical foundations and empirical support for the joint job-search behavior of couples. For Guler et al. [2012], a job offer will quickly be accepted by a dual-searcher couple (in relation to a single individual), because job search options are still open for the second spouse. The acceptance of a job by a searching spouse may then trigger the employed spouse to quit one job for a better one. This gives rise to a 'work-quit-search-work' dynamic within the couple. In my context, a male partner may encourage his wife to enter employment, before he quits one job and starts searching for a better one.

From an econometric point of view, this boils down to a reverse causality issue, where partners' decisions mutually influence each other. Since female participation can cause male unemployment, such lifetime arrangements should bias γ upward.[12] To address this possibility, I now turn to an instrumental variable strategy, where the unexpected component of the industry-by-occupation shock to male employment during the 2002 economic crisis serves as an instrument to male job loss.

3 Identification

My instrumental variable relies on a quasi-experiment, namely, the sudden collapse of the convertibility regime in January 2002 in Argentina. I start by providing key background information regarding the convertibility law, and show how a series of events built up to trigger the abrupt end of the convertibility regime. In particular, I provide a body of evidence pointing to the fact that the end of convertibility was largely unexpected by the public before autumn 2001. Then I show that this sudden event changed the exposure of individuals to a job loss, to a degree that they could not have predicted from the

12. This positive reverse causality can be counterbalanced by the presence of leisure complementaries if both spouses enjoy spending time together. Alternatively, the upward bias can be amplified if their leisure times are substitutes in their individual utilities.

employment trend of the preceding mild recession. I measure this exposure intensity at the industry-by-occupation level, with the aim of using it as an instrument for the actual job loss experienced by male partners between 2001 and 2002 (see Section 4, below). I end this section with a discussion about the identifying assumptions underlying the use of this instrument.

Background information on the convertibility law – The convertibility law was adopted in March 1991 under the presidency of Carlos Menem, with the objective of containing hyperinflation. The law provided that the newly created Argentinean peso would be pegged to the dollar, at a one-to-one exchange rate. In parallel, a series of structural reforms were undertaken in order to ensure the credibility and sustainability of this new exchange regime. In the spirit of the 'Washington Consensus' promoted by the IMF, these reforms included massive waves of privatization for publicly owned companies, as well as measures for the liberalization of trade, and of the labor market. These liberal reforms formed the backbone of the Argentinian economy during the presidency of Carlos Menem, or *menemato* (1989-1999). They contributed to restoring people's trust in the domestic currency: prices had been growing by 1300% between January and December 1990, but in April 1991, inflation shrank to 5.5% monthly, and by April 1992 the annual inflation rate was reduced to 25%. The Argentine economy grew at an annual rate of 5-10% over the period, and weathered the contagion of the Tequila crisis with success in 1995. Although the reforms came at the cost of growing unemployment [Cerrutti, 2000], the convertibility remained extremely popular among the middle class.

From a moderate recession to the sudden collapse of the convertibility regime – In 1999, Argentina entered a moderate recession because of a combination of international and domestic factors. The Russian financial crisis, the US monetary policy, and the currency devaluation in Brazil led to a first phase of destabilization [Fanelli, 2002]. Furthermore, 1999 was presidential election year in Argentina. A spending race between the incumbent president and his opponents in the Peronist party jeopardized fiscal austerity. The opposition (a center-left coalition) won the elections, but the newly elected president, Fernando de la Rúa, lacked the necessary political support [Corrales, 2002]. This combination of economic and political factors built up vulnerabilities in the convertibility regime. In spite of the fiscal deterioration in 1999 and 2000, however, Argentina was repeatedly backed by IMF-supported programs.

So when did people lose their trust in convertibility? Although Argentinians had been used to booms and busts since WWII, they had serious reasons to believe that this time was different. First, the convertibility was enshrined in law. Then, under the convertibility, the country had experienced an unusually prolonged period of sustained growth with no inflation. Finally, until shortly before the crisis, the country had been publicly praised by the IMF and the US Treasury for its achievements in stabilization, economic growth and liberal reforms. In spite of the recession, Argentina, like any other emergent economy, was predicted to recover rapidly [Outlook, 1999].

More serious doubts about the stability of the currency board did not surface until 2001 [Daseking et al., 2005a]. Early in 2001, the first signs of pessimism over the sustainability of the regime came from the capital markets, as the spread on Argentine over US bonds suddenly rose [IMF, 2003]. During the first half of 2001, the Argentine authorities reiterated their commitment to severe fiscal austerity and the convertibility regime with several unpopular measures, demonstrating that they were willing to save convertibility at all costs, including electoral ones. In these efforts, they were publicly backed by the IMF with massive loans. In particular, in July and August, the outflow of bank deposits increased as a consequence of a mixture of fears among savers and investors. The joint response of the government with the 'deficit zero' law (July 29th) and the IMF with a loan (August 22nd) temporarily restored trust: the main index of the Buenos Aires stock exchange rose, the investment risk rate in Argentina fell [Relea, August 23 2001], and in September, bank cash deposits were increasing again [Otaloa, October 3 2001]. Mid-term elections took place in October 2001. The government issued a statement saying that it would maintain its policy of economic austerity to maintain convertibility, regardless of the outcome of the mid-term elections [Illiano, October 3 2001]. When asked about economic policy, the main political opponent of the government coalition, Eduardo Duhalde, declared ' ... it is the model that is perverse' [Naranjo, October 5 2001], but ruled out dollarization and did not mention devaluation at all. On this occasion, voters harshly punished the political establishment, by abstaining or submitting blank or defaced ballot papers. The main sources of discontent were not convertibility per se, but corruption and discontent with the general course of public affairs, notably the management of the recession and the budgetary cuts [Corrales, 2002]. This debilitated even more the already fragile legitimacy of the incumbent government.

In November 2001, the situation rapidly escalated. A bank run started, and the government's response came in the first days of December, in the form of the highly unpopular *corralito* measure of freezing bank accounts. Shortly afterwards, the IMF announced the end of its support, due to Argentina's inability to meet the loan conditions on the zero deficit. This triggered the social and political collapse of December 2001, as well as the end of convertibility. Protests in the form of *cacerolazos* were so massive and violent that

President Fernando de la Rúa had to resign; he escaped from the presidential palace by helicopter on December 20, 2001. Note that even in the midst of these events, the IMF predicted a small 1.1% fall in output and deflation for 2002 [Outlook, 2001], suggesting that experts still contemplated the possibility that convertibility would survive the economic, social and institutional collapse of December 2001. But on January 6th, 2002, the newly appointed president Eduardo Duhalde repealed the convertibility law that had been in place for ten years. The inflation rate rose at once from 0 to 10% per month, the recession reached 10% of GDP in 2002, and the unemployment rate peaked at 22%.

From this body of evidence, I conclude that the depth of the 2002 shock was mostly unexpected by households before and during the October 2001 wave. I now turn to the second interesting feature of the 2002 crisis, namely, that it created a break in the sequence of male employment for many industry-by-occupation cells. This is important, because it suggests that the end of convertibility was not only unexpected in its timing, but also in its diverse consequences for male employment at the industry-by-occupation level.

Industry-by-occupation exposure – To capture the unexpected component of the industry-by-occupation male exposure to unemployment following the sudden collapse of convertibility, I rely on the full sample of men aged 16-64 between 1996 and 2002.[13] $E_j^{i,o}$ is a dummy equal to 1 if individual j is employed in industry i and occupation o, 0 otherwise.[14] For each industry-occupation pair, I estimate the following specification:

$$E_{j(r),t}^{i,o} = a^{i,o} Trend_t + b^{i,o} Post_t + \phi_r^{i,o} + \epsilon_{j(r),t}^{i,o} \tag{2}$$

where $\phi_r^{i,o}$ is a set of province fixed effects, controlling for differences in industrial composition across labor markets, $Trend$ is a linear time trend, and $Post$ is a dummy equal to 1 if individual j is observed after the collapse of convertibility and therefore either in May, or October 2002; it equals 0 otherwise. $a^{i,o}$ thus captures the structural evolution of male employment in the industry-occupation pair i, o, as anticipated by households. By contrast, $b^{i,o}$ measures how male employment in the industry-occupation pair i, o reacts to the sudden end of convertibility, and serves as an instrument for a male partner's job loss.

13. May 1996 saw the first available wave at the national level. In 2003, Argentina entered a strong recovery phase.

14. There are 22 industries, and 4 types of occupations in the present study. Section B in the Appendix provides additionnal information about the classifications used by the *EPH*, the definitions of industries and occupations; it also presents descriptive statistics on each industry-occupation pair.

Figures B1 and B2 in the Appendix display point estimates $\hat{a}^{i,o}$ and $\hat{b}^{i,o}$ for all industry-occupation pairs i, o, along with their 90% confidence interval. For instance, $\hat{b}^{12,3}$ (operators in the transportation sector) is -0.008, so the probability for men aged 16-64 of being employed in this particular industry-occupation relative to being employed in another industry, or unemployed, or inactive, decreases by 0.8 percentage points with the collapse of convertibility. In the years before the collapse, this particular industry-occupation pair had shown signs of relative expansion, as suggested by the positive $\hat{a}^{12,3}$ point estimate. As visible on Figure B1, before the sudden end of convertibility, men's employment in some sectors was already on a relatively upward trend – with the example of clerical workers in the private and public service sectors – or a downward trend – with the example of technicians and operators in the manufacturing sectors. As the convertibility regime collapsed in January 2002, one in every four industry-occupation pairs (weighting for 40% of total male employment) experienced a significantly negative unexpected employment shock. Employment was hardest hit in the construction and wholesale trade. This was expected, for these industries were the one affected by the greatest shocks to product demand: a 33.4% GDP recession in construction in 2002, and one of 18.4% in wholesale and retail trade [McKenzie, 2004]. Other industries were affected as the crisis spread. The only industries where the probability of employment increased in 2002 relative to other industries and unemployment or inactivity were industries where the rate of self-employment was high (such as repair for technicians, or domestic services for operators), or sectors where labor demand was supported by the state (such as local social services, or public administration for the least qualified).

$\hat{b}^{i,o}$ measures the severeness of the shock to male employment in different industry-occupation cells upon the collapse of the convertibility regime. Therefore, my instrumental variable takes the value of 0 for all couples observed before January 2002, and $-b^{i,o}$ for the May and October 2002 waves.[15]

Validity of the exclusion restriction – For the empirical strategy exposed above to be valid, the exposure of men to unemployment should affect women's participation only through the channel of their partner's job loss, and not through some other channel. In particular, the exclusion restriction loses its validity if their partners' exposures to a job loss are correlated.

An obvious case of correlation arises when both partners work in the same industry. Fig-

15. Note that I use the opposite of $\hat{b}^{i,o}$ rather than $\hat{b}^{i,o}$ itself to get a measure of the *intensity* of the shock to male employment. Since I instrument for unemployment, this allows for a more intuitive interpretation of the results: a higher shock to employment in a given industry-occupation pair implies a higher likelihood of being unemployed for men working in this particular industry-occupation pair.

ure B3 in the Appendix pictures the distribution of the partners' industry combinations for dual-earner partners in October 2001. Overall, partners are more likely to work in a different industry than in the same one, and, for most industries, the conditional distribution of industries which employ females is rather uniform. This confirms the view that industries are gender segregated, a stylized fact also documented in developed economies [Verdugo and Allègre, 2020]. Two instances stand out, namely, retail trade, reflecting the existence of family businesses, and to a lesser extent the public sectors (administration, education, and health), reflecting assortative matching. In the discussion, I show that my results are robust to the exclusion of couples where male partners work in retail trade, or in these public sectors.

Exposure to the end of convertibility can be correlated across partners, whether or not they work in the same industry. To check whether the exposure to the convertibility shock is correlated within couples, I start by estimating again Equation (2) for all industry-occupation pairs in the subsample of *women* aged 16-64. This allows me to assign to each dual-earner couple observed in October 2001 in a given industry and occupation combination its own combination of future shocks. Figure B4 plots all these combinations. The size of the circle represents the weight of a particular combination of male-female shocks over the total number of existing combinations. Most combinations are weighted below 0.1%, but a few combinations are weighted more than 1%: clerical workers in public administration with technicians in education (1.12%), couples of technicians in education (1.20%), couples of clerical workers in public administration (1.55%), couples with an elementary occupation in retail trade (2.88%), and construction operators in couple with unqualified domestic workers (3.3%). As reflected by the solid line in Figure B4, a simple weighted regression of the future shock experienced by woman on her partner's future shock suggests that shocks are slightly positively correlated within households: men working in the industry-occupation pairs most affected by the end of convertibility tend to be the partners of women who are also slightly more exposed than other women. However, this effect is essentially driven by the joint negative shock experienced by construction operators whose partners are unqualified domestic workers (represented by what is highlighted in black at the bottom left-hand corner of Figure B4). When I drop these dual-earner couples, the positive correlation disappears, as reflected by the dashed line also incorporated on Figure B4. In Section 5, I show that my core results are robust to the exclusion of couples where the men are construction workers.[16]

16. Note that in the presence of positively correlated shocks, my baseline results would likely measure the lower bound of the true parameter values. Indeed, in this case, a woman in a relationship with an exposed man would also have a higher probability than other women of being discouraged from entering

4 Results

In this section, I estimate women's participation and employment reactions to their partners' exogenous job loss (through the unanticipated, heterogeneous shock to male employment, fueled by the end of convertibility).

The two-stage empirical specification is given by:

$$\begin{aligned} Unemployment^{m}_{h(i,o,r),t} &= \pi Intensity^{m}_{i,o,t} && +\eta'_{i,o} + X_{h,t}\boldsymbol{\beta'} + c'_h + \mu'_{r,t} + \epsilon'_{h(i,o,r),t} \\ Y^{f}_{h(i,o,r),t} &= \gamma \widehat{Unemployment^{m}}_{h,t} && +\eta_{i,o} + X_{h,t}\boldsymbol{\beta} + c_h + \mu_{r,t} + \epsilon_{h(i,o,r),t} \end{aligned} \tag{3}$$

where $Intensity^{m}_{i,o}$ is the instrument defined in Section 3, and $\widehat{Unemployment^{m}}_{h,t}$ is the predicted value for male unemployment, as given by the first stage equation. To the covariates $X_{h,t}$ and fixed effects c_h, $\mu_{r,t}$ described above, I add a set of industry-occupation fixed effects $\eta_{i,o}$, identified in men changing their industry-occupation cell over time, because omitting these fixed effects could lead to biased estimates. For instance, during the six month window between October 2001 and May 2002, women may have entered the labor market, and men may have found a job in an other, less exposed industry. Failing to control for the switch would lead to a downward bias of the estimate, because a woman's participation would be mistakenly associated with a relatively low industry-occupation exposure for her partner. $\epsilon_{h(i,o,r),t}$ is clustered by industry-occupation pair, because the instrument varies at the industry-occupation level.

Consistently with my empirical strategy, I estimate the set of equations (3) on the subsample of partners observed at least once in or before October 2001, and once after the convertibility collapses, i.e. in May and/or October 2002. My sample consists of 20,071 observations of 6,268 unique households between October 2000 and October 2002.[17]

Table 3 displays the first stage estimates for the extensive and intensive margin estimations. As expected, the correlation between the intensity of the shock to male employment and mens' reported unemployment is significant and positive. As indicated in the bottom part of Table 3, the average shock is 0.001, with a standard deviation of 0.005. In concrete terms, a one standard deviation change in the shock intensity is associated with a (5.611 $\times$ 0.005 =) 2.8 percentage point increase in the likelihood for male partners to be unemployed. At the bottom, I report the KP F statistic as a test for weak instruments. At

the labor market.

17. I replicate the baseline specification of Equation (1) on this subsample, and get results that are very similar to Table 2, displayed in Appendix Table C1. In Section 5, I also discuss the possibility that my results are driven by a composition effect due to non-random attrition.

Table 3: Main Results – First Stage Estimates

	Extensive Margin (All Women)	Intensive Margin (Working Women)
	Male partner is unemployed	Male partner is unemployed
Shock Intensity	5.611***	4.441***
	(0.486)	(0.815)
Household FE	Yes	Yes
Region × Time FE	Yes	Yes
Industry × Occupation FE	Yes	Yes
Time-varying controls	Yes	Yes
KP	133.13	29.68
Mean(Intensity)	0.00153	0.00130
SD(Intensity)	0.00499	0.00464
Observations	20,071	7,720
Unique households	6,268	2,642

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by industry-occupation pairs are shown in parentheses. There are two first stage regressions: the first column reports the first stage results for all couples, at the extensive margin of participation and employment, and the second column reports the results on the restricted sample of working women, to document the intensive margin. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2. For the sake of completeness, Appendix Table B1 also displays the coefficients of the time-varying controls.

the extensive margin, the statistic is far above the critical values provided by Stock and Yogo [2005]. When I implement the first stage on the subsample of dual-earner couples, the statistic, albeit much smaller, also lies in an acceptable range.

Table 4 presents the results for the two-stage estimation. The response to the unexpected job loss of the partner is greater than the simple panel estimates suggest (see Section 2, as well as Table C1 in the Appendix). The probability of participating and being employed is multiplied by two. However, the probability of finding a full-time job does not increase with a partner's displacement. Working full-time may require anticipatory action from women who are also caregivers in their household. In addition, the low labor demand may also condemn women to underemployment. This latter explanation is supported by the results on the intensive margin. Working women do not work more hours after their partner loses his job, but they vehemently wish that they could.

Table 4: Main Results - 2SLS Estimates

	Extensive Margin (All Women)			Intensive Margin (Working Women)	
	Particip.	Empl.	Full T.	Wish	Hours
Male partner is unemployed	0.484***	0.492***	-0.0677	0.543**	-12.63
	(0.085)	(0.097)	(0.076)	(0.268)	(9.132)
Household FE	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes
Industry × Occupation FE	Yes	Yes	Yes	Yes	Yes
Time-varying controls	Yes	Yes	Yes	Yes	Yes
$\bar{Y}$ if Husb. Employed	0.475	0.426	0.296	0.280	33.587
Observations	20,071	20,071	20,071	7,720	7,720
Unique households	6,268	6,268	6,268	2,642	2,642

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by industry-occupation pairs are shown in parentheses. Particip. is a female participation dummy, Empl. is a female employment dummy, Full T. is a binary variable equal to 1 if women report working 20 hours or more, Wish is a binary variable equal to 1 if women declare that they are seeking to work longer hours, Hours is a continuous variable measuring the weekly working hours. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility. In the last two columns, the sample is restricted to couples where women report working at least once before and after January 2002. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2. For the sake of completeness, Appendix Table B2 also displays the coefficients of the time-varying controls.

5 Discussion

In this section, I discuss the robustness of my core results. I also investigate the mechanisms that may be driving these results.

Robustness – First, I provide additional supporting evidence of the validity of the exclusion restriction. I start by testing the sensitivity of my results to the exclusion of different male industries. To do so, I estimate again the set of equations (3), excluding one industry at a time. Figure C1 shows the coefficients obtained for these first stage and second stage estimations. Sub-figure C1a reports the female participation response to male unemployment on these different subsamples. Sub-figure C1b displays the corresponding first stage male unemployment response to the intensity of the convertibility shock. As can be seen in Figure C1, my main results are not driven by any particular male industry. Note that excluding construction introduces some noise, because this industry is weighted as of October 2001 for 11% of the employed men, and 40% of the unemployed men aged 16-64. Still, the sign and magnitude of the coefficients do not change upon the exclusion of this industry.

Then I check whether the shock intensity captured by my instrumental variable in 2002 is statistically related to a pre-trend in male unemployment before the convertibility collapse in 2002. To this end, I estimate again the first stage of Equation (3), augmented with interaction terms between year dummies and the 2002 shock intensity, on couples aged 18-64, observed between 1996 and 2002:

$$Unemployment^{m}_{h(i,o,r),t} = \sum_{y=1996}^{2002} \pi^{y} \left[Intensity^{m}_{i,o,2002} \times \mathbb{1}(\text{Year} = y)\right] + \eta'_{i,o} + X_{h,t}\boldsymbol{\beta'} + c'_{h} + \mu'_{r,t} + \epsilon'_{h(i,o,r),t} \tag{4}$$

Figure C2 plots the point estimates $\hat{\pi}^{y}$, together with their 90% confidence interval. Taking 1996 as a reference point, I find that the probability of being unemployed does not increase faster in the industry-occupation pairs that will be most hit in 2002. The only exception is 2001, when the difference in unemployment trends across industry-occupation pairs seems positively related to the intensity of the 2002 shock. In other words, unemployment in more exposed industry-occupations was already growing at a slightly faster rate a year before the actual end of convertibility.

I argue that this divergence does not pose a serious threat to identification. First, in absolute terms, the magnitude $\hat{\pi}^{2001}$ lies in the same range as $\hat{\pi}^{1997} - \hat{\pi}^{2000}$, and is comparable to $\hat{\pi}^{1998}$. Then, by contrast, the divergence in unemployment once the convertibility collapses, measured by $\hat{\pi}^{2002}$, is five times higher than the unemployment divergence captured by $\hat{\pi}^{2001}$. Therefore, $\hat{\pi}^{2002}$ stands out as the only wide divergence in unemployment trends.

I close this section by providing some evidence on another potential threat to validity, namely non-random panel attrition. I consider non-random panel attrition to be a potential issue in the present research context, for two reasons. First, by design, the survey follows dwellings rather than households, so couples moving out automatically generate panel attrition. It has been documented that changes in labor market opportunities are a key driver of mobility, so the probability of attrition could be positively correlated with the convertibility shock. Then, in May 2002, the national statistical institute recorded threats and assaults on *EPH* interviewers. These violent acts caused attrition, notably in the Great Buenos Aires area [INDEC, 2002], and are probably correlated with particularly bad labor market outcomes.

To what extent does the likelihood of dropping out of the sample depend on the shock intensity experienced by male partners in 2002? To answer this question, I first use the main sample described in Section 4, and estimate the following specification:

$$Attrit_{h(i,o,r),t} = \gamma Intensity^{m}_{i,o,2002} + X_{h,t}\boldsymbol{\beta} + c_h + \mu_{r,t} + \epsilon_{h(i,o,r),t} \tag{5}$$

where $Attrit_{h(i,o,r),t}$ is equal to 1 if household h living in region r in period t drops out of the survey prematurely at $t+1$, 0 otherwise, $Intensity^{m}_{i,o,2002}$ is the shock intensity measure calculated in Section 3 for the industry-occupation pair i, o of the male partner, c_h are household fixed effects, $\mu_{r,t}$ are province period fixed effects, and $X_{h,t}$ are the same covariates as in Equation (1). I thus identify γ on variations in the shock intensity *within* households, since a male partner may change industry and/or occupation. Second, for external validity, I test whether the probability of attrition varies *across* households, depending on their exposure to the 2002 shock. To this end, I estimate again Equation (5) using pooled OLS, where I replace household fixed effects with the basic demographic characteristics of both partners, i.e. their age and their level of education using four stages (none, primary, secondary and tertiary education). Columns (1) and (2) of Appendix Table C2 show that the likelihood of dropping out of the survey is not related to an increased exposure to the convertibility crisis *within* households, nor to a difference in exposure *across* households.[18] Finally, I take the attrition issue upstream. Indeed, my sample restriction imposes the condition that partners should be observed at least once in October 2001 or before, and once after. If the likelihood of being observed after October 2001 depends on the intensity of the shock, my results may be driven by a composition effect. Columns (3) and (4) replicate the estimations of Columns (1) and (2) on the sample of households surveyed between October 2000 and October 2002 before I imposed the restriction. I do not find any evidence that households about to be hit by a more intense shock have a greater tendency to leave the survey.

Mechanisms – I now explore the mechanisms behind my main results. I start by investigating the role played by a state-wide workfare policy, the program *Jefes y jefas*, in sustaining the demand for labor. Then I study whether the causal link running from male unemployment to female participation is driven more by job destruction, job creation, or equally by both. Finally, I extend my research question to other male labor outcomes, and provide complementary evidence on a woman's labor market response to a decline in her partner's monthly labor income.

The workfare program *Jefes y jefas* was first introduced in January 2002 under loan and

18. The coefficient γ is very small: in column (1), a one standard deviation change in the shock intensity is associated with a $(0.119 \times 0.005 =)$ 0.06 percentage point increase in the likelihood of dropping out prematurely from the survey.

technical assistance from the World Bank. It was extended after April 2002, and reached more than 2 million beneficiaries within a few months. Eligible for entry to the program were unemployed household heads with at least one child under 18. Though the program was universal, 20 weekly working hours were required as a counterpart for the 150AR$ (representing around half of the mean household income per capita in 2002), in order to target the poorest households, whose members have a lower reservation wage. In most cases, the local municipalities in charge of the program proposed part-time, unskilled positions in the public sector that were particularly attractive to women. Indeed, in their evaluation of the program, Galasso and Ravallion [2004] report that two thirds of the beneficiaries were women. Anecdotal evidence suggests that women preferred to enroll themselves on the program as unemployed household heads, while their unemployed partners kept searching for higher paid jobs. To what extent did the workfare program facilitate the response of female participation to male unemployment?

Table 5: Accounting for *Jefes y jefas* – 2SLS Estimates

	Extensive Margin (All Women)			Intensive Margin (Working Women)	
	Particip.	Empl.	Full T.	Wish	Hours
Male partner is unemployed	0.239**	0.177*	-0.0839	0.568**	-10.91
	(0.092)	(0.097)	(0.077)	(0.252)	(9.370)
Workfare Beneficiary	0.377***	0.484***	0.0248	-0.0256	-1.795
	(0.022)	(0.022)	(0.032)	(0.089)	(1.408)
Household FE	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes
Industry × Occupation FE	Yes	Yes	Yes	Yes	Yes
Time-varying controls	Yes	Yes	Yes	Yes	Yes
$\bar{Y}$ if Husb. Employed	0.475	0.426	0.296	0.280	33.587
Observations	20,071	20,071	20,071	7,720	7,720
Unique households	6,268	6,268	6,268	2,642	2,642

Note: Table 5 replicates Table 4, adding the female participation to the workfare program *Jefes y jefas* as a control. *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by industry-occupation pairs are shown in parentheses. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility. In the last two columns, the sample is restricted to couples where women report working at least once before and after January 2002. Particip. is a female participation dummy, Empl. is a female employment dummy, Full T. is a binary variable equal to 1 if women report working 20 hours or more, Wish is a binary variable equal to 1 if women declare that they are seeking to work longer hours, Hours is a continuous variable measuring the weekly working hours. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2.

To investigate this question, I estimate again the set of equations (3), adding female enrollment into *Jefes y jefas* as a binary control variable.[19] Table 5 displays the second stage results (first stage results are shown in Appendix Table C3). As expected, female enrollment closely correlates with both female participation and employment. Conditional on the program, the causal link from male unemployment to female participation and employment is still positive and significant, albeit divided by two. I conclude that the new jobs created by *Jefes y jefas* account only for roughly half of the increase in participation and in employment caused by the sudden displacement of male partners.[20]

Another interesting question is whether the causal positive link from male unemployment to female participation is driven by men *entering* or *exiting* unemployment. Indeed, in the context of a deep economic crisis, one could as well imagine that a discouraged woman would withdraw from the labor market as soon as her partner found a job in a relatively protected sector. To decide between these alternative interpretations, I estimate the following first-difference specification:

$$\Delta Y^f_{h(r),t} = \gamma^{entry} Entry^m_{h,t} + \gamma^{exit} Exit^m_{h,t} + \Delta X_{h,t}\boldsymbol{\beta} + \mu_{r,t} + \Delta\epsilon_{h(r),t} \tag{6}$$

where $Entry^m_{h,t}$ is a binary variable equal to 1 if the male partner *enters* unemployment ($\Delta Unemployment^m_{h,t} = 1$), 0 otherwise, and $Exit^m_{h,t}$ is a binary variable equal to 1 if he *exits* unemployment and finds a job ($\Delta Unemployment^m_{h,t} = -1$), 0 otherwise.[21] As summarized in Table 6, the causal link running from male unemployment to female participation is exclusively driven by the *loss* of a male job.[22]

19. The national statistical institute collected data on *Jefes y jefas* beneficiaries in a separate survey, starting in October 2002. I merge this data with my sample, and I use information on the date of entry in the program to generate retrospective information on the beneficiaries in May 2002. I find that 1.9% of female partners (resp. 9.9%) were beneficiaries in May (resp. October) 2002. Note that the program was mostly extended after April 2002; hence the difference. In October 2002, I am also able to document that 83.5% of the female beneficiaries actually complied with the working requirements.

20. In the context of the rapid scale-up of *Jefes y jefas*, one could imagine that the positive participation response to male unemployment is *indirectly* driven by the program. Inactive, non-beneficiary women with displaced partners would file for unemployment, with the sole objective of eventually enrolling on *Jefes y jefas*. Under this scenario, however, I should find that female *employment* does *not* respond to male unemployment. The results in Table 5 contradict this scenario; in fact, male displacement generates both a participation response *and* an employment response.

21. In other words, $\Delta Unemployment^m_{h,t}$ is a linear combination of $Exit^m_{h,t}$ and $Entry^m_{h,t}$: $\Delta Unemployment^m_{h,t} = Exit^m_{h,t} - Entry^m_{h,t}$. Note that I do not instrument for $Entry^m_{h,t}$ and $Exit^m_{h,t}$, because I have a single instrument.

22. Interestingly, women work 3 hours more in response to any change in their partner's status – irrespective of entry or exit. This is probably why I find no overall effect of male unemployment on female working hours in Tables 2 and C1.

Table 6: Asymmetric Effect of Male Employment on Participation – First Difference Estimates

	Extensive Margin (All Women)			Intensive Margin (Working Women)	
	ΔParticip.	ΔEmpl.	ΔFull T.	ΔWish	ΔHours
Partner enters unemployment	0.0685***	0.0414***	0.0158	0.164***	3.194***
	(0.017)	(0.016)	(0.014)	(0.031)	(0.959)
Partner exits unemployment	-0.00803	0.00168	-0.0121	0.0433	3.178***
	(0.020)	(0.019)	(0.016)	(0.036)	(1.152)
Region × Time FE	Yes	Yes	Yes	Yes	Yes
Δ Time-varying controls	Yes	Yes	Yes	Yes	Yes
Observations	12,902	12,902	12,902	5,549	5,549
Unique households	5,834	5,834	5,834	3,019	3,019

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by households are shown in parentheses. Δ indicates that outcomes and time-varying controls are measured in first difference. *Entry* is a binary variable equal to 1 if the male partner *enters* unemployment ($\Delta Unemployment^m = 1$), 0 otherwise, and *Exit* is a binary variable equal to 1 if he *exits* unemployment and finds a job ($\Delta Unemployment^m = -1$), 0 otherwise. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility. Note that by construction, the number of observations is inferior to that in Table 4, due to the first differentiation.

Besides the rise in unemployment, the sudden resurgence of inflation in 2002[23] had important consequences on other labor market outcomes, notably wages. McKenzie [2004] documents that 78% of the surveyed households experienced real income declines in 2002, and 63% suffered a real income fall of 20% or more. Given the deindexing of labor contracts during the 1990s, I conjecture that the magnitude of the fall in real wages in a given industry-occupation pair should be related to the variation in the demand for labor in this particular industry-occupation. In other words, the shock intensity on male employment at the industry-occupation level can be used as an exogenous source of variation for male partners' real labor income. To capture the causal effect of male labor income on female participation and employment, I therefore estimate the set of equations (3) on the sub-sample of couples with an employed male partner, where the variable of interest is not their unemployment status, but the logarithm of their real monthly labor income.

The first stage results in Table 7 confirm the negative impact of job destruction on wages. A one standard deviation increase in the intensity of the shock to male employment re-

23. Fiszbein et al. [2003] report that the consumer price inflation for the Greater Buenos Aires area was 41% for 2002.

Table 7: Female Participation and Male Income – First Stage & 2SLS Estimates

	Simple Model				Model with Interactions		
	First Stage	2SLS: Part	2SLS: Emp	2SLS: Full	2SLS: Part	2SLS: Emp	2SLS: Full
Shock Intensity	-11.21***						
	(2.462)						
Male Partner's Log Lab. Inc.		-0.246***	-0.215**	0.0381			
		(0.082)	(0.086)	(0.049)			
Male Partner's Log Lab. Inc. × Q1					-0.274***	-0.257***	0.0788
					(0.089)	(0.081)	(0.054)
Male Partner's Log Lab. Inc. × Q2					-0.126*	-0.0514	0.0962
					(0.063)	(0.066)	(0.058)
Male Partner's Log Lab. Inc. × Q3					0.0103	0.0178	0.124
					(0.052)	(0.067)	(0.120)
Male Partner's Log Lab. Inc. × Q4					0.0208	0.0504	0.134**
					(0.052)	(0.048)	(0.065)
Male Partner's Log Lab. Inc. × Q5					0.125**	0.200***	0.288***
					(0.058)	(0.060)	(0.100)
Household FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry × Occupation FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-varying controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Male Monthly Income before 2002	689.38						
Male Monthly Income in 2002	461.18						
$\bar{Y}$		0.473	0.425	0.296	0.473	0.425	0.296
Observations	17,027	17,027	17,027	17,027	17,027	17,027	17,027
Unique households	5,545	5,545	5,545	5,545	5,545	5,545	5,545

Note: The simple model measures the causal participation and employment responses to a negative income shocks affecting employed male partners. The first column displays the first stage estimates. The model with interaction disentangles the effect by quintile of the per capita household income distribution at baseline (first stage estimations not reported). *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by industry-occupation pairs are shown in parentheses. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility. Note that the number of observations is inferior to Table 4, since I documenting only men who are working over the period.

sults in a 5.6% decrease in male labor income.[24] The second stage results show that female participation and employment respond positively to a decline in real male labor income. The bottom part of Table 7 indicates that the average male labor income decreased by 33% between 2000-2001 and 2002, implying an 8 (resp. 7) percentage points increase in participation (resp. employment).[25] I further unpack this result, and investigate how this response varies across the pre-crisis income distribution. To this end, I interact my endogenous variable and my instrument with five binary variables, indicative of couples' household income quintile (per capita) before 2002. An effect concentrated

24. $-11.21 \times 0.005 \times 100 = -5.6$.
25. $((461 - 689)/689) \times 100 \times 0.00246$ $(0.00215) = -0.0814$ (-.0711)

on the poorest households would suggest that female participation is driven by necessity. By contrast, a homogeneous effect would rather stand in support of a ratchet effect, where households seek to maintain their pre-crisis standard of living.[26] The right panel of Table 7 supports the subsistence motive over the ratchet effect. The participation and employment responses stem exclusively from the first two quintiles of the household income distribution.

6 Dynamics of Female Participation

So far, I have documented the contemporaneous female participation and employment responses to male unemployment and income loss. I now propose two complementary approaches for studying the dynamics of female participation.

Delayed response – First, I rely on the panel dimension of my data, and investigate possible delays in female participation or employment response to male unemployment. In Turkey, using quarterly SILC employment data between 2007 and 2010, Ayhan [2018] finds that the positive female labor supply response appears with a lag of one quarter, and operates for two quarters only. In a similar vein, I estimate again Equation (1), adding lagged male unemployment as a control variable. In this way, I can capture whether female participation also increases with some delay, holding the contemporaneous response to male unemployment constant.

Table 8 reports the results. Note that because of the lagged term, the sample size decreased by 60%. Hence, for comparison purposes, I also report the effect measured without lagged male unemployment. Table 8 confirms that the contemporaneous participation and employment responses are large in size. In addition, I find that both female outcomes also react to male unemployment, with some delay.[27] Note that the ratio for female employment of delayed-to-contemporaneous response is higher than that of participation. This is because employment is more constrained by labor demand than the mere decision to participate.

Medium run response – When studying the dynamics of female labor supply, another important question is whether the increase in female participation during a crisis can be long-lasting. As soon as 2003, and until 2008, Argentina returned to a sustained 8-9%

26. Wealthier households may adopt other strategies, involving savings, or loans, to smooth their consumption.

27. The time span between two waves is a semester between 1996 and 2002, and a quarter for later waves.

Table 8: Delayed Female Labor Market Responses to Male Unemployment

	Extensive Margin (All Women)						Intensive Margin (Working Women)			
	Particip.	Particip.	Empl.	Empl.	Full T.	Full T.	Wish	Wish	Hours	Hours
Unemployed in t	0.0631***	0.0583***	0.0421***	0.0377***	0.0225***	0.0216***	0.0579***	0.0589***	0.0196	0.0878
	(0.007)	(0.007)	(0.007)	(0.007)	(0.006)	(0.006)	(0.014)	(0.014)	(0.396)	(0.388)
Unemployed in t-1	0.0181***		0.0167**		0.00334		-0.00418		-0.269	
	(0.007)		(0.007)		(0.006)		(0.013)		(0.383)	
Household FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-varying controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.797	0.797	0.822	0.822	0.798	0.798	0.639	0.639	0.803	0.803
$\bar{Y}$ if Husb. Employed	0.480	0.480	0.436	0.436	0.295	0.295	0.226	0.226	32.807	32.807
Observations	95,222	95,222	95,222	95,222	95,222	95,222	36,136	36,136	36,136	36,136
Unique households	39,278	39,278	39,278	39,278	39,278	39,278	15,546	15,546	15,546	15,546

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by households are shown in parentheses. In the 'Extensive margin' columns, the full sample of couples is used. In the 'Intensive margin' column, the sample is restricted to couples where women report working at least once before and after January 2002. Note that by construction, the number of observations is inferior to Table 2, because of the lagged independent term. Particip. is a female participation dummy, Empl. is a female employment dummy, Full T. is a binary variable equal to 1 if women report working 20 hours or more, Wish is a binary variable equal to 1 if women declare that they are seeking to work longer hours, Hours is a continuous variable measuring the weekly working hours. Unemployed in t is equal to 1 if the male partner is unemployed, and Unemployed in $t-1$ is equal to 1 if the male partner was unemployed during the previous period, 0 otherwise. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2.

growth rate each year. With the recovery, did new entrants withdraw from the labor market, or did they rather remain active, possibly revealing an attachment to the labor market?

To the best of my knowledge, the literature on developing economies is silent in this regard. In this paper, I am also constrained by data availability. First, in the *EPH*, households are followed over four waves, that is, for 18 months in total. This time span is insufficient to document any medium-term effect on female participation or employment. Even more critical, as explained in Section 2, the *EPH* survey methodology evolved dramatically in 2003, creating a break in the panel. Therefore, I cannot rely on the panel dimension of the data to measure how female labor market outcomes evolved with the economic recovery.

In what follows, I present descriptive evidence on female participation and employment responses in the medium term. To this end, I rely on the main sample presented in Section 2. First, I assign to each male partner observed between 1996 and 2007 a correspondingly intense shock for 2002, according to his industry-occupation pair observed in t. Then, to capture the medium term participation and employment responses, I interact my measure of exposure to the convertibility collapse with the actual period of observation t. The reduced form specification in cross-section is given by:

$$Y^f_{h(i,o,r),t} = \sum_{t=1996q1}^{2007q4} a^t \left[Intensity^m_{i,o,2002} \times \mathbb{1}(\text{Period} = t) \right] + X^f_{h,t}\boldsymbol{a^f} + X^m_{h,t}\boldsymbol{a^m} + X_{h,t}\boldsymbol{a} + \phi_{r,t} + u_{h(i,o,r),t} \tag{7}$$

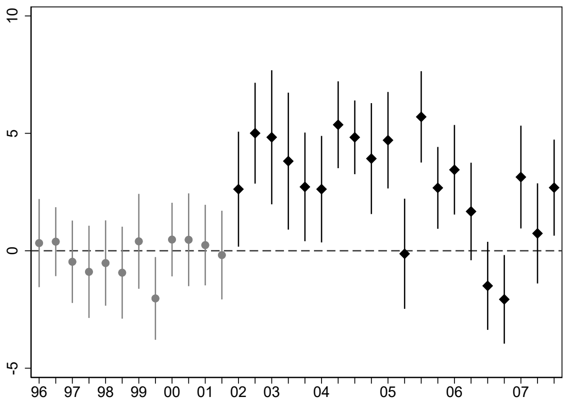
where $Y^f_{h(i,o,r),t}$ stands for both female participation and employment status, $Intensity^m_{i,o,2002}$ is the shock intensity assigned to the male partner in household h given his current industry and occupation (i,o) at period t, $\mathbb{1}(\text{Period} = t)$ is a dummy variable equal to 1 if the household is observed in t, 0 otherwise. $X^{m,f}$ are matrices with both partners' age (in years) and completed level of education (in four stages, as detailed above), $X_{h,t}$ are the usual time-varying household controls, $\phi_{r,t}$ are the province-by-period fixed effects, notably controlling for the evolution of the local labor market opportunities for men and women in time. $u_{h(i,o,r),t}$ is two-way clustered by household and by male industry-occupation pairs (because $Intensity^m_{i,o,2002}$ varies at the industry-occupation level).

Figure 1 plots $\hat{a}^t$. The first point estimate is the base effect of $Intensity^m_{i,o,2002}$ on female participation (1a) or employment (1b) in May 1996. At each period t, $\hat{a}^t$ captures the differential trend in women's participation, according to their male partners' degree of exposure to the 2002 shock, with respect to May 1996. When $t < 2002q2$, $\hat{a}^t$ is generally not significantly different from 0, confirming the absence of anticipation already discussed in the robustness section. $\hat{a}^{2002q2}$ and $\hat{a}^{2002q4}$ are positive, and measure the contemporaneous participation response to the shock.[28] In 2003 and beyond, a^t coefficients capture whether the female participation response lasts in the context of a very rapid economic recovery. I find this response to persist at least until 2005, and possibly even until the end of 2007, albeit smaller in size.

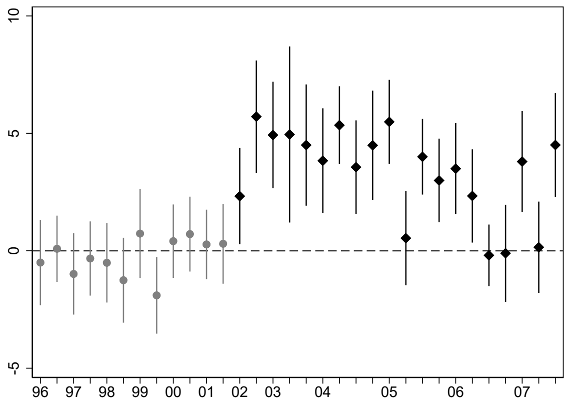
Discussion – An underlying hypothesis behind my interpretation of a^t is that I can rely on male partners' *contemporaneous* industry and occupation to infer the intensity of their exposure to the convertibility crisis in 2002. In other words, I suppose that the male partners observed in t are either working in exactly the same industry-occupation pair as they were when the convertibility collapsed in 2002, or in a very similarly exposed one.

28. For the sake of comparison, I also estimate the reduced form $Y^f_{h(i,o,r),t} = \rho Intensity^m_{i,o,2002} + \eta_{i,o} + X_{h,t}\boldsymbol{\beta} + c_h + \mu_{r,t} + \epsilon_{h(i,o,r),t}$ on the October 2000-October 2002 sample used in the main result section, where $Y^f_{h(i,o,r),t}$ is female participation. I find that $\hat{\rho} = 2.72$ (results not reported here). This gives additional credibility to the cross-sectional results displayed in Figure 1, as $\hat{\rho}$ lies between $\hat{a}^{2002q2} = 2.62$ and $\hat{a}^{2002q4} = 5.01$.

Figure 1: Medium Run Labor Market Response: Coefficients $\hat{a}^t$



(a) Participation



(b) Employment

Note: Figure (a) shows estimates from a single regression, where the dependent variable is female participation, and each coefficient and corresponding 90% confidence interval captures the deviation in participation explained by the different intensity of the shock experienced by male partners after the repeal of the convertibility law, taking May 1996 as a reference (Equation 7). Figure (b) presents the same results with regard to female employment. The sample used is the main sample described in Section 2. In the last two columns, the sample is restricted to couples where the women are working. I estimate the model using pooled OLS with period-by-province fixed effects, and control for all the demographic characteristics and time-varying controls listed in Table Table A1, as well as for the amounts of household income listed in Table A2. Standard errors are two-way clustered by household, and by male industry-occupation pairs, because the shock exposure varies at the industry-occupation level.

Although I cannot test this hypothesis directly, I find that the correlation between lagged and present male exposure in the panel is high and positive (0.66). Even so, a positive $\hat{a}^t$ beyond 2003 need not be interpretable as reflecting an *attachment* to the labor market. Women may simply increase their participation over a longer period of time in order to compensate for the deep income loss incurred in 2002. To shed additional light on the eventual compensation mechanism, I test whether the female participation response in 2003 and beyond is higher for women partnered with lower-income men, holding the 2002 shock intensity constant. If the persistence is due to compensation, rather than attachment, then a woman's participation should be higher when her partner's labor income in t is low, and this negative relation should increase over time, because only women partnered with the lowest earners would remain active. Adding to specification (7) a full set of interactions between the $\mathbb{1}(\text{Period} = t)$ dummies, $Intensity^m_{i,o,2002}$, and the real log labor income reported by the male partner, I did not find any evidence in support of compensation.

7 Conclusion

Employment is one of the most important elements in women's empowerment, and ultimately in economic development. In this paper, I find that subsistence is still an important motive for female participation, even in the context of an upper-middle-income country like Argentina, where fertility is declining and female education is expanding. Between 1996 and 2007, the correlation between partners' labor outcomes is close: when her partner becomes unemployed, a woman is 6 percentage points more likely than before to enter the labor market. Decomposing this effect, I find that one third of new female entrants find a full-time job, one third settles for a part-time job, and the remainder are unemployed. Then, to properly measure the causal participation response to a negative, unanticipated shock on male employment, I rely on the sudden repeal of the convertibility law in January 2002, which had been the foundation of the Argentine economic system since 1991. I demonstrate that this unexpected legal decision caused a specific change in male employment relative to the underlying employment trend, differing across the 22×4 industry-occupation pairs, which I used as an exogenous source of variation to instrument the actual job loss experienced by male partners around January 2002. In this context, I find that the female labor supply response is very high at the extensive margin: when unemployment hits her partner, a woman is twice as likely to enter the labor market and work. The rapid expansion of a poverty alleviation workfare program in May 2002 contributes to a relaxation of the constraints on the labor demand, and accounts for half of the effect. At the intensive margin, women do not actually increase their working time, despite the fact that they vociferously declare that they would rather work longer hours, signaling either institutional rigidity in working hours, or constraints on the demand side. Using a similar approach for couples where the male partner remains employed, I find that the female labor supply also responds positively to a negative shock on earnings. I do not find evidence that this negative income effect fades out over time between 1996 and 2007.

The other important contribution of this paper is to shed light on the potential impact of such a labor supply response in the medium term. First, this response is for the most part contemporaneous to the negative shock, but I also find a significant delayed reaction, suggesting that adjustments take time. Then, contrary to expectations, while women *enter* the labor market as their partner loses his job, I find no symmetric evidence that they *withdraw* from the labor market upon their partner's re-employment. Finally, consistently with this latter finding, the repeal of the convertibility law appears to have had persistent impacts on female participation. Therefore, even responses to temporary negative shocks probably have lasting consequences for women's labor force participation.

To conclude, the most recent contributions in the literature have made very important progress in the identification of the structural barriers to female employment, as well as in the design of innovative policies to address them (see Jayachandran [2020] for a review). I highlight an additional mechanism, where a woman's participation can be explained by subsistence, even in countries with intermediate level of income. My paper contributes to the broader understanding of female labor force participation by proposing an in-depth exploration of female participation and employment responses in the context of negative shocks, and by shedding light on their role in the dynamics of the female labor force in the longer term. Investigating the participation for the sake of subsistence is important, since it also highlights the potential negative consequences of the labor supply on women, such as reduced leisure, or the double burden of caring and work, often overlooked in the literature [Heath and Jayachandran, 2017]. This is an important direction for future research.

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A Appendix: Data & Baseline

Table A1: Summary Statistics: Demographics

	1996-1999	2000-2002	2003-2007
Female Demographics			
Age	39.02	39.23	39.49
	(9.75)	(9.83)	(9.90)
Level of education achieved			
No education	0.11	0.09	0.07
	(0.31)	(0.28)	(0.25)
Primary	0.50	0.48	0.42
	(0.50)	(0.50)	(0.49)
Secondary	0.27	0.28	0.32
	(0.44)	(0.45)	(0.47)
Tertiary	0.13	0.15	0.19
	(0.34)	(0.36)	(0.40)
Male Demographics			
Age	41.81	41.91	42.21
	(10.04)	(10.07)	(10.21)
Level of education achieved			
No education	0.11	0.09	0.07
	(0.31)	(0.29)	(0.26)
Primary	0.53	0.52	0.47
	(0.50)	(0.50)	(0.50)
Secondary	0.25	0.27	0.31
	(0.43)	(0.45)	(0.46)
Tertiary	0.11	0.12	0.15
	(0.31)	(0.32)	(0.36)
Household Demographics			
Number of children 0-2	0.18	0.17	0.15
	(0.40)	(0.39)	(0.37)
Number of children 3-5	0.29	0.27	0.25
	(0.53)	(0.52)	(0.49)
Number of children 6-15	0.97	0.99	0.88
	(1.13)	(1.17)	(1.08)
Number of children 15+ studying	0.36	0.40	0.38
	(0.68)	(0.72)	(0.71)
Number of adults 16-64	0.42	0.41	0.40
	(0.83)	(0.83)	(0.81)
Number of adults 65+	0.05	0.05	0.04
	(0.24)	(0.23)	(0.20)
Observations	89,475	53,751	100,014

Note: The sample used is the main sample described in Section 2, divided into three periods. The core analysis focuses on the period 2000-2002.

Table A2: Summary Statistics: Income Sources

	1996-1999	2000-2002	2003-2007
Household income			
Other Labor Income > 0	0.23	0.21	0.21
	(0.42)	(0.41)	(0.41)
Pensions > 0	0.06	0.06	0.06
	(0.25)	(0.23)	(0.23)
Capital > 0	0.01	0.01	0.02
	(0.11)	(0.11)	(0.15)
Unemployment > 0	0.01	0.01	0.01
	(0.10)	(0.12)	(0.10)
Remittances > 0	0.01	0.02	0.02
	(0.12)	(0.15)	(0.15)
Other > 0	0.01	0.03	0.08
	(0.12)	(0.18)	(0.27)
Observations	89,475	53,751	100,014
Amount when positive			
Other Labor Income	564.73	457.95	447.15
	(525.21)	(434.53)	(393.92)
Pensions	318.83	299.86	294.59
	(314.29)	(286.67)	(315.22)
Capital	580.36	510.62	495.64
	(657.85)	(906.54)	(2,325.10)
Unemployment	769.26	373.57	508.26
	(2,817.77)	(740.02)	(1,534.14)
Remittances	197.90	160.95	169.59
	(191.84)	(177.88)	(401.39)
Other	340.43	218.10	142.81
	(1,349.30)	(792.86)	(746.06)

Note: The sample used is the main sample described in Section 2, divided into three periods. The core analysis focuses on the period 2000-2002.

Table A3: Summary Statistics (Means) on Female Partners, by Employment Status

	All women		Working once or more	
	Not working	Working	Discontinuously	Continuously
Own Demographics				
Age	38.318	39.128***	37.972	39.454***
Level of education achieved				
No education	0.118	0.064***	0.109	0.049***
Primary	0.547	0.352***	0.525	0.299***
Secondary	0.280	0.309***	0.280	0.319***
Tertiary	0.054	0.275***	0.087	0.332***
Partner's Demographics				
Age	41.358	41.735***	40.901	41.968***
Level of education achieved				
No education	0.115	0.077***	0.113	0.065***
Primary	0.556	0.429***	0.556	0.389***
Secondary	0.258	0.327***	0.256	0.351***
Tertiary	0.072	0.166***	0.075	0.195***
Household composition				
N children 0-2	0.209	0.141***	0.175	0.131***
N children 3-5	0.341	0.252***	0.308	0.238***
N children 6-15	1.097	0.987***	1.131	0.940***
N children 15+ studying	0.372	0.398***	0.369	0.407***
N adults 16-64	0.455	0.348***	0.438	0.317***
N adults 65+	0.040	0.041	0.039	0.042**
Household income				
Other Labor Income > 0	0.206	0.189***	0.211	0.176***
Pensions > 0	0.074	0.051***	0.059	0.050***
Capital > 0	0.015	0.019***	0.014	0.020***
Unemployment > 0	0.010	0.007***	0.012	0.006***
Remittances > 0	0.019	0.014***	0.021	0.012***
Other > 0	0.057	0.040***	0.069	0.034***
Observations	131,138	112,102	58,385	83,815

Note: *** denotes significance at 1% of the difference in means between women working/not working (col 1-2), and then, using the panel dimension, between women continuously employed and women temporarily employed (col 3-4). ** denotes the difference at 5%, and * at 10%. The sample used is the main sample described in Section 2.

B Appendix: Instrument

Industries in the *EPH* – Over 1996–2007, *INDEC* used two different classifications to characterize industries in the *EPH*.

- Between 1996 and 2002, the survey relied on the 3-digit ISIC Rev.3 classification provided by the ILO.
- Between 2003 and 2007, the survey turned to a 4-digit structure and adopted the *CAES* Mercosur 2000 classification.

In this paper, I make use of the first 2 digits of both industry classifications. The main advantage is that they correspond to exactly the same broad industries across classifications, except those for the wholesale and retail trade, which are merged in the *CAES* Mercosur 2000 [INDEC, 2011]. Therefore, in the core of the paper, I consider the following 22 industries: (1) Agriculture, (2) Food industry, (3) Textile, (4) Chemistry, (5) Metal, (6) Other goods, (7) Energy/water, (8) Construction, (9) Wholesale trade, (10) Retail trade, (11) Accommodation/food services, (12) Transportation, (13) Communication, (14) Finance –Insurance, (15) Real estate, (16) Public administration, (17) Education, (18) Health, (19) Other social-local services, (20) Repair, (21) Domestic services, (22) Other services. When I turn to the dynamics of participation, I merge sectors (9) and (10) into a single 'Trade' sector.

Occupations in the *EPH* – Over 1996-2007, *INDEC* used three different classifications to define occupations, which differed in many ways from the international standard ISCO. All these classifications have in common that they are multidimensional, and notably characterize the qualifications for a given job. Qualification reflects the complexity of the tasks in each job. It is measured according to the combination of activities or actions performed, the instruments used, and the management of raw materials. It reflects the knowledge and skills required to perform the tasks, and therefore helps to assess the degree of complexity of an occupation. Since its origin, qualification in the *EPH* has been coded as a number between 1 and 4, and its significance has remained remarkably stable over time.

- Originally, the *Clasificador de Ocupaciones* of the *EPH* (*CO-EPH*) was *INDEC*'s own classificatory instrument for occupations. This two-digit classification was bidimensional and assessed the purpose of the work (character) and the complexity of the tasks performed (qualification). Qualification had four categories: professional, qualified, semi-qualified and unqualified [INDEC, 2000].

- Next, the *Clasificador Nacional de Ocupaciones 1991* (*CNO-91*) was introduced with a double objective: to update and unify the measurement of occupations and to ease international comparisons, notably with ISCO. The new classification combined four dimensions (general character, technology, hierarchy and complexity) in a three-digit structure. The last digit measured qualification in four categories: professional, technical, clerical/operational and elementary. The classification was gradually implemented until October 1997. In May 1996, the most populated provinces had already adopted the *CNO-91* instead of *CO-EPH*.[29] The remaining provinces opted for *CNO-91* between 1996 and 1997 [INDEC, 2000].[30]

- In 2001, the *Clasificador Nacional de Ocupaciones 2001* (*CNO-01*) further disentangled the types of occupations with a 5-digit structure [INDEC, 2001]. The last digit, qualification, kept the same definition as in *CNO-91*. All the provinces adopted the CNO-01 in 2003.

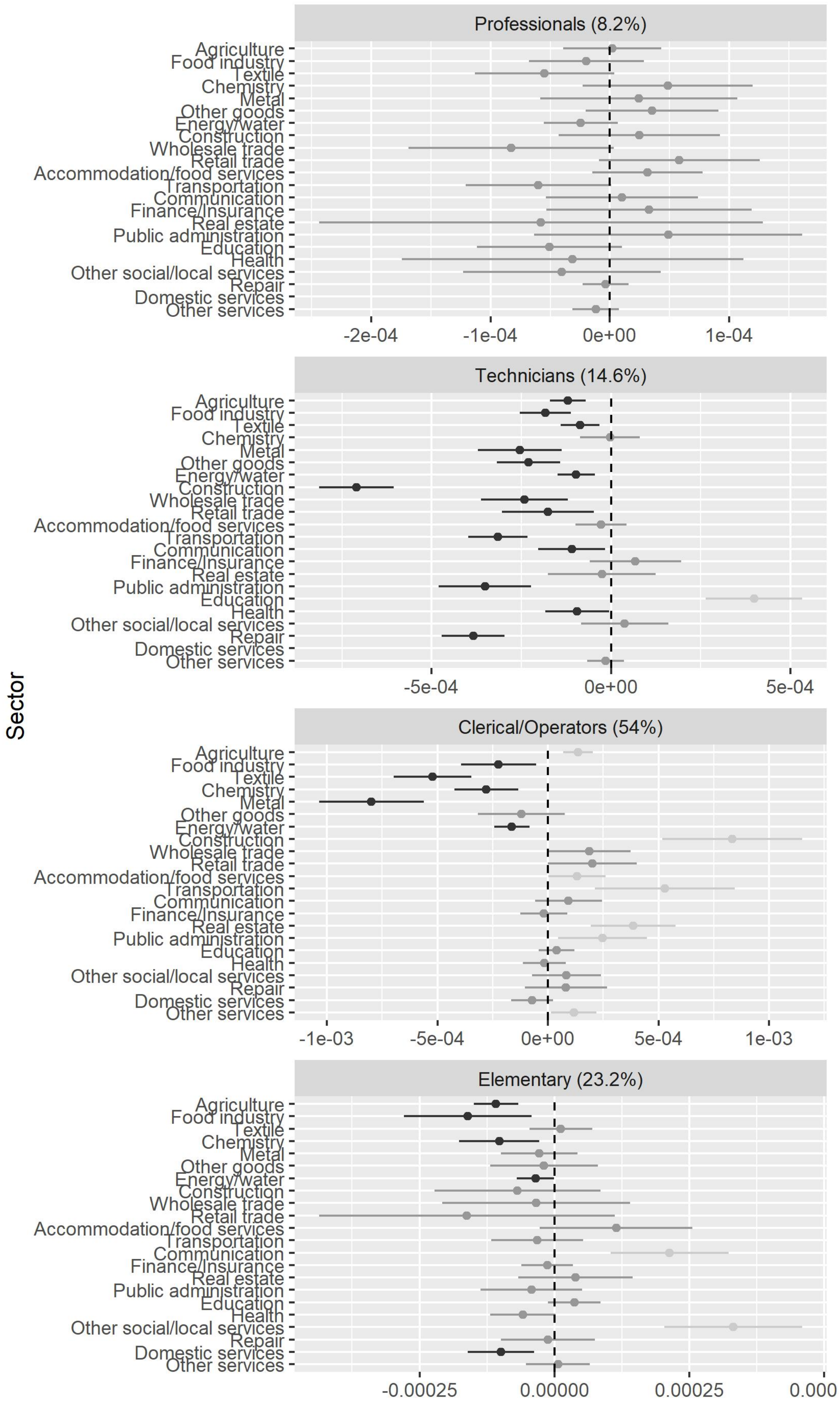
In this paper, following *INDEC*'s recommendations [INDEC, 2000], I adopt the following simple equivalence:

CO-EPH	*CNO-91*	*CNO-01*	Digit
Professional	Professional	Professional	1
Qualified	Technical	Technical	2
Semi-qualified	Clerical/operational	Clerical/operational	3
Unqualified	Elementary	Elementary	4

29. Ciudad de Buenos Aires, Partidos de Buenos Aires, Gran La Plata, Gran Rosario, Santa Fe, Paraná, Gran Mendoza, Gran Córdoba, Neuquén, Tucumán, Santa Rosa, Mar del Plata, Río Cuarto.

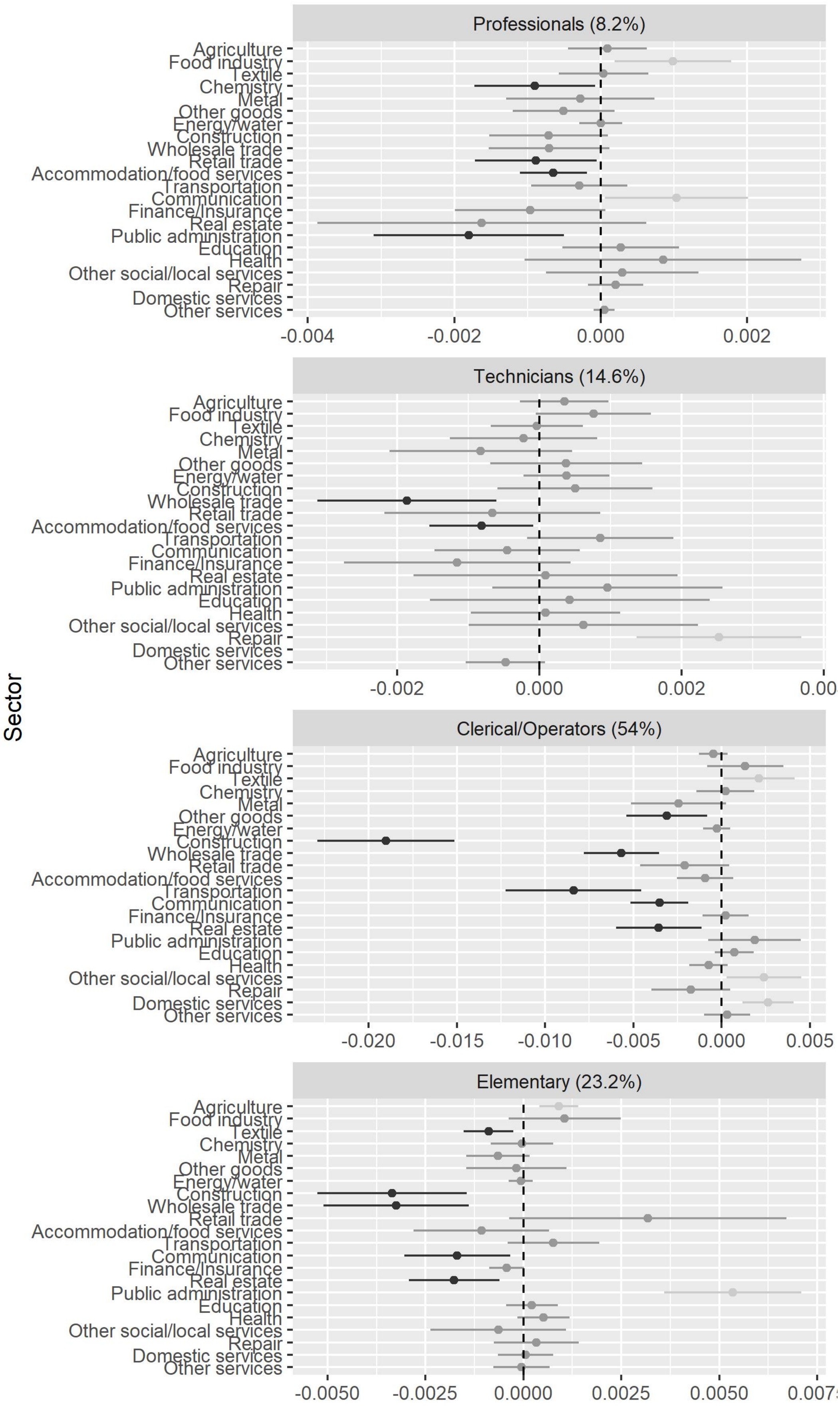
30. Santiago del Estero y La Banda, Jujuy y Palpalá, Gran Catamarca, La Rioja joined in October 1996 ; Gran Resistencia, Corrientes, Formosa, Salta in May 1997 ; Bahía Blanca, Posadas, Comodoro Rivadavia, Río Gallegos, San Luis y el Chorillo, Gran San Juan, Tierra del Fuego in October 1997.

Figure B1: Trend Coefficients $\hat{a}^{i,o}$



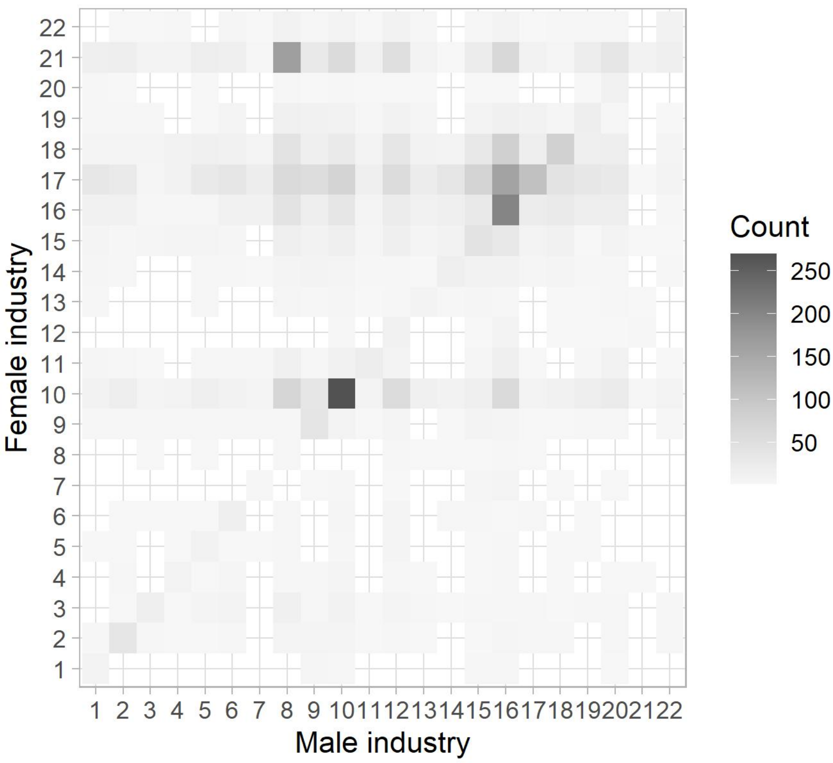
Note: The sample used consists of men of working age in urban Argentina between 1996 and 2002. For each industry-occupation pair, I run the following regression: $E^{i,o}_{j,r,t} = a^{i,o} Trend_{j,t} + b^{i,o} Post_t + \phi^{i,o}_r + \epsilon^{i,o}_{j,r,t}$. The figure displays $\hat{a}^{i,o}$ estimates from these separate regressions, and their corresponding 90% confidence interval. The significantly negative coefficients are displayed in black, the significantly positive coefficients in light grey, and the insignificant coefficients in grey.

Figure B2: Shock Coefficients $\hat{b}^{i,o}$



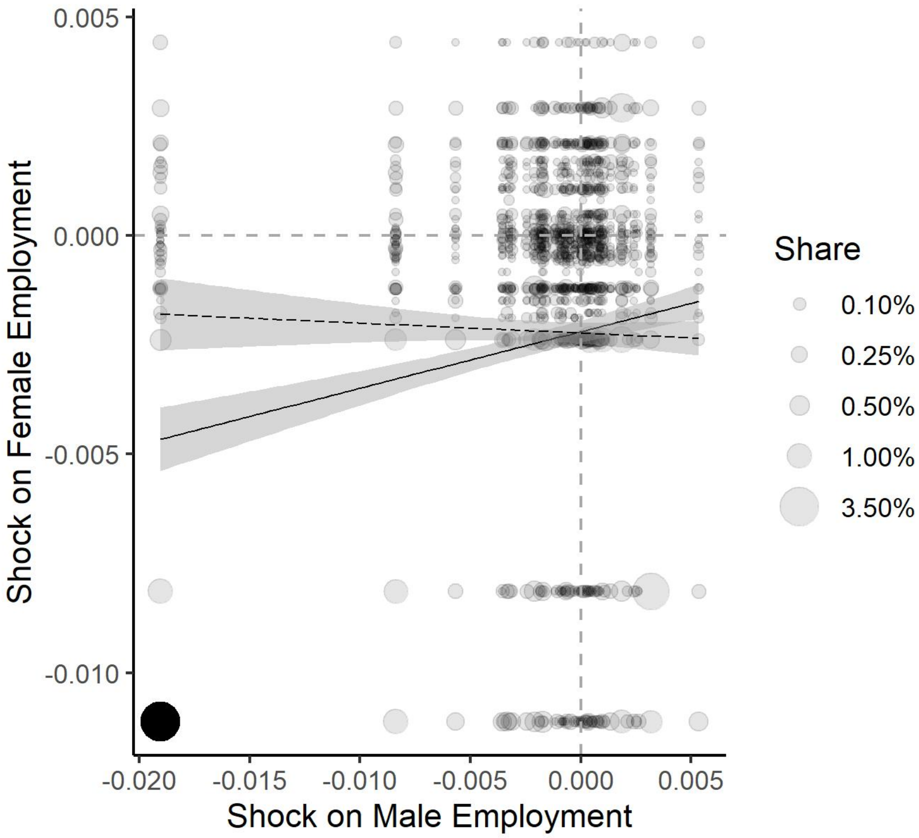
Note: The sample used consists of men of working age in urban Argentina between 1996 and 2002. For each industry-occupation pair, I run the following regression: $E^{i,o}_{j,r,t} = a^{i,o}Trend_{j,t} + b^{i,o}Post_t + \phi^{i,o}_r + \epsilon^{i,o}_{j,r,t}$. The figure displays $\hat{b}^{i,o}$ estimates from these separate regressions, and their corresponding 90% confidence interval. The significantly negative coefficients are displayed in black, the significantly positive coefficients in light grey, and the insignificant coefficients in grey.

Figure B3: Partners' Industry in October 2001



Note: Figure B3 displays the frequency of each possible combination of industry-occupation pairs for couples where both partners are active. Industries: (1) Agriculture, (2) Food industry, (3) Textile, (4) Chemistry, (5) Metal, (6) Other goods, (7) Energy/water, (8) Construction, (9) Wholesale trade, (10) Retail trade, (11) Accommodation/food services, (12) Transportation, (13) Communication, (14) Finance – Insurance, (15) Real estate, (16) Public administration, (17) Education, (18) Health, (19) Other social-local services, (20) Repair, (21) Domestic services, (22) Other services.

Figure B4: Partners' Future Shock given their Industry in October 2001



Note: Figure B4 plots the possible combinations of male-female employment shocks for couples where both partners are active, together with their share in my sample. The plain line displays the slope of a regression of women's shock intensity on their partner's shock intensity, weighted by their share. The dashed line displays the slope obtained from the same regression, but excluding the most prevalent combination – construction operators living with unqualified domestic workers.

Table B1: First Stage Estimates

	Extensive Margin (All Women)	Intensive Margin (Working Women)
	Male partner is unemployed	Male partner is unemployed
Shock Intensity	5.611***	4.441***
	(0.486)	(0.815)
Composition		
N children 0-2	0.00117	0.0338*
	(0.012)	(0.020)
N children 3-5	-0.0141	0.00374
	(0.016)	(0.036)
N children 6-15	-0.0109	0.00737
	(0.011)	(0.024)
N children 15+ studying	-0.00777	-0.00916
	(0.014)	(0.013)
N adult 16-64	-0.00182	-0.00677
	(0.011)	(0.015)
N adult 65+	0.0765***	0.0165
	(0.029)	(0.063)
Household Income		
Labor	-0.00426**	-0.00679**
	(0.002)	(0.003)
Pensions	0.00109	0.000818
	(0.005)	(0.007)
Capital	0.00835**	0.00920
	(0.004)	(0.008)
Unemployment	0.0697***	0.0742***
	(0.007)	(0.018)
Remittances	0.0446***	0.0113
	(0.006)	(0.010)
Other	0.0307***	0.0193***
	(0.005)	(0.005)
Household FE	Yes	Yes
Region × Time FE	Yes	Yes
Industry × Occupation FE	Yes	Yes
KP	133.13	29.68
Mean(Intensity)	0.00153	0.00130
SD(Intensity)	0.00499	0.00464
Observations	20,071	7,720
Unique households	6,268	2,642

Note: See Table 3.

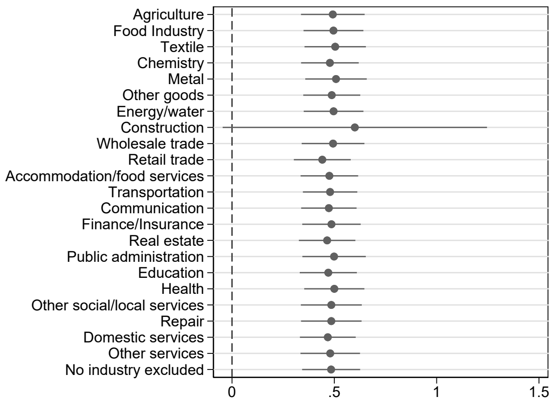
Table B2: 2SLS Estimates

	Extensive Margin (All Women)			Intensive Margin (Working Women)	
	Particip.	Empl.	Full T.	Wish	Hours
Male partner is unemployed	0.484***	0.492***	-0.0677	0.543**	-12.63
	(0.085)	(0.097)	(0.076)	(0.268)	(9.132)
Composition					
N children 0-2	0.0109	0.00996	0.0203*	0.0209	3.165***
	(0.014)	(0.012)	(0.012)	(0.027)	(1.059)
N children 3-5	0.0238	0.0249*	0.0246*	0.0248	2.286*
	(0.015)	(0.013)	(0.014)	(0.037)	(1.243)
N children 6-15	0.0252	0.0267*	0.0109	0.0172	0.137
	(0.016)	(0.014)	(0.010)	(0.027)	(1.071)
N children 15+ studying	0.0146	0.0252	0.00984	0.0373	0.420
	(0.018)	(0.016)	(0.011)	(0.023)	(0.923)
N adult 16-64	0.00761	0.0124	-0.00114	0.0564***	-0.484
	(0.012)	(0.010)	(0.008)	(0.019)	(0.835)
N adult 65+	0.0192	-0.0133	-0.0135	-0.0282	0.837
	(0.051)	(0.045)	(0.031)	(0.089)	(2.273)
Household Income					
Labor	0.0116***	0.0102***	0.00240	-0.00185	-0.104
	(0.002)	(0.003)	(0.002)	(0.004)	(0.187)
Pensions	-0.0181***	-0.0174***	-0.0142***	-0.00775	-0.482
	(0.005)	(0.006)	(0.004)	(0.012)	(0.384)
Capital	0.00686	0.00811*	0.00393	-0.0200**	-0.0181
	(0.006)	(0.005)	(0.004)	(0.010)	(0.351)
Unemployment	-0.0377***	-0.0461***	0.00101	-0.0611***	1.298*
	(0.007)	(0.007)	(0.008)	(0.023)	(0.669)
Remittances	-0.0230***	-0.0360***	-0.0126**	0.0143	-0.214
	(0.007)	(0.008)	(0.006)	(0.018)	(0.505)
Other	-0.0235***	-0.0273***	-0.00434	-0.00830	-0.118
	(0.006)	(0.008)	(0.004)	(0.009)	(0.293)
Household FE	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes
Industry × Occupation FE	Yes	Yes	Yes	Yes	Yes
$\bar{Y}$ if Husb. Employed	0.475	0.426	0.296	0.280	33.587
Observations	20,071	20,071	20,071	7,720	7,720
Unique households	6,268	6,268	6,268	2,642	2,642

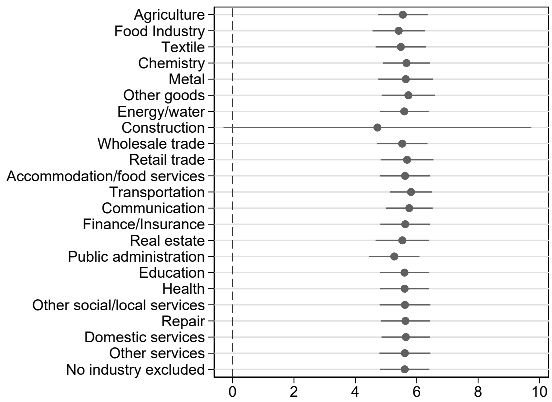
Note: See Table 4.

C Appendix: Robustness & Discussion

Figure C1: 2SLS Results, Excluding one Industry at a Time



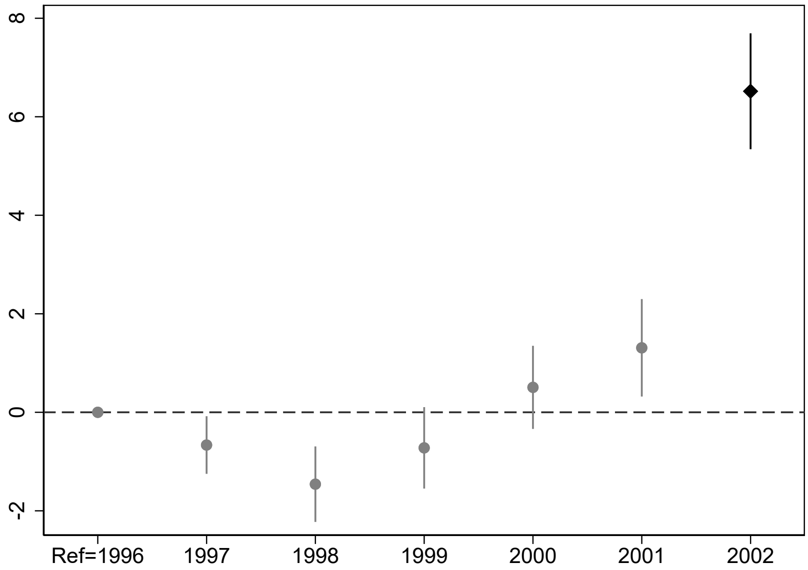
(a) Second Stage Results – Participation Response to Male Job Loss



(b) First Stage Results – Job Loss Response to Shock

Note: The figures show estimates from 23 different regressions. Each coefficient and its corresponding 90% confidence interval come from a separate regression, where I estimate the set of equations (3) excluding each of the 22 industries one by one. For each estimation, the excluded industry is displayed on the y-axis. As a benchmark, I also report the main estimates of Tables 3 and 4, where none of the industries is excluded. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility.

Figure C2: Response of Male Unemployment to the 2002 Shock Intensity: $\hat{\pi}^y$



Note: Figure C2 shows estimates from a single regression, where the dependent variable is male unemployment, and each coefficient and corresponding 90% confidence interval captures the deviation in unemployment explained by the different intensity of the shock experienced by male partners after the repeal of the convertibility law, taking May 1996 as a reference (Equation 4). Circles (diamonds) denote differential time effects before (after) January 2002. The specification controls for household fixed effects, period-by-province fixed effects, and industry-occupation fixed effects. S.e. clustered by industry-occupation pairs are shown in parentheses. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility.

Table C1: Baseline Specification – Replication

	Extensive Margin (All Women)			Intensive Margin (Working Women)	
	Particip.	Empl.	Full T.	Hours	Wish
Male partner is unemployed	0.0607***	0.0457***	0.0219**	0.659	0.0699***
	(0.012)	(0.011)	(0.010)	(0.611)	(0.023)
Composition					
N children 0-2	0.0133	0.0119	0.0208**	2.709***	0.0282
	(0.013)	(0.012)	(0.010)	(0.859)	(0.027)
N children 3-5	0.0165	0.0179	0.0256**	2.344**	0.0234
	(0.015)	(0.013)	(0.012)	(0.956)	(0.031)
N children 6-15	0.0196	0.0209*	0.0120	0.250	0.0178
	(0.013)	(0.012)	(0.011)	(0.863)	(0.026)
N children 15+ studying	0.0109	0.0215*	0.0103	0.685	0.0322
	(0.012)	(0.011)	(0.011)	(0.790)	(0.024)
N adult 16-64	0.00777	0.0122	-0.00189	-0.288	0.0547***
	(0.010)	(0.010)	(0.009)	(0.745)	(0.020)
N adult 65+	0.0507	0.0179	-0.0190	0.400	-0.0140
	(0.039)	(0.037)	(0.030)	(1.965)	(0.080)
Household Income					
Labor	0.00952***	0.00826***	0.00296	-0.0207	-0.00532
	(0.002)	(0.002)	(0.002)	(0.150)	(0.004)
Pensions	-0.0187***	-0.0174***	-0.0148***	-0.455	-0.00796
	(0.005)	(0.005)	(0.004)	(0.345)	(0.011)
Capital	0.0101*	0.0118**	0.00340	-0.147	-0.0168*
	(0.005)	(0.005)	(0.004)	(0.338)	(0.009)
Unemployment	-0.00786	-0.0144***	-0.00529	0.330	-0.0273***
	(0.005)	(0.005)	(0.005)	(0.318)	(0.010)
Remittances	-0.00457	-0.0163***	-0.0164***	-0.338	0.0169
	(0.006)	(0.005)	(0.004)	(0.426)	(0.012)
Other	-0.0102***	-0.0133***	-0.00696***	-0.313*	0.000238
	(0.003)	(0.003)	(0.002)	(0.172)	(0.007)
Household FE	Yes	Yes	Yes	Yes	Yes
Region × Time FE	Yes	Yes	Yes	Yes	Yes
R^2	0.731	0.764	0.760	0.767	0.597
$\bar{Y}$ if Husb. Employed	0.475	0.426	0.296	33.587	0.280
Observations	20,071	20,071	20,071	7,720	7,720
Unique households	6,268	6,268	6,268	2,642	2,642

Note: This table replicates Table 2 with the subsample of couples observed around the repeal of convertibility, used in the core results section (Section 4). *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by households are shown in parentheses.

Table C2: Panel Attrition

	Dep. var. = 1 if household drops out in t+1			
	(1)	(2)	(3)	(4)
2002 Shock intensity	0.119	-0.171	0.363	-0.208
	(0.454)	(0.250)	(0.484)	(0.301)
Composition				
N children 0-2	0.0151*	0.00511	0.0280***	-0.0127***
	(0.007)	(0.004)	(0.008)	(0.005)
N children 3-5	0.0226**	-0.00224	0.0463***	-0.0183***
	(0.010)	(0.003)	(0.011)	(0.004)
N children 6-15	0.0285***	-0.00484***	0.0381***	-0.0200***
	(0.008)	(0.001)	(0.008)	(0.002)
N children 15+ studying	0.0119*	-0.000650	0.00670	-0.0119***
	(0.006)	(0.002)	(0.006)	(0.004)
N adult 16-64	0.00486	-0.00323*	0.00214	-0.00502
	(0.005)	(0.002)	(0.007)	(0.003)
N adult 65+	-0.0239	-0.0181**	-0.0407	-0.0380***
	(0.022)	(0.008)	(0.024)	(0.008)
Household Income				
Labor	0.00219**	0.000649	0.000877	-0.00157
	(0.001)	(0.001)	(0.001)	(0.001)
Pensions	-0.00127	-0.000153	-0.00267	0.000134
	(0.003)	(0.001)	(0.003)	(0.001)
Capital	-0.00101	-0.000358	0.00296	0.00344
	(0.003)	(0.002)	(0.004)	(0.004)
Unemployment	-0.00123	-0.00265	-0.00309	0.0000378
	(0.004)	(0.004)	(0.002)	(0.003)
Remittances	0.00177	0.00210	-0.000147	0.00618**
	(0.003)	(0.003)	(0.003)	(0.003)
Other	-0.000864	-0.000182	0.000239	0.00185
	(0.001)	(0.001)	(0.002)	(0.001)
Region $\times$ Time FE	Yes	Yes	Yes	Yes
Household FE	Yes	No	Yes	No
Demog. Controls	No	Yes	No	Yes
$\bar{Y}$	0.062	0.062	0.085	0.179
Proportion of couples interviewed < 4	0.200	0.200	0.260	0.369
Observations	20,071	20,071	35,771	47,654

Note: Columns (1) and (2) of Table C2 display estimates based on the main sample described in Section 4. Column (1) presents the results of an estimation where the probability of dropping out of the sample depends on within-household variation in the exposure to the shock intensity measured in 2002, controlling for time-varying household characteristics, as well as household and province-by-period fixed effects. Column (2) presents the results of a pooled OLS regression, without the household fixed effects, but with standard demographic controls: the age of both spouses in years and their education split into 4 stages: none, primary, secondary, and tertiary. The probability of dropping out depends on the variation in the exposure to the shock intensity measured in 2002 across households. Columns (3) and (4) take the attrition issue upstream, and consider the sample of households surveyed between October 2000 and October 2002, before I imposed the condition that partners should be observed at least once in October 2001 or before, and once after. The difference in the number of observations comes from the fact that singleton households are excluded from the regression in Column (3), but not from that in Column (4).

Table C3: Accounting for *Jefes y jefas* – First Stage Estimates

	Extensive Margin (All Women)	Intensive Margin (Working Women)
	Male partner is unemployed	Male partner is unemployed
Shock Intensity	5.578***	4.563***
	(0.489)	(0.891)
Workfare Beneficiary	0.00902	-0.0286
	(0.019)	(0.038)
Household FE	Yes	Yes
Region × Time FE	Yes	Yes
Industry × Occupation FE	Yes	Yes
Time-varying controls	Yes	Yes
F-stat	130.02	26.24
Mean(Intensity)	0.00153	0.00130
SD(Intensity)	0.00499	0.00464
Observations	20,071	7,720
Unique households	6,268	2,642

Note: Table C3 replicates Table 3, adding the female participation to the workfare program *Jefes y jefas* as a control. *** denotes significance at 1%, ** at 5%, and * at 10%. S.e. clustered by industry-occupation pairs are shown in parentheses. The sample used is the sample described in Section 4, that is, the sample of couples observed at least once in October 2001 or before, and once after the repeal of convertibility. In the last two columns, the sample is restricted to couples where women report working at least once before and after January 2002. The time-varying controls are the household composition variables listed in the bottom panel of Table A1, as well as the amounts of household income listed in Table A2.