

EARLY LABOR MARKET PROSPECTS AND FAMILY FORMATION: ONLINE APPENDIX

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Online Appendix

Table of Contents

A	Policy Pilot	3
A.1	Educational slots	3
A.2	Spatial roll-out of the reform	3
B	Data and construction of variables	8
B.1	Labor-market and education	8
B.2	Partner’s GPA	8
C	Additional Descriptive Statistics	9
D	Additional Results	10
D.1	Short and long-term impact of unemployment rate at graduation on selected outcome	10
D.2	Comparing IV to alternative OLS specifications	11
E	Robustness checks	16
E.1	Alternative Definitions of Earnings and Employment	16
E.2	Dichotomized measures of partner’s quality based on GPA	16
E.3	Evaluating the Identifying Assumptions for the Bottom GPA Quintile	17
E.4	Robustness Checks on the Model and Sample	18
E.5	Sibling Fixed-Effects	21
E.6	Reduced Form of the Model	22
F	A stylized model of partnership choice	24
G	Results for Males	28
G.1	Males graduating in the recession	28
G.2	Partner pool	32

A Policy Pilot

A.1 Educational slots

Table A1: Number of available educational slots by program type and year.

Year of enrollment	1987	1988	1989	1990	1991 ^(a)
Electrical engineering	48	528	656	776	n/a
Health care	46	2 182	2 918	3 072	n/a
Heating, ventilation and sanitation	64	64	72	104	n/a
Industry	352	1 608	1 952	1 968	n/a
Business and services		210	660	990	n/a
Caring services: children and youth		256	420	420	n/a
Construction		296	408	432	n/a
Textile and clothing manufacturing		136	208	224	n/a
Transport and vehicle engineering		752	992	1 056	n/a
Use of natural resources		352	640	720	n/a
Constructional metalwork			56	56	n/a
Food manufacturing			224	256	n/a
Handicraft			32	64	n/a
Painting			56	88	n/a
Process technology			176	208	n/a
Restaurant			336	416	n/a
Wood technology			144	168	n/a
Graphic				112	n/a
Total	510	6 384	9 950	11 130	12 818

Notes: Table adapted from Grönqvist and Hall (2011). ^(a) There is no available statistics on the number of slots per program type for 1991. The total number of students, however, is reported in a report from Skolverket (1992).

A.2 Spatial roll-out of the reform

In Figure 1 we graphically display the roll-out of the policy pilot at the municipal level. In 1987 the number of participating municipalities was low (in comparison with later years). 22 out 284 municipalities came to participate. The participating municipalities were scattered across the country – from north to south and east to west. The following year (1988) 91 new municipalities was included in the policy scheme, once again, they

appear to be spread across all regions of Sweden. The only exception is the sparsely populated northern part of Sweden (Norrland), with a more limited number of participating municipalities. In 1989, 31 more municipalities became part of the policy pilot, and again, the new municipalities can be found in various parts of Sweden. The final two years of the pilot (1990 and 1991) very few new municipalities (8 in total) entered but just as in the early years of the pilot they are found across the country.

Another way of illustrating the geographical roll-out can be found in Figure 2 where blue areas represent all municipalities that entered the policy pilot between 1987-1991. Based on eyeballing our reading of the geographical roll-out is that it can be considered quasi-random. This pattern is also in line with the fact that "the Government stipulated that it [the pilot scheme] should be distributed among regions with different industry and population structure" (Hall (2009), p. 9).

More formally we test for the possibility of spatial correlation in the rollout of the pilot scheme by calculating Moran's I, a measure of spatial autocorrelation. We start out by creating a spatial-weighting matrix (contiguity matrix) W and an inverse distance matrix M based on a shape-file containing all Swedish municipalities. Next, we merge a data set containing a treatment indicator for all municipalities. The indicator is set to 1 if a municipality came to participate at any point during the policy pilot. In Table 1 (panel A) we present the result of this exercise. All test gives a p-values > 0.05 , indicating the tests were not significant. Similarly, we run the same test on the fraction of years that a municipality participated in the policy pilot between 1987-1991 (see panel B); comfortably the test remains insignificant.

The overall impression we get from these exercises is that there is little evidence that point in the direction of spatial correlations in the rollout of the pilot scheme is a threat to the validity of our research design.

Figure 1: Roll-out of reform by year of entry

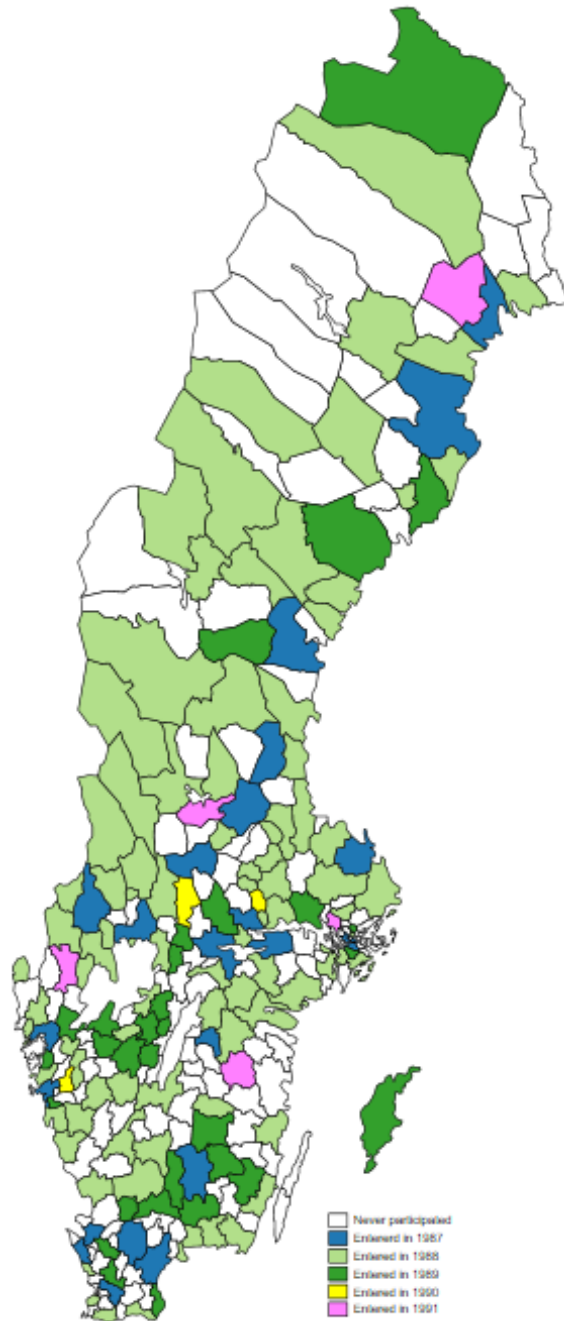


Figure 2: Roll-out of reform - ever participated

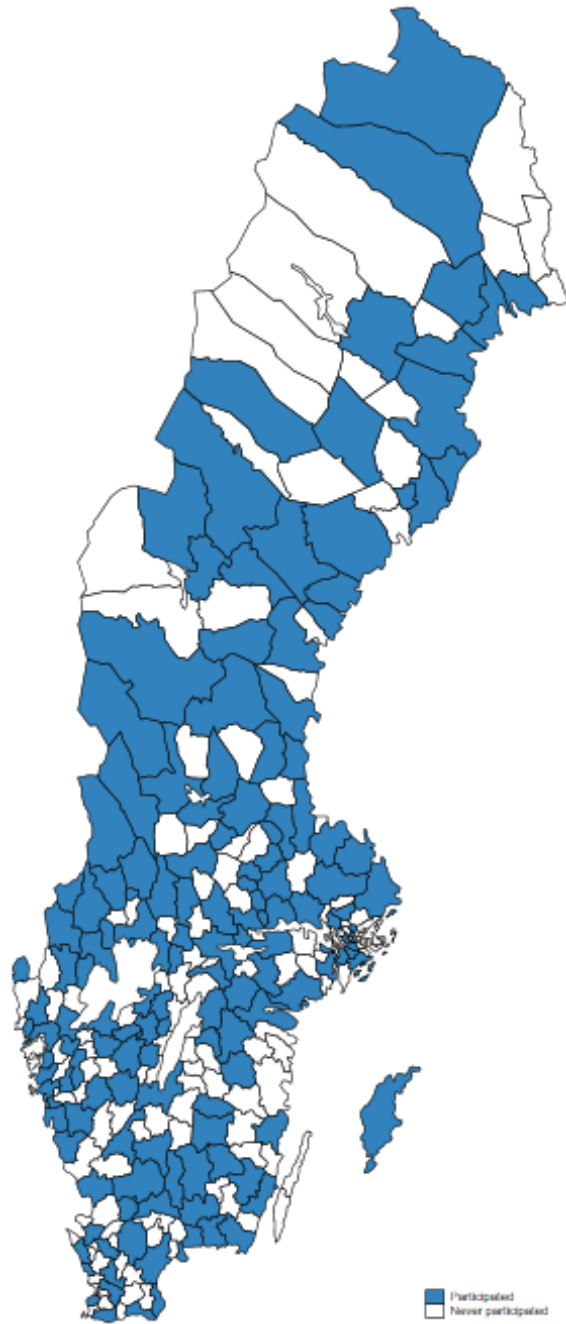


Table A2: Test of spatial correlation (Morans's I)

Panel A: Ever participated		
Matrix	chi2(1)	Prob > chi2
Spatial-weighting matrix (W)	2.40	0.1214
Inverse distance matrix (M)	0.04	0.8410
W and M (Joint test)	2.94	0.2305
Panel B: Fraction of years 1987-1991 that the municipality participated*		
Spatial-weighting matrix (W)	1.63	0.2015
Inverse distance matrix (M)	0.03	0.8699
W and M (Joint test)	2.00	0.3682

Note: *Fraction set to 1 if in trial 1987-1991, 4/5 if in trial 1988-1991, etc.

B Data and construction of variables

B.1 Labor-market and education

To measure labor market performance, we use data on annual labor earnings and translate them into one extensive-margin measure (*employment*) and one overall measure (*annual earnings*). We scale earnings in relationship to the monthly minimum wage.¹ The indicator for *employment* takes value one in years when annual earnings exceed the *de facto* monthly minimum wage and zero otherwise. Thus, being “employed” indicates that the person worked for at least one full-time month during the year. Similarly, our measure of *annual earnings* is scaled by the minimum wage, i.e. we divide the (annual) sum of gross labor earnings by a full year of monthly minimum wages so that we measure earnings in terms of shares of full-time employment at the monthly minimum wage.² None of these definitions are crucial for our conclusions (see online Appendix E.1). With regards to educational attainment, we translate information on the highest degree obtained into accumulated years of schooling.

B.2 Partner’s GPA

In each year, we link women to their partners (if in a relationship) to obtain information on their current partners’ characteristics, including GPA from compulsory school. For men graduating from 9th grade in 1988 or later we observe their grades in the graduation register; for men graduating before 1988 we rely on information from the application register to upper secondary school, if they applied. Thus, the coverage is lower for the earlier cohorts and is dependent on the individual having applied to upper secondary school. Our regressions using partner’s GPA are thus estimated on a smaller sample of women.

¹There is no formal minimum wage in Sweden as these are defined in collective agreements. Instead, we use the wages of janitors as our proxy since their wages consistently stay in the very lowest part of the wage distribution. In 2018, the monthly minimum wage as we construct it correspond to 23,000 SEK or 2,646 USD/ 2,242 EURO.

²In the year of graduation, we scale earnings by 6 instead of 12 months. The reason is that students typically graduate in June.

C Additional Descriptive Statistics

Table C1: Program types for the sample of female vocational students.

Program type	Share (%)		
	All	Low GPA	High GPA
Business and services	29.76	25.50	17.39
Health care ^(a)	27.22	30.24	33.01
Caring services: children and youth	12.56	6.71	24.65
Consumer studies ^(b)	11.73	21.18	3.81
Food manufacturing and restaurant ^(c)	9.26	5.13	11.01
Use of natural resources ^(d)	2.92	3.56	3.15
Textile and clothing manufacture	2.71	2.06	4.27
Electrical engineering	1.00	0.36	1.39
Transport and vehicle engineering	0.71	1.47	0.32
Construction ^(e)	0.59	0.62	0.35
Process technology	0.22	0.41	0.12
Operation and maintenance engineering ^(b)	0.21	0.14	0.22
Wood technology	0.20	0.30	0.12
Industry	0.91	2.32	0.20
Nb. of individuals	78,595	15,930	16,079

(i) The sample includes female vocational students observed in the year of (predicted) graduation – including dropouts. Low (high) GPA students refer to students in the bottom (top) quintile of the GPA distribution in 9th grade. (ii) Students in graphic/handicraft programs are excluded (no strict correspondence in two-year tracks). For a comparison with Table 1 note that: ^(a) The three-year pilot track Health Care corresponds to the two-year tracks Caring Services and Social Services. ^(b) This track does not directly correspond to any of the pilot tracks, but is included in the analysis as important elements appear to be present on one or more of the pilot tracks (see Hall (2009)). ^(c) The three-year pilot tracks Food Manufacturing and Restaurant correspond to the two-year track, Food Manufacturing. ^(d) The three-year pilot track Use of natural resources corresponds to the three separate two-year tracks, Agriculture, Forestry and Gardening. ^(e) The three-year pilot tracks Construction, Constructional metalwork, Heating, Ventilation and Sanitation correspond to one two-year track, construction.

D Additional Results

D.1 Short and long-term impact of unemployment rate at graduation on selected outcome

Table D1: The short and long-run effect of unemployment rate at graduation on selected outcomes, IV specification – All quintiles.

Dependent variable:	Full sample (1)	Bottom quintile (2)	Median quintile (3)	Top quintile (4)
<i>Panel A: Single mother on welfare at age 38</i>				
Effect of unemp. at grad.	0.002 (0.002)	0.017*** (0.006)	0.000 (0.003)	-0.002 (0.002)
Number of individuals	66,336	12,800	12,854	12,433
<i>Panel B: Employment</i>				
Effect of unemp. at grad. by years since (predicted) grad.				
0	-0.030*** (0.003)	-0.045*** (0.007)	-0.023*** (0.006)	-0.0159*** (0.005)
1-2	-0.014*** (0.003)	-0.027*** (0.007)	-0.008 (0.006)	-0.003 (0.005)
3-10	-0.000 (0.003)	-0.006 (0.007)	0.005 (0.006)	0.005 (0.005)
11-19	-0.001 (0.003)	-0.009 (0.007)	0.004 (0.006)	0.006 (0.005)
Nb. of observations	1,453,749	280,759	282,666	271,428
Track FE (instrumented)	✓	✓	✓	✓
Cohort fixed effects	✓	✓	✓	✓
Municipality fixed effects	✓	✓	✓	✓
Individual characteristics	✓	✓	✓	✓

Notes: (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) In Panel A, each model is estimated on the sample of women observed at age 38. In Panel B, we estimate each model using the balanced panel of individuals. The coefficients shown in Panel B are on the interaction of the instrumented unemployment rate at graduation and selected experience groups. (iv) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level.

D.2 Comparing IV to alternative OLS specifications

Our IV strategy handles all concerns related to cross cohort differences, endogenous graduation timing, and selective internal migration. To highlight how our IV estimates fixes problems in the literature, we start by comparing our IV model to the corresponding OLS estimates (obtained when estimating Equation 1 in the paper). The results are presented in columns (1) and (2) of Table D2 (for the full sample) and D3 (for the bottom quintile). In each Table, Panel A) shown the effect of unemployment rate at graduation on employment at different stages of the career, while Panel B) shows the effect of unemployment rate at graduation on the probability of being a single mother on welfare at the age of 38.

The results show rapidly declining effects for the OLS as well as the IV strategy. The power varies between the models but the patterns are similar. The result where the instrumentation do matter is when looking at the long run results where the IV-estimates are much larger than the OLS ones. Note that this OLS strategy is poorly identified as completely relies on endogenous choices of study duration for identification. Comparing the IV to the OLS results suggest that students who endogenously choose to study on a longer track later in periods when unemployment increased have lower propensities for welfare usage.

We then benchmark our model to two models commonly used in the literature, to which we refer to as the “national OLS Oreo” and “local OLS Oreo”. We then derive lessons w.r.t to the role of cohort unobservable characteristics, endogeneous graduation timing, intermunicipality (possibly selective) migration upon graduation, and the effect of national vs. local initial economic conditions.

National OLS

We start by estimating the “national OLS Oreo” specification. This specification is the one estimated in Oreopoulos et al. (2012). We regress annual outcomes on the unemployment rate at the year of graduation from vocational school, interacted with dummies for years of potential labor-market experience, plus experience fixed effects, a linear graduation cohort trend, and individual controls (grades, immigrant background, indicator for 2-3 year track). Note that this model captures the effect of initial conditions at the national level, but fails to control graduation cohort unobservable characteristics.

Formally, it writes as:

Table D2: The short and long-run effect of unemployment rate at graduation on selected outcomes – Full sample.

Dependent variable:	IV (1)	OLS (2)	National OLS Oreo (3)	Local OLS Oreo (4)
<i>Panel A: Employment</i>				
Effect of unemp. at grad. by years since (predicted) grad.				
0	-0.060*** (0.006)	-0.061*** (0.001)	-0.047*** (0.001)	-0.016*** (0.002)
1-2	-0.029*** (0.006)	-0.033*** (0.000)	-0.019*** (0.000)	-0.003** (0.001)
3-10	-0.001 (0.006)	-0.007*** (0.000)	0.007*** (0.000)	-0.000 (0.001)
11-19	-0.002 (0.006)	-0.009*** (0.000)	0.005*** (0.000)	0.004*** (0.001)
Nb. of observations	1,453,749	1,454,317	1,454,317	1,454,317
<i>Panel B: Single mother on welfare at age 38</i>				
Effect of unemp. at grad.				
	0.003 (0.004)	0.001*** (0.000)	0.0002 (0.0004)	0.0005 (0.0006)
Number of individuals	66,336	66,632	66,362	66,362

Notes: (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (ii) In Panel A, we estimate each model using the balanced panel of individuals. The coefficients shown in Panel A are on the interaction of the instrumented unemployment rate at graduation and selected experience groups. (iii) In Panel A, each model is estimated on the sample of women observed at age 38.

$$y_{it} = \alpha + \beta_e UR_i^0 + GraduationYear_i + \chi_{e(i,t)} + \zeta X_i + \epsilon_{it} \quad (1)$$

The unemployment rate UR_i^0 is measured at the time of graduation from vocational school. The main coefficients of interest β_e are allowed to vary with potential experience. $GraduationYear_i$ is the graduation cohort trend, $\chi_{e(i,t)}$ the experience fixed effect, and X_i refer to individual controls.

Local OLS

We then proceed to the “local OLS Oreo” specification. This specification is close to the one used by Oreopoulos et al. (2012) for their main results. We regress annual outcomes on the unemployment rate at the year of graduation in the municipality of residence at that

Table D3: The short and long-run effect of unemployment rate at graduation on selected outcomes – Bottom quintile.

Dependent variable:	IV (1)	OLS (2)	National OLS Oreo (3)	Local OLS Oreo (4)
<i>Panel A: Employment</i>				
Effect of unemp. at grad. by years since (predicted) grad.				
0	-0.087*** (0.014)	-0.074*** (0.001)	-0.061*** (0.001)	-0.022*** (0.004)
1-2	-0.051*** (0.014)	-0.044*** (0.001)	-0.031*** (0.001)	-0.013*** (0.004)
3-10	-0.009 (0.014)	-0.003*** (0.001)	0.010*** (0.001)	-0.007** (0.003)
11-19	-0.014 (0.014)	-0.008*** (0.001)	0.005*** (0.001)	-0.001 (0.003)
Nb. of observations	280,759	280,907	280,907	280,907
<i>Panel B: Single mother on welfare at age 38</i>				
Effect of unemp. at grad.				
	0.044*** (0.014)	0.002 (0.002)	-0.0007 (0.001)	0.007*** (0.002)
Number of individuals	12,800	12,805	12,805	12,805

Notes: (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (ii) In Panel A, we estimate each model using the balanced panel of individuals. The coefficients shown in Panel A are on the interaction of the instrumented unemployment rate at graduation and selected experience groups. (iii) In Panel A, each model is estimated on the sample of women observed at age 38.

time, interacted with dummies for each year of potential labor-market experience, municipality of residence fixed effects, experience fixed effects, year of graduation, and year fixed effects.³ We also include individual controls (grades, immigrant background, indicator for 2/3-year track). Standard errors clustered at the municipality-of-graduation*graduation-cohort level. Formally, it corresponds to the following model:

$$y_{it} = \alpha + \beta_e UR_{j(i)}^0 + \lambda_{c(i)} + \chi_{e(i,t)} + \delta_{j(i)} + \phi_{it} + \zeta X_i + \epsilon_{it} \quad (2)$$

The unemployment rate $UR_{j(i)}^0$ is measured at the time of graduation from vocational school and the municipality of residence at that time. The main coefficients of interest β_e

³As in Oreopoulos et al. (2012), we simply drop one additional cohort effect from the regression to identify separately cohort effects, potential experience effects, and year effects.

are allowed to vary with potential experience. $\lambda_{c(i)}$ is the graduation cohort fixed effect, $\chi_{e(i,t)}$ the experience fixed effect, $\delta_{j(i)}$ the municipality of graduation fixed effect, ϕ_{it} the year fixed effect and X_i refer to individual controls.

Results

The results are shown in columns (3) and (4) of Table D2 (for the full sample) and D3 (for the bottom quintile). Note that for the sake of comparison, we use annual unemployment rate at the country (municipality) level for the prime-age group across across specifications (1)-(4).⁴

We first note that the specification relying on national unemployment rate while accounting for a trend (column 3) delivers very peculiar results, suggesting that the trend is not sufficient to capture secular changes between cohorts. Turning to the more reliable column (4) which relies on local unemployment rate, we see that the long-run effects on unemployment are very similar to the estimates from the national IV. The main difference is instead in the short run, where the IV estimates are much larger. This may well be because a national recession may affects students more forcefully for natural reasons. The fact that our identifying IV-variation is different from the local unemployment rate (i.e. it only relies on the timing early in the recession) may be the reason for why this additional effect is not as lasting.

Turning to the long-run impact on welfare benefits, we see effects both from our national shock and from the local OLS specification in the low GPA group, but effects are much larger in the national IV. Thus, this supports our interpretation that it is the short-run effects that are driving the long-run impact on welfare benefit usage (as this is the key difference between the national IV and the local OLS).

To check whether endogeneous graduation timing or migration across municipalities upon graduation matter, we estimate an alternative version of Eq 2 where unemployment rate at graduation is measured in the predicted year of graduation from vocational school (based on cohort of enrollement + length of program in which s/he enrolled) and the municipality in which the individual resided in the last year of compulsory school. The fixed effects in Eq. 2 are modified accordingly. The estimates (shown upon request)

⁴In the paper, our IV specification uses sex-specific national youth unemployment rate (for age group 15-24). Because the measure of unemployment rate at the municipality level in the “local OLS Oreo” is not age/sex-specific, we use here prime-age unemployment rate for the sake of comparison.

closely match the ones presented above in column (4) of Tables D2 and D3. This is in line with Schwandt and von Wachter (2019), who find that endogenous graduation timing and interstate migration play a limited role in their estimates. This suggests that the difference between the “local OLS Oreo” and our IV specification mainly lies in the fact that our model captures the causal effect of national economic conditions instead of local ones. As we believe national estimates speak directly to much of the concerns that arise whenever economies are hit by very large shocks, we think the national estimates are of direct interest for policy. Indeed our IV strategy explicitly compares graduates that enter the labor market in different phases of a rapidly evolving national recession. We believe it speaks directly to much of the concerns that arise whenever economies are hit by very large shocks (“unlucky cohorts” or “scarred generations”), where conditions at labour-market entry change drastically from a year to another, for example in the context of a financial crisis or the current covid recession.

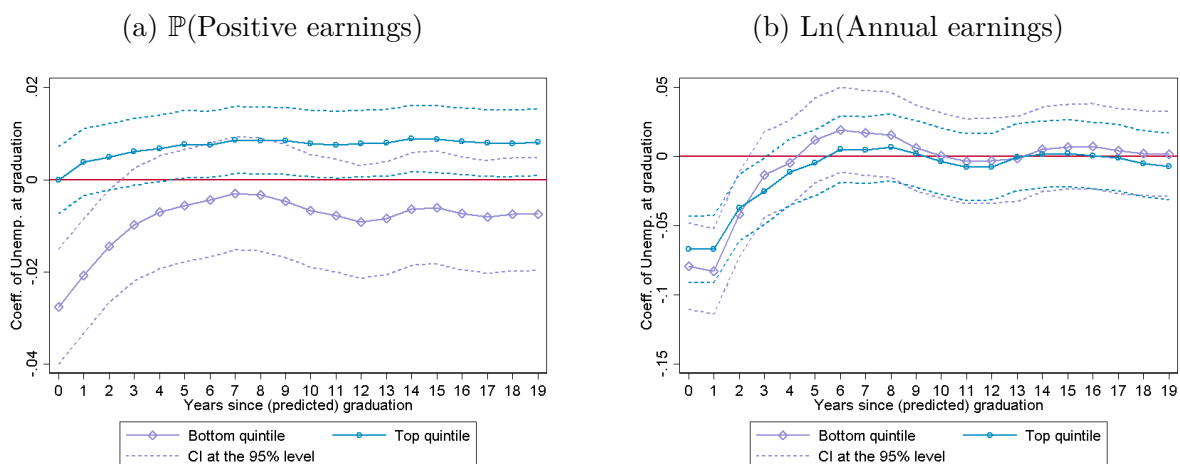
E Robustness checks

We here present a number of robustness checks. The exercises verify that our results are robust to a number of variations.

E.1 Alternative Definitions of Earnings and Employment

We check that that results for earnings and employment give an identical picture if we vary the functional form. For each year since (predicted) graduation, Figure E1 shows the effect of unemployment rate at graduation on the probability of having positive earnings (Figure E1a), and on log annual earnings (Figure E1b). Results are virtually unchanged when using these alternative definitions.

Figure E1: The effect of unemployment at graduation on the probability of having positive earnings, and on log(earnings)

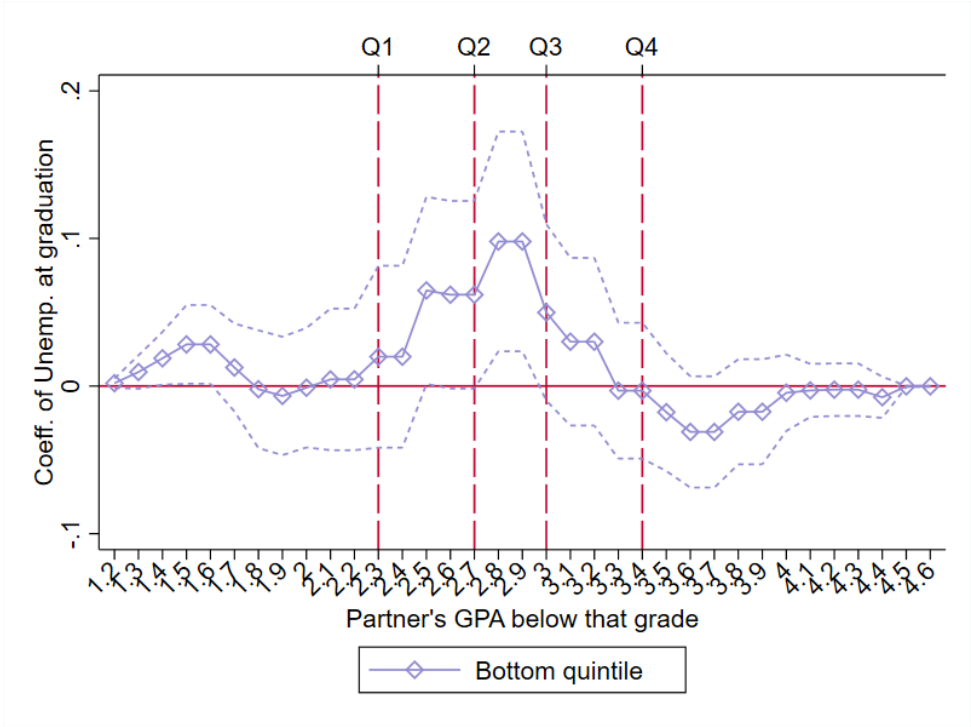


E.2 Dichotomized measures of partner's quality based on GPA

While our main results in Table 9 use partner's GPA as a continuous outcome variable, we use here dichotomized measures of partner's GPA (e.g. partner's GPA below median/first quintile etc.). We explore different thresholds for partner's GPA, as is shown in Figure E2. The Figure shows the effect of unemployment rate at graduation on the probability of mating with a partner whose GPA is below a given grade (with grades ranging from 1 to 5). Vertical (dashed) red lines show quintiles in the partners' GPA distribution. Overall, Figure E2 shows that low-GPA partners can be extensively defined as having a GPA below the median (GPA<2.9), the second quintile (GPA<2.7) or the third quintile

(GPA<3). Results for the first quintile (GPA <2.4) are not significant at conventional levels, suggesting that women who graduate in a recession are not more likely to draw partners from the very lower end of the GPA distribution. Although the functional form matters for whether the results are statistically significant or not at conventional levels, the overall impression is that there is a clear tendency for the women to marry lower GPA men.

Figure E2: The effect of unemployment rate at graduation on the probability of having a first partner with 9th grade GPA below grade “xx” at age 27 – IV specification.



Notes: Vertical (dashed) red lines represent quintiles in the partners' GPA distribution.

E.3 Evaluating the Identifying Assumptions for the Bottom GPA Quintile

We verify that the tests of identifying assumptions discussed in section 4.2 hold for the bottom quintile of the GPA distribution (see Table E1 below).

Table E1: Evaluating the identifying assumptions – Bottom quintile^(a). The effect of predicted unemployment rate at graduation (ϕ_{cj}) and pilot intensity (P_{cj}) on the individual’s choice of whether to enroll in upper secondary school (at all), whether to enroll in a vocational track (vs. an academic one), of delaying enrollment in vocational studies, of dropping out of vocational studies, and on the composition of dropouts.

	Dependent variable				
	$\mathbb{P}(\text{enrolls in upp. sec. school})$	$\mathbb{P}(\text{enrolls in voc. track})$	$\mathbb{P}(\text{delays enrollment in voc. track})$	$\mathbb{P}(\text{drops out from voc. track})$	GPA rank (for dropouts)
	(1)	(2)	(3)	(4)	(5)
Predicted UR grad. (ϕ_{cj})	-0.003 (0.006)	-0.006 (0.004)	-0.001 (0.003)	0.000 (0.005)	0.141 (0.176)
Pilot intensity (P_{cj})	0.047 (0.043)	0.044 (0.029)	-0.006 (0.018)	0.020 (0.039)	-2.119* (1.248)
GPA rank in 9 th grade	0.026*** (0.001)	-0.009*** (0.001)	-0.003*** (0.000)	-0.015*** (0.001)	- -
Born in Nordic country	-0.032*** (0.012)	0.022* (0.011)	-0.150*** (0.014)	0.022 (0.016)	0.089 (0.494)
Cohort fixed effects	✓	✓	✓	✓	✓
Municipality fixed effects	✓	✓	✓	✓	✓
Mean of dep. variable	0.62	0.92	0.05	0.26	8.97
Nb. of individuals	37,752	19,251	18,444	17,412	4,150

Notes : ^(a) Here, the bottom quintile refers to students whose GPA in 9th grade was below 2.6 (which corresponds to the maximum grade obtained by vocational students in the bottom quintile in our main sample). (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) Model (1) is estimated using the sample of students observed three years after graduation from compulsory school– including dropouts. Since not all students in compulsory school enrolled in upper secondary school, P_{cj} and ϕ_{cj} are measured at the municipality*cohort ($j \times c$) level, where j stands for municipality of residence at age 16 (as information on the municipality of residence during the last year of compulsory schooling is only available from the *Upper secondary application register*) and c stands for the year the individual finishes compulsory school. (iv) Model (2)-(5) are estimated using the sample of students observed in the year of (predicted) graduation – including dropouts. We estimate Model (2) conditional on enrolling in upper secondary school at age 16, Model (3) conditional on enrolling in vocational studies, Model (4) conditional on enrolling in vocational studies at age 16, and Model (5) conditional on enrolling in vocational studies at age 16 and dropping out. (v) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level.

E.4 Robustness Checks on the Model and Sample

Table E2 shows a number of robustness checks where we use alternative specifications or modify the sample. For simplicity, we report estimates for a selected set of outcomes (being a welfare-dependent single mother; employment) for the bottom quintile of vocational students. The table shows that results are unchanged when including program fixed effects (see Table A1 for list of programs), when controlling for current unemployment,

when excluding late graduates from vocational school, or when excluding graduates who engage in further education (i.e. who ever enrolled in university by the age of 38). Results using local (municipality-level) unemployment rates are quantitatively similar but smaller in magnitude, possibly because of domestic migration. Varying the clustering to the municipality level has no effects on the significance of the estimates.

Table E2: Robustness checks on the model and the sample. The effect of unemployment rate at graduation on selected outcomes, IV specification – Bottom quintile.

Dependent variable:	Baseline (1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>Single mother on welfare at age 38</i>							
	0.017***	0.018***	0.017***	0.017***	0.018**	0.008***	0.017**
	(0.006)	(0.006)	(0.006)	(0.006)	(0.007)	(0.003)	(0.007)
Nb. of individuals	12,800	12,800	12,800	11,527	10,853	12,800	12,800
<i>Employment</i>							
Years since (predicted) graduation							
0	-0.045***	-0.045***	-0.053***	-0.046***	-0.044***	-0.066**	-0.045**
	(0.007)	(0.007)	(0.015)	(0.007)	(0.008)	(0.004)	(0.007)
1-2	-0.027***	-0.027***	-0.019***	-0.027***	-0.027***	-0.036***	-0.027***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.004)	(0.007)
3-10	-0.006	-0.006	-0.009	-0.007	-0.006	0.004	-0.006
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.004)	(0.007)
11-19	-0.009	-0.008	-0.007	-0.009	-0.008	0.001	-0.009
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.003)	(0.007)
Nb. of observations	280,759	280,759	275,581	252,986	238,093	280,759	280,759
Program type		✓					
Current UR			✓				
No late graduates				✓			
No university studies					✓		
Local unemployment rate						✓	
Clustering at municipality							✓
Track FE (instrumented)	✓	✓	✓	✓	✓	✓	✓
Cohort fixed effects	✓	✓	✓	✓	✓	✓	✓
Municipality fixed effects	✓	✓	✓	✓	✓	✓	✓
Individual characteristics	✓	✓	✓	✓	✓	✓	✓

Notes: (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) In Panel A, each model is estimated on the sample of women observed at age 38. In Panel B, we estimate each model using the balanced panel of individuals. The coefficients shown in Panel B are on the interaction of the instrumented unemployment rate at graduation and selected experience groups. (iv) In Panel A, column (3), the effect of current national unemployment rate cannot be identified, as we compare women at the same age (hence same year). To circumvent that problem, we control instead for current unemployment rate at the *local* (municipality) level. (v) To estimate the model in column (6), we use the baseline sample (see column (1)) and replace the observed unemployment rate at graduation at the national level by the observed unemployment rate at graduation at the municipality level. We define the instrument as $\phi_{c,j} = P_{c,j} * UR_{c+3,j} + (1 - P_{c,j}) * UR_{c+2,j}$. (vi) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level.

E.5 Sibling Fixed-Effects

Causal interpretation of our results relies on the exogenous timing of labor-market entry. We argue that our IV strategy generates a setting where students are exogeneously pushed in the labor market. If that is true, our results should be robust to the inclusion of sibling fixed effects, which remove all unobserved heterogeneity that is common within a family. As shown in Table E3, our results are virtually identical if including family fixed effects.⁵

⁵Our sample uses the whole sample of boys and girls to increase the number of families where we can identify the fixed effects but all variables (except the sibling dummy) are interacted by sex. Results show that a one percentage point increase in unemployment rate at graduation increases the probability of being a welfare-dependent single mother at age 38 by 1.4 percentage points (10% significance level). A one percentage point increase in unemployment rate at graduation decreases by 3.2 percentage points the probability of being employed on the year of (predicted) graduation (significant at the 1% level), and by 1.6% the probability of being employed 1-2 years after graduation (significant at the 5% level).

Table E3: Sibling fixed-effects model, IV specification – Full sample of men and women, all quintiles. The effect of unemployment rate at graduation on selected outcomes, coefficient shown for women.

Dependent variable:	(1)
<i>Panel A: Single mother</i>	
<i>on welfare at age 38</i>	0.014* (0.008)
Nb. of individuals	23,615
<i>Panel B: Employment</i>	
Years since (predicted) graduation	
0	-0.032*** (0.007)
1-2	-0.016** (0.007)
3-10	-0.002 (0.007)
10-19	-0.002 (0.007)
Nb. of observations	3,570,609
Sibling FEs	✓
Track FEs (instrumented)	✓
Cohort fixed effects	✓
Municipality fixed effects	✓
Individual characteristics	✓

Notes: Our sample uses the whole sample of boys and girls to increase the number of families where we can identify the fixed effects but all variables (excepted the sibling dummy) are interacted by sex. For similar reasons, we use all vocational students, irrespective of GPA quintiles. (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) In Panel A, each model is estimated on the sample of individuals observed at age 38. In Panel B, we estimate each model using the balanced panel of individuals. The coefficients shown in Panel B are on the interaction of the instrumented unemployment rate at graduation and selected experience groups. (iv) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level.

E.6 Reduced Form of the Model

As discussed in the main text, the actual unemployment rate (UR_i^{Grad}) is unobserved for dropouts, and the latter are therefore *de facto* excluded from the analysis. We already verified in section 4.2 that both the *number* of drop outs and the *composition* of dropouts (in terms of, e.g., predetermined GPA) are unrelated to the instrument. These results suggest that endogenous dropout composition should not be biasing our results. However, to further assess the robustness, we also check that the results are unchanged if we re-

introduce dropout students in the sample and run a “reduced-form” version of the model (i.e. using the instrument directly without using the endogenous independent variable). In Table E4 we show reduced-form estimates including and excluding dropouts: the results are very similar. Note that the coefficients of interest in both cases are smaller in magnitudes than in our main analysis since we estimate the direct effect of the instrument (the “intention-to-treat” parameter).

Table E4: Reduced form of the model. The effect of ϕ_{jc} on selected outcomes – Bottom quintile.

Dependent variable	(1)	(2)
Panel A: Single mother on welfare		
At age 38	0.009*** (0.003)	0.006** (0.003)
Nb. of individuals	12,800	17,167
Panel B: Employment		
Years since (predicted) graduation		
0	-0.038*** (0.003)	-0.039*** (0.003)
1-2	-0.020*** (0.003)	-0.022*** (0.003)
3-10	-0.000 (0.003)	0.000 (0.003)
11-19	-0.002 (0.003)	-0.003 (0.003)
Nb. of observations.	280,782	377,179
Including dropouts		✓
P_{cj}	✓	✓
Cohort fixed effects	✓	✓
Municipality fixed effects	✓	✓
Individual characteristics	✓	✓

Notes: (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) In Panel A, each model is estimated on the sample of women observed at age 38. In Panel B, we estimate each model using the balanced panel of individuals. The coefficients shown in Panel B are on the interaction of the instrumented unemployment rate at graduation and selected experience groups. (iv) In column (1), we run the model for the bottom quintile of students, i.e. for students with GPA lower than 2.6 in 9th grade. In column (2) the estimation sample includes drop outs with GPA lower than 2.6 in 9th grade. (v) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level.

F A stylized model of partnership choice

Our empirical analysis focuses on how a transitory worsening of labor market conditions affect family formation outcomes of females in the short and long run. To fix ideas, consider a marriage market where agents, in each period, receive draws from a distribution of match qualities. Agents will accept the best offer in the period if it has an expected value that is better than the expected value of waiting until the next period. In such a setting, agents will accept matches of worse quality in a recession if (and only if) recessions reduce the value of staying single relative to the value of forming a marginal partnership.⁶ If the offer arrival rate is fixed, this will also lead to more early marriages. And if divorces are more likely to occur in cases where initial match quality is low, there will also be more divorces.

To make this intuition more precise, we define a stylized model of partnership choice. We do not model the equilibrium outcomes because we are interested in illustrating how agents on one side of the market change their behavior in a stable environment. Our empirical setting is one where we have variation in the severity of the income shock facing one side of the market (females), variation which is uncorrelated with the prospects of the other side of the market. This is because the relevant males (who, on average, are three years older) belong to other cohorts and their labor market prospects will be unrelated to the pilot intensity which provides our identifying variation. It is, however, clear that changes in the partnership thresholds of agents on one side of the market (here: women) should affect the marriage probability of (in particular, less attractive) agents on the other side (here: men) and we analyze this process empirically in online Appendix G. Here, however, we choose to abstract from feedback effects on the male side to keep the model simple, and to not cloud the intuition about how the relevant women's choices may change with economic conditions.

F.0.1 A stylized model of partnership choice

Consider an indirect period-specific utility function $V(I, q)$ where individuals derive utility from income I and from marital quality q in each period.⁷ The utility function is concave in earnings and marital quality. Individuals live for two periods (1 and 2). In period 1

⁶In theory, the acceptance threshold can also increase, but in line with our empirical results, we focus our narrative on the case where the acceptance threshold falls.

⁷The model lends elements from Boulier and Rosenzweig (1984) and Fernandez et al. (2005).

they draw a (best) marital offer of quality q and decide if they want to accept it or not. In period two, the formed marriages are dissolved with probability ϕ (which may be a decreasing function of q).

Individuals who remain single when entering period 2, draw marriages of fixed quality μ , whereas divorcees draw new partners of the same fixed quality μ but pay a utility penalty λ . Our forcing variable of interest is labor earnings in the first period I . Because we are interested in the impact of transitory shocks, we let earnings in the second period be fixed at J . The discount factor is denoted by β . Hence, in the first period, the individual trades off first-period marriage with expected (indirect) utility:

$$V(I, q) + \beta[(1 - \phi)V(J, q) + \phi(V(J, \mu) - \lambda)] \quad (3)$$

against staying single and deriving

$$V(I, s) + \beta V(J, \mu) \quad (4)$$

where s denotes the single-life substitute for marital happiness in period 1. We are interested in how variations in first-period income affect the willingness to enter partnerships in the first period. Changes in the willingness to mate operate through changes in the minimum acceptable match quality. Denote this threshold value by q^* and note that it satisfies equality between Equations (3) and (4). Simplify to:

$$V(I, q^*) - V(I, s) + \beta[(1 - \phi)V(J, q^*) - (1 - \phi)V(J, \mu) - \phi\lambda] = 0 \quad (5)$$

Treating ϕ as a constant, and denoting the derivatives of V by V'_I and V'_q respectively, we get the impact of first-period income on the partnership threshold as:

$$\frac{dq^*}{dI} = \frac{V'_I(I, s) - V'_I(I, q^*)}{V'_q(I, q^*) + \beta(1 - \phi)V'_q(J, q^*)}. \quad (6)$$

This implies that the partnership threshold will be a positive function of first-period earnings if the marginal utility of earnings is higher if single than if within a formed marginal partnership. One obvious reason for why this may be the case is if couples are pooling incomes. To make this explicit, we follow Fernandez et al. (2005) and let the utility function take the form :

$$V(I, q) = \ln\left(\frac{I + I^P}{2}\right) + q \quad (7)$$

if in a partnership, where I^P denotes partner income, and

$$V(I, s) = \ln(I) + s \quad (8)$$

if single. This simplifies Equation (6) to

$$\frac{dq^*}{dI} = \frac{\frac{I^P}{I(I+I^P)}}{1 + \beta(1 - \phi)} > 0. \quad (9)$$

Equation (9) implies that fewer and better marriages are formed if short-run labor market prospects improve. The higher is the partner's income share of total income, the larger is the effect, suggesting that female thresholds should respond more than male thresholds to own income shocks on average. Furthermore, females with a high baseline divorce rate (ϕ) or a low earnings potential should respond the most. In our empirical setting, both of these attributes are closely related low compulsory school grades. In a more general sense, the reaction should also be larger for females with less access to alternative sources of income, e.g., from within the family. This suggests that children of low-earning parents should respond more. In the paper we do not focus on parental income as such, but instead focus on the role of compulsory school GPA, which is correlated with both own and parental labor market outcomes.

The main intuition is valid as long as the potential partner's earnings is not perfectly correlated with the own income. This is typically the case if economic conditions are rapidly evolving, because of the age gap between the two sides of the market.

Letting the divorce probability ϕ be a function of match quality, e.g. by letting $\phi = 1 - q$, does not change the intuition. It induces additional terms in the denominator and generates a straightforward relationship between income and divorces:

$$\frac{dq^*}{dI} = \frac{\frac{I^P}{I(I+I^P)}}{1 + \beta 2q^* + \beta[\lambda - \mu]} > 0. \quad (10)$$

$$\frac{d\phi}{dI} = \frac{dq^*}{dI} * \frac{d\phi}{dq^*} < 0. \quad (11)$$

Thus, the agents accept to form partnerships where the dissolution risk is higher when

the labor market returns are low, which naturally leads to more divorces and associated outcomes. In the empiric we study a number of related outcomes such as the incidence of single-parenthood and welfare dependency.

G Results for Males

We present results for males to complete the picture. We perform a series of exercises to see i) how men who graduate during the recession fare in terms of labor and family outcomes, ii) how the female's partner pool is affected. Recall that because of the substantial age difference between partners (a three-year gap on average, with only 13% of partners being formed within a cohort), these two pools of men hardly overlap. These empirical exercises, unfortunately, suffer from a number of data and identification limitations. The results below should be seen as illustrative evidence that help us gain insights into the male side.

G.1 Males graduating in the recession

As a first exercise, we investigate how males respond when graduating in a recession. Recall that in the presence of military service, males have a much looser relationship between the graduation year and the time of labor market entry. This is because most men enlist during their 19th to 21st year, regardless of duration of upper secondary school. As we study males' labor market and family formation outcomes using the exact same empirical approach as we used for females, the results presented here are only suggestive. Results for employment and earnings are presented in Table G1 and Figure G1. Both the dynamics of the effects and the order of magnitude of the estimates are similar to females. Young males who enter the labor market in a deep recession suffer significant employment and earnings losses in the short-run, but these initial losses do not persist after four years. To some extent, this suggests that males are more affected by labor market conditions since estimates are similar, despite of the additional noise generated by the military service intermission. In stark contrast with our results for females, we do not find a permanent effect on welfare receipts (see Figure G1c).

Table G1: The impact of unemployment rate at graduation on **male** labor-market outcomes in the year of (predicted) graduation – IV specification.

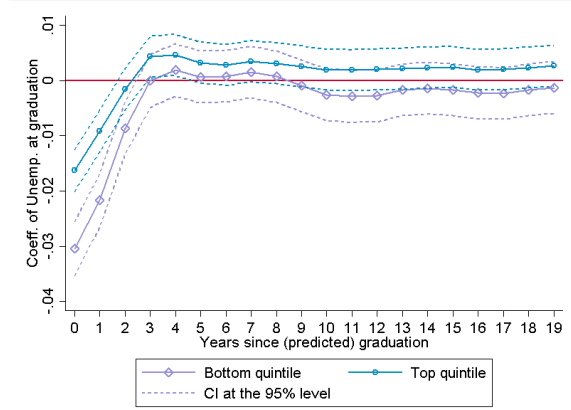
Dependent variable:	Full sample (1)	Bottom quintile (2)	Median quintile (3)	Top quintile (4)
Employment^(a)				
Effect of unemployment rate at graduation (instrumented)	-0.027*** (0.004)	-0.041*** (0.006)	-0.017** (0.006)	-0.019*** (0.006)
Number of individuals	100,674	19,531	17,657	18,370
Mean of dependent variable	0.716	0.662	0.726	0.762
Annual earnings^(b)				
Effect of unemployment rate at graduation (instrumented)	-0.024*** (0.003)	-0.026*** (0.005)	-0.018*** (0.005)	-0.019*** (0.005)
Number of individuals	100,674	19,31	17,657	18,370
Mean of dependent variable	0.512	0.479	0.516	0.536
Track FE (instrumented)	✓	✓	✓	✓
Cohort FE	✓	✓	✓	✓
Municipality FE	✓	✓	✓	✓
Individual characteristics	✓	✓	✓	✓

Notes: (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) Each model is estimated using the sample of students present in the year of (predicted) graduation. (iv) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level. ^(a)We assume than an individual is employed (or to some extent active on the labor-market) if his earnings in the current year are higher than the monthly minimum wage (see Section 3 for details) ^(b)Annual earnings are scaled by the annual minimum wage level (see Section 3 for details).

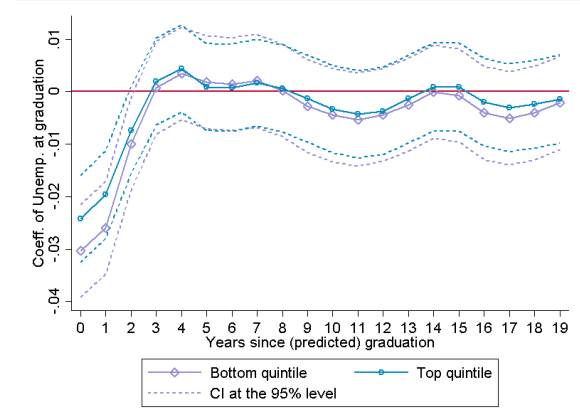
Overall, our results for males' family formation suggest opposite patterns for male than for females, at least in terms of timing. Men who graduate in a recession stay longer at their parent's home (up until seven years after enrollment – i.e. up until age 23, see Figure G2), and tend to delay partnership formation (see Panel A of Table G2). As it is the case for women, the effects are only significant for men in the bottom quintile of the GPA distribution. Partner characteristics do not seem to be significantly affected (not shown), and men do seem to divorce *less* when graduating in a recession (see Panel B of Table G2). These results are consistent with work and family formation not being substitutes for men, but also with the idea that men are less attractive on the marriage market when their labor market prospects deteriorate relative to women (as shown by Autor et al. (2019) and Blau et al. (2000)).

Figure G1: The effect of unemployment rate at graduation on **males'** labor-market outcomes and welfare receipt for each year since (predicted) graduation – IV specification.

(a) $\mathbb{P}(\text{Annual earnings} \geq \text{monthly minimum wage})$



(b) Annual earnings



(c) $\mathbb{P}(\text{Belongs to a household that received welfare benefits during that year})$

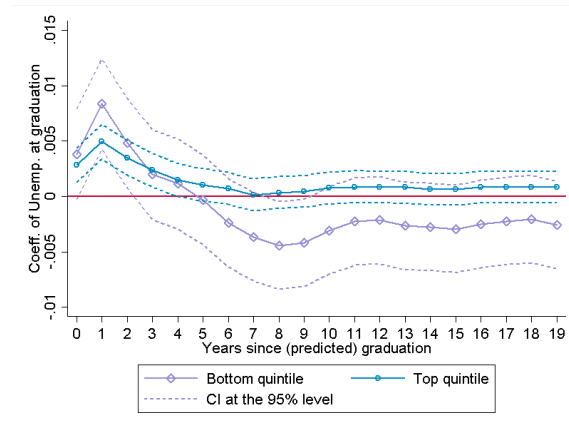


Figure G2: The effect of unemployment rate at graduation on **males'** probability to live at one parent's home for each year since enrollment – IV specification.

(a) $\mathbb{P}(\text{Lives at parent's home})$

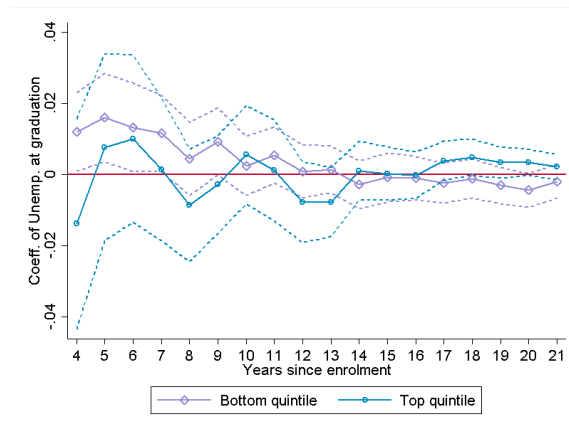


Table G2: The impact of unemployment rate at graduation on **male** family formation – IV specification.

Dependent variable:	Full sample (1)	Bottom quintile (2)	Median quintile (3)	Top quintile (4)
<i>Panel A: Timing of family formation</i>				
Age at first stable partnership^(a)				
Effect of unemployment rate at graduation (instrumented)	0.022 (0.029)	0.158** (0.062)	0.056 (0.079)	0.080 (0.053)
Number of individuals	74,619	13,704	13,013	14,205
Mean of dependent variable	29.232	28.676	29.398	29.486
Age at first birth				
Effect of unemployment rate at graduation (instrumented)	0.016 (0.028)	0.161** (0.066)	0.117 (0.082)	0.034 (0.052)
Number of individuals	72,589	13,477	12,649	13,725
Mean of dependent variable	29.327	28.460	29.493	29.860
<i>Panel B: Completion of family formation^(b)</i>				
Ever in a stable partnership				
Effect of unemployment rate at graduation (instrumented)	-0.001 (0.002)	-0.006 (0.005)	0.009 (0.006)	0.006 (0.004)
Number of individuals	97,927	18,972	17,170	17,880
Mean of dependent variable	0.762	0.722	0.758	0.795
Ever had a child				
Effect of unemployment rate at graduation (instrumented)	0.000 (0.002)	-0.002 (0.005)	0.011 (0.007)	0.007 (0.005)
Number of individuals	97,927	18,972	17,170	17,880
Mean of dependent variable	0.742	0.712	0.738	0.769
Total number of children (if ever had a child)				
Effect of unemployment rate at graduation (instrumented)	-0.001 (0.005)	-0.005 (0.012)	-0.007 (0.012)	-0.016 (0.011)
Number of individuals	72,489	13,463	12,621	13,705
Mean of dependent variable	1.989	1.999	1.977	2.013
<i>Panel D: Partnership dissolution^(c)</i>				
Ever ended the first stable partnership				
Effect of unemployment rate at graduation (instrumented)	-0.005* (0.003)	-0.015** (0.006)	-0.001 (0.007)	-0.001 (0.005)
Number of individuals	74,619	13,704	13,013	14,205
Mean of dependent variable	0.249	0.350	0.234	0.179
Track FE (instrumented)	✓	✓	✓	✓
Cohort FE	✓	✓	✓	✓
Municipality FE	✓	✓	✓	✓
Individual characteristics	✓	✓	✓	✓

Notes: ^(a)A stable partnership is defined as either being married or cohabitation with a partner and common child. ^(b)Completed family formation refers to family outcomes at age 38. ^(c)Partnership dissolution is measured by the age of 38. (i) Marginal effects are presented. (ii) Standard errors (in parentheses) are clustered at the municipality*cohort ($j \times c$) level. (iii) Each model is estimated using the sample of men observed at age 38. (iv) *** Significant at 1% level. ** Significant at 5% level. * Significant at 10% level.

G.2 Partner pool

As a second exercise, we investigate how the female's partner pool is affected. Our previous findings provide evidence that women accelerate their entry into partnerships through a lower quality threshold when their labor-market conditions worsen (relative to men). This leads to the reverse hypothesis that some men in their potential partner pool accelerate their entry into partnerships, forming partnerships that they could otherwise not afford with higher-quality women. We find support for this in the data. We define the female's partner pool as men in vocational education who are three years older.⁸ We run the reduced form of the 2SLS model for men in the partner pool thus defined, but use instead $P_{c+3,j}$ and $\phi_{c+3,j}$, the instruments defined for an enrollment cohort three years *ahead* (i.e. enrolling three years *after* males do), for their municipality of residence in the last year of compulsory school. In line with our hypothesis, our results suggest that an increase in $\phi_{c+3,j}$ – the predicted unemployment rate at graduation for females in their partner pool – leads males (those in the bottom quintile of the GPA distribution in particular) to form partnerships earlier and with higher-GPA women (see Figure G3). So, the relative attractiveness of males induces the early formation of partnerships that low-GPA men could otherwise not afford.⁹

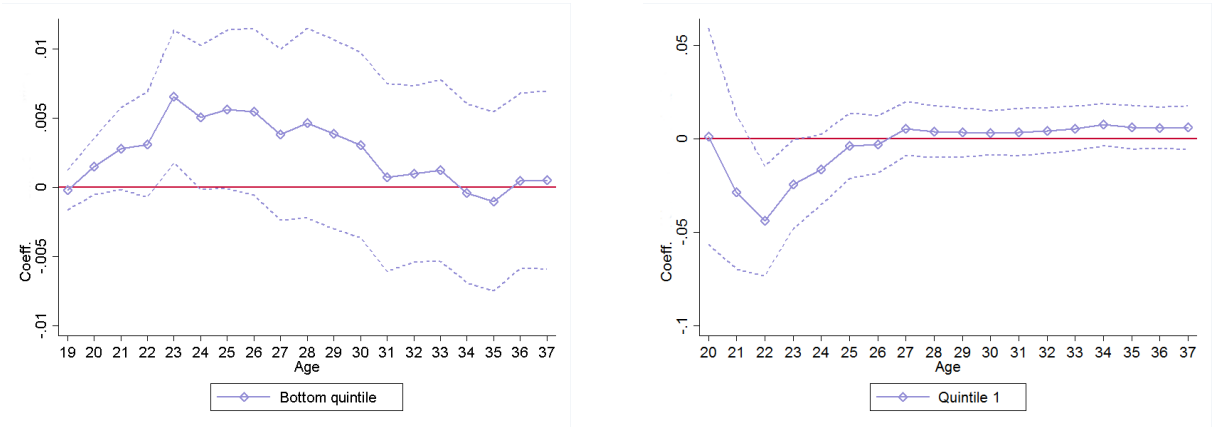
⁸This is of course an over-simplified definition of the female's partner pool, but nonetheless illustrative. This yields a sample of men enrolling in vocational studies in 1984-1988. As highlighted in the data section, we do not have data on compulsory school GPA before 1985, so we consider men enrolling in vocational studies in 1985-1988.

⁹Possibly at the expense of lower-quality girls, e.g. those who dropped out after compulsory school (about 10% of all 9th grade students).

Figure G3: The effect of the male's partner pool predicted unemployment rate at graduation ($\phi_{c+3,j}$) on **male** family formation – Reduced form specification – Bottom quintile.

(a) $\mathbb{P}(\text{Ever formed a first partnership at different ages})$

(b) $\mathbb{P}(\text{In a partnership with a low-GPA female (conditional on being in a partnership)})$



Note: The sample of men includes all men enrolling in vocational studies in 1985-1988.

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