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Understanding and Breaking the Intergenerational Transmission of Gender-Based Violence

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Abstract

Using administrative data from Sweden and the United States this paper provides the first large-scale estimates on the intergenerational transmission of gender-based violence and evidence on how this cycle can be broken. We show that sons of fathers suspected of gender-based violence are more than twice as likely to become suspected perpetrators, and daughters exposed to violent fathers are almost twice as likely to partner with violent men. These patterns persist after controlling for rich demographic, socioeconomic, and neighborhood characteristics. We then examine whether removing abusive fathers from the household disrupts this transmission. Although separations generate sizable and persistent income losses, differencing against comparable non-violent households reveals that father removal significantly reduces both sons' perpetration and daughters' victimization risks in adulthood. Family fixed-effects estimates further show that removal of a violent father before age 11 yields larger protective effects, particularly for sons. Finally, using a judge-IV design, we find that quasi-random removal of abusive fathers reduces the intergenerational cycle of violence by reducing the perpetration risk among sons separated from abusive fathers in their formative ages. Together, these results indicate that gender-based violence is a learned behavior developed over childhood and that reducing exposure to abusive fathers can meaningfully weaken its intergenerational persistence.

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1 Introduction

Violence against women is a widespread social problem with severe adverse consequences for the health and economic outcomes of victims and their children (Bhuller, Dahl, Løken and Mogstad, 2024; Adams, Huttunen, Nix and Zhang, 2024). Globally, 30% of women aged 15 and older have experienced physical or sexual violence by an intimate partner (Sardinha *et al.*, 2022). 29% of women report experiencing a sexual assault (Li *et al.*, 2023), with 6% of these incidents being perpetrated by a non-partner (Sardinha *et al.*, 2024). Data from the National Intimate Partner and Sexual Violence Survey indicate that 6.6% of women report experiencing IPV in the past year, with a lifetime prevalence of 37% (Smith *et al.*, 2017). The estimated costs of these crimes are staggering: lifetime health and economic costs of IPV in the U.S. are estimated to reach $4.1 trillion (in 2021$), with $1.5 trillion attributed to productivity losses and reductions in lifetime earnings (Peterson *et al.*, 2018). In Sweden, the annual costs are estimated at $3.85 billion (Regeringskansliet, 2015). These likely underestimate the true lifetime costs, as they neglect to accurately capture labor market effects (Adams *et al.*, 2024).

Yet despite the widespread prevalence and significant costs of gender-based violence, its underlying causes are still poorly understood. Is the likelihood of perpetrating gender-based violence an innate and immutable characteristic, or is it a learned behavior developed over childhood and adolescence? In this paper, we use rich administrative data from Sweden and the United States [in progress], and a series of complementary research designs to provide two new key insights. First, we provide evidence that children learn to either perpetrate or become victims of gender-based violence through childhood exposure. Second, we provide evidence that this transmission channel can be broken, drastically reducing the extent of gender-based violence perpetrated by young men, and causing young women to seek out partners less apt to commit violence against them.

To provide evidence on the transmission of gender-based violence, we use rich nationwide administrative data from Sweden and the United States [in progress] to estimate whether children exposed to fathers who perpetrate gender-based violence are more likely to become perpetra-

tors or victims themselves. Our analysis employs population-wide linked register data covering all individuals born between 1970 and 1995 and their biological or adoptive fathers. These data link demographic and socioeconomic records from Statistics Sweden with comprehensive criminal justice information from the Swedish Conviction Register and Crime Suspicion Register. Together, these data sources allow us to construct precise measures of fathers' involvement in gender-based violence—including suspicions of assault, unlawful threats, harassment, violations of restraining orders, and sexual violence—as well as children's subsequent perpetration or victimization in adulthood.

Using this linked and nationally-representative dataset, we provide the first large-scale descriptive estimates on the extent to which gender-based violence is transmitted across generations. Raw correlations indicate that sons who are exposed to fathers suspected of gender-based violence are 350% more likely to become suspected perpetrators themselves. Controlling for an extensive set of demographic, socioeconomic, and neighborhood characteristics reduces this estimate to 260%, and restricting the comparison to a matched sample further attenuates it to 230%, which implies that exposure to an abusive father at least doubles a son's likelihood of perpetrating abuse. Next, we document that daughters whose fathers were suspected of violence against women are 240% more likely to form relationships with partners suspected of committing gender-based violence. Conditioning on detailed demographic, socioeconomic, and neighborhood characteristics slightly reduces this likelihood to a 190% greater likelihood of victimization, and restricting the comparison to a matched sample lowers it further to 170%, suggesting that even among otherwise similar women, exposure to a violent father substantially increases the risk of cohabiting with a violent partner.

Next, we examine whether reduced exposure to fathers who commit gender-based violence can break this intergenerational cycle of gender-based violence. A priori, the effect of removing fathers on the intergenerational transmission of violence is theoretically ambiguous. On the one hand, the removal of abusive fathers may reduce children's exposure to harmful social learning mechanisms, wherein children witness and model their fathers' violent behavior toward women (Bandura, 1977; Ehrensaft and Langhinrichsen-Rohling, 2022). On the other hand, removing a

father from the household may lead to negative economic consequences. Evidence suggests that economic hardship during childhood is associated with adverse developmental outcomes, including increased aggression and antisocial behavior (Evans, 2004). If the father's removal results in a significant reduction in household income, the child may experience heightened economic stress, which could exacerbate behavioral problems and ultimately increase the risk of perpetrating violence in adulthood. Thus, while limiting exposure to abusive behavior may be protective, the economic strain resulting from such interventions could act as a competing mechanism that reinforces negative outcomes.

Estimating the impacts of father removal separately for violent and non-violent fathers reveals a common pattern: children who experience father removal at younger ages face elevated risks later in life, consistent with the large income losses triggered by separation. However, when we estimate a simple difference in difference model comparing boys whose violent fathers are removed to boys who also experience father removal, but whose fathers do not commit violence against women, we observe a clear pattern: sons whose violent fathers are removed experience a significantly lower likelihood of perpetrating gender-based violence in adulthood. This negative effect is robust to the inclusion of rich demographic, socioeconomic, and neighborhood controls, as well as to a matched comparison where non-violent fathers are observationally identical to violent fathers.

Turning to daughters, we estimate the same specifications, but estimate impacts on their likelihood of cohabiting with a partner suspected of gender-based violence as a proxy for victimization. As with sons, raw estimates show that daughters who experience father removal at younger ages face elevated risks later in life, consistent with the sizable income losses associated with separation. After differencing against comparable non-violent households, the estimates generally turn negative; however, they are much smaller and less precisely estimated.

To address concerns that cross-family differences may confound our estimates, we further exploit within-family variation by including family fixed effects and comparing siblings who experienced the removal of an abusive father at different ages. This design absorbs all time-invariant family characteristics, including parental preferences, parenting styles, and unobserved neigh-

borhood selection. The results show clear evidence of age-specific effects for sons: removal of an abusive father before roughly age 11 leads to statistically significant reductions in sonsâ likelihood of perpetrating gender-based violence in adulthood, with the effects attenuating and becoming indistinguishable from zero at later ages. For daughters, point estimates are generally negative but less precisely estimated. Overall, the family fixed-effects evidence reinforces the interpretation that exposure to abusive fathers during early childhood plays a particularly important role in shaping sonsâ later violent behavior, and that reducing such exposure can meaningfully weaken intergenerational transmission within abusive families.

While family fixed effects address unobserved heterogeneity across households, father removal may still be endogenous if separations reflect unobserved maternal capacity, resources, or willingness to exit abusive relationships. For example, a mother who has more confidence may be better able to leave an abusive partner, and this could also separately lead to better outcomes for their children. To overcome this concern, we implement a judge instrumental variables (IV) design that leverages quasi-random assignment of cases to judges who differ in their propensity to incarcerate and remove defendants from the household. This approach compares women who are observationally identical, and should also result in women who are unobservably identical as well, except one experiences a removal of their abusive partner due to a harsher judge.

Using a residualized, leave-out measure of judge stringency constructed at the court-by-year-by-randomization-strata level, we find a strong first-stage relationship between judge severity and father incarceration. Our IV estimates capture the local average treatment effect of quasi-random father removal for families at the margin of incarceration. Consistent with the reduced-form and family fixed-effects results, the judge-IV estimates indicate that quasi-random removal of abusive fathers significantly reduces sons' risk of perpetrating gender-based violence in adulthood, particularly when removal occurs during formative years (between ages 1 and 11). These findings provide causal evidence that the intergenerational transmission of gender-based violence is not merely correlational but is shaped by childhood exposure and that policy-induced reductions in exposure to abusive fathers can disrupt this cycle.

We contribute to the existing economics literature on gender-based violence by providing

the first large-scale evidence on intergenerational transmission of gender-based violence and estimating whether the removal of abusive fathers disrupts or reinforces this cycle. Previous studies have found mixed evidence on the intergenerational transmission of GBV, but have been limited to survey-based measures (Ehrensaft *et al.*, 2003; Fergusson *et al.*, 2008; Brown *et al.*, 2015; Ehrensaft and Langhinrichsen-Rohling, 2022). Considering the broader domestic violence literature, a limited set of papers has documented that parents who were abused during their childhood by their parents are much more likely to abuse and maltreat their own children (Erten and Keskin, 2020; Armfield *et al.*, 2021). These studies suffer from small samples, potentially inconsistent definitions of childhood exposure, and reporting biases, especially underreporting by perpetrators (Heckert and Gondolf, 2000). By leveraging longitudinal administrative records, our study overcomes these limitations. This approach allows us to construct objective measures of exposure to GBV, link them to long-term outcomes, and provide novel causal evidence on whether childhood exposure increases adult perpetration or victimization, and whether the exit of a violent father and/or criminal justice policy interventions can break this cycle.

Our study also relates to the broader literature on the intergenerational transmission of criminal behavior and empirical evidence on the causal effects of parental incarceration on children's outcomes. Using Swedish administrative data, Hjalmarsson and Lindquist (2012) show that children with criminal fathers are twice as likely to have a criminal conviction than those with non-criminal fathers. Using Swedish adoption data, Hjalmarsson and Lindquist (2013) find that both pre-birth factors and post-birth factors contribute to parent-son associations in crime.[1] Using administrative data from Ohio combined with a judge fixed effects design, Norris *et al.* (2021) find that parental incarceration reduces children's likelihood of future incarceration and improves their adult neighborhood quality. Using Norwegian longitudinal data, Bhuller *et al.* (2018) find that imprisonment does not reduce fathers' recidivism but substantially lowers their employment, however, findings show no evidence that paternal incarceration affects a child's criminal activity or school performance. Using administrative data from Sweden, Grönqvist *et al.* (2024)

[1] A meta-analysis of 23 samples shows that children of criminal parents face substantially higher risks of criminal behavior themselves (pooled odds ratio = 2.4), and underscores the need for policies that disrupt the intergenerational cycle of crime (Besemer *et al.*, 2017).

find that the incarceration of a parent in childhood increases the risk of teen crime and pregnancy and reduces early-life employment. We contribute to this literature by examining the degree to which intergenerational transmission of gender-based violence occurs, and examining whether removal of abusive fathers in various ways, including quasi-random exposure to incarceration, has any significant effects on reducing the likelihood of children's likelihood of perpetrating or being a victim of gender-based violence in adulthood.

2 Data

Sweden We use data from multiple administrative sources, including the individual-level register data maintained by Statistics Sweden and the National Council for Crime Prevention. The data from Statistics Sweden span the entire Swedish population born between 1972 and 2022, covering demographic characteristics such as birth year, country of origin, and gender. These records include annual information on place of residence, formal sector earnings and employment, educational attainment, and detailed household information, such as household composition and family linkages. The family linkages allow us to link parents with children.

We further link these data to information from the Swedish Conviction Register and Crime Suspicion Register, which are maintained by the National Council for Crime Prevention. The Conviction Register includes all criminal convictions in Sweden from 1973 onward for individuals aged 15 and older. It provides details on the type of crime, the date of the crime, and court-imposed sentences, with each conviction potentially encompassing multiple crimes. This allows us to construct a complete criminal history for individuals.

The Crime Suspicion Register contains information on individuals aged 15 or older identified as suspects following police or prosecutorial investigations from 1995 onward. Since only a very small proportion of cases involving violence against women result in criminal convictions, we primarily focus on data indicating criminal suspicions of violence against women. This approach allows us to capture a broader spectrum of cases that might not progress to convictions.

This register includes data on the type of crime, dates of crime and suspicion, and outcomes such as indictments or convictions. Additionally, since the introduction of "Gross Violation of a

Woman's Integrity" (*Kvinnofridskränkning*) to the Swedish Penal Code in 1998 and the mandatory recording of victim gender for certain crime types, the register provides data on the gender of victims. For specific crimes, it also includes information on the relationship between the victim and the perpetrator, as well as the location of the crime. This enables us to identify cases of violence against women.

Using these sources of information, we construct two large-scale administrative datasets linking children to their fathers. First, we link men born in Sweden between 1970 and 1995 to their registered fathers. After restricting the sample to cases where both father and son were ever registered residents in Sweden during the period 1995–2020, we obtain a sample of 1,446,816 unique father-son pairs. Fathers may appear more than once if they have multiple sons in the eligible birth cohort. In cases where information is available about both biological and adoptive fathers, we retain information for the adoptive father. Second, we link women between 1970 and 1995 to their registered fathers. We have a sample of 1,374,446 unique father-daughter pairs after dropping pairs where one individual was not a registered resident during 1995–2020. Fathers may appear multiple times in the sample if they have more than one daughter born 1970-1995.

Table 1 provides descriptive statistics for sons and daughters in our sample. We find that sons and daughters with fathers suspected of committing gender-based violence are negatively selected on demographic, socioeconomic, and neighborhood characteristics. Their parents are more likely to be foreign born, less likely to attain education levels higher than high school, are less likely to be employed prior to child birth, and earn less on average. They are also more likely to live in more violent neighborhoods compared to children with non-violent fathers.

The primary variable of interest is an indicator for whether an individual (father or son) was ever suspected of any offense involving violence against women. This includes gross violation of a woman's integrity; physical assault; attempted or completed murder of a woman; unlawful threats; harassment; violations of restraining orders; and sexual violence including rape and molestation. The categorization captures both intimate partner violence and violence against female acquaintances or strangers. Appendix Table A1 presents descriptive statistics on the prevalence and composition of gender-based violence suspicions among fathers and sons in our sample. It

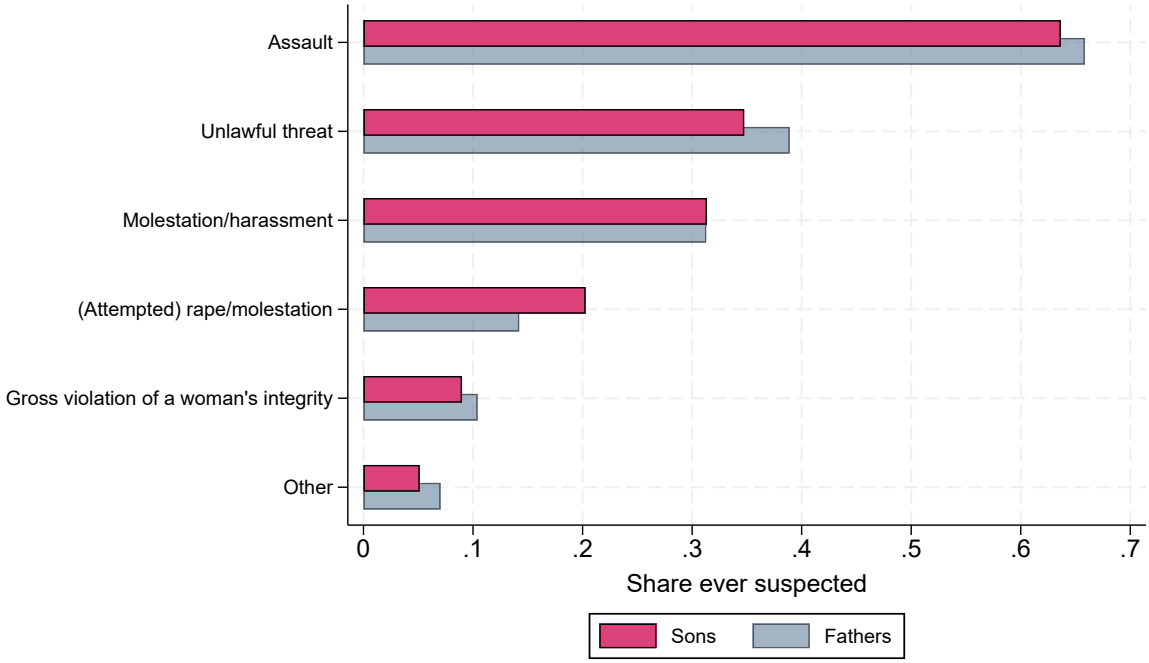
Table 1: Descriptive Statistics

	Sons		Daughters	
	Violent Father	Non-violent Father	Violent Father	Non-violent Father
Panel A: Demographics				
Mother age at birth	26.287	27.964	26.332	27.984
Mother aged -20 at birth	0.135	0.059	0.135	0.058
Mother aged 21–30 at birth	0.641	0.644	0.641	0.643
Mother aged 31–40 at birth	0.201	0.281	0.204	0.283
Mother aged 41+ at birth	0.009	0.012	0.010	0.012
Father age at birth	29.328	30.811	29.398	30.834
Father aged -20 at birth	0.044	0.014	0.042	0.014
Father aged 21–30 at birth	0.581	0.514	0.581	0.513
Father aged 31–40 at birth	0.323	0.408	0.324	0.409
Father aged 41+ at birth	0.053	0.064	0.053	0.064
Mother foreign born	0.307	0.157	0.311	0.154
Father foreign born	0.353	0.161	0.353	0.158
Child foreign born	0.136	0.075	0.138	0.075
Birth order father side	1.918	1.882	1.924	1.884
First child mother side	0.445	0.419	0.445	0.419
Second child mother side	0.315	0.366	0.315	0.365
Third child mother side	0.142	0.151	0.141	0.152
Fourth child+ mother side	0.085	0.061	0.087	0.061
Birth order mother side	1.964	1.906	1.973	1.908
First child father side	0.448	0.423	0.445	0.422
Second child father side	0.308	0.356	0.309	0.355
Third child father side	0.145	0.152	0.145	0.152
Fourth child+ father side	0.099	0.070	0.101	0.070
Live with father at birth	0.819	0.936	0.823	0.935
Panel B: Baseline socioeconomic status				
More than HS education mother	0.242	0.356	0.243	0.355
HS education mother	0.539	0.484	0.541	0.485
Less than HS education mother	0.210	0.153	0.207	0.153
More than HS education father	0.181	0.294	0.181	0.294
HS education father	0.521	0.462	0.520	0.461
Less than HS education father	0.290	0.237	0.291	0.238
Employment mother (year of birth)	0.790	0.809	0.791	0.808
Employment mother (year before birth)	0.781	0.827	0.782	0.827
Employment father (year of birth)	0.881	0.960	0.882	0.960
Employment father (year before birth)	0.869	0.952	0.873	0.952
Earnings mother (year of birth)	125682.287	133989.495	125705.598	134095.636
Earnings mother (year before birth)	126225.946	142629.488	126393.276	142645.483
Earnings father (year of birth)	190759.111	246721.517	191251.317	246804.468
Earnings father (year before birth)	183394.881	237158.164	184245.064	237207.693
Panel C: Baseline neighborhood characteristics				
Population	10776.691	9355.516	10749.696	9353.120
Share of immigrants	0.136	0.104	0.135	0.104
Employment rate	0.878	0.864	0.878	0.864
Mean earnings	195214.659	189776.419	195307.569	189737.462
More than high school	0.248	0.230	0.248	0.229
Less than high school	0.294	0.315	0.294	0.315
Offenses per 100,000	5374.958	4686.000	5356.621	4687.273
Offenders per 100,000	953.024	857.588	950.777	857.757
Assault rate per 100,000	83.994	67.783	83.357	67.794
Homicide rate per 100,000	1.077	0.800	1.042	0.800
Obs.	96,799	1,350,017	93,433	1,281,013

Note: This table reports the mean values of variables for sons and daughters born between 1970 and 1995. Columns (1)–(2) report statistics for sons with violent vs. non-violent fathers, while columns (3)–(4) report statistics for daughters with violent vs. non-violent fathers. The violence is defined as being suspected of committing gender-based violence.

reports the share of individuals ever suspected of specific offenses, separately for sons and fathers, and then conditions these distributions on having any criminal suspicion. Figure 1 plots these conditional distributions in a bar chart. These exhibits highlights two key patterns. First, across both generations, the most common categories of suspected offenses are assault of a woman, unlawful threats, and harassment, which together account for the majority of gender-based violence suspicions. Second, fathers have slightly higher prevalence rates than sons for most of the categories, which is consistent with their longer time at risk of perpetration since they are older.

Figure 1: Gender-Based Violence among Fathers and Sons Conditional on Suspicion



Notes: This figure shows the share of fathers and sons suspected of each category of gender-based violence, conditional on having been suspected of any crime. Sons re born between 1970 and 1995 in Sweden and matched with their fathers using registry information. All data come from Crime Suspicion Register of Sweden.

United States We supplement the Swedish data analysis with analysis of United States data. This analysis is currently in progress and has not been disclosed.

3 The Intergenerational Transmission of Gender-Based Violence

Children who witness violent behaviors towards women in their households may learn that violence is an acceptable response to interpersonal conflict, and that violence and dominance to-

wards women is normal. In this section, we leverage our data to provide the first large-scale descriptive estimates of the degree to which this intergenerational transmission of gender-based violence occurs. We do so in two parts. First, we document the degree to which sons who are exposed to fathers committing gender-based violence are more likely to commit gender-based violence themselves. In addition, we also estimate the degree to which daughters who are exposed to fathers who commit gender-based violence are more likely to choose cohabiting partners who commit gender-based violence.

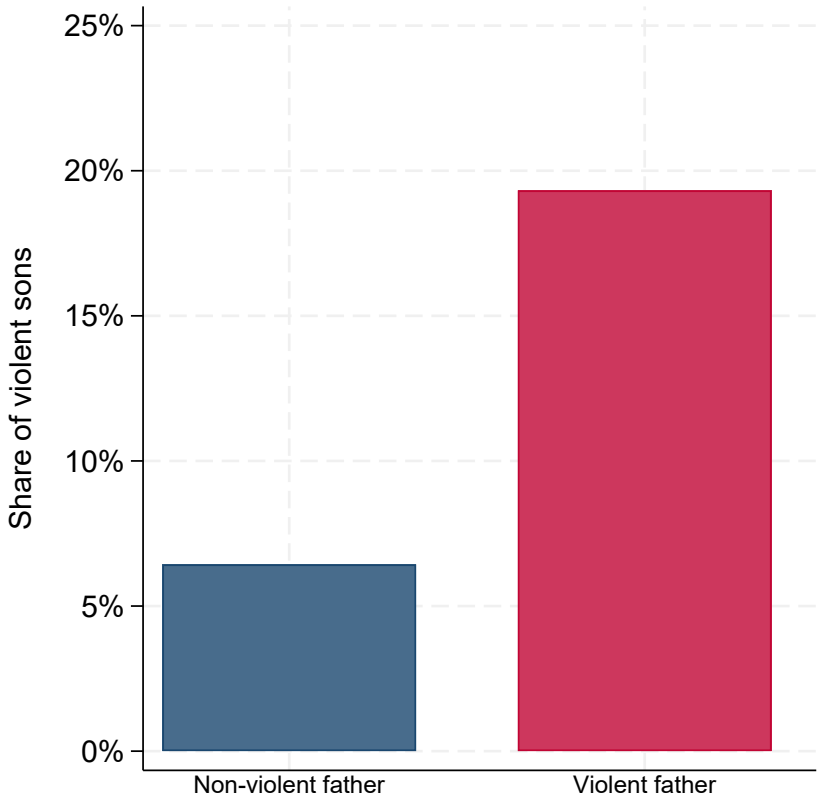
Intergenerational Transmission of Perpetration Figure 2 reports the raw correlation between having a father with a history of gender-based violence and whether the son commits gender-based violence when he becomes an adult. The incidence of violence against women among sons varies markedly by paternal history. In Sweden, 6.4% of sons whose fathers were not suspected of gender-based violence grow up to commit violence against a woman. In contrast, 19.3% of sons whose fathers have a history of violence against women also commit violence against women as adults. In the United States, we find [results in progress and not yet disclosed].

However, these raw correlations may simply capture selection on observables: as shown in Table 1, sons of fathers who become violent are born into very different families compared with the general population. These families are poorer, less educated, and much more likely not to be born in Sweden on average compared to other families in Sweden. To account for this, Figure 3 reports a series of linear probability model (LPM) capturing whether sons with a violent father have a higher likelihood of committing violence against women. In this figure, we progressively control for other potential drivers of the raw correlation shown in Figure 2.

Starting with the raw linear probability, consistent with Figure 2, the blue bar in Figure 3 shows that sons whose fathers were suspected of violent offenses towards women are 12.9 percentage points (p.p.) more likely to be suspected of violence against women themselves in Sweden.[2] As the row below the coefficient estimates shows, this implies that having a violent father increases the likelihood of perpetration of the son by 177 percent relative to sons without a vi-

[2] Appendix Table A3 reports the OLS estimates that correspond to the estimates plotted in Figure 3.

Figure 2: Raw Correlation for Sons' Gender-Based Violence Perpetration

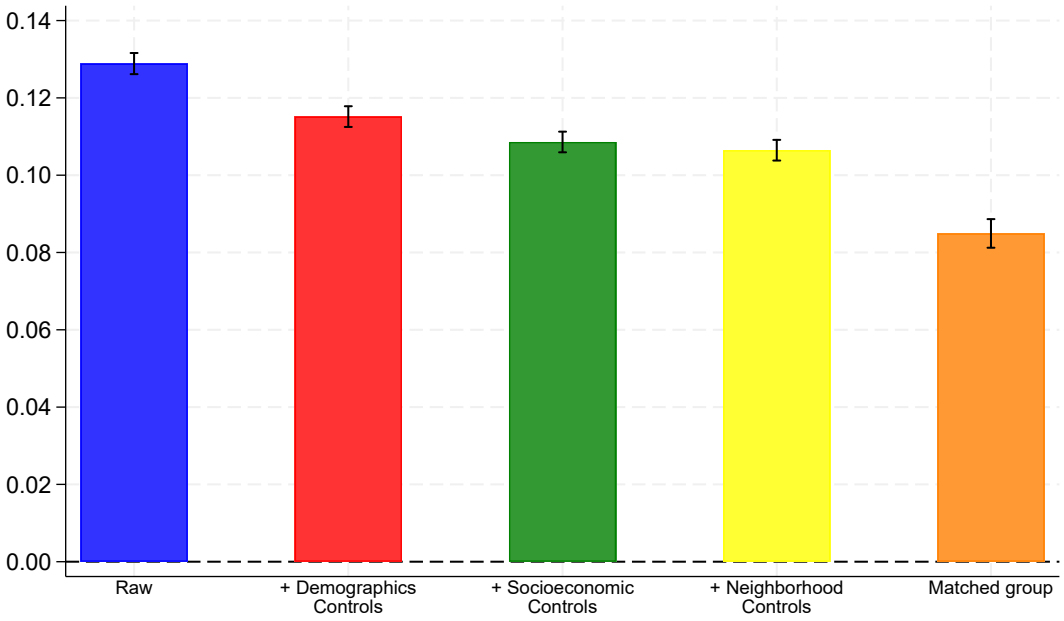


Notes: This figure shows the intergenerational transmission of violence perpetration against women in Sweden. The blue bar plots the share of sons suspected of committing violence against women among sons whose fathers were not suspected of such offenses. The red bar plots the corresponding share among sons whose fathers were suspected perpetrators.

olent father. Consistent with this raw correlation, the odds ratio estimates in Appendix Figure A1 shows that sons with GBV-suspected fathers are 3.5 times more likely to be suspected of such violence themselves as adults.

The red bar in Figure 3 adds family demographic controls, including father's and son's birth year fixed effects, indicators for whether the mother, father, or son is foreign born, and birth order indicators for both parents. The association decreases slightly but remains strong: having a GBV suspected father increases the likelihood of being suspected of GBV in adulthood by 11.5 percentage points. The green bar adds socioeconomic controls, including parental educational attainment, employment at birth, and adjusted parental earnings at birth, which reduces the association slightly to 10.9 percentage points. Adding neighborhood controls, including parish pop-

Figure 3: Intergenerational Transmission of GBV Perpetration: Linear Probability Estimates

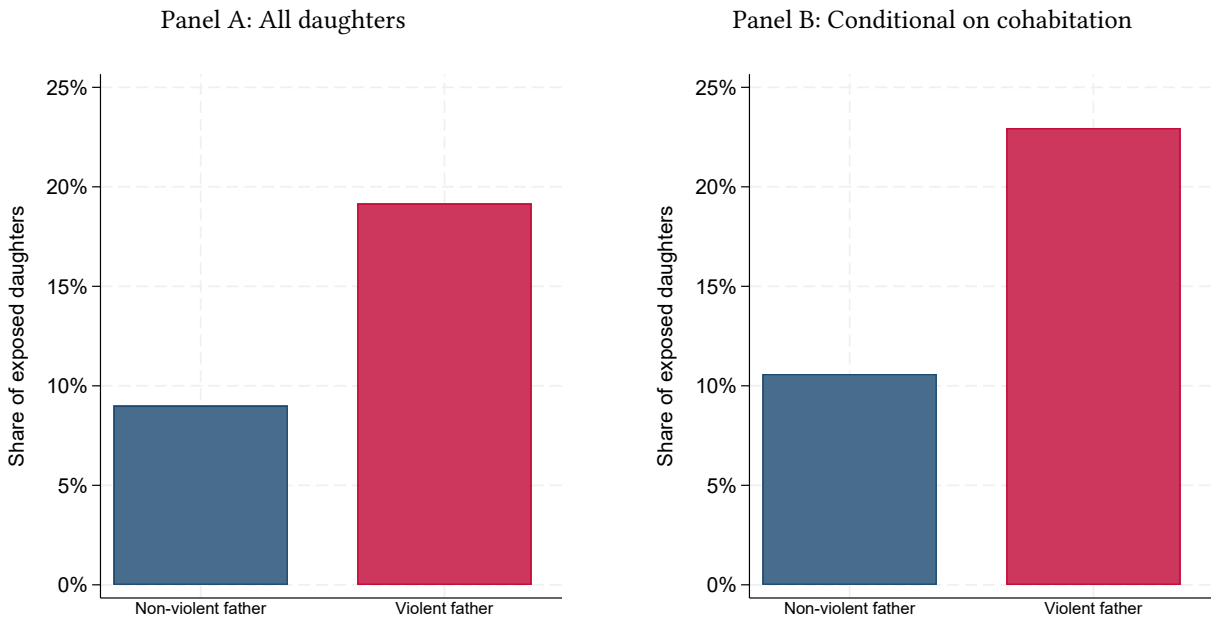


Notes: This figure illustrates the intergenerational transmission of gender-based violence in Sweden. The blue bar reports the coefficient from a linear probability model regressing an indicator for the son being suspected of GBV on an indicator for the father being suspected of GBV, without controls. The red bar adds demographic controls, including birth-year fixed effects for father and son, indicators for whether the mother, father, or son is foreign born, and birth-order indicators from both parents. The green bar adds socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. The yellow bar adds neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school), and reports the corresponding coefficient estimate. Standard errors are clustered at the father level. All bars display 95% confidence intervals.

ulation size, immigrant share, employment rate, average earnings, share of university-educated residents, and rates of offenders, assaults, and homicides per 100 thousand residents, does not meaningfully change the estimate, which remains at 10.6 percentage points. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles at birth, birth order, parental age in five-year bins, and parental education. Even in this matched sample, the intergenerational transmission remains substantial: sons with GBV suspected fathers are 8.5 percentage points more likely to be suspected of GBV in adulthood. This implies that having a violent father increases a son's likelihood of perpetrating violence by 116% relative to a matched group of sons with similar socioeconomic characteristics but without

a violent father. In this matched sample, the odds ratio estimates in Appendix Figure A1 show that sons with GBV suspected fathers are 2.2 times more likely to be suspected of such violence in adulthood compared to otherwise similar sons who did not have a violent father.[3]

Figure 4: Raw Correlation for Gender-Based Violence Victimization



Notes: This figure shows the intergenerational transmission of GBV victimization in Sweden. Panel A includes all daughters in our sample, and Panel B includes daughters who have ever cohabited with a partner. In both panels, the blue bars report the share of daughters who cohabit with a partner suspected of committing violence against women among those whose fathers were not suspected of such offenses. The red bars report the corresponding share for daughters whose fathers were suspected perpetrators.

Intergenerational Transmission of Victimization Figure 4 reports the share of daughters whose fathers were suspected of violence against women who later form relationships with partners who commit gender-based violence. We use being in a cohabiting or marital relationship with a man who has ever been suspected of gender-based violence as a proxy for potential victimization, and this outcome is measured over the years 1990 to 2020.

Panel A shows that among all women with no paternal history of gender-based violence, 9% are in relationships with a man who committed GBV. This figure rises substantially to 19.2% for women whose fathers were suspected of violence against women, constituting more than double the victimization risk. Panel B of Figure 4 shows that among women who have ever had a

[3] Appendix Table A6 provides the corresponding odds ratio estimates.

partner, 10.6% of women whose fathers were not suspected of gender-based violence have been in relationships with violent partners, compared to 22.9% of those whose fathers were suspected perpetrators.

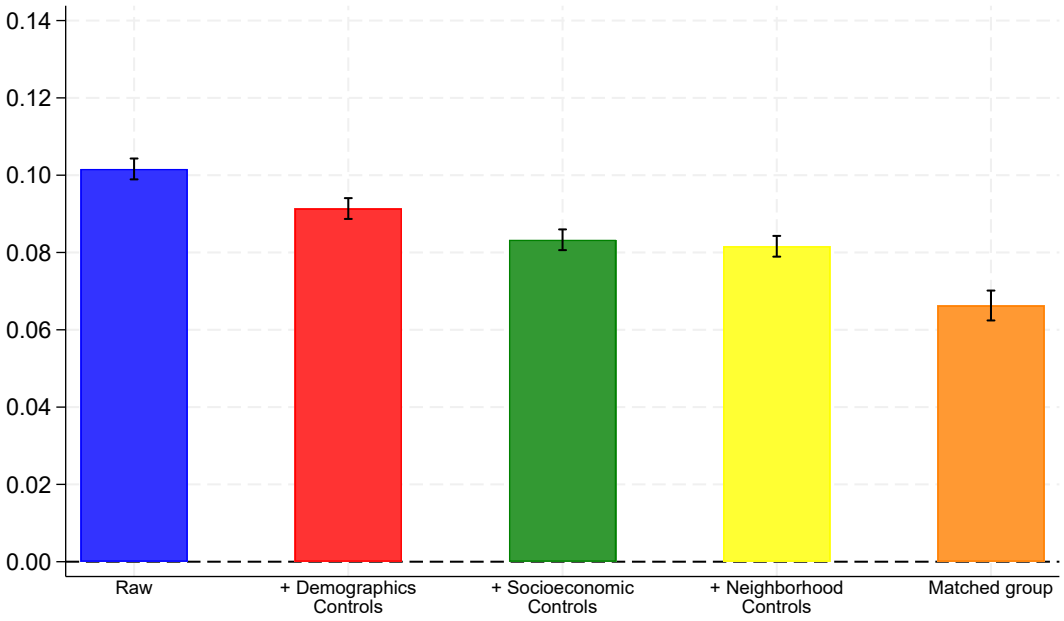
In Appendix Table A2, we test whether having a father suspected of gender-based violence is significantly associated with the daughters' probability of ever cohabiting with a partner. Roughly 85% of women in our sample have ever cohabited with a partner. Having a GBV-suspected father reduces this likelihood by 1.4 p.p. in OLS estimates with no controls added; however, the inclusion of controls in a sequential manner as well as the estimation based on a matched sample increases the probability of cohabitation by 1.5–1.8 p.p., which corresponds to 2% relative to the sample mean. Despite being small, the increase in cohabitation likelihood may increase potential exposure to violent partners, potentially biasing the OLS estimates for intergenerational transmission upwards in the conditional sample. For this reason, we focus our remaining analysis on the unconditional sample, presenting the conditional estimates as supplementary descriptive evidence in the appendix.

In Figure 5, we account for a richer set of covariates to examine whether the observed associations could be explained by other family, socioeconomic, or neighborhood characteristics (see Appendix Figure A4 for estimates conditional on cohabitation). Beginning with the raw linear probability, consistent with Figure 4, the blue bar in Figure 5 shows that daughters whose fathers were suspected of violent offenses towards women are 10.2 p.p. more likely to cohabit with partners who were suspected of gender-based violence in Sweden.[4] This implies that having a violent father increases the victimization likelihood of the daughter by 105 percent relative to daughters without a violent father. The odds ratio estimates in Appendix Figure A2 show that daughters with GBV-suspected fathers are 2.4 times more likely to live together with violent partners compared to those without violent fathers.[5]

[4] Appendix Table A4 reports the OLS estimates that correspond to the estimates plotted in Figure 5, which include the sample of women conditional on ever cohabiting with a partner. Appendix Table A5 reports the OLS estimates that correspond to the estimates plotted in Figure A4, which include all women regardless of cohabitation status.

[5] Appendix Table A7 provides the corresponding odds ratio estimates. We also provide the odds ratio estimates conditional on ever cohabiting with a partner in Appendix Figure A3 and Table A8.

Figure 5: Intergenerational Transmission of GBV Victimization: Linear Probability Estimates



Notes: This figure illustrates the intergenerational transmission of victimization based on gender-based violence in Sweden. The sample includes all women born between 1970 and 1995 in Sweden, including those who have not cohabited with a partner. The blue bar reports the coefficient from a linear probability model regressing an indicator for the daughter cohabiting with a partner who was suspected of GBV on an indicator for the father being suspected of GBV, without controls. The red bar adds demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. The green bar adds socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. The yellow bar adds neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school), and reports the corresponding coefficient estimate. Standard errors are clustered at the father level. All bars display 95% confidence intervals.

The remaining bars in Figure 5 gradually add demographic controls, socioeconomic controls, and neighborbood controls, with LPM estimates only slightly declining from 9.1 to 8.2 p.p. The orange bar on the far right shows the estimate from a matched sample where we select women for the control group women who are observationally identical on a wide range of variables: parental immigrant status, employment at birth, earnings quintiles at birth, birth order, parental age in five year bins, and parental education. The intergenerational association declines only slightly to 6.6 percentage points in the matched sample, implying a 68 percent increase in the likelihood of

experiencing GBV victimization from exposure to a GBV father. In this matched sample, the odds ratio estimates in Appendix Figure A2 show that daughters with GBV-suspected fathers are 1.7 times more likely to cohabit with violent partners in adulthood compared to otherwise similar daughters who did not have a violent father.

Taken together, these findings provide the first large-scale evidence that exposure to fathers who commit gender-based violence is a powerful predictor of both sons' and daughters' trajectories in adulthood. Sons who grow up with fathers suspected of gender-based violence are substantially more likely to commit gender-based violence themselves, even after accounting for a rich set of demographic, socioeconomic, and neighborhood characteristics and comparing them with observationally identical men.

Daughters exposed to paternal violence also face significantly elevated risks: they are far more likely to form cohabiting relationships with men who have histories of violent behavior, and this pattern persists even when comparing them to closely matched control groups. The consistency of these results across multiple specifications points to a robust learning mechanism through which gender-based violence is transmitted across generations, shaping both perpetration and victimization risks. These patterns highlight how early-life exposure to violent male role models can normalize abusive behavior, influencing the partners children choose and the ways in which they navigate conflict in adulthood.

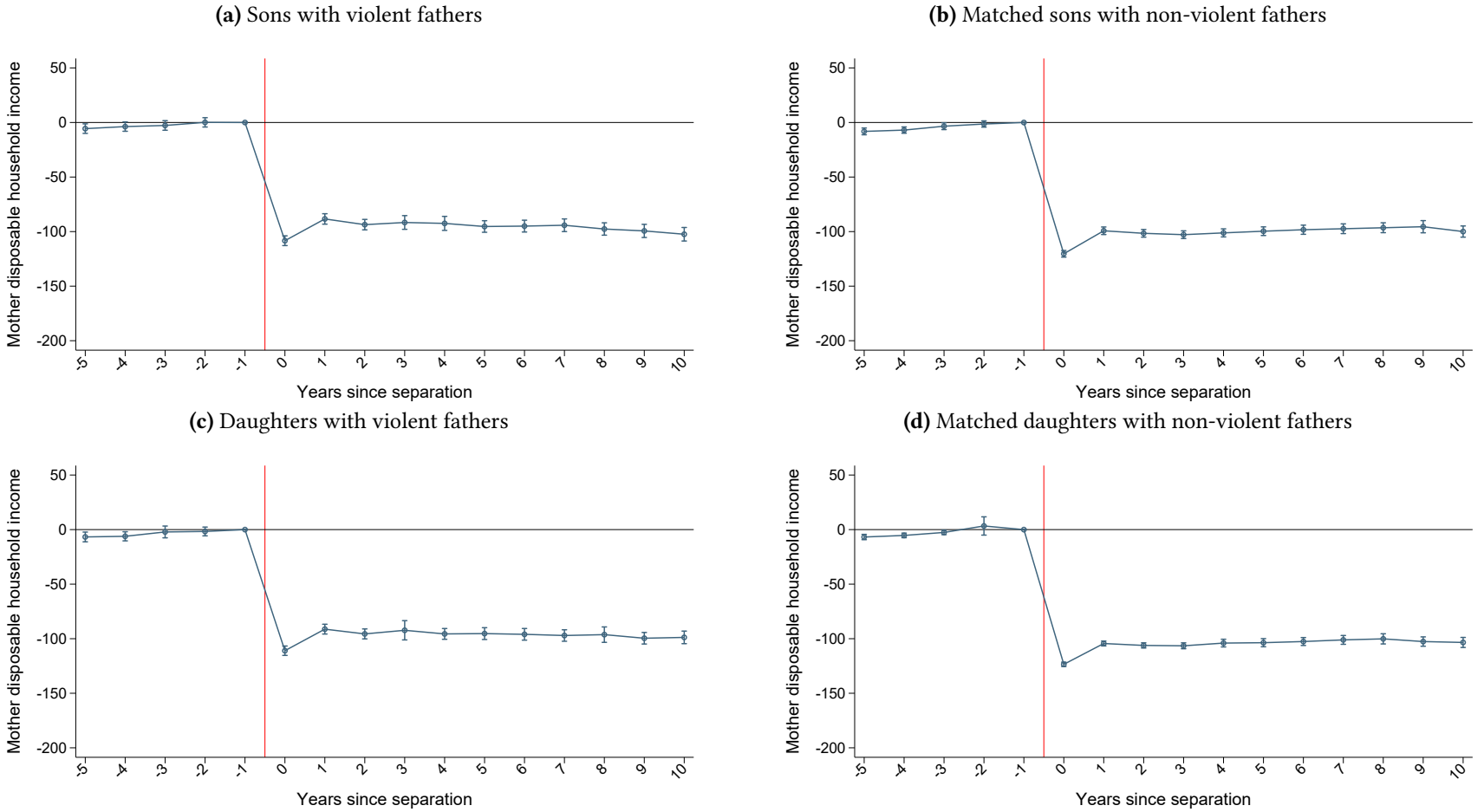
4 Breaking Transmission: Father Removal

In this section, we examine whether reducing exposure to fathers who commit gender-based violence can break the intergenerational cycle of gender-based violence documented in the previous section. *A priori*, the effect of removing fathers on the intergenerational transmission of violence is ambiguous. Removing an abusive father from the household–through separation, incarceration, or similar events–may reduce children's exposure to harmful social learning mechanisms, thereby weakening the channel through which violent behavior is modeled and reproduced across generations (Bandura, 1977; Ehrensaft and Langhinrichsen-Rohling, 2022; Kimber *et al.*, 2018).

However, removing the father likely also reduces household income. Prior research has shown that reductions in household income increase the risk of domestic violence (Bhalotra *et al.*, 2025; Erten *et al.*, 2022; Carr and Packham, 2021; Leslie and Wilson, 2020). As a result, the negative income shock from father removal could result in greater transmission of gender based violence.

To explore the degree to which income drops might be salient for these families, we first estimate the effects of father removal on household income. Figure 6 reports estimated changes in household income relative to the year prior to father removal for the sample of sons and daughters. Panel A presents results for sons whose fathers are suspected of committing gender-based violence, while Panel B shows the corresponding estimates for a matched control group of sons who are observationally similar but whose fathers are not suspected of such violence. The bottom panels present these results for the sample of daughters.

Figure 6: Household Income Around Father Removal



Notes: The figures track household income around the time of father's separation from the household for children born between 1970 and 1995. The outcome is annual disposable household income in 2020 Swedish kronas. The control groups in Panels B and D are matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). Panel A shows the changes in household income for sons with fathers suspected of gender-based violence, Panel B shows it for a matched control group of sons whose fathers are not suspected of gender-based violence, Panel C shows it for daughters with fathers suspected of gender-based violence, and Panel D shows it for a matched control group of daughters whose fathers are not suspected of gender-based violence. The x-axis represents years relative to separation with 0 marking the year of separation (indicated by the red vertical line). The omitted category is the year before separation. The regression includes calendar year fixed effects and child birth year fixed effects. The standard errors are clustered at the father level. Disposable income measure includes household income from labor, capital, and transfers minus taxes, plus non-taxable benefits, including child allowance, social assistance, social security benefits, and other transfers. The coefficient estimates and 95% confidence intervals are plotted.

The event-study results in Figure 6 reveal a sharp decline in household income following father removal for both sons and their matched control group, as well as for daughters and their corresponding matched controls. We track household income for mothers, who generally retain primary custody of children following separation. We include calendar year fixed effects and child birth cohort fixed effects to remove broader time trends and cohort-specific patterns in household income, and cluster standard errors at the father level. At the time of separation, household income drops by roughly SEK 150,000, amounting to about half of the average pre-separation income of SEK 300,000–350,000. Although there is a small rebound of about SEK 20,000 the year after, income remains persistently low thereafter. Longer term, these children experience income declines in their primary household of roughly SEK 130,000, or 40% below baseline levels, even a decade after separation. Although the decline in household income is slightly larger for the matched control groups, the difference relative to sons with abusive fathers is not statistically significant.

The large and persistent income losses documented in Figure 6 underscore why the effect of father removal on children's later-life GBV outcomes is theoretically ambiguous. Separation sharply reduces household resources and generates long-lasting financial strain, a mechanism that prior work shows can reduce adult earnings and increase the risk of incarceration, mortality, and teen births (Johnston *et al.*, 2025). At the same time, separation from an abusive father may disrupt the social learning channel through which violent behavior is modeled and transmitted across generations.

These household income event-study results motivate a careful estimation strategy to understand whether removing fathers can reduce the intergenerational transmission of GBV, after accounting for income effects. Specifically, we will compare families who both experience a father removal. However, in one family the father has committed GBV, while in the other family there is no report of GBV for the father. This approach could allow us to obtain the exposure effect of father separation net of the negative income shock, which Figure 6 demonstrates is similar across observationally identical families.

Main Estimates Our estimation strategy takes two steps. First, we estimate the effects of father removal on sons' probability of perpetrating violence against women during adulthood separately for sons of fathers who commit GBV and sons of fathers with no record of GBV. To identify whether there are periods of special sensitivity to father removal, we plot these estimates by son's age at separation, for each age between 1 and 19. Formally, we estimate:

$$Y_{ic} = \alpha_0 + \sum_{m=1,\neq 19}^{19} \beta_m R_{im} + \gamma_c + \varepsilon_{ic} \tag{1}$$

where Y_{ic} is an indicator for whether individual i in cohort c is suspected of committing violence against women at any point in their life; R_{im} is an indicator for whether individual i's father was removed or separated from the household when the child was age m; and γ_c is a birth cohort fixed effect. The coefficients β_m capture the effect of separating from the father at each developmental age from 1 to 19. We normalize relative to separation at age 19 by omitting that coefficient. For all estimates, we cluster standard errors at the father level.

To remove the income effect from father removal, we combine the two samples and estimate the effects of father removal for sons with GBV fathers relative to sons without GBV fathers using the following specification:

$$Y_{ic} = \alpha_0 + \sum_{m=1,\neq 19}^{19} \beta_m (R_{im} \times V_{ic}) + \delta X_{ic} + \gamma_c + \varepsilon_{ic} \tag{2}$$

where V_{ic} is an indicator for whether the father was suspected of committing violence against women, X_{ic} denote a vector of controls for birth-year fixed effects for father and son, indicators for whether the mother, father, or son is foreign born, and birth-order indicators from both parents, parental educational attainment (high school and university) at birth, employment status at birth, adjusted parental earnings at birth, parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. All other variables are defined as in Equation 1.

In a complementary identification strategy, we estimate the same specification using a matched sample, where we find sons who experience father removal at the exact same age and are obser-

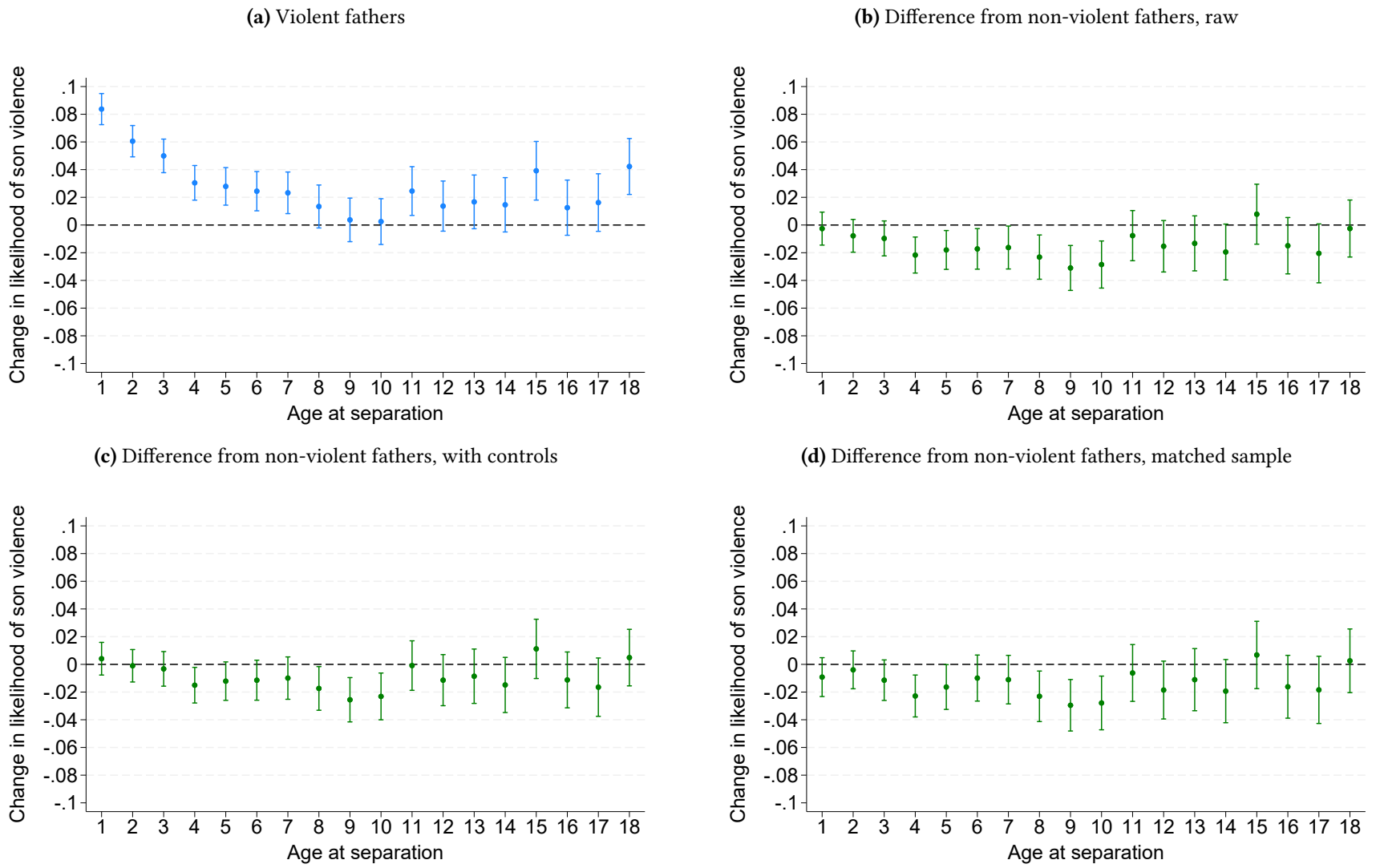
vationally identical based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). For all estimates, we cluster standard errors at the father level.

Figure 7 Panel A reports the effects of removing of a violent father from the household on the son's GBV risk in adulthood, i.e. the estimates from Equation 1. We observe that GBV likelihood increases for removals of the abusive father particularly at younger ages, consistent with the detrimental effects of income loss dominating the exposure channel. A similar age gradient is present for the sample of non-violent households as displayed in Appendix Figure A5.

To remove the detrimental effects of income losses that take place during father removal, we estimate Equation 2, and plot the coefficient estimates first without any controls. The estimates shown in Panel B indicate that removing the income channel results in a significant reduction in the probability that sons whose fathers commit GBV go on to commit GBV themselves. Adding controls to this specification only slightly reduces this negative effect as shown in Panel C. Finally, Panel D shows the estimates from a matched sample where the non-violent fathers are observationally identical to violent fathers in terms of demographic and socioeconomic characteristics. Even in this matched control sample we see a reduction in the likelihood that sons commit GBV in adulthood if their father is removed in childhood.

Figure A7 restricts the estimation sample to families that do not reunite after the initial separation, thereby eliminating any subsequent exposure to fathers following the first separation. The estimated effects on sons' later-life risk of perpetrating gender-based violence are larger in this sample, reinforcing the interpretation that fully removing exposure to abusive fathers more effectively disrupts the intergenerational transmission of gender-based violence.

Figure 7: Effects of Father Removal on Sons' GBV Perpetration by Age at Separation



Notes: The figure shows the estimates for the effects of father removal on sons' GBV perpetration by their age at separation. Panel A plots the estimates for the sample of sons with violent fathers by estimating Equation 1. Panel B plots the estimated differences between sons with violent fathers and those without violent fathers using Equation 2, but it does not include controls. Panel C adds controls for demographic, socioeconomic and neighborhood characteristics as specified in Equation 2. Panel D uses a matched sample to estimate Equation 2 where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

Table 2 shows that these estimates are also economically meaningful. We pool estimates for separation at younger ages (1-11) and older ages (12-18) to capture differences between early childhood and adolescence. Specifically, we estimate the following specification:

$$Y_{ic} = \alpha_0 + \beta_1\big(R_{i,1\text{–}11} \times V_{ic}\big) + \beta_2\big(R_{i,12\text{–}18} \times V_{ic}\big) + \beta_3 V_{ic} + \delta X_{ic} + \gamma_c + \varepsilon_{ic}. \tag{3}$$

where $R_{i,1\text{–}11}$ and $R_{i,12\text{–}18}$ are indicators for father removal occurring when the child is aged 1-11 or 12-18, respectively. Removal at age 19 serves as the omitted reference category. The indicator V_{ic} equals one if the father was suspected of committing violence against women. The coefficients β_1 and β_2 capture how the effect of having a violent father differs when removal occurs earlier in childhood or adolescence, relative to removal at age 19. All other variables are defined as in Equation 2.

Table 2 presents the results. The first column shows approximately a 1 p.p. decline in the probability that a son is suspected of gender-based violence when a violent father is removed during either early childhood (ages 1-11) or adolescence (ages 12-18), with no statistically significant difference between the two estimates. Adding demographic, socioeconomic, and neighborhood controls in column 2 attenuates the estimates and renders them statistically insignificant.[6] In contrast, column 3, which uses a matched sample, again yields effects of roughly 1 percentage point. In this matched specification, the estimated effect of father removal at ages 1-11 is about 30% larger than the corresponding estimate for ages 12-18 (albeit not statistically different), suggesting that reductions in exposure to violent fathers earlier in childhood may have larger effects than those occurring during adolescence.

[6] Appendix Table A9 presents the same results using a more granular division of age groups. These estimates suggest that the largest declines in perpetration risk are concentrated among sons separated from abusive fathers between ages 3 and 10, which remain robust to the inclusion of additional controls.

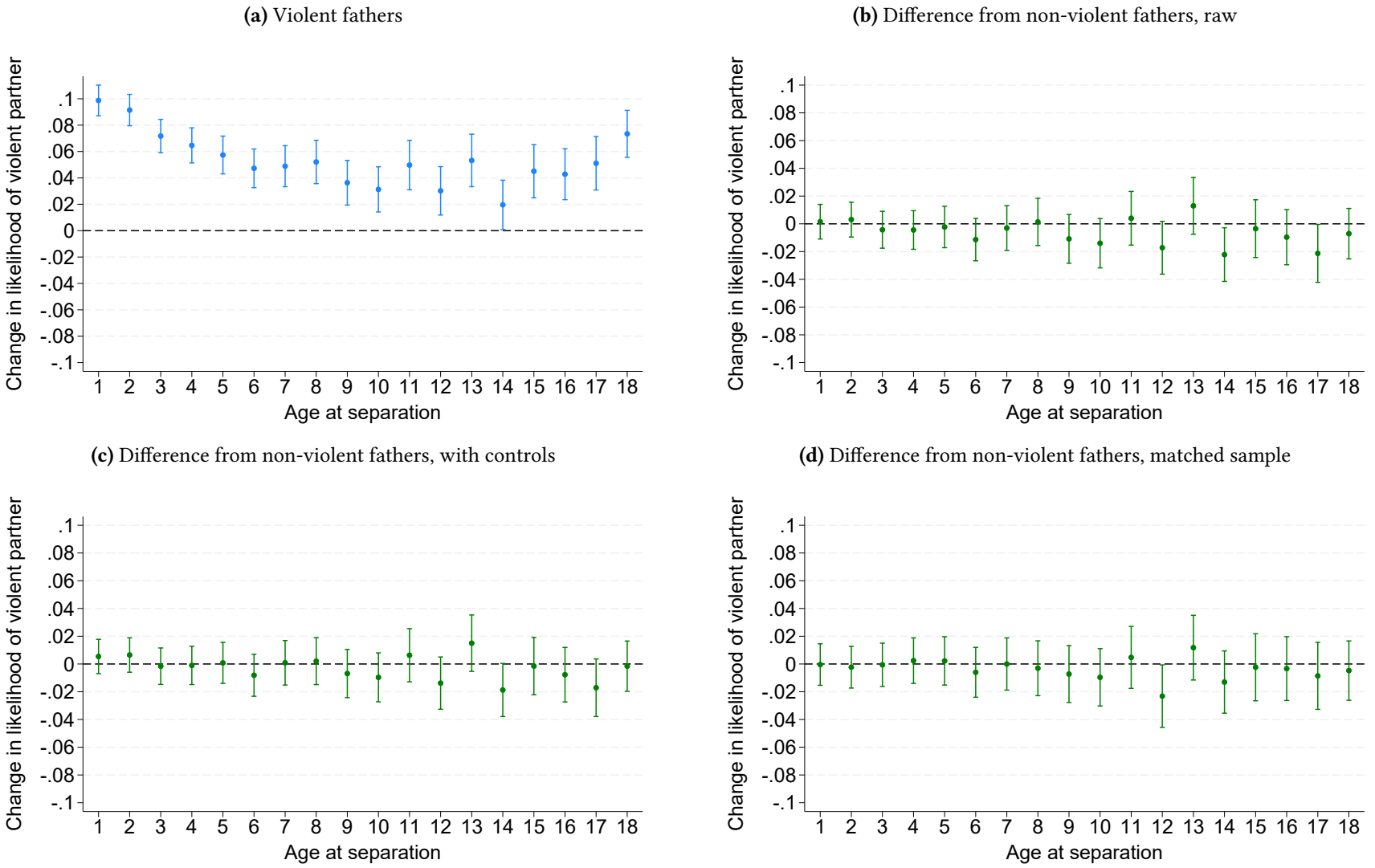
Table 2: Effects of Father Removal on Sons' GBV Perpetration by Age Groups

	Raw difference (1)	Controls added (2)	Matched sample (3)	Family fixed effects (4)
Separation age 1-11 X Violent father	-0.00977***	-0.00430	-0.0137***	-0.0221**
	(0.00336)	(0.00331)	(0.00381)	(0.0109)
Separation age 12-18 X Violent father	-0.0118**	-0.00711	-0.0107**	0.0184*
	(0.00463)	(0.00457)	(0.00520)	(0.0103)
Violent father	0.0956***	0.0817***	0.0835***	
	(0.00263)	(0.00259)	(0.00295)	
Separation age 1-11	0.0513***	0.0447***	0.0593***	0.0180***
	(0.000639)	(0.000641)	(0.00187)	(0.00206)
Separation age 12-18	0.0340***	0.0313***	0.0324***	0.0164***
	(0.000870)	(0.000864)	(0.00245)	(0.00207)
Constant	0.0409***	-0.105***	0.00439	0.0808***
	(0.000224)	(0.00787)	(0.00721)	(0.00289)
Observations	1,231,450	1,225,728	258,136	1,231,450
Mean dep. var.	0.0606	0.0606	0.110	0.0606
Number of unique fathers				893,221

Notes: The table shows the estimates for the effects of father removal on sons' GBV perpetration by their age groups. Column (1) shows the estimated differences between sons with violent fathers and those without violent fathers by separation age groups, but it does not include controls. Column (2) adds controls for demographic, socioeconomic and neighborhood characteristics. Column (3) uses a matched sample where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). Column (4) adds family fixed effects. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level and reported in parentheses. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Next, we use corollary specifications to Equations 1 and 2 to estimate the effects of removing abusive fathers from the household on daughters' risk of GBV victimization in adulthood. As a proxy for such victimization, we again focus on whether she cohabits with a partner who was suspected of committing GBV. Figure 8 presents the results. Panel A shows the estimates for the sample of daughters with violent fathers. We again see an elevated risk of exposure to GBV for daughters where the separation from the father occurred in earlier ages. However, these estimates also embody the negative income effect that takes comes with father removal.

Figure 8: Effects of Father Removal on Daughters' GBV Victimization by Age at Separation



Notes: The figure shows the estimates for the effects of father removal on daughters' GBV victimization proxied by their cohabitation with a partner suspected of committing GBV by their age at separation. Panel A plots the estimates for the sample of daughters with violent fathers by estimating a corollary specification to Equation 1. Panel B plots the estimated differences between daughters with violent fathers and those without violent fathers using Equation 2, but it does not include controls. Panel C adds controls for demographic, socioeconomic and neighborhood characteristics as specified in Equation 2. Panel D uses a matched sample to estimate a corollary specification to Equation 2 where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

Differencing out these estimates using the sample of non-violent fathers[7], we observe some negative estimates for later ages across Panels B (raw difference), Panel C (with controls), and Panel D (matched sample). However, most estimates hover around zero, and are not statistically significant, suggesting that the removal of abusive fathers from the household does not generate comparable effects for daughters' risk of cohabiting with an abusive father. The pooled estimates reported in Table 3 are also small and do not statistically differ from zero for most of them.

Table 3: Effects of Father Removal on Daughters' GBV Victimization by Age Groups

	Raw difference (1)	Controls added (2)	Matched sample (3)	Family fixed effects (4)
Separation age 1-11 X Violent father	0.00154	0.00396	-0.00143	-0.0130
	(0.00349)	(0.00346)	(0.00410)	(0.0122)
Separation age 12-18 X Violent father	-0.0131***	-0.00890*	-0.00628	-0.0138
	(0.00459)	(0.00455)	(0.00533)	(0.0109)
Violent father	0.0663***	0.0546***	0.0529***	
	(0.00271)	(0.00269)	(0.00317)	
Separation age 1-11	0.0660***	0.0588***	0.0712***	0.0276***
	(0.000747)	(0.000751)	(0.00220)	(0.00260)
Separation age 12-18	0.0594***	0.0536***	0.0509***	0.0342***
	(0.000960)	(0.000953)	(0.00277)	(0.00234)
Constant	0.0620***	-0.200***	-0.00153	0.104***
	(0.000286)	(0.00870)	(0.00871)	(0.00359)
Observations	1,170,215	1,164,773	249,664	1,170,215
Mean dep. var.	0.0886	0.0886	0.140	0.0886
Number of unique fathers				863,673

Notes: The table shows the estimates for the effects of father removal on daughters' GBV victimization by their age groups. Column (1) shows the estimated differences between daughters with violent fathers and those without violent fathers by separation age groups, but it does not include controls. Column (2) adds controls for demographic, socioeconomic and neighborhood characteristics. Column (3) uses a matched sample where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). Column (4) adds family fixed effects. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level and reported in parentheses. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

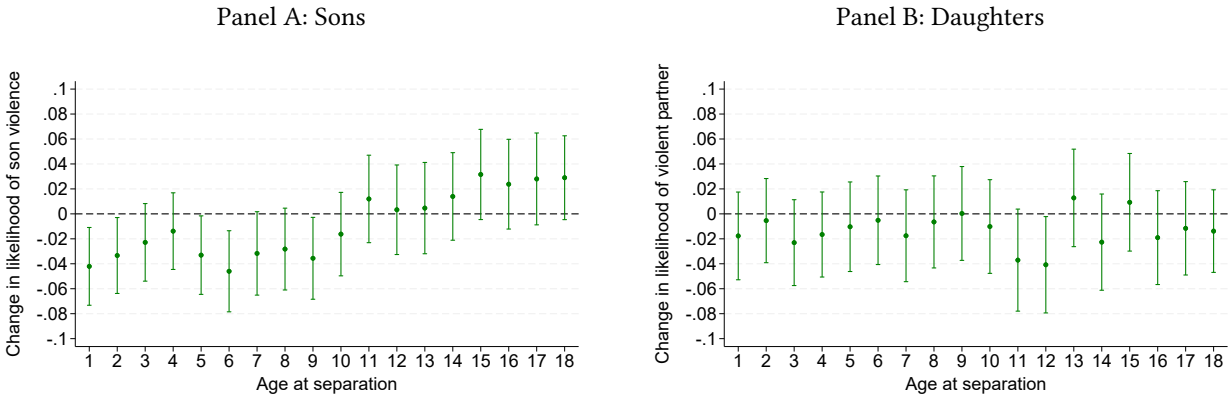
[7] Appendix Figure A6 provides the estimates for the sample of non-violent fathers based on a corollary of Equation 1.

Appendix Figure A8 restricts the analysis to families that remain permanently separated after the initial breakup, ensuring that daughters are not subsequently re-exposed to their fathers. In this subsample, the estimated effects on daughters' later-life risk of partnering with men suspected of gender-based violence appear larger; however, they are mostly imprecisely estimated for most ages of separation.

Within-Abusive Families Robustness One potential concern with the preceding analysis is that families with a violent father may differ in important ways from families without one. Although our matching strategy ensures that these families are observationally equivalent, unobservable differences may remain. For example, families where fathers are reported for GBV may have different parenting practices, neighborhood environments, or other unobservable parent characteristics that could account for reduction in GBV among sons where the father is removed from the household.

To address this concern, Figure 9 replicates Figures 7 and 8 by including family fixed effects. The inclusion of family fixed effects enables us to compare outcomes within abusive families, restricting identification to differences between siblings who were separated from an abusive father at different ages. The estimated coefficients are statistically significant and negative for sons whose abusive fathers left before the age of 11, dissipating for sons who were separated from abusive fathers at later ages. The pooled estimates in Appendix Table 2 show a large and statistically significant decline of 3.2 p.p. for sons who are separated from abusive fathers between ages 1-11. While we also observe mostly negative effects for daughters, they are much smaller and statistically insignificant for most ages. These findings suggest that the intergenerational transmission effects might be particularly stronger for boys when exposed to abusive fathers at younger ages.

Figure 9: Effects of Father Removal: Family Fixed Effects Estimates



Notes: The figure shows the estimates for the effects of father removal on sons' GBV perpetration in Panel (a) and daughters' GBV victimization proxied by their cohabitation with a partner suspected of committing GBV by their age at separation in Panel (b). Both plots show the estimated effects of father removal within abusive families, comparing siblings whose abusive father was removed at a younger age vs older age. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

5 Quasi-Random Removal of Fathers

The previous section suggests that separating from an abusive partner can materially reduce the incidence of GBV for sons when they grow up. However, a common refrain in this literature is that ânothing worksâ to encourage women to exit abusive relationships, reflecting the substantial and overlapping barriers that make separation extraordinarily difficult. A large body of research shows that women often face substantial economic dependence on their partners, limited access to affordable housing, and fears of financial instability following separation, all of which reduce the feasibility of leaving (Farmer and Tiefenthaler, 1997; Baker *et al.*, 2010; Adams *et al.*, 2024). Psychological factors, including trauma bonding, diminished self-efficacy, and manipulation by the abuser, can further impede exit (Dutton and Painter, 1993; Lahav, 2023). Stigma towards separation, cultural norms on gender relations, or legal barriers against divorce can also deter women from leaving abusive relationships (Stevenson and Wolfers, 2006).

This literature leads to two conculsions. First, it suggests that it may be difficult for policy levers to accomplish separations from abusive partners. Second, it suggests that women who are able to unilaterally initiate a divorce or separation may be unobservably different from women who are not. In particular, they may have more confidence, greater access to resources, or more

supportive social networks. These unobserved factors could also result in more positive outcomes for their children. If this is true, then the reductions in GBV from father removal identified in the previous section could fail to generalize to a broader sample of women.

To address both concerns, we implement a judge instrumental variables design that exploits the quasi-random assignment of criminal cases to judges who differ in their propensity to remove defendants from their households. This design isolates the causal effects of father removal by comparing women who are observably and unobservably similar, but who–solely due to judicial assignment–experience separation in one case and not in the other. The approach also directly addresses concerns about whether policy can meaningfully induce separation. In Sweden, prison sentences are typically short. This implies that any long-run effects on children's subsequent perpetration or victimization of GBV is likely only possible if incarceration facilitates a more durable separation than the sentence itself would mechanically generate. We test this possibility directly below, after reporting the impacts on GBV perpetration and victimization.

We closely follow the approach in Grönqvist *et al.* (2024), but restrict attention to fathers suspected of perpetrating gender-based violence and exploit quasi-random variation in the leniency of the judge assigned to their cases. Fathers whose cases are heard by less lenient judges face a higher likelihood of incarceration or court-ordered removal, whereas fathers assigned to more lenient judges are less likely to be removed. Comparing these otherwise similar fathers provides quasi-experimental variation in exposure to abusive fathers during childhood. The design identifies the local average treatment effect (LATE) of father removal for children whose fathers' household separation is sensitive to which judge hears the case, i.e., fathers at the margin of being removed.

We focus on children born between 1970 and 1995 whose fathers were suspected of committing gender-based violence. Let i index children, f their father, and c the criminal case involving the father, and consider a model that relates outcomes such as son's likelihood of being suspected of gender-based violence in adulthood to an indicator for whether the father was incarcerated

during childhood, $Prison_{fct}$:

$$Y_{ifct} = \beta_0 + \beta_1 Prison_{fct} + \beta_2 \mathbf{X}_{ifct} + \varepsilon_{ifct}, \tag{4}$$

where Y_{ifct} is a child's adulthood GBV outcome—perpetration for sons or partnering with a violent man for daughters; $Prison_{fct}$ indicates whether the father received a prison sentence (and was therefore removed from the household) following case c; and $\mathbf{X}_{ifct}$ is a vector of controls including the child's sex, birth order, migration status, and age at trial; the number of children in the family; the father's migration status, age, education, employment in the year before trial, and pretrial earnings; and the household's disposable income in the year before trial.

The main challenge for causal inference is that OLS estimates of Equation 4 are likely to be upward-biased due to unobserved characteristics transmitted across generations. For instance, gender-based violence could be correlated across generations due to unobservable factors common to fathers and children, including attaining lower levels of education or living in a high-crime neighborhood.[8]

Instrumental Variable We construct our instrument using the as-good-as-random assignment of cases to judges within district courts following Grönqvist *et al.* (2024). During our study period, Swedish courts used a computerized system that allocated cases randomly to judges, conditional on a small set of administrative strata. Judges differ substantially in their propensity to impose prison sentences, even after accounting for case composition. We construct a leave-out judge-year incarceration rate as the average probability that the judge sentenced *other* defendants to prison in the same year, excluding all fathers in our estimation sample. This ensures that the instrument is not mechanically correlated with the child outcomes we study and avoids precision inflation that can arise when instruments are constructed using estimation-sample outcomes.

More specifically, we construct the instrument as a residualized, leave-out measure of judge

[8]The direction of bias is, however, ambiguous since judges may view certain characteristics such as substance abuse or mental health issues as mitigating and thus opt for leniency, even though these characteristics are correlated with worse future outcomes (Agan *et al.*, 2023; Grönqvist *et al.*, 2024); conversely, judges may impose harsher sanctions on fathers perceived as particularly dangerous in gender-based violence cases. As a result, OLS comparisons cannot credibly disentangle the causal impact of paternal incarceration from these underlying differences.

stringency that aligns with the level at which random assignment occurs, i.e., court-by-year-by-randomization strata. Because cases were randomly assigned within age and crime-type groups, we partial out court-by-year-by-strata fixed effects when constructing the instrument (Grönqvist *et al.*, 2024; Agan *et al.*, 2023; Dobbie *et al.*, 2018). The randomization strata consist of five categories: juveniles under 18, juveniles aged 19-21, least serious crimes, most serious crimes, and other crimes.[9] This structure ensures that comparisons are restricted to defendants who could have been assigned to the same judges. As a result, the within-cell variation in our instrument captures differences in a quasi-randomly assigned judge's likelihood of imposing a prison sentence relative to otherwise similar cases heard in the same court and year.

First Stage Our first-stage regression examines the relationship between judge stringency Z_{fctj} and father incarceration $Prison_{fctj}$. We estimate the following specification for individual i with father f who is charged in case c and assigned to judge j using OLS:

$$Prison_{fctj} = \pi_0 + \pi_1 Z_{fctj} + \pi_2 \mathbf{X}_{ifct} + u_{ifctj}, \tag{5}$$

where Z_{fctj} is our leave-out judge incarceration stringency measure, and the vector $\mathbf{X}_{ifct}$ includes court-by-year-by-strata fixed effects and, in some specifications, control variables. The control variables include the child's sex, birth order, migration status, and age at trial; the number of children in the family; the father's migration status, age, education, employment in the year before trial, and pretrial earnings; and the household's disposable income in the year before trial. Robust standard errors are two-way clustered at the father and court section level.

[9] These crime categories were developed using information from interviews with court staff and a review of court records. Offenses classified as least serious are those carrying a maximum sentence of six months, whereas the most serious category includes crimes such as murder, kidnapping, arson, robbery, rape, aggravated assault, and domestic violence. Juveniles charged with these severe offenses are also placed in the serious-crime group. For roughly 5 percent of cases where the Convictions Register lacks offense details, we rely on the Suspicion Register to assign the appropriate crime category.

Table 4: First Stage Results

	Sample Mean	Judge Stringency	
	(1)	(2)	(3)
Father Incarceration	0.153	0.371***	0.375***
	(0.360)	(0.092)	(0.091)
Sanderson-Windmeijer F-stat			16.71
Cragg-Donald F-stat			35.80
Court x Year x Strata FE	–	Yes	Yes
Controls	–	No	Yes
Observations		18,561	18,561

Note: This table reports first stage results. Judge stringency is estimated using data from other cases assigned to a section in the same year. Column 1 reports the mean and standard deviation of the dependent variable (father incarceration). Column 2 reports estimates from an OLS regression of father incarceration on judge stringency controlling for only court-by-year-by-strata fixed effects. Column 3 adds controls, including the child's sex, birth order, migration status, and age at trial; the number of children in the family; the father's migration status, age, education, employment in the year before trial, and pretrial earnings; and the household's disposable income in the year before trial. All regressions are run at the child level. Standard errors are two-way clustered at the father and section levels. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table 4 presents the first-stage results from estimating Equation 5. Column 1 reports the average incarceration rate among fathers suspected of gender-based violence in our estimation sample. Column 2 presents first-stage results controlling only for court-by-year-by-strata fixed effects. Column 3 adds the baseline characteristics reported in the table notes and indicators for any missing baseline characteristics. The estimate including all controls (column 3) shows that having a father assigned to a judge who is 10 percentage points more likely to incarcerate increases the probability of incarceration by 3.75 percentage points. The Sanderson-Windmeijer and Cragg Donald F statistics are both high, suggesting that the instrument is a strong predictor.

Independence and Exclusion An additional condition required to interpret our two-stage least squares estimates as the LATE of incarcerating abusive fathers is that judge assignment affects child outcomes only through its impact on the probability that the father is incarcerated. Table 5 provides evidence consistent with this exclusion restriction. In Column 1, we estimate a linear probability model examining whether baseline characteristics of the child, their father, or

the case predict incarceration of the father. Column 2 tests whether these same characteristics predict our residualized judge stringency measure using the same specification. Both models include court-by-year-by-strata fixed effects and two-way clustered standard errors at the family and court-section level. As expected, observable characteristics strongly predict incarceration, but they are not systematically related to judge stringency. These results indicate that judges with different incarceration tendencies are assigned very similar cases, consistent with the randomized assignment process described earlier.

The exclusion restriction may still be violated if judge assignment affects child outcomes through channels other than the father's incarceration. This assumption is fundamentally untestable, and our estimates should be interpreted with this caveat in mind. Nonetheless, we argue that the exclusion restriction is plausible in our context. The Swedish criminal justice system places strict limits on judicial discretion: judges must choose from a narrow set of punishment types, and sentence lengths are guided by mandatory national sentencing guidelines issued by the Supreme Court and the Prosecutor Authority. These institutional features reduce the scope for judges to systematically affect defendants in ways unrelated to the incarceration margin.

Table 5: Test of Randomization

	Father Incarcerated	Judge Stringency
	(1)	(2)
Child is male	0.00434	0.00063
	(0.00632)	(0.00067)
Child is foreign born	-0.04200***	-0.00039
	(0.01519)	(0.00162)
Child age at trial	0.00051	-0.00013
	(0.00129)	(0.00013)
Child is a second child	0.00810	0.00057
	(0.00575)	(0.00057)
Child is a third child	0.01062	0.00024
	(0.00963)	(0.00099)
Child is a forth child or higher child	0.00288	0.00100
	(0.01453)	(0.00159)
Number of children in family	0.00061	0.00004
	(0.00492)	(0.00049)
Father is foreign born	0.02586**	0.00017
	(0.01171)	(0.00118)
Father age at trial	-0.00200**	0.00006
	(0.00079)	(0.00009)
Father has high school degree	-0.00715	0.00007
	(0.01213)	(0.00116)
Father has at least some college	-0.03913**	-0.00250
	(0.01583)	(0.00175)
Father employment in 1 year before trial	-0.07567***	-0.00048
	(0.01504)	(0.00157)
Family disposable income in 1 year before trial	-0.00000*	-0.00000
	(0.00000)	(0.00000)
Father earnings in 1 year before trial	0.00000	0.00000**
	(0.00000)	(0.00000)
Average processing time	0.00009	-0.00000
	(0.00012)	(0.00004)
Joint F-Test	[0.00000]	[0.61574]
Observations	18,561	18,561

Note: This table reports reduced-form tests of the random assignment of cases to court sections. Column 1 presents OLS estimates from regressing father incarceration on the listed covariates, controlling for court-by-year-by-strata fixed effects. Column 2 presents analogous OLS estimates where the dependent variable is judge stringency. Both regressions also include indicators for missing values. The p-values at the bottom of Columns 1-2 correspond to F-tests of the joint significance of all listed covariates and missing-value indicators. Standard errors are two-way clustered at the parent and section levels. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Furthermore, other courtroom actors such as prosecutors and lay jurors are assigned independently of the judge. This institutional separation makes it unlikely that judge assignment is correlated with the assignment of other actors who might independently affect defendants or their families. While we cannot completely rule out all alternative channels, such as small differences in how quickly judges process case paperwork, we view them as unlikely to meaningfully bias our two stage least squares estimates.

Table 6: Father Incarceration and Child GBV Outcomes in Adulthood

	Ages 1-18	Ages 1-11	Ages 12-18
Panel A: Sons	(1)	(2)	(3)
Suspected of any violence against women	-0.047	-0.632*	0.309
	(0.225)	(0.368)	(0.321)
Outcome mean	[0.250]	[0.214]	[0.300]
Complier control mean	{0.287}	{0.855}	{0.028}
Observations	9,919	5,765	4,154
Panel B: Daughters			
Cohabiting with a partner suspected of GBV	-0.063	-0.181	-0.174
	(0.260)	(0.625)	(0.418)
Outcome mean	[0.220]	[0.195]	[0.255]
Complier control mean	{0.191}	{0.368}	{0.149}
Observations	8,642	5,084	3,558

Note: This table reports two-stage least squares estimates of the impact of father incarceration on children's outcomes. Columns 1-3 reports two-stage least squares estimates that instrument for paternal incarceration using judge stringency. All regressions control for court-by-year-by-strata fixed effects and controls, including the child's sex, birth order, migration status, and age at trial; the number of children in the family; the father's migration status, age, education, employment in the year before trial, and pretrial earnings; and the household's disposable income in the year before trial. Dependent variable means are in square brackets. The complier control means (CCM) are in curly brackets. Standard errors are two-way clustered at the parent and section levels. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Results from Judge IV Table 6 presents two-stage least squares estimates of the impact of incarceration of fathers suspected of gender-based violence on our outcome variables: an indicator of the son being suspected of any violence against women (Panel A), and an indicator of the daughter of cohabiting with a partner suspected of gender-based violence (Panel B). Across all estimates, judge stringency is used as an instrument for incarceration of fathers suspected of

gender-based violence. The dependent variable means are reported in square parentheses under the coefficient estimates. Column 1 reports estimates pooling separations due to father incarceration occurring between ages 1 and 18. Column 2 restricts the sample to separations occurring between ages 1 and 11, while Column 3 focuses on separations occurring between ages 12 and 18. We focus on these age breaks given that the previous section suggests removal at formative early ages may be instrumental in reducing later GBV of sons.

The pooled two-stage least squares estimates reported in Table 6 are negative but imprecise for both sons and daughters (column 1). However, the column 2 estimate for sons indicates that being separated from abusive fathers between ages 1-11 due to father's incarceration reduces the probability of the son being suspected of any violence against women by 63.3 percentage points (74% compared to the control complier mean). Turning to the impact for sons at later ages, we find an insignificant but positive estimate. One plausible interpretation is that sons are more susceptible to imitation during early, formative years (ages 1-11), and removing the father in these early years allows sons to avoid learning GBV from their fathers. By adolescence, however, these behaviors may already be internalized, such that separation at later ages is less effective at disrupting the intergenerational transmission of gender-based violence, but does lead to lower income and higher violence. For daughters, we find no evidence of a significant impact of separating from abusive fathers for any age group.

These results are broadly consistent with the findings in Section 4, indicating that reducing exposure to fathers suspected of gender-based violence is particularly effective in weakening the intergenerational transmission of such violence to sons in early childhood. In contrast, we find little evidence that reducing daughters' exposure to abusive fathers lowers their subsequent risk of partnering with abusive men. Together, these findings suggest that early-life exposure to abusive fathers operates through distinct gender-specific channels, with direct behavioral transmission playing a larger role for sons than for daughters.

6 Conclusion

This paper provides the first large-scale administrative evidence that gender-based violence is transmitted across generations through childhood exposure. Using linked population-wide data from Sweden (and complementary evidence from the United States, which is in progress), we show that sons of fathers suspected of gender-based violence are more than twice as likely to commit gender-based violence themselves, and daughters exposed to violent fathers are almost twice as likely to partner with violent men. These patterns persist after controlling for rich demographic, socioeconomic, and neighborhood characteristics and when comparisons are restricted to closely matched families. The magnitude and robustness of these associations underscore a strong social learning mechanism through which violent behaviors are imitated and victimization risks are shaped early in life.

We then examine whether this intergenerational cycle can be disrupted by reducing children's exposure to abusive fathers. Although father removal is accompanied by large and persistent income losses that raise risks for children, for sons, abusive father removal significantly reduces the likelihood of perpetrating gender-based violence in adulthood, with the strongest effects arising when separation occurs in early childhood. Family fixed-effects estimates confirm this pattern within abusive families, showing that siblings separated from violent fathers at younger ages experience substantially lower risks later in life. For daughters, estimates are generally negative but smaller and less precisely estimated, suggesting that direct behavioral transmission effects may be more substantial between fathers and sons.

Finally, leveraging quasi-random assignment of cases to judges with differing incarceration propensities, we provide causal evidence that quasi-random removal of abusive fathers weakens the intergenerational transmission of gender-based violence. The judge-IV results show that incarceration of abusive fathers during children's formative years significantly reduces sons' subsequent perpetration risk. Taken together, our findings imply that gender-based violence is not an immutable trait but a learned behavior shaped by childhood environments. Policies that safely reduce children's exposure to abusive fathers, while mitigating the associated economic harms,

have the potential to meaningfully weaken the persistence of gender-based violence across generations.

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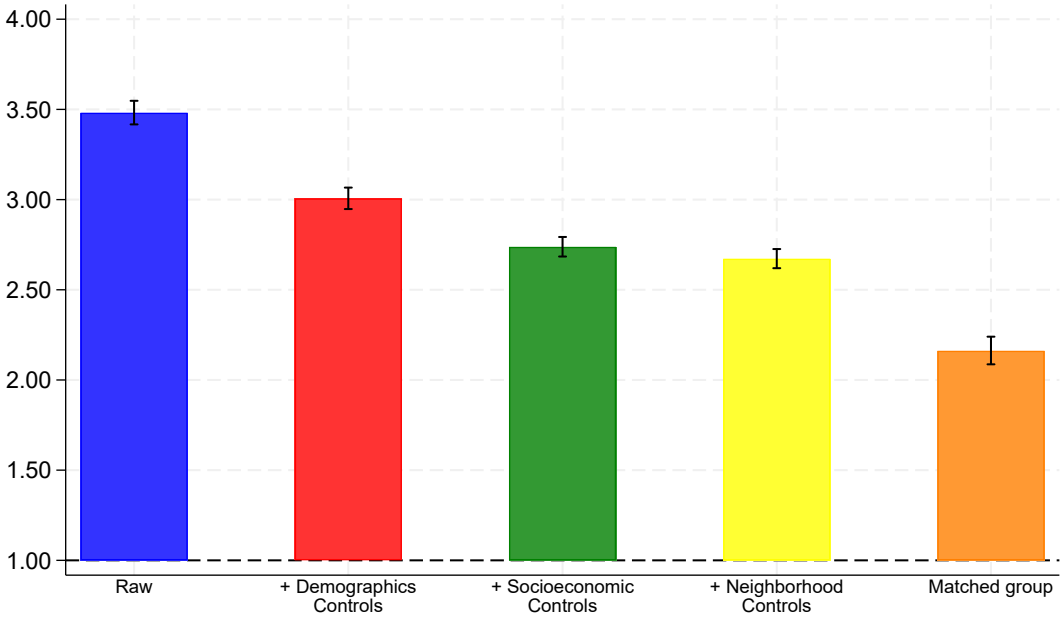
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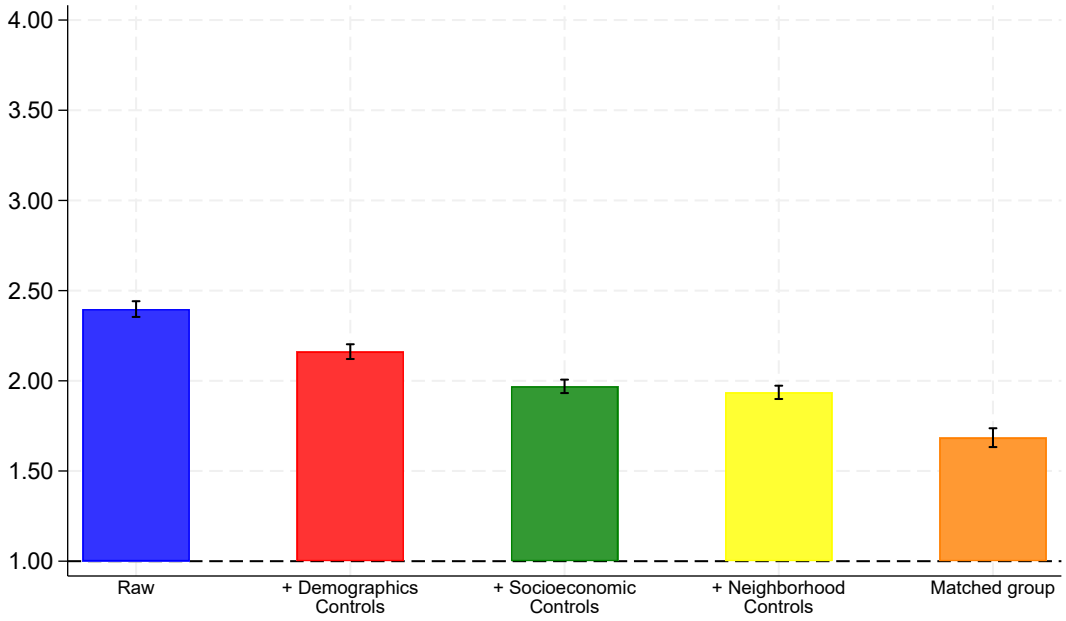
Appendix Figures and Tables

Figure A1: Odds Ratios for Sons' GBV Perpetration



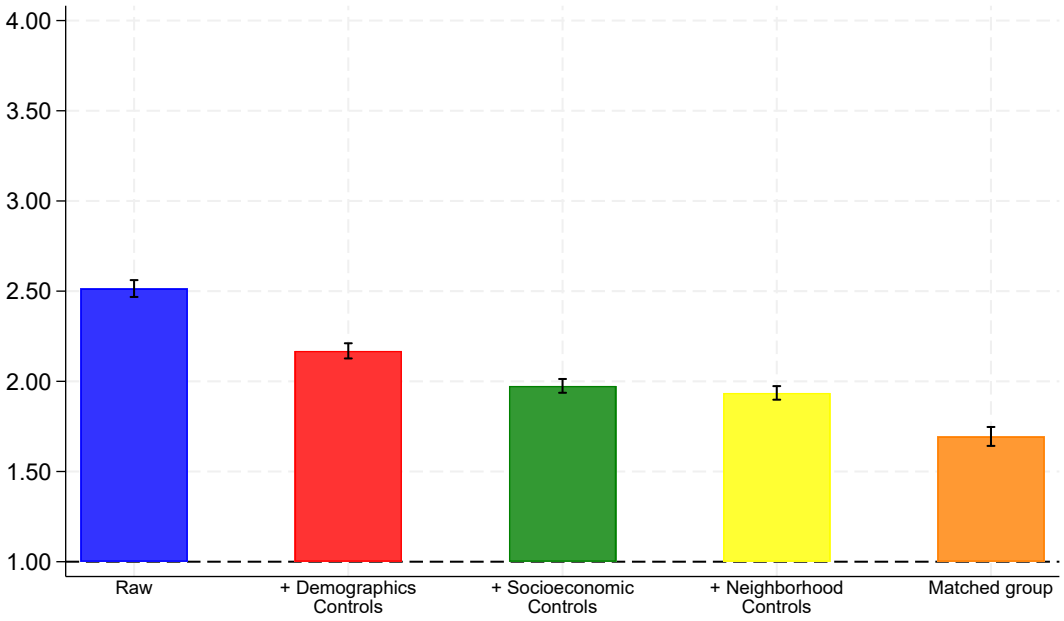
Notes: This figure illustrates the intergenerational transmission of gender-based violence perpetration in Sweden. The blue bar reports the odds ratio for an indicator for the son being suspected of GBV based on an indicator for the father being suspected of GBV, without controls. The red bar adds demographic controls, including birth-year fixed effects for father and son, indicators for whether the mother, father, or son is foreign born, and birth-order indicators from both parents. The green bar adds socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. The yellow bar adds neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school), and reports the corresponding coefficient estimate. All bars display 95% confidence intervals.

Figure A2: Odds Ratios for All Daughters' GBV Victimization



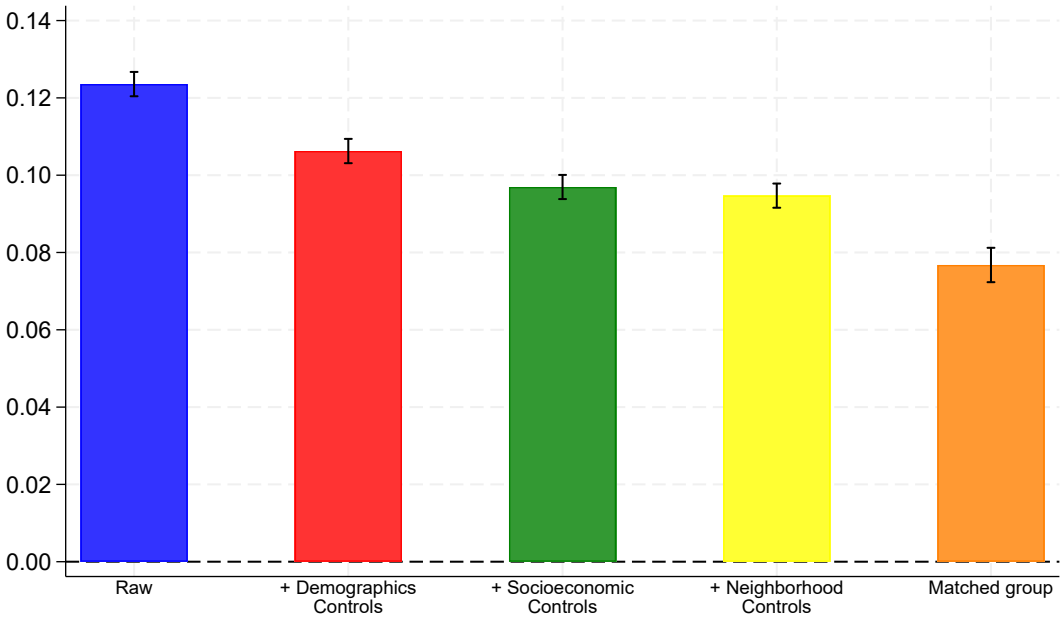
Notes: This figure illustrates the intergenerational transmission of gender-based violence victimization in Sweden. The sample includes all women born from 1970 to 1995 in Sweden. The blue bar reports the odds ratio for an indicator for the daughter cohabiting with a partner suspected of GBV based on an indicator for the father being suspected of GBV, without controls. The red bar adds demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. The green bar adds socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. The yellow bar adds neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school), and reports the corresponding coefficient estimate. All bars display 95% confidence intervals.

Figure A3: Odds Ratios for Daughters' GBV Victimization Conditional on Having Ever a Partner



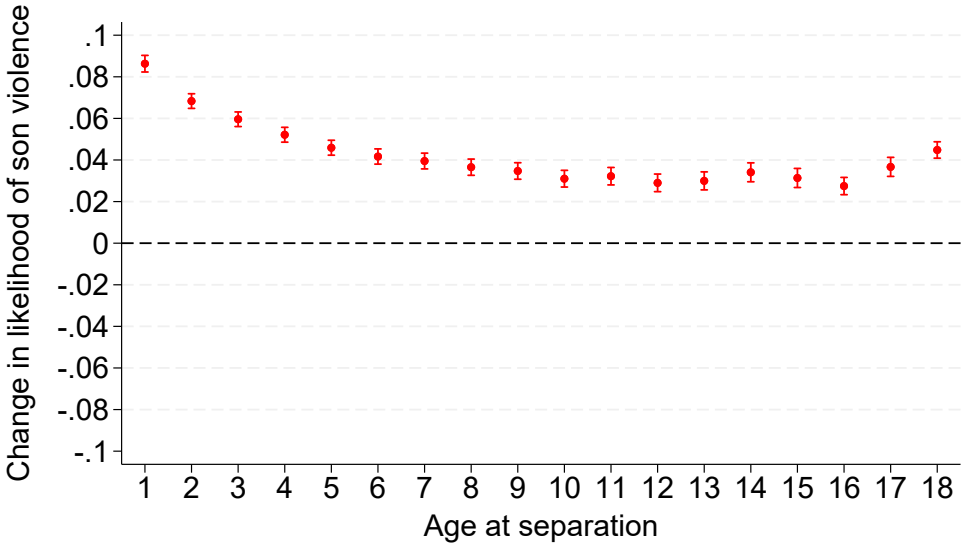
Notes: This figure illustrates the intergenerational transmission of gender-based violence victimization in Sweden. The sample includes women born from 1970 to 1995 in Sweden who ever cohabited with a partner. The blue bar reports the odds ratio for an indicator for the daughter cohabiting with a partner suspected of GBV based on an indicator for the father being suspected of GBV, without controls. The red bar adds demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. The green bar adds socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. The yellow bar adds neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school), and reports the corresponding coefficient estimate. All bars display 95% confidence intervals.

Figure A4: Intergenerational Transmission of GBV Victimization: Linear Probability Estimates



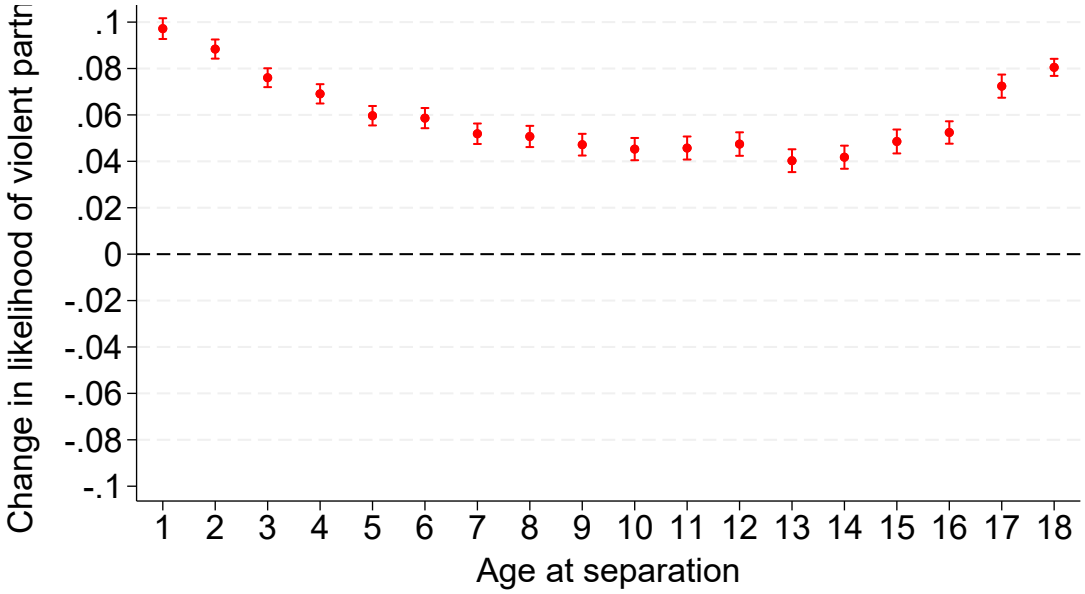
Notes: This figure illustrates the intergenerational transmission of victimization based on gender-based violence in Sweden. The sample is restricted to women who have ever cohabited with a partner. The blue bar reports the coefficient from a linear probability model regressing an indicator for the daughter cohabiting with a partner who was suspected of GBV on an indicator for the father being suspected of GBV, without controls. The red bar adds demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. The green bar adds socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. The yellow bar adds neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. The orange bar uses a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school), and reports the corresponding coefficient estimate. Standard errors are clustered at the father level. All bars display 95% confidence intervals.

Figure A5: Effects of Father Removal on Sons' GBV Perpetration: Sample of Non-Violent Fathers



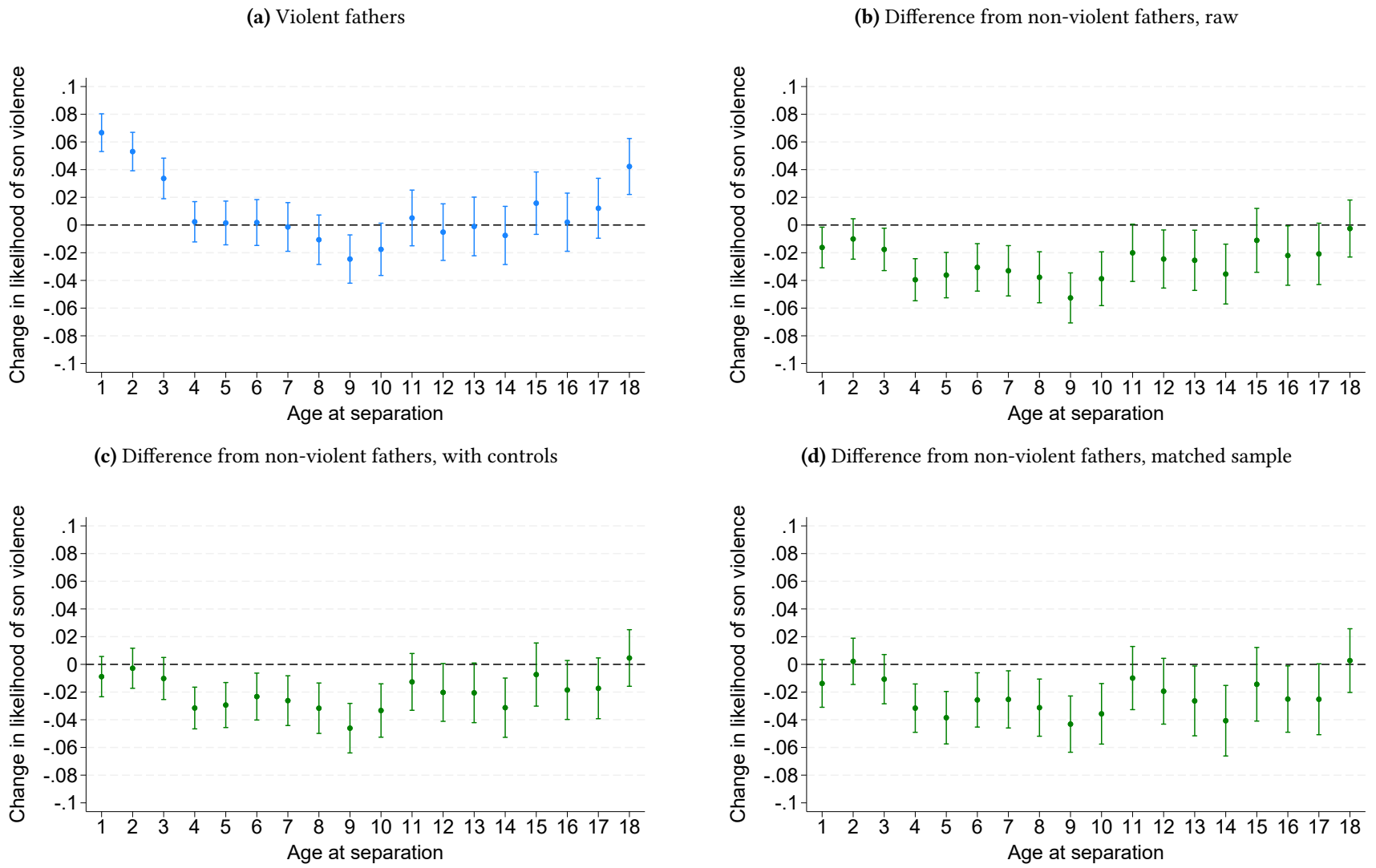
Notes: This figure shows the estimates for the effects of removing the father from the household for each age between 1 and 19 for the sample of non-violent fathers based on Equation 1. The outcome is sons' likelihood of suspected violence against women in adulthood. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

Figure A6: Effects of Father Removal on Daughters' GBV Victimization: Sample of Non-Violent Fathers



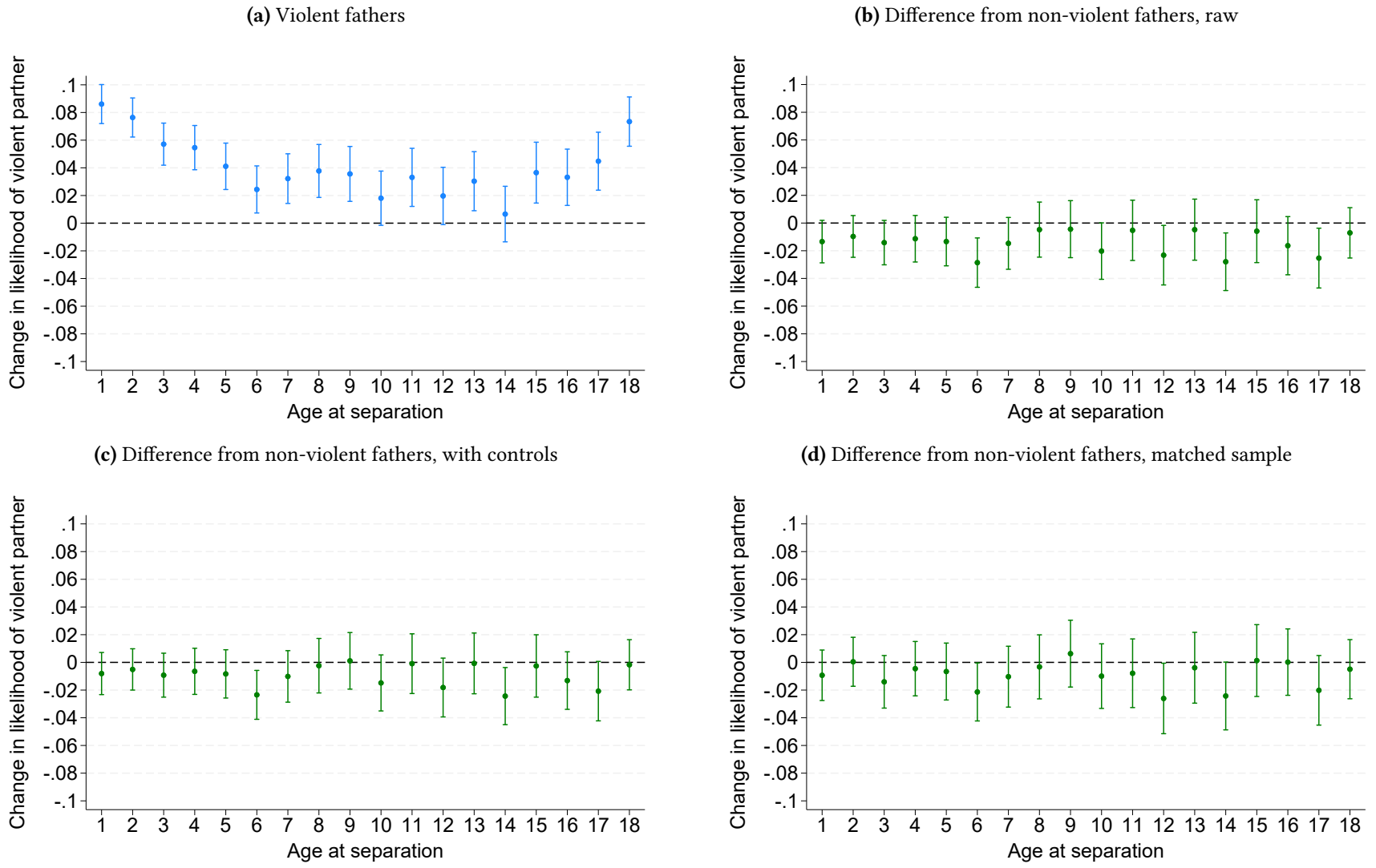
Notes: This figure shows the estimates for the effects of removing the father from the household for each age between 1 and 19 for the sample of non-violent fathers based on Equation 1. The outcome is daughters' likelihood of cohabiting with a partner who was suspected of violence against women. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

Figure A7: Effects of Father Removal on Sons' GBV Perpetration by Age at Separation: Excluding Families that Reunite



Notes: The figure shows the estimates for the effects of father removal on sons' GBV perpetration by their age at separation. The sample excludes families that reunite, i.e., where the father returns to the household in subsequent years. Panel A plots the estimates for the sample of sons with violent fathers by estimating Equation 1. Panel B plots the estimated differences between sons with violent fathers and those without violent fathers using Equation 2, but it does not include controls. Panel C adds controls for demographic, socioeconomic and neighborhood characteristics as specified in Equation 2. Panel D uses a matched sample to estimate Equation 2 where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

Figure A8: Effects of Father Removal on Daughters' GBV Victimization by Age at Separation: Excluding Families that Reunite



Notes: The figure shows the estimates for the effects of father removal on daughters' GBV victimization proxied by their cohabitation with a partner suspected of committing GBV by their age at separation. The sample excludes families that reunite, i.e., where the father returns to the household in subsequent years. Panel A plots the estimates for the sample of daughters with violent fathers by estimating a corollary specification to Equation 1. Panel B plots the estimated differences between daughters with violent fathers and those without violent fathers using Equation 2, but it does not include controls. Panel C adds controls for demographic, socioeconomic and neighborhood characteristics as specified in Equation 2. Panel D uses a matched sample to estimate a corollary specification to Equation 2 where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level. The plot includes coefficient estimates and 95% confidence intervals.

Table A1: Distribution of Crime Categories among Fathers and Sons

Crime category	% Suspected sons (1)	% Suspected fathers (2)	% Susp. sons |Any suspicion (3)	% Suspected fathers |Any suspicion (4)
Gross violation of woman's integrity	0.007	0.007	0.090	0.104
Assault of woman	0.046	0.044	0.637	0.659
(Attempted) murder of woman	0.001	0.001	0.012	0.011
Unlawful threat of woman	0.025	0.026	0.348	0.389
Molestation/harassment of woman	0.021	0.017	0.285	0.259
Molestation/harassment, unspecified gender	0.002	0.005	0.034	0.070
Violation of restraining order against woman	0.000	0.000	0.005	0.001
Violation of restraining order, unspecified gender	0.003	0.004	0.035	0.060
(Attempted) rape/molestation of woman	0.010	0.004	0.134	0.058
(Attempted) rape/molestation, unspecified gender	0.006	0.006	0.078	0.091
Stalking	0.000	0.000	0.002	0.001

Note: This table reports the distribution of gender-based violence across crime categories for fathers and sons in our sample. Sons were born between 1970 and 1995 in Sweden and matched with their fathers using registry information. Column 1 reports the proportion of suspected sons in each crime category, column 2 reports the proportion of suspected fathers in each crime category, column 3 reports the proportion of suspected sons conditional on any suspicion, and column 4 reports the proportion of suspected fathers conditional on any suspicion. All data come from Crime Suspicion Register of Sweden.

Table A2: Probability of Ever Cohabiting with a Partner

	Outcome mean	(1)	(2)	(3)	(4)	(5)
Violent father	0.848	-0.014***	0.015***	0.012***	0.015***	0.018***
	(0.359)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
% relative to outcome mean		[-2]	[2]	[1]	[2]	[2]
Demographic controls		–	Yes	Yes	Yes	–
Socioeconomic controls		–	–	Yes	Yes	–
Neighborhood controls		–	–	–	Yes	–
Matched sample		–	–	–	–	Yes
Obs.		1,374,446	1,374,446	1,374,446	1,374,446	306,404

Note: This table reports the probability of ever cohabiting with a partner in relation to having a father suspected of gender-based violence in Sweden. The outcome is an indicator for the daughter cohabiting with a partner. Column 1 reports the OLS coefficient from a linear probability model regressing an indicator for the daughter cohabiting with a partner on an indicator for the father being suspected of GBV. Column 2 reports OLS coefficient adding demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. Column 3 reports OLS coefficient adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports OLS coefficient adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports OLS estimate based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A3: Linear probability model estimates for sons

	Outcome mean	(1)	(2)	(3)	(4)	(5)
Violent father	0.073	0.129***	0.115***	0.109***	0.106***	0.085***
	(0.260)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
% relative to outcome mean		[177]	[158]	[149]	[146]	[116]
Demographic controls		–	Yes	Yes	Yes	–
Socioeconomic controls		–	–	Yes	Yes	–
Neighborhood controls		–	–	–	Yes	–
Matched sample		–	–	–	–	Yes
Obs.		1,446,816	1,446,816	1,446,816	1,446,816	317,672

Note: This table reports the intergenerational transmission of gender-based violence in Sweden. The outcome is an indicator for the son being suspected of GBV. Column 1 reports the OLS coefficient from a linear probability model regressing an indicator for the son being suspected of GBV on an indicator for the father being suspected of GBV, without controls. Column 2 reports OLS coefficient adding demographic controls, including birth-year fixed effects for father and son, indicators for whether the mother, father, or son is foreign born, and birth-order indicators from both parents. Column 3 reports OLS coefficient adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports OLS coefficient adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports OLS estimate based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A4: Linear probability model estimates for all daughters

	Outcome mean	(1)	(2)	(3)	(4)	(5)
Violent father	0.097	0.102***	0.091***	0.083***	0.082***	0.066***
	(0.296)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
% relative to outcome mean		[105]	[94]	[86]	[84]	[68]
Demographic controls		–	Yes	Yes	Yes	–
Socioeconomic controls		–	–	Yes	Yes	–
Neighborhood controls		–	–	–	Yes	–
Matched sample		–	–	–	–	Yes
Obs.		1,374,446	1,374,446	1,374,446	1,374,446	306,404

Note: This table reports the intergenerational transmission of victimization based on gender-based violence in Sweden. The outcome is an indicator for the daughter cohabiting with a partner who was suspected of GBV. Column 1 reports the OLS coefficient from a linear probability model regressing an indicator for the daughter cohabiting with a partner who was suspected of GBV on an indicator for the father being suspected of GBV. Column 2 reports OLS coefficient adding demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. Column 3 reports OLS coefficient adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports OLS coefficient adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports OLS estimate based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A5: Linear probability model estimates for daughters conditional on having a partner

	Outcome mean	(1)	(2)	(3)	(4)	(5)
Violent father	0.114	0.124***	0.106***	0.097***	0.095***	0.077***
	(0.318)	(0.002)	(0.002)	(0.002)	(0.002)	
% relative to outcome mean		[108]	[93]	[85]	[83]	[67]
Demographic controls		–	Yes	Yes	Yes	–
Socioeconomic controls		–	–	Yes	Yes	–
Neighborhood controls		–	–	–	Yes	–
Matched sample		–	–	–	–	Yes
Obs.		1,165,947	1,165,947	1,165,947	1,165,947	259,224

Note: This table reports the intergenerational transmission of victimization based on gender-based violence in Sweden. The outcome is an indicator for the daughter cohabiting with a partner who was suspected of GBV. Column 1 reports the OLS coefficient from a linear probability model regressing an indicator for the daughter cohabiting with a partner who was suspected of GBV on an indicator for the father being suspected of GBV. Column 2 reports OLS coefficient adding demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. Column 3 reports OLS coefficient adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports OLS coefficient adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports OLS estimate based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A6: Odds ratio estimates for sons

	(1)	(2)	(3)	(4)	(5)
Violent father	3.482***	3.006***	2.738***	2.672***	2.162***
	(0.010)	(0.010)	(0.010)	(0.010)	(0.018)
Demographic controls	–	Yes	Yes	Yes	–
Socioeconomic controls	–	–	Yes	Yes	–
Neighborhood controls	–	–	–	Yes	–
Matched sample	–	–	–	–	Yes
Obs.	1,446,816	1,446,816	1,446,816	1,446,816	317,672

Note: This table reports the intergenerational transmission of gender-based violence in Sweden. Column 1 reports the odds ratio obtained by regressing an indicator for the son being suspected of GBV on an indicator for the father being suspected of GBV, with no additional controls included. Column 2 reports the odds ratio adding demographic controls, including birth-year fixed effects for father and son, indicators for whether the mother, father, or son is foreign born, and birth-order indicators from both parents. Column 3 reports the odds ratio adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports the odds ratio adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports the odds ratio based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A7: Odds ratio estimates for all daughters

	(1)	(2)	(3)	(4)	(5)
Violent father	2.397***	2.161***	1.969***	1.936***	1.684***
	(0.009)	(0.010)	(0.010)	(0.010)	(0.016)
Demographic controls	–	Yes	Yes	Yes	–
Socioeconomic controls	–	–	Yes	Yes	–
Neighborhood controls	–	–	–	Yes	–
Matched sample	–	–	–	–	Yes
Obs.	1,374,446	1,374,446	1,374,446	1,374,446	306,404

Note: This table reports the intergenerational transmission of victimization for gender-based violence in Sweden. Column 1 reports the odds ratio obtained by regressing an indicator for the daughter cohabiting with a partner being suspected of GBV on an indicator for the father being suspected of GBV, with no additional controls included. Column 2 reports the odds ratio adding demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. Column 3 reports the odds ratio adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports the odds ratio adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports the odds ratio based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A8: Odds ratio estimates for daughters who ever had a partner

	(1)	(2)	(3)	(4)	(5)
Violent father	2.514***	2.169***	1.974***	1.935***	1.694***
	(0.009)	(0.010)	(0.010)	(0.010)	(0.016)
Demographic controls	–	Yes	Yes	Yes	–
Socioeconomic controls	–	–	Yes	Yes	–
Neighborhood controls	–	–	–	Yes	–
Matched sample	–	–	–	–	Yes
Obs.	1,165,947	1,165,947	1,165,947	1,165,947	259,224

Note: This table reports the intergenerational transmission of victimization for gender-based violence in Sweden. The sample is restricted to women who are born between 1970 and 1995 and who ever cohabited with a partner. Column 1 reports the odds ratio obtained by regressing an indicator for the daughter cohabiting with a partner being suspected of GBV on an indicator for the father being suspected of GBV, with no additional controls included. Column 2 reports the odds ratio adding demographic controls, including birth-year fixed effects for father and daughter, indicators for whether the mother, father, or daughter is foreign born, and birth-order indicators from both parents. Column 3 reports the odds ratio adding socioeconomic controls: parental educational attainment (high school and university) at birth, employment status at birth, and adjusted parental earnings at birth. Column 4 reports the odds ratio adding neighborhood controls, including parish population, immigrant share, employment rate, average earnings, share of university-educated residents, number of offenders per 100K, and assault and homicide rates. Column 5 reports the odds ratio based on a matched control group based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (less than high school, high school, more than high school). Standard errors are clustered at the father level. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A9: Effects of Father Removal on Sons' GBV Perpetration by Alternative Age Groups

	Raw difference (1)	Controls added (2)	Matched sample (3)	Family fixed effects (4)
Separation age 1-2 X Violent father	-0.00409	0.00248	-0.00675	-0.0353***
	(0.00468)	(0.00462)	(0.00544)	(0.0131)
Separation age 3-10 X Violent father	-0.0179***	-0.0121***	-0.0178***	-0.0237**
	(0.00365)	(0.00360)	(0.00414)	(0.0111)
Separation age 11-18 X Violent father	-0.0112**	-0.00619	-0.00994**	0.0198*
	(0.00439)	(0.00433)	(0.00493)	(0.0102)
Violent father	0.0956***	0.0820***	0.0834***	
	(0.00263)	(0.00259)	(0.00295)	
Separation age 1-2	0.0768***	0.0653***	0.0816***	0.0226***
	(0.00136)	(0.00136)	(0.00302)	(0.00313)
Separation age 3-10	0.0443***	0.0394***	0.0498***	0.0167***
	(0.000720)	(0.000721)	(0.00201)	(0.00219)
Separation age 11-18	0.0337***	0.0312***	0.0329***	0.0159***
	(0.000816)	(0.000811)	(0.00232)	(0.00203)
Constant	0.0409***	-0.104***	0.00507	0.0808***
	(0.000224)	(0.00786)	(0.00721)	(0.00289)
Observations	1,231,450	1,225,728	258,136	1,231,450
R-squared	0.020	0.034	0.025	0.004
Mean dep. var.	0.0606	0.0606	0.110	0.0606
Number of unique fathers				893,221

Notes: The table shows the estimates for the effects of father removal on sons' GBV perpetration by the age groups of 1-2, 3-10, and 11-18. Column (1) shows the estimated differences between sons with violent fathers and those without violent fathers by separation age groups, but it does not include controls. Column (2) adds controls for demographic, socioeconomic and neighborhood characteristics. Column (3) uses a matched sample where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). Column (4) adds family fixed effects. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level and reported in parentheses. ***, **, and * denote significance at the 1, 5, and 10 percent levels.

Table A10: Effects of Father Removal on Daughters' GBV Victimization by Alternative Age Groups

	Raw difference (1)	Controls added (2)	Matched sample (3)	Family fixed effects (4)
Separation age 1-2 X Violent father	0.00269	0.00624	-0.00132	-0.00525
	(0.00492)	(0.00487)	(0.00588)	(0.0148)
Separation age 3-10 X Violent father	-0.00434	-0.00143	-0.00187	-0.00783
	(0.00381)	(0.00377)	(0.00444)	(0.0125)
Separation age 11-18 X Violent father	-0.0112**	-0.00724*	-0.00480	-0.0169
	(0.00439)	(0.00435)	(0.00510)	(0.0108)
Violent father	0.0663***	0.0549***	0.0529***	
	(0.00271)	(0.00269)	(0.00317)	
Separation age 1-2	0.0925***	0.0802***	0.0969***	0.0302***
	(0.00156)	(0.00156)	(0.00349)	(0.00382)
Separation age 3-10	0.0591***	0.0537***	0.0608***	0.0267***
	(0.000848)	(0.000849)	(0.00236)	(0.00275)
Separation age 11-18	0.0579***	0.0524***	0.0497***	0.0338***
	(0.000909)	(0.000902)	(0.00266)	(0.00232)
Constant	0.0620***	-0.197***	-0.00130	0.104***
	(0.000286)	(0.00867)	(0.00868)	(0.00359)
Observations	1,170,215	1,164,773	249,664	1,170,215
R-squared	0.017	0.031	0.020	0.003
Mean dep. var.	0.0886	0.0886	0.140	0.0886
Number of unique fathers				863,673

Notes: The table shows the estimates for the effects of father removal on daughters' GBV victimization by the age groups of 1-2, 3-10, and 11-18. Column (1) shows the estimated differences between daughters with violent fathers and those without violent fathers by separation age groups, but it does not include controls. Column (2) adds controls for demographic, socioeconomic and neighborhood characteristics. Column (3) uses a matched sample where the control group is matched based on parental immigrant status, employment at birth, earnings quintiles, birth order, parental age (in 5-year bins), and parental education (indicators for less than high school, high school, more than high school). Column (4) adds family fixed effects. All estimated effects are relative to separation at age 19, which is the omitted reference group. Standard errors are clustered at the father level and reported in parentheses. ***, **, and * denote significance at the 1, 5, and 10 percent levels.