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# Fair Trade-Organic Coffee Cooperatives, Migration, and Secondary Schooling in Southern Mexico

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# Fair Trade-Organic Coffee Cooperatives, Migration, and Secondary Schooling in Southern Mexico

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ABSTRACT *We explore three trends in rural southern Mexico (Fair Trade coffee, migration, and conditional cash transfers) that could explain the rapid rise in education from 1995–2005 using survey data from 845 coffee farming households in Oaxaca and Chiapas, Mexico. Results from a household fixed-effects model show that household participation in a Fair Trade-organic cooperative contributed to about a 0.7 year increase in schooling for girls. US migration opportunities appear to have even stronger positive impacts on schooling for females. Although participation in Fair Trade-organic cooperatives appears also to have increased male schooling, increased migration opportunities have had an indeterminate effect for males.*



## I. Introduction

Recent research on secondary school outcomes in Latin America has explored three major changes that could affect the education investments of poor households: changing economic opportunities, migration, and conditional cash transfers. Southern Mexico has experienced all three of these trends. Changes in economic opportunities, ranging from coffee price movements to labour market trends to natural disasters, are shaping the child labour versus schooling choices (Levinson et al., 2001; de Janvry et al., 2006; Krueger, 2007; Gitter and Barham, 2007). Higher agricultural returns from expanding markets for non-traditional export crops, including in some areas Fair Trade[1] and/or organic certified products such as coffee, provide one potential new economic opportunity. Migration and remittances have also dramatically expanded over the last decade in southern Mexico (Barham et al., 2011) and several works have shown the potential for both to influence education outcomes (Edwards and Ureta, 2003; McKenzie and Rapoport, 2007a). The third trend is the growth of cash transfer programmes, such as Oportunidades in Mexico (Schultz,

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2004; Rawlings and Rubio, 2005). This article provides an integrated theoretical and empirical strategy for exploring the impacts of these three changes on secondary school attainment in southern Mexico using a random stratified sample of 845 coffee farming households in Oaxaca and Chiapas, surveyed in 2005–2006.

We open by discussing the expansion of Fair Trade-organic coffee cooperatives, migration, and Mexico's conditional cash transfer programme for education. Next, we structure the empirical analysis with a formal model of labour/schooling decision for an adolescent in a coffee producing household. This model provides testable hypotheses on how schooling attainment for young males and females could vary with household participation in Fair Trade-organic cooperatives, income transfers, and migration options. These hypotheses are then examined using a household fixed-effects econometric model. A robustness check using instrumental variables (IV) to address the potential endogeneity of household cooperative participation provides consistent, although slightly larger impacts of Fair Trade participation on schooling.[2] The estimation results are used in a decomposition analysis to identify the primary drivers of the dramatic change in educational attainment that has occurred especially among female youth in our sample.

Our core findings are as follows. First, we find that for females, their household's participation in a cooperative with access to Fair Trade/organic markets contributed about 0.7 years to increased schooling between the 16 and 25 year old cohorts of female youth. This effect is also positive for males but is less precisely estimated. Second, higher community migration rates to the United States interacted with family structure (our proxy variable for migration and remittances) contributes two years of additional schooling for females, or almost three times the effect of participation in a cooperative. The effect is again weaker for males, perhaps because they are more likely to work or migrate to the United States at younger ages. Third, the effect of Oportunidades (estimated in Schultz, 2004) on female educational attainment is similar in magnitude to the effect of household participation in a Fair Trade-organic cooperative. Together, these trends help explain the nearly two year increase in schooling between the 16 and 25 year old cohorts of female youth in our sample.



## II. Context: Coffee, Migration, and Cash Transfers

Especially as children move into their teens, school enrolment choices for poor rural households involve stark trade-offs between the potential returns to schooling versus the direct costs of schooling and the opportunity cost of forgone labour returns from their children (Basu, 1999). Households in many rural communities of southern Mexico and Central America earn much of their income by growing coffee, which is labour intensive; this makes it both a source of finance for education and an opportunity for generating family earnings from child labour (Krueger, 2007; Gitter and Barham, 2009).

Participation in Fair Trade-organic cooperatives is a relatively new phenomenon for the households in our sample. In 1995 only 9 per cent of households participated in a cooperative compared with 42 per cent in 2005.[3] Furthermore, almost all sampled cooperatives were certified to sell to the Fair Trade market by 2005 and most of their members held an organic certification. Both markets provide opportunities for higher prices. At the time of study, Fair Trade rules stipulated a minimum price of $1.26 per pound of coffee, or $0.05 per pound above the world price, whichever was higher. Coffee that was also certified organic received an additional $0.15 per pound (for a total minimum price of $1.41). This payment goes to the producer cooperative, which subtracts expenses before paying members. In our sample, coffee farmers participating in cooperatives earned about $150 per hectare more than non-participating farmers, which translates into $440 for the average sized cooperative farm, or 13 per cent of the sample-mean household income.

By raising returns to coffee production and decreasing producer exposure to price shocks (such as the severe price decline in 2000–2003), Fair Trade and organic coffee arrangements create

cross cutting effects for education decisions in rural households. On the positive side are the income effect[4] from price premiums and the risk reduction effect of more stable prices,[5] both of which should aid rural households' investments in education. On the negative side is the potential increased opportunity cost of the child's time in school, especially for poorer households that might not be able to finance hired labour (Krueger, 2007; Gitter and Barham, 2009). Furthermore, income and substitution effects may affect males and females differently. The substitution effect is likely to affect males more as they typically provide more year-round labour on coffee farms and are also more likely than girls to work in rural labour markets (Levinson et al., 2001).

Traditionally, labour opportunities for southern Mexican households depended primarily on local or perhaps regional activities. Since the late 1990s, however, migration to the United States has exploded, while intra-Mexico migration has also grown markedly. In the sample, 25 per cent of households had a member living in the United States in 2005 as compared to 3 per cent in 1998. Likewise, migration within Mexico expanded from just over 20 per cent of households in 1998 with a member living elsewhere in the country to close to 40 per cent in 2005. These trends reflect both lifecycle effects of families and the expansion in migration that occurred throughout the region during this era.

Migration opportunities, especially to the United States, can also affect household decisions to invest in education. Migration provides remittance flows to finance education of youth. Migration can also change the returns to schooling by opening new labour markets that may or may not reward marginal increases in child education (Borjas, 1987; Chiquiar and Hanson, 2005; McKenzie and Rapoport, 2007a). As Borjas and others document, the pull of US migration may decrease the incentive to complete high school, because of low returns to Mexican schooling earned in unskilled jobs in the United States. On the other hand, domestic migration opportunities could increase returns to education as jobs within Mexico value more education (Davis et al., 2002).

International migration typically provides more labour market opportunities for males than females. In the sample, about 25 per cent of males between the ages of 20 and 30 had migrated to the United States, three times the percentage of female migrants. Also, consistent with better local labour market opportunities for males, Amuedo-Dorantes and Pozo (2006) found that households that received remittances tended to reduce labour market participation more for female members than for males. Both of these effects could translate into higher educational attainment among adolescent girls.

Oportunidades (formerly called Progresa) has been in place in our study region in southern Mexico since 1998–1999. Since 1995, the Mexican government has been at the international forefront of promoting educational attainment for the poor (Fiszbein and Schady, 2009), with almost one in five Mexican households participating in Oportunidades (Rawlings and Rubio, 2005). Evaluations have verified its positive effects on schooling of younger children (Schultz, 2004; de Janvry et al., 2006). Oportunidades was present in every sampled community, and over 80 per cent of households participated, with the average annual payment approaching $500 per household, an amount similar to the average net coffee revenue.

## III. A Simple Model of Coffee Growing and Schooling

To guide our empirical analysis, we present a theoretical model with testable implications. Consider a household deciding how to allocate an adolescent's time between schooling ($s$) and coffee production ($1-s$). Coffee production is assumed to be a decreasing returns to scale technology $f(.)$ that for simplicity requires only labour. The household can produce coffee with its own labour or hired labour ($l$); the two are assumed to be perfect substitutes. For simplicity, we assume that adult household labour is fixed. Therefore, the amount of labour used in coffee production is $(1-s+1)$ and the fixed adult labour implicit in the

production function. The adolescent only works on-farm. If the household needs more labour, it can hire at wage $w$.

The wage is an increasing function of the price of coffee ($p$) and the community migration rate ($M_c$).[6] Intuitively, more migration means less labour supply and greater labour demand as remittances stimulate demand for other labour services. Indeed, in our sample US migration rates and community average remittances are positively correlated with community average wages.

The family receives $p$ for coffee if it is uncertified while it receives an additional per unit premium ($\pi$) if it is certified as Fair Trade-organic. A natural question is why some farmers are not certified. Organic certification is costly, and includes time spent participating in cooperative activities. Additionally, costs vary across households as some live far from the nearest cooperative and distance increases the cost of attending workshops important for fulfilling certification standards. More importantly, certification like other technologies (Feder, 1980) requires fixed costs leading to higher average certification costs for small farms; indeed, organic farms in the sample are on average 50 per cent larger than non-organic farms. To incorporate the decision to certify into the model we suppose that the household faces a certification cost of $\theta$. The farm will be certified if the increased revenue from the premium price exceeds certification cost: $(p+\pi)f(L^*)-pf(L')+w(L'-L^*) > \theta$,[7] where $L^*$and $L'$ are the profit maximising labour levels with and without the premium.

Remittances or government transfer payments enter into our framework by giving the household an exogenous income transfer y. Similar to Krueger (2007), the family derives utility from consumption $U(c)$ and schooling $V(s)$ where both $U$ and $V$ are strictly concave functions. Assuming a separable utility function, the farmer maximises $U(c) + V(s)$ subject to the constraint specified in Equations (1.1) or (1.2).

$$c = y + (p+\pi)f(1-s+1) - w(p, M_c)l - \theta \quad \text{if } (p+\pi)f(L^*) - pf(L') - w(L'-L^*) > \theta \tag{1.1}$$

$$c = y + pf(1-s+l) - w(p, M_c)l \quad \text{otherwise} \tag{1.2}$$

Substituting the expression for $c$ into $U(c)$ yields an unconstrained maximisation problem where the farmer chooses labour hired $l$ and schooling $s$. We consider the case of the certified farmer:

$$\underset{s,1}{\text{Max}}\ U(y + (p+\pi)f(1-s+l) - w(p, M_c)l - \theta) + V(s) \tag{2}$$

Assuming an interior solution, the first order conditions (FOCS) are

$$s^* : U'(p+\pi)f_1 = V' \tag{3.1}$$

$$l^* : (p+\pi)f_1 = w(p, M_c) \tag{3.2}$$

where $U'$ and $V'$ are marginal utilities and $f_1$ is the marginal productivity of agricultural labour. The FOCs are intuitive. Under optimality, the marginal utility of consumption multiplied by the marginal value product of allocating one more unit of adolescent labour to coffee farming must equal the marginal benefit from schooling. Because the household can hire in labour, the marginal value product of agricultural labour must equal the wage (Equation (3.2)). Combining the two conditions yields:

$$\frac{V'}{U'} = w(p, M_c) \tag{4}$$

From the FOCs summarised in Equation (4) we can derive relevant comparative statics. The premium $\pi$ and the cash transfer $y$, which could be from conditional cash transfers or

remittances,[8] increase schooling. By increasing income, they lower the marginal benefit of allocating the adolescent's labour to coffee farming.

$$\frac{\partial s^*}{\partial \pi} > 0 \tag{5.1}$$

$$\frac{\partial s^*}{\partial y} > 0 \tag{5.2}$$

Because a higher wage (either from a higher $p$ or a higher $M_c$) decreases profits from coffee farming and by extension income, it has the opposite effect on schooling.

$$\frac{\partial s^*}{\partial w} < 0 \tag{6}$$

The premium induces only an income effect because labour can be freely bought in our model. Large swings in world coffee prices undoubtedly affect employment and wage outcomes in coffee growing regions (Krueger, 2007). Fair Trade-organic premiums paid to a small proportion of growers for some of their coffee, however, are unlikely to affect wages, even though adoption of more labour-intensive organic production methods could increase labour demand. Because participation in Fair Trade-organic coffee production likely has a small effect on labour demand and wages (Barham et al., 2011), we make wages a function of only the world coffee price and not the premium.

If households were limited in how much labour they could hire, either because of shallow rural labour markets or liquidity constraints, then the opportunity cost of schooling would be a household specific shadow wage tied to agricultural returns. In this case, the premium would induce a substitution effect that would work against the income effect highlighted earlier. Krueger (2007) and Gitter and Barham (2009) show that the substitution effect can dominate the income effect resulting in a negative relationship between schooling and agricultural wages. Similarly, when labour markets are tight, general equilibrium effects may increase agricultural wages that could create a similar substitution effect if adolescents work off farm.

A higher wage decreases the effect of the premium on schooling. The same holds for community migration rates since migration increases the wage in the sending community.

$$\frac{\partial^2 s^*}{\partial \pi \partial w} < 0 \tag{7.1}$$

$$\frac{\partial^2 s^*}{\partial \pi \partial M_c} < 0 \tag{7.2}$$

Intuitively, the premium increases the marginal value product of labour, increasing labour demand, output, and income. With a higher $w$, the premium will have a smaller effect on income because expanding production is more costly. In short, a higher $w$ decreases the income effect of the premium. A similar result is that the cash transfer y also attenuates the income effect of the premium:

$$\frac{\partial^2 s^*}{\partial \pi \partial y} < 0 \tag{8}$$

At higher income levels, the increase in income and consumption caused by the premium has a smaller effect on the marginal utility of consumption since $U$ is strictly concave. The household, therefore, needs to withdraw a smaller amount of its own labour from coffee farming (and by extension increases schooling less) to satisfy Equation (4) after the introduction of the premium.

We incorporate gender into the model via labour productivity and the benefit derived from education $V(s)$. If we assume that the benefit to schooling $V(s)$ is the same for boys and girls, and that boys have higher agricultural labour productivity, then males should have lower schooling than females, since they would be more likely to work. Additionally, the certification premium will increase the schooling of girls more than for boys. Upon receiving the premium, the household replaces its own labour with hired labour (implicitly increasing schooling in the process) and because male labour is more productive than female labour, it takes a smaller decrease in the use of male family labour to once again satisfy Equation (4).

Our simple model of child labour allocation assumes well functioning labour markets and no general equilibrium effects of coffee price premiums on wages. We highlight four model predictions that guide the empirics.

1. Participation in Fair Trade-organic cooperatives should increase schooling, though more so for girls than boys.
2. Access to income transfers through remittances or government programmes should increase schooling.
3. Higher community migration rates should decrease schooling by increasing wages and decreasing income from coffee growing. Though outside of the model, higher community migration rates also improve migration networks which lower the cost of migrating. Both effects could decrease schooling especially for boys if they are more likely to migrate as adolescents. To the extent that adolescents work off-farm, an increase in local wages increases the opportunity cost of schooling which should reduce their education level but potentially raise that of younger siblings.
4. Higher community migration (and hence higher wages) should attenuate the positive effect of Fair Trade-organic cooperative participation on schooling.

The model does not explicitly explore liquidity constraints. However, coffee cooperatives may help alleviate some of the negative effects of credit constraints on educational attainment mentioned in the literature (Brown and Park, 2002; Beegle et al., 2003; Gitter and Barham, 2007). Some cooperatives in the study have associated organisations that extended credit to members. Credit could help households to smooth consumption and meet beginning of school year expenses such as books and uniforms. Additionally, access to more stable Fair Trade-organic coffee prices could increase schooling since households sometimes remove children *ex-ante* of negative shocks to insure consumption.

## IV. Data Description and Descriptive Statistics

This study is based on a survey conducted in southern Mexico between August 2005 and June 2006 (Lewis and Runsten, 2011). The survey included 845 coffee farming households in nine regions and two states; 640 households are from the state of Oaxaca and 205 from the state of Chiapas. While regions were chosen on the basis of having a Fair Trade-organic cooperative and being regionally representative of the coffee growing areas, households were selected as part of a random stratified sample. Coffee producers in each region were enumerated and divided into two basic groups: those that are organised and participate in Fair Trade-organic coffee production and those that do not. Each group was further stratified by prior information on migration history, and a random sample was drawn from the strata. Each coffee household was weighted in the descriptive analysis according to the selection probability of their particular stratum.[9] For a more information on the data, sample frame, and survey instrument see Lewis and Runsten (2011) and Barham et al. (2011).

### *Explosion in Educational Attainment*

Educational attainment has risen tremendously for young adults in our sample. The median number of years of school completed has risen almost two years in the last decade, compared to the United

States where it took 30 years (1950 to 1980) to increase schooling by one year (Hanushek, 1986). Table 1 shows that the percentage of women completing at least 9th grade (the final year of middle school in Mexico) has risen tremendously over the last 30 years. Also noteworthy is that in the youngest cohort (17–19) female educational attainment recently surpassed male attainment for the first time. This historic switch could be from both Oportunidades, which provides higher payments for girls' participation, and from better on- and off-farm earning and migration opportunities for males. As demand for schooling has increased there has also been an increase in the number of schools. At least three of the communities have built high schools in the last 10 years and around half had built middle schools.[10]

In Mexico, once children start in a level of school they typically finish it (Mehta and Villarreal, 2008). This trend is evident in our data, and justifies our focus on educational attainment of children of secondary school age. Table 2 shows the cumulative percentage of children who have completed at least the number of schooling years listed. Most children, especially those in the younger cohorts, complete primary school. Of those who continue to middle school, about 90 per cent complete it. Most youth drop out occurs between middle school and high school, with around 30–40 per cent ending their schooling after nine years. Because the key transition points are between the sixth and seventh year (when a child is 12 years old) and between the ninth and tenth year (when a child is 15), the independent variables in the econometric models of educational attainment outcomes correspond to the year when the child was 15. Similar results emerge in a parallel analysis of 12 year olds, as they do for the main variables of interest when we restrict our sample to 19–25 year olds to control for the fact that many 16–18 year olds are still in school.



### *The Household Economic Context: Income Sources and Labour Markets*

Mean household income is about 33,000 pesos (10 pesos:1 US$). Of sampled households, 85 per cent reported positive net revenue from coffee, but the average net coffee revenue was just over 6000 pesos or one-fifth of income. Subsidies including Oportunidades make up a similar proportion of income and are received by 96 per cent of the sample, while income from off-farm activities account for a quarter of total income. Finally, remittances dwarf all other sources of income with an average of over 11,000 pesos received, even though nearly two-thirds of households receive no remittances.

**Table 1.** Completed 9th grade by age cohort

| | Age | | | | | |
|---|---|---|---|---|---|---|
| | 17–19 | 20–22 | 23–25 | 26–30 | 31–40 | 41–50 |
| Male | 69% | 60% | 54% | 29% | 21% | 10% |
| Female | 77% | 55% | 44% | 21% | 10% | 5% |
| Total | 73% | 58% | 48% | 25% | 15% | 8% |

**Table 2.** Distribution of years of school completed for the sample

| Age | 16 | 17 | 18 | 19 | 20 | 21 | 22 | 23 | 24 |
|---|---|---|---|---|---|---|---|---|---|
| Less than 6 years | 100% | 100% | 100% | 100% | 100% | 100% | 100% | 100% | 100% |
| 6 years | 97% | 88% | 91% | 88% | 85% | 84% | 82% | 84% | 75% |
| 7–8 years (some middle school) | 88% | 76% | 76% | 67% | 62% | 53% | 49% | 50% | 49% |
| 9 years | 75% | 69% | 69% | 61% | 60% | 49% | 47% | 44% | 46% |
| 10–11 years | 27% | 31% | 28% | 26% | 23% | 17% | 13% | 16% | 7% |
| 12 years+ | 4% | 12% | 17% | 22% | 19% | 16% | 11% | 14% | 6% |

Most households (80%) buy or sell labour or do both. Conditional on working off-farm in agriculture, the median number of days working off-farm in agriculture is 46, which is associated with a median income of 2400 pesos ($5 per day). A little more than half of all households (54%) hire labour for coffee farming, mostly for harvesting. Barham et al. (2011) use the same sample and provide further evidence of active labour markets in the region.

Forty-two per cent of sample households participated in a coffee cooperative. Of these organised households, 99 per cent were certified organic or in transition to receiving certification; certification generally takes three years during which farmers are classified as in transition. Because small farmers receive organic certification through a cooperative, all households with organic certification were also organised. In addition, 89 per cent of organised households participated in cooperatives certified Fair Trade.

Prices paid to growers and total net revenue per hectare, defined as revenue minus cash costs, increase from the non-organised group to the transition group to organic farmers. Differences in prices and net revenue, however, may reflect differences between producers other than organic status or productivity improvements associated with organic practices (Barham et al., 2011). The median numbers in Table 3, nonetheless, display the potential income effects of organic premiums assumed in the model.



*Rapid Growth in Migration*

Migration within Mexico and to the United States has taken off in recent years (Table 4). The proportion of sample households with a US migrant increased from 1 per cent to 24 per cent from 1995 to 2005, while the number of households with a Mexican migrant increased from 17 to 39 per cent. While migration rates and organic certification rates have both increased over the last decade, we do not find evidence that the two are related. In 2005, the difference in the percentage of households with a US migrant was less than 1 percentage point between households with and without organic certification.

**Table 3.** Median prices and net revenue by group

| Group | Price received (Pesos/lb) | Net revenue per hectare (Pesos) |
|---|---|---|
| Non-organised | 6.71 | 1200 |
| Transition | 7.26 | 2146 |
| Organic certified | 8.47 | 2541 |

**Table 4.** Trends that may affect education

| | Percent of households with ... | | | # Regions received |
|---|---|---|---|---|
| Year | Organic certification | a US migrant | a Mexican migrant | Oportunidades |
| 1995 | 5% | 1% | 17% | 0 |
| 1996 | 7% | 2% | 17% | 0 |
| 1997 | 7% | 3% | 18% | 0 |
| 1998 | 11% | 3% | 21% | 8 |
| 1999 | 17% | 3% | 23% | 9 |
| 2000 | 24% | 6% | 27% | 9 |
| 2001 | 28% | 8% | 28% | 9 |
| 2002 | 33% | 12% | 30% | 9 |
| 2003 | 41% | 17% | 34% | 9 |
| 2004 | 45% | 19% | 36% | 9 |
| 2005 | 48% | 24% | 39% | 9 |

*Oportunidades*

The final trend that we highlight is the introduction of Oportunidades in eight of the nine regions in 1998 and the ninth in 1999. This lack of variation in the introduction of Oportunidades prevents us from directly estimating its impacts. We do, however, compare estimates of year effects in education outcomes with estimates produced elsewhere to gauge the programme's impact.

## V. Econometric Model of Educational Attainment

Our econometric analysis examines variation in educational attainment measured by years of schooling completed. Table 5 defines the main variables used, which were constructed either based on a single observation (for example, schooling) or on recall data (for example, *Mig Mex*).

The dependent variable of interest is years of schooling completed at the time of the survey. The sample includes household members, currently living in the household residence or not, ages of 16–25 at the time of the survey. We use 25 as the upper limit because individuals who were 25 in 2005 were 15 before the start of the trends in certification, migration, and Progresa/Oportunidades. Because of significant migration among people in their early twenties, it is important to include household members living elsewhere at the time of the survey because their schooling attainment is driven by household conditions at an earlier date.

The theoretical model and the literature guide the specification of our empirical model. *Organised* indicates if a household received a price premium, $\pi$, from participation in a Fair Trade/organic cooperative when the child was 15.[11] While only the premium is formally modelled, *Organised* captures additional benefits from cooperative participation like access to extension services and possibly credit.[12] A descriptive comparison shows that children whose households were organised by the time they were 15 had 1.2 years more schooling (8.7 years) than those that had not organised (7.4 years).[13]

The community migration rate, for US migration (*Mig US*) or Mexican Migration (*Mig Mex* ($M_c$ in the model)) captures the general equilibrium effect of migration on wages and by extension schooling. A comparison of households with and without US migrants at the time of the survey shows no difference in schooling. However, children in households with Mexican migrants at the time of survey completed a year more of schooling than those without migrants.[14] To allow the premium $\pi$ to affect schooling less in high migration/high wage communities, we interact *Organised* with the community migration rate to the United States.

The full effect of migration depends on whether a child benefits from remittances. We use household structure to proxy for the likelihood of a child being in a migrant or remittance household. Children with older siblings in high migration villages are more likely to benefit from

**Table 5.** Descriptive statistics

| Variable | Mean | SD | Definition |
|---|---|---|---|
| Schooling | 8.16 | 3.13 | Years of school completed at the time of the survey. |
| Organised | 0.44 | 0.5 | Indicates if the household participated in a cooperative when the child was 15 ($=1$ if organised). |
| Mig Mex | 0.28 | 0.13 | The community's within-Mexico migration rate when the child was 15. |
| Mig US | 0.1 | 0.12 | The community's US migration rate when the person was 15. |
| Older sibs | 2.49 | 2.19 | The number of older siblings. |
| Younger sibs | 2.61 | 1.83 | The number of younger siblings. |
| Age | 20.31 | 2.85 | Indicates that the person belongs to the cohort of age $j$ were $j$ ranges from 16 to 25. |
| Female | 0.49 | 0.5 | Indicates the child's gender ($=1$ if female). |

remittances. Conversely, the need for remittances may be greater if the child has many younger siblings. A higher expectation of US migration should decrease incentives to stay in school.

Household fixed effects, used in Models 1 through 3, control for time invariant differences across households. An age fixed effect ($\gamma$) controls for temporal factors affecting all children of the same cohort such as economic shocks (particularly to world coffee prices) or the introduction of Oportunidades.

For the household fixed effects models, estimating the relationship between participation in Fair Trade/organic cooperatives and schooling is possible because of temporal variation of a household's organisational status; from 1995 to 2005 the percentage of organised households in our sample increased from less than 10 per cent to almost 50 per cent. The fixed-effect specification compares educational attainment of a child who was 15 prior to the household becoming organised with that of a child who turned 15 after the household became organised. One potential concern involves households where the most educated child dropped out of school prior to turning 15, in which case the difference in schooling between the most educated child and other members may have been fixed before the most educated child turned 15. To avoid spurious correlation, we drop households where the most educated child has less than seven years of schooling. After doing so, the sample has a total of 981 people ages 16–25, of which 102 households had at least one child was 15 before the household became organised and one child who turned 15 afterwards.

Model 1 is our regression of educational attainment that we estimate separately for males and females. The dependent variable is the schooling of child $i$ in household $h$, living in community $c$. These distinctions are important as we control for individual, household, and community characteristics within the estimation. To clarify notation, $\alpha_{h(i)}$ is the household fixed effect for person $i$.

Model 1: Separate Regressions for Males and Females

$$\begin{aligned}\text{Schooling } i \;=\; & \theta_o + \theta_1 \text{Organized}_{h(i)} + \theta_2 \text{Mig US}_{c(i)} \;+\; \theta_3 \text{Mig Mex}_{c(i)} + \theta_4 \text{Older Sib}_{\text{S}i} \\ & + \theta_5 \text{Younger Sib}_{\text{S}i} \;+\; \theta_6(\text{Mig US}_{c(i)} * \text{Older Sib}_{\text{S}i}) + \theta_7(\text{Mig US}_{c(i)} * \text{Younger Sib}_{\text{S}i}) \\ & + \theta_8(\text{Mig Mex}_{c(i)} * \text{Older Sib}_{\text{S}i}) + \theta_9(\text{Mig Mex}_{c(i)} * \text{Younger Sib}_{\text{S}i}) \\ & + \theta_{10}(\text{Organized}_{h(i)} * \text{Mig US}_{c(i)}) + \alpha_{h(i)} + \gamma_{j(i)} + \varepsilon_i\end{aligned}$$

The second econometric model pools males and females. The pooled model includes all variables in the separate model (represented by the $k$ by 1 vector $x_i$ below) and several gender related variables. The variable *Female* captures gender differences like wage earning opportunities and returns to schooling while gender interaction terms allow for certain trends to affect girls and boys differently.

Model 2: Pooling Males and Females

$$\begin{aligned}\text{Schooling}_i \;=\; & \beta_o + \beta_1 x_i \;+\; \beta_2 \text{Female}_i \;+ \beta_3(\text{Organized}_{h(i)} * \text{Female}_i) \\ & + \beta_4(\text{Migration US}_{c(i)} * \text{Female}_i) + \beta_5(\text{Organized}_{h(i)} * \text{Migration US}_{c(i)} * \text{Female}_i) \\ & + \alpha_{h(i)} \;+\; \gamma_{j(i)} \;+\; \varepsilon_i\end{aligned}$$

The third model is a parsimonious version of Model 2 and excludes interaction terms involving gender and the household's organisational status.

Models 1 through 3 are estimated using ordinary least squares (OLS), assuming that the conditional mean of the error, $\varepsilon_i$, is zero. For all models we report robust standard errors clustered at the household level.[15] A final set of estimations provides a robustness check on the organised coefficient estimate by creating an instrument for the variable using community cooperative participation rates.

## VI. Results

The econometric estimates are shown in Table 6, and a decomposition of educational attainment increases for females is presented in Table 7. The first two sets of results in Table 6 present the separate regressions for males and females. The coefficient estimates in all of the regressions are of reasonable magnitude. Potential concerns about the robustness of the 'organised' measure are considered in a falsification test presented in Online Appendix I. The statistical significance of the coefficient on *Organised* requires the temporally accurate specification, and is not obtained in a falsification test where the timing of becoming organised is altered. The results for females are also robust to a non-linear specification – Online Appendix II includes the results from a censored ordered Probit model where the schooling variable is broken into discrete categories.

Model 1 shows that participation in Fair Trade-organic cooperatives and greater US migration opportunities are positively associated with female schooling. We also find age cohort effects consistent with positive impacts from Oportunidades. The impacts of organisation and migration on males are less clear, as the separate regressions show few statistically significant results for males. Both pooled models suggest that being in an organised household increases schooling for males and females, a result that holds when we instrument for *Organised*. A decomposition of the change in female schooling from 1995 to 2005 finds that household participation in Fair Trade-organic cooperatives contributed to a gain in schooling of 0.72 years while increased migration to the US contributed almost two years. Overall, the net gain in schooling between the youngest and oldest cohorts of females used in estimation was 1.6 years.

While most of these results are consistent with the hypotheses developed above, it appears that community migration captures a wage effect and the impact of remittances; hence the lack of a negative relationship between migration and schooling. Furthermore, the gender-differentiated effect of US migration on schooling suggests that migration, which is male dominated, may increase male wages relative to female wages.

### *A Closer Look at the Econometric Estimates*

As Table 6 shows, access to Fair Trade-organic price premiums (*Organised*) increases female schooling by about a year and a half. Migration to the US (*Migration US*) also increases schooling for females; a one standard deviation increase in the community migration rate (13%) is associated with about 2.5 years of schooling. This positive effect of migration decrease as females have additional younger siblings (*Migration US*Younger Siblings*), suggesting that migration opportunities may burden adolescent females with additional domestic responsibilities that impede schooling.

Looking at the age fixed effects for females, we find that the 20 and 21 year olds have about 1.7 and 2.3 more years of schooling than the 22 year olds. The 21 year olds were 15 in 1999 and were the first cohort to fully participate in Oportunidades, the conditional cash programme, which started late in 1998. The difference in age fixed effects is larger than Shultz's estimate that Progresa (now Oportunidades) increased years of schooling by 0.66. Finally, it is worth noting that many of the children in the youngest cohorts may still be in school, which would explain their smaller age fixed effects terms.

In the separate regression for males we find no statistically significant results though the coefficient on *Organised* is positive, and using an F-test we fail to reject at the 10 per cent level the hypothesis that the coefficients on *Organised* are different for females and males.[16] Interestingly, US community migration has no effect on male schooling in contrast to the positive relationship found for females – and, at the 10 per cent level, an F-test rejects the hypothesis of an equal effect for males and females. The community migration rate probably captures the effect of migration on wages and that of remittances on community wealth; hence the lack of a negative relationship between migration and schooling. Furthermore, males migrate more than females, so to the extent that male and female workers are imperfect substitutes, migration should increase wages in male dominated activities while

**Table 6.** Impact of organisation, migration and gender on years of school completed for ages 16–25, controlling for household fixed effects

| | Model 1 | | | | Model 2 | | Model 3 | |
|---|---|---|---|---|---|---|---|---|
| | Males | | Females | | Pooled | | Simplified pooled | |
| Variable | Coef. | SE | Coef. | SE | Coef. | SE | Coef. | SE |
| Organised | 0.68 | 0.69 | 1.54** | 0.70 | 0.94* | 0.49 | 0.99*** | 0.35 |
| Female | | | | | −0.85** | 0.35 | −0.34 | 0.21 |
| Migration US | −1.87 | 5.09 | 19.00** | 7.41 | 2.26 | 3.32 | 2.83 | 3.16 |
| Migration Mex | 1.34 | 8.14 | −9.25 | 8.40 | −4.04 | 4.67 | −4.61 | 4.66 |
| Older siblings | −0.04 | 1.10 | −0.58 | 0.75 | −0.85 | 0.67 | −0.78 | 0.67 |
| Younger siblings | 0.27 | 1.07 | −0.41 | 0.71 | −0.53 | 0.66 | −0.46 | 0.66 |
| Mig US*older siblings | 0.70 | 0.79 | −0.59 | 0.66 | −0.07 | 0.54 | 0 | 0.49 |
| Mig US*younger siblings | 0.15 | 1.15 | −3.43*** | 0.95 | −1.10* | 0.67 | −0.81 | 0.64 |
| Mig Mex*older siblings | 0.42 | 1.28 | 1.25 | 1.03 | 0.95 | 0.68 | 0.92 | 0.68 |
| Mig Mex*younger siblings | −0.03 | 1.33 | 0.85 | 1.29 | 0.02 | 0.79 | −0.05 | 0.80 |
| Migration US*female | | | | | 4.15 | 2.55 | | |
| Organised*female | | | | | 0.25 | 0.55 | | |
| Organised*Mig. US | −3.19 | 3.15 | −3.60 | 3.62 | −1.26 | 2.32 | | |
| Organised*Mig. US*female | | | | | 0.84 | 3.35 | | |
| Age 17 | 0.10 | 0.58 | 1.06 | 0.80 | 0.35 | 0.42 | 0.3 | 0.42 |
| Age 18 | 1.02 | 0.77 | 1.80** | 0.85 | 0.85* | 0.47 | 0.74 | 0.46 |
| Age 19 | −0.45 | 0.94 | 2.95** | 1.45 | 0.97 | 0.68 | 0.91 | 0.67 |
| Age 20 | 0.15 | 1.12 | 2.28** | 1.14 | 0.61 | 0.68 | 0.45 | 0.67 |
| Age 21 | 0.50 | 1.43 | 3.41* | 1.78 | 0.87 | 0.90 | 0.71 | 0.90 |
| Age 22 | 0.92 | 1.64 | 1.05 | 1.67 | 0.05 | 0.93 | −0.18 | 0.92 |
| Age 23 | 0.65 | 1.96 | 1.35 | 1.95 | −0.09 | 1.05 | −0.34 | 1.05 |
| Age 24 | −0.45 | 2.00 | 0.96 | 2.02 | −0.52 | 1.12 | −0.75 | 1.11 |
| Age 25 | −0.17 | 2.20 | 1.41 | 2.24 | −0.70 | 1.18 | −1.02 | 1.19 |
| Intercept | 7.02 | 6.088 | 9.12** | 4.628 | 12.08*** | 3.61 | 11.91*** | 3.652 |
| N/Households | 491 | 298 | 490 | 294 | 981 | 393 | 981 | 393 |

*Notes*: *,**,*** significant at the 10, 5, and 1 per cent levels.

**Table 7.** Impact of organisation, migration and gender on years of school completed for ages 16–25, omitting household fixed effects

| | Model 5 (females only) | | | | Model 6 (pooled) | | | |
|---|---|---|---|---|---|---|---|---|
| | OLS | | IV | | OLS | | IV | |
| | Coef. | SE | Coef. | SE | Coef. | SE | Coef. | SE |
| Organised | 0.67** | 0.33 | 6.28*** | 2.26 | 0.56** | 0.23 | 2.84* | 1.61 |
| Female | | | | | −0.17 | 0.18 | −0.25 | 0.19 |
| Migration US | 3.17 | 3.77 | 4.03 | 4.99 | 2.12 | 2.58 | 2.88 | 2.65 |
| Migration Mex | −3.87 | 4.33 | −1.07 | 5.99 | −4.09 | 2.91 | −3.46 | 3.25 |
| Older siblings | −0.25 | 0.19 | −0.42 | 0.29 | −0.36*** | 0.14 | −0.38** | 0.15 |
| Younger siblings | −0.34 | 0.22 | −0.47 | 0.34 | −0.51*** | 0.15 | −0.58*** | 0.17 |
| Mig US*older siblings | −0.27 | 0.45 | 0.17 | 0.67 | 0.03 | 0.34 | 0.11 | 0.35 |
| Mig US*younger siblings | 0.06 | 0.83 | 1.50 | 1.38 | 0.27 | 0.53 | 0.59 | 0.60 |
| Mig Mex*older siblings | −0.06 | 0.47 | 0.10 | 0.71 | 0.03 | 0.32 | 0.05 | 0.38 |
| Mig Mex*younger siblings | 0.38 | 0.74 | 0.39 | 1.14 | 0.55 | 0.48 | 0.78 | 0.55 |
| Age 17 | 0.54 | 0.42 | 0.99 | 0.66 | 0.09 | 0.29 | 0.21 | 0.33 |
| Age 18 | 0.57 | 0.52 | 0.68 | 0.72 | 0.42 | 0.36 | 0.61 | 0.41 |
| Age 19 | 0.75 | 0.78 | 1.46 | 1.06 | 0.42 | 0.52 | 0.82 | 0.63 |
| Age 20 | 0.34 | 0.76 | 2.40* | 1.27 | 0.04 | 0.55 | 0.91 | 0.86 |
| Age 21 | 0.27 | 1.02 | 3.46* | 1.81 | 0.13 | 0.70 | 1.29 | 1.09 |
| Age 22 | −1.44 | 0.92 | 1.00 | 1.47 | −0.62 | 0.71 | 0.59 | 1.12 |
| Age 23 | −1.02 | 1.12 | 3.51 | 2.32 | −0.53 | 0.80 | 1.17 | 1.52 |
| Age 24 | −1.30 | 1.08 | 2.55 | 2.19 | −1.03 | 0.83 | 0.56 | 1.45 |
| Age 25 | −1.47 | 1.11 | 2.87 | 2.29 | −1.22 | 0.87 | 0.57 | 1.60 |
| Lives in outskirts | −0.35 | 0.51 | −0.05 | 0.63 | −0.69* | 0.36 | −0.54 | 0.37 |
| Lives in countryside | −1.42** | 0.61 | −1.19 | 0.75 | −1.44*** | 0.53 | −1.31** | 0.56 |
| Coffee land (Ha) | 0.02 | 0.02 | −0.02 | 0.04 | 0.02 | 0.02 | 0.01 | 0.03 |
| Household size | 0.22** | 0.09 | 0.22 | 0.16 | 0.23*** | 0.08 | 0.19** | 0.09 |
| Intercept | 10.47*** | 2.41 | 6.11* | 3.61 | 10.77*** | 1.69 | 9.36*** | 2.11 |
| Controls for region | yes | | yes | | yes | | yes | |
| Observations | 490 | | 490 | | 981 | | 981 | |

*Notes*: *,**,*** significant at the 10, 5, and 1 per cent levels.

having less effect on female wages. A greater wage gap between males and females would favour schooling of females over males.

The pooled model yields several noteworthy results. The coefficient on the binary indicator for gender (*Female*) shows that females completed about four-fifths of a year of school less than their male counterparts, all else constant. Consistent with the separate regression results, being part of an organised household increases schooling by about a year. The pooled model also supports (though it is not statistically significant) the hypothesis that migration opportunities reduce the impact of coffee price premiums for males (US Migration*Organised).

### *A Robustness Check*

With the fixed-effects models, identification of the causal effect of *Organised* on schooling requires that time variant factors omitted from the analysis are uncorrelated with schooling and a household's organisational status. Also, because identification of the coefficient on *Organised* comes from only households with multiple children, the estimated effect is conditional on having more than one child in a certain age range. It is possible that the effect of cooperative participation on education is different for large and small households, meaning that if households with fewer than two children had several children, the effect of *Organised* on their children's schooling would be different to the effect on youth in sample households which actually have several children.

To test the sensitivity of our results to assumptions about time variant factors and differences between households with several children and those with two or fewer, we estimate Models 5 and 6 where we omit the household fixed effect. Model 5 corresponds to the female only version of Model 1 while Model 6 corresponds to Model 3. We estimate both models with and without instrumenting for the household's organised status. Since these models exclude a fixed effect, several time invariant household variables (household size and dwelling location, coffee land, region fixed effects) are added. Furthermore, to simplify IV estimation, we drop the interaction between *Organised* and *US Migration*, which is statistically insignificant in the previously estimated models.

A valid instrument must be correlated with the potentially endogenous variable that it is instrumenting for (in this case, *Organised*) but uncorrelated with the outcome of interest (an individual's years of schooling). Community level variables can often be good instruments because they can capture the influence of exogenous historical, social, or geographic factors on individual behaviour (for example, Taylor and Lopez-Feldman, 2010). To instrument for *Organised*, we use *Per cent Organised* – the percentage of households in the community that were organised (excluding the household in question) when the child was 15. At the most basic level, participation rates reflect the opportunity to participate in a cooperative in a given area since at some times and in some areas there may have been few or no cooperatives operating. Furthermore, having neighbours who are cooperative members makes it easier to learn about the costs and benefits of participation. Greater social pressure to imitate peers would also encourage participation.

As expected, the first stage of the IV regressions show that *Per cent Organised* is strongly and positively correlated with household cooperative participation (results of the first stage are provided in Online Appendix III). For the first stage of Model 5, the *F* test that the coefficient on *Per cent Organised* is zero is 11.55, while it is 20.73 for Model 6. The F stats for both models are greater than 10, which is the general rule of thumb for identifying weak instruments (Stock et al., 2002), suggesting that weak instrument bias may influence the results in Model 5 but not in Model 6 (the pooled model).

The OLS versions of both Models 5 and 6 suggest that being organised is associated with more than half a year of additional schooling. In both cases, the estimated effect is less than in the fixed-effects models. For Model 5 this difference may reflect the omission of the interaction between US community migration and *Organised*. However, a similar result for Model 6 suggests

that unobservable household variables may lower schooling outcomes more for organised households than for other households. Alternatively, it could mean that the effect for households with multiple children is larger than for single child households.

Instrumenting for *Organised* using community cooperative participation rates when the child was 15 increases the estimated effect of *Organised* on schooling relative to the household fixed-effects models and to the OLS versions of Models 5 and 6. The increased effect suggests that the correlation between cooperative participation and the error term, $\varepsilon_i$, may be negative. The coefficient on *Organised* in the IV version of Model 5 is very large and could reflect a weak instrument. The coefficient estimated in Model 6, where the instrument was substantially stronger, is more reasonable though at 2.8 years is still large. Together, the results in Table 7 support the results of the fixed-effect models in Table 6, which while not incontrovertible, are likely the most reliable since they control for time invariant unobservable household characteristics.



*Decomposing Changes in Schooling (1995–2005)*

The 16 year old female cohort has an average of 1.91 more years of schooling than the 25 year old cohort. For the subsample used in estimation – youth from households where the maximum education among children exceeded six years – the change was 1.61 years. We explore in Table 8 how distinct trends in southern Mexico contributed to this gain in schooling. Gains in schooling can come from changes in levels (families becoming wealthier) and from changes in coefficients (families of a given wealth having a greater propensity to invest in education). For the 16 and 25 year old cohorts we can calculate the predicted mean level of schooling

$$\bar{\hat{s}}^{16} = \hat{\theta}\bar{x}^{16} + \bar{\hat{\alpha}}^{16} + \hat{\gamma}^{16} \tag{9.1}$$

$$\bar{\hat{s}}^{25} = \hat{\theta}\bar{x}^{25} + \bar{\hat{\alpha}}^{25} + \hat{\gamma}^{25} \tag{9.2}$$

where $\hat{\theta}$ is the 1 by $k$ vector of coefficient estimates specified in the equation for the females only model, $\bar{\hat{\alpha}}^{16}$ is the mean household fixed effect for the cohort of 16 year old females and $\hat{\gamma}^{16}$ is the cohort's age fixed effect. Subtracting the predicted schooling of 16 year-old females from that of 26 year old females,

$$\bar{s}^{16} - \bar{s}^{25} = \hat{\theta}(\bar{x}^{16} - \bar{x}^{25}) + \bar{\hat{\alpha}}^{16} - \bar{\hat{\alpha}}^{25} + \hat{\gamma}^{16} - \hat{\gamma}^{25} \tag{10}$$

The contribution for a given covariate $x_1$ is $\hat{\theta}(\bar{x}_1^{16} - \bar{x}_1^{25})$ while the contribution for the cohort and household fixed effects are $\hat{\gamma}^{16} - \hat{\gamma}^{25}$ and $\bar{\hat{\alpha}}^{16} - \bar{\hat{\alpha}}^{25}$.

Our data do not permit us to allow all coefficients to change through time, hence changes in $\theta$ are not part of the changes in schooling between cohorts. Instead, the cohort fixed effect $\gamma$

**Table 8.** Decomposition of educational attainment growth

| Variable | Contribution |
|---|---|
| Organised | 0.72 |
| Migration US* | 1.99 |
| Migration Mex* | −0.76 |
| Siblings** | 0.02 |
| Cohort fixed effect | −1.41 |
| Household fixed effect | 1.05 |
| Total change | 1.61 |

*Notes*: **Migration US* and *Migration Mex* includes the contribution of migration interaction terms. ***Siblings* includes the contribution of *younger siblings* and *older siblings*, excluding interaction terms.

captures the effect of being a different age in the same household while the average household fixed effect reflects differences in households with 16 year olds and households with 25 year olds.

Household participation in Fair Trade-organic cooperatives contributes 0.58 years of the 1.61 year increase in schooling between 16 and 25 year old females. Migration to the United States, including the interaction terms, contributes almost 2 years, more than double that of participation in cooperatives, while migration within Mexico accounts for a decrease of −0.76 years, also including interactions. The contribution of age fixed effects is −1.47 years. It is not surprising that the 25 year olds have a greater cohort fixed effect since 16 year olds are still in school.

The contribution of household fixed effects is 1.05 years and is large relative to the total gain in education. This contribution reflects differences between younger and older household 'cohorts', to the extent that households with 16 year olds belong to a younger household 'cohort' than households with 25 year olds. Younger households may be more disposed to invest in education in part because of improved access to community schools. Though they may have less wealth than older households, younger households may also invest more in education because they themselves are better educated.



## VII. Conclusion

Rural communities in southern Mexico have undergone dramatic changes on several fronts in the past two decades including rapid growth in education. This article develops a formal model of education and child labour choices for a coffee producing household that yields testable hypotheses related to the effects of participation in Fair Trade-organic cooperatives, expanding US migration opportunities, and the implementation of Progresa/Oportunidades. We probe several hypotheses using a household fixed-effect econometric model. The estimation results allow a statistical decomposition of the factors driving educational attainment growth in the sample.

Our results confirm the following hypotheses. First, participation in Fair Trade-organic cooperatives increases schooling more for girls than boys, and accounts for a 0.7 year increase in schooling between the 16 and 25 year old cohorts of females. The stronger effect for females is consistent with greater rural labour market opportunities for males, so the improvement in income associated with cooperative participation is likely to favour females. Second, access to income transfers (and/or higher community wages) associated with US migration significantly improved female educational outcomes, and accounted for more than twice the increase in schooling associated with cooperative participation. These effects were more ambiguous for males, potentially reflecting their greater propensity to migrate and their more immediate trade-offs than females from increased labour market opportunities. The results also are consistent with previous findings that Oportunidades increased educational attainment for females. It is worth adding that those females with younger siblings benefit less from higher US migration in their community. It seems likely that they assume responsibilities for household activities, such as child care, when a household head migrates. Other than this trade-off, the higher income opportunities associated with participation in Fair Trade-organic cooperatives, US migration, and Oportunidades all contributed significantly to the major gains observed in female education over the past decade.

While we find positive effects of participation in Fair Trade-organic coffee cooperatives on schooling, the gains may be temporary. Barham et al. (2011), which uses the same sample as this article, find modest improvements in household welfare from participation in the Fair Trade-organic market, but that the most promising pathway out of poverty for sample households appears to be in greater integration with labour markets, not in intensifying coffee cultivation. Thus, while Fair Trade-organic coffee arrangements may improve welfare in the short term, the gains in the long term are unclear. In their study of snow pea production (a non-traditional export) in Guatemala, Carletto et al. (2009) find that households that adopted early and then

exited the market after a price boom saw the biggest improvement in long-term welfare, while those who never exited experienced only modest gains compared to non-adopters. If new market entrants reduce price premiums associated with Fair Trade-organic coffee, households that use the short term welfare gains to invest in education or international migration to exit from coffee may see the biggest welfare gains.

Our results suggest that the trends studied can help poor families improve what have historically been dismal educational outcomes. The results may be meaningful in other contexts, especially in Central America, where participation in Fair Trade-organic coffee grew in recent years, similar US migration dynamics are underway, and conditional cash transfer programmes are being piloted or implemented. Increased demand for education is likely to necessitate investments in the supply of education (schools and teachers) to avoid overcrowding in classrooms and shortening of school days to serve more students. In addition, because the associated migration and labour market opportunities tend to involve young males more than females, the historic pattern of greater schooling of males (by a year or so) may change and even reverse as it has in our sample in southern Mexico. This imbalance may reflect households foregoing long-term opportunities of certain male family members to achieve broader gains for other members, and as such may warrant further attention by researchers and policymakers.



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## Notes

1. This refers to coffee certified according to the norms of the Fair Trade Labelling Organizations International (FLO, 2009).
2. Household fixed effects control for wealth endowments that can affect migration. Massey et al. (1994) and McKenzie and Rapoport (2007b) show that middle-class households with the resources to send migrants, and larger potential gains from migration, are most likely to migrate.
3. Unless otherwise noted, descriptive statistics are weighted to account for oversampling.
4. Income does not affect schooling in the original Becker model (1964). If, however, borrowing constraints exist or if education has value as consumption, income will affect schooling decisions.
5. In the case of one large Fair Trade cooperative in Central America, De Janvry et al. (2010) estimate the Fair Trade premium for the years 1997 to 2009 and find that the Fair Trade market always paid higher prices and the premium was greatest during times of low prices in the world coffee market.
6. Note that the wage is not a function of the coffee price premium, which assumes no general equilibrium effects from premiums.
7. Additional benefits may include technical assistance and help with purchasing coffee production capital. To make the model tractable we do not include these benefits.
8. Conditional cash transfers may also include minimum schooling time, however, we do not include this to make the model tractable. Their inclusion would not change the basic results.
9. We do not reweight the econometric estimates because they are used for inference rather than developing representative descriptive statistics of the population.
10. Unfortunately we lack complete data on school construction so we are not able to include it in the analysis.
11. The *Organised* group includes both certified organic households and households in transition to certification. In a preliminary analysis we broke the *Organised* group into organic transition and certified organic households and found no evidence that the groups have different schooling patterns.
12. It is possible that cooperatives provide benefits to households that are unrelated to price premiums and that affect schooling. We do not isolate the premium or agricultural-returns effect apart from other possible effects captured by the *Organised* variable, though casual observations during data collection suggested that these alternative effects are likely negligible.

13. This result is significantly different from zero at the 1 per cent level.
14. This result is significantly different from zero at the 1 per cent level.
15. Because the survey design included geographically clustered samples, we ran all four models clustering the standard errors at the community level. The statistical significance of *Organised* did not change.
16. The test was conducted by estimating a pooled model with a full set of gender interaction terms and running a restricted (equal coefficients) and unrestricted (different coefficients) model.

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