FCND DP No. 129

FCND DISCUSSION PAPER NO. 129

LABOR MARKET SHOCKS AND THEIR IMPACTS ON WORK AND SCHOOLING: EVIDENCE FROM URBAN MEXICO

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March 2002

ABSTRACT

We use individual observations from a panel of families during the period of the peso crisis in Mexico to investigate whether and how labor market shocks, as proxied by changes in the gender- and age-specific unemployment rates in the metropolitan area of the household, affect the intertemporal time allocation of adult members and children.

Our findings suggest that significant added-worker effects are in operation, especially for adult females of poorer households and in some cases children. The same shocks also increase significantly the probability that children do not continue school in the next year. We also present evidence suggesting differential treatment based on the sex of children within families.

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ACKNOWLEDGMENTS

We thank Agnes Isnawangsih for excellent research assistance, and participants at seminars at the International Food Policy Research Institute, George Washington University, and the University of Pennsylvania for helpful and constructive comments. All errors remain our own.

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1. INTRODUCTION

In recent years there has been increasing concern that macroeconomic fluctuations may have an adverse effect not only on the accumulation of physical but also human capital. Dellas and Sakellaris (1995), for example, using data from the United States, find that the decision to enroll in college is related strongly to labor market conditions (measured by the state-level unemployment rate and earnings) and to the real interest rate. In related studies, Behrman, Duryea, and Szekely (1999), using cohort-specific data combined from 18 Latin American and Caribbean countries, find that macroeconomic stability, represented by the international terms of trade and GDP volatility, are the most significant determinants of schooling attainment to explore the causes of the slowdown in schooling accumulation in these countries since the 1980s debt crisis, while Flug, Spilimbergo, and Wachtenheim (1998), using aggregate cross-country panel data, find a significant negative correlation between secondary school enrollment rates and income or employment volatility. Despite the available evidence on the relationship between schooling and indicators of macroeconomic activity, there is little solid empirical evidence at the individual level on how labor market conditions affect the intertemporal use of time of both younger and older family members.

In this paper we attempt to shed some light on this issue. Invariably, macroeconomic crises lead to significant changes in the purchasing power of household income as salaries fail to keep up with the general inflation rate. Crises may also directly affect household income as production firms shut down their plants and lay off workers.

Even if household members are not directly displaced from their jobs, macroeconomic events and labor indicators can also affect household choices through (1) the information they impart about employment prospects and (2) the increased uncertainty associated with future returns to investments.

In response to these new economic conditions, households may choose to reallocate the utilization of their available resources. For example, during periods of relatively low wages, households may decide to allocate their time in nonwork activities (MaCurdy 1981). The lower opportunity cost of time, for example, may induce families to invest in children's schooling. On the other hand, the lower level of household income combined with the increased uncertainty about the future returns to investments may work in the opposite direction, decreasing the demand of families for children's education. In general, investments in the human capital of children are intricately related to the ability of families to insure through formal financial markets (borrowing or buying insurance) or informal arrangements (getting help from friends and relatives). Insofar as households have limited access to formal and informal insurance and are thus forced to use their children's time for income-earning activities (e.g., sending children to work or withdrawing them from school), then child school attainment may be lowered and, in effect, poverty may be intergenerationally transmitted.

Our analysis allows for the possibility that labor markets can play simultaneously a dual role for households as a source of risk and a means of insuring against some types of risk. Labor markets, for example, to the extent they are segmented by region, worker age, and gender, can also provide a means for insuring total family consumption and

investments in children's human capital against shocks. They do this by providing the opportunity for other family members to adjust their labor supply to offset the losses in family income (e.g., see Kochar 1999). Adjustments in the labor supply of adult females, for example, may be a means of protecting their investment in their children's human capital.

In recent years questions of adjustments in labor supply have received considerable attention. The literature on the added-worker effect, for example, examines the effects of unanticipated changes in husbands' earnings on women's hours of work (Maloney 1987; Hyslop 2001; Stephens 2001; Haurin 1989; Yeung and Hofferth 1998; Gruber and Cullen 1996). Much of this research, however, focused on the effects of shocks on the labor supply of the spouse of the displaced or unemployed worker. One critical question for developing countries where social insurance, safety nets, and financial markets are less developed is whether the adjustments in time use spill over to younger children in the household. The available evidence suggests that they do. Pitt and Rosenzweig (1990), for example, investigated the implications of infant morbidity on the time allocation of teenage boys and girls in Indonesia. Their findings indicate that these health shocks tended to reinforce the existing gender-based differences in the division of time between household, labor force, and schooling activities. Similarly, Jacoby and Skoufias (1997) examined the impact of unanticipated income shocks on the school attendance of children in rural India and found that households withdrew their children from school when experiencing shortfalls in crop income. This suggests that poor households, at least in the short term, use children's time as a form of insurance.

Macroeconomic shocks, even if short lived, may also have long-term consequences on school attainment and not only the school attendance of children. Duryea (1998), for example, using a rotating panel survey set from Brazil, found evidence that when fathers became unemployed during the school year, the probability of grade advancement for both boys and girls declined significantly.

Our analysis relies on the National Mexican Urban Employment Survey *(Encuesta Nacional de Empleo Urbano*—ENEU), a large, longitudinal survey in urban areas of Mexico. The ENEU contains repeated observations over a period of five quarters on the time use of individuals 12 years of age or older in six main activities for the week prior to the interview. The survey also contains demographic and socioeconomic information as well as a standard set of detailed questions on employment, wages, and job-search activities. Although the length of the panel is quite short, one of the compensating advantages was the inclusion of questions on both intrafamily and intertemporal behavior.

With few exceptions (e.g., Skoufias 1996) research in this area has been constrained by the scarcity of longitudinal data on hours devoted to domestic activities in addition to market work. Relying on hours of wage work only is likely to yield estimates that reflect substitution away from work at home as well as leisure (e.g., see Killingsworth and Heckman 1986). Estimates based on broader measures of work provide a more accurate estimate of the household preferences toward leisure. In addition, provided that an appropriate measure of time is available for all members of the

sample, the empirical analysis is not affected by problems arising from corner solutions and labor market entry and exit decisions.

Using the intertemporal model of labor supply (MaCurdy 1981; 1985) as the basis for our empirical analysis, we regress individual specific changes in time allocated to work and schooling on a set of variables describing changes in the value of time of household members, and shock variables such as the growth rate in the local unemployment rate corresponding to the age of the household head and his spouse. The analysis is conducted separately for adult (18–65 years) males and females and for boys and girls (12–17 years) to capture the potential differences in the effect of the shocks according to sex and age.

Our study is one of the first providing estimates of the cross effects on time allocation within an intertemporal model. We analyze two household panels. One panel was followed for five quarters during the peak of the economic crisis of 1995 in Mexico (1995:Q2–1996:Q2). The crisis began with a devaluation of the peso in December 1994, with GDP falling by 7 percent in real terms in 1995 and recovering in 1996. The other panel covers the period 1998:Q2–1999:Q2, which was characterized by low and decreasing unemployment and general economic recovery. We use two different periods in order to test whether the nature of the aggregate shocks in the economy interacted significantly with the potential insurance arrangements among households.

The paper is structured as follows. Section 2 presents a simple intertemporal model of family time allocation, and Section 3 outlines the empirical strategy adopted for

estimating the impact of labor market shocks on time allocation. Section 4 describes the data, and details the results of the analysis. Section 5 concludes.

2. MODEL

Our model is a simple extension of the standard intertemporal model of labor supply (MaCurdy 1981; 1985). To keep the presentation simple, we do not model household production explicitly, and instead we specify that the household utility function depends on the time allocated to each activity by each household member.[1] We assume for the moment that a household consists of an adult and a younger member, and we specify the preferences of the household as additively separable across time.[2] In our empirical analysis, we distinguish between four types of labor, that of boys, girls, adult males, and adult females. Utility at time t is given by

$$U(t) = U\left(C(t), L^A(t), H^A(t), L^C(t), H^C(t), S^C(t), X^*(t)\right)$$

where $C(t)$ denotes consumption of a composite commodity in period t, $L^A(t)$ is the hours of leisure of an adult member and $H^A(t)$ is her hours of work either at home or in

[1] Killingsworth and Heckman (1986) provide a useful illustration of the equivalence of these two approaches. For an equivalent dynamic model with investment in human capital, see Jacoby and Skoufias (1997).

[2] We also abstract from the possibility that time allocation decisions are made within a bargaining framework.

the market,[3] and $L^C(t), H^C(t), S^C(t)$ are the hours of leisure, work, and schooling by a younger household member in period *t*, and $X^*(t)$ is a vector of observable and unobservable factors affecting a household's preferences (as specified below). The function *U* is assumed to be strictly concave in its arguments, reflecting diminishing marginal utility of consumption or leisure.

The real wage rates expressed in units of the consumption good, are given by $W^A(t)$ and $W^C(t)$, for adults and children, respectively, and are assumed to be treated as fixed by individuals. We assume the existence of a perfectly competitive credit market that allows wealth to be transferred from period to period by holding an asset with a known and riskless real rate of return $r(t+1)$ payable at the beginning of period *t+1*.

Formally the maximization problem of the household is to choose values for

$$C(t), L^A(t), H^A(t),\ L^C(t), H^C(t), S^C(t), \text{ for } t = 1, \ldots, T,$$

to

$$MaxE_1 \sum_{t=1}^{T} \beta^{t-1} U\left(C(t), L^A(t), H^A(t), L^C(t), H^C(t), S^C(t), X^*(t)\right), \tag{1}$$

subject to the asset accumulation constraints

[3] Thus we assume that work at home and work in the market are perfect substitutes in the utility function. The assumption also has some convenient empirical implications, discussed in more detail below.

$$\overline{A}(t+1) = (1 + r(t+1))A(t) \tag{2}$$

$$A(1) \text{ constant, and } A(T+1) = 0 \tag{3}$$

and the time constraints

$$L^A(t) + H^A(t) = 1 \tag{4}$$

$$L^C(t) + H^C(t) + S^C(t) = 1 \tag{5}$$

where β is the subjective discount factor, E_t is the expectation operator conditional on the information set at period t, $\overline{A}(t)$ and $A(t)$ denote the value of real assets held at the beginning and at the end of period *t*, *V(t)* is unearned income and

$$A(t) = \overline{A}(t) + W^A(t) + W^C(t) + V(t) - \left(C(t) + W^A(t)L^A(t) + W^C(t)L^C(t) + W^C(t)S^C(t)\right).$$

Assuming interior solutions, we obtain the following first-order necessary conditions for a maximum:

$$U_C(t) = \lambda(t), \tag{6}$$

$$U_{L^A}(t) = \lambda(t)W^A(t), \tag{7}$$

$$U_{H^A}(t) = \lambda(t)W^A(t), \tag{8}$$

$$U_{L^C}(t) = \lambda(t)W^C(t), \tag{9}$$

$$U_{H^C}(t) = \lambda(t) W^C(t), \tag{10}$$

$$U_{S^C}(t) = \lambda(t) W^C(t), \tag{11}$$

$$\lambda(t) = (1 + r(t+1))\beta \; E_t(\lambda(t+1)), \tag{12}$$

where $\lambda(t)$ is the lagrangian multiplier associated with the period t assets accumulation constraint, and the subscripts of *U* denote partial derivatives. As Heckman and MaCurdy (1980) and MaCurdy (1981) note, $\lambda(t)$ represents the marginal utility of wealth in period *t*, which is a sufficient statistic summarizing all past and future information relevant to the current choices of the household. Specifically, $\lambda(t)$ is a function of the path of past, current, and expected wage rates of adults and children, initial assets, the vector of observable and unobservable factors affecting utility, and the parameters describing the household's preferences.

Equations (6)–(11) in combination with the "Euler equation" for the marginal utility of wealth $\lambda(t)$ provide a characterization of the optimal consumption and time allocation choices across time. For example, equations (10) and (12) may be expressed as

$$E_t\left(\frac{U_{H^C}(t+1)}{U_{H^C}(t)}\right) = \left(\frac{\beta^{-1}}{1 + r(t+1)}\right) E_t\left(\frac{W^C(t+1)}{W^C(t)}\right). \tag{13}$$

Holding the interest rate constant, and assuming a within-period utility function that is additively separable in each of its arguments, this condition implies that the time

allocated to work (schooling) between periods $t + 1$ and t is inversely (positively) related to the ratio of the child wage rates between period $t + 1$ and t. A similar result holds for the allocation of time of adult members in work activities across time.

3. EMPIRICAL STRATEGY

We estimate time allocation functions for adult males and females, and boys and girls, derived from the subset of first-order conditions above. As is the common practice in life-cycle labor supply research, we derive marginal utility of wealth or Frisch time allocation functions that decompose current decisions at any point in time into two components (MaCurdy 1981; Blundell and MaCurdy 1999). The first component is a set of variables observed in the current period, such as current period wage rates, prices, and factors influencing individual attitudes toward work; the second component is the marginal utility of wealth $\lambda(t)$ that summarizes the influence of all past events and the expectations about future events on current decisions. An essential requirement for the empirical specification is that the marginal utility of wealth of the household $\lambda(t)$ enters additively in the time allocation decision rules of all household members. With panel data, first differencing of the individual-specific observations allows us to eliminate the unobserved marginal utility of wealth $\lambda(t)$ from the equation to be estimated.

For our empirical model, we assume the utility function in each period is additively separable in consumption and labor supply but not necessarily additively separable in the remaining arguments; these assumptions thus allow for "cross-wage"

effects in the allocation of time across activities and age-gender type. Distinguishing a household by the letter *i* and each of its members by the letter *j*, we specify

$$X^*(i,j,t) = X(i,j,t) + \varepsilon(i,j,t),$$

where $X(i,j,t)$ is the vector of observable characteristics affecting household tastes, such as age and education, and $\varepsilon(i,j,t)$ summarizes the influence of all unobservable shocks to tastes. Using the functional form common in the intertemporal labor supply literature, the equation for the leisure time adult male in the family in period *t*, for example, may be written as

$$\begin{aligned}\ln H^m(i,t) = \alpha^m(t) + \beta^m \ln W^m(i,t) + \gamma^m \ln W^f(i,t) + \delta^m \ln W^c(i,t) \\ + \psi^m \ln \lambda(i,t) + \theta^m X(i,t) + \varepsilon^m(i,j,t),\end{aligned} \tag{14}$$

where $\alpha^m(t)$ is a time varying constant term containing the market interest rate; β^m, γ^m, and δ^m are parameters summarizing the elasticity of hours of work with respect to changes in the male wage rate $W^m(i,t)$, the female wage $W^f(i,t)$, and the child wage rate $W^c(i,t)$, respectively. ψ^m is the parameter summarizing the elasticity of hours of work to changes in the marginal utility of wealth of the household. A similar equation can be specified for the hours of work by the adult women and by the children in the

household, with the superscripts on the dependent variable *H*, the parameters $\alpha, \beta, \gamma, \delta, \theta$, and the error term $\varepsilon(i, j, t)$ changed to *f* and *c*, respectively.

After taking first differences between period *t* and *t* - 1, the equation above may be expressed as

$$\Delta \ln H^m(i,t) = \alpha^m(t) + \beta^m \Delta \ln W^m(i,t) + \gamma^m \Delta \ln W^f(i,t) + \delta^m \Delta \ln W^C(i,t) + \psi^m \eta(i,t) + \theta^m \Delta X(i,j,t) + \Delta \varepsilon^m(i,j,t), \quad (15)$$

where $\Delta Z(i,t) = Z(i,t) - Z(i,t-1)$ for $Z = \ln H$, $\ln W$, X, and ε, and the constant term $\alpha^m(t)$ is appropriately redefined.

In contrast to the numerous intertemporal labor supply studies that concentrate on the parameters β^m or γ^m and δ^m summarizing the "own" and "cross-wage" of changes in the wage rate, the primary focus of this analysis is the coefficient ψ^m of the shock $\eta(i,t)$ that reflects the difference between the one-period ahead expected value of the marginal utility of wealth of household *i* and its realized value (i.e., $\lambda(i,t) = E_{t-1}(\lambda(i,t)) + \eta(i,t)$).[4]

In an uncertain environment, as various unexpected events are realized in each period, households acquire new information about their current and future prospects and

[4] The main prediction derived from the intertemporal labor supply model is that $\beta > 0$. Although the signs of γ and δ are not restricted, the model does impose a number of equality restrictions across equations. For example, the basic theoretical model predicts that $\gamma^m = \beta^f$. Given the focus of this paper on the impact of shocks, we do not explicitly test any of these cross-equation restrictions.

respond by adjusting the value of their marginal utility of wealth according to the empirical analog of equation (12). The (unitary) life-cycle model of time allocation specified above predicts that the parameter ψ^m is positive, which implies that, ceteris paribus, a negative shock is likely to lead to an increase in $H^m(i,j,t)$ the hours of work in period *t*.[5] The incidence of a negative shock increases the realized value of marginal utility of wealth in period $t + 1$ relative to its expected level as of period *t*, i.e., $\lambda(i,t) > E_t(\lambda(i,t))$, which implies that $\eta(i,t)$ has a higher value. Combined with the positive value of ψ^m, a negative shock is thus supposed to increase total hours of work, provided that there are no binding constraints in the hours a male can devote to the wage market or to domestic activities (Ham 1986; Card 1994). Moreover, the unitary model outlined above suggests that the term $\eta(i,t)$ is an integral part of the time allocation decision rules of *every* household member. This implies that negative shocks to the household must also have a positive effect on the labor supply of all other household members, giving rise to "added-worker effects."[6] The incidence of a negative shock, for example, such as the loss of employment by the male household head between period $t - 1$ and *t*, is likely to increase the hours of work of his spouse as well as children.

[5] In fact, the functional forms used commonly in the intertemporal labor supply literature restrict the parameter ψ to be equal to the intertemporal elasticity of substitution β (e.g., MaCurdy 1981). We do not impose any such restrictions.

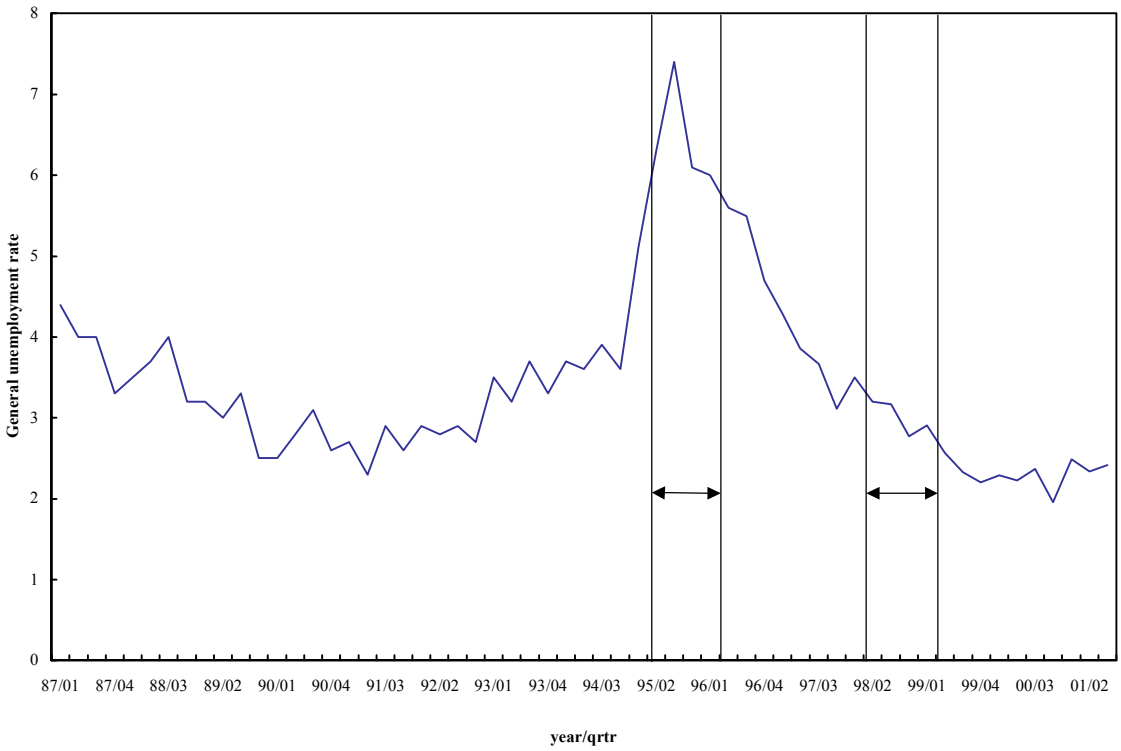
[6] For a paper on added-worker effects in the United States, based on a similar intertemporal model, see Stephens (2001).

DATA

The National Urban Employment Survey (Encuesta Nacional de Empleo Urbano or ENEU) has been undertaken quarterly since 1986 by the National Institute of Geographic Statistics and Information (INEGI). The sample and areas covered have been expanded over the years of the survey, and as of 1999, it includes 44 metropolitan centers and over 100,000 households. The ENEU includes information on time use for individuals aged 12 and over, education, family structure, dwelling characteristics, and a standard set of detailed questions on employment, unemployment, and labor market withdrawal. The time-use and labor force participation information include hours spent during the last week on school, housework, market work, and community activities.

The longitudinal data included in this research come from the 20 percent, five-quarter, rotating panel that is embedded in the design of the ENEU. Each household is interviewed every three months for a year, so that there are five observations for each household (and consequently all of the individuals in the household). The design is such that in any given cross-section of the ENEU, 20 percent of households are in their first interview, 20 percent are in their second interview, etc. Panels can be constructed by following each 20 percent of households over time. This paper uses two separate panels, one for the period from 1995:Q2 to 1996:Q2 and another from 1998:Q2 to 1999:Q2. Figure 1 displays the national unemployment rate in Mexico by quarter for the period 1987:Q1 to 2000:Q2. The two pairs of vertical band lines are drawn to highlight the time intervals covered by our two short panels in reference to the national unemployment rate.

Figure 1—National unemployment rate in Mexico



Our analysis is based on the repeated observations on individuals from households with an adult male head and a female spouse between ages 18 and 65 and at least one child between ages 12 and 17. While there is a fair amount of attrition in these short panels (about 27 percent have left the sample by the end of the fifth interview), our paper keeps all individuals in each panel who are observed at least twice. We eliminate individuals observed only once, which corresponds to about 8 percent of each panel. These restrictions left 4, 589 families in our sample.

DEPENDENT VARIABLES

Our dependent variables include hours spent during the last week on work (the sum of hours in market and household or domestic work) and school.[7] For the time allocation analysis, we divide the sample into four groups, boys aged 12 to 17, girls aged 12 to 17, men aged 18 to 65, and women aged 18 to 65. Overall educational attainment, while growing rapidly, remains low—the average level of education of men and women 15 and over was 7.5 and 7.0 years, respectively, in 1995. By the age of 17, the majority of children, even in urban areas, are no longer enrolled in school. Appendix Table 8 shows overall participation rates of and weekly hours spent by boys and girls aged 12 to 17 in school, market work, and domestic work. The table makes evident the large decreases in school attendance that begin by age 12 to 13 and the consequent increases in market work and household work for both girls and boys.

WAGES AND SHOCKS

One complication associated with the estimation of equation (14) is that the wage data for individuals who work at home but not in the market may not be available, as is the case for most of the adult women and children in our sample. To address the problem of missing wages, we adopt the assumption that market and domestic work are perfect substitutes in the utility function of the household. Then the market wage rate for labor of

[7] Prior to applying the logarithmic transformation, we transformed the hours spent on school or work by adding one to each of these variables. This way individual observations with zero hours, particularly on schooling, are not left out of the analysis.

each labor type provides a measure of the value of time of each family member irrespective of whether they work at home or in the market.[8]

Yet another problem is associated with finding suitable instruments for predicting the changes in wages. Even if the wage data were available, one needs to consider the possibility that the change in observed wages might contain unanticipated components that may, in turn, be correlated with the term $\eta(i,t)$, representing innovations in the marginal utility of wealth. In order to instrument individual wage changes, we follow the influential paper of MaCurdy (1981) and regress the period-to-period change in the logarithm of observed individual wages on human capital variables (such as years of education and marital status) and their interactions with age and age squared. There are two main advantages offered by this method. First, these variables are uncorrelated with the term $\eta(i,t)$, representing innovations in the marginal utility of wealth. Second, the differencing of the wage data removes all unobservable fixed effects from the error term of the wage regression. The latter minimizes the need to correct for the role of selection bias in the wage regression.[9] However, one shortcoming is that human capital variables alone are typically poor predictors of the changes in individual wages. One plausible way of improving the explanatory power of the first-stage regression is to include in the list of

[8] Under the alternative assumption of heterogeneity of preferences between market and domestic work, the opportunity cost of time is given by the "shadow wage rate," which is determined endogenously by the household's equilibrium choices (e.g., Skoufias 1994). Although in principle it is possible to estimate shadow wage rates from the estimates of a household or agricultural production function, this possibility is precluded by the urban context and the quarterly frequency of the data.

[9] Heckman and MaCurdy (1980), for example, report that correcting for selection bias in a wage regression, after removing individual fixed effects, did not yield any significant changes in the coefficients of a wage regression.

instruments variables proxying changes in the demand for labor, such as the *lagged* first differences in the logarithm of the unemployment rate in the local metropolitan area, and its interaction with the human capital variables (Ham 1986).[10]

Given our primary interest on the role of labor market shocks rather than the estimation of the intertemporal elasticity of substitution of labor supply, we chose to include the local unemployment rate directly as an explanatory variable in the individual time allocation regressions instead of using it (and its interactions with human capital variables) as a predictor of the changes in individual wages. Thus in contrast to the majority of the intertemporal labor supply studies that treat $\eta(i,t)$ as an unobservable, our approach consists of treating $\eta(i,t)$ as an observable in the time allocation regressions.

Although this alternative carries the cost of relying on rather weak instruments to predict individual wage changes, it also offers some compensating advantages. As long as unobserved taste factors are not correlated across individuals, the change in the local unemployment rate provides an "exogenous" measure of the shock to the marginal utility of wealth. In contrast, other "shock" variables, such as the loss of job by the household head or spouse, are likely to be correlated with an individual's unobserved components of tastes $\Delta\varepsilon(i,j,t)$ (e.g., Ham 1986). Workers who are laid off, for example, may also have

[10] The lagged first differences in the unemployment rate are used instead of the current first differences so to avoid any potential correlation with the $\eta(i,t)$ term.

stronger unobserved preferences for leisure (or distaste for work).[11] An additional advantage of this alternative is that it also allows us to examine whether individual time use is insured against idiosyncratic shocks (e.g., Ham and Jacobs 2000). In principle, the forecast errors $\eta(i,t)$ may contain both idiosyncratic as well as aggregate components that are common across the forecast errors of all households. According to the recent literature of complete risk-sharing and consumption insurance (Cochrane 1991; Townsend 1994), individuals or communities may devise a set of institutions or contracts, formal or otherwise, that allows them to fully diversify idiosyncratic risk. In these circumstances, it is only aggregate or uninsurable risk that matters in determining individual changes in time allocation (or consumption).[12] Put differently, with complete risk-sharing, household idiosyncratic shocks (whether anticipated or not) will have no significant effect on the time allocation of individuals within an insurance community. In terms of the notation used above, the complete markets hypothesis implies that the term $\eta(i,t)$ simply varies across time and not across households, i.e.,

$$\eta(i,t) = \eta(t),$$

[11] It should be noted that in the earlier version of this paper, we also experimented with constructing alternative measures of shocks, separately for household heads and their spouse/companion, defined according to whether they reported having lost their job within the past three months. Job loss includes those who report that the reason they are not working is due to (1) layoffs at their firm, (2) the firm moved, or (3) work was temporary. As pointed out by various reviewers, these shocks are likely to be correlated with unobserved tastes and are rather difficult to instrument, given that these are infrequent events in most samples.

[12] See Attanasio and Szekely (2001) for an investigation of the extent to which Mexican households are able to insure their consumption from fluctuations in their wages.

denoting that only aggregate shocks have an effect on changes in the time allocation of individual members of an insurance group.[13]

Given the preceding considerations, equation (15) was estimated in the following manner. We first instrumented the changes in real wages[14] for adult males, adult females, and children in our sample, using the parameters from a first-stage (or instrumental variable) regression that includes the years of education and marital status and the interactions of these two variables with age and age squared. We chose to use the initial rather than the current values for years of education or age or marital status for each individual in the sample in order to minimize the potential correlation of these variables with the unobservable variables affecting tastes in the time allocation equation (15).[15] The estimated coefficients of these first-stage regressions predicting real wage changes are presented in Appendix Table 9.[16] It should be noted that these regressions use the contemporaneous sample of individuals who were not in our panel but were interviewed at least twice over the same period.

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[13] See Altug and Miller (1990) for a formal derivation of this important result and Jacoby and Skoufias (1998) for a more detailed presentation of the differences between permanent income and complete markets hypotheses in explaining the consumption behavior of households.

[14] The nominal wage for each individual in any given quarter was deflated by the simple average of the value of the national consumer price index (CPI) prevailing in each the three months of the quarter. The monthly CPI was obtained from the INEGI website.

[15] We use the term "initial" for the values of these variables in the first observation for each individual. For children and younger adults, for example, it is possible that the years of education may change from one period to another as an individual completes a grade and moves on to the next. To minimize the potential correlation between the current level of education and the unobserved components of tastes summarized by the term $\Delta\varepsilon$ in the time allocation equation (15), we used the years of education of an individual in the first, rather than the current, round. We thank Hanan Jacoby for this suggestion.

[16] Note that we did not include any time dummies in the list of instruments.

Next, using the estimated coefficients of the wage functions for adult males, adult females, and children, we derived predicted values for wage changes for individuals in our sample. Our measure of shock to the marginal utility of wealth, summarized by $\eta(i,t)$, is the local metropolitan area growth rate in the male and female unemployment corresponding to the age of the male household head and his spouse. Specifically, INEGI calculates and makes publicly available the unemployment rates separately for males and females for different age groups (12–19, 20–24, 25–34, 35–44 and 45 and older), for each of the main metropolitan areas of Mexico. These unemployment rates are calculated on a quarterly basis using the full sample of households covered by the ENEU survey in each quarter and not just the subsample of households in our panel.[17] Using these data we constructed two variables characterizing household-specific (or idiosyncratic) shocks related to unemployment. One variable was based on the age of the male household head, and the second was based on the age of his spouse.[18] Unlike Ham and Jacobs (2000), who use the *current level* of the unemployment rate in the head's industry and the head's occupation, we use the *change in the logarithm* of the unemployment rate in the metropolitan area of the household, specific to the age and gender of the household head and spouse.

In addition to the growth in the local unemployment rate that captures idiosyncratic shocks to the household, we used a set of dummy variables for each quarter

[17] It should be noted that the calculation of the unemployment rate includes the incidence of job loss among individuals included in our panel, which constitutes 20 percent of the households surveyed in each quarter.

[18] In the rare cases where a male household head was less than 20 years of age, we used the unemployment rate for 20–24-year-old males.

in the panel to control for the role aggregate shocks in the sample. These time dummies also absorb potential changes in the real interest rate $r(t)$ faced by households. An implicit assumption maintained at least in the initial part of our empirical analysis is that households do not face any binding credit constraints. Given the high frequency of our time observations in the panel, this assumption may be more palatable. In the later part of the paper, the sensitivity of the results to the assumption of a perfect credit market are examined in more detail.

The standard errors reported were obtained using robust methods by taking into account the potential correlation of the unobserved components of tastes of individuals within a household as well as over time within a household.[19] Using White's correction for arbitrary forms of heteroskedasticity—or clustering on the metropolitan area of the household or on a variable identifying a individual rather than a household—did not lead to any remarkable differences from the standard errors and significant results reported here.

4. DISCUSSION OF RESULTS

Based on the panel of households observed between 1995:Q2 and 1996:Q2, a period of economic crisis and adjustment, Table 1 reports the estimated coefficients, standard errors, and associated test statistics of the idiosyncratic shocks on individual

[19] The robust standard errors were calculated by clustering on a variable identifying a household.

Table 1—The effects of shocks on individual hours of work and schooling during an economic crisis: 1995:Q2-1996:Q2

	1				2			
	Hours of work (home + market)				Hours of schooling			
	Coefficient	Standard error	t-value	p-value	Coefficient	Standard error	t-value	p-value
Adult males (18-65 years old)								
Change in:								
ln(local male unemployment rate)	-0.0309	0.020	-1.55	0.12				
ln(local female unemployment rate)	0.0172	0.009	1.91	0.06				
Adult females (18-65 years old)								
Change in:								
ln(local male unemployment rate)	0.0183	0.010	1.81	0.07				
ln(local female unemployment rate)	0.0131	0.005	2.82	0.01				
Boys (12-17 years old)								
Change in:								
ln(local male unemployment rate)	-0.0106	0.034	-0.31	0.757	0.0447	0.050	0.89	0.37
ln(local female unemployment rate)	0.0259	0.015	1.71	0.088	0.0110	0.025	0.44	0.66
Girls (12-17 years old)								
Change in:								
ln(local male unemployment rate)	0.0655	0.031	2.10	0.04	0.0008	0.054	0.01	0.99
ln(local female unemployment rate)	0.0054	0.014	0.38	0.70	0.0383	0.027	1.44	0.15
Nobs (NxT)		71,424				17,291		
R-squared		0.0040				0.2225		
F-value		3.99				156.89		
p-value		0.000				0.000		

Notes: The estimates in column 1 are based on the sample of adults and children from households with a male household head, a female spouse, and at least one child between 12 and 17 years of age. The estimates in column 2 are based on the previous sample children conditional on reporting initial positive hours of schooling. The standard errors reported are calculated using robust methods. See text for more details on additional regressors included and Appendix 3 for estimates of their coefficients.

hours of work (measured as the sum of hours in market and domestic activities) and hours devoted to school-related activities. The full set of estimated coefficients is presented in Appendix Table 10. Own-wage elasticities are rarely significantly different from zero, suggesting that the opportunity cost of time may play only a small role in the allocation of work across time, especially for adult males and children. Females are a notable exception, with an own-wage elasticity for labor supply of 0.85, which is closer to the high end of the range of elasticity estimates for females in the United States (Blundell and MaCurdy 1999).

The shock estimates reported are obtained by placing no restrictions on the cross-wage effects. We have also examined the sensitivity of the estimates by restricting the cross-wage effects for each labor type to zero and found that these restrictions made no difference in the results. The regressions for both adults and children include age and age squared and a dummy variable identifying whether the individual is a son or daughter of the household head, so as to distinguish any potential differences between individuals who simply coreside in the household (like nephews and nieces) and sons and daughters.

We begin with the estimates for adult males. The growth rate in the male unemployment rate appears to have a negative effect on the hours of work of adult males, while the growth rate in the female unemployment rate has a positive and significant effect. In combination, the significance of these coefficients implies that adult male labor supply is not insured from idiosyncratic shocks. As implied by the theoretical model presented earlier, in the absence of any binding constraints on work hours or any fixed costs associated with working, negative shocks on the marginal utility of wealth of the

household, represented here by the growth rate in the unemployment rate for males in the age group of the household head, should have a positive effect on labor supply. Although not significant (p-value = 0.12), the negative coefficient on this shock variable suggests that males are constrained in their total hours of work. Apparently, adult males are not able to fully compensate the loss of their time from wage-earning activities with work in other household activities, and as a result, they end-up consuming (unwillingly) more leisure than they would like. This provides further confirmation to the earlier finding by Ham (1986) for American adult males that unemployment is not a form of intertemporal labor supply behavior. In contrast, the significant and positive coefficient of the growth rate in female unemployment confirm that the "added-worker effect" is in operation as males attempt to make-up for this negative shock through increased hours of work.

The significantly positive coefficients of the unemployment shocks on adult female work hours imply that females are not insured from idiosyncratic shocks like increases in the local male and female unemployment rate. As predicted by the unitary model of intertemporal labor supply, increases in either the male or female unemployment rate lead to positive responses in female hours of work. In addition, the positive effect of the female unemployment rate on the hours of work of adult females suggests that, unlike adult males, females are able to more than compensate for lost hours in market work through work at home.

The significant coefficients of the unemployment variables on the hours of work of boys and girls suggest that their work hours are also not insured from idiosyncratic shocks. Specifically, the hours of work by boys appear to be unaffected by changes in the

male unemployment rate but positively affected by changes in the female unemployment rate. In contrast, a higher male unemployment rate leads to more hours of work by girls. As was the case for adults, the positive coefficients of these negative idiosyncratic shocks on hours of work of children are generally consistent with the prediction of the unitary model of family labor supply. Increases in unemployment for adult males (females) lead to added worker effects not only for adult females (males) but also for children.

To examine further the potential implications of these household idiosyncratic shocks on time children devote to schooling, we have also estimated a regression equation similar to equation (15) using the log of weekly hours devoted to schooling as the dependent variable. These estimates are presented in column 2 of Table 1 and are based on the subsample of children with positive reported hours of schooling in their first observation. Somewhat surprisingly, the unemployment variables do not appear to have a significant effect at conventional significance levels on the hours either boys or girls devote to schooling. These results suggest that children's school attendance, at least in the short-run, is insulated from idiosyncratic shocks to households. Thus the positive effect of these same shocks on the work hours of both boys and girls takes place at the expense of children's leisure rather than schooling. This also suggests that schooling and work activities, at least in the short run, are complementary rather that competing activities.[20]

[20] Interestingly, Skoufias and Parker (2001) find that school and work activities are competing activities for children in rural areas of Mexico.

One response to economic shocks may be adjustments in family size; for instance, a family member may migrate elsewhere looking for work. While our empirical model does not explicitly allow family size to be endogenous, we also made an effort to eliminate the role of adjustments in family size and composition from influencing estimates. We estimated labor market shock regressions on two different samples: first, for the full sample of individuals, and second, for the sample of individuals in households where there was no change in family size during the five quarters. Only about 15 percent of all individuals live in a household where there was a change in family size over the five quarters. To check whether our results so far are contaminated from ex-post adjustments in family size, we have also replicated the analysis restricting the sample to individual observations from households where there was no change in family size. The changes in the estimates results were not substantial enough to warrant any changes in the general conclusions and thus are not reported.

CREDIT MARKET CONSTRAINTS

While generally consistent with the predictions of the simple unitary model of intertemporal time allocation, the estimates presented so far were based on the assumption of a perfect credit market. In the presence of borrowing constraints, the Euler condition on the intertemporal allocation of consumption and hence leisure may be violated (Zeldes 1989). Poor households, for example, who presumably have fewer assets

and are more likely to be credit constrained, may exhibit larger responses to negative shocks affecting their marginal utility of wealth. [21]

To shed more light on the issue of differential access to credit, we have also examined whether the impact of the labor market shocks varies depending on the poverty status of the household. Given that our data do not allow us to construct a permanent income or consumption-based measure of poverty, we generated two different binary variables classifying households as poor or not. Both of these variables were constructed using one or more variables that are typically strongly correlated with long-term poverty.[22] Households with heads that have a lower level of education, for example, are also more likely to have a lower permanent income. Therefore, our first poverty status variable classifies a household as poor if the head of the household has, at most, nine years of education (equivalent to a junior high school level of education). The other poverty status variable is based on the quality of material used in the construction of the household residence and its access to basic services. Specifically, a household was classified as poor if either the roof, walls, or floor were made of low-quality material (such as a floor made of dirt, laminate walls or roof, etc.) or if the residence did not have plumbing or access to electricity, water, or telephone services.

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[21] As Zeldes (1989) demonstrates, the borrowing constraint has a shadow price that enters in the Euler equation for the marginal utility of consumption (equation 12), and acts as if it increases the effective interest rate faced by the household.

[22] We did not use the initial or current level of reported household income, since this is already likely to be negatively affected by the economic crisis that began in December 1994.

Next, we reestimated the regression equation for work hours of adults and children, allowing the coefficients of the two shock variables to differ between poor and nonpoor households. As mentioned above, in the presence of credit constraints, poorer families that are more likely to be affected by such constraints may respond to shocks differently than families where such constraints are not effective. These estimates are presented in Table 2. Although far from the ideal method of measuring poverty, our two poverty status variables yield results that are generally consistent with each other, and both are consistent with the role of effective credit constraints. For example, irrespective of the definition of poverty used, adult females in poorer households increase their work hours in response to a higher growth rate in the male unemployment rate. Moreover, their response to the shock from a higher growth rate in male unemployment is greater than that resulting from an increase in the adult female unemployment growth rate. In contrast, the male unemployment rate does not lead to a significant "added-worker effect" on the work hours of adult females in wealthier households. Yet, at the same time, in wealthier households, adult females do seem to increase their work hours when the female unemployment rate increases.

Overall the same general picture emerges from the differences in the coefficients between poorer and wealthier households in the work hours of adult males and children. In most cases, the coefficients of the shock variables are not significantly different from zero for the work hours of adult males and children in wealthier households, suggesting that their time allocation is insured from idiosyncratic shocks. Positive and significant

Table 2—The effects of shocks on individual hours of work (home + market), by household poverty status during an economic crisis: 1995:Q2-1996:Q2

	1				2			
	Household poverty status defined based on head's years of schooling				Household poverty status defined based on quality and access to services			
	Coefficient	Standard error	t-value	p-value	Coefficient	Standard error	t-value	p-value
Adult males (18-65 years old)								
Poor household	0.0022	0.012	0.18	0.88	-0.0035	0.012	-0.29	0.77
MUR x nonpoor household	-0.0009	0.048	-0.02	0.99	-0.0300	0.057	-0.52	0.60
FUR x nonpoor household	0.0153	0.022	0.69	0.49	-0.0177	0.026	-0.68	0.50
MUR x poor household	-0.0191	0.024	-0.79	0.43	-0.0138	0.023	-0.60	0.55
FUR x poor household	0.0145	0.010	1.43	0.15	0.0218	0.010	2.23	0.03
Adult females (18-65 years old)								
Poor household	0.0006	0.007	0.09	0.93	-0.0108	0.007	-1.57	0.12
MUR x nonpoor household	0.0347	0.026	1.35	0.18	0.0245	0.027	0.92	0.36
FUR x nonpoor household	0.0231	0.011	2.03	0.04	0.0104	0.011	0.96	0.34
MUR x poor household	0.0297	0.012	2.51	0.01	0.0336	0.012	2.91	0.00
FUR x poor household	0.0118	0.005	2.25	0.02	0.0146	0.005	2.75	0.01
Boys (12-17 years old)								
Poor household	0.0223	0.020	1.12	0.26	-0.0296	0.021	-1.40	0.16
MUR x nonpoor household	-0.0420	0.082	-0.51	0.61	0.0401	0.091	0.44	0.66
FUR x nonpoor household	0.0301	0.039	0.76	0.45	0.0081	0.041	0.20	0.84
MUR x poor household	0.0220	0.042	0.53	0.60	-0.0249	0.040	-0.63	0.53
FUR x poor household	0.0249	0.017	1.45	0.15	0.0259	0.017	1.54	0.12
Girls (12-17 years old)								
Poor household	0.0108	0.019	0.58	0.56	-0.0125	0.020	-0.63	0.53
MUR x nonpoor household	0.1224	0.098	1.25	0.21	0.1086	0.099	1.09	0.27
FUR x nonpoor household	0.0242	0.030	0.80	0.42	-0.0193	0.042	-0.45	0.65
MUR x poor household	0.0587	0.034	1.73	0.08	0.0403	0.034	1.17	0.24
FUR x poor household	-0.0017	0.017	-0.10	0.92	0.0069	0.016	0.44	0.66
Nobs (NxT)		71,424				71,424		
R-squared		0.0040				0.0041		
F-value		3.31				3.39		
p-value		0.000				0.000		

Notes: MUR denotes the growth rate (change in logarithm) of the local male unemployment rate; FUR denotes the growth rate (change in logarithm) of the local female unemployment rate.

coefficients are encountered for adult males' work hours in poorer households in response to increases in the female unemployment rate and for the work hours of girls in poorer households responding positively to an increase in the male unemployment. The absence of significant added worker effects in the work hours of boys in poorer households suggests that boys may be better protected from these shocks relative to girls.

We have also estimated similar regressions for the hours children devote to school and school-related activities. These estimates are reported in Table 3. Consistent with the estimates in Table 1, allowing the impact of the shock to differ according to the poverty status of the household did not result in any remarkable differences. In general, children from poorer households devoted fewer hours to schooling as evidenced by the negative and strongly significant coefficient of the dummy variable identifying poor households. As in Table 1, the hours children devote to school from wealthier as well as poorer households continue to be independent of the shocks, suggesting that school attendance is insured from idiosyncratic shocks even for children in poor households. Thus the increased work time by adult women in poor households appears to be adequate, at least in the short run, for preventing these shocks from having an adverse effect on the weekly school attendance of children.

Yet, the effects of these shocks on children's work hours raise doubts about whether this situation can be sustained for extensive periods.[23] Poorer families, for

[23] Hours devoted to schooling are an input in the human capital production function, and the shock may also have a direct effect on the stock of human capital independent of its effect on child school hours.

Table 3—The effects of shocks on children's schooling hours, by household poverty status during an economic crisis: 1995:Q2-1996:Q2

	1				2			
	Household poverty status defined based on head's years of schooling				Household poverty status defined based on quality and access to services			
	Coefficient	Standard error	t-value	p-value	Coefficient	Standard error	t-value	p-value
Boys (12-17 years old)								
Poor household	-0.0662	0.020	-3.37	0.00	-0.0810	0.019	-4.37	0.00
MUR x nonpoor household	0.0144	0.143	0.10	0.92	0.0082	0.176	0.05	0.96
FUR x nonpoor household	0.0391	0.052	0.75	0.46	-0.0073	0.065	-0.11	0.91
MUR x poor household	0.0569	0.059	0.96	0.34	0.0828	0.056	1.47	0.14
FUR x poor household	0.0200	0.030	0.68	0.50	0.0205	0.028	0.73	0.47
Girls (12-17 years old)								
Poor household	-0.0729	0.021	-3.46	0.00	-0.0744	0.021	-3.50	0.00
MUR x nonpoor household	-0.1458	0.154	-0.95	0.34	-0.0326	0.163	-0.20	0.84
FUR x nonpoor household	0.0191	0.065	0.30	0.77	0.0970	0.078	1.24	0.22
MUR x poor household	-0.0156	0.062	-0.25	0.80	0.0150	0.062	0.24	0.81
FUR x poor household	0.0411	0.031	1.32	0.19	0.0297	0.029	1.02	0.31
Nobs (NxT)	17,291				17,291			
R-squared	0.2228				0.2232			
F-value	126.65				1,326.48			
p-value	0.000				0.000			

Notes: MUR denotes the growth rate (change in logarithm) of the local male unemployment rate; FUR denotes the growth rate (change in logarithm) of the local female unemployment rate.

example, may manage to keep their children in school until they complete the current grade but be forced to withdraw the children at the start of the next academic year. For this reason, we also examine whether the same labor market shocks are associated with the continuation of schooling during the next academic year. Although the survey does not contain an explicit question on whether a child is currently enrolled in and attending school in each quarter, we use the time allocated to school in the second quarter of the calendar year and the second quarter of the next calendar year, thus spanning two school years, to construct a binary variable indicating whether the child reports positive hours in school in each of these quarters. Limiting our sample to children with positive hours of schooling in the first quarter, we then examine whether in this sample of children the shocks have a significant effect on the likelihood of dropping out. Unlike the earlier regressions, this regression is estimated as a reduced form using individual, household, and dwelling characteristics at the initial (second) quarter of the panel.

The main individual characteristics included the age (and age squared) of the child, while the household characteristics included the composition of the household by age and gender group, the age of the household head and his spouse, the education level of the household head, and the predicted wage rate for adult males, adult females, and children in the household.[24] The variables characterizing the household dwelling included the number of rooms in the residence, whether it is rented or not, whether the roof, floors,

[24] The predicted wages for adults and children were constructed from separate regressions (each one corrected for possible selection bias using Heckman's (1979) two-step method) for adult males, adult females, and children.

or walls were constructed with low-quality materials, and the availability of some basic services. We also included a set of dummy variables for each of the metropolitan areas of the household's residence to account for fixed effects at the metropolitan area level. These fixed effects also control for potential differences in the access and quality of schooling facilities across cities. Our measure of the idiosyncratic shock is the five-quarter average of the growth rate in the local unemployment rate corresponding to the age of the household head and his spouse. Given the earlier indications that credit constraints play an important role on the intertemporal time allocation of members in poorer households, we have also interacted our two shock variables with the years of education of the household head.

Table 4 contains the marginal effects of the explanatory variables estimated from a probit specification of the probability of dropping out of school.[25] As can be seen, the higher male and female unemployment rates increase the probability that a child will drop out of school. Thus, although the earlier results in Tables 1–3 imply that the school hours of both boys and girls are insulated from these shocks, we now have evidence that these efforts to protect children's investments in human capital are not completely effective when it comes to children continuing school next year. A similar finding for Mexican households is also reported by Attanasio and Szekely (2001), who find that in

[25] Another potential effect of shocks is that while they may not result in immediate dropout, they may result in higher repetition or failure rates, which are then likely to reduce overall lifetime educational achievement. To address this question, we also estimated probit models of the probability of advancing to the next grade and the probability of repeating the same grade using the sample of children in school in 1995:Q2 and in 1996:Q2. We found no impacts on these schooling attainment indicators.

response to temporary shocks, households contract consumption expenditures on health and education representing longer-run investment in human capital.

Table 4—The impact of shocks on the probability of dropping out of school during an economic crisis: 1995:Q2-1996:Q2

	Boys (12-17 years old)			Girls (12-17 years old)		
	Coefficient	z-value	p-value	Coefficient	z-value	p-value
Aver. growth in local unemp. for males (MUR)	0.4728	4.96	0.00	0.7080	6.50	0.000
Aver. growth in local unemp. for females (FUR)	0.0967	2.39	0.02	0.1274	3.17	0.002
MUR x years of schooling of household head	-0.0036	-1.32	0.19	-0.0081	-2.58	0.010
FUR x years of schooling of household head	-0.0015	-1.26	0.21	-0.0024	-1.81	0.071
Years of schooling of household head	-0.0091	-1.85	0.07	0.0042	0.82	0.415
Age of household head	-0.0010	-0.48	0.63	0.0021	0.94	0.345
Age of spouse of household head	-0.0019	-0.76	0.44	-0.0031	-1.24	0.214
Predicted adult male wage	0.0751	1.35	0.18	-0.0171	-0.30	0.766
Predicted adult female wage	-0.0271	-1.02	0.31	-0.0617	-2.17	0.030
Predicted child wage	-0.1495	-0.93	0.35	-0.1751	-1.07	0.284
Age of child	0.1200	1.10	0.27	0.0680	0.58	0.561
Age of child squared/100	-0.2663	-0.70	0.48	-0.1007	-0.25	0.802
No. of 18-65 year old males in the household	0.0208	1.23	0.22	0.0147	0.83	0.405
No. of 18-65 year old females in the household	0.0112	0.73	0.46	-0.0168	-1.05	0.292
No. of 66+ year old males in the household	-0.1827	-2.08	0.04	0.0436	0.49	0.624
No. of 66+ year old females in the household	-0.0247	-0.40	0.69	0.0860	1.27	0.204
No. of 12-17 year old males in the household	0.0046	0.29	0.77	0.0283	1.60	0.109
No. of 12-17 year old females in the household	0.0122	0.77	0.44	0.0169	0.98	0.329
No. of 6-11 year old boys in the household	0.0101	0.65	0.52	0.0170	1.02	0.310
No. of 6-11 year old girls in the household	0.0256	1.70	0.09	-0.0084	-0.53	0.599
No. of 0-5 year old boys in the household	0.0040	0.19	0.85	0.0327	1.53	0.126
No. of 0-5 year old girls in the household	0.0201	0.94	0.35	0.0101	0.46	0.645
Residence is rented? (1=yes, 0=no)	0.1758	4.83	0.00	0.2175	5.77	0.000
Residence has no kitchen? (1=yes, 0=no)	0.0235	0.66	0.51	0.0012	0.03	0.975
Number of rooms in residence	-0.0229	-3.27	0.00	-0.0177	-2.50	0.013
Roof made of low quality material? (1=yes, 0=no)	-0.1199	-1.49	0.14	-0.0371	-0.38	0.701
Wall made of low quality material? (1=yes, 0=no)	-0.0605	-1.79	0.07	0.0039	0.10	0.919
Floor made of low quality material? (1=yes, 0=no)	0.0048	0.19	0.85	-0.0395	-1.53	0.126
Residence has no electricity? (1=yes, 0=no)	0.0057	0.04	0.97	-0.0967	-0.64	0.524
Residence has no water? (1=yes, 0=no)	0.1003	1.84	0.07	-0.0414	-0.70	0.484
Residence has no plumbing? (1=yes, 0=no)	0.0142	0.29	0.77	0.1870	3.45	0.001
Residence has no telephone? (1=yes, 0=no)	0.0336	1.39	0.16	0.0540	2.10	0.036
Metropolitan area dummies included?		Yes			Yes	
Number of observations		2,416			2,265	
LR chi2 =		242.790			291.710	
Prob > chi2 =		0.000			0.000	
Pseudo R2 =		0.086			0.107	

Notes: The probit equations for Dropping Out of School were based on the sample of children with positive hours in school during the 1st quarter. Coefficients are the marginal effects of the respective shock on the probability.

The negative and significant coefficient in Table 4 suggests that children from households in which the head has more years of education are less likely not to be in school in the next school year. This finding is in accordance with that shown in Table 3, where it was reported that children in poorer households (where the head has low education) spend fewer hours in school. Significant differences are also apparent on the relative impact of the shocks on boys and girls. The marginal effects of either the male or female unemployment rate on the probability of dropping out of school is higher for girls than boys, suggesting that the shocks affect girls more than boys. The negative coefficients of the interactions of the male and female local unemployment growth rate with the years of education of the household head also suggest that the impact of the shocks on the probability of dropping out of school is smaller for girls in wealthier households. Thus, the adverse effects of the credit constraints on girls is greater in poorer households—as was also found earlier in Table 2 with respect to the work hours of girls in poor households. These findings suggest some differentiation within households based on the child's gender.

The results so far have been based on a panel of households from a period of economic crisis in Mexico. Depending on the general economic conditions prevailing, the same idiosyncratic shocks may convey very different information about the lifetime wealth of the household and therefore on the extent to which a household adjusts its marginal utility of wealth as a result of the shock experienced. In addition, credit constraints may be less "binding" during a period of economic growth. For this reason, we have also examined the potential impact of the same shocks during a period of

economic recovery. The estimates obtained by applying the exact same methods for constructing variables and predicting wages for the new panel of households during the 1998:Q2-1999:Q2 period, a period of economic recovery and lower levels of unemployment, are presented in Tables 5–7.

Overall these results are consistent with the interpretation that the information content of the same shock conveys different information about the lifetime opportunities of the households during periods of economic downturn and recovery. The strong added-worker effects observed on the work hours of adult females during the time of crisis seem to disappear during the period of recovery. This suggests that the same shock during a period of economic recovery does not lead to a significant updating of the marginal utility of wealth of the household. However, there continues to be weak evidence about the added-worker effect on the work hours of girls in poor households, as shown in Table 2. Also, the time allocated by children to school continues to be unaffected from these shocks, suggesting that children are insulated from economic shocks (compare Tables 3 and 6).

Finally an examination of the impact of the same shock variables on the probability that children drop out of school reveals that the same shocks now have a very different effect than at a time of crisis (see Table 7 and Table 5). Increases in the local female unemployment rate have a smaller marginal effect on the probability that boys drop out of school and no effect on whether girls drop out of school. At a time of crisis, the same shock had a significant and positive effect on the probability that girls drop out of school, and the marginal effect was higher for girls than boys.

Table 5—The effects of shocks on individual hours of work (home + market), by household poverty status during a period of economic growth: 1998:Q2-1999:Q2

	1				2			
	Household poverty status defined based on head's years of schooling				Household poverty status defined based on quality and access to services			
	Coefficient	Standard error	t-value	p-value	Coefficient	Standard error	t-value	p-value
Adult males (18-65 years old)								
Poor household	-0.0059	0.009	-0.65	0.52	0.0155	0.010	1.55	0.12
MUR x nonpoor household	-0.0042	0.020	-0.21	0.83	-0.0092	0.019	-0.48	0.64
FUR x nonpoor household	0.0056	0.014	0.39	0.69	0.0099	0.011	0.89	0.38
MUR x poor household	-0.0163	0.011	-1.49	0.14	-0.0073	0.022	-0.34	0.74
FUR x poor household	0.0001	0.007	0.01	0.99	-0.0130	0.013	-0.98	0.33
Adult females (18-65 years old)								
Poor household	-0.0059	0.006	-1.03	0.30	-0.0011	0.006	-0.17	0.86
MUR x nonpoor household	-0.0064	0.017	-0.38	0.71	0.0246	0.017	1.47	0.14
FUR x nonpoor household	0.0065	0.010	0.63	0.53	-0.0090	0.009	-1.04	0.30
MUR x poor household	0.0077	0.007	1.07	0.29	-0.0241	0.018	-1.34	0.18
FUR x poor household	-0.0024	0.004	-0.57	0.57	0.0111	0.010	1.16	0.25
Boys (12-17 years old)								
Poor household	0.0201	0.017	1.16	0.25	0.0291	0.019	1.52	0.13
MUR x nonpoor household	0.0229	0.054	0.42	0.67	-0.0154	0.044	-0.35	0.73
FUR x nonpoor household	-0.0189	0.024	-0.78	0.44	0.0132	0.025	0.53	0.60
MUR x poor household	-0.0024	0.022	-0.11	0.91	0.0205	0.049	0.42	0.68
FUR x poor household	0.0020	0.014	0.14	0.89	-0.0170	0.028	-0.60	0.55
Girls (12-17 years old)								
Poor household	-0.0064	0.015	-0.42	0.68	-0.0013	0.018	-0.08	0.94
MUR x nonpoor household	-0.0743	0.042	-1.78	0.08	-0.0117	0.037	-0.31	0.75
FUR x nonpoor household	-0.0182	0.022	-0.81	0.42	-0.0319	0.023	-1.38	0.17
MUR x poor household	-0.0015	0.018	-0.08	0.93	-0.0036	0.041	-0.09	0.93
FUR x poor household	0.0200	0.012	1.66	0.10	0.0506	0.026	1.97	0.05
Nobs (NxT)		70,275				70,275		
R-squared		0.0036				0.0036		
F-value		2.62				2.70		
p-value		0.000				0.000		

Notes: MUR denotes the growth rate (change in logarithm) of the local male unemployment rate; FUR denotes the growth rate (change in logarithm) of the local female unemployment rate.

Table 6—The effects of shocks on children's schooling hours, by household poverty status during a period of economic growth: 1998:Q2-1999:Q2

	1				2			
	Household poverty status defined based on head's years of schooling				Household poverty status defined based on quality and access to services			
	Coefficient	Standard error	t-value	p-value	Coefficient	Standard error	t-value	p-value
Boys (12-17 years old)								
Poor household	-0.1113	0.017	-6.56	0.00	-0.0829	0.019	-4.44	0.00
MUR x nonpoor household	0.0406	0.074	0.55	0.58	0.0870	0.054	1.60	0.11
FUR x nonpoor household	-0.0276	0.032	-0.87	0.38	-0.0270	0.030	-0.91	0.36
MUR x poor household	-0.0245	0.033	-0.74	0.46	-0.1258	0.063	-1.99	0.05
FUR x poor household	-0.0091	0.018	-0.50	0.62	0.0070	0.034	0.20	0.84
Girls (12-17 years old)								
Poor household	-0.0898	0.017	-5.18	0.00	-0.0723	0.019	-3.72	0.00
MUR x nonpoor household	0.0123	0.070	0.18	0.86	-0.0346	0.058	-0.60	0.55
FUR x nonpoor household	-0.0226	0.031	-0.73	0.47	0.0175	0.031	0.57	0.57
MUR x poor household	0.0463	0.036	1.29	0.20	0.0935	0.067	1.40	0.16
FUR x poor household	-0.0251	0.018	-1.42	0.16	-0.0511	0.035	-1.47	0.14
Nobs (NxT)	17,320				17,291			
R-squared	0.3131				0.2232			
F-value	132.26				1,326.48			
p-value	0.000				0.000			

Notes: MUR denotes the growth rate (change in logarithm) of the local male unemployment rate; FUR denotes the growth rate (change in logarithm) of the local female unemployment rate.

Table 7—The impact of shocks on the probability of dropping out of school and grade advancement during a period of economic growth: 1998:Q2-1999:Q2

	Boys (12-17 years old)			Girls (12-17 years old)		
	Coefficient	z-value	p-value	Coefficient	z-value	p-value
Growth in local unemployment for males (MUR)	-0.1598	-3.00	0.00	-0.1918	-3.87	0.00
Growth in local unemployment for females (FUR)	0.0638	2.19	0.03	0.0367	1.25	0.21
MUR x years of schooling of household head	0.0023	1.16	0.25	0.0006	0.30	0.77
FUR x years of schooling of household head	-0.0010	-1.27	0.20	0.0001	0.10	0.92
Years of schooling of household head	0.0017	0.34	0.73	0.0011	0.21	0.83
Age of household head	-0.0009	-0.41	0.68	0.0006	0.27	0.79
Age of spouse of household head	0.0017	0.69	0.49	-0.0015	-0.64	0.53
Predicted adult male wage	-0.0718	-1.27	0.20	-0.0513	-0.91	0.36
Predicted adult female wage	-0.0338	-1.38	0.17	0.0339	1.40	0.16
Predicted child wage	-0.0940	-1.07	0.29	0.0980	0.92	0.36
Age of child	0.1261	1.13	0.26	-0.0551	-0.50	0.62
Age of child squared/100	-0.2769	-0.72	0.47	0.3091	0.81	0.42
No. of 18-65 year old males in the household	0.0031	0.17	0.87	0.0172	1.00	0.32
No. of 18-65 year old females in the household	0.0190	1.16	0.24	0.0122	0.75	0.46
No. of 66+ year old males in the household	-0.1555	-1.67	0.10	-0.0031	-0.04	0.97
No. of 66+ year old females in the household	-0.0007	-0.01	0.99	-0.2034	-3.10	0.00
No. of 12-17 year old males in the household	0.0078	0.46	0.65	-0.0221	-1.38	0.17
No. of 12-17 year old females in the household	-0.0234	-1.24	0.22	0.0430	2.56	0.01
No. of 6-11 year old boys in the household	-0.0059	-0.35	0.72	0.0140	0.86	0.39
No. of 6-11 year old girls in the household	0.0020	0.12	0.90	0.0244	1.50	0.14
No. of 0-5 year old boys in the household	0.0551	2.63	0.01	0.0159	0.74	0.46
No. of 0-5 year old girls in the household	-0.0225	-0.97	0.33	0.0501	2.35	0.02
Residence is rented? (1=yes, 0=no)	0.2018	5.73	0.00	0.1071	3.14	0.00
Residence has no kitchen? (1=yes, 0=no)	-0.0312	-0.85	0.40	-0.0005	-0.01	0.99
Number of rooms in residence	-0.0090	-1.22	0.22	-0.0131	-1.74	0.08
Roof made of low quality material? (1=yes, 0=no)	0.1522	1.62	0.10	0.0484	0.53	0.60
Wall made of low quality material? (1=yes, 0=no)	-0.0844	-2.42	0.02	0.0457	1.22	0.22
Floor made of low quality material? (1=yes, 0=no)	0.0766	3.12	0.00	0.0133	0.56	0.57
Residence has no electricity? (1=yes, 0=no)	-0.2070	-1.56	0.12	0.0554	0.33	0.74
Residence has no water? (1=yes, 0=no)	0.0871	1.19	0.23	-0.0825	-1.47	0.14
Residence has no plumbing? (1=yes, 0=no)	0.1221	2.07	0.04	0.1454	2.40	0.02
Residence has no telephone? (1=yes, 0=no)	0.0002	0.01	0.99	0.0896	3.92	0.00
Metropolitan area dummies included?		Yes			Yes	
Number of observations		2,342			2,265	
LR chi2 =		225.570			291.710	
Prob > chi2 =		0.000			0.000	
Pseudo R2 =		0.082			0.107	

Notes: The probit equations for Dropping Out of School were based on the sample of children with positive hours in school during the 1st quarter. Coefficients are the marginal effects of the respective shock on the probability.

Also, the local unemployment rate does not have any significant interactions with the poverty status of the household (measured by the years of schooling of the household head), a finding consistent with the interpretation that credit constraints are less binding during a period of economic growth. In contrast to Table 4, the local male unemployment rate appears to have a negative effect on the probability of children dropping out of school. This is consistent with the countercyclical patterns observed by Dellas and Sakellaris (1995), but in combination with the effects observed during the period of crisis, it also suggests that the school enrollment of children may be countercyclical only during periods with a positive overall economic trend.

5. CONCLUDING REMARKS

Labor markets can simultaneously be a source of risk for households and a means of insuring against risk. In this paper we examined the extent to which labor market shocks such as changes in the local growth rate of male and female unemployment rates have an impact on the intertemporal allocation of time of adult and younger members in families. Labor market shocks affecting the marginal utility of wealth of a household can lead to increases in the work hours (added-worker effects) intended to compensate for the adverse economic conditions faced by the household. One critical question is whether these shocks and the accompanying added-worker responses take place only among adult members or also affect the hours children devote to work and school.

The evidence presented here, allowing for potential differences in responses due to household credit constraints, suggests that at a time of crisis. significant added-worker effects are in operation, especially for adult females of poorer or constrained households and in some cases children. Rather surprisingly, we find that the time that boys and girls devote to schooling is largely unaffected by market shocks. This finding implies that the increased work time by adult women in poor households appears to be adequate, at least in the short run, for preventing these shocks from having an adverse impact on the weekly school attendance of children. However, this does not turn out to be sufficient for protecting the human capital stock invested on children. When we examine the impact of the same shocks on whether children continue attending school in the next school year, we find that they increase significantly the probability that children do not continue school.

We also present evidence suggesting some differential treatment based on the sex of children with families in Mexico. It seems that efforts to protect family investments on children's human capital are more effective for boys than girls. Thus economic crises not only play a critical role in the intergenerational transmission of poverty but also tend to reinforce any gender-based preferences and inequalities.

Our results also highlight the fact that the same labor market shocks may have different impact depending on the macroeconomic conditions prevailing in the economy. Clearly, safety programs have a critical role to play and especially in periods of economic crisis. The evidence from our study suggests that safety net programs can become more

effective in protecting families if, by design, these programs are flexible enough to adjust based on the magnitude and the nature of the aggregate shock.

APPENDIX TABLES

Table 8—School, domestic, and market work, by age and gender

	Percent of individuals reporting nonzero hours last week in:				Mean hours (including zero values) devoted last week to:			
Age	**School**	**Domestic**	**Market**	**Work**	**School**	**Domestic**	**Market**	**Work**
1A Males								
12	84	64	7	67	27	6	2	8
13	81	63	10	68	27	7	2	9
14	76	60	16	69	26	7	5	11
15	66	60	23	72	23	7	8	15
16	56	57	30	75	19	7	12	18
17	48	52	41	77	16	6	17	23
18	41	50	48	79	14	6	20	26
19	36	49	55	82	12	6	24	29
20	30	47	62	84	10	5	27	33
21	28	49	65	85	10	6	29	34
22	22	47	70	87	8	5	31	36
23	19	46	75	89	6	5	34	40
24	14	47	78	91	5	6	36	42
25-34	5	49	87	94	1	6	41	47
35-44	2	48	88	95	0	6	42	47
45-65	1	47	76	91	0	6	35	41
1B: Females								
12	82	81	3	81	27	11	1	11
13	81	83	5	84	28	12	1	13
14	75	83	8	84	26	13	2	15
15	68	85	13	87	24	15	4	19
16	58	87	19	91	20	17	7	24
17	51	87	24	91	17	18	9	28
18	40	88	30	93	14	20	12	32
19	33	86	36	94	11	21	15	36
20	29	89	37	95	10	23	15	38
21	27	89	39	96	9	24	16	40
22	20	90	43	97	7	26	17	43
23	14	92	44	98	5	28	18	46
24	9	92	47	98	3	29	19	48
25-34	3	96	45	99	1	35	17	52
35-44	2	98	45	99	0	38	17	54
45-65	1	98	30	99	0	39	11	50

Source: ENEU 1995:Q2.
Note: Work includes domestic and market activities as well as other activities.

Table 9—Instrumental variable regressions for the changes in the log of real wages

	Adult males 18-65 years old		Adult females 18-65 years old		Children 12-17 years old	
	Coefficient	**t-value**	**Coefficient**	**t-value**	**Coefficient**	**t-value**
Years of education	0.0070	0.69	-0.0079	-0.60	-0.0366	-1.03
Age	-0.0017	-0.59	-0.0066	-1.81	-0.0518	-3.46
Age squared/100	0.0036	0.99	0.0075	1.51	0.0650	2.71
Education x Age	-0.0003	-0.51	0.0003	0.41	0.0031	1.07
Education x (age squared/100)	0.0005	0.71	-0.0002	-0.22	-0.0034	-0.72
Boy? (1=yes, 0=no)					-0.2203	-1.50
Boy x years of education					-0.0031	-0.38
Boy x age					0.0185	1.58
Boy x (age squared/100)					-0.0222	-1.17
Married? (1=yes, 0=no)	0.1198	2.72	-0.0458	-0.67		
Married x years of education	0.0017	0.60	-0.0010	-0.27		
Married x age	-0.0055	-2.27	0.0022	0.61		
Married x (age square/100)	0.0041	1.34	-0.0029	-0.63		
Constant term	-0.0074	-0.16	0.1063	1.78	0.6092	3.29
Dummies for metropolitan area included?	Yes		Yes		Yes	
Number of observations (NxT)	105,796		41,431		6,851	
F-value =	8.93		2.78		4.87	
Probability > F =	0.000		0.003		0.000	
R-squared =	0.0016		0.0019		0.0099	
Adjusted R-squared =	0.0011		0.0007		0.0027	
Root MSE =	0.707		0.665		0.611	

Notes: These regression estimates are based on the sample of individuals interviewed at least twice by the ENEU survey between 1995:Q2 and 1996:Q2.

Table 10—Estimates of the determinants of changes in work hours (equation 15)

	Adult males 18-65 years old		Adult females 18-65 years old		Boys 12-17 years old		Girls 12-17 years old	
	Coefficient	**t-value**	**Coefficient**	**t-value**	**Coefficient**	**t-value**	**Coefficient**	**t-value**
Constant term	-0.0399	-0.39	0.1140	1.95	0.2231	0.36	-0.5156	-0.96
Age	-0.0002	-0.04	-0.0046	-1.65	-0.0139	-0.16	0.1041	1.39
Age squared/100	0.0015	0.30	0.0057	1.64	0.0493	0.16	-0.3617	-1.40
Predicted change in log male wage	-0.3710	-0.73	0.0152	0.06	1.2816	1.72	-0.3576	-0.55
Predicted change in log female wage	0.2961	0.54	0.8525	2.14	-0.0908	-0.09	-0.5439	-0.48
Predicted change in log child wage	0.0435	0.27	-0.0082	-0.10	-0.5108	-1.34	-0.0280	-0.11
Local MUR	-0.0309	-1.55	0.0183	1.81	-0.0106	-0.31	0.0655	2.10
Local FUR	0.0172	1.91	0.0131	2.82	0.0259	1.71	0.0054	0.38
Dummy for 1995:Q4	0.0780	2.40	0.0126	0.71	-0.2047	-3.69	-0.2785	-5.95
Dummy for 1996:Q1	-0.0408	-1.56	0.0017	0.12	-0.2056	-4.78	-0.2013	-5.40
Dummy for 1996:Q2	0.0369	1.37	0.0111	0.80	-0.1531	-3.53	-0.1774	-4.84
Child of household head? (1=yes, 0=no)	0.0563	1.86	0.0050	0.34	0.0875	0.88	-0.0529	-1.09
Number of observations (NxT)	23,297		22,554		13,284		12,289	

Notes: The total number of households is 4,589. MUR (FUR) denote the growth rate (or change in log) of the local male (female) unemployment rate. T-values reported are based on robust standard errors (see text for details).

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