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ABSTRACT

This paper studies empirically the relationship between trade policy and individual income risk faced by workers, and uses the estimates of this empirical analysis to evaluate the welfare effect of trade reform. The analysis proceeds in three steps. First, longitudinal data on workers are used to estimate time-varying individual income risk parameters in various manufacturing sectors. Second, the estimated income risk parameters and data on trade barriers are used to analyze the relationship between trade policy and income risk. Finally, a simple dynamic incomplete-market model is used to assess the corresponding welfare costs. In the implementation of this methodology using Mexican data, we find that trade policy changes have a significant short run effect on income risk. Further, while the tariff level has an insignificant mean effect, it nevertheless changes the degree to which macroeconomic shocks affect income risk.

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I. Introduction

The recent years have seen an increased integration of countries into the world economy through trade and capital market liberalization. This has led to a parallel surge of interest in the academic and policy literature on the implications of increased "openness" of countries to cross-border trade in goods and factors.[1] The economic benefits and costs of openness are now being actively debated: While many economists have pointed to the gain in allocational efficiency that results from free international exchange, others have pointed out potential downsides, arguing that openness may lead to an increase in income inequality and, separately, income risk (income volatility). Although there is by now a large empirical literature analyzing the impact of trade openness on wage *levels* and the distribution of income,[2] an empirical analysis of the effect of trade openness on individual income *volatility* has so far been lacking. This paper conducts such an empirical investigation, and uses the empirical results in conjunction with a simple dynamic general equilibrium model to assess the corresponding welfare effects.

The theoretical literature has suggested various channels through which trade reform might affect individual income risk. For example, lowering trade barriers leads to an increase in foreign competition in the import-competing sectors and is likely to induce a reallocation of capital and labor across firms and sectors. In the short run, the resulting turbulence may raise individual labor income risk.[3] Rodrik (1997), going beyond the short term re-

[1]For a general discussion of the debate, see for instance, Rodrik (1997) and Bhagwati (2001).

[2]Early papers in this area include Lawrence and Slaughter (1993) and Borjas, Freeman and Katz (1992). See Feenstra and Hanson (2002) for a comprehensive survey treatment.

[3]See, for instance, the analysis of policy change by Fernandez and Rodrik (1991), in which ex-ante identical workers experience ex-post different outcomes since some workers retain their jobs while others are forced to move to other firms. More recently, Melitz (2003) has developed a formal framework in which trade policy changes affecting an entire sector lead to heterogeneous outcomes at the firm level.

allocational effects of trade reform on income risk, has additionally argued that increased foreign competition following trade reform will increase the elasticity of the goods and the derived labor demand functions. If a higher demand elasticity translates any given shock into larger variations in wages and employment, lower trade barriers may lead to increased individual income risk.[4] On the other hand, it has also been suggested that the world economy is likely to be less volatile than the economy of any single country, which leads to goods prices that are more stable worldwide than in any single autarkic economy. This opens up the possibility that greater openness may reduce the variance in individual incomes. Thus, theoretically, the openness-volatility relationship is ambiguous, that is, the theoretical literature does not offer a strong prior on the sign or magnitude of this relationship.[5]

In this paper, we study *empirically* the effects of trade policy on individual income risk using the following approach. For each industry (sector), we use longitudinal data on individual earnings to estimate time-varying parameters of individual income risk using a methodology that follows the approach taken by the extensive empirical literature on labor market risk.[6] More specifically, we focus on the variance of (unpredictable) changes of individual income as a measure of income risk, and carefully distinguish between transitory and persistent income shocks. The distinction between transitory and persistent income shock is important since workers can effectively "self-insure" against transitory shocks through borrowing or own savings, which implies that the effect of these types of shocks on workers' consumption and welfare are quite small (Aiyagari (1994), Heaton and Lucas (1996), Levine and Zame (2002)).

[4]While Rodrik (1997) appears to have in mind mostly aggregate volatility, it is easy to see that his arguments equally apply to individual income volatility if there are idiosyncratic shocks to firm-level productivity.

[5]Clearly, this sign-ambiguity does not extend to the short-term re-allocational effect of trade policy reforms which, as we have discussed above, are generally expected to raise income risk. However, we do not have strong priors on the magnitude of this relationship either.

[6]See, for example, Carroll and Samwick (1997), Gottschalk and Moffitt (1994), Gourinchas and Parker, (2002), Hubbard, Skinner, and Zeldes (1994), Meghir and Pistaferri (2004), and Storesletten, Telmer, and Yaron (2004).

In contrast, highly persistent or permanent income shocks have a substantial effect on the present value of future earnings, and therefore lead to significant changes in consumption even if workers can borrow or have own savings (Constantinides and Duffie (1996) and Krebs (2003a and 2004)). Thus, from a welfare point of view, persistent income shocks matter the most, and we therefore focus on the relationship between trade policy and the persistent component of income risk.[7] More specifically, after obtaining the estimates of the persistent component of income risk for each year and industry, we use these estimates in conjunction with tariff data (as a proxy for trade policy) to study empirically the effect of trade policy on income risk.

In addition to the empirical analysis of the relationship between trade policy and income risk, this paper also provides a quantitative evaluation of the welfare consequences of any changes in income risk that are brought about by changes in trade policy. If insurance markets and other institutional arrangements for sharing individual income risk are missing (incomplete markets), then changes in income risk will alter consumption volatility and therefore workers' welfare. To find out how income risk is linked to consumption volatility and welfare, we use a dynamic general equilibrium model with incomplete markets in which the consumption/saving choice of workers in the presence of idiosyncratic income risk is explicitly modeled. As is well known, general versions of such models are difficult to solve, and most work in the literature has therefore been computationally intensive (Aiyagari (1994), Huggett (1993), Krusell and Smith (1998), Rios-Rull (1996)). In contrast to this literature, we rely upon an extended version of the incomplete-markets model recently developed and analyzed by Constantinides and Duffie (1996) and Krebs (2004) that is highly tractable,

[7] To see the importance of this distinction more clearly, consider the example of a worker who loses his job due to plant closure or any other "exogenous" event. If the worker quickly finds a new job that pays him as well as the previous job, then the worker's consumption level is not likely to drop by too much either during or after the period of unemployment. If, on the other hand, the worker is forced to accept a job that pays him a permanently lower wage because, for example, firm- or occupation-specific human capital has been lost, then the worker's likely response is to reduce consumption.

but still rich enough to allow for a tight link between the econometric framework and the theoretical model. The welfare expressions that we derive theoretically can then be used to translate changes in individual income risk into welfare changes.

Our previous discussion highlights the need for longitudinal information on incomes at a disaggregated level (individual or household)[8] in countries that have undergone discernable (and ideally substantial) changes in their external regime. Unfortunately, countries that maintain detailed longitudinal records on individual incomes have rarely undertaken major trade reforms and countries that have undertaken extensive trade policy reforms have rarely collected data on individuals of requisite scope and quality. In this paper, however, we focus on one country that satisfies both criteria, namely Mexico. As it is well known, the Mexican economy experienced substantial changes in trade policy in the late 1980's and in the later half of the 1990s.

Our empirical results for the Mexican case can be summarized as follows. First, we find that trade policy *changes* have a significant short run effect on income risk for industries with high levels of import penetration, with a tariff reduction of five percent raising the standard deviation of the persistent shocks to income by about twenty five percent. In terms of welfare, we find that this increase in income risk is equivalent to a decrease in lifetime consumption by almost one percent (using a discount factor and degree of risk aversion that are standard in the macroeconomic literature, Cooley (1995)) for workers in the high import-penetration industries.[9] Second, the effect of the tariff *level* on income risk is insignificant. Third,

[8] It should be clear that our need for longitudinal data follows from our desire to study how trade policy impacts the magnitude and frequency of individual income shocks (changes). This is a quite distinct task from that of measuring the impact of trade policy on the distribution of income levels.

[9] Even though these are only short-run effects, the fact that we are dealing with permanent income shocks to individual workers means that in this relatively short period *some* of the workers get scarred for life. Thus, ex ante, workers are willing to give up a substantial amount of their expected lifetime consumption in return for the elimination of the risk of losing with a trade reform.

while the tariff *level* has an insignificant mean effect, it nevertheless changes the degree to which macroeconomic shocks affect income risk. For instance, we find that tariff reductions increase the cost of recessions substantially. More specifically, at a tariff level of ten percent a reduction in the growth rate of GDP of five percent is estimated to raise the standard deviation of persistent income shocks by twelve percent, whereas at a five percent tariff rate the same reduction in GDP growth increases income risk by twenty five percent. In terms of welfare, this amounts to an increase in the cost of recessions that is equivalent to almost half a percentage point of lifetime consumption. Notice, however, that our empirical estimates also indicate that tariff reductions decrease individual income risk during economic booms, so that the net welfare cost of tariff reforms due to this interaction effect is smaller than half a percentage point of lifetime consumption.[10]

At this stage, it is worth pointing out some of the limitations of our analysis. First, we focus exclusively on the link between trade policy and individual income risk, and therefore neglect other channels through which trade policy may affect the economy. More specifically, one would expect trade liberalization to have positive effects on the efficiency of resource allocation and economic growth (the mean of income changes), and these effects are important factors that any comprehensive welfare analysis of trade liberalization ought to take into account. Second, our welfare calculations do not allow for the possibility that an increase in income risk might lead to a simultaneous rise in insurance opportunities (endogenous market incompleteness).[11] Third, we follow a long-standing tradition in economics and measure risk by the variance (second moment) of the relevant distribution, which is justified if (as

[10]Because of space limitations, in this paper we do not attempt to find a precise estimate of this welfare cost taking into account both the increase in income risk during recessions and the decrease during economic booms. Such an estimate could be found by adopting the methodological approach used in the literature on the welfare cost of business cycles when markets are incomplete. See, for example, Krebs (2003b) and Lucas (2003) for more details.

[11]See, for example, Attanasio and Rios-Rull (2000) and Krueger and Perri (2002), for a formal analysis of this phenomenon in economies with limited commitment.

assumed in this paper) the economic variables of interest are (log)-normally distributed. Finally, the Mexican household survey we use to implement our general approach is a rotating panel that follows individual workers for five quarters over time, which means that the panel dimension of our income data is somewhat limited. Thus, our data do not allow us to assess with certainty the persistence of income shocks beyond five quarters. However, a comparison of our estimates of the income risk parameters with existing results that use data sets with a much longer panel dimension suggests that a large fraction of the income shocks we label "persistent" in this paper last indeed for many years (see Section II.5 for more details). In short, the welfare results presented here do not necessarily show that trade liberalization is costly, but they do provide strong evidence that any comprehensive welfare analysis of trade liberalization ought to take into account the cost of increased labor market risk.

In summary, in this paper we articulate a general framework that allows us to study empirically the impact of trade reform on individual income risk and to evaluate the corresponding welfare effects. We use this framework to study the Mexican economy, which, as we have argued above, seems well-suited for such an analysis. In our empirical implementation of this methodology using longitudinal data on Mexican workers, we find economically significant effects of trade policy on income risk.

We conclude this introduction with a brief comment on some of the earlier empirical literature on the relationship between trade policy and factors related to labor market risk. The impact of trade liberalization on short-run worker displacement has been investigated in the well-known papers of Currie and Harrison (1997), Gaston and Trefler (1994), Levinsohn (1999) and Revenga (1997), among others. More recently, in an innovative paper, Trefler (2004) has analyzed the short-run adjustment costs borne by displaced workers simultaneously with the long run benefits (of higher firm productivity and resource allocation) that accrued in the context of the trade agreement between United States and Canada. While these papers have provided us with very valuable analyses of the labor market impact of trade policy reforms, they do not focus *directly* on income risk, which is the primary topic of interest to the current paper. Specifically, none of the existing studies estimates an individual income process that allows one to gauge the severity and persistence of shocks to individual income (resulting, for instance, from job displacement following trade policy reform), which, as we have argued above, is crucial when thinking about the welfare consequences of trade reform. In a similar vein, while several scholars have commented upon the potential importance of the link between openness and aggregate volatility in the presence of market incompleteness,[12] empirical studies of the relationship between openness and *aggregate* volatility (Rodrik (1998)) have the drawback that the welfare effects of aggregate fluctuations are often found to be quite small (Lucas (2003)). In short, none of the previous studies has analyzed the link between openness and income risk in the manner and detail that we do here.

II. Income Risk

The first stage of our analysis concerns the estimation of individual income risk. Our estimation strategy follows earlier approaches in the literature estimating US labor income risk (Carroll and Samwick (1997), Hubbard et al (1994), Gourinchas and Parker (2002), Meghir and Pistaferri (2004), and Storesletten et al. (2004)) with some important differences which we discuss in detail below. As in these papers, we define income risk as the variance of (unpredictable) changes in individual income, and carefully distinguish between transitory and persistent income shocks. From a welfare point of view, this separation is essential for

[12]Early theoretical analyses of trade patterns and optimal trade policy with aggregate risk and incomplete markets include Eaton and Grossman (1985) and Helpman and Razin (1980), among others. An interesting and somewhat related theoretical literature on international production and trade patterns with incomplete *contracting* has been developed recently (see Antras (2004) and Helpman and Grossman (2002)), but it has not (yet) considered explicitly either aggregate or idiosyncratic risk in the economic environment.

two reasons. First, consumption smoothing through borrowing or own saving works well for transitory income shocks (Aiyagari (1994), Heaton and Lucas (1996), and Levine and Zame (2002)), but not when income shocks are highly persistent or permanent (Constantinides and Duffie (1996) and Krebs (2003a and 2004)). Thus, highly persistent income shocks have a large effect on consumption volatility and welfare, whereas the effect of transitory shocks is relatively small. Second, the transitory term in our econometric specification of the income process will absorb the measurement error in individual income, and therefore allows us to arrive at a better estimate of the true amount of individual income volatility. For these reasons, we eventually focus on persistent shocks and their relation to trade policy.

II.1. Data

In Mexico, the National Urban Employment Survey (ENEU) conducts extensive *quarterly* household interviews in the 16 major metropolitan areas and is available from 1987 (we use data from 1987-1998 in our study). The ENEU is structured so as to track a fifth of each sample across a five quarter period. The sample is selected to be geographically and socio-economically representative. The treatment of sample design, collection and data cleaning is careful. The survey questionnaire is extensive in scope and covers all standard elements such as participation in the labor market, earnings etc.[13]

We use information on labor market participants between the ages of 16 and 65. Individual panels were constructed by matching workers by their position in an identified household, level of education (years of schooling), age and sex. Questions referring to labor income refer to income earned in the previous quarter. Workers earnings include their overall earnings from fixed salary payments, hourly or daily wages, piece-meal work, commissions, tips and any entrepreneurial earnings (earned by the self-employed). Taken together, we have 44

[13]The actual surveys and documentation of methodology are available on request.

complete panels of 5 periods (i.e., quarters) each, spanning a total of 12 years (48 quarters).

Table I presents a summary description of the workers surveyed by the ENEU. Other aspects of our ENEU data − the evolution of the mean and variance of earnings and returns to education over time (not presented here but available on request) − matched the facts about earnings in the Mexican labor market reported by previous authors.[14]

Data on sectoral trade barriers and other sectoral and macroeconomic variables were obtained from the World Bank.

II.2. Specification

Our survey data provide us with earnings (wage rate times number of hours worked) of individuals. As in previous empirical work, we assume that the log of this labor income of individual i employed in industry j in period t, $log\, y_{ijt}$, is given by:

$$log\, y_{ijt} = \alpha_{jt} + \beta_t \cdot x_{ijt} + u_{ijt}\ . \tag{1}$$

In (1) α_{jt} and β_t denote time-varying coefficients, x_{ijt} is a vector of observable characteristics (such as age and education), and u_{it} is the stochastic component of earnings. The stochastic component u_{ijt} represents individual income changes that are *not* due to changes in the return to observable worker characteristics. For example, income changes that are caused by an increase in the skill (education) premium are not contained in u_{ijt}. In this sense, u_{ijt} measures the unpredictable part of changes in individual income. Notice that we allow the fixed effects α_{jt} to vary across sectors, but that the coefficient β_t is restricted to be equal across sectors. The latter assumption is made in order to ensure that the number of observations is large compared to the number of parameters to be estimated.

[14]See Hanson (2003) for a broad analysis of wage patterns in Mexico in the 1990s based on population census data.

We assume that the stochastic term is the sum of two (unobserved) components, a permanent component ω_{ijt} and a transitory component η_{ijt}:

$$u_{ijt} = \omega_{ijt} + \eta_{ijt} \ . \tag{2}$$

Permanent shocks to income are fully persistent in the sense that the permanent component follows a random walk:

$$\omega_{ij,t+1} = \omega_{ijt} + \epsilon_{ij,t+1} \ , \tag{3}$$

where the innovation terms, $\{\epsilon_{ijt}\}$, are independently distributed over time and identically distributed across households. Notice that we allow the parameters to depend on time t and industry j, but not on individual i. We further assume that $\epsilon_{ij,t+1} \sim N(0, \sigma^2_{\epsilon j,t+1})$. Transitory shocks have no persistence, that is, the random variables $\{\eta_{ijt}\}$ are independently distributed over time and identically distributed across households. Clearly, η_{ijt} captures both temporary income shocks and measurement error. We assume that they are normally distributed with zero mean and a variance that is independent of i, but may depend on time or industry: $\eta_{ijt} \sim N(0, \sigma^2_{\eta jt})$.

Our specification for the labor income process is in accordance with the empirical work on US labor income risk. For example, Carroll and Samwick (1997) and Gourinchas and Parker (2002) use exactly our specification. Hubbard, Skinner and Zeldes (1994) and Storesletten, Telmer and Yaron (2004) assume that the permanent component is an AR(1) process, but estimate an autocorrelation coefficient close to one (the random walk case). Finally, some papers have allowed for a third, MA(1), component. See, for example, Meghir and Pistaferri (2004). Notice also that with the exception of Meghir and Pistaferri (2004) and Storesletten et al. (2004), the previous literature has confined attention to the special case of time-independent variances (homoscedastic case). As we discuss in II.3, the introduction of time-variation in the parameters $\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$ makes the estimation of these parameters more challenging.

II.3. Estimation

Consider the change in the residual of income of individual i between period t and $t+n$:

$$\begin{aligned} \Delta_n u_{ijt} &= u_{ij,t+n} - u_{ijt} \\ &= \epsilon_{ij,t+1} + \ldots + \epsilon_{ij,t+n} + \eta_{ij,t+n} - \eta_{ijt} \,. \end{aligned} \tag{4}$$

We have the following expression for the variance of these income changes:

$$var[\Delta_n u_{ijt}] = \sigma^2_{\epsilon j,t+1} + \ldots \sigma^2_{\epsilon j,t+n} + \sigma^2_{\eta jt} + \sigma^2_{\eta j,t+n} \,. \tag{5}$$

We use the moment restrictions (5) to estimate the parameters $\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$ using GMM,[15] where the sample analogs to the moment conditions are formed by using the estimates of u_{ijt} obtained as residuals from regressions of labor income on observable characteristics as specified in (1) − an approach also used by Meghir and Pistaferri (2004), Storesletten et al. (2004) and Gourinchas and Parker (2002).[16] Specifically, the estimator is obtained by minimizing:

$$\sum_{t,n} \Big(var[\Delta_n u_{ijt}] - \Big(\sigma^2_{\epsilon j,t+1} + \ldots \sigma^2_{\epsilon j,t+n} + \sigma^2_{\eta jt} + \sigma^2_{\eta j,t+n} \Big) \Big)^2 \tag{6}$$

The first-order conditions corresponding to the parameters $\sigma^2_{\epsilon j,t}$ and $\sigma^2_{\eta j,t}$ are given by:

$$\begin{aligned} \forall t \; : \; \frac{\partial \sum}{\partial \sigma^2_{\epsilon j,t}} &= 0 \\ \forall t \; : \; \frac{\partial \sum}{\partial \sigma^2_{\eta j,t}} &= 0 \end{aligned} \tag{7}$$

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[15]More specifically, we follow the bulk of the literature and use the equally weighted minimum distance (EWMD) estimator. Altonji and Segal (1996) suggests that the EWMD estimator (identity weighting matrix) is superior to the two-stage GMM estimator (optimal weighting matrix) once small-sample bias is taken into account.

[16]Notice that Meghir and Pistaferri (2004) and Storesletten et al. (2004) exploit additional moment restrictions that follow from the autocovariance function of income changes.

Notice that in general there are many more moment conditions (5) than there are parameters to be estimated. More precisely, for each time period t and each industry j, there are two parameters ($\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$), but n moment conditions (5). For example, in our data set on Mexico, for each industry j we have $t = 48$ quarters and $n = 4$ quarters (individuals drop out of the sample after 5 quarters), and the number of parameters is therefore $2 * (48)$, whereas the number of moment conditions is approximately $4 * (48)$.[17] The system is thus over-identified.

Notice also that the objective function (6) is quadratic, which implies that the first-order conditions associated with the corresponding minimum-distance problem are linear in $\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$– a feature that facilitates the estimation substantially. Specifically, the first-order conditions can be organized into a linear equation system

$$\mathbf{A} \cdot \sigma = \mathbf{b} \tag{8}$$

where $\sigma = (\sigma^2_{\epsilon,2}....\sigma^2_{\epsilon,t}..\sigma^2_{\epsilon,T},\ \sigma^2_{\eta,2}..\sigma^2_{\epsilon,t}..\sigma^2_{\eta,T})'$ is a 2(T-1)-dimensional vector of income parameters (T being the total number of time periods). Estimates of these income parameters can then easily be obtained through matrix inversion: $\sigma = \mathbf{A^{-1}b}$.

Some intuition for the way in which our approach separates transitory from permanent income shocks can be obtained from the following simple example. Suppose that risk is time-invariant, $\sigma^2_{\epsilon jt} = \sigma^2_{\epsilon j}$ and $\sigma^2_{\eta jt} = \sigma^2_{\eta j}$, an assumption that has been made by most of the previous empirical literature on income risk. In this case, the moment restrictions (5) become the following:

$$var[\Delta_n u_{ijt}] = 2\sigma^2_{\eta j} + n\,\sigma^2_{\epsilon j} \tag{9}$$

Thus, the variance of observed n-period income changes is a linear function of n, where

[17] We say "approximate" because towards the very the end of the sample period, clearly fewer than $n = 4$ income changes are observed. In the penultimate quarter, for instance, only one income change is observed. However, this does not pose a problem for the estimation of any but the parameters of the very last quarter.

the slope coefficient is equal to $\sigma^2_{\epsilon j}$. The insight that the random walk component in income implies a linearly increasing income dispersion over time is the basis of the estimation method used by several authors. For example, Carroll and Samwick (1997) estimate σ^2_ϵ by performing OLS regressions of the left-hand-side of (9) on n. While the preceding example, with time-invariant parameters, serves to illustrate the intuition underlying the estimation procedure, it should be clear that our exercise is more general in the sense that it allows for arbitrary time-variation in the income risk parameters.

II.4. Estimation using ENEU Data

The preceding section provided a detailed description of a general econometric methodology that may be used to estimate time-variant income risk parameters given longitudinal data on individual incomes. We note here some additional issues that arise in applying this methodology to our data, with particular emphasis on the type of income risk accounted for by our estimation procedure.

In forming the sample analogs to the moment conditions (5), we use information on all individuals who are present in a given manufacturing industry in both time periods t and $t+n$ (with $n \leq 5$) regardless of their employment status in any intermediate period. In doing so, we pick up shocks to workers who retain their jobs but experience income changes due to changes in their wage rates or the number of hours worked. Moreover, we also account for changes in income experienced by workers who have lost their job in period t, but are re-employed in the same industry in some subsequent period $t+n$ (with $n \leq 5$), and this is true even if these workers are unemployed in any intermediate period. In particular, we do account for the long-term earnings losses of a large fraction of displaced workers, namely all those displaced workers who are re-employed in the same industry but have lost firm-or occupation-specific human capital.[18] In contrast, displaced workers who are reallocated to a different manufacturing industry are not taken into account.[19] However, in our data set, the exclusion of such workers is not expected to cause too much of an under-estimation of the income risk parameters since the fraction of displaced manufacturing workers who make the transition from one manufacturing sector to another is very small. Indeed, examining re-employment rates for workers who start in manufacturing and go through a period of unemployment suggests that only approximately ten percent of these displaced workers undergo a transition from one manufacturing sector to another. Note that this finding is consistent with observations from the United States that most job creation and destruction takes place within industries (see, for instance, Davis, Haltiwanger and Schuh (1996)).

Finally, our construction of the sample analogs to the moment conditions (5) could lead to an under-estimation of the *persistent* component of income risk due to the non-inclusion of workers undergoing prolonged spells of unemployment (specifically those workers who experience unemployment spells exceeding four quarters). However, this is not a severe problem here. One consequence of the lack of any government-provided unemployment insurance in Mexico and the very active informal labor market is that there are few labor force participants in our survey with extended unemployment durations. Specifically, of those workers looking for work, the proportion who had experienced unemployment durations of four quarters or more was extremely small (less than 0.05 percent of workers).

Finally, we should mention that the variability in income experienced by workers in our data set derives from both changes in the number of hours worked and changes in the real wage.

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[18] For the U.S., these long-term earnings losses have been estimated to be very large (on average 25% for high-tenure workers according to Jacobson, LaLonde, and Sullivan (1993)).

[19] This allows us to circumvent the extremely difficult problem of assigning industries (and thus trade policy) to individuals who transit to different industries. Including individuals who make transitions to the service (non-tradables) sector by using the procedure of counting them as belonging to the manufacturing sector in which they are first observed does not result in any qualitative difference in our reported results.

Real wage changes, in turn, can be positive or negative, and in our Mexican data substantial declines in the real wage are quite common. More specifically, Mexico experienced very high inflation rates during our sample period with annual declines in the *aggregate* real wage as high as 25 percent during this time (see, for instance, Hanson (2003)), implying that the wage rates of some individual workers declined by an even larger amount. Thus, despite the often cited downward rigidity of wages, our sample includes large numbers of workers whose *real* wages declined dramatically.

II.5. Results

As described before, we have individual income data for the time period 1987-1998 covering 21 different manufacturing sectors in Mexico. Using the methodology outlined above, we estimate the risk parameters σ_ϵ^2 and σ_η^2 for each quarter and each manufacturing sector. In Tables II and III we provide the average estimate of σ_ϵ^2 and σ_η^2 for each year (averaged across industries) and for each industry (averaged over time) respectively.[20] The mean value (across industries and over time) of the quarterly variance of the persistent shock, σ_ϵ^2, is estimated to be 0.008, or 0.032 annualized (i.e., σ_ϵ, is estimated to have a mean quarterly value of 0.09 and a mean annualized value of 0.18).[21] As expected, given the extent of measurement error in the income data (see our discussion in Section II), the estimated variances of transitory shocks are much larger in magnitude. More precisely, the mean value of the annualized variance of transitory shocks is 0.2 (an annual standard deviation of 45 percent), which is

[20]The averages presented in Tables II and III are merely summary descriptions and do not allow for any direct inferences regarding the relationship between trade policy and income risk.

[21]Given that in Section III we seek to uncover the relationship between trade policy and income risk using our estimates of the income risk parameters σ_ϵ, it is also interesting to investigate to what extent these estimates differ across industries and over time after making some adjustment for the fact that there is estimation error. To quantify this variation, we use the methodology of Krueger and Summers (1988). More specifically, we compute a measure of the "adjusted standard deviation" of the point estimates of the income risk parameters. In turns out that this number (0.018) is over twice the mean value of σ_ϵ in our sample – indicating that the variation in σ_ϵ across industries and over time is indeed significant in our exercise.

clearly too large to be a true measure of income volatility.

It seems informative to compare our estimates of the permanent component of income risk, σ_ϵ^2, with the estimates obtained by the extensive empirical literature on U.S. labor market risk using annual income data drawn from the PSID. Most of these studies find an average value of around .0225 for the annual variance σ_ϵ^2 (Carroll and Samwick (1997), Gourinchas and Parker (2002), Hubbard, Skinner and Zeldes (1994), and Storesletten, Telmer and Yaron (2004)), with a value of $\sigma_\epsilon^2 = .0324$ being the upper bound (Meghir and Pistaferri, 2004). Assuming that these income shocks are i.i.d. over time (the maintained random walk assumption), this means that these studies have found a quarterly variance of $\sigma_\epsilon^2 = .0056$, with one study estimating $\sigma_\epsilon^2 = .008$. Thus, the average value of our estimates of permanent income risk is in line with the estimates that have been obtained by the previous literature on U.S. labor market risk, although our estimates lie somewhat on the high end. Notice that our estimates are obtained using a five-quarter rotating panel, whereas Carroll and Samwick (1997), Gourinchas and Parker (2002), Hubbard, Skinner and Zeldes (1994), Meghir and Pistaferri (2004), and Storesletten, Telmer and Yaron (2004) use the PSID data with a panel dimension of many years. Thus, as long as Mexican workers face similar amounts of permanent labor income risk as U.S. workers (or more), this result suggests that most income shocks we label "permanent" in this paper indeed persist for a very long time.

III. Trade Policy and Income Risk

The procedure outlined in the previous section provides us with estimates of individual income risk, $\sigma_{\epsilon jt}^2$, for each industry (i.e., manufacturing sector) j and time period, i.e., quarter, t. We now use these time-varying, industry-specific estimates in conjunction with observations on trade policy, τ_{jt}, to estimate the relationship between income risk, $\sigma_{\epsilon jt}^2$, and openness, τ_{jt}, using a linear regression model. As mentioned before, in this paper we focus on permanent component of income risk, σ_ϵ^2, instead of the transitory component, σ_η^2, for two reasons: i) transitory income shocks are unlikely to generate substantial consumption volatility and ii) σ_η^2 is likely to contain a large amount of measurement error. Despite these theoretical arguments, it might still be of interest to study the relationship between trade policy and income risk using σ_η^2 as a measure of income risk. We therefore also conducted a similar regression analysis (not reported here) for transitory income-shock parameters, σ_η^2, but we did not find any statistically significant relationship between transitory shocks to income and trade policy.

III.1. Specification

We first consider a linear specification that allows for industry fixed effects and aggregate time effects:

$$\sigma_{\epsilon jt}^2 = \alpha_0 + \alpha_{1j} + \alpha_{2t} + \alpha_\tau \, \tau_{jt} + \alpha_{\delta 1} \, \Delta\tau_{jt} + \alpha_{\delta 2} \, \Delta\tau_{jt} D_{jt} + \nu_{jt} \; . \tag{10}$$

In (10) we have included on the right hand side the following variables: τ − the ad valorem sectoral tariff rate, $\Delta\tau$ − the change in the tariff over the preceding year, $\Delta\tau D$ − the tariff change over the preceding year interacted with an indicator variable that takes the value one if the import penetration ratio is greater than its sample median and zero otherwise,[22] α_j − an industry fixed- effect, and α_t − a time dummy that captures general macroeconomic trends in the economy.

The inclusion of industry dummies in the specification (10) allows us to control for any fixed industry-specific factors that may affect the level of riskiness of income in that industry. Moreover, the inclusion of time dummies controls for any changes in macroeconomic condi-

[22]Clearly, $\alpha_{\delta 1}$ measures the effect of a trade policy change in sectors that had lower than median import-penetration both before and after this change and $\alpha_{\delta 1} + \alpha_{\delta 2}$ correspondingly measures the effect of trade policy changes in sectors that had higher than median import-penetration both before and after the change. This is also true with specification (10') below.

tions that affect the level of income risk. While this ensures that our estimation results are not driven by changes in *macroeconomic* conditions (business cycle effects and/or long-run structural changes) unrelated to trade policy, it also means that identification of the relationship between $\sigma^2_{\epsilon jt}$ and τ_{jt} will have to be based on the differential rate of change in trade barriers across sectors over time (or the vector of observations on tariffs in the panel corresponding to (10) will be perfectly collinear with the time-dummy vector). This, however, does not pose problems for our estimation since trade barriers in Mexico and their changes over time do in fact do exhibit substantial cross-sectional variation.[23]

Specification (10) provides the starting point for our econometric analysis. An alternate specification is the following:

$$\sigma^2_{\epsilon jt} = \alpha_0 + \alpha_j + \alpha_\tau \, \tau_{jt} + \alpha_{\delta 1} \, \Delta\tau_{jt} + \alpha_{\delta 2} \, \Delta\tau_{jt} D_{jt} + \beta_e \Delta e_t + \beta_g g_t + \phi_e (1+\tau_{jt}) \Delta e_t + \phi_g (1+\tau_{jt}) g_t + \nu_{jt} \; . \tag{10'}$$

Specification (10') exploits the *within* industry variation in tariffs over time to a greater extent by dropping the time dummies and including instead the following two macroeconomic variables: Δe, the depreciation of the real exchange rate over the preceding year, and g, the GDP growth rate. Also included are the interaction terms $(1+\tau)\Delta e$ and $(1+\tau)g$, which measure the extent to which the relationship between income risk and these macroeconomic factors varies with trade policy.[24]

Several econometric issues arise in the estimation of equations (10) and (10') above, most of which we discuss in more detail below (sections III.3 and III.4). At this stage, we only

[23] For instance, in Mexico, tariffs varied between 80 and 20 percent prior to the trade reforms of 1987 and ranged between 20 and 10 percent by 1994 - implying a variation in tariff changes across sectors that is quite substantial.

[24] Note that the only variable that is interacted with the dummy variable D (representing greater-than-median import penetration) is the change in tariffs, $\Delta\tau_{jt}$. The remaining variables such as exchange rate depreciation, Δe_t, and growth rate of GDP, g_t, are already interacted with the tariff level (which itself has a quite strong *within* industry correlation with import penetration). Estimating (10') separately for industries with $D = 0$ and $D = 1$ gave results very similar to those reported here.

note the following. First, one concern is that the left-hand-side variable, income risk, is estimated and not observed. This is not a substantial problem by itself as it is well known that while "measurement error" in the dependent variable does reduce precision, it does not bias our estimates. Second, a concern arises from the fact that the estimates of $\sigma^2_{\epsilon jt}$ have different standard errors across industries, that is, the specification we have described above suffers from a heteroscedasticity problem. Third, since the industries all belong to the same macroeconomic environment, there is a possibility of contemporaneous correlation in their σ's even after controlling for observable macroeconomic factors as in (10'), i.e., $Cov(\nu_{jt}\nu_{j't}) \neq 0$. Finally, serial correlation in income volatility within an industry is a possibility, i.e., $Cov(\nu_{jt}\nu_{jt'}) \neq 0$. Given the possible presence of heteroscedasticity, spatial correlation and serial dependence, consistent estimates of the standard errors associated with the coefficient estimates in (10) and (10') above are obtained by using robust estimation techniques.

III.2. Results

In (10), the effect of the tariff level on income risk is given by the coefficient α_τ and the effect of tariff changes on income risk is given by the coefficient α_δ. The first column in Table IV presents the estimation results. We note first that the estimate of α_τ is insignificant and we are therefore unable to reject that the *mean* effect of the tariff level on income risk is zero. However, trade policy *changes*, in sectors with above-median level of import penetration ($D = 1$), have statistically and economically significant short run effect on income risk ($\hat{\alpha}_{\delta 1} + \hat{\alpha}_{\delta 2} =$ -0.125, with an estimated standard error of 0.05). This estimate indicates that lowering the tariff rate by five percent would, for a year, raise σ^2_ϵ by .00625 from, for example, .008 (its mean value) to .01425 . In terms of the standard deviation σ_ϵ, this amounts to an increase from .089 to .1193, that is, an increase by more than thirty percent − a substantial increase in income risk indeed.

Estimates from (10') are presented in the second column of Table IV. Note that tariff *changes* in high import-penetration sectors continue to have economically and statistically significant effects of magnitude quite similar to those obtained from estimation of (10) ($\hat{\alpha}_{\delta 1} + \hat{\alpha}_{\delta 2}$ = -0.092, with an estimated standard error of 0.045). More specifically, a five reduction in tariffs increases σ_ϵ^2 from a mean level of .008 to .0126, which in terms of the standard deviation σ_ϵ amounts to an increase from .089 to .1122 (a twenty five percent increase). Interestingly, the coefficient α_τ is now significant. However, the effect of the tariff *level* on income risk is now given by $(\alpha_\tau + \phi_e \Delta e + \phi_g g)$. After substituting in the mean values of Δe and g from the sample, this estimated sum revealed to be insignificantly different from zero ($\hat{\alpha_\tau} + \hat{\phi}_e \bar{\Delta} e + \hat{\phi}_g \bar{g}$ = 0.02, with an estimated standard error of 0.02). Thus, we are again unable to reject that the *mean* effect of the tariff level on income risk is zero.[25]

Consider now our estimates of how the tariff level alters the effect of macroeconomic variables on income risk. The coefficient on real exchange rate depreciation, β_e, is estimated negative and significant as is the coefficient on GDP growth, β_g, while the coefficients ϕ_e and ϕ_g relating to the interaction terms, $(1+\tau)\Delta e$ and $(1+\tau)g$, are both positive and significant. The extent to which the tariff level alters the effects of exchange rate changes on income risk is given by ϕ_e. As reported in Table IV, this parameter is estimated to have a mean value of 0.54 and an estimated standard error of 0.18. Consider a real exchange rate appreciation of ten percent under two scenarios − when the tariff rate is ten percent and when the tariff rate is five percent. If the tariff rate is ten percent, our estimates indicate that an exchange rate appreciation of ten percent (in the preceding year) raises σ_ϵ^2 from 0.008 to 0.0108 (an

[25]Our estimates of the timing and magnitude of the effect of trade policy changes on measured income shocks (i.e., large changes in the year following policy changes and zero mean effects) also indicate that our results are not being driven by other "unobserved" factors such as skill *and* sector biased technical changes that are possibly correlated with trade policy changes. More specifically, we would expect any such changes in technology to impact income in a gradual manner taking several years for its full impact to be realized. Note also that our own estimates of the returns to education suggest a striking similarity across manufacturing sectors in Mexico, which provides indirect evidence against the view that technological progress in Mexico during the relevant sample period was both skill and sector biased.

increase of just about thirty five percent). In contrast, if the tariff rate is five percent instead, the same appreciation implies an increase in income risk from 0.008 to 0.013 (an increase of over sixty percent). Similarly, if the growth rate of GDP, g, is lowered by five percent, σ_{ϵ}^2 is raised from 0.008 to 0.01 (an increase of over twenty five percent) when the tariff rate is ten percent, but the same change in g results in a short run increase in income risk from 0.008 to 0.013 (an increase of over sixty percent) when the tariff rate is at five percent. Of course, as noted earlier, our empirical estimates also indicate that tariff reductions lead to a corresponding reduction in individual income risk during economic booms. Overall, our estimates suggest that the magnitude of the (short run) effects of macroeconomic shocks on income risk is significantly altered by the tariff level.

The dependence of the income risk parameter σ_{ϵ}^2 on cyclical conditions is not only observed in Mexico, but has also been well documented for the United States (Meghir and Pistaferri (2004), Storesletten, Telmer and Yaron (2004)). However, this literature has not studied how trade policy affects this dependence of idiosyncratic risk on cyclical conditions. Thus, the estimation results reported in Table IV provide the first empirical evidence that trade liberalization increases the sensitivity of idiosyncratic risk to business cycle conditions. Theoretically, one might *speculate* that a mechanism similar to the one modeled by Newberry and Stiglitz (1984) is behind our empirical finding. More specifically, Newberry and Stiglitz (1984) argue that a negative productivity shock would have a smaller equilibrium effect on output and employment in a closed economy than an open one - as prices rise with a negative supply shock in the former but are constrained by world prices in the latter. With heterogeneous effects on firms and individuals, the link between macroeconomic downturns and idiosyncratic income risk may therefore also be amplified in more open economies. A more rigorous modeling of this idea within the context of a dynamic general equilibrium model with incomplete markets is an interesting topic for future research.

III.3. Endogeneity and Selection Bias

One concern that arises in our estimation of equations (10) and (10') is that tariff rates are not fully exogenous. Indeed, the theoretical literature on the political economy of trade policy has proposed several hypotheses concerning the endogenous determination of tariffs. Furthermore, a number of empirical studies have explained (partially) the cross industry variation in tariffs using a number of economic and political variables that vary across industries such as the lobbying strength and employment size of particular sectors.[26] While the literature has not studied (or indeed even suggested) income risk as a determinant of cross-sectional variation in trade policy, the possibility that it might be a relevant determinant of policy makes is potentially problematic. Consider, for instance, an economy in which raising the tariff rate in a sector would in fact lower income risk in that sector. Consider further that the government there is "equity" minded and chooses higher protection levels for those industries with intrinsically high levels of income risk − thereby eliminating cross-sectional variation in income risk. If such an economy were studied purely in the cross-section, it may appear that there is no relation between trade policy and income risk even though such a relationship does exist. This type of purely cross-sectional endogeneity, however, is not a problem for our empirical analysis since we follow industries over time. More precisely, the *within* estimator we use is formed by considering changes *within* industries in income risk and tariffs over time, and any endogeneity bias deriving from purely cross-sectionally varying political-economy determinants of trade policy is therefore eliminated.

Along the time dimension, estimation bias could arise if the government attempts to protect vulnerable industries by raising tariff rates for those industries that have experienced an increase in income risk. While such endogeneity bias is in principle a matter of concern, there are at least two facts that speak against this view. First, the trade policy changes that

[26]See, for instance, Trefler (1993). Gawande and Krishna (2003) provide a survey discussion.

we study here are changes that were undertaken during major policy reform episodes (both in the late 1980s and under NAFTA), and many observers have argued that the lowering of trade barriers was mainly used by the Mexican government to signal its commitment to overall policy reform (Tornell and Esquivel, 1995). Second, and somewhat related to the first point, in our data virtually no industry experienced a rollback of the liberalization effort once tariff rates had been reduced. Finally, we note that such pattern of endogeneity would only cause a bias against our reported findings. That is, if such bias exists, the true short-run effect of trade policy changes on income risk is even larger than what we report in this paper. However, it also means that our finding that trade liberalization has no long-run "level effect" could be the result of two opposing effects canceling each other out.[27]

Estimation bias could, of course, also arise if systematic changes in non-tariff barriers reversed the effects of tariff reductions, but these changes in non-tariff barriers were not taken into account by us. To ensure that this is not the case, we studied the patterns in the use of non-tariff barriers (NTBs) in Mexico in the years included in our sample. NTB use in Mexico primarily took the form of anti-dumping duties in these years and the anti-dumping duties were concentrated entirely in the 'Basic Metal Products', 'Chemicals' and 'Textiles' industries.[28] Studying the link between trade policy and income risk using data from the remaining industries did not alter qualitatively or quantitatively any of the reported estimates (see Table VIII).

Our estimation results could also be biased if there is unobserved heterogeneity among work-

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[27]Notice also that despite the work by Alesina and Drazen (1994) and others, major trade policy reforms are in general rather difficult to understand theoretically once policy is treated as being endogenous. The dominant theory of endogenous trade policy determination - the interest group theory - simply does not predict such dramatic changes in policy. Since the competing strengths of various interest groups are not expected to (and do not) change dramatically over the medium term, the theory predicts stickiness in trade policy over these horizons (consistent with what is observed most of the time). Lacking theoretical guidance, the choice of suitable "exogenous variables" to help with identification is even more difficult than usual.

[28]See the recent UNCTAD study, "Mexico's Experience with the use of Anti-dumping Measures," 2002.

ers *and* industries, and heterogenous workers select into different industries. Suppose, for example, that industries with high levels of protection (high tariff levels) are also industries with low job destruction rates.[29] Suppose further that there are two types of workers, good and bad, and that good workers quickly find a new job in the event of job displacement, but bad workers do not. Other things being equal, we would expect bad workers to move to high-protection industries. In this world, high tariff rates lower income risk because they reduce job destruction rates, but they also attract high-risk (bad) workers leading to a downward bias of our empirical estimates of the relationship between income risk and tariff levels (the coefficient α_τ in equation (10)). Thus, it is possible that our empirical finding that tariff levels have no effect on income risk is simply due to this type of selection bias.[30]

In general, it is difficult to deal with the type of selection bias we have just described. However, there is some evidence that in our case any effect due to selection bias is relatively moderate. More specifically, we would expect workers with low job finding rates be mainly low-ability workers. If we use years-of-schooling as a observable proxy for (unobserved) ability, then one implication of the type of selection bias described above is that years-of-schooling (human capital) and income risk should be negatively correlated across industries. However, in our data set, the correlation between average education levels and income risk across industries is very small (-0.06) and insignificant.

Clearly, there could be unobserved ability differences among workers that are uncorrelated with years-of-schooling, in which case selection bias might still be problematic even if the

[29] We thank a perceptive referee for suggesting this example. Note also that the selection bias we discuss here bears some resemblance with the type of lemons' problem discussed by Gibbons and Katz (1991).

[30] If trade liberalization mainly targets high-protection industries and high-risk workers leave industries that experience large tariff cuts, then this self-selection effect also causes a downward bias of our estimates of the relationship between tariff changes and income risk (the coefficients $\alpha_{\delta 1}$ and $\alpha_{\delta 2}$ in equation (10)) . Thus, the true short-run effect of trade liberalization might be even larger than the (already substantial) effect reported in Table IV.

cross-industry correlation between years-of-schooling and income risk is nil. However, even in this case we would expect any selection bias to manifest itself in unexplained wage differentials across sectors, at least as long as high-ability workers are paid higher wages. A casual examination of the data, however, suggests that such cross industry wage differentials are small (at least in relation to the differences in magnitudes of income risk across industries and our estimates of changes in these magnitudes following trade policy changes). More precisely, across the manufacturing sectors we study, the mean industry wages are highly correlated with mean educational attainment. That is, the R^2 of a simple cross sectional regression of average earnings on average worker characteristics is about 0.8 (see the data prtesented in Table V). Thus unobserved worker characteristics have very little influence on average earnings in an industry, suggesting little selectivity of workers of differing (unobserved) abilities into different manufacturing sectors in our data.

III.4. Robustness

We conducted a series of additional estimation exercises to study the robustness of the findings reported here. First, the effective rate of protection was computed (using the tariff series and input-output matrices for Mexico) and used in place of the raw tariff series in estimating (10'). As the results presented in Table VI indicate, this does not change the results in any significant quantitative or qualitative way. Second, given that many of the right hand side variables were only observed on an annual basis, (10') was estimated using annually averaged observations (on income risk as well as the right hand side variables). These results, presented in Table VII, are also very similar to the ones we have reported before. More precisely, we calculated the average quarterly σ_ϵ^2 for each year and used these averages as the left hand side variable in (10'). Since in this case averaging reduces to a greater extent the variation in the left hand side variable, the degree of fit is now higher. To ensure that the dramatic nominal exchange rate devaluation undertaken by the Mexican authorities at the end of 1994 did not drive our results, (10') was estimated by dropping observations from the years 1995 and 1996. These results are also reported in Table VIII. As is evident, dropping observations from the years immediately following the exchange rate crisis in Mexico does not alter our results.

An additional point concerns the lagged effects of policy changes. Note that we measure tariff changes as the change between the beginning-of-year tariffs of two subsequent *years*. The corresponding change in income risk measures the average effect over a total of a two year period. Thus, a tariff change implemented at the beginning of 1988 could affect income risk in the last quarter of 1989, and this change in income risk would still be taken into account in our specification (10'). Estimation results (not reported here but available upon request) with specifications in which we included lagged tariff changes (and other lagged independent variables) on the right-hand-side of (10') did not support the inclusion of such lags.

Finally, experimenting with other specifications with additional interactive and non-linear terms did not reveal any significant or systematic patterns in the data.

IV. Income Risk and Welfare

The preceding discussion has outlined our approach to estimating the relationship between trade policy and income risk. We now turn to the analysis of the link between income risk and welfare, which is provided by a simple dynamic model with incomplete markets along the lines of Constantinides and Duffie (1996) and Krebs (2004). The model extends the basic insights of the large literature on the permanent income hypothesis to a general-equilibrium setting with iso-elastic preferences and incomplete markets.[31] It remains tractable enough

[31]Deaton (1991) and Carroll (1997) provide a quantitative analysis of the consumption-saving problem

to permit closed-form solutions for equilibrium consumption and welfare, yet is rich enough to provide a tight link to the empirical analysis. Clearly, our goal here is not to provide a complete assessment of the effects of income risk on welfare taking into account all possible channels, but rather to articulate a simple framework that allows us to obtain *indicative* estimates of welfare change through the *income risk* channel.

The model features long-lived workers that make consumption/saving choices in the face of uninsurable income shocks. These income shocks are permanent, which implies that "self-insurance" through borrowing or own saving is an ineffective means to smooth out income fluctuations. In other words, the effect of permanent income shocks on consumption is substantial.[32] In accordance with Constantinides and Duffie (1996) and Krebs (2004), we consider an exchange economy. Thus, we rule out by assumption any effect of changes in income risk on aggregate output. In this section, we briefly discuss the basic assumptions of the model and state the main welfare results. All derivations are relegated to the Appendix.

IV.1. Model

Time is discrete and open ended. Income of worker i employed in industry j in period t is denoted by $\tilde{y}_{ijt}$. Income is random and defined by an initial level $\tilde{y}_{ij0}$ and the law of motion

$$\tilde{y}_{ij,t+1} = (1 + \mu_{j,t+1})(1 + \theta_{ij,t+1})\,\tilde{y}_{it} \ , \tag{11}$$

where $\mu_{j,t+1}$ is a mean growth-rate effect common across workers in the sector and $\theta_{ij,t+1}$ is

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with permanent income shocks in a partial equilibrium context (exogenous interest rate).

[32]Krebs (2003a) considers a production economy with only permanent income shocks, and shows again that self-insurance is highly ineffective. Thus, the result that self-insurance is not very effective does not depend on the zero aggregate saving feature of endowment economies, even though we will make it to simplify the analysis. Notice also that there are differences between the current analysis and the work by Constantinides and Duffie (1996) and Krebs (2004). First, Constantinides and Duffie (1996) and Krebs (2004) focus on the asset price implications of market incompleteness, whereas the current analysis explores the welfare effects. Second, Constantinides and Duffie (1996) and Krebs (2004) consider a one-sector economy. In contrast, the current model has multiple sectors (industries) that differ with respect to the amount of income risk households have to bear. Finally, we assume that households can save, but not borrow.

an individual-specific shock to the growth rate of income. We assume that $log(1+\theta_{ij,t+1})$ is normally distributed with time- and industry-dependent variance $\sigma^2_{j,t+1}$. Although the distribution of individual-specific shocks may change over time, the shocks are unpredictable in the sense that current and future shocks are uncorrelated. To ensure that workers are ex-ante identical, we also assume that the distribution of shocks is identical across workers.

Each worker begins life with no initial financial wealth. Workers have the opportunity to save at the common interest rate r_t, but they cannot borrow. Hence, the sequential budget constraint of worker i reads

$$\begin{aligned} a_{ij,t+1} &= (1+r_t)a_{ijt} + \tilde{y}_{ijt} - c_{ijt} \\ a_{ijt} &\geq 0 \quad , \quad a_{ij0} = 0 \ . \end{aligned} \tag{12}$$

Here c_{ijt} denotes consumption of worker i employed in industry j in period t and a_{ijt} his asset holdings at the beginning of period t (excluding interest payment in this period).

Workers have identical preferences that allow for a time-additive expected utility representation:

$$U(\{c_{ijt}\}) = E\left[\sum_{t=0}^{\infty}\beta^t u(c_{ijt})\right] \ . \tag{13}$$

Moreover, we assume that the one-period utility function, u, is given by $u(c) = \frac{c^{1-\gamma}}{1-\gamma}$, $\gamma \neq 1$, or $u(c) = log\, c$, that is, preferences exhibit constant degree of relative risk aversion γ.[33]

IV.2. Welfare

[33]The model can easily be extended to allow for an endogenous labor-supply decision. Suppose, for example, that $\tilde{y}_{ijt} = w_{jt}h_{ijt}l_{ijt}$, where w_{jt} is the wage rate per effective unit of labor, h_{ijt} is the stock of human capital (general and specific) of worker i, and l_{ijt} is the number of hours worked. Suppose further that h_{ijt} is stochastic and that idiosyncratic shocks to h_{ijt} are unpredictable (permanent) as in Krebs (2003a,b). Then a straightforward extension of the argument made in the appendix shows that the optimal labor choice, l_{ijt}, is independent of idiosyncratic shocks to h_{ijt} if preferences over consumption and leisure are homothetic with respect to consumption (as assumed above) and multiplicative in consumption and leisure. That is, permanent shocks to the hourly wage rate of workers will not change labor supply, and the welfare formula (13), respectively (14), is still valid.

In the Appendix, we derive an explicit formula for equilibrium welfare that depends on the preference parameters β and γ and the income parameters μ_{jt} and σ_{jt}^2, where σ_{jt}^2 is the variance of the log-normally distributed income shocks θ. We also show that the variance σ_{jt}^2 of the income process (11) can be identified with the variance $\sigma_{\epsilon jt}^2$ of the permanent component of our empirical specification (1). This provides a tight link between the empirical results obtained in Section II and the welfare analysis conducted in this section. We now briefly outline and discuss the main welfare results.

For simplicity, assume that the income parameters are time- and industry-independent: $\mu_{jt} = \mu$ and $\sigma_{\epsilon jt}^2 = \sigma_\epsilon^2$. Suppose further that trade reform changes the tariff rate from τ to $(1 + \Delta_\tau)\tau$ permanently, and that this change in the tariff rate leads to a corresponding permanent change in income risk from σ_ϵ^2 to $(1 + \Delta_\sigma)\sigma_\epsilon^2$. Clearly, the change in income risk $\Delta_\sigma \sigma_\epsilon^2$ corresponds to the long-run effect that is associated with the level term, $\alpha_\tau \tau$, on the right-hand-side of our regression equation (10). We can find the welfare effect of the change in risk, Δ_σ, by calculating the compensating variation in lifetime consumption, Δ_c. That is, we can ask by how much we have to change consumption in each period and state of the world to compensate the household for the change in income risk. In the appendix we show that this compensating differential, expressed as percent of lifetime consumption, is given by

$$
\begin{aligned}
\Delta_c &= \left(\frac{1 - \beta(1+\mu)^{1-\gamma} exp\left(.5\gamma(\gamma-1)(1+\Delta_\sigma)\sigma_\epsilon^2\right)}{1 - \beta(1+\mu)^{1-\gamma} exp\left(.5\gamma(\gamma-1)\sigma_\epsilon^2\right)} \right)^{\frac{1}{1-\gamma}} - 1 \quad if\ \gamma \neq 1 \\
\Delta_c &= exp\left(\frac{\beta}{(1-\beta)^2} \frac{\sigma_\epsilon^2 \Delta_\sigma}{2} \right) - 1 \quad if\ \gamma = 1\ . \qquad (14)
\end{aligned}
$$

Equation (14) shows how to translate long-run changes in labor income risk, Δ_σ, into equivalent changes in lifetime consumption, Δ_c. It provide the answer to the following question: how much lifetime consumption are risk averse workers willing to give up in return for not having to experience the increase in income risk that is caused by a change in trade policy. Notice that (14) is the result of an ex-ante welfare calculation under rational expectations.

More specifically, (14) assumes that workers do not know who will lose and who will gain from trade reform, but they know to what extent trade reform creates winners and losers (the effect of trade reform on the income risk parameters is known ex-ante).

The welfare expression (14) assumes that the change in σ_ϵ^2 is permanent. However,we are also interested in the welfare effect of an increase in income risk from σ_ϵ^2 to $(1+\Delta_\sigma)\sigma_\epsilon^2$ for n periods. In this case, the welfare effect is given by

$$\begin{aligned} \Delta_c &= \left[\left(\frac{1-x}{1-x'}\right)\left(1-x'^{n+1}\right)+xx'^{n}\right]^{\frac{1}{\gamma-1}}-1 \quad if\ \gamma\neq 1 \\ \Delta_c &= exp\left(\frac{\beta(1-\beta^n)}{2(1-\beta)^2}\sigma_\epsilon^2\,\Delta_\sigma\right)-1 \quad otherwise \end{aligned} \tag{15}$$

where we introduced the following notation:

$$\begin{aligned} x &= \beta(1+\mu)^{1-\gamma}exp\left(.5\gamma(\gamma-1)\sigma_\epsilon^2\right) \\ x' &= \beta(1+\mu)^{1-\gamma}exp\left(.5\gamma(\gamma-1)(1+\Delta_\sigma)\sigma_\epsilon^2\right)\ . \end{aligned}$$

The welfare expressions (14) and (15) have some intuitive properties. First, the welfare effect of a change in income risk is a nonlinear and increasing function of the initial level of income risk. Put differently, if workers are already exposed to a large amount of income risk, then increasing income risk hurts a lot. This property explains why the welfare effects we find in this paper (see below) are so much larger than the welfare cost of business cycles found in the macroeconomic literature (Lucas, 2003). Second, the welfare effects are increasing in the risk-aversion parameter γ: the more risk-averse the workers are, the stronger is the welfare effect of a change in income risk. Finally, the welfare effects are the same for all workers regardless of their wealth. This property is the result of the joint assumption of homothetic preferences and an income process defined as in (11).

IV.3. Results

The welfare expressions (14) and (15) form the basis for our quantitative welfare analysis of trade reform. In order to conduct such an analysis, we need information about the income parameters μ, σ_ϵ^2, and Δ_σ and the preferences parameters β and γ. Our empirical analysis provides estimates of the income parameters. We estimate an average value of a quarterly variance, σ_ϵ^2, of of .008 (averaged across industries and over time), and this is also the value we use in all welfare calculations reported below. Similarly, we choose a quarterly growth rate $\mu = .005$ to match the average growth rate in aggregate real income in Mexico over the relevant sample period. For the preference parameters, we choose a quarterly discount factor of $\beta = .99$ and a degree of risk aversion of $\gamma = 1$ (log-utility) for the baseline economy. These values for the preference parameters are in line with the values used in the macroeconomic literature (Cooley (1995)). However, we also report the welfare results for a higher degree of risk aversion ($\gamma = 2$).

We conduct the following exercises. Starting from a tariff level of $\tau = .10$, which is roughly the average tariff level in our data set, we consider the welfare consequences of reducing the tariff level to $\tau' = .05$. Our empirical analysis in section III suggests that this tariff reduction has two effects for industries with high import penetration. First, there is a short-run effect that leads to an increase in income risk for one year (four quarters), and in this section we evaluate the welfare cost of this short-run effect. Second, there is an "interaction effect", and we report the welfare cost corresponding to this effect as follows. We compute the welfare cost of a short-run increase in income risk following a real exchange rate appreciation of ten percent with the tariff level also at ten percent, and then compare this welfare cost of the same exchange rate appreciation with the cost that obtains when the prevalent tariff level were five percent instead. Finally, we consider the welfare effect of a change in income risk due to a downturn in the economy, with the growth rate of GDP lowered by five percent, and again see how this is altered if the tariff level were lowered by five percent.

Table IX reports the effects of a trade reform that lowers tariff rates from ten percent to five percent for industries with high levels of import penetration. As indicated in Table IX, this would raise σ_ϵ^2 for one year following the reform from a mean level of 0.08 to 0.013 (here we use our regression results from equation (10') reported in Table VI). The corresponding welfare cost of this change is calculated to be 0.98 percent of permanent consumption if the co-efficient of risk aversion is $\gamma = 1$, and this cost increases to 1.96 percent of lifetime consumption if we choose $\gamma = 2$ instead.[34] Now consider the indirect effects of trade policy as measured by the interaction terms in (10'). As noted above, an exchange rate appreciation of ten percent raises σ_ϵ^2 for a year from 0.008 to 0.011 if the tariff level is ten percent. This translates into a welfare cost of 0.59 percent of lifetime consumption if $\gamma = 1$ and 1.18 percent if $\gamma = 2$. If the tariff rate were lowered to five percent, however, σ_ϵ^2 rises to 0.014 and the corresponding welfare costs are 1.18 and 2.36 percent of lifetime consumption, respectively. Finally, if the tariff rate is ten percent, a cyclical downturn in the economy (a drop in g by five percent) raises σ_ϵ^2 for a year from 0.008 to 0.010, and the corresponding welfare cost is calculated to be 0.39 percent of lifetime consumption if $\gamma = 1$ and 0.78 percent with $\gamma = 2$. In contrast, if the tariff rate were lowered to five percent, σ_ϵ^2 rises to 0.013 instead, and the corresponding welfare costs are 0.98 and 1.96 percent of lifetime consumption, respectively. Thus, our calculation suggest that both the short-run direct effects of tariff reforms and the indirect effects of the level of the tariff in amplifying the effects of macroeconomic shocks are economically significant.

As we have mentioned before, the limited time series dimension of our income data might lead us to over-estimate the amount of permanent labor income risk Mexican workers face.

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[34] Although the welfare formula (15) is non-linear in γ, this non-linearity is not very pronounced for moderate degrees of risk aversion. For example, if $\gamma = 4$, then the welfare cost of this short-run change in σ_ϵ^2 is 4.16% of lifetime consumption. Notice also that the results reported in Table IX assume $n = 4$ since we use quarterly risk and preference parameters and the increase in income risk lasts for four quarters (one year).

Consequently, the welfare results reported in Table IX might overstate the true cost of trade liberalization. We therefore also calculate the welfare effects for an economy in which the average income risk, σ_ϵ^2, and all changes in income risk, $\Delta\sigma_\epsilon^2$, are scaled down by a factor of 0.7. The factor 0.7 is derived from the fact that the estimate of income risk, σ_ϵ^2, obtained by Carroll and Samwick (1997), Gourinchas and Parker (2002), Hubbard, Skinner and Zeldes (1994), and Storesletten, Telmer and Yaron (2004) using income data with a very long panel dimension is roughly 70 percent of our estimate of income risk using a much shorter panel dimension (see our discussion in Section II.5 for details). Using the scaled down values for income risk and income risk changes, we find the following welfare cost of a five percent tariff reduction for a degree of risk aversion of $\gamma = 1$. First, the one-year increase in income risk immediately following the tariff reduction is equivalent to a decrease in lifetime consumption by .68 percent. Second, a five percent decline in GDP growth leads to an increase in income risk that is equivalent to a loss of lifetime consumption of .27 percent before the tariff reduction, and this loss increases to .68 percent after the tariff reduction (that is, the difference is .41 percent). Thus, although the welfare cost of trade liberalization are somewhat smaller than for the baseline case, they are still quite substantial.

VI. Conclusions

This paper studies empirically the relationship between trade policy and *individual* income risk. The analysis proceeds in three steps. First, longitudinal data on Mexican workers are used to estimate individual income risk in various manufacturing sectors. Second, the variation in income risk and trade barriers – both over time and across sectors – is used to arrive at estimates of the relationship between trade policy and individual income risk. Finally, using the estimates of this relationship between trade policy and income risk, a simple dynamic general equilibrium model with incomplete markets is used to obtain estimates of the corresponding welfare effects.

Our findings can be summarized as follows. First, for industries with high levels of import penetration, trade policy changes have a significant short-run effect on income risk. Second, the effect of the tariff *level* on income risk is insignificant. Third, while the tariff *level* has an insignificant mean effect, it nevertheless changes the degree to which macroeconomic shocks affect income risk. Finally, the welfare costs associated with the estimated increases in income risk are substantial.

As we have pointed out before, the welfare results reported in this paper have to be interpreted with caution keeping in mind several limitations of our analysis. More specifically, we focus *exclusively* on the link between trade policy and individual income risk, and do not study how trade reform affects the mean of income growth. Second, our welfare calculations do not allow for the possibility that an increase in income risk might lead to a simultaneous rise in insurance opportunities (endogenous market incompleteness). Third, we follow a long-standing tradition in economics and measure risk by the variance (second moment) of the relevant distribution, which is justified if (as assumed in this paper) the economic variables of interest are (log)-normally distributed. Finally, the Mexican household survey we use to implement our general approach is a rotating panel that follows individual workers for five quarters over time, which means that the panel dimension of our income data is somewhat limited. In short, the welfare results presented here do not show that trade liberalization is necessarily costly, but they do provide strong evidence that any comprehensive welfare analysis of trade liberalization ought to take into account the cost of increased labor market risk.

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Appendix

In this appendix, we construct the equilibrium and derive the welfare expressions. Notice first that the Euler equation associated with the consumption/saving problem of household i reads

$$c_{ijt}^{-\gamma} \geq \beta(1+r_{t+1})E[c_{ij,t+1}^{-\gamma}|\mathcal{F}_{ijt}] \ , \tag{A1}$$

where $\mathcal{F}_{ijt}$ is the information that is available to household i in period t and (A1) holds with equality if $a_{ijt} > 0$. The Euler equation (A1) says that the utility cost of saving one more unit of consumption is greater than or equal to the expected utility gain of doing so. If we rule out international borrowing and lending, then the domestic interest rate is determined by the saving decisions of domestic households only.[35] In this case, domestic asset market clearing reads:

$$\sum_{i,j} a_{ijt} = 0 \ . \tag{A2}$$

Suppose the interest rate is

$$r_{t+1} = min_j \left\{ \frac{1}{\beta \left(1+\mu_{j,t+1}\right)^{-\gamma} E\left[\left(1+\theta_{ij,t+1}\right)^{-\gamma} |\mathcal{F}_{ijt}\right]} - 1 \right\} \ . \tag{A3}$$

Notice that the right-hand side of (A3) does not depend on i because of our assumption that the distribution of $\theta_{ij,t+1}$ is independent of i and $\mathcal{F}_{ijt}$. Clearly, at the interest rate (A3) the Euler equation (A1) holds for all households i if they all choose $a_{ijt} = 0$ and $c_{ijt} = \tilde{y}_{ijt}$. Moreover, a tedious but straightforward argument show that expected lifetime utility is finite and that a corresponding transversality condition holds if (Krebs, 2004)

$$\beta(1+\mu_{j,t+1})^{1-\gamma} E\left[\left(1+\theta_{ij,t+1}\right)^{1-\gamma}\right] < 1 \ . \tag{A4}$$

Thus, the plan $a_{ijt} = 0$ and $c_{ijt} = \tilde{y}_{ijt}$ is individually optimal for all households. Since $a_{ijt} = 0$ satisfies market clearing, we have found an equilibrium.

We now turn to the welfare analysis. For simplicity, suppose that tariff rates and income parameters are constant over time and equal across industries: $\tau_{jt} = \tau$, $\mu_{jt} = \mu$, and $\sigma_{jt}^2 = \sigma^2$. If $c_{ijt} = \tilde{y}_{ijt}$ and there are no aggregate fluctuations, then expected lifetime utility (13) becomes

$$\begin{aligned} U_i &= \frac{c_{i0}^{1-\gamma}}{(1-\gamma)\left(1-\beta(1+\mu)^{1-\gamma}E[(1+\theta)^{1-\gamma}]\right)} \quad if \quad \gamma \neq 1 \\ U_i &= \frac{1}{1-\beta} log c_{i0} + \frac{\beta}{(1-\beta)^2}\left(log(1+\mu) + E[log(1+\theta)]\right) \quad otherwise \end{aligned} \tag{A5}$$

[35] Clearly, an alternative interpretation is that the model describes a small open economy with exogenous interest rate that is at least as low as (A3).

where the expectation is taken over idiosyncratic shocks (over the random variable θ) and for simplicity we dropped the industry-index j on c_{ij0}. Using the assumption that $\theta \sim N(-.5\sigma^2, \sigma^2)$, integration over income shocks yields

$$
\begin{aligned}
U_i &= \frac{c_{i0}^{1-\gamma}}{(1-\gamma)\,(1-\beta(1+\mu)^{1-\gamma} exp\,(.5\gamma(\gamma-1)\sigma^2))} \quad if \quad \gamma \neq 1 \qquad (A6)\\
U_i &= \frac{1}{1-\beta} log c_{i0} + \frac{\beta}{(1-\beta)^2}\Big(log(1+\mu) - \sigma^2/2\Big) \quad otherwise\ .
\end{aligned}
$$

Equation (A6) shows how welfare depends on income risk, σ^2, which in turn depends on tariff rates, τ. Thus, the welfare expression (A6) can be used to calculate how trade reform affects welfare through its effect on income risk. Clearly, this change in income risk induced by trade reform corresponds to the long-run effect that is associated with the level term, τ_{jt}, on the right-hand-side of our regression equation (10) and (10'). In order to get numbers for these welfare changes with economically meaningful units, we calculate the percentage change in initial consumption, c_{i0}, that is necessary to compensate the worker for the change in risk. More precisely, for any c_{i0}, σ^2, and Δ_σ, we are searching for the percentage change in initial consumption, Δ_c solving

$$
U(c_{i0}, \sigma^2) = U\Big((1+\Delta_c)c_{i0}, (1+\Delta_\sigma)\sigma^2\Big) \qquad (A7)
$$

Notice that because of our random walk assumption, any increase in initial consumption, c_{i0}, amounts to an increase in consumption for all future dates and events (lifetime consumption). Using (A6) and (A7), we find the welfare expression(14). Notice that expression (14) is independent of c_{i0}, that is, the welfare change expressed in percentage changes of lifetime consumption is the same for all workers.

So far, we have calculated the welfare effect of a permanent increase in σ^2. However,we are also interested in the welfare effect of an increase in income risk from σ^2 to $(1+\Delta_\sigma)\sigma^2$ for n periods. In this case, expected lifetime utility of workers without the increase is still given by (A6), and expected lifetime utility with the increase is:

$$
\begin{aligned}
U_i' &= \sum_{t=0}^{n} \beta^t \frac{E[(c_{it}')^{1-\gamma}]}{1-\gamma} + \sum_{t=n+1}^{\infty} \beta^t \frac{E[(c_{it}')^{1-\gamma}]}{1-\gamma}\\
E[(c_{it}')^{1-\gamma}] &= \frac{c_{i0}^{1-\gamma}}{1-\gamma}(1+\mu)^{(1-\gamma)t}\Big(E[(1+\theta')^{1-\gamma}]\Big)^t \quad t = 0, 1, \ldots, n \qquad (A8)\\
E[(c_{it}')^{1-\gamma}] &= \frac{c_{i0}^{1-\gamma}}{1-\gamma}(1+\mu)^{(1-\gamma)t}\Big(E[(1+\theta')^{1-\gamma}]\Big)^n\Big(E[(1+\theta)^{1-\gamma}]\Big)^{(t-n)} \quad t = n+1, n+2, \ldots
\end{aligned}
$$

where $log(1+\theta) \sim N(-\sigma^2/2, \sigma^2)$ and $log(1+\theta') \sim N(-\sigma^2(1+\Delta_\sigma)/2, \sigma^2(1+\Delta_\sigma))$. A similar expression holds for the case of log utility. We define again the welfare cost of trade reform, Δ_c, as the increase in average consumption that is necessary to compensate workers for the (n-period) increase in income risk. Using this definition and evaluating the expression (A8), we find the welfare expression (15) in section IV.

Finally, let us discuss the link between the specification of the income process (1)-(3) in the Section II and the income process (11) used in the Section IV. Recall that we assume that $log(1+\theta)$ in (11) is normally distributed. More specifically, we assume $log(1+\theta_{ij,t+1}) \sim N(-\sigma^2_{j,t+1}/2, \sigma^2_{j,t+1})$. The term $-.5\sigma^2_{j,t+1}$ ensures that the mean of income growth is independent of $\sigma^2_{j,t+1}$, a property that is useful since it allows us to vary income risk without changing the mean growth rate. Notice that this type of specifying the distribution of income shocks is standard in the asset pricing and macroeconomic literature (Carroll, 1997, and Constantinides and Duffie, 1996). To understand the economic meaning of this assumption, notice that with this specification we have $E[\theta_{ij,t+1}] = 0$ and $var[\theta_{ij,t+1}] = e^{\sigma^2_{j,t+1}}(e^{\sigma^2_{j,t+1}} - 1)$ using the standard formula for log-normal distributions (see, for example, Campbell, Lo, and MacKinlay 1997). Thus, any increase in $\sigma^2_{j,t+1}$ increases $var[\theta_{ij,t+1}]$, but leaves $E[\theta_{ij,t+1}]$ unchanged. Taking the logarithm in (11), we find

$$log\,\tilde{y}_{ij,t+1} = log\,\tilde{y}_{ijt} + log(1+\mu_{j,t+1}) + log(1+\theta_{ij,t+1}) \qquad (A9)\ .$$

Thus, income follows a logarithmic random walk with drift $log(1+\mu_{j,t+1})$ and heteroscedastic error term $log(1+\theta_{ij,t+1})$. Comparison of (A9) with the econometric specification (1)-(3) suggests that we relate $\log(1+\theta_{ij,t+1})$ in (A9) with the innovation term of the permanent, unpredictable component of income changes in (1):

$$log(1+\theta_{ij,t+1}) = \epsilon_{ij,t+1} - \sigma^2_{j,t+1}/2 \qquad (A10)\ .$$

In (A10) we introduce the term $-\sigma^2_{j,t+1}/2$ to ensure that both random variables have the same mean. Taking the variance in (A10) we find

$$\sigma^2_{j,t+1} = \sigma^2_{\epsilon j,t+1} \qquad (A11)\ .$$

Thus, our empirical measure of income risk, σ^2_ϵ, coincides with our theoretical measure of income risk, σ^2. This shows that we can use our empirical estimates of σ^2_ϵ obtained in Section II when evaluating the welfare expressions (14) and (15) in Section IV.

Table I: ENEU Worker Survey - Summary (1987-1998)

Variables	
Mean Age	32
Mean Years of Education	8
Fraction High School and Above	17
Fraction Wage Earners	65
Fraction Self Employed	25

Table II: Estimates of Persistent and Transitory Income Shocks*
Annual Averages (1987-1998)

Year	σ_ϵ^2	σ_η^2	Sample Size
87	0.011	0.096	19136
	(0.003)	(0.002)	
88	0.005	0.101	35397
	(0.003)	(0.002)	
89	0.004	0.103	28203
	(0.002)	(0.001)	
90	0.014	0.098	35167
	(0.002)	(0.001)	
91	0.001	0.103	37344
	(0.002)	(0.001)	
92	0.006	0.106	54022
	(0.001)	(0.001)	
93	0.007	0.112	78741
	(0.001)	(0.001)	
94	0.006	0.110	121716
	(0.001)	(0.001)	
95	0.014	0.118	164212
	(0.001)	(0.001)	
96	0.000	0.107	172766
	(0.001)	(0.001)	
97	0.006	0.104	172870
	(0.001)	(0.001)	
98	0.008	0.097	158707
	(0.001)	(0.001)	

*Figures shown are annual averages (across industries and quarters) of the point estimates of the persistent shock σ_ϵ^2 and the transitory shock σ_η^2. The figures in parentheses are the averages of the corresponding standard errors. Sample size denotes the numbers of workers surveyed in the respective year.

Table III: Estimates of Persistent and Transitory Income Shocks*
Industry Averages (1987-1998)

Industry	σ^2_ϵ	σ^2_η	Industry	σ^2_ϵ	σ^2_η
311	0.013 (0.0004)	0.131 (0.0003)	352	0.020 (0.0025)	0.111 (0.0019)
313	0.012 (0.0007)	0.088 (0.0005)	353	0.002 (0.0009)	0.081 (0.0007)
321	0.005 (0.0006)	0.097 (0.0005)	356	0.006 (0.0016)	0.079 (0.0011)
322	0.012 (0.0008)	0.124 (0.0006)	369	0.011 (0.0014)	0.113 (0.0011)
323	0.008 (0.0022)	0.107 (0.0015)	371	0.003 (0.0031)	0.110 (0.0025)
324	0.004 (0.0002)	0.088 (0.0001)	381	0.006 (0.0006)	0.125 (0.0004)
331	0.004 (0.0027)	0.120 (0.0020)	382	-0.002 (0.0015)	0.098 (0.0011)
332	0.019 (0.0017)	0.121 (0.0013)	383	0.008 (0.0002)	0.056 (0.0002)
341	0.004 (0.0016)	0.102 (0.0012)	384	0.004 (0.0002)	0.073 (0.0001)
342	0.011 (0.0016)	0.134 (0.0012)	390	0.005 (0.0062)	0.143 (0.0047)
351	0.012 (0.0029)	0.107 (0.0023)			

*Figures shown are averages over time of the point estimates of the persistent shock σ^2_ϵ and the transitory shock σ^2_η for the respective industries. The figures in parentheses are the averages of the corresponding standard errors.

Table IV: Trade Policy and Income Risk - Panel Estimates [†]

Variables	σ_ϵ^2 vs	σ_ϵ^2 vs
τ	0.043 (0.060)	-0.140 (0.051)
$\Delta\tau$	-0.035 (0.044)	0.017 (0.031)
$\Delta\tau \cdot D_n$	-0.090 (0.047)	-0.109 (0.047)
Δe		-0.621 (0.207)
g		-1.208 (0.414)
$\tau \cdot \Delta e$		0.539 (0.184)
$\tau \cdot g$		1.055 (0.370)
Time Effects	Included	
Industry Fixed Effects	Included	Included
N	945	945
R^2	0.058	0.044

[†]Figures in parentheses are robust standard error estimates obtained by allowing for heteroscedasticity, contemporaneous correlation of errors across industries and serial correlation within industries.

Table V: Industry Average Characteristics (1997)*

Industry	Age	Education	Wage
311	32.11	7.98	14.52
313	31.45	9.76	24.80
321	33.31	8.69	17.09
322	30.02	8.44	13.50
323	29.76	7.82	17.42
324	29.55	7.14	15.66
331	30.83	8.77	14.40
332	30.99	8.31	17.44
341	30.05	8.69	18.31
342	31.68	10.77	23.55
351	34.41	11.93	50.63
352	32.75	11.22	30.06
353	38.54	11.83	41.58
356	30.27	9.16	19.43
369	33.98	7.79	19.27
371	36.31	11.07	47.89
381	32.20	8.85	18.51
382	30.98	10.50	25.91
383	28.81	9.60	23.19
384	29.40	10.12	24.90
390	29.93	9.05	13.92

*Age and education are average age and education of the labor force measured in years. Wage denotes the average monthly wage in thousands of Pesos.

Table VI: Trade Policy and Income Risk - Effective Rates of Protection†

Variables	σ^2_ϵ vs	σ^2_ϵ vs
τ	0.019 (0.043)	-0.109 (0.045)
$\Delta\tau$	-0.009 (0.032)	0.015 (0.026)
$\Delta\tau \cdot D_n$	-0.076 (0.042)	-0.098 (0.042)
Δe		-0.463 (0.179)
g		-0.935 (0.345)
$\tau \cdot \Delta e$		0.397 (0.157)
$\tau \cdot g$		0.807 (0.307)
Time Effects	Included	
Industry Fixed Effects	Included	Included
N	945	945
R^2	0.058	0.042

†Figures in parentheses are robust standard error estimates obtained by allowing for heteroscedasticity, contemporaneous correlation of errors across industries and serial correlation within industries.

Table VII: Trade Policy and Income Risk - Annual Estimates of σ_ϵ^2 [‡]

Variables	σ_ϵ^2 τ	σ_ϵ^2 ERP
τ	-0.132 (0.061)	-0.103 (0.056)
$\Delta\tau$	0.017 (0.038)	0.007 (0.028)
$\Delta\tau \cdot D_n$	-0.094 (0.035)	-0.081 (0.038)
Δe	-0.635 (0.229)	-0.485 (0.231)
g	-1.162 (0.537)	-0.910 (0.447)
$\tau \cdot \Delta e$	0.549 (0.204)	0.413 (0.200)
$\tau \cdot g$	1.010 (0.486)	0.781 (0.400)
Industry Fixed Effects	Included	Included
N	252	252
R^2	0.13	0.14

[‡] Figures in parentheses are robust standard error estimates obtained by allowing for heteroscedasticity, contemporaneous correlation of errors across industries and serial correlation within industries.

Table VIII: Trade Policy and Income Risk - Robustness[§]

Variables	σ_ϵ^2 AD Excluded	σ_ϵ^2 95-96 Excluded
τ	-0.133 (0.052)	-0.150 (0.055)
$\Delta\tau$	0.034 (0.031)	0.028 (0.032)
$\Delta\tau \cdot D_n$	-0.113 (0.048)	-0.116 (0.046)
Δe	-0.608 (0.212)	-0.540 (0.226)
g	-1.126 (0.425)	-1.303 (0.466)
$\tau \cdot \Delta e$	0.531 (0.188)	0.472 (0.199)
$\tau \cdot g$	0.985 (0.379)	1.123 (0.414)
Industry Fixed Effects	Included	Included
N	809	861
R^2	0.04	0.045

[§]Figures in parentheses are robust standard error estimates obtained by allowing for heteroscedasticity, contemporaneous correlation of errors across industries and serial correlation within industries. In the first column (marked 'AD Excluded'), observations from industries with high levels of anti-dumping protection were excluded. In the second column (marked '95-96 Excluded'), observations from the years 1995 and 1996 have been excluded. See Section VI for a detailed discussion.

Table IX: Welfare Effects[¶]

	Change in σ_ϵ^2 ($\bar{\sigma_\epsilon^2} = 0.008$)	Welfare Change $\gamma = 1$	Welfare Change $\gamma = 2$
Trade Reform			
τ reduced by five percent	0.005 (0.002)	0.98 (0.39)	1.96 (0.79)
Macroeconomic Factors (τ level = ten percent)			
g lower by five percent	0.002 (0.001)	0.39 (0.20)	0.78 (0.40)
e appreciation by ten percent	0.003 (0.001)	0.59 (0.20)	1.18 (0.39)
Macroeconomic Factors (τ level = five percent)			
g lower by five percent	0.005 (0.001)	0.98 (0.29)	1.95 (0.59)
e appreciation by ten percent	0.006 (0.002)	1.18 (0.40)	2.36 (0.80)

[¶]Welfare changes are measured in compensating variation terms and denote the change in lifetime consumption necessary to compensate agents for the short term (one year) increases in σ_ϵ^2 (relative to its sample mean of 0.008) that result under the exercises being considered. γ denotes the co-efficient of relative risk aversion. Standard errors for the estimated welfare effects were obtained by simulation.