Trade Policy, Income Risk, and Welfare*

Tom Krebs†
Brown University

Pravin Krishna
Brown University

William Maloney
World Bank

First Draft: March 2003

This Draft: April 2004

Abstract

This paper studies empirically the relationship between trade policy and individual income risk faced by workers. 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. The estimated income risk parameters and data on trade barriers are then used to analyze the relationship between trade policy and income risk. Finally, a simple dynamic general equilibrium model with incomplete markets is used 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.

JEL Classification: D 52, E21, F1, F13
Keywords: Trade Policy, Income Risk, Macroeconomics

---

*For helpful comments and suggestions, we would like to thank Pol Antras, Michele Boldrin, Herschel Grossman, Peter Howitt, Kishore Gawande, Bojan Jovanovic, Marc Melitz, Devashish Mitra, and David Weil. We are also grateful to seminar participants at Brown, Columbia, Georgetown, Harvard, NYU, Princeton, University of Pennsylvania, Rutgers and Syracuse for helpful comments, and Jungjin Lee for outstanding research assistance.

†Corresponding Author. Department of Economics, 64 Waterman Street, Providence, RI 02906. Email: Tom_Krebs@Brown.edu

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 asses 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 is likely to raise individual labor income risk.[3] Rodrik (1997), going beyond the short term

[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 recent survey treatment.

[3]See, for instance, Fernandez and Rodrik (1991) in which ex-ante identical workers experience heterogeneous outcomes following a trade policy change. See also the analysis of Melitz (2003) for an example of an aggregate policy shock affecting an entire sector leading to heterogeneous outcomes for individual firms within that sector.

reallocational 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. 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.[4]

In this paper, we study *empirically* the effects of trade policy on individual income risk using the following approach. First, for each industry (sector), we use longitudinal data on individual incomes to estimate time-varying parameters of individual income risk (defined as the variance of unpredictable changes in individual income). In this first step, we are careful to distinguish between transitory and persistent shocks to income since the two types of shocks have very different welfare implications. More specifically, workers can effectively self-insure against transitory shocks through saving, which implies that these type of shocks have only small effects on consumption and welfare.[5] Our focus in this paper is therefore on persistent shocks to income. Using the estimates of individual income risk thus obtained, we then investigate empirically the relationship between income risk and trade policy.

In addition to analyzing empirically the relationship between trade policy and income risk, this paper also provides a quantitative evaluation of the welfare consequences of any changes

[4]Clearly, this sign-ambiguity does not extend to the short-term reallocational 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.

[5]See, for instance, Aiyagari (1994) and Levine and Zame (2002).

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, and Krusell and Smith, 1998). 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, 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.

To study the link between trade policy and individual income risk empirically, it is necessary to have longitudinal information on incomes at a disaggregated level (individual or household)[6] 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 our empirical implementation, then, we focus on one country that satisfies both criteria, namely Mexico. As is well known, the Mexican economy experienced substantial changes in trade policy in the late 1980's and in the later half of the

[6]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.

1990s.[7] Moreover, as we discuss in detail later in this paper, the Mexican government, since the mid-1980's, has conducted quarterly longitudinal income surveys that comprehensively surveyed workers in all manufacturing sectors of the economy − providing the unique data source that we use in our study.

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, with a tariff reform (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).[8] 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. 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

[7] In an early wave of trade reforms in the late 1980s, tariffs were cut from an average of about 40 percent to about 15 percent.

[8] 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.

a percentage point of lifetime consumption.[9]

At this stage, it is worth pointing out that our welfare analysis focuses *exclusively* on the link between trade policy and individual income risk, and that other possible channels through which trade policy may affect the economy are not studied here. More specifically, we would expect trade reform to have positive effects on the efficiency of resource allocation and economic growth, and these effects are important factors that should be taken into account when evaluating the total costs and benefits of trade reform. Additionally, our welfare calculations are based on a simple theoretical model whose limitations include its neglect of the effect of income risk on labor supply and capital accumulation.[10] Thus, the welfare results presented in this paper have to be interpreted with caution keeping in mind our exclusive focus on the link between trade policy and income risk and the methodological limitations noted above.

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. It is worth emphasizing that the type of study we conduct here is the first of its kind. While several scholars have commented upon the potential importance of the link between openness and income risk, and while some attempts have

[9]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 (2003a) and Lucas (2003) for more details.

[10]See, for example, Aiyagari (1994) for physical capital accumulation and Krebs (2003b) for human capital accumulation.

been made to estimate the relationship between openness and *aggregate* volatility,[11] none has studied the relationship between openness and *individual* income risk in the manner or detail that we do here.

The rest of the paper proceeds as follows. Section II describes the estimation procedure and data that we use to estimate individual income risk. Section III discusses the empirical methodology we use in a second stage to find estimates of the relationship between income risk and trade policy. Section IV describes the theoretical framework that will be used to translate changes in income risk into changes in welfare. Section V presents our results. Section VI concludes.

II. Income Risk

The first stage of our analysis concerns the estimation of individual income risk, where income risk is defined as the variance of unpredictable changes in individual income. In this first stage, we will distinguish between transitory and persistent shocks to income. From a welfare point of view this separation is essential since self-insurance through saving works well for transitory income shocks, but not for persistent ones (Aiyagari, 1994, and Levine and Zame, 2002). For this and other reasons (to be discussed in detail below), 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 the mid-1980s (we use data from 1987-1998 in our study). The sample is selected to be geographically

[11] See, for example, Rodrik (1998).

and socio-economically representative. The survey questionnaire is extensive in scope and covers all standard elements such as participation in the labor market, wages, hours worked, etc. The treatment of sample design, collection and data cleaning is careful.[12] The ENEU is structured so as to track a fifth of each sample across a five quarter period. To construct the panels, workers were matched by position in an identified household, level of education, age and sex to ensure against generating spurious transitions. Using just the first variables to concatenate and following changes in sex across the panel led to mismatching (or misreporting) of under .5 percent. Taken together, we have 44 complete panels of 5 periods (i.e., quarters) each, spanning a total of 12 years (48 quarters). The number of individuals surveyed in any given calender year is approximately 100,000. Table I presents a summary description of the workers surveyed by the ENEU.[13] Data on sectoral trade barriers and other sectoral and macroeconomic variables were obtained from the World Bank.

II.2. Specification

As in previous empirical work, we assume that the log of labor income (earnings) 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. 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.

We assume that the stochastic term is the sum of two (unobserved) components, a permanent

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

[13]See also Hanson (2003) for an a broad analytical discussion of wage levels in Mexico in the 1990s.

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 η_{ijt} are independently distributed over time. We further assume that they are normally distributed with zero mean. Clearly, η_{ijt} captures both temporary income shocks and measurement error. We assume that the variance of η_{ijt} is independent of i, but allow for time and industry dependence: $\eta_{ijt} \sim N(0, \sigma^2_{\eta jt})$.

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 (2002) 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. (2002), the previous literature has confined attention to the special case of time-independent variances (homoscedastic case). Clearly, 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.

In principle, both $\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$ represent measures of individual income risk. In this paper, we will focus on $\sigma_{\epsilon jt}$ and its relationship to trade policy. This choice is motivated by the following two considerations. First, as mentioned before, transitory income shocks are unlikely to generate consumption volatility since self-insurance through own-saving is highly effective, and the welfare effects of these shocks are therefore small (Aiyagari, 1994, and Levine and Zame, 2002). Second, term $\sigma^2_{\eta jt}$ is likely to contain a large amount of measurement error, and therefore overstates the degree of transitory income risk.

II.3. Estimation

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

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

Thus, we have the following expression for the variance of 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,[14] 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. (2002) and Gourinchas and Parker (2002).[15] Notice that the restrictions are linear in the parameters $\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$, which implies that the first-order conditions associated with the

[14]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.

[15]Notice that Meghir and Pistaferri (2004) and Storesletten et al. (2002) exploit additional moment restrictions that follow from the autocovariance function of income changes.

corresponding minimum-distance problem are linear in $\sigma^2_{\epsilon jt}$ and $\sigma^2_{\eta jt}$ – a feature that facilitates the estimation substantially. Since, for each time period, there are two parameters to be estimated and one moment condition corresponding to *each* time interval into the future, there are, in general, many more moment conditions than there are parameters. The system is thus (over) identified. Specifically, in our data set on Mexico, where individuals drop out of the sample after 5 quarters and where we have data spanning a total of 48 quarters, the number of parameters to be estimated is 2*(48) and the number of moment conditions is approximately 4*(48).[16]

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:

$$var[\Delta_n u_{ijt}] = 2\sigma^2_{\eta j} + n\,\sigma^2_{\epsilon j} \tag{6}$$

Thus, the variance of observed n-period income changes is a linear function of n, where 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 (6) on n. While the preceding example, with time-

[16]We should note that in forming the sample analogs of the moment condition (5), we use only those individuals who are present in the given industry in both time periods t and $t+n$. This allows us to circumvent the extremely difficult problem of assignment of industries (and thus trade policy) to individuals who transit industries during the time period in which they are observed. Including individuals who make transitions to the service sector (but not to other manufacturing sectors) by using the *ad hoc* procedure of counting them among those in the manufacturing sector in which they are first observed does not result in any qualitative difference in our reported results. It should perhaps also be noted that since transition of individuals from one manufacturing sector to another were relatively rare in our data, the exclusion of these individuals should not be expected to cause too great an under-estimation of our income risk parameters.

invariant parameters, serves to illustrate the intuition underlying the estimation procedure, we should note that our exercise is more general in the sense that it allows for arbitrary time variation in income risk parameters.

III. Trade Reform and Income Risk

The procedure outlined in the previous section provides us with estimates of individual income risk, $\sigma^2_{\epsilon jt}$, for each industry (i. e., manufacturing sector) j and time period, i.e., quarter, t. These time-varying, industry-specific estimates in conjunction with observations on trade policy, τ_{jt}, allow us to estimate the relationship between income risk, $\sigma^2_{\epsilon jt}$, and openness, τ_{jt}. Consider the following linear specification allowing for industry fixed-effects and aggregate time effects:

$$\sigma^2_{\epsilon jt} = \alpha_0 + \alpha_{1j} + \alpha_{2t} + \alpha_\tau \, \tau_{jt} + \alpha_\delta \, \Delta\tau_{jt} + \nu_{jt} \ . \tag{7}$$

In (7) the coefficients α_{1j} capture the industry fixed-effects, the α_{2t}'s pick up aggregate trends, the coefficient α_τ measures the effect of openness on income risk and α_δ captures the effects of changes (in the preceding year, say) in trade policy, $\Delta\tau_{jt}$. The inclusion of industry dummies in the specification above 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 conditions 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 (7) 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.[17]

Specification (7) provides the starting point for our econometric analysis. An alternate specification, which exploits to a greater extent the time variation in trade policy *within* each industry in the estimation of α_τ, is obtained by dropping the time dummies but controlling for relevant macroeconomic factors affecting income risk, S_t, by directly including them on the right hand side of the estimating equation. Allowing further for the possibility that trade policy affects the response of the economy to these macroeconomic factors gives us the following specification:

$$\sigma^2_{\epsilon jt} = \alpha_0 + \alpha_j + \alpha_\tau \, \tau_{jt} + \alpha_\delta \, \Delta\tau_{jt} + \beta \cdot S_t + \phi \cdot S_t \tau_{jt} + \nu_{jt} \tag{7'}$$

where β captures the effect of macroeconomic factors and ϕ captures the extent to which trade policy changes the effect of macroeconomic factors on income risk.

Several econometric issues arise in the estimation of equations (7) and (7') above. 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. A concern arises, however, 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. Further, 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

[17]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.

observable macroeconomic factors as in (7'), 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 (7) and (7') above are obtained by using robust estimation techniques.

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,[18]. It remains tractable enough to permit closed-form solutions for equilibrium consumption and welfare which are simple and transparent. 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. The model structure and assumptions underlying our approach and the limitations of our methodology are discussed below in detail.

The model features long-lived households (workers) that make consumption/saving choices in the face of uninsurable income shocks. Income shocks are permanent, which implies that self-insurance is an ineffective means to smooth out income fluctuations. In other words,

[18]See, for example, Deaton (1992) for a survey of the literature on the permanent income hypothesis.

the effect of permanent income shocks on consumption is substantial.[19] In accordance with Constantinides and Duffie (1996) and Krebs (2004), we consider an exchange economy and do not model the labor-leisure choice.[20] 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 household i employed in industry j in period t is denoted by y_{ijt}. Income is random and defined by an initial level $\tilde{y}_{ij0}$ and the law of motion

$$\tilde{y}_{ij,t+1} = (1 + \mu_{j,t+1})(1 + \theta_{ij,t+1})\, \tilde{y}_{it} \ , \tag{8}$$

where $\mu_{j,t+1}$ is a mean growth-rate effect common across workers in the sector and $\theta_{ij,t+1}$ is 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 σ^2_{jt}. 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 household begins life with no initial financial wealth. Households have the opportunity

[19] Krebs (2003b) 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 assumption that can be interpreted as reflecting lending and borrowing rates that are sufficiently different.

[20] More specifically, the model disregards the possibility that workers react to changes in the wage rate by substituting labor supply over time. Notice, however, that empirical micro-studies tend to find small intertemporal elasticities of labor supply (Altonji, 1986). Moreover, there are theoretical reasons to expect this intertemporal substitution effect to be small when, as assumed in this paper, wage shocks are permanent.

to save, but not borrow, at the common risk-free rate r_t. Hence, the sequential budget constraint of worker i reads

$$\begin{aligned} a_{ij,t+1} &= (1+r_t)a_{ijt} + y_{ijt} - c_{ijt} \\ a_{ijt} &\geq 0 \ , \ a_{ij0} = 0 \ . \end{aligned} \tag{9}$$

Here c_{ijt} denotes consumption of household i in period t and a_{ijt} his asset holdings at the beginning of period t (excluding interest payment in this period). Notice that by assuming the non-negativity of a_{ijt}, we have automatically ruled out Ponzi schemes.

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

$$U(\{c_{ijt}\}) = E\left[\sum_{t=0}^{\infty} \beta^t u(c_{ijt})\right] \ . \tag{10}$$

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.

IV.2. Welfare

As described 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^2_{jt}, where σ^2_{jt} is the variance of the log-normally distributed income shocks η. We also show that the variance σ^2_{jt} of the income process (8) can be identified with the variance $\sigma^2_{\epsilon jt}$ 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.

For simplicity, assume that the income parameters are time-independent: $\mu_{jt} = \mu_j$ and $\sigma^2_{\epsilon jt} = \sigma^2_{\epsilon j}$. Suppose now that trade reform changes the tariff rate in a particular industry j from τ to $(1+\Delta_\tau)\tau$ permanently, and that this change was not expected by workers. Suppose also that the 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, 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 (7). We can find the welfare effect of the change in risk, Δ_σ, by calculating the compensating variation in lifetime consumption, Δ_c.[21] 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

$$\begin{aligned}
\Delta_c &= \left(\frac{1-\beta(1+\mu)^{1-\gamma}exp\left(.5\left((1-\gamma)^2-(1-\gamma)\right)(1+\Delta_\sigma)\sigma_\epsilon^2\right)}{1-\beta(1+\mu)^{1-\gamma}exp\left(.5\left((1-\gamma)^2-(1-\gamma)\right)\sigma_\epsilon^2\right)}\right)^{\frac{1}{1-\gamma}} - 1 \quad if\ \gamma \neq 1 \\
\Delta_c &= exp\left(\frac{\beta}{(1-\beta)^2}\,\frac{\sigma_\epsilon^2\Delta_\sigma}{2}\right) - 1 \quad if\ \gamma = 1\ . \qquad (11)
\end{aligned}$$

Equation (11) shows how to translate long-run changes in labor income risk, Δ_σ, into equivalent changes in average consumption, Δ_c. Notice that expression (11) is the same for all workers since workers are ex ante identical.

The welfare expression (11) 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

$$\begin{aligned}
\Delta_c &= \left[\left(\frac{1-x}{1-x'}\right)\left(1-x'^{n+1}\right)+xx'^{n}\right]^{\frac{1}{\gamma-1}} - 1 \quad if\ \gamma \neq 1 \qquad (12) \\
\Delta_c &= exp\left(\frac{\beta(1-\beta^n)}{2(1-\beta)^2}\sigma_\epsilon^2\,\Delta_\sigma\right) - 1 \quad otherwise
\end{aligned}$$

where we introduced the following notation:

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

[21]Notice that for the case considered here, this compensating variation is equal to the equivalent variation.

The welfare expressions (11) and (12) 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. For the preference parameters, we choose an annual discount factor of $\beta = .96$ and a degree of risk aversion of $\gamma = 1$ or $\gamma = 2$. These values for the preference parameters are in line with the values used in the macroeconomic literature (Cooley, 1995).

It is worth emphasizing that the welfare analysis described here focuses *exclusively* on the link between trade policy and individual income risk, and other possible channels through which trade policy may affect the economy are not studied here. More specifically, we would expect trade reform to have positive effects on the efficiency of resource allocation and economic growth, and such effects are important factors that ought to be taken into account when evaluating the total costs and benefits of trade reform. Additionally, our welfare calculations are based on a simple theoretical model whose limitations include its neglect of the effect of income risk on labor supply and capital accumulation. Moreover, our calculations do not take into account that the welfare cost of an increase in income risk might be partially offset by a rise in transfer payments from the government or firms. The welfare estimates obtained in this exercise should therefore be seen as indicative and should be considered keeping the methodological limitations we have just noted firmly in mind.

V. Results

V.1. Trade Policy and Income Risk

In the first step of our analysis, we use data on individual income changes from workers in 21 different manufacturing sectors in Mexico and the methodology outlined in section II to estimate quarterly income risk parameters in each of these sectors during the time period 1987-1998. The mean value (across industries and over time) of the quarterly variance of the persistent shock, σ^2_ϵ, is estimated to be 0.008, or 0.032 in annual variance (i.e., σ_ϵ, is estimated to have a mean quarterly value of 0.09 and a mean annual value of 0.18).[22]

We analyze next the relationship between σ^2_ϵ and trade policy using specifications of the type discussed in Section III. Our first specification is

$$\sigma^2_{\epsilon jt} = \alpha_0 + \alpha_{1j} + \alpha_{2t} + \alpha_\tau \, \tau_{jt} + \alpha_{\delta 1} \, \Delta\tau_{jt} + \alpha_{\delta 2} \, \Delta\tau_{jt} D_{jt} + \nu_{jt}. \tag{7}$$

In (7) 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,[23] α_j $-$ an industry fixed- effect, and α_t $-$ a time dummy that captures general macroeconomic trends in the economy.

In (7), 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 I 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.

[22] 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 (and are measured less precisely as well).

[23] 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 (7') below.

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 here, on average, lowering the tariff rate by five percent would, for a year, raise σ_ϵ from a mean level of 0.009 to 0.012 (i.e., by more than thirty percent)− a substantial increase in the risk to income faced by individuals.

Our second specification is

$$\sigma^2_{\epsilon jt} = \alpha_0 + \alpha_j + \alpha_\tau \, \tau_{jt} + \alpha_{\delta 1} \, \Delta\tau_{jt} + \alpha_{\delta 2} \, \Delta\tau_{jt} D_{jt} + \beta_e \Delta e_t + \beta_g g_t + \phi_e \Delta e \tau_{jt} + \phi_g g \tau_{jt} + \nu_{jt}, \quad (7')$$

which exploits the within-industry variation in tariffs over time to a greater extent by dropping the time dummies and including instead 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 $\tau\Delta e$ and τg which measure the extent to which the relationship between income risk and these macroeconomic factors varies with trade policy.[24]

Estimates from (7') are presented in the second column of Table II. 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 (7) ($\hat{\alpha}_{\delta 1} + \hat{\alpha}_{\delta 2} =$ -0.092, with an estimated standard error of 0.045 − implying a twenty five percent increase in σ_ϵ with a five percent reduction in tariffs). 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

[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 (7') separately for industries with $D = 0$ and $D = 1$ gave results very similar to those reported here.

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, $\tau\Delta e$, and τg, are both positive and significant. The extent to which the tariff level alters the effects of exchange rate depreciation on income risk is given by ϕ_e. As reported in Table II, 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 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

[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 indicates 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. As such, evidence of the sector bias of skill biased technological change and its corelation with trade policy is quite scant (if anything, our own estimates of the returns to education suggest a striking similarity across manufacturing sectors in Mexico). We would also expect any such changes in technology to only impact income levels in a more gradual manner.

estimates suggest that the magnitude of the (short run) effects of macroeconomic shocks on income risk is significantly altered by the tariff level.

V.2. Endogeneity

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 variety 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: while variations in tariffs are observed across sectors, there is no variation in income risk. This type of purely cross-sectional endogeneity, however, is not a problem for our empirical analysis since we follow industries over time. More specifically, 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 eliminated. Along the time series dimension, we should note that the trade policy changes that we have studied were changes undertaken during major policy reform episodes (both in the late 1980s and

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

under NAFTA). These factors, in combination, suggest that concerns regarding bias resulting from the endogenous determination of trade policy should be minimal in our context.[27]

Estimation bias could, of course, also arise if systematic changes in non-tariff barriers reversed the effects of tariff reductions and 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 antidumping duties in these years and the antidumping 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 V).

V.3. 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 (7'). As the results presented in Table III 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, (7') was estimated using annually averaged observations (on income risk as well as the right hand side variables).

[27]To explain this further, consider an economy which starts with some initial level of tariffs and undertakes tariff reductions in some (any) number of industries. Consider further that the magnitude of the tariff reductions varies across sectors due to, say varying strengths of the import competing lobbies in these sectors. Given that our "within" estimate of the relationship between trade policy and income risk is formed by evaluating the change in income risk within an industry given its tariff change (and then averaging this across sectors), it should be easy to see that the varying political strength of sectors does not bias this estimate.

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

These results, presented in Table IV, 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 (7'). 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, (7') was estimated by dropping observations from the years 1995 and 1996. These results are also reported in Table V. As is evident, dropping observations from the years immediately following the exchange rate crisis in Mexico does not alter our results. Finally, our estimation results (not reported here but available upon request) with specifications in which we experimented with lagged independent variables (such as lagged tariff changes) with lags longer those reported here did not support the inclusion of such lags.

V.4. Welfare Analysis

Table II presents illustrative welfare calculations using the theoretical results derived in section IV and the empirical estimates obtained from the estimation of (7'). We conduct the following exercises. First, we evaluate the welfare effect of the short run change in income risk brought about by a five percent reduction in tariffs in high import penetration sectors. Second, we evaluate the welfare effects of a short run change in income risk following a real exchange rate appreciation of ten percent with the tariff level also at ten percent and see how these costs are altered if the prevalent tariff level were five percent instead. Finally, we consider the welfare effects of a changes 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 lower by five percent.

Consider first a tariff reform which involves a lowering of the tariff level by five percent.

As indicated in Table VI, this would raise σ_ϵ^2 in the short run (i.e., for one year following the reform) from a mean level of 0.08 to 0.013 (i.e., σ_ϵ goes up from 0.089 to 0.114). The corresponding welfare cost of this change is calculated to be 0.98 percent of permanent consumption if the co-efficient of risk aversion $\gamma = 1$ and is calculated to be 1.96 percent of lifetime consumption if the $\gamma = 2$ instead (always using an annual discount factor of $\beta = .96$).

Now consider the indirect effects of trade policy as measured by the interaction terms in (7'). As noted above, an exchange rate appreciation of ten percent raises σ_ϵ for a year from 0.089 to 0.105 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 rises to 0.118 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 for a year from 0.089 to 0.100, 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 rises to 0.114 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.

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 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 welfare costs of the effects of trade policy on income risk.

Our findings can be summarized as follows. First, 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 cost associated with the estimated increases in income risk are substantial. However, it is worth pointing out that our welfare analysis here focuses *exclusively* on the link between trade policy and individual income risk, and other possible channels through which trade policy may affect the economy are not studied here. More specifically, we would expect trade reform to have positive effects on the efficiency of resource allocation and economic growth, and such effects are important factors that ought to be taken into account when evaluating the total costs and benefits of trade reform. Additionally, our welfare calculations are based on a simple theoretical model whose limitations include its neglect of the effect of income risk on labor supply and capital accumulation.[29] Moreover, our calculations do not take into account that the welfare cost of an increase in income risk might be partially

[29] See, for example, Aiyagari (1994) for physical capital accumulation and Krebs (2003b) for human capital accumulation.

offset by a rise in transfer payments from the government or firms.[30] Finally, while our estimates of income shocks were obtained using observations on individuals over a limited time period, our welfare analysis assumes that shocks that are highly persistent through our sample period are equally persistent beyond this period. Thus, the welfare results presented in this paper have to be interpreted with caution keeping in mind our exclusive focus on the link between trade policy and income risk and the methodological limitations noted above.

[30] Being that such transfers are provided by entities within the economy, they should perhaps nevertheless be counted as costs, even if the risk to workers is fully offset by these payments.

References

Altonji,J. (1986) "Intertemporal Substitution in Labor Supply: Evidence From Micro Data," *Journal of Political Economy* 94: 176-215.

Aiyagari, R., 1994, "Uninsured Idiosyncratic Risk and Aggregate Saving," *Quarterly Journal of Economics*, 109: 659-84.

Altonji, J., and L. Segal, 1996, "Small-Sample Bias in GMM estimation of covariance structures," *Journal of Business and Economic Statistics* 14: 353-366.

Alesina, A., and Rodrik, D., 1994, "Distributive Politics and Economic Growth," *Quarterly Journal of Economics*, Vol. 109 (2) pp. 465-90.

Antras, P., 2003, "Firms, Contracts, and Trade Structure," *Quarterly Journal of Economics*, Vol 118 (4) pp. 1375-1418

Bhagwati, J., 2001, *Free Trade Today*, Princeton University Press, Princeton, NJ.

Borjas, G., Freeman, R., and Katz, L., 1992, "On the Labor Market Effects of Immigration and Trade," In Borjas and Freeman eds., *Immigration and the Workforce*, Chicago: University of Chicago Press.

Campbell, J., Lo, A., and MacKinlay, C., *The Econometrics of Financial Markets*, Princeton University Press, 1997.

Carroll, C. and Samwick, A., 1997, "The Nature of Precautionary Wealth," *Journal of Monetary Economics* 40: 41-72.

Cooley, T., 1995, *Frontiers of Business Cycle Research*, Princeton University Press, Princeton, NJ.

Constantinides, G., and Duffie, D., 1996, "Asset Pricing with Heterogeneous Consumers," *Journal of Political Economy* 104: 219-240.

Deaton, A.,1992, *Understanding Consumption*, Oxford University Press, Oxford.

Eaton, J., and G. Grossman, 1985, "Tariffs as Insurance: Optimal Commercial Policy When Domestic Markets Are Incomplete," *Canadian Journal of Economics*

Fernandez, R., and Rodrik, D., 1991, "Resistance to Reform: Status Quo Bias in the Presence of Individual-Specific Uncertainty," *American Economic Review*, Vol. 81 (5) pp. 1146-55

Feenstra, R., and Hanson, G., 2002, "Global production and Inequality: A survey of Trade and

Wages," in Choi and Harrigan eds., *Handbook of International Trade*, Basil Blackwell.

Gawande, K., and Krishna, P., 2003, "The Political Economy of Trade Policy: Empirical Approaches," in Choi and Harrigan eds., *Handbook of International Trade*, Basil Blackwell.

Gourinchas, P., and Parker, J., "Consumption over the Life-Cycle", *Econometrica* 70: 47-89.

Hanson, G., 2003, "What Has Happened To Wages in Mexico since NAFTA? Implications for Hemispheric Free Trade," *National Bureau of Economic Research*, Working Paper Number 9563.

Helpman, E., and A. Razin, 1980, "Efficient Production Under Uncertainty," *American Economic Review*

Hubbard, G., Skinner, J., and S. Zeldes, 1994, "The Importance of Precautionary Motives in Explaining Individual and Aggregate Savings," *Carnegie-Rochster Conference Series on Public Policy* 40: 59-126.

Huggett, M., 1993, "The Risk-Free Rate in Heterogenous-Agent Incomplete-Market Economies," *Journal of Economic Dynamics and Control* 17: 953-969.

Krebs, T., 2003a, "Growth and Welfare effects of Business Cycles in Economies with Idiosyncratic Human Capital Risk," *Review of Economic Dynamics* 6: 846-868.

Krebs, T., 2003b, "Idiosyncratic Risk, Aggregate Saving, and Economic Growth," *Quarterly Journal of Economics* 11: 709-744.

Krebs, T., 2004, "Consumption-Based Asset Pricing Models with Incomplete Markets," *Journal of Mathematical Economics* 40: 191-206.

Krusell, P. and Smith, A., 1998, "Income and Wealth Heterogeneity in the Macroeconomy," *Journal of Political Economy*, 106: 867-896.

Lawrence, R., and Slaughter, M., 1993, "International Trade and American Wages in the 1980s: Giant Sucking Sound or Small Hiccup?" in Martin Neil Baily and Clifford Winston eds., *Brookings Papers on Economic Activity: Microeconomics 2*, 1993, pp. 161-211.

Levine, D. and Zame, W., 2002, "Does Market Incompleteness Matter?", *Econometrica* 71: 1695-1725.

Melitz, M., 2003, "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 71: 1695-1726

Meghir, C. and L. Pistaferri, 2004, "Income Variance Dynamics and Heterogeneity," *Econometrica* 72: 1-32.

Rodrik, D., 1997, *Has Globalization Gone Too Far?*, Institute for International Economics, Washington, DC.

Rodrik, D., 1998, "Why Do More Open Economies Have Bigger Governments?" *Journal of Political Economy*, 106(5).

Storesletten, K., Telmer, C., and A. Yaron, 2002, "Cyclical Dynamics in Idiosyncratic Labor Market Risk," Working Paper, Carnegie Mellon University, forthcoming *Journal of Political Economy.*

Trefler, D., 1993, "Trade Liberalization and the Theory of Endogenous Protection: An Econometric Study of U.S. Import Policy," Journal of Political Economy, Vol. 101 (1) pp. 138-60.

Appendix

In this appendix, we construct the equilibrium and derive the welfare expressions. The Euler equations associated with the consumption/saving problem of household i read

$$c_{ijt}^{-\gamma} \geq \beta(1+r_{t+1})E[c_{ij,t+1}^{-\gamma}|\mathcal{F}_{ijt}] \ , \tag{13}$$

where $\mathcal{F}_{ijt}$ is the information that is available to household i in period t. The Euler equation (13) says that the marginal utility cost of saving one more unit is greater or equal to the expected marginal utility gain of doing so. As long as the borrowing constraint is not binding, $a_{ijt} > 0$, equation (13) must hold with equality. Notice that any plan solving the Euler equation (13) and the budget constraint (9) also satisfies a corresponding transversality condition if the following condition is satisfied (Krebs 2004):

$$\beta E\left[(1+\mu_{jt})^{1-\gamma}(1+\theta_{ijt})^{1-\gamma}\right] < 1 \ . \tag{14}$$

Thus, we can focus on Euler equations when discussing optimal consumption/saving plans. Notice that this condition is automatically satisfied if $\gamma = 1$ (log-utility).

If we rule out international borrowing and lending, then the domestic interest rate is determined by the saving decisions of domestic households only.[31] In this case, domestic asset market clearing reads:

$$\sum_{i,j} a_{ijt} = 0 \ . \tag{15}$$

Suppose the interest rate is given by

$$r_{t+1} = \frac{1}{\beta\,(1+\mu_{\hat{j},t+1})^{-\gamma}\,E\left[\left(1+\theta_{i\hat{j},t+1}\right)^{-\gamma}|\mathcal{F}_{ijt}\right]} - 1 \ , \tag{16}$$

where $\hat{j}$ is the sector for which the right-hand side of (16) is maximal. Notice that the right-hand side of (16) 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 this interest rate, the Euler equation holds with equality if all households in sector $\hat{j}$ choose $a_{i\hat{j}t} = 0$ and $c_{i\hat{j}t} = \tilde{y}_{i\hat{j}t}$. Moreover, for all households in any sector $j \neq \hat{j}$, the Euler equation also holds if $a_{ijt} = 0$ and $c_{ijt} = \tilde{y}_{ijt}$, although for these households the Euler equation will in general hold as an inequality (the borrowing constraint binds). Since a corresponding transversality condition holds for the individual plan $a_{ijt} = 0$ and $c_{ijt} = \tilde{y}_{ijt}$, all

[31] Clearly, an alternative interpretation is that the model describes a small open economy with an exogenous interest rate that is low enough so that households do not want to save. In other words, any interest rate process for which the interest rate is lower than the interest rate defined in (16) supports the allocation as an equilibrium outcome.

households optimize. Since $a_{ijt} = 0$ satisfies market clearing, we have shown that $c_{ijt} = \tilde{y}_{ijt}$ is an equilibrium.

Let us now discuss the link between the specification of the income process (1)-(3) in the empirical section II and the income process used in the theoretical section IV. Recall that we assume that $log(1+\theta)$ 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 (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 (8), we find

$$log\,\tilde{y}_{ij,t+1} = log\,\tilde{y}_{ijt} + log(1+\mu_{j,t+1}) \; + \; log(1+\theta_{ij,t+1}) \; . \tag{17}$$

Thus, income follows a logarithmic random walk with drift $log(1+\mu_{j,t+1})$ and heteroscedastic error term $log(1+\theta_{i,t+1})$. Comparison of (17) with the econometric specification (3) suggests that we relate $\log(1+\theta_{ij,t+1})$ in (17) with the innovation term of the permanent, unpredictable component of income changes in (1):

$$log(1+\theta_{ij,t+1}) = \epsilon_{ij,t+1} - \sigma^2_{j,t+1}/2 \; . \tag{18}$$

In (18) we introduce the term $-\sigma^2_{j,t+1}/2$ to ensure that both random variables have the same mean. Taking the variance in (18) we find

$$\sigma^2_{j,t+1} = \sigma^2_{\epsilon j,t+1} \; . \tag{19}$$

Thus, our empirical measure of income risk, σ^2_ϵ, coincides with our theoretical measure of income risk, σ^2.

We now turn to the welfare analysis. Suppose that tariff rates and income parameters are constant over time: $\tau_{jt} = \tau_j$, $\mu_{jt} = \mu_j$, and $\sigma^2_{jt} = \sigma^2_j$. If $c_{ijt} = \tilde{y}_{ijt}$ and there are no aggregate fluctuations, then expected lifetime utility (10) is

$$\begin{aligned} U &= \frac{c_0^{1-\gamma}}{(1-\gamma)\,(1-\beta(1+\mu)^{1-\gamma}E[(1+\theta)^{1-\gamma}])} \quad if\;\gamma \neq 1 \\ U &= \frac{1}{1-\beta}logc_0 \; + \; \frac{\beta}{(1-\beta)^2}\left(log(1+\mu) + E[log(1+\theta)]\right) \quad otherwise \; , \end{aligned} \tag{20}$$

where the expectation is taken over idiosyncratic shocks (over the random variable θ). Notice that we dropped the indexes i and t because with the exception of initial consumption c_0, all terms in the expression (18) are household- and time-independent. To ease the exposition, we have also dropped the index j. Using the assumption that $\theta \sim N(-.5\sigma^2, \sigma^2)$, integration over income shocks yields

$$U \;\; = \;\; \frac{c_0^{1-\gamma}}{(1-\gamma)\;(1-\beta(1+\mu)^{1-\gamma}exp\left(.5\left((1-\gamma)^2-(1-\gamma)\right)\sigma^2\right))} \quad if\;\gamma \neq 1 \tag{21}$$

$$U = \frac{1}{1-\beta} log c_0 + \frac{\beta}{(1-\beta)^2}\left(log(1+\mu) - \sigma^2/2\right) \quad otherwise \ .$$

Equation (21) shows how welfare depends on income risk, σ^2, which in turn depends on tariff rates, τ. Thus, the welfare expression (21) 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 (7). In order to get numbers for these welfare changes with economically meaningful units, we calculate the change in initial consumption, c_0, that is necessary to compensate the worker for the change in risk.[32] More precisely, for any c_0, σ^2, and Δ_σ, we are searching for the percentage change in initial consumption, Δ_c solving

$$U(c_0, \sigma^2) = U\left((1+\Delta_c)c_0, (1+\Delta_\sigma)\sigma^2\right) \tag{22}$$

Notice that because of our random walk assumption, any increase in initial consumption, c_0, amounts to an increase in consumption for all future dates and events (lifetime consumption). Using (21) and (22), we find (11). Notice that expression (11) is independent of c_0, that is, the welfare change expressed in percentage changes of consumption levels 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 (11), and expected lifetime utility with the increase is:

$$\begin{aligned} U' &= \sum_{t=0}^{n} \beta^t \frac{E[(c_t')^{1-\gamma}]}{1-\gamma} + \sum_{t=n+1}^{\infty} \beta^t \frac{E[(c_t')^{1-\gamma}]}{1-\gamma} \qquad (23) \\ E[(c_t')^{1-\gamma}] &= \frac{c_0^{1-\gamma}}{1-\gamma}(1+\mu)^{(1-\gamma)t}\left(E[(1+\theta')^{1-\gamma}]\right)^t \quad t=0,1,\ldots,n \\ E[(c_t')^{1-\gamma}] &= \frac{c_0^{1-\gamma}}{1-\gamma}(1+\mu)^{(1-\gamma)t}\left(E[(1+\theta')^{1-\gamma}]\right)^n \left(E[(1+\theta)^{1-\gamma}]\right)^{(t-n)} \quad t=n+1,n+2,\ldots \end{aligned}$$

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 (23), we find equation (12) in section IV.

[32]Notice that for the case considered here, this compensating variation is equal to the equivalent variation.

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: Trade Policy and Income Risk - Panel Estimates *

Variables	σ^2_ϵ vs	σ^2_ϵ 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 III: 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 IV: Trade Policy and Income Risk - Annual Estimates of $\sigma_\epsilon^{2\ddagger}$

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 V: 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 antidumping 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 VI: Welfare Effects[¶]

	Change in σ_ϵ^2 ($\bar{\sigma_\epsilon^2} = 0.008$)	Welfare Change γ =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.