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INTERNATIONAL TRADE AND LABOR INCOME RISK IN THE UNITED STATES

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ABSTRACT

This paper studies empirically the links between international trade and labor income risk faced by workers in the United States. We use longitudinal data on workers to estimate time-varying individual income risk at the industry level. We then combine our estimates of persistent labor income risk with measures of exposure to international trade to analyze the relationship between trade and labor income risk. Importantly, by contrasting estimates from various sub-samples of workers, such as those who switched to a different industry (or sector) with those who remained in the same industry throughout the sample, we study the relative importance of the different channels through which international trade affects individual income risk. Finally, we use these estimates to conduct a welfare analysis evaluating the benefits or costs of trade through the income risk channel. We find import penetration to have a statistically significant association with labor income risk in the United States, with economically significant welfare effects.

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I. Introduction

A vast empirical literature has examined the effects that globalization may have on workers in the domestic economy with particular focus on the important question of how trade might affect, on average, the wages of workers in different human capital or occupational categories.[1] This impressive literature has uncovered many interesting findings regarding the "mean" effects of globalization on labor markets. However, for the most part, this literature has not addressed a broadly expressed public concern regarding another possible channel through which globalization might affect labor markets: Openness to international trade may expose workers to riskier economic environments causing greater volatility (variance) in their incomes.[2]

How might trade openness affect labor income volatility? The theoretical literature has suggested various channels through which exposure to international trade may increase labor income risk. One possibility derives from the fact that openness may expose import-competing sectors to a variable international economic environment. Here, changing international patterns of comparative advantage will induce reallocations of capital and labor across firms within and between sectors. To the extent that similar workers experience different outcomes during this reallocation process, openness will raise individual labor income risk. Another link between trade and income volatility has been argued by Rodrik (1997). Here, increased foreign competition, which increases the elasticity of the demand for goods, also raises the elasticity of the derived labor demand. This, in turn implies that shocks to labor demand may result in larger variations in wages and employment, and hence increase

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[1] Well-known papers in this literature include Lawrence and Slaughter (1993), Leamer (1996) and Feenstra and Hanson (1999) and Goldberg and Pavcnik (2005). Davidson and Matusz (2004), Feenstra and Hanson (2002), Goldberg and Pavcnik (2007) and Harrison (2007) provide excellent survey treatments. Recent analyses have emphasized the role of firm heterogeneity (as in Melitz (2003)), worker heterogeneity and labor market frictions in studying the labor market effects of globalization. Amiti and Davis (2008), Davis and Harrigan (2007), Egger and Kreickemeier (2006), explore labor market reallocation with trade openness in heterogeneous firm models, Davidson, Matusz and Shevchenko (2008) and Mitra and Ranjan (2009) study the employment impact of trade and offshoring (respectively) in the presence of labor market (search) frictions, Ohnsorge and Trefler (2007) explore the consequences of worker heterogeneity for trade and the domestic distribution of income and Helpman, Itskhoki and Redding (2008) combine firm heterogeneity and worker heterogeneity with labor market frictions to study the effects of trade openness on unemployment and inequality.

[2] Exceptions include Krebs, Krishna, and Maloney (2008), which studies Mexico, diGiovanni and Levchenko (2007), which provides interesting cross-country evidence regarding the links between trade and sectoral output volatility, and McLaren and Newman (2002), which studies how globalization may weaken domestic institutions for risk-sharing.

the volatility in the labor market. On the other hand, it has also been suggested that the world economy, by aggregating shocks across countries, may be less volatile than the economy of any single country, opening 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. The question is an empirical one.

This paper undertakes a detailed empirical analysis of the association between trade and labor income risk in the United States. We use longitudinal data on workers to estimate idiosyncratic individual income risk and to study the role of trade in explaining the variation in this risk across workers employed in different industries.[3] To estimate labor income risk (defined as the variance of unpredictable changes in earnings), we employ specifications of the labor income process that account for shocks to labor income that workers receive and that distinguish between transitory and persistent shocks to income. This distinction between transitory and persistent shocks is important. Workers can effectively "self-insure" against transitory shocks through borrowing or own savings, and the welfare effects of such shocks are quite small (Heaton and Lucas (1996), Levine and Zame (2002)). 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. Thus, from a welfare point of view, it is the persistent income shocks that matter the most and it is on these shocks that we focus our attention.

In our analysis, we combine industry-level, time-varying estimates of the persistent component of labor income risk with measures of industry exposure to international trade to estimate the relationship between labor income risk and trade. Importantly, we then repeat this analysis for different sub-samples of workers, such as those who switched to a different industry or sector, or those who remained in the same industry throughout the sample. This allows us to identify these separate components of risk faced by individuals and to evaluate the relative importance of the different channels

[3]We use the Survey of Income and Program Participation (SIPP) in our analysis. SIPP contains longitudinal panels on individuals (and households), with each panel ranging roughly three years in duration. We use data from 3 SIPP panels—the 1993–1995, 1996–1998 and 2001–2003 panels—in our study.

through which international trade can affect individual income risk. One strength of our methodological approach is that we are able to use these estimates to conduct a welfare analysis to evaluate benefits or costs of trade through the income risk channel.[4]

Our empirical results for the United States can be summarized as follows. First, we find those workers who switched industries (moving to a different manufacturing industry or to the non-manufacturing sector) experience higher income risk compared to those who stayed in the same manufacturing industry throughout the sample. Estimated risk for those who switched to the non-manufacturing sector is higher than those who switched within manufacturing. Second, we find that within-industry changes in income risk are strongly related to changes in import penetration over the corresponding time-periods. This relationship holds for the full sample of workers as well as various sub-samples we consider and is robust to controlling for other time varying industry specific factors (such as exports, skill-biased technological change, offshoring, unionization, productivity) that are potentially correlated with both income risk and import penetration. We also consider the implications of sample-selection in estimating this relationship but find no evidence of any selection in the data that could bias the estimates in our favor. Quantitatively, estimates from our preferred specification suggest that an increase in import penetration by ten percent is associated with an increase in the standard deviation of persistent income shocks of about 20 to 25 percent for the full sample of workers.[5] Our welfare calculations suggest that these effects are economically significant: Under standard parameter values for the inter-temporal discount rate ($\beta = 0.98$) and the coefficient of relative risk aversion ($\gamma = 2$), the increase in persistent income risk is (certainty) equivalent to

[4] In our study, we follow the methodological approach taken by Krebs, Krishna, and Maloney (2008) – with some important differences. First, SIPP panels have a much longer longitudinal dimension than the Mexican data used by Krebs, Krishna, and Maloney (2008). This allows for methodological improvements in the estimation of risk with more precise estimates of the magnitude of persistent shocks to income. Second, our data set is large enough to allow us to estimate risk separately for various sub-samples of workers such as those who switched to a different industry or those who remained in the same industry throughout. This allows us to evaluate the relative importance of the different channels through which international trade can affect individual income risk. Furthermore, as our study is based in the United States, data on a variety of industry characteristics ia available which we use as controls in our econometric exercises, as discussed later in the paper.

[5] The same increase in import penetration is associated with an increase in income risk of about 30 percent and 20 percent for workers who remained in the same manufacturing industry throughout and those who switched, respectively.

a reduction by roughly 4 % to 11 % percent of lifetime consumption.

We should emphasize that our analysis focuses *exclusively* on the link between trade and individual income risk. Hence, our results should be taken together with the findings of a large literature on international trade exploring the many ways in which trade may affect the economy positively, through improved resource allocation, access to greater varieties of intermediate and final goods, greater exploitation of external economies and by possibly raising growth rates, *inter alia*. Specifically, the results presented here should not be interpreted as suggesting that exposure to trade results in welfare reduction, but instead as evidence that the costs of increased labor income risk ought to be taken into account when evaluating the total costs and benefits of trade and trade policy reform.

II. Labor Income Risk

The first stage of our analysis concerns the estimation of individual income risk and its separation into transitory and persistent components. As we have discussed earlier, it is this focus on income risk that separates our analysis from much of the earlier literature that has examined instead the "mean" effect of trade on wages of workers in different skill and occupational categories. Figures I and II present heuristic illustrations to clarify this difference further.

Figure I depicts income paths for an ex-ante identical group of workers whose incomes in time period t are identical and equal to y_t. In time period t+1, we see that the average income for this group of workers drops to $\bar{y}_{t+1}$. However, around this mean drop in incomes there is a variance in individual outcomes. To the extent that individual outcomes are unpredictable ex-ante, the process is risky and risk-averse workers would find this to be costly. It is this variance around y_{t+1} that we are interested in – while the prior literature has largely examined the mean income gap ($y_t - \bar{y}_{t+1}$).[6]

[6] Note that under the expected utility hypothesis, the variance in outcomes would, in of itself, be seen as costly even if the mean income $\bar{y}_{t+1}$ was higher than the income in the earlier period (and even if all workers saw an increase in incomes in t+1). In our welfare calculations, we do not explore attitudes

Figure II illustrates the difference between transitory and persistent income shocks for a group consisting of two ex-ante identical individuals whose incomes in time period t are identical and equal to y_t. At t+1, they experience shocks to income (some part transitory and some part persistent) that separate their incomes as indicated. By t+2, the transitory components of the income changes they experienced at t+1 expire and incomes for both workers move closer to their initial levels and stay at these levels for the rest of time. In this case, the magnitude of the variance of the persistent shock experienced at t+1 is measured by the spread in incomes at t+2 (and beyond). The spread in incomes at t+1 measures the sum of the variance of the transitory shock as well as the permanent shock experienced at t+1.

The separation between transitory and persistent shocks is essential for multiple reasons. First, consumption smoothing through borrowing or own savings works well for transitory income shocks (Aiyagari (1994), Heaton and Lucas (1996), Levine and Zame (2002)), but not when income shocks are highly persistent or permanent (Constantinides and Duffie (1996) and Krebs (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. For these reasons, we will focus on persistent shocks and their relation to trade exposure. Our estimation strategy itself follows earlier approaches in the literature estimating US labor income risk (Carroll and Samwick (1997), Gourinchas and Parker (2002), Meghir and Pistaferri (2004)) with some important differences to account for the structure of our data that we discuss in detail below.

II.1. Data

In this paper, we use longitudinal data on individuals from the 1993–1995, 1996–1999 and 2001–2003 panels of the Survey of Income and Program Participation

towards risk that deviate from the expected utility hypothesis. Alternative preference specifications that treat variance in gains asymmetrically from variance in losses are outside of the scope of the present analysis.

(SIPP).[7] Each panel of the SIPP is designed to be a nationally representative sample of the US population and surveys thousands of workers. The interviews are conducted at four-month intervals over a period of three years for the 1993 panel, four years for the 1996 panel and three years again for the 2001 panel.[8] At each interview, data on earnings and labor force activity are collected for each of the preceding four months.

SIPP has several advantages over other commonly used individual-level datasets in that it includes monthly information on earnings and employment over a long panel period for a large sample. Although the Current Population Survey (CPS) provides a larger sample, individuals are only sampled for 8 months over a two-year period in comparison to 33 months in the SIPP. While the Panel Study of Income Dynamics (PSID) provides a much longer longitudinal panel, it has a significantly smaller sample size compared to the SIPP and therefore does not support the estimation of risk at the industry level.

In our analysis, we restrict the SIPP sample to respondents of age 16 to 65 who were not enrolled in school during a given month. Following previous literature, we exclude all observations for individuals whose earnings in any month were less than 5% or higher than 195% of the individual's average monthly earnings.[9] Table I presents a summary description of the workers surveyed in each panel. The summary statistics calculated for the first month of each panel are reported separately for the whole sample and for the manufacturing sector only. Workers earnings represent amounts actually received in wages and salary and/or from self-employment, before deductions for income and payroll taxes, union dues, Medicare premiums, etc.

[7] For the estimation of individual income risk, longitudinal data capturing individual income changes is desirable. It is generally not sufficient to have information on changes in the aggregate distribution of income to make inferences about the extent of income risk faced by individuals (although this is still possible under specific assumptions). For instance, the aggregate distribution of income may stay the same across different time periods but there may be stochastic (risky) transitions taking place underneath - with some individuals at the top of the distribution exchanging places with others at the bottom end of the distribution. To capture the risk in incomes faced by these individuals, longitudinal data tracking these individual transitions is therefore useful to have.

[8] We limit our main analysis to data from the first three years of the 1996 panel to ensure comparability of our risk estimates with the other two panels. As we discuss later, we do exploit the additional year of data in the 1996 panel in our analysis of robustness.

[9] This results in the omission of approximately 13% of the respondents of each panel from our sample.

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 labor income of individual i employed in industry j in time period (month) t, $log\ y_{ijt}$, is given by:

$$log\ y_{ijt} = \alpha_{jt} + \beta_t \cdot x_{ijt} + u_{ijt}\ . \qquad (1)$$

In (1) α_{jt} and β_t denote time-varying coefficients, x_{ijt} is a vector of observable characteristics (such as age, age-squared, education, marital status, occupation, race, gender and industry), and u_{ijt} is the stochastic component of earnings. Changes in 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 changes in u_{ijt}. In this sense, changes in u_{ijt} over time measure the unpredictable part of changes in individual income.[10]

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

$$u_{ijt} = \omega_{ijt} + \eta_{ijt}\ . \qquad (2)$$

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

$$\omega_{ij,t+1} = \omega_{ijt} + \varepsilon_{ij,t+1}\ , \qquad (3)$$

where the innovation terms, $\{\varepsilon_{ijt}\}$, are independently distributed over time and identically distributed across individuals, $\varepsilon_{ijt} \sim N(0, \sigma^2_{\varepsilon js})$, where s denotes the SIPP panel (i.e., one of the 1993–1995, 1996–1998 or 2001–2003 panels). In this basic specification, transitory shocks have no persistence, that is, the random variables

[10] Since income risk is calculated as the variance of unpredictable changes in earnings, it is understood that any time-invariant individual specific component of earnings will be purged out from our risk estimates. As such, inclusion of individual-fixed effects to specification (1) should not alter our risk estimates. Indeed, we find this to be the case when we pool all months in a panel and estimate the Mincer regression with month- and individual-fixed effects.

$\{\eta_{ijt}\}$ are independently distributed over time and identically distributed across individuals, $\eta_{ijt} \sim N(0, \sigma^2_{\eta\ js})$. Note that the parameters describing the magnitude of both transitory and persistent shocks are assumed to depend on the sector *j* and the SIPP panel *s*, but do not depend on *t*. That is to say, they are assumed to be constant within a SIPP panel, but allowed to vary across panels. Estimation of $\sigma^2_{\varepsilon js}$ and $\sigma^2_{\eta\ js}$ will therefore give us industry specific, time varying estimates of transitory and permanent income risk faced by individuals.

Notice that in (1), 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. However, in addition to specification (1), we also conduct our analysis using alternate specifications. As we have just discussed, (1) takes out any changes to income that may have occurred due to changes in returns to observable characteristics. Another possibility is to treat these changes as unpredictable by requiring the coefficients β to be time-invariant within a panel. In this case, estimated income risk will include any changes in the returns to observable characteristics that take place in reality. Which set of estimates to use will depend on whether we think of changes in the coefficients on observable worker characteristics to be predictable or not. While this an interesting conceptual issue, in practice, estimates of the parameters representing income risk do not seem to depend very much on whether the changes in returns to observable characteristics are accounted for by allowing β to be time varying, or not, in estimating (1).

Notice also that the inclusion of industry dummies in (1) filters out mean income changes in an industry but also filters out from our measure of individual risk any volatility in the changes of the mean industry earnings. Our risk estimates therefore measure idiosyncratic income risk (effectively individual variation around the industry mean, conditional on the other covariates in (1)) experienced by individuals.[11]

[11] While it is possible that trade may additionally affect workers (positively or negatively) by affecting the volatility of mean income growth in industries, in our analysis of the association between the

II.2.1 Filtering out Shocks of Short Duration

Our specification of the labor income process (Equations (1)–(3)) describes shocks to income to be either purely transitory or purely persistent and is in accordance with other empirical work on US labor income risk.[12] However, this specification does not capture shocks that have duration greater than one period (i.e., are not purely transitory) but that are also not permanent (i.e., last for a finite amount of time). Estimation of permanent income risk in this case requires us to filter out such shocks of longer duration (See Meghir and Pistaferri (2004)). To achieve this, we admit into the specification some moving average terms:

$$u_{ijt} = \omega_{ijt} + \sum_{k=0}^{K} \eta_{ijt-k} \,, \qquad (2')$$

with K indicating the number of moving average terms. In addition to the benchmark specification where transitory shocks have no persistence (K=0), we consider two alternative specifications of the labor income process that allow for transitory shocks that last up to six months (K=6) and, separately, up to a year (K=12). We denote the corresponding parameters estimating permanent income risk by $\sigma^2_{\varepsilon,k=0}$, $\sigma^2_{\varepsilon,k=6}$ and $\sigma^2_{\varepsilon,k=12}$, respectively. Note that we expect the estimates of permanent income risk to be smaller in magnitude when shocks of shorter duration have been filtered out; that is, we expect $\sigma^2_{\varepsilon,k=0} > \sigma^2_{\varepsilon,k=6} > \sigma^2_{\varepsilon,k=12}$ (See Meghir and Pistaferri (2004)).

While we report our results obtained for each value of K, we place greater emphasis on results from specification (2') with K=12. $\sigma^2_{\varepsilon,k=12}$ is our preferred risk estimate because we are interested in permanent income risk and this specification of the labor income process allows us to filter out transitory shocks of greater duration than the other two estimates do.

variance of changes in mean industry earnings and import penetration, we do not find evidence of any relationship between these two variables.

[12] For example, Carroll and Samwick (1997) and Gourinchas and Parker (2002) use the same specification that we do. 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).

Our intention is to estimate parameters measuring income risk and see how changes in these parameters over time (i.e., across panels) may be related to international trade. In order to do this, we first estimate the income risk parameters at the industry level separately for each panel (for each of the cases with K=0, K=6 and K=12). Estimation of the income process parameters is discussed next.[13]

II.3. Estimation

Consider the change in the residual of income of individual i between period t and $t+n$ (we drop the subscript s for notational convenience; it is understood that the estimation exercises are conducted separately for each panel):

$$\Delta_n u_{ijt} = u_{ij,t+n} - u_{ijt} = \varepsilon_{ij,t+1} + \ldots + \varepsilon_{ij,t+n} + \eta_{ij,t+n} - \eta_{ijt} \ . \qquad (4)$$

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

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

As noted earlier, the parameters $\sigma^2_{\varepsilon j}$ and $\sigma^2_{\eta j}$ are assumed to be constant within the period covered by a single SIPP panel (i.e., within each of the 1993–1995, 1996–1998 and 2001–2003 panels).

Given this constancy, (5) can be written as:

$$var[\Delta_n u_{ijt}] = 2\sigma^2_{\eta j} + n\,\sigma^2_{\varepsilon j} \qquad (6)$$

Thus, the variance of observed n-period income changes is a linear function of n, where the slope coefficient is equal to $\sigma^2_{\varepsilon j}$. This 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. Following Carroll and Samwick (1997), we estimate the parameters in (6) by regressing individual measures of $var[\Delta_n u_{ijt}]$, the square of the individual deviation from mean income difference

[13] We discuss below the estimation of the parameters of (2) and (3). The estimation of income risk parameters when K>0 as in (2') is entirely analogous.

over the n periods, on n. (6) is estimated separately for each industry and panel. As is well recognized in the literature, the transitory term in the specification of the income process will absorb the measurement error in individual income. Given this and the fact that the welfare effects of transitory shocks to income are much smaller (as we have discussed), we will focus on persistent shocks and their relation to trade.

II.4. Data and Implementation of Estimation Methodology with the SIPP data

Since trade data is only available for the manufacturing sector, we restrict our sample to those workers employed in the manufacturing sector during the first month of each panel. We assign individuals to those industries in which they were initially observed, and maintain this industry assignment throughout.

The risk estimates from this sample account for both the shocks to workers who experience income changes due to changes in their wage rates or the number of hours in a given job and the shocks to workers who change jobs within or between industries, allowing for intermediate periods of unemployment.[14] Specifically, the sample analogs to $var[\Delta_n u_{ijt}]$ are formed by considering income differences for workers between time periods t and t+n regardless of their employment status in any intermediate period. While losing a worker from the data set due to unemployment in intermediate periods between t and t+n will bias the estimate of transitory income shocks, it will not bias our estimate of the magnitude of permanent income risk as long as the individual does not remain unemployed for the remainder of the duration of the panel. In the event that individuals are simply lost from the data set because of unemployment, we would indeed underestimate the magnitude of shocks to income. However, this is not a severe problem here since less that 2% of the individuals in our sample are unemployed as of the last month they were surveyed and the average duration of unemployment for our sample is less than 2 months in all three panels.[15]

[14] One issue that arises from assigning industries the way we described in constructing our baseline sample above is that individuals may experience shocks to income due to some changes in trade exposure in the subsequent industry of employment, but this income change will be included in estimation of income risk in the initial industry of employment instead. However, the vast majority of displacements in our sample are within the same industry or to the non-manufacturing sector. This is consistent with the well-known findings of Davis, Haltiwanger and Schuh (1996) that most job creation and destruction in the United States takes place within industries.

[15] We also find that the change in attrition rates between panels is not correlated with change in import

II.5. Results

The preceding section provided a detailed description of the general econometric methodology that we use to estimate income risk given longitudinal data on individual incomes. Using this methodology, we estimate the risk parameters, σ_{ε}^{2} and σ_{η}^{2}, separately for the three SIPP panels and 18 manufacturing industries in the United States.[16] In this section, we report these risk estimates and note some additional issues that arise in applying this methodology to our data.

Table II describes the estimates obtained using our benchmark specification, where transitory shocks are purely transitory and have no persistence at all ($\sigma_{\varepsilon,k=0}^{2}$) as well as when we allow for transitory shocks of longer duration ($\sigma_{\varepsilon,k=6}^{2}$ and $\sigma_{\varepsilon,k=12}^{2}$). As indicated earlier, $\sigma_{\varepsilon,k=12}^{2}$, obtained after we filter out shocks lasting up to a year, is our preferred estimate. [17,18]

As indicated in Table II, the mean value of the monthly variance of the persistent shock, $\sigma_{\varepsilon,k=0}^{2}$, for the 1993 panel is estimated to be 0.0033 (or 0.0396 annualized). For the 1996 and 2001 panels, the corresponding estimates for monthly $\sigma_{\varepsilon,k=0}^{2}$ are 0.0043 (or 0.0516 annualized) and 0.0052 (or 0.0624 annualized), respectively. The corresponding annualized standard deviations of permanent income growth (calculated as $(12*\sigma_{\varepsilon,k=0}^{2})^{1/2}$) are 0.20, 0.23 and 0.25 for the 1993, 1996 and 2001 panels, respectively. Clearly, income risk is rising over time: On average, $\sigma_{\varepsilon,k=0}^{2}$ rose by 30 percent between the 1993 and 1996 panels and by a further 20 percent between

penetration in our sample. This suggests that attrition due to non-response or to unemployment is not likely to bias our main results on the relationship between income risk and import penetration.

[16] Tobacco Products (SIC 21) and Petroleum Refining (SIC 29) are omitted from our analysis due to an insufficient number of observations on individuals within these industries.

[17] We also use the additional year of data for the 1996 panel to explore the implications of filtering out shocks of even longer duration (18 months and 24 months) from our estimates of income risk. We find that the estimates are relatively stable after K=12 and do not decrease as much when we filter out shocks of longer durations, suggesting that shocks that last up to a year are mostly permanent.

[18] As described in Section II.2, an alternative to specification (1) is to estimate income risk by treating the changes in returns to observable worker characteristics as unpredictable. We explore this alternative by pooling all months, and estimating the Mincer regression for each panel with month-fixed effects. We also estimate specification (1) by including individual fixed effects. The risk estimates from these two time invariant Mincer specifications differ very little from those reported in Table II.

the 1996 and 2001 panels.

Table II also reports the summary statistics for the estimates of $\sigma^2_{\varepsilon,k=6}$ and $\sigma^2_{\varepsilon,k=12}$. As expected, allowing for shocks of greater duration, but which are not permanent, lowers our estimates of risk: The mean estimate of the monthly value of $\sigma^2_{\varepsilon,k=12}$ is 0.0014, 0.0025 and 0.0031 for the 1993, the 1996 and the 2001 panels (with corresponding annualized values of 0.0168, 0.03 and 0.0372), respectively. The annualized standard deviations of the reported estimates of $\sigma^2_{\varepsilon,k=12}$ are 0.13, 0.17 and 0.19 for the 1993, 1996 and 2001 panels, respectively.

Since our estimates for $\sigma^2_{\varepsilon,k=6}$ are simply intermediate in magnitude to the estimates of $\sigma^2_{\varepsilon,k=0}$ and $\sigma^2_{\varepsilon,k=12}$, we simply focus on this latter sets of estimates throughout the rest of the paper. Greater detail on $\sigma^2_{\varepsilon,k=0}$ and $\sigma^2_{\varepsilon,k=12}$ is provided in Table III, which lists the industry level estimates of these parameters for each of the three SIPP panels.

It is informative to compare our estimates of the permanent component of income risk, σ^2_{ε}, with the estimates obtained by the extensive empirical literature on US labor market risk using annual income data drawn from the PSID. Most of these studies find an average value of around 0.0225 for the annual variance σ^2_{ε} (Carroll and Samwick (1997), Gourinchas and Parker (2002), Hubbard, Skinner and Zeldes (1994), and Storesletten, Telmer and Yaron (2004)), with a value of $\sigma^2_{\varepsilon} = .0324$ being the upper bound (Meghir and Pistaferri, 2004). Thus, the average values of our estimates of permanent income risk, especially those that allow for transitory shocks of longer duration, are in line with the estimates that have been obtained by the previous literature on US labor market risk. Note that our results are obtained using SIPP, a three-year panel for the United States, instead of the PSID data with a panel dimension of many years used in previous literature. The similarity of the estimates from the two datasets suggests that most income shocks we label "permanent" in this paper indeed persist for a very long time.

II.6. Income Risk in Sub-Samples

Our dataset is sufficiently large enough to identify separate components of risk faced by different sub-samples of workers, allowing us to evaluate the relative importance of different channels through which international trade can affect individual income risk. In this section, we provide a description of our income risk estimates for these different sub-samples with particular emphasis on the type of risk we account for.

Our first sub-sample is constructed by including only the individuals who were employed in the same manufacturing industry each month they were employed (and surveyed). [19] This sample (denoted STAY-IND) includes workers who remained in the same job as well as those who switched jobs within the same industry (thereby possibly losing returns to firm or occupation specific human capital). Displaced workers who move to a different manufacturing or non-manufacturing industry are excluded from this sample and are instead grouped together in a different sample (SWITCH-ALL).

We then analyze further the importance of switching industries on income risk using two additional sub-samples. First, we construct a sample that includes individuals who stayed in the manufacturing sector throughout, but may have worked in a different industry within manufacturing than their original industry at some point (STAY-MANUF). Second, we consider those individuals who switched to the non-manufacturing sector for at least one period in the panel (SWITCH-NON-MANUF). Comparing income risk experienced by workers in these four different sub-samples (STAY-IND, STAY-MANUF, SWITCH-ALL and SWITCH-NON-MANUF) will allow us to study the costs of switching industries both within and outside manufacturing.

Table IV provides a summary description of our estimates of income risk for the sub-samples described above for each panel.[20] As Table IV indicates, $\sigma^2_{\varepsilon,k=0}$ continues to

[19] In constructing these sub-samples, an industry is defined according to the Census of Population Industry Classification System, which includes 235 industry categories, 82 of which are in the manufacturing scctor.

[20] Due to sample size restrictions, income risk for these sub-samples are estimated at the 2-digit SIC

be greater than $\sigma^2_{\varepsilon,k=12}$ in each of the sub-samples. Note that income risk for those who stayed in the same manufacturing industry throughout the sample (STAY-IND) is the lowest, as these workers continue to earn returns on their industry-specific skills (even if they switch jobs within the sector). The mean estimate of the monthly value of $\sigma^2_{\varepsilon,k=12}$ for this sub-sample increases from 0.0008 to 0.0021 between the 1993 and 1996 panels, and then rises to 0.0025 in the 2001 panel. The corresponding annualized standard deviations are 0.098, 0.159 and 0.173 for the 1993, 1996 and 2001 panels, respectively. The risk faced by workers in STAY-MANUF, who stay within manufacturing throughout but may have switched from one industry to a different industry at some point in time, are close to (but in almost all cases smaller than) the risk faced by workers in STAY-IND. Workers in SWITCH-ALL who have switched to jobs in either a different industry within the manufacturing sector or to the non-manufacturing sector, face higher levels of risk. As Table IV indicates, the monthly variances for this group are 0.0029, 0.0030 and 0.0033 (with corresponding annualized standard deviations of 0.187, 0.19 and 0.199) for the 1993, 1996 and 2001 panels, respectively. Finally, the group with the highest risk are workers in SWITCH-NON-MANUF who switch out of the manufacturing sector altogether.[21]

III. Trade and Income Risk

The procedure outlined in the previous section provides us with estimates of individual income risk, $\sigma^2_{\varepsilon js}$, for each industry j and SIPP panel s. We now use these time-varying, industry-specific estimates in conjunction with observations on trade exposure to examine the relationship between income risk, $\sigma^2_{\varepsilon js}$, and import penetration, M_{js}.[22] In Figures III-A. and III-B, we plot the changes in estimated permanent income risk, $\Delta\sigma^2_{\varepsilon,k=0}$ and $\Delta\sigma^2_{\varepsilon,k=12}$, against changes in import penetration calculated at the beginning of each panel. More specifically, we plot differences in risk and import penetration between the 1993 and 1996 panels and between the 1996 and 2001 panels. In each case, for both K=0 and K=12, the relationship appears to be

level which is more aggregated than the Census classification used in constructing the sub-samples.

[21] This is true for all the specifications and panels except for the K=12 specification for the 1993 panel.

[22] Import penetration is defined as Imports/(Shipments - Exports + Imports).

strongly positive, suggesting that an increase in import penetration is associated with an increase in income risk for the workers in that industry.

III.1. Specification

More formally, we examine the relationship between income risk, $\sigma^2_{\varepsilon js}$, and import penetration, M_{js}, using a linear regression specification that includes industry fixed effects and time fixed effects:

$$\sigma^2_{\varepsilon js} = \alpha_s + \alpha_j + \alpha_M M_{js} + \nu_{js} . \qquad (7)$$

In (7), the inclusion of industry dummies, α_j, in the specification allows us to control for any time invariant industry-specific factors that may affect the level of riskiness of income in that industry. Similarly, the time dummy, α_s, 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 (such as business cycle effects and/or long-run structural changes) unrelated to trade, it also means that identification of the relationship between $\sigma^2_{\varepsilon js}$ and M_{js} will have to be based on the differential rate of change in import penetration across sectors over time. This, however, does not pose problems for our estimation since changes in import penetration over time do in fact exhibit substantial cross-sectional variation. For instance, the change in import penetration between 1993 and 1996 (1996 and 2001) varies between -0.03 and 0.08 (0.004 and 0.09), with a standard deviation of 0.025 (0.0026).

The estimates from (7) for our baseline sample reflect the impact of trade exposure on risk faced by individuals in the manufacturing sector. By repeating this analysis for various sub-samples described in Section II.6, we will be able to evaluate the relative magnitudes of the different channels through which international trade could affect individual income risk.

III.2. Endogeneity and Selection Bias

One potential concern with our estimation of equation (7), which relates trade to income risk, is that import penetration may not be fully exogenous to income risk. One possible reason for this is the endogenous choice of trade policies. While the large theoretical and empirical literature on the political economy of trade policy has not directly studied income risk as a determinant of cross-sectional variation in trade policy,[23] it is possible that trade policy, which affects import penetration, may itself be endogenously determined by income risk in the sector. Consider an "equity" minded government that uses trade policy to reach its goal of equalizing welfare across individuals in this economy. This government will choose high (low) protection levels for those industries with intrinsically high (low) levels of income risk, in order to say, increase (decrease) the mean level of wages in these industries. Nevertheless, our fixed-effects estimates of α_M, identified by within-industry variation, will not be biased due to such cross-sectional variation in the determinants of trade policy. But it is also plausible that this government could increase (decrease) protection and lower (raise) import penetration in industries that experience an increase (decrease) in income risk. If this is the case, such endogeneity of policy will bias our estimates of the relationship between income risk and import penetration (α_M) downwards (i.e., towards not finding a positive relationship between trade and risk) and therefore strengthen the results presented in this paper.

Another potential concern relates to the possibility that workers of different types may self-select into particular industries. Suppose, for example, that industries with high levels of import penetration are also industries with high job destruction rates. Suppose further that there are two types of workers, Type I and Type II, and that Type I workers quickly find a new job in the event of job displacement, but Type II workers do not. Other things being equal, we would expect Type II workers to move to low import penetration industries (or, over time, to industries in which import penetration has increased to a smaller extent relative to other industries). This type of self-selection, if present, would again bias our results against finding a positive

[23] See however, Davidson, Magee and Matusz (2005) for an interesting study of how the trade policy preferences of different economic groups within an industry may be shaped by the extent of job turnover rates in that industry.

association between income risk and import penetration. Nevertheless, we consider the importance of such selection and find that this concern is greatly mitigated for the following reasons. First, we examine industries over time, so any fixed differences across industries in worker characteristics are taken into account by our fixed effects estimation. Furthermore, we test whether the distribution of workers within an industry is related to change in import penetration in our data. We find that changes in the share of each education, occupation, age, gender or race group within a sector are completely uncorrelated with changes in import penetration across the span of the three SIPP panels. Finally, we examine the possibility that selection is based on unobserved ability differences across workers that are uncorrelated with educational attainment. In this case, we would expect selection to be reflected in unexplained wage differentials across sectors, as long as high-ability workers are paid higher wages. Our data suggests that such unobserved 'ability' differentials (that are uncorrelated with observable characteristics) across industries are small: About 80% of the cross-sectional variation in mean earnings across industries is explained by educational attainment alone. Moreover, we find no evidence of workers with different unobserved abilities selecting into sectors of different trade exposure. In our data, changes in unexplained portion of industry average wages are uncorrelated with changes in import penetration, further mitigating our concern regarding selection bias of this nature. In any event, we must note that a selection of the nature we have discussed would again bias our results against finding a positive relationship between exposure to trade and income risk (and thus would only strengthen our results reported below).[24]

III.3. Results: Full Sample

The results estimated for our full sample of workers using the specification described above are reported in Table V. We estimate two separate regressions described by (7), including, separately, import penetration at the beginning of each panel (i.e., for 1993, 1996 and 2001) and import penetration lagged one year (i.e., for 1992, 1995 and 2000). For each specification, the dependent variable is income risk measured either

[24] Another source of endogeneity in equation (7) is the omission of any time varying industry specific factors that are correlated with both income risk and import penetration simultaneously. We will explore this possibility in detail in Section III.5.

by filtering out purely transitory shocks ($\sigma^2_{\varepsilon,k=0}$) or by filtering out transitory shocks that last up to a year ($\sigma^2_{\varepsilon,k=12}$). Since the dependent variable is estimated, we adjust the standard errors for heteroscedasticity using a White correction.[25]

We find that import penetration is significantly associated with income risk in each of the specifications we examine. When only purely transitory shocks are filtered out, the coefficient on import penetration (measured at the beginning of each panel) is estimated to be $\hat{\alpha}_M$ = 0.022. This estimate indicates that an increase in import penetration by 10% of its initial (1993) level would raise $\sigma_{\varepsilon,k=0}$ by a little over 5%. In our preferred specification, when transitory shocks of duration up to a year are filtered out, the coefficient estimate is larger, $\hat{\alpha}_M$ = 0.045. This corresponds to an increase in $\sigma_{\varepsilon,k=12}$ by about 23%. Our estimates change very little when we instead include lagged values of import penetration as the independent variable.[26]

III.4. Results: Sub-Samples

In order to evaluate the effects of international trade on workers in different sub-groups, we next repeat the analysis described above for various sub-samples described in Section II.5. We estimate specification (7) separately for each sub-sample and as before, we include import penetration both as of the beginning of each panel and one year lagged. The results from specifications with $\sigma^2_{\varepsilon,k=0}$ and $\sigma^2_{\varepsilon,k=12}$ as the dependent variable are reported in Table VI-A and Table VI-B, respectively.

The first two columns of Table VI-A report the results using income risk estimates $\sigma^2_{\varepsilon,k=0}$ for the sub-sample STAY-IND as the dependent variable. When values of import penetration at the beginning of the panel are used as the explanatory variable, our estimates suggest that for workers who stayed in the same industry throughout,

[25] We also use weighted least squares (WLS) to correct for a heteroscedastic error structure, as suggested by Saxonhouse (1976). This correction has little effect on the magnitude or the significance of the coefficients on import penetration reported here.
[26] The coefficient on import penetration remains significant and positive with little change in its magnitude, when the dependent variable in specification (7) is replaced with the risk estimates from the Mincer specifications with time invariant coefficients described earlier.

$\sigma_{\varepsilon,k=0}$ would increase by 5% as a result of a 10% increase in import penetration over its initial value. When $\sigma^2_{\varepsilon,k=12}$ is the dependent variable, we find that the same increase in import penetration would result in an increase in $\sigma_{\varepsilon,k=12}$ of about 27% percent. The next two columns in Table VI report results for workers in the sub-sample STAY-MANUF, which includes workers who stay within the manufacturing sector (in the same industry or moving to another industry within manufacturing). Our estimates suggest that for this group, a 10% increase in import penetration is associated with an increase in $\sigma_{\varepsilon,k=0}$ and $\sigma_{\varepsilon,k=12}$ by about 5% and 22%, respectively.

Next, we focus exclusively on workers who switch industries. For the two sub-samples we consider here (SWITCH-ALL and SWITCH-NON-MANUF), the estimated coefficient on import penetration is positive in each specification but significant only when $\sigma^2_{\varepsilon,k=12}$ is the dependent variable.[27] The finding that We find that a 10% increase in import penetration leads to an increase in $\sigma_{\varepsilon,k=12}$ of 18% for workers who switch sectors (either within or outside the manufacturing sector) and of 22% for workers who switch to the non-manufacturing sector.

III.5. Robustness

All the specifications reported in Tables V and VI include both industry and year fixed effects in addition to import penetration (measured at the beginning of each panel and one-year lagged). These estimates will be biased if there are time varying industry specific factors that are correlated with both income risk and import penetration simultaneously. In the analysis that follows, we include additional explanatory variables to explore this possibility.

Specifically, we explore the following possibilities. First, we include share of exports

[27] Here, we are, in effect, asking whether workers who switch from sectors with bigger increases in import penetration face higher income risk. Since we are examining income risk "conditional on switching", we do not have a strong prior that the coefficient on import penetration should be different from zero – even taking as given the result that greater import penetration is associated with higher income risk for the full sample of workers. The positive coefficient on import penetration that we obtain implies that risk is indeed higher when switching from sectors with higher increases in import penetration – possibly because congestion with larger number of workers with similar skills leaving the industry at the same time leads to a greater variance in outcomes.

in total sales. If the risk faced by individuals employed in the export sector is lower, and exporting industries face lower import competition, then omission of this variable could lead to an overestimation of the coefficient on import competition. A second concern is that industries with high levels of final good imports tend to import high levels of intermediate inputs. Increased imports of intermediate inputs could lead to an increase in income risk due to an increased elasticity of labor demand (Rodrik, 1997). On the other hand, off-shoring could insulate domestic workers from output volatility by shifting the non-core activities of an industry abroad and hence decreasing risk for those who remain (Bergin, Feenstra and Hanson, 2009). To address this issue, we include share of imported intermediate inputs as a measure of off-shoring. Third, if industries respond to increased import competition by investing in information and communication technologies (ICT) and if such technology increases the risk faced by workers (for example, by increasing their substitutability with machines), this would lead to an upward bias in our coefficient of interest. Fourth, we include labor productivity against the possibility that a negative productivity shock in an industry could simultaneously lead to an increase in both import penetration and in income risk. Finally, omission of union density could bias our estimates if union density changes in response to increased import competition and if higher unionization rates are associated with lower levels of risk. In Table VII, we report the summary statistics for each of these variables calculated at the beginning of each panel.

We report our estimation results in Table VIII. As before, each specification reported includes industry and year fixed effects. All explanatory variables are measured as of the first year of each panel (columns 7-11) and in one-year lags (columns 1-6). For brevity, we report the results with our preferred income risk estimates (allowing for transitory shocks that last up to a year) as the dependent variable. In columns (2) and (8), we include share of exports in addition to share of imports. The coefficient of import penetration remains significant and positive with little change in its magnitude. The coefficient of exports is insignificant. Inclusion of offshoring leads to an increase in the coefficient of import penetration. In the specifications reported here, the offshoring variable is significant and negative, suggesting that an increase in offshoring in an industry is associated with a decline in income risk in that industry. Inclusion of ICT, labor productivity and union density does not affect the coefficient on import penetration.[28] [29]

Another robustness check we consider is to allow income risk to vary by individual characteristics within an industry. More specifically, we estimate risk separately for each age group and education level within an industry for each panel and estimate equation (7) by including dummy variables for age and education in addition to import penetration and time and industry fixed effects. In both cases, the coefficient on import penetration remains significant and positive, with little change in its magnitude.[30]

IV. Welfare

The preceding sections have focused on estimating the relationship between trade exposure and income risk. We now turn to the analysis of the link between income risk and welfare using a simple dynamic model with incomplete markets and (exclusively) permanent income shocks, developed by Krebs (2004) and implemented in Krebs, Krishna and Maloney (2008). The model is tractable enough to permit closed-form solutions for equilibrium consumption and welfare, yet rich enough to provide a tight link to the empirical analysis we have outlined. Clearly, our goal here is not to provide a complete assessment of the effects of international trade on welfare, taking into account all possible channels, but rather to obtain suggestive estimates of welfare change exclusively through the income risk channel.[31]

[28] In specifications not reported here, we also consider the effect of including the share of foreign multinationals (MNE) in total industry employment. Exclusion of this variable could lead to an upward bias in the magnitude of the coefficient of import competition if an increase in MNE share is associated with a decrease in imports in that industry and if employment in such firms is more stable than that of domestic firms. Since the MNE measure comparable across time is available until 1996, we check the robustness of our results to the inclusion of this variable for only the 1993 and 1996 panels. We find that the coefficient on import penetration remains positive and significant, while the coefficient on MNE share is insignificant.

[29] We also estimate Equation (7) by including each additional explanatory variable one-by-one along with import penetration. In each of these specifications, the coefficient on import penetration remains significant with little or no change in its magnitude.

[30] These results are available from the authors upon request.

[31] While our focus in this paper is on the welfare effects of international trade solely through the income risk channel, we have also explored the relationship between mean growth rates of (residual) income and import penetration (using econometric specifications like (7) – with income growth on the left hand side rather than income risk). However, we do not find any consistent relationship between mean growth rates of (residual) income and import penetration.

The specific thought experiment that the theoretical structure allows us to answer is the following one (Krebs (2004)): Imagine a group of ex-ante identical workers with Constant Relative Risk Aversion (CRRA) preferences facing an income process with variance of permanent income risk σ_s^2. Assume that workers are unable to insure themselves against permanent shocks to their labor income (market incompleteness),[32] and that they can only use their own savings to smooth consumption. Consider now an increase in permanent income risk measured by Δ_σ, so that $\sigma_{s'}^2 = (1+\Delta_\sigma)\sigma_s^2$ is now the risk to income that they face forever going forward. What is the welfare effect of this increase in risk, in compensating variation terms?

It can be shown (Krebs (2004)) that the percent change in consumption Δ_c, in each period and each state of the world, required to compensate the individual for the change in risk Δ_σ is given by:[33]

$$\Delta_c = \left(\frac{1-\beta(1+\mu)^{1-\gamma}\exp(.5\gamma(\gamma-1)(1+\Delta_\sigma)\sigma_\varepsilon^2)}{1-\beta(1+\mu)^{1-\gamma}\exp(0.5\gamma(\gamma-1)\sigma_\varepsilon^2)} \right)^{\frac{1}{1-\gamma}} -1, \text{ if } \gamma \neq 1$$

and $$\Delta_c = \left(\frac{\beta\Delta_\sigma\sigma_\varepsilon^2)}{(1-\beta)^2 2} \right) -1, \text{ if } \gamma = 1 \qquad (9)$$

where β is the pure discount factor, γ the coefficient of relative risk aversion, μ the mean growth rate of income and σ_ε^2 the estimated variance of the permanent component of labor income shocks.

The welfare expression (9) has standard properties. With $\gamma > 0$, individuals are risk averse and risk is costly. That is, an increase in risk, $\Delta_\sigma > 0$, requires positive compensation, $\Delta_c > 0$, for the individual to be just as well off as before. The

[32] We should note that not allowing insurance against permanent labor income shocks is not particularly restrictive. As a practical matter, direct insurance against labor income shocks is generally not available to workers. More importantly, our results concerning either the estimates of permanent income risk or its links with trade do not change in the slightest when total income (including any capital earnings and transfers) instead of labor earnings are used as our income measure.

[33] The interested reader is referred to Krebs, Krishna and Maloney (2008) for a detailed derivation and discussion.

magnitude of this compensation is increasing in the degree of risk aversion, γ. Using (9) along with estimates of change in risk associated with trade, Δ_σ (from Sections III.3 and III.4), and standard values for the parameters β and γ, we could obtain suggestive estimates of the benefits or costs of trade through the income risk channel.

The welfare expression (9) is derived under the assumption that increase in permanent income risk, Δ_σ, associated with the increase in import penetration lasts forever. Similarly, specification (7) is a "long-run" specification associating the level of import penetration with the level of income risk.[34] However, since our data spans only a 10 year period (between 1993-2003), our estimates, strictly speaking, do not allow us to reject the hypothesis that changes in income risk associated with changes in import penetration do not last longer than 10 years. We therefore conduct the quantitative welfare analysis by allowing for income risk to be higher with higher import penetration for a period of T = 10 years, while also reporting calculations for T = 5 (shorter duration) and 15 years (longer duration).[35]

The welfare change corresponding to a change in the variance of the permanent income shocks (income risk) for T years is given by (Krebs (2004)):

$$\Delta c = [(1-x)(1-x'^{T+1})/(1-x') + xx'^{T}]^{1/(\gamma-1)} - 1, \text{ if } \gamma \neq 1 \qquad \text{and} \qquad (10)$$

$$\Delta_c = \left(\frac{\beta(1-\beta^T)\Delta_\sigma \sigma_\varepsilon^2)}{(1-\beta)^2 2} \right) - 1, \text{ if } \gamma = 1 \text{ where,}$$

$$x = \beta(1+\mu)^{1-\gamma} \exp(0.5\gamma(\gamma-1)\sigma_\varepsilon^{\ 2}) \text{ and}$$

$$x' = \beta(1+\mu)^{1-\gamma} \exp(0.5\gamma(\gamma-1)(1+\Delta_\sigma)\sigma_\varepsilon^{\ 2})$$

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[34]To ensure that the increases in income risk we estimate in (7) are indeed "long-run" changes, we have also estimated variants of specification (7) by including changes in import penetration in preceding periods on the right hand side. We find that while the coefficient on the level of import penetration remains unchanged, the lagged (1 and 2 year) changes in import penetration, capturing purely "short-run" effects, are not significant.

[35] Note that even when the increase in permanent income risk with greater import penetration lasts only for a temporary period of T years, any shocks to worker incomes $\varepsilon_{ij,t}$ have permanent effects. Specifically, when permanent income risk rises for a duration of time T, workers draw their permanent income innovation terms $\varepsilon_{ij,t}$ in (3) from a bin with greater variance $\sigma_{\varepsilon'}^2$ than before, for duration T, before returning to a bin with the original level of σ_ε^2.

Table IX provides welfare calculations using our preferred income risk estimates, $\sigma^2_{\varepsilon,k=12}$ as well as results when income risk is estimated assuming K=0 ($\sigma^2_{\varepsilon,k=0}$). Results are provided separately for parameter values for the coefficient of risk aversion at γ = 1 and γ = 2 and for durations of T = 5, 10 and 15 years. All of the calculations use a discount factor β = 0.98. With γ = 2, for our central set of risk estimates with K=12, the increase in persistent income risk associated with a 10% increase in import penetration is certainty equivalent to a reduction in lifetime consumption in the range of 4% to 11%. On the other hand, with γ = 2 and K=0, the welfare cost is estimated instead to be between 2% and 6% reduction in lifetime consumption. In Table IX, we also report welfare estimates corresponding to a lower level of risk aversion, γ = 1. As expected, welfare costs are smaller when individuals are less risk averse. However, in both cases, the welfare costs associated with the income risk channel are economically quite significant.

We emphasize here again that our analysis has focused exclusively on the link between trade and income risk. Our results should be considered alongside the findings of a large literature on international trade, which has explored the many ways in which exposure to trade may positively affect the economy. Our finding of economically significant negative effects through the income risk channel does not suggest that the gains from trade are negative overall. It indicates instead that the income risk channel should be considered seriously in exercises evaluating the gains from trade.

V. Conclusions

This paper studies the links between international trade and individual income risk using longitudinal earnings data on workers in the United States. We find increased import penetration to have a statistically and economically significant effect on labor income risk in US manufacturing. We find evidence of increased income risk for workers who stayed within the same manufacturing industry throughout as well as for those who either change manufacturing industries or move out of manufacturing altogether. The welfare effects of the increased income risk are economically significant. For our central set of estimates, the welfare cost of the increase in risk associated with a 10 percent increase in import penetration is in the range of 4 % to 11 % of lifetime consumption. Our analysis has focused exclusively on the links between trade exposure and income risk. Our finding of economically significant negative effects through the income risk channel does not suggest that the gains from trade are negative overall. It indicates instead that the income risk channel should be considered seriously in exercises evaluating the overall gains from trade.

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Table I. Summary Statistics

	1993		1996		2001	
Variable	*Mean (All)*	*Mean (Manuf.)*	*Mean (All)*	*Mean (Manuf)*	*Mean (All)*	*Mean (Manuf)*
Log (Real Earnings)	7.34	7.64	7.37	7.61	7.46	7.67
Age	35.39	37.51	36.62	37.97	37.40	39.34
Variable	*Percent (All)*	*Percent (Manuf)*	*Percent (All)*	*Percent (Manuf)*	*Percent (All)*	*Percent (Manuf)*
High school drop out	17.53	19.55	11.49	14.77	11.55	13.78
High school graduate	38.1	43.86	36.37	43.51	33.87	41.07
Some college	21.92	19.26	29.76	26.07	30.11	27.06
College graduate	12.73	10.96	15.51	11.77	16.69	13.25
More than college	9.72	6.37	6.87	3.88	7.79	4.85
Female	48.32	32.72	49.04	35.63	48.68	32.76
Married	56.99	64.35	57.75	62.87	56.32	62.44
White	78.37	78.35	73.05	73.33	69.72	69.97
N	24,998	4,471	41,008	7,270	37,579	5,647

Table II. Risk Estimates

	Mean	**Median**	**Std. Dev.**
1993-1995			
$\sigma^2_{\varepsilon,k=0}$	0.0033	0.0031	0.0016
$\sigma^2_{\varepsilon,k=6}$	0.0018	0.0015	0.0016
$\sigma^2_{\varepsilon,k=12}$	0.0014	0.0014	0.0019
1996-1998			
$\sigma^2_{\varepsilon,k=0}$	0.0043	0.0042	0.0013
$\sigma^2_{\varepsilon,k=6}$	0.0024	0.0023	0.0014
$\sigma^2_{\varepsilon,k=12}$	0.0025	0.0026	0.0018
2001-2003			
$\sigma^2_{\varepsilon,k=0}$	0.0052	0.0051	0.0016
$\sigma^2_{\varepsilon,k=6}$	0.0033	0.0034	0.0019
$\sigma^2_{\varepsilon,k=12}$	0.0031	0.0032	0.0025

Reported mean, median and standard deviations are calculated across point estimates for eighteen 2-digit SIC industries.

Table III Risk Estimates by Industry for each Panel ($\sigma^2_{\varepsilon,k=0}$ and $\sigma^2_{\varepsilon,k=12}$)

SIC	$\sigma^2_{\varepsilon,k=0}$			$\sigma^2_{\varepsilon,k=12}$		
	1993-1995	**1996-1998**	**2001-2003**	**1993-1996**	**1996-1998**	**2001-2003**
20	0.004*** (0.0002)	0.004*** (0.0001)	0.005*** (0.0002)	0.003*** (0.0005)	0.000 (0.0004)	0.004*** (0.0005)
22	0.006*** (0.0003)	0.003*** (0.0002)	0.004*** (0.0003)	0.005*** (0.0008)	-0.000 (0.0006)	0.004*** (0.0009)
23	0.003*** (0.0002)	0.005*** (0.0002)	0.009*** (0.0004)	0.001** (0.0006)	0.004*** (0.0007)	0.010*** (0.0011)
24	0.004*** (0.0003)	0.005*** (0.0003)	0.004*** (0.0002)	0.003*** (0.0008)	0.006*** (0.0008)	0.002*** (0.0006)
25	0.003*** (0.0003)	0.003*** (0.0003)	0.002*** (0.0003)	-0.000 (0.0010)	0.001 (0.0008)	0.002*** (0.0007)
26	0.003*** (0.0002)	0.004*** (0.0002)	0.005*** (0.0003)	0.001** (0.0006)	0.004*** (0.0006)	-0.000 (0.0008)
27	0.005*** (0.0002)	0.004*** (0.0002)	0.005*** (0.0002)	0.002*** (0.0005)	0.003*** (0.0005)	0.001 (0.0006)
28	0.003*** (0.0002)	0.004*** (0.0002)	0.006*** (0.0002)	0.001* (0.0005)	0.001** (0.0005)	0.003*** (0.0005)
30	0.002*** (0.0003)	0.003*** (0.0002)	0.007*** (0.0003)	-0.002*** (0.0007)	0.000 (0.0005)	0.006*** (0.0008)
31	-0.000 (0.0007)	0.003*** (0.0005)	0.004*** (0.0004)	-0.001 (0.0016)	0.003** (0.0012)	0.006*** (0.0010)
32	0.005*** (0.0003)	0.004*** (0.0002)	0.006*** (0.0003)	0.004*** (0.0010)	0.003*** (0.0007)	0.004*** (0.0009)
33	0.002*** (0.0002)	0.004*** (0.0002)	0.005*** (0.0003)	-0.001*** (0.0006)	0.001** (0.0005)	-0.001 (0.0008)
34	0.004*** (0.0001)	0.003*** (0.0001)	0.004*** (0.0002)	0.003*** (0.0004)	0.002*** (0.0004)	0.001** (0.0004)
35	0.002*** (0.0001)	0.004*** (0.0001)	0.005*** (0.0001)	0.001*** (0.0003)	0.002*** (0.0004)	0.002*** (0.0005)
36	0.003*** (0.0001)	0.003*** (0.0001)	0.005*** (0.0002)	0.002*** (0.0003)	0.002*** (0.0003)	0.002*** (0.0005)
37	0.003*** (0.0001)	0.005*** (0.0001)	0.005*** (0.0001)	0.002*** (0.0004)	0.003*** (0.0004)	0.003*** (0.0004)
38	0.002*** (0.0002)	0.005*** (0.0002)	0.006*** (0.0003)	0.001** (0.0006)	0.005*** (0.0006)	0.003*** (0.0008)
39	0.006*** (0.0004)	0.008*** (0.0004)	0.007*** (0.0004)	0.001 (0.0014)	0.005*** (0.0012)	0.005*** (0.0012)

Robust standard errors in parantheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table IV. Income Risk in Sub-Samples

	$\sigma^2_{\varepsilon,k=0}$		$\sigma^2_{\varepsilon,k=12}$	
1993-1995	Mean	Std. Dev	Mean	Std. Dev.
SWITCH_NON-MANUF	0.0063	0.0033	0.0026	0.0053
SWITCH_ALL	0.0059	0.0029	0.0029	0.0050
STAY_MANUF	0.0027	0.0014	0.0011	0.0019
STAY_IND	0.0024	0.0012	0.0008	0.0016
1996-1998				
SWITCH_NON-MANUF	0.0082	0.0031	0.0036	0.0055
SWITCH_ALL	0.0067	0.0026	0.0030	0.0043
STAY_MANUF	0.0033	0.0010	0.0021	0.0017
STAY_IND	0.0031	0.0008	0.0021	0.0015
2001-2003				
SWITCH_NON-MANUF	0.0090	0.0032	0.0039	0.0057
SWITCH_ALL	0.0081	0.0026	0.0033	0.0036
STAY_MANUF	0.0039	0.0017	0.0024	0.0023
STAY_IND	0.0037	0.0017	0.0025	0.0025

Reported mean, median and standard deviations are calculated across point estimates for eighteen 2-digit SIC industries.

Table V. International Trade and Income Risk: Full Sample

	$\sigma^2_{\varepsilon,k=0}$		$\sigma^2_{\varepsilon,k=12}$	
Import penetration (Lagged)	0.023** (0.009)		0.042*** (0.014)	
Import penetration		0.022** (0.010)		0.045*** (0.013)
Constant	0.003*** (0.000)	0.003*** (0.000)	0.001 (0.001)	0.001 (0.001)
R-squared	0.71	0.70	0.58	0.60
N	54	54	54	54

Robust standard errors in parantheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table VI-A International Trade and Income Risk: Sub-Samples ($\sigma^2_{\varepsilon,k=0}$)

	STAY IND		STAY MANUF		SWITCH ALL		SWITCH NON-MANUF	
Import Penetration (Lagged)	0.017** (0.0084)		0.019* (0.0097)		0.028* (0.0157)		0.023 (0.0201)	
Import Penetration		0.015* (0.0089)		0.017* (0.010)		0.027 (0.0169)		0.024 (0.0218)
Constant	0.002*** (0.0002)	0.002*** (0.0003)	0.002*** (0.0003)	0.002*** (0.0007)	0.007*** (0.001)	0.007*** (0.0011)	0.008*** (0.001)	0.008*** (0.0014)
R-squared	0.59	0.58	0.62	0.61	0.61	0.61	0.55	0.55
N	54	54	54	54	54	54	54	54

Robust standard errors in parantheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table VI-B International Trade and Income Risk: Sub-Samples ($\sigma^2_{\varepsilon,k=12}$)

	STAY IND		STAY MANUF		SWITCH ALL		SWITCH NON-MANUF	
Import Penetration (Lagged)	0.028* (0.0158)		0.031* (0.0159)		0.070*** (0.0240)		0.081** (0.0330)	
Import Penetration		0.031* (0.0157)		0.034** (0.0157)		0.070*** (0.0251)		0.081** (0.0344)
Constant	0.000 (0.0008)	0.000 (0.0009)	0.001 (0.0008)	0.000 (0.0009)	0.003 (0.0027)	0.003 (0.0028)	0.002 (0.0033)	0.002 (0.0034)
R-squared	0.50	0.51	0.51	0.53	0.50	0.49	0.42	0.42
N	54	54	54	54	54	54	54	54

Robust standard errors in parantheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table VII Summary Statistics: Explanatory Variables

Variable	Mean	Std. Dev.	Min	Max
1993				
Import Penetration	0.169	0.140	0.014	0.561
Share of Exports	0.101	0.063	0.022	0.235
Offshoring	0.148	0.082	0.039	0.324
Share of ICT	0.080	0.058	0.029	0.225
$(\text{Labor Productivity})_{t-1}$	1.098	0.112	0.981	1.474
Union Density	0.188	0.105	0.072	0.398
1996				
Import Penetration	0.192	0.158	0.015	0.638
Share of Exports	0.122	0.080	0.022	0.282
Offshoring	0.160	0.080	0.047	0.352
Share of ICT	0.082	0.057	0.028	0.219
$(\text{Labor Productivity})_{t-1}$	1.232	0.343	0.963	2.464
Union Density	0.171	0.105	0.036	0.391
2001				
Import Penetration	0.234	0.178	0.019	0.717
Share of Exports	0.138	0.092	0.023	0.320
Offshoring	0.192	0.097	0.054	0.393
Share of ICT	0.082	0.057	0.024	0.222
$(\text{Labor Productivity})_{t-1}$	1.769	1.519	1.076	7.464
Union Density	0.148	0.085	0.043	0.317

These summary statistics are calculated at the beginning of each panel, except labor productivity. Since this variable is not available after 2000, summary statistics for one year lags are reported.
Import Penetration=Imports/Shipments-exports+imports
Share of Exports=Exports/Shipments

$$\text{Offshoring} = \sum_j \left[\frac{\text{purchases of input } j \text{ by industry } i \text{ at time } t}{\text{total non} - \text{energy inputs used by industry } i \text{ at time } t} \right] * \left[\frac{\text{imports of input } j \text{ at time } t}{\text{production}_j + \text{imports}_j - \text{exports}_j \text{ at time } t} \right]$$

Source: Federal Reserve Bank of New York
Share of ICT= (Software+Computers and peripheral equipment+Communication equipment + Photocopy and related equipment+Instruments)/K. *Source: BEA, NIPA*
Labor productivity=Output/Hours. Base year: 1987. Aggregated to 2-digit SIC using employment shares as of 1992 as weights. *Source: BLS*
Union Density= (Union Members)/Employment. *Source:* Hirsch and Macpherson (2003)

Table VIII Robustness ($\sigma^2_{\varepsilon,k=12}$)

Full Sample	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Import Penetration (lagged)	0.042*** (0.014)	0.044** (0.020)	0.050** (0.019)	0.049** (0.020)	0.050** (0.019)	0.053*** (0.019)					
Share of exports (lagged)		-0.005 (0.018)	-0.002 (0.019)	-0.001 (0.020)	0.004 (0.022)	-0.009 (0.023)					
Offshoring (lagged)			-0.023* (0.011)	-0.022* (0.012)	-0.024* (0.013)	-0.021 (0.013)					
Share of ICT (lagged)				0.020 (0.036)	-0.014 (0.040)	-0.038 (0.046)					
Labor Productivity (lagged)					-0.001 (0.000)	0.000 (0.000)					
Union Density (lagged)						0.027 (0.016)					
Import Penetration							0.045*** (0.013)	0.045** (0.019)	0.057*** (0.020)	0.056** (0.021)	0.055** (0.020)
Share of exports								0.000 (0.016)	0.001 (0.017)	0.004 (0.017)	0.001 (0.017)
Offshoring									-0.044** (0.018)	-0.043** (0.019)	-0.039* (0.021)
Share of ICT										0.031 (0.034)	0.023 (0.038)
Union Density											0.013 (0.014)
Constant	0.001 (0.001)	0.001 (0.001)	0.001 (0.001)	0.000 (0.002)	0.002 (0.003)	-0.003 (0.004)	0.001 (0.001)	0.001 (0.001)	0.002 (0.001)	0.000 (0.002)	-0.003 (0.004)
R-squared	0.58	0.58	0.61	0.61	0.62	0.66	0.60	0.60	0.63	0.64	0.65
N	54	54	54	54	54	54	54	54	54	54	54

Robust standard errors in parantheses. * significant at 10%; ** significant at 5%; *** significant at 1%. Since comparable data for labor productivity is not available after 2000, the estimates from the specification including productivity as of the beginning of the panel are not included in this table.

Table IX. Welfare Effects (Percent of Lifetime Consumption)

		T=5	T=10	T=15
K=0	$\gamma = 1$	1.06	2.02	2.90
	$\gamma = 2$	2.19	4.32	6.39
K=12	$\gamma = 1$	2.18	4.18	6.03
	$\gamma = 2$	4.24	8.00	11.34

Figure I. Variance in Wage outcomes

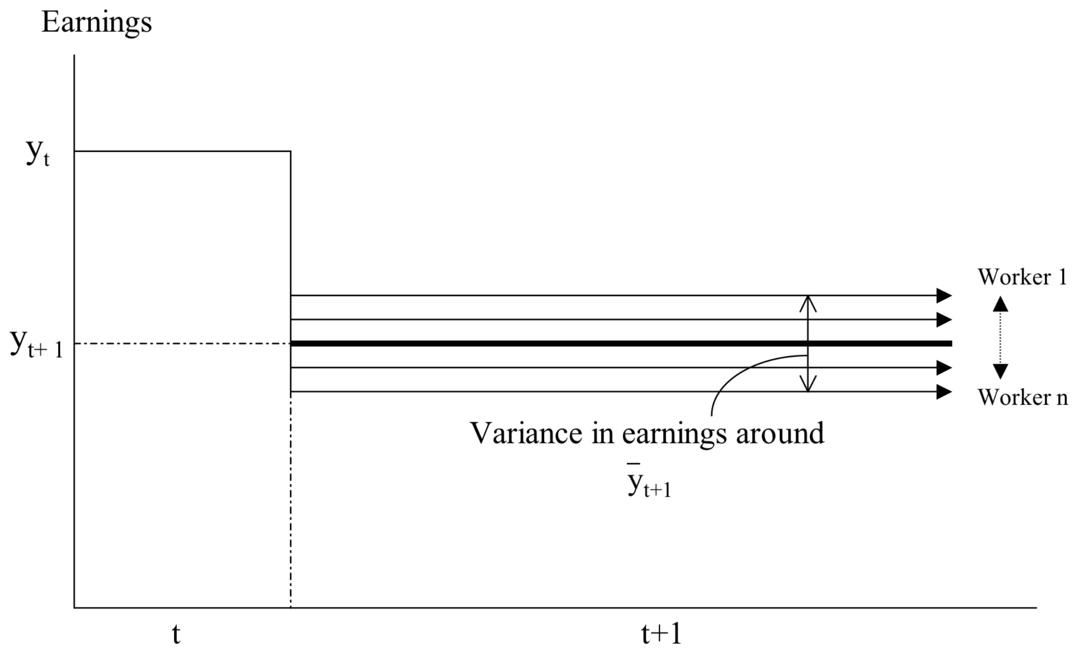


Figure II. Transitory versus Permanent Shocks

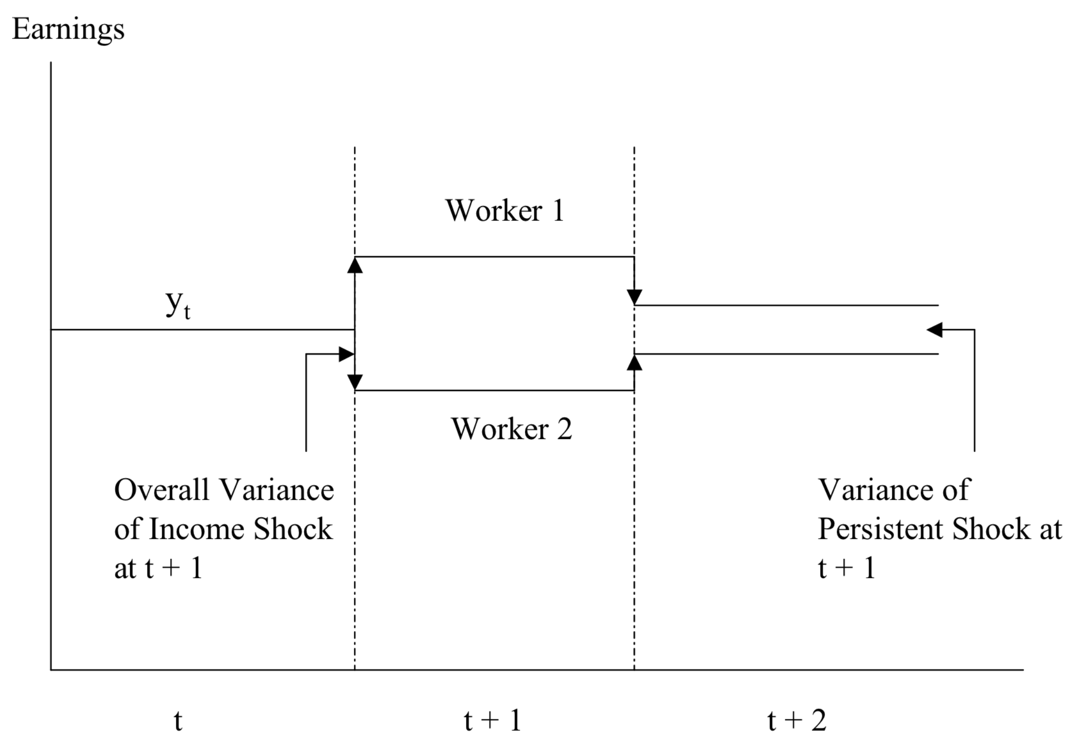
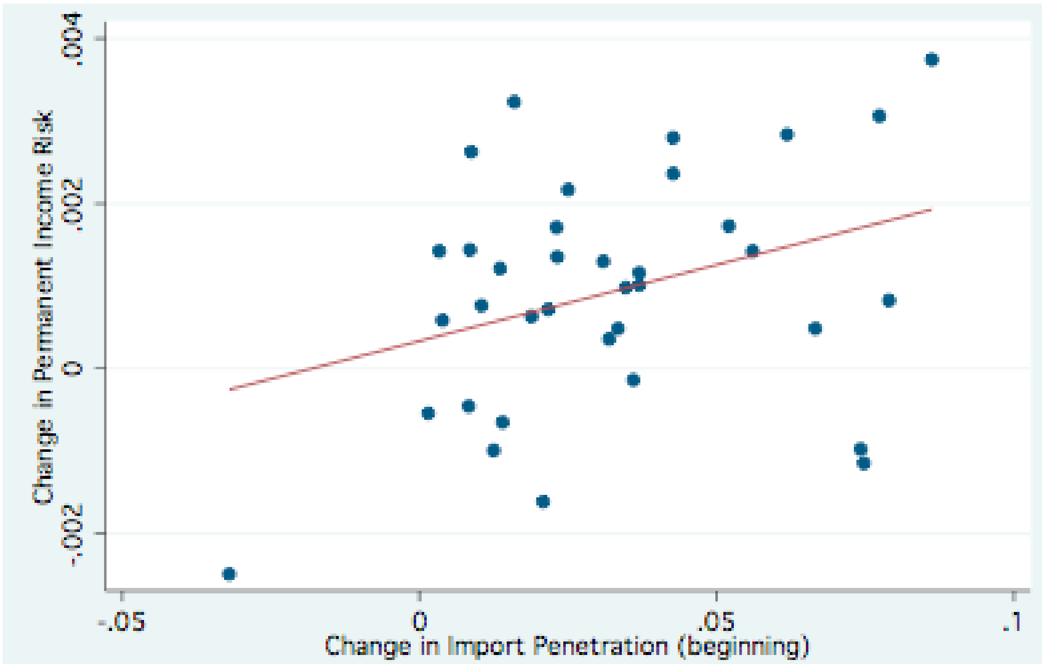


Figure III. Changes in Permanent Income Risk and Changes in Import Penetration

A. $\sigma^2_{\varepsilon,k=0}$



B. $\sigma^2_{\varepsilon,k=12}$