

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. We also study risk estimates from various sub-samples of workers, such as those who switched to a different industry (or sector) and those who remained in the same industry throughout. 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. Our welfare calculations suggest that these effects are economically significant and that a sizeable social safety net would be necessary to insure workers against this risk.

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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. 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 with greater volatility (variance) in their incomes.¹

How might trade openness affect labor income volatility? The theoretical literature has suggested various channels through which exposure to international trade can affect labor income risk. Rodrik (1997) has argued that this may happen because 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 will result in larger variations in wages and employment, and hence increase the volatility in the labor market. Similarly, Newberry and Stiglitz (1984) have argued that while in a closed economy, domestic price adjustments insulate producers against supply shocks, producers are constrained by world prices in an open economy. This implies that productivity shocks will have a smaller equilibrium effect on output and employment in a closed economy compared to an open one. Furthermore, openness can increase labor income risk by exposing import-competing sectors to a variable international economic environment. Thus, changing patterns of comparative advantage will induce continuous reallocations of capital and labor across tasks within firms and across firms within and between sectors.² This process may not be an orderly or costless one. To the extent that similar workers experience different outcomes in the process, they are exposed to labor income risk (defined as the variance of unpredictable changes in income).

¹ 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.

² See, for instance, the recent study of Menezes-Filho and Meundler (2007) which comprehensively documents the patterns of worker reallocation in response to trade policy changes in Brazil.

We should note at the outset that labor income risk, which measures the variance of income *changes* is a distinct concept from wage inequality which has been the focus of a large theoretical and empirical literature in international trade. A number of researchers have examined the implications of the theoretical “Stolper-Samuelson” prediction that trade openness will lead to an increase in earnings of abundant factors and a reduction in the earnings of scarce factors (see for instance, Lawrence and Slaughter (1993), Leamer (1996), Feenstra and Hanson (1999) and Goldberg and Pavcnik (2005)).³ More recently, a number of papers have developed and empirically tested predictions concerning the effects of globalization on wages, wage inequality and employment in the presence of firm heterogeneity and labor market frictions (see, for instance, Amiti and Davis (2008), Davis and Harrigan (2007), Egger and Kreickemeier (2009), Davidson, Matusz and Shevchenko (2008) and Mitra and Ranjan (2009)). Ohnsorge and Trefler (2007) have explored the consequences of worker heterogeneity for trade and the domestic distribution of income. Finally, the recent papers by Helpman, Itskhoki and Redding (2008a, 2008b) and Helpman and Itskhoki (2007) have developed a rich set of predictions concerning the effects of openness on wage inequality and unemployment in a world with heterogeneity in both workers and firms. While the links between trade, wage levels and wage inequality are clearly highly important issues to study, our focus is on a different dimension of the labor market experience – the variability (risk) in incomes experienced by workers.

This paper conducts an empirical analysis of the links between trade and individual labor income risk in the United States. We use longitudinal data on workers to estimate idiosyncratic labor income risk and to study the role of trade in explaining the variation in risk across workers employed in different industries.⁴ In estimating labor income risk, 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

³ Feenstra and Hanson (2002), Davidson and Matusz (2004), Goldberg and Pavcnik (2007) and Harrison (2007) provide excellent survey treatments.

⁴We use the Survey of Income and Program Participation (SIPP) in our analysis. SIPP contains longitudinal panels on individuals, 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.

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. We also 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. Finally, we use our empirical estimates to conduct a simple welfare analysis to obtain indicative estimates of the benefits or costs of trade through the income risk channel.⁵

Our empirical results for the United States can be summarized as follows. First, we find that income risk is increasing over time for both the full sample of workers as well as workers in each sub-sample. Second, we find that 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 industry throughout the sample; among switchers, risk for those who switched to the non-manufacturing sector is higher than those who switched within manufacturing.⁶ Finally, and most importantly, 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

⁵ 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 providing more precise estimates of the magnitude of persistent shocks to income. Second, we estimate risk faced by various sub-samples of workers and study the differential impact of trade exposure on workers in these groups. Furthermore, we use the greater availability of data in the United States on a variety of industry characteristics to include necessary controls in our econometric analysis, as discussed later in the paper.

⁶ As we will discuss in detail later in the paper, the estimates of income risk for the different groups of workers reflect the differences in worker characteristics and the endogenous actions that place workers in these different sub-samples, and should be interpreted with this qualification in mind.

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. 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.⁷ Our welfare calculations suggest that these effects are economically significant and that a sizeable social safety net would be necessary to insure workers against this risk.

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

⁷ 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.

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 ($y_t - \bar{y}_{t+1}$).⁸ To see the distinction between income risk and income inequality, consider a second time period, $t+2$ (not illustrated in Figure I), in which the distribution of incomes stays the same as in time period $t+1$, but where workers (stochastically) exchange positions with each other under the same income distribution. Here, while income inequality stays the same between $t+1$ and $t+2$, workers experience risky income changes, which will be captured in our estimates of labor income risk.

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 but not when income shocks are highly persistent or permanent. Thus, highly persistent income shocks have a large effect on consumption

⁸ 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 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.

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.

II.1. Data

As we have discussed, while the aggregate distribution of earnings may stay the same across different time periods, we may see stochastic (risky) transitions taking place underneath – with some individuals at the top of the distribution possibly exchanging places with others at the bottom end of the distribution. These income changes reflect the risk that workers are exposed to and would be captured as income risk, but would not be reflected in estimates of the change in the variance of income. Therefore, to estimate the risk in incomes faced by these individuals, longitudinal data tracking these individual transitions is useful to have.

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 (SIPP).⁹ 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.¹⁰ 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

¹⁰ We limit our main analysis to data from the first three years of the 1996 panel to ensure comparability of our risk estimates from 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.

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.¹¹ 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. Worker's 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.

II.2. Specification

To estimate labor income risk, we follow the approach taken in previous empirical work on this topic (see for instance, Carroll and Samwick (1997), Gourinchas and Parker (2002) and Meghir and Pistaferri (2004)).¹² Our survey data provide us with earnings (wage rate times number of hours worked) of individuals. 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:

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

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} represent 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

¹¹ This results in the omission of approximately 13% of the respondents of each panel from our sample.

¹² We should note that these papers have pursued the empirical estimation of labor income largely at the "macroeconomic" level. None has examined the variation of income risk across industries nor has focused on the relationship between labor income risk and international trade, as we do in this paper.

contained in changes in u_{ijt} . In this sense, changes in u_{ijt} over time measure the unpredictable part of changes in individual income.¹³

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

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

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

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

where the innovation terms, $\{\varepsilon_{ijt}\}$, are independently distributed over time and identically distributed across individuals, $\varepsilon_{ijt} \sim N(0, \sigma_{\varepsilon js}^2)$, 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 $\{\eta_{ijt}\}$ are independently distributed over time and identically distributed across individuals, $\eta_{ijt} \sim N(0, \sigma_{\eta js}^2)$. 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 $\sigma_{\varepsilon js}^2$ and $\sigma_{\eta js}^2$ will therefore give us industry specific, time varying estimates of transitory and persistent 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

¹³ 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, the inclusion of individual-fixed effects in specification (1) should not and do not alter our risk estimates.

our analysis using alternate specifications. As we have just discussed, (1) takes out any changes in 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 is 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).

Finally, notice that the inclusion of industry dummies in (1) filters out mean income changes in an industry and thus any volatility in the changes of the mean industry earnings from our measure of individual risk. Our risk estimates therefore measure idiosyncratic income risk (effectively individual variation around the industry mean, conditional on the other covariates in (1)).¹⁴

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. 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:

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

¹⁴ 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 data we do not find evidence of any relationship between the variance of changes in mean industry earnings and import penetration.

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 $\sigma_{\varepsilon,k=0}^2$, $\sigma_{\varepsilon,k=6}^2$ and $\sigma_{\varepsilon,k=12}^2$, 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 $\sigma_{\varepsilon,k=0}^2 > \sigma_{\varepsilon,k=6}^2 > \sigma_{\varepsilon,k=12}^2$ (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 $K=12$. $\sigma_{\varepsilon,k=12}^2$ 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.

Our intention is to estimate parameters measuring income risk and see how changes in these parameters over time (i.e., across panels) are 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.¹⁵

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):

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

¹⁵ 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.

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

$$\text{var}[\Delta_n u_{ijt}] = \sigma_{\varepsilon_j,t+1}^2 + \dots + \sigma_{\varepsilon_j,t+n}^2 + \sigma_{\eta_{jt}}^2 + \sigma_{\eta_{j,t+n}}^2. \quad (5)$$

As noted earlier, the parameters $\sigma_{\varepsilon_j}^2$ and $\sigma_{\eta_j}^2$ 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:

$$\text{var}[\Delta_n u_{ijt}] = 2\sigma_{\eta_j}^2 + n\sigma_{\varepsilon_j}^2 \quad (6)$$

Thus, the variance of observed n -period income changes is a linear function of n , where the slope coefficient is equal to $\sigma_{\varepsilon_j}^2$. 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 $\text{var}[\Delta_n u_{ijt}]$, the square of the individual deviation from mean income difference over the n periods, on n . (6) is estimated separately for each industry and panel.

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 this 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. Specifically, the sample analogs to $\text{var}[\Delta_n u_{ijt}]$ are formed by estimating (residual) 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 than 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.¹⁶

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.¹⁷ 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.¹⁸

¹⁶ We also find that the change in attrition rates between panels is not correlated with change in import 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.

¹⁷ 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.

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

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 the 1996 and 2001 panels.

Table II also reports the summary statistics for the estimates of $\sigma_{\varepsilon,k=6}^2$ and $\sigma_{\varepsilon,k=12}^2$. 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 $\sigma_{\varepsilon,k=12}^2$ 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 $\sigma_{\varepsilon,k=12}^2$ are 0.13, 0.17 and 0.19 for the 1993, 1996 and 2001 panels, respectively.

Since our estimates for $\sigma_{\varepsilon,k=6}^2$ are simply intermediate in magnitude to the estimates of $\sigma_{\varepsilon,k=0}^2$ and $\sigma_{\varepsilon,k=12}^2$, we simply focus on this latter sets of estimates throughout the rest of the paper. Greater detail on $\sigma_{\varepsilon,k=0}^2$ and $\sigma_{\varepsilon,k=12}^2$ 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_{ε}^2 , with the estimates obtained by the extensive empirical literature on US labor market risk using annual income data drawn from the PSID. Note that our results are

estimates from these two time invariant Mincer specifications differ very little from those reported in Table II.

estimated using SIPP, a three-year panel for the United States, instead of the PSID data with a time dimension of many years. Most of these studies find an average value of around 0.0225 for the annual variance σ_ε^2 (Carroll and Samwick (1997), Gourinchas and Parker (2002), Hubbard, Skinner and Zeldes (1994), and Storesletten, Telmer and Yaron (2004)), with a value of $\sigma_\varepsilon^2 = 0.0324$ 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. Furthermore, we 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 significantly when we filter out shocks of longer durations. Taken together, these factors suggest that shocks that last longer than a year and that we label as “permanent” in this paper indeed persist for a very long time.

II.6. Income Risk in Sub-Samples

Our dataset is sufficiently large to estimate risk faced by sub-samples of workers who experience a range of different outcomes in the labor market. Our first sub-sample is constructed by including only the individuals who were employed in the same manufacturing industry in each month that 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 restrict the SWITCH_ALL sample to those individuals who switched to the non-manufacturing sector for at least one period in the panel to construct the SWITCH-NON-MANUF sample. Our last sub-sample includes individuals who stayed in the

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 sector.

manufacturing sector throughout, but may have worked in a different industry within manufacturing than their original industry at some point (STAY-MANUF).

Differences in income risk experienced by workers in these four different sub-samples (STAY-IND, STAY-MANUF, SWITCH-ALL and SWITCH-NON-MAN) reflect the costs of switching industries both within and outside manufacturing. However, the interpretation of the differences in risk estimates across different sub-samples is subject to an important qualification. In principle, whether a worker remains in the same manufacturing industry or switches to another industry is an endogenous decision. The allocation may be non-random and may very well reflect differences in worker characteristics such as their level of human capital. Therefore, we cannot infer that a worker who did not switch sectors (i.e., a worker in STAY-IND) would face the level of income risk estimated for the sub-sample of workers who did switch sectors (say SWITCH ALL) if he was exogenously forced to switch. The estimates of income risk for the different groups of workers are conditional estimates, reflecting the differences in worker characteristics and endogenous actions that place workers in these different sub-samples, and should be interpreted with this qualification in mind.²⁰

Table IV provides a summary description of our estimates of income risk for each sub-sample and panel.²¹ As Table IV indicates, $\sigma_{\varepsilon,k=0}^2$ continues to be greater than $\sigma_{\varepsilon,k=12}^2$ 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

²⁰ In practice, there is little systematic difference in the observable educational characteristics of switchers relative to non-switchers in our data. The fraction of switchers who fall in different educational categories (high school dropout, high school graduate, some college, college graduate, more than college) is nearly identical across switchers and non-switchers. While, these two samples reflect some differences in other observable characteristics, these differences are rather small (switchers are slightly younger with an average age of 36 years compared to 39, and slightly less likely to be married, 59% versus 64% married compared to non switchers). These empirical facts somewhat mitigate our concern regarding the comparison of income risk estimates across different sub samples. Nevertheless, we cannot rule out the possibility that the workers in these sub-samples are different in terms of unobservable characteristics and therefore, the estimates we discuss below are still subject to the qualification regarding the non-random allocation of workers discussed above.

²¹ Due to sample size restrictions, income risk for these sub-samples are estimated at the 2-digit SIC level which is more aggregated than the Census classification used in constructing the sub-samples.

switch jobs within the sector). The mean estimate of the monthly value of $\sigma_{\varepsilon,k=12}^2$ 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 higher 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 0.19, 0.19 and 0.20) for the 1993, 1996 and 2001 panels, respectively. Table IV also provides estimates of income risk experienced by workers who switch out of manufacturing (SWITCH-NON-MANUF). The variance in shocks to permanent income experienced by these workers is significantly larger (at least fifty percent higher) than those who stayed in the same industry throughout.

III. Trade and Income Risk

The procedure outlined in the previous section provides us with estimates of individual income risk, $\sigma_{\varepsilon js}^2$, 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, $\sigma_{\varepsilon js}^2$, and import penetration, M_{js} .²² In Figures III-A. and III-B, we plot the changes in estimated permanent income risk, $\Delta\sigma_{\varepsilon,k=0}^2$ and $\Delta\sigma_{\varepsilon,k=12}^2$, 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 strongly positive, suggesting that an increase in import penetration is associated with

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

an increase in income risk for the workers in that industry.

III.1. Specification

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

$$\sigma_{\varepsilon_{js}}^2 = \alpha_s + \alpha_j + \alpha_M M_{js} + v_{js}. \quad (7)$$

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 $\sigma_{\varepsilon_{js}}^2$ 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). Finally, since the dependent variable is estimated rather than measured, we adjust the standard errors for heteroscedasticity using a White correction.²³

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.

²³ 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 in the paper.

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,²⁴ 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 association between income risk and import penetration.

Nevertheless, we consider the possibility of such selection in the data and find that our concerns regarding selection bias are mitigated for the following reasons. First,

²⁴ 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.

we examine industries over time, so any fixed differences across industries in the composition of the workforce or worker characteristics are taken into account by our fixed effects estimation. Second, 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 share of each *educational* category, share of each *occupational* category, share of each *race* group, share of each *gender* and, finally, the average *age* within a sector are each uncorrelated with changes in import penetration across the span of the three SIPP panels. In addition, we find that the changes in the *variance* of years of education and age within a sector across panels are also uncorrelated with changes in import penetration. Third, even when we allow income risk to vary with age and education within industries in a specification analogous to (7) (see details in section III.5), we continue to find the coefficient on import penetration to be significant. This suggests that controlling for the (potentially changing) age and educational composition of sectors does not alter our findings regarding the link between trade and income risk.

Finally, we examine the possibility that selection is based on unobserved ability differences across workers. In this case, we would expect selection to be reflected in unexplained wage differentials across industries, as long as high-ability workers are paid higher wages. Our data suggests that any bias due to such unobserved ‘ability’ differentials (that are uncorrelated with observable characteristics) across industries is likely to be small. Specifically, we find that changes in unexplained portion of industry average wages are uncorrelated with changes in import penetration in our data. While the preceding discussion does not entirely eliminate our concerns regarding selection bias, we believe it substantially mitigates them.²⁵

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,

²⁵ 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.

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 by filtering out purely transitory shocks ($\sigma_{\varepsilon,k=0}^2$) or by filtering out transitory shocks that last up to a year ($\sigma_{\varepsilon,k=12}^2$). As noted earlier, since the dependent variable is estimated rather than measured, we adjust the standard errors for heteroscedasticity using a White correction.

We find that import penetration is significantly and positively 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

These findings of a positive and economically significant association between import penetration and income risk stand in some contrast to the findings of the literature on trade and wage inequality, where the broad conclusion has been that empirical work has only “small effects” (see Goldberg and Pavcnik (2004)).

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. As discussed earlier, the interpretation of the differences in risk estimates across different sub-samples is subject to the important qualification

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.

that the switching decision itself is an endogenous one and the estimates of income risk for the different groups of workers are conditional estimates, reflecting the endogenous actions that place workers in the different sub-samples. Nevertheless, this exercise is useful in understanding the relationship between international trade and income risk as experienced by different segments of the labor market.

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 $\sigma_{\varepsilon,k=0}^2$ and $\sigma_{\varepsilon,k=12}^2$ 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 $\sigma_{\varepsilon,k=0}^2$ 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, $\sigma_{\varepsilon,k=0}$ would increase by 5% as a result of a 10% increase in import penetration over its initial value. When $\sigma_{\varepsilon,k=12}^2$ 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 $\sigma_{\varepsilon,k=12}^2$ is the dependent variable.²⁷ We find that a 10% increase

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

in import penetration is associated with 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 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 be associated with an increase in income risk due to an increased elasticity of labor demand (Rodrik, 1997). On the other hand, offshoring 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

income risk for the full sample of workers. The positive estimated coefficient on import penetration implies that risk is indeed higher when switching from sectors with greater increases in import penetration. We hypothesize that this could be due to congestion as larger number of workers with similar skills leave these industries at the same time and experience greater variance in outcomes as a result.

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}

The “additional” variables discussed above are mostly variables that relate to the “supply side” of the economy. It could be argued that demand shocks may also potentially bias the estimate of the relationship between trade and income risk. Thus, there may be say, positive shocks to product demand in a sector that lead to greater import flows and raise both the mean and the variance of income in that sector. While this is indeed a theoretical possibility (for instance, if innovations to permanent income are related to industry prices), we do not find support for this argument in our

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.

data. We find the correlation between mean income changes and income risk to be insignificantly different from zero. Furthermore both the correlations between mean income changes and import penetration and between sectoral prices and income risk are not significantly different from zero, mitigating this as an empirical concern.

As a final robustness check, we allow income risk to vary by individual characteristics within an industry. More specifically, we first estimate risk separately for each age group and education level within an industry for each panel.³⁰ Then we estimate equation (7) by including dummy variables for each age group in addition to import penetration and time and industry fixed effects. We then repeat this analysis for workers with different education levels. In both cases, the coefficient on import penetration remains significant and positive, with little change in its magnitude.³¹

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.³²

The specific thought experiment that the theoretical structure allows us to address is the following one (Krebs (2004)): Imagine a group of ex-ante identical workers with

³⁰ We estimate risk by industry for five age groups (less than 30, between 30 and 40, between 40 and 50, and above 50) and for five education categories (high school drop-outs, high school graduates, college drop-outs, college graduates).

³¹ These results are available from the authors upon request.

³² 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 (raw and residual) income and import penetration (using econometric specifications like (7), with income growth on the left hand side rather than income risk). However, we did not find any consistent relationship between mean growth rates of (raw and residual) income and import penetration.

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),³³ 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:³⁴

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

and $\Delta_c = \left(\frac{\beta\Delta_\sigma\sigma_\varepsilon^2}{(1 - \beta)^2 2} \right) - 1, \text{ if } \gamma = 1$ (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 magnitude of this compensation is increasing in the degree of risk aversion, γ . Using

³³ 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 significantly when total income (including any capital earnings and transfers) instead of labor earnings is used as our income measure.

³⁴ The interested reader is referred to Krebs, Krishna and Maloney (2008) for a detailed derivation and discussion.

(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.³⁵ 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).³⁶

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)):

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

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

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

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

³⁵To 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.

³⁶Note 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 T years, workers draw their *permanent* income innovation terms $\varepsilon_{ij,t}$ in (3) from a bin with greater variance σ_ε^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, $\sigma_{\varepsilon, k=12}^2$ as well as results when income risk is estimated assuming $K=0$ ($\sigma_{\varepsilon, k=0}^2$). Results are provided separately for parameter values for the coefficient of risk aversion at $\gamma = 1$ and $\gamma = 2$ and for durations of $T = 5, 10$ and 15 years. All of the calculations use a discount factor $\beta = 0.98$. With $\gamma = 2$, 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 $\gamma = 2$ 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, $\gamma = 1$. As expected, welfare costs are smaller when individuals are less risk averse and are in the range of 1% to 3%. However, in both cases, the welfare costs associated with the income risk channel are economically quite significant and points to the need for a sizeable social safety net to insure workers against this risk.

The welfare analysis we have presented may be qualified along several different dimensions. First, we may ask about the extent to which the additional risk being borne by workers is risk that they “seek” in exchange for higher mean compensation that they receive. To address this issue, we examine whether mean income changes are positively correlated with changes in import penetration. We find that the two are uncorrelated in our data, a finding that we additionally confirm with regression analysis (with time and industry dummies included on the right hand side). Furthermore, mean wages are uncorrelated with income risk itself. This suggests that the risk we measure is indeed “borne” by workers rather than being sought by them in exchange for greater mean compensation.

Second, we may ask about the plausibility of the quantitative estimates of welfare costs delivered by our welfare-theoretic framework. As discussed above, our welfare estimates are simply measures of the “willingness to pay” to avoid the higher risk associated with greater exposure to trade. Seen in this light, and given our estimates regarding the magnitude of the association between income risk and trade, our estimates seem quantitatively plausible especially when computed with lower

magnitudes of parameters γ and T (i.e., when they are in the range of 1% to 3 % of consumption). Furthermore, it should be clear that these estimates only provide an indicative sense of what our risk estimates translate into in welfare terms since we ignore any mechanisms that agents may use to smooth consumption other than savings. Specifically, we have ignored several potential dimensions of “adjustment” which may allow workers to lower the cost of risk, such as the labor-leisure choice or intra-household diversification of sector and occupation.³⁷ The inclusion of these channels is outside of the scope of the present analysis and is left for future research.

Finally, we should emphasize 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. Our results suggest that increased import penetration has a statistically and economically significant effect on labor income risk in US manufacturing. We find that within-industry changes in income risk are strongly related to changes in import penetration for the full sample of workers as well as various sub-samples we consider, such as workers who stayed within the same manufacturing industry throughout and those who switched industries within or outside the manufacturing sector. Our welfare analyses suggest that the welfare cost of increased income risk associated with increased trade exposure is economically significant and that a sizeable social safety net is necessary to insure workers against this risk.

³⁷ However, we have considered the possibility that workers may make asset ownership choices that allow them to reduce risk by considering “total” earnings rather than just labor income. As noted earlier, this does not make any significant difference to our results.

We emphasize that 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

Variable	1993		1996		2001	
	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			
$\sigma_{\varepsilon,k=0}^2$	0.0033	0.0031	0.0016
$\sigma_{\varepsilon,k=6}^2$	0.0018	0.0015	0.0016
$\sigma_{\varepsilon,k=12}^2$	0.0014	0.0014	0.0019
1996-1998			
$\sigma_{\varepsilon,k=0}^2$	0.0043	0.0042	0.0013
$\sigma_{\varepsilon,k=6}^2$	0.0024	0.0023	0.0014
$\sigma_{\varepsilon,k=12}^2$	0.0025	0.0026	0.0018
2001-2003			
$\sigma_{\varepsilon,k=0}^2$	0.0052	0.0051	0.0016
$\sigma_{\varepsilon,k=6}^2$	0.0033	0.0034	0.0019
$\sigma_{\varepsilon,k=12}^2$	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 ($\sigma_{\varepsilon,k=0}^2$ and $\sigma_{\varepsilon,k=12}^2$)

SIC	$\sigma_{\varepsilon,k=0}^2$			$\sigma_{\varepsilon,k=12}^2$		
	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

	$\sigma_{\varepsilon,k=0}^2$		$\sigma_{\varepsilon,k=12}^2$	
	Mean	Std. Dev.	Mean	Std. Dev.
1993-1995				
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

	$\sigma_{\varepsilon,k=0}^2$		$\sigma_{\varepsilon,k=12}^2$	
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 ($\sigma_{\varepsilon,k=0}^2$)

	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 ($\sigma_{\varepsilon,k=12}^2$)

	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
(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
(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
(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

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

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 ($\sigma_{\varepsilon, k=12}^2$)

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

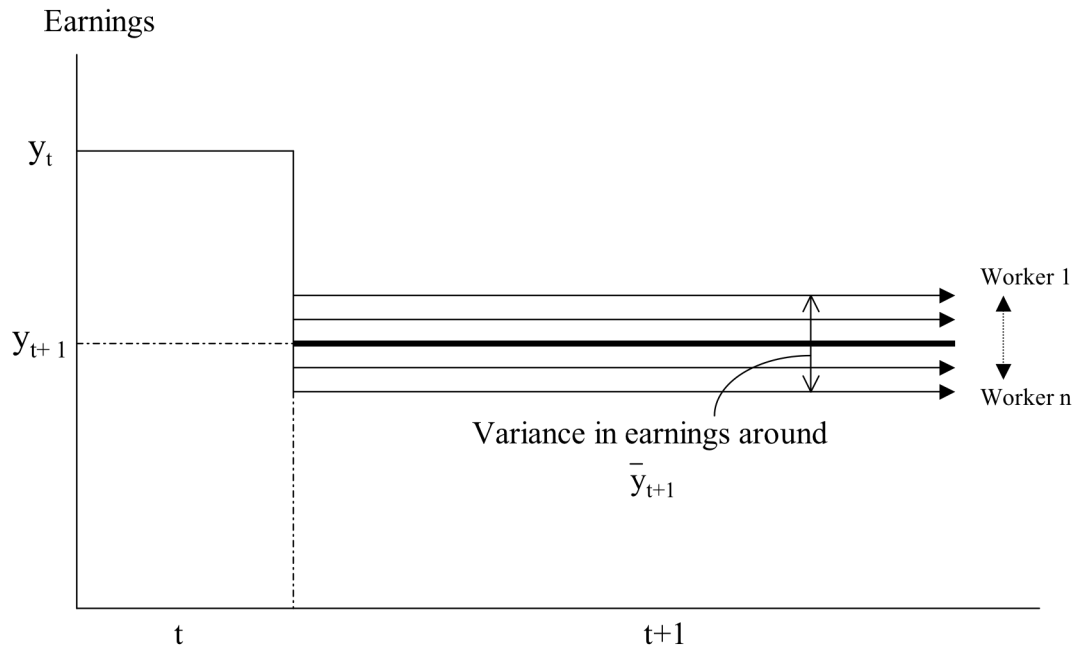


Figure II. Transitory versus Permanent Shocks

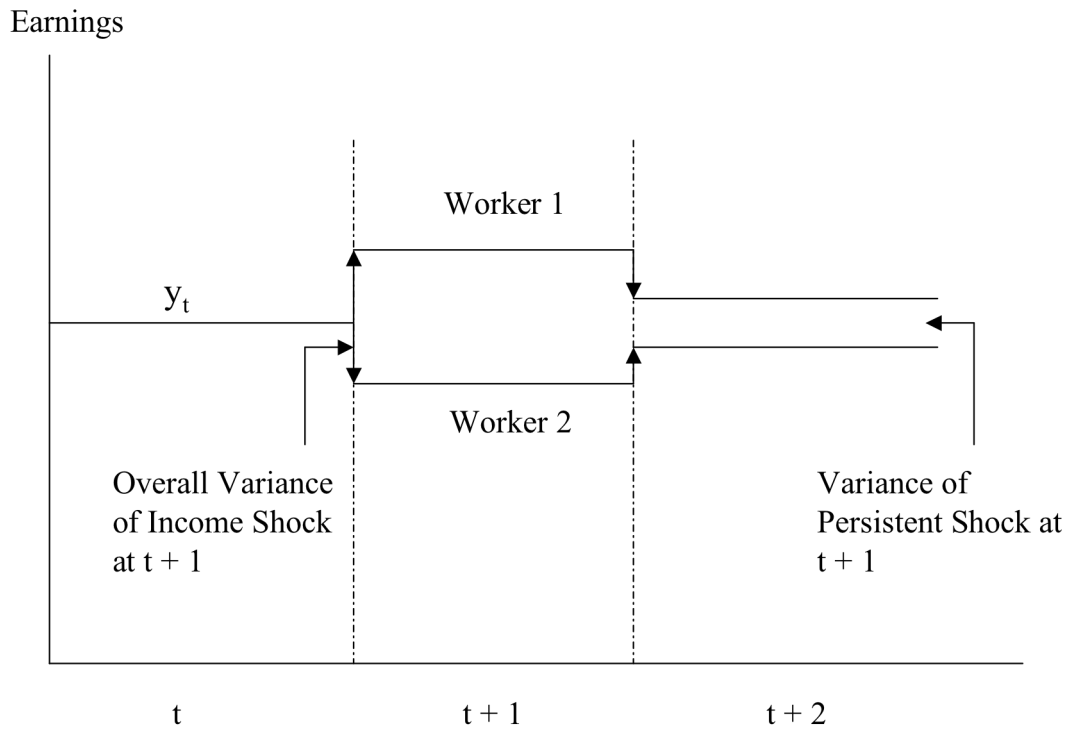
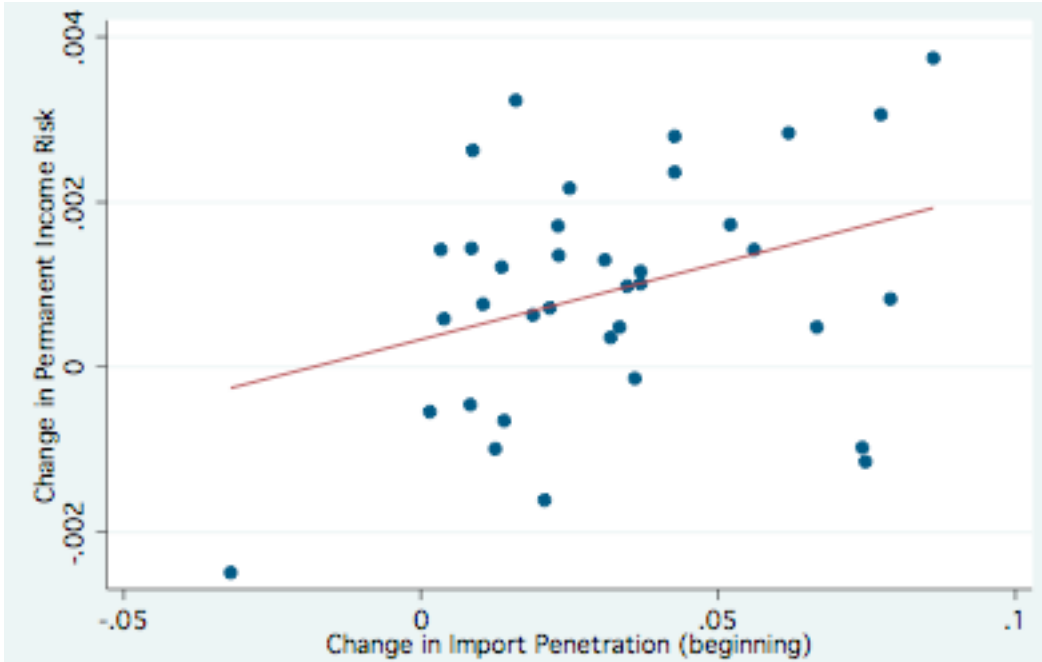


Figure III. Changes in Permanent Income Risk and Changes in Import Penetration

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



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

